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Fiscal Sustainability, Sovereign Credit Risk and Fiscal Policy

Pallara Kevin

UNIL | Université de Lausanne

FACULTÉ DES HAUTES ÉTUDES COMMERCIALES

DÉPARTEMENT D'ÉCONOMIE

Fiscal Sustainability, Sovereign Credit Risk and Fiscal Policy

THÈSE DE DOCTORAT

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pour l'obtention du grade de
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I M P R I M A T U R

Sans se prononcer sur les opinions de l'auteur, la Faculté des Hautes Etudes Commerciales de l'Université de Lausanne autorise l'impression de la thèse de Monsieur Kevin PALLARA, titulaire d’un bachelor et d’un master en sciences économiques de l’Université de Rome Tor Vergata, en vue de l'obtention du grade de docteur ès Sciences économiques, mention « économie politique ».

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Fiscal Sustainability, Sovereign Credit Risk and Fiscal Policy

By
Kevin Pallara

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Introduction

Since the Great Recession and, then, with the euro-area sovereign debt crisis and the COVID pandemic, fiscal sustainability has started to be a concern not only for emerging economies but also for advanced ones. Given that rising default risk might have serious economic implications, the study of sovereign credit risk has grown in relevance. At the same time, the ability of the government budget to be resilient to crises has become crucial. Thus, the measurement of the space of manoeuvre available to governments for the implementation of discretionary fiscal policy — dubbed as fiscal space — has increasingly drawn attention.

Sovereign bond prices depend on the view of investors on fiscal sustainability. And sovereign bond data is abundant, with large cross sections (for varying maturities). This data is however underused for investigating sovereign debt sustainability. Arguably, this is due to the lack of flexible models capturing the joint dynamics of debt and the cross-section of sovereign bond prices. In the first chapter of this thesis, coauthored with Jean-Paul Renne from the University of Lausanne, we aim to fill this gap. We develop a novel sovereign credit risk model that accounts for public debt dynamics, bond prices, and the fiscal limit, which is the maximum outstanding debt that a country could credibly sustain. Our model strikes a unique balance between completely reduced-form credit risk approaches and structural macroeconomic models. Indeed, our framework features a small-scale macroeconomic model that encapsulates the debt accumulation process and where risk-averse investors take sovereign default into account when it comes to price financial assets exposed to government credit risk, such as credit default swaps (CDS). A key ingredient of the model is that the probability of sovereign default explicitly depends on fiscal space, namely the distance of sovereign debt from the fiscal limit: as long as this distance is not exhausted, default has zero probability of materializing. Additionally, our model features a feedback effect of default on the macroeconomy and a stochastic discount factor deriving from actual preferences. We estimate our single-country model for eight countries comprising both advanced (Canada, Japan, United Kingdom and the United States) and emerging economies (Brazil, China, India and Russia), over the period 2007-2021. Given that fiscal limits are unobserved, we exploit the information contained in the term-structure of sovereign credit spreads to provide time-varying estimates of fiscal limits and fiscal spaces, which is a novelty in the literature. Our estimates for fiscal limits feature substantial time variation and comove with economic policy uncertainty. Moreover, our framework captures the nonlinear sensitivities of credit spreads to fiscal conditions. Interesting byproducts of the model are credit risk premiums — the components of sovereign spreads that would not exist if agents were risk neutral — that arise from the computed large discrepancies between risk-adjusted and non risk-adjusted probabilities of default.

Due to the turmoil linked to the euro-debt crisis, the study of fiscal sustainability in the euro-area has become more and more relevant. The second chapter of this thesis —- joint work with Jean-Paul Renne from the University of Lausanne — focuses on the sustainability and the issuance of public debt in the euro-area. Specifically, we price debt instruments jointly issued by countries part of the currency union, which represents a novelty in the literature. We also study the impact of joint-issuance on the cost of debt. In the context of the euro-area, bonds that are jointly issued by a group of countries are dubbed as Eurobonds. We focus on two types of bonds: the first is backed by several and joint (SJG) guarantees, and the second features several but not joint (SNJG) guarantees. In the first instance, we assume that the countries participating in the issuance of the joint bond pool fiscal revenues and public debts together. Thus, each country is liable also for other states' eventual unpaid contributions. In the second instance, each country participating in the issuance is liable only for the debt service and principal redemption corresponding to its share of the bond. Eurobonds are more or less explicitly advocated as an effective way to offer debt service relief to peripheral member states. Moreover, if issued on large scale, Eurobonds could increase bond market liquidity to make the euro-area bond market compete with the US one. Importantly, the issuance of joint debt instruments might address the growing demand for safe assets of financial institutions. We build a model similar to the one proposed in the first chapter (even though we need to simplify the machinery at play to ensure tractability in the context of a multi-country framework). We use data on national bond prices and public debts for the six largest euro-area economies to estimate the multi-country model over the period 2008-2021. We are able to estimate time-varying fiscal limits that are used to deduce counterfactual Eurobond prices under both the SJG and SNJG issuance schemes. For the 5-year maturity, we find that SNJG bond yield spreads are almost three times larger than SJG ones over the estimation sample. Thus, the issuance of SJG bonds in the euro-area context could yield significant aggregate gains. Given the concerns of moral hazard associated with joint debt issuance, we envision post-issuance redistribution schemes whereby the yield gains arising from SJG bond issuance are shared among the participating countries. This way, the reduction in market discipline is diminished. With this study, we show that aggregate fiscal sustainability in the euro-area might be improved by issuing jointly guaranteed debt instruments.

The study of fiscal sustainability is key also to understand the role of fiscal policy in light of the large spending plans implemented in many advanced economies during the Great Recession, especially in the US. Indeed, the role of fiscal policy has gained traction as discretionary fiscal measures have started afresh to serve as policy tools in advanced economies. Sizeable stimulus packages translated into growing deficits that piled up into unprecedented levels of public debt. The latter, together with stagnant growth, raised attention on the sustainability of public finances. This called into question whether the effects of fiscal policy might as well depend on fiscal sustainability itself. According to this view, expansionary fiscal policy can be an effective tool in certain situations while not in others. For instance, an expansion in the public budget associated with a weak fiscal position can even produce harmful effects, while, on the opposite, the same fiscal shock implemented when public finances are sound generates expansionary effects. In the third chapter of this thesis, jointly written with Luca Metelli from the Bank of Italy, we empirically investigate such hypothesis. To do so, as a measure of fiscal sustainability, we focus on fiscal space. The latter can also be defined as the *"room in a government's budget that allows it to provide resources for a desired purpose without jeopardizing the sustainability of its financial position or the stability of the economy"* (Heller, 2005). Hence, we measure the historical evolution of fiscal space in the US according to various indicators. We then use these fiscal space indicators to define periods under which the fiscal position in the US is weak or strong. Thus, we estimate the effects of government spending shocks in the US according to the level of fiscal space over the period 1929-2015. We find that the state of fiscal space is key for the effectiveness of government spending expansions: the fiscal multiplier is above one when fiscal space is large, while it is below one when fiscal space is tight. This result is robust across different specifications for the fiscal space indicator, identifications of the government spending shock and estimation samples. By computing consumption and investment multipliers, we observe crowding out of the private sector when fiscal space is tight, while private consumption is boosted when fiscal space is large.

The idea of a differential effect of fiscal policy according to the fiscal position fits in the more general debate on fiscal policy, which has established a consensus on the fact that there is no such thing as a unique fiscal multiplier. Indeed, the effects of fiscal shocks are likely to be state dependent. Thus, it is relevant to focus also on other state dependencies and different fiscal policy measures. In the context of soaring public debts and unsustainable fiscal paths, policy circles have recently advocated for fiscal adjustments to curb debt dynamics. Especially in the European Union, under the *Stability and Growth Pact*, the response to large increases in deficit and debt consists in strengthening the budget balance via fiscal consolidations. The latter are policy measures explicitly aimed at reducing government deficits and debt accumulation. These fiscal measures can be divided into two categories depending on whether they mainly hinge on tax hikes or spending cuts. Based on the type of fiscal adjustment that is implemented, namely *how* the consolidation is carried out, the effects on the economy can vary. In the fourth chapter of this thesis, I study how tax-based (TB) and expenditure-based (EB) consolidations affect the macroeconomy depending on the initial state of the economy — *when* the fiscal adjustment is announced. The empirical analysis is carried out over the period 1978-2013 on 13 European economies. The principal aim of the study is to disentangle the effects of fiscal consolidations across monetary policy regimes and fiscal position states as these state dependencies are partially neglected in the literature. I find that an EB announcement of consolidation measures is contractionary when monetary policy is constrained. On the other hand, I observe mild expansionary effects of EB consolidations when monetary policy is free to adjust. To capture the state of the fiscal position, I use fiscal space indicators as in the third chapter. I find that both EB and TB announcements of consolidation measures yield different effects on the economy depending on the state of fiscal space. TB consolidations are contractionary (non-recessionary) and EB ones are non-recessionary (expansionary) when the economy starts in a weak (strong) fiscal position.

The primary objective of this thesis is to propose a new methodology to assess sovereign debt sustainability and measure credit risk. The first two chapters provide estimates of the maximum outstanding sovereign debt that can be credibly sustained — the fiscal limit — and develop flexible pricing frameworks for different types of sovereign bonds. Furthermore, the third chapter of this thesis takes an additional step and analyses how the evolution of fiscal sustainability — captured through the concept of fiscal space — affects fiscal policy. In the last chapter, more generally, I explore how the effectiveness of fiscal adjustments hinges on the state of the economy, which remains a topical subject to be further investigated.

Chapter 1

Fiscal Limits and Sovereign Credit Spreads[1]

1.1 Introduction

Before the Great Financial Crisis that began in 2007, it was widely assumed that developed-countries sovereign bonds were perfectly safe. That is, they were believed to provide the same payoff at any point in time, and in any state of the world. This belief has, however, been undermined since then, and in particular since the inception of the euro-area sovereign debt crisis in the early 2010s. Because government defaults have severe economic implications, and against the backdrop of soaring public indebtedness, the measurement of sovereign credit risk has been drawing particular attention over the last decade.[2]

Unsustainable fiscal paths lead to sovereign defaults. Accordingly, sovereign bond prices depend on investors' perception of public debt sustainability. Data on sovereign bond prices are abundant: they are available at high frequency and for large cross-sections of maturities. The richness of these data is however underused in the literature investigating sovereign debt sustainability. Arguably, this underuse can be accounted for by the lack of a modeling framework that explicitly incorporates the debt dynamics while being flexible enough to capture the time- and cross-section variability of sovereign-bond prices. The objective of this paper is to fill this gap.

We develop a small-scale macroeconomic model that encapsulates the debt accumulation process and where risk-averse investors take sovereign default into account when it comes to price financial assets exposed to government credit risk, such as credit default swaps (CDS).

[1] This chapter is coauthored with Jean-Paul Renne from the University of Lausanne.

[2] On the literature discussing the disruptive economic effects of sovereign default, see, e.g., Panizza et al. (2009), Reinhart and Rogoff (2011), and Mendoza and Yue (2012).

A sovereign default takes place when the level of debt (d_t) exceeds the *fiscal limit* (ℓ_t), that is the maximum outstanding debt that can credibly be covered by future primary budget surpluses. Because agents' expectations regarding the present value of future surpluses are state-dependent, the fiscal limit is subject to time variation. This limit is not directly observed by the econometrician.[3] However, since the model predicts how financial prices depend on current expectations regarding the fiscal limit, estimates of the latter can be backed out from credit spreads.

Our model offers a unique balance between macroeconomic structure and fit of observed credit derivative prices. Its fitting property hinges on the existence of approximate formulas for valuing credit derivatives, which in turn depend on econometric modeling choices. In our case, the structure of the model dictates the type of relationships between default probabilities and macroeconomic variables—the latter being typically real-valued, i.e. with positive or negative values. Standard tractable (affine) processes used in credit-risk models can not handle this general situation. Indeed, to accommodate (nonnegative) default-intensities, affine credit-risk settings typically entail only nonnegative and positively-correlated pricing factors, which is too restrictive in our context.[4] We address this issue by borrowing some modeling ingredients from the shadow-rate interest-rate literature that builds on Black (1995).[5] In our framework, the default intensity is a nonlinear function of the fiscal space $d_t - \ell_t$: it is equal to zero as long as debt (d_t) is lower than the fiscal limit (ℓ_t), and strictly positive otherwise. More specifically, it is of the form $\max(0, \alpha(d_t - \ell_t))$. Hence, up to a negative multiplicative factor (α), the fiscal space here takes the place of the shadow rate used in credit-risk-free yield curve models. The resulting asset-pricing framework is flexible and tractable. While default-intensities remain non-negative, they can positively/negatively depend on positive/negative macroeconomic variables, that are positively/negatively correlated with each other.

[3] In what follows, fiscal limits do not represent political or institutional constraints on the level of public debt, such as the Maastricht debt criteria for eurozone countries, or the federal debt ceiling in the US. As noted in Hall and Sargent (2015), US debt ceilings do not affect surpluses/deficits and do not represent a structural criterion to steer fiscal action.

[4] Default intensities are typically expressed as linear combinations of variables following multivariate square-root, "CIR," diffusions (Cox, Ingersoll, and Ross, 1985). In these frameworks, the unconditional correlations between the components of the multivariate process are positive (Dai and Singleton, 2000), which generates restrictions in the credit-risk modeling context (Duffie and Singleton, 1999, Subsection 2.2). Doshi et al. (2013) develop an affine framework where default intensities depend on observable covariates while remaining positive; their quadratic specification however has the disadvantage of implying large probabilities of default when any observables has either large positive or large negative values.

[5] In shadow-rate models, the nominal short-term interest rate is taken equal to the maximum between a real-valued shadow rate and zero, thereby guaranteeing the non-negativitiy of nominal yields (see, e.g., Christensen and Rudebusch, 2013; Kim and Singleton, 2012; Kim and Priebsch, 2013; Krippner, 2013; Wu and Xia, 2016; Coroneo and Pastorello, 2020).

Our empirical analysis focuses on eight large countries, four of which being advanced economies (Canada, Japan, the U.K., and the U.S.), and four being emerging markets (Brazil, China, India, and Russia).[6] We choose these countries because most of their sovereign debt is owned domestically, which helps make the data consistent with our closed-economy framework. The results indicate that fiscal limits of these countries feature a substantial degree of time variation. Following the Great Financial Crisis, we observe a large drop (about 10 percentage points of GDP) in fiscal space estimates on average across advanced countries. By contrast, starting from the end of 2009 and until 2020, estimated average fiscal spaces have increased by roughly 35 percentage points of GDP across countries. Moreover, our fiscal space estimates negatively correlate to economic policy uncertainty (EPU) indexes.

The estimated models predict a non-linear influence of fiscal conditions on credit spreads, in line with the findings of a wide body of empirical studies.[7] Compared to regression-based analysis, our framework provides richer predictions of the changes in the term structure of credit spreads that can be expected from fiscal deterioration. Simulation exercises highlight in particular that the sensitivity of CDS spreads to deficits crucially depends (i) on the size of the shock and (ii) on the initial debt level. The increase in spreads can for instance be several times larger when (i) the deficit goes from 9% to 10% of GDP, compared to when it goes from 0% to 1%, say, and (ii) when the initial (pre-shock) debt-to-GDP is larger.

Finally, we investigate the wedge between non-risk-adjusted CDS-based probabilities of default and risk-adjusted, or physical, probabilities of default. This wedge is driven by sovereign credit risk premiums, which are those components of sovereign credit spreads that would not exist if investors were not risk-averse. If agents were risk-neutral, CDS spreads would be approximately equal to expected credit losses, i.e., the products of the loss-given default multiplied by the probabilities of default. But with risk-averse agents, and if sovereign defaults tend to take place in "bad states," i.e. high-marginal-utility states, then protection sellers are willing to enter the credit swap only if the CDS spread is larger than the expected credit loss (see, e.g., Chen, Collin-Dufresne, and Goldstein, 2009; Gabaix, 2012, Subsection III.D). In our model, two channels imply that sovereign defaults are expected to be more frequent in bad states, i.e. during recessions: first, the debt-to-GDP ratio mechanically soars as GDP (the denomina-

[6] We estimate one model per country. In a companion paper (Pallara and Renne, 2022), we develop a multi-country model to jointly price sovereign bonds issued by different euro-area countries. The multi-country dimension comes at the cost of having a more reduced-form model. In particular, there is no debt-accumulation process in Pallara and Renne (2022).

[7] See e.g. Haugh, Ollivaud, and Turner (2009), Bernoth, von Hagen, and Schuknecht (2012), Alessi, Balduzzi, and Savona (2020).

tor) falls; second, the model accounts for the recessionary effect of the default event itself—in a manner akin to catastrophe modeling in the disaster-risk literature. Together, these two mechanisms underlie sizeable risk premiums, translating into substantial differences between risk-premium-adjusted sovereign probabilities of default and the unadjusted ones stemming from basic models like in Litterman and Iben (1991). The latter are extensively used by market practitioners, who refer to them as "market-implied default probabilities;" our results suggest that they overestimate the physical probabilities of default by a factor of two.

The remaining of this paper is organized as follows. Section 1.2 reviews related literature. The model is developed in Section 1.3. Section 1.4 describes the estimation strategy. Section 1.5 discusses the results. Section 1.6 summarizes our findings and makes concluding remarks. The appendix gathers technical results, supplementary details, proofs and additional findings.

1.2 Related literature

This paper contributes to the growing literature on sovereign credit risk and its pricing. Over the last decades, sovereign credit risk has been studied both from the macroeconomic and the financial points of view, but in somewhat separate ways.

1.2.1 Reduced-form approaches and sovereign risk premiums

In finance, different models entailing realistic default probabilities and default risk premiums have been proposed.[8] In most of these studies, the basic ingredient is a reduced-form default intensity (Duffie and Singleton, 1999). Duffie, Pedersen, and Singleton (2003) employ such a reduced-form approach to model the term structure of Russian credit spreads. Pan and Singleton (2008) estimate intensity-based models using sovereign CDSs. Ang and Longstaff (2013) consider multi-factor affine models allowing for both systemic and sovereign-specific credit shocks to price the term structures of US states and Eurozone Member States. Longstaff, Pan, Pedersen, and Singleton (2011) estimate default intensities for 26 countries; they find that, on average, the risk premium represents about a third of credit spreads. Allowing for both credit

[8]"Risk premium" refers here to the part of a credit spread that would not exist if investors were risk-neutral. This premium corresponds to the excess return (beyond expected credit losses) asked by the investors to be compensated for the fact that defaults tend to take place in "bad states" of the world, i.e. in states of high marginal utility (see, e.g., Chen, Collin-Dufresne, and Goldstein, 2009; Gabaix, 2012, Subsection III.D). Such risk premiums are often seen as explanations to the so-called credit spread puzzle (D'Amato and Remolona, 2003).

and liquidity effects—modeled through credit and liquidity intensities—Monfort and Renne (2014) find substantial sovereign risk premiums in euro-area sovereign spreads.

These studies generally involve latent factors and present a close fit of sovereign bond yields and spreads; they also provide useful estimates of sovereign risk premiums. But they are silent about the economic forces that drive the movements of the sovereign default probabilities. Borgy, Laubach, Mésonnier, and Renne (2011) and Hördahl and Tristani (2013) propose sovereign credit risk frameworks that also involve reduced-form relationships but where default intensities explicitly depend on fiscal variables, thereby allowing the fiscal environment to capture part of the fluctuations of sovereign CDS spreads. Augustin and Tedongap (2016) and Augustin et al. (2021) provide a more structural approach and value, respectively, Eurozone and US CDSs from the perspective of an Epstein-Zin agent. Without explicitly basing it on a government budget constraint, they posit a function connecting the sovereign default probability to expected consumption growth, macro volatility, and government expenditures, among other variables. The model developed in the present paper also uses a default intensity. However, compared to the papers mentioned above, the default intensity is more structural. Indeed, it directly depends on the fiscal space, which is itself affected by debt, the accumulation process of which is captured by the model.

1.2.2 Theory of sovereign defaults and fiscal limits

Early studies on sovereign credit risk focus on the strategic aspect of such defaults. Following the seminal contribution of Eaton and Gersovitz (1981) several studies have modeled sovereign defaults as strategic decisions of governments balancing the gains from stopping repaying debt against the costs of exclusion from international credit markets (influential studies include Aguiar and Gopinath, 2006; Arellano, 2008; Arellano and Ramanarayanan, 2012; Mendoza and Yue, 2012). These models predict that the probability of default increases in debt level. In most instances, these models are solved in the context of risk-neutral investors, ruling out the existence of risk premiums in credit spreads.[9]

Recently, another line of work, that we will refer to as the "fiscal limit" literature, has emerged. This literature relates to Bohn (1998), who provides evidence of fiscal corrective action. More precisely, Bohn (1998) finds that the US primary surplus is an increasing function of the debt-to-GDP ratio (see Mendoza and Ostry, 2008; Ghosh, Ostry, and Qureshi, 2013, for

[9] An exception is Verdelhan and Borri (2010) who consider risk-sensitive lenders buying emerging-market sovereign bonds.

more recent evidence). If the government is committed to raising fiscal surplus in response to rising debt levels, then it can guarantee intertemporal solvency as long as (i) tax rates are below the revenue-maximizing level and (ii) tax rates can be freely adjusted. Papers belonging to the fiscal limit literature depart from that intertemporal solvency situation by assuming that the government is not—or cannot be—committed to such a policy. In Bi (2012), Leeper (2013), Bi and Leeper (2013), Bi and Traum (2012), Bi and Traum (2014), the fiscal limit corresponds to the discounted present value of future maximum primary surpluses. These maximum surpluses can be seen as peak points of the Laffer curve (Trabandt and Uhlig, 2011). Ghosh, Kim, Mendoza, Ostry, and Qureshi (2013) estimate the responses of primary surpluses to debt levels for 23 advanced economies and observe that the responses are weaker at higher levels of debt—a phenomenon the authors dub "fiscal fatigue." After having introduced their estimated parametric reaction function in a model of debt accumulation, Ghosh et al. (2013) show that there is a point—akin to the fiscal limit—where the primary balance cannot realistically keep pace with the rising interest burden as debt increases. Beyond this point, debt dynamics becomes explosive and the government becomes unable to fully meet its obligations. Collard, Habib, and Rochet (2015) also exploit the idea of a maximum primary surplus to derive a measure of debt limit. But contrary to the previous studies, Collard et al. (2015)'s approach is not based on the computation of the discounted present value of future maximum primary surpluses; instead, their notion of maximum sustainable debt derives from the maximum amount that can be issued on each date (that itself depends on the maximum budget surplus). More recently, Mehrotra and Sergeyev (2020) combine disaster risk and fiscal fatigue. In their framework, as in Lorenzoni and Werning (2013), debt dynamics are subject to a tipping point situation: in some instances, the public debt can be on an unsustainable path without immediately triggering default.

For tractability, the models used in the fiscal limit literature generally assume that the government issues short-term (one-period) bonds only, which may alter the assessment of sovereign credit risk (Arellano and Ramanarayanan, 2012). An exception is the study by Chernov, Schmid, and Schneider (2020), where the government issues both short- and long-term bonds. In Chernov et al. (2020)'s model, increases in tax rates have negative effects on output, which also implies that there is a point where taxes cannot be raised further without reducing future tax revenues, in the spirit of the Laffer curve. In this context, the default probability tends to be higher in recessions, translating into large sovereign risk premiums, that is into sovereign spreads being larger than expected credit losses.

Solving the previous structural fiscal-limit-related models is challenging as soon as several shocks and state variables are considered (see, e.g., the Appendix of Chernov et al., 2020). Typically, in most of these models, the risk-free rates are considered to be constant and the term-structure of credit spreads is not discussed. Because of the challenging solution procedures, the models are essentially calibrated—i.e. the parameters are not econometrically estimated—and the ability of the model to capture the data, in the time and/or maturity dimensions is not examined.

1.3 Model

1.3.1 Overview

We consider an economy populated by a representative risk-averse agent that prices instruments whose payoffs are exposed to the default event of the government. The government default status is denoted by a binary variable $\mathcal{D}_t$, with $\mathcal{D}_t = 1$ if the government has defaulted at or before t, and $\mathcal{D}_t = 0$ otherwise. On date t, the investor observes $\mathcal{D}_t$ as well as consumption (C_t), the gross domestic product (Y_t), a price index (P_t), and the government budget surplus (S_t). The (log) growth rates of consumption, GDP, and the price index are respectively denoted by Δc_t, Δy_t, and π_t. (Hence, π_t is the inflation rate.) These three growth rates are themselves driven by persistent processes gathered in a n_w-dimensional vector w_t, which is also observed by the representative agent. We denote by $\mathbb{E}_t$ the conditional expectation given the information at time t, that is $\mathcal{I}_t = \{X_t, X_{t-1}, \dots\}$, where $X_t = \{\mathcal{D}_t, \Delta c_t, \Delta y_t, \pi_t, S_t, w_t\}$.

The remainder of the present section is organized as follows. Subsections 1.3.2 to 1.3.5 present different modeling ingredients, including agents' preferences, the type of debt instruments issued by the government, and the resulting debt accumulation process. Subsections 1.3.6 and 1.3.7 respectively specify the fiscal limit and the sovereign default probability. Subsection 1.3.8 provides a synthetic representation of the state vector. Finally, Subsection 1.3.9 introduces Credit Default Swaps (CDSs), whose pricing is pivotal for our estimation approach.

1.3.2 Consumption, GDP and inflation

Consumption growth , output growth , and inflation jointly depend on (a) vector w_t, (b) the binary default indicator $\mathcal{D}_t$ and (c) a n_η-dimensional vector of i.i.d. normal shocks η_t:

$$\begin{bmatrix} \Delta c_t \\ \Delta y_t \\ \pi_t \end{bmatrix} = \begin{bmatrix} \mu_c \\ \mu_y \\ \mu_\pi \end{bmatrix} + \begin{bmatrix} \Lambda_c' \\ \Lambda_y' \\ \Lambda_\pi' \end{bmatrix} w_t - \begin{bmatrix} b_c \\ b_y \\ b_\pi \end{bmatrix} (\mathcal{D}_t - \mathcal{D}_{t-1}) + \begin{bmatrix} \sigma_c' \\ \sigma_y' \\ \sigma_\pi' \end{bmatrix} \eta_t, \tag{1.1}$$

where $\eta_t \sim i.i.d.\mathcal{N}(0, I)$.

Parameters b_c, b_y and b_π capture the impact of a sovereign default on the macroeconomy. Eq. (1.1) is inspired by studies concerned with the asset-pricing influence of disasters (e.g. Barro, 2006, eq. 7, Arellano, 2008, eq. 3, Barro and Jin, 2011, eq. 1, Gabaix, 2012, eq. 1, Arellano and Ramanarayanan, 2012, last equation of Section III, Wachter, 2013, eq. 1).

Moreover, we posit an exogenous Gaussian vector auto-regressive (VAR) process for w_t.[10] Specifically:

$$w_t = \Phi w_{t-1} + \varepsilon_t, \tag{1.2}$$

where $\varepsilon_t \sim i.i.d.\mathcal{N}(0, I)$. The exogenous vectors w_t and η_t are components of vector x_t (see Subsection 1.3.1).

It remains to specify the dynamics of the default indicator $\mathcal{D}_t$. As detailed in Subsection 1.3.7, the probability of default will depend on the fiscal space, that is the difference between the fiscal limit and the debt level. The dynamics of the latter two variables notably hinges on the specification of agent attitude towards risk, which is done in the next subsection.

1.3.3 Investors' preferences and s.d.f.

The representative agent derives utility from consumption relative to an external reference level (Garcia, Renault, and Semenov, 2006; Bekaert, Engstrom, and Xu, 2022). Formally, the intertemporal utility is given by:

$$V_t = \mathbb{E}_t \left[\frac{1}{1-\gamma} \sum_{h=0}^{\infty} \delta^h \left(\frac{C_{t+h}}{S_{t+h}} \right)^{1-\gamma} \right], \tag{1.3}$$

[10]Because it follows a Gaussian VAR model, w_t is an affine process. As is well-known, this implies in particular that the conditional expectations of exponential affine transformations of future values of w_t can be computed in closed-form (see eq. a.1.7 in Appendix 1.B.1).

where S_{t+h} is a deterministic reference scenario defined by a constant growth rate equal to μ_c, that is: $S_t = \exp(t\mu_c)$. The resulting stochastic discount factor between dates t and $t+1$ is:

$$\mathcal{M}_{t,t+1} = \delta \exp[-\gamma(\Delta c_{t+1} - \mu_c) - \mu_c], \tag{1.4}$$

where $\gamma \geq 0$ is the coefficient of relative risk aversion and $\delta \geq 0$ is the rate of time preference.

Because consumption growth is affected by changes in $\mathcal{D}_t$ (see eq. 1.1), the s.d.f. jumps upon sovereign default. This has important implications in terms of pricing, by giving rise to specific risk premiums—called credit-event premiums—in the prices of financial instruments whose payoffs depend on the government default status, such as CDSs (Driessen, 2005; Gouriéroux, Monfort, and Renne, 2014; Bai, Collin-Dufresne, Goldstein, and Helwege, 2015).[11]

1.3.4 Government debt issuances

Following, among others Leland (1998), Hatchondo and Martinez (2009) or Arellano and Ramanarayanan (2012), we adopt the simplifying assumption that the government issues perpetuity contracts with nominal coupon payments that decay geometrically at rate χ. The closer χ to one, the larger the duration of the debt instrument.[12,13]

In the present model, government-issued perpetuities feature credit risk. Denoting by RR the recovery rate, an investor having purchased the perpetuity on date t receives the following

[11] These risk premiums help, in particular, fit short-term credit spreads by allowing the "risk-neutral" default intensity of the considered entity to deviate from its historical default intensity. Specifically, Appendix 1.I (eq. a.1.51) shows that the relationship between the physical ($\mathbb{P}$) and risk-neutral ($\mathbb{Q}$) default intensities is:

$$\underline{\lambda}_t^{\mathbb{Q}} = \underline{\lambda}_t + \log(\exp(\gamma b_c)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t)),$$

where $\underline{\lambda}_t^{\mathbb{Q}}$ satisfies $\exp(-\underline{\lambda}_t^{\mathbb{Q}}) \equiv \mathbb{Q}(\mathcal{D}_t = 0|\mathcal{D}_{t-1} = 0, w_t, \eta_t, \mathcal{I}_{t-1})$, and where $\mathbb{Q}$ is a measure equivalent to the physical one ($\mathbb{P}$) defined through the Radon-Nikodym derivatives $d\mathbb{Q}/d\mathbb{P}|_{t,t+1} = \mathcal{M}_{t,t+1}/\mathbb{E}_t(\mathcal{M}_{t,t+1})$ (Subsection 1.5.3 elaborates further on the risk-neutral measure.) In particular, the previous equation implies that if $\gamma b_c > 0$, then $\underline{\lambda}_t^{\mathbb{Q}} > \underline{\lambda}_t$. Moreover, in this context, the price of a sovereign maturity-h zero-recovery-rate bond is not given by the standard formula $\mathbb{E}_t^{\mathbb{Q}}(\exp(-r_t - \underline{\lambda}_{t+1}^{\mathbb{Q}} - \dots - r_{t+h-1} - \underline{\lambda}_{t+h}^{\mathbb{Q}}))$ because the "no-jump" condition is not verified for this bond (the previous conditional expectation jumps on the default date; see Duffie, Schroder, and Skiadas, 1996, Kusuoka, 1999, Collin-Dufresne, Goldstein, and Hugonnier, 2004).

[12] As noted by Hatchondo and Martinez (2009), this coupon structure can be interpreted as if the debt issued by the government consisted of a portfolio of zero-coupon bonds of different maturities, with portfolio weights that decline geometrically with maturity. This assumption allows to synthesize the repayment schedule, on any date, by means of a single number (the sum of future repayments, for instance). When this is not the case, all past issuances have to enter the state vector, which dramatically complicates the model solution.

[13] The modified duration of such an instrument has the specificity to be equal to its price. Indeed, we have $\mathcal{P} = 1/(1 + q - \chi)$, where q is the perpetuity's yield-to-maturity (eq. 1.9), which implies $(\partial\mathcal{P}/\partial q)/\mathcal{P} = -\mathcal{P}$.

nominal amount on date $t+h$:

$$\chi^{h-1}[1 \times (1-\mathcal{D}_{t+h}) + RR \times \mathcal{D}_{t+h}].$$

The price of the perpetuity therefore is:

$$\mathcal{P}_t = \sum_{h=1}^{\infty} \chi^{h-1} B_{t,h}, \tag{1.5}$$

where the $B_{t,h}$'s are prices of binary zero-coupon providing the payoffs $(1-\mathcal{D}_{t+h}) + RR\mathcal{D}_{t+h}$ on date $t+h$. Appendix 1.B.1 shows that, under the assumption that

$$RR = \exp(-\gamma b_c - b_\pi), \tag{1.6}$$

we have:

$$B_{t,h} = \exp(B_h + A_h' w_t), \tag{1.7}$$

where the B_h's and the A_h's derive from simple Riccati's equations (see eqs. a.1.6, a.1.7 and a.1.8 in Appendix 1.B.1). The continuously-compounded yield-to-maturity associated with a government zero-coupon of maturity h is therefore given by

$$r_{t,h} = -\frac{1}{h}(B_h + A_h' w_t). \tag{1.8}$$

Though *ad hoc*, the recovery-rate assumption (1.6) brings essential simplification in our model. It avoids solving a fixed-point problem to determine the price of the perpetuity, which is needed to derive the debt accumulation process (next subsection). Importantly, as will be shown in Subsection 1.4.2, the recovery-rate assumption (1.6) is satisfied for reasonable parameter values. Let us stress that while this assumption dramatically simplifies the pricing of defaultable bonds (eq. 1.7) and leads to government yields that are affine in w_t (eq. 1.8), it is not the case for risk-free yields (see, e.g., Appendix 1.H.1). But the model is much more tractable when government yields are affine—instead of risk-free ones—because government yields are the ones appearing in the debt accumulation process.

Denote by q_t the perpetuity's yield-to-maturity. It satisfies:

$$\mathcal{P}_t = \sum_{h=1}^{\infty} \frac{\chi^{h-1}}{(1+q_t)^h} = \frac{1}{1+q_t-\chi}. \tag{1.9}$$

Together, eqs. (1.5) and (1.9) determine q_t. Thanks to eq. (1.7), the solution for q_t is explicit. This exact solution is not exactly affine in w_t, but it is not far from being so. Indeed, as shown by eq. (1.5) the perpetuity can be seen as a weighted sum of zero-coupon bonds whose yields-to-maturity are affine in w_t (see eq. 1.8). We therefore expect q_t to be close to one of these $r_{t,h}$'s and, therefore, to be approximately affine in w_t. Appendix 1.H.2 details how we proceed to select a maturity h^* satisfying:

$$q_t \approx r_{t,h^*} = -\frac{1}{h^*}(B_{h^*} + A'_{h^*} w_t). \tag{1.10}$$

We exploit the previous representation to obtain an affine debt accumulation process, as shown in the next subsection.

1.3.5 Debt dynamics

When the government issues the perpetual bonds presented in the previous subsection, Appendix 1.B shows that the apparent interest rate approximately takes the form of a weighted sum of the past perpetuity's yields-to-maturity, with weights decaying geometrically at rate χ. That is:

$$\frac{R_{t+1}}{D_t} \approx (1-\chi)q_t + \chi \frac{R_t}{D_{t-1}}, \tag{1.11}$$

where R_t denotes the date-t nominal debt service and where D_t denotes the face value of the debt on date t.[14] Consistently with international debt accounting standards—on which our data are based—the concept of debt valuation we opt for is that of "nominal valuation of debt securities," where the debt outstanding covers the sum of funds originally advanced, plus any subsequent advances, less any repayments, plus any accrued interest.[15,16]

Let us introduce the following notations:

$$rd_t = \frac{R_t}{D_{t-1}} - \bar{q} \qquad sd_t = \frac{S_t}{D_{t-1}} - \overline{sd}, \tag{1.12}$$

[14] Using the vocabulary of Mauro and Zhou (2020), while R_t is the effective cost of servicing debt, q_t reflects the marginal funding cost (Subsection 2.1 of Mauro and Zhou, 2020).

[15] See International Monetary Fund, Bank for International Settlements and European Central Bank (2015). Although such a precision is innocuous in the context of models considering only short-term issuances, it is not in the present context, where the government issues long-dated debt instruments. This accounting concept is used, specifically, in eq. (a.1.9) of Appendix 1.B.2.

[16] This debt concept does not coincide with the market value of debt. The latter, denoted by $\mathscr{D}_t$ in Appendix 1.E, is the concept that one uses to derive the standard relationship between debt and the net present values of future primary surpluses. See also Jiang, Lustig, Nieuwerburgh, and Xiaolan (2019).

where $\bar{q}$ and $\overline{sd}$ are the respective unconditional means of the perpetuity yield and of the surplus-to-debt ratio (S_t/D_{t-1}), and where S_t denotes the nominal value of the date-t primary surplus, whose dynamics is discussed in the next subsection. The unconditional means of both rd_t and sd_t are zero. With this notation, eq. (1.11) rewrites:

$$rd_{t+1} \approx (1-\chi)(q_t - \bar{q}) + \chi rd_t. \tag{1.13}$$

If $rd_t - sd_t$ is small, and in the absence of default on date t, Appendix 1.B.2 shows that we get the following approximated law of motion for d_t, the logarithm of the debt-to-GDP ratio:

$$d_t \approx d_{t-1} - \Delta y_t - \pi_t + \log\left(1 + \bar{q} - \overline{sd}\right) + \frac{1}{1 + \bar{q} - \overline{sd}}(rd_t - sd_t). \tag{1.14}$$

The previous equation is notably consistent with the fact that an increase in nominal GDP growth—coming either from higher real growth Δy_t or from higher inflation π_t—results in a decrease in the debt-to-GDP ratio.

The deviation between the surplus-to-debt ratio and its unconditional mean, denoted by sd_t, follows:

$$sd_t = \gamma_d(d_{t-1} - \bar{d}) + \Lambda_s' w_t + \sigma_s' \eta_t, \tag{1.15}$$

where $\bar{d}$ denotes the unconditional mean of d_t. The term $\gamma_d(d_{t-1} - \bar{d})$ captures the reaction of the primary surplus to the debt level. If $\gamma_d > 0$, the government increases the primary surplus in response to rising public debt; this mechanism is consistent with the empirical evidence provided by Bohn (1998) and Mendoza and Ostry (2008).

1.3.6 Fiscal limit

In the spirit of Bi (2012), the fiscal limit is defined as the sum of present values of maximum future budget surplus (see Appendix 1.E). Let us denote by $\mu_{s,t}^*$ the maximum primary surplus, expressed as a fraction of the nominal GDP (i.e. $Y_t \times P_t$). The maximum primary surplus can be understood as the surplus implicit in the peak of the Laffer curve—the reverse bell-shaped relationship between the average tax rate and government revenues. We assume that $\mu_{s,t}^*$ is i.i.d. distributed, of mean $\mu_s^* > 0$.

We have:

$$\begin{aligned}\exp(\ell_t) &= \frac{1}{Y_t P_t}\mathbb{E}_t\left(\sum_{h=1}^{+\infty}\mathcal{M}^n_{t,t+h}\left[\mu^*_{s,t}Y_{t+h}P_{t+h}\right]|\mathcal{D}\equiv 0\right)\\ &= \mu^*_s\mathbb{E}_t\left(\sum_{h=1}^{+\infty}\mathcal{M}_{t,t+h}\exp(\Delta y_{t+1}+\cdots+\Delta y_{t+h})|\mathcal{D}\equiv 0\right),\end{aligned}$$

or

$$\ell_t = \log(\mu^*_s) + \log\left(\sum_{h=1}^{+\infty}\mathbb{E}_t\left[\mathcal{M}_{t,t+h}\exp(\Delta y_{t+1}+\cdots+\Delta y_{t+h})|\mathcal{D}\equiv 0\right]\right), \tag{1.16}$$

where Δy_t and π_t are the (log) growth rates of the real GDP and of the deflator, respectively, and where $\mathcal{M}_{t,t+h}$ and $\mathcal{M}^n_{t,t+h}$ respectively denote the real and nominal stochastic discount factors, or s.d.f., between dates t and $t+h$.[17]

To get insight into eq. (1.16), consider the deterministic case where $\mathcal{M}_{t,t+1} = \exp(-r_t)$, r_t being the real risk-free rate. Assume further that $\Delta y_{t+1} - r_t \geq \varepsilon$ for all dates t, where ε is strictly positive. In this situation, discussed by Blanchard (2019a), ℓ_t is infinite. But this is not necessarily the case in general.[18]

Our framework offers approximations of ℓ_t that are simple to implement numerically. Eq. (1.16) shows that the computation of the fiscal limit is deduced from that of conditional expectations of linear-exponential affine transformations of Δy_t and Δc_t (as $\mathcal{M}_{t-1,t}$ depends on Δc_{t+1}, see eq. 1.4). Moreover, eq. (1.1) implies that, conditionally on $\mathcal{D}_t \equiv 0$, the latter two variables linearly depend on $[w'_t, \eta'_t]'$, which itself follows a Gaussian vector auto-regressive process (VAR)—since η_t is i.i.d. normal and w_t follows a VAR(1) (eq. 1.2). This guarantees the existence of closed-form formulas to obtain each of the conditional expectations entering eq. (1.16). Computational details are provided in Appendix 1.C.

1.3.7 Sovereign default probability

The last modeling ingredient pertains to the conditional default probability of the government. This probability is a decreasing function of the fiscal space $\ell_t - d_t$. A possibility is to have

[17]The two types of s.d.f. are linked through $\mathcal{M}^n_{t,t+k} = \mathcal{M}_{t,t+h}\exp(-\pi_{t+1} - \cdots - \pi_{t+k})$ (see, e.g., Campbell and Viceira, 2001, eq. 10). Note also that the multi-period s.d.f. $\mathcal{M}_{t,t+h}$ is the product of the one-period s.d.f., that is $\mathcal{M}_{t,t+h} = \prod_{k=1}^{h}\mathcal{M}_{t+k-1,t+k}$, where $\mathcal{M}_{t+k-1,t+k}$ is given by (1.4) (the same holds true for $\mathcal{M}^n_{t,t+h}$).

[18]If both $\log\mathcal{M}_{t,t+1}$ and Δy_t are Gaussian, as is the case in the present model (in the absence of default), then a necessary condition for the fiscal limit to be finite is that $\mathbb{E}(\Delta y - r) < 0$, where $r = -\log\mathbb{E}(\mathcal{M}_{t,t+1})$.

a default probability of zero (respectively one) when $\ell_t > d_t$ (respectively $\ell_t \leq d_t$). In this case, the fiscal limit is "strict," in the sense that default is automatically triggered when the limit is breached. However, in order to capture non-modeled factors that may precipitate or delay default—e.g. political factors (Hatchondo and Martinez, 2010)—we introduce a Gaussian white noise ν_t, of variance σ_ν^2, and assume that the probability of default depends on $\ell_t - d_t + \nu_{t+1}$, which is a noisy measure of the fiscal space. Specifically, the conditional probability of observing a sovereign default on date $t+1$ is of the form:

$$\mathbb{P}(\mathcal{D}_{t+1} = 1 | \mathcal{D}_t = 0, \mathcal{I}_t, \nu_{t+1}) = \mathcal{F}(\ell_t - d_t + \nu_{t+1}), \tag{1.17}$$

where $\mathcal{F}$ is a function valued in $[0,1]$. Function $\mathcal{F}$ is such that $\mathcal{F}(u) = 0$ for $u \geq 0$, implying that the default probability is equal to zero as long as the (noisy) fiscal space debt is nonnegative. Moreover, function $\mathcal{F}$ is increasing: the larger the distance between debt and the fiscal limit, the higher the probability of default. In the following, we employ the following specification for $\mathcal{F}$:

$$\mathcal{F}(d_t - \ell_t - \nu_{t+1}) = 1 - \exp(-\underbrace{\max[0, \alpha(d_t - \ell_t - \nu_{t+1})]}_{=\underline{\lambda}_{t+1}}), \tag{1.18}$$

with $\alpha > 0$. Hereinafter, we refer to $\underline{\lambda}_{t+1} \equiv \alpha \max(0, d_t - \ell_t - \nu_{t+1})$ as the default intensity. According to eqs. (1.17) and (1.18), when it is small, the default intensity $\underline{\lambda}_{t+1}$ is close to the conditional probability of default $\mathbb{P}(\mathcal{D}_{t+1} = 1 | \mathcal{D}_t = 0, \mathcal{I}_t, \nu_{t+1})$.

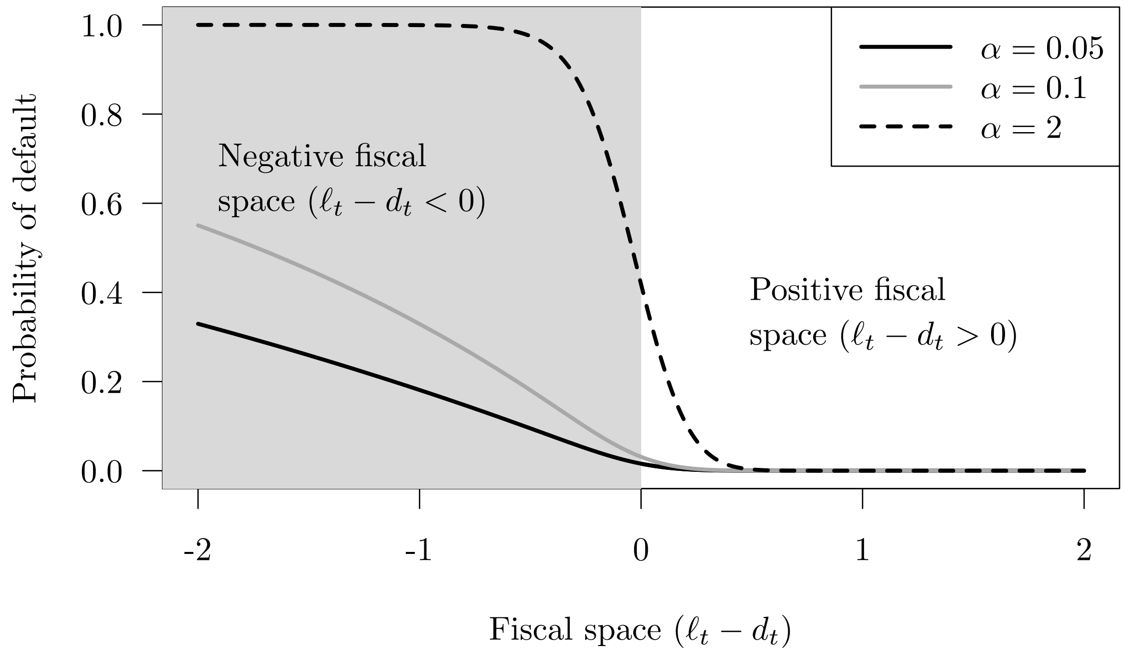
The parameter α is another modeling ingredient—on top of ν_t—that makes it possible to control for the strictness of the fiscal limit. This is illustrated by Figure 1.1, that displays the probabilities of observing a default over a year, conditional on a given level of the (log) fiscal space $d_t - \ell_t$, and for different values of α. If α is large, the fiscal limit is strict, in the sense that default is likely to happen as soon as $d_t > \ell_t$. By contrast, if α is small, the fiscal limit is softer, in the sense that, for the same value $d_t - \ell_t > 0$, a sovereign default on date t is possible, but less likely.

1.3.8 The dynamics of the full state vector

Let us define x_t as follows:

$$x_t = [w_t', d_t, rd_t, q_t, w_{t-1}, d_{t-1}, \nu_t]'.$$

Figure 1.1: Annual probability of default with respect to fiscal space ($\ell_t - d_t$)



Note: This figure shows the posited relationship between sovereign default probability and fiscal space ($\ell_t - d_t$). More precisely, it shows $\mathbb{P}(\mathcal{D}_{t+4} = 1 | \mathcal{D}_t = 0, \ell_{t+i} - d_{t+i} = u, i = 0, \dots, 3)$. According to eqs. (1.17) and (1.18), conditionally on the fiscal space $\ell - d$ and on the noise ν (with $\nu \sim i.i.d.\mathcal{N}(0, \sigma_\nu^2)$, and $\sigma_\nu = 0.2$ here), the (one-quarter) probability of default is $1 - \exp(-\underline{\lambda})$, where $\underline{\lambda} = \max(0, d - \ell - \nu)$. The probability of default is therefore strictly positive only when the noisy fiscal space ($\ell + \nu - d$) is strictly negative, and null otherwise. Standard results on truncated normal distributions allow to compute the default probability conditional on the fiscal space only ($\ell - d$)—i.e. integrating out over all possible values of the noise ν. Formally: $\mathbb{P}(\mathcal{D}_{t+1} = 1 | \mathcal{D}_t = 0, \ell_t - d_t = u) = 1 - f(u)$, with $f(u) = \Phi\left(u/\sigma_\nu\right) + \exp\left(\alpha u + \alpha^2\sigma_\nu^2/2\right)\left(1 - \Phi\left(u/\sigma_\nu + \sigma_\nu\alpha\right)\right)$. The one-year default probabilities—displayed on the figure—are given by $1 - f(u)^4$.

The fact that lagged values of w_t and d_t are included in x_t allows to express the default intensity $\underline{\lambda}_t$ as $\max(0, a + b'x_t)$, consistently with (1.18).[19]

The model described above is such that, as long as there is no default until date t, the dynamics of x_t satisfies

$$x_t \approx \mu_x + \Phi_x x_{t-1} + \Sigma_x \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}, \tag{1.19}$$

where matrices μ_x, Φ_x and Σ_x are deduced from eqs. (1.2), (1.10), (1.13), (1.14) and (1.15). These matrices are detailed in Appendix 1.K.

Remark that d_t depends on Δy_t and π_t (eq. 1.14), and these two variables themselves depend on $\mathcal{D}_t$ (eq. 1.1). Hence, if a default occurs on date t, (1.19) is not valid any longer, as a term in $\mathcal{D}_t$

[19] This formulation is required in the formulas presented in Appendix 1.F.2.

would then appear on the right-hand side. However, given that the date-t default probability depends on d_{t-1} (see eq. 1.17), and since the CDS payoff do not depend on what happens after the default date (next subsection), there is no need to know x_t's dynamics after the default date to price CDSs.

1.3.9 Credit Default Swaps

Our estimation approach employs Credit Default Swap (CDS) prices. A CDS is an agreement between a protection buyer and a protection seller, whereby the buyer pays a periodic fee in return for a contingent payment by the seller upon a credit event—such as bankruptcy or failure to pay—of a reference entity. The contingent payment usually replicates the loss incurred by a creditor of the reference entity in the event of its default (see, e.g., Duffie, 1999).

More precisely, a CDS works as follows: the protection buyer pays a regular premium to the so-called protection seller. These payments end either after a given period of time—the maturity of the CDS, that we denote by h—or at default of the reference entity. Upon the default of the debtor, the protection seller compensates the protection buyer for the loss incurred, assuming the latter was holding defaulted bonds. Following the "Recovery of Treasury" (RT) convention of Duffie and Singleton (1999), we assume that the bond recovery payment, upon default, is a fraction RR of the price of a risk-free zero-coupon bond of equivalent residual maturity. If t is the inception date of a maturity-h CDS, the amount paid on date $t+k$ (with $0 < k \leq h$) by the protection seller to the protection buyer is:

$$(\mathcal{D}_{t+k} - \mathcal{D}_{t+k-1})(1 - RR)\mathbb{E}_{t+k}(\mathcal{M}^n_{t+k,h-k}),$$

where $\mathbb{E}_{t+k}(\mathcal{M}^n_{t+k,h-k})$ is the price, as of date $t+k$, of a nominal risk-free bond of residual maturity $h-k$. Conversely, on date $t+k$, the protection buyer pays

$$S^{cds}_{t,h}(1 - \mathcal{D}_{t+k})$$

to the protection seller, where $S^{cds}_{t,h}$ denotes the CDS premium—as negotiated on date t—expressed in percentage of the notional.

At inception of the CDS contract (date t), there is no cash-flow exchanged between both parties; that is, the CDS spread $S^{cds}_{t,h}$ is determined so as to equalize the present discounted values of the payments promised by each of them. Assuming the reference entity has not defaulted before date t, i.e. $\mathcal{D}_t = 0$, we have:

$$\underbrace{\mathbb{E}_t \left\{ \sum_{k=1}^{h} \mathcal{M}^n_{t,t+k}(\mathcal{D}_{t+k} - \mathcal{D}_{t+k-1})(1 - RR)\mathbb{E}_{t+k}(\mathcal{M}^n_{t+k,h-k}) \right\}}_{\text{Protection leg}} = \underbrace{S^{cds}_{t,h}\mathbb{E}_t \left\{ \sum_{k=1}^{h} \mathcal{M}^n_{t,t+k}(1 - \mathcal{D}_{t+k}) \right\}}_{\text{Premium leg}}, \tag{1.20}$$

which gives, after some algebra (see Appendix 1.F):

$$S^{cds}_{t,h} = (1 - RR)\frac{\mathbb{E}_t \left\{ \mathcal{M}^n_{t,t+h} D_{t+h} \right\}}{\mathbb{E}_t \left\{ \sum_{k=1}^{h} \mathcal{M}^n_{t,t+k}(1 - \mathcal{D}_{t+k}) \right\}}. \tag{1.21}$$

Therefore, the CDS spread $S^{cds}_{t,h}$ is deduced from the knowledge of the following conditional expectations: $\mathbb{E}_t[\mathcal{M}^n_{t,t+k}\mathcal{D}_{t+k}]$ for $k \in \{1, \dots, h\}$, and $\mathbb{E}_t[\mathcal{M}^n_{t,t+k}(1 - \mathcal{D}_{t+k})]$. The former (respectively latter) expectation corresponds to the date-t price of a zero-coupon bond that provides a unit payoff on date $t + k$ if the reference entity has defaulted before date $t + k$ (respectively has not defaulted before date $t + k$), and zero otherwise.

Appendix 1.A shows that these two conditional expectations can be rewritten as expectations of exponential linear combinations of future values of $[x'_t, \underline{\lambda}_t]'$, with $\underline{\lambda}_t = \max(0, \lambda_t)$, where $\lambda_t \equiv \alpha(d_{t-1} - \ell_{t-1} + \nu_t)$. Moreover, the model developed above is such that λ_t is approximately affine in x_t, i.e., of the form $a + b'x_t$.[20] Therefore, in our framework, pricing CDSs amounts to computing conditional expectations of (exponential) linear combinations of $[x'_t, \max(0, a + b'x_t)]'$. This problem is reminiscent of that arising in the context of shadow-rate models *à la* Black (1995).[21] Shadow-rate models have attracted a lot of interest over the last decade. The reason is that these models accommodate the existence of a lower bound for nominal interest rates, a welcome feature in a context of extremely low yields. Though simple and intuitive, this framework does not offer closed-form bond pricing formulas because of the non-linearity stemming from the max operator. Different approaches have however been introduced to approximate bond prices in shadow-rate models. Wu and Xia (2016) have notably proposed a particularly simple and accurate approximation. An adaptation of this approach to the present context is detailed in Appendix 1.F.2.[22]

[20] The approximation stems from the fact that the log fiscal limit ℓ_t, defined in (1.16), is not affine in x_t. Appendix 1.C shows how it can be approximated by an affine transformation of x_t.

[21] In shadow-rate models, nominal zero-coupon bond prices are given by $\mathbb{E}(\exp(-i_t - \dots - i_{t+h-1}))$, where the short-term nominal interest rate i_t is equal to $\max(0, s_t)$, s_t ($\in \mathbb{R}$) being the shadow rate.

[22] Our approach shares some similarities with the Black-Scholes-Merton model (Black and Scholes, 1973; Merton, 1974) (and its numerous extensions) in that it also features a default threshold. As noted by Duffie and Single-

1.4 Estimation

Bringing the model to the data amounts to estimating two types of objects: model parameters and latent variables, i.e. the components of w_t. To reduce the number of free parameters, we calibrate some of them (Subsection 1.4.2). The estimation of the remaining parameters and of the latent variables is based on maximum-likelihood (ML) techniques (Subsection 1.4.3). The next subsection briefly describes the data, additional details are gathered in Appendix 1.M.

1.4.1 Data

We consider eight countries: Brazil, Canada, China, India, Japan, Russia, the United Kingdom, and the United States. According to the IMF sovereign debt investor bases for advanced economies and emerging markets, more than three quarters of sovereign debts are held domestically in these eight countries (IMF sovereign debt investor bases build on Arslanalp and Tsuda, 2014). The fact that most bondholders are domestic helps make the data consistent with our closed-economy framework.

The data are quarterly. Estimation samples vary across countries; on average, they cover the last 13 years. Given that sovereign CDSs for India have been traded only for a few years, we proxy for these CDSs with those written on the State Bank of India, which is standard practice (see, e.g., de Boyrie and Pavlova, 2016). Financial data come from CMA and Refinitiv. For each country, we use government yields of three maturities (2, 5, and 10 years), and CDS spreads of 5 maturities (1, 2, 3, 5, and 10 years). These data are respectively gathered in a 3-dimensional vector yds_t and a 5-dimensional vector CDS_t. We refer to Tables from 1.M1 to 1.M8 in the Appendix for further details on the data.

1.4.2 Calibrated parameters

Table 1.1 reports calibrated and estimated parameters. This subsection presents the calibrated ones, that are those that are not obtained by the numerical maximization of the likelihood function.

ton (2003, Subsection 3.2.2), the tractability of the Black-Scholes-Merton model rapidly declines as one allows for a time-varying default threshold. Although our framework features a time-varying debt threshold, tractability is preserved thanks to the use of the approximation formulas developed in the shadow-rate context.

We set the annual rate of time preference to 0.96 and 0.98 for emerging and advanced economies, respectively.[23] The coefficient of relative risk aversion is set to 4, a value used for instance by Barro (2006) and Gabaix (2012).

We assume that a sovereign default results in a consumption drop of 20% (i.e. $b_c = 0.2$ in eq. 1.1), which broadly corresponds to the average disaster magnitude documented by e.g. Barro and Ursua (2011). Though larger than the average recessionary effect associated with sovereign defaults documented by Mendoza and Yue (2012) (5%), or by Reinhart and Rogoff (2011) (7%), an output fall of 20% is commensurate to estimated losses resulting from a sovereign default combined with a banking crisis (De Paoli, Hoggarth, and Saporta, 2006) or from "hard defaults"—defined by Trebesch and Zabel (2017) as those defaults resulting in large losses for investors. As regards the inflationary effect of a default, we use $b_\pi = -2.1\%$, which is the average inflationary effect of a disaster used by Gabaix (2012). In our framework, b_π is used only to calibrate the recovery rate, which is discussed in the next paragraph.

As explained in Subsection 1.3.4, the tractability of the model notably hinges on the assumption formulated in (1.6), i.e., $RR = \exp(-\gamma b_c - b_\pi)$. Given the calibrated values of γ, b_c and b_π, this gives $RR = 46\%$, which turns out to be a reasonable value: according to sovereign defaults data collected by Moody's (2019, Exhibit 20), the value-weighted (respectively issuer-weighted) average recovery rate is of 41% (resp. 55%) over the last 35 years.

The decay rate of the perpetuity's coupons χ is set to 0.92. Implementing the approach mentioned at the end of Subsection 1.3.4, and detailed in Appendix 1.H.2, this choice leads to average durations (h^*) of the perpetuities varying between about 3 and 5 years across countries (ninth line of Table 1.1). These values are roughly in line with average maturities of sovereign issuances.

[23] Setting the rate of time preference to 0.98 for emerging economies implies that $\mathbb{E}(\Delta y) > r$ for these countries, leading to infinite fiscal limits (see Footnote 18). This difference is supported, for instance, by the empirical findings of Falk et al. (2018): using the Global Preference Survey (GPS), an experimentally validated survey data set of time preference and risk preference from 80000 people in 76 countries, Falk et al. (2018) find that inhabitants of low-income and emerging economies are less patient than ones living in advanced economies. Using data from six European advanced economies, De Lipsis (2021) estimates an average discount rate equal to 0.02 across income quantiles, which translates into a rate of time preference equal to 0.98 (in line with our calibration for advanced economies). Furthermore, Arellano (2008) sets the rate of time preference to 0.953 in a small open economy model of endogenous default replicating the characteristics of Argentina, in line with our calibration for emerging markets.

The dimension of w_t, as that of η_t, is set to four. We further posit that Φ (eq. 1.2) is of the form:

$$\Phi = \begin{bmatrix} \phi_{1,1} & \phi_{1,2} & 0 & 0 \\ 0 & \phi_{2,2} & 0 & 0 \\ 0 & 0 & \phi_{3,3} & \phi_{3,4} \\ 0 & 0 & 0 & \phi_{4,4} \end{bmatrix}. \tag{1.22}$$

This specification implies that $w_{2,t}$ and $w_{4,t}$ are autonomous autoregressive processes of order one. These two processes respectively Granger-cause $w_{1,t}$ and $w_{3,t}$. Together, the four components of w_t account for the persistent fluctuations of Δc_t and Δy_t (see eq. 1.1), as well as sd_t (eq. 1.15). For parsimony, inflation is assumed not to depend on w_t, that is, $\Lambda_\pi = 0$.

We assume that the persistent parts of Δc_t and Δy_t, namely $\mu_c + \Lambda_c' w_t$ and $\mu_y + \Lambda_y' w_t$, are identical (i.e. $\mu_c = \mu_y$ and $\Lambda_c = \Lambda_y$). In addition, Λ_c and Λ_y only load on $w_{1,t}$, the first component of w_t (i.e., $\Lambda_c = \Lambda_y = [\Lambda_{c,1}, 0, 0, 0]'$). Therefore, given the shape of Φ (eq. 1.22), Δc_t and Δy_t are driven by $w_{1,t}$, and also by $w_{2,t}$ (through $w_{1,t}$), but not by $w_{3,t}$ and $w_{4,t}$. The latter two processes affect the budget surplus sd_t. In order to capture the procyclicality of budget surplus, sd_t also depends on $w_{1,t}$. More precisely, eq. (1.15) takes the form:

$$sd_t = \gamma_d(d_{t-1} - \bar{d}) + \Lambda_{s,1} w_{1,t} + \Lambda_{s,3} w_{3,t} + \sigma_s' \eta_t. \tag{1.23}$$

Additional restrictions on the components of Λ_i are presented in Appendix 1.D. These restrictions impose that some model-implied moments of the macroeconomic variables match sample counterparts. These restrictions also involve the components of the σ_i vectors, which determine the covariances of the volatile components of the macroeconomic variables.

The estimation of γ_d, the effect of debt on surplus (eq. 1.15), proved to make the optimization of the likelihood function numerically unstable. (Indeed debt then turns out to be explosive for many sets of parameters considered by the numerical routines used to optimize the likelihood function.) To address this issue, we set γ_d to an arbitrary low value of 0.01. This approach, which appears to have only mild effects on the results, suffices to solve the numerical problem.

1.4.3 Maximum Likelihood (ML) estimation strategy

For each country, the remaining parameters are estimated by ML techniques.

The model can be cast into a state-space form, with (i) transition equations describing the dynamics of the latent vector w_t (this is eq. 1.2) and (ii) measurement equations describing the relationships between w_t and observed financial market data—CDSs and yield spreads—as well as macroeconomic variables (Δc_t, Δy_t, π_t, and sd_t). Formally, the state-space model is of the form:

$$(i) \qquad w_t = \Phi w_{t-1} + \varepsilon_t, \quad \text{(this is eq. 1.2)}$$

$$(ii) \quad \begin{bmatrix} CDS_t \\ yds_t \\ \Delta c_t \\ \Delta y_t \\ \pi_t \\ sd_t \end{bmatrix} = \begin{bmatrix} \mathcal{S}(w_t, d_t, rd_t, sd_t) \\ A + Bw_t \\ \Lambda_c' w_t \\ \Lambda_y' w_t \\ \Lambda_\pi' w_t \\ \Lambda_s' w_t \end{bmatrix} + \begin{bmatrix} \sigma_{CDS} I_{5\times 5} & 0 & 0 \\ 0 & \sigma_{yds} I_{3\times 3} & 0 \\ 0 & 0 & \sigma_c' \\ 0 & 0 & \sigma_y' \\ 0 & 0 & \sigma_\pi' \\ 0 & 0 & \sigma_s' \end{bmatrix} \begin{bmatrix} \xi_{CDS,t} \\ \xi_{yds,t} \\ \eta_t \end{bmatrix}, \tag{1.24}$$

where $\mathcal{S}$ is a function computing model-implied CDS spreads (see Subsection 1.3.9), and where vector A and matrix B define the model-implied affine relationship between government zero-coupon yields and w_t (see eq. 1.7). The vectors $\xi_{CDS,t}$, $\xi_{yds,t}$, ε_t, and η_t contains i.i.d. zero-mean, unit-variance Gaussian shocks; the first two of these vectors are measurement errors.

Because function $\mathcal{S}$ is nonlinear in w_t, we resort to the extended Kalman filter to get an approximation to the likelihood function. Maximizing this function with respect to $\phi_{1,1}$, $\phi_{1,2}$, $\phi_{2,2}$, $\phi_{3,3}$, $\phi_{3,4}$, $\phi_{4,4}$, $\Lambda_{s,1}$, μ_s^*, α, and σ_ν provides us with Quasi-Maximum Likelihood (QML) estimates of these model parameters. In order to discipline and facilitate the numerical optimization, we impose bounds on the last three parameters. Specifically, we impose that: (i) the average maximum budget surplus (μ_s^*) is larger than 0.5%, (ii) that the standard deviation fiscal space noise is lower than 20% of debt-to-GDP (i.e., $\sigma_\nu < 0.2$), and (iii) that the model-implied standard deviation of the persistent component of sd_t (that is $\Lambda_{s,1} w_{1,t} + \Lambda_{s,3} w_{3,t}$, see eq. 1.23) is smaller than twice is sample counterpart, (iv) α is larger than 0.05. The latter restriction avoids having a too soft concept of fiscal limits. Bear in mind that when $d_t > \ell_t$, the larger α, the higher the default probability (eq. 1.18). In other words, for a given level of default probability (or CDS), the larger α, the higher the fiscal limit estimate. When allowing for the estimation of α, the data would call for small values for many countries, implying low values of the fiscal limit—often below the actual debt level—which is at odds with our intended interpretation of this limit. A minimal value of 0.05 for α offers a good compromise between data fitting performances and interpretability of the fiscal limit.

We report the resulting model parametrizations in Table 1.1.[24] The estimated (filtered) time series of the latent factors $w_{i,t}$, $i = 1, \ldots, 4$ are shown on Figure 1.N2 of Appendix 1.N. It is worth mentioning that the estimation of the eight models—one for each country—takes a handful of minutes on a standard laptop. We perform numerical computations using C++, via the Rcpp library of R. This way, one evaluation of the likelihood function, which involves solving the model, computing prices, and running the Kalman filter algorithm, takes a fraction of a second.

Table 1.1: Models' parameterization

Notation	Mult.	US	CA	UK	JP	BR	RU	IN	CN
RR (eq. 1.21)	$\times 10^2$	46	46	46	46	46	46	46	46
χ (Sub. 1.3.5)	$\times 10^2$	92	92	92	92	92	92	92	92
γ (eq. 1.4)		4	4	4	4	4	4	4	4
$b_c = b_y$ (eq. 1.1)	$\times 10^2$	20	20	20	20	20	20	20	20
$-b_\pi$ (eq. 1.1)	$\times 10^2$	2.1	2.1	2.1	2.1	2.1	2.1	2.1	2.1
$\exp(\bar{d})$ (eq. 1.15)	$\times 10^2$	98	111	76	218	69	40	47	40
α (eq. 1.18)	$\times 10^2$	5	5	15	5	5	5	7	5
σ_ν (eq. 1.18)	$\times 10^2$	20	20	10	12	20	20	13	10
h^* (eq. 1.10)		4.47	4.54	4.21	3.50	3.18	3.35	2.33	3.25
μ_{s^*} (eq. 1.16)	$\times 10^2$	0.51	1.64	0.90	3.52	1.55	0.50	0.50	0.50
sd_t (eq. 1.15)									
γ_d	$\times 10^2$	1.0	1.0	1.0	1.0	1.0	1.0	1.0	1.0
$\overline{sd}$ (eq. 1.12)	$\times 10^2$	0.47	0.48	0.46	0.47	0.92	0.94	0.53	0.83
$\Lambda_{s,3}$	$\times 10^2$	2.06	1.08	1.97	0.26	1.20	0.69	0.76	0.35
$\Vert\sigma_s\Vert$	$\times 10^2$	0.46	0.24	0.45	0.12	0.57	1.41	0.77	0.82
c_t (eq. 1.1)									
μ_c	$\times 10^2$	0.38	0.34	0.07	0.11	0.43	0.50	1.65	2.06
$\Lambda_{c,1}$	$\times 10^3$	1.13	1.96	4.70	4.56	4.92	7.59	5.95	3.76
$\Vert\sigma_c\Vert$	$\times 10^2$	0.56	0.40	0.82	0.67	0.93	0.32	2.24	1.29
y_t (eq. 1.1)									
μ_y	$\times 10^2$	0.38	0.34	0.07	0.11	0.43	0.50	1.65	2.06
$\Lambda_{y,1}$	$\times 10^3$	1.13	1.96	4.70	4.56	4.92	7.59	5.95	3.76
$\Vert\sigma_y\Vert$	$\times 10^2$	0.55	0.67	0.66	1.01	1.11	0.87	1.09	0.48

Cont'd on next page

[24] As regards the key parameter α, it is important to mention that the above-described restriction on the elasticity of the conditional probability of default to fiscal space turns out to be binding for all countries but the UK and India, for which this parameter is estimated.

Table 1.1 – *Cont'd from prev. page*

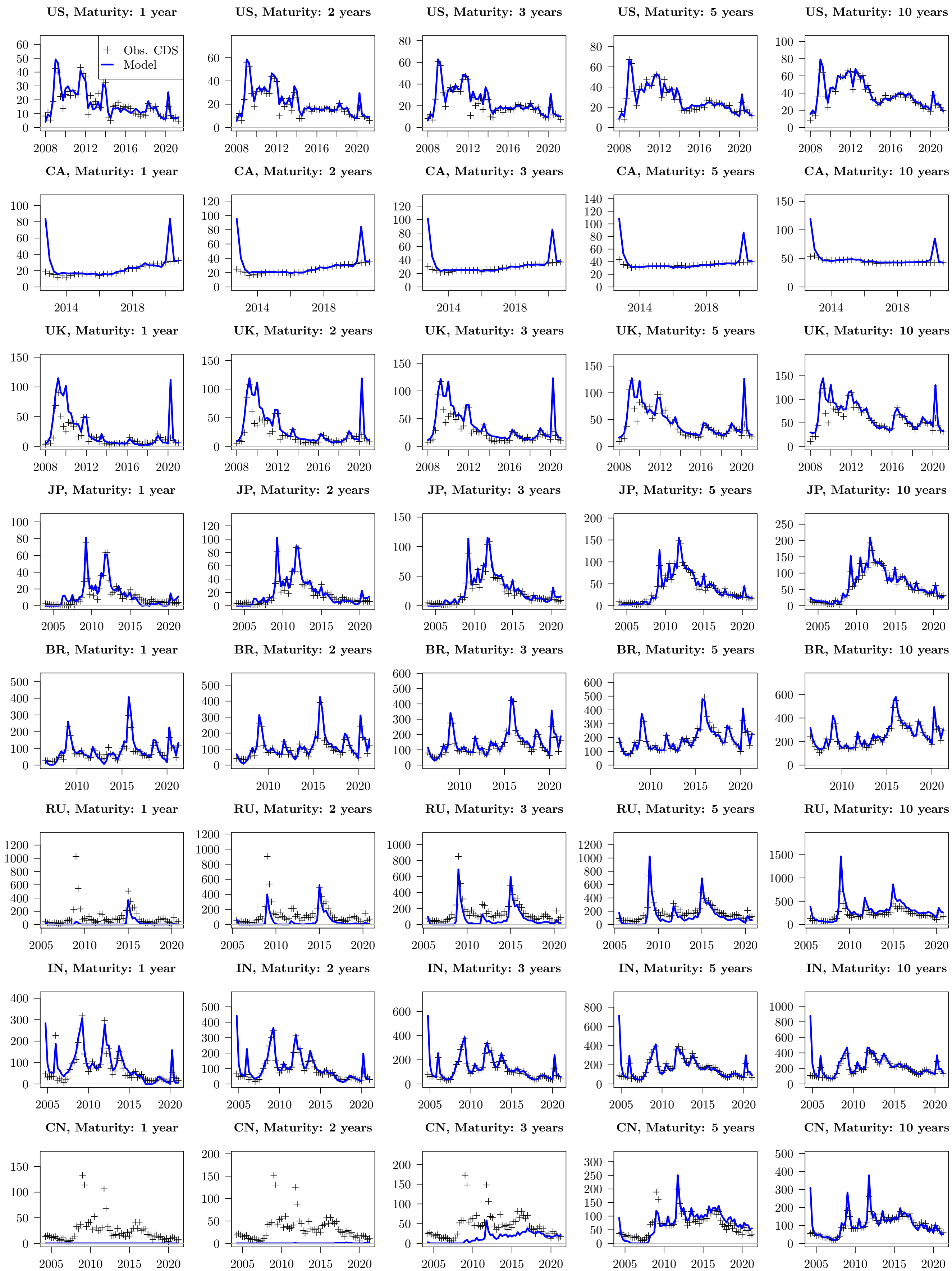
Notation	Mult.	US	CA	UK	JP	BR	RU	IN	CN
π_t (eq. 1.1)									
μ_π	$\times 10^2$	0.43	0.36	0.53	-0.07	1.74	2.10	1.33	0.88
$\|\sigma_\pi\|$	$\times 10^2$	0.23	0.68	0.54	0.35	0.83	2.21	0.89	0.86
w_t (eqs. 1.2 and 1.22)									
$\Phi_{1,1}$		0.954	0.406	0.000	0.389	0.658	0.000	0.402	0.328
$\Phi_{1,2}$		0.012	0.065	0.019	0.011	0.059	0.049	0.012	0.131
$\Phi_{2,2}$		0.983	0.984	0.994	0.992	0.974	0.971	0.990	0.981
$\Phi_{3,3}$		0.460	0.336	0.042	0.834	0.211	0.603	0.714	0.937
$\Phi_{3,4}$		0.097	0.041	0.036	0.192	0.318	1.010	0.234	0.014
$\Phi_{4,4}$		0.959	0.984	0.990	0.895	0.879	0.650	0.822	0.993

Note: This table presents the models' parameterizations. Part of the parameters are calibrated. The remaining ones are estimated by maximizing the log-likelihood associated with the model (see Subsections 1.4.2 and 1.4.3). RR denotes the recovery rate (eq. 1.21); χ is the decay rate of the perpetuities' payoffs (Sub. 1.3.5); γ stands for the risk aversion parameter (eq. 1.4); b_c is the consumption fall upon default (eq. 1.1); b_π is the inflation increase upon default (eq. 1.1); $\exp(\bar{d})$ is the average debt-to-GDP ratio (eq. 1.15); $\overline{sd}$ represents the average surplus-to-debt ratio (eq. 1.12); α is the elasticity of PD to $d_t - \ell_t$ (eq. 1.18); h^* is the duration of perpetuities, in years (eq. 1.10); μ_i, Λ_i, σ_i are the parameters employed in the specification of the exogenous macro block, $i = c, y, \pi$ (eq. 1.1); In the state-space model, the standard deviations of the measurement errors associated with yields and CDS spreads (i.e., σ_{yds} and σ_{CDS} in eq.1.24) are respectively set to 10% of the sample standard deviations of yields and CDS spreads.

1.4.4 Model fit

Figure 1.2 shows the fit of CDS spreads. In spite of the fact that the model relies on a framework that is more structural—and therefore more constrained—than in the term-structure sovereign spread literature (see references in Subsection 1.2.1), the fit is comparable with that found in intensity-based studies. It proves however challenging for the model to track CDS spreads in Canada and in the UK during the peak of the COVID-19 crisis (2020Q1). Also, the fit of the short-end of the Chinese CDS term structure is not as good as for other countries. Figure 1.N1 (in Appendix 1.N) shows that the model also captures a substantial share of the fluctuations in sovereign bond yields.

Figure 1.2: Observed vs model-implied CDS



Note: This figure compares observed CDS (crosses) with their model-implied counterparts (solid lines). Model-implied CDS spreads result from eq. (1.21).

1.5 Results

This section starts with a presentation of our fiscal limit and fiscal space estimates (Subsection 1.5.1). Next, Subsection 1.5.2 discusses the model implications in terms of spreads' sensitivity. Subsection 1.5.3 elaborates on the influence of risk premiums on model-implied default probabilities.

1.5.1 Fiscal limit estimates

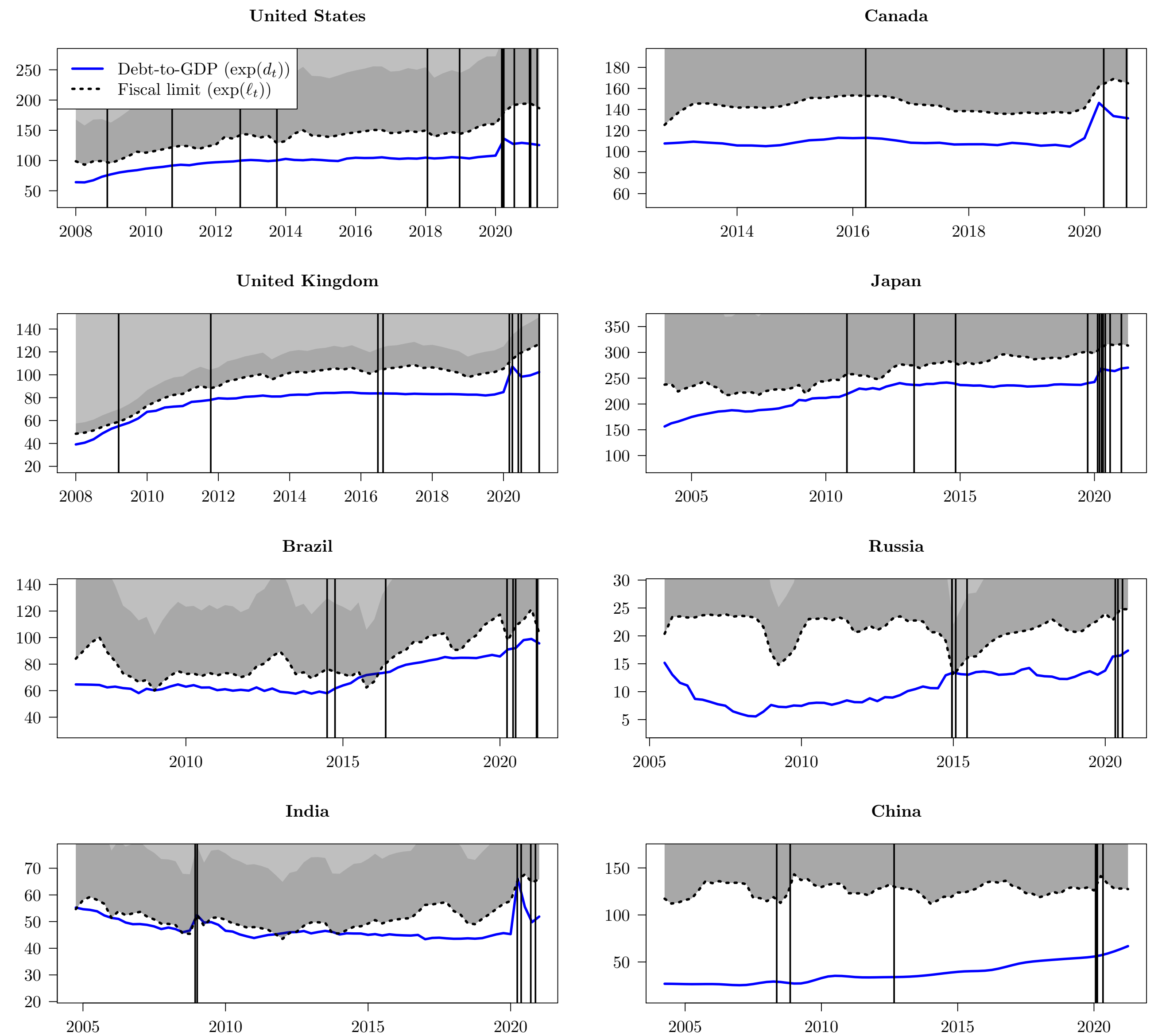
Even though policy makers and economic analysts have been increasingly focusing on fiscal sustainability, the literature building and studying measures of fiscal space/limits has been scarce. Kose, Kurlat, Ohnsorge, and Sugawara (2017) introduce an extensive multi-country dataset collecting variables that relate to the availability of budgetary resources for a government to service its financial obligations, but these measures are not directly interpretable as fiscal space measures. As explained in the literature review (Subsection 1.2.2), Ostry et al. (2010), Ghosh et al. (2013), Ostry et al. (2015) and Collard et al. (2015) compute debt limits based on the existence of a maximum primary surplus. The latter approaches however deliver static debt limit estimates. To the best of our knowledge, the present study is the first to propose time-varying estimates of fiscal limits. We will nevertheless compare our average estimates with static fiscal limits existing in the literature (Table 1.4).

Figures 1.3 and 1.4 respectively display fiscal limit and fiscal space estimates, expressed in percent of GDP. For both figures, vertical bars indicate key policy decisions and pivotal events; details regarding these dates are given in the caption of Figure 1.4. On a given quarter, if debt-to-GDP is higher than the fiscal limit ($\exp(\ell_t)$, dotted line), then the probability of default is strictly positive (see eqs. 1.17 and 1.18, with a noise $\nu = 0$). The shaded areas further give a sense of the strictness of the limit: everything else equal, if debt-to-GDP stays in the dark-shaded (respectively light-shaded) area for four quarters in a row, then the annual default probability of the considered country is larger than 10% (respectively between 0 and 10%). For what follows, and unless differently specified, our numbers refer to the threshold fiscal limit and fiscal space estimates, namely the dotted lines in Figures 1.3 and 1.4. Let us stress that the use of CDS data in the estimation is crucial to obtain series of fiscal limits and fiscal spaces that incorporate forward-looking elements. Indeed, sovereign credit spreads capture a timely assessment of debt sustainability dynamics by market participants that is translated into our fiscal limit and space estimates. Moreover, referring to eq. 1.16 in Subsection 1.3.6, fiscal limit (ℓ_t) at time t is a function of the (mean) maximum primary surplus and the present value of future GDP growth rates. Bear in mind that the GDP growth rate (Δy_t) is affine in the first two latent factors $w_{1,t}$ and $w_{2,t}$ (see Subsection 1.4.2). In our model, higher future and current GDP growths translate into a lower current fiscal limit: by implying higher interest rates and, thus, lower discount factors, a higher GDP growth path leads to a lower fiscal limit through the net present value computation. This interesting dynamics can be uncovered by the fact that the model-implied unconditional correlation existing between GDP growth (Δy_t) and the fiscal limit (ℓ_t) is negative.[25,26]

[25]The unconditional correlation is computed using $\mathbb{C}or(\Delta y_t, \ell_t) = \mathbb{C}ov(\Delta y_t, \ell_t)[\mathbb{V}ar(\Delta y_t)\mathbb{V}ar(\ell_t)]^{-\frac{1}{2}}$; where $\mathbb{C}ov(\Delta y_t, \ell_t) = \Lambda_y' \Omega_w a_\ell$ is the unconditional covariance between Δy_t and ℓ_t, $\mathbb{V}ar(\Delta y_t) = \Lambda_y' \Omega_w \Lambda_y + \sigma_y' \sigma_y$ is the unconditional variance of Δy_t, $\mathbb{V}ar(\ell_t) = a_\ell' \Omega_w a_\ell$ is the unconditional variance of ℓ_t and Ω_w is the unconditional variance of the unobserved factors w_t. This computation stems from the fact that we are able to compute the unconditional variance of the state vector in closed form and that the variables of interest are linear in the state vector.

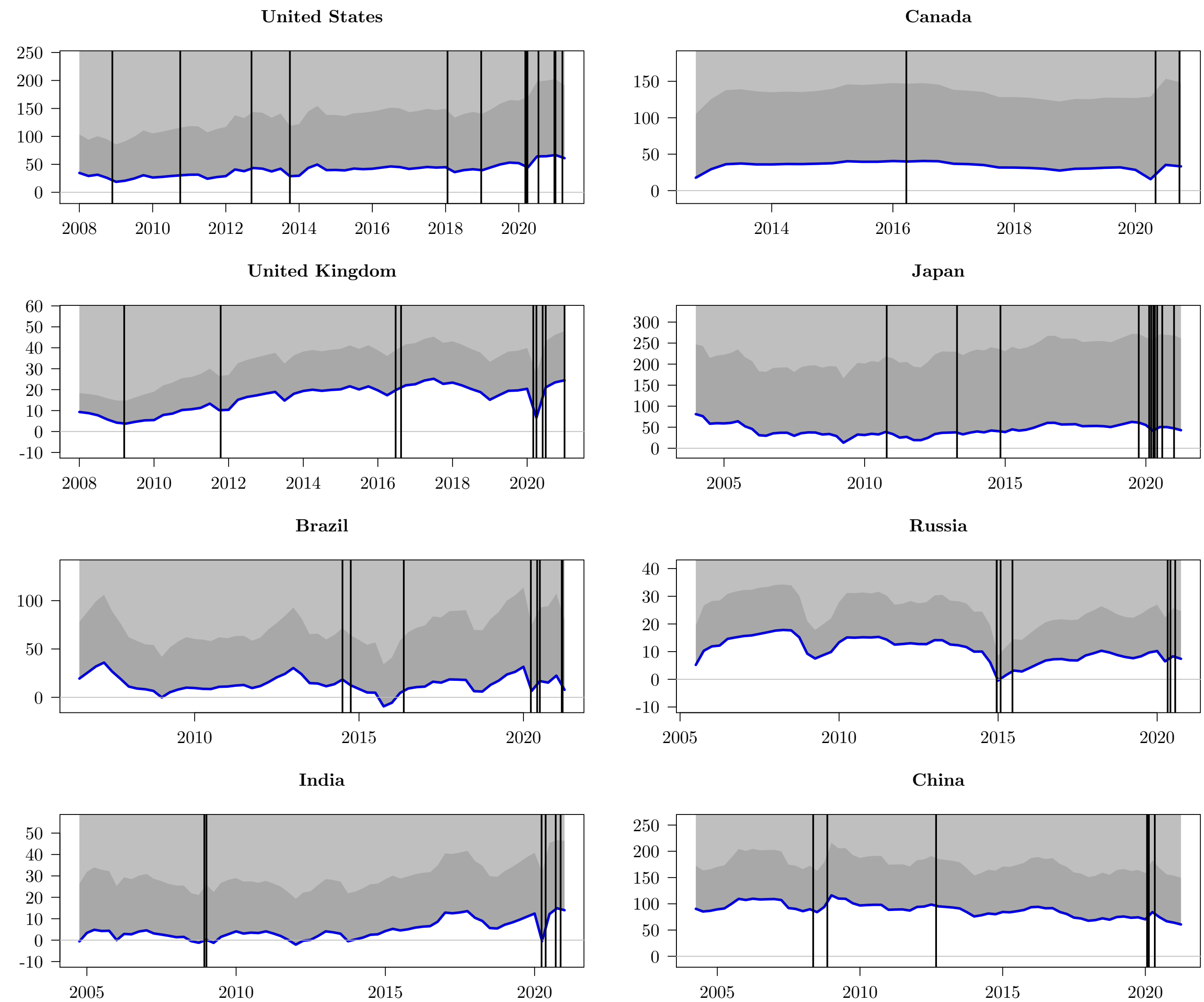
[26]The model-implied unconditional correlation between GDP growth (Δy_t) and the fiscal limit (ℓ_t) for Brasil, Canada, China, India, Japan, Russia, United Kingdom and the United States is equal to -0.308, -0.172, -0.506, -0.106, -0.083, -0.097 and -0.465, respectively.

Figure 1.3: Fiscal limits estimates



Note: These plots display estimated fiscal limits (exp(ℓ_t)), expressed in % of GDP. Everything else equal, if debt-to-GDP stayed in the darker-shaded (respectively lighter-shaded) area for four quarters in a row, then the annual default probability of the considered country would be larger than 10% (respectively in $]0\%, 10\%]$). The delimitation of these areas is based on the formula given at the end of the caption of Figure 1.1. On each plot, the vertical bars indicate noteworthy pivotal economic events (see caption of Figure 1.4 for details regarding these dates).

Figure 1.4: Fiscal space estimates



Note: These plots show the estimates of the fiscal space expressed in % of GDP, that is $100 \times (\exp(\ell_t) - \exp(d_t))$. See caption of Figure 1.3 for the interpretation of the two shaded areas. On each plot, the vertical bars indicate important dates, including key monetary-policy decisions and/or noteworthy pivotal economic events: **Canada**–22/03/2016: Stimulus of more than $47bn over 5 years; 05/2020 and 01/2021: COVID-19 waves' peaks; 22/09/2020 and 12/2020: Major fiscal stimulus package announcements. **Japan**–10/2010: Announcement of QE; 04/2013 and 31/10/2014: QE expansions; 01/10/2019: Consumption tax hike; 13/02/2020, 10/03/2020, 07/04/2020 and 27/05/2020: COVID-19 emergency response package announcements (stimulus plans); 08/2020, 01/2021: COVID-19 wave's peak; 27/04/2020: Ramp-up of QE. **UK**–03/2009: Announcement of Asset Purchase Program; 10/2011: New round of QE; 23/06/2016: Brexit vote; 08/2016: BoE announces bond purchases to tackle uncertainty over Brexit; 03/2020 and 07/2020: Major fiscal stimulus package announcements; 04/2020 and 01/2021: COVID-19 waves' peaks. **US**–25/11/2008: Announcement of QE1 (purchase of GSE bonds and MBS); 3/11/2010: Announcement of QE2 (purchase of additional $600bn of Treasury securities); 13/09/2012: Announcement of QE3; 01/10/2013, 20/01/2018 and 22/12/2018: Federal government shutdowns (debt ceiling); 06/03/2020, 18/03/2020, 27/03/2020, 21/12/2020 and 11/03/2021: Stimulus and Relief Package (COVID-19); 07/2020 and 01/2021: COVID-19 waves' peaks. **Brazil**–01/07/2014: Onset of the Brazil recession; 01/10/2014: Petrobras loses 60% of market value; 12/05/2016: Resignation of Dilma Rousseff (President) amid corruption scandal; 07/2020 and 03/2021: COVID-19 waves' peaks; 23/03/2020, 06/2020 and 11/03/2021: Stimulus package announcements (COVID-19). **China**–12/05/2008: Sichuan earthquake; 09/11/2008: Announcement of China Stimulus Plan; 01/09/2012: China infrastructure plan ($156.6bn); 23/01/2020: Chinese authorities isolate Wuhan Province (COVID-19); 30/01/2020: The W.H.O. declared a global health emergency; 02/2020: Peak of COVID-19 pandemic; 05/2020: Stimulus package announcement (COVID-19). **India**–08/12/2008 and 03/01/2009: Economic stimulus package announcements; 09/2020: COVID-19 wave's peak; 26/03/2020, 15/05/2020 and 14/11/2020: Economic stimulus package announcements (COVID-19). **Russia**–15-16/12/2014: During the Russian financial crisis, Russian central bank spends $2bn in foreign reserves and increases interest rate from 10.5% to 17% to stop declining ruble; 30/01/2015: Interest rate cut (−2%); 15/06/2015: Interest rate cut (−1%); 05/2020 and 01/2021: COVID-19 waves' peaks; 02/06/2020: Economic stimulus announcement (COVID-19); 27/07/2020: Interest rate cut (COVID-19);

According to our estimates reported in Figures 1.3 and 1.4, the global financial crisis of 2008 translated into falls of fiscal space in advanced economies. On average across Canada, Japan, the UK, and the US, fiscal space decreased by 10 percent of GDP from 2008Q1 to 2009Q1. This drop in fiscal space may be seen as a consequence of transfers from private to public debts through explicit channels (bank bailouts) or implicit ones (debt and deposit guarantees), along the logic of the so-called sovereign-bank nexus (see, e.g., Acharya, Drechsler, and Schnabl, 2014; Jordà, Schularick, and Taylor, 2016).

Fiscal spaces for the US and the UK were tightest at the beginning of 2009, then reaching respective minimums of 18% and 4% of GDP. After 2009, the US fiscal limit (respectively the fiscal space) has steadily increased, from 115% in 2009Q4 to 195% of GDP at the beginning of 2021 (resp. from 30% to 65% of GDP). While the situation was similar in the UK before 2016, the UK fiscal space decreased by 5 percentage points (p.p.) from the end of 2015 until the midst of the following year, against the Brexit vote backdrop. Interestingly, in the US, fiscal space decreased around the government shutdown crises in 2013 and 2018 (debt ceiling crises). Japan features the largest average fiscal limit, being equal to 265% of GDP. After 2009, the Japanese fiscal limit shows a similar trend to that of the debt-to-GDP ratio, implying a fairly stable fiscal space at the end of 2015 close to 40% of GDP. As regards Canada, the fiscal limit oscillates around 140% of GDP, featuring minor time variation. (Note however that the Canadian sample is shorter, starting only in 2012 due to CDS data availability.)

In Brazil, following the recession that started in 2014, the fiscal space decreased from 2014Q2 until 2015Q4.[27] Notably, during the same period, fiscal space turned slightly negative. Subsequent to the impeachment and resignation of President Dilma Rousseff that began at the end 2015, fiscal space in Brazil showed a positive trend.[28] Amid COVID-19-related uncertainty at the beginning of 2020, the Brazilian fiscal space decreased by 15 p.p. In Russia, fiscal space diminished at the onset of the Russian financial crisis in 2014—this crisis, which was partly due to a fall in oil prices and to geopolitical factors, was marked by a collapse of the ruble. India features a very tight fiscal space throughout the whole estimation sample. In particular, amid the global financial crisis and the COVID-19 pandemic, fiscal space turned slightly negative: the fiscal limit for India was equal to 45% in 2008Q4 and to 65% in 2020Q1, implying in both cases a negative fiscal space of about 1% of GDP.

In China, fiscal space exhibits a decreasing trend over the estimation sample, from about 100% of GDP between 2005 and 2010 to around 60% in the beginning of 2021. Notably, fol-

[27] At the end of 2014, in particular, the largest state-owned Brazilian multinational corporation in the petroleum industry (Petrobras) lost 60% of its market value.

[28] President Dilma Rousseff was accused of criminal administrative misconduct and disregard for the federal budget law.

lowing the infrastructure spending plan (totalling $156.6bn) at the end of 2012, fiscal space decreased by 10 p.p. throughout 2013.

Table 1.2: Fiscal space and economic policy uncertainty - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.860^{***}	0.849^{***}	0.985^{***}
	(0.029)	(0.028)	(0.009)
EPU_t	-0.008^{***}	-0.005^{**}	-0.002
	(0.002)	(0.002)	(0.002)
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.802^{***}	0.848^{***}	0.928^{***}
	(0.047)	(0.038)	(0.027)
EPU_t	-0.005^{*}	0.002	-0.007^{**}
	(0.003)	(0.003)	(0.003)
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.824^{***}	0.812^{***}	0.988^{***}
	(0.044)	(0.045)	(0.008)
EPU_t	-0.012^{***}	-0.011^{***}	-0.003
	(0.004)	(0.003)	(0.003)

Note: This table reports the results of panel regressions of fiscal space (FS) estimates on the Economic Policy Uncertainty (EPU) indices. The estimation sample goes from 2004Q1 to 2021Q1. See text for more details. FE. stands for Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

The previous discussion suggests that economic policy uncertainty contributes to fluctuations in fiscal space. Tables 1.2 shows the results of panel regressions where the fiscal space is accounted for by its first lag and by the economic policy uncertainty (EPU) index computed by Baker, Bloom, and Davis (2016).[29] Including both country and time fixed effects, we obtain significant estimates for the EPU index. As expected, higher uncertainty causes fiscal space to

[29] This index is constructed from three types of underlying components: (i) newspaper coverage of policy-related economic uncertainty, (ii) number of federal tax code provisions set to expire in future years, (iii) disagreement among economic forecasters as a proxy for uncertainty. The Japanese and Chinese EPU indexes are respectively obtained from Arbatli, Davis, Ito, and Miake (2017) and Davis, Liu, and Sheng (2019).

Table 1.3: Fiscal space and economic policy uncertainty - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.486***	0.508***	0.883***
	(0.069)	(0.088)	(0.034)
EPU_t	−0.203***	−0.059	−0.008
	(0.077)	(0.050)	(0.072)
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.710***	0.829***	0.925***
	(0.049)	(0.043)	(0.028)
EPU_t	0.054	0.030	−0.043
	(0.039)	(0.041)	(0.029)
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	FS_t	FS_t	FS_t
FS_{t-1}	0.412***	0.415***	0.877***
	(0.077)	(0.109)	(0.035)
EPU_t	−0.348***	−0.144**	0.033
	(0.112)	(0.071)	(0.114)

Note: This table reports the results of panel regressions of the log of fiscal space (FS) estimates on the log of the Economic Policy Uncertainty (EPU) indices. The estimation sample goes from 2004Q1 to 2021Q1. See text for. more details. FE stands for Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

decline (see Panel A of Table 1.2). Comparable results hold true for the specification including only either country fixed effects or time fixed effects. Similar conclusions arise if we run our panel regressions on countries' subsets (Panel B and C of Table 1.2 include only advanced and emerging economies, respectively).[30]

In Table 1.4, we compare the sample averages of our fiscal limits with point estimates of the debt limit derived from both Ghosh, Kim, Mendoza, Ostry, and Qureshi (2013) and Collard, Habib, and Rochet (2015). As both studies exclude emerging economies, we can only compare advanced economies' fiscal limits. With respect to the point estimates provided in Ghosh et al. (2013), our mean fiscal limits tend to be be more conservative. (Note that, in Ghosh et al. (2013), debt limits for Japan failed to be computed.) The "maximum sustainable debt" figures computed in Collard et al. (2015) are broadly comparable to our mean estimates across all countries. Notably, debt limits in Collard et al. (2015) arising from a constant 5% maximum primary surplus-to-GDP (column "5% **MPS**" in Table 1.4) and from the computation involving a maximum recovery rate (column "**MRR**" in Table 1.4) are the closest to our average fiscal limits. Furthermore, there is a substantial overlap between our US fiscal estimates, that varies between 93% and 194% of GDP over the sample, and the values computed by Mehrotra and Sergeyev (2020), that lay between 150 and 220 percentage points (for different model calibrations).

[30] In Table 1.3, we report results from the same panel regressions specifying the variables in logarithms and comparable conclusions are reached, with the exception of advanced economies (Panel B) for which the results are less robust.

Table 1.4: Fiscal limit estimates comparison

Ctry	This paper				Ghosh et al. (2013)		Collard et al. (2015)					
	$\overline{\mathbf{FL}}$	**SD**(**FL**)	min(**FL**)	max(**FL**)	**Hist.**	**Proj.**	5% **MPS**	**MRR**	**TVR**	**CATA**	4% **MPS**	**h. MPS**
US	138.4	23.9	93.0	194.3	183.3	160.5	120.9	123.2	110.3	79.5	96.7	123.0
CA	144.9	9.1	125.4	169.0	152.3	181.1	120.8	123.0	109.9	79.8	96.6	242.7
UK	93.7	18.9	48.5	126.8	182.0	166.5	126.2	128.6	113.7	82.0	101.0	159.4
JP	263.4	29.7	217.3	316.1	–	–	–	–	–	–	–	–
BR	84.2	15.3	60.1	121.4	–	–	–	–	–	–	–	–
RU	21.3	2.7	13.0	24.8	–	–	–	–	–	–	–	–
IN	51.8	5.2	43.5	67.7	–	–	–	–	–	–	–	–
CN	126.6	7.4	112.1	143.1	–	–	–	–	–	–	–	–

Note: All estimates are reported in percent of GDP. **This paper –** $\overline{\mathbf{FL}}$: Sample mean of the fiscal limit estimates. **SD**(**FL**): Standard deviation of the fiscal limit estimates. min(**FL**): absolute minimum for fiscal limit estimates. max(**FL**): absolute maximum for fiscal limit estimates. **Estimates of Ghosh et al. (2013) –** Debt limits (fiscal limits in our terminology) are statically estimated through the interest payment schedule for the period 1985-2007. **Hist.**: Estimates are based on the average interest rate / growth differential of 1998-2007, using the implied interest rate on public debt; **Proj.**: The interest rate / growth differential is based on the long term government bond yield (average for 2010-2014, IMF projections as of 2010). **Estimates of Collard et al. (2015) –** The computation of maximum sustainable debts (fiscal limits in our terminology) exploits the idea of a maximum primary surplus (MPS). In the model, there is a maximum amount that can be issued on each date (that itself depends on the MPS). 5% **MPS**: Case where the MPS is set to to 5%; **MRR**: The computation involves a maximum recovery rate; **TVR**: The model features a time-varying interest rate; **CATA**: The model features catastrophes; 4% **MPS**: The MPS is set to 5%; **h. MPS**: The MPS is set to the historical peak of primary surplus-to-GDP.

1.5.2 Sensitivity of CDS spreads to fiscal conditions

A large body of empirical studies estimates the relationship between fiscal variables and government funding costs. Either focusing on a given country or based on panel data, regressions usually detect a positive relationship between sovereign interest rates on the one hand and deficits or debt levels on the other hand (see, e.g., the literature reviews in Gale and Orszag, 2004; Haugh, Ollivaud, and Turner, 2009; Bernoth, von Hagen, and Schuknecht, 2012). Moreover, a wide range of evidence points to a time-varying dependence of interest rates' sensitivity to fiscal conditions.[31] In particular, several studies highlight the role of public indebtedness itself as a relevant determinant of this sensitivity.[32]

Our framework is qualitatively consistent with this regression-based evidence, but provides richer predictions regarding the changes in credit spreads that can be expected from fiscal deterioration. In the model, CDS spreads are nonlinear functions of the state of the economy. The influence of increases in the deficit, say, on spreads is therefore inherently time-dependent (it depends on x_t) and nonlinear (the marginal effect of the shock depends on its magnitude). It is worth noting that this complexity makes it challenging to provide a comprehensive picture of spreads' sensitivity in our model. Typically, because it depends on all components of x_t, the (nonlinear) relationship between spreads and the economic variables entering the model cannot be perfectly summarized by a small set of numbers.

We nevertheless attempt to illustrate the sensitivity of credit spreads to deficits by means of the following exercise. We produce a one-year-long benchmark macroeconomic scenario with initial conditions coinciding with our last estimation period. To this purpose, we simply employ the vectorial auto-regressive process of x_t (eq. 1.19), setting the shocks to zero. We then construct alternative scenarios featuring larger deficits. Specifically, we evenly spread increases in the deficit over the four simulated quarters to reach a desired fiscal shock; we consider three sizes of deficit shocks, consisting of 1 p.p. marginal increases in deficit: from 0 to 1%, from 4 to 5%, and from 9 to 10% of GDP. We then use the Kalman filter to derive expected values of the different components of the state vector, conditional on the adverse deficit trajectory, and

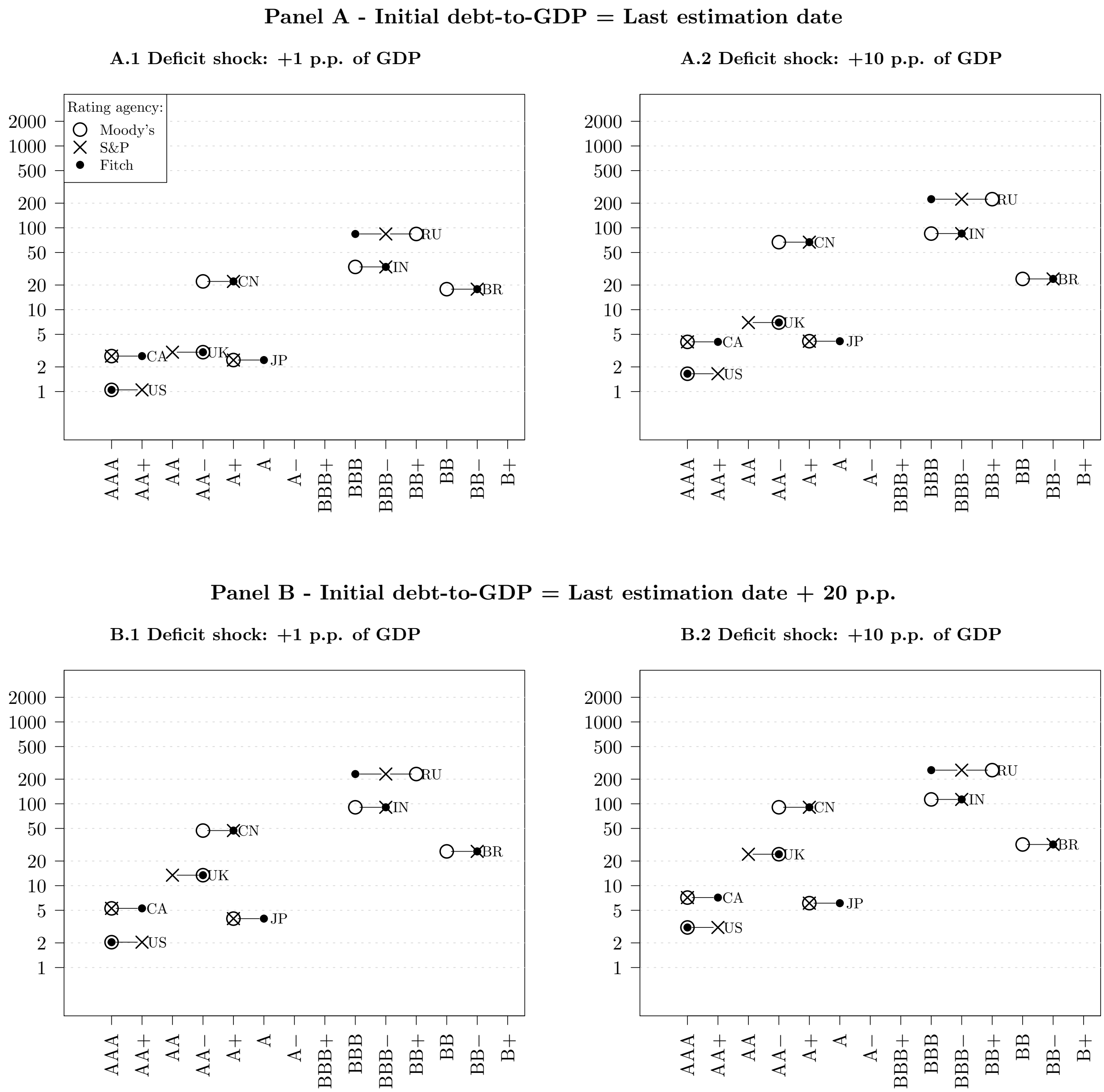
[31] Using rolling regressions, Laubach (2010, Subsection III.A) highlights time-variation in the the spread sensitivity to the debt-to-GDP ratio; stressed periods are associated with larger sensitivities. This result is notably in line with a recent contribution by Alessi, Balduzzi, and Savona (2020) who resort to machine-learning-based regressions to identify different sensitivity regimes.

[32] Caggiano and Greco (2012) and López-Espinosa, Moreno, Rubia, and Valderrama (2017) provide evidence of threshold effects, the spread sensitivity being larger when debt-to-GDP is above 100%. Numerous studies find that adding the square of the debt-to-GDP ratio among the explanatory variables of a regression improves the fit of credit spreads (among many others Baldacci and Kumar, 2010; Bernoth, von Hagen, and Schuknecht, 2012; De Grauwe and Ji, 2013; Delatte, Fouquau, and Portes, 2014).

keeping GDP growth as under the benchmark scenario.[33] We finally examine the deviations between the CDS spreads computed under the three alternative scenarios, on the one hand, and those of the benchmark one, on the other hand. The results of this first exercise are documented by Table 1.5 (Panel A) and Figure 1.5, which focus on the 10-year maturity. Results for the 2-year and 5-year maturities are reported in Tables 1.N1 and 1.N2 of Appendix 1.N.

[33]The underlying state-space model features two measurement equations, that state that the deficit and GDP growth have to coincide with those defined in the considered scenario. The state vector is x_t, and the the transition equations are therefore given by x_t's dynamics (eq. 1.19). In the alternative scenarios, we keep GDP growth as under the benchmark one in order to isolate the effect of a deficit shock.

Figure 1.5: Sensitivity of 10-year CDS spreads to increases in deficits



Note: These plots display the model-predicted sensitivities of 10-year CDS spreads to deficits. Effects are measured in basis points. We consider two magnitudes of fiscal shocks (increases in primary deficits by 1% and 10% of GDP). The shocks are spread on four quarters and we compare the levels of the CDS spreads, after these four quarters, with those obtained in a benchmark scenario (with no fiscal shock). The initial conditions of the scenarios underlying Panel A are those observed on the last date of the estimation sample; the same initial conditions are used to generate the results of Panel B, except that, in the latter case, the initial debt-to-GDP ratio is increased by 20 percentage points. See text for more details regarding the construction of scenarios. Reported figures correspond to the marginal influence of an additional unit increase in the deficit (i.e. from 0% to 1% of GDP for left-hand-side plots and from 9% to 10% for right-hand-side plots). It appears that CDS sensitivities are stronger for larger deficits (the effects shown on the right-hand-side plots are larger), which illustrates the nonlinear relationship between fiscal conditions and credit spreads. The credit ratings are those observed in August 2021.

