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ESSAYS IN INTERNATIONAL MACROECONOMICS AND FINANCE

GRAZIANO Marco

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FACULTÉ DES HAUTES ÉTUDES COMMERCIALES

DÉPARTEMENT D'ÉCONOMIE

ESSAYS IN INTERNATIONAL MACROECONOMICS AND FINANCE

THÈSE DE DOCTORAT

présentée à la

Faculté des Hautes Études Commerciales
de l'Université de Lausanne

pour l'obtention du grade de

Doctorat en économie

par

Marco GRAZIANO

Directeur de thèse
Prof. Philippe BACCHETTA

Jury
Prof. Valérie Chavez-Demoulin, présidente
Prof. Kenza Benhima, experte interne
Prof. Steven Ongena, expert externe
Prof. Cédric Tille, expert externe

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Marco GRAZIANO

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Lausanne, le 04.09.2024

Professeure Marianne Schmid Mast, Doyenne

Thesis committee

Philippe Bacchetta
Professor, HEC Lausanne
Thesis supervisor

Kenza Benhima
Professor, HEC Lausanne
Internal expert

Steven Ongena
Professor, University of Zurich
External expert

Cédric Tille
Professor, Graduate Institute for International and Development Studies
External expert

University of Lausanne
Faculty of Business and Economics

PhD in Economics

I hereby certify that I have examined the doctoral thesis of

Marco GRAZIANO

and have found it to meet the requirements for a doctoral thesis.
All revisions that I or committee members
made during the doctoral colloquium
have been addressed to my entire satisfaction.

Signature: _____ _ Date: _______26.7.2024_____

Prof. Philippe BACCHETTA
Thesis supervisor

University of Lausanne
Faculty of Business and Economics

PhD in Economics

I hereby certify that I have examined the doctoral thesis of

Marco GRAZIANO

and have found it to meet the requirements for a doctoral thesis.
All revisions that I or committee members
made during the doctoral colloquium
have been addressed to my entire satisfaction.

Signature: ______________________ Date: 19/07/2024

Prof. Kenza BENHIMA
Internal expert

University of Lausanne
Faculty of Business and Economics

PhD in Economics

I hereby certify that I have examined the doctoral thesis of

Marco GRAZIANO

and have found it to meet the requirements for a doctoral thesis.
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made during the doctoral colloquium
have been addressed to my entire satisfaction.

Signature: ______________________________ Date: _____17 July 2024______

Prof. Steven ONGENA
External expert

University of Lausanne
Faculty of Business and Economics

PhD in Economics

I hereby certify that I have examined the doctoral thesis of

Marco GRAZIANO

and have found it to meet the requirements for a doctoral thesis.
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have been addressed to my entire satisfaction.

Signature: Date: __July 17, 2024_

Prof. Cédric TILLE
External expert

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Introduction

The central topic of this thesis is the demand for and supply of safe assets, which play a critical role in the economy as store of value and mechanism of insurance. Many pivotal issues in international macroeconomics can be seen through the lens of the interplay between demand and supply of safe assets, especially those denominated in dollars, by far the most important international currency. The last two decades have been characterised by an increase in the global demand for safe assets, driven partially by the high growth of East Asian economies in what has been dubbed the 'saving glut' phenomenon by Bernanke (2005). In turn, this allowed the United States to sustain a negative current account position by being long risky foreign assets and short safe dollar assets, earning an extra return ('exorbitant privilege') and absorbing the global demand for safe assets due to their higher risk tolerance, essentially providing insurance and transferring wealth abroad during global crises ('exorbitant duty') (Gourinchas et al., 2010).

US Treasuries are the dollar-denominated safe asset of choice for the world economy, playing a key role in the portfolio of a diverse set of worldwide investors, and bucking the general trend towards home bias (Maggiori et al., 2020). The reasons behind their special position in the international financial system are complex. Firstly, they display particularly desirable safety and liquidity properties, which makes them attractive as a store of value and as a close substitute for money (Nagel, 2016), for example by easing transaction costs (Bansal and Coleman, 1996). The global prominence of Treasuries as safe assets is also tightly linked to the role of the dollar as an international invoicing currency in trade and finance (Gopinath et al., 2020), as agents seek to match their trade in debt contracts to the most liquid currency available (Coppola et al., 2023). The hedging properties of US government bonds also contribute to their special status: in line with the 'exorbitant duty' argument, recent evidence showed that US Treasuries tend to pay a particularly low yield when the covariance of Treasury returns with the aggregate stock market is low (Acharya and Laarits, 2023); and provide a safe haven in global recessions that coincide with dollar appreciation (Jiang et al., 2023a; Jiang, 2024).

Taken together, these features make Treasuries especially valuable for investors, who are willing to pay a premium, the convenience yield, compared to both comparable foreign government bonds (Du et al., 2018), and domestic securities with similar safety or liquidity attributes (Krishnamurthy and Vissing-Jorgensen, 2012). As a result, deviations in international asset pricing conditions such as the covered (CIP) and uncovered (UIP) interest parity conditions have arisen. Another fundamental factor contributing to the failure of these conditions is intermediation of foreign exchange flows by leverage-constrained financiers, limiting the extent of arbitrage in forex markets (Gabaix and Maggiori, 2015; Rime et al., 2022).

Regardless of their origin, wedges in international asset pricing conditions have momentous consequences for our understanding of exchange rates. Several recent papers (Engel and Wu, 2018; Itskhoki and Mukhin, 2021; Jiang et al., 2023b, 2024) show that such wedges can rationalise several long-standing exchange rate puzzles, including disconnect and excess volatility with respect to macroeconomic fundamentals (Meese and Rogoff, 1983), weak correlation with relative consumption (Backus and Smith, 1993), and the forward premium puzzle (Fama, 1984). The first three chapters of the thesis explore different aspects and ramifications of these deviations.

In the first chapter, *Convenience yields and the foreign demand for US Treasuries*, I analyse the implications of the special features of US Treasuries for the portfolio choice of foreign investors, whose large and stable demand underpins the ability of the US government to borrow at a competitive rate, hence supporting the burgeoning public debt (Jiang et al., 2019). I quantify the importance of convenience yields in explaining the cross-sector differences in the demand elasticity for Treasuries reported by previous studies (Eren et al., 2023; Fang et al., 2023). I model the portfolio choice between domestic (euro area) and foreign (US) government bonds in a market with two types of investors (banks and insurance companies) that derive utility directly from US Treasuries. The investors differ both in their risk aversion and in the weight on Treasuries in their preferences, allowing me to disentangle the effects of these two features. The model predicts that investors choose to hold US Treasuries even when they offer negative excess returns and positive correlation with income risk, in accordance with the data. Furthermore, thier portfolios display a lower sensitivity to the mean and variance of excess returns. Finally, in the presence of convenience yield investors, equilibrium excess returns are lower, and react more strongly to changes in debt supply.

I then estimate the model via two-stage least squares on data from European banks and insurers, showing that banks have a higher sensitivity to excess returns than insurances, in line with previous findings. I use the restrictions on first- and second-stage coefficients implied by the theoretical model to back out risk aversion and the parameters regulating investors' preferences for Treasuries. Counterfactuals on the structural parameters reveal that, absent convenience yields, the sensitivity to the mean and variance of excess returns would be three times as high for banks, and nine times as high for insurance companies. According to the model, cross-sector differences in sensitivty to excess returns can be explained almost entirely by heterogeneity in the weight that investors place on US Treasuries in their preferences. Convenience yields have a substantial impact on interest rates as well, as they reduce Treasury excess returns by 79 basis points on average. The results imply that the sustainability of US public debt is reliant on the special status of US Treasuries as the global safe asset, and in turn that it is vulnerable to the loss of this special status.

The second chapter of my thesis, *Treasury supply, convenience yields and exchange rates*, joint with Maxime Phillot, is concerned with exploring the interplay between convenience yields and fiscal policy from the point of view of the US government, the largest issuer of safe assets, and the consequences for the properties of exchange rates. Previous studies already highlighted the connection between the US fiscal cycle and the dollar through the channel of the global demand for safe assets (Jiang, 2021, 2022).

We argue that accounting for the endogenous optimal choice of government debt supply is crucial to correctly estimate the elasticity of convenience yields and exchange rates to debt supply. Previous studies offer only correlational evidence of the negative relationship between debt supply, Treasury premia and exchange rates, biased downward by the confusion of demand and supply shocks in the data. We demonstrate how this bias arises in a stylised model with convenience yield for Treasuries and endogenous government debt supply. The predicitons of the model are then testedby identifying Treasury supply shocks through a novel instrumental variable approach introduced in Phillot (2024). The drop in convenience yields and the exchange rate depreciation caused by an increase in debt supply is up to three times higher than previously estimated. Due to this high sensitivity to shocks in debt supply, this study provides an additional piece of evidence that the funding advantage enjoyed by the US government is more fragile than it might appear at first glance.

In the third chapter, *Mutual funds and safe government bonds: do returns matter?*, co-authored with Maurizio Habib, we focus on the investment of US mutual funds in both domestic and foreign safe government bonds. Non-bank financial institutions are a large and increasingly important player in sovereign debt markets, and studying the reaction of their portfolio shares to yields can help build a broader picture of the demand for safe assets in general. Since US mutual funds are based in the country that issues the global safe asset, they stand in a peculiar position to take advantage of deviations in pricing conditions related to the special status of US Treasuries. The deviations in UIP and CIP for the dollar differ in size and sign. Across currencies, aggregate hedging flows intermediated by leverage-constrained institutions can put upward or downward pressure on forward rates, depending on the net dollar exposure of their financial sectors. As a result, dollar CIP deviations are positive for some countries, like the yen, and negative for others, like the Australian dollar (Borio et al., 2016). On the other hand, a given currency might display opposite signs of deviations for covered and uncovered interest parity due to the combination of limited CIP arbitrage and safe asset status, which affects the UIP too (Bacchetta et al., 2023). As a consequence, dollar investors who observe these wedges have different incentives to respond to unhedged or hedged excess returns in safe assets across currencies.

We find that US mutual funds do actively rebalance their portfolio shares in response to unhedged excess returns for currencies that display a positive dollar UIP deviation, like the euro. On the other hand, we observe portfolio rebalancng mainly in response to hedged excess returns for currencies with a positive dollar CIP deviaiton, like the yen. Our results imply that capital flows between the United States and other major currency areas are sensitive not only to push factors, as shown in the search for yield literature (Ammer et al., 2018), but also to pull factors like excess returns, even in the usually yield-insensitive context of safe assets. Furthermore, the differences in sensitivity to excess returns across destination currencies suggest intriguingly that institutional investors actively exploit failures in arbitrage conditions, which might in turn substantially affect capital flows.

The fourth chapter of the thesis, *Spillovers of LSAPs through US Treasuries on foreign balance sheets*, with Marius Koechlin and Andreas Tischbirek, studies the consequences of the role of US Treasuries as global safe assets for the international transmission of the Federal Reserve's monetary policy. Much of the existing literature on the spillovers of monetary policy through the international role of the dollar focused on emerging markets, where firms and financial intermediaries tend to issue US dollar liabilities in excess of their dollar assets, making them vulnerable to a dollar appreciation (Eichengreen et al., 2007; Aoki et al., 2016; Akinci and Queraltó, 2019). Instead, we propose a novel transmission channel that operates through US Treasuries on banks' balance sheets, hence exposing them to dollar depreciations. This channel is more relevant for advanced economies, as their financial systems tend to hold dollar assets in excess of liabilities.

We use balance sheet data for European banks to quantify a novel channel of spillovers from the Federal Reserve's asset purchase programmes to the European economy. The Treasury holdings of European banks, concentrated in large institutions, expose them to two opposing valuation effects after a quantitative easing shock: the dollar price of long-maturity Treasuries increases, boosting their balance sheet; while the dollar depreciates against the euro, with a negative effect on net worth. We show that the exchange rate effect dominates, leading to a lower net worth and a substantial drop in credit, consistently with the predictions of a financial accelerator model *à la* Karadi and Nakov (2021). This chapter offers a new perspective on the spillovers of the Fed's quantitative easing (QE) to credit conditions in Europe. Contrary to the positive spillovers highlighted by previous studies, our results demonstrate that the dollar depreciation of US Treasuries on banks' balance sheets in response to QE leads to a contraction in lending. On the flipside, they suggest that the ongoing quantiative tightening policies might instead have a *positive* effect on the European economy through a dollar appreciation.