Table 1.5: 10-year CDS sensitivity to deficits

Panel A - Initial state for simulations = Last estimation period						
Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	1.0	[1.0]	5.8	[1.3]	13.3	[1.6]
Canada	2.7	[2.7]	14.9	[3.3]	33.5	[4.0]
United Kingdom	3.0	[3.0]	18.5	[4.4]	47.8	[7.0]
Japan	2.4	[2.4]	13.8	[3.1]	32.3	[4.1]
Brazil	17.8	[17.8]	95.5	[20.4]	207.6	[23.8]
Russia	84.1	[84.1]	598.6	[156.9]	1.6×10^3	[0.2×10^3]
India	33.3	[33.3]	224.3	[56.9]	596.1	[85.0]
China	22.1	[22.1]	155.2	[40.6]	437.1	[66.8]

Panel B - Lower initial debt-to-GDP ratio (-10 p.p. of GDP w.r.t. Panel A)						
Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	0.7	[0.7]	4.0	[0.9]	9.1	[1.1]
Canada	1.7	[1.7]	9.7	[2.1]	22.1	[2.7]
United Kingdom	1.2	[1.2]	7.2	[1.7]	19.0	[2.8]
Japan	1.8	[1.8]	10.5	[2.4]	24.9	[3.2]
Brazil	13.4	[13.4]	72.9	[15.8]	161.0	[18.9]
Russia	11.2	[11.2]	100.3	[31.0]	401.7	[84.8]
India	9.7	[9.7]	76.6	[21.9]	258.3	[47.3]
China	12.1	[12.1]	92.9	[25.9]	294.0	[50.4]

Panel C - Larger initial debt-to-GDP ratio ($+20$ p.p. of GDP w.r.t. Panel A)						
Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	2.0	[2.0]	11.2	[2.5]	25.4	[3.1]
Canada	5.3	[5.3]	28.4	[6.1]	62.0	[7.1]
United Kingdom	13.4	[13.4]	77.8	[17.8]	185.7	[24.2]
Japan	4.0	[4.0]	22.0	[4.9]	50.0	[6.1]
Brazil	26.2	[26.2]	137.6	[28.8]	290.8	[31.8]
Russia	230.9	[230.9]	1.2×10^3	[0.2×10^3]	2.5×10^3	[0.3×10^3]
India	90.6	[90.6]	488.9	[104.1]	1.0×10^3	[0.1×10^3]
China	47.1	[47.1]	290.2	[68.9]	703.9	[90.6]

Note: This table documents the sensitivity of the 10-year CDS spreads to fiscal conditions. We consider three sizes of fiscal shocks (increases in primary deficits by 1%, 5% and 10% of GDP). The shocks are spread on four quarters and we compare the levels of the CDS spreads, after these four quarters, with those obtained in a benchmark scenario (with no fiscal shock). See text for more details regarding the construction of scenarios. The reported figures are in basis points. The number in square brackets correspond to the marginal influence of an additional unit increase in the deficit.

Let us first consider higher-rated countries: the US, the UK, Canada and Japan, that are countries with credit ratings going from A to AAA (see Figure 1.5). For this group of countries, an increase in deficit by 1 percentage point (p.p.) of GDP causes an average increase of 2 basis points (b.p.) in the 10-year CDS spreads (see left columns of Panel A). The middle and right columns of Panel A show that marginal effects increase with the magnitude of the shock: the right columns show, in particular, that the 10-year CDS spreads increase by 5 b.p., on average across advanced economies, when the deficit goes from +9 to +10 p.p. of GDP (see numbers in square brackets). Hence, the increase in CDS spreads is twice higher when the deficit goes from +9 to +10 p.p. than when it goes for 0 to +1 p.p. of GDP; this illustrates the non-linear sensitivity of the CDS in this model.

The nonlinearity is also pronounced for Brazil, India and Russia, countries with credit ratings ranging from BB^- to BBB. While a one-unit deficit shock results in respective augmentations of the spreads of roughly 20, 35 and 85 b.p., a marginal deficit increase from +9 p.p. to +10 p.p. of GDP expands the spreads by about 25, 85 and 200 basis points. Even though China's credit rating is comparable to that of an advanced economy such as Japan, the impact of deficit shocks in China is similar to that observed in the other emerging economies. Indeed, the effect of a 1 p.p. increase in deficit yields a 25 b.p. increase in the 10-year spread, and this spread increases by 70 b.p. when the deficit goes from +9 to +10 p.p. of GDP. We obtain similar results analyzing the 2 and 5-year CDS spreads.

We further examine the influence of the debt level on spreads' sensitivities. In this purpose, we replicate the previous simulation exercise, decreasing by 10 percentage points (Panel B of Table 1.5) and increasing by 20 percentage points (Panel C) the initial debt-to-GDP ratios. Simulations starting with lower (respectively higher) debt-to-GDP ratios yield substantially milder (resp. stronger) responses in spreads compared to the baseline case (Panel A). For instance, in Russia, while the marginal effect of a one-unit deficit shock leads to a 85 b.p. increase in the 10-year CDS spread (Panel A of Table 1.5), the model suggests the effect would be 12 b.p. (resp. 230 b.p.) if the debt-to-GDP was initially 10 p.p. lower (resp. 20 p.p. larger).

As stressed above, the above simulation exercises cannot completely summarize the sensitivity of CDSs to fiscal conditions—let alone the sensitivity to general economic conditions. They nevertheless give a sense of the nonlinearity of CDS spreads. Such effects may explain why regression-based analysis fail to obtain specifications providing a good fit across time, which sometimes leads the authors to conclude that sovereign spreads are disconnected from fundamentals.

1.5.3 Risk premiums and sovereign probabilities of default

Because our model is founded on a representative risk-averse agent, it can be used to examine how risk premiums affect the pricing of sovereign credit risk. This subsection shows that these risk premiums translate into substantial discrepancies between non-risk-adjusted CDS-based probabilities of default and risk-adjusted, or physical, probabilities of default.

Generally speaking, risk premiums are defined as those components of asset returns that would not exist if investors were not risk-averse. Consider a CDS contract. If agents were risk-neutral, the CDS spread would be approximately equal to the expected credit loss, i.e. the product of the loss-given default multiplied by the probability of default. However, if agents are risk-averse and if sovereign defaults tend to take place in bad states of nature—i.e. states of high marginal utility—then protection sellers are willing to enter the credit swap only if the CDS spread is larger than the expected credit loss.

To explore the model implications regarding this matter, it will prove convenient to introduce the risk-neutral measure $\mathbb{Q}$. This probability measure can be understood here as a convenient mathematical tool aimed at facilitating the presentation and interpretation of those results pertaining to risk premiums. It is defined through the following change of density, or pricing kernel:[34]

$$\left.\frac{d\mathbb{Q}^h}{d\mathbb{P}}\right|_{t,t+h} = \frac{\mathcal{M}^n_{t,t+h}}{\mathbb{E}_t(\mathcal{M}^n_{t,t+h})}. \tag{1.25}$$

This definition notably implies that the (forward) price, decided on date-t but settled on date $t+h$, of a future nominal payoff Π_{t+h}, is given by $\mathbb{E}^{\mathbb{Q}^h}_t(\Pi_{t+h})$.[35] It is easily seen that if agents were risk-neutral ($\gamma = 0$), then the s.d.f. would be deterministic (see eq. 1.4), and this forward price would be equal to $\mathbb{E}_t(\Pi_{t+h})$. (The so-called "expectation hypothesis" would then hold true.) The pricing kernel (1.25) reflects how physical probabilities are distorted when it comes to price uncertain future payoffs; it implies that those assets that provide relatively higher payoffs when the s.d.f. is high—that is when consumption is low—have larger prices than under the expectation hypothesis.

To illustrate this notion in the present credit risk context, consider a forward contract providing $\mathcal{D}_{t+h}$ on date $t+h$, with payment deferred to date $t+h$. (We assume that the government has not defaulted before the current date, t, i.e., $\mathcal{D}_t = 0$.) If investors were risk-neutral, they

[34] Eq. (1.25) defines the h-forward risk-neutral measure $\mathbb{Q}^h$. For $h > 1$, this measure is equivalent to—but does not coincide with—the risk-neutral measure $\mathbb{Q}^1$ (usually called $\mathbb{Q}$). Under $\mathbb{Q}^h$, the numeraire is a zero-coupon bond of maturity h (while it is the money market under $\mathbb{Q}$). The following Radon-Nikodym derivative characterizes the relationship between these two two risk-neutral measures: $\frac{d\mathbb{Q}^h}{d\mathbb{Q}}|_{t,t+h} = \exp(-i_t - \dots - i_{t+h-1})/\mathbb{E}_t(\mathcal{M}^n_{t,t+h})$, where i_t is the short-term nominal interest rate between dates t and $t+1$ (see, e.g., Jamshidian, 1989).

[35] This price indeed is $\frac{1}{\mathbb{E}_t(\mathcal{M}^n_{t,t+h})}\mathbb{E}_t\left(\mathcal{M}^n_{t,t+h}\Pi_{t+h}\right) = \mathbb{E}_t\left(\frac{\mathcal{M}^n_{t,t+h}}{\mathbb{E}_t(\mathcal{M}^n_{t,t+h})}\Pi_{t+h}\right)$, which is equal to $\mathbb{E}^{\mathbb{Q}^h}_t(\Pi_{t+h})$ given the change of density from $\mathbb{P}$ to $\mathbb{Q}^h$ (eq. 1.25). More generally, a risk-neutral probability measure is defined so that the price of any asset is equal to the discounted $\mathbb{Q}$-expectation of this asset's payoffs.

would be willing to enter this contract if its (forward) price was equal to $\mathbb{E}_t(\mathcal{D}_{t+h})$, that is to the physical probability that the government defaults before date $t+h$. But agents are risk-averse, and the associated forward price is $\mathbb{E}_t^{\mathbb{Q}^h}(\mathcal{D}_{t+h})$, that is the $\mathbb{Q}^h$-probability of default. By virtue of the pricing kernel definition (eq. 1.25), it can be seen that the difference between the $\mathbb{P}$ and $\mathbb{Q}^h$ probabilities of default is:

$$\underbrace{\mathbb{E}_t^{\mathbb{Q}^h}(\mathcal{D}_{t+h}) - \mathbb{E}_t(\mathcal{D}_{t+h})}_{\text{credit-risk premium}} = \mathbb{C}ov_t\left(\frac{\mathcal{M}^n_{t,t+h}}{\mathbb{E}_t(\mathcal{M}^n_{t,t+h})}, \mathcal{D}_{t+h}\right). \tag{1.26}$$

Hence the credit-risk premium, that is—by definition—the difference between the forward price and the price that would prevail under the expectation hypothesis, is equal to the covariance between the pricing kernel and the default event indicator. Because the pricing kernel negatively depends on consumption, and since defaults tend to happen in recessions (when consumption is low), one expects the covariance term to be positive.[36] Hence, the existence of risk premiums associated with sovereign credit risk implies that physical ($\mathbb{P}$) probabilities of default differ from their risk-neutral counterparts. Yet, the latter, derived from basic credit-risk models like in Litterman and Iben (1991), are extensively used by market practitioners, who refer to them as market-implied default probabilities (see, e.g., Hull, Predescu, and White, 2005).

Our approach makes it possible to quantify and compare physical and risk-neutral default probabilities. The latter are indeed given by:

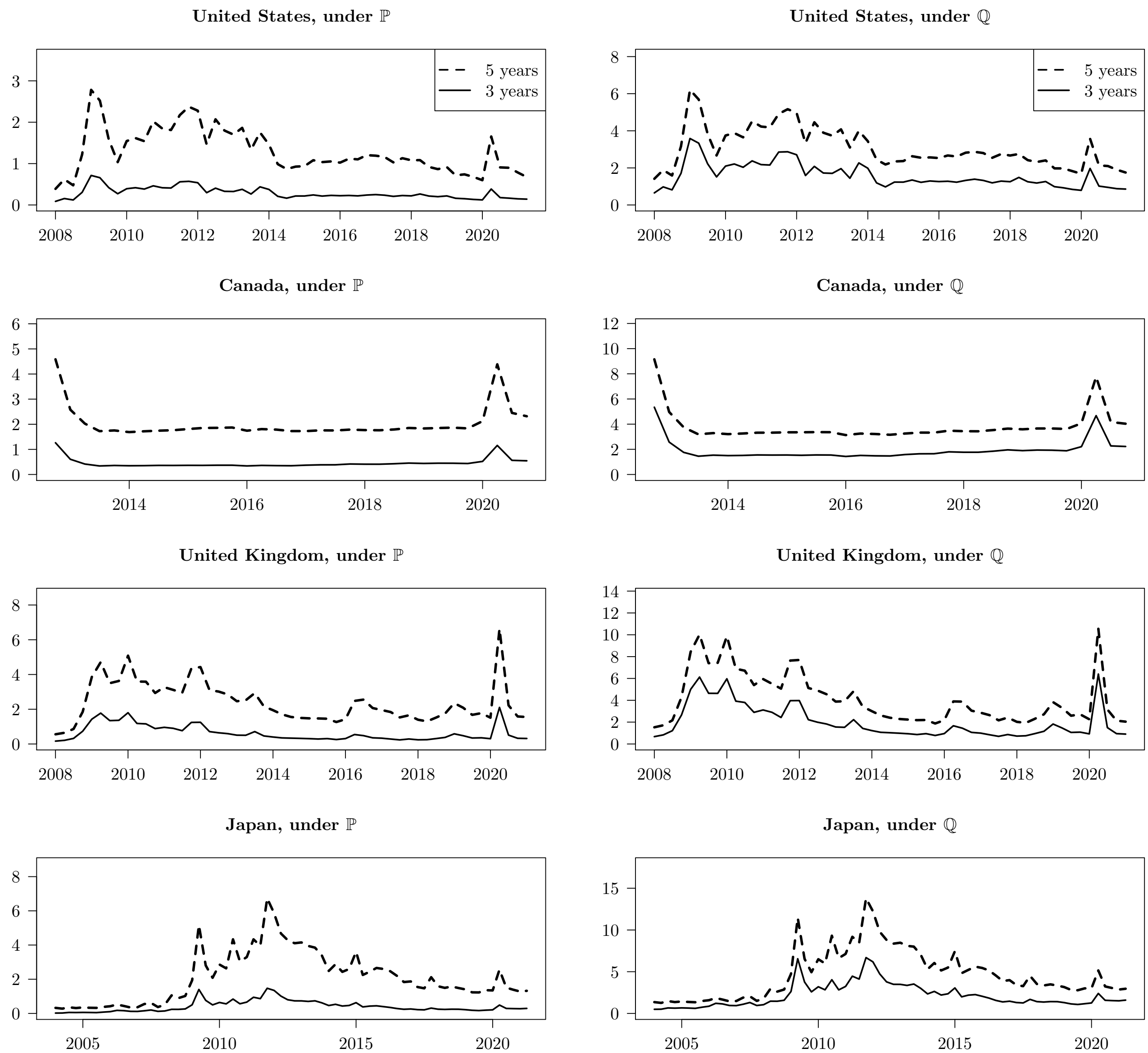
$$\mathbb{E}_t^{\mathbb{Q}^h}(\mathcal{D}_{t+h}) = \frac{\mathbb{E}_t\left[\mathcal{M}_{t,t+h}\mathcal{D}_{t+h}\right]}{\mathbb{E}_t\left(\mathcal{M}_{t,t+h}\right)}, \tag{1.27}$$

and Appendix 1.A explains how to get approximations to the two conditional expectations appearing on the right-hand side of (1.27). The same type of formula can be used to compute physical probabilities of default, i.e., $\mathbb{E}_t(\mathcal{D}_{t+h})$ for different horizons h.[37] Physical and "risk-adjusted," probabilities of default respectively appear on the left-, and right-hand-side of Figures 1.6 and 1.7.

[36]Two channels account for the negative correlation between consumption and the default indicator in our model. To fix ideas, consider a recession. First, debt-to-GDP soars as GDP plunges. Second, upon default, consumption is expected to experiment a fall of magnitude b_c (see eq. 1.1).

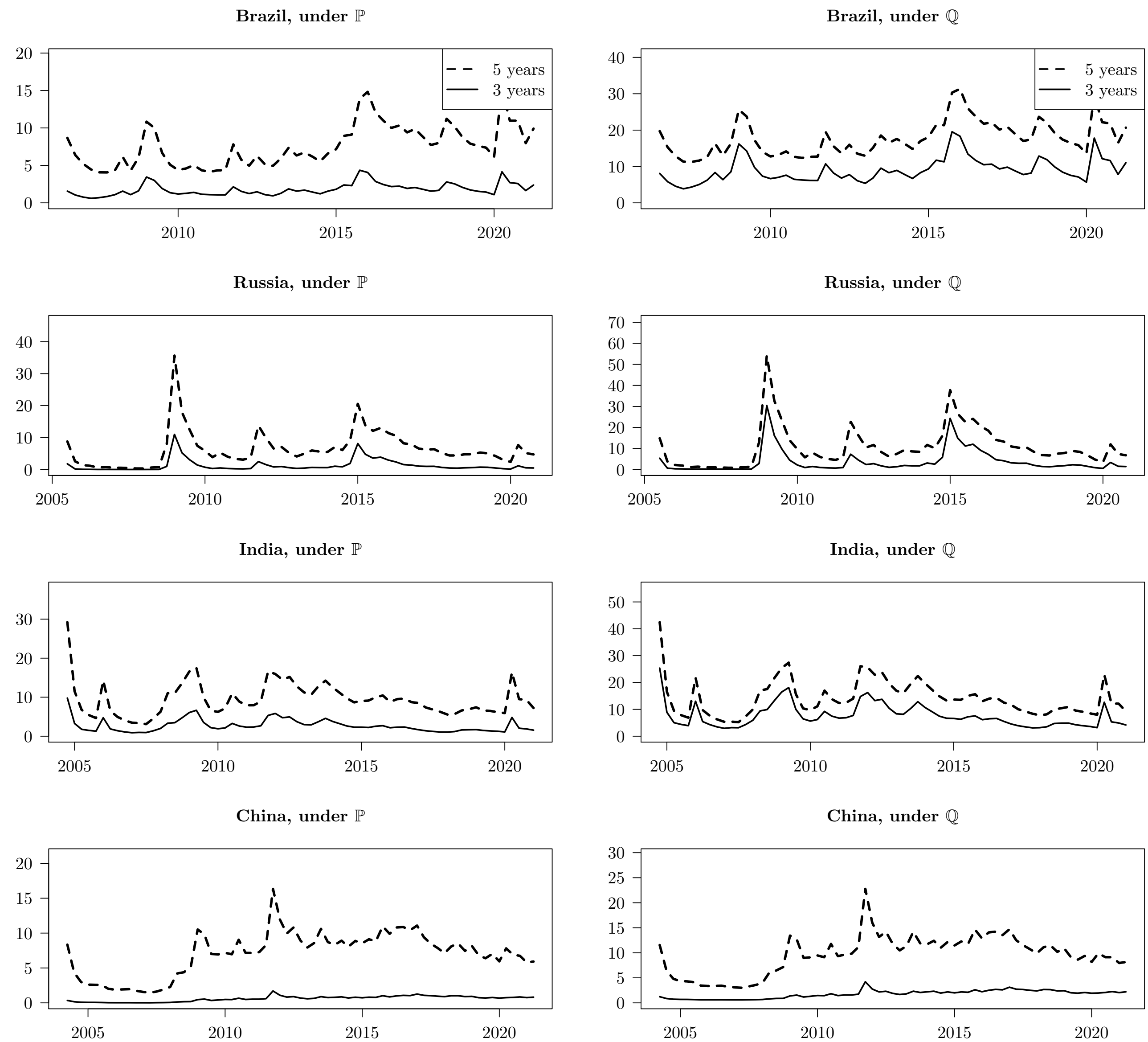
[37]Practically, we replace γ with 0 in eq. (1.4), This is simply done by making the s.d.f. $\mathcal{M}_{t,t+1h}$ deterministic, replacing γ with 0 in eq. (1.4).

Figure 1.6: Risk-adjusted probabilities of default



Note: This figures displays physical (left-hand-side plot) and risk-neutral (right-hand-side plot) sovereign probabilities of default at different horizons. Probabilities are expressed in percentage points. Formally, for each date, we compute $\mathbb{E}_t(\mathbb{1}_{\{\mathcal{D}_{t+h}=1\}})$ (left-hand-side plots) and $\mathbb{E}_t^{\mathbb{Q}^h}(\mathbb{1}_{\{\mathcal{D}_{t+h}=1\}})$ (right-hand-side plots), for different values of h (with $\mathcal{D}_t = 0$). See Subsection 1.5.3 for details regarding the computation of these probabilities.

Figure 1.7: Risk-adjusted probabilities of default (cont'd)



Note: This figures displays physical (left-hand-side plot) and risk-neutral (right-hand-side plot) sovereign probabilities of default at different horizons. Probabilities are expressed in percentage points. Formally, for each date, we compute $\mathbb{E}_t(\mathbb{1}_{\{\mathcal{D}_{t+h}=1\}})$ (left-hand-side plots) and $\mathbb{E}_t^{\mathbb{Q}^h}(\mathbb{1}_{\{\mathcal{D}_{t+h}=1\}})$ (right-hand-side plots), for different values of h (with $\mathcal{D}_t = 0$). See Subsection 1.5.3 for details regarding the computation of these probabilities.

According to eq. 1.26, the observable differences between the two types of probabilities reflect the existence of credit risk premiums, which are sizeable. Risk-neutral probabilities of default appear to be about twice larger than their physical counterpart, meaning that credit risk premiums represent a large share of spreads consistently with sovereign credit risk studies based on reduced-form intensity approaches (e.g., Pan and Singleton, 2008; Longstaff et al., 2011). We investigate how credit risk premiums relate to uncertainty and volatility indices, and also to our fiscal space estimates. Note that, in our approach, the covariates of the physical probabilities of default are the same as of the credit risk premiums given that the latent factors driving both computed objects are the same.[38] Indeed, the correlation between model-implied physical probabilities of default and credit risk premiums is close to one for all countries. This being said, we explore how credit risk premiums comove with the country-specific EPU indices, the Chicago Board Options Exchange (Cboe) VIX index as a measure for global volatility, the Emerging Markets Volatility Index (VXEEM) and the $S\&P$ 500 Index 9-Day Volatility Index as a measure for short-term volatility.[39] In Tables 1.L1 and 1.L2 (Appendix 1.L), we report the output of simple fixed effects panel regressions of credit risk premiums (3- and 5-year maturity) on their first lags together with the above-mentioned uncertainty and volatility indices. We find that credit risk premiums positively comove with the EPU index and with short-term volatility; while the coefficient for the VIX is negative, pointing out that a rise in global volatility reduces the gap between physical and risk-neutral probabilities of default. Moreover, in Tables 1.L3 and 1.L4 (Appendix 1.L), by means of fixed effects panel regressions, we show that there exists a negative relationship between credit risk premiums and fiscal space: the worse the fiscal position the higher the risk premiums demanded by risk-averse investors.

1.6 Concluding remarks

The present research attempts at estimating time-varying fiscal limits. The fiscal limit is defined as the maximum outstanding debt a government can sustain via budget surpluses in the future (as, e.g., in Bi, 2012; Leeper, 2013). The estimation is based on a novel sovereign credit-risk model that assumes that probabilities of default leave zero only when debt-to-GDP breaches the fiscal limit. The model offers tractable formulas to price credit-sensitive financial instruments.

[38] Our approach is different from the one employed in Gilchrist and Zakrajšek (2012) for corporate credit spreads. The authors use an empirical credit spread pricing framework to disentangle the spread into distinct components separating the drivers of the expected credit loss and the risk premiums (excess bond premiums in their jargon). According to their analysis, the shocks to credit risk premiums that are orthogonal to the current state of the economy cause economic activity contractions, disinflation, a plunge in the risk-free rates and a fall in the stock market.

[39] The Cboe VIX index is a widely used proxy to measure global expected volatility. Specifically, the VIX index is calculated to produce a measure of constant, 30-day expected volatility of the U.S. stock market, derived from real-time, mid-quote prices of $S\&P$ 500 index (SPX) call and put options. According to the Cboe website, on a global basis, the VIX index is one of the most recognized measures of volatility – widely reported by financial media and closely followed by a variety of market participants as a daily market indicator. For more details on the VIX and the other volatility measures used in the regressions, visit https://www.cboe.com/indices/.

These formulas are exploited to get fiscal limit estimates: since credit spreads are observed, the pricing formulas can be inverted to recover the fiscal limit prevailing on each sample date. To the best of our knowledge, this paper is the first to provide time-varying estimates of fiscal limits. Our application considers four advanced and four emerging economies: Canada, Japan, the UK, the US, and Brazil, China, India, Russia. Fiscal limit estimates show ample fluctuations over time. Moreover, the model succeeds in providing a good fit of sovereign Credit Default Swaps (CDSs), both in the time and maturity dimensions.

The estimated models predict a non-linear influence of fiscal conditions on credit spreads, in line with the findings of a wide body of empirical studies. Compared to standard regression-based analysis, our framework provides richer predictions of the changes in the term structure of credit spreads that can be expected from fiscal deterioration.

Because our model entails risk-averse investors, our approach provides us with estimates of credit risk premiums. From a quantitative point of view, we observe that a substantial part of the credit risk spreads is accounted for by credit risk premiums, in line with the findings of the purely-reduced-form default-intensity literature. Such hefty risk premiums translate into large discrepancies between the default probabilities adjusted for risk and the ones that are not.

Appendix

1.A CDS and bond pricing

As explained in Subsection 1.3.9, pricing CDS (i.e. solving for $S^{cds}_{t,h}$ in eq. 1.21) requires the computation of the following two conditional expectations: $\mathbb{E}_t[\mathcal{M}^n_{t,t+k}\mathcal{D}_{t+k}]$ and $\mathbb{E}_t[\mathcal{M}^n_{t,t+k}(1-\mathcal{D}_{t+k})]$. Appendix 1.F.1 shows that, in a model where: (i) the nominal s.d.f. between dates t and $t+1$ is if the form $\mathcal{M}^n_{t,t+1} = \exp(\varphi_0 + \varphi_1' w_{t+1} + \varphi_2(\mathcal{D}_{t+1} - \mathcal{D}_t))$, where $\mathcal{D}_t$ does not Granger-cause w_t, and (ii) $\underline{\lambda}_t$ is the default intensity defined through eq. (1.18), we have (when $\mathcal{D}_t = 0$):

$$\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h})\right] = \exp(h\varphi_0)K_{t,h}, \tag{a.1.1}$$

$$\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}\mathcal{D}_{t+h}\right] = \exp(h\varphi_0 + \varphi_2)\Big(\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1} + \cdots + w_{t+h})\}\right] - K_{t,h}\Big), \tag{a.1.2}$$

where

$$K_{t,h} \equiv \mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1} + \cdots + w_{t+h}) - \underline{\lambda}_{t+1} - \cdots - \underline{\lambda}_{t+h}\}\right].$$

Using the following notation:

$$f_{n-1,n} = -\log K_{t,n} + \log K_{t,n-1},$$

we have:

$$K_{t,h} = exp(f_{0,1} + \cdots + f_{h-1,h}) \tag{a.1.3}$$

While $\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1} + \cdots + w_{t+h})\}\right]$ (in a.1.2) is computed through recursive formulas (see a.1.7 and a.1.8), Appendix 1.F.2 proposes approximations to the $f_{h-1,h}$'s (building on Wu and Xia, 2016). These approximations further allow to compute approximated values of $K_{t,h}$ (using eq. a.1.3) and further, of $\mathbb{E}_t[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h})]$ and $\mathbb{E}_t[\mathcal{M}^n_{t,t+h}\mathcal{D}_{t+h}]$, using (a.1.1) and (a.1.2), respectively.

1.B Perpetuities and debt accumulation process

1.B.1 Perpetuities

The date-t price of this perpetuity described in Subsection 1.3.4 is given by eq. (1.5), that is: $P_t = \sum_{h=1}^{\infty} \chi^{h-1} B_{t,h}$, where $B_{t,h}$ is the date-t price of a generic zero-coupon bond providing the nominal payoff $1-(1-RR)\mathcal{D}_{t+h}$ on date $t+h$. We have:

$$B_{t,h} = \mathbb{E}_t\Big(\mathcal{M}^n_{t,t+1} \times \cdots \times \mathcal{M}^n_{t+h-1,t+h}[(1-(1-RR)\mathcal{D}_{t+h}]\Big),$$

where the nominal s.d.f. $\mathcal{M}^n_{t,t+1} \equiv \mathcal{M}_{t,t+1}\exp(-\pi_{t+1})$ is given by (combining eqs. 1.1 and 1.4):

$$\exp\left\{\log(\delta) - \mu_c - \mu_\pi - (\gamma\Lambda_c + \Lambda_\pi)' w_{t+1} + (\gamma b_c + b_\pi)(\mathcal{D}_{t+1} - \mathcal{D}_t) - (\gamma\sigma_c + \sigma_\pi)\eta_{t+1}\right\}. \tag{a.1.4}$$

Because the η_t's are exogenous i.i.d. shocks of covariance matrix I, and using the following notations:

$$\varphi_0 = \log(\delta) - \mu_c - \mu_\pi + \frac{1}{2}(\gamma\sigma_c + \sigma_\pi)'(\gamma\sigma_c + \sigma_\pi), \quad \text{and} \quad \varphi_1 = -(\gamma\Lambda_c + \Lambda_\pi), \tag{a.1.5}$$

we obtain:

$$\begin{aligned} B_{t,h} &= \mathbb{E}_t\Big\{\exp[h\varphi_0 + \varphi_1'(w_{t+1}+\dots+w_{t+h}) + (\gamma b_c + b_\pi)\mathcal{D}_{t+h}][(1-(1-RR)\mathcal{D}_{t+h}]\Big\} \\ &= \mathbb{E}_t\Big\{\exp[h\varphi_0 + \varphi_1'(w_{t+1}+\dots+w_{t+h})]\{(1-[1-RR\exp(\gamma b_c + b_\pi)]\mathcal{D}_{t+h}\}\Big\}. \end{aligned}$$

Therefore, under (1.6), i.e., if $RR = \exp[-(\gamma b_c + b_\pi)]$, we get:

$$B_{t,h} = \mathbb{E}_t\Big\{\exp[h\varphi_0 + \varphi_1'(w_{t+1}+\dots+w_{t+h})]\Big\}. \quad \text{(a.1.6)}$$

The conditional expectation appearing on the right-hand side of the previous equation is easily computed recursively. Indeed, as shown in Appendix 1.G, if w_t follows a Gaussian VAR (as in eq. 1.2), we have:

$$\mathbb{E}_t\left[\exp\{u'(w_{t+1}+\dots+w_{t+h})\}\right] = \exp\Big(b_h(u) + a_h(u)'w_t\Big), \quad \text{(a.1.7)}$$

where functions $a_h(\bullet)$ and $b_h(\bullet)$ satisfy the following recursive equations:

$$\left\{\begin{array}{lcl} a_h(u) &=& \Phi'(a_{h-1}(u)+u) \\ b_h(u) &=& b_{h-1}(u) + \frac{1}{2}(a_{h-1}(u)+u)'(a_{h-1}(u)+u), \end{array}\right. \quad \text{(a.1.8)}$$

with $a_0(u) = 0$ and $b_0(u) = 0$.

1.B.2 Debt accumulation process

Let us denote by I_t the proceeds of date-t issuances and by X_t the resulting first payments (settled on date $t+1$). We have:

$$I_t = \sum_{j=1}^{\infty} \frac{\chi^{j-1}X_t}{(1+q_t)^j} = \frac{X_t}{1+q_t-\chi}.$$

Consider the date-t (residual) face value of those issuances that took place on date $t-h$. This face value is computed as the sum of future associated payoffs $\chi^{h+1}X_{t-h}, \chi^{h+2}X_{t-h}, \dots$, discounted using the issuance yield-to-maturity that materialized on date $t-h$, that is q_{t-h}. It is easily seen that it is equal to $\chi^h I_{t-h}$. As a consequence, and because current debt D_t is the sum of the (residual) face values of all past issuances, we obtain:[40]

$$D_t \equiv I_t + \chi I_{t-1} + \chi^2 I_{t-2} + \dots = I_t + \chi D_{t-1}. \quad \text{(a.1.9)}$$

Using $X_t = (1+q_t-\chi)I_t = (1+q_t-\chi)(D_t - \chi D_{t-1})$, past debt issuances give rise to the following debt payments at date $t+1$:

$$\begin{aligned} & CF_{t+1} = X_t + \chi X_{t-1} + \chi^2 X_{t-2} + \dots \\ = \; & (1+q_t-\chi)(D_t-\chi D_{t-1}) + \\ & \chi(1+q_{t-1}-\chi)(D_{t-1}-\chi D_{t-2}) + \chi^2(1+q_{t-2}-\chi)(D_{t-2}-\chi D_{t-3}) + \dots \\ = \; & D_t - \chi D_t + q_t(D_t-\chi D_{t-1}) + \chi q_{t-1}(D_{t-1}-\chi D_{t-2}) + \chi^2 q_{t-2}(D_{t-2}-\chi D_{t-3}) + \dots \quad \text{(a.1.10)} \end{aligned}$$

[40] This computation is based on the so-called "nominal valuation of debt securities," a standard international debt accounting principle (see Subsection 1.3.5, and in particular Footnote 15).

On date t, the sum of the issuance proceeds (I_t) and of the primary budget surplus (S_t) has to equate date-t payments associated with previous issuances (CF_t). That is: $I_t = CF_t - S_t$. Using eq. (a.1.9), we get:

$$D_{t+1} - \chi D_t = CF_{t+1} - S_{t+1}. \tag{a.1.11}$$

Substituting for CF_t (eq. a.1.10) into eq. (a.1.11), we have:

$$\begin{aligned} D_{t+1} &= D_t - S_{t+1} + \\ & \underbrace{q_t(D_t - \chi D_{t-1}) + \chi q_{t-1}(D_{t-1} - \chi D_{t-2}) + \chi^2 q_{t-2}(D_{t-2} - \chi D_{t-3}) + \dots}_{\text{interest payments on date } t+1 \equiv R_{t+1}} \end{aligned} \tag{a.1.12}$$

Denote real GDP by Y_t and GDP price deflator by P_t. The previous equation rewrites:

$$\frac{D_t}{Y_t P_t} = \frac{D_{t-1}}{Y_{t-1}P_{t-1}} \frac{Y_{t-1}P_{t-1}}{Y_t P_t} \left(1 + \frac{R_t}{D_{t-1}} - \frac{S_t}{D_{t-1}}\right).$$

Introducing the log debt-to-GDP ratio d_t, we obtain:

$$d_t = d_{t-1} - \Delta y_t - \pi_t + \log\left(1 + \frac{R_t}{D_{t-1}} - \frac{S_t}{D_{t-1}}\right). \tag{a.1.13}$$

Appendix 1.B.3 shows that the unconditional mean of the apparent debt interest rate R_t/D_{t-1} is equal to that of q_t, that we denote by $\bar{q}$. Moreover, let us denote by $\overline{sd}$ the unconditional mean of S_t/D_{t-1}. The previous equation can be reformulated as follows:

$$d_t = d_{t-1} - \Delta y_t - \pi_t + \log\left(1 + \bar{q} - \overline{sd}\right) + \log\left(1 + \frac{rd_t - sd_t}{1 + \bar{q} - \overline{sd}}\right),$$

where rd_t and sd_t are defined in (1.12). The approximated law of motion for d_t (eq. 1.14) is obtained by considering first-order approximations of the last two terms around $rd_t - sd_t = 0$.

1.B.3 Interest payment dynamics

Assuming $D_t \approx D_{t-1}$, we obtain the following recursive approximation for the interest payments (see eq. a.1.12):

$$R_{t+1} \approx D_t(1-\chi)(q_t + \chi q_{t-1} + \chi^2 q_{t-2} + \dots), \tag{a.1.14}$$

which gives

$$\frac{R_{t+1}}{D_t} \approx (1-\chi)q_t + \chi \frac{R_t}{D_{t-1}}. \tag{a.1.15}$$

Hence, the apparent interest rate is given by an exponential smoothing of the yield-to-maturities associated with past debt issuances. This implies in particular that, when the apparent debt interest rate R_t/D_{t-1} is stationary, then its unconditional mean is equal to that of q_t, i.e. $\mathbb{E}(R_t/D_{t-1}) = \mathbb{E}(q_t) = \bar{q}$. We therefore use:

$$rd_{t+1} \approx (1-\chi)(q_t - \bar{q}) + \chi rd_t. \tag{a.1.16}$$

1.C Approximation to the fiscal limit

This appendix explains how ℓ_t (defined in eq. 1.16) is approximated as an affine function of w_t (and hence of x_t since w_t is included in vector x_t). Specifically, we look for the vector a^ℓ and the scalar b^ℓ that are such that $\exp(\ell_t) \approx \exp(a^{\ell\prime} w_t + b^\ell)$. This is done by solving the following system:

$$\left\{\begin{array}{rcl} \mathbb{E}(\exp(\ell_t)) & = & \mathbb{E}(\exp(a^{\ell\prime} w_t + b^\ell)) \\ \mathbb{E}\left(\dfrac{\partial}{\partial w_{k,t}} \exp(\ell_t)\right) & = & \mathbb{E}\left(\dfrac{\partial}{\partial w_{k,t}} \exp(a^{\ell\prime} w_t + b^\ell)\right), \quad k \in \{1, \dots, n_w\}. \end{array}\right. \tag{a.1.17}$$

Appendix 1.J shows that System (a.1.17) is satisfied when, for all $k \in \{1, \dots, n_w\}$:

$$a_k^\ell = \frac{\sum_{h=1}^{+\infty} a_{h,k}^\ell \mathbb{E}\left(\exp\left(a_h^{\ell\prime} w_t + b_h^\ell\right)\right)}{\sum_{h=1}^{+\infty} \mathbb{E}\left(\exp\left(a_h^{\ell\prime} w_t + b_h^\ell\right)\right)}, \tag{a.1.18}$$

where vectors a_h^ℓ and scalars b_h^ℓ satisfy:

$$\exp\left(a_h^{\ell\prime} w_t + b_h^\ell\right) = \mathbb{E}_t\left[\mathcal{M}_{t,t+h} \exp(\Delta y_{t+1} + \dots + \Delta y_{t+h}) | \mathcal{D} \equiv 0\right]. \tag{a.1.19}$$

Appendix 1.J shows that these a_h^ℓ's and b_h^ℓ's are given by:

$$\begin{aligned} a_h^\ell &= (I - \Phi'^h)\kappa_0 \\ b_h^\ell &= h\kappa_1 + (h-1)\kappa_0'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}(h-1)\kappa_0'\kappa_0 - \kappa_0'(I-\Phi)^{-1}(\Phi - \Phi^h)(\Lambda_y - \gamma\Lambda_c + \kappa_0) + \\ &\quad + \frac{1}{2}\kappa_0'\left(\sum_{k=0}^{\infty} \Phi^k \Phi'^k - I - \Phi^h\left[\sum_{k=0}^{\infty} \Phi^k \Phi'^k\right]\Phi'^h\right)\kappa_0, \end{aligned}$$

where

$$\left\{\begin{array}{rcl} \kappa_0 & = & (I - \Phi')^{-1}\Phi'(\Lambda_y - \gamma\Lambda_c) \\ \kappa_1 & = & \log(\delta) - \gamma\mu_c + \mu_y + \frac{1}{2}(\sigma_y - \gamma\sigma_c)'(\sigma_y - \gamma\sigma_c) + \frac{1}{2}(\Lambda_y - \gamma\Lambda_c)'(\Lambda_y - \gamma\Lambda_c), \end{array}\right.$$

and with $vec\left(\sum_{k=0}^{\infty} \Phi^k \Phi'^k\right) = \left(I_{n_w^2} - \Phi \otimes \Phi\right)^{-1} vec(I_{n_w^2})$.

Once the a_k^ℓ's, have been computed (using eq. a.1.18), we compute b^ℓ as follows:

$$b^\ell = \log \mathbb{E}\left(\exp(\ell_t)\right) - \frac{1}{2} a^{\ell\prime} \Omega_w a^\ell,$$

where $\mathbb{E}\left(\exp(\ell_t)\right)$ is the denominator of the fraction appearing on the right-hand side of eq. (a.1.18).

1.D Constraints on Λ_i's and σ_i's

The restrictions presented in Subsection 1.4.2 imply that the loadings Λ are defined by three parameters: $\Lambda_{c,1} = \Lambda_{y,1}$, $\Lambda_{s,1}$, and $\Lambda_{s,3}$. We define two additional restrictions on these parameters to discipline and facilitate the estimation. These restrictions pertain to second-order moments, namely the unconditional variance of the persistent component of Δy_t, and the correlation between the persistent components of Δy_t and sd_t. (The latter measures the procyclicality of the budget surplus.) The model parametrization is forced to replicate these sample moments.[41] To compute the latter, we need proxies of the persistent components of Δy_t and sd_t. We employ the Hodrick-Prescott filter on the observed series of Δy_t and sd_t to get such proxies (with a parameter of 100).

The HP filter is also involved in the calibration of vectors σ_c, σ_y, σ_π, and σ_s, which characterize the volatile components of Δc_t, Δy_t, π_t, and sd_t, respectively (see eqs. 1.1 and 1.15): we get initial proxies for the four shocks $\sigma_c'\eta_t$, $\sigma_y'\eta_t$, $\sigma_s'\eta_t$, and $\sigma_\pi'\eta_t$ by subtracting HP-based trends from Δc_t, Δy_t, π_t, and sd_t. The Cholesky decomposition of the sample covariance matrix of the shocks' proxies then provides us with vectors σ_c, σ_y, σ_π, and σ_s that are such that the model-implied covariance matrix of the four shocks is the same as that of their HP-based proxies.[42] Let us stress that the HP-based proxies are only used to help calibrate the model; the Kalman filter will ultimately estimate the latent factors w_t and the shocks η_t.

The constant μ_y and μ_π are set to the samples averages of Δy_t and π_t. The unconditional average of the log debt-to-GDP ratio ($\bar{d} = \mathbb{E}(d_t)$) is taken equal to the maximum between its sample mean and the logarithm of 40% (which is binding for Russia and China, see sixth line of Table 1.1). An internal-consistency constraint weighs on $\bar{d}$ and $\overline{sd}$, the latter being the unconditional average of sd_t (see eq. 1.12). Indeed, as can be seen from the expanded expression of μ_x (Appendix 1.K), this vector depends on both $\bar{d}$ and $\overline{sd}$. As a result, for an arbitrary pair $(\bar{d}, \overline{sd})$, the third to last component of $\mathbb{E}(x_t) = (I - \Phi_x)^{-1}\mu_x$ does not coincide with $\bar{d}$. In other words, $\bar{d}$ and $\overline{sd}$ cannot be set independently. We address this issue by numerically determining $\overline{sd}$ for a given value of $\bar{d}$. (This fixed point problem is simply solved by the Gauss-Newton algorithm, which converges in a few iterations.)

1.E Derivation of the fiscal limit

1.E.1 Derivation of the government debt value

We start with the derivation of the market value of government debt. Let us denote by $\mathscr{D}_t$ the date-t market value of the outstanding debt, by H ($\leq \infty$) the maximum maturity of the bonds issued by the government, and by $\omega_{t,h}$ the number of unit-face-value zero-coupon bonds of residual maturity h in the government portfolio at the end of date t. We recall that, on date t, the value of a maturity-h zero-coupon bonds of unit face value is $B_{t,h}$ (see eq. 1.7). In this context, we have:

$$\mathscr{D}_t = \sum_{h=1}^{H} \omega_{t,h} B_{t,h}.$$

Assume that the government repays all its debt at the beginning of date $t+1$; it then has to repay $\sum_{h=1}^{H} \omega_{t,h} B_{t+1,h-1}$. This is financed by new issuances $\{\omega_{t+1,1}, \dots, \omega_{t+1,H}\}$, whose proceeds are $\sum_{h=1}^{H} \omega_{t,h} B_{t+1,h-1}$,

[41] We remove the largest two absolute values of the series before computing the sample variances. This is to correct for the extreme volatility associated with the COVID-19 period.

[42] The covariance matrix of the shocks' proxies is based on trimmed data: the two quarters with the largest absolute values are removed. These values correspond to the peak of the COVID pandemic (2020Q2 and 2020Q3).

as well as by primary budget surpluses S_{t+1}, yielding the following government budget constraint:

$$\begin{aligned} & \sum_{h=1}^{H} \omega_{t,h} B_{t+1,h-1} = S_{t+1} + \sum_{h=1}^{H} \omega_{t+1,h} B_{t+1,h} \\ \Leftrightarrow \quad & \sum_{h=1}^{H} \omega_{t,h} (B_{t+1,h-1} - \exp(i_t) B_{t,h}) + \exp(i_t) \sum_{h=1}^{H} \omega_{t,h} B_{t,h} = S_{t+1} + \sum_{h=1}^{H} \omega_{t+1,h} B_{t+1,h}. \end{aligned} \tag{a.1.20}$$

Multiply both sides of the previous equation by the nominal stochastic discount factor (s.d.f.) $\mathcal{M}^n_{t,t+1}$ and take expectations conditional on the information available on date t. Then, consider one of the terms of the first sum:

$$\begin{aligned} \mathbb{E}_t(\mathcal{M}^n_{t,t+1} B_{t+1,h-1} - \mathcal{M}^n_{t,t+1} \exp(i_t) B_{t,h}) &= \mathbb{E}_t(\mathcal{M}^n_{t,t+1} \mathbb{E}_{t+1}(\mathcal{M}^n_{t+1,t+h})) - \exp(i_t) B_{t,h} \mathbb{E}_t(\mathcal{M}^n_{t,t+1}) \\ &= \mathbb{E}_t(\mathcal{M}^n_{t,t+h}) - B_{t,h} = 0, \end{aligned}$$

where we have used $\mathbb{E}_t(\mathcal{M}^n_{t,t+1}) = \exp(-i_t)$ and $B_{t,h} \equiv \mathbb{E}_t(\mathcal{M}^n_{t,t+h})$. eq. (a.1.20) then becomes:

$$\underbrace{\sum_{h=1}^{H} w_{t,h} B_{t,h}}_{\mathscr{D}_t} = \mathbb{E}_t \left(\mathcal{M}^n_{t,t+1} \left[S_{t+1} + \underbrace{\sum_{h=1}^{H} w_{t+1,h} B_{t+1,h}}_{\mathscr{D}_{t+1}} \right] \right), \tag{a.1.21}$$

and, under the transversality condition $\lim_{k \to +\infty} \mathbb{E}_t(\mathcal{M}^n_{t,t+k} \mathscr{D}_{t+k}) = 0$ (ruling out government debt bubbles), it comes that:

$$\mathscr{D}_t = \mathbb{E}_t \left(\sum_{h=1}^{+\infty} \mathcal{M}^n_{t,t+h} S_{t+h} \right). \tag{a.1.22}$$

1.E.2 Fiscal limit

On each date, there is a maximum surplus S^*_t that can be generated by the government. This surplus is implicit in the peak of the Laffer curve, which represents the reverse bell-shaped relationship between the average tax rate and government revenues. Since $S_t \leq S^*_t$ for all date t, and because the s.d.f. $\mathcal{M}^n_{t,t+1}$ is strictly positive, eq. (a.1.21) implies that:

$$\mathscr{D}_t \leq \mathbb{E}_t(\mathcal{M}^n_{t,t+1}[S^*_{t+1} + \mathscr{D}_{t+1}]),$$

which leads to:

$$\mathscr{D}_t \leq \mathbb{E}_t \left(\sum_{h=1}^{+\infty} \mathcal{M}^n_{t,t+h} S^*_{t+h} \right). \tag{a.1.23}$$

1.F Approximate CDS pricing formula

The date-t value of the protection premium is given by

$$\begin{aligned} & \mathbb{E}_t\left\{\sum_{k=1}^{h}\mathcal{M}^n_{t,t+k}(\mathcal{D}_{t+k}-\mathcal{D}_{t+k-1})(1-RR)\mathbb{E}_{t+k}(\mathcal{M}^n_{t+k,t+h})\right\} \\ = & \mathbb{E}_t\left\{\sum_{k=1}^{h}\mathcal{M}^n_{t,t+h}(\mathcal{D}_{t+k}-\mathcal{D}_{t+k-1})(1-RR)\right\} \\ = & (1-RR)\mathbb{E}_t\left\{\mathcal{M}^n_{t,t+h}(D_{t+h}-D_t)\right\}, \end{aligned}$$

where we have used $\mathcal{M}^n_{t,t+h}=\mathcal{M}^n_{t,t+k}\mathcal{M}^n_{t+k,t+h}$, as well as the law of iterated expectations.

Using the previous expression in (1.20), we obtain eq. (1.21), that is:

$$S^{cds}_{t,h}=(1-RR)\frac{\mathbb{E}_t\left\{\mathcal{M}^n_{t,t+h}D_{t+h}\right\}}{\mathbb{E}_t\left\{\sum_{k=1}^{h}\mathcal{M}^n_{t,t+k}(1-\mathcal{D}_{t+k})\right\}}. \tag{a.1.24}$$

As a consequence, the computation of the CDS spread $S^{cds}_{t,h}$ necessitates the knowledge of the following two conditional expectations: $\mathbb{E}_t[\mathcal{M}^n_{t,t+h}\mathcal{D}_{t+h-1}]$ and $\mathbb{E}_t[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h})]$, which can be seen as "binary CDSs" in the sense that they correspond to date-t prices of instruments providing a binary payoff (0 or 1) depending on the default status of the government on date $t+h$.

Subsection 1.F.1 shows that these two prices are given by combinations of conditional exponential expectations of future values of $(x'_t,\underline{\lambda}_t)'$. Subsection 1.F.2 and 1.F.3 explain how to approximate these conditional expectations.

1.F.1 Prices of binary CDSs

We consider the following situation:

(a) The s.d.f. $\mathcal{M}^n_{t,t+1}$ (see eq. 1.4, or eq. a.1.4) is of the form:

$$\mathcal{M}^n_{t,t+1}=\exp\left(\varphi_0+\varphi_1'w_{t+1}+\varphi_2(\mathcal{D}_{t+1}-\mathcal{D}_t)-(\gamma\sigma_c+\sigma_\pi)'\eta_{t+1}-\frac{1}{2}(\gamma\sigma_c+\sigma_\pi)'(\gamma\sigma_c+\sigma_\pi)\right), \tag{a.1.25}$$

where $\varphi_2=\gamma b_c+b_\pi$, and where we use the notations introduced in (a.1.5):

$$\varphi_0=\log(\delta)-\mu_c-\mu_\pi+\frac{1}{2}(\gamma\sigma_c+\sigma_\pi)'(\gamma\sigma_c+\sigma_\pi),\quad \varphi_1=-\gamma\Lambda_c-\Lambda_\pi. \tag{a.1.26}$$

(b) $\underline{\lambda}_t$ is the default intensity, defined as a nonlinear function of x_t (see eq. 1.18):

$$\underline{\lambda}_t=\max[0,\alpha(d_{t-1}-\ell_{t-1}-\nu_t)]. \tag{a.1.27}$$

Obviously, $\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h-1})\right]$ and $\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h})\right]$ are equal to zero if $\mathcal{D}_t=1$. In the following, we proceed under the assumption that $\mathcal{D}_t=0$.

- Computation of $\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}(1-\mathcal{D}_{t+h})\right]$. We have:

$$\begin{aligned} & \mathbb{E}_t\left[\mathcal{M}^n_{t,t+1}\times\cdots\times\mathcal{M}^n_{t+h-1,t+h}(1-\mathcal{D}_{t+h})\right] \\ = & \exp\left(h\varphi_0\right)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})+\varphi_2\mathcal{D}_{t+h}\}(1-\mathcal{D}_{t+h})\right] \\ = & \exp(h\varphi_0)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})\}\mathbb{1}_{\{\mathcal{D}_{t+h}=0\}}\right] \\ = & \exp(h\varphi_0)\mathbb{E}_t\left[\mathbb{E}_t\left(\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})\}\mathbb{1}_{\{\mathcal{D}_{t+h}=0\}}|\underline{w_{t+h}}\right)\right] \\ = & \exp(h\varphi_0)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})-\underline{\lambda}_{t+1}-\cdots-\underline{\lambda}_{t+h}\}\right], \end{aligned} \quad \text{(a.1.28)}$$

where the last equality results from the fact that $\mathcal{D}_t$ does not Granger-cause w_t, combined with the fact that Granger's and Sims' types of causality are equivalent (Sims, 1972).

- Computation of $\mathbb{E}_t\left[\mathcal{M}^n_{t,t+h}\mathcal{D}_{t+h}\right]$. We have:

$$\begin{aligned} & \mathbb{E}_t\left[\mathcal{M}^n_{t,t+1}\times\cdots\times\mathcal{M}^n_{t+h-1,t+h}\mathcal{D}_{t+h}\right] \\ = & \exp(h\varphi_0)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})+\varphi_2\mathcal{D}_{t+h}\}\mathcal{D}_{t+h}\right] \\ = & \exp(h\varphi_0+\varphi_2)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})\}\mathbb{1}_{\{\mathcal{D}_{t+h}=1\}}\right] \\ = & \exp(h\varphi_0+\varphi_2)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})\}\left(1-\mathbb{1}_{\{\mathcal{D}_{t+h}=0\}}\right)\right] \\ = & \exp(h\varphi_0+\varphi_2)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})\}\right] \\ & -\exp(h\varphi_0+\varphi_2)\mathbb{E}_t\left[\exp\{\varphi_1'(w_{t+1}+\cdots+w_{t+h})-\underline{\lambda}_{t+1}-\cdots-\underline{\lambda}_{t+h}\}\right], \end{aligned} \quad \text{(a.1.29)}$$

using the same reasons as the ones invoked to obtain (a.1.28).

The CDS spread $S^{cds}_{t,h}$ is then obtained by substituting for the conditional expectations in (a.1.24), using (a.1.28) and (a.1.29). In Subsection 1.F.2, we explain how to approximate these conditional expectations.

1.F.2 Approximating the conditional expectations appearing in eq. (a.1.28), (a.1.29), and (a.1.44)

These conditional expectations are of the form:

$$K_{t,n} \equiv \mathbb{E}_t\left[\exp(\varphi'(x_{t+1}+\cdots+x_{t+n})-(\underline{\lambda}_{t+1}+\cdots+\underline{\lambda}_{t+n})\right], \quad \text{(a.1.30)}$$

where $x_t=[w_t',d_t,rd_t,q_t,w_{t-1},d_{t-1},\nu_t]'$. (We set $\varphi=[\varphi_1',0,0,0,0',0,0]'$.)

Using the notation:

$$f_{n-1,n} = -\log K_{t,n}+\log K_{t,n-1}, \quad \text{(a.1.31)}$$

we have:

$$K_{t,n} = \exp(f_{0,1}+\cdots+f_{n-1,n}). \quad \text{(a.1.32)}$$

Following Wu and Xia (2016), we approximate $K_{t,n}$ by, first, determining approximations to the $f_{h-1,h}$'s, and, second, substituting for the $f_{h-1,h}$s' into (a.1.32).

Using, in (a.1.30), that, for any random variable Z, we have $\log \mathbb{E}[\exp(Z)] \approx \mathbb{E}(Z) + 1/2\mathbb{V}ar(Z)$ (the approximation being exact in the Gaussian case), and substituting for $K_{t,n}$ and $K_{t,n-1}$ in (a.1.31) yields:

$$\begin{aligned} f_{n-1,n} \approx\ & \mathbb{E}_t(-\varphi' x_{t+n} + \underline{\lambda}_{t+n}) \\ & -\frac{1}{2}\mathbb{V}ar_t\left(-\varphi' x_{t+n} + \underline{\lambda}_{t+n}\right) - \mathbb{C}ov_t\left(-\varphi' x_{t+n} + \underline{\lambda}_{t+n}, \sum_{i=1}^{n-1}(-\varphi' x_{t+i} + \underline{\lambda}_{t+i})\right). \end{aligned} \quad \text{(a.1.33)}$$

Let us introduce the following notations:

$$\lambda_t = \alpha(d_{t-1} - \ell_{t-1} + \nu_t), \quad \text{(a.1.34)}$$

which implies $\underline{\lambda}_t = \max(0, \lambda_t)$ (according to eqs. 1.18 or a.1.27), and

$$p_{t,n} = \mathbb{P}_t(d_{t+n} > \ell_{t+n}).$$

In the spirit of Wu and Xia (2016), exploiting the fact that λ_t is a expected to be a persistent process, we have, for $0 < n$ and $0 \le j \le n$:

$$\mathbb{C}ov_t(-\varphi' x_{t+n}, \underline{\lambda}_{t+n-j}) \approx p_{t,n-j}\mathbb{C}ov_t\left[-\varphi' x_{t+n}, \lambda_{t+n-j}\right], \quad \text{(a.1.35)}$$
$$\mathbb{C}ov_t(\underline{\lambda}_{t+n}, \underline{\lambda}_{t+n-j}) \approx p_{t,n-j}\mathbb{C}ov_t\left[\lambda_{t+n}, \lambda_{t+n-j}\right]. \quad \text{(a.1.36)}$$

Using the last two equations, we obtain an approximation to (a.1.33):

$$\begin{aligned} f_{n-1,n,t} \approx\ & \mathbb{E}_t\left[-\varphi' x_{t+n} + \underline{\lambda}_{t+n}\right] \quad \text{(a.1.37)} \\ & -\frac{1}{2}\left(p_{t,n}\mathbb{V}ar_t\left[-\varphi' x_{t+n} + \lambda_{t+n}\right] + (1-p_{t,n})\mathbb{V}ar_t(-\varphi' x_{t+n})\right) \\ & -\sum_{j=1}^{n-1}\left\{p_{t,j}\mathbb{C}ov_t\left[-\varphi' x_{t+n} + \lambda_{t+n}, -\varphi' x_{t+j} + \lambda_{t+j}\right] + (1-p_{t,j})\mathbb{C}ov_t(-\varphi' x_{t+n}, -\varphi' x_{t+j})\right\}. \end{aligned}$$

Introducing the following notations:[43]

$$\lambda_t =: a + b' x_t \quad \text{and} \quad \dot{b} := -\varphi_1, \quad \text{(a.1.38)}$$
$$\mu_{t,n} := \mathbb{E}_t(x_{t+n}), \quad \text{and} \quad \Gamma_{n,j} := \mathbb{C}ov_t(x_{t+n}, x_{t+n-j}), \quad \text{(a.1.39)}$$
$$\mu_{\lambda,t,n} := \mathbb{E}_t(\lambda_{t+n}) = a + b'\mu_{t,n}, \quad \text{and} \quad \sigma_{\lambda,n} := \sqrt{\mathbb{V}ar_t(\lambda_{t+n})} = \sqrt{b'\Gamma_{n,0}b}, \quad \text{(a.1.40)}$$

[43]Using (a.1.34), and since $x_t = [w_t', d_t, rd_t, q_t, w_{t-1}, d_{t-1}, \nu_t]'$ (see Subsection 1.3.8), we have, in particular: $a = -\alpha b^{\ell}$ and $b = [0', 0, 0, 0, -\alpha a^{\ell'}, \alpha, 1]'$, where $\ell_t = a^{\ell'} w_t + b^{\ell}$ (see Appendices 1.C and 1.J for details regarding the derivation of the affine expression of the fiscal limit ℓ_t).

Eq. (a.1.37) rewrites:

$$\begin{aligned} f_{n-1,n,t} \approx\ & \dot{b}'\mu_{t,n} + \Phi\left(\mu_{\lambda,t,n}/\sigma_{\lambda,n}\right)\mu_{\lambda,t,n} + \phi(-\mu_{\lambda,t,n}/\sigma_{\lambda,n})\sigma_{\lambda,n} \\ & -\frac{1}{2}\left(p_{t,n}\left[\dot{b}+b\right]'\Gamma_{n,0}\left[\dot{b}+b\right] + \left[1-p_{t,n}\right]\dot{b}'\Gamma_{n,0}\dot{b}\right) \\ & -\sum_{j=1}^{n-1}\left\{p_{t,n-j}\left[\dot{b}+b\right]'\Gamma_{n,j}\left[\dot{b}+b\right] + \left[1-p_{t,n-j}\right]\dot{b}'\Gamma_{n,j}\dot{b}\right\}. \end{aligned} \tag{a.1.41}$$

(Note that $\Gamma_{n,0} = \mathbb{V}ar_t(x_{t+n})$.)

Appendix 1.F.3 proposes a quick way to compute $\mu_{t,n}$ and $\Gamma_{n,j}$.

1.F.3 Computation of $\mu_{t,n}$ and $\Gamma_{n,j}$

Recall x_t's law of motion (eq. 1.19):

$$x_t = \mu_x + \Phi_x x_{t-1} + \Sigma_x \varepsilon_{x,t}, \quad \varepsilon_{x,t} = [\varepsilon_t', \eta_t']' \sim i.i.d.\mathcal{N}(0, I),$$

where the expanded expressions of μ_x, Φ_x, and Σ_x are given in Appendix 1.K.

Using the notation $\Omega = \Sigma_x\Sigma_x'$, we have:

$$\begin{cases} \mu_{t,n} = \mathbb{E}_t(x_{t+n}) = (I-\Phi_x)^{-1}(I-\Phi_x^n)\mu_x + \Phi_x^n x_t, \\ \Gamma_{n,0} = \mathbb{V}ar_t(x_{t+n}) = \Omega + \Phi_x\Gamma_{n-1,0}\Phi_x', \quad \text{with} \quad \Gamma_{1,0} = \Omega \\ \qquad = \Omega + \Phi_x\Omega\Phi_x' + \dots + \Phi_x^{n-1}\Omega\Phi_x^{n-1\prime}, \\ \Gamma_{n,j} = \mathbb{C}ov_t(x_{t+n}, x_{t+n-j}) = \Phi_x^j\Gamma_{n-j,0} \quad \text{if} \quad n-j>0. \end{cases}$$

The estimation involves a large number of computations of the $\Gamma_{n,j}$'s. In order to speed up the computation, one can employ the following approach.

Consider a vector β of dimension n_x, that is the dimension of x_t, and let us denote by ξ_i^β the vector defined by $\xi_i^\beta = (\Phi_x^i)'\beta$ (β will typically be b, or $(b+\dot{b})$, see eq. a.1.41).

Because we have $\Gamma_{n,j} = \Phi_x^j\Omega + \Phi_x^{j+1}\Omega\Phi_x' + \dots + \Phi_x^{n-1}\Omega\Phi_x^{n-1-j\prime}$, it comes that:

$$\beta'\Gamma_{n,j}\beta = \xi_j^{\beta\prime}\Omega\xi_0^\beta + \xi_{j+1}^{\beta}{}'\Omega\xi_1^\beta + \dots + \xi_{n-1}^{\beta}{}'\Omega\xi_{n-1-j}^\beta. \tag{a.1.42}$$

Let us consider a maximal value for n, say H, and let us denote by Ξ_β the $n_x \times (H+1)$ matrix whose i^{th} column is ξ_{i-1}^β. It can then be seen that the (j,k) entry of $\Psi^\beta := \Xi_\beta{}'\Omega\Xi_\beta$—which is a matrix of dimension $(H+1)\times(H+1)$—is equal to $\xi_{j-1}^{\beta}{}'\Omega\xi_{k-1}^\beta$. The sum of the entries $(j+1,1)$, $(j+2,2)$, …, $(j+k,k)$ of Ψ^β therefore is

$$\xi_j^{\beta\prime}\Omega\xi_0^\beta + \xi_{j+1}^{\beta}{}'\Omega\xi_1^\beta + \dots + \xi_{j+k-1}^{\beta}{}'\Omega\xi_{k-1}^\beta,$$

which is equal to $\beta'\Gamma_{j+k,j}\beta$ according to (a.1.42). Equivalently, $\beta'\Gamma_{n,j}\beta$ is the sum of the entries $(j+1,1)$, $(j+2,2)$, …, $(n, n-j)$ of Ψ^β.

In particular, the entry $(1,1)$ of Ψ^β is equal to $\beta'\Gamma_{1,0}\beta$, the sum of the entries $(1,1)$ and $(2,2)$ is equal to $\beta'\Omega\beta + \beta'\Phi_x\Omega\Phi_x'\beta = \beta'\Gamma_{2,0}\beta$, and, more generally, the sum of the entries $(1,1), \dots, (n-1,n-1)$ of Ψ^β is equal to $\beta'\Gamma_{n,0}\beta$.

1.G Multi-horizon Laplace-transform in the context of a Gaussian VAR

If w_t follows a Gaussian VAR, that is if

$$w_t = \Phi w_{t-1} + \varepsilon_t, \tag{a.1.43}$$

where $\varepsilon_t \sim i.i.d.\mathcal{N}(0, I)$, then we have:

$$\mathbb{E}_t\left[\exp\{u'(w_{t+1} + \cdots + w_{t+h})\}\right] = \exp(b_h(u) + a_h(u)'w_t),$$

where functions $a_h(\bullet)$ and $b_h(\bullet)$ recursively satisfy:

$$\left\{\begin{array}{lcl} a_h(u) & = & \Phi'(a_{h-1}(u) + u) \\ b_h(u) & = & b_{h-1}(u) + \frac{1}{2}(a_{h-1}(u) + u)'(a_{h-1}(u) + u), \end{array}\right.$$

with $a_0(u) = 0$ and $b_0(u) = 0$.

Proof. If $\mathbb{E}_t\left[\exp\{u'(w_{t+1} + \cdots + w_{t+h-1})\}\right] = \exp(b_{h-1}(u) + a_{h-1}(u)'w_t)$ holds for any vector u, then:

$$\begin{aligned} & \mathbb{E}_t\left[\exp\{u'(w_{t+1} + \cdots + w_{t+h})\}\right] \\ = \; & \mathbb{E}_t\left[\exp\{u'w_{t+1}\}\mathbb{E}_{t+1}\left[\exp\{u'(w_{t+2} + \cdots + w_{t+h})\}\right]\right] \\ = \; & \mathbb{E}_t[\exp\{u'w_{t+1} + b_{h-1}(u) + a_{h-1}(u)'w_{t+1}\}] \quad \text{(using the recursive assumption)} \\ = \; & \mathbb{E}_t[\exp\{(a_{h-1}(u) + u)'w_{t+1} + b_{h-1}(u)\}] \\ = \; & \mathbb{E}_t\left[\exp\left\{b_{h-1}(u) + [\Phi'(a_{h-1}(u) + u)]'w_t + \frac{1}{2}(a_{h-1}(u) + u)'(a_{h-1}(u) + u)\right\}\right], \end{aligned}$$

where the last equality results from eq. (a.1.43), that is w_t's law of motion. □

1.H Risk-free versus perpetuities yields

1.H.1 Risk-free yields

Considering the same situation as the one described in Appendix 1.F.1, let us derive the date-t price of a maturity-h nominal risk-free zero-coupon bond:

$$\begin{aligned} & \mathbb{E}_t\left[\mathcal{M}^n_{t,t+1} \times \cdots \times \mathcal{M}^n_{t+h-1,t+h}\right] \\ = \; & \exp(h\varphi_0)\mathbb{E}_t\left[\exp\{\varphi_1'(x_{t+1} + \cdots + x_{t+h}) + \varphi_2\mathcal{D}_{t+h}\}\right] \\ = \; & \exp(h\varphi_0)\mathbb{E}_t\left[\exp\{\varphi_1'(x_{t+1} + \cdots + x_{t+h})\}\left(\exp(\varphi_2) + \mathbb{1}_{\{\mathcal{D}_{t+h}=0\}}(1 - \exp(\varphi_2))\right)\right] \\ = \; & \exp(h\varphi_0 + \varphi_2)\mathbb{E}_t\left[\exp\{\varphi_1'(x_{t+1} + \cdots + x_{t+h})\}\right] + \\ & \exp(h\varphi_0)(1 - \exp(\varphi_2))\mathbb{E}_t\left[\exp\{\varphi_1'(x_{t+1} + \cdots + x_{t+h}) - \underline{\lambda}_{t+1} - \cdots - \underline{\lambda}_{t+h}\}\right], \end{aligned} \tag{a.1.44}$$

where the last equality makes use eq. (a.1.28).

The fact that $\underline{\lambda}_t$ appears in the risk-free bond pricing formula implies that risk-free yields have to rely on the approximate formula presented in Appendix 1.F.2 to be computed. As a result, risk-free yields are non-linear functions of x_t, contrary to the defaultable zero-coupon bonds issued by the government (see eq. a.1.6 in Appendix 1.B.1).

1.H.2 Pricing the decaying-coupon perpetuity

Subsection 1.3.4 (and more precisely eqs. 1.5 and 1.7) shows that the price of the perpetuity is of the form:

$$\mathcal{P}_t \equiv \sum_{i=1}^{\infty} \chi^{i-1} \exp[B_i + A_i' w_t], \tag{a.1.45}$$

where $B_i = i\bar{\varphi}_0 + b_i(\bar{\varphi}_1)$ and $A_i' w_t = a_i(\bar{\varphi}_1)' w_t$, where $\bar{\varphi}_0$, $\bar{\varphi}_1$, as well as functions $a_i(\bullet)$ and $b_i(\bullet)$ are defined in Appendix 1.B.1—the proof is given in Appendix 1.G. By definition, the yield-to-maturity of the perpetuity, denoted by q_t, satisfies:

$$\mathcal{P}_t = \sum_{h=1}^{\infty} \frac{\chi^{h-1}}{(1+q_t)^h}.$$

The right-hand-side sum of the previous expression is equal to

$$\mathcal{P}(q_t) \equiv \frac{1}{1+q_t-\chi}. \tag{a.1.46}$$

Therefore, the yield-to-maturity q_t of the perpetuity is the solution of the following equation $\mathcal{P}_t = \mathcal{P}(q_t)$, where $\mathcal{P}_t$ is given by eq. (a.1.45). Solving for q_t is straightforward and leads to:

$$q_t = \frac{1}{\sum_{i=1}^{\infty} \chi^{i-1} \exp[B_i + A_i' w_t]} - (1-\chi),$$

which shows that q_t is not an affine function of w_t (and therefore of x_t). However, because the perpetuity is a collection of zero-coupons of price $B_{t,h}$ (with geometrically-decaying weights, see eq. 1.5), the yield-to-maturity of the perpetuity is expected to be close to the yield of an "average" zero-coupon, that is to one of the $r_{t,h}$'s, where $r_{t,h} = -\frac{1}{h}B_h - \frac{1}{h}A_h' w_t$. Practically, we look for the maturity $h \in \mathbb{N}$ that minimizes the deviation between $\mathbb{V}ar(\mathcal{P}_t)$ and $\mathbb{V}ar(\mathcal{P}(r_{t,h}))$ (where function $\mathcal{P}(\bullet)$ is defined in eq. a.1.46). Formally, we use the following approximation:

$$\begin{aligned} q_t &\approx a_{h^*}' x_t + b_{h^*}, \\ \text{where } h^* &= \underset{h\in\mathbb{N}}{\operatorname{argmin}} |\mathbb{V}ar(\mathcal{P}_t) - \mathbb{V}ar(\mathcal{P}(r_{t,h}))|. \end{aligned} \tag{a.1.47}$$

It remains to explain how $\mathbb{V}ar(\mathcal{P}_t)$ and $\mathbb{V}ar(\mathcal{P}(r_{t,h}))$ are computed.

- The approximation of $\mathbb{V}ar(\mathcal{P}(r_{t,h}))$ is based on Taylor expansions of $\mathcal{P}(q)$. Specifically, a fourth-order Taylor expansion of $q \mapsto \mathcal{P}(q) = \frac{1}{1+q-\chi}$ around q_0 gives $\mathcal{P}(q) = \sum_{i=0}^{4} \frac{(q-q_0)^i}{(1+q_0-\chi)^{i+1}} + o((q-q_0)^4)$, leading to the following approximation of $\mathbb{E}(\mathcal{P}(q))$:

$$\frac{1}{1+\mathbb{E}(q)-\chi} + \frac{\mathbb{V}ar(q)}{(1+\mathbb{E}(q)-\chi)^3} + \frac{Skew(q)\mathbb{V}ar(q)^{3/2}}{(1+\mathbb{E}(q)-\chi)^3} + \frac{Kurt(q)\mathbb{V}ar(q)^2}{(1+\mathbb{E}(q)-\chi)^5}.$$

By the same token, using a second-order Taylor expansion of $q \mapsto \mathcal{P}(q)^2 = \frac{1}{(1+q-\chi)^2}$, we get the following approximation of $\mathbb{E}(\mathcal{P}(q)^2)$:

$$\frac{1}{(1+\mathbb{E}(q)-\chi)^2} + 3\frac{\mathbb{V}ar(q)}{(1+\mathbb{E}(q)-\chi)^4} - 4\frac{Skew(q)\mathbb{V}ar(q)^{3/2}}{(1+\mathbb{E}(q)-\chi)^5} + 5\frac{Kurt(q)\mathbb{V}ar(q)^2}{(1+\mathbb{E}(q)-\chi)^6}.$$

An approximation of $\mathbb{V}ar[\mathcal{P}(r_{t,h})] = \mathbb{E}[\mathcal{P}(r_{t,h})^2] - \mathbb{E}[\mathcal{P}(r_{t,h})]^2$ can then be obtained by employing the last two approximations of $\mathbb{E}[\mathcal{P}(r_{t,h})^2]$ and $\mathbb{E}[\mathcal{P}(r_{t,h})]$, replacing $\mathbb{E}(q)$ by $\mathbb{E}(r_{t,h}) = b_h$, and $\mathbb{V}ar(q)$ by $\mathbb{V}ar(r_{t,h}) = a_h'\Sigma_x a_h$ and—considering a Gaussian distribution for $r_{t,h}$—using $Skew(q) = 0$ and $Kurt(q) = 3$.

- Let us turn to the computation of $\mathbb{V}ar(\mathcal{P}_t)$, where $\mathcal{P}_t$ is given in eq. (a.1.45). We compute the variance in a recursive fashion. For this purpose, let us introduce the following notation:

$$\mathcal{P}_{t,h} \equiv \sum_{i=1}^{h} \chi^{i-1} \exp[B_i + A_i' w_t] \underset{h\to\infty}{\rightarrow} \mathcal{P}_t.$$

The variance of $\mathcal{P}_t$ can be approximated by $\mathbb{V}ar(\mathcal{P}_{t,H})$ for a sufficiently large H. The variance of $\mathcal{P}_{t,H}$ is computed recursively: We have $\mathbb{V}ar(\mathcal{P}_{t,0}) = 0$ and, $\mathbb{V}ar(\mathcal{P}_{t,h+1})$, $h \geq 1$, is given by:

$$\begin{aligned} & \mathbb{V}ar(\mathcal{P}_{t,h}) + \mathbb{V}ar(\chi^h \exp[B_{h+1} + A_{h+1}' w_t]) + 2\sum_{i=1}^{h} \mathbb{C}ov\left\{\chi^{i-1}\exp[B_i + A_i' w_t], \chi^h \exp[B_{h+1} + A_{h+1}' w_t]\right\} \\ = \; & \mathbb{V}ar(\mathcal{P}_{t,h}) + \chi^{2h}\exp(2B_{h+1})\left[\exp(2A_{h+1}'\mathbb{V}(w_t)A_{h+1}) - \exp(A_{h+1}'\mathbb{V}(w_t)A_{h+1})\right] + \\ & +2\chi^h \exp\left(B_{h+1} + \frac{1}{2}A_{h+1}'\mathbb{V}(w_t)A_{h+1}\right)\sum_{i=1}^{h}\chi^{i-1}\exp\left(B_i + \frac{1}{2}A_i'\mathbb{V}(w_t)A_i\right)\left[\exp\left(A_i'\mathbb{V}(w_t)A_{h+1}\right) - 1\right]. \end{aligned}$$

1.I Relationship between λ_t and $\lambda_t^{\mathbb{Q}}$

Let us recall the notations introduced at the beginning of Section 1.3, when presenting the information set up. On each date t, the representative agent observes the new information X_t, with $X_t = \{\mathcal{D}_t, \Delta c_t, \Delta y_t, \pi_t, S_t, w_t\}$. Hence, on date t, the total agent's information is $\mathcal{I}_t = \{X_t, X_{t-1}, \ldots\}$. Let us denote by X_t^* the partial information $\{w_t, \eta_t\}$. Using eqs. (1.1) and (1.15), it appears that $X_t = \{\mathcal{D}_t, X_t^*\}$.

By Bayes, we have:

$$f^{\mathbb{Q}}(\mathcal{D}_t|X_t^*, \mathcal{I}_{t-1}) = \frac{f^{\mathbb{Q}}(\mathcal{D}_t, X_t^*|\mathcal{I}_{t-1})}{f^{\mathbb{Q}}(X_t^*|\mathcal{I}_{t-1})}. \tag{a.1.48}$$

In our context, the s.d.f. $\mathcal{M}_{t,t+1}^n$ (see eq. 1.4, or equivalently eqs. a.1.4 and a.1.25) is of the form:

$$\mathcal{M}_{t,t+1}^n = \exp\left(\varphi_0 + \varphi_1' w_{t+1} + \varphi_2(\mathcal{D}_{t+1} - \mathcal{D}_t) - \varphi_3'\eta_{t+1} - \frac{1}{2}\varphi_3'\varphi_3\right),$$

where $\varphi_3 = \gamma\sigma_c + \sigma_\pi$.

Assume $\mathcal{D}_{t-1} = 0$. We have:

$$\begin{aligned} f^{\mathbb{Q}}(\mathcal{D}_t, X_t^* | \mathcal{I}_{t-1}) &= \frac{\mathcal{M}_{t-1,t}^n}{\mathbb{E}(\mathcal{M}_{t-1,t}^n | \mathcal{I}_{t-1})} f^{\mathbb{P}}(\mathcal{D}_t, X_t^* | \mathcal{I}_{t-1}) \\ &= \frac{\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t - \varphi_3' \eta_{t+1} - \frac{1}{2}\varphi_3'\varphi_3)}{\mathbb{E}[\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t)|\mathcal{I}_{t-1}]} f^{\mathbb{P}}(\mathcal{D}_t, X_t^* | \mathcal{I}_{t-1}) \\ &= \frac{\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t - \varphi_3' \eta_{t+1} - \frac{1}{2}\varphi_3'\varphi_3)}{\mathbb{E}[\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t)|\mathcal{I}_{t-1}]} f^{\mathbb{P}}(\mathcal{D}_t | X_t^*, \mathcal{I}_{t-1}) f^{\mathbb{P}}(X_t^* | X_{t-1}^*) \\ &= \frac{\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t - \varphi_3' \eta_{t+1} - \frac{1}{2}\varphi_3'\varphi_3)}{\mathbb{E}[\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t)|\mathcal{I}_{t-1}]} \times \\ &\quad (\mathcal{D}_t\{1 - \exp(-\underline{\lambda}_t)\} + (1 - \mathcal{D}_t)\{\exp(-\underline{\lambda}_t)\}) f^{\mathbb{P}}(X_t^* | X_{t-1}^*). \end{aligned} \quad \text{(a.1.49)}$$

Integrating both sides w.r.t. $\mathcal{D}_t$, we obtain:

$$f^{\mathbb{Q}}(X_t^* | \mathcal{I}_{t-1}) = \exp\left(\varphi_1' w_t - \varphi_3' \eta_{t+1} - \frac{1}{2}\varphi_3'\varphi_3\right) \frac{\exp(\varphi_2)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t)}{\mathbb{E}[\exp(\varphi_1' w_t + \varphi_2 \mathcal{D}_t)|\mathcal{I}_{t-1}]} f^{\mathbb{P}}(X_t^* | X_{t-1}^*). \quad \text{(a.1.50)}$$

Using eqs. (a.1.49) and (a.1.50) to substitute in the two conditional expectations appearing on the right-hand-side of eq. (a.1.48) leads to:

$$f^{\mathbb{Q}}(\mathcal{D}_t | X_t^*, \mathcal{I}_{t-1}) = \frac{\exp(\varphi_2 \mathcal{D}_t)\,(\mathcal{D}_t\{1 - \exp(-\underline{\lambda}_t)\} + (1 - \mathcal{D}_t)\{\exp(-\underline{\lambda}_t)\})}{\exp(\varphi_2)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t)},$$

which implies that:

$$\exp(-\underline{\lambda}_t^{\mathbb{Q}}) \equiv \mathbb{Q}(\mathcal{D}_t = 0 | \mathcal{D}_t = 0, X_t^*, \mathcal{I}_{t-1}) = \frac{\exp(-\underline{\lambda}_t)}{\exp(\varphi_2)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t)},$$

that is:

$$\underline{\lambda}_t^{\mathbb{Q}} = \underline{\lambda}_t + \log(\exp(\varphi_2)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t)). \quad \text{(a.1.51)}$$

If $\varphi_2 > 0$, we have $\exp(\varphi_2)\{1 - \exp(-\underline{\lambda}_t)\} + \exp(-\underline{\lambda}_t) > 1$, and therefore $\underline{\lambda}_t^{\mathbb{Q}} > \underline{\lambda}_t$.

1.J Approximating the fiscal limit

In this appendix, we explain how eq. (1.16) can be approximated. More precisely, we want to find an approximated representation of ℓ_t that is affine in w_t.

Because Δy_t and $\log(\mathcal{M}_{t,t+h}^n)$ are affine in w_t (up to i.i.d. shocks η_t, see eqs. 1.1 and a.1.25), and since the latter follows an affine process (1.2), it comes that there exist vectors and scalars, respectively denoted by a_h^ℓ and b_h^ℓ, that are such that:

$$\exp\left(a_h^{\ell\prime} w_t + b_h^\ell\right) = \mathbb{E}_t\left[\mathcal{M}_{t,t+h}^n \exp(\Delta y_{t+1} + \cdots + \Delta y_{t+h}) | \mathcal{D} \equiv 0\right]. \quad \text{(a.1.52)}$$

With these notations, eq. (1.16) becomes:

$$\exp(\ell_t) = \mu_{s^*} \sum_{h=1}^{+\infty} \exp\left(a_h^{\ell\prime} w_t + b_h^{\ell}\right), \tag{a.1.53}$$

We want to find a^{ℓ} and b^{ℓ} that are such that $\exp(\ell_t) \approx \exp(a^{\ell\prime} w_t + b^{\ell})$. This is done by solving the following system:

$$\left\{\begin{array}{rcl} \mathbb{E}(\exp(\ell_t)) & = & \mathbb{E}\left(\exp(a^{\ell\prime} w_t + b^{\ell})\right), \\ \mathbb{E}\left(\dfrac{\partial}{\partial w_{k,t}} \exp(\ell_t)\right) & = & \mathbb{E}\left(\dfrac{\partial}{\partial w_{k,t}} \exp(a^{\ell\prime} w_t + b^{\ell})\right), \quad k \in \{1, \dots, n_w\}. \end{array}\right. \tag{a.1.54}$$

Because the marginal distribution of w_t is Gaussian, we have:

$$\mathbb{E}\left(\exp(a^{\ell\prime} w_t + b^{\ell})\right) = \exp\left(\frac{1}{2} a^{\ell\prime} \Omega_w a^{\ell} + b^{\ell}\right), \tag{a.1.55}$$

$$\mathbb{E}\left(\frac{\partial}{\partial w_{k,t}} \exp(a^{\ell\prime} w_t + b^{\ell})\right) = a_k^{\ell} \exp\left(\frac{1}{2} a^{\ell\prime} \Omega_w a^{\ell} + b^{\ell}\right). \tag{a.1.56}$$

Moreover, using eq. (a.1.53), we have:

$$\begin{aligned} \mathbb{E}\left(\exp(\ell_t)\right) &= \mu_{s^*} \sum_{h=1}^{+\infty} \mathbb{E}\left(\exp\left(a_h^{\ell\prime} w_t + b_h^{\ell}\right)\right) \\ &= \mu_{s^*} \sum_{h=1}^{+\infty} \exp\left(\frac{1}{2} a_h^{\ell\prime} \Omega_w a_h^{\ell} + b_h^{\ell}\right), \end{aligned} \tag{a.1.57}$$

$$\mathbb{E}\left(\frac{\partial}{\partial w_{k,t}} \exp(\ell_t)\right) = \mu_{s^*} \sum_{h=1}^{+\infty} a_{1,h,k}^{\ell} \exp\left(\frac{1}{2} a_h^{\ell\prime} \Omega_w a_h^{\ell} + b_h^{\ell}\right). \tag{a.1.58}$$

System (a.1.54) implies that the result of the division of (a.1.56) by (a.1.55)—that is a_k^{ℓ}—should be equal to that of (a.1.58) by (a.1.57). Once the a_k^{ℓ}'s, are obtained (by a.1.58/a.1.57), we compute b^{ℓ} as follows:

$$b^{\ell} = \log \mathbb{E}\left(\exp(\ell_t)\right) - \frac{1}{2} a^{\ell\prime} \Omega_w a^{\ell},$$

where $\mathbb{E}\left(\exp(\ell_t)\right)$ is given by eq. (a.1.57).

Let us now explain how to compute the a_h^{ℓ}'s and b_h^{ℓ}'s, as defined in eq. a.1.52). For $h = 1$, we have:

$$\left\{\begin{array}{rcl} a_1^{\ell} & = & \Phi'(\Lambda_y - \gamma\Lambda_c) \\ b_1^{\ell} & = & \log(\delta) - \mu_c + \mu_y + \frac{1}{2}(\sigma_y - \gamma\sigma_c)'(\sigma_y - \gamma\sigma_c) \\ & & +\frac{1}{2}(\Lambda_y - \gamma\Lambda_c)'(\Lambda_y - \gamma\Lambda_c). \end{array}\right. \tag{a.1.59}$$

For $h > 0$, we have

$$\begin{array}{rl} & \exp\left(a_{h+1}^{\ell}{}' w_t + b_{h+1}^{\ell}\right) \\ = & \mathbb{E}_t\left[\mathcal{M}_{t,t+1}\exp\left(\Delta y_{t+1} + a_h^{\ell}{}' w_{t+1} + b_h^{\ell}\right)\right] \quad \text{(by the law of iterated expectations)} \\ = & \exp\left(\log(\delta) - \mu_c + \mu_y + \frac{1}{2}(\sigma_y - \gamma\sigma_c)'(\sigma_y - \gamma\sigma_c) + b_h^{\ell}\right)\mathbb{E}_t\left[\exp\left(\{\Lambda_y - \gamma\Lambda_c + a_h^{\ell}\}' w_{t+1}\right)\right]. \end{array}$$

Hence, for $h > 0$, we have:

$$\left\{\begin{array}{rcl} a_{h+1}^{\ell} & = & \Phi'(\Lambda_y - \gamma\Lambda_c + a_h^{\ell}) \\ b_{h+1}^{\ell} & = & \log(\delta) - \mu_c + \mu_y + \frac{1}{2}(\sigma_y - \gamma\sigma_c)'(\sigma_y - \gamma\sigma_c) + b_h^{\ell} \\ & & +\frac{1}{2}(\Lambda_y - \gamma\Lambda_c + a_h^{\ell})'(\Lambda_y + \Lambda_\pi + a_h^{\ell}), \end{array}\right. \tag{a.1.60}$$

with $a_0^{\ell} = 0$ and $b_0^{\ell} = 0$.

By iterating, we obtain, for $h \geq 1$:

$$\begin{array}{rcl} a_h^{\ell} & = & \Phi'(\Lambda_y - \gamma\Lambda_c + a_{h-1}^{\ell}) = \Phi'(\Lambda_y - \gamma\Lambda_c) + \Phi'^2(\Lambda_y - \gamma\Lambda_c + a_{h-2}^{\ell}) \\ & = & (I + \Phi' + \dots + \Phi'^{h-1})\Phi'(\Lambda_y - \gamma\Lambda_c) \\ & = & (I - \Phi'^h)\underbrace{(I - \Phi')^{-1}\Phi'(\Lambda_y - \gamma\Lambda_c)}_{=\kappa_0} = (I - \Phi'^h)\kappa_0. \end{array} \tag{a.1.61}$$

Moreover, for $h > 0$:

$$\begin{array}{rcl} b_h^{\ell} & = & \underbrace{\log(\delta) - \mu_c + \mu_y + \frac{1}{2}(\sigma_y - \gamma\sigma_c)'(\sigma_y - \gamma\sigma_c) + \frac{1}{2}(\Lambda_y - \gamma\Lambda_c)'(\Lambda_y - \gamma\Lambda_c)}_{=\kappa_1} \\ & & +b_{h-1}^{\ell} + a_{h-1}^{\ell}{}'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}a_{h-1}^{\ell}{}' a_{h-1}^{\ell} \\ & = & \left\{\kappa_1 + a_{h-1}^{\ell}{}'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}a_{h-1}^{\ell}{}' a_{h-1}^{\ell}\right\} + \dots + \left\{\kappa_1 + a_1^{\ell}{}'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}a_1^{\ell}{}' a_1^{\ell}\right\}. \end{array}$$

Using eq. (a.1.61) for $k > 0$, we obtain:

$$\begin{array}{rcl} b_h^{\ell} & = & h\kappa_1 + \sum_{k=1}^{h-1}\left(\kappa_0 - \Phi'^k\kappa_0\right)'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}\sum_{k=1}^{h-1}\left(\kappa_0 - \Phi'^k\kappa_0\right)'\left(\kappa_0 - \Phi'^k\kappa_0\right) \\ & = & h\kappa_1 + (h-1)\kappa_0'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}(h-1)\kappa_0'\kappa_0 - \kappa_0'\sum_{k=1}^{h-1}\Phi^k(\Lambda_y - \gamma\Lambda_c + \kappa_0) + \frac{1}{2}\kappa_0'\left(\sum_{k=1}^{h-1}\Phi^k\Phi'^k\right)\kappa_0 \\ & = & h\kappa_1 + (h-1)\kappa_0'(\Lambda_y - \gamma\Lambda_c) + \frac{1}{2}(h-1)\kappa_0'\kappa_0 - \kappa_0'(I - \Phi)^{-1}(\Phi - \Phi^h)(\Lambda_y - \gamma\Lambda_c + \kappa_0) + \\ & & +\frac{1}{2}\kappa_0'\left(\sum_{k=0}^{\infty}\Phi^k\Phi'^k - I - \Phi^h\left[\sum_{k=0}^{\infty}\Phi^k\Phi'^k\right]\Phi'^h\right)\kappa_0, \end{array}$$

with

$$vec\left(\sum_{k=0}^{\infty}\Phi^k\Phi'^k\right) = \left(I_{n_w^2} - \Phi\otimes\Phi\right)^{-1} vec(I_{n_w\times n_w}).$$

1.K VAR(1) dynamics of the state vector x_t

The dynamics of the state vector x_t (eq. 1.19) approximately is:

$$x_t = \begin{bmatrix} w_t \\ d_t \\ rd_t \\ q_t \\ w_{t-1} \\ d_{t-1} \\ \nu_t \end{bmatrix} \approx \mu_x + \Phi_x x_{t-1} + \Sigma_x \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix},$$

with

$$\mu_x = \begin{bmatrix} 0 \\ (-\mu_y - \mu_\pi + \log(1+\bar{q}-\overline{sd}) - \Psi(1-\chi)\bar{q} + \Psi\gamma_d\bar{d}) \\ -(1-\chi)\bar{q} \\ b_{h^*} \\ 0 \\ 0 \\ 0 \end{bmatrix},$$

$$\Phi_x = \begin{bmatrix} (\Phi) & 0 & 0 & 0 & 0 & 0 & 0 \\ -(\Lambda_y'\Phi + \Lambda_\pi'\Phi + \Psi\Lambda_s'\Phi) & (1-\Psi\gamma_d) & (\Psi\chi) & (\Psi(1-\chi)) & 0 & 0 & 0 \\ 0 & 0 & (\chi) & (1-\chi) & 0 & 0 & 0 \\ (a_{h^*}'\Phi) & 0 & 0 & 0 & 0 & 0 & 0 \\ I & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix},$$

and

$$\Sigma_x = \begin{bmatrix} I & 0 \\ (-(\Lambda_\pi + \Lambda_y + \Psi\Lambda_s)') & (-(\sigma_\pi + \sigma_y + \Psi\sigma_s)') \\ 0 & 0 \\ a_{h^*}' & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \end{bmatrix},$$

where $\Psi = 1/(1+\bar{q}-\overline{sd})$.

1.L Credit risk premiums, uncertainty, volatility and fiscal space

In this section, we report the output of fixed effects panel regressions of credit risk premiums arising from eq. 1.26 on uncertainty and volatility indices, and also on our fiscal space estimates. In Tables 1.L1 and 1.L2, we report the fixed effects panel regressions of, respectively, the 3- and 5-year credit risk premiums on their first lags, country-specific Economic Policy Uncertainty (*EPU*) indices, the Cboe VIX, the Cboe Emerging Markets Volatility Index (VXEEM) and the Cboe Short-term Volatility Index (VXSTV).

The estimation sample goes from 2004Q1 to 2021Q1. We observe that credit risk premiums positively comove with EPU and the short-term volatility index, while they negatively relate to the VIX. Moreover, in Tables 1.L3 and 1.L4, we report the fixed effects panel regressions of, respectively, the 3- and 5-year credit risk premiums on their first lags and our fiscal space estimates. We observe a negative relationship between fiscal space and credit risk premiums, which highlights that a worsening of the fiscal position corresponds to higher risk premiums demanded by risk-averse investors. In all tables discussed above, we report in Panel B and C the same type of regressions only for a subset of countries including advanced and emerging economies, respectively.

Table 1.L1: Credit risk premium (3-year maturity), uncertainty and volatility indices - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.689***	0.695***	0.915***
	(0.060)	(0.071)	(0.031)
EPU_t	0.002***	0.002***	0.0003
	(0.001)	(0.001)	(0.0004)
VIX_t		−0.148***	
		(0.043)	
$VXEEM_t$		0.015	
		(0.016)	
$VXSTV_t$		0.103***	
		(0.037)	
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.525***	0.698***	0.716***
	(0.066)	(0.073)	(0.052)
EPU_t	0.0005	0.001*	−0.0003
	(0.0003)	(0.0004)	(0.0002)
VIX_t		−0.057**	
		(0.026)	
$VXEEM_t$		0.010	
		(0.010)	
$VXSTV_t$		0.032	
		(0.022)	
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.688***	0.694***	0.894***
	(0.065)	(0.076)	(0.037)
EPU_t	0.002*	0.003**	0.001
	(0.001)	(0.001)	(0.001)
VIX_t		−0.229***	
		(0.079)	
$VXEEM_t$		0.025	
		(0.028)	
$VXSTV_t$		0.161**	
		(0.067)	

Note: This table reports the results of panel regressions of the 3-year maturity credit risk premium (*CRP*) estimates on country-specific Economic Policy Uncertainty (*EPU*) indices, the Cboe VIX, the Cboe Emerging Markets Volatility Index (VXEEM) and the Cboe Short-term Volatility Index (VXSTV). The estimation sample goes from 2004$Q1$ to 2021$Q1$. See text for more details.. FE stands for Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

Table 1.L2: Credit risk premium (5-year maturity), uncertainty and volatility indices - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.736***	0.751***	0.956***
	(0.055)	(0.067)	(0.022)
EPU_t	0.002***	0.002***	0.001
	(0.001)	(0.001)	(0.0005)
VIX_t		−0.214***	
		(0.057)	
$VXEEM_t$		0.014	
		(0.022)	
$VXSTV_t$		0.152***	
		(0.048)	
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.652***	0.784***	0.843***
	(0.061)	(0.065)	(0.042)
EPU_t	0.001*	0.001*	−0.0004
	(0.0004)	(0.001)	(0.0003)
VIX_t		−0.095**	
		(0.038)	
$VXEEM_t$		0.013	
		(0.014)	
$VXSTV_t$		0.056*	
		(0.032)	
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.740***	0.751***	0.936***
	(0.060)	(0.076)	(0.028)
EPU_t	0.003**	0.003**	0.001
	(0.001)	(0.001)	(0.001)
VIX_t		−0.321***	
		(0.103)	
$VXEEM_t$		0.017	
		(0.039)	
$VXSTV_t$		0.234***	
		(0.087)	

Note: This table reports the results of panel regressions of the 5-year credit risk premium (*CRP*) estimates on country-specific Economic Policy Uncertainty (*EPU*) indices, the Cboe VIX, the Cboe Emerging Markets Volatility Index (VXEEM) and the Cboe Short-term Volatility Index (VXSTV). The estimation sample goes from 2004$Q1$ to 2021$Q1$. See text for more details. FE stands. for Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

Table 1.L3: Credit risk premium (3-year maturity) and fiscal space - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.558***	0.542***	0.813***
	(0.052)	(0.056)	(0.031)
FS_t	−0.025***	−0.038***	−0.006***
	(0.004)	(0.004)	(0.002)
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.500***	0.496***	0.732***
	(0.057)	(0.052)	(0.047)
FS_t	−0.021***	−0.028***	−0.001
	(0.003)	(0.003)	(0.002)
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.523***	0.535***	0.760***
	(0.054)	(0.060)	(0.043)
FS_t	−0.044***	−0.053***	−0.007***
	(0.008)	(0.007)	(0.002)

Note: This table reports the results of panel regressions of the 3-year maturity credit risk premium (CRP) on fiscal space (FS) estimates. The estimation sample goes from 2004Q1 to 2021Q1. See text for more details. FE stands for. Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

Table 1.L4: Credit risk premium (5-year maturity) and fiscal space - Panel regression results

	Panel A - *All countries*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.651^{***}	0.635^{***}	0.906^{***}
	(0.048)	(0.052)	(0.021)
FS_t	-0.027^{***}	-0.049^{***}	-0.003
	(0.006)	(0.005)	(0.002)
	Panel B - *Advanced Economies (US, UK, JP, CA)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.591^{***}	0.602^{***}	0.831^{***}
	(0.052)	(0.049)	(0.041)
FS_t	-0.030^{***}	-0.038^{***}	0.001
	(0.006)	(0.005)	(0.003)
	Panel C - *Emerging Economies (BR, CN, IN, RU)*		
	Country & Time FE	Country FE	Time FE
	CRP_t	CRP_t	CRP_t
CRP_{t-1}	0.633^{***}	0.634^{***}	0.877^{***}
	(0.050)	(0.057)	(0.028)
FS_t	-0.043^{***}	-0.062^{***}	-0.004
	(0.010)	(0.010)	(0.003)

Note: This table reports the results of panel regressions of the 5-year maturity credit risk premium (CRP) on fiscal space (FS) estimates. The estimation sample goes from 2004Q1 to 2021Q1. See text for more details. FE stands for. Fixed Effects. *p<0.1; **p<0.05; ***p<0.01

1.M Data sources

CDS spreads and bond yields are extracted from CMA and Refinitiv Eikon Datastream. For the US, the original source of the yields is the Federal Reserve. For the other countries, we take zero-coupon bond yields bootstrapped by Refinitiv Eikon Datastream from government bond prices. Inflation is based on the GDP price deflator.

For the US, macroeconomic variables are drawn from the FRED database (Federal Reserve of St. Louis) and from the Bureau of Economic Analysis.

For the UK, GDP, consumption and gross government debt are collected from the British Office for National Statistics. The series for gross and net government interest payments and primary surplus/deficit are gathered from the OECD Economic Outlook; the same holds true for Japan. For the latter country, GDP and consumption variables are acquired from the Cabinet Office database (Government of Japan). Gross government debt for Japan is drawn from the Bank of Japan.

As regards Canada, all relevant series for model estimation are fetched from the respective official national statistics bureaus.

For Brazil, macroeconomic variables are drawn from OECD databases (Quarterly National Accounts and Main Economic Indicators). Public finance statistics (public debt and primary deficit/surplus) are collected from Banco Central do Brasil.

As regards China, we draw GDP at constant prices and the implicit price deflator from the National Bureau of Statistics of China; while, GDP at current prices is drawn from OECD Quarterly National Accounts. Private consumption, primary balance and interest payments are extracted from Oxford Economics. Public debt is drawn from the IMF Fiscal Monitor.

For India, we draw macroeconomic variables from OECD Quarterly National Accounts. Public debt is fetched from the Ministry of Finance (Government of India). Other public finance statistics are gathered from the Controller General of Accounts.

As regards Russia, GDP series are retrieved from the Federal State Statistics Service. Public debt, interest payments and primary surplus/deficit series are taken from the Ministry of Finance of the Russian Federation. Consumption expenditure and the implicit GDP deflator are collected from OECD national database.

Table 1.M1: Data Panel: United States of America

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. of Obs.**
Senior CDS	1 Year	CMA	2008Q1-2021Q2	54
	2 Years	CMA	2008Q1-2021Q2	54
	3 Years	CMA	2008Q1-2021Q2	54
	5 Years	CMA	2008Q1-2021Q2	54
	10 Years	CMA	2008Q1-2021Q2	54
Yields	1 Year	Federal Reserve, US	2008Q1-2021Q2	54
	2 Years	Federal Reserve, US	2008Q1-2021Q2	54
	3 Years	Federal Reserve, US	2008Q1-2021Q2	54
	5 Years	Federal Reserve, US	2008Q1-2021Q2	54
	10 Years	Federal Reserve, US	2008Q1-2021Q2	54
GDP, market constant prices (CHND 2012)	-	BEA - US Dep. of Commerce[a]	2008Q1-2021Q2	54
GDP, market current prices	-	BEA - US Dep. of Commerce	2008Q1-2021Q2	54
Final Consumption Expenditure, Services	-	FRED (ST. LOUIS FED)	2008Q1-2021Q2	54
Final Consumption Expenditure, Non-Durables	-	FRED (ST. LOUIS FED)	2008Q1-2021Q2	54
GDP Implicit Price Deflator (Index 2012=100)	-	BEA - US Dep. of Commerce	2008Q1-2021Q2	54
Gross Federal Government Debt, Current Prices	-	FRED (ST. LOUIS FED)	2008Q1-2021Q2	54
Net Government Interest Payments, Current Prices	-	FRED (ST. LOUIS FED)	2008Q1-2021Q2	54
Government Primary Surplus/Deficit, Current Prices	-	FRED (ST. LOUIS FED)	2008Q1-2021Q2	54

Note: [a] Bureau of Economic Analysis.

Table 1.M2: Data Panel: United Kingdom

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. of Obs.**
Senior CDS	1 Year	CMA	2008Q1-2021Q1	53
	2 Years	CMA	2008Q1-2021Q1	53
	3 Years	CMA	2008Q1-2021Q1	53
	5 Years	CMA	2008Q1-2021Q1	53
	10 Years	CMA	2008Q1-2021Q1	53
Yields	1 Year	Refinitiv Eikon Datastream	2008Q1-2021Q1	53
	2 Years	Refinitiv Eikon Datastream	2008Q1-2021Q1	53
	3 Years	Refinitiv Eikon Datastream	2008Q1-2021Q1	53
	5 Years	Refinitiv Eikon Datastream	2008Q1-2021Q1	53
	10 Years	Refinitiv Eikon Datastream	2008Q1-2021Q1	53
GDP, market constant prices (CHND 2016)	-	Office for National Statistics	2008Q1-2021Q1	53
GDP, market current prices	-	Office for National Statistics	2008Q1-2021Q1	53
Final Consumption Expenditure, Services	-	Office for National Statistics	2008Q1-2021Q1	53
Final Consumption Expenditure, Non-Durables	-	Office for National Statistics	2008Q1-2021Q1	53
GDP Implicit Price Deflator (Index 2016=100)	-	Office for National Statistics	2008Q1-2021Q1	53
Gross Government Debt, Total, Current Prices	-	Office for National Statistics	2008Q1-2021Q1	53
Gross Government Interest Payments, Current Prices	-	OECD Economic Outlook	2008Q1-2021Q1	53
Net Government Interest Payments, Current Prices	-	OECD Economic Outlook	2008Q1-2021Q1	53
Government Primary Surplus/Deficit, Current Prices	-	OECD Economic Outlook	2008Q1-2021Q1	53

Table 1.M3: Data Panel: Japan

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. degree of Obs.**
Senior CDS	1 Year	CMA	2004Q1-2021Q2	70
	2 Years	CMA	2004Q1-2021Q2	70
	3 Years	CMA	2004Q1-2021Q2	70
	5 Years	CMA	2004Q1-2021Q2	70
	10 Years	CMA	2004Q1-2021Q2	70
Yields	1 Year	Refinitiv Eikon Datastream	2004Q1-2021Q2	70
	2 Years	Refinitiv Eikon Datastream	2004Q1-2021Q2	70
	3 Years	Refinitiv Eikon Datastream	2004Q1-2021Q2	70
	5 Years	Refinitiv Eikon Datastream	2004Q1-2021Q2	70
	10 Years	Refinitiv Eikon Datastream	2004Q1-2021Q2	70
GDP, market constant prices (CHND 2011)	-	Cabinet Office (Gov. of Japan)	2004Q1-2021Q2	70
GDP, market current prices	-	Cabinet Office (Gov. of Japan)	2004Q1-2021Q2	70
Final Consumption Expenditure, Services	-	Cabinet Office (Gov. of Japan)	2004Q1-2021Q2	70
Final Consumption Expenditure, Non-Durables	-	Cabinet Office (Gov. of Japan)	2004Q1-2021Q2	70
GDP Implicit Price Deflator (Index 2011=100)	-	Thomson Reuters	2004Q1-2021Q2	70
Gross Government Debt, Total, Current Prices	-	Bank of Japan	2004Q1-2021Q2	70
Gross Government Interest Payments, Current Prices	-	OECD Economic Outlook	2004Q1-2021Q2	70
Net Government Interest Payments, Current Prices	-	OECD Economic Outlook	2004Q1-2021Q2	70
Government Primary Surplus/Deficit, Current Prices	-	OECD Economic Outlook	2004Q1-2021Q2	70

Table 1.M4: Data Panel: Canada

Variable	Horizon / Maturity	Source	Period	N. of Obs.
Senior CDS	1 Year	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	2 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	3 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	5 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	10 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
Yields	1 Year	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	2 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	3 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	5 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
	10 Years	Refinitiv Eikon Datastream	2012Q4-2020Q4	33
GDP, market constant prices (CHND 2012)	-	CANSIM[a]	2012Q4-2020Q4	33
GDP, market current prices	-	CANSIM	2012Q4-2020Q4	33
Final Consumption Expenditure, Services	-	CANSIM	2012Q4-2020Q4	33
Final Consumption Expenditure, Non-Durables	-	CANSIM	2012Q4-2020Q4	33
GDP Implicit Price Deflator (Index 2012=100)	-	CANSIM	2012Q4-2020Q4	33
Gross Central Government Debt, Total, Current Prices	-	CANSIM	2012Q4-2020Q4	33
Gross Government Interest Payments, Current Prices	-	CANSIM	2012Q4-2020Q4	33
Government Primary Surplus/Deficit, Current Prices	-	CANSIM	2012Q4-2020Q4	33

Note: [a] Canadian Socio-Economic Information Management System (Statistics Canada).

Table 1.M5: Data Panel: Brazil

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. of Obs.**
Senior CDS	1 Year	CMA	2006Q3-2021Q2	60
	2 Years	CMA	2006Q3-2021Q2	60
	3 Years	CMA	2006Q3-2021Q2	60
	5 Years	CMA	2006Q3-2021Q2	60
	10 Years	CMA	2006Q3-2021Q2	60
Yields	1 Year	Refinitiv Eikon Datastream	2006Q3-2021Q2	60
	2 Years	Refinitiv Eikon Datastream	2006Q3-2021Q2	60
	3 Years	Refinitiv Eikon Datastream	2006Q3-2021Q2	60
	5 Years	Refinitiv Eikon Datastream	2006Q3-2021Q2	60
	10 Years	Refinitiv Eikon Datastream	2006Q3-2021Q2	60
GDP, market constant prices (1995 prices)	-	Quarterly National Accounts, OECD	2006Q3-2021Q2	60
GDP, market current prices	-	Quarterly National Accounts, OECD	2006Q3-2021Q2	60
Final Consumption Expenditure	-	Quarterly National Accounts, OECD	2006Q3-2021Q2	60
GDP Implicit Price Deflator (Index 2015=100)	-	OECD Main Economic Indicators	2006Q3-2021Q2	60
Public Debt, General Government, Gross, Domestic, Current Prices	-	Banco Central do Brasil	2006Q3-2021Q2	60
Public Debt, General Government, Gross, External, Current Prices	-	Banco Central do Brasil	2006Q3-2021Q2	60
Primary Deficit/Surplus, Consolidated Public Sector, Current Prices	-	Banco Central do Brasil	2006Q3-2021Q2	60

Table 1.M6: Data Panel: China

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. of Obs.**
Senior CDS	1 Year	CMA	2004Q2-2021Q2	69
	2 Years	CMA	2004Q2-2021Q2	69
	3 Years	CMA	2004Q2-2021Q2	69
	5 Years	CMA	2004Q2-2021Q2	69
	10 Years	CMA	2004Q2-2021Q2	69
Yields	1 Year	Refinitiv Eikon Datastream	2004Q2-2021Q2	69
	2 Years	Refinitiv Eikon Datastream	2004Q2-2021Q2	69
	3 Years	Refinitiv Eikon Datastream	2004Q2-2021Q2	69
	5 Years	Refinitiv Eikon Datastream	2004Q2-2021Q2	69
	10 Years	Refinitiv Eikon Datastream	2004Q2-2021Q2	69
GDP, market constant prices (2020 prices)	-	National Bureau of Statistics of China/Refinitiv	2004Q2-2021Q2	69
GDP, market current prices	-	Quarterly National Accounts, OECD	2004Q2-2021Q2	69
Private Consumption	-	Oxford Economics	2004Q2-2021Q2	69
GDP Implicit Price Deflator (Index 2010=100)	-	National Bureau of Statistics of China/Refinitiv	2004Q2-2021Q2	69
Public Debt, Current Prices	-	IMF - Fiscal Monitor	2004Q2-2021Q2	69
Gross Government Interest Payments, Current Prices	-	Oxford Economics	2004Q2-2021Q2	69
Primary Balance, Gross, Current Prices	-	Oxford Economics	2004Q2-2021Q2	69

Table 1.M7: Data Panel: India

Variable	Horizon / Maturity	Source	Period	N. of Obs.
State Bank of India CDS	1 Year	CMA	2004Q4-2013Q3	36
	2 Years	CMA	2004Q4-2013Q3	36
	3 Years	CMA	2004Q4-2013Q3	36
	5 Years	CMA	2004Q4-2013Q3	36
	10 Years	CMA	2004Q4-2013Q3	36
Senior CDS	1 Year	CMA	2013Q4-2021Q1	30
	2 Years	CMA	2013Q4-2021Q1	30
	3 Years	CMA	2013Q4-2021Q1	30
	5 Years	CMA	2013Q4-2021Q1	30
	10 Years	CMA	2013Q4-2021Q1	30
Yields	1 Year	Refinitiv Eikon Datastream	2004Q4-2021Q1	66
	2 Years	Refinitiv Eikon Datastream	2004Q4-2021Q1	66
	3 Years	Refinitiv Eikon Datastream	2004Q4-2021Q1	66
	5 Years	Refinitiv Eikon Datastream	2004Q4-2021Q1	66
	10 Years	Refinitiv Eikon Datastream	2004Q4-2021Q1	66
GDP, market constant prices (2011-2012 prices)	-	Quarterly National Accounts, OECD	2004Q4-2021Q1	66
GDP, market current prices	-	Quarterly National Accounts, OECD	2004Q4-2021Q1	66
Private Final Consumption Expenditure	-	Quarterly National Accounts, OECD	2004Q4-2021Q1	66
GDP Implicit Price Deflator (Index 2015=100)	-	Quarterly National Accounts, OECD	2004Q4-2021Q1	66
Central Government Debt, Overall, Current Prices	-	Ministry of Finance, Government of India	2004Q4-2021Q1	66
Central Government, Interest Payments, Current Prices	-	Controller General of Accounts, India	2004Q4-2021Q1	66
Central Government, Deficit/Surplus, Primary	-	Controller General of Accounts, India	2004Q4-2021Q1	66

Table 1.M8: Data Panel: Russia

Variable	**Horizon / Maturity**	**Source**	**Period**	**N. of Obs.**
Senior CDS	1 Year	CMA	2005Q3-2020Q4	62
	2 Years	CMA	2005Q3-2020Q4	62
	3 Years	CMA	2005Q3-2020Q4	62
	5 Years	CMA	2005Q3-2020Q4	62
	10 Years	CMA	2005Q3-2020Q4	62
Yields	1 Year	Refinitiv Eikon Datastream	2005Q3-2020Q4	62
	2 Years	Refinitiv Eikon Datastream	2005Q3-2020Q4	62
	3 Years	Refinitiv Eikon Datastream	2005Q3-2020Q4	62
	5 Years	Refinitiv Eikon Datastream	2005Q3-2020Q4	62
	10 Years	Refinitiv Eikon Datastream	2005Q3-2020Q4	62
GDP, market constant prices (CHND 2016)	-	Federal State Statistics Service	2005Q3-2020Q4	62
GDP, market current prices	-	IMF-IFS[a]	2005Q3-2020Q4	62
Final Consumption Expenditure	-	Quarterly National Accounts, OECD	2005Q3-2020Q4	62
GDP Implicit Price Deflator (Index 2015=100)	-	OECD Main Economic Indicators	2005Q3-2020Q4	62
Public Debt, Total, Current Prices	-	World Bank QPSD[b]	2005Q3-2020Q4	62
Gross Government Interest Payments, Current Prices	-	Ministry of Finance of the Russian Federation	2005Q3-2020Q4	62
Government Primary Surplus/Deficit, Current Prices	-	Ministry of Finance of the Russian Federation	2005Q3-2020Q4	62

Note: [a] International Monetary Fund - International Financial Statistics; [b] World Bank Quarterly Public Sector Debt.