In conclusion, this thesis contributes to the international macroeconomics and finance literature on safe assets by providing evidence of the multi-faceted implications of safe assets, and especially the special circumstances surrounding their pricing, on capital flows and portfolio choice, as well as both fiscal and monetary policy.

Chapter 1

Convenience yields and the foreign demand for US Treasuries

This paper investigates the role of convenience yields in determining the yield sensitivity of the foreign demand for US Treasuries and the equilibrium interest rates of government bonds. A portfolio choice model featuring two sectors (banks and insurers) with heterogeneous risk aversion and preferences for holding US Treasuries shows that convenience yields reduce the sensitivity of portfolio shares to the mean and variance of excess returns. In equilibrium, excess returns for Treasuries are lower, and decrease more strongly in response to an increase in foreign debt supply. Structural parameters recovered from an estimation of the model on data from European banks and insurers reveal that convenience yields reduce the return sensitivity by 3 times for banks, and by 9 times for insurers. Convenience yields also explain virtually all of the difference in the sensitivity to excess returns across sectors, and they reduce Treasury excess returns by 79 basis points on average. The results imply that the sustainability of US public debt is reliant on the special status of US Treasuries as the global safe asset, and in turn that it is vulnerable to the loss of this special status.

1.1 Introduction

US Treasuries are the premier global safe asset, and their special role affords them a premium, or convenience yield, reflected in lower returns compared to other advanced economy sovereign debt (Du et al., 2018) and other dollar-denominated assets with similar safety and liquidity features (Longstaff, 2004; Krishnamurthy and Vissing-Jorgensen, 2012).

The convenience yield of US Treasuries is crucial for the sustainability of the burgeoning US government debt, as it allows the US government to borrow more cheaply than sovereigns with comparable credit rating; and it can explain the gap between the market value of US debt and projected government deficits (Jiang et al., 2019). This funding advantage is driven in large part by foreign investors, who are willing to accept a lower yield to meet their need for safe and liquid dollar-denominated assets, and display relatively yield-inelastic demand (Krishnamurthy and Vissing-Jorgensen, 2012; Jiang et al., 2022).

At the same time, there is substantial heterogeneity in the yield elasticity of safe asset demand across investor categories. Tabova and Warnock (2022) singles out the foreign official sector as particularly inelastic in its demand for Treasuries, while foreign private investors are especially sensitive to yields. Within private investors, Fang et al. (2023) finds that non-banks, including insurance companies and pension funds(ICPF), absorb a large amount of sovereign debt issuance, and display a particularly low yield elasticity for advanced-economy debt. In the context of corporate bonds, Bretscher et al. (2020) finds that insurers are inelastic to returns and prefer bonds by high-quality issuers.

On the contrary, the banking sector is generally more responsive to yields than insurers. (Timmer, 2018). This difference persists for US sovereign debt specifically: Eren et al. (2023) break down the ICPF sector into pension funds and insurers, finding a slightly larger yield elasticity than commercial banks for the former, but a much lower and not statistically significant response for the latter. Similarly, Koijen et al. (2021) report that, in the euro area, the yield elasticity of banks' demand for European government bonds is amongst the highest across sectors, while European ICPFs even have a *negative* elasticity. Since these sectors play a large role in absorbing new issuance of government debt (Fang et al., 2023), from the perspective of fiscal policy it is important to understand what drives the difference in their behaviour, and how changes in such features can affect the cost of government funding.

The literature explains the differences in demand elasticity mostly in terms of risk management practices (Eren et al., 2023), regulatory framework (Faia et al., 2022), or market-making versus speculative roles (Abbassi et al., 2016; Timmer, 2018). This paper quantifies the relative importance of risk aversion and special preferences for US Treasuries in explaining the cross-sectoral heterogeneity in the senstivity of demand, and the consequences of convenience yields on Treasury excess returns. I zoom in on the difference between insurers and banks because there is a well-established difference in their respective yield-sensitivity; and because ICPFs are likely more risk-averse due to their business model, so that the role of US Treasury preferences is not overstated by construction.

The theoretical framework consists of a simple mean-variance model of portfolio choice between US and domestic-currency government bonds in which investors have a preference for the special features of US Treasuries, modelled as an additional term in the investor's objective function following the approach of Krishnamurthy and Vissing-Jorgensen (2012), among others. Investors differ both in their risk aversion and in the degree of preference for Treasuries, so that the heterogeneity in yield sensitivity across sectors can be apportioned between these two features.

The model predicts that investors are less sensitive to the mean and the variance of excess returns on Treasuries than they otherwise would in the absence of convenience yields. Furthermore, they would be willing to hold a non-zero amount of US Treasuries even if they paid a lower return than domestic bonds, and did not provide a good hedge for income risk. This feature emerges uniquely from the convenience yield mechanism and cannot be explained by risk aversion: investors that value solely monetary payoffs hold assets only if they deliver an excess returns, or if they are a good hedge.

Equilibrium excess returns depend on the relative supply of US and domestic government debt, and they can be decomposed in a risk premium term and a convenience yield term. Preferences for Treasuries beyond their risk-return profile drive down excess returns and makes them more sensitive to changes in debt supply, in line with the safe asset supply channel of quantitative easing highlighted in Jiang et al. (2024) and Christensen et al. (2023).

The portfolio equations for the two sectors, jointly with the expression for equilibrium returns, imply restrictions that allow to calculate the structural parameters regulating risk aversion and the preference for Treasuries from the estimated regression coefficients of a linear version of the model. The structural parameters are key in disentangling and quantifying the role of risk aversion and convenience yields, as the predictions of the model on demand sensitivity and excess returns concern counterfactuals.

Understanding the rationale for differences in demand sensitivity is crucial to assess the capacity of markets to absorb additional US government debt. If the yield-insensitive demand by foreigners is due to risk aversion, the elasticity is heavily dependent on contingent market developments as encapsulated by the variance of returns. Therefore, events such as a temporary uncertainty on fiscal sustainability, due for example to negotiations in Congress over the debt ceiling, could jeopardise the ability of the US government to fund its debt cheaply. Conversely, convenience yields are tightly linked to the status of the US dollar as reserve currency, and of US Treasuries as global safe assets. These are much more persistent phenomena (Coppola et al., 2023), liable to evolve only in the face of extreme events such as a default (Choi et al., 2024) or major geopolitical upheaval (Eichengreen and Flandreau, 2009). Therefore, the more is low sensitivity driven by convenience yields, the more likely is it to be stable, and hence reliable from the point of view of the US government.

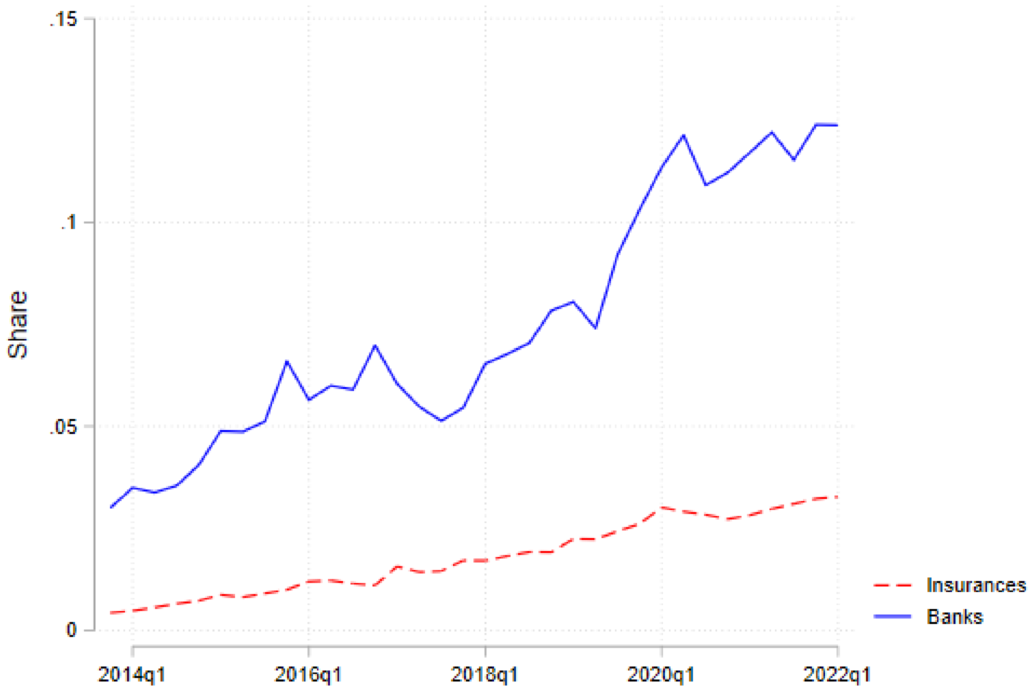
I estimate the model on data from the banking and insurance sector in the eurozone. The reason for this geographical focus is twofold. Firstly, the model implies that changes in debt supply are a valid instrument for excess returns in the portfolio equations using a two-stage least squares (2SLS) procedure. However, debt supply itself is likely endogenous to portfolio choice through general equilibrium effects, so it is necessary to isolate an exogenous component of debt supply to estimate the model. The Public Sector Purchase Programme (PSPP) by the European Central Bank (ECB) generates exogenous variation in the supply of sovereign debt issued by euro zone countries, solely as a function of the ECB's Capital Key and of the maturity structure of outstanding government bonds under the principle of market neutrality.[1] This structure provides an ideal laboratory to study responses to excess returns driven by exogenous changes in the relative supply of safe assets, thanks to an instrumental variables approach that matches the sets of equations derived from the theoretical model. The same identification strategy is exploited by Koijen et al. (2021) in the setting of demand for European government bonds.

Secondly, the very similar regulatory framework for banks and insurers in the realm of sovereign bonds and exposure to foreign exchange risk removes a potential alternative explanation for cross-sector differences in demand elasticity, thus sharpening the focus on differences in preferences.

[1]The Capital Key is the share of the ECB's capital held by each of the eurozone's national central banks.

The approach of estimating a mean-variance portfolio model through instrumental variables is related to the demand system asset pricing framework laid out in Koijen and Yogo (2019a) and adopted by a rapidly growing literature (Gabaix and Koijen, 2020; Bretscher et al., 2020; Haddad et al., 2021; Gabaix et al., 2022; Nenova, 2024). Differently from this methodology, I do not specify the full demand system but rather focus solely on the choice between US and domestic government bonds. This simpler approach allows to go beyond taking estimated elasticity as primitive parameters, but rather to back out directly the underlying preference parameters, and hence makes a step in the direction of understanding the nature of demand heterogeintiy at the core of the Koijen and Yogo (2019a) model. In this respect, the paper is related to the emerging literature that studies the theoretical foundations of demand-based asset pricing by endogenising heterogenoeus tastes (Fuchs et al., 2023).

Figure 1.1. US sovereign portfolio share for banks and insurers



Share of US Treasury holdings in a portfolio including US Treasuries and euro area government bonds for all banks (solid blue lines) and insurance companies and pension funds (dashed red line) domiciled in the euro area Source: European Central Bank Securities Holdings Statistics database.

Figure 1.1 plots the share of US Treasuries in a portfolio including euro area government bonds for banks (solid blue line) and ICPFs (dashed red line) domiciled in the euro area. It is evident at first glance that banks' share is much more volatile. While the comparison of unconditional volatility is not enough to draw any conclusions, it is certainly suggestive as to the plausible lower sensitivity of insurers' Treasury holdings to excess returns.

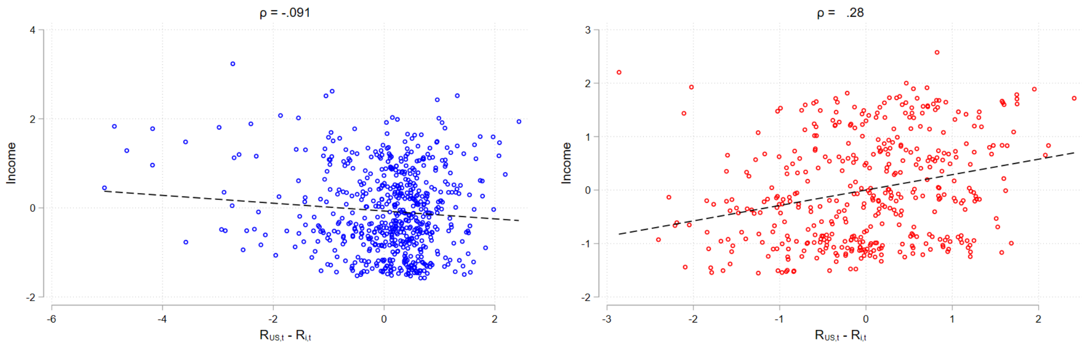
Figure 1.2 shows the correlation between the difference in the yield of Treasuries with respect to bonds issued by a given eurozone country, a rough proxy for excess returns, and the total balance sheet revenue of banks and ICPFs resident in the same country. The correlation is negative for banks and positive for insurance companies. Therefore, by this measure US government bonds are not a good hedge for the income risk of insurers. In the period between 2011 Q4 and 2023 Q3 over which the graph is constructed, Treasuries offer on average *negative* excess returns of about a quarter of a percentage point. [2] Therefore, insurance companies would have no incentive to include Treasuries in their portfolio under standard preferences that value assets solely for the balance of risks and rewards in monetary returns. This apparent contradiction can be resolved by the model presented in this paper: the presence of non-monetary payoffs motivates insurers to hold US Treasuries even in the face of a poor risk-return trade-off. Therefore, the observation in Acharya and Laarits (2023) that US Treasuries earn convenience yields because of their hedging properties againt stock market risk does not appear to extend to the case of income risk for European insurers.

Estimates of the portfolio equations via 2SLS reveal that banks increase their US Treasury portfolio share by 35.9 percentage points in response to a one percentage point increase in the excess returns of US Treasuries brought about by exogenous changes in the supply of eurozone governmnet debt. In contrast, ICPFs increase their portfolio share by only 6.24 percentage points.These findings are in line with existing evidence of lower sensitivity to excess returns of insurers' demand for government debt.

The structural parameters recovered from this estimation procedure imply that

[2] Here it is important to clarify the notion of excess returns used. In the model, US Treasury excess returns arise solely due to exogenous, stochastic fluctuations in exchange rates. Therefore, in this empirical proxy I account for the exchange rate expectations term in excess returns by adjusting for forward rates. This proxy is nonetheless imperfect due to the documented deviations in Covered Interest Parity stemming from frictions in FX markets (Borio et al., 2016; Rime et al., 2022). Other sources of variation that are disregarded in the model but likely affect the data, like sovereign credit risk, are allowed to influence this measure. In the empirical estimation, I will also correct yield differentials for credit risk to match the theoretical model.

Figure 1.2. Correlation between Treasury excess returns and sectoral income



(a) Banks

(b) Insurers

Correlation between excess returns on US Treasuries compared to country j's government bonds, and the income of banks (left-hand side panel) or insurance companies (right-hand side panel) domiciled in country j. The correlation is calculated over the country-quarter distribution on data from 2011 Q4 to 2023 Q2 for all euro area countries excluding Greece. Excess returns are averaged over the 1,2,3,5, and 10 year maturities and over quarters, and are adjusted for exchange rate forward premia as a market-implied measure of expected changes in the exchange rate of the euro *vis à vis* the dollar. Income is calculated as total income for banks, and total income from premia for insurers. Sources: Refinitiv Eikon, European Central Bank Consolidated Banking database, and EIOPA Insurance Statistics.

ICPFs are about 1.5 times more risk averse than banks, and that preferences for Treasuries have a 75% weight in both sectors' utility functions, albeit slightly larger for insurers. To understand the implications of these parameters for both portfolio choice and equilibrium interest rates, I perform four counterfactual experiments in the model.

First, I show that the slight difference in the weight of Treasury preferences across sectors translates to a very large effect on elasticities, do to the high estimated curvature of the Treasury preference component. Absent convenience yields, banks would be 3 times as sensitive to the mean and variance of excess returns, while insurers would be 9 times as sensitive. Then, I calculate that virtually all of the difference in sensitivity between banks and insurers is attributable to preferences for Treasuries.

I then analyse the effect of convenience yields on equilibrium excess returns. A time-series decomposition of the model-implied excess returns reveals that the convenience yield term is large, about 0.35 basis points on average, and much more volatile than the excess return component, which fluctuates around 11 basis points. Throughout the whole sample period, the convenience yield is large enough to turn the excess return negative, matching the empirical proxy for excess returns, as well as the low returns on Treasuries observed for foreign investors in particular (Jiang et al., 2022). Finally, I show that the model-implied excess returns are steeply increasing in the weight of the Treasury preference parameter for both investors, suggesting that the erosion of special status of US Treasuries can cost the government up to 79 basis points in higher interest rates on its debt from the loss of the convenience yield alone.

Therefore, the structural parameters recovered by taking the theoretical model to the data reveal that the convenience yield is a quantitatively important determinant of the portfolio choice for foreign banks and insurance companies, and it can explain almost all of the observed cross-section difference in the yield sensitivity of sovereign portfolio shares. Furthermore, the returns required to hold US Treasuries are substantially reduced by the presence of investors that value them because of their special status. On the other hand, Treasury returns would also rise very sharply should this special status be lost. Although limited to the context of banks and insurers in Europe, this result confirms that convenience yields are fundamental for the sustainability of US public debt, and at the same time casts a warning that the loss of credibility of US Treasuries as a global safe asset can be very costly from a fiscal perspective.

The remainder of the paper is organised as follows. Section 1.2 lays out a simple mean-variance portfolio problem with convenience yields, and derives propositions on portfolio shares and equilibrium excess returns. Section 1.3 estimates the coefficients of a linearised version of the model via 2SLS on data from eurozone banks and ICPFs. Section 1.4 recovers the structural parameters on investors' preferences from the estimated coefficients, and runs counterfactual experiments within the model. Section 1.5 concludes.

1.2 A model of portfolio choice with convenience yields

In this section, I build a simple model of sovereign portfolio choice with cross-sectoral heterogeneity in risk aversion and preference for US Treasuries, and derive the implications for asset pricing and the yield sensitivity of demand.

I model the static choice between euro-denominated government bonds issued by country j, offering a deterministic return, and US Treasuries, whose payoff is stochastic due to exogenous exchange rate fluctuations that are not modelled directly. Therefore, I implicitly assume no hedging of exchange rate risk. Importantly, the model also abstracts from sovereign credit risk, as the focus is on differences in returns that are motivated by convenience yields. Since credit risk is non-negligible for several countries in my sample, I control for it in the empirical analysis via credit default swaps (CDS) rates and country fixed effects. Investors also receive a stochastic income, which represents revenues from all other activities, for example loans for banks and premia from insurance. This assumption aims to capture succinctly other sources of income that are outside of the scope of this model of government bond portfolios, but nonetheless affect the investment choice by virtue of their correlation with sovereign returns.

I model the choice between US Treasuries and each country j's government bonds separately, to match the empirical setup that uses a panel of destination countries in the eurozone. In addition, note that the model abstracts from features like credit risk and home bias that would differentiate euro denominated assets. Therefore, all euro area sovereign bonds would be fungible and their optimal portfolio share would be indeterminate if I extended the analysis to the allocation of the entire portfolio jointly.

Two investor sectors populate the model: banks and insurance companies. They derive utility from wealth and additionally from holding US government bonds. This approach for modelling convenience yields is standard in the literature, and can be motivated by liquidity services(Nagel, 2016), for example by reducing transaction costs (Bansal and Coleman, 1996; Bansal et al., 2010), or by the desire to hold safe assets denominated in dollars, the dominant currency in both trade and financial markets (Maggiori et al., 2020; Coppola et al., 2023).[3]

1.2.1 Portfolio choice

Investors in sector k choose to allocate their initial wealth W_0^k between euro-denominated government bonds issued by country j, $b_{j,k}$, and US Treasuries $b_{US,k}$. By casting the choice in terms of shares $s_{j,k}$ and $s_{US,k}$ of fixed initial wealth $W_{0,k}$, the investor's problem is

[3] A non-exhaustive list of papers that adopt the bonds in the utility function approach includes Krishnamurthy and Vissing-Jorgensen (2012), Engel (2016), Engel and Wu (2018), Valchev (2020), Jiang et al. (2024), Nagel (2016), Jiang et al. (2021), and Bodenstein et al. (2023).

$$
\begin{aligned}
&\max_{s_{US,k}} \mathbb{E}[W_{0,k}] - 0.5\gamma_k \mathbb{V}[W_k] + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma} \\
\text{s.t.} \quad & W_k = W_{0,k}(R_j + (R_{US} - R_j)s_{US,k}) + Y_k \\
& s_{j,k} + s_{US,k} = 1,
\end{aligned}
$$

where $\gamma_k > 0$ is the investor's risk aversion parameter, W_k is their final wealth, Y_k is their stochastic income, $\psi_k > 0$ regulates the weight of the utility derived from holding Treasuries in the objective function, and $\sigma > 0$ regulates the curvature of the Treasury term in their preferences. Note that sectors are allowed to differ in their risk aversion and in the importance of US Treasuries in investors' preferences, but not in the curvature of the Treasury term in their utility. Thanks to this assumption, there are two structural parameters for each sector, which I then back out from the intercept and slope of each sector's estimated linearised portfolio equations. The common parameter σ is instead calculated from the slope of the equilibrium excess returns equation, which is also estimated in linear form. These structural parameters allow to perform counterfactual experiments on the relative importance of risk aversion and convenience yields for the sensitivity of portfolio shares and equilibrium excess returns.

The objective function is isomporphic to standard Markowitz (1952) mean-variance preferences, and can be derived from exponential utility over wealth and Treasury holdings, as shown in Appendix 1.B .

The first-order condition for $s_{US,k}$ is

$$
\mathbb{E}\left[R_{US} - R_j\right] - \gamma_k W_{0,k} \mathbb{V}\left[R_{US} - R_j\right] s_{US,k} - \gamma_k Cov\left[R_{US} - R_j, Y_k\right] + \psi_k \left(W_{0,1} s_{US,k}\right)^{-\sigma} = 0 \tag{1.1}
$$

This condition is analogous to that of the standard mean-variance portfolio problem, save for the additional term in ψ_k. Since investors also derive utility from holding Treasuries directly, this term implies that they potentially choose a positive portfolio share even if Treasuries offer a disadvantageous risk-return profile ($\mathbb{E}\left[R_{US} - R_j\right] < 0$ and $Cov\left[R_{US} - R_j, Y_k\right] > 0$).

To solve for the optimal portfolio share, I linearise the first-order condition around $s_{US,k} = \bar{s}$, $\mathbb{E}\left[R_{US} - R_j\right] = \bar{e}$, $\mathbb{V}\left[R_{US} - R_j\right] = \bar{v}$, and $Cov\left[R_{US} - R_j, Y_k\right] = \bar{c}$. In the special case of log utility for Treasuries with $\sigma \to 1$, an analytical solution for $s_{US,k}$ exists. In Appendix 1.C, I show how the results derived in this section extend to a nonlinear setting in the logarithmic case.