1.N Additional tables and figures

Table 1.N1: 2-year CDS sensitivity to deficits

Panel A - Initial state for simulations = Last estimation period

Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	0.6	[0.6]	3.7	[0.9]	9.1	[1.3]
Canada	2.7	[2.7]	15.4	[3.4]	35.6	[4.5]
United Kingdom	1.8	[1.8]	13.2	[3.6]	41.6	[7.4]
Japan	1.7	[1.7]	9.9	[2.3]	23.6	[3.1]
Brazil	12.4	[12.4]	67.6	[14.7]	149.7	[17.6]
Russia	17.0	[17.0]	234.4	[83.1]	917.1	[162.4]
India	23.0	[23.0]	167.9	[44.6]	464.8	[67.7]
China	6.5	[6.5]	46.2	[12.3]	136.0	[22.0]

Panel B - Lower initial debt-to-GDP ratio (-10 p.p. of GDP w.r.t. Panel A)

Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	0.3	[0.3]	1.8	[0.4]	4.6	[0.7]
Canada	1.5	[1.5]	8.7	[2.0]	20.7	[2.7]
United Kingdom	0.3	[0.3]	2.1	[0.6]	7.6	[1.5]
Japan	1.1	[1.1]	6.4	[1.5]	15.6	[2.1]
Brazil	7.9	[7.9]	44.4	[9.9]	101.9	[12.6]
Russia	0.0	[0.0]	0.4	[0.2]	23.3	[11.4]
India	2.4	[2.4]	27.9	[9.6]	132.6	[30.1]
China	2.2	[2.2]	18.0	[5.2]	62.7	[11.8]

Panel C - Larger initial debt-to-GDP ratio ($+20$ p.p. of GDP w.r.t. Panel A)

Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	1.9	[1.9]	11.0	[2.5]	26.1	[3.4]
Canada	6.3	[6.3]	33.9	[7.3]	74.7	[8.7]
United Kingdom	19.2	[19.2]	115.8	[27.2]	284.1	[38.0]
Japan	3.5	[3.5]	19.6	[4.3]	44.4	[5.4]
Brazil	20.8	[20.8]	109.0	[22.8]	229.8	[25.0]
Russia	173.6	[173.6]	897.7	[182.9]	1.8×10^3	[0.2×10^3]
India	76.6	[76.6]	396.1	[81.1]	806.5	[82.4]
China	22.5	[22.5]	136.1	[31.9]	327.6	[42.2]

Note: This table documents the sensitivity of the 2-year CDS spreads to fiscal conditions. We consider three sizes of fiscal shocks (increases in primary deficits by 1%, 5% and 10% of GDP). The shocks are spread on four quarters and we compare the levels of the CDS spreads, after these four quarters, with those obtained in a benchmark scenario (with no fiscal shock). See text for more details regarding the construction of scenarios. The reported figures are in basis points. The number in square brackets correspond to the marginal influence of an additional unit increase in the deficit.

Table 1.N2: 5-year CDS sensitivity to deficits

Panel A - Initial state for simulations = Last estimation period

Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	0.8	[0.8]	4.5	[1.0]	10.7	[1.4]
Canada	2.8	[2.8]	15.5	[3.4]	35.3	[4.3]
United Kingdom	2.6	[2.6]	16.7	[4.2]	46.8	[7.4]
Japan	2.2	[2.2]	12.9	[3.0]	30.9	[4.1]
Brazil	15.2	[15.2]	82.4	[17.7]	180.4	[20.9]
Russia	50.8	[50.8]	441.2	[128.1]	1.3×10^3	[0.2×10^3]
India	30.1	[30.1]	209.8	[54.1]	561.8	[79.7]
China	20.0	[20.0]	134.7	[34.2]	360.4	[52.0]

Panel B - Lower initial debt-to-GDP ratio (-10 p.p. of GDP w.r.t. Panel A)

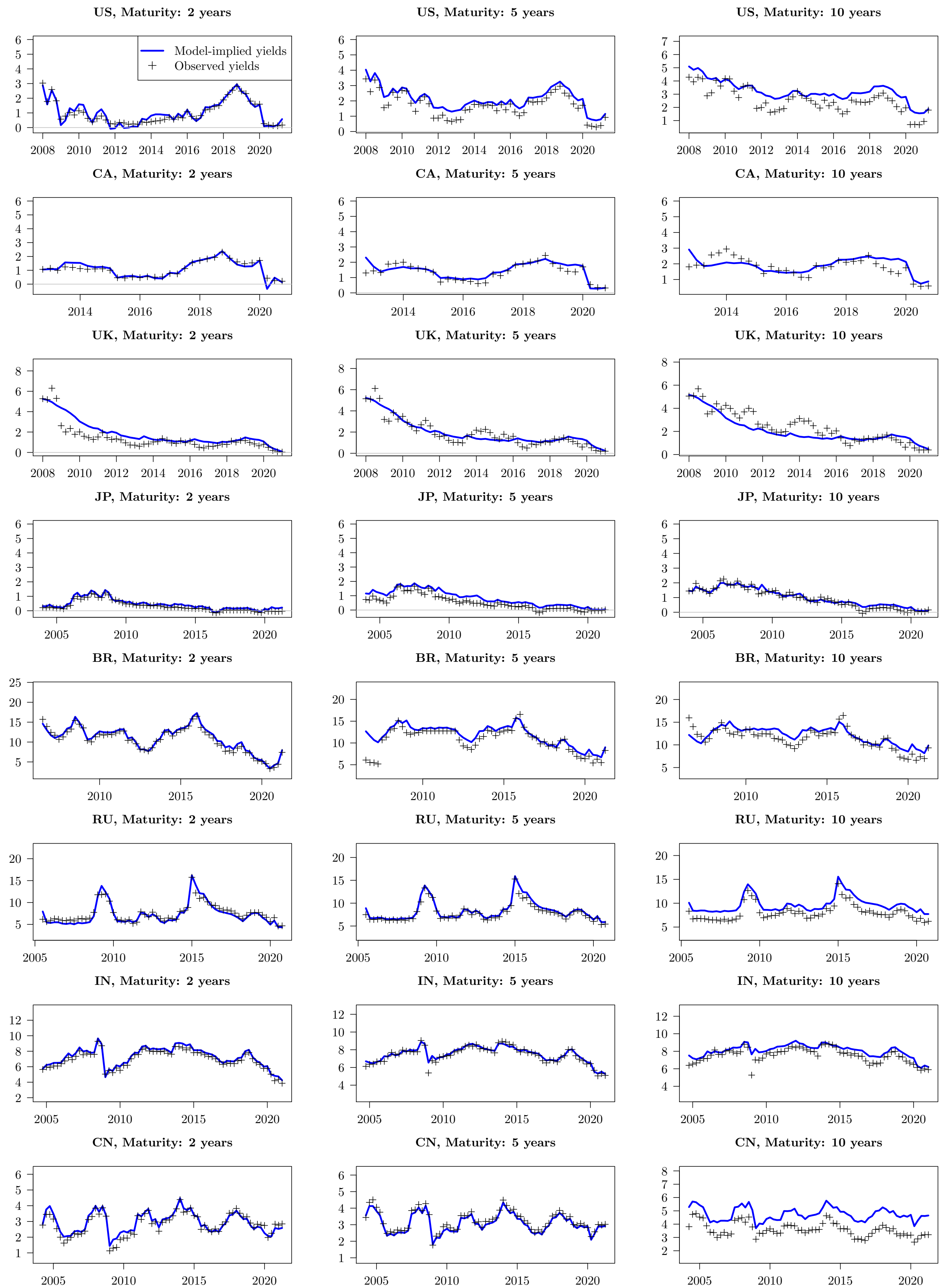
Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	0.4	[0.4]	2.6	[0.6]	6.3	[0.8]
Canada	1.7	[1.7]	9.4	[2.1]	22.0	[2.8]
United Kingdom	0.7	[0.7]	4.6	[1.2]	13.6	[2.3]
Japan	1.5	[1.5]	9.0	[2.1]	22.2	[3.0]
Brazil	10.8	[10.8]	59.7	[13.1]	134.0	[16.1]
Russia	0.9	[0.9]	18.3	[7.9]	163.4	[50.4]
India	6.2	[6.2]	56.7	[17.6]	217.0	[43.0]
China	10.5	[10.5]	78.3	[21.3]	236.8	[38.8]

Panel C - Larger initial debt-to-GDP ratio ($+20$ p.p. of GDP w.r.t. Panel A)

Fiscal shock:	+1 p.p. of GDP		+5 p.p. of GDP		+10 p.p. of GDP	
United States	1.9	[1.9]	10.8	[2.4]	25.0	[3.2]
Canada	5.8	[5.8]	31.5	[6.8]	69.1	[8.0]
United Kingdom	16.3	[16.3]	95.9	[22.2]	231.5	[30.5]
Japan	4.1	[4.1]	22.9	[5.1]	52.5	[6.5]
Brazil	23.5	[23.5]	123.2	[25.8]	260.1	[28.4]
Russia	202.2	[202.2]	1.0×10^3	[0.2×10^3]	2.1×10^3	[0.2×10^3]
India	85.9	[85.9]	452.5	[94.2]	935.6	[97.6]
China	43.0	[43.0]	252.8	[57.7]	581.5	[70.0]

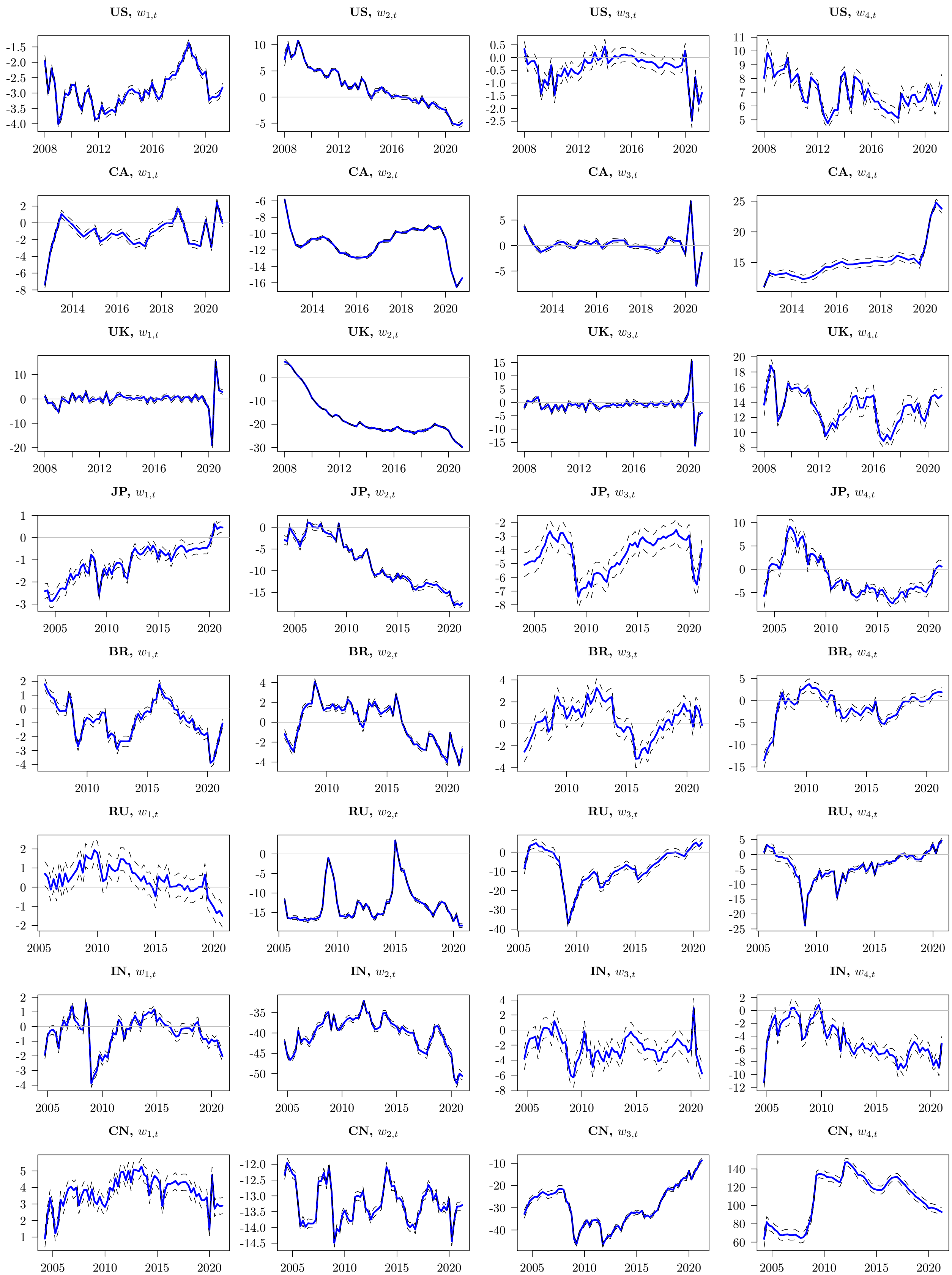
Note: This table documents the sensitivity of the 5-year CDS spreads to fiscal conditions. We consider three sizes of fiscal shocks (increases in primary deficits by 1%, 5% and 10% of GDP). The shocks are spread on four quarters and we compare the levels of the CDS spreads, after these four quarters, with those obtained in a benchmark scenario (with no fiscal shock). See text for more details regarding the construction of scenarios. The reported figures are in basis points. The number in square brackets correspond to the marginal influence of an additional unit increase in the deficit.

Figure 1.N1: Observed vs model-implied yields



Note: This figure compares model-implied and observed quarterly yields of zero-coupon government yields. The computation of model-implied yields is based on eq. (a.1.7) of Appendix 1.B.1 (the maturity-h yield is given by $-\frac{1}{h}\log B_{t,h}$, where $B_{t,h}$ is the date-t price of a zero-coupon bond of maturity h).

Figure 1.N2: Estimated factors



Note: This figure displays filtered factors $w_{i,t}$, $i = 1, \ldots, 4$, for each country. These estimates result from the Extended Kalman Filter (see Subsection 1.4.3). The dashed lines delineate the 95% confidence interval (accounting for filtering uncertainty).

Chapter 2

Fiscal limits and the Pricing of Eurobonds[1]

2.1 Introduction

Following the last financial crisis and the COVID-19 pandemic, sovereign debts across the euro area have risen to levels unprecedented since the Second World War. In this context, the sustainability of fiscal positions—especially in the peripheral Member States—has been called into question. Against this backdrop, numerous academics, policymakers, and analysts have discussed proposals for issuing common bonds—often referred to as Eurobonds. The rationale behind such common bonds is most often, and more or less explicitly, a debt service relief for peripheral member states (Beetsma and Mavromatis, 2014; Favero and Missale, 2012). An ulterior motive backing common issuances is to ensure financial stability, notably by addressing the demand of financial institutions for safe assets (Brunnermeier et al., 2017).[2] Moreover, if issued on a large scale, a joint debt instrument is advocated as a useful device to increase bond market liquidity in the euro area (Hellwig and Philippon, 2011).

Surprisingly, the different proposals for common debt issuance seldom come with pricing attempts.[3] Arguably, this shortage of quantitative analysis may have contributed to the lack of support for common bond issuances. This paper offers a way to explore the pricing of joint sovereign debt instruments.

[1]This chapter is coauthored with Jean-Paul Renne from the University of Lausanne.

[2]Although Eurobonds may constitute a way to guide the euro area towards financial stability, the objectives of Eurobond proposals do not fully overlap with those of the European Financial Stability Facility's (EFSF) and the European Stability Mechanism's (ESM) programs. Typically, the objective of the ESM is to provide financial assistance to euro-area countries experiencing, or threatened by, severe financing problems; this would complement joint issuance in times of financial distress, but goes beyond the preventive intention of a common euro-area bond.

[3]The evaluation of price effects remains merely speculative in this literature (Claessens, Mody, and Vallée, 2012, end of Section IV.B).

Guarantees play a significant role in the pricing of joint debt instruments. Our analysis focuses on two polar cases: (a) the case of several and joint guarantees (SJG) whereby all countries are jointly liable for each other's default through the common debt instrument, and (b) the case of several but not joint guarantees (SNJG) whereby each debtor is responsible only for a percentage contribution to each redemption. In the former case, participating European countries are responsible not only for their own percentage contribution to the bond, but also for covering any other state's unpaid contributions. In the latter case (SNJG bonds), each participant is liable only for the debt service and principal redemption corresponding to its share of the bond. In both cases, the joint debt instrument would trade as a single bond; it could be issued by an independent debt agency, with funds raised, and obligations divided between participating issuers in fixed shares (see, e.g., De Grauwe and Moesen, 2009; European Commission, 2011; Delivorias and Stamegna, 2020).

In the present paper, we propose a multi-country credit-risk model where both standard and common sovereign bonds—featuring one of the two polar types of guarantees discussed above—can be priced. The model estimation relies on national bond prices; the sample covers the period from 2008Q2 to 2021Q2.[4] We focus on the six largest euro-area economies: Germany, France, Italy, Spain, Netherlands, and Belgium (which account for almost 90% of the total Eurozone GDP). Once the model is estimated, we compute counterfactual Eurobond prices over the same period.

In the model, the probability of default depends on the considered entity's fiscal space, which can be a single country or a group of countries. The fiscal space is the distance between public debt and the so-called fiscal limit; this limit, in turn, represents the maximum outstanding debt that can credibly be covered by future primary budget surpluses (Bi, 2012; Bi and Leeper, 2013). The probability of default gets strictly positive only if public debt breaks the fiscal limit, that is, if the fiscal space is negative. In this framework, a natural way to conceive a SJG bond is to consider that it is issued by an entity for which both underlying debtors' fiscal revenues and debts are pooled. By contrast, a SNJG bond is equivalent to a combination of national bonds weighted by their participation share in the debt instrument.

Estimating the model involves the estimation of both the model parameters and the time series of national fiscal limits. These two tasks are jointly carried out by resorting to an adaptation of the so-called "inversion technique" *à la* Chen and Scott (1993). For a given model

[4]Some bonds issued by European institutions can be seen as proxies for Eurobonds (see end of Subsection 2.6.1, where we compare our model-implied SJG prices with the latter). However, for the time being, there are not enough of these bonds to determine constant-maturity interest rates on a sufficiently long sample.

parametrization, formulas for the sovereign bond yield spreads are inverted to get fiscal limit estimates.[5] The maximum likelihood function can then be computed; it is maximized to yield the estimated model parametrization.

Our model features a good fit of the observed term-structure of bond yield spreads across all countries; this fit is comparable to the one obtained in term-structure studies where default intensities are purely latent and have no macro-finance interpretation. We also obtain sizeable estimates of sovereign credit risk premiums, defined as those components of sovereign spreads that would not exist if agents were risk-neutral. Moreover, to the best of our knowledge, this paper is the first to provide time-varying estimates of fiscal limits for euro-area countries.[6]

Our counterfactual analysis results highlight the importance of guarantees on Eurobond pricing. By design, yield spreads associated with Eurobonds featuring several but not joint guarantees (SNJG) are close to the (issuance-weighted) average of country-specific spreads. By contrast, common bonds with several and joint guarantees (SJG) benefit from fiscal diversification effects resulting in a sizeable credit spread compression: across the estimation sample and different maturities, the SNJG bond yield spread was about three times larger than the SJG one. However, the model also predicts that SJG advantages diminish when expected fiscal spaces reduce at the euro-area scale, up to potential inversion; this turned out to be the case for two quarters in our sample—2011Q4 and 2012Q1, the peak of the euro-area debt crisis—and for the longer maturity only. Hence, except for strongly adverse fiscal conditions, raising funds through a joint liability debt instrument—the SJG bond—may reduce *aggregate* debt service in the presence of heterogenous fiscal conditions. Interestingly, for shorter maturities, the yield spread associated with the SJG bond is, at times, lower than the German bond one. (The German bonds, called Bunds, are considered the safest bond in the euro area.) Even when this is not the case, i.e. when SJG yields are higher than those of the bonds issued by the best-rated countries, one can envision post-issuance redistribution schemes under which all countries eventually benefit from common issuances. One such scheme is to distribute the overall gains in such a way as to achieve a reduction in "post-redistribution yields" that is the same in all countries

[5] We posit reduced-form dynamics for national debts and fiscal limits and derive resulting bond pricing. Our approach shares some similarities with the Black-Scholes-Merton model (Black and Scholes, 1973; Merton, 1974, and its numerous extensions) in that it also features a default threshold. As noted by Duffie and Singleton (2003, Subsection 3.2.2), the tractability of the Black-Scholes-Merton model rapidly declines as one allows for a time-varying default threshold. Although our framework features a time-varying debt threshold, tractability is preserved thanks to approximation formulas—presented in Appendix 2.B—that build on the literature on shadow-rate term-structure models (see, e.g., Krippner, 2015; Wu and Xia, 2016).

[6] Pallara and Renne (2021) provide time-varying estimates of fiscal limits for eight non-euro-area economies; therein, each country is considered independently from the others.

(w.r.t. the yield on their respective national bonds). For the 10-year maturity, this reduction would have been about 30 basis points on average over the estimation sample.

The main concern associated with common debt issuance usually pertains to moral hazard. Under several and joint guarantees issuance schemes, some countries might be tempted to issue more debt given that the interest rate on jointly-guaranteed debt is less sensitive to an individual debt increase than non-guaranteed debt. Although our reduced-form modeling framework does not deal with moral hazard in a structural way, our findings remain valid under the conditions that (i) the amount of debt issued under the SJG scheme is relatively small or that (ii) some form of ex-post redistribution of the yield gains applies. First, as long as a sizable share of countries' funding needs are met with the issuance of national bonds, the overall debt service remains sensitive to countries' indebtedness. Thus, in the absence of redistribution schemes (case (i)), a necessary condition for market discipline to remain effective is to limit the issuance of Eurobonds. In our calculation, we typically envision that jointly-issued debt does not exceed 5% of total consolidated GDP. Second, we show that some ex-post yield gains' redistribution schemes may dampen moral hazard effects (case (ii)). For instance, considering the above-mentioned scheme—in which the issuance of SJG bonds ultimately translates into the same yield reduction for all involved countries—the funding costs of the different countries remain sensitive to the national fiscal conditions, thereby alleviating concerns of reduced fiscal discipline stemming from the issuance of common bonds. More precisely, for this scheme, we obtain that the slope of the curve relating post-redistribution yields to indebtedness is similar to that associated with national bonds (but, for each country, the former curve is below the latter as long as the issuance of SJG bonds is associated with aggregate gains).

The rest of this paper is organized as follows. Section 2.2 reviews related literature. The model is developed in Sections 2.3 (stylized version) and 2.4 (full-fledged version). Section 2.5 describes the estimation strategy. Section 2.6 discusses the results. Section 2.7 summarizes our findings and makes concluding remarks. The appendix gathers technical results, supplementary details, proofs and additional findings.

2.2 Related Literature

This paper contributes to the growing literature on sovereign credit risk and its pricing. Specifically, this paper is among the first to provide a quantitative assessment of Eurobonds' pricing.

To do so, we develop a novel credit risk model where default intensities explicitly depend on fiscal variables.

2.2.1 Eurobonds

Various policy-oriented papers discuss pros and cons of common bond issuance in the euro area, and propose different forms of common bonds. Several of these studies stress that, if issued in large scale, a joint debt instrument could reduce market fragmentation and compete, in terms of size and liquidity, with the US bond market (Giovannini, 2000; Hellwig and Philippon, 2011). De Grauwe and Moesen (2009) and Claessens et al. (2012) argue that joint debt issuance can reduce borrowing costs for stressed sovereigns, allowing for gains at the aggregate level. Following the Great Financial Crisis and the euro-debt crisis, common debt issuance has been advocated by several policy-oriented studies as a device to enhance financial stability, notably because such a safe asset could break the "bank-sovereign doom loop" (European Commission, 2011; Brunnermeier et al., 2017; Delivorias and Stamegna, 2020). The challenges associated with joint debt issuances include coordination issues, political hurdle in transferring sovereignty to the EU level, and the removal of incentives for sound budgetary policies under the current fiscal discipline methods (Claessens et al., 2012). According, among others, to Delpla and Von Weizsacker (2010) and Claessens et al. (2012), common debt issuance calls for enhanced institutional frameworks and ex-ante surveillance to strengthen fiscal discipline.

In Table 2.1, we review the features of some prominent proposals for a European joint debt instrument. Three proposals involve joint guarantees, but with varying proportions: the "Stability bond" approach no. 1 of the European Commission (2011) considers a full replacement of standard national issuances by those of an SJG bond; only short-term debt instruments, amounting to 10% of GDP, would benefit from joint guarantees under the "Eurobills" scheme proposed by Hellwig and Philippon (2011); under the blue/red scheme of Delpla and Von Weizsacker (2010), European countries would pool their public debt up to the Maastricht Treaty threshold—60% of GDP—under joint and several liability as senior ("blue") debt, while debt above this threshold would be issued as junior ("red") debt.

Other schemes depart from joint liability and consist of the partial substitution of European Member States' national issuance with several but not joint guarantees (SNJG) bonds. This is for instance the case of the "Stabilitity bond" approach no. 3 of the European Commission (2011).[7]

[7] Issuance of bonds with several but not joint guarantees can be centralized (e.g., joint debt agency, European Commission, 2011; Delivorias and Stamegna, 2020) or left decentralized (De Grauwe and Moesen, 2009).

Table 2.1: Eurobond proposals: main features

Features	**Joint bond denomination**				
	Stability bonds[a]		**Euro-bills**[b]	**Blue/Red bonds**[c]	**ESBies/EJBies**[d]
	Approach no. 1	**Approach no. 3**			
Guarantees	SJG[e]	SNJG[f]	SJG (10% of GDP)	Only blue: SJG (60% of GDP)	
Tranching		✓		✓	✓
Pooling[g]	✓	✓	✓	✓	✓
New issuance[h]	✓	✓ (partial)	✓ (partial)	✓ (partial)	
Risk of moral hazard	✓	✓	✓	✓	
Coordinated revenue management	✓			✓	
Coordinated debt management	✓	✓	✓	✓	
Pricing attempt in the study					✓ (partial and incomplete)

Notes: This table shows key features of some prominent euro-area joint debt instrument proposals in the literature. *a*: European Commission (2011); *b*: Hellwig and Philippon (2011); *c*: Delpla and Von Weizsacker (2010); *d*: Brunnermeier et al. (2017); *e*: Joint and several guarantees; *f*: Several but not joint guarantees; *g*: with "Pooling" we mean the pooling or common issuance of sovereign debts (either *ex ante* or *ex post* via pooling a portfolio of sovereign debts); *h*: with "New Issuance" we mean the issuance of a new debt instrument replacing totally or partially national bond issuance.

In this scheme, Member States would retain liability for their respective share of "Stabilitity bond" issuance—as well as for their national issuances, naturally.[8] Due to the several but not joint guarantees, moral hazard would be mitigated.

The absence of joint guarantees also underlies Brunnermeier et al. (2017) proposal. Differently from the "Stability bond" approach no. 3 (European Commission, 2011), their proposal does not imply any substitution of national issuance. In their scheme, two synthetic tranches would be created out of a portfolio of (standard) national sovereign bonds, the senior and the junior tranche being respectively dubbed "European Safe Bonds" (ESBies) and "European junior bonds" (EJBies). As safe and liquid assets, ESBies would help limit financial institutions'

[8]The credit quality of a "Stability bond" underpinned by several but not joint guarantees would be close to the weighted average of the credit qualities of the euro-area Member States.

exposure to sovereign credit risk, and thereby break the so-called sovereign-bank doom loop. Brunnermeier et al. (2017) simulate the loss given default of ESBies and EJBies under different tranching scenarios, thereby providing a partial pricing attempt for their instruments.

A few theoretical studies focus on Eurobonds. Tirole (2015) studies the effect of common bonds' issuance, focusing on the moral hazard implications. He distinguishes between two forms of solidarity in a finite-horizon two-country setup: ex-post (spontaneous), e.g., bailouts, and ex-ante (contractual), e.g., joint-bond issuance. Given that one country's default imposes collateral damage on the other country, Tirole (2015) finds that ex-ante (respectively ex-post) solidarity is optimal when both countries exhibit a similar (resp. different) risk profile. Tsiropoulos (2019) builds a two-country general-equilibrium model of sovereign default and finds that welfare consequences of introducing SJG bonds hinge critically on the timing of their introduction. Lastly, Dávila and Weymuller (2016) study the optimal design of flexible joint borrowing agreements between a safe and a risky country; they find gains under joint liability schemes.

2.2.2 Reduced-form approaches and sovereign risk premiums

The present study draws extensively from the reduced-form approaches for pricing sovereign credit risk. Ang and Longstaff (2013) consider multi-factor affine models allowing for both systemic and sovereign-specific credit shocks to price the term structures of US states and Eurozone Member States. Estimating the default intensities for 26 countries, Longstaff, Pan, Pedersen, and Singleton (2011) find that the risk premium represents about a third of credit spreads on average. Monfort and Renne (2014) also estimate substantial sovereign risk premiums in euro-area sovereign spreads, employing a model allowing for both credit and liquidity effects. These studies show a close fit of sovereign bond yields/spreads and provide useful estimates of sovereign risk premiums. However, they do not explicitly account for the economic forces driving the movements of the sovereign default probabilities. By contrast, Borgy, Laubach, Mésonnier, and Renne (2011) and Hördahl and Tristani (2013) propose sovereign credit risk frameworks where default intensities explicitly depend on fiscal variables, and demonstrate that the fiscal environment is able to capture part of the fluctuations of sovereign credit spreads.

2.2.3 Theory of fiscal limits

Our paper relates to the literature studying the concept of *fiscal limit*, namely the maximum outstanding debt that a country could credibly sustain. In Bi (2012), Leeper (2013), Bi and Leeper

(2013), Bi and Traum (2012), the concept of fiscal limit corresponds to the net present value of future maximum primary surpluses.[9] These maximum surpluses represent those surpluses implicit in the peak of the Laffer curve (Trabandt and Uhlig, 2011). After having introduced an estimated parametric reaction function of primary surpluses in a model of debt accumulation, Ghosh et al. (2013) show that there is a point—akin to the fiscal limit—where the primary balance cannot keep pace with the rising interest burden as debt increases. Beyond this point, debt dynamics becomes explosive and the government becomes unable to fully meet its obligations. Collard, Habib, and Rochet (2015) also exploit the idea of a maximum primary surplus to derive a measure of debt limit. More recently, Mehrotra and Sergeyev (2020) combine disaster risk and fiscal fatigue. In their framework, as in Lorenzoni and Werning (2013), debt dynamics are subject to a tipping point situation: in some instances, the public debt can be on an unsustainable path without immediately triggering default.

In the present paper, we do not make the maximum surplus explicit and we rely on a reduced-form approach. Assuming that the default intensity becomes strictly positive when the effective (observed) debt is higher than the (unobserved) fiscal limit, we infer the latter from bond prices.

2.3 Stylized model

As mentioned above, a crucial ingredient of our modelling framework is the relationship between the fiscal space—the difference between the fiscal limit and debt—and the sovereign probability of default. The parametric function we retain to model this relationship is presented in Subsection 2.3.1. Before incorporating this ingredient in a standard asset pricing model (in Section 2.4), we present a stylized model in Subsection 2.3.2. In Subsection 2.3.3, we elaborate on the pricing of SJG and SNJG common bonds in this simplified framework; and we discuss resulting asset-pricing mechanisms in Subsection 2.3.4.

2.3.1 Sovereign default probability

On each date t, we assume that the default probability of country j ($j = A, B$) is given by

$$1 - \exp(-\underline{\lambda}_{j,t}), \tag{2.1}$$

[9] We refer to Aguiar and Amador (2014) or Yue and Wei (2019) for a general presentation of the theory of sovereign debt.

where the default intensity $\underline{\lambda}_{j,t}$ is assumed to negatively depend on the fiscal space, defined as the distance between fiscal limit-to-GDP ($\ell_{j,t}$) and debt-to-GDP ($d_{j,t}$). Specifically:[10]

$$\underline{\lambda}_{j,t} = \alpha \max(0, d_{j,t} - \ell_{j,t}). \tag{2.2}$$

The previous formulation implies that the probability of default is strictly positive only if the fiscal space is negative, i.e. if debt stands above the fiscal limit. Parameter α characterizes the nature of the fiscal limit: if α is large, the fiscal limit is "strict", as the probability of default becomes large as soon as debt breaches the fiscal limit; for lower values of α, the fiscal limit is "soft", as negative fiscal spaces then do not necessarily trigger default.

The notion of soft fiscal limit is consistent with the widespread idea that it is difficult to assess sovereign debt sustainability (e.g., Warmedinger et al., 2017; Debrun et al., 2019), which gives rise to "grey areas" where default becomes likely but can also be avoided. The World Bank and the IMF themselves reckon that, alongside quantitative approaches, the use of judgment is needed to assess sovereign debt sustainability (IMF and World Bank, 2021).

In the rest of the present section, we consider the case of $\alpha = 1$, implying a relatively soft concept of fiscal limit: for a fiscal space of -1% of GDP, the probability of default is 1%.[11]

2.3.2 Assumptions of the stylized model

Investors are risk-neutral and risk-free interest rates are zero. In this context, the date-t price of a one-period zero-coupon zero-recovery-rate bond issued by j is simply given by:

$$P_{t,1}^{(j)} = \mathbb{E}_t \exp(-\max[0, d_{j,t+1} - \ell_{j,t+1}]), \tag{2.3}$$

where $\mathbb{E}_t$ denotes the expectation conditional on the information available to the investor as of date t.

[10] It can be seen that we have $\underline{\lambda}_{j,t} = \max(0, \lambda_{j,t})$, with $\lambda_{j,t} = \alpha \times (d_{j,t} - \ell_{j,t})$. Using the vocabulary introduced by Black (1995), $\lambda_{j,t}$ can be interpreted as a "shadow default intensity." Alternatively, to have a non-negative intensity, $\underline{\lambda}_{j,t}$ could be modeled as a quadratic function of the fiscal space (see, e.g., Doshi et al., 2013). However, it is impossible to have a monotonous relationship between the (non-negative) default intensity and the covariates in a quadratic framework (while such a monotonous relationship is expected to hold in the present context). Coroneo and Pastorello (2020) also employ the shadow-rate approach to price sovereign bonds issued by different countries; contrary to the present paper though, sovereign default probabilities (or default intensities) are not explicitly modeled in their yields-only reduced-form framework. Therefore, the framework of Coroneo and Pastorello (2020) does not allow to recover sovereign probabilities of default, and cannot preclude negative default probabilities.

[11] Low values of α allow for approximate pricing formulas (Appendix 2.B) that are intensively used in our empirical analysis (Section 2.4). As shown by Footnote 12, these approximate formulas are not needed in the context of the stylized model.

For each country, the fiscal limit-to-GDP ($\ell_{j,t}$) is constant, fixed at $\overline{\ell}_j$, and the debt-to-GDP ratios are i.i.d. Gaussian:

$$\begin{bmatrix} d_{A,t} \\ d_{B,t} \end{bmatrix} \sim \mathcal{N}\left(\begin{bmatrix} \overline{d}_A \\ \overline{d}_B \end{bmatrix}, \begin{bmatrix} \sigma^2 & \rho\sigma^2 \\ \rho\sigma^2 & \sigma^2 \end{bmatrix}\right). \tag{2.4}$$

In this context, the prices of zero-coupon bonds (see eq. 2.3) admit closed-form solutions deduced from standard results on truncated normal distributions.[12]

2.3.3 Common bonds

We consider two types of common bonds: the first is backed by several and joint (SJG) guarantees, whereby each issuing country guarantees the totality of the obligations, and the second features several but not joint (SNJG) guarantees, whereby each issuing country guarantees only its share of the joint instrument.

A natural way to conceive the SJG bond is to consider that it is issued by a synthetic area where both fiscal revenues and debts are pooled, and to assume that this area also features a probability of default of the form of (2.1). Denoting by ω the vector of GDP weights, the price of SJG bond is given by:[13]

$$P_{t,1}^{(SJG)} = \mathbb{E}_t \exp(-\max[0, \omega \cdot d_{t+1} - \omega \cdot \overline{\ell}]), \tag{2.5}$$

where $\omega \cdot d_{t+1} = \omega_A d_{A,t+1} + \omega_B d_{B,t+1}$ and $\omega \cdot \overline{\ell} = \omega_A \overline{\ell}_{A,t+1} + \omega_B \overline{\ell}_{B,t+1}$ are, respectively, the GDP-weighted debt-to-GDP ratio and the GDP-weighted fiscal limit.

Regarding the SNJG bond, the absence of joint guarantee implies that the payoff of this bond is of the form $\omega \cdot (1 - \mathcal{D}_{t+1})$, where $\mathcal{D}_{t+1} = [\mathcal{D}_{A,t+1}, \mathcal{D}_{B,t+1}]$ is the vector of default indicators—a default indicator being equal to 1 in the case of default, and to 0 otherwise.[14] In other words, the payoff is equal to 1 if none of the countries default on date $t+1$, ω_A (respectively ω_B) if

[12]Formally, $P_{t,1}^{(j)}$ is given by:

$$\Phi\left(\frac{\overline{\ell}_j - \overline{d}_j}{\sigma}\right) + \left(1 - \Phi\left(\frac{\overline{\ell}_j - \overline{d}_j}{\sigma}\right)\right) \exp\left(\alpha(\overline{\ell}_j - \overline{d}_j) + \frac{\alpha^2\sigma^2}{2}\right) \left\{1 - \Phi\left(\frac{\overline{\ell}_j - \overline{d}_j}{\sigma} + \alpha\sigma\right)\right\} \Big/ \left\{1 - \Phi\left(\frac{\overline{\ell}_j - \overline{d}_j}{\sigma}\right)\right\}.$$

[13]ω is such that $\omega = [\omega_A, 1 - \omega_A]$, with $\omega_A = Y_A/(Y_A + Y_B)$, where Y_j is country j's GDP.

[14]We conceive state 1 (default) as an absorbing case. Given that the default state is a stopping time—in the sense that, in a case of default, the last payoff is on the default date—we can make this assumption without loss of generality (even when we will consider longer-term bonds).

only B (resp. A) defaults on date $t+1$, and 0 if both countries default. This implies that the price of a SNJG bond is given by:

$$P_{t,1}^{(SNJG)} = \omega_A \mathbb{E}_t(1 - \mathcal{D}_{A,t+1}) + \omega_B \mathbb{E}_t(1 - \mathcal{D}_{B,t+1}) = \omega \cdot P_{t,1}, \tag{2.6}$$

with $P_{t,1} = [P_{t,1}^{(A)}, P_{t,1}^{(B)}]$.

2.3.4 Calibration and resulting yields

The different calibrations used in this section are summarized in Table 2.2. In our baseline case, we set the average fiscal spaces of both countries to 20% ($= \bar{\ell}_j - \bar{d}_j = 100\% - 80\%$), and the two countries are alike in all respects. In particular, they have the same (GDP) size, i.e. $\omega_A = \omega_B = 50\%$, and the correlation between debts is set to 50%. In this baseline case, the yields on one-year national bonds are equal to 28 basis points.[15] In this baseline context, where both countries are similar, it also comes that SNJG bond prices are equal to those of country-specific bonds (see eq. 2.6, with $P_{t,1}^{(A)} = P_{t,1}^{(B)}$); the SNJG bond yield is therefore also equal to 28 basis points. By contrast, the price of the SJG bond is higher, the SJG bond yield being of 13 basis points. This results from the fact that, for the synthetic "pooled" area, the probability to have an (average) debt-to-GDP larger than the (average) fiscal limit is lower than for a single country. Formally:

$$\mathbb{P}(\omega_A d_{A,t} + \omega_B d_{B,t} \geq \bar{\ell}) < \mathbb{P}(d_{j,t} \geq \bar{\ell}), \quad j = A, B, \tag{2.7}$$

which is true as long as the correlation between the two debt-to-GDP ratios is strictly lower than 1. The fact that the SJG bond yield is lower than national bond yields implies that both countries would reduce their debt service through the issuance of joint-liability bonds.

The baseline situation discussed above is represented by a vertical grey bar in the first row of plots in Figure 2.1. These plots further show how the SNJG and SJG yields are affected with respect to: (Panel A.1) changes in the fiscal spaces of the two countries, (Panel A.2) changes in the correlation across debts, and (Panel A.3) changes in the relative size of country A (in terms of GDP).

[15] Since, in the stylized model described in this section, risk-free yields are taken equal to zero, bond yields essentially correspond to credit spreads. In addition, since the recovery rate is also zero, yields here coincide with probabilities of default. These restrictions are relaxed in the extended model (Section 2.4).

Table 2.2: Calibrations of stylized models

	Baseline	**Symmetric case** (A and B are alike)			**Asymmetric case** (B's fiscal space $\leq$ A's fiscal space)		
		A.1	**A.2**	**A.3**	**B.1**	**B.2**	**B.3**
σ	12.5%						
$\bar{\ell}_{A,B}$	100%						
$\bar{d}_B$	80%	75%→95%			95%	95%	95%
$\bar{d}_A$	80%	75%→95%			75%→95%		
ρ	50%		0%→100%			0%→100%	
ω_A	50%			0%→100%			0%→100%

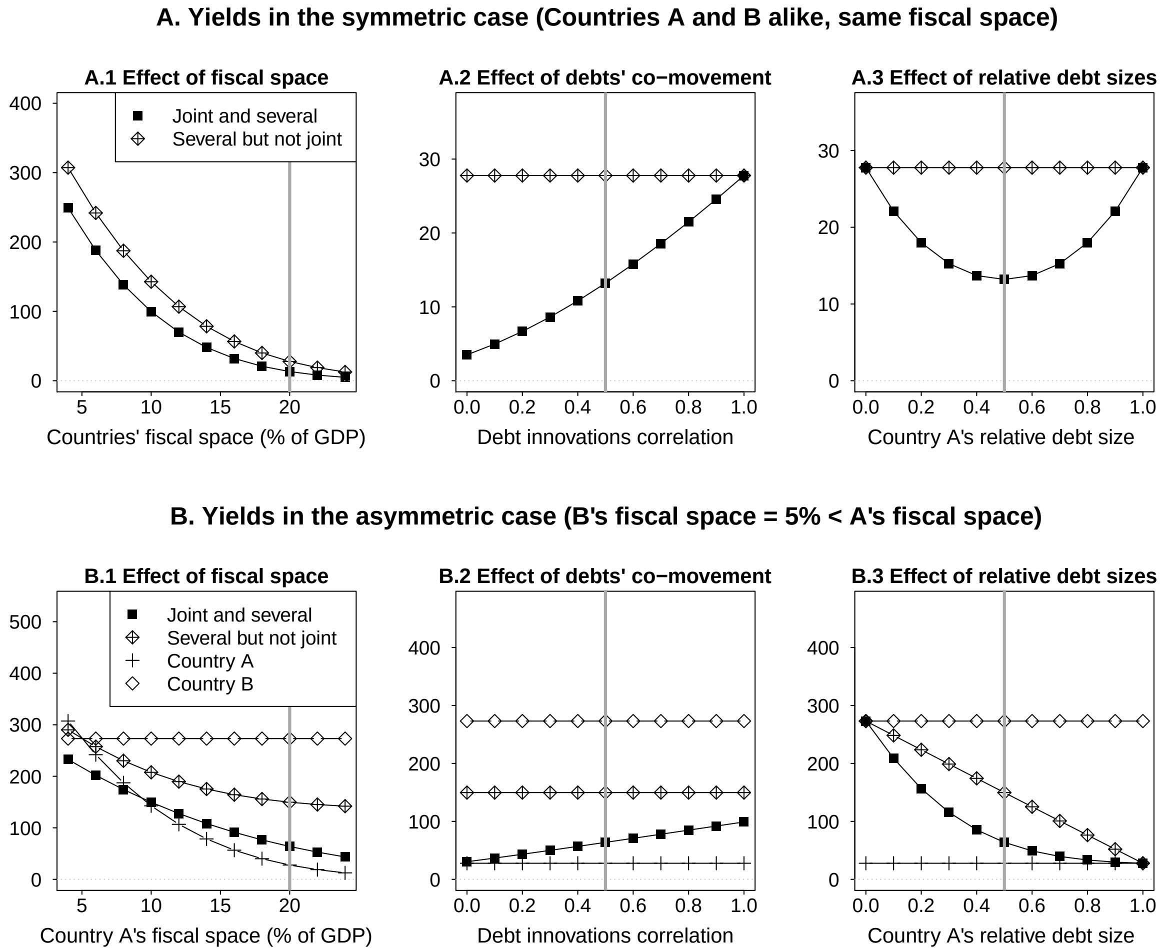
Notes: This table summarizes the calibrations used in our stylized model. The first column shows the calibration of the baseline case (represented by a vertical grey line in the first row of plots of Figure 2.1). The average fiscal space of country j corresponds to $\bar{\ell}_j - \bar{d}_j$, and ω_A denotes the relative GDP size of country A (such that $\omega_B = 1 - \omega_A$).

Panel A.1 shows that both SJG and SNJG bond yields nonlinearly decrease when fiscal spaces increase. It also shows that SJG bond yields are consistently lower than those of SNJG bonds. Panel A.2 illustrates the importance of debt co-movements to account for the yield reduction resulting from joint guarantees: while the SNJG bond yield is not affected by changes in debts' correlations, the yield of a SJG bond is reduced by a factor of 8 when the correlation decreases from 100%—in which case all bonds are equivalent—to 0%. Panel A.3 focuses on the effect of the two countries' relative sizes. In the extremes, when the relative size of country A is either 0 or 1, there is no difference between SJG and SNJG bonds. As in the case of debt co-movement, and because we consider two equally-risky countries for the time being, the relative size of country A has no effect on the SNJG yield. But it has on the SJG yield; the effect is maximum when the two countries are equally large, corresponding to a situation where diversification effects are maximum.[16]

The second row of plots in Figure 2.1 displays results obtained in an asymmetric situation, where country B is riskier than country A. We fix the fiscal space of country B to 5%, keeping A's one at 20%. National bond yields are now different for the two countries, and we add them to each plot. Up to small convexity effects, it can be checked that SNJG yields are equal to the GDP-weighted averages of the two national bond yields. In particular, in Panel B.3, where we modify the relative size of country A from 0 to 1, the SNJG bond yield goes from the (higher) country-B yield to the (lower) country-A yield. Regarding the difference between SNJG and SJG yields, an interesting situation is captured by Panel B.1: for low values of country A's fiscal space, not only is the SJG bond yield below the SNJG one (i.e. the average of the two national

[16]Formally, this is because the variance of the aggregate debt-to-GDP ratio—that is $\sigma^2(\omega_A^2 + (1 - \omega_A^2) + 2\rho\omega_A(1 - \omega_A))$—admits a minimum for $\omega_A = \frac{1}{2}$.

Figure 2.1: Two-country stylized model mechanisms



Notes: These plots show the yields-to-maturity, expressed in basis points, of different types of one-period bonds; it also shows how these yields are affected by changes in the calibration of the stylized model (see Table 2.2 for details regarding the baseline calibration and the alternative calibrations underlying Panels A.1 to B.3 of this figure). Three types of bonds are considered: national, or country-specific, bonds issued by countries A and B; a bond with several and joint guarantees (SJG); and a bond with several but not joint guarantees (SNJG). See Subsection 2.3.4 for more details. On each row of plots, the vertical grey line represents the same situation—the "baseline" case of Table 2.2.

bond yields), it is also lower than the safer country's bond yields. Finally, Panel B.2 shows that when the two countries do not have the same average fiscal space, a correlation of 1 across debts does not imply that the SJG and the SNJG bonds are equivalent. In this extreme case, and contrary to the symmetric case, diversification effects are still at play in the SJG bond pricing: the SJG bond yield is 1.5 times lower than the SNJG one.

Moreover, it is interesting to note that, in this framework, diversification mechanisms can have adverse effects on SJG bonds prices when expected fiscal spaces are negative enough. Intuitively, when this is the case, the distribution of the joint fiscal space (and therefore of the default intensity, see eq. 2.2) turns out to be more concentrated on the "wrong side" of zero, yielding to lower prices for SJG bonds. We discuss this situation in greater details in Appendix 2.D.

To sum up, the main drivers of the wedge between the SNJG and the SJG bond yields — namely, the aggregate yield gains — are the comovement of fiscal positions across issuing countries and the expected joint fiscal space. As long as either fiscal positions across countries are not perfectly correlating or the expected joint fiscal space is not largely negative, there is scope for diversification of risk across issuing countries and, thus, aggregate yield gains arise.

2.4 Model

In this section, we enrich the stylized model to make it amenable to the data. We consider N countries. While the conditional probabilities of default remain as in Subsection 2.3.1—with default intensities that depend on fiscal spaces—debt-to-GDP ratios and fiscal limits are now time-varying; in addition, the state vector is augmented with a stochastic short term interest rate (Subsection 2.4.1). The representative investor is now risk averse, her risk preferences being captured by a reduced-form stochastic discount factor (Subsection 2.4.2). After having derived prices of zero-coupon risk-free bonds, we discuss the pricing of zero-coupon bonds with non-zero recovery rates and bond yield spreads (Subsection 2.4.3). The ability to swiftly price risk-free bonds and yield spreads is crucial to estimate the model (Section 2.5).

2.4.1 Dynamics of the state vector

Fiscal limits follow autoregressive processes of order one. For country j:

$$\ell_{j,t} = (1-\rho_\ell)\overline{\ell}_j + \rho_\ell \ell_{j,t-1} + \varepsilon_{\ell,j,t}, \tag{2.8}$$

where the $\varepsilon_{\ell,j,t}$'s are Gaussian white noise shocks.

The formulation of debt-to-GDP dynamics is inspired by standard debt accumulation processes, where debt-to-GDP depends on its first lag and on the budget surplus. Specifically:[17]

$$d_{j,t} = \rho_d d_{j,t-1} + \{\overline{\gamma}_j + \gamma_{j,t}\}, \tag{2.9}$$

[17] Standard debt-to-GDP accumulation processes read $d_t = \frac{1+r_t}{1+g_t} d_{t-1} + \tilde{\gamma}_t$, where r_t denotes the apparent interest rate (i.e., debt service over debt outstanding), g_t denotes GDP growth, and $\tilde{\gamma}_t$ is the primary deficit.

where $\overline{\gamma}_j + \gamma_{j,t}$ proxies for country j's primary deficit (expressed as a fraction of GDP). The cyclical part of the deficit, $\gamma_{j,t}$, is assumed to follow an autoregressive process of order one:

$$\gamma_{j,t} = \rho_\gamma \gamma_{j,t-1} + \varepsilon_{d,j,t}. \tag{2.10}$$

Three remarks are in order. First, since $\gamma_{j,t}$ is of mean zero, eq. (2.9) implies that the unconditional mean of $d_{j,t}$ is given by $\overline{d}_j = \overline{\gamma}_j/(1-\rho_d)$. Second, from eqs. (2.9) and (2.10), it comes that $d_{j,t}$ follows an autoregressive process of order two; one can indeed easily check that $d_t = (1-\rho_\gamma)(1-\rho_d)\overline{d}_j + (\rho_d + \rho_\gamma)d_{t-1} - (\rho_d\rho_\gamma)d_{t-2} + \varepsilon_{d,j,t}$. Third, considering that both ρ_d and ρ_γ are in $[0,1[$, the debt process is stationary. Since investors use the previous processes to price government bonds, our framework implicitly excludes snowball effects and related multiple equilibria (that would give rise to non-stationary processes). This can be attributed to investors' limited rationality.

Fiscal-limit shocks ($\varepsilon_{\ell,j,t}$) and debt shocks ($\varepsilon_{d,j,t}$) can be correlated. Formally, using the notations $\varepsilon_{d,t} = [\varepsilon_{d,1,t}, \dots, \varepsilon_{d,N,t}]'$ and $\varepsilon_{\ell,t} = [\varepsilon_{\ell,1,t}, \dots, \varepsilon_{\ell,N,t}]'$, we set:

$$\begin{bmatrix} \varepsilon_{d,t} \\ \varepsilon_{\ell,t} \end{bmatrix} \sim i.i.d.\,\mathcal{N}(0,\Omega), \quad \text{with } \Omega = \begin{bmatrix} \Omega_d & \Omega'_{\ell,d} \\ \Omega_{\ell,d} & \Omega_\ell \end{bmatrix}.$$

The structures of Ω_d, Ω_ℓ, and $\Omega_{\ell,d}$ will be explained below, in Subsection 2.5.3, which details the estimation strategy and the parameter constraints.

The short-term risk-free interest rate also follows an auto-regressive process:[18]

$$i_t = (1-\rho_i)\bar{i} + \rho_i i_{t-1} + \sigma_i \eta_{i,t}, \quad \eta_{i,t} \sim i.i.d.\,\mathcal{N}(0,1). \tag{2.11}$$

Let us denote by d_t and ℓ_t two N-dimensional vectors gathering countries' debt-to-GDP ratios and fiscal limits, respectively. Under the previous assumptions, it is easily seen that the state vector $X_t = [i_t, i_{t-1}, d'_t, d'_{t-1}, \ell'_t]'$ follows a vector autoregressive process of order one.[19]

[18]This Gaussian process does not exclude negative nominal interest rates. Hence, this model is not consistent with the existence of the Zero Lower Bound (ZLB). This simple model however conveniently accommodates the period during which risk-free European nominal rates have been negative (indicating that the Effective Lower Bound, ELB, is lower than zero).

[19]We include the lagged short-term interest rate, i_{t-1}, in the state vector because the defaultable-bond pricing formulas are easier to derive if i_{t-1} can be expressed as a linear combination of X_t (see the notation below eq. a.2.7). Moreover, X_t includes d_{t-1} because, as mentioned above, d_t follows an auto-regressive process of order two under eqs. (2.9) and (2.10).

That is:

$$X_t = \mu + \Phi X_{t-1} + \Sigma \eta_t, \tag{2.12}$$

where $\eta_t \sim i.i.d.\mathcal{N}(0, I)$, and where μ, Φ, and Σ (with $\Omega = \Sigma\Sigma'$) are detailed in Appendix 2.A.

2.4.2 Stochastic discount factor and the term structure of risk-free rates

We assume that arbitrage opportunities do not exist, which ensures the existence of a positive stochastic discount factor (s.d.f.). Following Ang and Piazzesi (2003), we posit a reduced-form exponential affine s.d.f. between dates t and $t+1$:

$$\mathcal{M}_{t,t+1} = \exp(-i_t)\frac{\xi_{t+1}}{\xi_t}, \tag{2.13}$$

where ξ_{t+1} follows:

$$\xi_{t+1} = \xi_t \exp\left(-\frac{1}{2}\psi_t'\psi_t - \psi_t'\eta_{t+1}\right), \tag{2.14}$$

ψ_t being a vector of prices of risk that linearly depends on X_t:

$$\psi_t = \psi_0 + \psi_1 X_t. \tag{2.15}$$

In this context, it is well-known that risk-free bond prices admit closed-form recursive solutions. Specifically, the date-t price of a risk-free zero-coupon bond of maturity h is given by (proof in Appendix 2.F):

$$B_{t,h} = \exp(A_h + B_h X_t), \tag{2.16}$$

where $A_1 = 0$ and $B_1 = [-1, 0, \dots]'$ and, for $h > 1$:

$$\left\{ \begin{array}{lcl} A_h & = & A_{h-1} + B_{h-1}'(\mu - \Sigma\psi) + \frac{1}{2}B_{h-1}'\Sigma\Sigma' B_{h-1} \\ B_h & = & B_1 + \Phi' B_{h-1}. \end{array} \right. \tag{2.17}$$

Equivalently, the yield of a risk-free zero-coupon bond of maturity h is given by:

$$i_{t,h}^0 = -1/h(A_h - B_h' X_t). \tag{2.18}$$

2.4.3 Zero-coupon bonds with non-zero recovery rates and sovereign bond yield spreads

Consider a zero-coupon bond of maturity h issued by country j. Our recovery payoff assumption is based on the "Recovery of Treasury" (RT) convention of Duffie and Singleton (1999): on date $t+k$, with $0 < k \leq h$, the payoff of the considered bond is zero, unless the country defaults on date $t+k$, in which case the bond payoff is assumed to be the fraction RR (recovery rate) of the price of a risk-free zero-coupon bond of equivalent residual maturity, i.e. $\exp[-(h-k)i^0_{t+k,h-k}]$. Formally, the payoffs of this bond are of the form:[20]

$$\begin{cases} RR \times \exp(-(h-k)i^0_{t+k,h-k}) \times (\mathcal{D}_{j,t+k} - \mathcal{D}_{j,t+k-1}) & \text{if} \quad 0 < k < h, \\ 1 - \mathcal{D}_{j,t+k} + RR \times (\mathcal{D}_{j,t+k} - \mathcal{D}_{j,t+k-1}) & \text{if} \quad k = h. \end{cases}$$

Denoting by $\mathcal{M}_{t,t+k}$ the stochastic discount factor between dates t and $t+k$ (i.e., $\mathcal{M}_{t,t+k} = \mathcal{M}_{t,t+1} \times \cdots \times \mathcal{M}_{t+k-1,t+k}$) and after some algebra (Appendix 2.I), one obtains the following expression for the price of this bond:

$$\mathcal{P}^{(j)}_{t,h} \quad = \quad (1-RR) \times \mathbb{E}_t\left(\mathcal{M}_{t,t+h}(1-\mathcal{D}_{j,t+h})\right) + RR \times B_{t,h}, \tag{2.19}$$

where $B_{t,h}$, the price of the risk-free bond (Subsection 2.4.2), is equal to $\mathbb{E}_t(\mathcal{M}_{t,t+h})$, and the conditional expectation $\mathbb{E}_t\left(\mathcal{M}_{t,t+h}(1-\mathcal{D}_{j,t+h})\right)$ corresponds to the date-t price of a zero-coupon zero-recovery-rate bond of maturity h providing a payoff of 1 on date $t+h$ if country j has not defaulted before $t+h$, and zero otherwise. Appendix 2.B details the computation of the latter conditional expectation.

Sovereign bond yields for country j are given by:

$$i^{(j)}_{t,h} = -\log(\mathcal{P}^{(j)}_{t,h})/h, \tag{2.20}$$

and sovereign spreads are computed as follows ($i^0_{t,h}$ being given by eq. 2.18):

$$s^{(j)}_{t,h} = i^{(j)}_{t,h} - i^0_{t,h}. \tag{2.21}$$

[20] Note that $\mathcal{D}_{j,t}$ is valued in $\{0,1\}$, state 1 being the default state, which is absorbing. As a result, $\mathcal{D}_{j,t+k} - \mathcal{D}_{j,t+k-1}$ is equal to zero, except once, on the default date, where it is equal to 1. In reality, the default state is not absorbing. However, given that the default state is a stopping time—in the sense that, in a case of default, the last payoff is on the default date—we can make this assumption without loss of generality.

2.5 Estimation

2.5.1 Data

We consider six European countries: Germany, France, Italy, Spain, Netherlands, and Belgium. These countries' GDPs account for close to 90% of euro-area's GDP. The data are quarterly and span the period from 2008Q2 to 2021Q2. Sovereign yields and the 3-month Overnight Indexed Swap (OIS) interest rate—our short-term risk-free rate—are extracted from Thomson Reuters Datastream. Following Monfort and Renne (2014), risk-free yields of maturities of 2, 3, 5, and 10 years are proxied for by the difference between German bond yields and German CDSs of matching maturities. Observations of sovereign spreads ($s_{t,h}^{(j)}$'s in eq. 2.21) are computed as the difference between national bond yields and these risk-free yields. We consider three maturities of bond yield spreads: 3, 5, and 10 years. Time series of gross government debts and GDPs are collected from the Eurostat ESA2010 database.

2.5.2 Estimation approach

The model can be cast into a state-space form, with (i) transition equations describing the dynamics of the state variables (this is eq. 2.12) and (ii) measurement equations describing the relationships between observed financial market data—prices and yield spreads—and the state vector. Let us denote by Θ the set of model parameters,[21] the state-space model is of the form:

$$\begin{array}{llll} (i) & X_t & = & \mathcal{F}(X_{t-1}, \eta_t; \Theta), \quad \text{(reformulation of eq. 2.12)} \\ (ii) & Y_t & = & \mathcal{G}(X_t; \Theta) + \xi_t, \end{array}$$

where $X_t = [i_t, i_{t-1}, d_t', d_{t-1}', \ell_t']'$ is the state vector, Y_t denotes the vector of financial market data (gathering risk-free yields and sovereign spreads), and ξ_t is a vector of i.i.d. Gaussian measurement errors. Function $\mathcal{G}$ stands for pricing formulas, associating the state X_t to risk-free yields and sovereign spreads. While the risk-free rates are affine in X_t (see eq. 2.18), this is not the case for sovereign spreads because of the nonlinearity of the default intensity (resulting from the "max" operator in eq. 2.2). The vector of state variables X_t is only partially observed by the econometrician since the N national fiscal limits (ℓ_t) are latent. We therefore face two types of unknowns: the model parameters and the fiscal limits. We address this problem by employing "inversion techniques". These techniques, originally introduced by Chen and Scott (1993) in

[21] We have $\Theta = \{\bar{i}, \rho_i, \sigma_i, \bar{d}_1, \dots, \bar{d}_N, \rho_d, \bar{\ell}_1, \dots, \bar{\ell}_N, \rho_\ell, \rho_\gamma, \Omega_d, \Omega_\ell, \Omega_{\ell,d}, \psi_0, \psi_1\}$.

the term structure literature, consist in estimating the latent pricing factors by inverting a non-singular system relating prices to latent factors. This system results from the assumption that some of the observed prices are modeled without errors. In the present case, we assume that, for each country, the averages of the three sovereign spreads (with maturities 3, 5, and 10 years) are perfectly priced. Under this assumption, we can recover the fiscal limits and, simultaneously, compute the likelihood function associated with the considered model parametrization.[22] This opens the door to maximum-likelihood estimation. Appendix 2.K provides computational details.

2.5.3 Parameter constraints and estimates

To facilitate the estimation and ensure plausible fiscal limit estimates, some parameters are calibrated or restricted to lay in pre-specified intervals.

For all countries, we set the stationary debt-to-GDP ratio $\bar{d}_j$ and the unconditional mean of the short term rate $\bar{i}$ to their respective sample averages. We restrict the unconditional standard deviation of the short-term rate to be larger than 0.8%, which is slightly lower than the sample standard deviation. To favor numerical stability, we impose upper bounds, of 0.999, to all autoregressive parameters. We set the bounds for the unconditional mean of the fiscal limit ($\bar{\ell}$ in Table 2.3) to lie in between 0% and 300% of GDP.[23] We restrict the autoregressive parameter of the $\gamma_{j,t}$'s—the proxies for the cyclical components of primary surpluses—to be larger than 0.7, given that the cross-country average of the autocorrelations of primary balance is 0.8. The maximum Sharpe ratio, that characterizes the pricing of risk, is supposed to lie between 0.5 and 1.5 (it is set to one in Cochrane and Saa-Requejo, 2000).[24] The standard deviations of the measurement errors associated with yields and sovereign spreads are respectively set to 10 basis points and to 10% of the country-wise sample standard errors of sovereign spreads.

[22] The likelihood then involves an adjustment term corresponding to the determinant of the Jacobian matrix associated with the non-singular system; this adjustment results from the transformation of the observables to the latent components (see e.g. Ang and Piazzesi, 2003, Appendix B). Computational details are given in Appendix 2.K.

[23] The bounds on the unconditional mean of fiscal limits is based on the observation that the average of estimates obtained by Ghosh et al. (2013) and Collard et al. (2015) fall within the same interval (and are never above 220% of GDP for the set of countries here analyzed). These (static) estimates are reported in Table 2.O1 in Appendix 2.O.

[24] Setting bounds on the maximum Sharpe ratio is an approach that is employed by, e.g., Jiang et al. (2019). A maximum Sharpe ratio below 0.5 would be inconsistent with the empirical evidence, as Sharpe ratios above 0.5 are frequent (e.g., Lettau and Ludvigson, 2010; Hong and Linton, 2020). Note that, in our framework, the maximum Sharpe ratio is not constant since it depends on the short term rate i_t dynamics (see Appendix 2.C). The maximum Sharpe ratio that we effectively constrain is the one evaluated at the average of the state vector, which is close to its sample average given that i_t is fairly constant over the sample. More details are provided in Appendix 2.C.

The model parameters include the conditional covariance matrix of X_t, i.e., $\Omega = \Sigma\Sigma'$ (see eq. 2.12). Freely estimating all the parameters of this matrix would be numerically challenging. Instead, we design an approach that, while capturing the sample correlation structure of debt shocks, remains parsimonious. It works as follows:

(i) For a considered parametrization of ρ_d, ρ_γ, and of the $\bar{d}_j$'s (see eqs. 2.9 and 2.10), one can recover estimates of the $\varepsilon_{d,j,t}$'s. We perform a PCA analysis of the resulting shocks, and we denote the resulting standardized PCAs by $\eta_{d,k,t}$'s $(k =, 1, \dots, N)$. At that stage, we have:

$$\varepsilon_{d,t} = \Gamma_d \eta_{d,t}, \tag{2.22}$$

where Γ_d is the matrix of PCA weights, that is such that $\mathbb{V}ar(\varepsilon_{d,t}) = \Gamma_d \Gamma_d'$ (using that $\mathbb{V}ar(\eta_{d,t}) = I$). Note that the parameters of the first column of Γ_d are larger as they correspond to the first PCA. In other words, $\eta_{d,1,t}$ accounts for the largest common variance of the $\varepsilon_{d,j,t}$'s $(k =, 1, \dots, N)$.

(ii) We further assume that the fiscal-limit shocks admit the same structure, up to a multiplicative factor. More precisely, we assume that:

$$\varepsilon_{\ell,t} = \Gamma_\ell \eta_{\ell,t}, \tag{2.23}$$

with $\Gamma_\ell = \zeta \Gamma_d$ and $\mathbb{V}ar(\eta_{\ell,t}) = I$. Again, by construction, the first column of Γ_ℓ contains larger parameters. That is, $\eta_{\ell,1,t}$ is the main common driver of the fiscal-limit shocks. In the estimation, we restrict ζ to be between 0.5 and 1.5.

(iii) In order to allow for correlation between debts and fiscal limits in a parsimonious way, we assume that the two "main common shocks," namely $\eta_{d,1,t}$ and $\eta_{\ell,1,t}$, are correlated. Specifically, we assume that these shocks admit the following decomposition:

$$\eta_{d,1,t} = \sqrt{1-\rho_{d,\ell}}\,\tilde{\eta}_{d,1,t} + \sqrt{\rho_{d,\ell}}\,\tilde{\eta}_{d,\ell,t} \tag{2.24}$$

$$\eta_{\ell,1,t} = \sqrt{1-\rho_{d,\ell}}\,\tilde{\eta}_{\ell,1,t} + \sqrt{\rho_{d,\ell}}\,\tilde{\eta}_{d,\ell,t}, \tag{2.25}$$

where $\tilde{\eta}_{d,\ell,t}$, $\tilde{\eta}_{d,1,t}$, and $\tilde{\eta}_{\ell,1,t}$ are independent standard Gaussian shocks. Together, eqs. (2.24) and (2.25) imply that $\rho_{d,\ell}$ is the correlation between $\eta_{d,1,t}$ and $\eta_{\ell,1,t}$.

Hence, the complete vector of independent shocks affecting the system (eq. 2.12) is:

$$\eta_t = \left[\begin{array}{cc} \eta_{i,t} & \tilde{\eta}_t' \end{array}\right]' \sim i.i.d.\mathcal{N}(0, I), \quad \text{with } \tilde{\eta}_t = \left[\begin{array}{ccc} \tilde{\eta}_{d,\ell,t} & \tilde{\eta}_{d,t}' & \tilde{\eta}_{\ell,t}' \end{array}\right]', \tag{2.26}$$

where $\tilde{\eta}_{d,t} = [\tilde{\eta}_{d,1,t}, \eta_{d,2,t}, \ldots, \eta_{d,N,t}]'$ and $\tilde{\eta}_{\ell,t} = [\tilde{\eta}_{\ell,1,t}, \eta_{\ell,2,t}, \ldots, \eta_{\ell,N,t}]'$. Appendix 2.A details the shape of matrices Σ and Ω (with $\Omega = \Sigma\Sigma'$) that results from these assumptions.

To discipline the estimation further, we adopt a parsimonious specification for the prices of risk (ψ_t in eqs 2.14 and 2.15). First, we assume that only the first entry of ψ_t—that corresponds to interest-rate risk—is time-varying. Specifically, we have: $\psi_{1,t} = \psi_{i,0} + \psi_{i,1} i_t$. As a result, matrix ψ_1 (eq. 2.15) is filled with zeros, except its $(1,1)$ entry, which is equal to $\psi_{i,1}$.[25] Second, as regards debt and fiscal-limit shocks ($\tilde{\eta}_t$), we assume that the s.d.f. depends only on the main common shocks, namely $\tilde{\eta}_{d,1,t}$ and $\tilde{\eta}_{\ell,1,t}$. Formally, we posit:

$$\psi_0 = [\psi_{i,0}, 0, -\nu, \mathbf{0}_{1\times(N-1)}, \nu, \mathbf{0}_{1\times(N-1)}]'. \tag{2.27}$$

If $\nu > 0$, this specification ensures that the s.d.f.—that can be seen as the ratio of marginal utilities—goes up when there is an increase in the main common debt shock $\tilde{\eta}_{d,1,t}$ or a decrease in the main common fiscal-limit shock $\tilde{\eta}_{\ell,1,t}$ (see eqs. 2.13 and 2.14).

The resulting model parametrization is given in Table 2.3. Several of the restrictions described above turn out to be binding, which we indicate by "†" in the table. We find, in particular, that the unconditional average of the fiscal limit, namely $\bar{\ell}$, hits the upper bound of 300%. However, as we find that ρ_ℓ is close to one, this implies that ℓ_t almost follows a random walk process and, thus, $\bar{\ell}$ is only weakly identified.

2.5.4 Sovereign spreads fit and credit risk premiums

Figure 2.2 shows the fit of sovereign spreads. The fit is comparable to the one obtained in term-structure studies where default intensities are purely latent and have no macro-finance interpretation.

On Figure 2.2, model-implied spreads (dotted black lines) result from eq. (2.21), which involves formulas using the stochastic discount factor $\mathcal{M}_{t,t+1}$ that itself depends on prices of risk ψ (eq. 2.14). The black solid lines represent the (model-implied) spreads that would be observed if agents were not risk averse; these spreads are obtained by implementing the formulas implicit in eq. (2.21) after having set the prices of risk to zero. The differences between the two types of model-implied spreads correspond to credit-risk premiums. Our results indicate that these risk premiums are sizeable. The ratio between the two types of spreads, which reflects the

[25] This specification for $\psi_{1,t}$ implies that the risk-neutral dynamics of the short-term rate i_t is $i_t = (1 - \rho_i \bar{i} - \sigma_i \psi_{i,0}) + (\rho_i - \sigma_i \psi_{i,1}) i_{t-1} + \sigma_i \eta^*_{i,t}$, where $\eta^*_{i,t} \sim i.i.d. \mathcal{N}(0,1)$ under the risk-neutral measure (see Appendix 2.E).

Table 2.3: Model parametrization

Param.	Value	Param.	Value
ρ_d	0.970	$\rho_{d,\ell}$	0.745
ρ_ℓ	0.997	$\overline{\ell}$	3.000†
ρ_γ	0.700†	ζ	1.500†
ρ_i	0.938	$\sqrt{\mathbb{V}ar(i_t)}$	0.008†
$\rho_i^{\mathbb{Q}}$	0.999†	$maxSR$	0.500†
α	0.083	ν	0.321
$\overline{d}_{DE}$	0.714‡	$\overline{d}_{FR}$	0.938‡
$\overline{d}_{IT}$	1.311‡	$\overline{d}_{ES}$	0.862‡
$\overline{d}_{NL}$	0.591‡	$\overline{d}_{BE}$	1.050‡
$\overline{i}$	27.11‡		

Notes: The subscript † indicates parameters for which the restrictions described in 2.5.3 turn out to be binding in the context of the constrained maximum likelihood estimation. The subscript ‡ indicates parameters that are calibrated: $\overline{d_j}$ is set to the observed sample mean of debt-to-GDP of country j; $\overline{i}$ is set to the sample mean of the short-term rate (3-month OIS rate), it is expressed in basis points (annualized). $maxSR$ and ν determine the vector of prices of risk ψ (see eq. 2.14 and Appendix 2.C). We have the following relationship between ρ_i and $\rho_i^{\mathbb{Q}}$: $\rho_i^{\mathbb{Q}} = \rho_i - \sigma_i \psi_{i,1}$, where $\psi_{i,1}$ is the $(1,1)$ entry of matrix ψ_1 (see Appendix 2.E). Parameter $\rho_{d,\ell}$ is the correlation between the two "main common components" of debts and fiscal limits (see eqs. 2.24 and 2.25). Parameter ζ determines the covariance matrix of fiscal-limit shocks (see eq. 2.23). Parameter α is the elasticity of the probability of default to the fiscal space (see eq. 2.2).

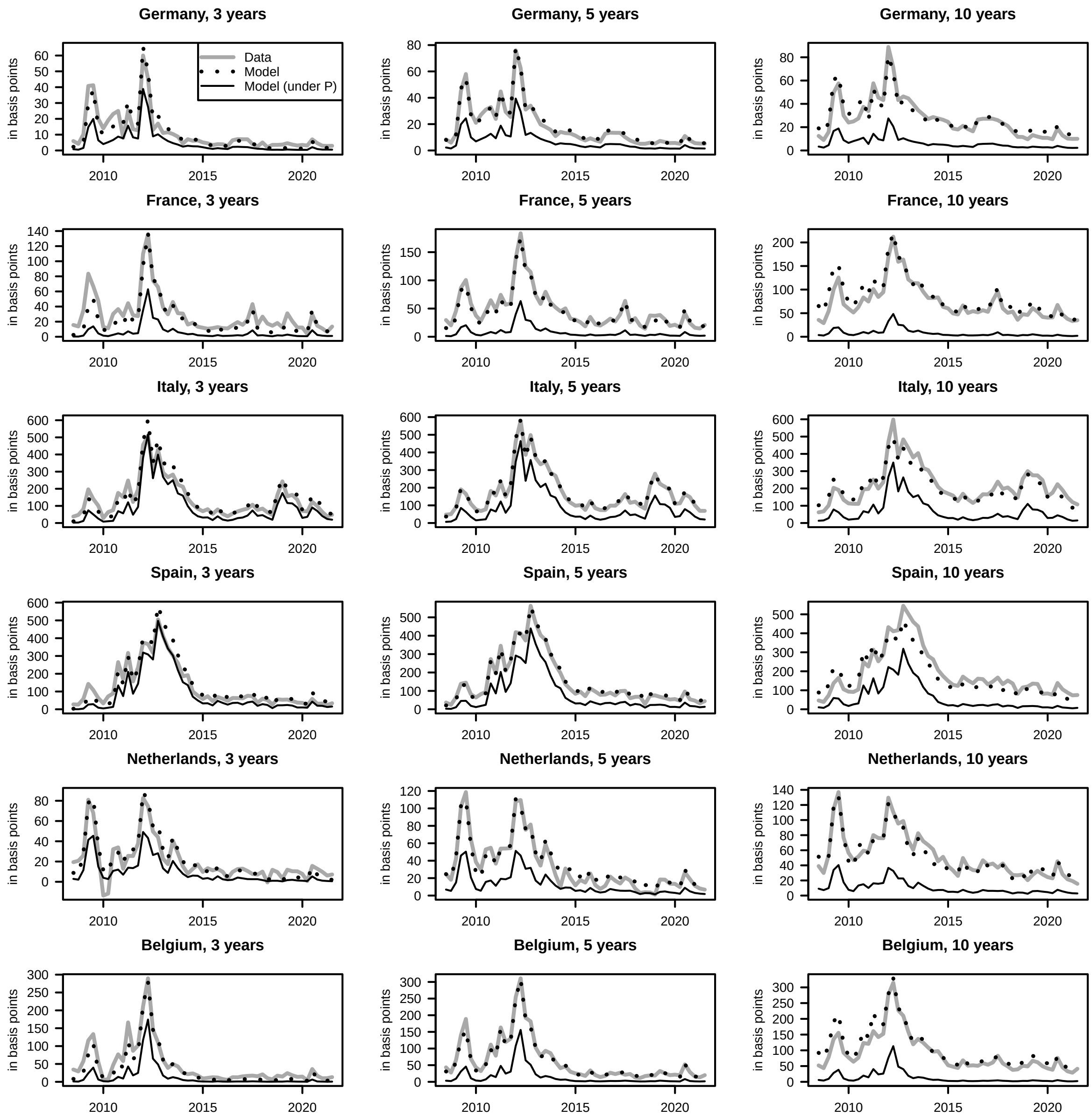
importance of risk premiums, is broadly comparable to the ones found in sovereign credit-risk studies based on reduced-form intensity approaches (e.g. Pan and Singleton, 2008; Longstaff et al., 2011; Monfort and Renne, 2014; Monfort et al., 2020).

Let us stress that, in the present model, credit risk premiums are time-varying even if the prices of risk associated with debt and fiscal-limit shocks are constant. This stems from the conditional heteroskedasticity of the default intensity inherent to our model.[26]

Lastly, Figure 2.3 shows that the model captures a substantial share of the fluctuations of risk-free rates across all maturities.

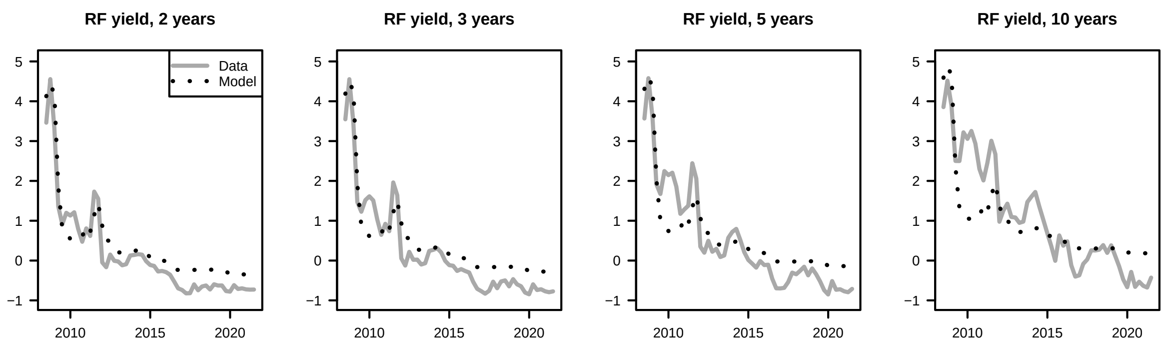
[26] Intuitively, a risk premium can be seen as the product of a price of risk times a risk quantity. Hence, the risk premium is time-varying if at least one of its two multiplicative constituents (prices of risk or the risk quantity) also is. In standard Gaussian affine term-structure models, prices of risk are time-varying but the risk quantity—that is the conditional variance of the factors—is constant. The opposite is true In the present model: prices of risk are constant (except for the short-term risk-free interest rate), but the conditional variance of the default intensity, i.e., $\alpha \max[0, (d_t - \ell_t)]$ (see eq. 2.2) is time-varying (because of the max operator). See Appendix 2.J for additional explanation.

Figure 2.2: Model fit of sovereign bond yield spreads



Notes: Model-implied sovereign spreads result from eq. (2.21). Dashed lines represent the (model-implied) spreads that would be observed if agents were not risk averse (obtained also by eq. 2.21, but after having set the prices of risk, that are the components of ψ, to zero). The differences between the two types of model-implied spreads (dotted and solid lines) correspond to credit-risk premiums.

Figure 2.3: Model fit of risk-free yields



Notes: The model implied risk-free yields (grey solid line) result from eq. (2.18). Interest rates are annualized, and expressed in percentage points.

2.5.5 Fiscal limit estimates

To the best of our knowledge, the present paper is the first to propose time-varying estimates of fiscal limits (together with Pallara and Renne, 2021). These estimates, expressed in percent of GDP, are displayed in Figure 2.4. On a given quarter, if debt-to-GDP ($d_{j,t}$, black solid line) is higher than the fiscal limit ($\ell_{j,t}$, grey solid line), then, the probability of default is strictly positive (see eq. 2.1). Everything else equal, if debt-to-GDP stays above the black dotted line (respectively in the grey-shaded area) for four quarters in a row, then the annual default probability of the considered country would be larger than 10% (respectively in $]0\%, 10\%]$). For what follows, and unless differently specified, our numbers refer to the threshold fiscal limit estimates, namely the grey solid lines in Figure 2.4. According to our estimates, the global financial crisis of 2008 translated into a decrease of the fiscal limits. On average, fiscal limits decreased by 10 percent of GDP between 2008 and 2009.[27] From the beginning of 2010 to early 2012, amid the European sovereign debt crisis, fiscal limits recorded an average decrease close to 20 percent of GDP. Notably, the "*whatever it takes*" statement by Draghi (2012, July) and the European Central Bank (ECB) announcement of the Outright Monetary Transactions (OMT) were followed by a 5 p.p. jump in the average fiscal limit (from 2012Q2 to 2012Q4).[28] After the euro-debt crisis, and until the onset of the COVID-19 pandemic, fiscal limits across countries show an increasing trend on average, translating into a widening of fiscal space in Europe. Fiscal limits decrease

[27] This may be seen as a consequence of transfers from private to public debts through explicit channels (bank bailouts) or implicit ones (debt and deposit guarantees), along the logic of the so-called sovereign-bank nexus (see e.g. Jordà, Schularick, and Taylor, 2016).

[28] The OMT represents a mechanism aimed to "safeguard an appropriate monetary policy transmission and the singleness of the monetary policy" (2012, August).

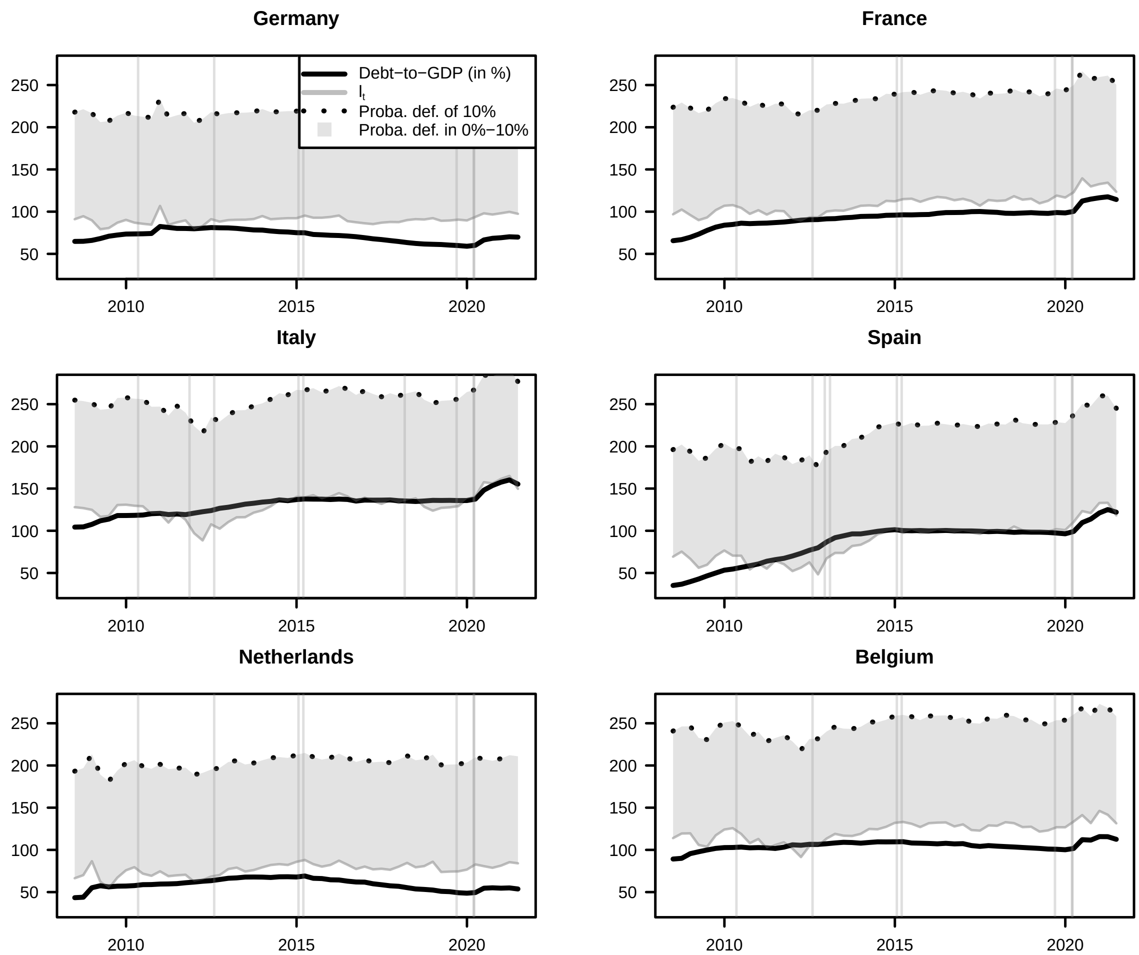
by 5 p.p. on average across countries during the pandemic, from mid-2020 until the end of the estimation sample. In Appendix 2.Q, we report a set of sensitivity analyses based on varying key parameters in the model. Specifically, in Figure 2.Q1 (Appendix 2.Q), we show the fiscal limit estimates across different model specifications.

2.6 Results

2.6.1 Pricing Eurobonds

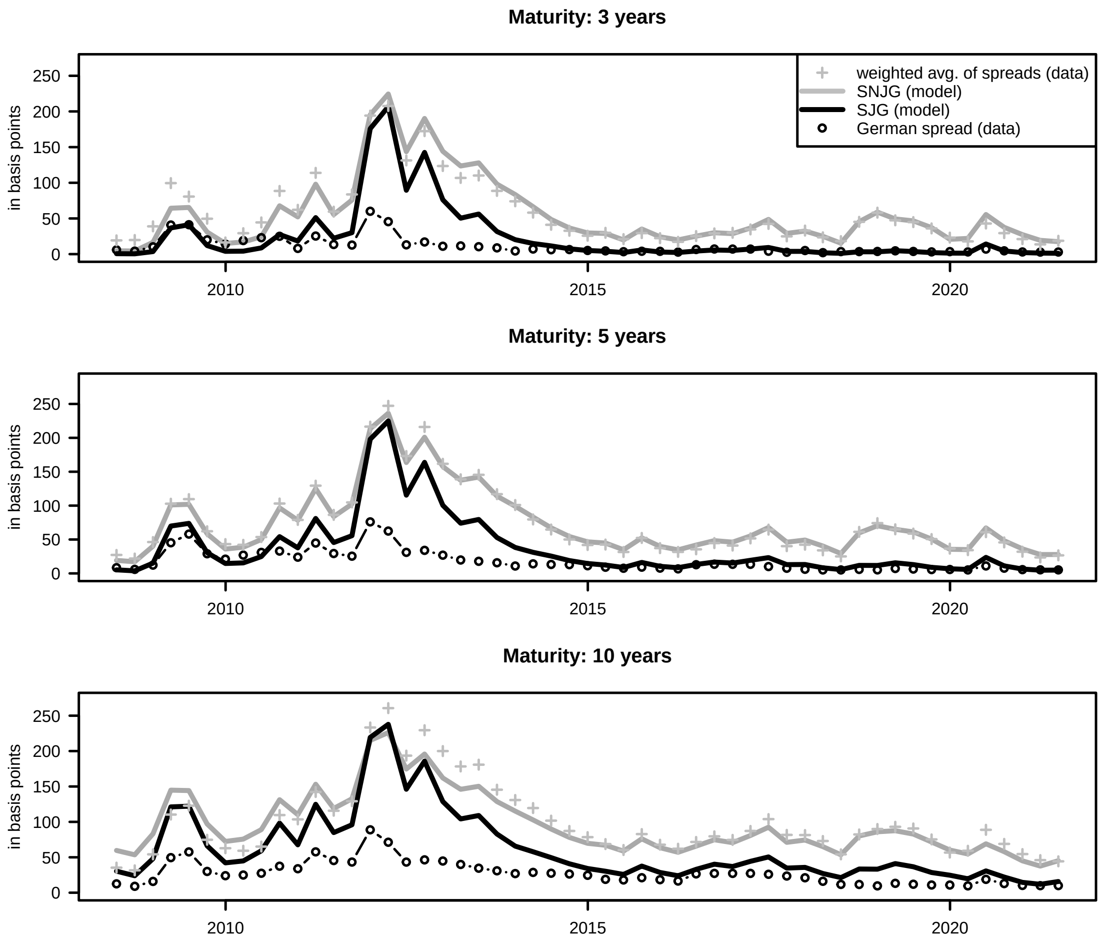
In Figure 2.5, we compare counterfactual yield spreads associated with common bonds benefitting from several and joint guarantees (SJG) and bonds with several but not joint guarantees (SNJG). By design, the latter is close to the debt-weighted average of country-specific observed sovereign spreads. The difference between SNJG and SJG is positive and sizeable across the estimation sample. This result suggests that raising funds through a joint liability debt instrument—the SJG bond—may substantially reduce debt service in the presence of heterogenous fiscal conditions. This is due to the associated diversification of fiscal risks across countries: as long as the fiscal positions across countries are not perfectly correlated, one can expect gains from common bond issuance in the presence of joint and several guarantees (SJG) w.r.t. several but not joint guarantees (SNJG). On average across the estimation sample, the ratio of SNJG bond yield spread on the SJG one is approximately equal to 7, 3.5 and 2 for the 3-, 5- and 10-year maturities, respectively. Over the estimation sample and maturities, the wedge between SJG and SNJG bond yields is equal, on average, to approximately 35 basis points. Notably, for the 10-year maturity, the SJG bond yield spread is higher than the SNJG one by roughly 10 basis points between 2011Q4 and 2012Q1. This implies that aggregate gains would have been slightly negative under issuance with joint and several guarantees during these two quarters. This finding parallels the discussion presented at the end of Section 2.3 concerning the possibility of a reversion of gains arising from SJG bond issuance compared to SNJG one. Given the turmoil jointly faced by European member states during the euro-debt crisis, several debt-to-GDP ratios prove larger than fiscal limits (see Figure 2.4 in Subsection 2.5.5), this leads to detrimental diversification effects. Such effects revert the probability of default that is larger in the SJG bond case compared to the national bond cases causing negative yield gains (see Appendix 2.D for further discussion on the detrimental diversification effects).

Figure 2.4: Fiscal limits



Notes: These plots display estimated fiscal limits ($\ell_{j,t}$) and observed public debts ($d_{j,t}$), both expressed in % of GDP. On a given quarter, if debt-to-GDP is higher than the fiscal limit (grey solid line), then the probability of default is strictly positive (see eq. 2.1). Everything else equal, if debt-to-GDP (black solid line) stayed above the black dotted line (respectively in the grey-shaded area) for four quarters in a row, then the annual default probability of the considered country is larger than 10% (respectively in $]0\%, 10\%]$). On each plot, the vertical bars indicate important dates (monetary-policy decisions and/or noteworthy pivotal economic events): **All countries**—10/05/2010: Announcement of Securities Market Program (SMP); 02/08/2012: ECB announces it may undertake outright transactions in sovereign bond markets (OMT); 22/01/2015: ECB announces expanded asset purchase programme to include bonds issued by euro area central governments, agencies and European institutions (combined monthly asset purchases to amount to €60bn); 04/03/2015: Announcement of the Public Sector Purchase Programme (PSPP); 12/09/2019: Announcement that net purchases will be restarted under the Governing Council's asset purchase programme (APP) at a monthly pace of €20bn as from 1 November 2019. **Italy**—12/11/2011: Berlusconi resigns from office (BTP/Bund spread is over 550 bps); 04/03/2018: Populist parties (M5S and Lega) win the majority of votes in Italian government elections. **Spain**—11/12/2012: ESM (European Stability Mechanism) disburses €39.5bn for recapitalisation of banking sector; 05/03/2013: ESM disburses €1.9bn.

Figure 2.5: Counterfactual bond yield spreads



Notes: This figure compares counterfactual yield spreads (versus risk-free interest rates) associated with common bonds benefitting from several and joint guarantees (SJG) and bonds with several but not joint guarantees (SNJG). For the sake of comparison, we also add German bond yield spreads (circles).

For the sake of comparison, we add the German bond yield spreads in Figure 2.5 (black circles). Interestingly, during the Great Financial Crisis, the baseline SJG spread lays below the German yield spread for the 3-year and 5-year maturity. Hence, diversification effects underlying the SJG bond pricing might, at times, prove beneficial also for fiscally virtuous countries in the euro area—and not only for the peripheral Member States. Notwithstanding, even in the scenarios under which SJG bond yields are higher than Bunds' ones, one can design post-issuance redistribution schemes translating into gains to all countries. This is discussed in Subsections 2.6.2 and 2.6.3.

The magnitudes of our model-implied SJG and SNJG bond spreads are broadly in line with those pertaining to observed proxies of (SJG) Eurobonds. We consider as Eurobond proxies those bonds issued by the following European institutions: the European Investment Bank

(EIB), the European Financial Stability Facility (EFSF), the European Stability Mechanism (ESM), and the European Commission itself, which, against the backdrop of the COVID-19 crisis, has initiated large-scale issuance programs.[29] These bonds benefit from various types of guarantees, which makes them close to SJG bonds.[30] Figure 2.6 shows the spreads between such bonds and the German benchmark bond (the Bund) of equivalent residual maturity. It also displays, in grey, proxies of SNJG spreads, computed as GDP-weighted averages of 10-year national spreads versus the Bund. It appears that the prices of the different SJG Eurobond proxies are close to each other. The red dots indicate the model-implied SJG and SNJG bond spreads (versus Germany). The plot shows that the model captures a substantial amount of the fluctuations of observed spreads.

Our framework also offers the possibility to consider "partial" SNJG and SJG bonds whose emission is circumscribed to a smaller set of countries excluding, for instance, either "super" core member states (Germany and Netherlands) or peripheral ones (Italy and Spain). The results of such counterfactual exercises are reported in Appendix 2.P. The main finding is that the wedge between "partial" SJG and SNJG bonds is smaller compared to the baseline scenario (under which all countries participate in the joint issuance), which reflects enhanced diversification effects in the latter case.

Appendix 2.Q reports the results of analyses where we study the sensitivity of SJG bond yield spreads to changes in several important parameters, or in the way these parameters are constrained within the estimation (see Subsection 2.5.3). The order of magnitude of the spread between SJG and SNJG bonds appears to be robust to these changes.

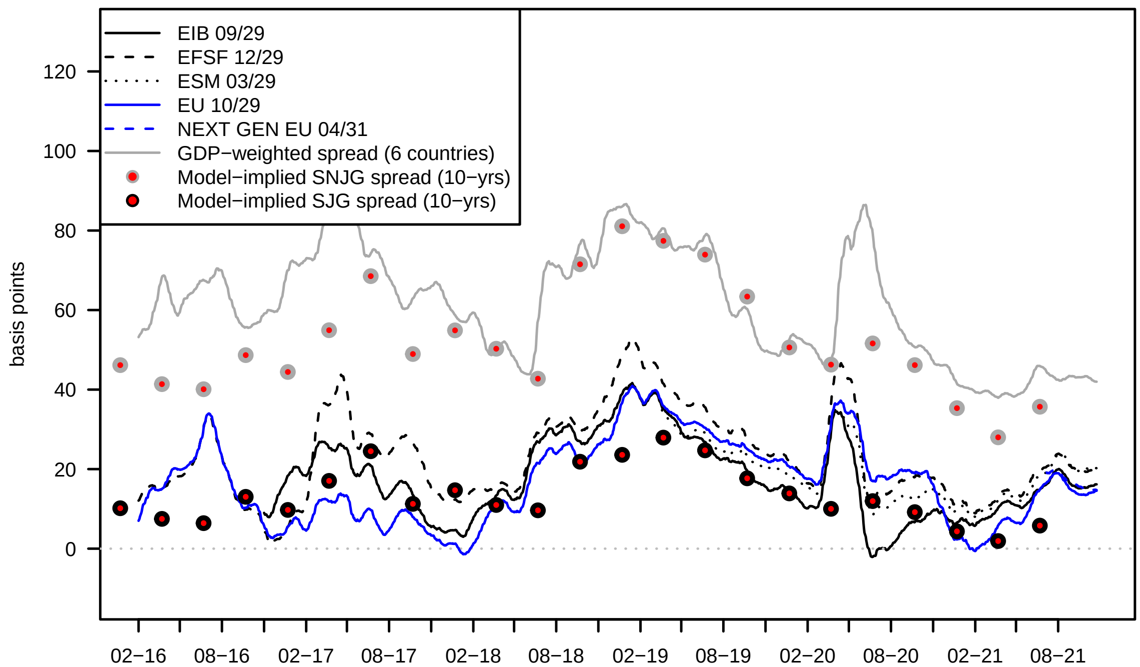
2.6.2 Aggregate gains and redistribution

In Subsection 2.6.1, we have seen that the price of a common debt instrument might be lower than the German one (equivalently, Eurobond yields are higher than Bund ones). However, SJG bond prices are usually higher than SNJG ones. Since the latter correspond to a weighted

[29] These programs notably include the SURE program (for "Support to mitigate Unemployment Risks in an Emergency") and the Next-Generation-EU program. See, e.g., the investor presentation of the European Commission (12 March 2021), available at https://ec.europa.eu/info/sites/default/files/about_the_european_commission/eu_budget/ip_07.2021.pdf. The EU already had issued some bonds before 2020, in particular in the context of the Euratom loans.

[30] To justify Moody's top rating (Aaa) for the EU's bond programs, the rating agency points out, for example, that "the multiple layers of debt service protection, including explicit recourse to extraordinary support [...] creates the equivalent of a joint and several undertaking and obligation on the part of EU member states to provide financial support to the EU" (https://www.moodys.com/research/Moodys-affirms-the-European-Unions-Aaa-rating-outlook-stable--PR_430731).

Figure 2.6: Observed proxies of common bond spreads versus 10-year German benchmark



Notes: This figure shows bond yield spreads w.r.t. the German 10-year benchmark bond. Black and blue (respectively grey) lines correspond to proxies for SJG bonds (resp. SNJG bonds). We consider bonds issued by the European Investment Bank (EIB), the European Financial Stability Facility (EFSF), the European Stability Mechanism (ESM), the European Union (EU, NEXT GEN EU). The SNJG proxy (grey lines) is computed as a GDP-weighted average of national-bond spreads (versus Germany). The data are at the daily frequency (20-day moving averages); they span the period from February, 9 2016 to November 1, 2021. The dates reported in the legend of the figure correspond to maturity dates (2029 or 2031) of the specific bonds. The spreads are computed as the differences in asset swap spreads w.r.t. to the Bund; (see Appendix 2.N for more details). As of November 2021, the credit ratings of the considered European institutions were as follows (Moody's/S&P/Fitch): EIB (Aaa/AAA/AAA), EFSF (Aa1/AA/AA+), ESM (Aa1, AAA/AAA), and EU (Aaa/AA/AAA).

average of national bond prices, replacing national bonds with SJG bonds results in aggregate gains. These gains could be redistributed ex-post—i.e. after issuance—across all countries. In that case, and considering only strictly positive redistribution weights, the issuance of SJG bonds would eventually result in a reduction in funding costs for all countries (w.r.t. the issuance of national bonds).[31] Naturally, the number of redistribution schemes is infinite. In this subsection, we focus on three situations. In the first one (Scheme A), countries pay the same yield (i.e., there is no redistribution); in the second one (Scheme B), gains are distributed in proportion to GDP; in the third one (Scheme C), gains are distributed in such a way that the interest rate reduction—relative to the respective national bond rates—is the same for all countries. Formulas used to perform these exercises are detailed in Appendix 2.L.[32]

Table 2.4 shows the results of these counterfactual exercises. We focus on 5-year bonds (5 years roughly being to the average issuance maturity in the euro area), and three periods: beginning of the estimation sample (2008Q2), midst of the euro-debt crisis (2011Q4), and end of the estimation sample (2021Q2). The three upper panels (A, B and C) of Table 2.4 correspond to the three SJG-based schemes described above. For the sake of comparison, the lower panel (Panel D) shows results for the SNJG case, for which there are no aggregated gains. For this latter case (Scheme D), we consider only the situation in which all countries pay the same interest rate (i.e. the SNJG issuance yield). Table 2.4 also reports post-redistribution yields, which are the differences between national bond yields and reductions in the funding costs (or "yield gains") resulting from the considered schemes. In addition, we show redistribution weights; these weights indicate how aggregate gains are shared across countries.

Let us stress that the reported reduction in the funding cost (or yield gain) pertains to one given bond, and not to the whole debt outstanding. To be sure: a yield gain of 100 basis points (say) would effectively translate into a reduction of yearly aggregate funding costs of €1bn if an outstanding amount of €100bn of SJG bonds was issued. This being said, to give an idea of the amounts potentially involved, the top part of Table 2.4 indicates the aggregate gains that would have resulted from the issuance of the equivalent of approximately 5% of the euro-area GDP (€500bn) during the three considered quarters. For instance, for the same face value (€500bn), issuing SJG bonds instead of SNJG bonds in 2008Q2 would have increased the issuance proceeds by €2.78bn. For 2011Q4 and 2021Q2, the gains would have been €3.3bn and €5.73bn, respectively.

[31] In some sense, any scheme involving strictly positive weights can be seen as Pareto-improving.

[32] This Appendix also reports results of schemes where the funding costs of Germany and France are left unchanged (see Appendix 2.L.5).

Panel A of Table 2.4 characterizes the scheme where there is no redistribution of the aggregated gains (Scheme A). As illustrated by our results, this scheme can result in negative "gains" for some countries: funding costs for Germany, France and the Netherlands get considerably higher in 2011Q4. Italy and Spain are the countries that benefit the most out of the SJG issuance scheme in 2011Q4: the spread between post-redistribution and national yields is equal to 280 basis points for Italy and 224 basis points for Spain.

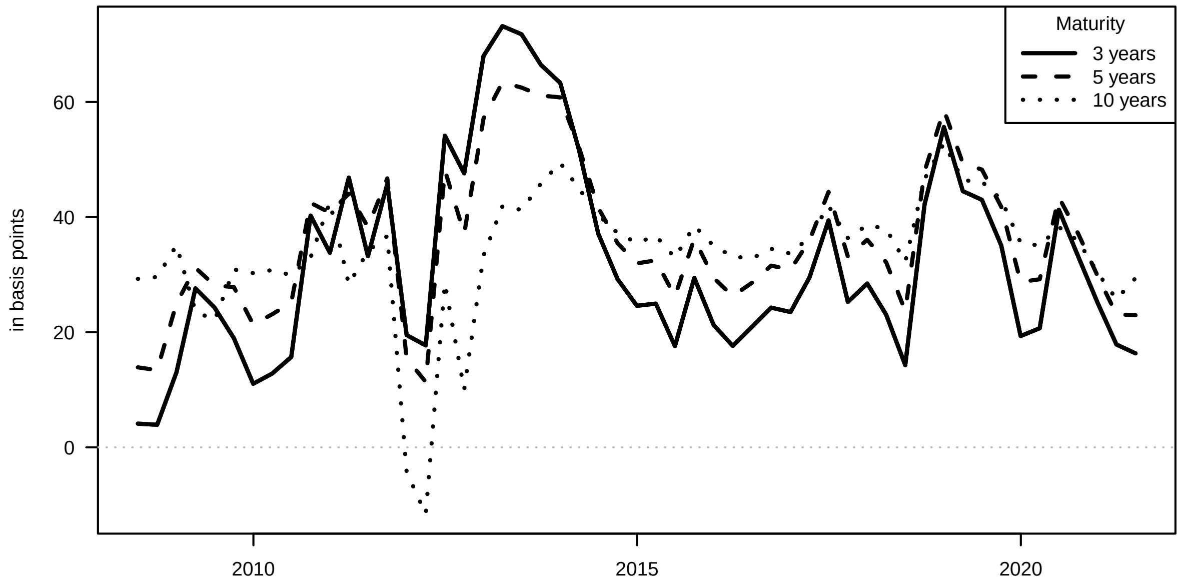
By contrast, Schemes B and C are such that all countries mechanically benefit from the issuance of SJG bonds. These two schemes deliver similar results (see Panels B and C of Table 2.4). While yield reductions are modest before and during the sovereign debt crisis period (around 15 basis points), they become more sizeable at the end of the estimation sample (about 25 basis points in 2021Q2).

Figure 2.7 displays the time series of yield gains associated with Scheme C. We consider three maturities: 3, 5 and 10 years. For the 3- and 5-year maturity, yield gains peak at the end of the euro-debt crisis, between 2012Q4 and 2013Q1, reaching approximately 75 and 65 basis points, respectively. As regards the 10-year maturity, yield gains associated with scheme C revolve around 35 basis points before and after the euro-debt crisis, while, between 2011Q4 and 2012Q1, they turn out to be negative (around -10 basis points). For further details on this finding concerning negative yield gains, we reference to the previous subsection and the discussion at the end of Subsection 2.3.

In Appendix 2.Q (Figure 2.Q3), we show the yield gains associated with Scheme C across different sensitivity exercises. The order of magnitude of these gains appears to be fairly robust to the considered changes in the model parametrization.

It is important to mention that our results do not take into account potential higher-order effects. The mechanisms underlying such effects would be as follows: if the average funding cost of a government decreases—because part of its funding needs are met with Eurobonds—then expected future debt would decrease because of lower debt service. (For this to hold, one has to assume, however, that the decrease in future debt service will not be compensated by higher primary deficits.) If agents effectively expect lower future debt levels, then bond prices move. That is, the initial funding cost effects are followed by second-order ones. This, in turn, reduces future debt service, and so on. This issue is complicated to handle, especially in the context of a reduced-form approach like ours. Nevertheless, in Appendix 2.M, we propose an iterative approach aimed at gauging the potential impacts of such higher-order effects. For

Figure 2.7: Yield gains associated with redistribution scheme with same yield gains across countries



Notes: This figure shows yield gains associated with redistribution Scheme C (same yield gains across countries) throughout the whole estimation sample and for different maturities. See Subsection 2.6.2 for details regarding this redistribution scheme. Yield gains are expressed in basis points.

moderate levels of SJG bond issuance—we consider, therein, that 20% of the euro-area debt is issued in the form of SJG bonds—our results point to relatively mild higher-order effects.

2.6.3 Moral hazard and redistribution schemes

Usual concerns associated with common debt issuance pertain to moral hazard (see, e.g., Claessens et al., 2012; Favero and Missale, 2012; Tirole, 2015; Dávila and Weymuller, 2016): knowing that part of their debt is guaranteed by other countries, some countries may be tempted to increase their spending—and start issuing more debt—since the interest rate on jointly-guaranteed debt is less sensitive to an individual debt increase than non-guaranteed debt.

Although our reduced-form modeling framework does not allow to explore such mechanisms in a structural way, it can illustrate how market discipline would be impaired by massive issuance of SJG bonds. Specifically, we perform counterfactual exercises in which Italy and Spain decide to deviate from their current debt level, all else being equal. We then observe the changes in spreads induced by these modifications. We consider two dates: 2011*Q*4 (euro-area debt crisis) and 2021*Q*2 (end of the estimation sample). Figure 2.8 shows the results. For each date and each country, large increases in the debt-to-GDP ratio result in modest increases in

Table 2.4: Effect of redistribution schemes on funding costs

	2008-06-30			2011-12-31			2021-06-30		
				SJG					
A.G.[a]	€2.78 bn			€3.3 bn			€5.73 bn		
Panel A: SJG, Same funding costs (i.e. no ex-post redistribution)									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	7%	436	3	-262%	306	-119	1%	-10	1
FR	18%	436	10	-99%	306	-62	16%	-10	15
IT	43%	436	31	325%	306	280	55%	-10	67
ES	14%	436	16	165%	306	224	25%	-10	47
NL	11%	436	19	-44%	306	-87	1%	-10	3
BE	8%	436	26	15%	306	52	2%	-10	13
Panel B: SJG, Redistribution based on GDP weights									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	33%	425	14	33%	172	14	33%	-32	23
FR	24%	433	14	24%	228	15	24%	-18	23
IT	19%	453	14	19%	568	18	19%	34	23
ES	12%	438	14	12%	512	17	12%	14	23
NL	8%	442	14	8%	204	15	8%	-29	23
BE	4%	448	14	4%	342	16	4%	-19	23
Panel C: SJG, Same yield gains across countries									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	33%	425	14	35%	171	15	33%	-32	23
FR	24%	433	14	25%	228	15	24%	-18	23
IT	19%	453	14	17%	570	15	19%	35	23
ES	12%	438	14	11%	514	15	12%	15	23
NL	8%	442	14	8%	203	15	8%	-29	23
BE	4%	448	14	4%	342	15	4%	-20	23
				SNJG					
A.G.	€0 bn			€0 bn			€0 bn		
Panel D: SNJG, Same funding costs									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	–	450	-11	–	321	-135	–	13	-22
FR	–	450	-4	–	321	-78	–	13	-8
IT	–	450	17	–	321	264	–	13	44
ES	–	450	2	–	321	208	–	13	24
NL	–	450	5	–	321	-103	–	13	-20
BE	–	450	12	–	321	36	–	13	-10

[a]: Aggregate gains. *Notes*: This table compares post-redistribution funding costs across countries under the two issuance schemes (SJG and SNJG) and under different redistribution schemes described in Subsection 2.6.2. We focus on the 5-year maturity and on three periods: beginning of the estimation sample (2008Q2), midst of the euro-debt crisis (2011Q4) and end of the estimation sample (2021Q2). Yields are expressed in basis points. Aggregate gains (reported at the top of the table) are computed under the assumption that total issuance is equal to 5% of aggregate GDP. In each panel, for all countries and dates, we show the redistribution weights, the post-redistribution yields, and the spread between national yields and the post-redistribution yields (that are the yield gains). Under SNJG (Panel D), redistribution weights are unnecessary since there are no aggregated gains. See Appendix 2.L for computational details.

SJG and SNJG Eurobond spreads (see, respectively, the grey and black solid lines).[33] These increases are far lower than those of national bond yields (grey dashed line). This illustrates the moral hazard issue: under the issuance of common bonds, and if each country pays the issuance SJG/SNJG yield (i.e., under Schemes A or D), then the ability of financial markets to restore fiscal discipline via rising interest rates is hampered. Let us stress that the strength of this hampering effect depends on the extent to which national issuances would be replaced with eurobonds: as long as a sizable share of countries' funding needs are met with the issuance of national bonds, the overall debt service remains sensitive to countries' indebtedness. In other words, under Schemes A or D, a necessary condition for market discipline to remain effective is to limit the issuance of eurobonds (as suggested by von Weizsäcker and Delpla, 2010; Hellwig and Philippon, 2011). The simulation results suggest that moral hazard effects are dampened under Schemes B and C (see black dashed lines in Figure 2.8); these schemes indeed imply that post-redistribution funding costs remain sensitive to countries' indebtedness showing a similar slope as national bond yields (grey dashed line). Moreover, these post-redistribution yields remain lower than national bond yields as long as aggregate gains are positive.[34]

2.7 Concluding remarks

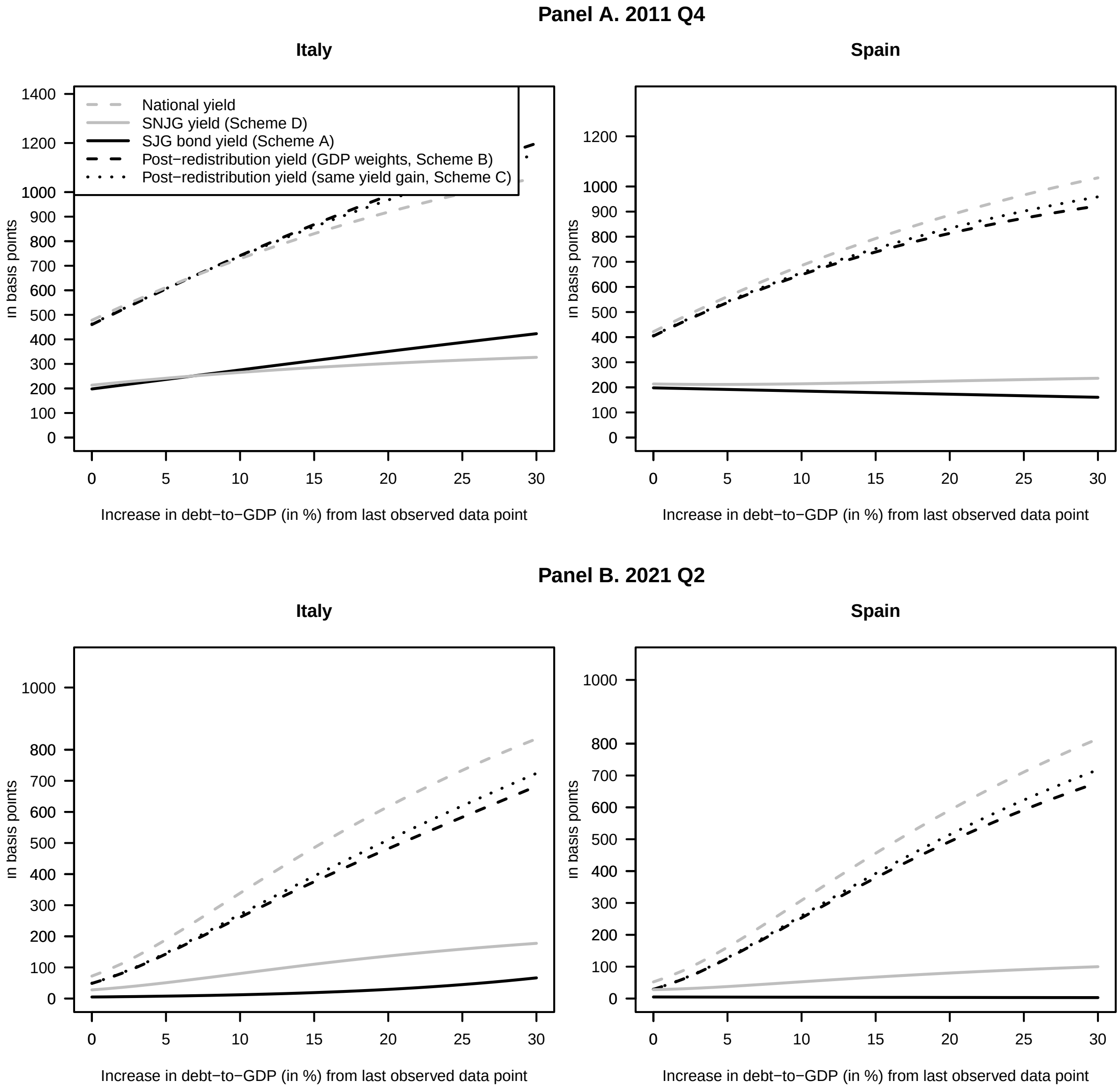
This paper aims at pricing bonds jointly issued by a group of countries. Our focus is on Eurobonds, which are debt instruments jointly issued by euro-area countries. We consider two types of common bonds: the first features joint and several guarantees (SJG bond); the second is characterized by several but not joint guarantees (SNJG bond). To price these two types of common bonds, we develop a novel multi-country sovereign credit risk framework. Our model captures the joint dynamics of national bond prices, sovereign debt, and the fiscal limit—the level of debt beyond which the risk of default is no longer zero.

Estimating the model involves both determining the model parameterization and countries' fiscal limits. Thanks to the tractability of our asset-pricing framework, these two tasks are oper-

[33]The SJG bond yield proves higher than the SNJG one under a sizeable rise in Italian indebtedness in 2011Q4 (top left panel of Figure 2.8). This stems from the fact that diversification effects become detrimental when expected joint fiscal limits are overcome by pooled debts (see discussion at the end of Section 2.3 and, also, Subsection 2.6.1). In this scenario, the probability of default is reversed (larger for SJG than for national bonds, on average), causing negative yield gains (see Appendix 2.D for further details on detrimental diversification effects).

[34]Post-redistribution yields under Scheme B and C are above the Italian national bond yield in 2011Q4 when Italian debt-to-GDP ratio considerably grows (top left panel in Figure 2.8). As mentioned in Footnote 33, this finding of negative aggregate yield gains arise from the reversal of diversification effects under periods of particular turmoil.

Figure 2.8: Moral hazard risk and redistribution: counterfactual exercise



Notes: This figure shows the increase in different bond spreads (w.r.t. to risk-free rates) resulting from counterfactual increases in Italian indebtedness (left column of plots) or Spanish indebtedness (right column of plots), all else being equal. The two rows correspond to different periods, namely 2011*Q*4 (euro-area sovereign debt crisis) and 2021*Q*2 (end of the estimation sample). The different schemes (A to D) are described in Subsection 2.6.2.

ated jointly. Our estimation sample comprises data associated with the sixth largest euro-area economies over the period 2008-2021. The estimated model fits observed sovereign spreads across maturities and countries. To the best of our knowledge, this paper is the first to provide time-varying estimates of fiscal limits for the euro area.

The estimated model is exploited to examine the pricing of (counterfactual) SJG and SNJG bonds. In most instances, yields associated with SNJG bonds are higher than those associated with their SJG equivalents. Notably, across the estimation sample and maturities, the SNJG bond yield spread, w.r.t. a risk-free rate, is three times larger than the SJG one. Interestingly, our model shows also that aggregate gains associated with SJG bond issuance might considerably decrease when expected fiscal spaces reduce at the euro-area scale, up to potential inversion. Therefore, in the presence of heterogeneous and not too adverse fiscal conditions, raising funds through SJG bonds may lower aggregate debt service (w.r.t. situations where only national bonds and/or SNJG bonds are issued). We discuss potential ex-post redistributions of such aggregate gains, and we show that some of these redistribution schemes may alleviate the reduction in market discipline resulting from joint bond issuances.

Appendix

2.A X_t's dynamics

Denote by $\bar{d}$ and $\bar{\ell}$ the unconditional means of vectors d_t and ℓ_t, respectively. The state vector $X_t = [i_t, i_{t-1}, 0, d_t', d_{t-1}', \ell_t']'$, follows the vector autoregressive process of order one given in eq. (2.12), with:

$$\mu = \begin{bmatrix} (1-\rho_i)\bar{i} \\ 0 \\ (1-\rho_d)\bar{d} \\ \mathbf{0}_{N\times 1} \\ (1-\rho_\ell)\bar{\ell} \end{bmatrix}, \; \Phi = \begin{bmatrix} \rho_i & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & \rho_d \mathbf{Id}_{N\times N} & 0 & 0 \\ 0 & 0 & \mathbf{Id}_{N\times N} & 0 & 0 \\ 0 & 0 & 0 & 0 & \rho_\ell \mathbf{Id}_{N\times N} \end{bmatrix}, \; \Sigma\Sigma' = \begin{bmatrix} \sigma_i^2 & 0 & \cdots & & \\ 0 & 0 & \cdots & & \\ \vdots & \vdots & \Omega_d & \mathbf{0}_{N\times N} & \Omega_{\ell,d}' \\ & & \mathbf{0}_{N\times N} & \mathbf{0}_{N\times N} & \mathbf{0}_{N\times N} \\ & & \Omega_{\ell,d} & \mathbf{0}_{N\times N} & \Omega_\ell \end{bmatrix}.$$

Let us detail the parametrization of matrices Ω_d, Ω_ℓ, and $\Omega_{\ell,d}$. The structure exposed in Subsection 2.5.3 implies that we have, for shocks $\varepsilon_{d,t}$ and $\varepsilon_{\ell,t}$ (appearing in eqs. 2.8 and 2.10):[35]

$$\begin{aligned} \varepsilon_{d,t} &= \Gamma_d \widetilde{\Gamma}_d \tilde{\eta}_t && \text{(a.2.1)} \\ \varepsilon_{\ell,t} &= \Gamma_\ell \widetilde{\Gamma}_\ell \tilde{\eta}_t, && \text{(a.2.2)} \end{aligned}$$

where (i) $\tilde{\eta}_t = [\tilde{\eta}_{d,\ell,t}, \tilde{\eta}_{d,t}', \tilde{\eta}_{\ell,t}']' \sim i.i.d.\mathcal{N}(0, I)$ (hence, the dimension of $\tilde{\eta}_t$ is $(1+2N)\times 1$) , (ii) Γ_d and Γ_ℓ are based on PCA (see Subsection 2.5.3), and (iii) with

$$\begin{aligned} \widetilde{\Gamma}_d &= \begin{bmatrix} \sqrt{\rho_{d,\ell}} & \sqrt{1-\rho_{d,\ell}} & 0_{1\times(N-1)} & 0_{1\times N} \\ 0_{(N-1)\times 1} & 0_{(N-1)\times 1} & I_{N-1} & 0_{(N-1)\times N} \end{bmatrix} && \text{(a.2.3)} \\ \widetilde{\Gamma}_\ell &= \begin{bmatrix} \sqrt{\rho_{d,\ell}} & 0_{1\times N} & \sqrt{1-\rho_{d,\ell}} & 0_{1\times(N-1)} \\ 0_{(N-1)\times 1} & 0_{(N-1)\times N} & 0_{(N-1)\times 1} & I_{N-1} \end{bmatrix}. && \text{(a.2.4)} \end{aligned}$$

With these notations, we have:

$$\Sigma = \begin{bmatrix} \sigma_i & \mathbf{0}_{1\times(2N+1)} \\ 0 & \mathbf{0}_{1\times(2N+1)} \\ \mathbf{0}_{N\times 1} & \Gamma_d\widetilde{\Gamma}_d \\ \mathbf{0}_{N\times 1} & \mathbf{0}_{N\times(2N+1)} \\ \mathbf{0}_{N\times 1} & \Gamma_\ell\widetilde{\Gamma}_\ell \end{bmatrix}.$$

Therefore, noting that $\widetilde{\Gamma}_\ell\widetilde{\Gamma}_\ell' = \widetilde{\Gamma}_d\widetilde{\Gamma}_d' = I$, we have $\Omega_d = \Gamma_d\Gamma_d'$, $\Omega_\ell = \Gamma_\ell\Gamma_\ell'$, and $\Omega_{\ell,d} = \Gamma_\ell\widetilde{\Gamma}_\ell(\Gamma_d\widetilde{\Gamma}_d)'$.

[35] Note that $\eta_t = [\eta_{i,t}, \tilde{\eta}_t']'$ (see eq. (2.26)).

2.B Pricing of zero-coupon zero-recovery risky bonds

Denote by $P^{(j)}_{t,h}$ the date-t price of a zero-coupon bond providing a payoff of 1 on date $t+h$ if country j has not defaulted before $t+h$, and zero otherwise. We have:

$$
\begin{aligned}
P^{(j)}_{t,h} &= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) = \mathbb{E}^{\mathbb{Q}}_t\left\{\exp(-i_t-\cdots-i_{t+h-1})(1-\mathcal{D}_{j,t+h})\right\}\\
&= \mathbb{E}^{\mathbb{Q}}_t\left\{\mathbb{E}^{\mathbb{Q}}_t\left\{\exp(-i_t-\cdots-i_{t+h-1})(1-\mathcal{D}_{j,t+h})|X_{t+h},X_{t+h-1},\dots\right\}\right\}\\
&= \mathbb{E}^{\mathbb{Q}}_t\left\{\exp(-i_t-\cdots-i_{t+h-1}-\underline{\lambda}_{j,t+1}-\cdots-\underline{\lambda}_{j,t+h})\right\}, \qquad \text{(a.2.5)}
\end{aligned}
$$

where the last equality is obtained under the assumption that $\mathcal{D}_t$ does not cause X_t in the Granger's or Sims' sense (Monfort and Renne, 2013, Proposition 3). Note here that the risk-neutral dynamics of X_t ($\mathbb{Q}$) is easily deduced from the physical one, characterized by eq. 2.12 (eq. a.2.10 in Appendix 2.E).

Because the default intensities $\underline{\lambda}_{j,t}$ involve a max operator (see eq. 2.2), eq. (a.2.5) does not admit closed-form solutions. We follow Wu and Xia (2016) and look for an approximation for the following "forward" rate:

$$
p_{j,h-1,h} = -\log(P^{(j)}_{t,h}) + \log(P^{(j)}_{t,h-1}). \qquad \text{(a.2.6)}
$$

Then, we get an approximation to $P^{(j)}_{t,h}$ by taking the exponential of the cumulated forward rates. The approximation is essentially based on $\log\mathbb{E}[\exp(Z)] \approx \mathbb{E}(Z)+\frac{1}{2}\mathbb{V}(Z)$, which is exact when Z is Gaussian, but not if it is truncated Gaussian, as is the case here.

As detailed in Appendix 2.G, we get:

$$
\begin{aligned}
p_{j,k-1,k} \approx\; & \delta'\mu^{\mathbb{Q}}_{t,k} + \Phi\left(\frac{\mu^{\mathbb{Q}}_{j,t,k}}{\sigma^{\mathbb{Q}}_{j,k}}\right)\mu^{\mathbb{Q}}_{j,t,k} + \phi\left(-\frac{\mu^{\mathbb{Q}}_{j,t,k}}{\sigma^{\mathbb{Q}}_{j,k}}\right)\sigma^{\mathbb{Q}}_{j,k} - \frac{1}{2}\left(\mathfrak{q}_{j,t,k}(\delta+\delta_j)'\Gamma^{\mathbb{Q}}_{k,0}(\delta+a_j) + (1-\mathfrak{q}_{j,t,k})\delta'\Gamma^{\mathbb{Q}}_{k,0}\delta\right)\\
& - \sum_{i=1}^{k-1}\left(\mathfrak{q}_{j,t,k-i}(\delta+\delta_j)'\Gamma^{\mathbb{Q}}_{k,i}(\delta+\delta_j) + (1-\mathfrak{q}_{j,t,k-i})\delta'\Gamma^{\mathbb{Q}}_{k,i}\delta\right), \qquad \text{(a.2.7)}
\end{aligned}
$$

where $\delta = [0,1,0,\dots]'$ (in such a way that $i_{t-1} = \delta' X_t$), and where $\delta_j = [0,0,\alpha e_j', \mathbf{0}_{1\times N}, -\alpha e_j']'$ (e_j denoting the j^{th} column vector of the $N\times N$ identity matrix), $\mathfrak{q}_{j,t,k} = \Phi\left(\mu^{\mathbb{Q}}_{t,k}/\sigma^{\mathbb{Q}}_{j,k}\right)$, and

$$
\left\{
\begin{array}{lllll}
\mu^{\mathbb{Q}}_{t,k} &=& \mathbb{E}^{\mathbb{Q}}_t(X_{t+k}) &=& (Id-\Phi^{\mathbb{Q}})^{-1}(Id-\Phi^{\mathbb{Q}^k})\mu^{\mathbb{Q}} + \Phi^{\mathbb{Q}^k}X_t,\\
\Gamma^{\mathbb{Q}}_{k,0} &=& \mathbb{V}^{\mathbb{Q}}_t(X_{t+k}) &=& \Omega + \Phi^{\mathbb{Q}}\Gamma^{\mathbb{Q}}_{k-1,0}\Phi^{\mathbb{Q}'}, \quad \text{with} \quad \Gamma^{\mathbb{Q}}_{1,0} = \Omega\\
&&&=& \Omega + \Phi^{\mathbb{Q}}\Omega\Phi^{\mathbb{Q}'} + \cdots + \Phi^{\mathbb{Q}^{k-1}}\Omega\Phi^{\mathbb{Q}^{k-1}\prime},\\
\Gamma^{\mathbb{Q}}_{k,i} &=& \mathbb{C}ov^{\mathbb{Q}}_t(X_{t+k},X_{t+k-i}) &=& \Phi^{\mathbb{Q}^i}\Gamma^{\mathbb{Q}}_{k-i,0} \quad \text{if} \quad k-i>0,
\end{array}
\right.
$$

where $\mu^{\mathbb{Q}} = \mu - \Sigma\psi_0$ and $\Phi^{\mathbb{Q}} = \Phi - \Sigma\psi_1$ (see Appendix 2.E).

2.C Maximum Sharpe Ratio

The maximum Sharpe ratio for a one-period investment is given by (Hansen and Jagannathan, 1991):

$$maxSR_t = \frac{\sqrt{\mathbb{V}ar_t(\mathcal{M}_{t,t+1})}}{\mathbb{E}_t(\mathcal{M}_{t,t+1})}.$$

In the present context, the exponential affine form of our s.d.f. (2.13) implies that:

$$maxSR_t = \frac{\sqrt{\mathbb{V}ar_t \exp(-\psi'\varepsilon_{t+1})}}{\mathbb{E}_t \exp(-\psi'\varepsilon_{t+1})} = \sqrt{\exp(\psi_t'\psi_t) - 1}.$$

Since ψ_1 is a matrix of zeros, except the (1,1) entry that is equal to $\psi_{i,1}$, and using the specification of ψ_0 given in eq. (2.27), we obtain $maxSR_t = \sqrt{\exp([\psi_{i,0} + \psi_{i,1}i_t]^2 + 2\nu^2) - 1}$. Denoting by $maxSR$ the value of $maxSR_t$ obtained when the state vector is at its average value, we get:

$$maxSR = \sqrt{\exp([\psi_{i,0} + \psi_{i,1}\bar{i}]^2 + 2\nu^2) - 1}. \tag{a.2.8}$$

The short-term interest rate i_t remained constant for most of our estimation sample (from 2012 to 2021), it was therefore often close to $\bar{i}$ ($\approx$ 0.3%). As a result, we have $maxSR_t \approx maxSR$ for most of the sample.

When calibrating the model, it is convenient to set constraints on $maxSR$, rather than on $\psi_{i,0}$, say, because the literature provides us with priors regarding $maxSR$. Accordingly, we choose to put $maxSR$ among the parameters and to use (a.2.8) to get $\psi_{i,0}$ (that is therefore removed from the list of degrees of freedom). Specifically, (a.2.8) gives:

$$\psi_{i,0} = -\psi_{i,1}\bar{i} \pm \sqrt{\log(maxSR^2 + 1) - 2\nu^2}.$$

We keep the solution that gives a positive average slope of the risk-free yield curve (that is the one for which $\pm$ is replaced by the minus sign in the previous expression).

2.D Negative expected fiscal spaces in the stylized model

This appendix discusses the effects of negative fiscal spaces on the prices of jointly-issued bonds in our framework. For that, we use the stylized situation described in Section 2.3. For the sake of simplicity, we focus on the symmetrical situation (Countries A and B are alike). Moreover, we set to zero the correlation between the fiscal spaces of A and B, i.e., $\rho = 0$ in eq. 2.4. (Mechanisms are more evident in this case.)

The three panels of Figure 2.D1 correspond to three situations. In the top panel, expected fiscal spaces are positive; expected fiscal spaces are null in the middle panel; they are negative in the third one. In all three cases, the distribution associated with the joint area (blue line) is narrower than the national ones (red line). But the implications of these diversification effects, in terms of default intensities, are different. The distributions of the default intensities are represented by the shaded areas: bluish for the joint area (or SJG bond) and reddish for the single countries (or, approximately, SNJG bonds). Note that, in addition to these shaded areas, the distributions of the default intensities also include (unrepresented) Dirac masses located at zero.

In the first situation (top panel), we see that the default probability is far smaller for the SJG bond (bluish area) than for the national bonds (reddish area). In the second case, where the average fiscal spaces are zero, we see that diversification effects are still at play: the bluish area is more concentrated towards zero. Finally, the third plot shows that when debts are large enough compared to fiscal limits, then diversification effects are reversed: the probability of default is larger in the SJG bond case than in the national bond cases (the reddish area is relatively more concentrated towards zero).

This is further illustrated by Figure 2.D2, that shows how the yields of SJG and SNJG bonds behave when the average debt is larger than the average fiscal limit. In the left-hand side panel of the figure, we plot SJG and SNJG bond yields when negative fiscal space (% of GDP) varies between -1% and -40%. The right-hand side panel of the figure shows the difference between the two yields. It appears that when fiscal space is large and negative, the spread between SJG and SNJG bond yields turns positive.

2.E $\mathbb{P}$ to $\mathbb{Q}$ dynamics

The risk-neutral measure is defined with respect to the physical measure through the following Radon-Nikodym derivative:

$$\left.\frac{d\mathbb{Q}}{d\mathbb{P}}\right|_{t,t+1} = \frac{\mathcal{M}_{t,t+1}}{\mathbb{E}_t(\mathcal{M}_{t,t+1})} = \exp\left(-\frac{1}{2}\psi_t'\psi_t - \psi_t'\eta_{t+1}\right),$$

where the vector of prices of risk ψ_t is given in eq. (2.15). Under the physical measure, the conditional Laplace transform of X_t is given by:

$$\mathbb{E}_t(\exp(u'X_{t+1})) = \exp\left(u'\mu + u'\Phi X_t + \frac{1}{2}u'\Sigma\Sigma'u\right). \tag{a.2.9}$$

Let us now compute the conditional Laplace transform of X_t under the risk-neutral measure:

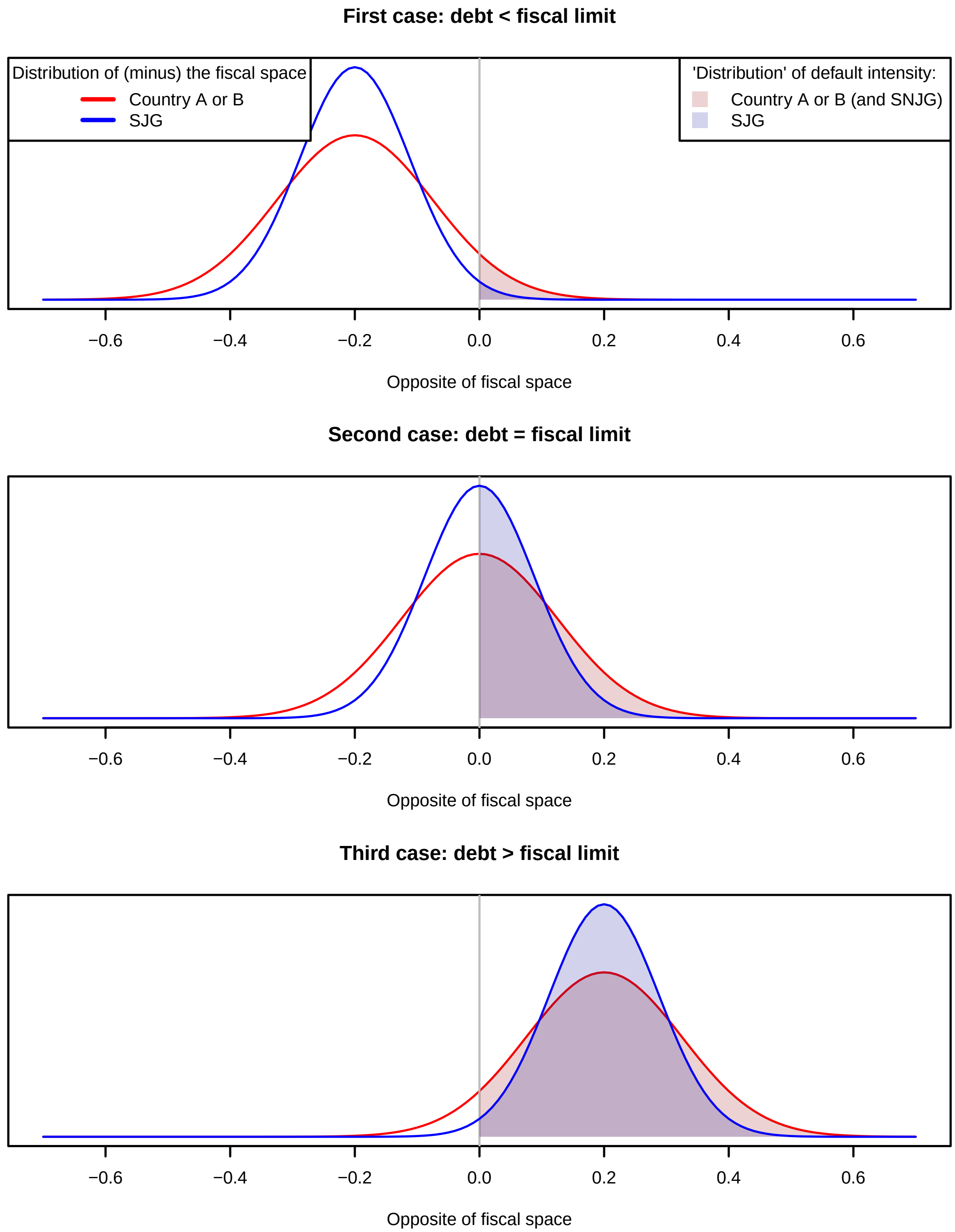
$$\begin{aligned}
\mathbb{E}_t^{\mathbb{Q}}(\exp(u'X_{t+1})) &= \mathbb{E}_t\Big(\exp\left(-\frac{1}{2}\psi_t'\psi_t - \psi_t'\eta_{t+1}\right)\exp(u'X_{t+1})\Big)\\
&= \mathbb{E}_t\Big(\exp\Big(-\frac{1}{2}(\psi_0+\psi_1X_t)'(\psi_0+\psi_1X_t) + u'\mu + u'\Phi X_t + (\Sigma'u-\psi_0-\psi_1X_t)'\eta_{t+1}\Big)\Big)\\
&= \mathbb{E}_t\Big(\exp\Big(-\frac{1}{2}(\psi_0+\psi_1X_t)'(\psi_0+\psi_1X_t) + u'\mu + u'\Phi X_t +\\
&\qquad \frac{1}{2}(\Sigma'u-\psi_0-\psi_1X_t)'(\Sigma'u-\psi_0-\psi_1X_t)\Big)\Big)\\
&= \exp\left(u'(\mu-\Sigma\psi_0) + u'(\Phi-\Sigma\psi_1)X_t + \frac{1}{2}u'\Sigma\Sigma'u\right).
\end{aligned}$$

By analogy with (a.2.9), it comes that the risk-neutral dynamics of X_t reads:

$$X_t = \mu^{\mathbb{Q}} + \Phi^{\mathbb{Q}}X_{t-1} + \Sigma\eta_t^{\mathbb{Q}}, \quad \eta_t^{\mathbb{Q}} \sim i.i.d.\,\mathcal{N}(0,I), \tag{a.2.10}$$

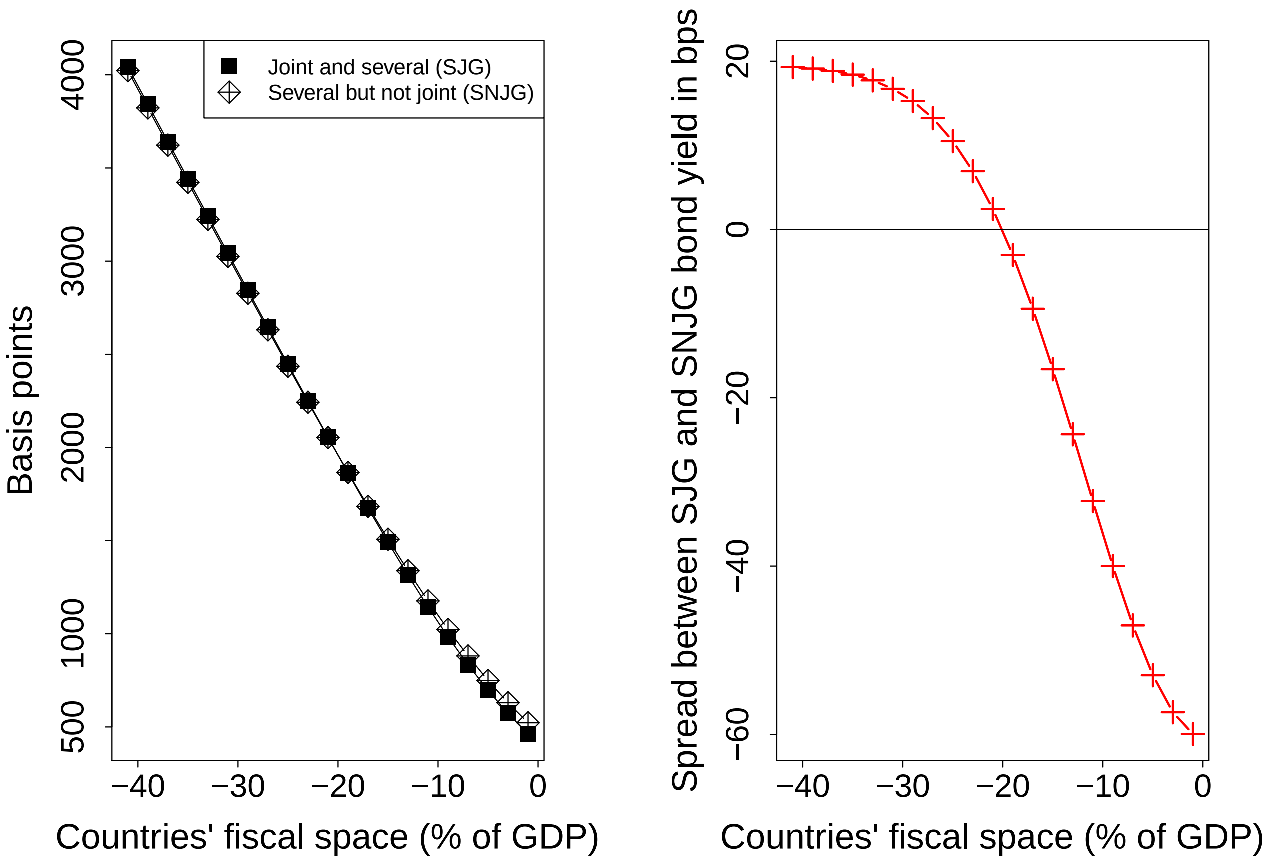
where $\mu^{\mathbb{Q}} = \mu - \Sigma\psi_0$, and $\Phi^{\mathbb{Q}} = \Phi - \Sigma\psi_1$.

Figure 2.D1: SJG and SNJG Bonds under negative fiscal space ($\bar{d} > \bar{\ell}$) - default intensity



This figure illustrates the influence of the sign of expected fiscal spaces on diversification effects. The context is the one of the stylized model presented in Section 2.3. Blue (respectively red) elements correspond to the joint area (respectively to single countries, or to SNJG bonds). The first panel corresponds to the conventional situation, where the expected fiscal space is positive. In that case, the default intensity associated with the SJG bond (joint area) is more concentrated towards zero than for the national default intensities. (To see that, compare the bluish and reddish areas, which represent the distributions of the default intensities—excluding the Dirac mass at zero.) The diversification effect is still at play when the fiscal space expectation is zero (middle plot); and it reverts when the expectation of the fiscal space is negative (third and last plot). In the latter case, the distribution of the joint-area default intensity is relatively more concentrated on the right-hand side of zero than for national default intensities. The calibration is as in the baseline situation described in Section 2.3 (stylized model), except that the correlation between debts (ρ in eq. 2.4) is set to zero.

Figure 2.D2: SJG and SNJG Bond yields under negative fiscal space ($\bar{d} > \bar{\ell}$)



This figure shows how the yields of SJG and SNJG bonds behave when the average debt is larger than the average fiscal limit. The left-hand side panel of the figure displays SJG and SNJG bond yields. The right-hand side panel reports the difference between the two types of yields. The calibration is as in the baseline situation described in Section 3 of the paper (stylized model), except that, for expository purpose, the correlation between debts is set to zero.

2.F Pricing of risk-free bonds

By definition of the state vector $X_t = [i_t, i_{t-1}, d_t, d_{t-1}, \ell_t]$, eq. (2.18) is satisfied for $h = 1$, with:

$$A_1 = 0, \quad \text{and} \quad B_1 = -[1, 0, \dots]'.$$

Let us assume that eq. (2.18) holds for a maturity $h - 1$, with $h > 1$ (and for any date t). Then, the price of a risk-free zero-coupon bond of maturity $h - 1$ is given by

$$P_{t,h-1} = \exp(A_{h-1} + B'_{h-1} X_t). \tag{a.2.11}$$

Let us then express the price of a risk-free zero-coupon bond of maturity h:

$$\begin{aligned} P_{t,h} &= \mathbb{E}_t(\mathcal{M}_{t,t+1}P_{t,h-1}) = \exp(-i_t)\mathbb{E}_t^{\mathbb{Q}}(\exp(A_{h-1}+B'_{h-1}X_{t+1})) \quad \text{using (a.2.11)} \\ &= \exp(B'_1X_t)\mathbb{E}_t^{\mathbb{Q}}(\exp(A_{h-1}+B'_{h-1}[\mu^{\mathbb{Q}}+\Phi^{\mathbb{Q}}X_t+\Sigma\eta_{t+1}])) \\ &= \exp\left(A_{h-1}+B'_{h-1}\mu^{\mathbb{Q}}+\frac{1}{2}B'_{h-1}\Sigma\Sigma'B_{h-1}+[B_1+\Phi^{\mathbb{Q}\prime}B_{h-1}]'X_t\right), \end{aligned}$$

which leads to eq. (2.17), using the definitions of $\mu^{\mathbb{Q}}$ and $\Phi^{\mathbb{Q}}$ given in (a.2.10).

2.G Approximate formula for zero-coupon risky bond

This appendix details the approximation to the price $P_{t,h}^{(j)}$ (this price being defined though eq. a.2.5); the resulting formula is given in Appendix 2.B.

Since $X_t = [i_t, i_{t-1}, d_{1,t}, \cdots, d_{n,t}, \ell_{1,t}, \cdots, \ell_{n,t}]'$, we have

$$i_{t-1} = \delta' X_t, \tag{a.2.12}$$

where $\delta = [0, 1, 0, \ldots, 0]'$. Moreover, we also introduce the following notation:

$$\lambda_{j,t} = \delta_j' X_t,$$

where $\delta_j = [0, 0, \alpha e_j', \mathbf{0}_{1\times N}, -\alpha e_j']'$, e_j denoting the j^{th} column vector of the $N \times N$ identity matrix. With these notations, eq. (2.2) rewrites:

$$\underline{\lambda}_{j,t} = \max(0, \lambda_{j,t}),$$

that is, $\lambda_{j,t}$ can be seen as a "shadow intensity". With these notations, we can rewrite eq. (a.2.5) as:

$$P_{t,h}^{(j)} = \mathbb{E}_t^{\mathbb{Q}}[\exp(-\delta' X_{t+1} - \max(0, \lambda_{j,t+1}) - \cdots - \delta' X_{t+h} - \max(0, \lambda_{j,t+h}))]. \tag{a.2.13}$$

Let us recall the notation introduced in Appendix 2.B:

$$p_{j,h-1,h} = -\log(P_{t,h}^{(j)}) + \log(P_{t,h-1}^{(j)}). \tag{a.2.14}$$

In the spirit of Wu and Xia (2016), we determine approximations to $p_{j,h-1,h}$ that we further use to get approximations to $P_{t,h}^{(j)}$, using:

$$P_{t,h}^{(j)} = \exp(p_{j,0,1} + p_{j,1,2} + \cdots + p_{j,h-1,h}). \tag{a.2.15}$$

The approximation to $p_{j,h-1,h}$ is essentially based on $\log \mathbb{E}[\exp(Z)] \approx \mathbb{E}(Z) + \frac{1}{2}\mathbb{V}(Z)$, which is exact when Z is Gaussian, but not if it is truncated Gaussian, as is the case here. This gives:

$$\begin{aligned} p_{j,k-1,k} &= \mathbb{E}_t^{\mathbb{Q}}(\delta' X_{t+k} + \underline{\lambda}_{j,t+k}) - \frac{1}{2}\mathbb{V}_t^{\mathbb{Q}}(\delta' X_{t+k} + \underline{\lambda}_{j,t+k}) - \\ &\quad -\mathbb{C}ov_t^{\mathbb{Q}}\left(\delta' X_{t+k} + \underline{\lambda}_{j,t+k}, \sum_{i=1}^{k-1}(\delta' X_{t+i} + \underline{\lambda}_{j,t+i})\right) \end{aligned} \tag{a.2.16}$$

Following Wu and Xia (2016), considering that $\lambda_{j,t}$ is a persistent process and introducing the following notation:

$$\mathfrak{q}_{j,t,k} = \mathbb{P}^{\mathbb{Q}}_t(d_{j,t+k} > \ell_{j,t+k}),$$

we have, for $k > 0$ and $0 \leq i \leq k$:

$$\mathbb{C}ov^{\mathbb{Q}}_t(\underbrace{i_{t-1+k}}_{\delta' X_{t+k}}, \underline{\lambda}_{j,t+k-i}) \approx \mathfrak{q}_{j,t,k-i}\mathbb{C}ov^{\mathbb{Q}}_t(\underbrace{i_{t-1+k}}_{\delta' X_{t+k}}, \lambda_{j,t+k-i}) \tag{a.2.17}$$

$$\mathbb{C}ov^{\mathbb{Q}}_t(\underline{\lambda}_{j,t+k}, \underline{\lambda}_{j,t+k-i}) \approx \mathfrak{q}_{j,t,k-i}\mathbb{C}ov^{\mathbb{Q}}_t(\lambda_{j,t+k}, \lambda_{j,t+k-i}) \tag{a.2.18}$$

Using the last two equations, we can rewrite eq. (a.2.16) as follows:

$$\begin{aligned} p_{j,k-1,k} \approx\ & \mathbb{E}^{\mathbb{Q}}_t(\delta' X_{t+k} + \underline{\lambda}_{j,t+k}) - \\ & -\frac{1}{2}\Big(\mathfrak{q}_{j,t,k}\mathbb{V}^{\mathbb{Q}}_t(\delta' X_{t+k} + \lambda_{j,t+k}) + (1-\mathfrak{q}_{j,t,k})\mathbb{V}^{\mathbb{Q}}_t(\delta' X_{t+k})\Big) - \\ & -\sum_{i=1}^{k-1}\Big(\mathfrak{q}_{j,t,i}\mathbb{C}ov^{\mathbb{Q}}_t\Big(\delta' X_{t+k} + \lambda_{j,t+k}, \delta' X_{t+i} + \lambda_{j,t+i}\Big) + \\ & +(1-\mathfrak{q}_{j,t,i})\mathbb{C}ov^{\mathbb{Q}}_t\Big(\delta' X_{t+k}, \delta' X_{t+i}\Big)\Big). \end{aligned} \tag{a.2.19}$$

Posing

$$\begin{aligned} \mu^{\mathbb{Q}}_{t,k} &= \mathbb{E}^{\mathbb{Q}}_t(X_{t+k}), & \mu^{\mathbb{Q}}_{j,t,k} &= \mathbb{E}^{\mathbb{Q}}_t(\lambda_{j,t+k}), \\ \sigma^{\mathbb{Q}}_{j,k} &= \sqrt{\mathbb{V}^{\mathbb{Q}}_t(\lambda_{j,t+k})}, & \Gamma^{\mathbb{Q}}_{k,i} &= \mathbb{C}ov^{\mathbb{Q}}_t(X_{t+k}, X_{t+k-i}), \end{aligned}$$

and using $\lambda_{j,t} = \delta'_j X_t$, we finally obtain

$$\begin{aligned} p_{j,k-1,k} \approx\ & \delta'\mu^{\mathbb{Q}}_{t,k} + \Phi(\mu^{\mathbb{Q}}_{j,t,k}/\sigma^{\mathbb{Q}}_{j,k})\mu^{\mathbb{Q}}_{j,t,k} + \phi(-\mu^{\mathbb{Q}}_{j,t,k}/\sigma^{\mathbb{Q}}_{j,k})\sigma^{\mathbb{Q}}_{j,k} - \\ & -\frac{1}{2}\Big(\mathfrak{q}_{j,t,k}(\delta+\delta_j)'\Gamma^{\mathbb{Q}}_{k,0}(\delta+a_j) + (1-\mathfrak{q}_{j,t,k})\delta'\Gamma^{\mathbb{Q}}_{k,0}\delta\Big) - \\ & -\sum_{i=1}^{k-1}\Big(\mathfrak{q}_{j,t,k-i}(\delta+\delta_j)'\Gamma^{\mathbb{Q}}_{k,i}(\delta+\delta_j) + (1-\mathfrak{q}_{j,t,k-i})\delta'\Gamma^{\mathbb{Q}}_{k,i}\delta\Big), \end{aligned} \tag{a.2.20}$$

with

$$\mathfrak{q}_{j,t,k} = \Phi\left(\mu^{\mathbb{Q}}_{t,k}/\sigma^{\mathbb{Q}}_{j,k}\right).$$

The next appendix details a fast (coding-oriented) approach to compute the $\mu^{\mathbb{Q}}_{t,k}$s and $\Gamma^{\mathbb{Q}}_{k,j}$s.

2.H Computation of $\mu^{\mathbb{Q}}_{t,k}$ and $\Gamma^{\mathbb{Q}}_{k,j}$

Recall X_t's law of motion (eq. 2.12):

$$X_t = \mu^{\mathbb{Q}} + \Phi^{\mathbb{Q}} x^{\mathbb{Q}}_{t-1} + \Sigma\varepsilon^{\mathbb{Q}}_{x,t}, \quad \varepsilon_{x,t} \sim i.i.d.\mathcal{N}(0, Id).$$

Using the notation $\Omega = \Sigma\Sigma'$, we have:

$$\left\{\begin{array}{lllll}\mu^{\mathbb{Q}}_{t,k} & = & \mathbb{E}^{\mathbb{Q}}_t(X_{t+k}) & = & (Id-\Phi^{\mathbb{Q}})^{-1}(Id-\Phi^{\mathbb{Q}^k})\mu^{\mathbb{Q}} + \Phi^{\mathbb{Q}^k}X_t,\\ \Gamma^{\mathbb{Q}}_{k,0} & = & \mathbb{V}^{\mathbb{Q}}_t(X_{t+k}) & = & \Omega + \Phi^{\mathbb{Q}}\Gamma^{\mathbb{Q}}_{k-1,0}\Phi^{\mathbb{Q}'}, \quad \text{with} \quad \Gamma^{\mathbb{Q}}_{1,0} = \Omega\\ & & & = & \Omega + \Phi^{\mathbb{Q}}\Omega\Phi^{\mathbb{Q}'} + \dots + \Phi^{\mathbb{Q}^{k-1}}\Omega\Phi^{\mathbb{Q}^{k-1'}},\\ \Gamma^{\mathbb{Q}}_{k,i} & = & \mathbb{C}ov^{\mathbb{Q}}_t(X_{t+k}, X_{t+k-i}) & = & \Phi^{\mathbb{Q}^i}\Gamma^{\mathbb{Q}}_{k-i,0} \quad \text{if} \quad k-i>0.\end{array}\right.$$

The estimation involves a large number of computations of the $\Gamma^{\mathbb{Q}}_{k,j}$'s. In order to speed up the computation, one can employ the following approach.

Consider a vector β of dimension n_x, that is the dimension of X_t, and let us denote by ξ^{β}_i the vector defined by $\xi^{\beta}_i = (\Phi^{\mathbb{Q}^i}_x)'\beta$ (β will typically be δ_j, or $(\delta_j+\delta)$). Because we have $\Gamma^{\mathbb{Q}}_{k,i} = \Phi^{\mathbb{Q}^i}_x\Omega + \Phi^{\mathbb{Q}^{i+1}}_x\Omega\Phi'_x + \dots + \Phi^{k-1}_x\Omega\Phi^{k-1-i'}_x$, it comes that:

$$\beta'\Gamma_{k,j}\beta = \xi^{\beta'}_i\Omega\xi^{\beta}_0 + \xi^{\beta}_{i+1}{}'\Omega\xi^{\beta}_1 + \dots + \xi^{\beta}_{k-1}{}'\Omega\xi^{\beta}_{k-1-i}. \tag{a.2.21}$$

Let us consider a maximal value for k, say H, and let us denote by Ξ_β the $n_x \times (H+1)$ matrix whose w^{th} column is ξ^{β}_{w-1}. It can then be seen that the (i,k) entry of $\Psi^{\beta} := \Xi_\beta{}'\Omega\Xi_\beta$ – which is a matrix of dimension $(H+1)\times(H+1)$ – is equal to $\xi^{\beta}_{i-1}{}'\Omega\xi^{\beta}_{k-1}$. The sum of the entries $(i+1,1), (i+2,2), \dots, (i+k,k)$ of Ψ^{β} therefore is

$$\xi^{\beta'}_j\Omega\xi^{\beta}_0 + \xi^{\beta}_{i+1}{}'\Omega\xi^{\beta}_1 + \dots + \xi^{\beta}_{i+k-1}{}'\Omega\xi^{\beta}_{k-1},$$

which is equal to $\beta'\Gamma^{\mathbb{Q}}_{i+k,i}\beta$ according to (a.2.21). Equivalently, $\beta'\Gamma^{\mathbb{Q}}_{k,i}\beta$ is the sum of the entries $(i+1,1)$, $(i+2,2), \dots, (k,k-i)$ of Ψ^{β}.

In particular, the entry $(1,1)$ of Ψ^{β} is equal to $\beta'\Gamma_{1,0}\beta$, the sum of the entries $(1,1)$ and $(2,2)$ is equal to $\beta'\Omega\beta + \beta'\Phi_x\Omega\Phi'_x\beta = \beta'\Gamma_{2,0}\beta$, and, more generally, the sum of the entries $(1,1),\dots,(n-1,n-1)$ of Ψ^{β} is equal to $\beta'\Gamma_{n,0}\beta$.

2.I Pricing zero-coupon bonds with non-zero recovery rates

Consider a zero-coupon bond of maturity h issued by country j. Assume the recovery rate is RR. On date $t+k$, with $0<k\leq h$, the payoff of this bond is zero, unless the country defaults on date $t+k$, in which case the bond payoff is assumed to be the fraction RR of the price of a risk-free zero-coupon bond of equivalent residual maturity, i.e. $\exp[-(h-k)i_{t+k,h-k}]$ (this is the Recovery of Treasury convention—RT—of Duffie and Singleton, 1999). Hence, the payoffs of this bond are of the form:

$$\left\{\begin{array}{ll} RR\times\exp(-(h-k)i_{t+k,h-k})\times(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1}) & \text{if} \quad 0<k<h,\\ 1-\mathcal{D}_{j,t+k}+RR\times(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1}) & \text{if} \quad k=h.\end{array}\right.$$

As a result, denoting by $\Lambda_{t,t+k}$ the (non stochastic) discount factor $\exp(-i_t - \cdots - i_{t+k-1})$, the price of this bond is given by:

$$\begin{aligned}
\mathcal{P}^{(j)}_{t,h} &= \mathbb{E}^{\mathbb{Q}}_t\left(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h}) + RR\sum_{k=1}^{h}\Lambda_{t,t+k}\exp(-(h-k)i_{t+k,h-k})(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1})\right)\\
&= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) + RR\sum_{k=1}^{h}\mathbb{E}^{\mathbb{Q}}_t\left[\Lambda_{t,t+k}\mathbb{E}^{\mathbb{Q}}_{t+k}\left\{\exp(-i_{t+k}-\cdots-i_{t+h-1})\right\}(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1})\right]\\
&= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) + RR\sum_{k=1}^{h}\mathbb{E}^{\mathbb{Q}}_t\left[\Lambda_{t,t+h}(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1})\right] \quad \text{(by the law of iterated expectations)}\\
&= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) + RR\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h})\sum_{k=1}^{h}\mathbb{E}^{\mathbb{Q}}_t\left[(\mathcal{D}_{j,t+k}-\mathcal{D}_{j,t+k-1})\right],
\end{aligned}$$

where the conditional expectation $\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h}))$ represents the date-t price of a zero-coupon zero-recovery risky bond of maturity h providing a payoff of 1 on date $t+h$ if country j has not defaulted before $t+h$, and zero otherwise (see Appendices 2.B for an approximation of this price). Moreover, $\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}\mathcal{D}_{j,t+k}) = \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h})\mathbb{E}^{\mathbb{Q}}_t(\mathcal{D}_{j,t+h})$ results from the fact that, under our assumptions regarding the s.d.f., $\mathcal{D}_t$ and i_t are independent under the risk-neutral measure $\mathbb{Q}$ (as they are under $\mathbb{P}$). Therefore:

$$\begin{aligned}
\mathcal{P}^{(j)}_{t,h} &= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) + RR\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}\mathcal{D}_{j,t+h})\\
&= \mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) - RR\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h}(1-\mathcal{D}_{j,t+h})) + RR\mathbb{E}^{\mathbb{Q}}_t(\Lambda_{t,t+h})\\
&= (1-RR)P^{(j)}_{t,h} + RR\exp(-hi^{0}_{t,h}),
\end{aligned}$$

where approximation formulas for $P^{(j)}_{t,h}$ are given in Appendix 2.B (computation details are given in Appendices 2.G and 2.H).

2.J Time variability of credit risk premiums

This appendix explains why the present framework accommodates time-varying credit risk premiums in spite of featuring constant prices of risk associated with debt and fiscal-limit shocks (that drive default risk).

Loosely speaking, risk premiums can be seen as the product of a (a) price of risk (ψ_t in our framework, see eq. 2.15) and (b) a quantity of risk, characterized by the amount of randomness in the system, and measured by conditional variances. We obtain time-varying risk premiums as soon as (a) or (b) varies. In our model, the conditional variance associated with the default intensity varies through time because of the non-linearity implied by the max operator, as detailed below.

To simplify, consider a situation where risk-free interest rates are null. The conditional probability of default is given by

$$\mathbb{P}(\mathcal{D}_t = 1|\mathcal{D}_{t-1} = 0, s_t) = 1 - \exp(-\max(-s_t, 0)), \tag{a.2.22}$$

where s_t is the fiscal space. (Hence, the probability of default is null if $s_t \geq 0$, and strictly positive otherwise.) The fiscal space follows a random walk:

$$s_t = s_{t-1} + \sigma\eta_t,$$

where $\eta_t \sim i.i.d.\mathcal{N}(0,1)$.

The s.d.f. is given by:

$$\mathcal{M}_{t,t+1} = \exp\left(-\nu\eta_{t+1} - \frac{1}{2}\nu^2\right).$$

With $\nu > 0$, this model implies that the s.d.f. is higher when η_{t+1} is negative, i.e., when the fiscal space diminishes.

In this context, and with a recovery rate, the price of a one-period defaultable bond is:

$$\begin{aligned} P_{t,1} &= \mathbb{E}_t^{\mathbb{Q}}\{(1-\mathcal{D}_{t+1})\} = \mathbb{E}_t^{\mathbb{Q}}\{\exp(-\max[-s_{t+1},0])\} \\ &= \mathbb{E}_t\left\{\exp(-\max[-s_{t+1},0])\exp\left(-\nu\eta_{t+1} - \frac{1}{2}\nu^2\right)\right\}. \end{aligned}$$

Denoting by $P_{t,1}^*$ the price of the bond that would be observed under the expectation hypothesis ($\nu = 0$), the credit risk premium is given by:

$$-\log P_{t,1} + \log P_{t,1}^* = -\log\left(\frac{\mathbb{E}_t^{\mathbb{Q}}\{(1-\mathcal{D}_{t+1})\}}{\mathbb{E}_t\{(1-\mathcal{D}_{t+1})\}}\right),$$

with

$$\left\{\begin{array}{lcl} P_{t,1} &=& \mathbb{E}_t\left\{\exp\left(-\max[-s_t - \sigma\eta_{t+1},0] - \nu\eta_{t+1} - \frac{1}{2}\nu^2\right)\right\} \\ P_{t,1}^* &=& \mathbb{E}_t\left\{\exp(-\max[-s_t - \sigma\eta_{t+1},0])\right\}. \end{array}\right.$$

Consider two polar cases:

- When s_t is large and positive (e.g., $s_t > 4\sigma$), it is extremely likely that $\max[-s_t - \sigma\eta_{t+1}, 0]$ will be equal to zero and, accordingly, we will have $P_{t,1} \approx P_{t,1}^* = 1$. The credit risk premium is therefore essentially zero.

- When s_t is large and negative, it is extremely likely that $\max[-s_t - \sigma\eta_{t+1}, 0] = -s_t - \sigma\eta_{t+1}$, and, as a result:

  $$\left\{\begin{array}{lcl} P_{t,1} &\approx& \mathbb{E}_t\left\{\exp\left(s_t + \sigma\eta_{t+1} - \nu\eta_{t+1} - \frac{1}{2}\nu^2\right)\right\} \\ P_{t,1}^* &\approx& \mathbb{E}_t\left\{\exp(s_t + \sigma\eta_{t+1})\right\}, \end{array}\right.$$

  which leads, after simple algebra, to:

  $$-\log P_{t,1} + \log P_{t,1}^* = \nu\sigma.$$

For intermediate values of s_t, the credit risk premium will vary between 0 and $\nu\sigma$.

Without the non-linearity stemming from the max() operator—and keeping the ψ_1 entries associated with those shocks affecting the fiscal space (i.e., the η_d's and η_ℓ's) at zero—then the credit risk premiums would be constant. In the previous example, this would correspond to the situation where we would remove the max() operator in eq. (a.2.22) above.[36] (The credit risk premium would then be equal to $\nu\sigma$ for any value of the only state variable considered in the present example, namely s_t.)

2.K Inversion technique

This appendix describes the computation of the likelihood function (see Subsection 2.5.2 for a general description of our estimation approach).

We consider the following decomposition of the state vector $X_t = [i_t, i_{t-1}, d_t', \ell_t']'$:

$$\underbrace{X_t}_{m\times 1} = \begin{bmatrix} \underbrace{\widetilde{X}_t}_{(m-N)\times 1} \\ \underbrace{\ell_t}_{N\times 1} \end{bmatrix},$$

where $\widetilde{X}_t$ are the observable components of X_t.

The state vector follows a vector autoregressive process of order one (eq. 2.12).

The vector of observed financial data is organized as follows:

$$Y_t = \begin{bmatrix} \underbrace{Y_t^{(y)}}_{n_y\times 1} \\ \underbrace{Y_{1,t}^{(YS)}}_{n_1\times 1} \\ \underbrace{Y_{2,t}^{(YS)}}_{N\times 1} \end{bmatrix},$$

where $Y_t^{(y)}$ is a vector of risk-free yields (of maturities 2, 3, 5 and 10 years), $Y_{1,t}^{(YS)}$ is a vector of imperfectly-fitted bond yield spreads (e.g. maturities 2 and 10 yrs) and $Y_{2,t}^{(YS)}$ is a $N \times 1$ vector of perfectly-fitted bond yield spreads (in our case, the average of bond yield spreads of maturities 2, 5 and 10 years). These yields and spreads are given by:

$$\left\{ \begin{array}{lcl} Y_t^{(y)} & = & A_y + B_y' \widetilde{X}_t + \xi_t^{(y)} \\ Y_{1,t}^{(YS)} & = & f_1(\widetilde{X}_t, \ell_t) + \xi_t^{(YS)} \\ Y_{2,t}^{(YS)} & = & f_2(\widetilde{X}_t, \ell_t) \quad \text{(these spreads are perfectly fitted).} \end{array} \right. \tag{a.2.23}$$

We assume that the components of $\xi_t^{(y)}$ and $\xi_t^{(YS)}$ are i.i.d. normally-distributed measurement errors. The variance of each component of $\xi_t^{(y)}$ is σ_y^2. The variance of the i^{th} component of $\xi_t^{(YS)}$ is $\sigma_{YS,i}^2$.

System (a.2.23) can be rewritten:

$$\left\{ \begin{array}{lcl} Y_t^{(y)} & = & A_y + B_y' \widetilde{X}_t + \xi_t^{(y)} \\ Y_{1,t}^{(YS)} & = & f_1(\widetilde{X}_t, f_2^*(\widetilde{X}_t, Y_{2,t}^{(YS)})) + \xi_t^{(YS)}, \end{array} \right. \tag{a.2.24}$$

[36] Such a purely-Gaussian specification would not be consistent with the fact that the probability of default cannot be negative. It has however been sometimes used in the literature, e.g., by Liu, Longstaff, and Mandell (2006).

where function f_2^* represents the inversion of the pricing of $Y_{2,t}^{(YS)}$, i.e.:

$$Y_{2,t}^{(YS)} = f_2\left(\widetilde{X}_t, \ell_t\right) \Leftrightarrow \ell_t = f_2^*\left(\widetilde{X}_t, Y_{2,t}^{(YS)}\right).$$

Let us use the following notations:

$$W_t = \begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ \widetilde{X}_t \\ Y_{2,t}^{(YS)} \end{bmatrix} \quad \text{and} \quad Z_t = \begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ \widetilde{X}_t \\ \ell_t \end{bmatrix} = \begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ X_t \end{bmatrix}.$$

Under the assumption that $Y_{2,t}^{(YS)}$ is perfectly fitted by the model, the information contained in Z_t is the same as that contained in W_t. But the p.d.f. of Z_t, conditional on W_{t-1} (or, equivalently, conditional on Z_{t-1}), is easier tom derive than that of W_t.

Indeed, we have:

$$\begin{aligned}
&\log f_{Z_t|Z_{t-1}}(Z_t) = \\
&-\frac{n_y}{2}\log(2\pi) - n_y \log\sigma_y - \frac{1}{2\sigma_y^2}\left(Y_t^{(y)} - A_y - B_y'\widetilde{X}_t\right)'\left(Y_t^{(y)} - A_y - B_y'\widetilde{X}_t\right) \\
&-\frac{n_1}{2}\log(2\pi) - \sum_{i=1}^{n_1}\log\sigma_{YS,i} - \frac{1}{2}\left(Y_{1,t}^{(YS)} - f_1(\widetilde{X}_t, \ell_t)\right)'\operatorname{diag}(1/\sigma_{YS}^2)\left(Y_{1,t}^{(YS)} - f_1(\widetilde{X}_t, \ell_t)\right) \\
&-\frac{m}{2}\log(2\pi) - \frac{1}{2}\log(|\Sigma\Sigma'|) - \frac{1}{2}\left(X_t - \mu - \Phi X_{t-1}\right)'(\Sigma\Sigma')^{-1}\left(X_t - \mu - \Phi X_{t-1}\right), \qquad \text{(a.2.25)}
\end{aligned}$$

where $\operatorname{diag}(1/\sigma_{YS}^2)$ is a diagonal matrix whose i^{th} diagonal entry is $1/\sigma_{YS,i}^2$.

Remark that this does not provide us with the likelihood associated with observed data since ℓ_t is not directly observed.

We have:

$$W_t = g(Z_t),$$

with

$$g\left(\begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ \widetilde{X}_t \\ \ell_t \end{bmatrix}\right) = \begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ \widetilde{X}_t \\ Y_{2,t}^{(YS)} \end{bmatrix} = \begin{bmatrix} Y_t^{(y)} \\ Y_{1,t}^{(YS)} \\ \widetilde{X}_t \\ f_2(\widetilde{X}_t, \ell_t) \end{bmatrix}.$$

In general, we have:

$$f_{W_t|W_{t-1}}(W_t) = \left|\frac{\partial g^{-1}(W_t)}{\partial W'}\right| f_{Z_t|Z_{t-1}}(g^{-1}(W_t)), \qquad \text{(a.2.26)}$$

and, therefore:

$$\log f_{W_t|W_{t-1}}(W_t) = \underbrace{\log\left|\frac{\partial g^{-1}(W_t)}{\partial W'}\right|}_{\text{calculated using eq. (a.2.28)}} + \underbrace{\log f_{Z_t|Z_{t-1}}(g^{-1}(W_t))}_{\text{calculated using eq. (a.2.25)}}, \qquad \text{(a.2.27)}$$

where, using the inverse function theorem and the fact that $\left|\frac{\partial g^{-1}(W_t)}{\partial W'}\right|$ is diagonal:

$$\log\left|\frac{\partial g^{-1}(W_t)}{\partial W'}\right| = -\sum_{i=1}^{N}\log\frac{\partial f_2(\widetilde{X}_t,\ell_t)}{\partial \ell_{i,t}}. \tag{a.2.28}$$

In practice, in (a.2.25) and (a.2.28), we replace ℓ_t by $f_2^*\left(\widetilde{X}_t, Y_{2,t}^{(YS)}\right)$ —that is the fiscal limit recovered by the inversion technique.

The vector of observed variables can be extended to include $\mathcal{D}_t$. Using the notation $W_t^* = [W_t', \mathcal{D}_t']'$ and exploiting the fact that $\mathcal{D}_t$ does not Granger-cause W_t, we have:

$$\log f_{W_t^*|W_{t-1}^*}(W_t,\mathcal{D}_t) = \underbrace{\log f_{W_t|W_{t-1}}(W_t)}_{\text{calculated using eq. (a.2.27)}} + \underbrace{\log f_{\mathcal{D}_t|W_t}(\mathcal{D}_t).}_{\text{calculated using eq. (a.2.30)}} \tag{a.2.29}$$

In particular, if all the components of $\mathcal{D}_t$ are zero (absence of default), we have:

$$\log f_{W_t^*|W_{t-1}^*}(W_t,\mathcal{D}_t=0) = \log f_{W_t|W_{t-1}}(W_t) + \sum_{j=1}^{N}\log\left[1-\mathcal{F}(d_{j,t}-\ell_{j,t})\right]. \tag{a.2.30}$$

2.L Redistribution schemes: formulas and additional schemes

This appendix details the formulas underlying Section 2.6.2 of the paper.

2.L.1 General formulas

Assume that, on date t, a European debt agency issues common bonds with maturity h and face value F (it repays F at date $t+h$). The proceeds of the issuance are $P_{t,h}^e F$, with $e \in \{SJG, SNJG\}$, depending on the type of common bond that is issued. The proceeds are allocated across countries proportionally to GDPs. Recalling that GDP weights are denoted by ω_j, country j receives $\omega_j P_{t,h}^e F$. If country j had issued national bonds with the same face value ($\omega_j F$), it would have obtained $P_{t,h}^{(j)}\omega_j F$ on date t. Therefore, at the euro-area level, the gains are:

$$G_{t,h}F = P_{t,h}^e F - \left(\omega'\mathbf{P}_{t,h}F\right), \tag{a.2.31}$$

where $\mathbf{P}_{t,h}$ represent the N-dimensional vector of national prices and ω stands for the N-dimensional vector of GDP weights. (It can be seen from the previous formula that the aggregate gains are null when $e = SNJG$.)

Now, denote by ω_G the redistribution weights of the gains (with $\sum_j \omega_{G,j} = 1$). The after-gain-redistribution proceeds are:

$$\omega'\mathbf{P}_{t,h}F + G_{t,h}\omega_G F,$$

which is of the form $\omega'\mathbf{P}_{e,t,h}(\omega_G)F$, with

$$\mathbf{P}_{e,t,h}(\omega_G) = \mathbf{P}_{t,h} + G_{t,h}\frac{\omega_G}{\omega}, \tag{a.2.32}$$

where, by abuse of notation, $\frac{\omega_G}{\omega}$ denotes the vector whose j^{th} entry is $\omega_{G,j}/\omega_j$. $\mathbf{P}_{e,t,h}(\omega_G)$ can be interpreted as the pseudo issuance N-dimensional vector of prices after redistribution. The post-redistribution yields faced by the different countries are given by the following N-dimensional vector:

$$\mathbf{i}_{e,t,h}(\omega_G) = -\frac{1}{h}\log \mathbf{P}_{e,t,h}(\omega_G), \tag{a.2.33}$$

where, by abuse of notation, the log operator is applied element-wise.

Below, we describe the different after-gain redistribution schemes that we propose. Given that aggregate gains for the SNJG bond issuance scheme are nil, for the latter, we only focus on the scheme in which all countries face the same funding costs.

2.L.2 Scheme where countries face the same funding costs

In this scheme, the after-redistribution issuance price faced by all countries is the eurobond price. That is:

$$\mathbf{P}_{e,t,h}(\omega_G) = P^e_{t,h}.$$

Using $G_{t,h} = P^e_{t,h} - \omega'\mathbf{P}_{t,h}$ together with (a.2.32) then gives:

$$\omega_G = \omega \odot \frac{P^e_{t,h}\mathbf{1} - \mathbf{P}_{t,h}}{P^e_{t,h} - \omega'\mathbf{P}_{t,h}},$$

where $\odot$ is the element-wise product. Note that the sign of each country's redistribution weight $\omega_{G,j}$ depends on $P^e_{t,h} - P^{(j)}_{t,h}$. Therefore, this scheme implies negative "gains" for those countries j whose national bond prices are higher than that of the considered eurobond.

2.L.3 Scheme with GDP weights

In this case, the redistribution weights (ω_G) are equal to the GDP weights (ω). Using $G_{t,h} = P^e_{t,h} - \omega'\mathbf{P}_{t,h}$ (i.e., Eq. (a.2.31)), Eq. (a.2.32) gives:

$$\mathbf{P}_{e,t,h}(\omega_G) = \mathbf{P}_{t,h} + (P^e_{t,h} - \omega'\mathbf{P}_{t,h})\mathbf{1}.$$

2.L.4 Scheme with the same yield gains across countries

Under this scheme, all countries benefit from the same yield gain, denoted by Δi_t. Denote by $\mathbf{i}_{t,h}$ the N-dimensional vector of national bond yields. We want to have $\mathbf{P}_{e,t,h}(\omega_G) = \exp(-h(\mathbf{i}_{t,h} - \Delta i_t))$. Using (a.2.32), we get:

$$\mathbf{P}_{t,h} + G_{t,h}\frac{\omega_G}{\omega} = \exp(-h(\mathbf{i}_{t,h} - \Delta i_t)),$$

where, by abuse of notation, $\frac{\omega_G}{\omega}$ denotes the vector whose j^{th} entry is $\omega_{G,j}/\omega_j$. This gives:

$$\omega_G = \frac{1}{G_{t,h}}\omega \odot \left[\exp(-h(\mathbf{i}_{t,h} - \Delta i_t)) - \mathbf{P}_{t,h}\right],$$

where $\odot$ is the element-wise product. Since the components of ω_G have to sum to one, we have:

$$1 = \mathbf{1}'\left(\frac{1}{G_{t,h}}\omega \odot \left[\exp(-h(\mathbf{i}_{t,h} - \Delta i_t)) - \mathbf{P}_{t,h}\right]\right),$$

or, using that $\exp(-h\mathbf{i}_{t,h}) = \mathbf{P}_{t,h}$:

$$G_{t,h} = (\exp(h\Delta i_t) - 1)\mathbf{1}'(\omega \odot \mathbf{P}_{t,h}).$$

This further gives:

$$1 + \frac{G_{t,h}}{\mathbf{1}'\omega \odot \mathbf{P}_{t,h}} = \exp(h\Delta i_t),$$

and, finally:

$$\Delta i_t = \frac{1}{h} \log \left(1 + \frac{G_{t,h}}{\mathbf{1}'(\omega \odot \mathbf{P}_{t,h})} \right).$$

2.L.5 Scheme with no change in funding costs for Germany and France

Table 2.L1 complements the analysis developed in Subsection 2.6.2 with two additional schemes. In the first scheme (respectively second scheme), Germany (resp. both Germany and France) faces the same funding costs it would have faced under national issuance. Moreover, the aggregate gains are shared among the other countries on the base of their relative GDP size.

2.M Higher-order effects

This appendix proposes an analysis of potential higher-order effects associated with debt-service relief. The mechanisms underlying such effects would be as follows: if the average funding cost of a government decreases—because part of its funding needs is met with Eurobonds—then expected future debt decreases (through lower debt service), which further decreases national bond yields (through lower credit spreads) which, in turn, reduces again future debt service, and so on.

Assume that the government of country j issues bonds of maturity h. For notational simplicity, let us drop the subscript h. That is, denote by $y_{j,t}$ the yields associated with these bonds, and by $y_{j,t}^{(SJG)}$ the post-redistribution yields associated with the issuance of SJG bonds.

Consider a change in the funding strategy: while the whole debt was only funded through national bonds before date t, a fraction θ_j of the debt gets funded by SJG bonds after that date. Note that θ_j depends on countries given that the proceeds of a Eurobond issuance are supposed to be allocated according to GDPs (eq. 2.7). Specifically, we have:[37]

$$\theta_j = \theta \frac{\omega_j}{\omega_j^D}, \tag{a.2.34}$$

where the ω_j^Ds denotes debt weights (while the ω_js are GDP weights). The previous expression shows that, for countries whose indebtedness is larger than that prevailing at the euro-area aggregate level (which corresponds to $\omega_j^D > \omega_j$), then the share of debt issued in the form of Eurobonds is lower (since $\theta_j < \theta$).

[37] Indeed, denoting by D_j the debt outstanding of country j (i.e., $D_j = d_j Y_j$), we must have $\theta_j D_j / (\theta \sum_i D_i) = \omega_j$. Hence, $\theta_j = \omega_j \theta (\sum_i D_i) / D_j$.

Table 2.L1: Effect of redistribution schemes on funding costs (additional schemes)

	2008-06-30			2011-12-31			2021-06-30		
				SJG					
Panel A: No change in German funding cost									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	0%	439	0	0%	186	0	0%	-9	0
FR	36%	426	21	36%	221	22	36%	-29	34
IT	28%	447	21	28%	559	26	28%	23	35
ES	18%	432	21	18%	504	25	18%	3	34
NL	11%	435	21	11%	197	22	11%	-40	34
BE	7%	442	21	7%	334	23	7%	-30	34
Panel B: No change in German and French funding cost									
	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain	redist. weigth	post redist. yield	yield gain
DE	0%	439	0	0%	186	0	0%	-9	0
FR	0%	446	0	0%	243	0	0%	5	0
IT	44%	435	32	44%	545	41	44%	4	54
ES	28%	420	32	28%	490	39	28%	-16	53
NL	18%	423	32	18%	185	34	18%	-59	52
BE	10%	430	32	10%	321	36	10%	-49	52

Notes: This table compares post-redistribution funding costs across countries under the two issuance schemes (SJG and SNJG) and under the redistribution schemes described in 2.L.5. We focus on the 5-year maturity and on three periods: beginning of the estimation sample (2008Q2), midst of the euro-debt crisis (2011Q4) and end of the estimation sample (2021Q2). Yields are expressed in basis points. Aggregate gains are computed under the assumption that total issuance is equal to 5% of aggregate GDP. In each panel, for all countries and dates, we show the redistribution weights $\omega_{G,j}$, the post-redistribution yields and the yield gains, that are the differences between national bond yields and post-redistribution yields.

For newly issued debt, the apparent yield then becomes $\theta_j y_{j,t}^{(SJG)} + (1-\theta_j)y_{j,t}$, which is lower than $y_{j,t}$ if $y_{j,t}^{(SJG)} < y_{j,t}$. All else being equal, this should give rise to a decrease in debt payments, and hence in debt. Using that the maturity of newly-issued debt duration is h, it comes that the total reduction in debt payments, after h years, will be of the order of magnitude of:[38]

$$h \times \left[y_{j,t} - \left(\theta_j y_{j,t}^{(SJG)} + (1-\theta_j) y_{j,t} \right) \right]. \tag{a.2.35}$$

On date t, investors may take into account that debt-reduction effect when pricing bonds. (If they consider, in particular, that this debt reduction will not be substituted with higher primary surpluses.) In that case, in terms of funding cost, the first-round effect:

$$\theta_j y_{j,t}^{(SJG)} + (1-\theta_j) y_{j,t} < y_{j,t}. \tag{a.2.36}$$

[38]Note that this is only an approximation as the exact number would also depend on future yields (since the debt is completely renewed in h dates).

would be reinforced by second-round effects resulting from lower future debt payments. And, in turn, higher-round effects may follow. Investigating these effects, therefore, involves solving a fixed-point problem.

We proceed as follows. We start by computing national and SJG yields using our pricing formulas and a given value of the state vector X (that we take equal to its sample average value). Representing our pricing formulas by function f, this first step formally is:

$$\mathcal{Y} = f(X),$$

where $\mathcal{Y}$ is a vector gathering the relevant national yields and SJG yields. More precisely:

$$\mathcal{Y} = \begin{bmatrix} y_1 \\ \vdots \\ y_N \\ y_{SJG} \end{bmatrix}.$$

Next, we use these yields to compute the debt reduction (a.2.35). We then modifiy the state vector X in such a way that, over the next h periods, the average expected debt will indeed be reduced by this amount. (More precisely, we modify the debt trend, that is $\gamma = d - (1-\rho_d)\overline{d} - \rho_d d_{-1}$ (eq. 2.9), to achieve that.[39]) This provides us with $X^{(1)}$, which we use to compute new yields:

$$\mathcal{Y}^{(1)} = f(X^{(1)}).$$

The superscript (1) indicates that the resulting yields result from the first iteration. The yields in $\mathcal{Y}^{(1)}$ are going to be different from those in $\mathcal{Y}$ (because expected debt is lower). Hence, the debt reduction resulting from the partial issuance of SJG bonds is higher than what is suggested by (a.2.35). For each country, we then compute a novel debt reduction by using:

$$h \times \left[y_j - \left(\theta y^{(1)}_{SJG} + (1-\theta) y^{(1)}_j \right) \right],$$

which we use to construct a new state vector $X^{(2)}$, which gives a new vector of yields $\mathcal{Y}^{(2)}$, and so on until convergence.

[39]The model described in Section 2.4 is such that, for each country:

$$\begin{bmatrix} d_t \\ d_{t-1} \end{bmatrix} = \underbrace{\begin{bmatrix} \overline{d}(1-\rho_d)(1-\rho_\gamma) \\ 0 \end{bmatrix}}_{=c} + \underbrace{\begin{bmatrix} \rho_d + \rho_\gamma & -\rho_\gamma \rho_d \\ 1 & 0 \end{bmatrix}}_{=F} \begin{bmatrix} d_{t-1} \\ d_{t-2} \end{bmatrix} + \begin{bmatrix} \varepsilon_{d,t} \\ 0 \end{bmatrix}.$$

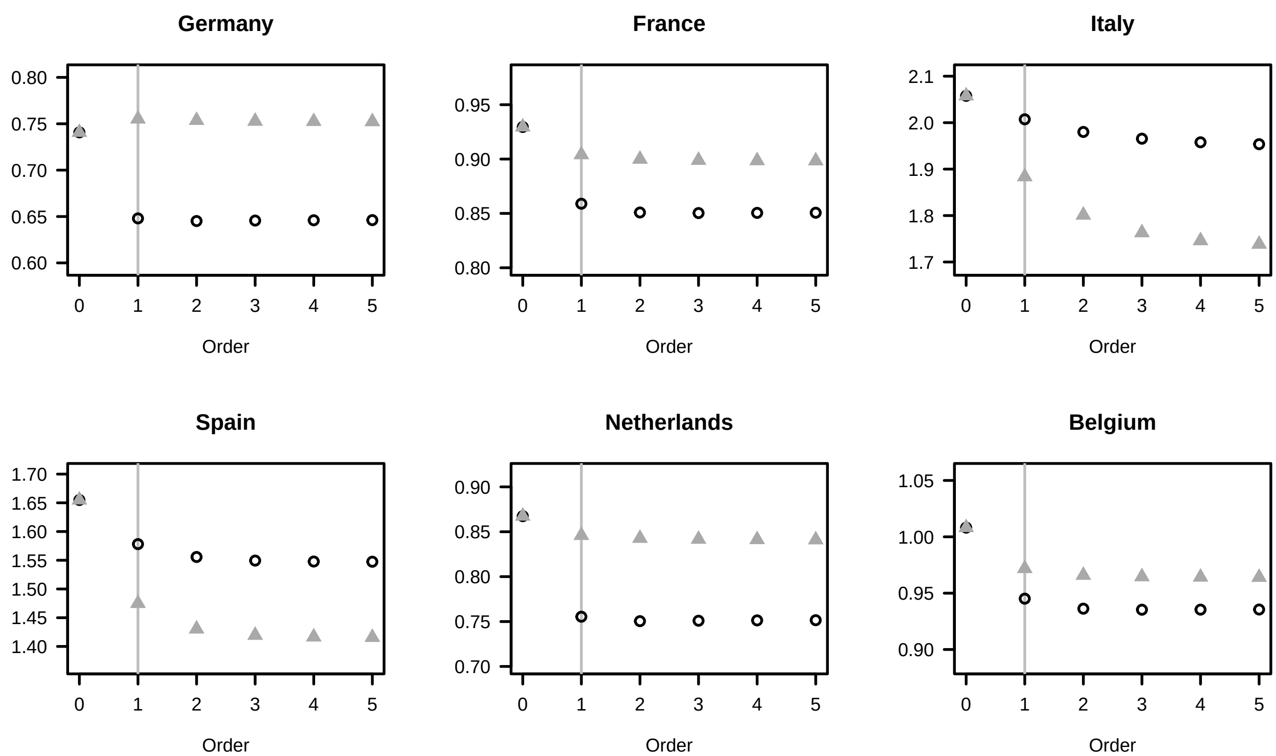
This implies that:

$$\begin{aligned} & \mathbb{E}_t \left[\frac{1}{h} \left(d_{t+1} + \dots + d_{t+h} \right) \right] \\ = \; & \frac{1}{h}(I-F)^{-1} \left[(h+1)I - (I-F)^{-1}(I - F^{h+1}) \right] c + \frac{1}{h}[(I-F)^{-1}(I-F^{h+1}) - I] \begin{bmatrix} d_t \\ d_{t-1} \end{bmatrix}. \end{aligned}$$

We use the previous formula to look for the value of d_{t-1} that results in the desired change in the expected average debt, i.e., $\mathbb{E}_t \left[\frac{1}{h} \left(d_{t+1} + \dots + d_{t+h} \right) \right]$. (This amounts to a change in $\gamma_t = d_t - (1-\rho_d)\overline{d} - \rho_d d_{t-1}$.)

Figure 2.M1 implements this approach, with a share of euro-area debt issued in the form of SJG bonds equal to $\theta = 20\%$. We consider two redistribution schemes (see Subsection 2.6.2), namely Scheme A (no redistribution after the issuance of SJG bonds) and Scheme B (where the redistribution of SJG aggregate gains is based on GDP weights). At "Order 0", the state vector X_t is set to its sample mean.

Figure 2.M1: Higher-order effects



Notes: These plots illustrate potential higher-order effects stemming from the (partial) issuance of SJG bonds. Specifically, we consider that $\theta = 20\%$ of the euro-area debt is issued in the form of SJG bonds. The grey triangles and black circles respectively correspond to Schemes A and B. (In Scheme A, there is no post-issuance redistribution when SJG bonds are issued; in Scheme B, aggregate gains associated with the issuance of SJG bonds are redistributed according to GDP weights). The first points of the plots ("Order" = 0) give the model-implied average 5-year yields associated with the different countries. If a fraction of the government funding needs is met by issuing Eurobonds, then the average funding cost is modified (eq. a.2.36). This gives the second point, of the charts ("Order" = 1), which is highlighted by a vertical grey line given that it represents the first-round effect of issuing Eurobonds (namely, the effects presented in the main findings in Section 2.6). Changes in funding costs affect expected debt trajectories, which, in turn, modify bond yields (for national and SJG bonds). The resulting funding costs are represented by the third set of points ("Order" = 2). The following points, for higher orders, are obtained by using the same steps, in an iterative fashion. Yields are annualized, and expressed in percentage points.

2.N Bonds used in Figure 2.6

Table 2.N1: Bonds used in Figure 2.6

Issuer	Eikon ticker	Coupon (in percent)	Maturity date
France	FRGV	2.50	25-May-2030
Belgium	BEGV	0.55	04-Mar-2029
Portugal	PTGV	3.875	15-Feb-2030
ESM	ESM	0.50	05-Mar-2029
Spain	ESGV	1.95	30-Jul-2030
Netherlands	NLGV	0.25	15-Jul-2029
Germany	DEGV IO Str	0	04-Jul-2030
NEXTGENEU	EUUNI	0	04-Jul-2031
EFSF	EFSFC	2.75	03-Dec-2029
EU	EUUNI	1.375	04-Oct-2029
Italy	ITGV	3.50	01-Mar-2030
EIB	EIB	0.25	14-Sep-2029

Notes: This table lists the bonds used in Figure 2.6. Asset swap spreads (ASW) are computed by Refinitiv Eikon.

2.O Alternative (static) fiscal limit estimates

Table 2.O1: Fiscal limit static estimates in the literature

Country	Ghosh et al. (2013)		Collard et al. (2015)					
	Hist.	**Proj.**	5% **MPS**	**MRR**	**TVR**	**CATA**	4% **MPS**	**hist. MPS**
DE	154.1	175.8	130.1	132.3	114.6	85.5	104.1	112.9
FR	170.9	176.1	146.6	148.6	119.8	97.8	117.2	40.0
IT	–	–	113.2	115.6	106.8	74.2	90.6	147.5
ES	218.3	153.9	144.2	146.2	119.3	95.8	115.3	115.6

Notes: All estimates are reported in percent of GDP. **Estimates of Ghosh et al. (2013) –** Debt limits (fiscal limits in our terminology) are statically estimated through the interest payment schedule for the period 1985-2007. **Hist.**: Estimates are based on the average interest rate / growth differential of 1998-2007, using the implied interest rate on public debt; **Proj.**: The interest rate / growth differential is based on the long term government bond yield (average for 2010-2014, IMF projections as of 2010). **Estimates of Collard et al. (2015) –** The computation of maximum sustainable debts (fiscal limits in our terminology) exploits the idea of a maximum primary surplus (MPS). In the model, there is a maximum amount that can be issued on each date (that itself depends on the MPS). 5% **MPS**: Case where the MPS is set to to 5%; **MRR**: The computation involves a maximum recovery rate; **TVR**: The model features a time-varying interest rate; **CATA**: The model features catastrophes; 4% **MPS**: The MPS is set to 5%; **hist. MPS**: The MPS is set to the historical peak of primary surplus-to-GDP.

2.P Partial Eurobonds

Our framework can also be used to study "partial" SNJG and SJG bonds, defined as bonds jointly issued by a subset of countries. Specifically, we focus on the computation of SNJG and SJG bonds issued by four countries (out of the six we consider) either excluding "super" core member states (Germany and Netherlands), or excluding peripheral countries (Italy and Spain). These prices are computed under the baseline estimated model, the only parameters that have to be adjusted to perform this analysis are the weights (ω) defining the groups of issuing countries (see Subsection 2.3.3).

Figure 2.P1 shows the SNJG and SJG bond yield spreads across different maturities (i) under the baseline scenario, where all countries participate in the emission, as in the main results presented in Section 2.2.1 (solid lines); (ii) under the scenario in which Germany and Netherlands do not participate in the issuance (dashed lines); and (iii) under the scenario in which peripheral member states are excluded from the joint issuance program (dotted lines). Not surprisingly, when Italy and Spain are excluded from the joint issuance program, yield spreads are below those obtained in the baseline scenario; and when Germany and Netherlands do not participate, yield spreads are higher. The spread between "partial" SJG and SNJG bonds—which reflects the aggregate yield gains—is smaller in the "partial" scenarios than when all countries participate in the program, as diversification effects are magnified in the latter case.

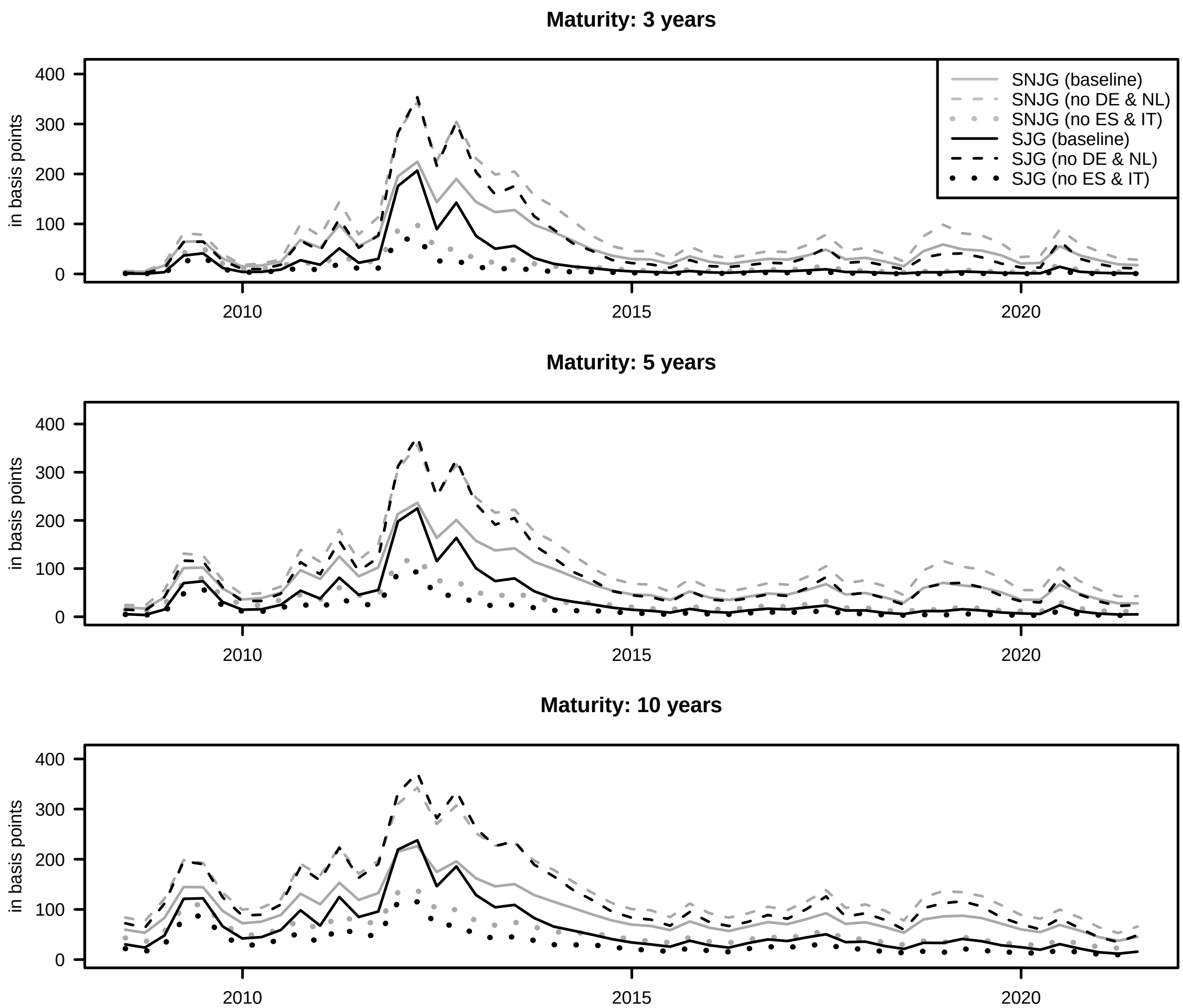
2.Q Sensitivity analysis

This appendix presents the results of sensitivity analyses performed to assess the robustness of our main baseline results. Specifically, we modify the model by changing the bounds or imposing a specific value on some key parameters, one at the time, and run the complete estimation. More precisely:

- We exclude the COVID period (after 2020Q1) from the PCA analysis of the recovered estimates for $\epsilon_{d,j,t}$'s so that also $\mathbb{V}ar(\varepsilon_{d,t}) = \Gamma_d \Gamma_d'$ is modified. (Γ_d represents the matrix of PCA weights, see Subsection 2.5.3 for details regarding Γ_d.)
- Considering that $maxSR$ is constrained at the lower bound, we relax such bound by reducing it to 0.25, instead of 0.5.
- Given that ρ_γ is constrained at the lower bound, we relax such bound by shifting it to 0.5, instead of 0.7.
- We impose a higher value on α (even if estimated), equal to 0.2, which corresponds to more than double the estimated value (see Table 2.3 in Sec. 2.5).
- We set $\rho_{d,\ell}$, the correlation between the two "main common shocks" ($\eta_{d,1,t}$ and $\eta_{\ell,1,t}$) to zero.
- Considering that the parameter ζ is constrained at its upper bound, we relax such bound by increasing it by 0.5 (from 1.5 to 2). Note that the parameter ζ is defined as the multiplicative factor disciplining the relation between Γ_d and Γ_ℓ ($\Gamma_\ell = \zeta \Gamma_d$) and, thus, it is pivotal for $\mathbb{V}ar(\varepsilon_{\ell,t}) = \Gamma_\ell \Gamma_\ell'$ (see Subsection 2.5.3).

Figure 2.Q1 shows the fiscal limit estimates under the baseline parametrization (grey thick solid lines) and across the above-described sensitivity exercises, together with debt-to-GDP ratios (black solid lines).

Figure 2.P1: Partial SJG Eurobonds: baseline, excluding "super" core (Germany and Netherlands) and excluding periphery (Italy and Spain).



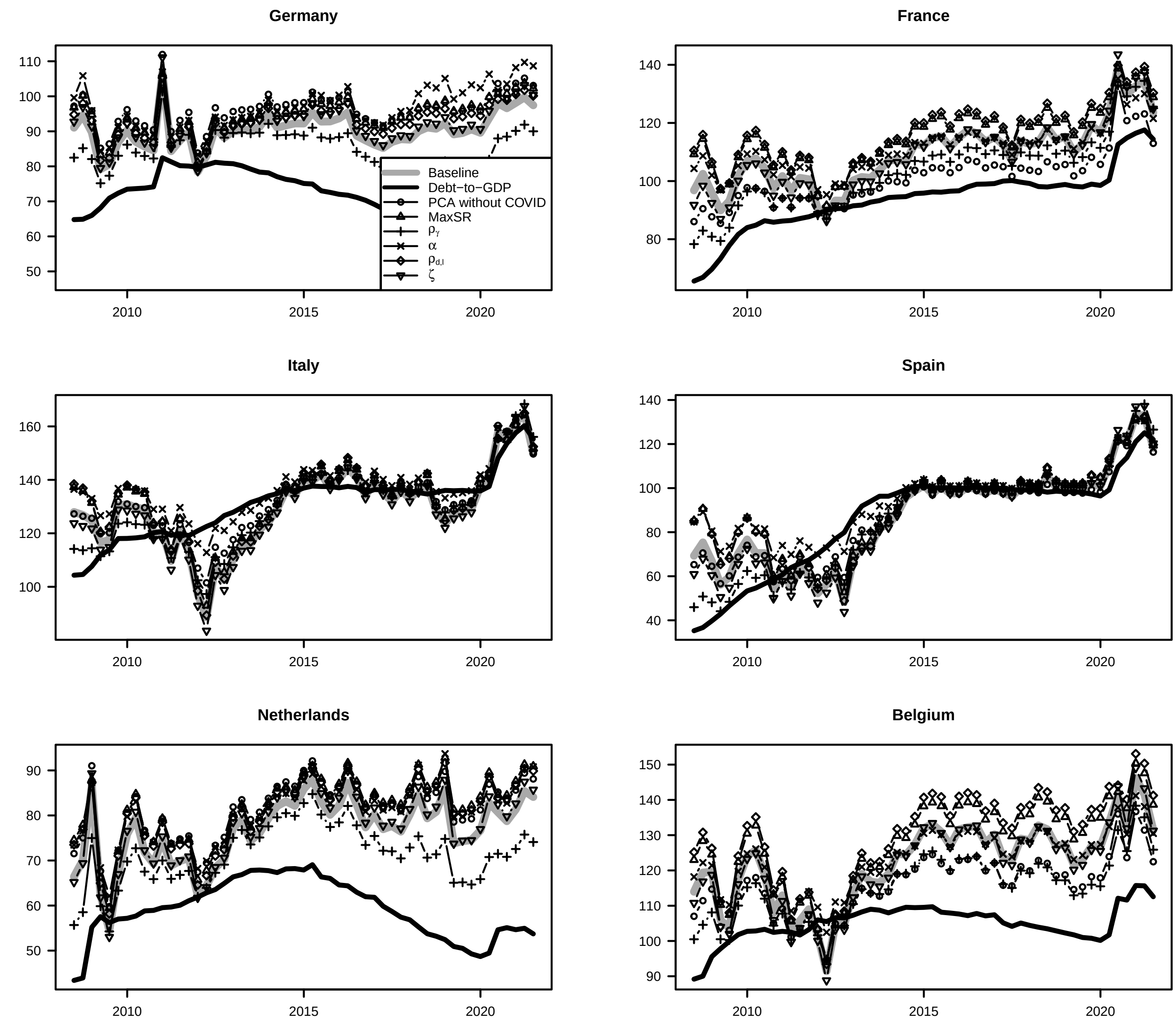
This figure shows the yield spreads in basis points across different maturities (3-, 5- and 10-year maturity) associated with SNJG (grey lines) and SJG (black lines) bonds under three different scenarios: (i) under the baseline scenario, where all countries participate in the emission, as in the main results presented in Section 2.2.1 (solid lines); (ii) under the scenario in which Germany and Netherlands do not participate in the issuance (dashed lines); and (iii) under the scenario in which peripheral member states are excluded from the joint issuance program (dotted lines).

Units are in percent of GDP. While the different parametrizations tend to result in shifts in the estimated fiscal limits, it appears that the fluctuations are fairly consistent across the different specifications.

Figure 2.Q2 displays the yield spreads of SJG bonds across the sensitivity exercises and for the baseline estimation (grey thick solid line). The three panels correspond to different maturities: 3, 5, and 10 years. The blue line corresponds to a model-free approximation of the SNJG bond spread, computed as the GDP-weighted average of the national bond spreads. We observe that the order of magnitude of the SJG-vs-SNJG spreads is fairly robust under different model parametrizations. This is confirmed by Fig-

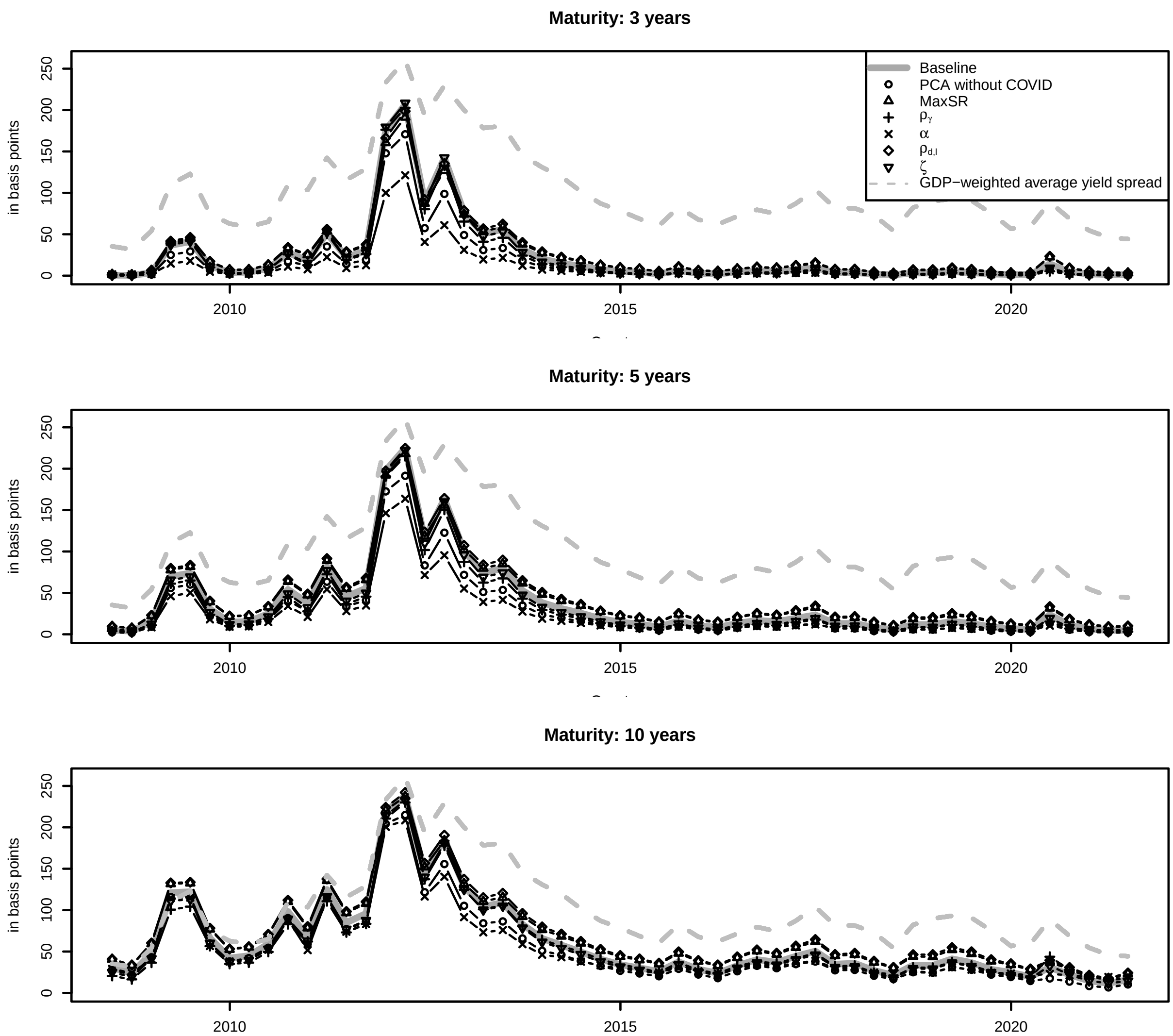
ure 2.Q3, that shows the 3-, 5-, and 10-year maturity yield gains associated with redistribution Scheme C (same yield gain across countries, see Subsection 2.6.2).

Figure 2.Q1: Fiscal limit estimates - Sensitivity analysis



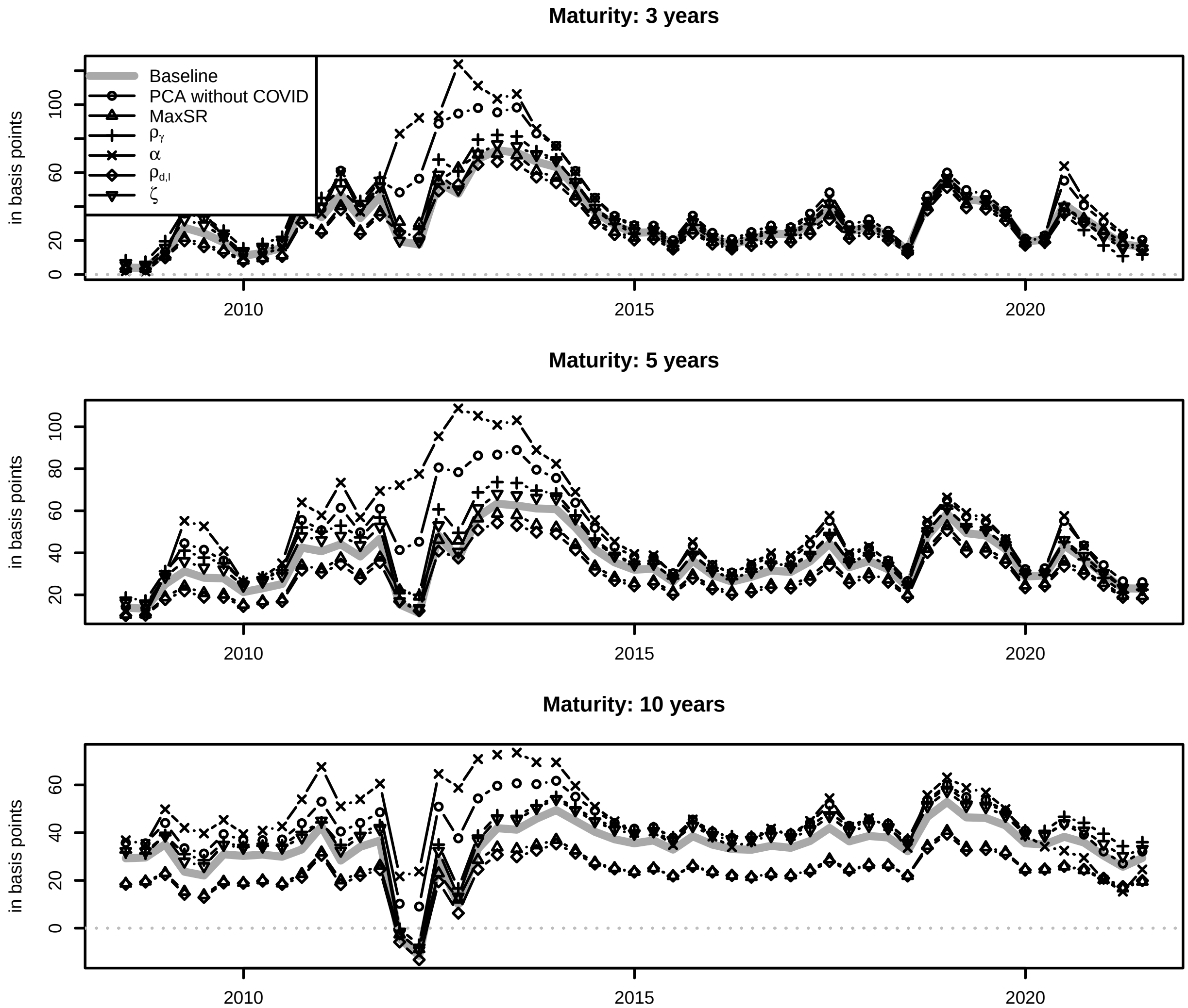
This figure shows the fiscal limit estimates under the baseline parametrization (grey thick solid lines) and across different sensitivity exercises: (i) exclusion of COVID period (after 2020Q1) from the PCA analysis of the estimated $\epsilon_{d,j,t}$'s (black line with circles) so that $\mathbb{V}ar(\varepsilon_{d,t}) = \Gamma_d \Gamma_d'$ is modified (Γ_d represents the matrix of PCA weights, see Subsection 2.5.3 for details regarding Γ_d.); (ii) the lower bound for $maxSR$ is set to 0.25 (black line with upward-facing triangles), considering that this bound is binding under the baseline model; (iii) the lower bound for ρ_γ is shifted to 0.5, instead of 0.7 (black crossed line), given that the lower bound is binding in the baseline parametrization; (iv) α is set to 0.2 (black line with "x" marks), which corresponds to more than double the estimated value (see Table 2.3 in Sec. 2.5); (v) the correlation between the two "main common shocks" ($\eta_{d,1,t}$ and $\eta_{\ell,1,t}$), $\rho_{d,\ell}$, is set to zero (black line with rhombuses); (vi) the upper bound for ζ is shifted from 1.5 to 2 (black line with downward-facing triangles) given that this parameter is constrained at its upper bound under the baseline parametrization (for more details on ζ, see Subsection 2.5.3). Debt-to-GDP ratios for each country are also plotted (black solid lines). Units are in percent of GDP.

Figure 2.Q2: SJG bond yield spreads - Sensitivity analysis



This figure shows SJG bond yield spreads for the 3-, 5- and 10-year maturity under the baseline parametrization (grey thick solid lines) and across different sensitivity exercises: (i) exclusion of COVID period (after 2020Q1) from the PCA analysis of the estimated $\epsilon_{d,j,t}$'s (black line with circles) so that $\mathbb{V}ar(\varepsilon_{d,t}) = \Gamma_d \Gamma_d'$ is modified (Γ_d represents the matrix of PCA weights, see Subsection 2.5.3 for details regarding Γ_d.); (ii) the lower bound for $maxSR$ is set to 0.25 (black line with upward-facing triangles), considering that this bound is binding under the baseline model; (iii) the lower bound for ρ_γ is shifted to 0.5, instead of 0.7 (black crossed line), given that the lower bound is binding in the baseline parametrization; (iv) α is set to 0.2 (black line with "x" marks), which corresponds to more than double the estimated value (see Table 2.3 in Sec. 2.5); (v) the correlation between the two "main common shocks" ($\eta_{d,1,t}$ and $\eta_{\ell,1,t}$), $\rho_{d,\ell}$, is set to zero (black line with rhombuses); (vi) the upper bound for ζ is shifted from 1.5 to 2 (black line with downward-facing triangles) given that this parameter is constrained at its upper bound under the baseline parametrization (for more details on ζ, see Subsection 2.5.3). The figure also reports the GDP-weighted average of the observed yield spreads across maturities (grey dashed lines), which are close to SNJG bond yield spreads (see Figure 2.5 in Subsection 2.6.1). Units are in basis points.

Figure 2.Q3: Yield gains associated with redistribution scheme with same yield gains across countries - Sensitivity analysis



This figure shows yield gains associated with redistribution scheme with same yield gains across countries for the 3-, 5- and 10-year maturity (redistribution scheme C as described in Subsection 2.6.2) under the baseline parametrization (grey thick solid lines) and across different sensitivity exercises: (i) exclusion of COVID period (after 2020Q1) from the PCA analysis of the estimated $\epsilon_{d,j,t}$'s (black line with circles) so that $\mathbb{V}ar(\varepsilon_{d,t}) = \Gamma_d \Gamma_d'$ is modified (Γ_d represents the matrix of PCA weights, see Subsection 2.5.3 for details regarding Γ_d.); (ii) the lower bound for $maxSR$ is set to 0.25 (black line with upward-facing triangles), considering that this bound is binding under the baseline model; (iii) the lower bound for ρ_γ is shifted to 0.5, instead of 0.7 (black crossed line), given that the lower bound is binding in the baseline parametrization; (iv) α is set to 0.2 (black line with "x" marks), which corresponds to more than double the estimated value (see Table 2.3 in Sec. 2.5); (v) the correlation between the two "main common shocks" ($\eta_{d,1,t}$ and $\eta_{\ell,1,t}$), $\rho_{d,\ell}$, is set to zero (black line with rhombuses); (vi) the upper bound for ζ is shifted from 1.5 to 2 (black line with downward-facing triangles) given that this parameter is constrained at its upper bound under the baseline parametrization (for more details on ζ, see Subsection 2.5.3). Units are in basis points.

Chapter 3

Fiscal Space and the size of the Fiscal Multiplier[1]

3.1 Introduction

Since the Great Recession, the debate on the role of fiscal policy has gained traction, as discretionary fiscal measures have started afresh to serve as policy tools in advanced economies. Large spending plans have been implemented in many advanced economies and especially in the United States. However, growing deficits piled up into unprecedented levels of public debt. The latter, together with stagnant growth and low inflation, raised the attention on the sustainability of public finances and called into question whether the effects of fiscal policy as well might depend on fiscal sustainability considerations. According to this view, fiscal policy can prove to be a powerful tool in certain situations while not in others. In particular, an expansion in the public budget associated with a weak fiscal position can even produce harmful effects, while, at the opposite, the same fiscal shock implemented when public finances are sound generates expansionary effects. In this paper, we investigate such hypothesis empirically. Addressing this conjecture is key also in light of the effects of the COVID-19 crisis. Governments around the world have expanded massively their budget deficits in an effort to mitigate the detrimental consequences of the pandemic. While these policies were certainly necessary to face the emergency, they will generate a strong deterioration in public finance sustainability in the medium-term, which might affect the effectiveness of fiscal policy in the future. Indeed, seminal contribution from Perotti (1999) already pointed out that shocks to government expen-

[1] This chapter is coauthored with Luca Metelli from the Bank of Italy.

diture in times of fiscal stress have very different effects on the economy than in normal times.[2]

This empirical work studies whether the transmission mechanism of fiscal policy is affected by the state of fiscal sustainability. In order to take into account fiscal sustainability we refer to the notion of fiscal space. Heller (2005) provides the following definition for fiscal space:[3]

> *Room in a government's budget that allows it to provide resources for a desired purpose without jeopardizing the sustainability of its financial position or the stability of the economy.*

However, there is no agreement on how to translate such a loose notion into a proper measure. In the definition provided above, the link to the concept of fiscal sustainability is explicit. This relates to the ability of the government to fund its preferred spending programs, while being able to service its obligations and to ensure solvency. On the contrary, Bi (2012), Bi and Leeper (2013) and Ghosh et al. (2013) regard fiscal space as the distance between the level of current debt-to-GDP ratio and a country specific debt limit, which represents the maximum amount of debt that an economy can credibly sustain. Additionally, Perotti (2007) delineates the concept of fiscal space as a different approach in setting up the intertemporal budget constraint. In front of these multiple interpretations regarding fiscal space, we propose different methods to track its evolution over time, each relating to different underlying theories, as we clarify later.