The optimal portfolio share for the linearised model is

$$s_{US,k} = \bar{s} - \gamma_k \frac{W_{0,k}\bar{s}\left(\mathbb{V}\left[R_{US} - R_j\right] - \bar{v}\right) + Cov\left[R_{US} - R_j, Y_k\right] - \bar{c}}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}} + \frac{1}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}}\left(\mathbb{E}\left[R_{US} - R_j\right] - \bar{e}\right) \tag{1.2}$$

Then, the derivative with respect to expected excess returns is

$$\frac{\partial s_{US,k}}{\partial \mathbb{E}\left[R_{US} - R_j\right]} = \frac{1}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}}. \tag{1.3}$$

Note that the derivative is higher for $\psi_k = 0$, where it collapses back to the standard mean-variance case. Therefore, convenience yields results in a *lower* sensitivity of investors' portfolio shares to excess returns compared to standard mean-variance preferences: as investors have a further motive to hold Treasuries beyond excess returns, they are less sensitive to changes in the latter. In general, both a higher risk aversion parameter γ_k and a higher Treasury preference parameter ψ_k would imply a lower sensitivity to excess returns, so that the simple comparison of elasticities across sector does not suffice to identify the effect of these two factors. In order to quantify their relative importance, it is then crucial to recover the structural parameters from the estimates of the intercept and slope of Equation 1.2 for both sectors.

Convenience yields also have implications for the sensitivity of Treasury demand to market voaltility. The derivative with respect to the variance of excess returns is

$$\frac{\partial s_{US,k}}{\partial \mathbb{V}\left[R_{US} - R_j\right]} = -\frac{\gamma_k W_{0,k}\bar{s}}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}}. \tag{1.4}$$

It is smaller in absolute value compared to the standard case without preference for US Treasuries. Therefore, investors rebalance away from Treasuries less intensely for any given increase in variance when they hold Treasuries for reasons other than their risk-return profile. To the extent that events that affect the variance of Treasury returns, such as monetary policy decisions or negotiations over the US debt limit, do not endanger the underlying special status of US Treasuries, represented by a lower ψ_k in the model, convenience yields also insulate Treasury demand from market volatility, resulting in a more stable source of funding for the government.

1.2.2 Equilibrium and pricing

The previous section analysed the response of investors' US Treasury portfolio share to changes in the mean and variance of excess returns, while remaining agnostic on the source of the latter. In this section, I derive excess returns in equilibrium as a function of the relative supply of euro area and US bonds. Thus, I obtain a theoretical counterpart for the empirical identificaiton strategy, which exploits exogenous changes in the supply of euro area government securities.

Equilibrium excess returns

The market clearing conditions for euro area and US government bonds, respectively, are

$$\sum_k b_{j,k} + b_{j,O} = B_j$$
$$\sum_k b_{US,k} + b_{US,O} = B_{US},$$

where B_j is the supply of euro-denominated bonds issued by country j, and likewise B_{US} is the supply of US Treasuries. The market clearing conditions also include holdings of country j and US government bonds held by other investors, respectively $b_{j,O}$ and $b_{US,O}$. While European banks and insurers are large players in the market for euro area sovereign bonds, their combined positions in US Treasuries add up to a maximum of 2.5 % of the total supply. Therefore, it is crucial to account for holdings of other investors to obtain realistic market-clearing conditions. These holdings are defined residually and maintained as exogenous throughout the model.

The equilibrium is a set of portfolio allocations $\{b_{j,k}, b_{US,k}\}$ for $k = \{B, I\}$ and Treasury excess returns $\mathbb{E}[R_{US} - R_j]$ such that the first-order conditions of banks and insurers hold, and the markets for euro area and US government bonds clear.

To derive equilibrium expected excess returns, sum the first-order conditions of both investors, defining for ease of exposition $\tau_k := \frac{1}{\gamma_k}$, the risk tolerance parameter of investor sector k.

$$\mathbb{E}[R_{US} - R_j] = \underbrace{\frac{\sum_k \left(\mathbb{V}[R_{US} - R_j] b_{US,k} + Cov\left[R_{US} - R_j, Y_k\right]\right)}{\sum_k \tau_k}}_{\text{Risk premium} \coloneqq RP} \\ \underbrace{- \frac{\sum_k \psi_k \tau_k b_{US,k}^{-\sigma}}{\sum_k \tau_k}}_{\text{Convenience yield} \coloneqq \phi} \tag{1.5}$$

Equilibrium excess returns of US Treasuries comprises two parts that can be interpreted intuitively. The first is a standard risk premium term: increasing in the volatility of excess returns and in the covariance between excess returns and income; and decreasing in the investors' risk tolerance. The second is specific to this model, and it can be interpreted as a convenience yield. The higher ψ_k, the weight of Treasuries in investors' preferences, the lower the equilibrium excess returns *ceteris paribus*. Since investors derive utility from holding Treasuries beyond their risk-return profile, they are willing to accept a lower monetary return, and this is is reflected in a convenience yield in equilibrium. This mechanism is well-studied in the literature on US Treasury pricing, since at least Krishnamurthy and Vissing-Jorgensen (2012), and it can explain the observed premium on US Treasuries (Du et al., 2018). Note that the convenience yield term is the average marginal benefit from Treasuries across investors, weighted by their risk tolerance. A higher risk tolerance for investor sector k implies a larger weight of their preferences on the equilibrium excess returns.

Much like the deviations from interest parity arising in open-economy macroeconomic models that incorporate convenience yields, the presence of Treasury holdings in the payoff function introduces a wedge in the pricing equation (Engel and Wu, 2018; Jiang et al., 2021; Valchev, 2020). However, in this model Treasuries carry exchange rate risk from the perspective of European investors, so the usual interest parity condition does not generally hold even in the absence of Treasuries in the utility function. As a consequence, the observed negative excess returns of Treasuries could be explained by a strongly negative correlation between excess returns combined with a relatively low excess return volatility. However, in the data we observe a low $Cov\left[R_{US} - R_j, Y_k\right]$ for banks and insurers alike. Therefore, from the perspective of the asset pricing equation implied by the model, the risk-return profile of US Treasuries for European investors is not likely to be a convincing explanation for negative excess returns.

The effects of debt supply

Equilibrium excess returns depend on debt supply through the risk premium component: the higher the amount of risky asset B_{US}, the higher the risk premium required for investors to absorb it. However, quantities enter equation 1.5 through the convenience yield term too. Due to the diminishing marginal utility of US government bonds, the weight of the convenience yield term is decreasing in the amount investors hold. This result is discussed in Jiang et al. (2024), which shows in a general equilibrium model how central bank quantitative easing affects asset prices also by altering the relative supply of safe assets with non-monetary payoffs.

To highlight this mechanism, re-write Equation 1.5 as a function of the supply of country j governemnt bonds B_j, using the market clearing conditions and the budget constraints of both agents.

$$\begin{aligned}\mathbb{E}[R_{US} - R_j] =& \frac{\sum_k \mathbb{V}[R_{US} - R_j] \left(\sum_k W_{0,k} - B_j + b_{j,O}\right) + \sum_k Cov\left[R_{US} - R_j, Y_k\right]}{\sum_k \tau_k} \\ &- \frac{\sum_k \psi_k \tau_k \left(W_{0,k} - B_j + b_{j,O} + \sum_{l \neq k} b_{j,l}\right)^{-\sigma}}{\sum_k \tau_k}.\end{aligned}$$

Excess returns then depend on the supply of euro-denominated government bonds B_j through both the risk-premium and the convenience yield terms. A change in B_j alters not only the relative amount of safe versus risky assets on the market, hence affecting the risk premium; but also the relative amount of US Treasuries that investors have to absorb, hence affecting the equilibrium convenience yield.

In the empirical estimation, I exploit exogenous changes in the supply of European government bonds due to the implementation structure of the PSPP programme by the European Central Bank. I use PSPP holdings as an instrument for changes in excess returns in a two-stage least squares setup. To understand the model-implied sign of the coefficient on PSPP amounts in the first-stage regression, I analyse the derivative of $\mathbb{E}[R_{US} - R]$ with respect to B_j. Furthermore, this derivative implies restrictions on ψ_k, γ_k and σ that I exploit to back out these structural parameters.

Proposition 1 (Reaction of excess returns to euro area debt supply)**.**

$$\frac{\partial \mathbb{E}[R_{US} - R_j]}{\partial B_j} = \frac{\mathbb{V}[R_{US} - R_j]\left(\frac{\partial b_{j,O}}{\partial B_j} - 1\right)}{\sum_k \tau_k \left(1 - \frac{\sigma \psi_k b_{US,k}^{-\sigma-1}}{\mathbb{V}[R_{US} - R_j]/\tau_k + \sigma \psi_k b_{US,k}^{-\sigma-1}}\right)} < 0. \tag{1.6}$$

Proof. In Appendix 1.D.1. □

Note that the derivative depends on $\frac{\partial b_{j,O}}{\partial B_j} \in [0,1]$, the absorption rate of government debt by other investors. It is an exogenous object in the model, and I estimate it empitically in the calibration of the model. Proposition 1 predicts a compression in US Treasuries excess returns in response to a higher supply of country j government bonds, through both the risk premium and the convenience yield components. An increase in the supply of euro government bonds reduces the relative amount of US Treasuries that investors have to absorb in equilibrium. Since Treasuries are a risky asset, the relative reduction in supply leads to a lower risk premium. At the same time, due the concavity of the Treasury component of investors' utility, they are now willing to accept a lower return *ceteris paribus* on Treasuries as their relative supply decreased. [4].

The presence of convenience yield preferences represented by ψ_k makes the fall in excess returns larger than it would be with $\psi_k = 0 \,\forall\, k$, as the second term in the denominator vanishes. Therefore, for any given reduction in the relative size of US government debt, the funding cost of US government debt falls more strongly if investors derive utility from holding Treasuries.

Linearisation

In order to estimate the reaction of equilibrium excess return to debt supply as the first stage of a two-stage least squares system with portfolio equations as second stage, I sart by linearising equation 1.5. The approximation points are $\mathbb{E}\left[R_{US} - R_j\right] = \bar{e}$, $Cov[Y_k, R_{US} - R_j] = \bar{c}$, $\mathbb{V}\left[R_{US} - R_j\right] = \bar{v}$, $B_j = \bar{B}_j$, $b_{j,O} = \bar{b}_{j,O}$ and $b_{US,k} = \bar{b}_{US,k} = W_{0,k}\bar{s}$. Note that I treat initial wealth $W_{0,k}$ as a fixed parameter, as the model makes predictions on portfolio shares rather than quantities. The linearised excess returns equation is

$$\mathbb{E}[R_{US} - R_j] = \bar{e} + \frac{1}{\sum_k \tau_k}\left(\left(\sum_k W_{0,k} - \bar{B}_j + \bar{b}_{j,O}\right)(\mathbb{V}[R_{US} - R_j] - \bar{v}) - \right.$$
$$\left. \bar{v}(B_j - \bar{B}_j + \bar{v}(b_{j,O} - \bar{b}_{j,O})) + \sum_k \left(Cov\left[R_{US} - R_j, Y_k\right] - \bar{c}\right) + \sigma \bar{b}_{US,k}^{-\sigma-1} \sum_k \tau_k \psi k (b_{US,k} - \bar{b}_{US,k})\right). \tag{1.7}$$

Following the same steps as in the proof of proposition 1 for the nonlinear case, the derivative with respect to euro-denominated debt supply B_j is

[4] I implicitly assume that changes in B_j have no effect on variances and covariances, which are treated as fixed parameters.

$$\frac{\partial \mathbb{E}[R_{US} - R_j]}{\partial B_j} = \frac{\bar{v}\left(\frac{\partial b_{j,O}}{\partial B_j} - 1\right)}{\sum_k \frac{1}{\gamma_k}\left(1 - \frac{\sigma \bar{b}_{US,k}^{-\sigma-1}\psi_k}{\gamma_k \bar{v} + \sigma\psi_k \bar{b}_{US,k}^{-\sigma-1}}\right)} \tag{1.8}$$

In the empirical model, this equation will provide restrictions on structural parameters that, together with those imposed by the intercept and slope of linearised portfolio equations, allow me to identify γ_k, ψ_k, and σ.

1.3 Estimation

The model makes three key predictions on the role of convenience yields for the sensitivity of portfolio shares to excess returns, and on equilibrium excess returns. First, investors increase their US Treasury sovereign portfolio shares in response to higher expected excess returns, but by a smaller amount than they would if they did not derive a convenience yield from holding Treasuries. Second, the reduction in US Treasury portfolio shares in response to an increase in the variance of returns is smaller in absolute value than it would be absent convenience yields. Third, an increase in the supply of euro area government bonds induces a decline in equilibrium excess returns for Treasuries. These predictions concern unobserved counterfactuals, so I need to estimate the structural parameters γ_k, ψ_k and σ to test it within the model. Furthermore, the structural parameters are informative on the extent to which cross-sector differences in sensitivity to excess returns are attributable to heterogeneity in risk aversion versus preference for US Treasuries. I recover the structural parameters by relating the theoretical restricions implied by the model to estimable objects in a two-stage least square system applied to data from euro area banks and insurers.

1.3.1 Mapping the model to the data

In order to map the regression coefficients of the 2SLS system to the structural parameters directly, I use the linearised portfolio equation 1.2 for each sector, and an equation that models excess returns directly as a linear function of B_j, using the restriction on the linearised derivative implied by equation 1.8. In addition to recovering structural parameters, the estimates of this system can also confirm the predictions of the model on the sign of portfolio share and excess returns derivatives.

Estimation equations

The estimation equations in the 2SLS model are the following, disregarding time subscripts for notational ease:

$$s_{US,k} = \alpha_k + \beta_k \mathbb{E}[R_{US} - R_j] + \varepsilon_{j,k} \text{ for } k = \{B, I\}, \tag{1.9}$$

$$\mathbb{E}[R_{US} - R_j] = \iota + \pi B_j + \nu_j. \tag{1.10}$$

The first equation corresponds to the linearised portfolio share 1.2, while the second one specifies excess returns directly as a linear function of country j debt supply. As a result, the coefficients of the empirical model as a function of structural parameters are

$$\alpha_k = \bar{s} - \gamma_k \frac{W_{0,k}\bar{s}\left(\mathbb{V}\left[R_{US} - R_j\right] - \bar{v}\right) + Cov\left[R_{US} - R_j, Y_k\right] - \bar{c}}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}}, \tag{1.11}$$

$$\beta_k = \frac{1}{\gamma_k W_{0,k}\bar{v} + \sigma\psi_k W_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}} > 0, \tag{1.12}$$

$$\pi = \frac{\bar{v}\left(\frac{\partial b_{j,O}}{\partial B_j} - 1\right)}{\sum_k \frac{1}{\gamma_k}\left(1 - \frac{\sigma \bar{b}_{US,k}^{-\sigma-1}\psi_k}{\gamma_k\bar{v} + \sigma\psi_k\bar{b}_{US,k}^{-\sigma-1}}\right)} < 0. \tag{1.13}$$

The linearised portfolio share for each sector has two empirical parameters, α_k and β_k, providing two equations per sector. The linearised excess returns provides only one additional equation, because only the parameter π, the derivative of excess returns to country j sovereign debt supply in a linear model, has a direct theoretical equivalent in equation 1.8. This feature arises because the empirical model expresses $\mathbb{E}[R_{US} - R_j]$ directly as a function of B_j, while in the theoretical model the derivative to B_j exploits the property that in equilibrium $s_{US,k}$ depends on B_j through excess returns. Therefore, there is no theoretical restriction on ι, so it is not used in the recovery of structural parameters. In total, the estimation of this system results in five equations, which allow to solve for the five structural parameters: γ_k and ψ_k for $k = \{B, I\}$, and σ.

I estimate β_B, β_I, and π via a 2SLS procedure with equation 1.10 as first stage, and equations 1.9 as second stage for both sectors. I identify β_B and β_I using exogenous changes in B_j as an instrument for $\mathbb{E}[R_{US} - R_j]$ in 1.9. The theoretical equivalent of parameter π in equation 1.10 is obtained by linearising $\mathbb{E}[R_{US} - R_j]$ as a function of B_j and $b_{j,US}$, and differentiating with respect to B_j while taking into account that $b_{j,US}$, and so $s_{j,US}$, is a function of $\mathbb{E}[R_{US} - R_j]$ only through

B_j. As a result, the linearised excess return implies that B_j satisfies the exclusion restriction as an instrument for $\mathbb{E}[R_{US} - R_j]$ in a model of s^B_{US}. The model also clearly implies that excess returns depend on B_j in equilibrium, so the requirement of instrument relevance is satisfied too.

1.3.2 Data

I estimate the equations outlined in the previous section on data from banks and ICPFs resident in the euro area, sourced from the publicly available Securities Holdings Statistics (SHS) dataset by the European Central Bank. The European setting provides an ideal context to study the role of convenience yields for the demand of US Treasuries by different types of investors because of several useful features.

First, the available data allows to observe the sovereign portfolios of different sectors operating under essentially the same regulatory regime. Existing literature on the demand for government bonds mainly relies on global data that does not provide a breakdown of government bond holdings by both country and sector. [5] While cross-sector differences in preferences are plausibly constant across jurisdictions, the regulatory regimes for banks relative to insurers might not be. Therefore, differences in the sensitivity of the demand for safe assets estimated in previous studies might confound heterogeneity in both preferences and regulation. In the European Union, banks and insurance companies are subject to very similar rules concerning investment in sovereign debt. Both Article 351 of the Capital Requirement Regulation (*EU Regulation No 575/2013*), applying to banks; and Article 180 of the Solvency II regulation(*EU Regulation No 35/2013*), applying to insurers, assign, with almost identical language, a zero weight for capital requirements to bonds that either have a high sovereign rating (like US Treasuries) or are denominated in euros. Therefore, from the regulatory standpoint, both banks and insurers are free to adjust the relative portfolio share of US and euro area sovereign bonds in response to returns without affecting their stock of risk-weighted assets relevant for capital buffers.[6] As a result, any observed discrepancy in the demand

[5] Some exceptions include Tabova and Warnock (2022), Eren et al. (2023)., and Fang et al. (2023)

[6] The only material difference in the regulatory treatment of euro area and US government bonds concerns exchange rate risk. Exposure to US Treasuries, if unhedged, counts against regulatory limits for foreign exchange risk exposure. Faia et al. (2022) uses this divergence, together with a different regime of capital requirements between insurance companies and mutual funds, to motivate differences in demand elasticity for these two sectors and deviations from covered interest parity. In this paper, I abstract from limits to foreign exchange exposure, which

sensitivity to convenience yields across these sectors is plausibly attributable to preferences. Thanks to this regulatory design, I can zoom in on differences in preferences only, using the estimates of structural parameters to disentangle the role of risk aversion and convenience yields.

Furthermore, the use of global data for different investor classes would introduce complications in mapping the model to the data. The theoretical model in this paper analyses the simple choice between US and domestic-currency government bonds. While euro-denominated sovereign bonds are a natural choice of domestic asset when focusing on the eurozone, this would not be the case when using global data.[7] Extending the exercise of this paper to global data would require either a more complex model of the whole sovereign portfolio with multiple assets, or the construction of a synthetic "domestic" asset for global foreign investors in US Treasuries from the data.

The rationale for comparing the sensitivity of demand of the banking and ICPF sector specifically also lies in the plausibly large difference in risk aversion due to their business models. As insurance companies are likely more risk-averse than banks in general, the structural model is less liable to overstate any difference in preferences for Treasuries due to the fact that it accounts only for these two parameters.

Finally, the peculiar structure of purchases under the PSPP quantiative easing policy by the ECB generates exogenous variation in the relative supply of government bonds across euro area countries. This policy intervention provides an ideal instrument to identify exogenous changes in US Treasury excess returns relative to euro area sovereign bonds. Therefore, I can estimate equations 1.9 and 1.10 with a panel 2SLS strategy.

1.3.3 Identification strategy

According to the model, changes in the supply of country j government bonds are a valid and relevant instrument for excess returns in estimating the portfolio equation 1.9, because the latter depends on B_j only through excess returns. However, the very stylised partial equilibrium model does not take into account that debt supply is likely endogenous to the portfolio choice of financial intermediaries

would affect both banks and insurers equally.

[7] I use data on the portfolios of the aggregate banking and insurance sector in the eurozone, so the domestic asset is defined at the currency rather than country level. This approach is consistent with the assumption of risk premia stemming only from exchange rate fluctuations in the model.

through general equilibrium effects. Even considering only changes in debt supply due to unconventional monetary policy is not enough to allay concerns of endogeneity, as these policies are adopted in response to highly endogenous macroeconomic conditions. This argument has particular bite for European banks, as they tend to load up on domestic government bonds in precisely the same turbulent times that motivate quantiative easing policies, either in a "gambling for resurrection" strategy (Acharya and Steffen, 2015) or due to "moral suasion" by their governments (De Marco and Macchiavelli, 2016; Ongena et al., 2019).

In order to obtain changes that are truly exogenous to investors' portfolio choice, I exploit the characteristics of the PSPP, implemented by the ECB starting in January 2015. The ECB bought government bonds issued by all countries with a credit rating of at least BBB-, and with maturities from 2 to 30 years.[8] The purchases are apportioned according to a scheme that aims at a market-neutral approach. They are proportional to each country's Capital Key, and they mirror as closely as possible the maturity structure of outstanding bonds.

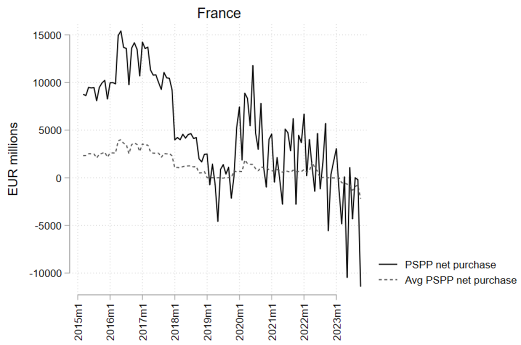
The Capital Key for each country is the equal-weighted average of its share of the euro zone's population and GDP. It is updated every five years, and whenever the membership of the European Union (EU) changes. In my sample, running from 2015 to 2022, the Capital Key changed twice: in 2019 due to a five-yearly update, and in 2020 due to the withdrawal of the United Kingdom from the EU. Since country size is plausibly independent of portfolio choice, and updates related to GDP are slow-moving, changes in Capital Key are likely exogenous. The other source of variation is the cross-country difference between the extant maturity structure of PSPP holdings, and that of the country's outstanding government bonds. Since this difference depends only on the governments' choice on the maturity of issuance and on the pre-existing PSPP term structure, it probably satisfies the exclusion restriction as well. Koijen et al. (2021) also uses PSPP purchases as an instrument for debt supply, but it relies on purchases predicted by the Capital Key rather than the actual amounts.

One step is missing to be able to claim PSPP holdings are a valid instrument for excess returns: while it can be argued that the *cross-sectional* variation in PSPP holdings is exogenous to investors' portfolio choice, this is not the case for the *time-series* dimension. Changes in total purchases over time track the overall size of the quantiative easing programme, which is obviously correlated to investors'

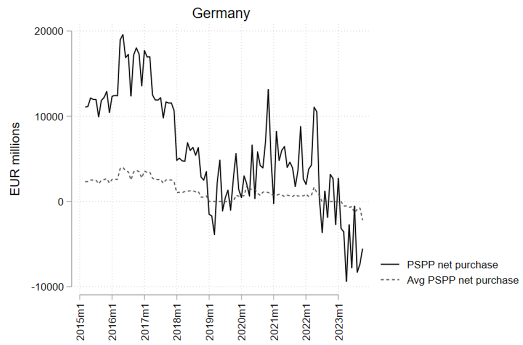
[8]The restriction on credit rating resulted in the exclusion of Greek bonds. In the context of this paper, the exclusion of Greece helps in ensuring that the empirical proxy for excess returns is driven by currency risk and not default risk, which reflects the assumption made in the model.

portfolios through macroeconomic fluctuations. To account for this issue, I use time fixed effects that soak up trends in the average amount of PSPP purchases across countries. Figure 1.3 illustrates the source of variation exploited for identification. The solid line represents purchases for each of the four largest euro area country, while the dashed line depicts average purchases across countries. By using time fixed effects, I rely only on the informtion contained in the differences between the solid and the dotted line for each country.