While the literature focused on debt-to-GDP ratio to capture the role of fiscal sustainability in the transmission of fiscal policy, we focus on fiscal space.[4] The key characteristic of fiscal space is to incorporate considerations on the ability of the government to service its debt, which relates, in turn, to the dynamics of macroeconomic variables like the interest rate together with perceived sovereign risk, the GDP growth rate, the amount of primary surplus and the ratio of debt-to-GDP. Although correlated with debt-to-GDP, fiscal space encompasses other crucial aspects. First of all, it measures the overall ability of the government to service its obligations, which depends only in part on the level of public debt to be repaid. Second, it considers public finance aggregates jointly with other key macroeconomic variables, taking also into account the fundamental debt capacity of the economy. Lastly, the forward-looking nature of fiscal space contrasts with the path-dependent nature of a stock variable like debt-to-GDP. Indeed

[2] Previous studies (Blanchard, 1990; Sutherland, 1997) showed that fiscal consolidation in the form of tax increases leads to non-Keynesian effects in times of weak fiscal position (namely, high debt or deficits).

[3] A similar definition can be found in Ley (2009) and Escolano (2010).

[4] E.g., Huidrom et al. (2020),Ilzetzki et al. (2013),Auerbach and Gorodnichenko (2017)

fiscal space varies with market and economic conditions, which often change abruptly. For instance, a productive fiscal stimulus could improve the economic outlook in a country and consequently free-up fiscal space while, instead, the dynamic of debt-to-GDP ratio could still follow an upward trend in the short run; on the contrary, a harmful fiscal tightening could ameliorate a country's indebtedness while worsening growth and, thus, reducing the perceived fiscal room.

The idea of a differential effect of fiscal policy according to the fiscal position fits in the more general debate on fiscal policy, which has established a consensus on the fact that there is no such thing as a unique fiscal multiplier, but, more likely, the effects of fiscal shocks are state dependent. This literature, however, focused mainly on studying how fiscal policy's effects vary with the business cycle, differentiating in particular recession versus expansion periods, and with the monetary policy stance, with a particular reference to periods in which this is constrained by the zero lower bound. By contrast, there are few studies considering fiscal sustainability as a state variable in the transmission of fiscal policy. In this paper, we investigate this form of state dependency and we aim at answering the following questions. How can we measure the evolution of fiscal space over time? Do the effects of fiscal shocks depend upon the level of fiscal space? If so, what is the rationale behind such differential effect?

The paper addresses the aforementioned questions in the following way. In the first part, we build four different indicators to measure the evolution of fiscal space over time, using data for the US. Our preferred indicator relates to the concept of primary surplus sustainability gap as in Kose et al. (2017) and we calculate, at each point in time, the primary surplus needed to stabilize the trajectory of public debt. Such simple yet effective indicator captures well periods of high debt velocity and inherent inability to contain debt roll-over needs via primary surpluses. The second indicator draws from the theoretical literature on fiscal limits as in Bi (2012) and Bi and Leeper (2013) and represents its empirical counterpart. The third one relates to the idea of fiscal imbalances of Auerbach (1997), while the last one builds on Aizenman and Jinjarak (2010). We refer to Section 3.3 for further details on the computation of these indicators. After the construction of such fiscal space indicators, we show that they correlate well among each other, reinforcing the idea that, although derived from different underlying theories, they capture slightly different aspects of the same phenomenon we want to measure, i.e. the evolution over time of fiscal sustainability. We also provide evidence that our measures do not confound with other cyclical indicators, like the economic cycle or ZLB periods. In the second part of the paper, we estimate empirically the effects of fiscal policy in a state-dependent fashion, differentiating periods in which fiscal space is tight and when fiscal space is ample. We estimate the effects of government spending shocks in the US for the period 1929-2015, using two different identification methods, the one of Ramey (2011b,a) and the one of Blanchard and Perotti (2002). We then employ the Local-Projections method developed by Jordà (2005) to estimate the state-dependent effect of fiscal policy, using our four indicators to define, for each proxy, a tight fiscal space state when the proxy is above the median and a large fiscal space state when it is below. We quantify the impact of fiscal policy in the large and tight fiscal space state calculating the cumulative fiscal multiplier, as in Ramey and Zubairy (2018). We also investigate the mechanism behind our results, analyzing the effects on other variables other than output, like private consumption, private investment, interest rates and debt-to-GDP ratio. In this regard, we calculate multipliers also for consumption and investment.

The main results of the paper are the following. First, we find that fiscal policy is much more effective when implemented in periods associated with large fiscal space. The corresponding fiscal multiplier is above one, while by contrast, when fiscal space is tight, the fiscal multiplier is smaller than one, with a difference in the two cases always statistically significant. This result is particularly relevant in light of the findings by other studies on non-linearities in fiscal policy. While Ramey and Zubairy (2018) conclude there is no difference in the effects of fiscal policy between expansion and recession and find only minor dissimilarities when monetary policy is at the ZLB, our paper identifies a major distinction in the size of fiscal multiplier across different fiscal space regimes, suggesting that such non-linearity might be the relevant one. Second, we show that our result occurs independently of the identification method adopted and is robust across different samples and different empirical specifications. More importantly, the result is strikingly similar across the four fiscal space indicators we construct, signaling that our indicators capture the same phenomenon. We also implement the estimations using debt-to-GDP ratio as the state variable. In this case, we do not find difference in the two states, suggesting the importance of looking at specific indicators of fiscal space, in contrast to other variables, when studying fiscal sustainability issues. Finally, the paper represents a first step to investigate the rationale behind the differential response of fiscal policy in the two states. We show that private consumption and investment follow a very different behaviour across states. Indeed, in case of ample fiscal space, government spending shock does not generate the standard Ricardian effect as in the case of weak fiscal space but, on the contrary, produces an increase in private consumption and does not crowd out investment.

This paper contributes to the literature in the following ways. First, we provide a historical time series for fiscal space in the US starting from 1929 up to 2015, at the quarterly frequency, according to four different indicators. The only database containing time-varying fiscal space measures, Kose et al. (2017), is at annual frequency and starts from 2001. We extend backwards some of such measures and we provide other fiscal space proxies drawing from multiple notions, while documenting the variability over time of fiscal space in the US. Second, this is the first paper to investigate empirically the effects of fiscal policy according to fiscal space conditions. While the literature estimated fiscal multipliers according to different levels of public debt, we do so using fiscal space. Finally, we also delve into an investigation of the transmission on other key variables, highlighting the striking difference in the response of private consumption and investment across states of fiscal space.

The rest of the paper is organized as follows. Section 3.2 provides a literature review, while Section 3.3 describes carefully the procedure we follow to construct our fiscal space indicators and their properties. Section 3.4 provides details on the empirical methodology, the data and the identification method adopted. Section 3.5 presents the empirical results, together with a robustness section. Finally, Section 3.6 concludes.

3.2 Literature Review

Our paper is related to different strands of the literature. First of all, we relate to the literature, both empirical and theoretical, aiming at investigating and measuring the concept of fiscal space. While no unique definition of fiscal space exists, its core aspect is to measure the debt service capacity of a country. The latter hinges on many dimensions: budget position, financing needs, spending and revenue prospects, resilience to contingent liabilities and access to markets. As correctly reported in Botev et al. (2016), fiscal space can be measured either in terms of losing market access or achieving long-term sustainability.

According to the former dimension, fiscal space is deemed as the distance between the actual debt and the debt limit for which the government would be unable to roll-over debt and, thus, lose market access. Such interpretation is more suited for a theoretical approach, given the need to define and derive the debt limit notion. Moreover, the lack of credit events in advanced economies complicates the empirical estimation of fiscal space in terms of the market access dimension, making this approach viable mainly through a theoretical framework.[5] Ostry, Ghosh, Kim, and Qureshi (2010), Ghosh, Kim, Mendoza, Ostry, and Qureshi (2013) and Ostry, Ghosh, and Espinoza (2015) compute static estimates for debt limits based on the observation that the higher the levels of debt, the weaker the reaction of primary surpluses ("fiscal fatigue").[6,7] On the other hand, in Bi (2012), Leeper (2013), Bi and Leeper (2013), Bi and Traum (2012) and Bi and Traum (2014), the theoretical fiscal limit corresponds to the discounted present value of future maximum primary surpluses.[8] Finally, Collard, Habib, and Rochet (2015) also exploit the idea of a maximum primary surplus to derive a static measure of debt limit. We relate to this literature by exploiting the concept of maximum primary surplus to construct one of our indicators.

According to the long-term sustainability dimension, instead, fiscal space is measured by the degree of public finance sustainability a country features. Kose et al. (2017) propose a large cross-sectional database covering the core aspects of government debt sustainability, like perceived sovereign risk, market access, balance sheet composition, external and private debt considerations. Among other indicators, Kose et al. (2017) reports also the so-called *de facto* fiscal space as derived in Aizenman and Jinjarak (2010) and Aizenman et al. (2013), which delineates the government's ability to raise tax revenues to contain public debt and deficits. We relate to Kose et al. (2017) and Aizenman and Jinjarak (2010) for the construction of two of our proxies for fiscal space in the US. Moreover, in Auerbach (1997), but also in Gale and Auerbach (2009) and Auerbach and Gorodnichenko (2017), the authors approximate fiscal room in the US by measuring the size of fiscal distress via the government intertemporal budget constraint.[9] We draw from this approach for building another measure of fiscal space.

Our paper is also related to the literature studying how the fiscal position affects the transmission of fiscal policy. A few papers have analyzed this aspect. On the theoretical side, in

[5] As far as we know, the only paper that manages to exploit the time-variation of sovereign credit data for advanced economies in order to estimate both debt limits and fiscal space in terms of the market access dimension is Pallara and Renne (2020).

[6] Their approach provides static debt limit and fiscal space estimates. These studies consider two calibration periods: from 1970 to 2007 and from 1985 to 2007. Moreover, updated fiscal space static estimates are regularly reported in Moody's (2011).

[7] Note that these debt limits cannot be retrieved by a model-free estimation.

[8] Maximum surpluses arise if the government can steer tax revenues at the peak points of the Laffer curve (Trabandt and Uhlig, 2011).

[9] More details are provided in Section 3.3.3.

the seminal contribution from Perotti (1999), the author builds a simple model where government expenditure shocks have a positive, non-Keynesian correlation in fiscal stress times. Symmetrically, tax shocks have a negative, Keynesian correlation in normal times and a positive, non-Keynesian correlation in fiscal stress times.[10] This study finds a strong evidence that expenditure shocks have Keynesian effects when the level of public debt or deficits is low, and non-Keynesian effects in the opposite circumstances. In short, the more burdensome the state of fiscal stress the more a positive spending shock will lead to a steeper path of future expected tax changes and, thus, to a lower present value of consumers' wealth. On the empirical side, Huidrom et al. (2020) is the first paper to systematically show that government spending multipliers decrease with the worsening of the fiscal position, as proxied by the level of debt-to-GDP. Using a panel of countries going from 1980 until 2014, the authors estimate an Interacted Panel VAR including 33 countries and find multipliers around zero when the fiscal position is weak. We see this contemporaneous work of Huidrom et al. (2020) as the closest to ours. The main difference is that we refine the idea of fiscal position using the notion of fiscal space and we focus on the United States. We also show that, at least in the case of the US, if we were to identify fiscal position through the debt-to-GDP metric, we would not obtain the dichotomic result in terms of fiscal multipliers that we instead find adopting the concept of fiscal space.[11] Moreover, Huidrom et al. (2020) adopt a narrow definition of weak and large fiscal position, defined respectively as the 90th and 10th percentile of debt-to-GDP ratios, documenting the differences in the transmission of fiscal policy only in such extreme cases, while we focus on a more ample range of scenarios. Another important paper is Ilzetzki et al. (2013), who investigated earlier the effects of debt in the transmission of fiscal policy. The authors estimate fiscal multipliers according to various dimensions in a large panel of countries spanning from the '60s until before the Great Financial Crisis. Finally, Auerbach and Gorodnichenko (2017) employ a local projection model for 25 OECD countries to study the multiplier in different government debt states. Using a sample spanning from 2003 until 2017, they estimate significant government spending multipliers above (below) 0 in low (high) public debt states. All of the aforementioned papers aim at studying the interaction between fiscal sustainability and fiscal policy transmission. However, they all use the narrower measure of debt-to-GDP ratio as a state variable, which cannot convey enough information on the available fiscal room. To the

[10] Two other existing models (Blanchard, 1990; Sutherland, 1997) formalize the non-Keynesian effects of tax hikes at high levels of public debt, but the model presented in Perotti (1999) allows for both tax and government spending shocks to have non-Keynesian effects on private consumption via the expectations channel.

[11] See the results in the Appendix, Tables 3.H1 - 3.H4.

best of our knowledge, our paper is the first to adopt a broader concept of fiscal space to study the transmission of fiscal shocks.

More in general, our paper relates to the literature studying state-dependency in fiscal policy both empirically and theoretically. Such literature, however, has focused mainly on investigating the role of the business cycle as a state. The two main empirical contributions on this topic are, on the one hand, Auerbach and Gorodnichenko (2012) and Auerbach and Gorodnichenko (2013) and, on the other hand, Ramey and Zubairy (2018).[12] Such papers deliver very different conclusions and do not reach a consensus in empirical research regarding the effects of expansionary fiscal shocks under slacks and booms. Indeed, using a regime-switching VAR approach, Auerbach and Gorodnichenko (2012) and Auerbach and Gorodnichenko (2013) find large differences between multipliers in recessions and expansions.[13] However, the adopted econometric model requires to assume for how many quarters the impulse response function should remain in each state of the economy. This could lead to distorted results in favor of an artificially higher multiplier in recession. By contrast, Ramey and Zubairy (2018) conclude that there is no difference in the size of multipliers across the business cycle. Using the Jordà (2005) local projection approach and high unemployment rate as proxy for recession, they show that government spending multipliers range between 0.3 and 0.8 no matter the state of the business cycle. The authors also find mixed evidence on the size of the fiscal multiplier at the zero lower bound.[14] Our paper draws from the approach developed in Ramey and Zubairy (2018) to estimate fiscal multipliers, but we instead find a major role for state dependency, suggesting that fiscal space could be the relevant state for the transmission of fiscal shocks.

[12]In macroeconomic theory, very few papers (e.g., Michaillat, 2014; Albertini et al., 2019) concentrated on how recessions and expansions affect the size of the fiscal multipliers. The study of business cycle-dependent fiscal multipliers parallels Keynesian theory, where government expenditure shocks have stronger expansionary outcomes during recessions as crowding out of private spending and investment is attenuated by the slack state.

[13]Indeed, the authors observe that the multiplier is much higher in recessions rather than in expansions. They report the multiplier to be as high as 2.5.

[14]In Ramey and Zubairy (2018), ZLB state is defined as the quarters in which the T-bill rate is equal to or below 50 basis points. When the authors use the full sample spanning from 1889:Q1 until 2015:Q4, the multiplier is not higher at the zero lower bound; while, excluding the World War II, they found a multiplier as high as 1.5 in the ZLB state.

3.3 Fiscal Space approximations

In this section, we describe our four historical indicators of fiscal space for the US. Two of them are based on model-free estimates of fiscal space, drawing from Kose et al. (2017), Aizenman and Jinjarak (2010) and Aizenman et al. (2013). Such measures focus on fiscal sustainability, the revenue capacity of the government, fiscal policy stance and, indirectly, market access. One additional fiscal space proxy derives from the concept of fiscal imbalance as in Auerbach (1997). One final measure draws from the concept of fiscal (or debt) limit described in Bi (2012), Bi and Leeper (2013) and Ostry et al. (2010).

All of our indicators, as we clarify in the following sections, depend, among other key factors, on public debt, the surplus/deficit and, more in general, on government finance variables, all series that are highly correlated with expansions and recessions. Given that this is a crucial concern to avoid confounding effects when we carry the empirical investigation, we need to adjust these series to purify the government's fiscal position from business cycle fluctuations. Therefore, we cyclically adjust the fiscal variables involved in the computation of our fiscal space indicators.[15] This correction ensures that we capture the discretionary dimension of fiscal room in the US. Moreover, one could think of the proposed indicators as resulting from different specifications of the government fiscal reaction function (Bohn, 1998) and of the target horizon of the government budget constraint.[16] Recall the debt accumulation accounting equation:

$$\Delta \frac{B_t}{Y_t} \approx \frac{i_t - \gamma_t}{1+\gamma_t} \frac{B_{t-1}}{Y_{t-1}} - s_t, \tag{3.1}$$

[15] Following the method implemented by the World Bank and reported in Kose et al. (2017), we cyclically adjust the government finance statistics variables by multiplying them by $(1+\tilde{y})^{-(\epsilon_x - 1)}$ where $\tilde{y}$ is the difference between the actual GDP and the CBO potential output as % of potential output; ϵ_x stands for the output gap elasticity of x for $x =$[Revenues, Government spending, Federal debt]. We use World Bank (see also Kose et al., 2017) elasticities for revenues and government spending, respectively equal to 1 and 0.1. For what concerns federal debt, we estimate the elasticity to be not significantly different from 0 and, thus, we assume it to be equal to 0. Note that the elasticities for spending and revenues proposed here are not distant from the estimates in Girouard and André (2006).

[16] A fiscal reaction function describes how governments react to the accumulation of debt and how corrective measures are taken. In the following discussion, rather than estimating the fiscal reaction function (e.g., Bohn, 1998), we focus on the rule chosen by the government on the stance of its fiscal reaction function.

where $\frac{B_t}{Y_t}$ is the debt-to-GDP, γ_t is the growth rate of nominal GDP, i_t is the nominal interest rate and s_t is the primary balance over GDP.[17] By solving forward, we obtain the following:

$$\mathbb{E}_t \frac{B_{t+H}}{Y_{t+H}} \approx \prod_{h=0}^{H} \frac{1+i_{t+h}}{1+\gamma_{t+h}} \frac{B_{t-1}}{Y_{t-1}} - \sum_{h=0}^{H} \prod_{j=1}^{h} \frac{1+i_{t+j}}{1+\gamma_{t+j}} s_{t+h}. \tag{3.2}$$

Indeed, different specifications of equation 3.2 in terms of government fiscal reaction s_{t+h} and horizon H suggest different fiscal space indicators.

3.3.1 Baseline indicator: primary surplus sustainability gap

Our first indicator draws from Kose et al. (2017). By setting to 1 the horizon H in equation 3.2, we obtain the simple debt accumulation equation (eq. 3.1). Using this equation, we calculate the level of the primary surplus required, in each quarter, to stabilize public debt, i.e. to make $\Delta \frac{B_t}{Y_t} = 0$. Thus, we assume that the optimal fiscal reaction function of the government is to stabilize debt at each horizon. We then define our fiscal space indicator as the distance between such primary surplus and the realized one. Using equation 3.1 and cyclically adjusting the variables as described in the previous section, our proxy is given by the following equation:

$$FS_{1,t} = \left(\frac{i_t - \tilde{\gamma}_t}{1+\tilde{\gamma}_t} \right) d_{t-1}^{c.a.} - s_t^{c.a.}, \tag{3.3}$$

where $s_t^{c.a.}$ is the cyclically adjusted primary surplus over potential GDP, $d_t^{c.a.}$ is the cyclically adjusted debt over potential GDP, $\tilde{\gamma}_t$ is the nominal potential GDP growth and i_t is the historical interest rate on 10-year maturity US government bonds.[18,19] This gap incorporates information on the fiscal stance, public debt acceleration and the difference between the interest rate and GDP growth. This last measure represents the first nod to the study of fiscal sustainability in macroeconomics and contributes in a relevant way to the dynamics of our fiscal space indica-

[17] Eq. 3.1 should include also the stock-flow adjustments not to hold as an approximation. Stock-flow adjustments comprise of factors that affect debt but are not included in the budget balance (such as acquisitions or sales of financial assets). For sake of simplicity, we focus on the "snowball-effect" side of the debt accumulation equation and on the government budget balance.

[18] As potential GDP we take the CBO potential GDP estimates. Additionally, we also implemented sensitivity analysis using different potential GDP measures (e.g., sixth-degree polynomial for the logarithm of GDP) that leaves unaffected both the dynamics of the fiscal space measure and the econometric results of the paper.

[19] We acknowledge that the average debt maturity for the US is around 6 years and using the 5-year maturity yield would be more precise. However, we have only historical data points for the 10-year maturity yield, which correlates more than 95% with the 5-year maturity yield.

tor.[20] We regard this measure as the benchmark indicator of fiscal space since it summarizes the many features a proxy for fiscal space should contain: considerations of fiscal sustainability, debt dynamics, interest rate, output growth and fiscal policy stance. In particular, this fiscal space measure highlights times of rapid debt accumulation due to inherent inability to roll-over debt via primary surpluses, crucial characteristics of the fiscal position of the government. We estimate FS_1 from 1889Q1 to explore the history of fiscal room in the United States. Figure 3.1 reports the results. The figure shows that fiscal space was especially tight over the depression of the early 20s, the Second World War and started to worsen from 2001 onward and, in particular, during the Great Financial Crisis.

3.3.2 Indicator II: Laffer curve peak-implied surplus gap

According to the literature on fiscal limits, fiscal space is defined by the distance between the actual debt and the maximum amount of debt the government can sustain, i.e. the debt limit. In Bi (2012) and Bi and Leeper (2013), this limit is defined by the discounted projected path of maximum primary surpluses implicit in the peak of the Laffer curve.[21,22] We exploit the concept of debt limit as described in Bi (2012) and Collard et al. (2015), but we apply this intuition on a quarter-by-quarter perspective. Indeed, we define and calculate below, for each quarter, the maximum primary surplus attainable in the US. This amounts into setting again to 1 the horizon H in equation 3.2 and assuming that the optimal reaction function of the government is to attain the peak of the Laffer curve, namely maximum surplus, in each horizon. Thus, we define fiscal

[20] According to Blanchard (2019b), as long as the yields are lower than the GDP growth rate, even in the current low-growth environment, countries have fiscal space (see also Mauro and Zhou, 2019). However, there is a growing consensus that such argument is incomplete. For instance, the simple measure of $i - \gamma$ does not consider the evolution of the primary balance itself and the stock-flow adjustments. Moreover Jiang et al. (2019) find that the discount factor on government debt is decoupled from the yields on bonds, which would nuance the claims in Blanchard (2019b).

[21] In Bi (2012) and Bi and Leeper (2013), fiscal limit is defined as follows,

$$\ell_t = \mathbb{E}_t \left(\sum_{j=1}^{+\infty} \exp(\gamma_t - i_t + \cdots + \gamma_{t+j-1} - i_{t+j-1}) s^*_{t+j} \right), \tag{3.4}$$

where ℓ_t is the fiscal limit-to-GDP, γ_t stands for nominal growth, i_t is the risk-free rate and s^*_t is the maximum surplus over GDP. The maximum primary surplus s^*_t is the surplus implicit at the peak of the Laffer curve.

[22] The Laffer curve represents the reverse bell-shaped relationship between the average tax rate and government revenues.

Figure 3.1: FS_1 (1889:1-2015:4): Primary surplus sustainability gap

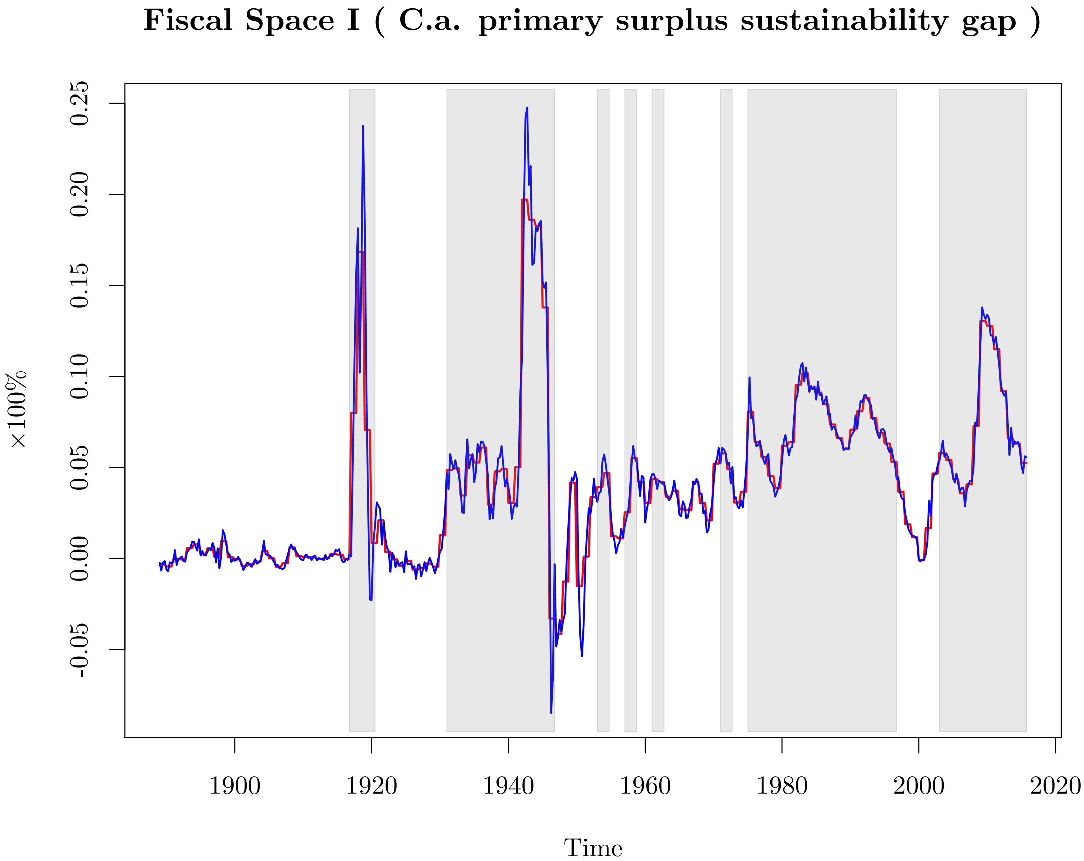


Figure 3.1 shows the estimates for indicator FS_1 (in blue) and its 1-year moving average (in red). The series are expressed in percent of potential output. Shaded regions indicate periods when fiscal space is tight (1-year $FS >$ median).

space as the difference between such maximum surplus and the realized primary deficit.[23] This measure captures how far is the government from revenue maximization.

We calculate maximum government revenues using Laffer curve-peak tax rates estimates as in Trabandt and Uhlig (2011) for labour, capital and consumption:

$$\hat{T}_t = \tau_K^\star TB_t^K + \tau_L^\star TB_t^L + \tau_c^\star TB_t^c, \tag{3.5}$$

where $\tau_K^\star$, $\tau_L^\star$ and $\tau_c^\star$ stand for the peak tax rates.[24] $\hat{T}_t$ represents the maximum revenues and TB_t^i ($i = K, L, c$) are the tax bases. Then, we compute the (cyclically adjusted) maximum surplus as follows:[25]

$$\hat{s}_t = (\hat{t}_t^{c.a.} - g_t^{c.a.}), \tag{3.6}$$

where $\hat{t}^{c.a.}$ and $g_t^{c.a.}$ are the cyclically adjusted maximum revenues and government spending, respectively. Finally, our fiscal space indicator is defined as:

$$FS_{2,t} = \hat{s}_t - s_t^{c.a.}, \tag{3.7}$$

Figure 3.A1 in Appendix 3.A reports the estimates for FS_2 expressed in percent of potential output.

3.3.3 Indicator III: fiscal imbalance

Auerbach (1997) first proposed a measure to quantify US fiscal imbalances by taking into account the impact of future government expenditure (such as contingent liabilities/implicit spending) and revenues on the intertemporal government budget constraint. The same approach has been employed in Gale and Auerbach (2009) and Auerbach and Gorodnichenko (2017). We adopt the idea of fiscal imbalance as described in Auerbach (1997) and we calculate, in each period, by what constant fraction of GDP taxes (revenues) would have to be increased

[23] Indeed, we can think of the debt-to-GDP ratio, $\frac{B_t}{Y_t}$, as the cumulative discounted stream of past deficits. Therefore, the distance between fiscal limit-to-GDP and $\frac{B_t}{Y_t}$ gives the size of the available fiscal space of the country.

[24] $\tau_K^\star$, $\tau_L^\star$ and $\tau_c^\star$ are respectively equal to 0.6, 0.52 and 0.05. These represent the tax rates estimated in the benchmark model in Trabandt and Uhlig (2011) for the US. $\tau_L^\star$ and $\tau_k^\star$ correspond to the benchmark Laffer curve model with Frisch elasticity equal to 3 and intertemporal elasticity of substitution equal to 2. For more details on the tax rates we refer to Appendix 3.D.2. Moreover, We decided to use ones of the lowest estimates because there is no explicit mention of compliance neither in the present work nor in Trabandt and Uhlig (2011), namely the higher the tax rate the more tax evasion and avoidance become relevant phenomena. Compliance is actually a meaningful issue concerning tax rates and revenues (Pappada and Zylberberg, 2017).

[25] See Section 3.3 for details on the cyclical adjustment.

for the government budget constraint to be satisfied when the dynamics of future spending is considered. In our approach, we take a 10-years horizon for the government budget constraint. Thus, this translates into setting to 10 years the horizon H in equation 3.2 and assuming that the optimal government fiscal reaction function is to balance future contingent liabilities over that projection horizon. Hence, we estimate the fiscal imbalance measuring the fiscal adjustment needed to satisfy the government budget constraint over a 10 years horizon:

$$B_t = (1+i_t)^{-[(t+H)-t]}\left(\frac{B_t}{GDP_t}\right)GDP_{t+H} - \sum_{k=0}^{t+H}(1+i_t)^{-(k+1-t)}(S_k + \Delta_t GDP_k), \tag{3.8}$$

where B_t is the total nominal government debt, S_t is surplus, i_t is the interest rate on the ten years maturity government bond and H is the last horizon (10 years, namely 40 quarters). Δ_t represents the quarterly fiscal imbalance as a percentage of GDP. The government budget constraint implies a projected path for purchases, revenues and income. These projections account for the foreseen dynamics in implicit spending for healthcare and the social security system. For the out-of-sample forecasts, we use CBO projections that consider spending for health and pensions under current and anticipated regulations.[26] The advantage of this fiscal space measure lies in its forward-looking nature, as it considers the projected path of the government budget. Figure 3.A2 in Appendix 3.A reports the estimates for FS_3 expressed in percent of potential output. The chart highlights the periods of major fiscal distress for the US government.

3.3.4 Indicator IV: *de facto* fiscal space

Aizenman and Jinjarak (2010) and Aizenman et al. (2013) build a measure to capture fiscal room defined as *de facto* fiscal space. Such measure is inversely related to the tax-years necessary to repay public debt or deficits and it is defined as the ratio of either public debt or the deficit over GDP to the *de facto* tax base. We build on the concept of *de facto* fiscal space to construct an indicator that uses current cyclically adjusted revenues and deficits run by the government. In our approach, we define *de facto* fiscal space in the following way:

$$FS_{4,t} = \frac{(deficit_t^{c.a.})}{(receipts_t^{c.a.})}, \tag{3.9}$$

[26] Since CBO projections are only publicly available from 2006 in electronic format, for the in-sample projections, we take the realized values of the considered variables assuming perfect foresight.

where $deficit_t^{c.a.}$ stands for the cyclically adjusted deficit-to-potential GDP and $receipts_t^{c.a.}$ represents the total realized government tax receipts over potential GDP.[27] The advantage of this measure is to provide insights on the actual tax capacity of a country to balance current deficits. In our framework, the fiscal reaction function subsumed in this measure has the objective of maximizing the revenue collection ability of the government over all horizons of the intertemporal budget constraint (eq. 3.2). FS_4 highlights periods of high deficit overhangs with respect to the government inability to raise revenues via tax collection. Figure 3.A3 in Appendix 3.A reports the estimates for FS_4.

3.3.5 Properties of the fiscal space indicators

For each fiscal space indicator, we generate a dummy variable equal to 1 (0) when our FS measure is above (below) its median value, meaning that fiscal space is tight (loose). In Table 3.A1 in Appendix 3.A, we calculate the correlations among such dummies to show how our indicators relate to each other. Dummy indicators FS_1 and FS_4 show the highest correlation (equal to 0.73 in median). FS_2 and FS_4 also display significant correlation (equal to 0.64 in median) since government revenues are key in their respective fiscal space definitions. Dummies FS_2 and FS_3 show very low median correlation (equal to 0.01) and we cannot reject that they are uncorrelated. The remaining cross-correlations among fiscal space dummies are above 30% in median. These results suggest that our indicators are mutually consistent while capturing different aspects of the evolution of fiscal space. Table 3.A2 in Appendix 3.A reports the correlations between fiscal space dummies and a broad set of relevant macroeconomic indicators, in particular NBER recession dates, zero lower bound (ZLB) dates and high/low federal debt-to-GDP ratio periods. The correlation between all four fiscal space measures and the NBER recession dates is approximately null, consistent with the fact that our proxies are purified from the transitory effects of the business cycle.[28] Our fiscal space measures are rather linked to the discretionary dimension of government finance variables and to medium/long run economic phenomena.[29] Tight fiscal space states also partially relate to ZLB periods.[30] This is mainly

[27] We use the latest CBO potential output estimates and we also estimated the fiscal space using a sixth-degree polynomial for the logarithm of GDP as real trend GDP as a robustness, which did not lead to changes in the dynamics of the estimates neither the interpretation of the fiscal space size over the sample.

[28] This is evident also from Figure 3.E1 in Appendix 3.E, which reports tight fiscal space periods across our FS dummies and recession events over time.

[29] In Table 3.E1 in Appendix 3.E, we also report the correlations of fiscal space series with unemployment and potential output.

[30] ZLB dummy indicates the state when the interest rate is at the zero lower bound or the FED is being very accommodative of fiscal policy (1932Q1-1951Q1, 2008Q4-2015Q4).

due to the Second World War and the Great Recession, in which both fiscal distress and low interest rates coexisted. However, although such correlations are positive, they are not particularly high. Table 3.A2 also shows that our measures are related to high/low government debt-to-GDP periods.[31] Such correlations range between 0.24 and 0.7, highlighting the role of public debt in the evolution of fiscal space.[32] However, debt-to-GDP is not sufficiently informative to capture the whole dynamics of fiscal space: other key factors such as output trends, the interest rate and the fiscal policy stance are also of crucial importance for identifying the level of fiscal room in the economy. Finally, in Table 3.A2 we also calculate the correlations of our indicators with a dummy variable equal to one when the US was involved in a major war and with another dummy variable capturing the party of the US president. None of these correlations are relevant, suggesting that our fiscal space indicators are not driven neither by military spending nor by the political cycle. Given that Bernardini and Peersman (2018) and Broner et al. (2021) find significant low government spending multipliers respectively under private debt overhangs and low share of public debt's foreign holding, we compare the latter states and periods of tight fiscal space. We find that there is no relevant overlapping aside from the Great Depression and the Great Financial Crisis for private debt overhangs only.

In Figure 3.2, we plot the periods in which fiscal space is identified as tight, according to each FS dummy. The panel at the center of the figure reports a similar information, however using debt-to-GDP, which is the standard indicator used by the literature to identify periods of fiscal distress. Two main results emerge from Figure 3.2. First, our measures are well related with each other, as most of the periods identified are common across the four indicators. This is especially the case for FS_1 and FS_4. Second, tight fiscal space periods identified by debt-to-GDP do not coincide with those identified by our method. Indeed, the concept of fiscal space refers to a broader notion of fiscal sustainability as opposed to the debt-to-GDP ratio. In fact, fiscal space takes into account the dynamics of other key macroeconomic variables and the fundamental debt capacity of the economy.

Finally, to further validate our fiscal space indicators, we also provide a narrative behind their evolution over time. Both dummies FS_1 and FS_4 indicate that the fiscal space was tight

[31] High (low) federal debt-to-GDP ratio means that the federal debt-to-GDP ratio is above (below) its median, which is equal to 40%.

[32] Additionally, in Table 3.E1 in Appendix 3.E, we can see that the fiscal space series are consistently correlated with debt-to-GDP series and the change in public debt.

Figure 3.2: FS dummies and High Debt-to-GDP ratio.

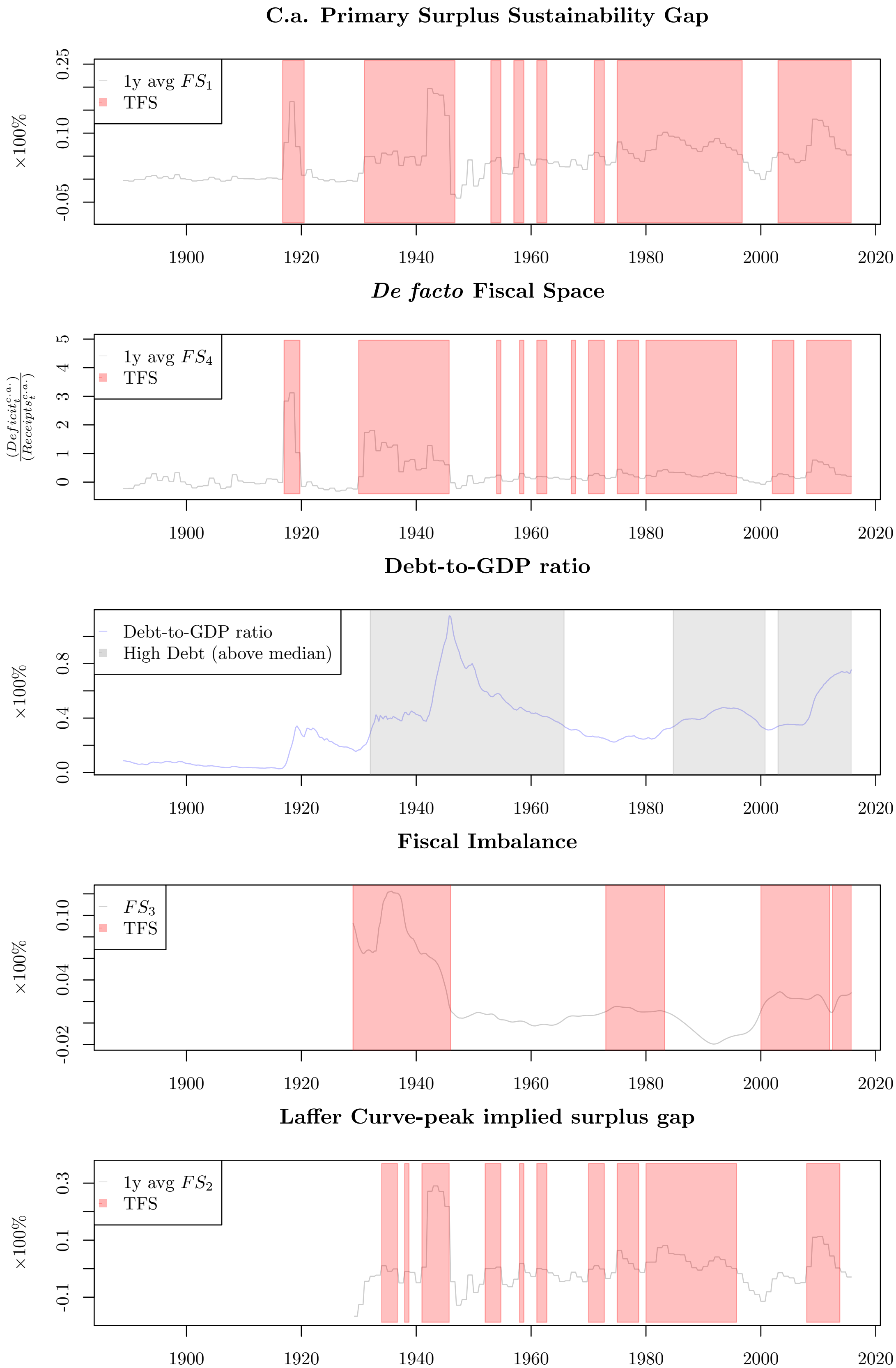


Figure 3.2 shows periods of tight fiscal space (in light red) as indicated by FS_1 (first panel), FS_4 (second panel), FS_3 (fourth panel) and FS_2 (fifth panel). The panel in the middle plots periods of high debt-to-GDP ratio (in grey), identified whenever debt-to-GDP ratio is above its median.

between 1917 and the end of 1920 (see Figures 3.1 and 3.A3). This is due to the large wartime increase in the debt-to-GDP ratio, which rose from a 3% level up to 30% on average. Moreover, this tight fiscal space state partially overlaps with the depression of 1920-21 characterized by extreme deflation and whose key factor was the erroneous tightening stance by the FED.[33] Not surprisingly, all indicators identify the Great Depression as periods of tight fiscal space. Indeed, these periods are characterized by sluggish growth, high real interest rates, increasing debt-to-GDP ratios and deficits, all ingredients generating a reduction in fiscal space. FS_1, FS_2 and FS_4 dummies correctly signal the military build-ups due to the Korean war in 1953-54 (see Figures 3.1, 3.A1 and 3.A3). Additionally, virtually all fiscal space dummies (Figures 3.1, 3.A1, 3.A2 and 3.A3) identify a long-lasting tight fiscal space state starting with the 1973 Oil crisis and continuing through the 1979 energy crisis, the fiscal expansionary policies during the Reagan administration and the Gulf war. Lastly, as already mentioned, the FS dummies characterize the Great Financial Crisis and its aftermath as tight fiscal space periods, given the hefty rise in public spending, the low-growth and low-inflation environment.[34]

Notably, in Figure 3.2, we observe that indicators FS_1 and FS_4 report periods of tight fiscal space between the end of 1916 and the beginning of 1920, while debt-to-GDP was low (and below its historical median). Indeed, in this time-span, deficit over GDP rose from being close to inexistent to oscillating around 15% until 1919. Our indicators righteously pick up the unsustainable fiscal path of government finances of those years compared to the erroneous signal of the debt stock. Similarly, from the end of the '70s until the beginning of the '80s, our indicators signal a period of tight fiscal space while debt-to-GDP was at low levels (around 30%). During this period, the distance between the cost of financing debt (namely, the yield on government bonds) and the growth rate of output (see discussion in Sec. 3.3.1) oscillated around 5% (capping in 1982 at 10%). This points out both that the economy was slowing down, due to the energy crisis, and that the United States experienced its most volatile money growth rates in the post-war era, which translated into higher yields. Additionally, in the late '90s and early 2000*s*, while debt-to-GDP was higher than 60% (and than its historical median), our indicators denote periods of loose fiscal space. This is due to the evolution of deficit over GDP, which

[33] For details, see Friedman and Schwartz (2008)

[34] Public spending rose following the fiscal stimulus packages enacted from 2008 onwards. First, the Economic Stimulus Act of 2008 (enacted February 13, 2008) was an Act of Congress providing for several kinds of economic stimuli intended to boost the United States economy in 2008 and to avert a recession, or ameliorate economic conditions. Second, the American Recovery and Reinvestment Act of 2009 was a stimulus package signed into law by President Barack Obama in February 2009. The approximate cost of this stimulus package was estimated to be $ 831 billion.

decreased from an average 5% in the early ’90s to being mildly negative (surplus) at the end of the decade before rising again from 2002 onwards to early ’90s levels.

3.4 Methodology

In this section, we present the methodology employed to estimate the fiscal multipliers depending on the state of fiscal space. We then describe the identification method adopted and we outline the specification of the empirical model.

3.4.1 State-dependent Local Projection

Local Projections - introduced in Jordà (2005) - are becoming an increasingly popular estimation strategy for Impulse Response Functions (IRF) as opposed to more standard methods like structural VARs. A wide range of estimation procedures can in principle be applied to estimate LPs, and our approach hinges on a standard IV strategy to identify the relevant IRFs. Nevertheless, the discussion to follow is general enough to be applied to other estimation procedures. In a general form, the kind of linear Local Projections we are interested in estimating can be written as

$$y_{t+h} = \alpha_h + \beta_h g_t + \psi_h(L)\mathbf{X}_{t-1} + \varepsilon_{t+h} \quad h = 0, 1, 2, \ldots, H, \tag{3.10}$$

where y_t is the variable whose dynamic response we want to track, g_t is the endogenous variable we want to shock (government spending in our application), and $\mathbf{X}_t$ is a vector of control variables. Estimation is performed *separately* for each horizon and for each dependent variable with two-stage least squares. Generally speaking, IRFs are defined by the sequence $\{\beta_h\}_{h=0}^{H}$, and inference is performed with Newey-West standard errors.

The focus of this paper is on state-dependent responses of macroeconomic variables to fiscal policy. The non-linearity we add is a very simple one, i.e. we investigate the extent to which fiscal policy is transmitted differently under two different regimes, and we separate those two states with a simple indicator variable. Specifically,

$$\begin{aligned} y_{t+h} =& \mathcal{S}_{t-1} \left[\alpha_{TFS,h} + \beta_{TFS,h} g_t + \psi_{TFS,h}(L)\mathbf{X}_{t-1}\right] + \\ &+ (1 - \mathcal{S}_{t-1}) \left[\alpha_{LFS,h} + \beta_{LFS,s} g_t + \psi_{LFS,h}(L)\mathbf{X}_{t-1}\right] + \varepsilon_{t+h} \\ & h = 0, 1, 2, \ldots, H. \end{aligned} \tag{3.11}$$

The state dependency is given by the lagged dummy variable $\mathcal{S}_{t-1}$ that indicates the fiscal space status. Taking as baseline proxy FS_1 (Eq. 3.3 in Section 3.3.1), we define the fiscal space state as tight (large) whenever the 1 year moving average proxy is above (below) its historical median.[35,36] TFS and LFS as subscripts of the parameters in eq. 3.11 indicate *tight* and *large fiscal space*, respectively. This kind of non-linearity is conceptually the same as the one used in (e.g.) Ramey and Zubairy (2018).[37] Other authors - e.g. Tenreyro and Thwaites (2016) - have opted for smooth transition local projections, which allow parameters to smoothly switch between the two regimes, instead of letting them change abruptly around a threshold. While a smooth transition is desirable, for this model - first developed in Granger et al. (1993) - to be employed one needs to calibrate two key curvature and location parameters, whose choice turns out to be quite important in terms of the final set of IRFs that are obtained. In principle, those parameters could be estimated, but in order to do so reliably the researcher would need a lot of data around the transition of the state variable, something that is virtually never the case in macroeconomic applications.[38] We therefore decided to stick with the easier to interpret (and more robust) discrete indicator variable, which nonetheless yields a cleaner interpretation of the coefficients as exact average causal effects within a given state.

3.4.2 Model specification and identification

In our approach, we estimate the LP model as in equation 3.11. We implement an IV approach using two different shock series to instrument government spending g_t in equation 3.11. The first shock series that we use as IV is Ramey news. Ramey (2011b,a) builds a series of estimated changes in expected present value of government purchases caused by military events: the so-called Ramey news shock series. The second series that we consider as instrument for government spending is the Blanchard and Perotti shock. Blanchard and Perotti (2002) pro-

[35]We take the 1 year averaged series so to have a smooth enough series and make sure that the fiscal space state is persistent and lasts at least one year. The average quarters spent in tight fiscal space are approximately equal to 20.

[36]The same definition for fiscal space state goes for FS_2 and FS_4 indicators proposed in Section 3.3; while, given the intrinsic smoothness of the series, for FS_3 (see Section 3.3) there is no need to take the 1-year average. Then, the derived fiscal space state series are used to support the results obtained in Section 3.5.

[37]The underlying assumption in this framework is that, once we calculate the impulse response function in each state, the IRFs remain in that same state through the whole horizon of the estimation. Such assumption seems plausible in our analysis given the persistent behaviour of our fiscal space indicators. Indeed, the average duration of periods characterized by tight fiscal space is, respectively for FS_1, FS_2, FS_3, FS_4, of 31.5, 17.6, 43.5, 23.3 quarters, suggesting the slow moving behaviour of fiscal room. In Section 3.5, we also calculate the impulse response of the state itself, showing how the dynamics of the states over time are not particularly affected by the shock in the short term.

[38]Teräsvirta (1994) discusses those estimation issues in detail.

vide identification for both government spending and tax shocks in a structural VAR, where government spending is ordered first. The identification is based on short-run restrictions and on the automatic stabilizers of fiscal policy to economic activity. A similar approach has been employed also by Fatás et al. (2001) and Galí et al. (2007) among others.

We follow the Ramey and Zubairy (2018) approach to compute fiscal multipliers.[39] The authors first propose to scale output, government spending and the shock series by trend GDP.[40,41] Then, they estimate integral multipliers in one-step.[42] Specifically, to estimate the cumulative output, consumption and investment multipliers, we adopt a one-step IV estimation of

$$\begin{aligned}\sum_{j=0}^{h} y_{t+j} =& \mathcal{S}_{t-1}[\alpha_{TFS,h} + m_{TFS,h}\sum_{j=0}^{h} g_{t+j} + \psi_{TFS,h}(L)\mathbf{X}_{t-1}] + \\ &+ (1-\mathcal{S}_{t-1})[\alpha_{LFS,h} + m_{LFS,h}\sum_{j=0}^{h} g_{t+j} + \psi_{LFS,h}(L)\mathbf{X}_{t-1}] + u_{t+h},\end{aligned} \tag{3.12}$$

instrumenting government spending with the Ramey news and Blanchard and Perotti shock series. Under this approach, we can estimate the integral state-dependent multiplier $m_{i,h}$ ($i = TFS, LFS$) in one step. This allows us to calculate directly the standard errors of the multipliers and, therefore, to implement statistical inference. We select a lag-order of 4 as in Ramey and Zubairy (2018).[43] An advantage of adopting LP methods compared to VARs is the possibility to include a wider set of controls at a lower cost in terms of degrees of freedom. Hence, we include an extensive set of controls (in $\mathbf{X}_{t-1}$ in equations 3.11 and 3.12) that include the average marginal tax rate as in Barro and Redlick (2011) and Bernardini and Peersman (2018), the nominal interest rate on 10 years maturity government bonds, the logarithm of the implicit GDP deflator, real consumption and real investment scaled by trend GDP, the ratio of federal debt to lagged GDP, the ratio of current government deficits on GDP and, lastly, the corporate bond spread (AAA Moody's - Y10). In the bag of controls, lags of output, government spending and the shock series (scaled by trend GDP) are included. Details on the data used can be found in Appendix 3.D.

[39] As observed in Ramey and Zubairy (2018), in post-WWII, the average output-to-spending ratio is equal to 5; while, from 1890 the average output-to-spending ratio is equal to 8. This might lead to biased estimates if multipliers are calculated using log transformed variables (as in most VAR analyses).

[40] And, by analogy, any real variables whose multiplier is of interest for the researcher.

[41] The real GDP time trend is estimated as a sixth-degree polynomial for the logarithm of GDP.

[42] We also prefer integral multipliers (e.g., Mountford and Uhlig, 2009) rather than peak multipliers as in Blanchard and Perotti (2002) or average multipliers given the initial shock as in Auerbach and Gorodnichenko (2012).

[43] Most of the literature studying empirically positive government spending shocks (e.g., Fatás et al., 2001; Blanchard and Perotti, 2002; Galí et al., 2007, among others) use between 2 and 4 lags.

3.5 Results

This section presents the results, reporting the state-dependent effects of government spending shocks. First, we provide the main result of the paper obtained using FS_1 as the baseline state variable for measuring fiscal space. However, later we provide additional evidence using the remaining fiscal space proxies to prove the stability of our results, together with a further robustness section regarding the sample size.

3.5.1 IRF and fiscal multipliers

Figure 3.B1 in Appendix 3.B reports the impulse response function for government spending and economic activity when fiscal shock is estimated using the Ramey news, both in the linear case (top panel) and in the two states of large and tight fiscal space (bottom panel). First of all, we note that the evolution of the two variables of interest in the linear case is fairly standard and in line with that reported in Ramey and Zubairy (2018). Consistent with the dynamics of a news shock, actual government consumption slowly increases and peaks around 10 quarters after the initial impulse. Economic activity follows a comparable pattern. Turning to the non-linear case, we observe a similar dynamic in the two states, even if the same reaction in GDP in the large fiscal space case is determined by a less pronounced increase in spending. However, in order to evaluate quantitatively the effects of fiscal policy, both in general and in particular in non-linear cases, we need to take into account the evolution of the instrumented variable (government consumption in this case). In line with the literature (Ramey and Zubairy, 2018; Mountford and Uhlig, 2009), we calculate the cumulative fiscal multipliers and we henceforth concentrate on this measure to quantify the impact of spending shocks. Tables 3.1 and 3.2 report the fiscal multipliers at each horizon, respectively for the Ramey shock and the Blanchard-Perotti shock. The first column of the tables presents the value of the fiscal multiplier in the linear case, while the second and the third column in large and tight fiscal space. Finally, the last column tests the statistical significance of the difference between the multipliers in the two states, reporting the p-value of the test. Table 3.1 shows that, while the linear case presents multipliers smaller than one, this average effect is very different once disentangled between our two states. Indeed, in tight fiscal space the multiplier averages around 0.6, while in large fiscal space is around 1.5. Such difference is present at each horizon and it is always statistically

significant.[44] A similar picture emerges from Table 3.2, which finds slightly larger fiscal multipliers in the two states. However, the main takeaway remains valid and represents the principal result of the paper, being the fiscal multiplier smaller than one in the tight fiscal space state and larger than one in the opposite case. Such result is particularly important as we draw very different conclusions regarding the state dependent nature of fiscal policy, adopting a methodology which follows closely that of Ramey and Zubairy (2018). We also note that such results are not driven by an unbalanced distribution of shocks in the two states. Indeed, as we show in Appendix 3.G, shocks are equally distributed between periods of tight fiscal space and periods of large fiscal space.[45] Tables 3.C1 and 3.C2 in Appendix 3.C calculate fiscal multipliers when we change the way we define our two states. Section 3.3.5 clarified that, for each of our measures, we define fiscal space as tight (large) when the indicator is above (below) the median. Tables 3.C1 and 3.C2 in Appendix 3.C report the results when instead we concentrate on extreme episodes, meaning that we define fiscal space as tight when the underlying indicator (FS_1 in this case) is above the 80th percentile and as large when it is below the 20th. We do so to investigate whether our results depend upon the threshold adopted to distinguish between the two states. Both tables show results consistent with the main takeaway of the paper. Indeed, the effects are even more pronounced, in line with what one would expect looking at the extreme tails of the distribution. In particular, while the fiscal multiplier becomes larger as there is more fiscal room available, the multiplier shrinks when we consider periods of very tight fiscal space.

[44] However, as the relevance of Ramey instrument is lower for the first horizons of the impulse response, we concentrate our attention on the period two-years after the shock

[45] We cannot instead exclude the possibility of composition effects occurring if shocks to government consumption and investment are distributed unevenly in periods of tight and large fiscal space. Unfortunately, there is not a proper way to check for this possibility. Ramey news have been constructed through a careful work consisting of analyzing weekly newspapers and magazines in the search of military spending news, without distinguishing whether the future increase/decrease in military spending regards consumption or investments. Moreover, the standard practice when adopting this identification method is to instrument general government spending, which comprises both government consumption and investment. While we acknowledge that this strategy could confound two possibly different types of shocks, such problem is common to every paper adopting this identification scheme. Moreover, as government investment shocks are much less frequent than consumption shocks, we believe that the possible confounding effect should not alter the estimates by a too large factor.

Table 3.1: Fiscal Space-dependent Fiscal Multiplier: Baseline - FS_1 - Ramey News Shock (1929-2015)

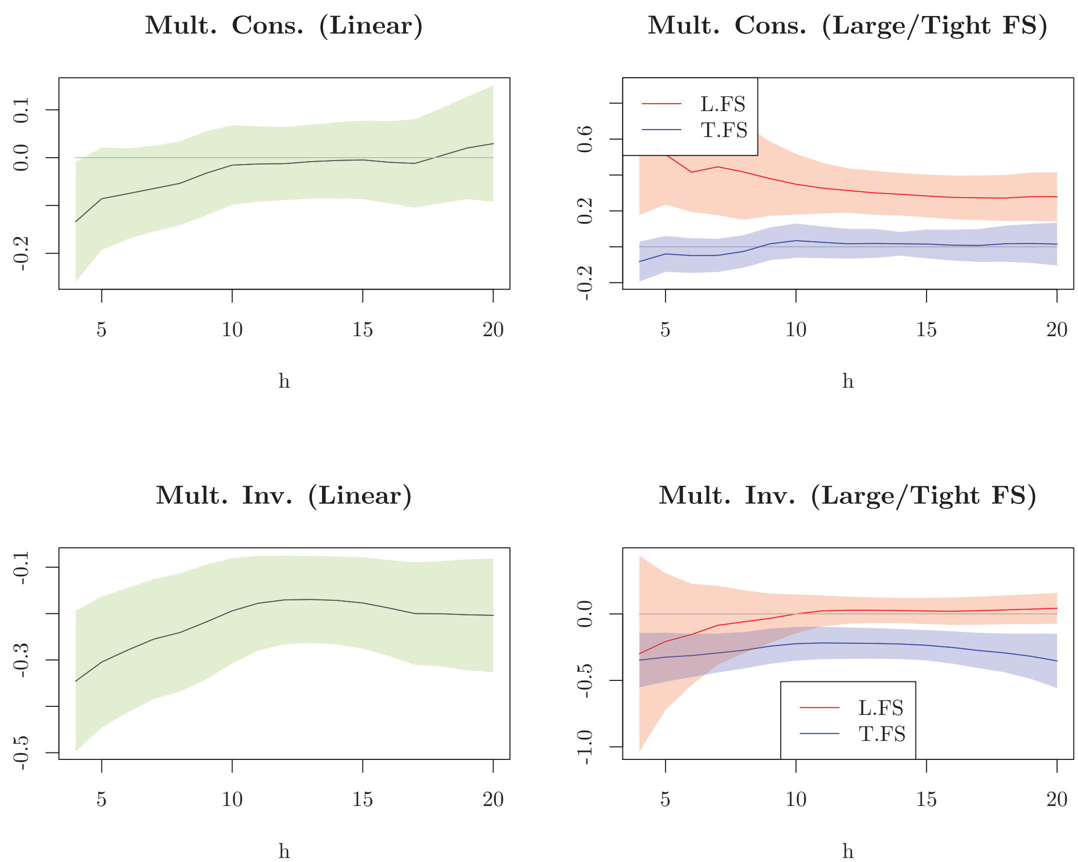
Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	2.63	0.26	0.002***	0.062*
	(0.16)	(0.83)	(0.25)		
2-years	0.74	1.80	0.51	0.001***	0.017**
	(0.15)	(0.36)	(0.18)		
3-years	0.85	1.46	0.62	<0.001***	0.035**
	(0.11)	(0.13)	(0.14)		
4-years	0.88	1.28	0.63	<0.001***	0.044**
	(0.11)	(0.14)	(0.15)		
5-years	0.98	1.30	0.62	0.001***	0.050**
	(0.14)	(0.15)	(0.20)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

The table shows our estimates for the cumulative fiscal multipliers. In the second column, we report the fiscal multiplier derived from the linear case; in the third and fourth columns, we report the multiplier estimates derived from the state-dependent case, respectively under large and tight fiscal space. The fifth column shows the p-values testing the difference, at each horizon, between the multipliers in the two states; the sixth column reports the same piece of information as of column 5 but calculating the p-values using Anderson-Rubin (AR) confidence intervals, to take into account the possibility that the Ramey news series is a weak instrument.

3.5.2 Consumption and Investment

In this subsection, we investigate the transmission on consumption and investment. Figure 3.B2 in Appendix 3.B reports the impulse response of private consumption and investment. While in the linear case they both fall in response to a positive government spending shock, when state-dependencies are considered we observe an opposite behavior in the two states. When fiscal space is tight, private consumption and investment decrease; by contrast, when fiscal space is large they both increase, giving rise to non-Ricardian effects. This result is made clear in Figure 3.3, which reports the consumption and investment multipliers, together with the associated error bands. The consumption multiplier in the large fiscal space state (red line) is around 0.5 one-year after the shock and slowly decays over time; in the tight fiscal space state, it is negative, around -0.1, and slowly reverts towards zero. A similar pattern, although less pronounced, holds for private investment. When fiscal space is large the multiplier is basically zero, given the wide uncertainty surrounding the estimates; when fiscal space is tight, instead, private investment multiplier is negative and significant.

Figure 3.3: Cons. and Inv. multipliers - FS_1 - Ramey News (1929-2015)



The top panels of Figure 3.3 shows the median value of the cumulative multiplier for real consumption, together with its 90% confidence band, both for the linear case (left) and for the non-linear one (right). The bottom panels reports the same information for real total private investment.

Table 3.2: Fiscal Space-dependent Fiscal Multiplier: Baseline - FS_1 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	1.14	0.60	0.209	0.288
	(0.08)	(0.36)	(0.08)		
2-years	0.73	1.18	0.82	0.338	0.372
	(0.11)	(0.32)	(0.12)		
3-years	0.77	1.53	0.87	0.001^{***}	0.094^{*}
	(0.12)	(0.15)	(0.15)		
4-years	0.83	1.58	0.92	$<0.001^{***}$	0.072^{*}
	(0.15)	(0.18)	(0.18)		
5-years	0.81	1.80	0.92	0.031^{**}	0.043^{**}
	(0.19)	(0.45)	(0.18)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

The table shows our estimates for the cumulative fiscal multipliers. In the second column, we report the fiscal multiplier derived from the linear case; in the third and fourth columns, we report the multiplier estimates derived from the state-dependent case, respectively under large and tight fiscal space. The fifth column shows the p-values testing the difference, at each horizon, between the multipliers in the two states; the sixth column reports the same piece of information as of column 5 but calculating the p-values using Anderson-Rubin (AR) confidence intervals, to take into account the possibility that the Blanchard-Perotti shock is a weak instrument.

The differential effects highlighted in Figure 3.3 are at the root of the difference in the output multipliers reported in the previous section. The mechanism we have in mind to rationalize this empirical evidence relates to Perotti (1999), which shows how the degree of public finance sustainability alters the transmission of fiscal policy.[46] Indeed, a deficit-financed increase in government spending generates an increase in future taxation, needed to repay the cost of the fiscal expansion. However, in an environment with distortionary taxation, such tightening will be more pronounced when fiscal space is already tight, because of the convexity in tax distortions. This steeper path of future taxes, internalized by agents, produces a larger negative wealth effect and therefore a subdued reaction of private consumption. Although we believe the aforementioned channel is at the heart of the differential evolution of private consumption in the two states, it is not straightforward to obtain supporting empirical evidence, given the forward-looking nature of the variables involved, in particular expectations regarding future taxation.[47]

[46] In his seminal contribution Perotti (1999) refers to public debt as the variable defining the state of public finance, as opposed to fiscal space. Moreover, Perotti (1999) studies an economy populated both by unconstrained and constrained individuals. However, the main intuition of the paper applies irrespective of such choices.

[47] Another potential way to interpret our results is that, in an economy with finite-lived agents, in the tight fiscal space state, we obtain standard Ricardian effects on consumption. At the opposite, in the large fiscal space state,

3.5.3 Effects on other variables

For the sake of completeness, we also analyze the response of all control variables employed in regression 3.11. Figures 3.B4 and 3.B5 in Appendix 3.B report the results. The average marginal tax rate (AMTR), at least for a large part of the horizon of the IRF, does not display significant different behaviour in the two states. Although such result would suggest that tax do not play a major role in explaining the dichotomy in fiscal multipliers, we stress that the horizon of our IRF is at most five years, too short to observe changes in future tax rates, also given the length of the political cycle[48]. The figure also shows the evolution of debt and deficit ratios, which both increase much more in the tight fiscal space. Such behavior suggests that economic growth and the evolution of budget variables, in the case of large fiscal space, contribute positively to repay for the initial fiscal stimulus and avoid to generating a large increase in the long-run overall level of debt-to-GDP. On the contrary, in the tight fiscal space state, debt and deficit are magnified by the less favourable environment following the shock. This is clear also from Figure 3.B3 in Appendix 3.B, where we report the estimates for the debt multiplier; the latter proves to be significantly higher in the tight state with respect to large fiscal space over the impulse response horizon.[49] Turning to other variables of interest for the transmission of fiscal shock, the most important one is the interest rate. Consistently with the puzzling behavior of the linear case, already shown in the empirical literature by Mountford and Uhlig (2009), Fisher and Peters (2010) and Ramey (2011b), the nominal interest rate slightly decreases in both states.[50,51] As the focus of this study is on times of fiscal distress, it would be interesting to focus on the response of the credit risk component following a government spending shock, which could show a different

the negative wealth effect following the fiscal shock is small as consumers perceive they will not have to repay all of the government spending during their finite lifetime, therefore pushing consumption up.

[48]Results from Perotti (1999) link explicitly the effect of government spending to the political cycle: a lower probability of survival of the policy maker implies a steeper path of future expected taxation, which is distortionary, and, thus, a larger negative wealth effect following a public expenditure shock.

[49]In Appendix 3.F, we show the adjustment needed for the correct computation of the debt multiplier.

[50]Mountford and Uhlig (2009) observe that a government spending shock reduces investment, although interestingly not via higher interest rates (that are moderately falling).

[51]Referring to Fisher and Peters (2010) and Ramey (2011b), Murphy and Walsh (2016) report that one possible explanation for the fall in interest rates is an endogenous response of monetary policy to government spending shocks.

behaviour with respect to the risk-free rate.[52,53] Unfortunately, due to historical data unavailability, we cannot analyze solely how sovereign credit risk comoves with the state of fiscal space following a public spending news shock by exploiting data on sovereign CDS spreads. Thus, focusing on private sector investment, I exploit the response of the corporate bond spread. In Figure 3.B5 in Appendix 3.B, we observe that the corporate bond spread shows a significant and positive response 3 years after the shock in tight fiscal space, while it decreases on average in large fiscal space. This finding might represent a relevant driver for our empirical results concerning the response of the private sector (see Subsection 3.5.2). Finally, the response of the price deflator falls slightly in the tight fiscal space state, while it remains constant in the other state.[54]

3.5.4 Additional results and robustness

This section proposes a series of additional results and robustness checks. First of all, we show that all of our results are robust to different definitions of fiscal space. In order to do so, we perform the same estimations as those provided in the previous section, however adopting the remaining measures, FS_2, FS_3 and FS_4, to identify periods of tight and large fiscal space. In all these cases, fiscal multipliers are calculated using both Ramey shocks and Blanchard-Perotti ones. Tables 3.C3 and 3.C4 in Appendix 3.C summarize, respectively for the two identification methods, the fiscal multiplier when fiscal space is given by measure FS_2. The multipliers in the two states are statistically significant, with those in the large fiscal space state being consistently

[52]In times of uncertainty, as periods of tight fiscal space could prove to be, risk-averse investors might want to increase their demand for safe assets causing a drop in the risk-free yields. At the same time, investors would expect higher returns from bonds bearing increased risk. This way, risk-free yields may fall, while credit spreads increase. Huidrom et al. (2020), in a cross-country framework, find that a positive government spending shock is associated with a rise in sovereign CDS spreads when public debt is high. Moreover, Oliveira et al. (2012) study the determinants of sovereign spreads in the Eurozone. They find a negative relation between sovereign credit spreads and the German yields (regarded as close to risk-free rates), while credit spreads positively comove with government consumption. Using a panel of 74 countries over the 2001-2006 period, Jeanneret (2018) shows that sovereign CDS spreads decrease with government effectiveness.

[53]For instance, using a VAR, Gilchrist and Zakrajšek (2012) show that a rise in corporate credit spreads that is orthogonal to the economic conditions implies a plunge in both the short-term and long-term risk-free rates. Moreover, Blanco et al. (2005) regress CDS prices and credit spread on changes in the long-term interest rate (10-year Treasury bond yield) and find a negative reaction for both. By regressing default swap spreads on firm leverage, volatility and the risk-free rates, Ericsson et al. (2009) find that the coefficient associated with the risk-free rate is consistently negative.

[54]Figure 3.B6 in Appendix 3.B reports the IRFs of the state variables itself, i.e. when $FS_{1,t}$ is considered as dependent variable. The figure shows that the responses are not very significant in large fiscal space, while fiscal space narrows even more in the tight state. This suggests that our states do not change regime over the impulse response horizon, validating the econometric framework adopted to investigate the state-dependency (as described in Section 3.4.1).

above one and those in the alternative state below one. A similar narrative emerges from Tables 3.C5 and 3.C6, which adopt FS_3 as fiscal space indicator. According to these estimates, the fiscal multiplier in the tight fiscal space is basically zero at each horizon given the large error bands associated, with a point-estimate which in some cases turns even negative. The multiplier in large fiscal space is instead statistically significant and greater than one. Tables 3.C7 and 3.C8 repeat the same exercise using FS_4 and once again confirm the main result of the paper. Additionally, using as fiscal space state the first principal component arising from the PCA of the four indicators at hand, we confirm the bottomline findings on fiscal multipliers.[55] All in all, these results substantiate the importance of considering fiscal space for the transmission of fiscal policy. Moreover, they re-assure that the methods employed to measure the evolution of fiscal space over time are consistent, as different indicators in the end produce very similar results. Finally, in this robustness section, we study how a different sample size affects our results. We first show the results when the estimation is performed only in the post-WWII period and after we reconsider the full sample (1929-2015) once we exclude the global financial crisis of the late 2000s. Tables 3.E2, 3.E3, 3.E4 and 3.E5 in Appendix 3.E report the fiscal multiplier, respectively for FS_1, FS_2, FS_3 and FS_4, when the estimation is performed over the period 1947-2015. The government spending shock is estimated only using Blanchard-Perotti shocks, as it is not possible to use the Ramey shock on such a short sample. As Ramey and Zubairy (2018) show, such shocks do not have enough variation to be relevant when instrumenting government spending in the more recent period. The aforementioned tables provide the same univocal picture. Indeed, when considering only the post-war period the difference among multipliers in the two states is magnified, with multipliers in the tight fiscal space shrinking towards zero and instead becoming larger than 1.5 when public finances are sound. Finally, tables from 3.E6 to 3.E13 in Appendix 3.E calculate fiscal multipliers over the full sample for each fiscal space indicator and each identification method, omitting the Great Financial Crisis. In order to do so, we exclude from the sample the period 2007:Q4 - 2010:Q4. Results clearly show that the financial crisis does not play a role in determining the size of fiscal multipliers, as results remain basically unchanged with respect to the baseline. In Appendix 3.H, we report additional estimates of fiscal multipliers to further validate our findings.[56]

[55] For both government spending shock identification methods, in Tables 3.E14- 3.E17 in Appendix 3.E, we report the multipliers using the principal component as indicator both for the full sample and excluding the global financial crisis.

[56] In Appendix 3.H, we report estimates of the fiscal multiplier under periods of High/Low Federal Debt-to-GDP (Tables 3.H1 and 3.H2) and High/Low Federal Debt-to-GDP velocity (Tables 3.H3 and 3.H4) finding no sig-

3.6 Conclusions

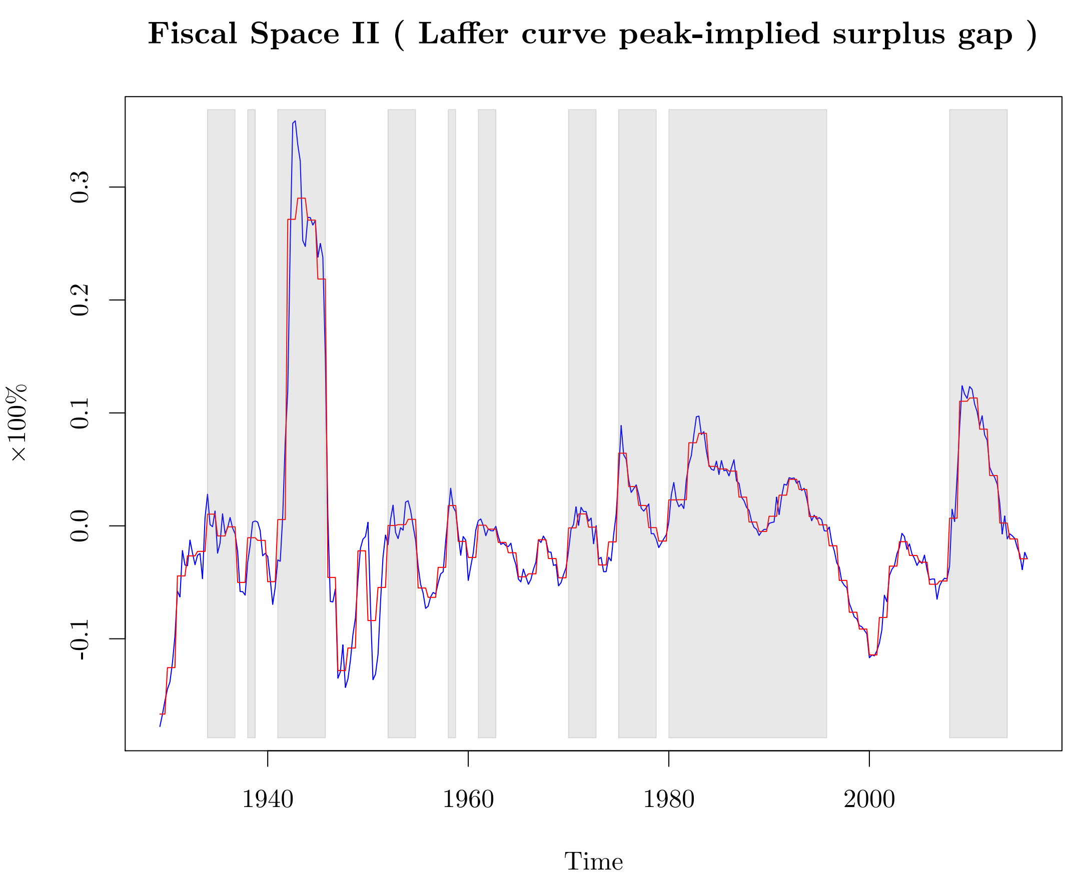
The paper investigated the state-dependent effects of fiscal policy, once the dynamics of fiscal sustainability and fiscal room are jointly considered. Drawing from different strands of the literature we developed several indicators of fiscal space and we measured its evolution over time. The main result highlighted by this paper is that fiscal space matters for the transmission of fiscal policy, as fiscal multipliers are much larger (smaller) when fiscal shocks are implemented in periods of loose (tight) fiscal space. Such a result appears important mainly in two respects. First of all, it stresses the importance of state-dependency in the study of fiscal policy and in particular in relation to fiscal sustainability. Indeed, while the recent literature adopting the identification method of Ramey (2011b) has found only minor differences in fiscal multipliers across business cycle and monetary policy regimes, our paper finds that, by contrast, fiscal space matters a lot. Second, the paper shows, especially from a policy perspective, that fiscal policy can be a very powerful tool in stimulating the economy, but this is not always the case. Particular attention needs to be paid to the economic conditions in which fiscal policy is implemented, as weak public finances could hamper the transmission of fiscal shocks and, in extreme cases, even produce detrimental effects. This latter aspect seems particularly important in light of the COVID-19 crisis, which required an unprecedented support from governments to the economy. Such massive fiscal spending, although necessary, is likely to weigh on the future state of public finance, in particular in those cases and in those countries where the actual fiscal space is already tight.

nificant difference across states. We also show the results on the fiscal multiplier interacting our baseline tight fiscal space state with the zero lower bound without getting any relevant result (see Tables 3.H5 and 3.H6).

Appendix

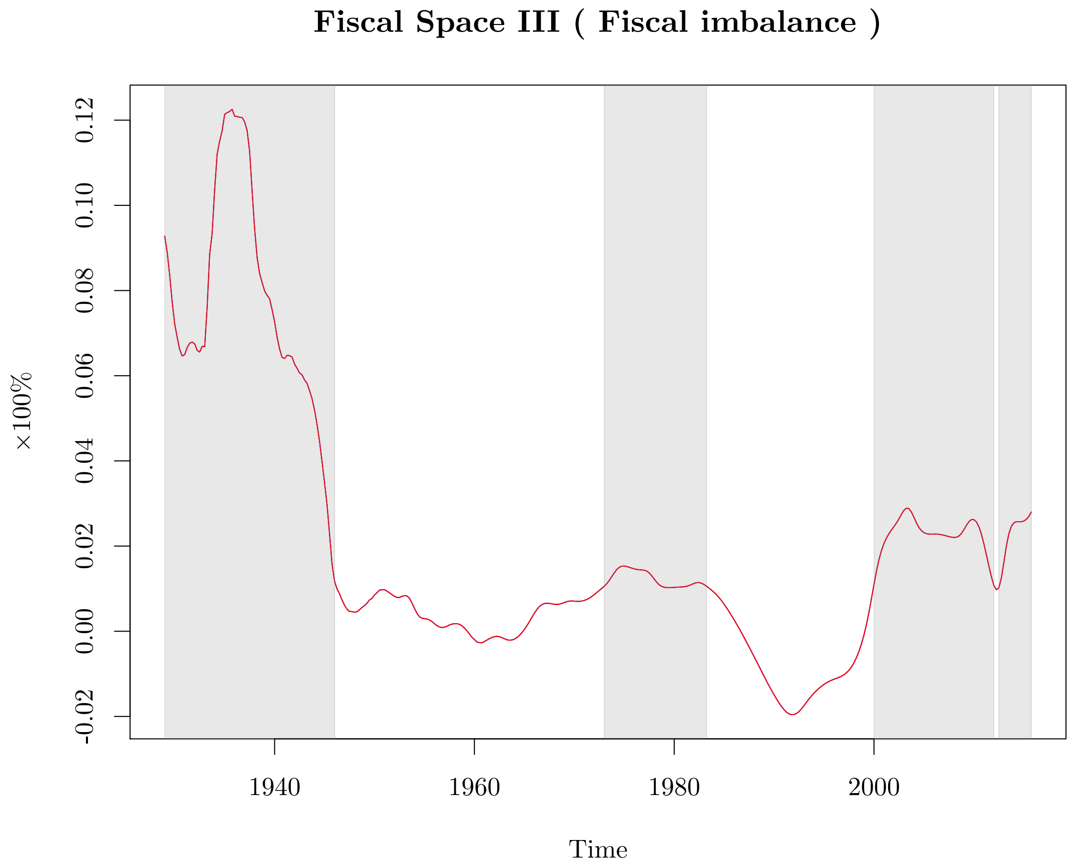
3.A Fiscal space indicators: figures and tables

Figure 3.A1: FS_2 (1929:2-2015:4): Laffer curve peak-implied surplus gap



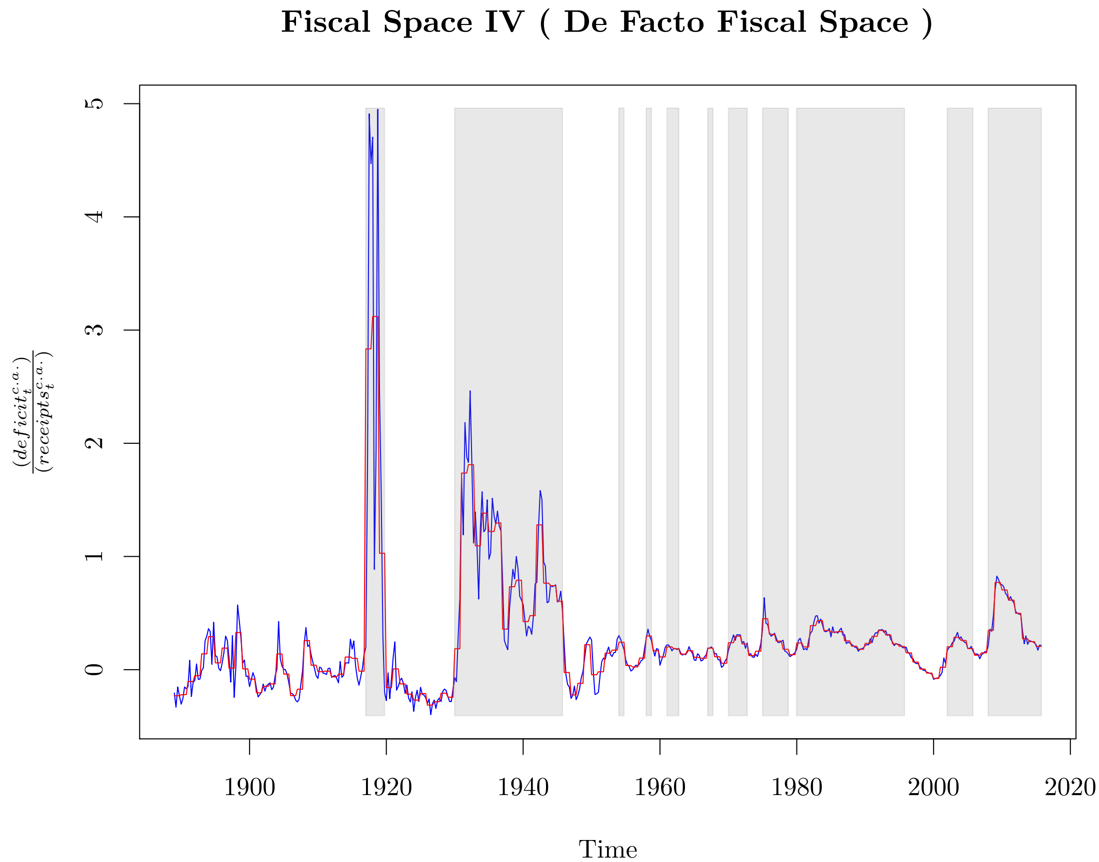
The figure plots indicator FS_2 (in blue) and its 1-year moving average (in red). The series are demeaned and expressed in percent of potential output at quarterly frequency. Shaded regions indicate the periods when fiscal space is tight (1-year $FS >$ median).

Figure 3.A2: FS_3 (1929:1-2015:4): Fiscal Imbalance *á la* Auerbach



The figure plots indicator FS_3, i.e. the fiscal imbalance measure *á la* Auerbach (1997). The series is expressed in percent of potential output at quarterly frequency. Shaded regions indicate the periods when fiscal space is tight ($FS >$ median).

Figure 3.A3: FS_4 (1889:1-2015:4): De Facto fiscal space



The figure plots indicator FS_4 (in blue), i.e. the *de facto* fiscal space *á la* Aizenman and Jinjarak (2010) and its 1-year moving average (in red). The series expresses the relative size of deficit with respect to government revenues. Shaded regions indicate the periods when fiscal space is tight (1-year $FS >$ median).

Table 3.A1: Bootstrapped correlations (95 % confidence interval) among Fiscal space dummies.

	Median	Lower bound	Upper bound
$\mathbb{C}orr(FS1^d_t, FS2^d_t)$	0.597	0.520	0.673
$\mathbb{C}orr(FS1^d_t, FS3^d_t)$	0.344	0.248	0.439
$\mathbb{C}orr(FS1^d_t, FS4^d_t)$	0.732	0.670	0.798
$\mathbb{C}orr(FS2^d_t, FS3^d_t)$	0.014	-0.090	0.121
$\mathbb{C}orr(FS2^d_t, FS4^d_t)$	0.636	0.564	0.707
$\mathbb{C}orr(FS3^d_t, FS4^d_t)$	0.326	0.234	0.424

The table shows the estimates for median non-parametric bootstrapped correlation coefficients and their intervals at the 95% confidence level. Intervals are calculated using the normal approximation. FSj^d_t $(j = 1, 2, 3, 4)$ stand for the fiscal space dummies.

Table 3.A2: Bootstrapped correlations (95 % confidence interval) between Fiscal space dummies and NBER recession dates, High/Low Debt-to-GDP states, ZLB periods, War dates and US political cycle.

	Median	Lower bound	Upper bound
$\mathbb{C}orr(FS1_t^d, Rec_t^d)$	-0.165	-0.247	-0.083
$\mathbb{C}orr(FS2_t^d, Rec_t^d)$	-0.074	-0.112	0.096
$\mathbb{C}orr(FS3_t^d, Rec_t^d)$	0.170	0.071	0.273
$\mathbb{C}orr(FS4_t^d, Rec_t^d)$	-0.070	-0.158	0.024
$\mathbb{C}orr(FS1_t^d, \frac{B}{GDP}_t^d)$	0.487	0.407	0.565
$\mathbb{C}orr(FS2_t^d, \frac{B}{GDP}_t^d)$	0.545	0.425	0.665
$\mathbb{C}orr(FS3_t^d, \frac{B}{GDP}_t^d)$	0.243	0.108	0.371
$\mathbb{C}orr(FS4_t^d, \frac{B}{GDP}_t^d)$	0.701	0.603	0.810
$\mathbb{C}orr(FS1_t^d, ZLB_t^d)$	0.349	0.277	0.420
$\mathbb{C}orr(FS2_t^d, ZLB_t^d)$	0.047	-0.053	0.150
$\mathbb{C}orr(FS3_t^d, ZLB_t^d)$	0.394	0.302	0.484
$\mathbb{C}orr(FS4_t^d, ZLB_t^d)$	0.341	0.268	0.417
$\mathbb{C}orr(FS1_t^d, War_t^d)$	0.089	0.004	0.172
$\mathbb{C}orr(FS2_t^d, War_t^d)$	-0.047	-0.158	0.060
$\mathbb{C}orr(FS3_t^d, War_t^d)$	0.130	0.026	0.227
$\mathbb{C}orr(FS4_t^d, War_t^d)$	0.110	0.026	0.197
$\mathbb{C}orr(FS1_t^d, Dem_t^d)$	0.169	0.083	0.255
$\mathbb{C}orr(FS2_t^d, Dem_t^d)$	-0.027	-0.133	0.076
$\mathbb{C}orr(FS3_t^d, Dem_t^d)$	0.081	-0.025	0.185
$\mathbb{C}orr(FS4_t^d, Dem_t^d)$	0.146	0.061	0.227

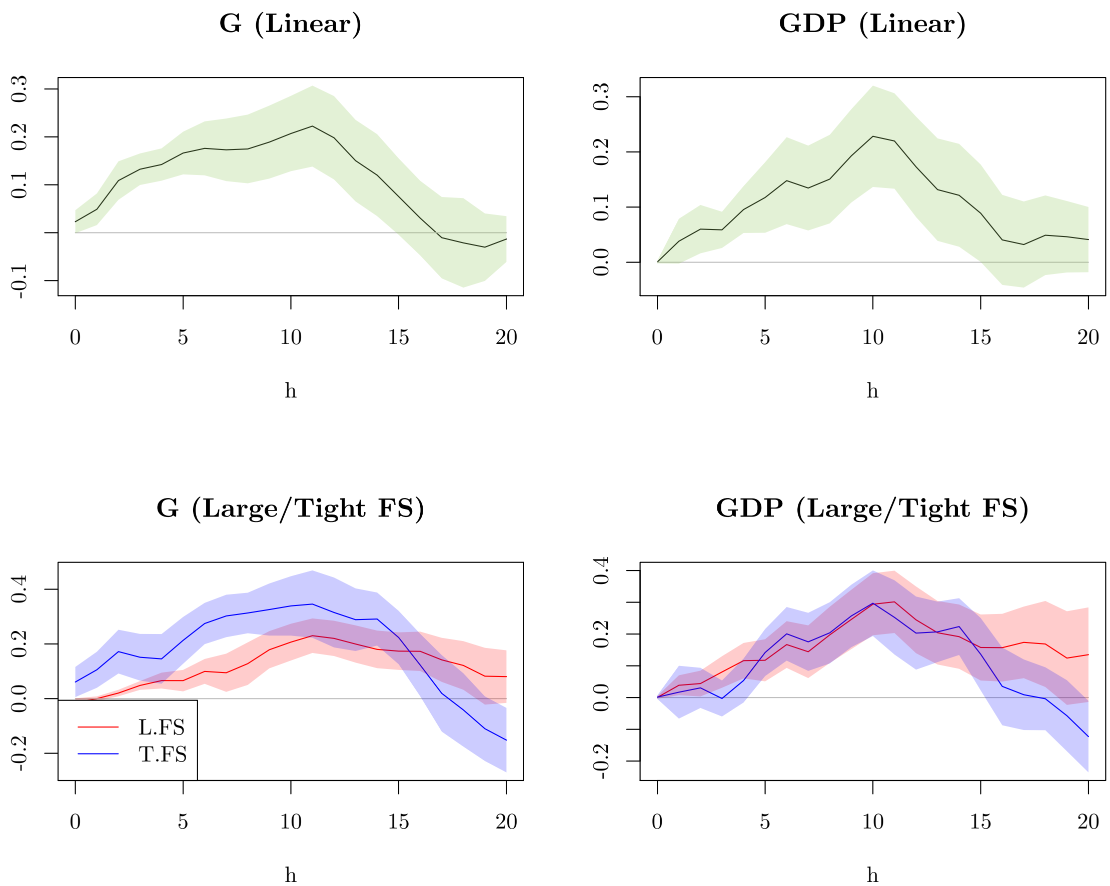
The table shows the estimates for median non-parametric bootstrapped correlation coefficients and their intervals at the 95% confidence level. Intervals are calculated using the normal approximation. FSj_t^d $(j = 1, 2, 3, 4)$ are the fiscal space dummies, Rec_t^d is the dummy for NBER recession dates $\frac{B}{GDP}_t^d$ is the dummy for high-low federal debt-to-GDP states[a], ZLB_t^d is the dummy for the zero lower bound state, War_t^d is the dummy indicating US involvement in major wars.[b] Lastly, Dem_t^d stands for the dummy indicating the party of the US president in the office each quarter.

[a]High (low) federal debt status is considered as above (below) the historical median.

[b]We consider as major wars involving the US the following conflicts: Spanish-American War (1898), Philippine-American War (1899-1902), World War I (1914-1918), World War II (1939-1945), Korean War (1950-1953), Vietnam War (1965-1973), Gulf War (1990-1991), Afghanistan War (started in 2001), Iraq War (2003-2011), American-led intervention in Syria and Iraq (2014-present).

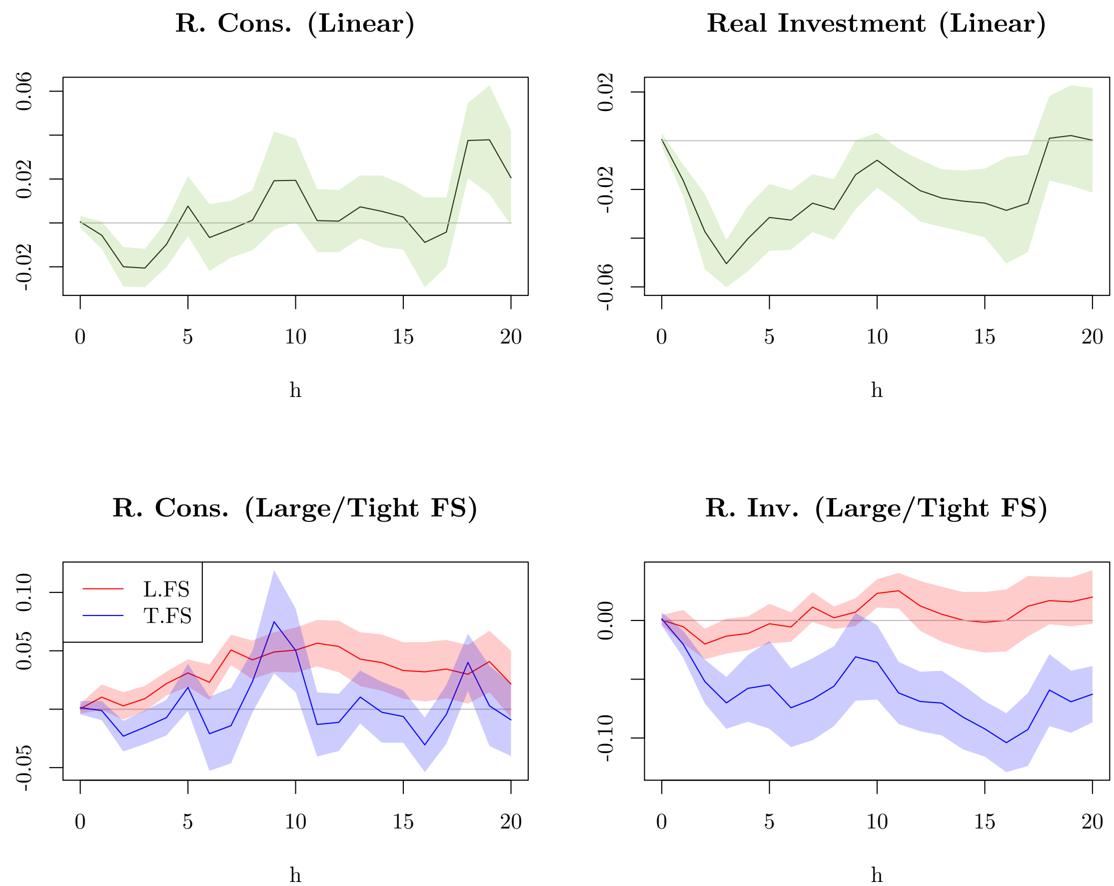
3.B Results: IRFs

Figure 3.B1: G and GDP IRFs - FS_1 - Ramey News (1929-2015)



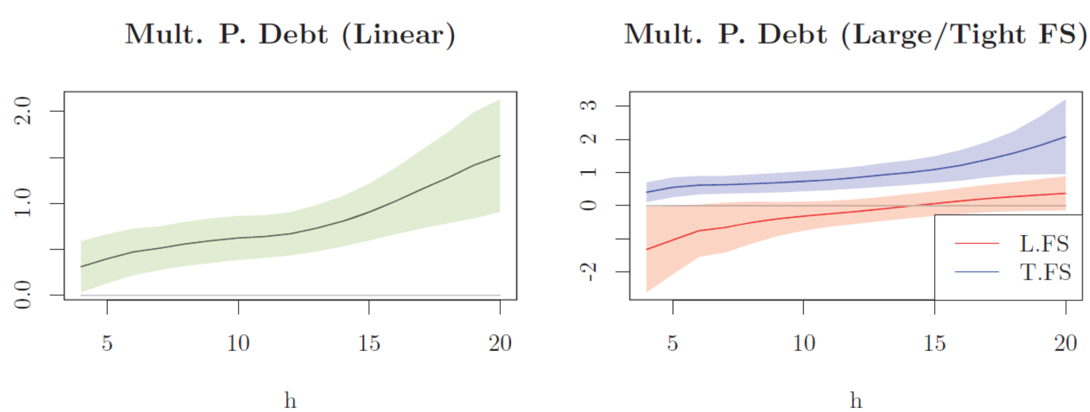
The figure shows the IRFs of real government spending (left panels) and real GDP (right panels) following a news shock equal to 1 percent of GDP identified using the Ramey series, both for the linear case and non-linear cases. Variables are scaled by trend GDP and IRFs are in percentage.

Figure 3.B2: C and I IRFs - FS_1 - Ramey News (1929-2015)



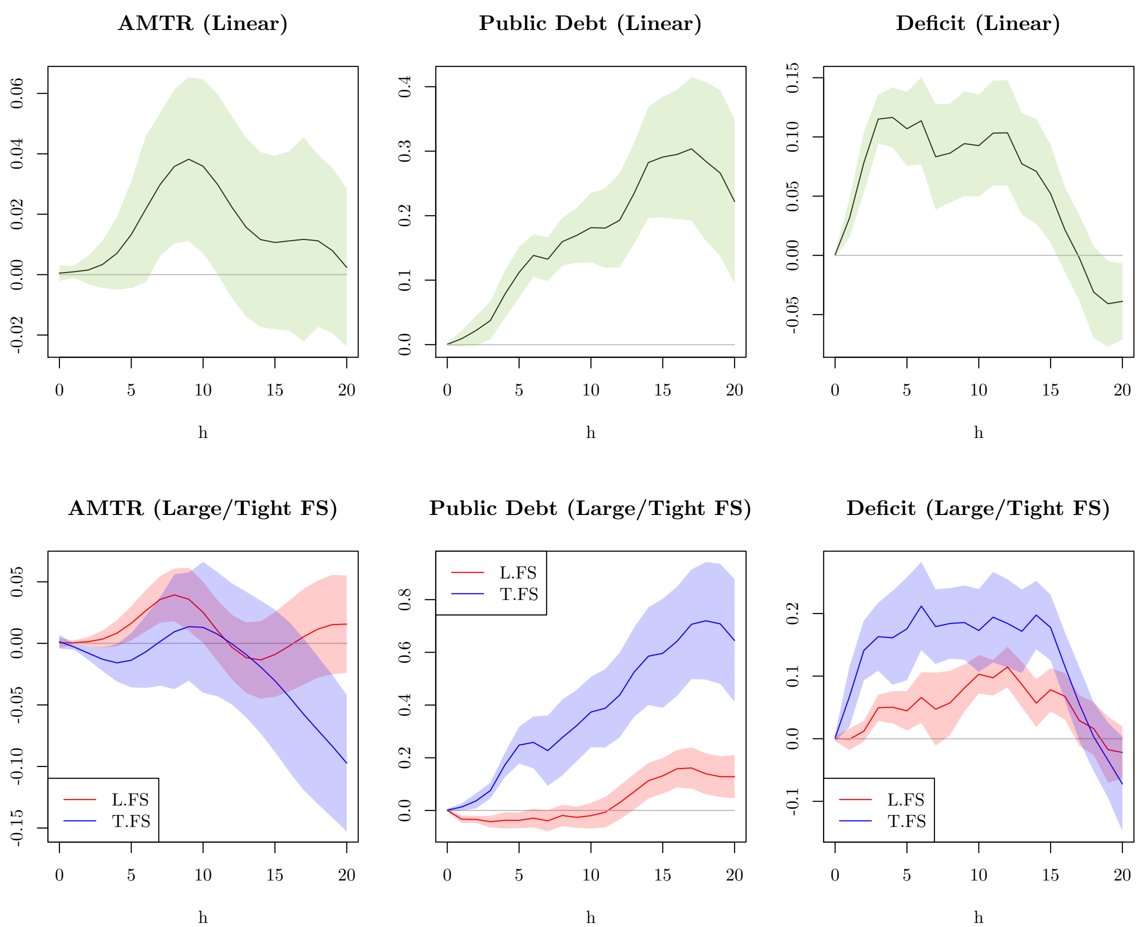
The figure shows the IRFs of real consumption (C) and real total private investment (I), following a news shock equal to 1 percent of GDP identified using the Ramey series, both for the linear case and non-linear cases. Variables are scaled by trend GDP and IRFs are in percentage.

Figure 3.B3: Public debt multiplier - $FS1$ - Ramey News (1929-2015)



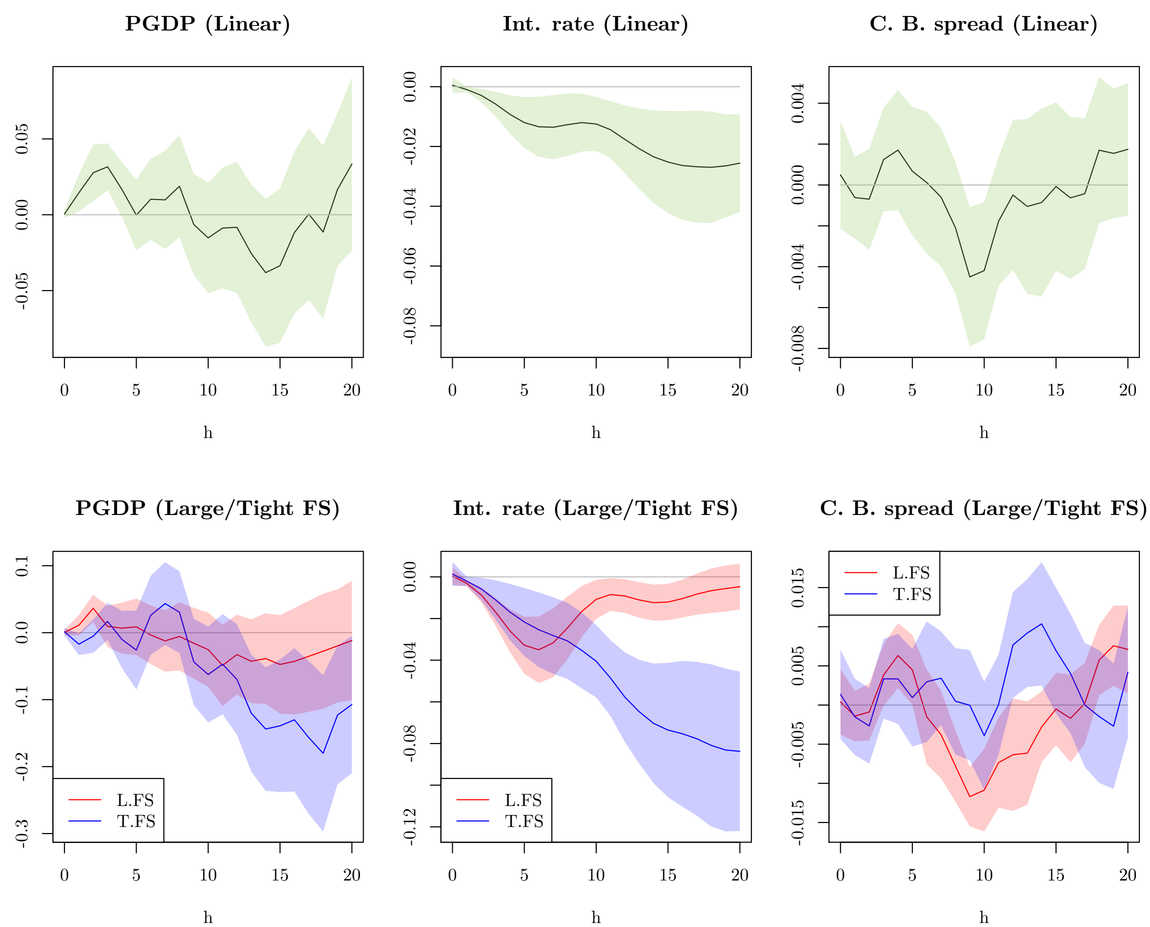
The figure shows the estimates for the integral multiplier of federal debt. The correct multiplier is derived by adjusting the resulting 2SLS, see Appendix 3.F for details.

Figure 3.B4: $AMTR$, P. Debt, deficit IRFs - FS_1 - Ramey News (1929-2015)



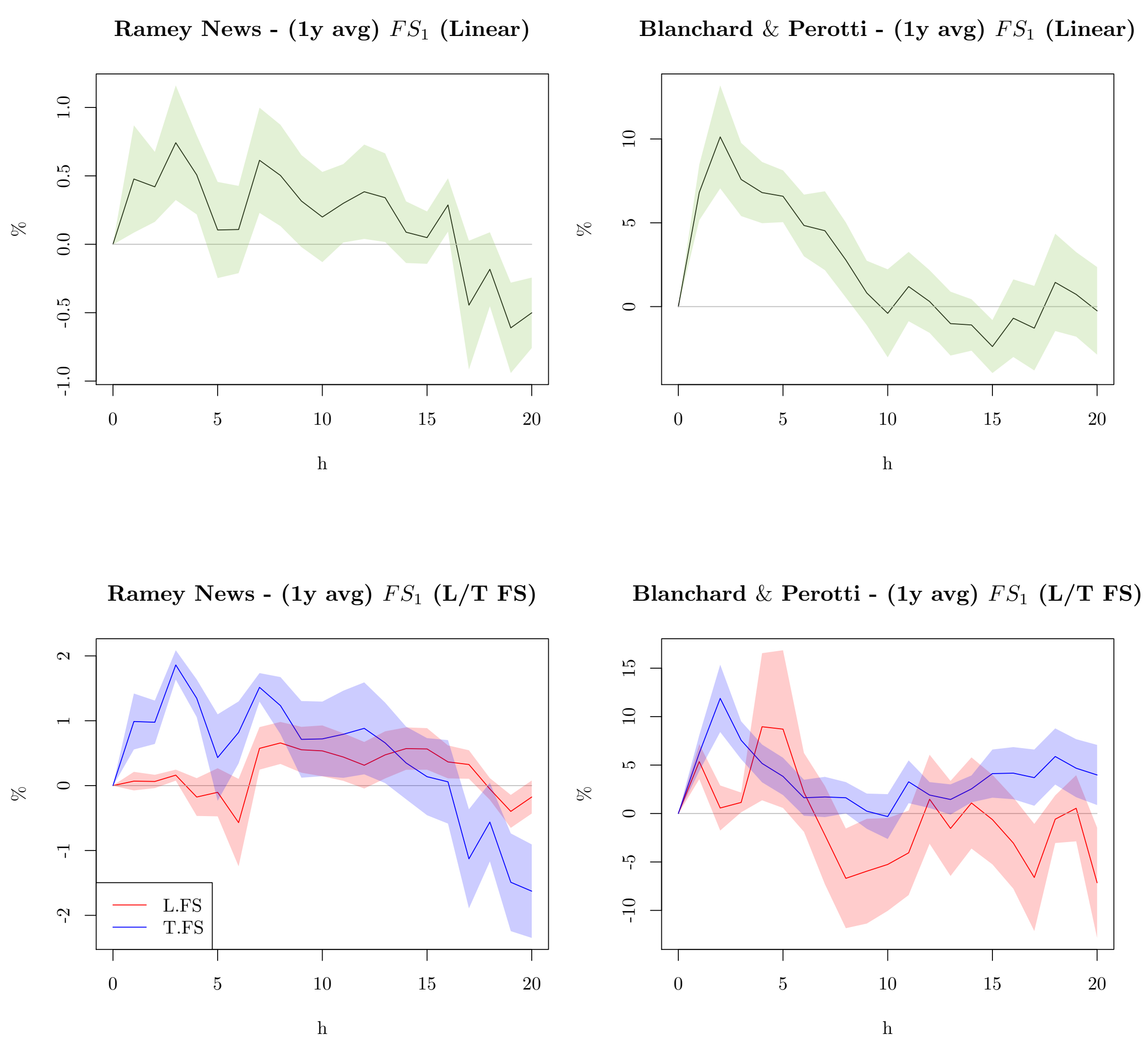
The figure shows the IRFs of the average marginal tax rate ($AMTR$), federal debt-to-lagged GDP and deficit-to-GDP, following a news shock equal to 1 percent of GDP identified using the Ramey series, both for the linear case and non-linear cases. IRFs are in percentage.

Figure 3.B5: $log(IPGDP)$, $Y10$, C.B. spread IRFs - FS_1 - Ramey News (1929-2015)



The figure shows the IRFs of the logarithm of the implicit price deflator ($log(IPGDP)$), 10-years government bond yield ($Y10$), corporate bond spread (C.B. spread) following a news shock equal to 1 percent of GDP identified using the Ramey series, both for the linear case and non-linear cases. IRFs are in percentage.

Figure 3.B6: (1 year average) FS_1 (in real terms) IRFs - Ramey news and Blanchard & Perotti shock (1929-2015).



The figure shows the IRFs of FS_1 indicator (1 year average) following a Ramey news shock (left panels) and Blanchard & Perotti shock (right panels).

3.C Results: Tables

Table 3.C1: Fiscal Space-dependent Fiscal Multiplier: Extremes - FS_1 - Ramey News Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	2.58	-2.63	0.125	0.004^{***}
	(0.15)	(0.47)	(3.55)		
2-years	0.74	1.62	0.16	0.002^{***}	0.015^{**}
	(0.14)	(0.40)	(0.28)		
3-years	0.85	1.82	0.65	$<0.001^{***}$	0.075^{*}
	(0.11)	(0.25)	(0.11)		
4-years	0.88	1.70	0.76	$<0.001^{***}$	0.034^{**}
	(0.12)	(0.19)	(0.08)		
5-years	0.98	1.71	0.74	$<0.001^{***}$	0.074^{*}
	(0.12)	(0.12)	(0.13)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

The table reports the multiplier estimastes across extreme fiscal space states for the whole sample, using Ramey news as instrument. Extreme tight (large) fiscal space is defined as periods where the one year average fiscal space indicator (FS_1) is above (below) its in-sample 80^{th} (20^{th}) percentile. The p-value reported derives from testing the difference between the median multipliers across the two states.

Table 3.C2: Fiscal Space-dependent Fiscal Multiplier: Extremes - FS_1 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	1.76	0.20	$<0.001^{***}$	0.026^{**}
	(0.08)	(0.30)	(0.21)		
2-years	0.73	2.21	0.40	$<0.001^{***}$	0.030^{**}
	(0.11)	(0.23)	(0.21)		
3-years	0.77	2.41	0.55	$<0.001^{***}$	0.029^{**}
	(0.12)	(0.15)	(0.14)		
4-years	0.83	2.27	0.53	$<0.001^{***}$	0.034^{**}
	(0.15)	(0.09)	(0.11)		
5-years	0.81	3.67	0.28	0.068^{*}	0.047^{**}
	(0.19)	(1.55)	(0.21)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

The table reports the multiplier estimastes across extreme fiscal space states for the whole sample, using Blanchard and Perotti as instrument. Extreme tight (large) fiscal space is defined as periods where the one year average fiscal space indicator (FS_1) is above (below) its in-sample 80^{th} (20^{th}) percentile. The p-value reported derives from testing the difference between the median multipliers across the two states.