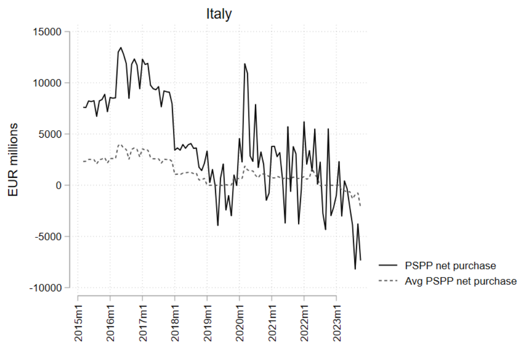
Figure 1.3. PSPP purchases



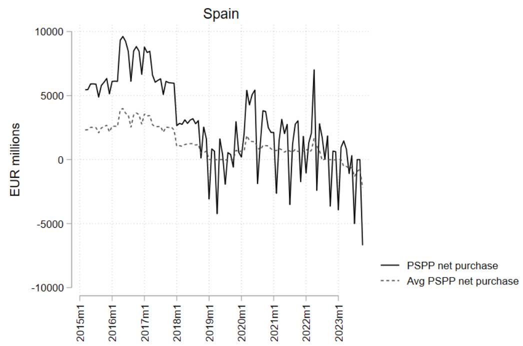
(a) France



(b) Germany



(c) Italy



(d) Spain

Monthly net purchases of sovereign debt under the Public Sector Purchase Programme, all maturities. The solid line depicts monthly net purchases for a country, the dashed line represents the cross-sectional average of monthly net purchases across all eligible countries over quarters. Source: European Central Bank

1.3.4 Estimation via two-stage least squares

In this section, I lay out the estimating equations for the empirical model and report the results. The starting point is the set of equations 1.9 and 1.10 derived from the linearised theoretical model. While they can be estimated directly as written above, I modify them to account for complications and nuances in the data that the model fails to capture. The baseline first-stage regression is

$$er_{j,t} = \iota_j + \iota_t - \pi PSPP_{j,t} + \lambda' \mathbf{V_{j,t}} + \kappa' \mathbf{W_{j,t}} + \nu_{j,t},$$

where $PSPP_{j,t}$ are PSPP holdings of country j government debt in quarter t, which proxy the exogenous component of B_j.[9] Note tht the coefficient π enters the equation with a minus sign because an increase in PSPP holdings corresponds to a *decrease* in the amount of country j's sovereign debt available to investors. The second-stage regression for $k = \{B, I\}$ is

$$s_{US,j,k,t} = \alpha_{k,j} + \alpha_{k,t} + \beta_k er_{j,t} + \delta_k' \mathbf{V_{j,t}} + \eta_k' \mathbf{W_{j,t}} + \varepsilon_{j,k,t}.$$

One difference from the theoretical model is due to the panel structure of the data. I rely on quarterly observations of sovereign holdings, so $s_{US,j,k,t}$ is the quarter t share of US Treasuries in a portfolio comprised of country j's government bond and US Treasuries for all euro area investors in sector k. Likewise, $er_{j,t}$ is a proxy for the excess returns of US Treasuries with respect to country j's government bonds, on average for quarter t. Expected excess returns in the model depend on a number of unobservable factors, such as investor's expectations on the future path of asset prices and exchange rates, and their investment horizion. Therefore, I follow the methodology in Koijen et al. (2021) and proxy $\mathbb{E}[R_{US} - R_j]$ as follows:

$$er_{j,t} = y_{US,t} - y_{j,t} + \rho_t - d_{j,t}$$

where $y_{US,t}$ and $y_{j,t}$ are the yields of US and country j government bonds; ρ_t is the market-implied forward premium for the euro against the dollar; and $d_{j,t} := CDS_{US,t} - CDS_{j,t}$ is the difference in sovereign CDS rates between the US and country j. All components are averaged over the 1,2,3,5, and 10 year maturities and over quarter t, weighted by the maturity structure of outstanding government debt for country j. This approach relies on using ρ_t as a measure of market-based expectations for exchange rates, and on controlling for differences in credit risk through $d_{j,t}$. The latter is pivotal in estimating the theoretical model, as it is assumed that excess returns arise only through currency fluctuations, while in the data sovereign risk is likely to play an important role, especially in a sample of

[9] Note that section 1.3.3 discusses the identification strategy in terms of PSPP purchases to build intuition, but here I use PSPP *holdings* because in the theoretical model excess returns depend on the *level* of government debt supply.

European government bonds. The macroeconomic and financial controls included as time fixed effects and in $\mathbf{V_{j,t}}$ and $\mathbf{W_{j,t}}$ complement the strategy by serving as potential predictors of excess returns, akin to the factor models popular in modern empirical asset pricing (Koijen and Yogo, 2019a,b).

The regressions also include fixed effects at both the country ($\alpha_{k,j}$ and $\iota_{k,j}$) and quarter ($\alpha_{k,t}$ and $\iota_{k,t}$) level. As explained in section 1.3.3, time fixed effects aid the identification strategy by isolating cross-country differences in PSPP holdings. Furthermore, they control for any global determinants of the demand and supply of safe assets. Country fixed effects account for time-invariant featrues such as idiosyncratic country risk, abstracted away in the theoretical model. Controlling for country risk is particularly important for bonds issued by distressed sovereign such as Italy, Ireland, Portugal and Spain. Given the "bank-sovereign nexus", whereby investment in risky sovereign bonds by domestic banks is often driven by political considerations (Andreeva and Vlassopoulos, 2016; Ongena et al., 2019; Saka, 2020), country-specific risk is especialy relevant in the model for bank portfolios. Country fixed effects also take care of any potential cross-country pattern in the correlation and variance of excess returns, which are kept as fixed parameters in the theoretical model.

I also augment the model with two sets of country-quarter level controls. The first set, $\mathbf{V_{j,t}}$ accounts for changes in the portfolio shares due to valuation effects. The theoretical model is cast in real terms, so in the data I need to control for valuation effects due to both bond prices and exchange rates, in order to isolate actual portfolio rebalancing. In the baseline specification with time and country fixed effects, $\mathbf{V_{j,t}}$ includes quartely changes in the all-maturity price index for country j's government bonds $\Delta BI_{j,t}$. Note that changes in the EUR/USD exchange rate $\Delta e_t^{EUR/USD}$ and in the dollar price of US bonds $\Delta BI_{US,t}$, which also affect portfolio shares, are subsumed in the quarterly fixed effects. They are included in $\mathbf{V_{j,t}}$ for specifications whose fixed effect structure allows it.

The second set of controls $\mathbf{W_{j,t}}$ include CPI inflation, real GDP growth and the ratio of government debt to GDP for country j in quarter t. The first two variables are included to succinctly capture macroeconomic factors that affect investment in country j, which might be correlated with the maturity choice of government debt issuance, in turn driving variation in PSPP holdings. The latter captures both a time-varying factor of country risk, and possible concerns of residual discretionality in PSPP purchases tilted towards highly indebted countries.

I estimate the model on country-quarter observations from 2015 Q1 to 2022 Q1 for all eurozone countries except Greece, as it is excluded from the PSPP; and Estonia and Luxembourg, due to data availability.

Summary statistics

Table 1.1 reports summary statistics for the variables used in the regression model. The shares of US Treasuries to country j government bonds are large, around 50% for ICPFs and more than 70% for banks on average. The empirical proxy for excess return is on average negative for Treasuries, at -26 basis points. Coupled with the positive correlation between Treasury excess returns and the income of insurance companies shown in Figure 1.2b, this feature suggests that the risk-return profile of Treasuries is not sufficient to explain ICPF's holdings of US Treasuries and convenience yields might indeed be at play.

In order to benchmark the convenience yield implied by the estimation of structural parameters, I report an empirical proxy of the convenience yield component of excess returns.[10] The empirical convenience yield is negative, as predicted by the model, with an average of -25 basis points. It is on average very close to excess returns, implying a very small risk premium in the model. Therefore, we would expect convenience yields to also explain the lion's share of the excess returns implied by the estimated structural parameters.

First-stage regression

Table 1.2 displays results from the first-stage regression. The coefficient on $PSPP_{i,t}$ in the column 3, the preferred specification including both time and country fixed effects, reports a highly statistically significant increase of 1.47 percentage points in US Treasury excess returns in response to a one standard deviation increase in PSPP holdings, equivalent to about \$ 153 billion. The size of the coefficient is also significant, almost three times larger than the 0.64 percentage point standard deviation of the empirical proxy for excess returns.

The sign of the coefficient is consistent with the prediction of the model, as an increase in PSPP holdings corresponds to a decrease in the supply of country j government debt on the market. The increase in US Treasury excess returns, possibly through the convenience yield component, echoes the findings of Jiang et al. (2024), which highlights the change in the relative supply of safe assets and convenience yields as an additional channel through which quantitative easing affects

[10] I follow Du et al. (2018) in estimating convenience yields as $\phi_{j,t} = y_{US,t} - y_{j,t} + \rho_t - bs_{j,t} - l_{i,t}$, where $bs_{j,t}$ is the EUR/USD cross-currency basis swap, a measure of CIP deviations in interbank rates that purges the measure of convenience yields of FX market frictions.

Table 1.1. Summary statistics

	N	Mean	SD	Min	P25	P50	P75	Max
A. Portfolio shares								
$s_{US,B,t}$	442	71.2	26.8	14.8	50.9	80.9	95.3	99.1
$s_{US,I,t}$	425	48.96	33.19	2.59	16.71	46.82	80.59	98.83
B. Financial variables								
$er_{j,t}$	463	-0.26	0.64	-5.21	-0.49	-0.16	0.08	0.77
$\phi_{j,t}$	387	-0.25	0.24	-2.04	-0.35	-0.25	-0.15	1.65
$\Delta BI_{j,t}$	371	0.20	4.39	-10.69	-1.89	0.42	3.10	14.09
$\Delta BI_{US,t}$	371	0.5	2.4	-4.7	-0.5	0.6	1.5	7.2
$\Delta e_t^{EUR/USD}$	463	0.4	3.9	-5.6	-1.8	0.1	2.5	13.7
C. Macroeconomic variables								
$Debt/GDP_{j,t}$	372	87.1	30.0	36.3	62.3	83.3	108.4	158.9
$\Delta CPI_{i,t}$	372	1.4	1.9	-2.2	0.2	1.1	2.0	11.7
$\Delta GDP_{i,t}$	343	0.6	3.5	-17.6	0.2	0.5	0.9	21.4

Summary statistics calculated on the data in which observations for PSPP holdings are non-empty. All variables in percentage points.

equilibrium interest rates.

Columns 1 and 2 show the results of models with no fixed effects, and with time fixed effects only, respectively. The inclusion of time fixed effects is crucial in my identification strategy as it allows to single out exogenous variation in relative debt supply that is independent of the ECB's overall unconventional monetary polcy stance. Column 2 shows how time fixed effects raise the F stat to 70.63, implying a stronger instrument as well as an arguably more valid one. Column 3 shows that the inclusion of country fixed effects does reduce the F statistic, albeit to a still high value of 23.86. Therefore, there is a tradeoff between having a strong instrument and controlling for important country-specific factors such as credit risk and the sovereign-bank nexus.

Second-stage regressions

Tables 1.3 and 1.4 report the estimation results of the second-stage equations for banks and ICPFs, respectively.

The preferred 2SLS specification with time and country fixed effects shows an in-

Table 1.2. First-stage regression

	(1)	(2)	(3)
$PSPP_{j,t}$	0.75***	1.37***	1.47***
	(0.10)	(0.16)	(0.30)
$\Delta BI_{j,t}$	-0.01*	-0.01*	-0.01
	(0.01)	(0.01)	(0.01)
$\Delta BI_{US,t}$	0.02**		
	(0.01)		
$\Delta e_t^{EUR/USD}$	-0.00		
	(0.00)		
N	309	309	309
F stat	58.71	70.63	23.86
Macro controls	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes
Country fixed effects	No	No	Yes

Coefficients from regression model $er_{j,t} = \iota_j + \iota_t - \pi PSPP_{j,t} + \lambda' \mathbf{V_{j,t}} + \kappa' \mathbf{W_{j,t}} + \nu_{j,t}$. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

crease of the US Treasury portfolio shares in response to higher excess returns. The coefficient is statistically significant at the 1% level for both banks and insurers. Therefore, the prediction of a positive reaction to excess returns is verified for both sectors.

The comparison of coefficients on $er_{j,t}$ estimated via OLS (columns 1 to 3) and 2SLS (columns 4 to 6) suggests that the instrumental variable strategy purges the coefficients from the bias due to the endogeneity of supply and demand. The coefficient β represents the sensitivity of portfolio shares, an equilibrium quantity, to excess returns, tightly connected to equilibrium asset prices. Therefore, the observed price-quantity data points are likely driven by both demand and supply shocks. The former introduce a negative correlation between $s_{US,j,k,t}$ and $er_{j,t}$, while the latter a positive correlation. Therefore, failing to isolate supply shocks would result in a bias toward zero for β. PSPP-induced exogenous changes in B_j act as a supply shock, so by using them as instrument for $er_{j,t}$ the slope of the demand curve can be recovered. The larger and more significant coefficients across the board in the 2SLS models indicate that this strategy is indeed successful.

In order to understand what global variables are accounted for by time fixed effects, models in columns 1 and 4 replace them with the VIX, a key determinant of the demand for safe assets (Miranda-Agrippino and Rey, 2022); and with the debt/GDP ratio in the US to control for the supply of US Treasuries. The estimated β in columns 1 and 4 are very similar to those in columns 2 and 5, which replace global controls with quarterly fixed effects. Therefore, time fixed effects appear to capture well the role of global drivers of the demand and supply of safe assets. The lack of time fixed effects also allows to augment the vector of valuation effect controls $\mathbf{V_{j,t}}$ with changes in the exchange rate and US bond prices, which vary only in the time dimension. However, none of the valuation effects are statistically significant even at the 10% level, possibly because of relatively small quarter-on-quarter variation.

The models estimated in columns 2 and 5 include time fixed effects, but not country fixed effects. Comparing them to the coefficients in columns 3 and 6 reveals the imporance of controlling for country-specific characteristics, especially for banks, as argued in the previous section. In fact, the eclusion of country fixed effects in the bank regressions results in negative coefficients that do not seem particularly credible. The lower sensitivity of the models for insurance companies to country fixed effects corroborates the hypothesis that the political economy factors affecting investment of European banks in distressed sovereign bonds contribute to biasing the estimates. However, it is still important to include country fixed effects in the model for insurers. The reasons lie both in consitency with the estimates for banks, and in accounting for the time-invariant portion of country-specific credit risk, which the theoretical model abstracts away.

Figure 1.4 compares the coefficients on $er_{j,t}$ from column 6 of the models for banks and insurers. The US Treasury portfolio share of banks increases by 35.94 percentage points in response to a one percentage point increase in the convenience yield component of excess returns, while the portfolio share of insurers increases by 6.24 percentage points. The coefficients are statistically different from each other at the 1% significance level. The muted reaction of insurers' portfolio shares to excess returns is consistent with results in the literature (Timmer, 2018; Eren et al., 2023; Koijen et al., 2021), and suggests there is enough of a discrepancy in behaviour between the two sectors that more than risk aversion might be at play. However, it is not enough to test the predictions of the model on the effects of convenience yields on portfolio share sensitivity, as they concern a counterfactuals. In order to quantify the relative importance of risk aversion and preferences for Treasuries, it is necessary to recover the structural parameters γ_k, ψ_k and σ from the estimates of the empirical model. I turn to this task in the next section.

Table 1.3. Second-stage regression for banks

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
$er_{j,t}$	-7.44	-15.20*	4.75**	22.86	-3.43	35.94***
	(6.98)	(8.79)	(2.32)	(24.19)	(25.08)	(9.94)
$\Delta BI_{j,t}$	0.01	0.89	-0.22	0.32	1.05	0.03
	(0.82)	(1.37)	(0.17)	(0.88)	(1.33)	(0.23)
$\Delta BI_{US,t}$	0.34			-0.47		
	(1.09)			(1.28)		
$\Delta e_t^{EUR/USD}$	0.19			0.35		
	(0.74)			(0.75)		
N	309	309	309	309	309	309
Macro controls	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes	No	Yes	Yes
Country fixed effects	No	No	Yes	No	No	Yes
Underid test p-value				0.00	0.00	0.00
Weak id test stat				58.71	70.63	23.86

Coefficients from regression model $s_{US,j,B,t} = \alpha_{B,i} + \alpha_{B,t} + \beta_B er_{j,t} + \delta'_B \mathbf{V_{j,t}} + \eta'_B \mathbf{W_{j,t}} + \varepsilon_{j,B,t}$ estimated via OLS, or via 2SLS using $PSPP_{j,t}$ as an instrument for $er_{j,t}$. The underidentification test uses the Kleibergen and Paap (2006) LM statistic. The weak identification test uses the Kleibergen and Paap (2006) Wald F statistic. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

1.4 Structural parameters and model experiments

1.4.1 Recovery of structural parameters

After obtaining the coefficient estimates $\hat{\alpha}_k$, $\hat{\beta}_k$ for $k = \{B, I\}$, and $-\hat{\pi}$ from the 2SLS model, I can back out values for the structural parameters by solving the system of equations 1.11, 1.12 and 1.13 for γ_k, ψ_k, and σ.

Equation 1.13 depends on $\partial b_{j,O}/\partial B_j$, the absorption rate of country B_j's government bonds by other investors. I estimate this parameter through absorption regressions that decompose total outstanding amounts of country j government bonds into sectoral holdings, and I replace the absorprion rate of other investors with its estimate $\hat{\theta}_O$.[11] Since it is an estimated parameter, I will have to account

[11] More details on this procedure, as well as the full regression results, can be found in Appendix 1.E

Table 1.4. Second-stage regression for insurers

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
$er_{j,t}$	-2.46	-4.99*	-0.45*	17.75***	10.03	6.24***
	(1.97)	(2.56)	(0.24)	(6.25)	(6.73)	(2.02)
$\Delta BI_{j,t}$	0.74	2.14	-0.17	1.72	3.15*	0.10
	(0.92)	(1.59)	(0.14)	(1.15)	(1.69)	(0.24)
$\Delta BI_{US,t}$	0.16			-2.56		
	(1.26)			(1.72)		
$\Delta e_t^{EUR/USD}$	0.49			0.98		
	(0.79)			(0.92)		
N	307	307	307	307	307	307
Macro controls	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes	No	Yes	Yes
Country fixed effects	No	No	Yes	No	No	Yes
Underid test p-value				0.00	0.00	0.00
Weak id test stat				58.21	69.86	23.65

Coefficients from regression model $s_{US,j,I,t} = \alpha_{I,i} + \alpha_{I,t} + \beta_I er_{j,t} + \delta_I' \mathbf{V_{j,t}} + \eta_I' \mathbf{W_{j,t}} + \varepsilon_{j,I,t}$ estimated via OLS, or via 2SLS using $PSPP_{j,t}$ as an instrument for $er_{j,t}$. The underidentification test uses the Kleibergen and Paap (2006) LM statistic. The weak identification test uses the Kleibergen and Paap (2006) Wald F statistic. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

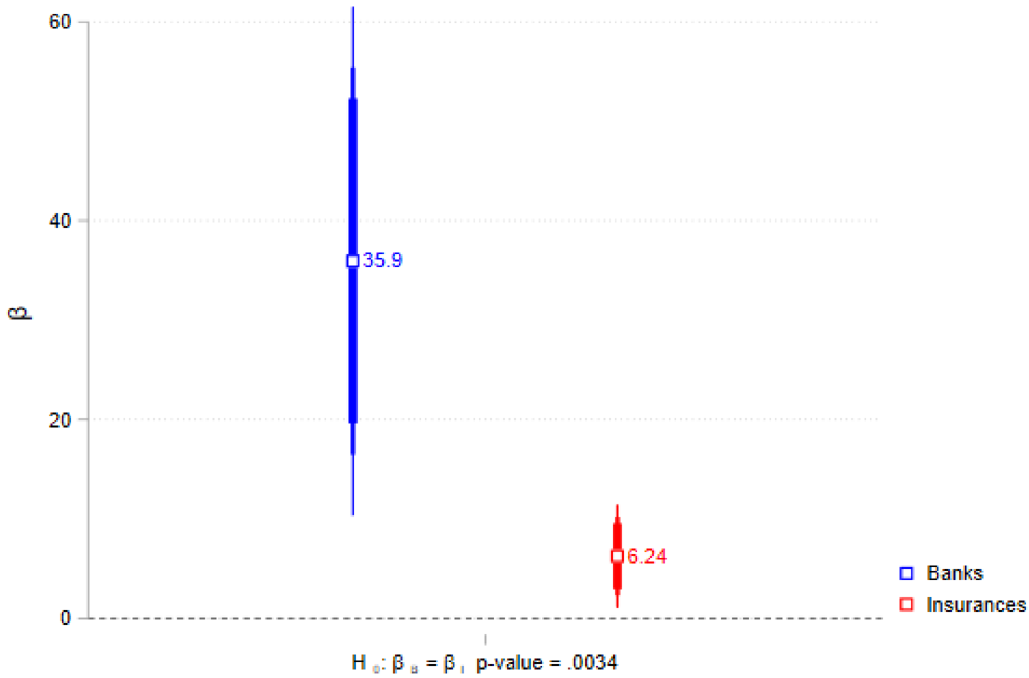
for estimation uncertainty in the simulation of confidence intervals for the structural parameters.

I replace α_k, β_k and π with their estimates from the model with country and quarter fixed effects, and choose approximation points $\bar{e} = 0$, $\bar{c} = 0$, $\bar{v} = 1$, and $\bar{s} = 0.5$. [12] Computational convenience guided the choice of these points, and the summary statistics in Table 1.1 show hat they are reasonably close to their sample counterparts.

Finally, I calibrate the other parameters in the theoretical model to sample means. Note that this approach is equivalent to maximum likelihood estimation of these parameters, given the normality of the limiting distribution of estimated regression coefficients. $\mathbb{V}[R_{US} - R_j]$ and $Cov[R_{US} - R_j, Y_k]$ are replaced with their sample

[12] I calculate the intercept coefficient α_k as the average of estimated fixed effects $\hat{\alpha}_{k,j}$ and $\hat{\alpha}_{k,t}$ over the country-quarter distribution.

Figure 1.4. Comparison of β for banks and insurers



The blue-bordered dot represnets the coefficient β_B, and the red-bordered dot represents the coefficients β_I. Both coefficients are estimated via 2SLS in the model $s_{US,k,t} = \alpha_{k,i} + \alpha_{k,t} + \beta_k er_{j,t} + \delta_k' \mathbf{V_{j,t}} + \eta_k' \mathbf{W_{j,t}} + \varepsilon_{j,k,t}$ via 2SLS using $PSPP_{j,t}$ as an instrument for $er_{j,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. The p-value on the hypothesis $H_0 : \beta_B = \beta_I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

counterparts calculated over the full country-quarter distribution. Likewise, I use the average amount of the total holdings of US and country j government bonds over the country-quarter distribution for each sector as an estimate for $W_{0,k}$. Since the model places no restrictions on the unit of measure of initial wealth, I calibrate it to match the order of magnitude of the empirical proxy of excess returns. [13]

Table 1.5 reports the distribution of recovered structural parameters, simulated from the asymptotic distribution of the vector of estimated coefficients $\boldsymbol{\lambda} := (\hat{\alpha}_B, \hat{\beta}_B, \hat{\alpha}_I, \hat{\beta}_I, \hat{\pi}, \hat{\theta}_O)$. Risk aversion is 1.5 times higher for ICPFs, with a mean of 0.37 compared to 0.24 for banks. This result is intuitively appealing due to the

[13] Appendix 1.E details the procedure to solve the system and simulate the distribution of structural parameters.