Table 3.C3: Fiscal Space-dependent Fiscal Multiplier: FS_2 - Ramey News Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	1.88	0.44	0.061*	0.063*
	(0.16)	(0.77)	(0.10)		
2-years	0.74	1.74	0.76	0.036**	0.083*
	(0.15)	(0.45)	(0.07)		
3-years	0.85	1.77	0.85	0.031**	0.068*
	(0.11)	(0.44)	(0.05)		
4-years	0.88	1.59	0.86	0.021**	0.038**
	(0.11)	(0.32)	(0.04)		
5-years	0.98	1.48	0.92	0.011**	0.023**
	(0.14)	(0.21)	(0.04)		
Signif. codes	≤ 0.01***	≤ 0.05**	≤ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.C4: Fiscal Space-dependent Fiscal Multiplier: FS_2 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	0.78	0.60	0.584	0.594
	(0.08)	(0.33)	(0.06)		
2-years	0.73	1.42	0.83	0.045*	0.138
	(0.11)	(0.30)	(0.06)		
3-years	0.77	2.01	0.89	0.051*	0.022**
	(0.12)	(0.48)	(0.06)		
4-years	0.83	2.23	0.91	0.114	0.023**
	(0.15)	(0.69)	(0.06)		
	(0.17)	(0.76)	(0.05)		
5-years	0.81	2.43	0.84	0.149	0.027**
	(0.19)	(1.07)	(0.05)		
Signif. codes	≤ 0.01***	≤ 0.05**	≤ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.C5: Fiscal Space-dependent Fiscal Multiplier: FS_3 - Ramey News Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	0.27	-1.26	0.392	0.294
	(0.16)	(0.60)	(1.60)		
2-years	0.74	1.16	-0.09	0.016**	0.055*
	(0.15)	(0.24)	(0.43)		
3-years	0.85	1.61	0.46	0.001***	0.054*
	(0.11)	(0.20)	(0.28)		
4-years	0.88	1.65	0.58	0.045**	0.061*
	(0.11)	(0.30)	(0.44)		
5-years	0.98	1.48	0.84	0.438	0.391
	(0.14)	(0.41)	(0.70)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.C6: Fiscal Space-dependent Fiscal Multiplier: FS_3 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	0.80	0.48	0.176	0.261
	(0.08)	(0.24)	(0.09)		
2-years	0.73	1.10	0.69	0.214	0.280
	(0.11)	(0.24)	(0.24)		
3-years	0.77	1.46	0.89	0.197	0.285
	(0.12)	(0.24)	(0.38)		
4-years	0.83	1.50	1.25	0.642	0.669
	(0.15)	(0.28)	(0.46)		
5-years	0.81	1.60	1.10	0.163	0.267
	(0.19)	(0.34)	(0.13)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.C7: Fiscal Space-dependent Fiscal Multiplier: FS_4 - Ramey News Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	3.77	0.31	0.022^{**}	0.103
	(0.16)	(1.47)	(0.23)		
2-years	0.74	1.96	0.62	0.087^{*}	0.159
	(0.15)	(0.74)	(0.15)		
3-years	0.85	1.78	0.75	0.100^{*}	0.197
	(0.11)	(0.60)	(0.12)		
4-years	0.88	1.76	0.76	0.089^{*}	0.211
	(0.11)	(0.55)	(0.13)		
5-years	0.98	1.88	0.82	0.046^{**}	0.196
	(0.14)	(0.50)	(0.14)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.C8: Fiscal Space-dependent Fiscal Multiplier: FS_4 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	1.28	0.62	0.062	0.150
	(0.08)	(0.39)	(0.08)		
2-years	0.73	2.09	0.80	0.005^{***}	0.060^{*}
	(0.11)	(0.50)	(0.11)		
3-years	0.77	2.80	0.82	0.003^{***}	0.046^{**}
	(0.12)	(0.66)	(0.13)		
4-years	0.83	2.98	0.88	0.001^{***}	0.042^{**}
	(0.15)	(0.64)	(0.16)		
5-years	0.81	2.87	0.85	$<0.001^{***}$	0.042^{**}
	(0.19)	(0.56)	(0.16)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

3.D Data

3.D.1 Data table

In Table 3.D1, we show the time series used both for the estimation of our fiscal space proxies in Section 3.3 and the empirical analysis described in Section 3.4. We provide also the sources where the data are retrieved with relative samples, the sections where the series are used and the transformation applied to them for both Sections 3.3 and 3.4.

Table 3.D1: Data table

Variable	Sample	Source	Transformation & Usage
Government Spending (Real and Nominal)	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4 : scaled by real trend GDP
Government Revenues (Real and Nominal)	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4 : nominal over nom. GDP
Nominal GDP	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4
Real GDP	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4
Real Consumption	1919:Q1-1946:Q4	Gordon and Krenn (2010)	Sec. 3.4 : growth rate
Real Consumption	1947:Q1-2015:Q4	FRED[a]	Sec. 3.4
Real Investment	1919:Q1-1946:Q4	Gordon and Krenn (2010)	Sec. 3.3; Sec. 3.4 : growth rate
Real Investment	1947:Q1-2015:Q4	FRED	Sec. 3.3; Sec. 3.4
T-bill rate	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4
10-y Gov. bond yield	1889:Q1-1952:Q4	Shiller (1992)	Sec. 3.3; Sec. 3.4
10-y Gov. bond yield	1953:Q1-2015:Q4	Bloomberg	Sec. 3.3; Sec. 3.4
Federal debt	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4 : nominal over lag nom. GDP
Deficit	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4 : nominal over nom. GDP
Implicit GDP deflator	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.3; Sec. 3.4 : in logarithm
Real Potential Output	1889:Q1-2015:Q4	CBO[b]	Sec. 3.3
Nominal Potential Output	1889:Q1-2015:Q4	CBO	Sec. 3.3
Government spending forecasts	2006:Q1-2015:Q4	CBO	Sec. 3.3
Government revenues forecasts	2006:Q1-2015:Q4	CBO	Sec. 3.3
Federal debt forecasts	2006:Q1-2015:Q4	CBO	Sec. 3.3
Real trend GDP[c]	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.4
Average marginal tax rate	1919:Q1-1949:Q4	Barro and Redlick (2011)	Sec. 3.4 : *Federal individual income tax*, stacking
Average marginal tax rate	1950:Q1-2013:Q4	Mertens and Ravn (2013)	Sec. 3.4 : *All tax units (series 1)*
NBER recession dates	1919:Q1-2015:Q4	FRED	Sec. 3.3
Unemployment rate	1889:Q1-2015:Q4	Ramey and Zubairy (2018)	Sec. 3.4 ; Sec. 3.3
Dividends	1929:Q1-1946:Q4	HSUS[d]	Sec. 3.3: cubic spline interpolation
Dividends	1947:Q1-2015:Q4	FRED	Sec. 3.3
Corporate profits before taxes	1929:Q1-1946:Q4	HSUS	Sec. 3.3: cubic spline interpolation
Corporate profits before taxes	1947:Q1-2015:Q4	FRED	Sec. 3.3
Gross wages and salaries	1929:Q1-1946:Q4	HSUS	Sec. 3.3: cubic spline interpolation
Gross wages and salaries	1947:Q1-2015:Q4	FRED	Sec. 3.3
Moody's Seasoned Aaa Corporate Bond Yield	1929:Q1-2015:Q4	FRED	Sec. 3.4
Moody's Seasoned Baa Corporate Bond Yield	1929:Q1-2015:Q4	FRED	Sec. 3.4

[a] Federal Reserve Economic Data (St. Louis FED)

[b] Congressional Budget Office

[c] The real GDP time trend is estimated as a sixth-degree polynomial for the logarithm of GDP, from 1889Q1 through 2015Q4.

[d] Historical Statistics of the United States. Series from HSUS are at annual frequencies and they are interpolated to quarterly frequencies by cubic spline.

3.D.2 FS_2: Maximum tax rates and tax base series

In order to compute approximated government maximum revenues, we use the peak tax rates as derived in Trabandt and Uhlig (2011) to compute an approximation of the government maximum revenues. In their paper, the authors characterize the Laffer curves for capital and labour quantitatively for the US and several EU countries by comparing the balanced growth paths of a neoclassical growth model with constant Frisch elasticity preferences. Moreover, the authors implement a dynamic scoring analysis to explore how tax revenues and production adjust when labour and/or capital income taxes change and which portion of labour and/or capital tax cuts is self-financing. The Laffer curve for consumption taxes does not have a peak and is always increasing (approaching a tax rate of infinity). Hence, we replicate their results using an intertemporal elasticity of substitution (η in Table 3.D2) equal to 2 and a Frisch elasticity (φ in Table 3.D2) equal to 3. For what concerns the consumption tax rate, we take the nearest half-point maximum rate among the tax rates reported in Table 3.D3 as it appears in Trabandt and Uhlig (2011). The maximum tax rates for labour (τ_n), capital (τ_k) are reported in Table 3.D2. Lastly, we take dividends and corporate profits before taxes, wages and salaries before taxes and the portion of disposable income not destined to savings as proxies for the tax base series for capital, labour and consumption respectively. The data sources used for the tax base series are the Historical Database for the United States (HSUS) and FRED (see Table 3.D1 in Appendix 3.D for details).

Table 3.D2: Characterization of US Laffer Curves for capital and labour (Dynamic Scoring at steady state, $\eta = 2$ and $\varphi = 3$).

% self-fin.		max. τ_k		max. add. tax rev.	
same	varied	same	varied	same	varied
60	56	60	65	4	5
% self-fin.		max. τ_n		max. add. tax rev.	
same	varied	same	varied	same	varied
49	47	52	53	14	16

Table 3.D3: Consumption tax rates in % across years (Trabandt and Uhlig, 2011).

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007
USA	5.1	5.1	5.0	5.0	4.9	4.8	4.6	4.5	4.5	4.4	4.5	4.4	4.2

3.E Robustness

Table 3.E1: Bootstrapped correlations (95 % confidence interval) among Fiscal space, potential CBO output, unemployment rate, federal debt-to-GDP, change in debt, growth rate of debt series.

	Median	Lower bound	Upper bound
$\mathbb{C}orr(FS1_t, \tilde{y}_t)$	0.396	0.304	0.482
$\mathbb{C}orr(FS2_t, \tilde{y}_t)$	0.002	-0.112	0.112
$\mathbb{C}orr(FS3_t, \tilde{y}_t)$	-0.347	-0.416	-0.279
$\mathbb{C}orr(FS4_t, \tilde{y}_t)$	-0.002	-0.078	0.072
$\mathbb{C}orr(FS1_t, U_t)$	0.079	-0.0003	0.158
$\mathbb{C}orr(FS2_t, U_t)$	-0.042	-0.141	0.055
$\mathbb{C}orr(FS3_t, U_t)$	0.685	0.605	0.770
$\mathbb{C}orr(FS4_t, U_t)$	0.427	0.269	0.579
$\mathbb{C}orr(FS1_t, \frac{B}{GDP}_t)$	0.413	0.296	0.577
$\mathbb{C}orr(FS2_t, \frac{B}{GDP}_t)$	0.244	0.090	0.395
$\mathbb{C}orr(FS3_t, \frac{B}{GDP}_t)$	-0.084	-0.145	-0.026
$\mathbb{C}orr(FS4_t, \frac{B}{GDP}_t)$	0.131	0.039	0.222
$\mathbb{C}orr(FS1_t, \Delta B_t)$	0.394	0.312	0.470
$\mathbb{C}orr(FS2_t, \Delta B_t)$	0.242	0.140	0.337
$\mathbb{C}orr(FS3_t, \Delta B_t)$	-0.055	-0.102	-0.009
$\mathbb{C}orr(FS4_t, \Delta B_t)$	0.117	0.053	0.176

The table shows the estimates for median non-parametric bootstrapped correlation coefficients and their intervals at the 95% confidence level. Intervals are calculated using the normal approximation. FSj_t ($j = 1, 2, 3, 4$) are the fiscal space proxies, $\tilde{y}_t$ is the potential output (latest estimates from CBO), U_t is the unemployment rate, $\frac{B}{GDP}_t$ is the federal debt-to-GDP ratio and ΔB_t represents its change.

Figure 3.E1: FS dummies and NBER recession dates.

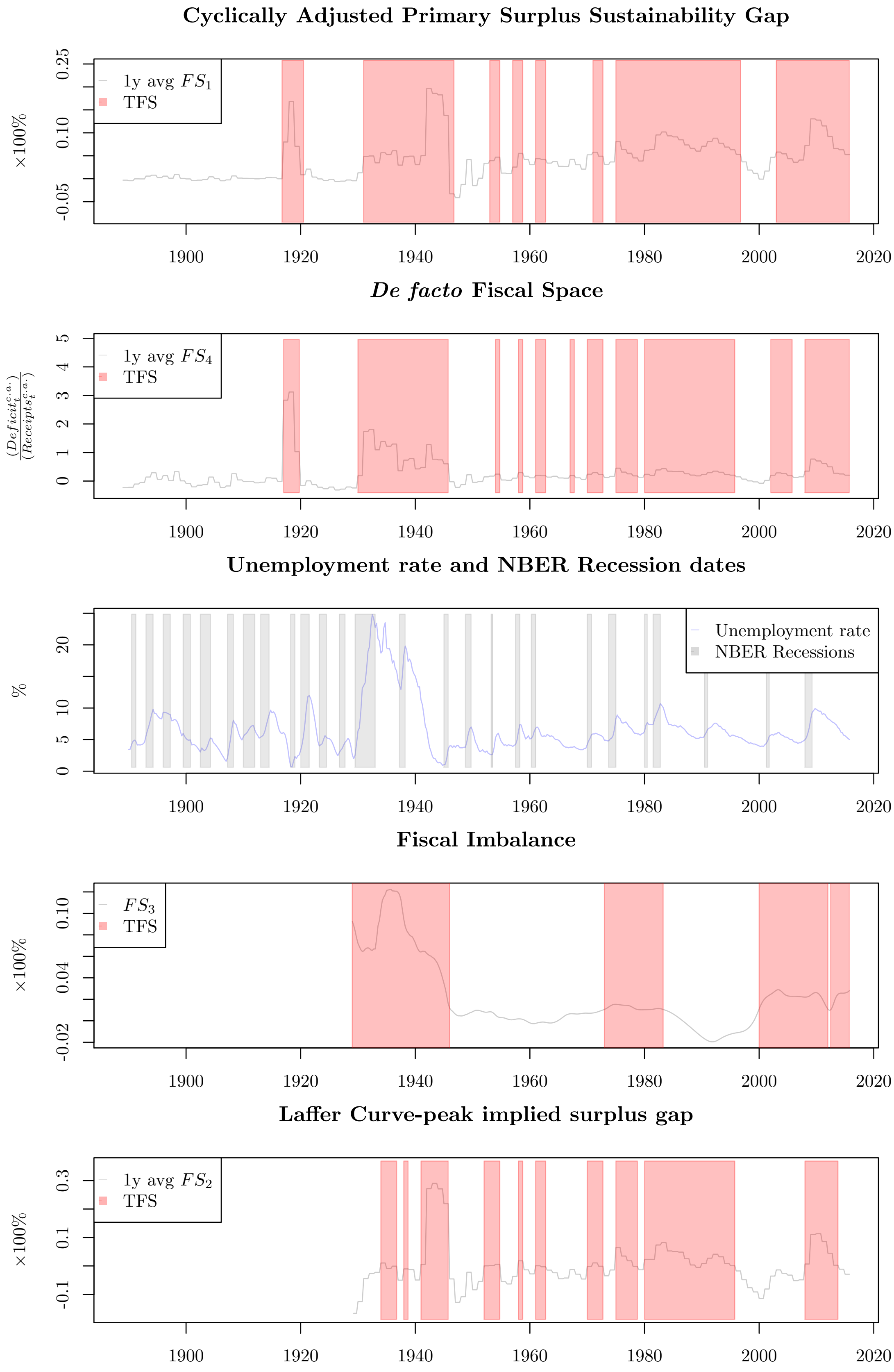


Figure 3.E1 shows periods of tight fiscal space (in light red) as indicated by FS_1 (first panel), FS_4 (second panel), FS_3 (fourth panel) and FS_2 (fifth panel). The panel in the middle plots periods of recessions, as identified by NBER recession dates.

Table 3.E2: Fiscal Space-dependent Fiscal Multiplier: Baseline - FS_1 - Blanchard and Perotti Shock (1947-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.79	1.61	0.03	<0.001***	0.017**
	(0.25)	(0.29)	(0.43)		
2-years	0.85	1.61	-0.22	0.003***	0.014**
	(0.25)	(0.25)	(0.62)		
3-years	0.98	1.70	-0.72	0.007***	0.013**
	(0.25)	(0.20)	(0.88)		
4-years	1.02	1.97	-0.68	0.028**	0.019**
	(0.27)	(0.21)	(1.19)		
5-years	1.08	2.21	-0.30	0.032**	0.012**
	(0.26)	(0.19)	(1.15)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E3: Fiscal Space-dependent Fiscal Multiplier: FS_2 - Blanchard and Perotti Shock (1947-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-years	0.79	1.10	0.80	0.473	0.481
	(0.25)	(0.30)	(0.32)		
2-years	0.85	1.83	0.90	0.021**	0.061*
	(0.25)	(0.33)	(0.28)		
3-years	0.98	1.85	0.93	0.012**	0.035**
	(0.25)	(0.32)	(0.28)		
4-years	1.02	1.76	0.89	0.010***	0.031**
	(0.27)	(0.21)	(0.29)		
5-years	1.08	1.83	0.57	0.003***	0.014**
	(0.26)	(0.15)	(0.39)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E4: Fiscal Space-dependent Fiscal Multiplier: FS_3 - Blanchard and Perotti Shock (1947-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.79	1.29	-1.11	0.025**	0.124
	(0.25)	(0.29)	(1.02)		
2-years	0.85	1.43	-1.93	<0.001***	0.073*
	(0.25)	(0.36)	(0.76)		
3-years	0.98	1.61	-3.63	<0.001***	0.050**
	(0.25)	(0.34)	(0.73)		
4-years	1.02	1.62	-5.96	<0.001***	0.052*
	(0.27)	(0.47)	(1.21)		
5-years	1.08	1.57	22.49	0.186	0.061*
	(0.26)	(0.52)	(15.90)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E5: Fiscal Space-dependent Fiscal Multiplier: FS_4 - Blanchard and Perotti Shock (1947-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.79	1.57	0.03	0.004***	0.029**
	(0.25)	(0.35)	(0.41)		
2-years	0.85	1.80	-0.21	0.003***	0.030**
	(0.25)	(0.35)	(0.56)		
3-years	0.98	2.01	-0.71	0.001***	0.017**
	(0.25)	(0.31)	(0.71)		
4-years	1.02	2.02	-0.81	0.004***	0.028**
	(0.27)	(0.31)	(0.83)		
5-years	1.08	2.05	-0.62	0.016**	0.041**
	(0.26)	(0.33)	(0.91)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E6: Fiscal Space-dependent Fiscal Multiplier: Baseline FS_1 Ramey News shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.64	2.56	0.07	0.001***	0.056*
	(0.14)	(0.72)	(0.27)		
2-years	0.74	1.74	0.36	0.001***	0.036**
	(0.15)	(0.35)	(0.17)		
3-years	0.85	1.41	0.47	<0.001***	0.032**
	(0.09)	(0.13)	(0.14)		
4-years	0.88	1.28	0.41	<0.001***	0.026**
	(0.11)	(0.13)	(0.23)		
5-years	0.98	1.30	0.30	0.011**	0.023**
	(0.12)	(0.15)	(0.43)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E7: Fiscal Space-dependent Fiscal Multiplier: Baseline FS_1 Blanchard and Perotti shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.55	1.14	0.53	0.160	0.233
	(0.08)	(0.40)	(0.08)		
2-years	0.67	1.22	0.70	0.224	0.296
	(0.11)	(0.34)	(0.14)		
3-years	0.70	1.63	0.72	<0.001***	0.116
	(0.14)	(0.18)	(0.14)		
4-years	0.76	1.59	0.79	<0.001***	0.077*
	(0.16)	(0.18)	(0.18)		
5-years	0.72	1.80	0.78	0.006***	0.057*
	(0.25)	(0.45)	(0.22)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E8: Fiscal Space-dependent Fiscal Multiplier: FS_2 Ramey News shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.64	1.94	0.45	0.015^{**}	0.028^{**}
	(0.14)	(0.59)	(0.14)		
2-years	0.74	1.86	0.72	0.005^{***}	0.035^{**}
	(0.15)	(0.40)	(0.07)		
3-years	0.85	1.80	0.80	0.054^{*}	0.048^{**}
	(0.09)	(0.50)	(0.05)		
4-years	0.88	1.66	0.85	0.018^{**}	0.025^{**}
	(0.11)	(0.33)	(0.04)		
5-years	0.98	1.55	0.92	0.022^{**}	0.025^{**}
	(0.12)	(0.27)	(0.05)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E9: Fiscal Space-dependent Fiscal Multiplier: FS_2 Blanchard and Perotti shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-years	0.55	0.77	0.50	0.420	0.448
	(0.08)	(0.33)	(0.05)		
2-years	0.67	1.26	0.71	0.071^{*}	0.196
	(0.11)	(0.31)	(0.05)		
3-years	0.70	1.64	0.79	0.003^{***}	0.088^{*}
	(0.14)	(0.28)	(0.04)		
4-years	0.76	2.00	0.80	0.008^{***}	0.034^{**}
	(0.16)	(0.45)	(0.05)		
5-years	0.72	3.54	0.75	0.532	0.094^{*}
	(0.25)	(4.49)	(0.07)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E10: Fiscal Space-dependent Fiscal Multiplier: FS_3 Ramey News shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.64	0.27	-1.89	0.421	0.226
	(0.14)	(0.60)	(2.53)		
2-years	0.74	1.16	-0.32	0.007***	0.050**
	(0.15)	(0.24)	(0.46)		
3-years	0.85	1.61	0.38	<0.001***	0.057*
	(0.09)	(0.20)	(0.22)		
4-years	0.88	1.65	0.33	0.008***	0.056*
	(0.11)	(0.30)	(0.39)		
5-years	0.98	1.48	-1.15	0.640	0.121
	(0.12)	(0.41)	(5.61)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E11: Fiscal Space-dependent Fiscal Multiplier: FS_3 Blanchard and Perotti shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.55	0.80	0.39	0.092*	0.186
	(0.08)	(0.24)	(0.12)		
2-years	0.67	1.10	0.44	0.035**	0.109
	(0.11)	(0.24)	(0.22)		
3-years	0.70	1.46	0.44	0.009***	0.050**
	(0.14)	(0.24)	(0.29)		
4-years	0.76	1.50	1.07	0.171	0.276
	(0.16)	(0.28)	(0.18)		
5-years	0.72	1.60	1.13	0.272	0.366
	(0.25)	(0.34)	(0.28)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E12: Fiscal Space-dependent Fiscal Multiplier: FS_4 Ramey News shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.64	4.85	0.14	0.070*	0.057*
	(0.14)	(2.58)	(0.23)		
2-years	0.74	2.38	0.47	0.005***	0.036**
	(0.15)	(0.72)	(0.12)		
3-years	0.85	2.07	0.63	0.020**	0.132
	(0.09)	(0.61)	(0.14)		
4-years	0.88	1.96	0.61	0.011**	0.129
	(0.11)	(0.51)	(0.15)		
5-years	0.98	2.07	0.59	0.002***	0.125
	(0.12)	(0.45)	(0.22)		
Signif. codes	≤ 0.01***	≤ 0.05**	≤ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E13: Fiscal Space-dependent Fiscal Multiplier: FS_4 Blanchard and Perotti shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.55	1.15	0.56	0.155	0.257
	(0.08)	(0.41)	(0.09)		
2-years	0.67	1.80	0.70	0.025**	0.122
	(0.11)	(0.49)	(0.13)		
3-years	0.70	2.38	0.69	0.002***	0.062*
	(0.14)	(0.55)	(0.17)		
4-years	0.76	2.87	0.75	<0.001***	0.051*
	(0.16)	(0.59)	(0.20)		
5-years	0.72	2.81	0.72	<0.001***	0.042**
	(0.25)	(0.53)	(0.23)		
Signif. codes	≤ 0.01***	≤ 0.05**	≤ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E14: Fiscal Space-dependent Fiscal Multiplier: Principal Component (FS)- Ramey News Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.65	2.93	0.13	0.201	0.202
	(0.16)	(2.11)	(0.44)		
2-years	0.74	1.83	0.53	0.079*	0.131
	(0.15)	(0.70)	(0.22)		
3-years	0.85	1.84	0.66	0.053*	0.138
	(0.11)	(0.56)	(0.16)		
4-years	0.88	1.87	0.68	0.049**	0.144
	(0.11)	(0.54)	(0.19)		
5-years	0.98	1.95	0.72	0.041**	0.133
	(0.14)	(0.50)	(0.24)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E15: Fiscal Space-dependent Fiscal Multiplier: Principal Component (FS) - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.58	1.00	0.62	0.180	0.221
	(0.08)	(0.29)	(0.10)		
2-years	0.73	1.50	0.85	0.077*	0.114
	(0.11)	(0.26)	(0.16)		
3-years	0.77	1.78	0.88	0.009***	0.072*
	(0.12)	(0.24)	(0.21)		
4-years	0.83	1.77	1.03	0.071*	0.144
	(0.15)	(0.26)	(0.28)		
5-years	0.81	1.74	0.95	0.047**	0.093*
	(0.19)	(0.30)	(0.21)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E16: Fiscal Space-dependent Fiscal Multiplier: Principal Component (FS)- Ramey News Shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.64	2.93	-0.17	0.159	0.170
	(0.14)	(2.11)	(0.50)		
2-years	0.74	1.83	0.24	0.031**	0.096*
	(0.15)	(0.70)	(0.19)		
3-years	0.85	1.84	0.45	0.020**	0.107
	(0.09)	(0.56)	(0.15)		
4-years	0.88	1.87	0.27	0.013**	0.078*
	(0.11)	(0.50)	(0.31)		
5-years	0.98	1.95	-0.91	0.340	0.024**
	(0.12)	(0.45)	(2.59)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.E17: Fiscal Space-dependent Fiscal Multiplier: Principal Component (FS) - Blanchard and Perotti Shock - Omit Crisis

Horizon	Linear	Large FS	Tight FS	p -value Diff.	AR p -value Diff.
1-year	0.55	1.00	0.56	0.144	0.180
	(0.08)	(0.29)	(0.12)		
2-years	0.67	1.50	0.64	0.005***	0.057*
	(0.11)	(0.26)	(0.20)		
3-years	0.70	1.78	0.51	<0.001***	0.030**
	(0.14)	(0.24)	(0.22)		
4-years	0.76	1.77	0.79	0.038**	0.103
	(0.16)	(0.27)	(0.34)		
5-years	0.72	1.74	0.66	0.062*	0.056**
	(0.25)	(0.35)	(0.35)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

3.F Debt multiplier adjustment

Since we compute the multipliers for output, consumption, investment and public debt using the technique described in Section 3.4, the latter multiplier needs an *ex-post* adjustment given that public debt is not scaled by trend GDP (real). First, we assume that

$$\sum_{h=1}^{H} \frac{Debt^N_{t+h}}{GDP^N_{t-1+h}} \approx \sum_{h=1}^{H} \frac{Debt^R_{t+h}}{GDP^R_{t-1+h}}, \tag{Ass. I}$$

where the superscripts N and R indicate nominal and real variables, respectively. Second, we assume that the ratio of real GDP (but also its lag) and trend GDP is approximately constant,

$$\frac{GDP^R_{t-1+j}}{Trend^{GDP^R}_{t+j}} \approx \kappa_h,\ j = 0, \dots, h \tag{Ass. II}$$

Indeed, κ_h oscillates around 1 over the impulse response horizon with very little variation. Thus, following the strategy described in Section 3.4, we regress cumulative debt over lagged GDP (nominal) on cumulative government spending scaled by trend output (real) as follows:

$$\sum_{j=0}^{h} \frac{Debt^N_{t+j}}{GDP^N_{t-1+j}} = \alpha_h + m_h \sum_{j=0}^{h} g_{t+j} + \psi_h(L)\mathbf{X}_{t-1} + u_{t+h}, \tag{a.3.1}$$

where m_h represents the one-step cumulative multiplier for debt for each h. However, we are interested in finding the multiplier $\hat{m}_h$ such that, for each h,

$$\sum_{j=0}^{h} \frac{Debt^R_{t+j}}{Trend^{GDP^R}_{t-1+j}} = \hat{\alpha}_h + \hat{m}_h \sum_{j=0}^{h} g_{t+j} + \hat{\psi}_h(L)\mathbf{X}_{t-1} + \hat{u}_{t+h}. \tag{a.3.2}$$

We know that

$$\hat{m}_h = \frac{\partial \sum_{j=0}^{h} \frac{Debt^R_{t+j}}{Trend^{GDP^R}_{t+j}}}{\partial \sum_{j=0}^{h} g_{t+j}} \tag{a.3.3}$$

and that

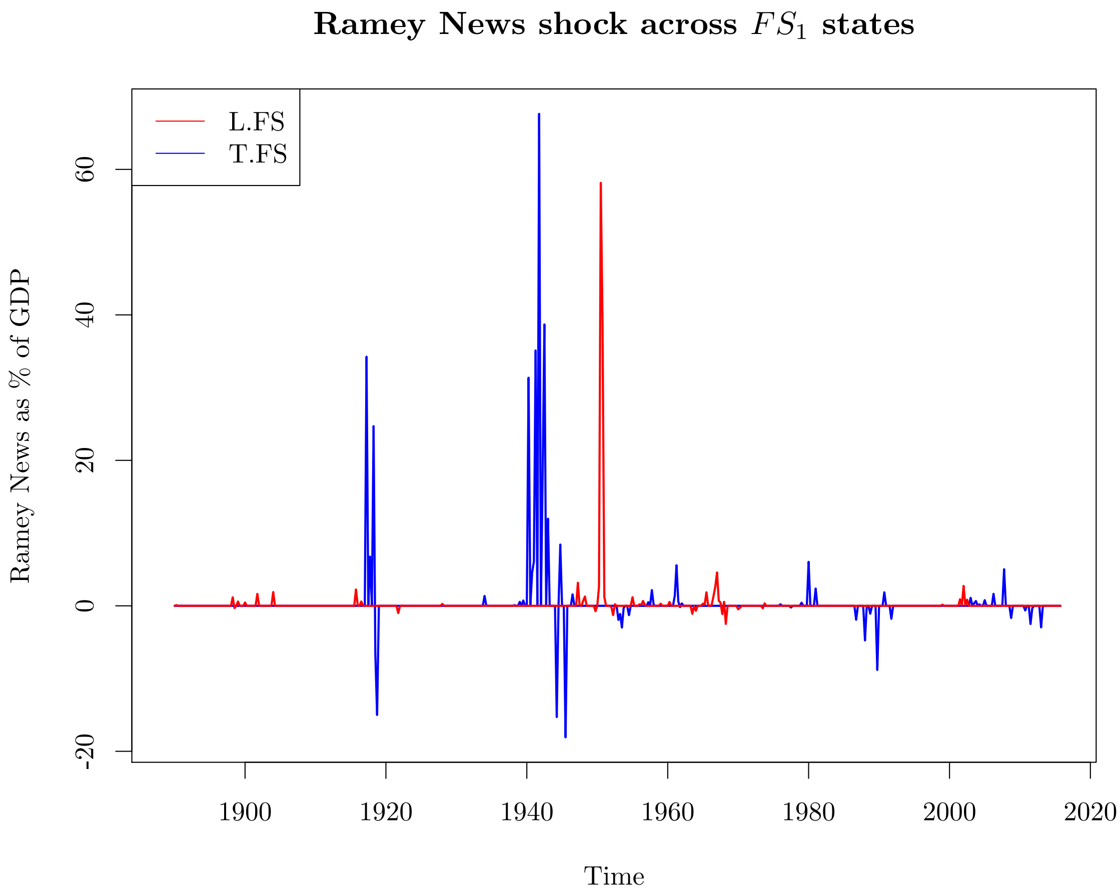
$$m_h = \frac{\partial \sum_{j=0}^{h} \frac{Debt^N_{t+j}}{GDP^N_{t-1+j}}}{\partial \sum_{j=0}^{h} g_{t+j}} \tag{a.3.4}$$

thus, we can rewrite Eq. a.3.3 as follows,

$$\hat{m}_h = \frac{\partial \sum_{j=0}^{h} \frac{Debt^R_{t+j}}{Trend^{GDP^R}_{t+j}}}{\partial \sum_{j=0}^{h} g_{t+j}} = \frac{\partial \sum_{j=0}^{h} \frac{Debt^R_{t+j}}{GDP^R_{t-1+j}} \frac{GDP^R_{t-1+j}}{Trend^{GDP^R}_{t+j}}}{\partial \sum_{j=0}^{h} g_{t+j}} \approx m_h \cdot \kappa_h \tag{a.3.5}$$

given Ass. I and Ass. II. Therefore, we adjust the debt multiplier following Eq. a.3.5. Note that the adjustment holds true also for the state-dependent case. Moreover, we adjust the standard error of the debt multiplier using standard delta methods.

Figure 3.G1: Ramey news shock - distribution across states



This figure plots the Ramey news series (scaled by nominal GDP) while distinguishing periods of large fiscal space (red) and periods of tight fiscal space (blue), as defined by indicator FS_1. The picture shows that the shock is roughly balanced across the two fiscal space regimes.

3.G Distribution of fiscal shocks across states

Figure 3.G1 below reports the ratio between the Ramey instrument, used to identify exogenous fiscal expansions/contractions, and GDP, distinguishing between shocks in tight fiscal space (blue) and large fiscal space (blue). The picture shows that shocks are roughly balanced across regimes, in the sense that there is similar mass of fiscal policy shocks in periods of tight fiscal space and in periods of large fiscal space. Looking at the two biggest shocks recorded by Ramey, one is in a period of fiscal space, while the other coincides with the large fiscal space period. The graph also shows that shocks are balanced between negative and positive signs. Shocks are instead less balanced in terms of magnitude. However, this is a well-known characteristic regarding shocks identified á la Ramey. Even in the linear analysis, WWII and the Korean war represent major fiscal policy shocks while the rest are much smaller in size. While we acknowledge this limitation, such problem is common to every paper employing the Ramey instrument, which is nonetheless one of the most common identification methods in fiscal policy.

3.H Additional results on the fiscal multiplier

3.H.1 Federal Debt-to-GDP

Tables 3.H1 and 3.H2 show the estimates for the fiscal multiplier according to the level of federal debt-to-GDP ratios. We define the state as *High Debt* (*Low Debt*) when the federal debt-to-GDP ratio is above (below) its median.[57] Consistently with Auerbach and Gorodnichenko (2017) and Huidrom et al. (2020), we find a multiplier that is higher in a low debt state. We do not find evidence to support a multiplier close to zero – or even negative – as in Ilzetzki et al. (2013). The multiplier is not significantly different across the two states for most horizons using both instruments. Notably, when the shock series is instrumented with Ramey news, the median multiplier is puzzlingly higher under tight fiscal space state.

Table 3.H1: Fiscal Space-dependent Fiscal Multiplier: Federal Debt-to-GDP - Ramey News Shock (1929-2015)

Horizon	Linear	Low debt	High debt	p -value Diff.	AR p -value Diff.
1-year	0.65	0.27	0.70	0.153	0.163
	(0.16)	(0.25)	(0.19)		
2-years	0.74	0.59	0.93	0.341	0.314
	(0.15)	(0.09)	(0.35)		
3-years	0.85	0.77	1.35	0.079*	0.160
	(0.11)	(0.04)	(0.28)		
4-years	0.88	0.82	1.64	0.280	0.168
	(0.11)	(0.03)	(0.76)		
5-years	0.98	0.85	1.97	0.529	0.091*
	(0.14)	(0.04)	(1.76)		
Signif. codes	≤ 0.01***	≤ 0.05**	≤ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

3.H.2 Δ Debt

Tables 3.H3 and 3.H4 show the estimates for the fiscal multiplier according to the change in federal debt-to-GDP ratio. We define the state as *High Δ Debt* (*Low Δ Debt*) when the change in federal debt-to-GDP ratio is above (below) zero.[58] We find a multiplier that is higher in a decelerating debt state using Ramey news only for the 4 and 5-years horizon. We find no statistically significant difference across states employing the Blanchard and Perotti shock.

[57] The median for federal debt-to-GDP ratio is equal to 40%.

[58] This value represents also the historical median other than the turning point between accelerating and decelerating debt.

Table 3.H2: Fiscal Space-dependent Fiscal Multiplier: Federal Debt-to-GDP - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Low debt	High debt	p -value Diff.	AR p -value Diff.
1-year	0.58	0.75	0.48	0.067*	0.193
	(0.08)	(0.13)	(0.07)		
2-years	0.73	0.89	0.68	0.482	0.505
	(0.11)	(0.16)	(0.22)		
3-years	0.77	0.82	0.64	0.896	0.880
	(0.12)	(0.08)	(1.36)		
4-years	0.83	0.79	0.75	0.931	0.927
	(0.15)	(0.11)	(0.54)		
5-years	0.81	0.78	0.96	0.515	0.605
	(0.19)	(0.22)	(0.23)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

3.H.3 Interaction of ZLB and Fiscal Space state

The two most remarkable periods of tight fiscal space are the Second World War and the Great Financial Crisis jointly with its aftermath. These periods coincide with the most long-lasting periods under the zero lower bound. Thus, we estimate the multiplier depending on the interaction of the zero lower bound and tight fiscal space periods. We do so to see whether the presence of the zero lower bound biased our estimates. Tables 3.H5 and 3.H6 show no difference across states in the multiplier for most horizons further validating our baseline results.

Table 3.H3: Fiscal Space-dependent Fiscal Multiplier: Change in Federal Debt-to-GDP - Ramey News Shock (1929-2015)

Horizon	Linear	Low Δ debt	High Δ debt	p -value Diff.	AR p -value Diff.
1-year	0.65	0.75	0.61	0.626	0.646
	(0.16)	(0.23)	(0.18)		
2-years	0.74	0.93	0.78	0.476	0.503
	(0.15)	(0.15)	(0.15)		
3-years	0.85	1.08	0.81	0.104	0.163
	(0.11)	(0.11)	(0.13)		
4-years	0.88	1.15	0.82	0.063*	0.156
	(0.11)	(0.12)	(0.14)		
5-years	0.98	1.32	0.84	0.040**	0.107
	(0.14)	(0.17)	(0.16)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.H4: Fiscal Space-dependent Fiscal Multiplier: Change in Federal Debt-to-GDP - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	Low Δ debt	High Δ debt	p -value Diff.	AR p -value Diff.
1-year	0.58	0.95	0.61	0.026**	0.070
	(0.08)	(0.12)	(0.06)		
2-years	0.73	1.08	0.78	0.121	0.123
	(0.11)	(0.16)	(0.10)		
3-years	0.77	1.14	0.82	0.113	0.140
	(0.12)	(0.17)	(0.11)		
4-years	0.83	1.17	0.87	0.348	0.404
	(0.15)	(0.30)	(0.13)		
5-years	0.81	2.90	0.85	0.933	0.644
	(0.19)	(26.80)	(0.14)		
Signif. codes	$\leq$ 0.01***	$\leq$ 0.05**	$\leq$ 0.1*		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.H5: Fiscal Space-dependent Fiscal Multiplier: Interaction ZLB and FS_1 - Ramey News Shock (1929-2015)

Horizon	Linear	LFS/no-ZLB	TFS/ZLB	p -value Diff.	AR p -value Diff.
1-year	0.65	2.08	1.02	0.217	0.297
	(0.16)	(0.87)	(0.41)		
2-years	0.74	1.44	0.56	0.162	0.250
	(0.15)	(0.47)	(0.30)		
3-years	0.85	1.28	0.75	0.090^{*}	0.062^{*}
	(0.11)	(0.24)	(0.11)		
4-years	0.88	1.19	0.11	0.783	0.523
	(0.11)	(0.22)	(3.45)		
5-years	0.98	1.21	0.96	0.225	0.025^{**}
	(0.14)	(0.21)	(0.01)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Table 3.H6: Fiscal Space-dependent Fiscal Multiplier: Interaction ZLB and FS_1 - Blanchard and Perotti Shock (1929-2015)

Horizon	Linear	LFS/no-ZLB	TFS/ZLB	p -value Diff.	AR p -value Diff.
1-year	0.58	0.60	0.37	0.547	0.519
	(0.08)	(0.25)	(0.26)		
2-years	0.73	0.78	4.78	0.673	0.279
	(0.11)	(0.20)	(8.40)		
3-years	0.77	0.83	1.65	0.303	0.112
	(0.12)	(0.19)	(0.79)		
4-years	0.83	0.81	0.99	0.538	0.612
	(0.15)	(0.20)	(0.22)		
5-years	0.81	0.78	2.73	0.877	0.189
	(0.19)	(0.26)	(12.56)		
Signif. codes	$\leq 0.01^{***}$	$\leq 0.05^{**}$	$\leq 0.1^{*}$		

For further details, see Tables 3.1 and 3.2 in Section 3.5.

Chapter 4

Fiscal Consolidations in Good Times and in Bad

4.1 Introduction

Following the Great Financial Crisis and the COVID-19 pandemic, European economies implemented large stimulus plans that translated into government budget deficits. The latter piled up into levels of debt-to-GDP that were unprecedented since the Second World War. This fragile fiscal position combined with sluggish growth raised concerns on fiscal sustainability. As shown in Figure 4.1, general government debt rose at the onset of the euro-debt crisis and during the COVID-19 pandemic, while primary balance plunged due to the large rolled-out fiscal stimulus packages. Especially in core European Union (EU) countries, policy circles have nowadays started to advocate for the need to tighten fiscal policy. Notably, the German Council of Economic Experts stated in its *2021/22 Annual Report* that sustainability and resilience of public finances to crises should be strengthened again.[1,2] Also the European Commission recommended in its *Fiscal Policy Guidance for 2023* for the implementation of multi-year fiscal adjustments to curb debt dynamics.[3] Given the fiscal rule framework conceived in the EU *Stability and Growth Pact*, the response to large increases in deficit and debt consists in strengthening the

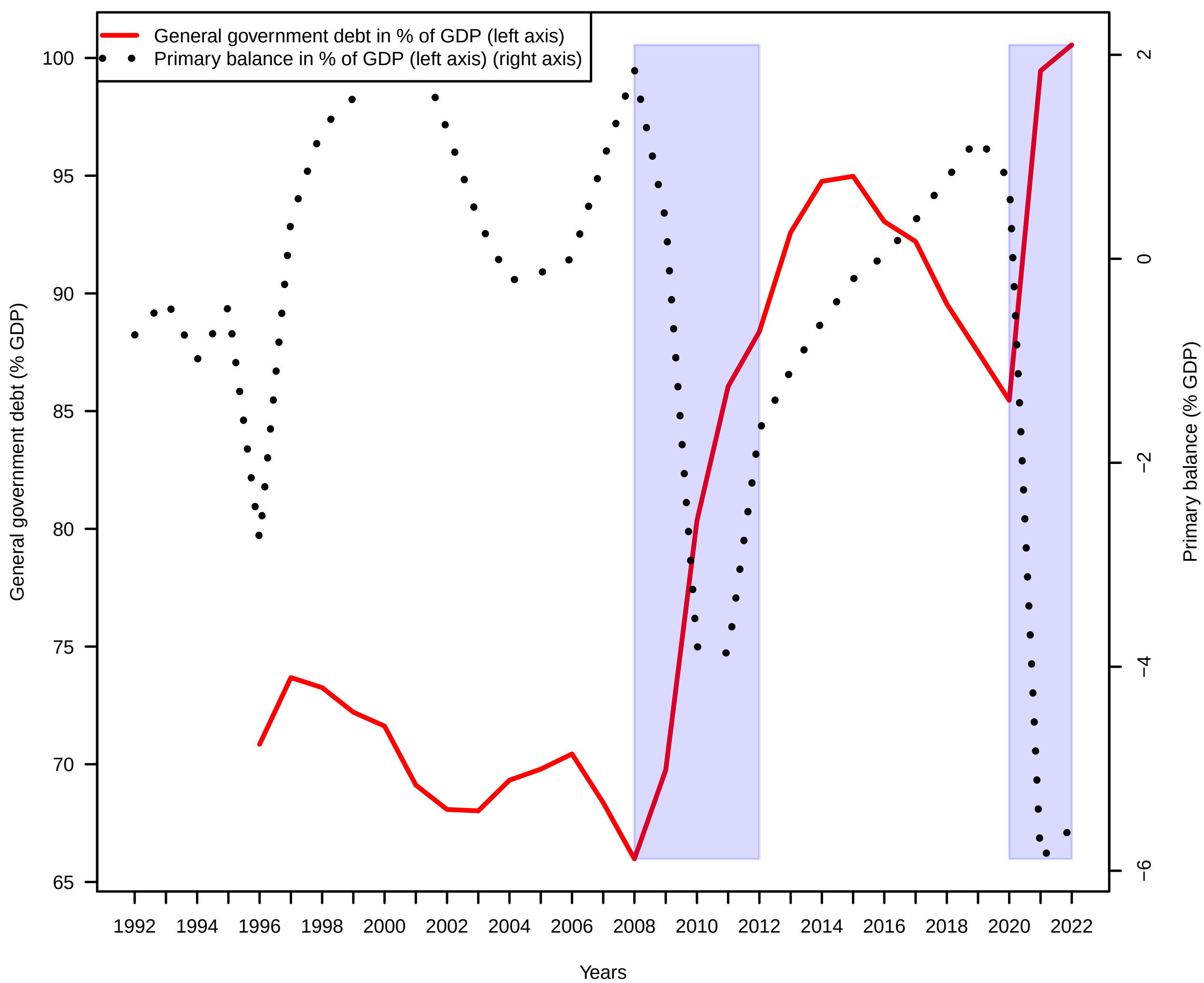
[1] The complete *2021/22 Annual Report* of the German Council of Economic Experts can be found here: https://www.sachverstaendigenrat-wirtschaft.de/en/annualreport-2021.html.

[2] Fiscal contractions are advocated for and being rolled out also in other advanced economies. The latest Hutchins Center Fiscal Impact Measure for the United States projects that fiscal policy is turning contractionary (https://www.brookings.edu/interactives/hutchins-center-fiscal-impact-measure/).

[3] The Communication *Fiscal Policy Guidance for 2023* published by the European Commission can be accessed at: https://ec.europa.eu/info/sites/default/files/economy-finance/com_2022_85_1_en_act_en.pdf.

budget balance via fiscal consolidations. For instance, following the peak of the euro-debt crisis, Member States consolidated their budget balance via austerity plans. In the period 2012-2019, the debt burden in Europe was reduced by 10 p.p. while achieving budget surpluses on average (see Fig. 4.1).

Figure 4.1: General Government debt and primary balance in % of GDP –
Euro Area (17 countries) – OECD Economic Outlook



This figure shows general government debt (red solid line) and primary balance (black dotted line) for 17 OECD Euro Area countries[a] as reported in the OECD Economic Outlook database. The blue-shaded areas represent the Great Financial Crisis (2008-2010), the euro-debt crisis (2010-2012) and the COVID-19 pandemic (2020-2022). On the left-hand side vertical axis, values for general government debt are reported in % of GDP. On the right-hand side, values for primary balance are reported in % of GDP.

[a] 17 OECD Euro Area countries: Austria, Belgium, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Latvia, Lithuania, The Netherlands, Portugal, Slovak Republic, Slovenia, Spain.

Fiscal consolidations are policy measures aimed at reducing government deficits and debt accumulation. These fiscal measures can be divided into two categories depending on whether they mainly hinge on tax hikes or spending cuts. Based on the type of fiscal adjustment that is implemented, namely *how* the consolidation is carried out, the effects on the economy can vary. Moreover, a recent growing strand of the literature called into question whether the effects of fiscal policy depend on the initial state of the economy. The focus on *when* a fiscal adjustment is implemented highlighted that fiscal policy can prove effective in certain situations while not in others. This paper studies the macroeconomic effects of both tax-based (TB) and expenditure-based (EB) fiscal consolidation announcements across different initial states of the economy. The present work focuses on the constrainedness of the monetary policy stance and the strength of the fiscal position as key state-dependencies, which are partially neglected in previous empirical studies.

I exploit a novel quarterly dataset of fiscal consolidation announcements constructed by Beetsma et al. (2021). The use of the exact moment of announcement enables to control for anticipation effects of both the legislative and implementation phase of the policy. The empirical analysis is carried out over the period 1978:Q1-2013:Q4 on 13 European economies, including Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal, Spain, Sweden and the UK. I use local projection (LP) methods developed by Jordà (2005) to estimate the state-dependent impulse response functions (IRFs) following an announcement of fiscal consolidation measures. LP methods are more flexible in dealing with the non-linearity imposed by state-dependencies compared to Vector Autoregression (VAR) models. I separate different states of the economy using an indicator variable. Thus, I construct state dummies for monetary policy stance, fiscal position and business cycle regimes. To identify states where monetary policy is constrained, I focus on constructing an indicator signalling periods in which monetary policy is near the zero lower bound (ZLB). To proxy the fiscal position, I use fiscal space indicators as in Metelli and Pallara (2020).[4] Fiscal space measures offer a broader assessment of the strength or weakness of the fiscal position than the simple debt-to-GDP ratio. Adopting Ramey and Zubairy (2018) approach, I compute cumulative multipliers for output, consumption and investment across states of the economy. This opens the door of both studying the effectiveness of fiscal consolidation announcements and testing the difference in the fiscal multipliers across states of the economy.

[4] In the definition of Heller (2005), fiscal space represents the room available in the budget of a government to provide resources for a desired purpose without jeopardizing the sustainability of its financial position or the stability of the economy.

The main findings of this study are that (i) an EB announcement of consolidation measures is contractionary when monetary policy is constrained differently from times when it is unconstrained and that, (ii) independently of the type of fiscal adjustment, consolidation announcements are not harmful when the economy is in a strong fiscal position.

Specifically, near the ZLB, an EB consolidation is recessionary contrary to the instances in which monetary policy is free to adjust. EB consolidations might be counterproductive if the ZLB is binding since they extend the length of the liquidity trap, which leads to an adverse impact on output at the margin. I find that the output multiplier is positive under *normal times* for the monetary policy stance, while the one near the ZLB is negative, being estimated to be close to -0.2. To avoid confounding effects that might arise from the high correlation between periods near the ZLB and the global financial crisis, I study whether similar findings on the effects of EB consolidations apply when the constraint on monetary policy is due to a currency union (the Eurozone in the present case) and, indeed, the bottomline results are substantiated.

As regards the fiscal position, I find that both EB and TB announcements of consolidation measures yield different effects on the economy depending on the state of fiscal space. Real GDP, consumption and investment plunge following a TB announcement when the fiscal position is weak, while real interest rate rises; on the contrary, output and the private sector do not contract when the fiscal position is strong. Following an EB consolidation announcement, under *large fiscal space*, I observe an expansionary effect on output and the private sector whose response is triggered by a large decrease in the real interest rate ($\approx -0.3\%$ on average). In contrast, under *tight fiscal space*, responses of macroeconomic variables are non-recessionary. These differences are reflected in the estimated fiscal multipliers. Focusing on TB announcements of fiscal consolidation measures, the fiscal multiplier is large and negative in *tight fiscal space*, while it is close to zero when the fiscal position is strong. In particular, one year after the announcement, the investment multiplier is as low as -3 when the fiscal position is weak. The stronger contraction under *tight fiscal space* can be explained through the expectation channel. Fiscal adjustments via tax hikes lead to sluggish growth and higher debt-to-GDP ratios. This is likely to enforce further fiscal adjustments that are distortionary. Ultimately, economic agents foresee a persistent fiscal adjustment bound to worsen economic growth for several years without ameliorating the fiscal position.

Moreover, consistently with recent empirical studies, my results highlight that TB consolidations are more contractionary than EB ones. The latter also prove to be expansionary at times, which is in line with neoclassical theory. Lastly, I observe that business cycle states are partially irrelevant for the transmission of fiscal consolidation shocks, while, as discussed, the strength of the fiscal position and the constrainedness of monetary policy prove to be crucial. Thus, there are some instances in which the initial state of the economy — *when* the consolidation is carried out — might have no influence on the propagation of the fiscal adjustment

I explore possible mechanisms driving the results across monetary policy regimes and fiscal space states by looking at the response of debt-to-GDP and confidence indicators for both private spending and investment. I find that a TB announcement of consolidation measures is unable to stabilize debt. Conversely, an EB consolidation manages to halt debt accumulation unless the economy is near the ZLB. Thus, spending-based fiscal adjustments are more effective at stabilizing debt-to-GDP without jeopardizing economic growth. Moreover, TB consolidations considerably hinder consumer confidence unless the economy is in a strong fiscal position. I also study the impact of both types of consolidations on income inequality. I show that income distribution becomes more unequal following a revenue-based fiscal adjustment due to the contraction in economic activity.

The rest of the paper is organized as follows. Section 4.2 provides a literature review, while Section 4.3 describes the dataset used to carry out the empirical analysis. Section 4.4 provides details on the empirical methodology. Section 4.5 presents and discusses the empirical results. Finally, Section 4.6 concludes.

4.2 Literature Review

Table 4.1 summarizes recent empirical studies on fiscal consolidations, reporting the methodology, sample and countries analyzed. I briefly outline also the main findings of each reported paper. The last entry of the table offers a concise summary of the present study.

This paper relates to the literature identifying fiscal consolidations and studying their effects on the economy. As regards the identification of fiscal consolidation episodes, Devries et al. (2011) build a database of fiscal consolidation measures for 17 OECD countries from 1978 to 2009 at annual frequency by following the narrative and historical approach of Ramey and Shapiro (1998) and Romer and Romer (2010).[5]

[5]The authors analyze contemporaneous policy documents to identify fiscal adjustments exclusively motivated by budget balance improvements. The historical sources examined by Devries et al. (2011) include Budget Reports, Budget Speeches, Central Bank reports, Convergence and Stability Programs submitted to the European Commis-

Table 4.1: Fiscal Consolidations: empirical studies

Paper	**Methodology**	**Sample & Countries**	**State-dependency**	**Results**
Devries et al. (2011)		1978:2009 – 17 OECD		*News*-based novel dataset.
Alesina et al. (2015)	Simulation of multi-year plans, multi-country system of MA[a], T & C FE[b]	1981:2007 – 16 OECD		TB: recessionary; EB: less costly; Private investment main driver.
Beetsma et al. (2015)	Panel Regression T & C FE, Event study	1978:2009 – 17 OECD		TB: confidence ↓; EB: confidence ≈ 0
Beetsma et al. (2021)	Panel Bayesian VAR	1978Q1:2013Q4 – 13 EU		Novel dataset of announcements; TB: recessionary, private sector crowding-out; EB: multiplier ≈ 0
Alesina et al. (2018)	Simulation of multi-year plans, Smooth-Transition VAR	1981:2014 – 16 OECD	Expansion-Recession	EB less costly than TB; Higher negative multiplier in expansion than in recession.
Fotiou (2020)	Interacted Smooth-Transition VAR	1980:2014 – 13 OECD	Expansion-Recession (joint with high-low debt)	TB recessionary & self-defeating under high-debt (no difference across exp/rec[c]); EB less costly, debt-stabilizing, contractionary under recession & high debt.
This paper	Local Projection	1978Q1:2013Q4 – 13 EU	Expansion-Recession; MP regimes; Fiscal space.	TB recessionary, private sector crowding-out, self-defeating & hinder confidence, less costly under large fiscal space; EB less costly, (multiplier ≈ 0), debt-stabilizing, contractionary near the ZLB, expansionary in large fiscal space.

a: Moving Average; *b*: Time and Country Fixed Effects; *c*: Expansion/recession.
This table summarizes recent empirical studies on fiscal consolidations, reporting the methodology, sample and countries analyzed. Also the main findings of the reported papers are briefly outlined.

sion, IMF reports and OECD Economic Surveys. In addition, the authors use country-specific sources (e.g., *Journal Officiel de la Republique Francaise* for France).

Using the information in Devries et al. (2011), Alesina et al. (2015) simulate the effects of multi-year fiscal adjustment plans differentiating between TB and EB consolidations for 16 OECD economies from 1981 to 2007. The authors find that TB fiscal corrections are recessionary; while, EB consolidations have negligible effects on output, which, on average, can be explained by small output costs and mild expansionary effects. The present work reports similar findings. Moreover, the main driver of the different results across types of consolidations in Alesina et al. (2015) is private investment, which is also pivotal in the differential effects observed in my results between TB and EB announcements.

Beetsma et al. (2015) find that fiscal consolidations have detrimental effects on confidence for both private consumption and investment using standard fixed effects annual panel regression techniques. Furthermore, using monthly consolidation announcements and an event study approach, the authors show that TB adjustments deteriorate confidence, while EB ones have negligible effects. In this paper, I observe similar responses of consumer and business confidence indicators around TB and EB announcements of consolidation measures. Confidence is considerably hindered following an EB announcement only when the economy is near the ZLB, while it is boosted when the fiscal position is strong independently of the type of fiscal adjustment.

Beetsma et al. (2021) provide evidence that there is a discrepancy in the follow-up consolidation plans by comparing the annual narratively identified plans of Devries et al. (2011) and Alesina et al. (2015) with the OECD fiscal data observed ex-post. The authors show systematic shortfalls especially in the EB measures.[6] Therefore, Beetsma et al. (2021) deem necessary to carry out an in-depth empirical analysis on the differential effects of EB and TB consolidations. In this regard, the authors build a novel quarterly narrative dataset of fiscal consolidation announcements for 13 European countries over the period 1978:Q1-2013:Q4. I adopt the same dataset of fiscal consolidation announcements in the present study. The above discussion further motivates my interest in evaluating the state-dependent effects of fiscal consolidations considering that recent empirical studies are based on Devries et al. (2011) and Alesina et al. (2015) datasets. Beetsma et al. (2021) study the effects of fiscal consolidation announcements and compute fiscal multipliers by means of Bayesian VAR techniques. Their findings confirm that TB consolidations are recessionary, leading to a large crowding-out of the private sector, while EB announcements lead to negligible effects on output and private spending. The latter

[6] For more details, see Sec. 3 of Beetsma et al. (2021).

results are very robust across different empirical analyses in the literature and consistent with the findings of the present work.

To my knowlegde, only two papers systematically analyze the state-dependent effects of fiscal consolidations: Alesina et al. (2018) and Fotiou (2020). The main limitation of both studies is the use of data at the annual frequency. This prevents from exploring various relevant state-dependencies by own admission of the authors (e.g., monetary policy regimes). By taking advantage of the higher frequency of Beetsma et al. (2021) consolidation announcements, I am able to explore the effects of fiscal adjustments close to the ZLB and also across fiscal space states other than business cycle regimes. Moreover, the present study focuses on the impact of policy announcements, purged from anticipation effects, rather than multi-year plans as explored in Alesina et al. (2018). In contrast to smooth-transition VAR as adopted in Alesina et al. (2018) and Fotiou (2020), the use of LP methods enables the researcher to locally approximate the responses of macroeconomic variables while remaining agnostic on the true data generating process (DGP). Another limitation linked to smooth-transition VAR concerns the need of imposing assumptions on the IRFs, which could lead to artificially higher multiplier depending on the initial state of the economy. Note that Ramey and Zubairy (2018) is the first study to move the above critique in reference to the work of Auerbach and Gorodnichenko (2012). The latter study employs smooth-transition VAR to study the effects of government spending shocks across business cycle states, finding a large multiplier in recession; while, using local projection, Ramey and Zubairy (2018) find no difference in government spending multipliers across expansions and recessions.

Using annual data for 16 OECD countries from 1981 to 2014, Alesina et al. (2018) is the first study analyzing both the effects of *how* and *when* fiscal consolidations are implemented. Employing the multi-year annual consolidation plans developed in Alesina et al. (2015), the authors study how the business cycle affects the response of the economy to TB and EB consolidations. One limitation of this study is the assumption of fully credible consolidation plans. The authors find strong evidence for a different output effect across TB and EB plans; the former yield large negative multipliers, while the latter have a small negative impact on economic activity. The authors observe minor evidence for larger contractionary effects of consolidations carried out under expansions than under recessions. Using a different empirical approach and dataset, my results do not highlight large differences across business cycle states, while findings on the asymmetry of effects across types of consolidation are in line with Alesina et al. (2018).

Fotiou (2020) uses an interacted smooth-transition VAR to study simultaneously the impact of the business cycle and states of high/low public debt on the transmission of TB and EB consolidations. Using narratively identified annual consolidation shocks and a panel of 13 OECD economies from 1980 to 2014, Fotiou (2020) finds that states of business cycles matter for the effects of fiscal consolidations under conditioning on debt-to-GDP regimes (high/low). On one hand, the author finds that TB consolidations are strongly recessionary and *self-defeating* when debt is high, while observing no difference across states of expansion and recession.[7] My findings lead to similar conclusions around TB announcements of consolidation measures. On the other hand, Fotiou (2020) reports that EB fiscal adjustments are less costly in terms of output losses and do stabilize public debt, proving to be largely contractionary only when the economy is both under recession and high debt. Differently from Fotiou (2020), I focus on a broader proxy for the fiscal position and use fiscal space indicators to study the state-dependent effects of fiscal consolidations. I find noticeable differences in the effects of fiscal consolidations conditioning on fiscal space indicators as measures for the strength or weakness of the fiscal position.[8]

Even though theoretical literature studying fiscal consolidations is vast, the strand comparing across TB and EB fiscal adjustments is scarce. According to Alesina et al. (2020), under a neoclassical framework, an EB consolidation adjusts the budget balance immediately leading to a positive wealth effect on private spending associated with lower future taxes; while a TB adjustment is associated with further fiscal corrections and distortions. This supports the empirical results of this paper, particularly under states of weak fiscal position. Moreover, debt stabilization is more likely under EB consolidations given that TB ones do not necessarily imply lower spending and yield a negative demand effect due to expected future taxes (Alesina et al., 2020). These findings are substantiated by various empirical studies (e.g., Attinasi and Metelli, 2017) and by the responses of debt-to-GDP shown in this paper following TB and EB consolidation announcements. Using a DSGE model, Erceg and Lindé (2012, 2013) report that

[7] *Self-defeating* is a term used to characterize fiscal consolidations when they fail to achieve their primary objective of stabilizing the stock of debt and containing deficit growth.

[8] When estimation is carried out using debt-to-GDP ratio as the state variable, I do not find difference in the two states, suggesting the importance of looking at specific indicators of fiscal space, in contrast to other variables, when studying fiscal sustainability issues. Although correlated with debt-to-GDP, fiscal space encompasses other crucial aspects: overall ability of the government to service its obligations and dynamics of public finance aggregates jointly with other key macroeconomic variables. Moreover, forward-looking nature of fiscal space contrasts with the path-dependent nature of a stock variable like debt-to-GDP. Indeed, fiscal space varies with market and economic conditions, which often change abruptly. Therefore, I avoid focusing on a single metric like the debt-to-GDP ratio, and we rely on multi-faceted indicators measuring the dynamic concept of fiscal space drawing from Metelli and Pallara (2020).

the effects of government spending cuts on output are smaller when a country conducts an independent monetary policy than when constrained by membership in a currency union or near the ZLB. These results are supported by my empirical findings near the ZLB and when monetary policy is constrained by a currency union (Euro-area).

4.3 Data

In this study, I use the dataset on consolidation announcements constructed by Beetsma et al. (2021) and, in this section, I provide details on these novel series and its pros compared to previous consolidation datasets. I also present the panel dataset and the state dummies employed to carry out the empirical analysis.

4.3.1 Fiscal consolidation announcements

In the present study, I employ the novel quarterly dataset on fiscal austerity announcements implemented by Beetsma et al. (2021), which builds partially on Devries et al. (2011) and Alesina et al. (2015). Under some circumstances, Devries et al. (2011) already report the date of announcement and, thus, Beetsma et al. (2021) use them. However, in other cases, the authors collect fiscal adjustment measures and map them into moments of announcement, namely when the policies are firstly mentioned either in the press or by the government.[9] The frequency of the announcements collected by Beetsma et al. (2021) is monthly. The authors also quantify the size of announcement by using various official documents (e.g., OECD Economic Surveys) and newspapers.[10] Beetsma et al. (2021) dataset contains 114 EB and 61 TB consolidation announcements and their average size is 1.42 and 1.14 % of GDP, respectively. On average across countries, the horizon of the consolidations extends for 1.8 years.

The announcement series are aggregated at the quarterly frequency to match the availability of fiscal macroeconomic variables and to avoid potential anticipation effects linked to information available before the policy announcement.[11] The annual consolidation datasets of Devries

[9]Further details are provided in Appendix 4.A and can be found in Sec. 4 and in the Data Construction Appendix of Beetsma et al. (2021).

[10]The magnitude of announcement for each measure is computed based on the information of the projected effects of the policy, namely the sum of the primary balance impact in % of GDP over the extension horizon of the plan.

[11]Moreover, similarly to Ramey (2011a), an announcement recorded in the first month of a quarter is assigned to the previous quarter to further avoid anticipation effects.

et al. (2011) and Alesina et al. (2015) do not take into account the impact of both implementation and legislative lags.[12] In contrast to previous studies, by using the exact moment of announcement as in Beetsma et al. (2021), I can control for anticipation effects of the legislative and implementation phase.[13] It is indeed crucial for the estimation of the true causal effects to identify the moment of announcement, namely fiscal *news*, because expected movements of budget variables induce economic agents to anticipate the response to fiscal adjustments.[14]

4.3.2 Panel dataset

In this study, I employ a quarterly panel dataset for 13 European countries over the period 1978:Q1-2013:Q4 including Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal, Spain, Sweden and the UK. The countries and the time-span in the database are the same as in Beetsma et al. (2021) to match the dataset of fiscal consolidation announcements described in Subsec. 4.3.1. Macroeconomic, financial and government budget variables are drawn from the OECD Economic Outlook, Eurostat and the IMF International Financial Statistics database. Income inequality indices are gathered using the Standardized World Income Inequality Database. Further details are provided in Appendix 4.A.

4.3.3 States of the economy

The focus of this study is on the state-dependent effects of fiscal consolidations, especially across regimes of monetary policy and fiscal position. First, I present the indicator for business cycle regimes that is mainly exploited to carry out a comparison with previous studies and shed new light on this state-dependency. Second, I show how the states for ZLB and fiscal space are constructed. Further details on the data used to construct the state dummies are provided in Appendix 4.A.

Expansion/recession

Following Auerbach and Gorodnichenko (2012) and Alesina et al. (2018), I use the seven quarters moving average (MA) of the real GDP growth as main indicator for business cycle states. In

[12]In example, Devries et al. (2011) assign the fiscal adjustment measure to the year of implementation, which opens the door to anticipation effects.

[13]As regards consolidation plans, both the legislative and implementation lags tend to be short amounting to a few months (Leeper et al., 2013). Usually, the official announcement of a fiscal consolidation coincides with the new budget presentation.

[14]This way, the shocks are not fundamental and cannot be regarded as the true shocks (Forni et al., 2014).

my case, I define states of *expansion* (*recession*) as an indicator variable taking value 1 (0) when the MA of real GDP growth is above (below) the median. In Fig. 4.B1 in Appendix 4.B, I plot the MA of GDP growth and shaded areas indicating *recession*. Moreover, as in Fotiou (2020), I employ the OECD recession dates as robustness indicator.[15]

Close-to-ZLB/normal times

To identify states where monetary policy is constrained by the ZLB, I use the short-term nominal interest rate as drawn from the OECD Economic Outlook. Specifically, I define that the economy is *close-to-ZLB* whenever the short-term nominal rate is below 75 basis points (dummy is equal to 1), otherwise it is under *normal times* (dummy takes null values). Similarly, Ramey and Zubairy (2018) define a ZLB regime for the US when the short-term interest rate is below 50 basis points. However, in my case, this stricter definition would reduce the number of observations in the *close-to-ZLB* regime, creating issues for the estimation of causal effects. In Fig. 4.B2 in Appendix 4.B, I plot the short-term rate and shaded areas indicating *close-to-ZLB* periods. Given that a limitation of the *close-to-ZLB* periods is to be highly correlated with the global financial crisis, I also construct a dummy that takes values equal to 1 after 1999:Q1 and when countries are part of the Eurozone to conduct further robustness on the effects of EB measures under states of constrained monetary policy (see Sec. 4.5).

Large/tight fiscal space

To construct the state dummies to measure the strength or weakness of the fiscal position, I employ fiscal space indicators as in Metelli and Pallara (2020).

The baseline indicator draws from Kose et al. (2017), which is simply derived from the debt accumulation accounting equation for each country i:

$$\Delta \frac{B_{i,t}}{Y_{i,t}} \approx \frac{r_{i,t} - \gamma_{i,t}}{1 + \gamma_{i,t}} \frac{B_{i,t-1}}{Y_{i,t-1}} - s_{i,t}, \tag{4.1}$$

where $\frac{B_{i,t}}{Y_{i,t}}$ is the debt-to-GDP, $\gamma_{i,t}$ is the growth rate of real GDP, $r_{i,t}$ is the long-term real interest rate and $s_{i,t}$ is the primary balance over GDP.[16] The level of primary surplus that would stabilize public debt (i.e., $\Delta \frac{B_{i,t}}{Y_{i,t}} = 0$) is calculated simply from Eq. 4.1. Then, the fiscal space indicator is

[15] As a further robustness, similarly to Ramey and Zubairy (2018), I use the unemployment rate. I construct a dummy that takes value 1, signalling *recession*, when unemployment rate is above its median.

[16] Eq. 4.1 should include also the stock-flow adjustments not to hold as an approximation. Stock-flow adjustments comprise factors that affect debt but are not included in the budget balance (such as acquisitions or sales of

defined as the distance between such primary surplus and the realized one. Thus, by ciclycally adjusting the terms in Eq. 4.1, the indicator is given by the following equation:

$$FS1_{i,t} = \left(\frac{r_{i,t} - \tilde{\gamma}_{i,t}}{1 + \tilde{\gamma}_{i,t}}\right) d^{c.a.}_{i,t-1} - s^{c.a.}_{i,t}, \tag{4.2}$$

where $s^{c.a.}_{i,t}$ is the cyclically adjusted primary surplus over potential GDP, $d^{c.a.}_{i,t}$ is the cyclically adjusted debt-to-GDP, $\tilde{\gamma}_{i,t}$ is the real potential GDP growth and $r_{i,t}$ is the long-term interest rate. This measure represents my benchmark indicator of fiscal space since it summarizes the many features a proxy for fiscal space should contain: considerations of fiscal sustainability, debt dynamics, interest rate, output growth and fiscal policy stance.[17] The state dummy is constructed so that equals 1 (0) when fiscal space is *tight* (*large*), namely $FS1$ is above (below) its median. In Fig. 4.B3 in Appendix 4.B, I plot $FS1$ and shaded areas indicating *tight fiscal* space periods.

For robustness, I also adopt an alternative measure to capture fiscal room defined as *de facto* fiscal space originally built by Aizenman and Jinjarak (2010) and Aizenman et al. (2013). This measure is computed as the ratio of deficit-to-GDP over the *de facto* tax base and is inversely related to the tax-years needed to compensate deficits. Using cyclically adjusted variables, I define *de facto* fiscal space as follows:

$$FS2_{i,t} = \frac{(def^{c.a.}_{i,t})}{(rev^{c.a.}_{i,t})}, \tag{4.3}$$

where $def^{c.a.}_{i,t}$ stands for the cyclically adjusted deficit-to-potential GDP and $rev^{c.a.}_{i,t}$ represents the cyclically adjusted government revenues over potential GDP. This measure outlines the tax capacity of a country to weigh deficits.[18] The state dummy is constructed so that equals 1 (0) when fiscal space is *tight* (*large*), namely $FS2$ is above (below) its median.

financial assets). For sake of simplicity, we focus on the "snowball-effect" side of the debt accumulation equation and on the government budget balance for the construction of $FS1$.

[17] In particular, this fiscal space measure highlights times of rapid debt accumulation due to inherent inability to roll-over debt via primary surpluses, crucial characteristics of the fiscal position of the government. Indeed, a simpler measure, such as the distance between the government funding cost and output growth, does not consider the evolution of the primary balance itself and the stock-flow adjustments. Even though according to Blanchard (2019b), as long as the yields are lower than the GDP growth rate countries have fiscal space (see also Mauro and Zhou, 2019). However, there is a growing consensus that such argument is incomplete. Moreover, Jiang et al. (2019) find that the discount factor on government debt is decoupled from the yields on bonds, which would nuance the claims in Blanchard (2019b).

[18] Indeed, $FS2$ highlights periods of high deficit overhangs with respect to the government inability to raise revenues via tax collection.

4.4 Empirical methodology

4.4.1 Local Projection

Local projection (LP) methods à la Jordà (2005) have become a prominent methodology to estimate IRFs. The difference between LP and VAR resides in LP not assuming any data generating process (DGP) for the data at hand, making them a less parametric tool than VAR that instead are fully parametric. As a direct consequence, if one believes the economy to be structurally well characterized by a set of stochastic equations as in a VAR, then estimating IRFs with a VAR will yield more reliable and more efficient estimates. However, if the researcher does not have a strong belief for the data to be generated by a VAR, then a case exists for estimating IRFs with nonparametric methods as LP. The latter are "local" in the sense that they target the relevant IRF at each horizon, while VAR models globally approximate endogenous variables' responses. The latter represent the true IRFs if the VAR is the true DGP, while, when using LP, locally approximating the IRF at each horizon remains agnostic on the true DGP.[19] Therefore, when model uncertainty is a concern, LP possibly represents a better option as opposed to VAR that cannot account for uncertainty in the DGP by construction, but only for uncertainty in parameters conditional on a given DGP. LP allows to use different variables for every equation and for every horizon, while still estimating the model with OLS. VAR models are instead very rigid because if the researcher wants to add restrictions on the specification then OLS estimaton is precluded, and more complicated algorithms have to be applied.

4.4.2 State-dependent Local Projection and model specification

The aim of this study is to estimate the state-dependent responses of macroeconomic variables to fiscal consolidation announcements. The non-linearity adopted to study how fiscal policy is transmitted across two different states of the economy is very simple: the two regimes are separated using an indicator variable. Note that this type of non-linearity used is equivalent to the one employed by Ramey and Zubairy (2018). Other studies use smooth-transition LP

[19] Recently, Plagborg-Møller and Wolf (2021) find that *linear* LP and VAR estimate (in theory) the same IRFs. However, the authors state that "*the relative mean-square error of the two methods [...] necessarily depends on assumptions about the data generating process (DGP). VAR estimators are optimal if the true DGP is exactly a finite-order VAR, but this is rarely the case in theory or practice*". In addition, "[...] *we only explore linear estimators. The equivalence of VAR and LP estimators does not apply if we augment the regressions with non-linear terms*". Finally, this equivalence result holds true only when no constraint is imposed on the lag structure, meaning that only IRFs from linear VAR and linear LP with an infinite number of lags coincide.

models (Tenreyro and Thwaites, 2016). In such models, parameters can smoothly switch across the two regimes instead of hastily changing from one state to the other (Granger and Terasvirta, 1993). Even though a smooth transition across regimes might be desirable, one needs to calibrate curvature and location parameters that are key to the estimation and, ultimately, affect the set of resulting IRFs. In principle, these pivotal parameters could be estimated. To do so, the researcher would need to collect a lot of data around the transition of the regime, which represents a highly unlikely scenario in macroeconomic applications.[20,21] Therefore, I employ the more simple and robust approach of using a dummy variable. This strategy yields a cleaner interpretation of the coefficients as exact average causal effects within a given state.[22]

Specifically, I am interested in estimating the following regression:

$$\begin{aligned}(y_{i,t+h} - y_{i,t-1}) =& \mathcal{S}_{i,t-1}\left[\alpha_{A,h,i} + \delta_{A,h,i}t + \beta_{A,h}\eta_{i,t} + \Phi_{A,h}(L)\mathbf{X}_{i,t}\right] + \\ &+ (1-\mathcal{S}_{i,t-1})\left[\alpha_{B,h,i} + \delta_{B,h,i}t + \beta_{B,h}\eta_{i,t} + \Phi_{B,h}(L)\mathbf{X}_{i,t}\right] + \varepsilon_{i,t+h} \\ & h = 0,1,2,\ldots,H\end{aligned} \tag{4.4}$$

The state dependency is given by the lagged dummy variable $\mathcal{S}_{i,t-1}$ that indicates the state of the economy in $t-1$ as described in Sec. 4.3, while subscripts $j = [A, B]$ denote the two states of the economy that are analyzed (e.g., *tight* and *large fiscal space*). At each horizon h, the estimated parameter $\beta_{j,h}$ in Eq. 4.4 measures the impact on the interest variable, $(y_{i,t+h} - y_{i,t-1})$, of a change in $\eta_{i,t}$, which stands for the TB or EB fiscal consolidation announcement.[23,24] Hence, the sequence of parameters $\{\beta_{j,h}\}_{h=0}^{H}$ represents the estimated points of the IRF across all horizons. In Eq. 4.4, $\alpha_{j,h,i}$ represents the country fixed effects. I also include country-specific time trend $(\delta_{j,h,i}t)$.[25] In the baseline specification, the bag of controls $(\Phi_{j,h}(L)\mathbf{X}_{i,t})$ consists of lagged values of log-change in real output (*gdp*), real consumption (*cons*) and real investment (*inv*), change

[20] Teräsvirta (1994) discuss those estimation issues in detail.

[21] Similar critique can be moved to studies using smooth-transition VAR models such as Auerbach and Gorodnichenko (2012), Alesina et al. (2018) and Fotiou (2020).

[22] The Threshold VAR would represent the VAR counterpart of the model described. However, as anticipated in Subsec. 4.4.1, when degrees of freedom are a concern, doubling the number of parameters to be estimated on an already over-parametrized model might be undesirable. Therefore, LPs are more robust to non-linearities in the sense that one can model almost any non-linear function, something that cannot be said when dealing with a VAR, where the non-linearities that can be introduced are far less and generate significantly more problems in estimation.

[23] The size of the announcement shock is equal to 1% of GDP.

[24] As shown in Ramey and Zubairy (2018) for the Blanchard and Perotti (2002) shock, this way is equivalent of identifying the relevant shock via Choleski ordering in a VAR. Indeed, the latter identification strategy is also carried out by Beetsma et al. (2021) in the context of a Bayesian VAR and consolidation announcement shocks.

[25] My findings are robust to the exclusion of the country-specific time trend. In Appendix 4.D, specifically in Figs. 4.D8- 4.D13, I report the results excluding the country-specific time trend from the specification.

in government revenues (*rev*) and spending (g) in % of GDP, change in real long-term interest rate in percentage points (r) and consolidation announcements in % of GDP (η).[26,27] Note that the left-hand side in Eq. 4.4, namely $(y_{i,t+h} - y_{i,t-1})$, implies that I estimate cumulative IRFs. Given that an intrinsic issue of LP is autocorrelation in the measurement errors, clustering at the country level is common practice (Jordà et al., 2015). This way, standard errors are robust to heteroskedasticity and autocorrelation. Given the interconnectedness of the countries included in my panel (see Sec. 4.3), I am concerned about potential correlation across countries during the same time period and, thus, I also cluster the standard errors at the quarter level.[28,29]

4.4.3 Fiscal multipliers

I compute cumulative fiscal multipliers as described in Beetsma et al. (2021). Specifically, multipliers are computed as the cumulative percent change of output divided by the cumulative increase (decrease) in revenues (spending) in percent of GDP over a horizon of h periods following a TB (EB) consolidation announcement. Same logic applies to consumption and investment multipliers. I also compute *primary balance*-based multipliers, for which the denominator is the cumulative change in primary surplus in percent of GDP. As regards the estimation method, I follow the Ramey and Zubairy (2018) approach to estimate fiscal multipliers. I regress cumulative changes in output, consumption and investment on cumulative changes in government spending and revenues in % of GDP ($-\sum_{j=0}^{h} \Delta g_{t+j}$ and $\sum_{j=0}^{h} \Delta rev_{t+j}$) instrumented by the EB and TB consolidation announcements, respectively.[30] This approach has the advantage of opening the door to direct inference. In particular, it is possible to easily test differences in the multipliers across states.

[26] The set of controls is large and also lagged values of the announcements are included to reduce concerns linked to exogeneity of the announcements constructed by Beetsma et al. (2021). A more extensive discussion on this point is provided in Appendix 4.F.

[27] I include 2 lags of the controls in the regression, which is in line with lags adopted in previous empirical studies on fiscal consolidation effects that range between 1 and 4 (e.g., Alesina et al., 2018; Beetsma et al., 2021, among others). Moreover, by fitting a VAR under the baseline specification and using the Bayesian information criterion (BIC), the optimal number of lags results to be equal to 2. Therefore, including 2 lags is both parsimonious and in line with the literature. However, different number of lags lead to similar results. In Appendix 4.D, specifically in Figs. 4.D14- 4.D19, I report the results obtained by including 4 lags of the controls in the regression.

[28] However, results are robust if standard errors are clustered only at the country level.

[29] Additionally, results are robust if I allow for standard errors to be clustered only at the country level and country-specific time trends are dropped.

[30] In the case of *primary balance*-based multipliers, cumulative changes in output, consumption and investment are regressed on cumulative changes in primary balance in % of GDP ($\sum_{j=0}^{h} \Delta rev_{t+j} - \sum_{j=0}^{h} \Delta g_{t+j}$) instrumented by the EB and TB consolidation announcements.

4.5 Results

This section presents the results, reporting the state-dependent effects of fiscal consolidation announcements. The main findings of the present study prove that both the constrainedness of monetary policy and the strength of the fiscal position are key for the transmission of an announcement of consolidation measures.
This section is organized as follows. Subsection 4.5.1 briefly shows the effects of TB and EB consolidations under the linear case and sheds new light on the impact of consolidation announcements across states of expansion and recession. Subsections 4.5.2 and 4.5.3 focus on the main findings of the paper: the effects of TB and EB consolidations across monetary policy regimes and fiscal space states. Robustness checks are reported within the results section. Lastly, I present the impact of consolidation announcements on debt-to-GDP, confidence and income inequality.

In Table 4.2, I provide an outline for the results of this paper and summarize its salient findings to help the reader navigate through this section and to easily locate the relevant figures and tables. Specifically, Figures 4.2, 4.3, 4.4, 4.5, 4.6 and 4.7 present the cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB and an EB consolidation announcement. The projection horizon extends over 4 years after announcement (measured in quarters). The top of each figure features the responses under the linear case, namely in absence of state-dependencies, using grey-shaded confidence intervals (C.I.). The mid- and bottom-part of the graphs report the IRFs under two different states of the economy whose C.I. are shaded in red and in blue.[31] Tables 4.3, 4.5 and 4.7 report cumulative output multipliers. As for the IRFs, the projection horizon extends over 4 years. The left-hand side panels of the tables report the fiscal multipliers following a TB consolidation, while, the right-hand side ones present the multipliers following an EB one. In each panel of the tables, the first column reports the multiplier under the linear case. The second and third columns show the multipliers under two different states of the economy. In each panel of the tables, the last two columns report the p-values arising from testing the difference between multipliers in

[31] Confidence intervals are reported at the 90% significance level. Note also that the state-dependent IRFs arise from the estimation of Eq. 4.4.

the two states using respectively clustered standard errors and Anderson-Rubin (AR) C.I.[32] Tables 4.4, 4.6 and 4.8 report the output multipliers for which the ratio is taken over the cumulative response of primary balance (*pb*). The estimates for consumption and investment multipliers are shown in Tables from 4.C1 to 4.C12 in Appendix 4.C.

[32] Anderson-Rubin C.I. are used to account for the possibility that the announcement series is a weak instrument.

Table 4.2: Results: outline and summary

Panel A - Linear (Sec. 4.5.1)

	TB[a]	**EB**[b]
IRFs	Fig. 4.2: Recessionary; GDP, Cons., Inv. ↓; LT int. rate[c] ↑	Fig. 4.3: Non-contractionary; mildly expansionary; LT int. rate ↓
Multiplier, Tabs. 4.3, 4.4	-0.5; *pb*[d]: -0.3	0.1; *pb*: 0.08

Panel B - Expansion/Recession (Sec. 4.5.1)

	Expansion		**Recession**	
	TB	**EB**	**TB**	**EB**
IRFs	Fig. 4.2: Similar to *Linear*[e]; LT int. rate ≈ 0	Fig. 4.3: Similar to *Linear*; Expansionary at end of *h*[f]	Fig. 4.2: Similar to *Linear*	Fig. 4.3: Negligible effects
Multiplier, Tabs. 4.3, 4.4	-0.5; *pb*: -0.2	0.15; *pb*: 0.09	-0.5; *pb*: -0.3	0.05; *pb*: 0.03

Summary: *No large difference across states (statistically significant difference for EB fiscal consolidation announcements only at the end of the projection).*

Panel C - Normal times/Close-to-ZLB (Sec. 4.5.2)

	Normal times		**Close-to-ZLB**	
	TB	**EB**	**TB**	**EB**
IRFs	Fig. 4.4: Recessionary; Short-lived contraction in Inv.	Fig. 4.5: Expansionary; GDP, Cons., Inv. ↑; LT int. rate ↓	Fig. 4.4: Recessionary; Persistent contraction in Inv.; LT int. rate ↑	Fig. 4.5: Recessionary; GDP, Cons., Inv. ↓
Multiplier, Tabs. 4.5, 4.6	-0.8; *pb*: -0.2	0.3; *pb*: 0.15	-0.7; *pb*: -0.4	-0.2; *pb*: -0.2

Summary: *Statistically significant difference across states following an EB fiscal consolidation announcement; only minor differences after a TB consolidation.*

Panel D - Large/Tight Fiscal Space (Sec. 4.5.3)

	Large FS[g]		**Tight FS**	
	TB	**EB**	**TB**	**EB**
IRFs	Fig. 4.6: Non-contractionary; LT int. rate ≈ 0	Fig. 4.7: Mildly expans.; LT int. rate ↓	Fig. 4.6: Recessionary; GDP, Cons., Inv. ↓; LT int. rate ↑	Fig. 4.7: Negligible eff.
Mult., Tabs. 4.7, 4.8	0.3; *pb*: 0.1	0.25; *pb*: 0.15	-1.3; *pb*: -0.5	0.08; *pb*: 0.06

Summary: *Statistically significant difference across states, particularly following a TB fiscal consolidation announcement.*

a: Tax-Based consolidation announcement; *b*: Expenditure-Based consolidation announcement; *c*: LT int. rate stands for the long-term real interest rate; *d*: *pb* denotes that the average multiplier reported is the one based on *primary-balance*; *e*: *Linear* refers to the initial state of the economy without any state dependencies whose results are summarized in **Panel A**; *f*: *h* stands for the projection horizon; *g*: FS stands for fiscal space.

This table provides an outline of the main results and it summarizes the salient findings of the paper to help the reader navigate through Sec. 4.5 and to easily locate the relevant figures and tables. The fiscal multipliers reported in each panel represent the average point estimates of output multipliers across projection horizons (excluding outliers). The *Summary* at the bottom of each panel summarizes the comparison of findings across states of the economy.

4.5.1 Linear case & Expansion/Recession

In Panel A of Table 4.2, I summarize the main findings under the linear case, namely in absence of state-dependencies: TB consolidations show large contractionary effects on the economy, while EB ones have no detrimental effects. I take Figures 4.2 and 4.3 to describe the IRFs under the linear case whose C.I. are shaded in grey. In Figure 4.2, government revenues exhibit a hump-shaped positive response peaking at around 1% of GDP after 1 year following a TB consolidation announcement, while government spending does not significantly respond. The effect of a TB consolidation yields a negative and persistent effect on output and consumption. Moreover, private investment shows a strong negative response peaking at around −2.5% after 2 years. I also observe an increase in the long-term real interest rate, which triggers the crowding out effect of private spending and investment. On the contrary, following an EB consolidation announcement and a consequent plunge in government spending, Figure 4.3 shows a negligible impact on output, consumption and investment, while real interest rate decreases.[33] These findings are also reflected in the estimated output multipliers (Tables 4.3 and 4.4). After a TB consolidation announcement, the output multiplier is negative, but smaller than 1, being approximately equal to −0.5 on average across projection horizons (as reported also in Panel A of Table 4.2).[34] The estimated output multiplier following an EB consolidation in Table 4.3 is close to zero as in Beetsma et al. (2021). However, towards the end of the projection horizon, the fiscal multiplier is positive. Results under both types of consolidations are also supported by the estimates for consumption and investment multipliers in Tables 4.C1-4.C4 in Appendix 4.C.[35] Under the linear case, I find that TB consolidations are more contractionary than EB ones, which actually feature a negligible impact on GDP.[36] This result is consistent with Alesina et al. (2020), which show that, under a neoclassical framework, an EB consolidation adjusts the budget balance immediately leading to a positive wealth effect on private

[33] Similar response of the long-term real interest rate is observed in Beetsma et al. (2015), which describe the effect of a spending cut as an effective way to restore sovereign confidence.

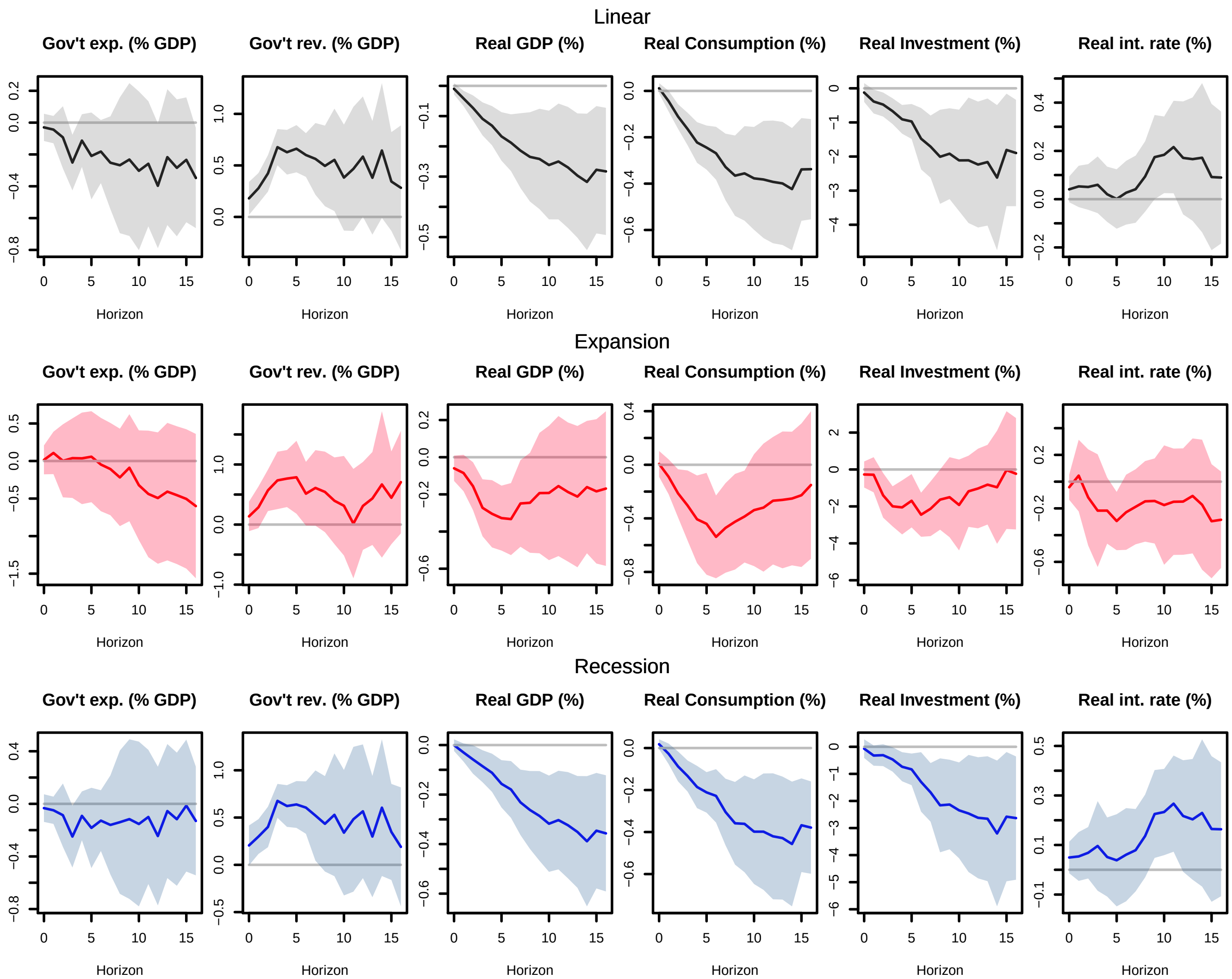
[34] I find a smaller negative output multiplier compared to recent studies. Nonetheless, this result is consistent with previous empirical evidence showing that positive revenue shocks are contractionary with an estimated output multiplier equal to −0.5 (Barro and Redlick, 2011).

[35] In particular, I find that a TB consolidation strongly crowds out private investment, which shows a negative multiplier as low as −4 after 2 years (Table 4.C3).

[36] As correctly explained in Fotiou (2020), a cut in expenditure is not distortionary, while tax hikes might imply distortions. Thus, the effect on GDP growth is less pronounced. Moreover, note also that government spending is often wasteful.

spending associated with lower future taxes; while a TB adjustment is associated with further fiscal corrections and distortions.[37]

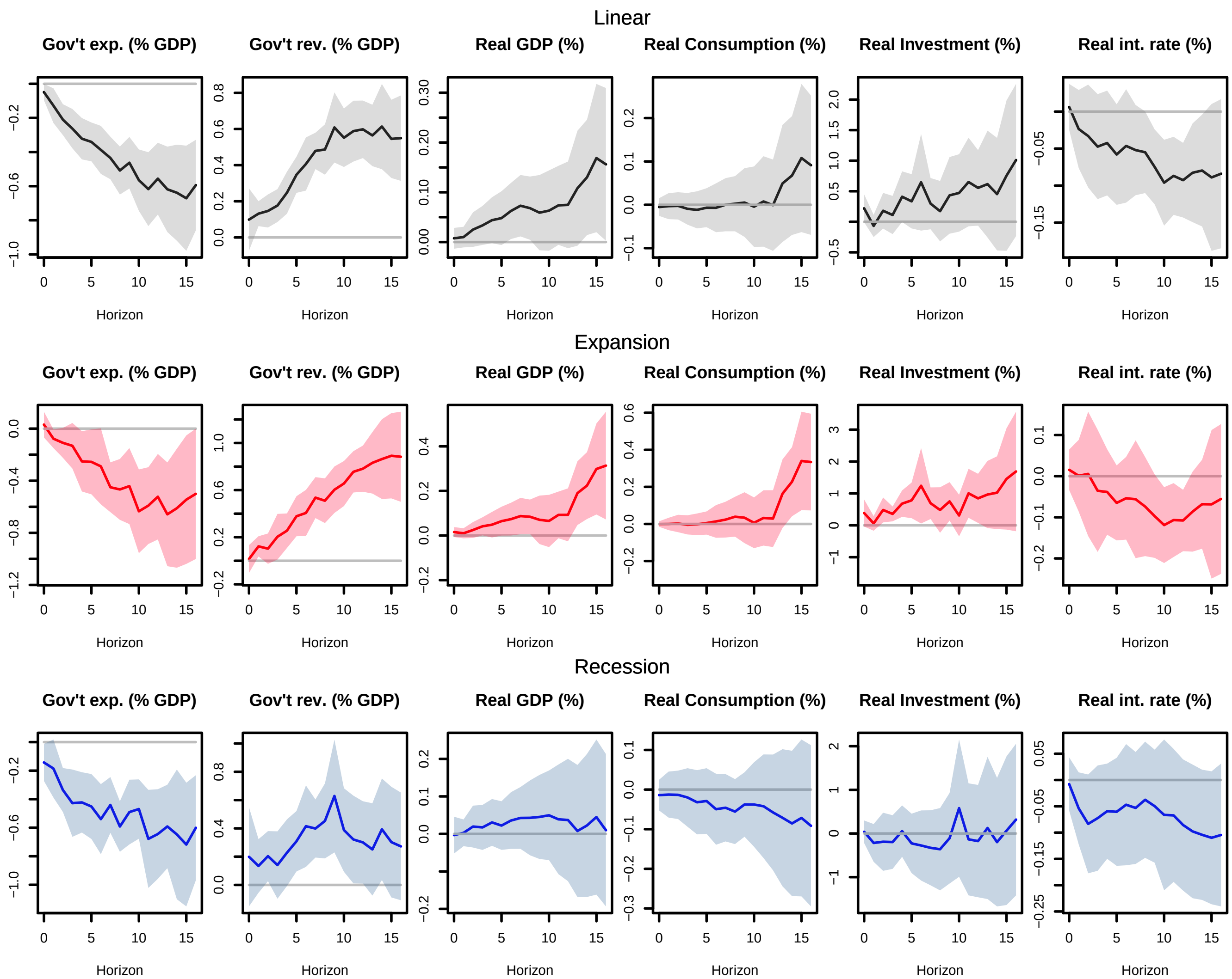
Figure 4.2: IRFs: TB consolidation announcement - *Expansion/Recession*



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

[37]Even under a new-Keynesian framework, Alesina et al. (2017) reach similar conclusions if the fiscal adjustments are long-lasting. A persistent government spending cut implies enduring higher transfers, which raise private consumption and partially compensate lower government spending. Thus, to a high persistence of government cuts corresponds a diminished adverse impact on aggregate demand. Due to price rigidities, firms must reduce their labor demand and, when the wealth effect on aggregate demand increases, output falls less. Conversely, in the case of a tax increase, the output effect is purely explained by shifts in aggregate supply that imply labor distortions.

Figure 4.3: IRFs: EB consolidation announcement - *Expansion/Recession*



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

In Panel B of Table 4.2, I summarize the main findings under states of *expansion* and *recession*: results are similar to the linear case both following a TB and an EB fiscal consolidation announcement, showing no large difference across business cycle states. However, in *expansion* and at the end of the projection horizon, EB consolidations have expansionary effects. First, Fig. 4.2 shows that a TB announcement yields a similar contractionary impact across states of *expansion* (C.I. in red) and *recession* (C.I. in blue) on real GDP, consumption and investment in the first two years. Second, Fig. 4.3 does not highlight differences in the IRFs across business cycle states following an EB announcement. Only at the end of the projection horizon, in *expansion*, responses of output and private sector exhibit an expansionary behaviour (in correspondence of a decrease in the real rate).

I also compare fiscal multipliers in Tables 4.3 and 4.4 across states of business cycle. The left panels of the tables highlight the absence of statistical difference across states of business cycles following a TB consolidation. Looking at the right-hand side panels, similar finding applies to EB consolidations aside from the last year of the projection horizon: under *expansion*, the output multiplier is positive and statistically different from the one in *recession*. This is illustrated by the resulting p-values from testing the difference across state-dependent multipliers (see the last two columns of the right-hand panels of Tables 4.3 and 4.4). This finding corroborates theory of *expansionary austerity* of Giavazzi and Pagano (1990); Alesina and Ardagna (2010) and neoclassical theory.[38],[39] Results are confirmed also by consumption and investment multipliers in Tables 4.C1- 4.C4 in Appendix 4.C.[40]

[38] A strand of the empirical literature on the effects of fiscal consolidations finds evidence for expansionary effects of EB fiscal consolidations (Giavazzi and Pagano, 1990; Blanchard, 1990; Alesina and Ardagna, 2010), which are dubbed as *expansionary austerity* episodes.

[39] The expansionary effect of an EB consolidation is even more evident by looking at private investment multipliers in Tables 4.C3,4.C4 in Appendix 4.C.

[40] As robustness, I use also OECD recession dates and unemployment as proxies for the state of the business cycle (see Sec. 4.3). Indeed, Tables 4.E1- 4.E6 in Appendix 4.E confirm the conclusions on the fiscal multipliers. Same applies to the response of the endogenous variables as shown in Figures 4.D1- 4.D4 in Appendix 4.D.

Table 4.3: Output multiplier under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					Expenditure-Based Fiscal Consolidation				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.15***	-0.28	-0.10*	0.436	0.400	0.08	-0.07	0.08	0.730	0.791
	(0.06)	(0.23)	(0.06)			(0.11)	(0.24)	(0.11)		
2	-0.17***	-0.27**	-0.15**	0.356	0.396	0.12	-0.01	0.11	0.987	0.987
	(0.06)	(0.12)	(0.07)			(0.10)	(0.47)	(0.09)		
3	-0.16***	-0.37**	-0.13***	0.046	0.123	0.13	0.07	0.10	0.918	0.915
	(0.04)	(0.15)	(0.03)			(0.09)	(0.19)	(0.08)		
4	-0.21***	-0.39**	-0.19***	0.127	0.213	0.14*	0.06	0.10	0.983	0.983
	(0.06)	(0.16)	(0.06)			(0.08)	(0.22)	(0.07)		
5	-0.25***	-0.40**	-0.26***	0.319	0.324	0.14	0.20	0.07	0.525	0.474
	(0.06)	(0.18)	(0.07)			(0.09)	(0.26)	(0.08)		
6	-0.31***	-0.61*	-0.32***	0.369	0.355	0.16*	0.20	0.08	0.374	0.335
	(0.08)	(0.33)	(0.11)			(0.08)	(0.15)	(0.08)		
7	-0.38**	-0.39	-0.47**	0.934	0.934	0.17**	0.26	0.07	0.247	0.173
	(0.15)	(0.33)	(0.23)			(0.07)	(0.18)	(0.06)		
8	-0.47**	-0.43	-0.61*	0.770	0.780	0.13**	0.18*	0.04	0.200	0.209
	(0.22)	(0.45)	(0.34)			(0.06)	(0.10)	(0.05)		
9	-0.44*	-0.45	-0.54	0.941	0.942	0.13	0.17	0.02	0.166	0.166
	(0.24)	(0.61)	(0.35)			(0.08)	(0.11)	(0.06)		
10	-0.69	-0.55	-0.94	0.817	0.834	0.11	0.12	0.02	0.122	0.155
	(0.51)	(0.97)	(0.84)			(0.07)	(0.08)	(0.04)		
11	-0.54	-3.07	-0.58	0.925	0.696	0.12*	0.16*	0.02	0.105	0.105
	(0.37)	(24.64)	(0.36)			(0.07)	(0.09)	(0.04)		
12	-0.46*	-0.55	-0.52**	0.973	0.973	0.13	0.20**	0.00	0.036	0.037
	(0.26)	(0.93)	(0.26)			(0.09)	(0.08)	(0.05)		
13	-0.78	-0.45	-0.99	0.508	0.472	0.17	0.21***	0.01	0.025	0.035
	(0.62)	(0.70)	(0.77)			(0.11)	(0.07)	(0.07)		
14	-0.49**	-0.22	-0.55***	0.339	0.549	0.20**	0.24***	0.01	0.022	0.030
	(0.21)	(0.38)	(0.18)			(0.10)	(0.07)	(0.06)		
15	-0.81	-0.37	-0.75**	0.317	0.389	0.25*	0.24***	0.06	0.155	0.215
	(0.63)	(0.59)	(0.33)			(0.13)	(0.06)	(0.09)		
16	-1.00	-0.22	-1.26	0.511	0.116	0.26*	0.25***	0.04	0.178	0.256
	(1.24)	(0.37)	(1.54)			(0.15)	(0.07)	(0.11)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.4: Output mult. (*pb*) under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.13***	-0.34	-0.09*	0.436	0.330	0.04	-0.02	0.05	0.270	0.393
	(0.04)	(0.34)	(0.05)			(0.05)	(0.06)	(0.05)		
2	-0.14**	-0.23	-0.12	0.478	0.472	0.07	-0.00	0.08	0.411	0.490
	(0.06)	(0.16)	(0.08)			(0.06)	(0.08)	(0.07)		
3	-0.12***	-0.33**	-0.09***	0.090	0.123	0.08	0.02	0.08	0.363	0.384
	(0.03)	(0.16)	(0.02)			(0.05)	(0.05)	(0.06)		
4	-0.18***	-0.33**	-0.16***	0.190	0.201	0.08*	0.02	0.08*	0.489	0.515
	(0.04)	(0.16)	(0.05)			(0.04)	(0.07)	(0.05)		
5	-0.19***	-0.34*	-0.20***	0.314	0.259	0.07	0.07	0.04	0.659	0.659
	(0.05)	(0.19)	(0.06)			(0.04)	(0.08)	(0.05)		
6	-0.24***	-0.40	-0.25***	0.498	0.409	0.08*	0.08	0.05	0.523	0.537
	(0.06)	(0.30)	(0.08)			(0.04)	(0.06)	(0.05)		
7	-0.26***	-0.24	-0.32***	0.987	0.987	0.08**	0.12*	0.04	0.222	0.234
	(0.08)	(0.23)	(0.12)			(0.03)	(0.07)	(0.04)		
8	-0.31**	-0.22	-0.40**	0.763	0.782	0.07**	0.10*	0.03	0.220	0.246
	(0.12)	(0.23)	(0.19)			(0.03)	(0.06)	(0.04)		
9	-0.31**	-0.23	-0.38**	0.955	0.956	0.06	0.10	0.01	0.163	0.180
	(0.12)	(0.34)	(0.18)			(0.04)	(0.07)	(0.03)		
10	-0.38**	-0.18	-0.52**	0.638	0.688	0.06	0.08	0.01	0.147	0.165
	(0.16)	(0.26)	(0.24)			(0.04)	(0.06)	(0.03)		
11	-0.35***	-0.16	-0.39***	0.854	0.869	0.06*	0.11*	0.02	0.096	0.119
	(0.13)	(0.30)	(0.11)			(0.04)	(0.06)	(0.02)		
12	-0.27***	-0.14	-0.31***	0.704	0.732	0.06*	0.13**	0.00	0.022	0.042
	(0.10)	(0.21)	(0.08)			(0.04)	(0.06)	(0.03)		
13	-0.50**	-0.15	-0.59***	0.335	0.403	0.09*	0.14***	0.00	0.015	0.026
	(0.21)	(0.20)	(0.21)			(0.05)	(0.05)	(0.05)		
14	-0.34***	-0.09	-0.36***	0.266	0.394	0.10**	0.16***	0.01	0.017	0.025
	(0.13)	(0.14)	(0.10)			(0.05)	(0.05)	(0.04)		
15	-0.48	-0.11	-0.48**	0.228	0.184	0.14**	0.17***	0.04	0.158	0.161
	(0.29)	(0.16)	(0.22)			(0.07)	(0.05)	(0.06)		
16	-0.45*	-0.08	-0.52	0.302	0.123	0.14**	0.19***	0.03	0.126	0.110
	(0.26)	(0.14)	(0.34)			(0.07)	(0.06)	(0.07)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

4.5.2 Normal times/Close-to-ZLB

In Panel C of Table 4.2, I summarize the salient findings under states of unconstrained and constrained monetary policy: EB consolidations are recessionary when the economy is near the ZLB, while they have expansionary effects otherwise; the effects of TB consolidations show no large difference across states of monetary policy constrainedness. Specifically, Figs. 4.4 and 4.5 report IRFs across states of monetary policy regimes, *normal times* (C.I. shaded in red) and *close-to-ZLB* (C.I. shaded in blue), following a TB and an EB consolidation announcement, respectively. Differences across states are minor following a TB consolidation. Fig. 4.4 shows a mild rise in the real long-term interest rate and a more persistent contraction in private investment near the ZLB, which, on the contrary, is short-lived during a normal monetary policy regime. However, Fig. 4.5 shows striking differences across monetary policy regimes following an EB announcement. Real GDP, consumption and investment exhibit a large contraction under the *close-to-ZLB* state, while they rise under *normal times*.[41] This difference is noticeable also in the estimates for output multipliers in Tables 4.5 and 4.6. On the right-hand side panels of the tables, across all projection horizons, the multiplier under *normal times* is positive and significant, while the one under the *close-to-ZLB* state is negative and significant.[42] The former is around 0.3 on average across horizons, while the latter is estimated to be close to -0.2 (see also Panel C of Table 4.2). Tests for the difference in the multipliers across states yield p-values confirming this finding (see the last two columns of the right-hand panels in Tables 4.5 and 4.6). Equivalent differences are found in the estimation for consumption and investment multipliers in Tables 4.C5- 4.C8 in Appendix 4.C. Comparing multipliers following a TB consolidation, I do not observe significant differences across monetary policy regimes.[43]

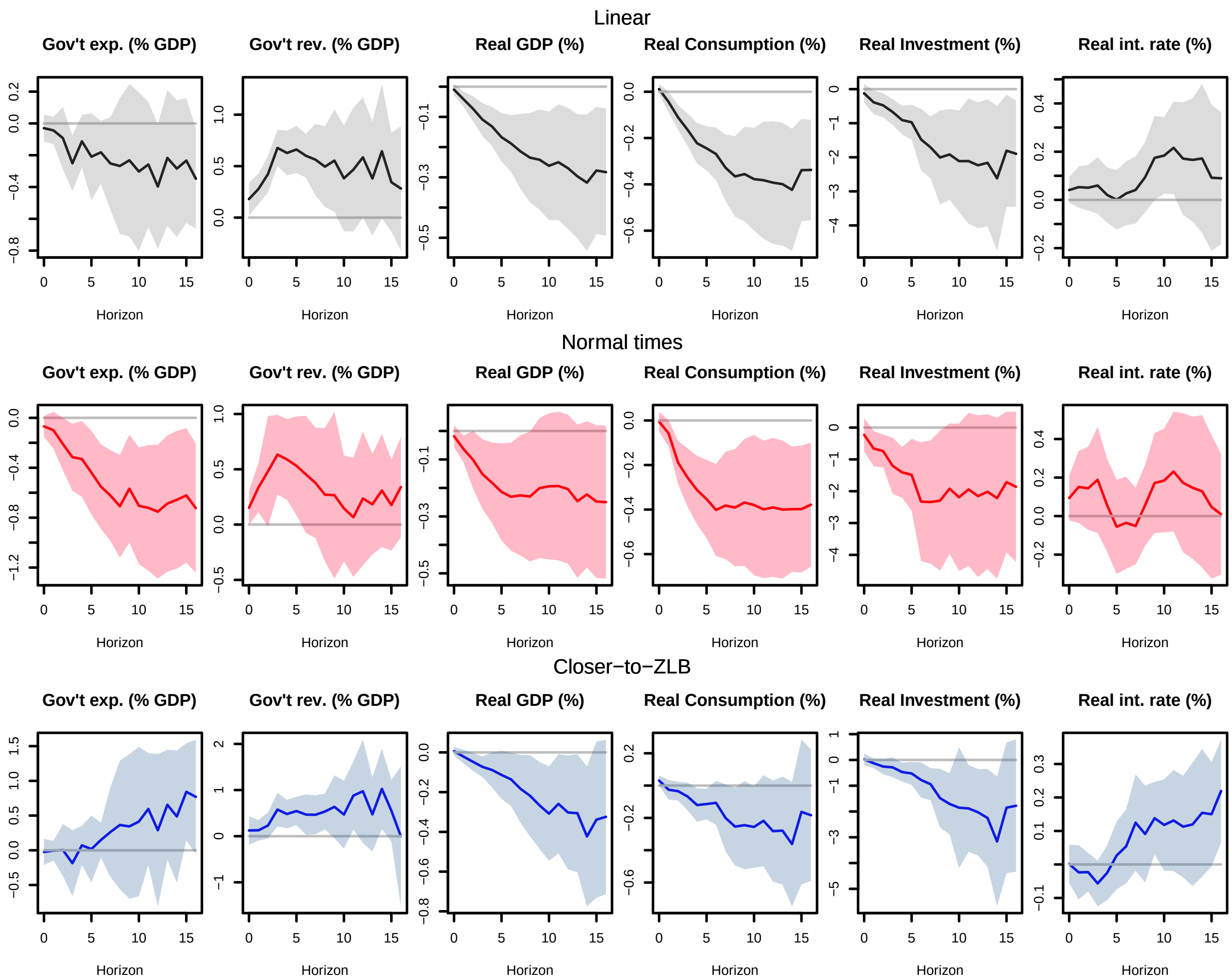
Considering that states of *close-to-ZLB* are highly correlated with the global financial crisis in my sample, I am interested in investigating further the effects of EB consolidations under periods of constrained monetary policy using the Eurozone currency union. Specifically, the state-dummy signaling periods of constrained monetary policy is constructed such that it takes values equal to 1 when a country is in the Eurozone and quarters are past 1999:Q1 (more details are provided in Sec. 4.3). I find that the point estimates for fiscal multipliers are

[41] Note that the drop in government spending following an EB consolidation announcement is similar across states ($\approx -0.5\%$ on average across horizons).

[42] Results of a positive multiplier in *normal times* further corroborates theory of *expansionary austerity* as discussed in Subsec. 4.5.1.

[43] The only statistically significant difference is observed in the consumption multipliers: around two years after announcement, the multiplier is not statistically different from zero under *normal times*, while it is negative and significant (≈ -0.3) near the ZLB (see Tables 4.C5).

Figure 4.4: IRFs: TB consolidation announcement - *Normal Times/Close-to-ZLB*

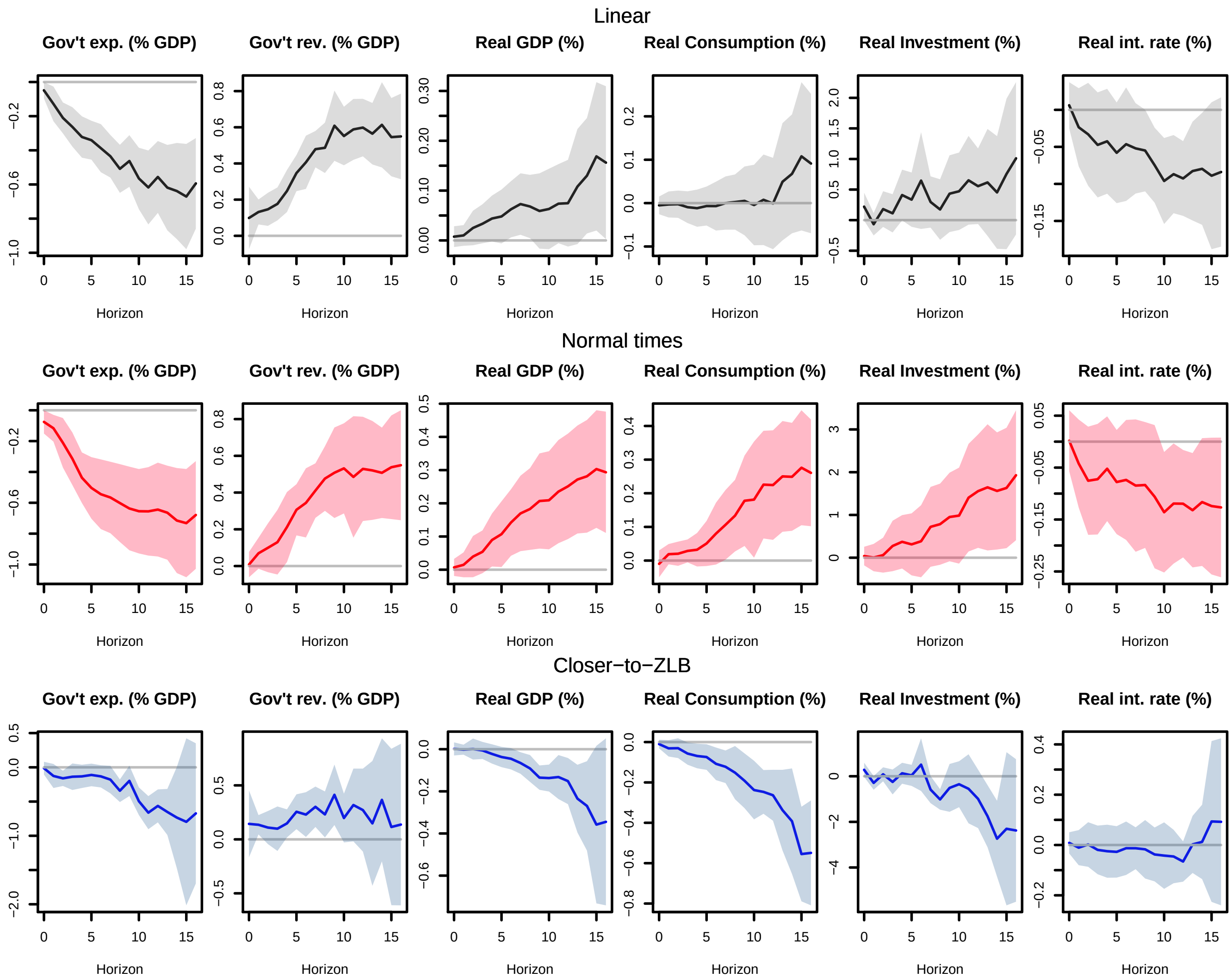


This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

negative when an EB consolidation takes place in a currency union (see Figure 4.D5 and Tables 4.E7- 4.E9 in Appendices 4.D and 4.E).[44] These findings substantiate the above-discussed results under constrained monetary policy (*close-to-ZLB*). These empirical findings parallel the theoretical conclusions of Erceg and Lindé (2012, 2013), which show that an EB consolidation may be counterproductive if monetary policy is constrained by a currency union since it de-

[44]For further details on this robustness check, see Appendices 4.D and 4.E.

Figure 4.5: IRFs: EB consolidation announcement - *Normal Times/Close-to-ZLB*



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

presses output given that the central bank cannot provide enough accommodation. Moreover, fixed exchange rates make spending cuts more contractionary than under unconstrained monetary policy, while causing tax hikes to be somewhat less contractionary by preventing the real appreciation that would occur when monetary policy is independent. Thus, the effects of TB consolidations are less sensitive to the degree of monetary accommodation. In light of my findings, this indeed implies that the contractionary effects of a TB consolidation do not necessarily need to be different between constrained and unconstrained monetary policy regimes. When also the ZLB is binding, Erceg and Lindé (2013) show that the duration of the liquidity trap is lengthened by spending cuts, which yields a stronger adverse impact on output at the margin. Also, by flipping the argument used in Christiano et al. (2011) for a rise in government spending, near the ZLB, an EB consolidation should lead to a fall in expected inflation, which translates in a rise in the real interest rate (as observed in the median response of long-term real rate under ZLB in my findings), and, then, to a large drop in private spending and output. Thus, the decrease of government consumption exacerbates the deflationary spiral associated with the zero-bound state.

In Subsec. 4.5.4, I explore other mechanisms driving the results near the ZLB by looking at the response of debt-to-GDP and confidence indicators.

Table 4.5: Output mult. under Normal times and Close-to-ZLB following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					Expenditure-Based Fiscal Consolidation				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-0.15***	-0.19**	-0.24	0.950	0.948	0.08	0.11	-0.03	0.423	0.496
	(0.06)	(0.08)	(0.23)			(0.11)	(0.17)	(0.07)		
2	-0.17***	-0.21*	-0.27***	0.842	0.846	0.12	0.17	-0.01	0.434	0.470
	(0.06)	(0.12)	(0.05)			(0.10)	(0.18)	(0.12)		
3	-0.16***	-0.24**	-0.15***	0.384	0.358	0.13	0.16	-0.05	0.178	0.261
	(0.04)	(0.12)	(0.04)			(0.09)	(0.12)	(0.11)		
4	-0.21***	-0.31**	-0.23**	0.437	0.451	0.14*	0.19**	-0.16	0.020	0.115
	(0.06)	(0.13)	(0.10)			(0.08)	(0.09)	(0.13)		
5	-0.25***	-0.41*	-0.26***	0.442	0.431	0.14	0.20**	-0.25*	0.019	0.097
	(0.06)	(0.21)	(0.10)			(0.09)	(0.10)	(0.15)		
6	-0.31***	-0.52	-0.33***	0.527	0.444	0.16*	0.25***	-0.24***	0.004	0.068
	(0.08)	(0.34)	(0.09)			(0.08)	(0.09)	(0.09)		
7	-0.38**	-0.62	-0.50***	0.712	0.671	0.17**	0.29***	-0.26*	0.004	0.078
	(0.15)	(0.50)	(0.16)			(0.07)	(0.10)	(0.14)		
8	-0.47**	-0.88	-0.48***	0.646	0.486	0.13**	0.29***	-0.21***	0.001	0.021
	(0.22)	(1.04)	(0.16)			(0.06)	(0.10)	(0.08)		
9	-0.44**	-0.79	-0.49*	0.729	0.635	0.13	0.32***	-0.41***	0.001	0.047
	(0.24)	(1.07)	(0.28)			(0.08)	(0.12)	(0.15)		
10	-0.69	-1.37	-0.78	0.766	0.573	0.11	0.31***	-0.21***	0.000	0.039
	(0.51)	(2.63)	(0.56)			(0.07)	(0.11)	(0.04)		
11	-0.54	-2.53	-0.29**	0.801	0.273	0.12	0.35***	-0.16**	0.000	0.011
	(0.37)	(8.79)	(0.14)			(0.07)	(0.12)	(0.07)		
12	-0.46*	-0.82	-0.32**	0.590	0.384	0.13	0.38***	-0.19**	0.000	0.015
	(0.26)	(1.06)	(0.16)			(0.09)	(0.12)	(0.08)		
13	-0.78	-1.24	-0.65	0.648	0.451	0.17	0.40***	-0.26***	0.000	0.016
	(0.62)	(1.63)	(0.52)			(0.11)	(0.12)	(0.09)		
14	-0.49**	-0.68	-0.38*	0.672	0.629	0.20**	0.38***	-0.25*	0.000	0.031
	(0.21)	(0.64)	(0.22)			(0.10)	(0.10)	(0.15)		
15	-0.81	-1.24	-0.49	0.674	0.585	0.25*	0.41***	-0.29	0.009	0.080
	(0.63)	(1.50)	(0.41)			(0.13)	(0.10)	(0.25)		
16	-1.00	-0.71	-4.65	0.933	0.484	0.26*	0.42***	-0.30	0.012	0.097
	(1.24)	(0.65)	(45.58)			(0.15)	(0.11)	(0.25)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.6: Output mult. (*pb*) under Normal times and Close-to-ZLB following a Fiscal Consolidation ann.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-0.13***	-0.14**	-0.19	0.833	0.799	0.04	0.07	-0.02	0.466	0.505
	(0.04)	(0.06)	(0.20)			(0.05)	(0.10)	(0.05)		
2	-0.14**	-0.14*	-0.23	0.773	0.676	0.07	0.11	-0.01	0.439	0.473
	(0.06)	(0.08)	(0.30)			(0.06)	(0.12)	(0.10)		
3	-0.12***	-0.16**	-0.10***	0.454	0.394	0.08	0.11	-0.05	0.204	0.266
	(0.03)	(0.08)	(0.01)			(0.05)	(0.09)	(0.11)		
4	-0.18***	-0.20**	-0.24***	0.670	0.662	0.08*	0.13*	-0.13	0.103	0.098
	(0.04)	(0.08)	(0.09)			(0.04)	(0.07)	(0.14)		
5	-0.19***	-0.22**	-0.24*	0.894	0.890	0.07	0.13**	-0.15**	0.011	0.054
	(0.05)	(0.10)	(0.13)			(0.04)	(0.06)	(0.08)		
6	-0.24***	-0.23*	-0.38***	0.270	0.264	0.08*	0.16***	-0.19**	0.002	0.031
	(0.06)	(0.12)	(0.10)			(0.04)	(0.06)	(0.08)		
7	-0.26***	-0.22*	-0.69	0.553	0.223	0.08**	0.17***	-0.19**	0.001	0.035
	(0.08)	(0.13)	(0.78)			(0.03)	(0.05)	(0.08)		
8	-0.31**	-0.23	-0.73	0.646	0.252	0.07**	0.17***	-0.20***	0.000	0.018
	(0.12)	(0.14)	(1.10)			(0.03)	(0.05)	(0.08)		
9	-0.31**	-0.23	-0.58	0.613	0.367	0.06	0.18***	-0.26***	0.000	0.030
	(0.12)	(0.19)	(0.72)			(0.04)	(0.05)	(0.03)		
10	-0.38**	-0.22	-1.03	0.618	0.128	0.06	0.17***	-0.23***	0.000	0.048
	(0.16)	(0.18)	(1.56)			(0.04)	(0.05)	(0.07)		
11	-0.35***	-0.23	-0.36	0.488	0.363	0.06*	0.20***	-0.15**	0.000	0.012
	(0.13)	(0.20)	(0.24)			(0.04)	(0.05)	(0.07)		
12	-0.27***	-0.20	-0.26*	0.556	0.521	0.06*	0.21***	-0.20**	0.000	0.013
	(0.10)	(0.16)	(0.14)			(0.04)	(0.05)	(0.09)		
13	-0.50**	-0.27	-0.80	0.277	0.034	0.09*	0.22***	-0.30***	0.000	0.012
	(0.21)	(0.19)	(0.57)			(0.05)	(0.05)	(0.09)		
14	-0.34***	-0.22	-0.30	0.475	0.340	0.10**	0.22***	-0.22*	0.002	0.025
	(0.13)	(0.15)	(0.19)			(0.05)	(0.05)	(0.13)		
15	-0.48	-0.29	-0.66	0.274	0.054	0.14**	0.23***	-0.31	0.041	0.083
	(0.29)	(0.21)	(0.48)			(0.07)	(0.05)	(0.29)		
16	-0.45*	-0.23	356.25	0.943	0.035	0.14*	0.23***	-0.32	0.066	0.094
	(0.26)	(0.15)	(3e+05)			(0.07)	(0.05)	(0.31)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

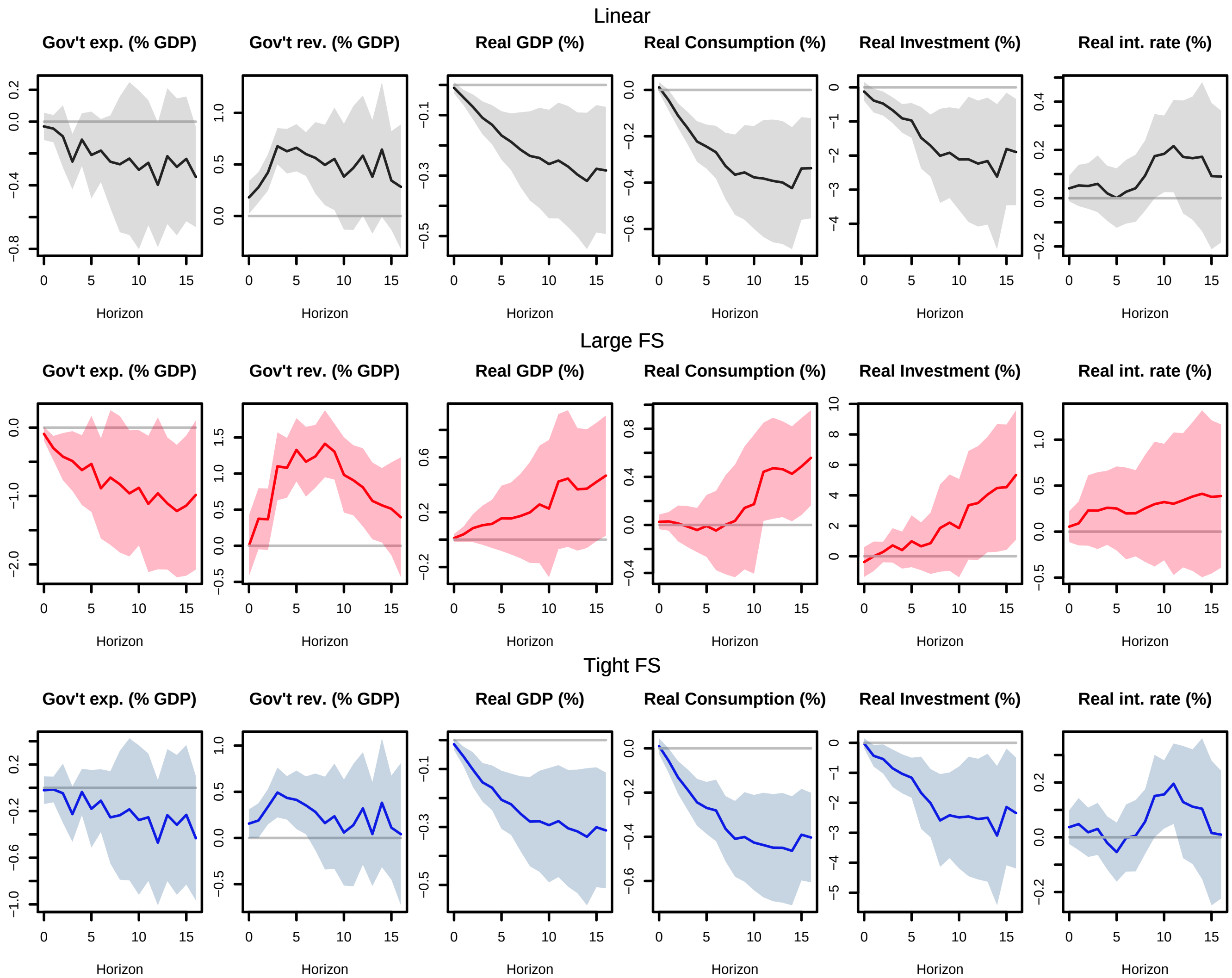
4.5.3 Large/Tight fiscal space

In Panel D of Table 4.2, I summarize the main findings under states of *large* and *tight fiscal space*: TB consolidations are contractionary (non-recessionary) and EB ones are non-recessionary (expansionary) when the economy starts in a weak (strong) fiscal position. Figs. 4.6 and 4.7 report IRFs across fiscal position states using an indicator for fiscal space ($FS1$) as described in Sec. 4.3. The C.I. for the IRFs under state of *large fiscal space* (strong fiscal position) are shaded in red, while they are shaded in blue under *tight fiscal space* (weak fiscal position). Both following a TB and an EB consolidation, my findings highlight differences across states of fiscal space. In particular, Fig. 4.6 shows that real GDP, consumption and investment plunge following a TB announcement when the fiscal position is weak, while real interest rate rises; on the contrary, output and the private sector do not contract when the fiscal position is strong. Following an EB consolidation announcement, under *large fiscal space*, I observe an expansionary effect on output and the private sector whose response is triggered by a large decrease in the real interest rate ($\approx -0.3\%$ on average across projection horizons). In contrast, under *tight fiscal space*, responses of macroeconomic variables are non-recessionary as observed in previous subsections.

This asymmetry across states of fiscal space can be observed also in the estimates for the fiscal multipliers in Tables 4.7 and 4.8. The left-hand side panels of the tables feature a statistically significant difference across output multipliers in the two states around a TB announcement: the fiscal multiplier is large and negative in *tight fiscal space*, while it is close to zero when the fiscal position is strong. The p-values testing the difference in the multiplier across states confirm such divergence also in the estimates of consumption and investment multipliers (especially for the *primary balance*-based ones), which are reported in Tables 4.C9- 4.C12 in Appendix 4.C (see the last two columns of the left-hand side panels of the tables). The dissimilarity between output multipliers following an EB consolidation is only minor and confirmed around 3 years after the shock (see p-values in the last two columns in the right-hand panel of Table 4.7). However, the investment multiplier is large and positive under *large fiscal space*, while, in the opposite state, it is close to zero. Such contrast is corroborated by results of testing the difference in the multipliers across states after year 2 (see p-values in the last two columns in the right-hand panels of Tables 4.C11 and 4.C12 in Appendix 4.C).[45] This confirms also that private investment

[45]Note also that consumption multipliers exhibit statistically significant differences across states around one year after the announcement, being positive under *large fiscal space* and mildly negative in median under *tight fiscal space* (see right-hand panels of Tables 4.C9 and 4.C10).

Figure 4.6: IRFs: TB consolidation announcement - *Large/Tight fiscal space*

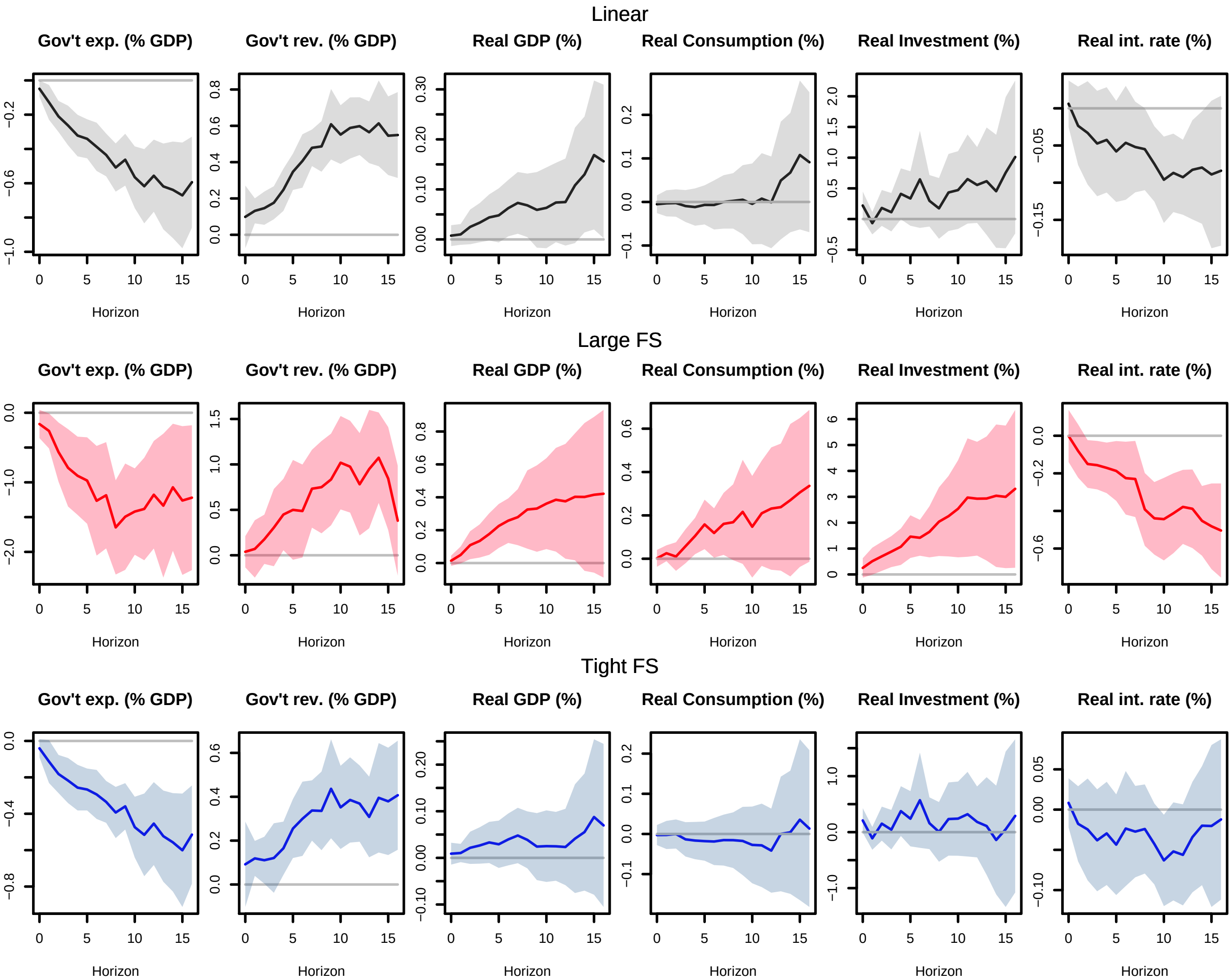


This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

is particularly sensitive to fiscal adjustments as observed in previous studies (e.g., Alesina et al., 2015).[46]

[46]Note that, following a TB announcement, the *primary balance*-based investment multiplier is as low as -5 in *tight fiscal space* (see left panel of Table 4.C12 in Appendix 4.C).

Figure 4.7: IRFs: EB consolidation announcement - *Large/Tight fiscal space*



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Table 4.7: Output multiplier under Large and Tight Fiscal Space following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.15***	0.10	-0.29*	0.019	0.031	0.08	0.20	0.09	0.590	0.605
	(0.06)	(0.08)	(0.16)			(0.11)	(0.14)	(0.12)		
2	-0.17***	0.22*	-0.30**	0.002	0.064	0.12	0.20*	0.11	0.646	0.646
	(0.06)	(0.13)	(0.12)			(0.10)	(0.11)	(0.11)		
3	-0.16***	0.10	-0.30***	0.000	0.016	0.13	0.17***	0.11	0.731	0.732
	(0.04)	(0.08)	(0.08)			(0.09)	(0.06)	(0.11)		
4	-0.21***	0.11	-0.39***	0.000	0.012	0.14*	0.20***	0.12	0.602	0.576
	(0.06)	(0.09)	(0.12)			(0.08)	(0.07)	(0.09)		
5	-0.25***	0.12	-0.51***	0.002	0.008	0.14	0.24***	0.10	0.362	0.320
	(0.06)	(0.11)	(0.19)			(0.09)	(0.07)	(0.10)		
6	-0.31***	0.13	-0.63**	0.004	0.026	0.16*	0.21***	0.12	0.585	0.569
	(0.08)	(0.13)	(0.30)			(0.08)	(0.06)	(0.10)		
7	-0.38**	0.14	-0.88	0.100	0.013	0.17**	0.24***	0.12	0.483	0.426
	(0.15)	(0.13)	(0.69)			(0.07)	(0.08)	(0.08)		
8	-0.47**	0.14	-1.50	0.406	0.012	0.13**	0.20***	0.08	0.358	0.342
	(0.22)	(0.15)	(2.09)			(0.06)	(0.06)	(0.07)		
9	-0.44*	0.20	-1.07	0.241	0.016	0.13	0.22***	0.05	0.249	0.238
	(0.24)	(0.17)	(1.21)			(0.08)	(0.06)	(0.09)		
10	-0.69	0.23	-3.09	0.727	0.022	0.11	0.25***	0.04	0.100	0.096
	(0.51)	(0.29)	(9.69)			(0.07)	(0.08)	(0.07)		
11	-0.54	0.46	-1.41	0.324	0.045	0.12*	0.27***	0.04	0.064	0.080
	(0.37)	(0.33)	(2.28)			(0.07)	(0.08)	(0.06)		
12	-0.46*	0.55	-0.79	0.000	0.103	0.13	0.30***	0.04	0.116	0.121
	(0.26)	(0.38)	(0.62)			(0.09)	(0.10)	(0.07)		
13	-0.78	0.60	-2.60	0.545	0.073	0.17	0.29***	0.06	0.136	0.170
	(0.62)	(0.49)	(5.97)			(0.11)	(0.08)	(0.10)		
14	-0.49**	0.64	-0.74	0.032	0.126	0.20**	0.35***	0.07	0.185	0.199
	(0.21)	(0.67)	(0.46)			(0.10)	(0.12)	(0.10)		
15	-0.81	0.77	-1.82	0.244	0.038	0.25*	0.30***	0.10	0.392	0.404
	(0.63)	(0.96)	(2.95)			(0.13)	(0.09)	(0.12)		
16	-1.00	1.04	-3.71	0.785	0.297	0.26*	0.31***	0.09	0.390	0.406
	(1.24)	(1.75)	(18.05)			(0.15)	(0.10)	(0.14)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.8: Output mult. (*pb*) under Large and Tight Fiscal Space following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.13***	0.05	-0.26	0.070	0.015	0.04	0.14*	0.05	0.298	0.211
	(0.04)	(0.03)	(0.17)			(0.05)	(0.08)	(0.06)		
2	-0.14**	0.10**	-0.25	0.052	0.026	0.07	0.14**	0.08	0.450	0.483
	(0.06)	(0.04)	(0.17)			(0.06)	(0.06)	(0.08)		
3	-0.12***	0.06	-0.20***	0.000	0.009	0.08	0.12***	0.09	0.599	0.614
	(0.03)	(0.04)	(0.05)			(0.05)	(0.03)	(0.08)		
4	-0.18***	0.06	-0.35***	0.000	0.006	0.08*	0.13***	0.09	0.472	0.454
	(0.04)	(0.05)	(0.09)			(0.04)	(0.04)	(0.07)		
5	-0.19***	0.08	-0.35***	0.001	0.006	0.07	0.15***	0.06	0.163	0.136
	(0.05)	(0.06)	(0.13)			(0.04)	(0.04)	(0.06)		
6	-0.24***	0.07	-0.47***	0.000	0.007	0.08**	0.14***	0.07	0.214	0.226
	(0.06)	(0.05)	(0.14)			(0.04)	(0.03)	(0.06)		
7	-0.26***	0.08	-0.46***	0.003	0.012	0.08**	0.14***	0.08	0.242	0.209
	(0.08)	(0.05)	(0.18)			(0.03)	(0.04)	(0.05)		
8	-0.31**	0.08	-0.65*	0.047	0.011	0.07**	0.13***	0.06	0.179	0.163
	(0.12)	(0.07)	(0.39)			(0.03)	(0.04)	(0.05)		
9	-0.31**	0.10	-0.62	0.055	0.015	0.06	0.14 ***	0.03	0.136	0.127
	(0.12)	(0.07)	(0.39)			(0.04)	(0.05)	(0.05)		
10	-0.38**	0.10	-0.79	0.112	0.023	0.06	0.14**	0.03	0.131	0.107
	(0.16)	(0.11)	(0.56)			(0.04)	(0.06)	(0.05)		
11	-0.35***	0.17**	-0.58***	0.001	0.020	0.06*	0.15**	0.03	0.089	0.078
	(0.13)	(0.09)	(0.22)			(0.04)	(0.06)	(0.04)		
12	-0.27***	0.20**	-0.34***	0.001	0.075	0.06*	0.18*	0.03	0.153	0.102
	(0.10)	(0.08)	(0.12)			(0.04)	(0.09)	(0.05)		
13	-0.50**	0.17*	-0.83*	0.075	0.021	0.09*	0.16**	0.05	0.211	0.194
	(0.21)	(0.09)	(0.50)			(0.05)	(0.07)	(0.08)		
14	-0.34***	0.16**	-0.40***	0.006	0.030	0.10**	0.17**	0.05	0.307	0.285
	(0.13)	(0.08)	(0.15)			(0.05)	(0.09)	(0.07)		
15	-0.48	0.19**	-0.60*	0.035	0.014	0.14**	0.18**	0.08	0.437	0.425
	(0.29)	(0.09)	(0.33)			(0.07)	(0.09)	(0.09)		
16	-0.45*	0.23**	-0.50	0.057	0.014	0.14*	0.23*	0.07	0.305	0.272
	(0.26)	(0.10)	(0.36)			(0.07)	(0.12)	(0.10)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq$ 0.01***,$\leq$ 0.05**,$\leq$ 0.1*

The strong contraction observed under *tight fiscal space* following a TB consolidation can be explained via the expectation channel: fiscal adjustments through a tax hike lead to a lower GDP growth and to a higher debt-to-GDP ratio that worsens the fiscal position. The latter effect is likely to enforce further (potentially distortionary) fiscal adjustments and, ultimately, it drives towards even weaker growth.[47] Hence, economic agents foresee a persistent fiscal adjustment bound to worsen economic growth for several years without ameliorating the fiscal position.[48] When (and given that) the fiscal position is strong, economic agents expect a short-lived consolidation that does not dampen the economic outlook. This way, it is possible to explain also the differential effects of EB consolidations across states of fiscal space.[49] This mechanism is also reflected in the dynamics of the long-term real interest rate (see Figs. 4.6 and 4.7). In Subsec. 4.5.4, I investigate such mechanism by looking at the impact on debt-to-GDP and confidence indicators, which can proxy well expectations of private consumers and investors.

4.5.4 Effect on Debt-to-GDP, Confidence indicators and Income inequality

In Table 4.9, I summarize the main findings on the impact of fiscal consolidation announcements on debt-to-GDP, confidence indicators and income inequality across different initial states of the economy. These results are shown in Figs. 4.8 and 4.9, which report the responses of debt-to-GDP (%), consumer confidence (Δ), business confidence indicator (Δ) and the Gini index for (gross) income inequality (%) following a TB and an EB consolidation announcement, respectively. Starting from the top of the graphs, I first report the IRFs under the linear case; second, under *expansion* and *recession*; third, under *normal times* and *close-to-ZLB*; last, under *tight* and *large fiscal space*. The C.I. for the responses in the *good* state (i.e., *normal times*) are shaded in red, while they are shaded in blue in the *bad* state (i.e., *close-to-ZLB*). The IRFs in the linear case are shaded in grey. The projection horizon extends over 4 years (measured in quarters). The IRFs are obtained via the estimation of Eq. 4.4 in Sec. 4.4 by extending the set of endogenous vari-

[47] Under a neoclassical framework, Alesina et al. (2020) find that TB consolidations lead to a negative demand effect due to expected future taxes.

[48] As shown in Alesina et al. (2015), the more persistent is the fiscal adjustment the larger the detrimental effect on the economic outlook.

[49] The results reported here for a cut in government spending are symmetric to the findings in Metelli and Pallara (2020). The latter study reports that the output multiplier is lower under *tight fiscal space* following a positive government spending shock.

Table 4.9: Results on debt-to-GDP, confidence and income inequality: summary

State	Type	IRF Figures	Effect
Linear	TB[a]	4.8	Debt-to-GDP, Income inequality ↑; Confidence ↓
	EB[b]	4.9	Debt-to-GDP stabilized; Business confidence, inequality ≈ 0; Consumer confidence ↑ in last projection year
Expansion	TB	4.8	Similar to *Linear*; Business confidence ↑
	EB	4.9	Similar to *Linear*
Recession	TB	4.8	Similar to *Linear*
	EB	4.9	Similar to *Linear*; Debt-to-GDP↑
Normal times	TB	4.8	Similar to *Linear*; Business confidence ↑; Income inequality ↑
	EB	4.9	Debt-to-GDP stabilized; Consumer confidence ↑
Close-to-ZLB	TB	4.8	Similar to *Linear*; Business confidence ↓; Income inequality ≈ 0
	EB	4.9	Debt-to-GDP ↑; Consumer confidence ↓
Large FS[c]	TB	4.8	Debt-to-GDP ↓; Consumer confidence ↑ in last projection year
	EB	4.9	Debt-to-GDP stabilized; Confidence ↑
Tight FS	TB	4.8	Debt-to-GDP ↑; Income inequality ↑; Confidence ↓
	EB	4.9	Confidence, Income inequality ≈ 0; Mild rise in Debt-to-GDP

a: Tax-Based consolidation announcement; *b*: Expenditure-Based consolidation announcement; *c*: FS stands for fiscal space.

This table outlines the results on the effects of both TB and EB fiscal consolidation announcements on debt-to-GDP, confidence and income inequality across different initial states of the economy. Specifically, Sec. 4.5.4 describes in details these findings and IRFs are reported in Figs. 4.8, 4.9.

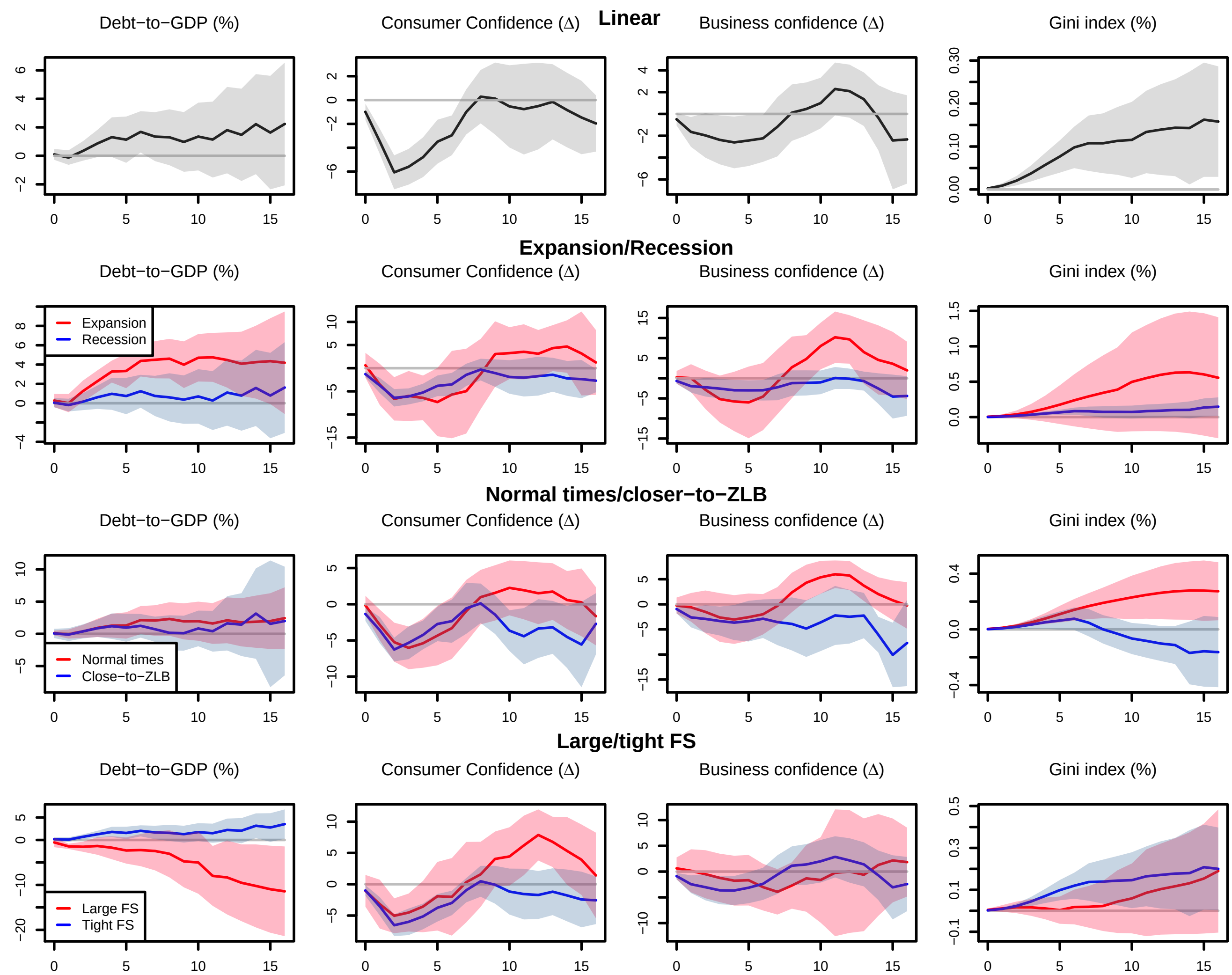
ables. In turn, I include in the specification debt-to-GDP, confidence indicators and the income inequality Gini indices (for gross and net income).[50]

Debt-to-GDP

Fig. 4.8 reports that debt-to-GDP rises following a TB consolidation (significant response is close to 2% after 6 quarters), which is consistent with the findings of Beetsma et al. (2021). This response also confirms that TB consolidations are often *self-defeating* (Attinasi and Metelli, 2017).

[50] By augmenting the set of endogenous variables and, thus, of controls, I am also able to perform additional robustness checks on the baseline results.

Figure 4.8: TB consolidation announcement

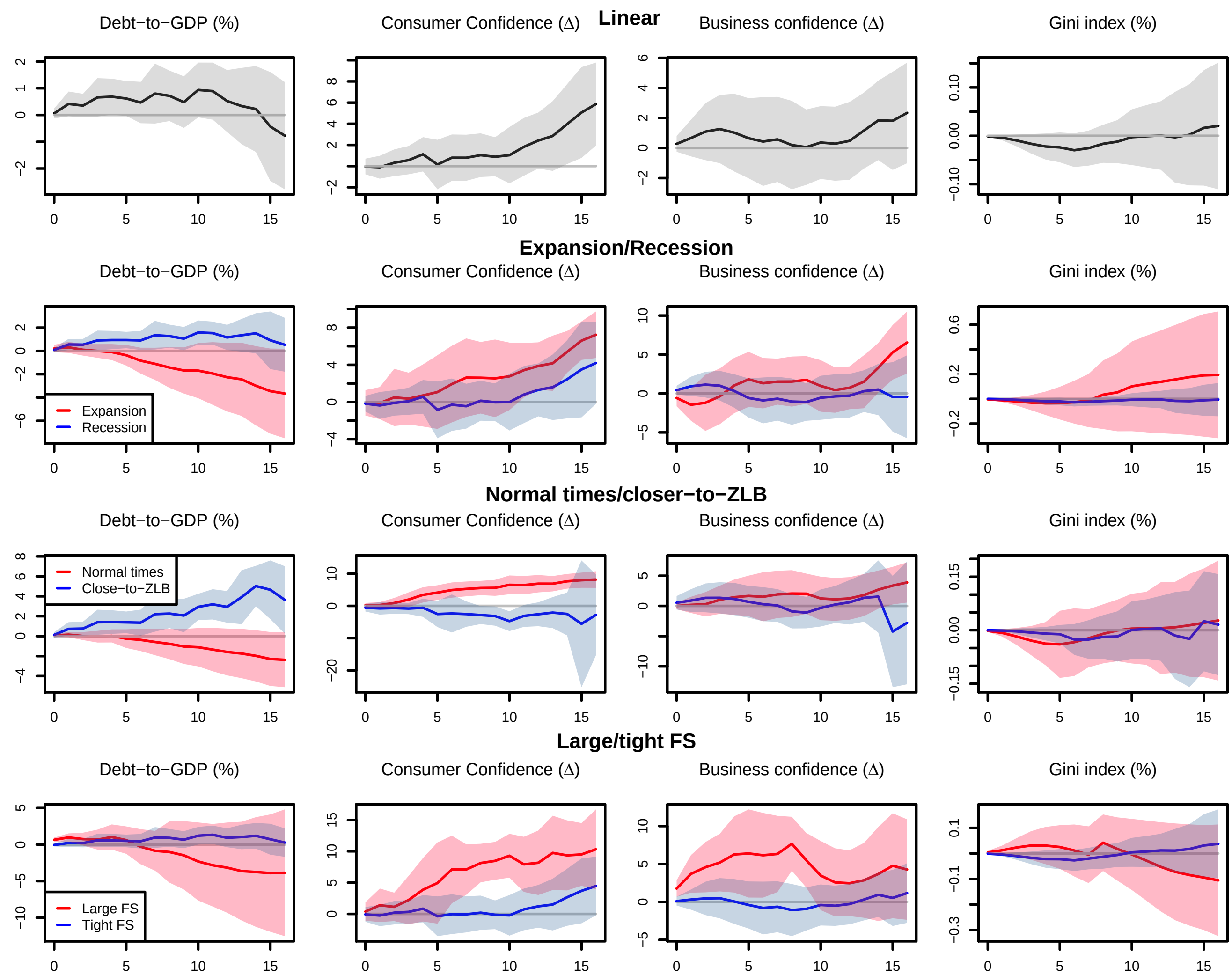


This figure shows the responses of debt-to-GDP (%), consumer confidence (Δ), business confidence indicator (Δ) and the Gini index for (gross) income inequality (%) following a TB consolidation announcement. Starting from the top of the graphs, I first report the IRFs under the linear case (C.I. shaded in grey), second, under *expansion* and *recession*, third, under *normal times* and *close-to-ZLB* and, last, under *tight* and *large fiscal space*. The C.I. of the responses in the *good* state (i.e., *normal times*) are shaded in red, while they are shaded in blue in the *bad* state (i.e., *close-to-ZLB*). The projection horizon extends over 4 years measured in quarters. The IRFs are obtained via the estimation of Eq. 4.4 in Sec. 4.4 by extending the set of endogenous variables in turn.

Conversely, EB consolidations seem to stabilize the stock of debt given that the cumulative response of debt-to-GDP is null and output is not dampened (see linear case in Fig. 4.9). Thus, spending cuts seem to be more effective at achieving fiscal sustainability without jeopardizing economic growth (Fotiou, 2020).[51] I do not observe significant overall differences in the re-

[51] Under a neoclassical framework, Alesina et al. (2020) find that debt stabilization is more likely under EB consolidations given that TB ones do not necessarily imply lower government expenditure and crowd out private spending due to expected future taxes.

Figure 4.9: EB consolidation announcement



This figure shows the responses of debt-to-GDP (%), consumer confidence (Δ), business confidence indicator (Δ) and the Gini index for (gross) income inequality (%) following an EB consolidation announcement. Starting from the top of the graphs, I first report the IRFs under the linear case (C.I. shaded in grey), second, under *expansion* and *recession*, third, under *normal times* and *close-to-ZLB* and, last, under *tight* and *large fiscal space*. The C.I. of the responses in the *good* state (i.e., *normal times*) are shaded in red, while they are shaded in blue in the *bad* state (i.e., *close-to-ZLB*). The projection horizon extends over 4 years measured in quarters. The IRFs are obtained via the estimation of Eq. 4.4 in Sec. 4.4 by extending the set of endogenous variables in turn.

sponse of debt-to-GDP across states of the economy. Nonetheless, when the economy is near the ZLB, debt-to-GDP shows a sizeable increase (≈ 3% on average across horizons) around EB announcements of consolidation measures. This highlights that, under certain states of the economy, spending-based consolidations are *self-defeating*. This finding parallels the recessionary effects of an EB consolidation observed near the ZLB (see Fig. 4.5). Following a TB consolidation announcement, debt-to-GDP plunges when the fiscal position is strong (*large fiscal* *space*), while it grows under *tight fiscal space*. The latter finding corroborates the mechanism described in Subsec. 4.5.3: in a weak fiscal position, revenue-based consolidations dampen economic growth to the point to be strongly *self-defeating*. On the contrary, revenue-based adjustments are neither recessionary nor *self-defeating* when the fiscal position is strong.

Confidence indicators

Consumer confidence falls around TB announcements of consolidation measures in line with Beetsma et al. (2015) (Fig. 4.8). Following an EB consolidation, I observe a mild increase in consumer confidence at the end of the projection horizon (Fig. 4.9). Contemporaneously, economic activity positively responds to spending cuts. EB consolidations cause no significant response in the business confidence indicator. TB announcements of consolidation measures lead to a decrease in the business confidence indicator for about 2 years. Thus, announcements of consolidations significantly deteriorate confidence only when the fiscal adjustment is revenue-based. Nonetheless, consumer confidence is boosted both around TB and EB announcements of consolidation measures when the economy is in a strong fiscal position (*large fiscal space*). This substantiates the mechanism described in Subsec. 4.5.3 for which economic agents foresee that consolidations in *large fiscal space* are likely to be short-lived and non-recessionary. Moreover, a positive response of the business confidence indicator is observed around an EB announcement when the economy starts in a strong fiscal position. Near the ZLB, consumer confidence is hindered following a spending-based consolidation. On the contrary, in *normal times*, consumer confidence is boosted. This supports the observed expansionary effects under unconstrained monetary policy (see Fig. 4.5). Business confidence is undermined around revenue-based announcements of consolidation measures when the economy is close to the ZLB.

Income inequality

Gini index for gross income inequality steadily increases following a TB consolidation announcement (Fig. 4.8). The cumulative response peaks at 0.15% at the end of the projection horizon. The observed rise in income inequality can be rationalized by the contraction in economic activity, which inevitably leads to an increment in unemployment (Furceri et al., 2015). The effect on inequality is stronger when interest rates are free to be adjusted (*normal times*) and when fiscal space is tight. When the fiscal position is weak, the revenue-based consolidation is expected to be very persistent and distortionary, which translates in a long-lasting recession.

Conversely, I find no significant effects on the Gini index for gross income inequality following an EB consolidation, even across different states of the economy (Fig. 4.9).[52]

4.6 Conclusions

This paper investigates the state-dependent effects of fiscal consolidations using local projection methods. Using fiscal consolidation announcements, I find that both the *when*, namely the initial state of the economy, and the *how*, the type of fiscal adjustment, matter for the transmission of austerity plans on the economy. I also compute and compare cumulative multipliers for output, consumption and investment across states of the economy. Consistently with previous studies, I find that TB consolidations are more contractionary than EB ones. The latter also prove to be expansionary at times, which is consistent with neoclassical theory.

The main findings of the paper highlight that both the strength of the fiscal position and the constrainedness of monetary policy are crucial for the transmission of fiscal consolidation shocks. Indeed, near the ZLB, an EB announcement of consolidation measures is contractionary differently from times when monetary policy is unconstrained. Moreover, both EB and TB announcement of consolidation measures yield different effects on the economy depending on the state of fiscal space: TB consolidations are contractionary (non-recessionary) and EB ones are non-recessionary (expansionary) when the economy starts in a weak (strong) fiscal position. Indeed, economic agents expect long-lasting (short-lived) fiscal adjustments when fiscal space is tight (large).

I also find that a TB announcement of consolidation measures is *self-defeating*, while an EB one is not unless the economy is near the ZLB. Thus, spending-based fiscal adjustments are more effective at stabilizing debt-to-GDP without jeopardizing economic growth. Additionally, TB consolidations considerably hinder consumer confidence unless the economy is in a strong fiscal position. Lastly, income distribution becomes more unequal following a revenue-based fiscal adjustment due to the contraction in economic activity.

[52]Similar results are found for the Gini index for net income inequality and reported in Fig. 4.B4 in Appendix 4.B.

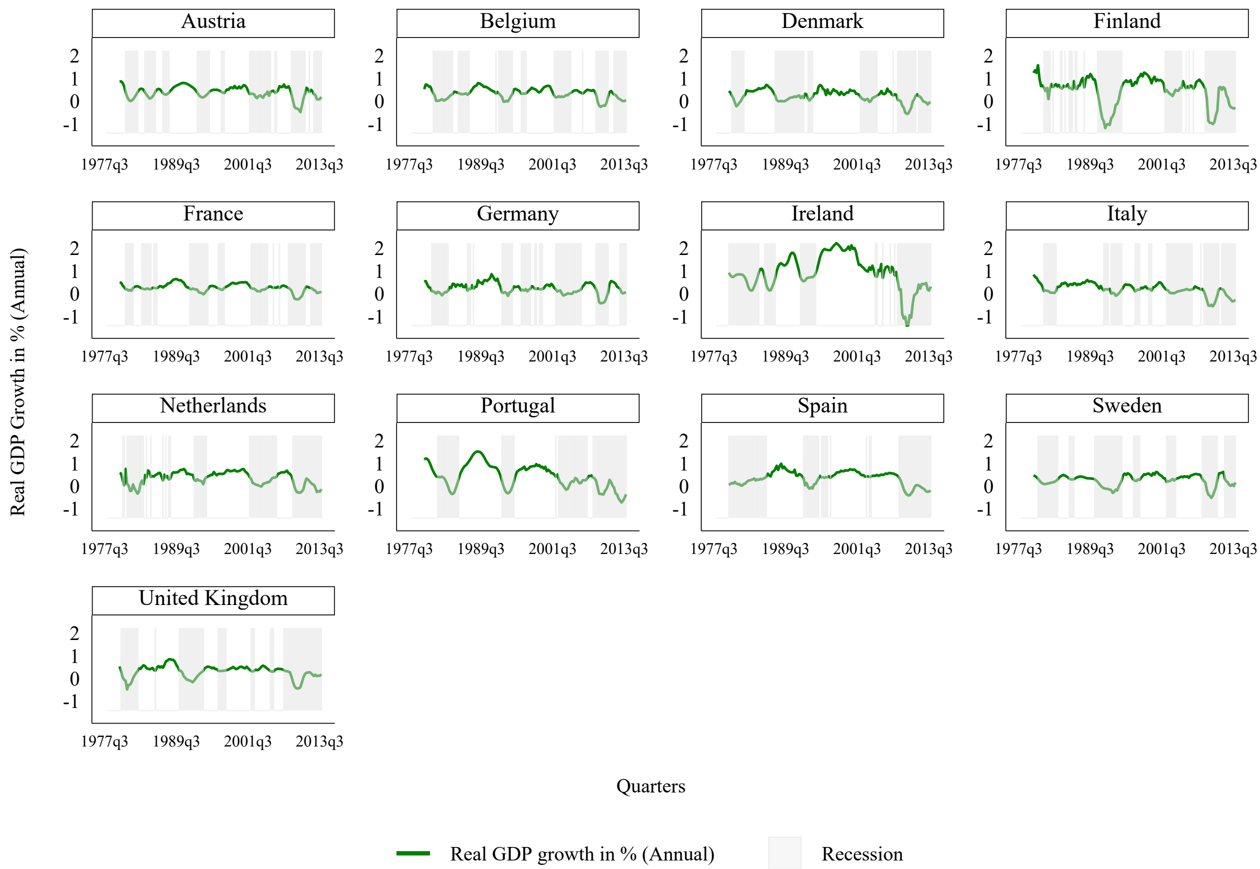
Appendix

4.A Data

Most of the macroeconomic variables are collected from the OECD Economic Outlook database. This includes nominal GDP, real private consumption, consumer price index, GDP deflator, both the consumer and the business confidence indicators. Private investment is taken from the IMF International Financial Statistics database. Moreover, the short-term and the long-term interest rates are also drawn from the OECD Economic Outlook dataset. The former is based on three-month money market rates where available, while the latter refers to government bonds maturing in ten years. Both rates are generally averages of daily rates, measured as a percentage. I use the series of government spending and government revenues as constructed in Beetsma et al. (2021). These series are originally drawn from Eurostat and the OECD Economic Outlook. The former includes government final consumption expenditure, government fixed capital formation and social security benefits paid by the government. The latter comprehends proceeds from total direct taxes, indirect taxes and social security contributions received by the government. Public debt series are taken from Eurostat and the OECD Economic Outlook. Note that all variables are seasonally adjusted and converted into euros. The measures for income inequality, namely the Gini indices for both gross and net income (ex-post redistribution of taxes and transfers), are gathered from the Standardized World Income Inequality Database (SWIID) at the annual frequency and interpolated at the quarterly frequency using cubic spline. As regards the construction of the dummies for the state of the economy, the data for the OECD Recession dates are based on the OECD Composite Leading Indicator (CLI), drawn at the monthly frequency from the Federal Reserve Bank of Saint Louis and aggregated at the quarterly frequency. The series of the CLI is based on the growth cycle approach, where business cycles and turning points are identified through a deviation from the trend method. OECD recession dates take value 1 if the economy is under recession, 0 otherwise. A specific quarter is defined under recession ($= 1$) if for at least two months the monthly OECD Recession dates indicator is equal to 1. Moreover, unemployment rates are seasonally adjusted and drawn from the OECD Economic Outlook. Only the unemployment rate for Germany is taken from the International Labour Organization (ILO) database. In reference to the fiscal space indicators, potential GDP is drawn from the OECD Economic Outlook. Moreover, following the method implemented by the World Bank and reported in Kose et al. (2017), cyclical adjustment of the government finance statistics variables is obtained by multiplying them by $(1+\tilde{y})^{-(\epsilon_x-1)}$ where $\tilde{y}$ is the difference between the actual GDP and the potential output as % of potential output; ϵ_x stands for the output gap elasticity of x for $x =$[revenues, spending, debt]. We use World Bank (see also Kose et al., 2017) elasticities for revenues and government spending, respectively equal to 1 and 0.1. For what concerns public debt, the estimated elasticity is not significantly different from 0 and, thus, assumed to be equal to 0. Note that the elasticities for spending and revenues proposed here are not distant from the estimates in Girouard and André (2006).

4.B Additional Figures

Figure 4.B1: Real GDP growth and periods of Expansion/Recession.

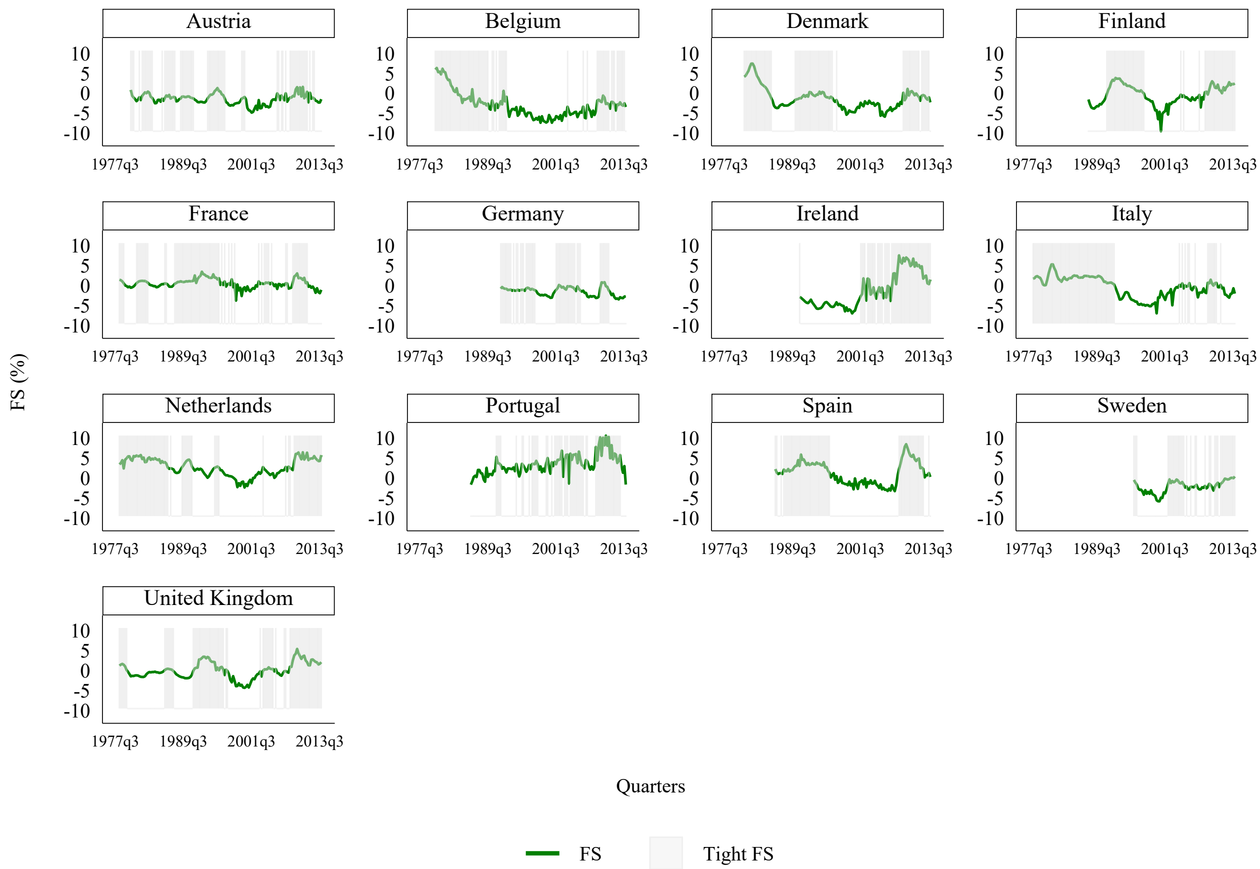


In this figure, real GDP growth in % (Annual) is plotted together with shaded areas indicating periods of recession. Shaded areas represent periods in which the state dummy used for *Expansion/Recession* is equal to 1, namely whenever real GDP growth is below its median.

Figure 4.B2: Short-term interest rate and periods close-to-ZLB.

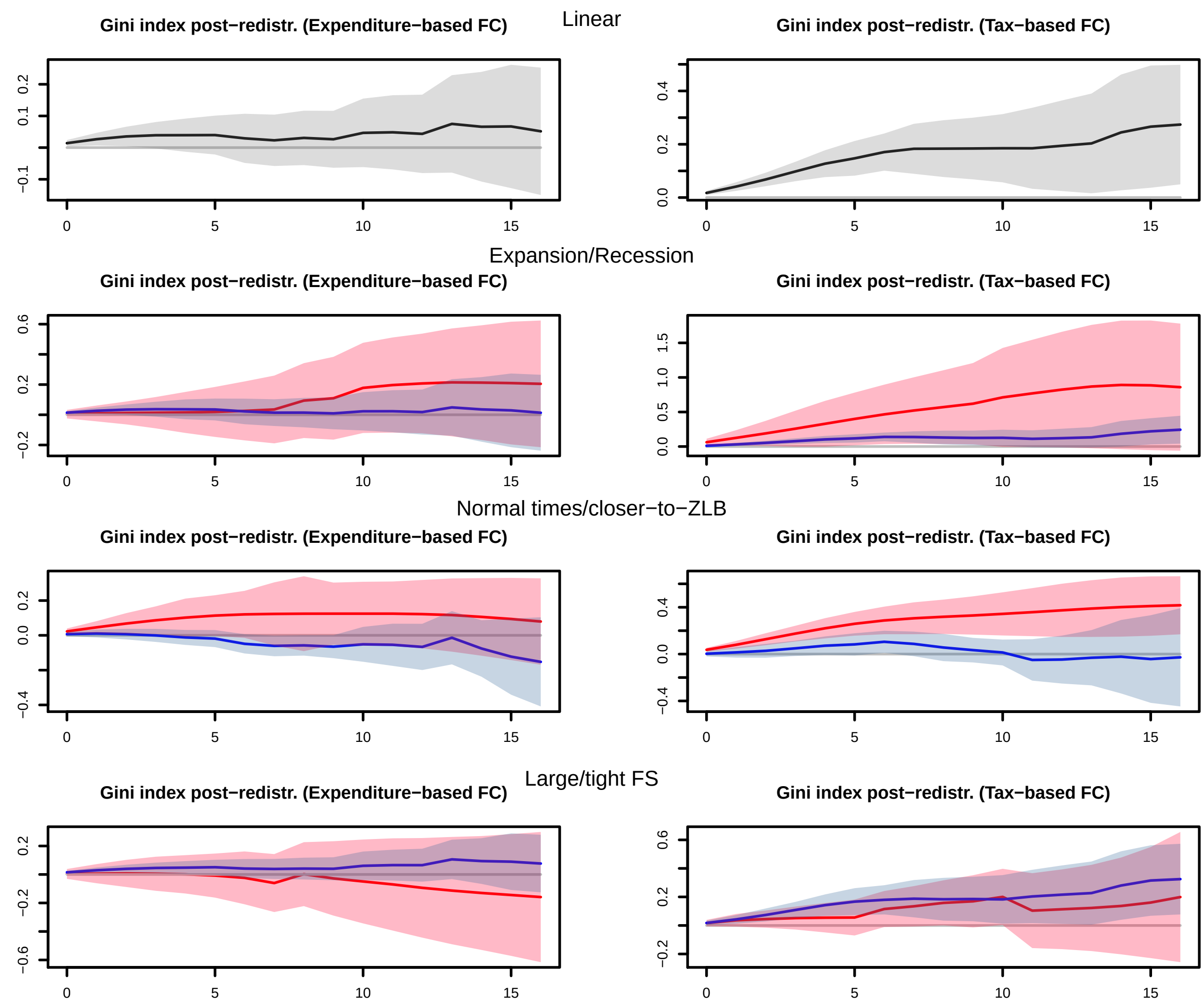
In this figure, the short-term nominal interest rate in basis points (bps) is plotted together with shaded areas indicating periods close-to-ZLB. Shaded areas represent periods in which the state dummy used for *Normal times/Close-to-ZLB* is equal to 1, namely whenever the short-term rate is below 75 bps.

Figure 4.B3: $FS1$ and periods of tight fiscal space.



In this figure, the indicator of fiscal space ($FS1$) in % is plotted together with shaded areas indicating periods of tight fiscal space. Shaded areas represent periods in which the state dummy used for *Large/Tight fiscal space* is equal to 1, namely whenever $FS1$ is above its median.

Figure 4.B4: IRFs of Gini index (net income inequality): TB and EB consolidation announcement.



This figure shows the responses of the Gini index for net income inequality (%) following a TB and an EB consolidation announcement. Starting from the top of the graphs, I first report the IRFs under the linear case (C.I. shaded in grey); second, under *expansion* and *recession*; third, under *normal times* and *close-to-ZLB*; last, under *tight* and *large fiscal space*. The C.I. of the responses in the *good* state (i.e., *normal times*) are shaded in red, while they are shaded in blue in the *bad* state (i.e., *close-to-ZLB*). The projection horizon extends over 4 years measured in quarters. The IRFs are obtained via the estimation of Eq. 4.4 in Sec. 4.4 by extending the set of endogenous variables with the inclusion of the Gini indices for income inequality.

4.C Results: Tables

In this section, tables for consumption and investment multipliers are reported, including those whose ratio is taken over the cumulative response of primary balance (*pb*). These estimates enrich and complement the results discussed in Sec. 4.5. In Tables 4.C1- 4.C4, I illustrate estimates for consumption and investment multipliers across *expansion* and *recession*. In Tables 4.C5- 4.C8, I report multipliers for consumption and investment across *normal times* and *close-to-ZLB*. In Tables 4.C9- 4.C12, estimates for consumption and investment multipliers across *large* and *tight fiscal space* are presented.

Table 4.C1: Consumption mult. under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.16	-0.30	-0.09	0.553	0.451	-0.02	0.53	-0.06	0.364	0.215
	(0.10)	(0.36)	(0.09)			(0.12)	(0.65)	(0.11)		
2	-0.26***	-0.37**	-0.22**	0.407	0.410	-0.01	0.22	-0.03	0.741	0.680
	(0.10)	(0.17)	(0.11)			(0.08)	(0.67)	(0.08)		
3	-0.24***	-0.41*	-0.20***	0.279	0.323	-0.04	0.16	-0.07	0.376	0.394
	(0.06)	(0.22)	(0.06)			(0.08)	(0.22)	(0.07)		
4	-0.35***	-0.52**	-0.32***	0.286	0.359	-0.04	0.13	-0.06	0.429	0.453
	(0.09)	(0.24)	(0.09)			(0.07)	(0.21)	(0.07)		
5	-0.37***	-0.54*	-0.35***	0.431	0.452	-0.02	0.15	-0.07	0.349	0.394
	(0.07)	(0.29)	(0.07)			(0.07)	(0.21)	(0.07)		
6	-0.45***	-0.99***	-0.41***	0.131	0.201	-0.02	0.04	-0.06	0.538	0.568
	(0.13)	(0.36)	(0.15)			(0.08)	(0.13)	(0.08)		
7	-0.58**	-0.73	-0.61*	0.784	0.786	0.00	0.14	-0.07	0.226	0.283
	(0.25)	(0.50)	(0.34)			(0.08)	(0.15)	(0.08)		
8	-0.74**	-0.75	-0.83*	0.952	0.952	0.00	0.09	-0.05	0.239	0.302
	(0.33)	(0.58)	(0.48)			(0.07)	(0.09)	(0.07)		
9	-0.64*	-0.91	-0.69	0.796	0.781	0.01	0.17*	-0.08	0.053	0.105
	(0.36)	(0.89)	(0.46)			(0.09)	(0.10)	(0.09)		
10	-0.99	-0.98	-1.17	0.963	0.964	-0.01	0.14*	-0.09	0.019	0.070
	(0.77)	(1.39)	(1.12)			(0.09)	(0.08)	(0.07)		
11	-0.82	-6.37	-0.76	0.918	0.546	0.01	0.17**	-0.07	0.010	0.054
	(0.57)	(49.73)	(0.48)			(0.09)	(0.07)	(0.07)		
12	-0.67*	-0.79	-0.68*	0.920	0.917	-0.00	0.18**	-0.10	0.006	0.043
	(0.37)	(1.21)	(0.35)			(0.10)	(0.08)	(0.07)		
13	-1.05	-0.56	-1.21	0.549	0.499	0.08	0.19***	-0.06	0.014	0.037
	(0.86)	(0.90)	(0.98)			(0.13)	(0.06)	(0.09)		
14	-0.66**	-0.34	-0.65***	0.528	0.634	0.11	0.19***	-0.05	0.044	0.062
	(0.30)	(0.54)	(0.24)			(0.12)	(0.07)	(0.09)		
15	-0.99	-0.46	-0.79**	0.479	0.502	0.16	0.17***	0.02	0.302	0.362
	(0.78)	(0.75)	(0.37)			(0.15)	(0.06)	(0.12)		
16	-1.19	-0.19	-1.33	0.509	0.137	0.15	0.19***	-0.02	0.229	0.292
	(1.53)	(0.44)	(1.66)			(0.15)	(0.06)	(0.13)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.C2: Consumption mult. (*pb*) under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.14	-0.36	-0.08	0.515	0.358	-0.01	0.13**	-0.04	0.102	0.186
	(0.09)	(0.46)	(0.08)			(0.06)	(0.06)	(0.07)		
2	-0.21**	-0.32*	-0.18*	0.419	0.408	-0.01	0.04	-0.02	0.602	0.606
	(0.10)	(0.19)	(0.11)			(0.05)	(0.09)	(0.06)		
3	-0.18***	-0.37**	-0.14***	0.106	0.216	-0.02	0.05	-0.05	0.217	0.271
	(0.04)	(0.17)	(0.03)			(0.05)	(0.06)	(0.05)		
4	-0.30***	-0.45**	-0.27***	0.178	0.246	-0.02	0.04	-0.05	0.278	0.320
	(0.06)	(0.19)	(0.07)			(0.04)	(0.07)	(0.05)		
5	-0.28***	-0.46*	-0.26***	0.330	0.325	-0.01	0.05	-0.04	0.237	0.291
	(0.06)	(0.27)	(0.06)			(0.04)	(0.07)	(0.05)		
6	-0.34***	-0.65*	-0.31***	0.326	0.242	-0.01	0.02	-0.04	0.417	0.456
	(0.07)	(0.39)	(0.09)			(0.04)	(0.05)	(0.05)		
7	-0.40***	-0.46	-0.42***	0.756	0.742	0.00	0.06	-0.04	0.177	0.236
	(0.11)	(0.37)	(0.15)			(0.04)	(0.06)	(0.05)		
8	-0.48***	-0.39	-0.55**	0.913	0.915	0.00	0.05	-0.03	0.226	0.292
	(0.15)	(0.33)	(0.22)			(0.04)	(0.05)	(0.05)		
9	-0.45***	-0.45	-0.48***	0.843	0.828	0.00	0.10*	-0.05	0.043	0.107
	(0.13)	(0.51)	(0.18)			(0.04)	(0.06)	(0.05)		
10	-0.55***	-0.32	-0.65**	0.778	0.797	-0.00	0.10*	-0.06	0.019	0.069
	(0.18)	(0.37)	(0.27)			(0.04)	(0.05)	(0.05)		
11	-0.53***	-0.34	-0.51***	0.989	0.989	0.01	0.12**	-0.05	0.008	0.055
	(0.17)	(0.47)	(0.13)			(0.05)	(0.05)	(0.05)		
12	-0.40***	-0.20	-0.41***	0.750	0.769	-0.00	0.12**	-0.07	0.006	0.044
	(0.12)	(0.28)	(0.10)			(0.05)	(0.05)	(0.05)		
13	-0.67**	-0.18	-0.72***	0.323	0.399	0.04	0.13***	-0.04	0.012	0.035
	(0.27)	(0.27)	(0.26)			(0.06)	(0.04)	(0.07)		
14	-0.46***	-0.14	-0.42***	0.340	0.460	0.05	0.13***	-0.03	0.037	0.057
	(0.14)	(0.20)	(0.10)			(0.06)	(0.04)	(0.06)		
15	-0.59*	-0.14	-0.51**	0.211	0.221	0.09	0.12***	0.01	0.298	0.316
	(0.32)	(0.21)	(0.20)			(0.08)	(0.04)	(0.09)		
16	-0.54*	-0.08	-0.56	0.268	0.122	0.08	0.14***	-0.01	0.190	0.190
	(0.31)	(0.17)	(0.34)			(0.08)	(0.04)	(0.09)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C3: Investment mult. under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-1.39*	-0.93	-1.08	0.965	0.964	-0.53	6.36	-1.08	0.373	0.323
	(0.83)	(1.69)	(0.75)			(0.76)	(8.23)	(0.90)		
2	-1.13**	-2.43**	-0.78	0.252	0.338	0.86	5.27	0.52	0.531	0.140
	(0.53)	(0.98)	(0.66)			(0.78)	(7.74)	(0.76)		
3	-0.99***	-2.73***	-0.70*	0.073	0.145	0.42	1.75	0.14	0.297	0.253
	(0.33)	(1.01)	(0.39)			(0.68)	(1.55)	(0.67)		
4	-1.45***	-2.64**	-1.26**	0.205	0.287	1.27	2.25	1.05	0.479	0.437
	(0.48)	(1.11)	(0.50)			(0.87)	(2.09)	(0.96)		
5	-1.47***	-2.08*	-1.40***	0.553	0.557	0.98	1.23	0.77	0.663	0.644
	(0.39)	(1.24)	(0.44)			(0.81)	(1.72)	(0.86)		
6	-2.46***	-4.52*	-2.33*	0.433	0.425	1.67	0.63	1.47	0.785	0.787
	(0.88)	(2.36)	(1.21)			(1.40)	(1.07)	(1.57)		
7	-3.04**	-3.31	-3.38*	0.946	0.946	0.68	1.34	0.21	0.401	0.370
	(1.39)	(2.40)	(1.99)			(0.48)	(1.35)	(0.50)		
8	-4.05**	-2.85	-5.01	0.640	0.651	0.34	0.97	-0.09	0.279	0.310
	(2.08)	(2.74)	(3.07)			(0.52)	(0.96)	(0.41)		
9	-3.46*	-3.53	-4.05	0.966	0.967	0.94	1.01	0.43	0.637	0.639
	(1.93)	(4.42)	(2.67)			(0.75)	(1.01)	(0.80)		
10	-5.54	-5.56	-6.91	0.948	0.949	0.84	0.75	0.41	0.655	0.670
	(4.18)	(7.75)	(6.46)			(0.54)	(0.89)	(0.50)		
11	-4.53*	-23.33	-4.68*	0.925	0.683	1.05*	1.17	0.50	0.454	0.460
	(2.74)	(184.26)	(2.73)			(0.60)	(0.88)	(0.58)		
12	-3.82**	-2.99	-4.25**	0.831	0.848	1.00*	1.67*	0.15	0.116	0.167
	(1.90)	(5.21)	(2.11)			(0.55)	(0.87)	(0.48)		
13	-5.69	-1.72	-7.53	0.417	0.367	1.00	1.36*	0.03	0.090	0.177
	(4.54)	(3.37)	(6.21)			(0.73)	(0.79)	(0.38)		
14	-4.06***	-1.30	-4.54***	0.159	0.448	0.71	2.11**	-0.54	0.013	0.086
	(1.36)	(2.62)	(1.11)			(0.73)	(0.93)	(0.55)		
15	-5.26	-0.02	-5.57*	0.152	0.056	1.13	1.93**	-0.10	0.042	0.091
	(3.69)	(3.49)	(2.95)			(0.95)	(0.83)	(0.65)		
16	-6.71	-0.27	-9.28	0.516	0.146	1.70	2.16***	0.22	0.043	0.063
	(9.08)	(2.15)	(12.99)			(1.05)	(0.80)	(0.64)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C4: Investment mult. (*pb*) under Expansion and Recession following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-1.20*	-1.11	-0.92	0.860	0.855	-0.26	1.53	-0.61	0.189	0.259
	(0.69)	(2.43)	(0.64)			(0.41)	(1.31)	(0.46)		
2	-0.92*	-2.08	-0.64	0.401	0.410	0.51	0.91*	0.38	0.198	0.253
	(0.52)	(1.49)	(0.63)			(0.44)	(0.47)	(0.52)		
3	-0.72***	-2.44**	-0.51*	0.127	0.144	0.25	0.52	0.11	0.421	0.435
	(0.22)	(1.16)	(0.29)			(0.39)	(0.36)	(0.50)		
4	-1.23***	-2.26*	-1.06***	0.290	0.291	0.72	0.77	0.78	0.862	0.863
	(0.30)	(1.23)	(0.37)			(0.46)	(0.62)	(0.67)		
5	-1.12***	-1.77	-1.04***	0.523	0.482	0.49	0.42	0.50	0.988	0.988
	(0.24)	(1.39)	(0.27)			(0.37)	(0.54)	(0.50)		
6	-1.89***	-2.97	-1.79**	0.525	0.467	0.81	0.26	0.96	0.525	0.532
	(0.55)	(2.16)	(0.77)			(0.62)	(0.42)	(0.93)		
7	-2.10***	-2.08	-2.32***	0.891	0.888	0.32	0.59	0.13	0.392	0.412
	(0.56)	(1.74)	(0.84)			(0.23)	(0.54)	(0.30)		
8	-2.63***	-1.49	-3.31***	0.607	0.666	0.17	0.55	-0.06	0.272	0.309
	(0.78)	(1.43)	(1.12)			(0.27)	(0.56)	(0.27)		
9	-2.44***	-1.76	-2.86***	0.971	0.972	0.40	0.61	0.24	0.586	0.596
	(0.68)	(2.36)	(0.94)			(0.32)	(0.63)	(0.44)		
10	-3.09***	-1.80	-3.85***	0.745	0.773	0.42	0.51	0.27	0.657	0.664
	(0.87)	(1.95)	(1.23)			(0.28)	(0.64)	(0.32)		
11	-2.91***	-1.25	-3.13***	0.809	0.835	0.54*	0.80	0.33	0.441	0.451
	(0.74)	(1.72)	(0.68)			(0.29)	(0.66)	(0.37)		
12	-2.28***	-0.77	-2.53***	0.519	0.581	0.48*	1.11*	0.10	0.127	0.174
	(0.58)	(1.13)	(0.59)			(0.25)	(0.66)	(0.31)		
13	-3.62**	-0.57	-4.44**	0.278	0.293	0.52	0.93	0.02	0.109	0.187
	(1.65)	(0.98)	(2.09)			(0.37)	(0.59)	(0.26)		
14	-2.82***	-0.54	-2.96***	0.179	0.323	0.36	1.42*	-0.35	0.035	0.085
	(0.83)	(1.08)	(0.79)			(0.38)	(0.74)	(0.35)		
15	-3.13*	-0.00	-3.57*	0.176	0.127	0.62	1.40**	-0.07	0.085	0.117
	(1.71)	(1.08)	(2.11)			(0.53)	(0.71)	(0.46)		
16	-3.01	-0.10	-3.88	0.339	0.147	0.88	1.61**	0.15	0.065	0.096
	(2.08)	(0.84)	(3.23)			(0.55)	(0.73)	(0.43)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C5: Consumption mult. under Normal times and Close-to-ZLB following a Fiscal Consolidation ann.

	Tax-Based Fiscal Consolidation					Expenditure-Based Fiscal Consolidation				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-0.16	-0.17	-0.26	0.880	0.867	-0.02	0.15	-0.24***	0.038	0.089
	(0.10)	(0.12)	(0.44)			(0.12)	(0.14)	(0.09)		
2	-0.26***	-0.39**	-0.18***	0.060	0.008	-0.01	0.09	-0.18	0.083	0.182
	(0.10)	(0.16)	(0.05)			(0.08)	(0.10)	(0.11)		
3	-0.24***	-0.41**	-0.13*	0.082	0.069	-0.04	0.09	-0.37*	0.031	0.078
	(0.06)	(0.17)	(0.08)			(0.08)	(0.06)	(0.22)		
4	-0.35***	-0.53***	-0.30**	0.107	0.166	-0.04	0.08	-0.41*	0.033	0.100
	(0.09)	(0.18)	(0.13)			(0.07)	(0.05)	(0.23)		
5	-0.37***	-0.66***	-0.25***	0.043	0.042	-0.02	0.10*	-0.46	0.053	0.109
	(0.07)	(0.25)	(0.07)			(0.07)	(0.06)	(0.29)		
6	-0.45***	-0.89*	-0.26**	0.141	0.055	-0.02	0.15**	-0.51**	0.011	0.072
	(0.13)	(0.48)	(0.11)			(0.08)	(0.07)	(0.23)		
7	-0.58**	-1.03	-0.53*	0.381	0.238	0.00	0.19**	-0.45*	0.026	0.069
	(0.25)	(0.76)	(0.29)			(0.08)	(0.08)	(0.26)		
8	-0.74**	-1.47	-0.55***	0.483	0.032	0.00	0.23***	-0.33***	0.000	0.049
	(0.33)	(1.61)	(0.17)			(0.07)	(0.07)	(0.09)		
9	-0.64*	-1.43	-0.44*	0.509	0.090	0.01	0.29***	-0.56**	0.005	0.059
	(0.36)	(1.76)	(0.25)			(0.09)	(0.08)	(0.26)		
10	-0.99	-2.66	-0.65	0.642	0.070	-0.01	0.29***	-0.35***	0.000	0.039
	(0.77)	(4.89)	(0.51)			(0.09)	(0.10)	(0.09)		
11	-0.82	-5.17	-0.25**	0.785	0.060	0.01	0.35***	-0.28***	0.000	0.023
	(0.57)	(17.83)	(0.11)			(0.09)	(0.11)	(0.08)		
12	-0.67*	-1.56	-0.30**	0.462	0.089	-0.00	0.35***	-0.33***	0.000	0.026
	(0.37)	(1.86)	(0.14)			(0.10)	(0.10)	(0.11)		
13	-1.05	-2.00	-0.62	0.532	0.142	0.08	0.38***	-0.36***	0.000	0.022
	(0.86)	(2.60)	(0.56)			(0.13)	(0.10)	(0.13)		
14	-0.66**	-1.21	-0.34*	0.435	0.246	0.11	0.35***	-0.35*	0.003	0.033
	(0.30)	(1.10)	(0.18)			(0.12)	(0.08)	(0.21)		
15	-0.99	-1.99	-0.24	0.472	0.232	0.16	0.38***	-0.44	0.011	0.069
	(0.78)	(2.33)	(0.34)			(0.15)	(0.09)	(0.30)		
16	-1.19	-1.08	-2.69	0.943	0.781	0.15	0.38***	-0.47	0.009	0.080
	(1.53)	(0.94)	(26.75)			(0.15)	(0.08)	(0.29)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.C6: Consumption mult. (*pb*) under Normal times and Close-to-ZLB following a Fiscal Cons. announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-0.14	-0.13	-0.21	0.810	0.769	-0.01	0.09	-0.15**	0.013	0.100
	(0.09)	(0.09)	(0.35)			(0.06)	(0.08)	(0.06)		
2	-0.21**	-0.27***	-0.16	0.631	0.771	-0.01	0.06	-0.15*	0.046	0.174
	(0.10)	(0.09)	(0.18)			(0.05)	(0.06)	(0.08)		
3	-0.18***	-0.27***	-0.09***	0.012	1.000	-0.02	0.06	-0.36**	0.009	0.083
	(0.04)	(0.09)	(0.03)			(0.05)	(0.04)	(0.17)		
4	-0.30***	-0.33***	-0.31***	0.934	0.935	-0.02	0.05	-0.35	0.085	0.096
	(0.06)	(0.09)	(0.12)			(0.04)	(0.04)	(0.23)		
5	-0.28***	-0.35***	-0.23**	0.506	0.590	-0.01	0.06*	-0.27***	0.004	0.096
	(0.06)	(0.10)	(0.10)			(0.04)	(0.04)	(0.10)		
6	-0.34***	-0.39***	-0.30**	0.947	0.948	-0.01	0.09**	-0.41**	0.011	0.071
	(0.07)	(0.15)	(0.13)			(0.04)	(0.04)	(0.17)		
7	-0.40***	-0.37**	-0.73	0.610	0.334	0.00	0.11***	-0.33**	0.003	0.056
	(0.11)	(0.17)	(0.79)			(0.04)	(0.04)	(0.13)		
8	-0.48***	-0.39**	-0.83	0.665	0.289	0.00	0.13***	-0.31***	0.000	0.052
	(0.15)	(0.18)	(1.14)			(0.04)	(0.03)	(0.08)		
9	-0.45***	-0.43*	-0.53	0.707	0.572	0.00	0.16***	-0.36***	0.000	0.045
	(0.13)	(0.24)	(0.56)			(0.04)	(0.03)	(0.07)		
10	-0.55***	-0.42*	-0.86	0.665	0.274	-0.00	0.16***	-0.38***	0.000	0.049
	(0.18)	(0.23)	(1.21)			(0.04)	(0.04)	(0.13)		
11	-0.53***	-0.47*	-0.31***	0.692	0.645	0.01	0.20***	-0.27***	0.000	0.027
	(0.17)	(0.28)	(0.08)			(0.05)	(0.04)	(0.09)		
12	-0.40***	-0.37*	-0.25***	0.873	0.871	-0.00	0.19***	-0.33**	0.000	0.029
	(0.12)	(0.21)	(0.08)			(0.05)	(0.04)	(0.14)		
13	-0.67**	-0.43*	-0.76*	0.266	0.070	0.04	0.21***	-0.41***	0.000	0.023
	(0.27)	(0.24)	(0.46)			(0.06)	(0.03)	(0.15)		
14	-0.46***	-0.39*	-0.27**	0.702	0.652	0.05	0.20***	-0.31*	0.008	0.035
	(0.14)	(0.20)	(0.13)			(0.06)	(0.03)	(0.18)		
15	-0.59*	-0.46*	-0.33	0.377	0.165	0.09	0.21***	-0.47	0.045	0.086
	(0.32)	(0.26)	(0.35)			(0.08)	(0.04)	(0.35)		
16	-0.54*	-0.34**	205.76	0.942	0.102	0.08	0.21***	-0.50	0.063	0.092
	(0.31)	(0.17)	(1.8e+05)			(0.08)	(0.03)	(0.38)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C7: Investment mult. under Normal times and Close-to-ZLB following a Fiscal Consolidation ann.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-1.39***	-1.92**	-1.18	0.145	0.521	-0.53	0.11	-2.29**	0.220	0.291
	(0.83)	(0.89)	(0.90)			(0.76)	(1.41)	(1.14)		
2	-1.13**	-1.53**	-1.27*	0.628	0.665	0.86	0.28	0.53	0.832	0.829
	(0.53)	(0.64)	(0.77)			(0.78)	(1.02)	(0.97)		
3	-0.99***	-1.91**	-0.52***	0.061	0.000	0.42	0.84	-1.64	0.089	0.248
	(0.33)	(0.80)	(0.12)			(0.68)	(0.90)	(1.18)		
4	-1.45***	-2.41***	-1.11***	0.109	0.157	1.27	0.79	0.76	0.931	0.930
	(0.48)	(0.93)	(0.34)			(0.87)	(0.62)	(1.80)		
5	-1.47***	-2.79**	-1.10***	0.184	0.199	0.98	0.57	0.23	0.933	0.933
	(0.39)	(1.36)	(0.32)			(0.81)	(0.64)	(1.60)		
6	-2.46***	-5.16	-1.75**	0.263	0.167	1.67	0.64	2.42	0.581	0.587
	(0.88)	(3.29)	(0.72)			(1.40)	(0.63)	(3.59)		
7	-3.04**	-6.34	-2.26***	0.351	0.198	0.68	1.22**	-2.13	0.052	0.041
	(1.39)	(4.90)	(0.64)			(0.48)	(0.62)	(1.44)		
8	-4.05**	-8.66	-2.95***	0.533	0.268	0.34	1.27**	-2.18***	0.000	0.042
	(2.08)	(10.34)	(1.12)			(0.52)	(0.62)	(0.33)		
9	-3.46*	-7.45	-2.86**	0.606	0.372	0.94	1.44**	-1.45	0.259	0.135
	(1.93)	(10.09)	(1.45)			(0.75)	(0.61)	(1.93)		
10	-5.54	-15.33	-4.38	0.651	0.141	0.84	1.44**	-0.54	0.061	0.094
	(4.18)	(27.61)	(2.81)			(0.54)	(0.70)	(0.76)		
11	-4.53*	-25.14	-2.01**	0.788	0.204	1.05*	2.03***	-0.66	0.034	0.080
	(2.74)	(84.81)	(0.83)			(0.60)	(0.72)	(0.92)		
12	-3.82**	-8.58	-2.03***	0.468	0.216	1.00*	2.31***	-1.25*	0.001	0.026
	(1.90)	(9.80)	(0.67)			(0.55)	(0.74)	(0.72)		
13	-5.69	-10.05	-4.60	0.574	0.268	1.00	2.39***	-1.86***	0.000	0.003
	(4.54)	(12.48)	(3.39)			(0.73)	(0.83)	(0.54)		
14	-4.06***	-6.72	-2.78**	0.549	0.480	0.71	2.11***	-2.44*	0.005	0.040
	(1.36)	(5.93)	(1.24)			(0.73)	(0.65)	(1.30)		
15	-5.26	-8.57	-2.62	0.564	0.440	1.13	2.14***	-1.83	0.097	0.100
	(3.69)	(9.30)	(2.34)			(0.95)	(0.67)	(2.01)		
16	-6.71	-5.31	-24.50	0.936	0.556	1.70	2.71***	-2.04	0.057	0.115
	(9.08)	(4.83)	(238.37)			(1.05)	(0.69)	(2.10)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C8: Investment mult. (*pb*) under Normal times and Close-to-ZLB following a Fiscal Cons. announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Normal**	**ZLB**	p-value	AR p-value	Linear	**Normal**	**ZLB**	p-value	AR p-value
1	-1.20*	-1.45**	-0.94	0.722	0.789	-0.26	0.06	-1.38*	0.351	0.399
	(0.69)	(0.67)	(0.99)			(0.41)	(0.83)	(0.75)		
2	-0.92*	-1.04**	-1.12	0.945	0.939	0.51	0.18	0.43	0.877	0.877
	(0.52)	(0.44)	(1.68)			(0.44)	(0.70)	(0.77)		
3	-0.72***	-1.25**	-0.36***	0.136	0.225	0.25	0.58	-1.63	0.195	0.245
	(0.22)	(0.51)	(0.13)			(0.39)	(0.68)	(1.32)		
4	-1.23***	-1.50***	-1.15***	0.751	0.762	0.72	0.53	0.65	0.924	0.925
	(0.30)	(0.56)	(0.40)			(0.46)	(0.45)	(1.56)		
5	-1.12***	-1.49**	-0.99*	0.714	0.739	0.49	0.36	0.13	0.704	0.693
	(0.24)	(0.72)	(0.60)			(0.37)	(0.43)	(0.94)		
6	-1.89***	-2.25*	-2.01**	0.916	0.915	0.81	0.40	1.98	0.616	0.620
	(0.55)	(1.15)	(1.02)			(0.62)	(0.42)	(2.98)		
7	-2.10***	-2.29**	-3.13	0.714	0.600	0.32	0.72*	-1.58*	0.022	0.042
	(0.56)	(1.18)	(3.60)			(0.23)	(0.43)	(0.84)		
8	-2.63***	-2.28*	-4.49	0.675	0.400	0.17	0.72*	-2.09***	0.000	0.042
	(0.78)	(1.35)	(5.97)			(0.27)	(0.39)	(0.47)		
9	-2.44***	-2.22	-3.43	0.685	0.553	0.40	0.82**	-0.94	0.068	0.098
	(0.68)	(1.60)	(3.86)			(0.32)	(0.39)	(1.06)		
10	-3.09***	-2.45	-5.79	0.661	0.316	0.42	0.80*	-0.59	0.098	0.097
	(0.87)	(1.56)	(8.62)			(0.28)	(0.41)	(0.88)		
11	-2.91***	-2.28	-2.45*	0.592	0.540	0.54*	1.16***	-0.64	0.048	0.075
	(0.74)	(1.73)	(1.37)			(0.29)	(0.43)	(0.92)		
12	-2.28***	-2.05	-1.67**	0.810	0.804	0.48*	1.26***	-1.27	0.007	0.017
	(0.58)	(1.43)	(0.74)			(0.25)	(0.45)	(0.85)		
13	-3.62**	-2.15	-5.66	0.253	0.024	0.52	1.32***	-2.13***	0.000	0.001
	(1.65)	(1.54)	(3.83)			(0.37)	(0.50)	(0.51)		
14	-2.82***	-2.14	-2.20*	0.582	0.514	0.36	1.22***	-2.17*	0.012	0.038
	(0.83)	(1.39)	(1.16)			(0.38)	(0.43)	(1.17)		
15	-3.13*	-1.99	-3.50	0.349	0.085	0.62	1.21***	-1.95	0.171	0.100
	(1.71)	(1.38)	(3.02)			(0.53)	(0.44)	(2.27)		
16	-3.01	-1.68	1876.63	0.942	0.036	0.88	1.49**	-2.19	0.150	0.102
	(2.08)	(1.11)	(1.6e+06)			(0.55)	(0.45)	(2.56)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C9: Consumption multiplier under Large and Tight FS following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.16	0.07	-0.28	0.119	0.088	-0.02	0.10	-0.02	0.516	0.605
	(0.10)	(0.09)	(0.22)			(0.12)	(0.07)	(0.15)		
2	-0.26***	0.04	-0.38**	0.053	0.203	-0.01	0.02	-0.00	0.888	0.888
	(0.10)	(0.16)	(0.16)			(0.08)	(0.06)	(0.10)		
3	-0.24***	-0.01	-0.38***	0.013	0.045	-0.04	0.08*	-0.06	0.337	0.356
	(0.06)	(0.07)	(0.13)			(0.08)	(0.04)	(0.10)		
4	-0.35***	-0.04	-0.57***	0.001	0.017	-0.04	0.12*	-0.06	0.191	0.147
	(0.09)	(0.08)	(0.18)			(0.07)	(0.05)	(0.09)		
5	-0.37***	-0.00	-0.66***	0.003	0.013	-0.02	0.17***	-0.06	0.099	0.061
	(0.07)	(0.09)	(0.22)			(0.07)	(0.05)	(0.09)		
6	-0.45***	-0.04	-0.80**	0.040	0.074	-0.02	0.10**	-0.06	0.272	0.237
	(0.13)	(0.14)	(0.40)			(0.08)	(0.05)	(0.10)		
7	-0.58**	0.00	-1.26	0.172	0.023	0.00	0.14**	-0.04	0.225	0.178
	(0.25)	(0.17)	(1.03)			(0.08)	(0.07)	(0.09)		
8	-0.74**	0.03	-2.18	0.444	0.015	0.00	0.10**	-0.03	0.278	0.272
	(0.33)	(0.17)	(3.00)			(0.07)	(0.05)	(0.08)		
9	-0.64*	0.11	-1.53	0.301	0.018	0.01	0.15**	-0.04	0.258	0.250
	(0.36)	(0.20)	(1.74)			(0.09)	(0.06)	(0.10)		
10	-0.99	0.17	-4.48	0.741	0.021	-0.01	0.10	-0.05	0.296	0.303
	(0.77)	(0.32)	(14.25)			(0.09)	(0.08)	(0.09)		
11	-0.82	0.48*	-2.20	0.415	0.037	0.01	0.15**	-0.05	0.164	0.206
	(0.57)	(0.27)	(3.67)			(0.09)	(0.07)	(0.08)		
12	-0.67*	0.58*	-1.16	0.002	0.077	-0.00	0.19**	-0.07	0.115	0.161
	(0.37)	(0.34)	(0.94)			(0.10)	(0.09)	(0.09)		
13	-1.05	0.76	-3.71	0.573	0.103	0.08	0.17**	-0.00	0.341	0.360
	(0.86)	(0.47)	(8.75)			(0.13)	(0.07)	(0.11)		
14	-0.66**	0.74	-1.03	0.010	0.109	0.11	0.24**	0.00	0.267	0.289
	(0.30)	(0.67)	(0.72)			(0.12)	(0.10)	(0.11)		
15	-0.99	0.89	-2.36	0.312	0.019	0.16	0.22***	0.04	0.418	0.421
	(0.78)	(1.01)	(4.05)			(0.15)	(0.07)	(0.14)		
16	-1.19	1.25	-4.79	0.793	0.312	0.15	0.25***	0.02	0.318	0.323
	(1.53)	(1.90)	(23.70)			(0.15)	(0.07)	(0.14)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C10: Consumption mult. (*pb*) under Large and Tight Fiscal Space following a Fiscal Cons. announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.14	0.04	-0.25	0.206	0.057	-0.01	0.07	-0.01	0.397	0.455
	(0.09)	(0.05)	(0.23)			(0.06)	(0.05)	(0.09)		
2	-0.21**	0.02	-0.33	0.139	0.061	-0.01	0.01	-0.00	0.870	0.871
	(0.10)	(0.07)	(0.22)			(0.05)	(0.04)	(0.07)		
3	-0.18***	-0.01	-0.26***	0.003	0.029	-0.02	0.05*	-0.04	0.310	0.328
	(0.04)	(0.05)	(0.07)			(0.05)	(0.03)	(0.08)		
4	-0.30***	-0.02	-0.53***	0.000	0.005	-0.02	0.08**	-0.04	0.149	0.117
	(0.06)	(0.05)	(0.13)			(0.04)	(0.03)	(0.07)		
5	-0.28***	-0.00	-0.46***	0.001	0.005	-0.01	0.11***	-0.04	0.052	0.042
	(0.06)	(0.06)	(0.13)			(0.04)	(0.03)	(0.05)		
6	-0.34***	-0.02	-0.60***	0.000	0.009	-0.01	0.07**	-0.03	0.188	0.171
	(0.07)	(0.08)	(0.15)			(0.04)	(0.03)	(0.06)		
7	-0.40***	0.00	-0.66***	0.004	0.015	0.00	0.08**	-0.03	0.163	0.139
	(0.11)	(0.09)	(0.21)			(0.04)	(0.04)	(0.05)		
8	-0.48***	0.02	-0.95**	0.030	0.009	0.00	0.07**	-0.02	0.231	0.229
	(0.15)	(0.10)	(0.46)			(0.04)	(0.03)	(0.05)		
9	-0.45***	0.06	-0.89**	0.027	0.013	0.00	0.09**	-0.02	0.198	0.201
	(0.13)	(0.09)	(0.43)			(0.04)	(0.05)	(0.06)		
10	-0.55***	0.08	-1.14*	0.064	0.015	-0.00	0.06	-0.04	0.300	0.308
	(0.18)	(0.13)	(0.65)			(0.04)	(0.05)	(0.06)		
11	-0.53***	0.18***	-0.91***	0.000	0.013	0.01	0.09*	-0.03	0.180	0.209
	(0.17)	(0.06)	(0.27)			(0.05)	(0.05)	(0.06)		
12	-0.40***	0.21***	-0.50***	0.000	0.049	-0.00	0.11	-0.05	0.139	0.160
	(0.12)	(0.05)	(0.13)			(0.05)	(0.07)	(0.06)		
13	-0.67**	0.21***	-1.18*	0.034	0.011	0.04	0.10**	-0.00	0.372	0.379
	(0.27)	(0.07)	(0.62)			(0.06)	(0.05)	(0.09)		
14	-0.46***	0.18***	-0.55***	0.000	0.017	0.05	0.12*	0.00	0.328	0.336
	(0.14)	(0.06)	(0.17)			(0.06)	(0.06)	(0.08)		
15	-0.59*	0.22***	-0.77**	0.006	0.009	0.09	0.13**	0.03	0.444	0.446
	(0.32)	(0.07)	(0.34)			(0.08)	(0.06)	(0.10)		
16	-0.54*	0.28***	-0.65	0.040	0.015	0.08	0.18**	0.01	0.248	0.254
	(0.31)	(0.08)	(0.46)			(0.08)	(0.08)	(0.10)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.C11: Investment multiplier under Large and Tight FS following a Fiscal Consolidation announcement.

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-1.39*	-0.00	-2.19	0.199	0.166	-0.53	2.11	-0.92	0.107	0.053
	(0.83)	(1.28)	(1.60)			(0.76)	(1.36)	(0.87)		
2	-1.13**	0.73	-1.56*	0.024	0.066	0.86	1.28*	0.78	0.700	0.692
	(0.53)	(0.87)	(0.88)			(0.78)	(0.69)	(0.86)		
3	-0.99***	0.65	-1.74**	0.012	0.027	0.42	1.15***	0.18	0.392	0.400
	(0.33)	(0.65)	(0.77)			(0.68)	(0.32)	(0.85)		
4	-1.45***	0.38	-2.43**	0.004	0.022	1.27	1.20***	1.36	0.833	0.834
	(0.48)	(0.57)	(0.99)			(0.87)	(0.40)	(1.10)		
5	-1.47***	0.74	-2.86***	0.001	0.012	0.98	1.53***	0.80	0.615	0.601
	(0.39)	(0.77)	(1.08)			(0.81)	(0.41)	(1.01)		
6	-2.46***	0.55	-4.77*	0.042	0.055	1.67	1.12***	1.69	0.672	0.675
	(0.88)	(0.73)	(2.75)			(1.40)	(0.36)	(1.71)		
7	-3.04**	0.69	-7.00	0.158	0.030	0.68	1.40***	0.41	0.318	0.214
	(1.39)	(0.79)	(5.83)			(0.48)	(0.48)	(0.61)		
8	-4.05*	1.29	-13.83	0.402	0.032	0.34	1.24***	-0.00	0.139	0.082
	(2.08)	(1.15)	(18.97)			(0.52)	(0.27)	(0.62)		
9	-3.46*	1.68	-9.27	0.235	0.028	0.94	1.50***	0.53	0.412	0.335
	(1.93)	(1.19)	(10.30)			(0.75)	(0.33)	(0.81)		
10	-5.54	1.84	-26.19	0.729	0.040	0.84	1.77***	0.41	0.064	0.010
	(4.18)	(1.72)	(82.04)			(0.54)	(0.41)	(0.54)		
11	-4.53*	3.67	-12.34	0.318	0.109	1.05*	2.08***	0.50	0.116	0.065
	(2.74)	(2.32)	(18.88)			(0.60)	(0.48)	(0.61)		
12	-3.82**	4.30	-6.57	0.001	0.060	1.00*	2.38***	0.30	0.070	0.029
	(1.90)	(2.84)	(4.80)			(0.55)	(0.59)	(0.52)		
13	-5.69	6.62	-20.63	0.495	0.228	1.00	2.08***	0.15	0.029	0.025
	(4.54)	(4.08)	(46.74)			(0.73)	(0.42)	(0.63)		
14	-4.06***	7.76	-6.85*	0.022	0.117	0.71	2.65***	-0.21	0.016	0.020
	(1.36)	(6.58)	(3.87)			(0.73)	(0.61)	(0.63)		
15	-5.26	8.30	-12.95	0.126	0.249	1.13	2.20***	0.05	0.097	0.070
	(3.69)	(9.01)	(20.02)			(0.95)	(0.44)	(0.84)		
16	-6.71	11.88	-27.88	0.753	0.488	1.70	2.44***	0.37	0.148	0.128
	(9.08)	(18.22)	(138.11)			(1.05)	(0.36)	(0.88)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.C12: Investment mult. (*pb*) under Large and Tight Fiscal Space following a Fiscal Cons. announcement.

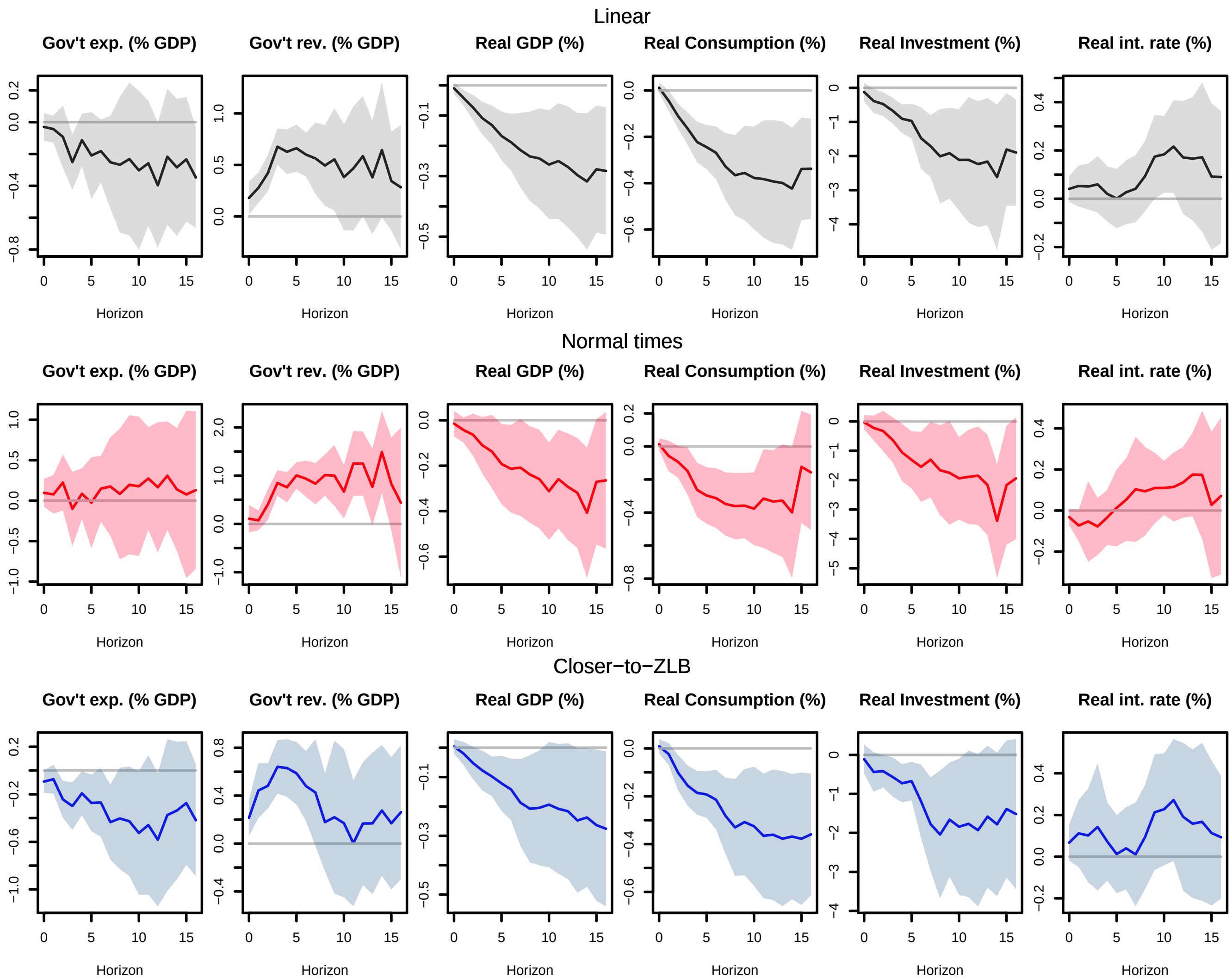
	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-1.20*	-0.00	-1.97	0.227	0.072	-0.26	1.47	-0.52	0.154	0.022
	(0.69)	(0.69)	(1.58)			(0.41)	(1.21)	(0.54)		
2	-0.92*	0.33	-1.34	0.153	0.046	0.51	0.91**	0.55	0.465	0.437
	(0.52)	(0.34)	(1.09)			(0.44)	(0.39)	(0.56)		
3	-0.72***	0.43	-1.18**	0.006	0.026	0.25	0.78***	0.14	0.307	0.323
	(0.22)	(0.36)	(0.51)			(0.39)	(0.21)	(0.64)		
4	-1.23***	0.23	-2.23***	0.004	0.012	0.72	0.77***	0.98	0.891	0.891
	(0.30)	(0.32)	(0.75)			(0.46)	(0.21)	(0.75)		
5	-1.12***	0.49	-1.98***	0.000	0.008	0.49	0.96***	0.50	0.330	0.293
	(0.24)	(0.42)	(0.63)			(0.37)	(0.25)	(0.58)		
6	-1.89***	0.29	-3.57***	0.011	0.030	0.81	0.78***	1.03	0.851	0.854
	(0.55)	(0.33)	(1.33)			(0.62)	(0.16)	(0.93)		
7	-2.10***	0.39	-3.67***	0.010	0.022	0.32	0.84***	0.25	0.110	0.053
	(0.56)	(0.36)	(1.38)			(0.23)	(0.22)	(0.37)		
8	-2.63***	0.74	-6.00**	0.011	0.029	0.17	0.83***	-0.00	0.046	0.020
	(0.78)	(0.47)	(2.71)			(0.27)	(0.19)	(0.41)		
9	-2.44***	0.86*	-5.38**	0.022	0.028	0.40	0.94***	0.31	0.131	0.082
	(0.68)	(0.45)	(2.72)			(0.32)	(0.23)	(0.46)		
10	-3.09***	0.83	-6.67*	0.074	0.037	0.42	1.00***	0.30	0.040	0.009
	(0.87)	(0.56)	(3.95)			(0.28)	(0.35)	(0.38)		
11	-2.91***	1.37**	-5.11***	0.001	0.053	0.54*	1.19***	0.36	0.102	0.033
	(0.74)	(0.54)	(1.51)			(0.29)	(0.42)	(0.42)		
12	-2.28***	1.58***	-2.86***	0.002	0.061	0.48*	1.38**	0.22	0.043	0.017
	(0.58)	(0.56)	(0.97)			(0.25)	(0.54)	(0.37)		
13	-3.62**	1.83***	-6.56	0.094	0.047	0.52	1.20***	0.12	0.022	0.012
	(1.65)	(0.65)	(4.56)			(0.37)	(0.38)	(0.49)		
14	-2.82***	1.94***	-3.69***	0.001	0.036	0.36	1.31***	-0.15	0.029	0.019
	(0.83)	(0.60)	(1.34)			(0.38)	(0.47)	(0.46)		
15	-3.13*	2.03***	-4.24	0.043	0.049	0.62	1.30***	0.04	0.094	0.071
	(1.71)	(0.62)	(2.69)			(0.53)	(0.40)	(0.64)		
16	-3.01	2.65***	-3.78	0.066	0.029	0.88	1.78***	0.27	0.071	0.047
	(2.08)	(0.82)	(3.34)			(0.55)	(0.61)	(0.64)		

The table shows cumulative *primary balance*-based multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the mult. after a TB (EB) announcement. In each panel, columns 1, 2, 3 report mult. under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between mult. in the two states using respectively clustered S.E. and Anderson-Rubin (AR) C.I. (to account for the possibility that announcement series is a weak instrument). *p*-value $\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

4.D Robustness: Figures

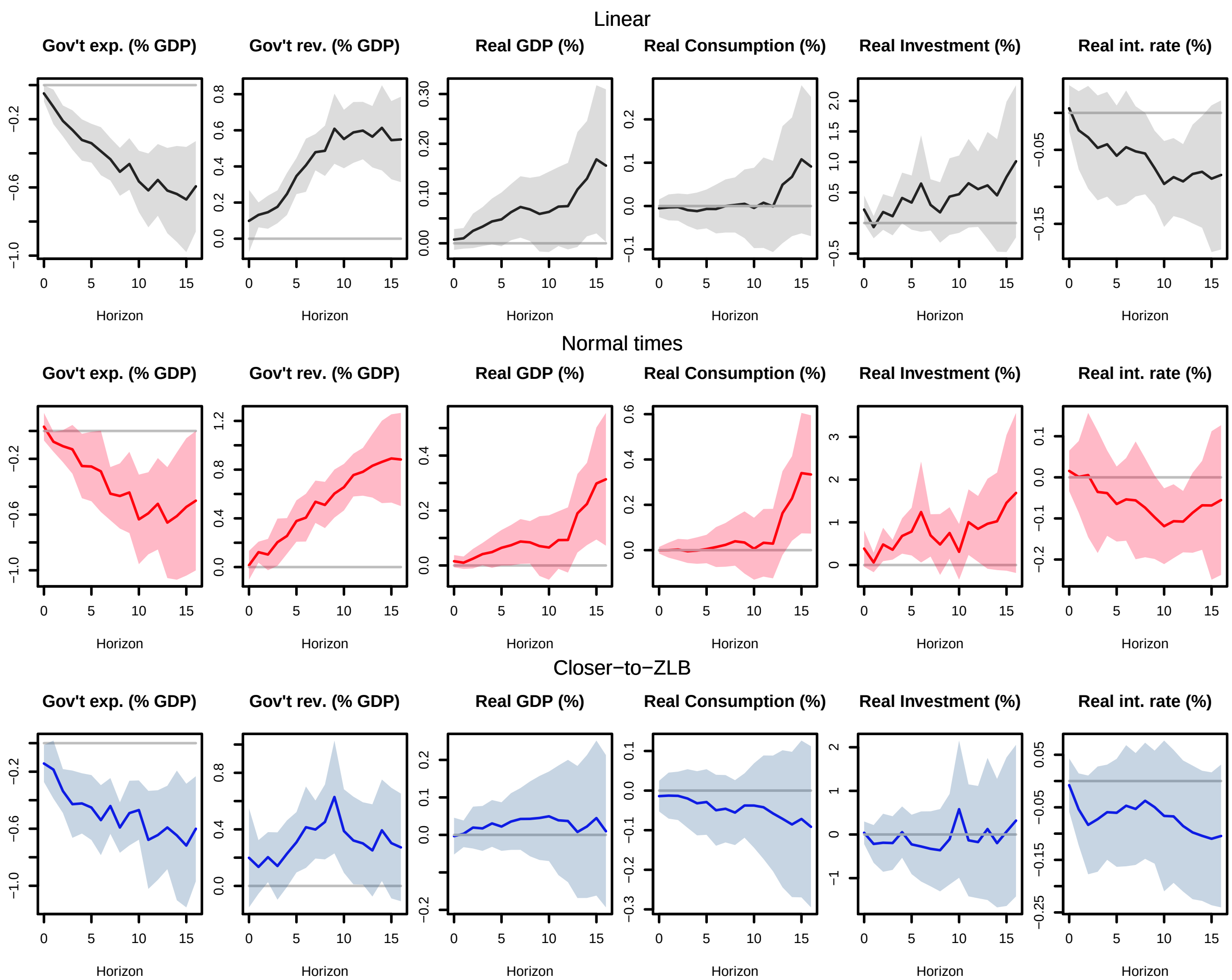
In this section, IRFs from various robustness checks are reported. Figures 4.D1 and 4.D2 report the IRFs using OECD Recession dates as indicator for business cycle states (see Sec. 4.3). Figures 4.D3 and 4.D4 report the IRFs using unemployment as indicator for business cycle states (see Sec. 4.3). Figure 4.D5 reports the IRFs following an EB consolidation announcement using as state a dummy taking values equal to 1 if the country is part of the Eurozone and quarters are past 1999:Q1, which means that monetary policy is constrained (see Sec. 4.3). Figures 4.D6 and 4.D7 report IRFs using *de facto* fiscal space, namely FS2, as indicator for the fiscal position (see Sec. 4.3). Moreover, Figures 4.D8- 4.D13 show the IRFs resulting from the specification of Eq. 4.4 without any country-specific time trend. Lastly, Figures 4.D14- 4.D19 report the IRFs obtained under the specification of Eq. 4.4 in which I include 4 lags.

Figure 4.D1: IRFs: TB consolidation announcement - *Expansion/Recession* (OECD Recession dates).



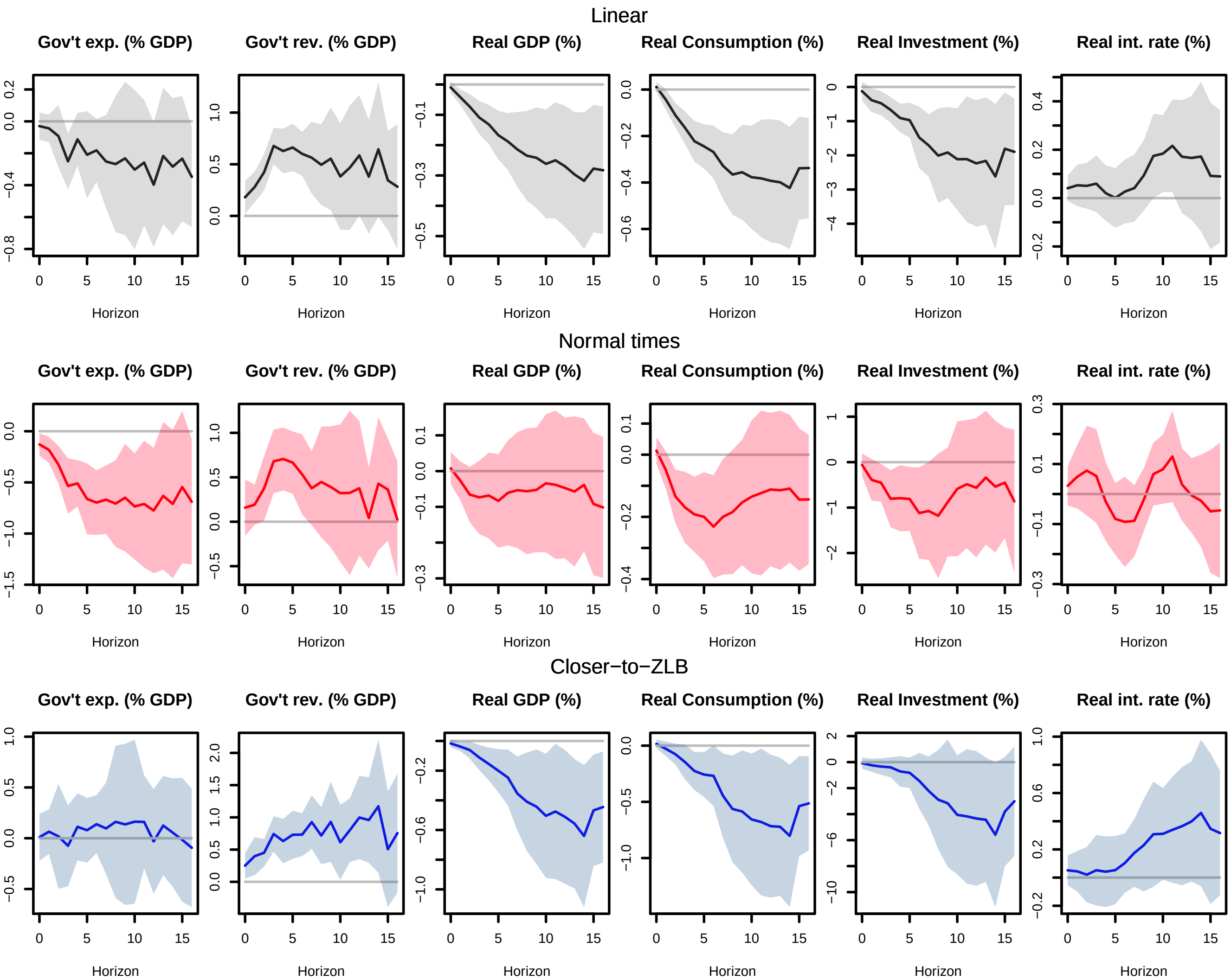
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D2: IRFs: EB consolidation announcement - *Expansion/Recession* (OECD Recession dates).



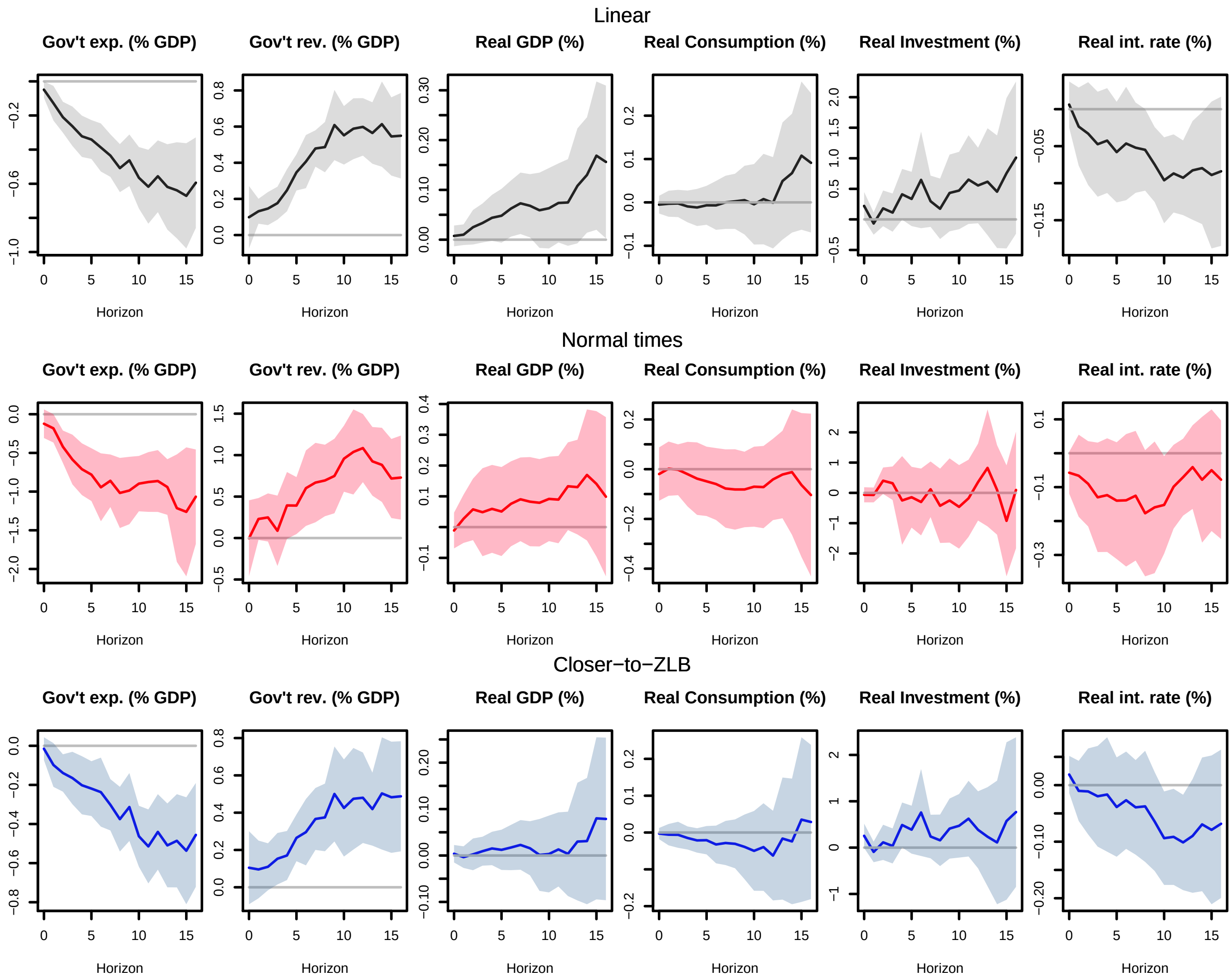
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D3: IRFs: TB consolidation announcement - *Expansion/Recession* (Unemployment).



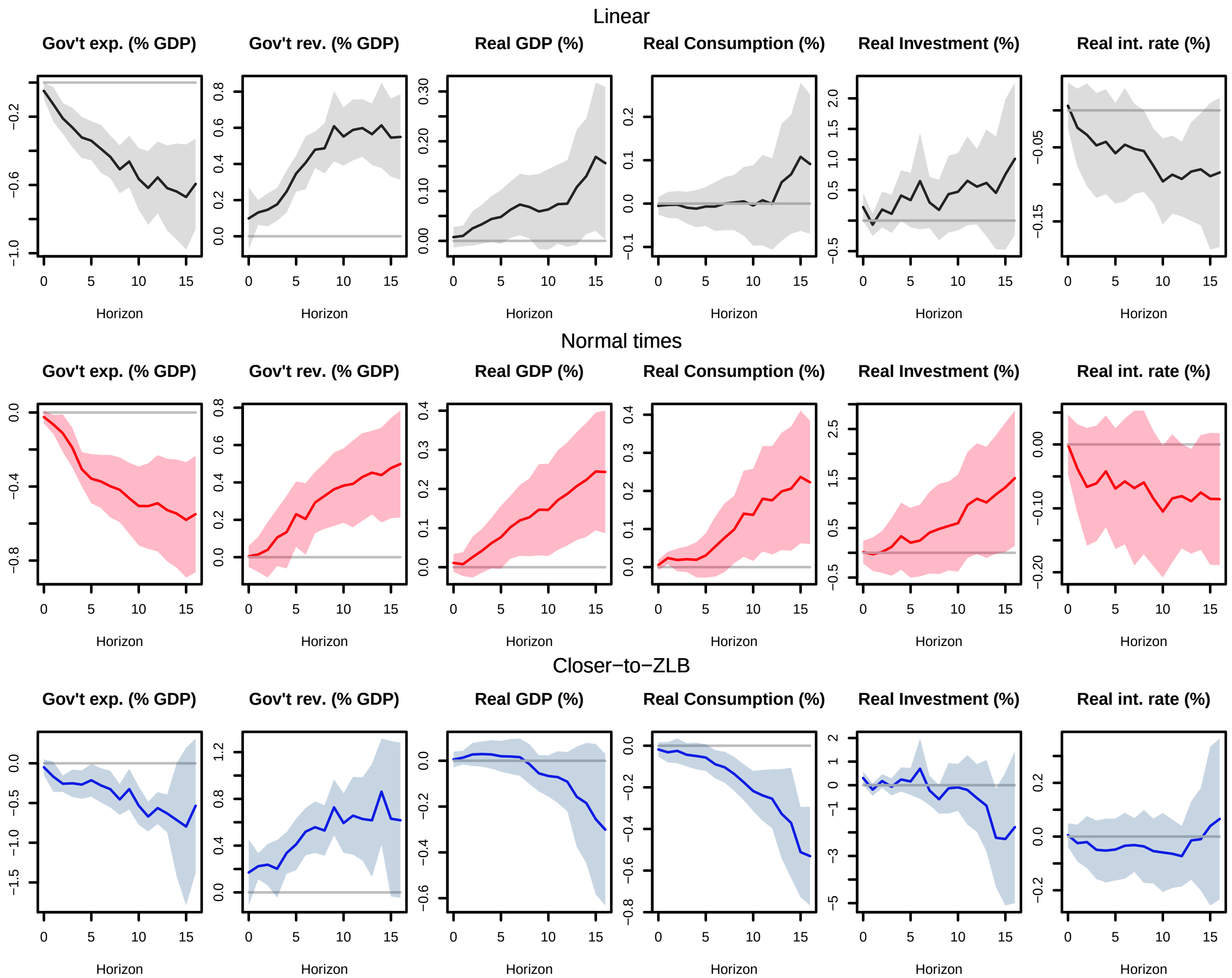
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D4: IRFs: EB consolidation announcement - *Expansion/Recession* (Unemployment).



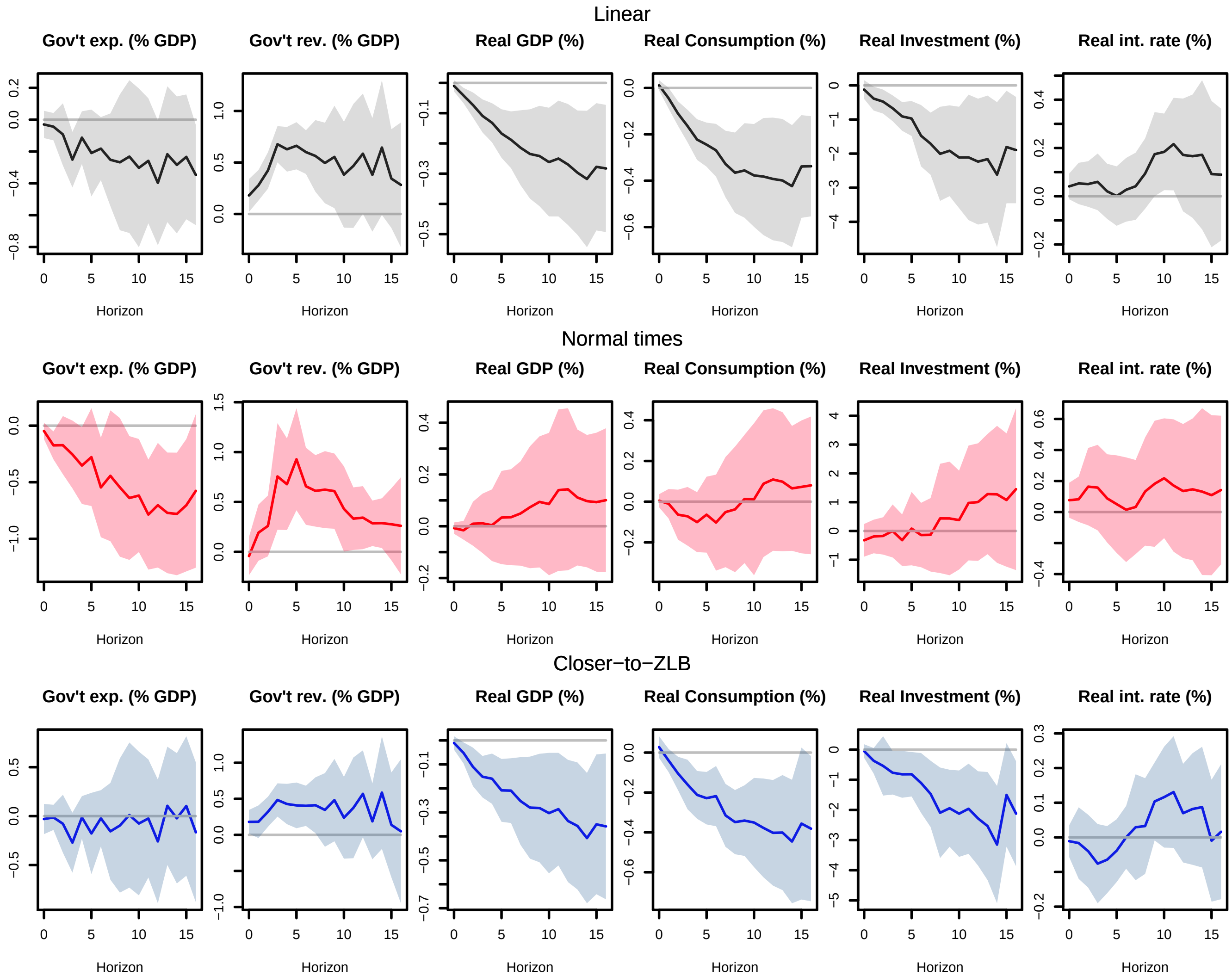
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D5: IRFs: EB consolidation announcement - *Unconstrained/Constrained MP* (EZ countries, t >1999:Q1).



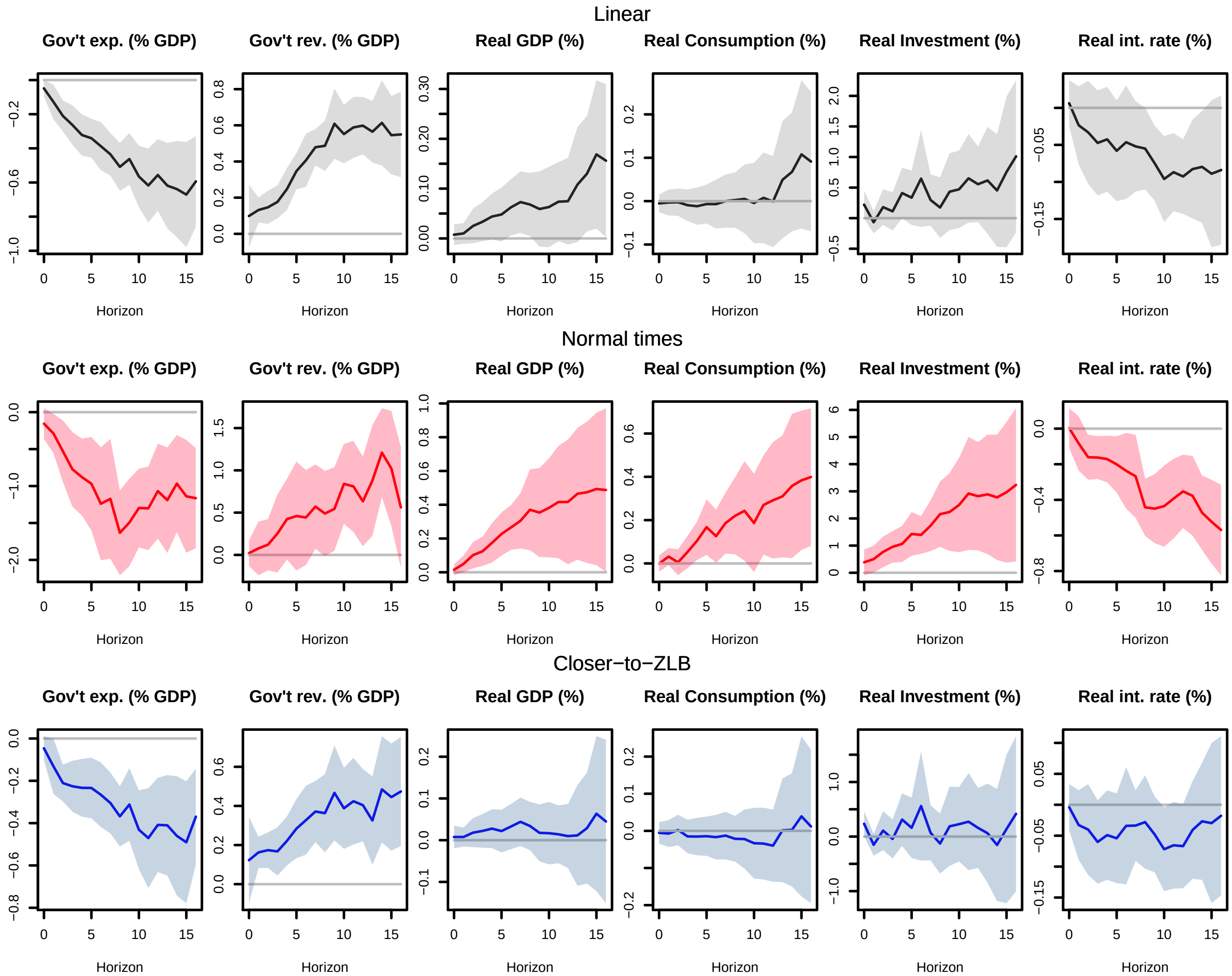
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *unconstrained monetary policy* (confidence interval in red); while, at the bottom, are shown the IRFs under *constrained monetary policy* (confidence interval in blue).

Figure 4.D6: IRFs: TB consolidation announcement - *Large/Tight FS* (*de facto* fiscal space).



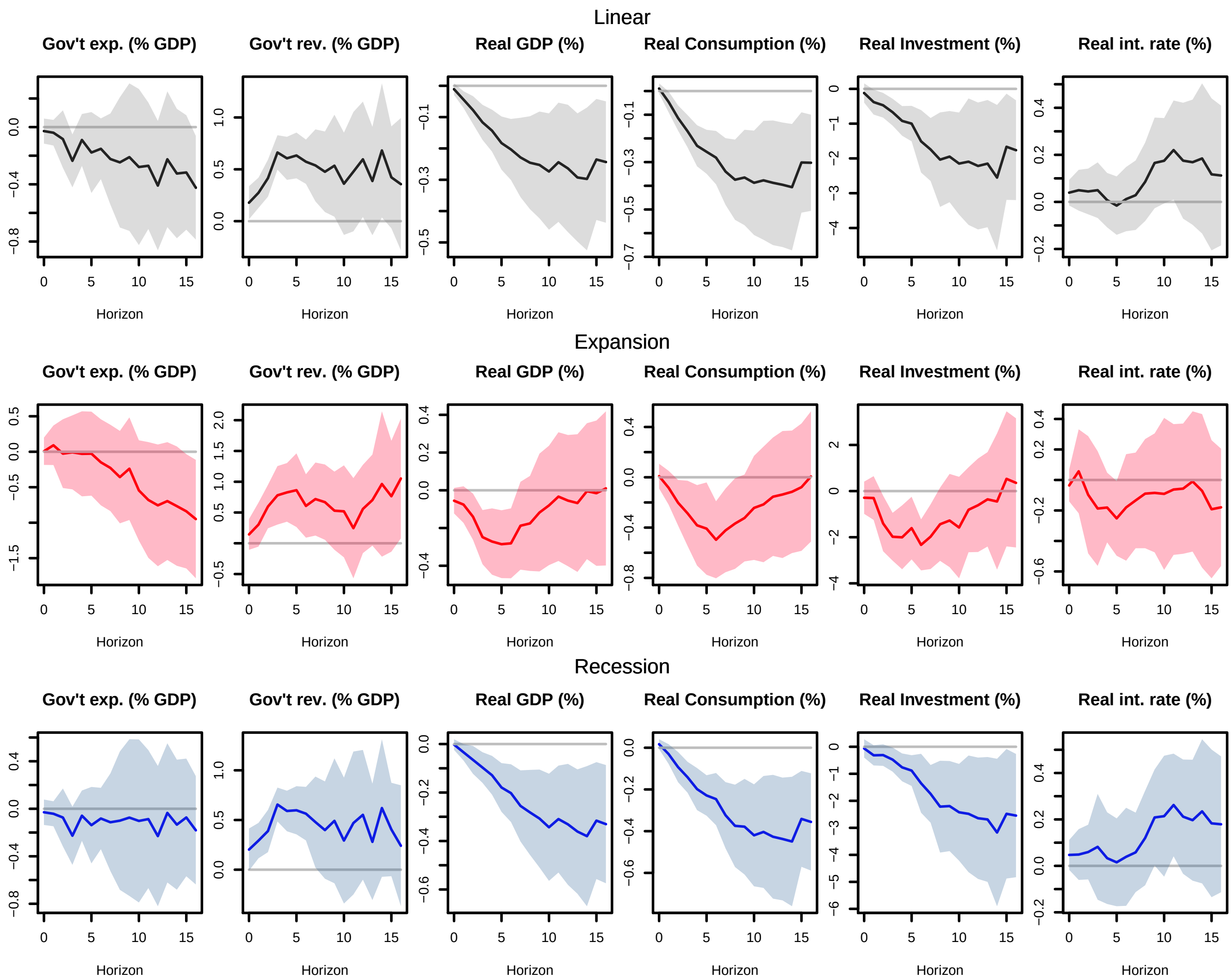
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Figure 4.D7: IRFs: EB consolidation announcement - *Large/Tight FS* (*de facto* fiscal space).



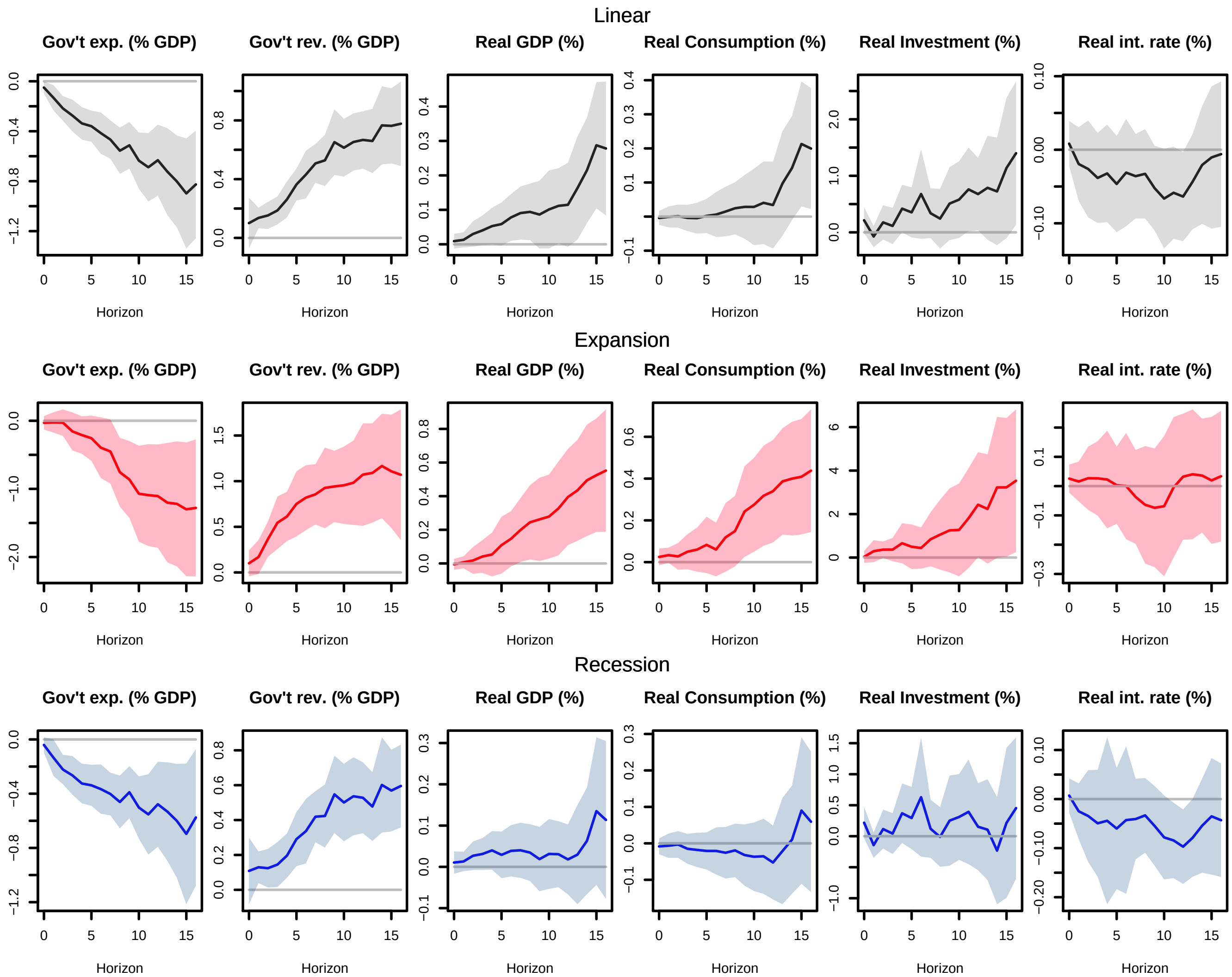
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Figure 4.D8: IRFs: TB consolidation announcement - *Expansion/Recession* (no trend).



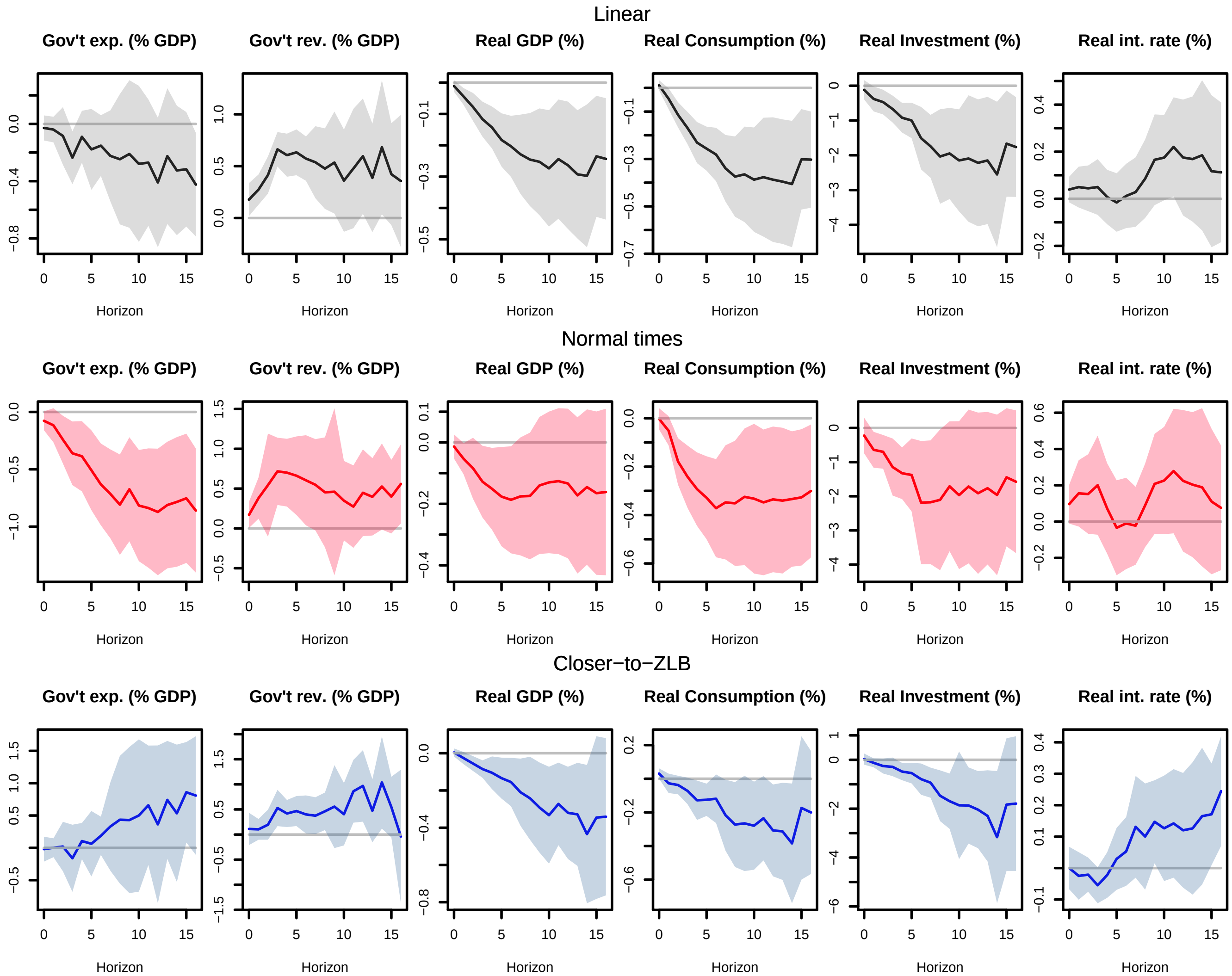
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D9: IRFs: EB consolidation announcement - *Expansion/Recession* (no trend).



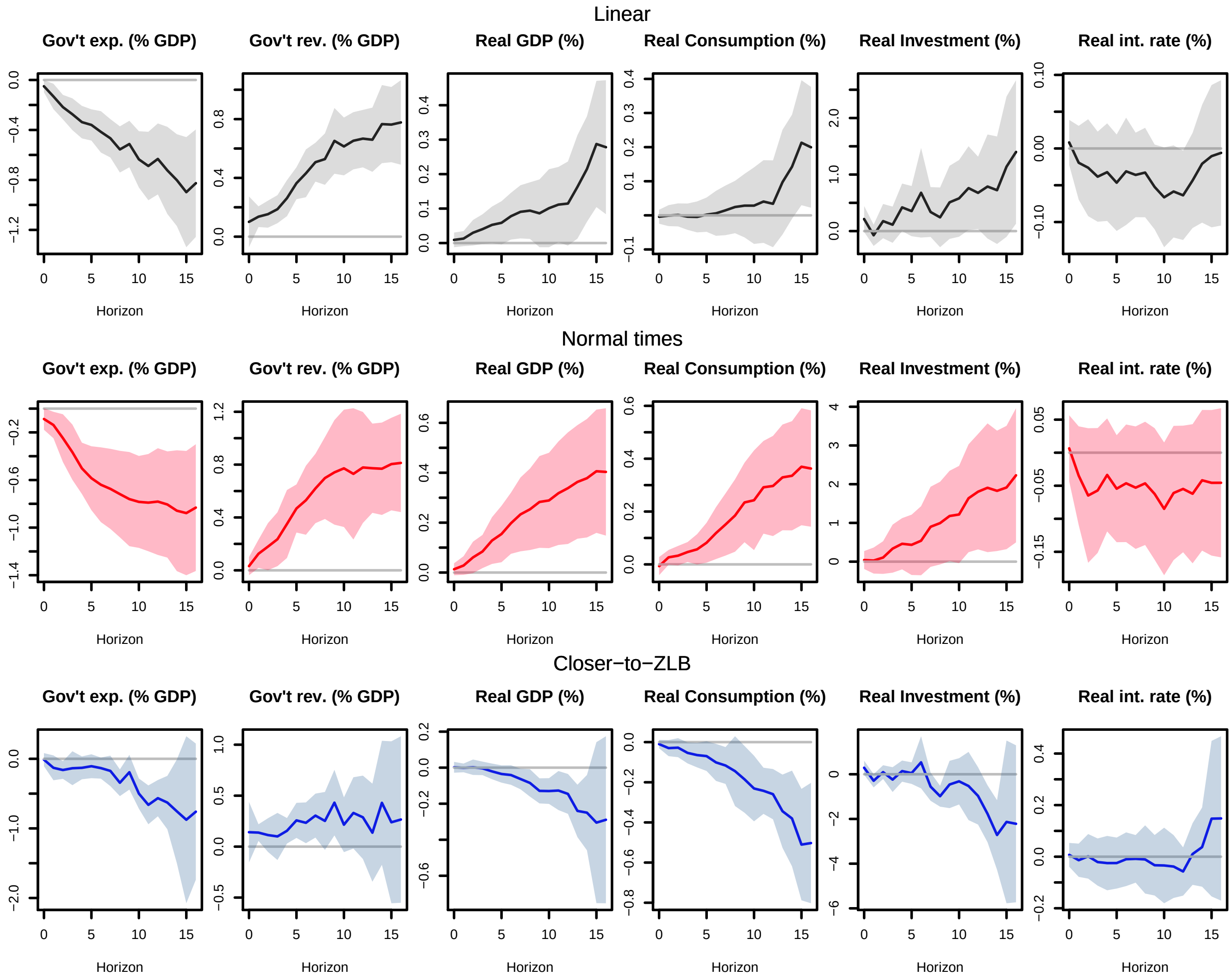
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D10: IRFs: TB consolidation announcement - *Normal Times/Close-to-ZLB* (no trend).



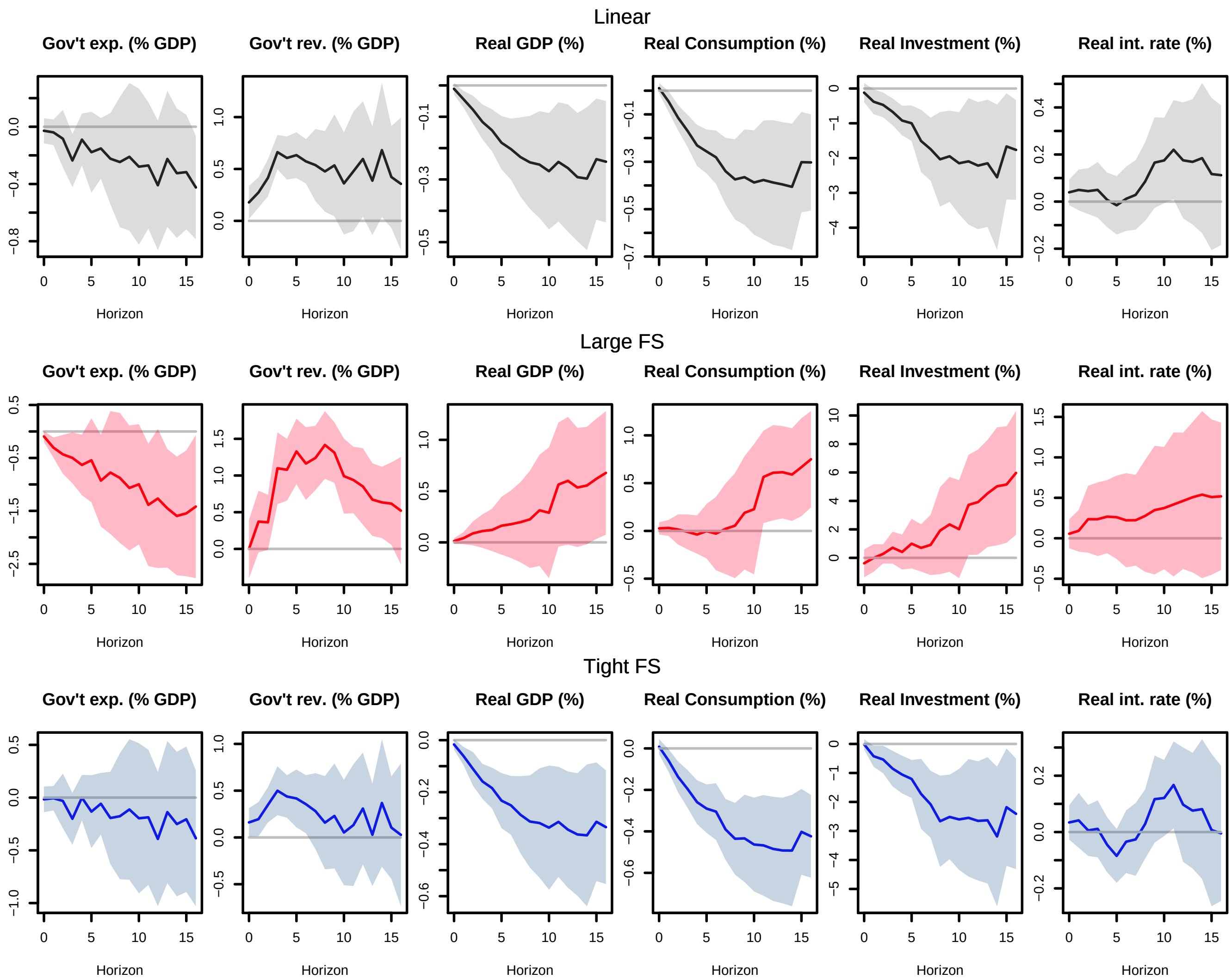
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

Figure 4.D11: IRFs: EB consolidation announcement - *Normal Times/Close-to-ZLB* (no trend).



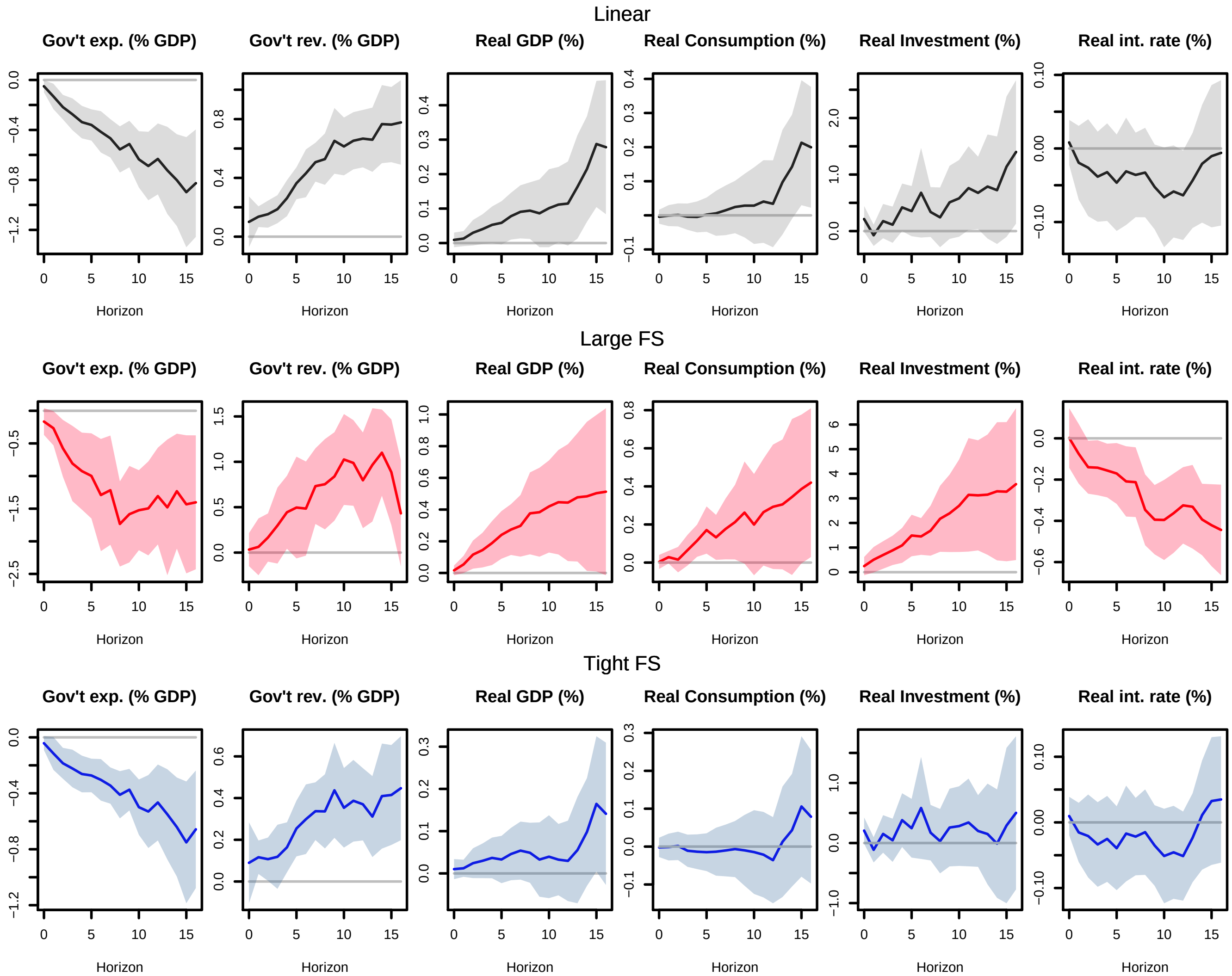
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

Figure 4.D12: IRFs: TB consolidation announcement - *Large/Tight fiscal space* (no trend).



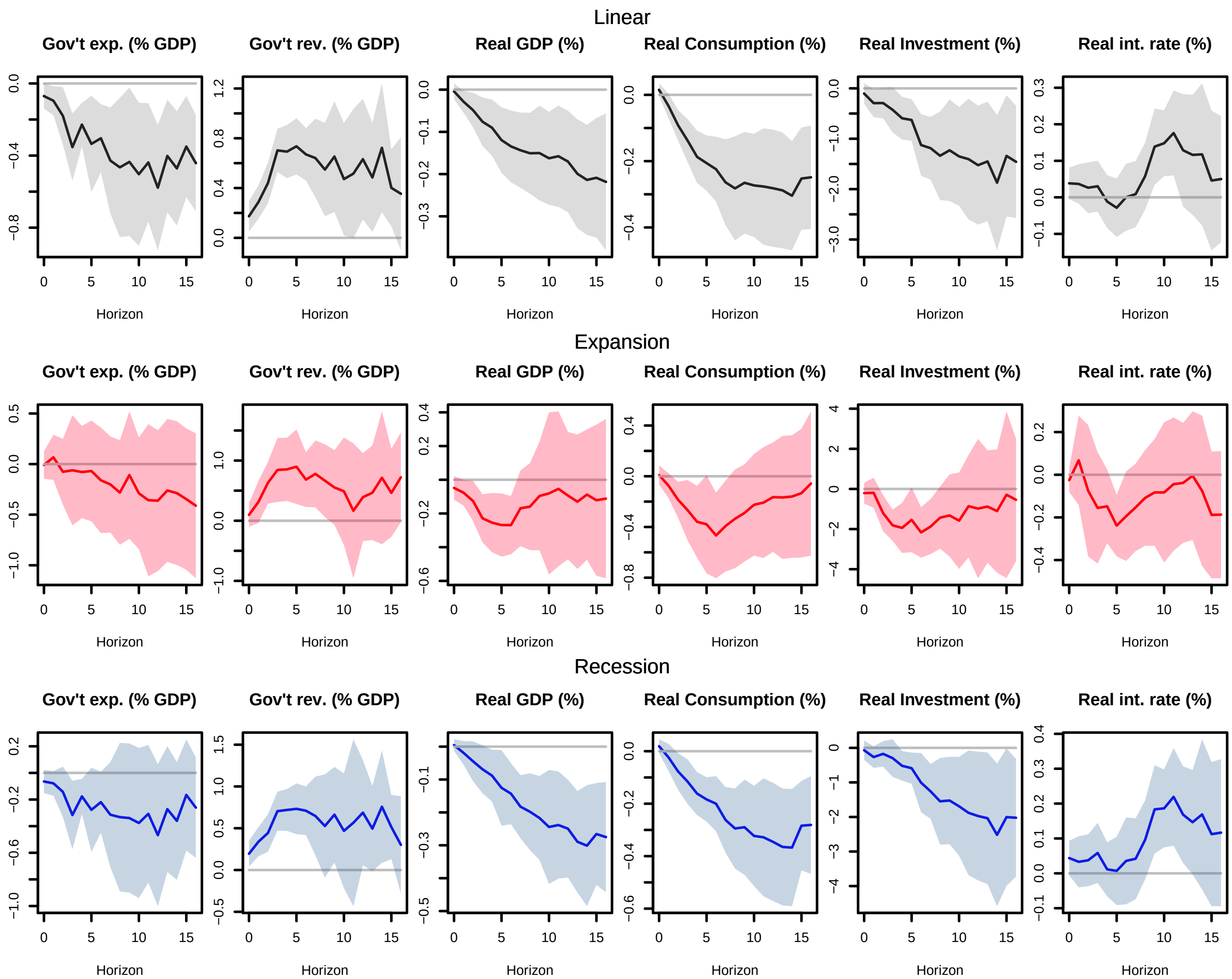
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Figure 4.D13: IRFs: EB consolidation announcement - *Large/Tight fiscal space* (no trend).



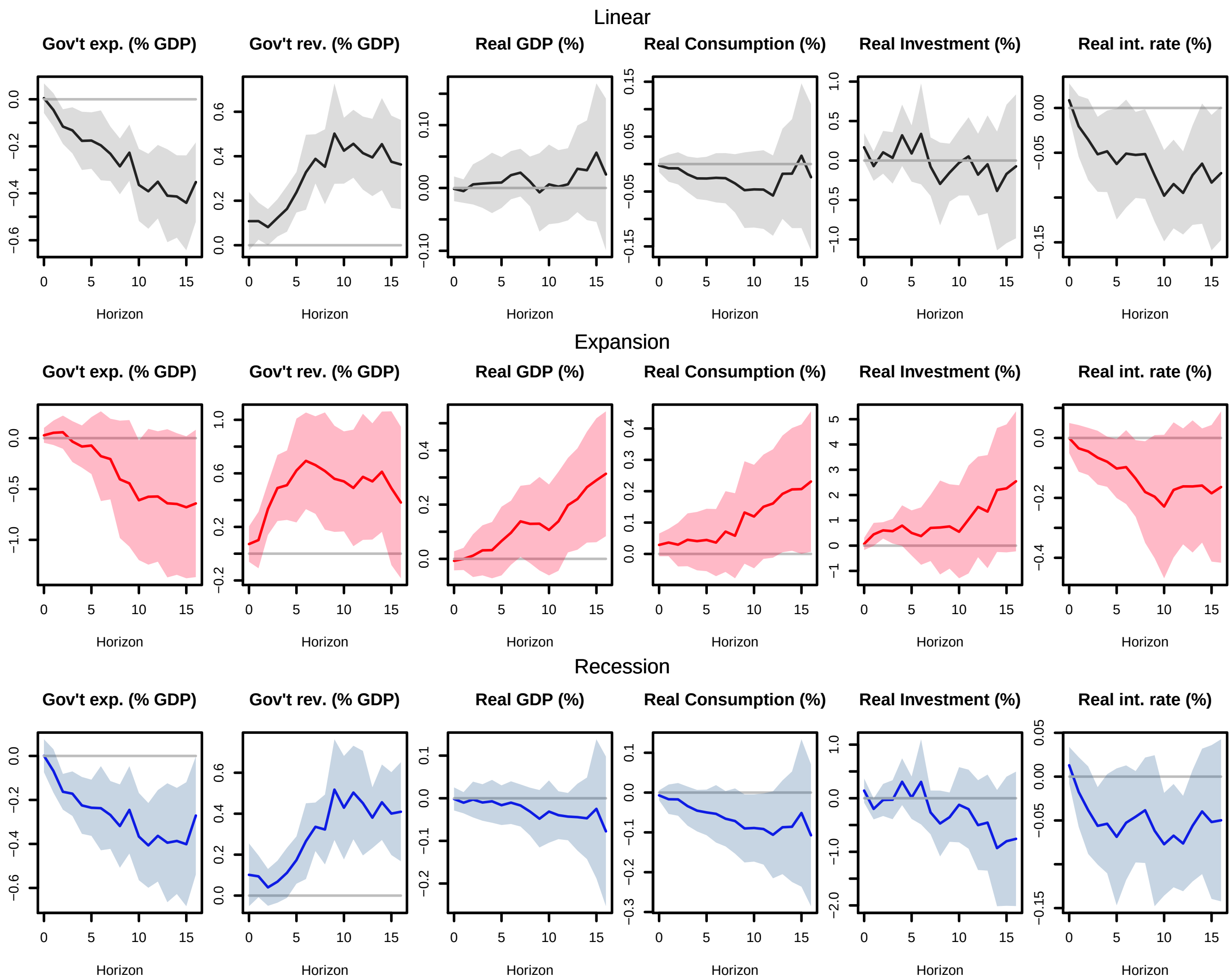
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Figure 4.D14: IRFs: TB consolidation announcement - *Expansion/Recession* (4 lags).



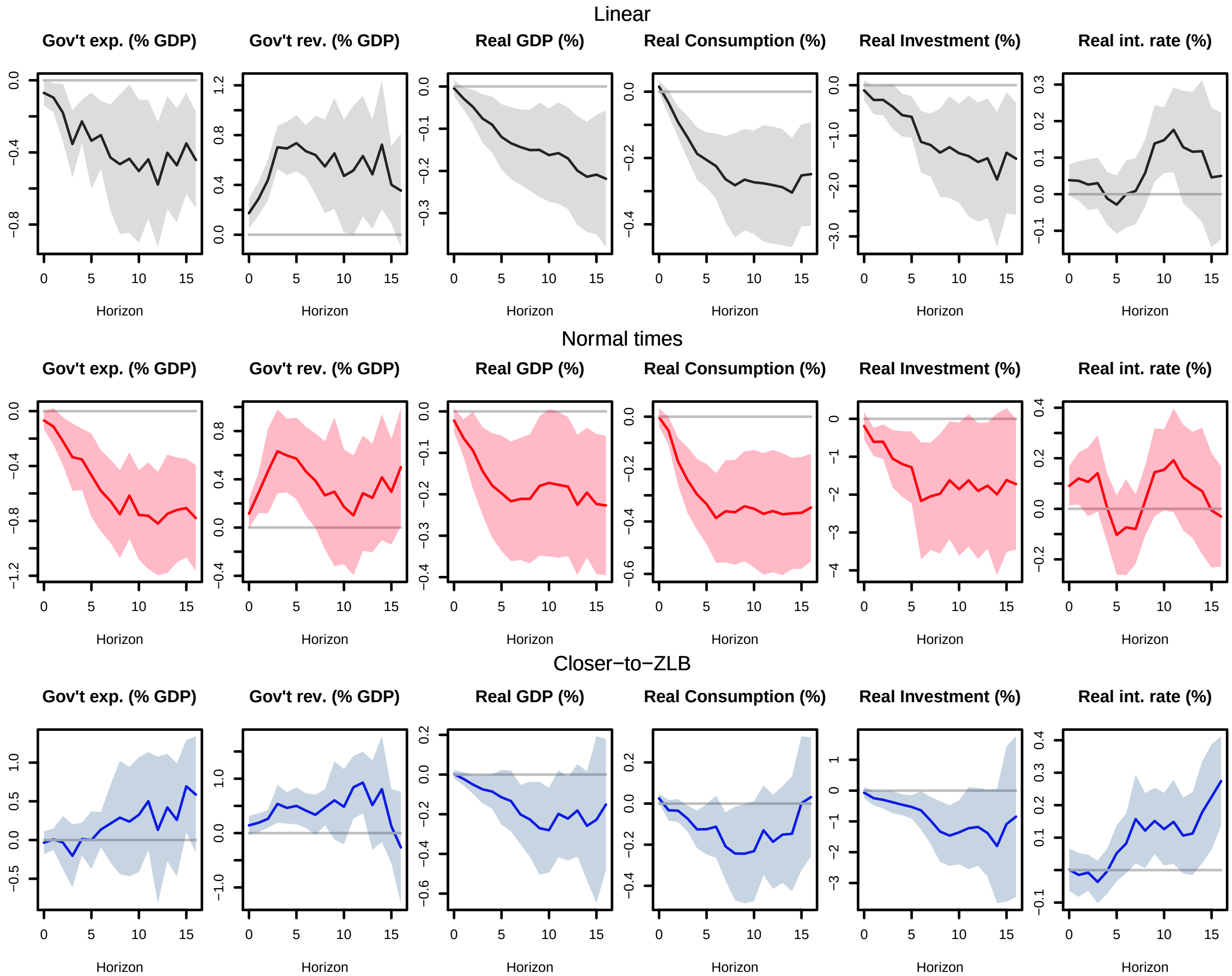
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D15: IRFs: EB consolidation announcement - *Expansion/Recession* (4 lags).



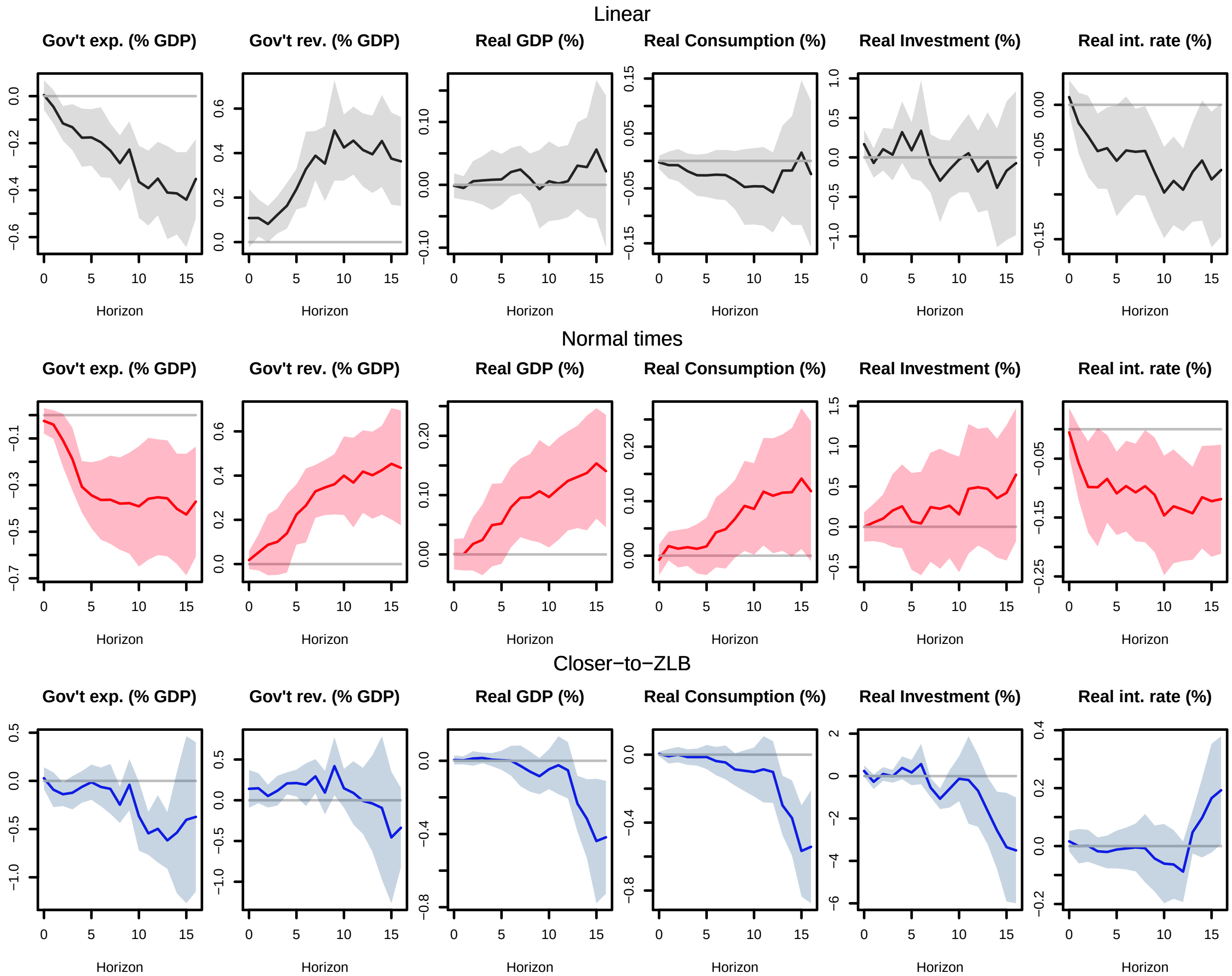
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *expansion* (confidence interval in red); while, at the bottom, are shown the IRFs under *recession* (confidence interval in blue).

Figure 4.D16: IRFs: TB consolidation announcement - *Normal Times/Close-to-ZLB* (4 lags).



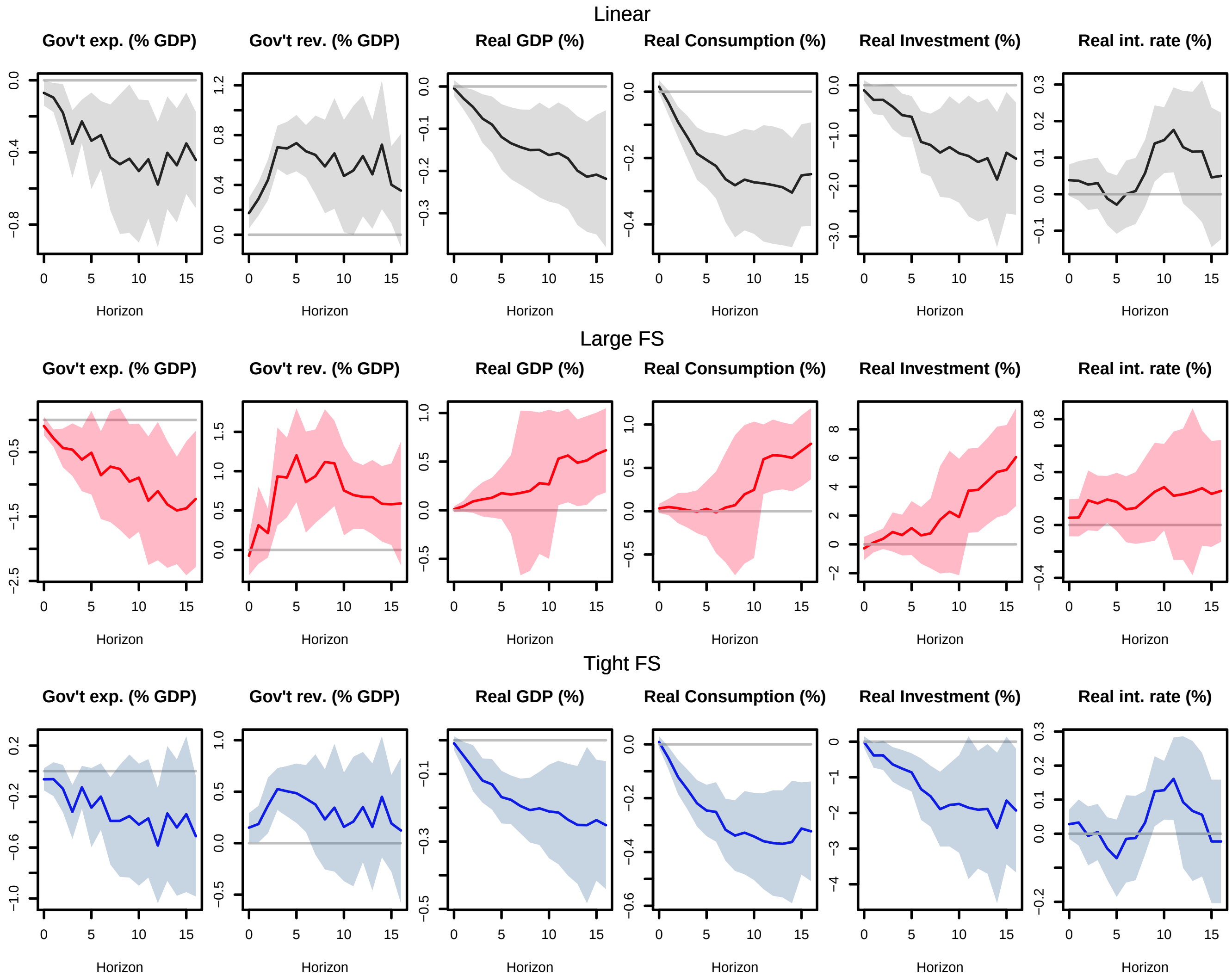
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

Figure 4.D17: IRFs: EB consolidation announcement - *Normal Times/Close-to-ZLB* (4 lags).



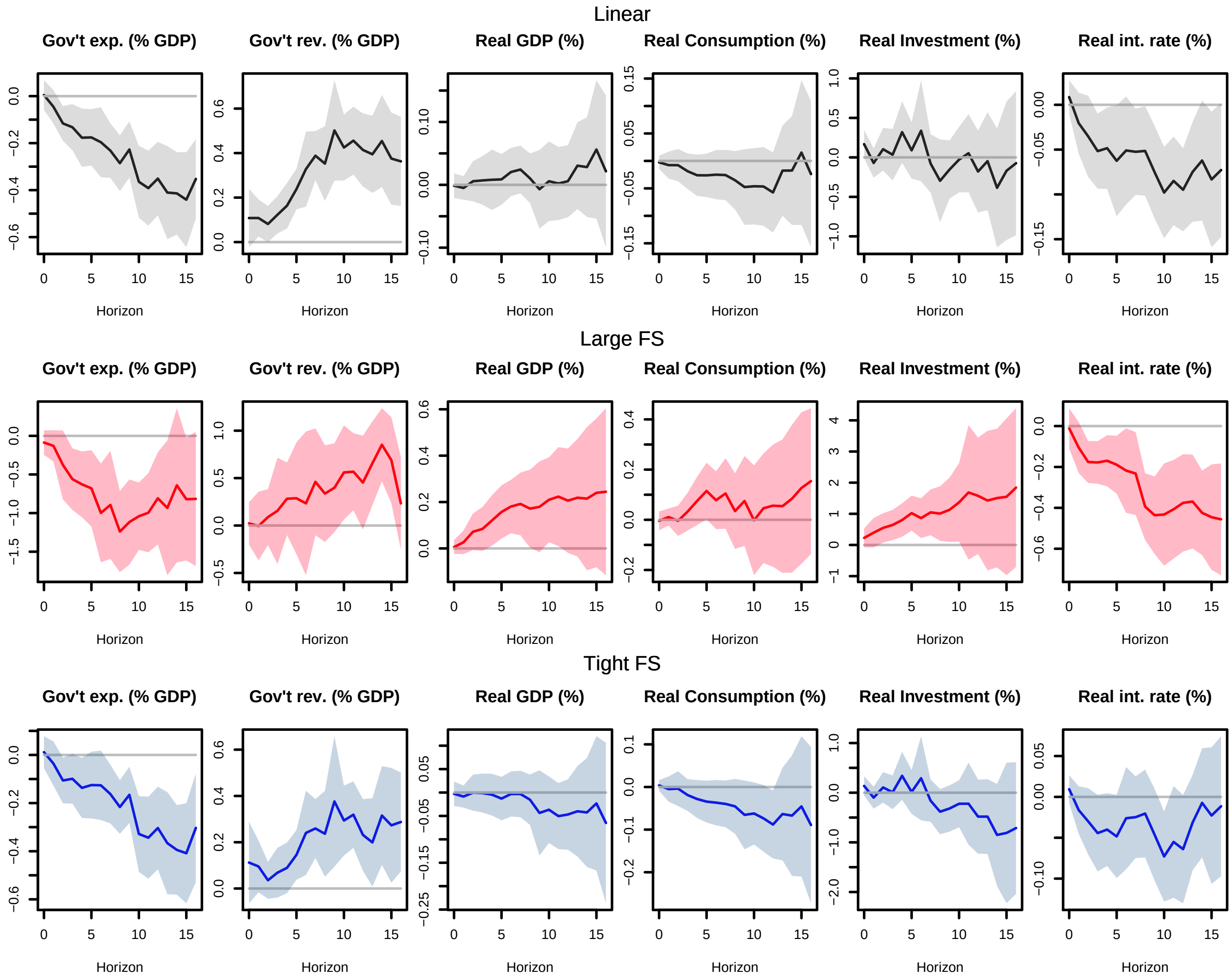
This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *normal times* (confidence interval in red); while, at the bottom, are shown the IRFs under *close-to-ZLB* (confidence interval in blue).

Figure 4.D18: IRFs: TB consolidation announcement - *Large/Tight fiscal space* (4 lags).



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following a TB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

Figure 4.D19: IRFs: EB consolidation announcement - *Large/Tight fiscal space* (4 lags).



This figure reports cumulative IRFs for government expenditure (% of GDP), government revenues (% of GDP), real GDP (%), real consumption (%), real investment (%) and real long-term interest rate (%) following an EB consolidation announcement. Horizons (x-axis) are expressed in quarters and the maximum projection horizon is 16 quarters (4 years). Confidence intervals are reported at the 90% significance level. At the top of the graph, the linear case (confidence interval in grey) is illustrated. In the middle, I report the responses under *large fiscal space* (confidence interval in red); while, at the bottom, are shown the IRFs under *tight fiscal space* (confidence interval in blue).

4.E Robustness: Tables

In this section, output, consumption and investment multipliers from various robustness checks are reported. Tables 4.E1 - 4.E3 report the fiscal multipliers using OECD Recession dates as indicator for business cycle states (see Sec. 4.3). Tables 4.E4 - 4.E6 report the fiscal multipliers using unemployment as indicator for business cycle states (see Sec. 4.3). Tables 4.E7 - 4.E9 reports the fiscal multipliers following an EB consolidation announcement using as state a dummy taking values equal to 1 if the country is part of the Eurozone and quarters are past 1999:Q1, which means that monetary policy is constrained (see Sec. 4.3). Tables 4.E10 - 4.E12 report IRFs using *de facto* fiscal space, namely FS2, as indicator for the fiscal position (see Sec. 4.3).

Table 4.E1: Output multiplier under Expansion and Recession following a FC ann. (OECD Recession dates).

	Tax-Based Fiscal Consolidation					Expenditure-Based Fiscal Consolidation				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.15***	-0.59	-0.05	0.173	0.192	0.08	0.11	0.01	0.484	0.472
	(0.06)	(0.42)	(0.04)			(0.11)	(0.16)	(0.11)		
2	-0.17***	-0.16	-0.11***	0.738	0.730	0.12	0.18	0.05	0.463	0.388
	(0.06)	(0.15)	(0.04)			(0.10)	(0.20)	(0.08)		
3	-0.16***	-0.13*	-0.13***	0.899	0.899	0.13	0.24	0.04	0.224	0.171
	(0.04)	(0.07)	(0.04)			(0.09)	(0.17)	(0.07)		
4	-0.21***	-0.19	-0.17***	0.813	0.809	0.14*	0.15	0.07	0.350	0.352
	(0.06)	(0.13)	(0.04)			(0.08)	(0.10)	(0.07)		
5	-0.25***	-0.19*	-0.23***	0.876	0.877	0.14	0.19*	0.04	0.131	0.159
	(0.06)	(0.10)	(0.07)			(0.09)	(0.11)	(0.07)		
6	-0.31***	-0.23*	-0.33***	0.731	0.733	0.16*	0.19*	0.06	0.201	0.207
	(0.08)	(0.14)	(0.13)			(0.08)	(0.10)	(0.06)		
7	-0.38**	-0.26*	-0.48	0.557	0.486	0.17***	0.15**	0.08	0.420	0.425
	(0.15)	(0.13)	(0.30)			(0.07)	(0.07)	(0.07)		
8	-0.47**	-0.25**	-1.24	0.541	0.207	0.13**	0.14**	0.06	0.313	0.326
	(0.22)	(0.11)	(1.57)			(0.06)	(0.06)	(0.06)		
9	-0.44*	-0.27*	-0.94	0.631	0.322	0.13	0.12	0.07	0.601	0.594
	(0.24)	(0.14)	(1.35)			(0.08)	(0.09)	(0.07)		
10	-0.69	-0.52**	-1.06	0.742	0.575	0.11	0.08	0.07	0.981	0.982
	(0.51)	(0.24)	(1.77)			(0.07)	(0.07)	(0.07)		
11	-0.54	-0.21**	-7.10	0.900	0.157	0.12	0.12*	0.04	0.441	0.436
	(0.37)	(0.09)	(57.38)			(0.07)	(0.07)	(0.08)		
12	-0.46*	-0.25**	-1.06	0.510	0.216	0.13	0.13	0.04	0.440	0.434
	(0.26)	(0.13)	(1.23)			(0.09)	(0.08)	(0.08)		
13	-0.78	-0.48	-1.13	0.623	0.438	0.17	0.21**	0.00	0.167	0.211
	(0.62)	(0.38)	(1.43)			(0.11)	(0.09)	(0.10)		
14	-0.49**	-0.31**	-0.67	0.477	0.351	0.20**	0.24***	0.02	0.112	0.128
	(0.21)	(0.13)	(0.58)			(0.10)	(0.09)	(0.10)		
15	-0.81	-0.38	-1.00	0.503	0.360	0.25*	0.33***	0.04	0.042	0.078
	(0.63)	(0.42)	(1.02)			(0.13)	(0.11)	(0.09)		
16	-1.00	-0.86	-0.75	0.768	0.839	0.26*	0.36***	0.00	0.033	0.072
	(1.24)	(2.68)	(0.53)			(0.15)	(0.13)	(0.10)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E2: Consumption mult. under Expansion and Recession following a FC ann. (OECD Recession dates).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.16	-0.79	-0.05	0.232	0.343	-0.02	0.00	-0.07	0.732	0.739
	(0.10)	(0.64)	(0.04)			(0.12)	(0.18)	(0.11)		
2	-0.26***	-0.23	-0.22***	0.865	0.865	-0.01	0.02	-0.04	0.797	0.793
	(0.10)	(0.14)	(0.08)			(0.08)	(0.17)	(0.10)		
3	-0.24***	-0.18**	-0.26***	0.503	0.539	-0.04	-0.02	-0.05	0.966	0.966
	(0.06)	(0.08)	(0.07)			(0.08)	(0.15)	(0.08)		
4	-0.35***	-0.36**	-0.32***	0.731	0.721	-0.04	-0.00	-0.07	0.647	0.660
	(0.09)	(0.17)	(0.08)			(0.07)	(0.08)	(0.09)		
5	-0.37***	-0.30***	-0.36***	0.798	0.805	-0.02	0.02	-0.06	0.556	0.585
	(0.07)	(0.11)	(0.08)			(0.07)	(0.08)	(0.08)		
6	-0.4***5	-0.34**	-0.50**	0.636	0.623	-0.02	0.04	-0.08	0.441	0.475
	(0.13)	(0.14)	(0.20)			(0.08)	(0.10)	(0.08)		
7	-0.58**	-0.43***	-0.73*	0.556	0.473	0.00	0.04	-0.09	0.246	0.298
	(0.25)	(0.16)	(0.43)			(0.08)	(0.08)	(0.09)		
8	-0.74**	-0.37***	-1.97	0.498	0.139	0.00	0.07	-0.08	0.201	0.234
	(0.33)	(0.09)	(2.30)			(0.07)	(0.08)	(0.06)		
9	-0.64*	-0.38***	-1.41	0.606	0.212	0.01	0.06	-0.06	0.337	0.361
	(0.36)	(0.14)	(1.99)			(0.09)	(0.10)	(0.06)		
10	-0.99	-0.63***	-1.78	0.678	0.353	-0.01	0.01	-0.06	0.497	0.502
	(0.77)	(0.24)	(2.89)			(0.09)	(0.09)	(0.07)		
11	-0.82	-0.26***	-12.47	0.899	0.096	0.01	0.04	-0.05	0.432	0.437
	(0.57)	(0.07)	(100.79)			(0.09)	(0.10)	(0.06)		
12	-0.67*	-0.28**	-1.77	0.462	0.120	-0.00	0.04	-0.07	0.401	0.411
	(0.37)	(0.11)	(2.02)			(0.10)	(0.10)	(0.08)		
13	-1.05	-0.49	-1.72	0.543	0.258	0.08	0.18	-0.09	0.090	0.139
	(0.86)	(0.40)	(2.10)			(0.13)	(0.12)	(0.09)		
14	-0.66**	-0.30***	-1.05	0.346	0.197	0.11	0.25**	-0.11	0.036	0.058
	(0.30)	(0.11)	(0.83)			(0.12)	(0.12)	(0.11)		
15	-0.99	-0.17	-1.43	0.323	0.147	0.16	0.38***	-0.08	0.012	0.050
	(0.78)	(0.34)	(1.34)			(0.15)	(0.15)	(0.09)		
16	-1.19	-0.50	-0.97	0.369	0.679	0.15	0.38***	-0.11	0.010	0.052
	(1.53)	(1.77)	(0.65)			(0.15)	(0.14)	(0.11)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E3: Investment mult. under Expansion and Recession following a FC ann. (OECD Recession dates).

	Tax-Based Fiscal Consolidation					Expenditure-Based Fiscal Consolidation				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-1.39*	-3.02	-1.01	0.250	0.396	-0.53	0.74	-1.15	0.387	0.392
	(0.83)	(2.32)	(0.67)			(0.76)	(1.52)	(1.13)		
2	-1.13**	-0.80	-0.92	0.945	0.945	0.86	3.62	-0.60	0.214	0.170
	(0.53)	(0.74)	(0.58)			(0.78)	(3.08)	(1.08)		
3	-0.99***	-0.76*	-0.94**	0.805	0.807	0.42	2.10	-0.47	0.133	0.077
	(0.33)	(0.40)	(0.46)			(0.68)	(1.55)	(0.69)		
4	-1.45***	-1.46*	-1.24**	0.765	0.760	1.27	2.15	0.07	0.173	0.143
	(0.48)	(0.77)	(0.51)			(0.87)	(1.42)	(0.72)		
5	-1.47***	-1.33**	-1.25**	0.842	0.841	0.98	2.32	-0.52	0.125	0.102
	(0.39)	(0.54)	(0.55)			(0.81)	(1.65)	(0.71)		
6	-2.46***	-1.70**	-2.77*	0.640	0.612	1.67	3.22	-0.53	0.198	0.125
	(0.88)	(0.84)	(1.62)			(1.40)	(2.75)	(0.73)		
7	-3.04**	-1.61**	-4.59	0.384	0.220	0.68	1.25**	-0.68	0.063	0.140
	(1.39)	(0.74)	(3.12)			(0.48)	(0.50)	(0.85)		
8	-4.05*	-1.71**	-12.17	0.496	0.145	0.34	0.83	-0.53	0.167	0.199
	(2.08)	(0.70)	(15.00)			(0.52)	(0.58)	(0.72)		
9	-3.46*	-1.84**	-7.61	0.601	0.258	0.94	1.30**	-0.22	0.131	0.175
	(1.93)	(0.77)	(10.80)			(0.75)	(0.62)	(0.91)		
10	-5.54	-3.24***	-10.10	0.674	0.374	0.84	0.37	0.89	0.709	0.716
	(4.18)	(1.21)	(16.77)			(0.54)	(0.38)	(1.35)		
11	-4.53*	-1.56***	-60.25	0.899	0.147	1.05*	1.32**	-0.21	0.115	0.174
	(2.74)	(0.51)	(486.70)			(0.60)	(0.54)	(0.75)		
12	-3.82**	-1.57***	-9.45	0.477	0.152	1.00*	1.20**	-0.25	0.154	0.217
	(1.90)	(0.42)	(11.08)			(0.55)	(0.51)	(0.71)		
13	-5.69	-3.24	-7.20	0.630	0.457	1.00	1.07**	0.11	0.435	0.462
	(4.54)	(2.25)	(8.99)			(0.73)	(0.48)	(0.99)		
14	-4.06***	-2.54***	-5.06	0.482	0.405	0.71	1.13**	-0.26	0.165	0.195
	(1.36)	(0.77)	(4.02)			(0.73)	(0.48)	(0.83)		
15	-5.26	-3.03*	-5.25	0.563	0.447	1.13	1.62**	0.02	0.118	0.154
	(3.69)	(1.84)	(5.12)			(0.95)	(0.71)	(0.83)		
16	-6.71	-6.31	-4.10	0.871	0.894	1.70	1.90**	0.30	0.166	0.195
	(9.08)	(17.04)	(3.16)			(1.05)	(0.74)	(0.93)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E4: Output multiplier under Expansion and Recession following a FC ann. (Unemployment).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.15***	-0.13	-0.10***	0.785	0.779	0.08	0.13	-0.05	0.548	0.504
	(0.06)	(0.13)	(0.03)			(0.11)	(0.21)	(0.15)		
2	-0.17***	-0.17	-0.14**	0.806	0.803	0.12	0.13	0.02	0.511	0.500
	(0.06)	(0.11)	(0.07)			(0.10)	(0.12)	(0.15)		
3	-0.16***	-0.10	-0.16***	0.689	0.695	0.13	0.08	0.06	0.861	0.858
	(0.04)	(0.07)	(0.05)			(0.09)	(0.13)	(0.10)		
4	-0.21***	-0.09	-0.27**	0.351	0.372	0.14*	0.08	0.07	0.937	0.937
	(0.06)	(0.08)	(0.11)			(0.08)	(0.10)	(0.08)		
5	-0.25***	-0.12	-0.31***	0.342	0.419	0.14	0.06	0.05	0.941	0.941
	(0.06)	(0.10)	(0.08)			(0.09)	(0.09)	(0.10)		
6	-0.31***	-0.10	-0.38***	0.386	0.453	0.16*	0.08	0.06	0.951	0.951
	(0.08)	(0.14)	(0.11)			(0.08)	(0.07)	(0.10)		
7	-0.38**	-0.12	-0.42**	0.572	0.599	0.17**	0.10	0.06	0.772	0.767
	(0.15)	(0.22)	(0.18)			(0.07)	(0.08)	(0.08)		
8	-0.47**	-0.11	-0.61	0.477	0.535	0.13**	0.08	0.03	0.707	0.704
	(0.22)	(0.23)	(0.38)			(0.06)	(0.07)	(0.07)		
9	-0.44*	-0.11	-0.51	0.594	0.638	0.13	0.08	0.00	0.579	0.569
	(0.24)	(0.25)	(0.34)			(0.08)	(0.07)	(0.10)		
10	-0.69	-0.09	-0.93	0.529	0.586	0.11	0.10	0.00	0.393	0.421
	(0.51)	(0.31)	(0.68)			(0.07)	(0.07)	(0.08)		
11	-0.54	-0.09	-0.63*	0.583	0.699	0.12	0.10	0.02	0.483	0.503
	(0.37)	(0.35)	(0.34)			(0.07)	(0.07)	(0.07)		
12	-0.46*	-0.10	-0.55**	0.527	0.633	0.13	0.15**	0.00	0.248	0.286
	(0.26)	(0.27)	(0.26)			(0.09)	(0.07)	(0.08)		
13	-0.78	-0.48	-0.63**	0.908	0.884	0.17	0.13**	0.04	0.583	0.580
	(0.62)	(1.50)	(0.25)			(0.11)	(0.07)	(0.10)		
14	-0.49**	-0.07	-0.58***	0.213	0.460	0.20**	0.12*	0.04	0.657	0.648
	(0.21)	(0.19)	(0.12)			(0.10)	(0.07)	(0.11)		
15	-0.81	-0.20	-0.95**	0.366	0.454	0.25**	0.09	0.11	0.842	0.845
	(0.63)	(0.25)	(0.39)			(0.13)	(0.07)	(0.13)		
16	-1.00	-0.87	-0.61***	0.858	0.685	0.26*	0.07	0.11	0.727	0.738
	(1.24)	(2.98)	(0.20)			(0.15)	(0.09)	(0.14)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E5: Consumption multiplier under Expansion and Recession following a FC ann. (Unemployment).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-0.16	-0.24	-0.07	0.370	0.321	-0.02	0.01	-0.07	0.824	0.818
	(0.10)	(0.20)	(0.07)			(0.12)	(0.28)	(0.13)		
2	-0.26***	-0.34	-0.18*	0.508	0.484	-0.01	-0.00	-0.05	0.777	0.775
	(0.10)	(0.22)	(0.09)			(0.08)	(0.12)	(0.12)		
3	-0.24***	-0.24**	-0.21**	0.740	0.745	-0.04	-0.03	-0.10	0.553	0.557
	(0.06)	(0.10)	(0.10)			(0.08)	(0.10)	(0.10)		
4	-0.35***	-0.26**	-0.40**	0.720	0.716	-0.04	-0.05	-0.10	0.565	0.568
	(0.09)	(0.12)	(0.16)			(0.07)	(0.09)	(0.10)		
5	-0.37***	-0.29**	-0.39***	0.824	0.826	-0.02	-0.06	-0.09	0.753	0.746
	(0.07)	(0.14)	(0.10)			(0.07)	(0.08)	(0.09)		
6	-0.45***	-0.40	-0.41**	0.871	0.870	-0.02	-0.06	-0.12	0.635	0.616
	(0.13)	(0.25)	(0.17)			(0.08)	(0.07)	(0.12)		
7	-0.58**	-0.48	-0.53*	0.942	0.941	0.00	-0.08	-0.08	0.978	0.978
	(0.25)	(0.39)	(0.28)			(0.08)	(0.08)	(0.10)		
8	-0.74**	-0.37	-0.84	0.703	0.719	0.00	-0.08	-0.07	0.988	0.988
	(0.33)	(0.40)	(0.55)			(0.07)	(0.07)	(0.08)		
9	-0.64*	-0.34	-0.67	0.812	0.822	0.01	-0.08	-0.10	0.832	0.830
	(0.36)	(0.45)	(0.49)			(0.09)	(0.07)	(0.13)		
10	-0.99	-0.35	-1.20	0.710	0.734	-0.01	-0.07	-0.09	0.858	0.859
	(0.77)	(0.63)	(0.95)			(0.09)	(0.08)	(0.10)		
11	-0.82	-0.31	-0.90*	0.753	0.801	0.01	-0.08	-0.06	0.975	0.975
	(0.57)	(0.68)	(0.48)			(0.09)	(0.08)	(0.10)		
12	-0.67*	-0.25	-0.77**	0.665	0.723	-0.00	-0.04	-0.11	0.639	0.645
	(0.37)	(0.45)	(0.35)			(0.10)	(0.08)	(0.11)		
13	-1.05	-0.98	-0.81**	0.862	0.797	0.08	-0.02	-0.02	0.966	0.966
	(0.86)	(2.90)	(0.33)			(0.13)	(0.08)	(0.13)		
14	-0.66**	-0.21	-0.73***	0.499	0.618	0.11	-0.01	-0.04	0.874	0.873
	(0.30)	(0.33)	(0.15)			(0.12)	(0.09)	(0.14)		
15	-0.99	-0.32	-1.09**	0.529	0.562	0.16	-0.05	0.05	0.684	0.693
	(0.78)	(0.35)	(0.48)			(0.15)	(0.10)	(0.16)		
16	-1.19	-1.24	-0.71***	0.848	0.647	0.15	-0.08	0.04	0.620	0.638
	(1.53)	(4.31)	(0.27)			(0.15)	(0.12)	(0.16)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.E6: Investment multiplier under Expansion and Recession following a FC ann. (Unemployment).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Expansion**	**Recession**	p-value	AR p-value	Linear	**Expansion**	**Recession**	p-value	AR p-value
1	-1.39*	-1.87*	-0.62	0.116	0.116	-0.53	-0.31	-1.13	0.609	0.655
	(0.83)	(1.08)	(0.65)			(0.76)	(0.58)	(1.28)		
2	-1.13**	-1.15**	-0.79	0.645	0.662	0.86	0.88	0.88	0.936	0.935
	(0.53)	(0.55)	(0.76)			(0.78)	(0.58)	(1.68)		
3	-0.99***	-1.15**	-0.57	0.454	0.477	0.42	0.51	0.22	0.797	0.795
	(0.33)	(0.48)	(0.53)			(0.68)	(0.39)	(1.31)		
4	-1.45***	-1.09*	-1.26	0.995	0.995	1.27	-0.36	2.41	0.111	0.165
	(0.48)	(0.60)	(0.78)			(0.87)	(0.53)	(1.70)		
5	-1.47***	-1.18**	-1.24*	0.941	0.940	0.98	-0.19	1.71	0.137	0.238
	(0.39)	(0.57)	(0.75)			(0.81)	(0.51)	(1.32)		
6	-2.46***	-1.95*	-2.21	0.953	0.953	1.67	-0.34	2.86	0.227	0.247
	(0.88)	(1.15)	(1.55)			(1.40)	(0.53)	(2.58)		
7	-3.04**	-2.57	-2.63	0.900	0.900	0.68	0.12	0.69	0.341	0.426
	(1.39)	(1.72)	(2.01)			(0.48)	(0.48)	(0.62)		
8	-4.05**	-2.36	-4.31	0.825	0.829	0.34	-0.41	0.37	0.293	0.370
	(2.08)	(2.36)	(4.04)			(0.52)	(0.53)	(0.63)		
9	-3.46*	-1.95	-3.65	0.843	0.847	0.94	-0.27	1.06	0.101	0.232
	(1.93)	(2.17)	(3.34)			(0.75)	(0.56)	(0.86)		
10	-5.54	-1.52	-7.45	0.644	0.665	0.84	-0.52	0.85	0.036	0.139
	(4.18)	(2.70)	(6.72)			(0.54)	(0.62)	(0.52)		
11	-4.53*	-1.21	-5.54	0.641	0.716	1.05*	-0.23	1.00	0.061	0.147
	(2.74)	(2.48)	(3.79)			(0.60)	(0.63)	(0.68)		
12	-3.82**	-1.23	-4.67	0.634	0.687	1.00*	0.37	0.67	0.596	0.605
	(1.90)	(2.17)	(2.89)			(0.55)	(0.65)	(0.66)		
13	-5.69	-2.84	-5.00*	0.934	0.924	1.00	0.83	0.34	0.746	0.749
	(4.54)	(7.99)	(2.66)			(0.73)	(0.95)	(0.75)		
14	-4.06***	-1.01	-5.05***	0.350	0.523	0.71	0.07	0.17	0.925	0.926
	(1.36)	(1.52)	(1.25)			(0.73)	(0.47)	(1.00)		
15	-5.26	-0.98	-7.71*	0.332	0.336	1.13	-0.66*	0.79	0.278	0.350
	(3.69)	(1.24)	(4.01)			(0.95)	(0.36)	(1.18)		
16	-6.71	-7.41	-4.15	0.840	0.505	1.70	0.04	1.13	0.428	0.471
	(9.08)	(24.98)	(2.76)			(1.05)	(0.72)	(1.26)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E7: Output multiplier under Unconstrained/Constrained MP following a Fiscal Consolidation announcement.

				Test for difference btw. multipliers	
h	Linear	Unconstr. MP	Constr. MP	p-value	AR p-value
1	0.08	0.11	0.07	0.895	0.897
	(0.11)	(0.26)	(0.11)		
2	0.12	0.22	0.09	0.706	0.703
	(0.10)	(0.30)	(0.09)		
3	0.13	0.21	0.09	0.629	0.646
	(0.09)	(0.18)	(0.09)		
4	0.14*	0.19*	0.08	0.523	0.560
	(0.08)	(0.10)	(0.08)		
5	0.14	0.20*	0.06	0.502	0.536
	(0.09)	(0.11)	(0.10)		
6	0.16*	0.26**	0.04	0.254	0.319
	(0.08)	(0.10)	(0.09)		
7	0.17**	0.28***	0.03	0.135	0.214
	(0.07)	(0.11)	(0.08)		
8	0.13**	0.29***	-0.03	0.034	0.111
	(0.06)	(0.11)	(0.07)		
9	0.13	0.31**	-0.11	0.015	0.070
	(0.08)	(0.13)	(0.09)		
10	0.11	0.28**	-0.09	0.010	0.070
	(0.07)	(0.12)	(0.07)		
11	0.12	0.33**	-0.08	0.005	0.045
	(0.07)	(0.14)	(0.06)		
12	0.13	0.38***	-0.11	0.001	0.034
	(0.09)	(0.15)	(0.08)		
13	0.17	0.40***	-0.16	0.000	0.035
	(0.11)	(0.14)	(0.11)		
14	0.20**	0.41***	-0.16	0.000	0.039
	(0.10)	(0.13)	(0.14)		
15	0.25*	0.42***	-0.19	0.000	0.074
	(0.13)	(0.13)	(0.18)		
16	0.26*	0.44***	-0.27	0.000	0.117
	(0.15)	(0.15)	(0.21)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. Columns 1, 2, 3 report multipliers under the linear case and the two states: *unconstrained* and *constrained monetary policy*; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$.

Table 4.E8: Consumption multiplier under Unconstrained/Constrained MP following a Fiscal Consolidation announcement.

				Test for difference btw. multipliers	
h	Linear	Unconstr. MP	Constr. MP	p-value	AR p-value
1	-0.02	0.36**	-0.17***	0.006	0.057
	(0.12)	(0.16)	(0.07)		
2	-0.01	0.17	-0.09	0.206	0.261
	(0.08)	(0.16)	(0.08)		
3	-0.04	0.11	-0.14	0.072	0.176
	(0.08)	(0.08)	(0.10)		
4	-0.04	0.06	-0.14	0.091	0.186
	(0.07)	(0.07)	(0.10)		
5	-0.02	0.08	-0.18	0.076	0.158
	(0.07)	(0.07)	(0.13)		
6	-0.02	0.14	-0.22*	0.023	0.105
	(0.08)	(0.09)	(0.12)		
7	0.00	0.19*	-0.22*	0.008	0.072
	(0.08)	(0.10)	(0.12)		
8	0.00	0.23**	-0.21**	0.000	0.045
	(0.07)	(0.09)	(0.09)		
9	0.01	0.30***	-0.33**	0.000	0.033
	(0.09)	(0.11)	(0.14)		
10	-0.01	0.27**	-0.29***	0.000	0.028
	(0.09)	(0.13)	(0.09)		
11	0.01	0.36**	-0.26***	0.000	0.013
	(0.09)	(0.16)	(0.07)		
12	-0.00	0.36**	-0.31***	0.000	0.016
	(0.10)	(0.16)	(0.10)		
13	0.08	0.39**	-0.32***	0.000	0.016
	(0.13)	(0.16)	(0.11)		
14	0.11	0.38***	-0.32*	0.000	0.024
	(0.12)	(0.15)	(0.18)		
15	0.16	0.41***	-0.38*	0.000	0.043
	(0.15)	(0.15)	(0.20)		
16	0.15	0.41***	-0.47**	0.000	0.076
	(0.15)	(0.14)	(0.23)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. Columns 1, 2, 3 report multipliers under the linear case and the two states: *unconstrained* and *constrained monetary policy*; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$.

Table 4.E9: Investment multiplier under Unconstrained/Constrained MP following a Fiscal Consolidation announcement.

				Test for difference btw. multipliers	
h	Linear	Unconstr. MP	Constr. MP	p-value	AR p-value
1	-0.53	-0.38	-1.00	0.826	0.825
	(0.76)	(2.74)	(0.72)		
2	0.86	0.11	0.60	0.798	0.807
	(0.78)	(2.08)	(0.50)		
3	0.42	0.61	-0.22	0.630	0.636
	(0.68)	(1.67)	(0.57)		
4	1.27	1.01	0.69	0.906	0.907
	(0.87)	(1.07)	(0.93)		
5	0.98	0.53	0.50	0.917	0.917
	(0.81)	(0.97)	(0.99)		
6	1.67	0.59	1.72	0.539	0.555
	(1.40)	(0.91)	(2.02)		
7	0.68	0.98	-0.48	0.246	0.307
	(0.48)	(0.91)	(0.75)		
8	0.34	1.13	-0.95*	0.027	0.120
	(0.52)	(0.98)	(0.49)		
9	0.94	1.13	-0.27	0.393	0.388
	(0.75)	(0.86)	(1.10)		
10	0.84	1.15	-0.14	0.220	0.291
	(0.54)	(0.86)	(0.65)		
11	1.05*	1.87**	-0.24	0.094	0.174
	(0.60)	(0.89)	(0.88)		
12	1.00*	2.21**	-0.68	0.012	0.074
	(0.55)	(0.91)	(0.86)		
13	1.00	1.95**	-0.87	0.023	0.100
	(0.73)	(0.80)	(0.98)		
14	0.71	2.19**	-1.91	0.001	0.042
	(0.73)	(0.88)	(1.29)		
15	1.13	2.27**	-1.69	0.005	0.071
	(0.95)	(0.96)	(1.50)		
16	1.70	2.71***	-1.59	0.048	0.123
	(1.05)	(1.02)	(1.99)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. Columns 1, 2, 3 report multipliers under the linear case and the two states: *unconstrained* and *constrained monetary policy*; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$.

Table 4.E10: Output multiplier under Large and Tight FS following a FC ann. (*de facto* fiscal space – FS2).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.15***	-0.06	-0.27	0.319	0.303	0.08	0.18	0.05	0.442	0.476
	(0.06)	(0.07)	(0.17)			(0.11)	(0.12)	(0.09)		
2	-0.17***	0.03	-0.33**	0.063	0.124	0.12	0.19*	0.08	0.418	0.383
	(0.06)	(0.13)	(0.14)			(0.10)	(0.10)	(0.08)		
3	-0.16***	0.02	-0.32***	0.001	0.009	0.13	0.16***	0.08	0.569	0.575
	(0.04)	(0.07)	(0.08)			(0.09)	(0.04)	(0.08)		
4	-0.21***	0.01	-0.39***	0.011	0.013	0.14	0.20***	0.10	0.524	0.513
	(0.06)	(0.10)	(0.15)			(0.08)	(0.06)	(0.09)		
5	-0.25***	0.04	-0.53**	0.019	0.010	0.14	0.23***	0.07	0.301	0.309
	(0.06)	(0.09)	(0.24)			(0.09)	(0.06)	(0.10)		
6	-0.31***	0.05	-0.56**	0.035	0.051	0.16*	0.21***	0.09	0.474	0.481
	(0.08)	(0.13)	(0.28)			(0.08)	(0.05)	(0.09)		
7	-0.38**	0.07	-0.67*	0.065	0.049	0.17**	0.26***	0.11	0.396	0.349
	(0.15)	(0.14)	(0.40)			(0.07)	(0.08)	(0.07)		
8	-0.47**	0.10	-0.82	0.219	0.038	0.13**	0.22***	0.07	0.214	0.210
	(0.22)	(0.17)	(0.79)			(0.06)	(0.06)	(0.06)		
9	-0.44*	0.14	-0.62	0.122	0.058	0.13	0.23***	0.04	0.217	0.251
	(0.24)	(0.17)	(0.50)			(0.08)	(0.07)	(0.07)		
10	-0.69	0.16	-1.37	0.434	0.057	0.11	0.28**	0.02	0.088	0.109
	(0.51)	(0.27)	(2.14)			(0.07)	(0.11)	(0.06)		
11	-0.54	0.32	-0.73	0.176	0.084	0.12	0.29**	0.02	0.058	0.087
	(0.37)	(0.34)	(0.83)			(0.07)	(0.12)	(0.04)		
12	-0.46*	0.32	-0.57	0.067	0.157	0.13	0.35**	0.01	0.083	0.095
	(0.26)	(0.35)	(0.43)			(0.09)	(0.16)	(0.05)		
13	-0.78	0.30	-1.64	0.289	0.020	0.17	0.35***	0.01	0.072	0.137
	(0.62)	(0.40)	(2.21)			(0.11)	(0.12)	(0.09)		
14	-0.49**	0.25	-0.65	0.098	0.119	0.20**	0.42***	0.03	0.091	0.122
	(0.21)	(0.39)	(0.47)			(0.10)	(0.15)	(0.08)		
15	-0.81	0.24	-2.19	0.595	0.025	0.25*	0.37***	0.07	0.283	0.321
	(0.63)	(0.46)	(5.03)			(0.13)	(0.13)	(0.12)		
16	-1.00	0.26	-6.04	0.910	0.224	0.26*	0.35***	0.05	0.389	0.429
	(1.24)	(0.57)	(55.21)			(0.15)	(0.13)	(0.14)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

Table 4.E11: Consumption multiplier under Large and Tight FS following a FC ann. (*de facto* fiscal space – FS2).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-0.16	-0.04	-0.21	0.582	0.562	-0.02	0.11	-0.05	0.285	0.465
	(0.10)	(0.16)	(0.20)			(0.12)	(0.08)	(0.12)		
2	-0.26***	-0.21	-0.32***	0.815	0.816	-0.01	0.01	0.01	0.997	0.997
	(0.10)	(0.27)	(0.11)			(0.08)	(0.06)	(0.09)		
3	-0.24***	-0.09	-0.34***	0.121	0.110	-0.04	0.07**	-0.06	0.325	0.335
	(0.06)	(0.10)	(0.12)			(0.08)	(0.04)	(0.10)		
4	-0.35***	-0.14	-0.51***	0.045	0.040	-0.04	0.12**	-0.06	0.216	0.187
	(0.09)	(0.13)	(0.17)			(0.07)	(0.05)	(0.10)		
5	-0.37***	-0.06	-0.58**	0.032	0.017	-0.02	0.17***	-0.05	0.148	0.131
	(0.07)	(0.11)	(0.23)			(0.07)	(0.06)	(0.10)		
6	-0.45***	-0.14	-0.59**	0.211	0.150	-0.02	0.10**	-0.05	0.306	0.285
	(0.13)	(0.17)	(0.30)			(0.08)	(0.04)	(0.10)		
7	-0.58**	-0.07	-0.83	0.148	0.074	0.00	0.16**	-0.03	0.221	0.183
	(0.25)	(0.21)	(0.51)			(0.08)	(0.06)	(0.09)		
8	-0.74**	-0.05	-1.02	0.264	0.050	0.00	0.13***	-0.05	0.110	0.119
	(0.33)	(0.23)	(0.88)			(0.07)	(0.04)	(0.06)		
9	-0.64*	0.02	-0.74	0.180	0.073	0.01	0.16***	-0.05	0.183	0.201
	(0.36)	(0.24)	(0.57)			(0.09)	(0.06)	(0.09)		
10	-0.99	0.02	-1.59	0.476	0.093	-0.01	0.14**	-0.06	0.149	0.192
	(0.77)	(0.37)	(2.43)			(0.09)	(0.07)	(0.08)		
11	-0.82	0.20	-0.97	0.250	0.076	0.01	0.19***	-0.05	0.047	0.107
	(0.57)	(0.43)	(1.09)			(0.09)	(0.06)	(0.06)		
12	-0.67*	0.25	-0.68	0.101	0.155	-0.00	0.25***	-0.07	0.041	0.100
	(0.37)	(0.42)	(0.52)			(0.10)	(0.08)	(0.07)		
13	-1.05	0.26	-1.84	0.308	0.023	0.08	0.23***	-0.00	0.231	0.251
	(0.86)	(0.54)	(2.53)			(0.13)	(0.05)	(0.10)		
14	-0.66**	0.17	-0.71	0.177	0.091	0.11	0.32***	0.00	0.119	0.151
	(0.30)	(0.48)	(0.57)			(0.12)	(0.09)	(0.10)		
15	-0.99	0.19	-2.22	0.618	0.075	0.16	0.29***	0.04	0.364	0.352
	(0.78)	(0.56)	(5.24)			(0.15)	(0.05)	(0.14)		
16	-1.19	0.21	-6.42	0.911	0.209	0.15	0.29***	0.01	0.344	0.354
	(1.53)	(0.65)	(58.53)			(0.15)	(0.05)	(0.14)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show *p*-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). *p*-value$\leq 0.01^{***}$,$\leq 0.05^{**}$,$\leq 0.1^{*}$

Table 4.E12: Investment multiplier under Large and Tight FS following a FC ann. (*de facto* fiscal space – FS2).

	Tax-Based Fiscal Consolidation					**Expenditure-Based Fiscal Consolidation**				
h	Linear	**Large FS**	**Tight FS**	p-value	AR p-value	Linear	**Large FS**	**Tight FS**	p-value	AR p-value
1	-1.39*	-0.80	-1.90	0.632	0.596	-0.53	1.83	-1.05	0.073	0.025
	(0.83)	(1.19)	(1.71)			(0.76)	(1.29)	(0.75)		
2	-1.13**	-0.56	-1.63*	0.445	0.425	0.86	1.47*	0.45	0.339	0.301
	(0.53)	(0.86)	(0.92)			(0.78)	(0.81)	(0.77)		
3	-0.99***	0.01	-1.61**	0.107	0.060	0.42	1.25***	-0.18	0.111	0.175
	(0.33)	(0.59)	(0.81)			(0.68)	(0.30)	(0.71)		
4	-1.45***	-0.43	-1.98*	0.142	0.095	1.27	1.19***	1.12	0.929	0.929
	(0.48)	(0.65)	(1.03)			(0.87)	(0.30)	(1.19)		
5	-1.47***	0.09	-2.06*	0.054	0.020	0.98	1.46***	0.53	0.562	0.563
	(0.39)	(0.70)	(1.19)			(0.81)	(0.35)	(1.13)		
6	-2.46***	-0.18	-2.97	0.171	0.075	1.67	1.10***	1.60	0.691	0.689
	(0.88)	(0.78)	(1.88)			(1.40)	(0.29)	(1.96)		
7	-3.04**	-0.18	-3.87	0.173	0.064	0.68	1.47***	0.16	0.217	0.149
	(1.39)	(0.90)	(2.61)			(0.48)	(0.43)	(0.68)		
8	-4.05*	0.62	-6.14	0.220	0.039	0.34	1.30***	-0.29	0.061	0.050
	(2.08)	(1.44)	(5.73)			(0.52)	(0.24)	(0.63)		
9	-3.46*	0.63	-4.24	0.105	0.046	0.94	1.44***	0.40	0.477	0.446
	(1.93)	(1.41)	(3.20)			(0.75)	(0.30)	(0.92)		
10	-5.54	0.71	-9.63	0.434	0.040	0.84	1.81***	0.36	0.093	0.070
	(4.18)	(1.62)	(14.58)			(0.54)	(0.52)	(0.60)		
11	-4.53*	2.24	-5.03	0.122	0.045	1.05*	2.06***	0.40	0.156	0.144
	(2.74)	(2.28)	(5.23)			(0.60)	(0.61)	(0.71)		
12	-3.82**	2.26	-3.85	0.050	0.090	1.00*	2.35***	0.24	0.079	0.067
	(1.90)	(2.49)	(2.55)			(0.55)	(0.70)	(0.61)		
13	-5.69	3.42	-11.64	0.222	0.029	1.00	2.15***	0.07	0.051	0.075
	(4.54)	(3.33)	(15.05)			(0.73)	(0.57)	(0.69)		
14	-4.06***	3.24	-5.05	0.120	0.081	0.71	2.47***	-0.23	0.036	0.050
	(1.36)	(4.17)	(3.40)			(0.73)	(0.65)	(0.64)		
15	-5.26	2.76	-9.39	0.476	0.109	1.13	2.22***	0.14	0.164	0.184
	(3.69)	(4.60)	(20.93)			(0.95)	(0.66)	(0.81)		
16	-6.71	3.80	-35.76	0.901	0.122	1.70	2.32***	0.52	0.402	0.411
	(9.08)	(7.04)	(323.79)			(1.05)	(0.58)	(0.92)		

The table shows cumulative multipliers. In column 1, projection horizons (quarters) are reported. The left (right) panel of the table presents the multipliers after a TB (EB) announcement. In each panel, columns 1, 2, 3 report multipliers under the linear case and the two states; columns 4 and 5 show p-values testing the difference between multipliers in the two states using respectively clustered S.E. and Anderson-Rubin (AR) confidence intervals (to account for the possibility that announcement series is a weak instrument). p-value$\leq 0.01^{***}, \leq 0.05^{**}, \leq 0.1^{*}$

4.F Exogeneity

Jordà and Taylor (2016) and De Cos and Moral-Benito (2016) argue that Devries et al. (2011) narrative measures for fiscal consolidations are relevant and valid instrument, but might suffer from lack of exogeneity if predetermined and exogenous controls are omitted from the specification.[53] De Cos and Moral-Benito (2016) find that output, budget balance variables, long-term interest rate, inflation and investment are relevant controls and should be included in the specification to avoid endogeneity issues. Indeed, as suggested by De Cos and Moral-Benito (2016), in the baseline specification, I include output, government revenues, government spending, private consumption and investment, real long-term interest rate and lags of the fiscal announcement measures. Moreover, Alesina et al. (2015) show that Devries et al. (2011) episodes appear to be predicted by past values of output growth only if the fiscal adjustment measures are transformed into mere dummy variables (0/1) as carried out in Jordà and Taylor (2016), while Devries et al. (2011) original episodes respect both *weak* and *strong* exogeneity.[54] Moreover, in more recent work by Alesina et al. (2018), the authors explain that the sources of identification of narrative adjustments are both the *timing* and the *size* of fiscal corrections. Thus, transforming the narrative measures into dummies disregards the central role of the size of the fiscal adjustment in the identification. Minor evidence of predictability of the *timing* of narrative adjustments does not translate into predictability of the *size*, which cannot be predicted as shown in Alesina et al. (2018).[55] Hence, in a VAR or in a LP framework, the estimates for the narrative measures' coefficients quantify the impact on output of the component of this measure that is orthogonal to lagged values of controls and, thus, the computed multipliers are not affected by any predictability. The most recent work on state-dependent effects of fiscal consolidations by Fotiou (2020) employs the narrative measures of Devries et al. (2011) and Alesina et al. (2015). Fotiou (2020) investigates exogeneity by using a simple Granger causality test by regressing the narrative adjustments on lagged values of both output growth and narrative measures themselves. The results of the Granger causality tests show that the past variables do not predict the narrative measures. In the same fashion, using the narrative adjustments developed by Beetsma et al. (2021), for each country and each type of measure, I regress the fiscal announcements on their own lagged values and lagged output growth. The results of the Granger causality test report that lagged values of announcements and output growth do not Granger-cause the announcements of Beetsma et al. (2021).

[53] Note that this potential endogeneity issue arises only for Devries et al. (2011) EB fiscal consolidations, while there is no evidence supporting lack of exogeneity for TB measures.

[54] By means of simple regressions, Alesina et al. (2015) show that the transformation into 0/1 dummies and the consequent loss of information open the door to correlation with lagged values of output growth.

[55] Quoting Alesina et al. (2018): *"It is useful to remember that fiscal policy is different from a medical treatment in which a group of patients are given the same dose of a medicine"*.

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