Table 1.5. Structural parameters

Structural parameter	Mean	95% CI lower bound	95% CI upper bound
A. Banks			
γ_B	0.24	0.02	0.2
ψ_B	3.34	1.0	978.64
B. Insurance companies			
γ_I	0.37	0.13	1.0
ψ_I	3.63	0.13	149.34
C. Common parameters			
σ	2.97	0.51	101.0

Confidence intervals are obtained by drawing 100,000 times from the joint asymptotic distribution of estimated parameters in the empirical model, solving for structural parameters for each joint draw, and computing the 5th and 95th percentiles of the simulated distribution.

intrinsic differences in business models for the two sectors.

The size of the parameter ψ reveals that convenience yields play a non-negligible role in the preferences of both investors. The mean values are above 3, corresponding to a weight of about 75% in their objective function. This result is striking as dollar-denominated government bonds carry exchange rate risk for European investors, highlighting the special role of US Treasuries in the internataional financial system even beyond their safety properties, and echoing Kaldorf and Röttger (2023)'s discussion of "convenient but risky" sovereign debt in the context of peripheral eurozone countries. Insurers appear to assign a slightly higher importance to Treasuries, with a mean of 3.63 for ψ, compared to 3.34 for banks, translating into a 1.5 percentage point difference in the weight of the US Treasuries component of their preferences.

Note that the mean value of 2.97 for the parameter σ implies a very high curvature of the Treasury term of the objective function. Therefore, the marginal benefit of holding Treasuries is steep, and even the seemigly small difference in the estimates of ψ between the two sectors can potentially translate in a large impact on the sensitivty of demand, and in turn on equilibrium excess returns.

However, the mere comparision of the size of estimated parameters is not enough to pin down the relative importance of the two facets of preferences analysed in the model. In the next section, I perform experiments on γ_k and ψ_k to investigate how parameter values translate into the relative strength of risk aversion and convenience yields in determining both portfolio choice and equilibrium interest rates.

1.4.2 Model experiments

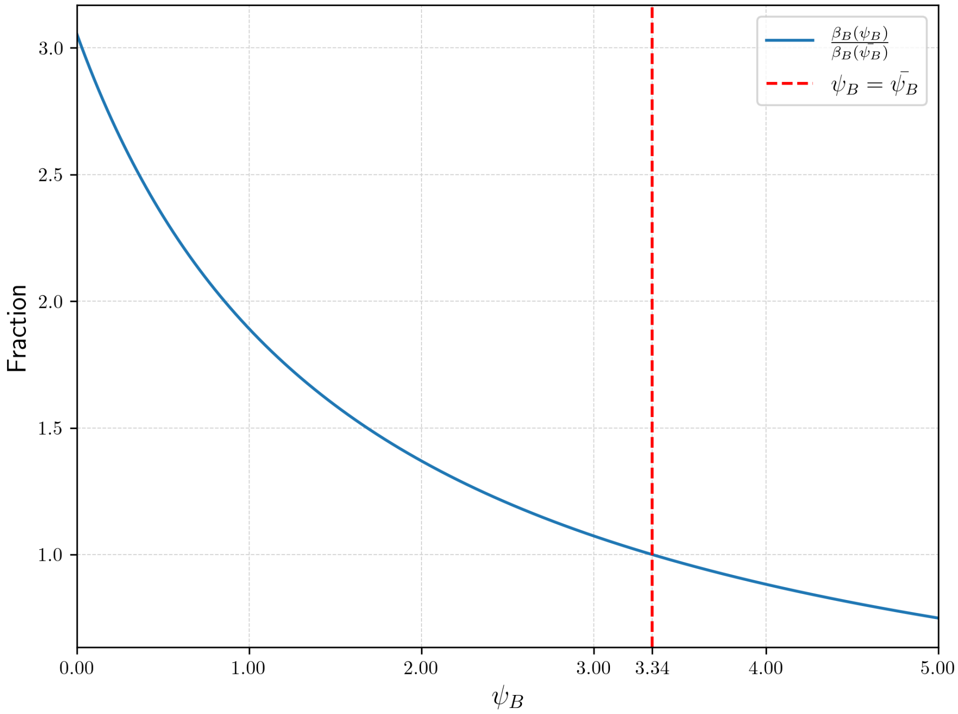
First, I perform two distinct but related exercises to quantify the importance of risk aversion versus convenience yields in explaining the magnitude of investors' portfolio share sensitivity to the mean and variance of excess returns, and their differences across sectors. Then, I investigate the effect of convenience yield investors on equilibrium excess returns.

Portfolio share sensitivity as a function of ψ_k

Equations 1.3 and 1.4 claim that the presence of convenience yield preference for Treasuries reduces investors' reaction to both the mean and the variance of excess returns compared to the case of $\psi_k = 0$. To test and quantify this prediction, I calculate the counterfactual values of β_k for both investors as a function of ψ_k, fixing risk aversion γ_k at its mean. To get a sense of the magnitude of the effect of ψ_k on the sensitivity to excess returns, I divide $\beta_k(\psi_k)$ by $\beta_k(\bar{\psi}_k)$, its value at the mean for ψ_k. Figures 1.5 and 1.6 plot this function against ψ_k (blue line). For values lower than the mean $\bar{\psi}_k$ (red dashed line), the function is positive, meaning that the corresponding β_k is larger. The sensitivity coefficient β_k at $\psi = 0$, in the absence of Treasuries in the utility function, is 3 times larger than at the mean value for banks, and more than 9 times larger for ICPFs. The estimated parameters imply that the Treasury component of preferences has a sizeable impact on the yield sensitivity, with large differences in the impact across sector. This gap is due to the curvature of the Treasury component of preferences magnifying the small differences in ψ_k.

The exact same conclusions of this counterfactual exercise hold for the sensitivity to the variance of excess return. As evident from equations 1.3 and 1.4, the two derivatives differ only by the numerator, while ψ_k affects only the denominator. Therefore, the ratios plotted are exactly the same for the model-implied reaction to the variance of excess returns. The sensitivity of investors' demand to volatility in the Treasuries market is then substantially lower then it would be absent convenience yields, implying that special preferences for US Treasuries make the funding of US government debt more stable even in the face of turbulent periods in the markets.

Figure 1.5. Percentage change in β_B as a function of ψ_B



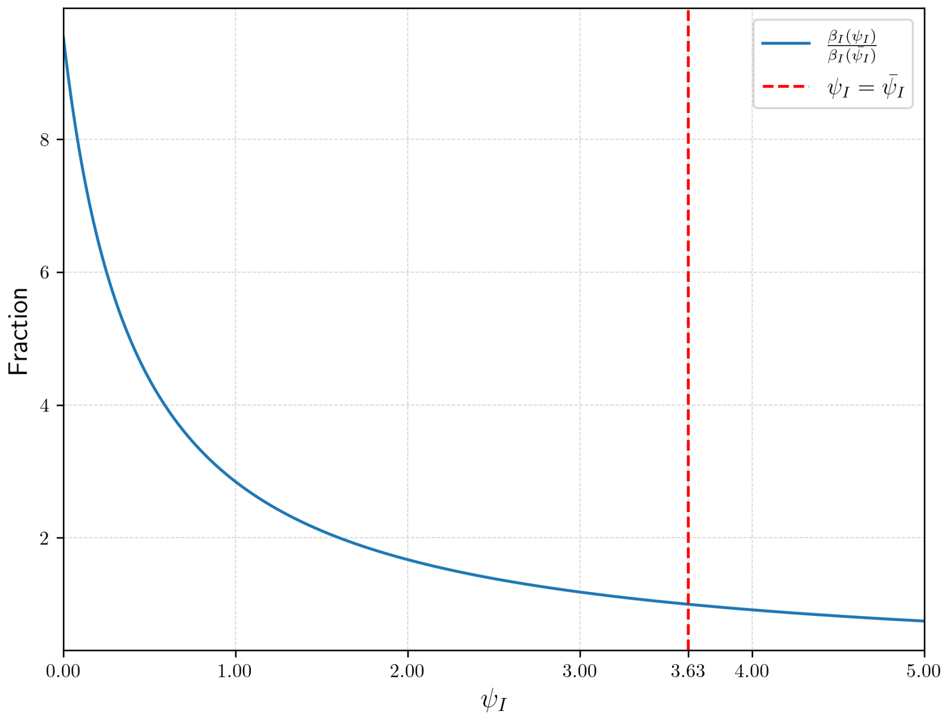
$\beta_B(\psi)$ is calculated using the means of parameters γ_B and σ drawn from the simulated distribution, letting ψ_B vary, and using the calibrated values displayed in Table 1.8 for other parameters. $\beta_B(\bar{\psi}_B)$ is calcualted using the same parameters as $\beta_B(\psi_B)$, but using the mean of ψ_B from the simulated distribution, defined as $\bar{\psi}_B$.

Difference in sensitivity across sectors as a function of γ_I

This exercise is aimed at quantifying the percentage of the difference in portfolio share sensitivity between banks and insurers that can be attributed to convenience yields versus risk aversion. The results of the previous experiment suggest that this percentage might be quite large, given the difference in the impact of ψ_k on sensitivity across sectors.

I compute the difference in coefficients $\beta_B - \beta_I$ as a function of γ_I, which influences only β_I. I then divide it by the same difference evaluated at the mean for both parameters, to obtain a fraction. Figure 1.7 plots this function against γ_I (blue line). The function is increasing because β^I is decreasing in γ_I: all else

Figure 1.6. Percentage change in β_I as a function of ψ_I

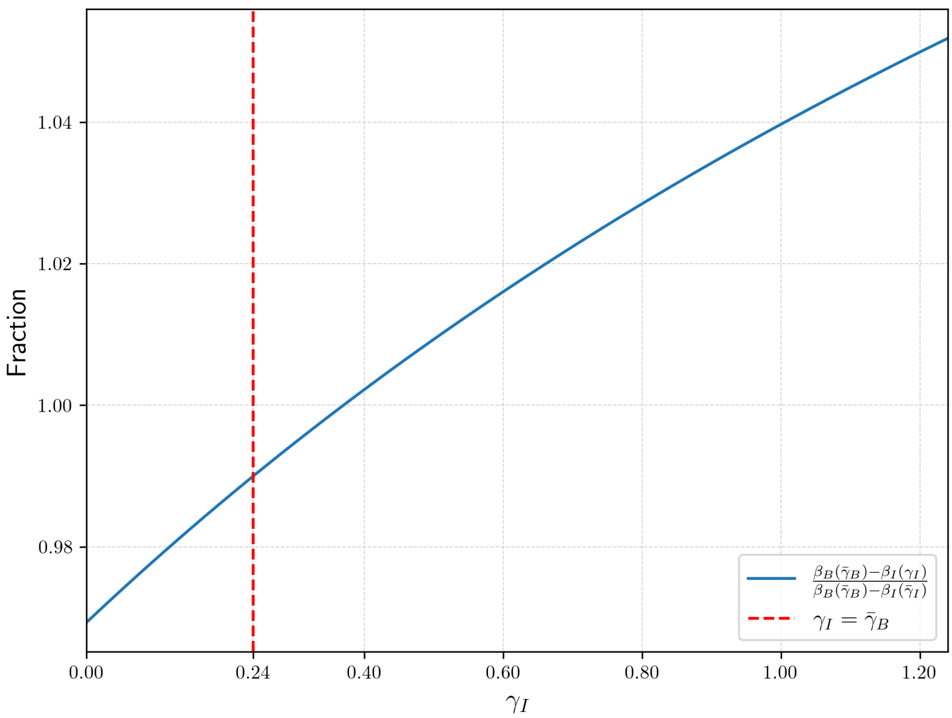


$\beta_I(\psi)$ is calculated using the means of parameters γ_I and σ drawn from the simulated distribution, letting ψ_I vary, and using the calibrated values displayed in Table 1.8 for other parameters. $\beta_I(\bar{\psi}_I)$ is calcualted using the same parameters as $\beta_I(\psi_I)$, but using the mean of ψ_I from the simulated distribution, defined as $\bar{\psi}_I$.

equal, a higher risk aversion translates into a weaker reaction to excess returns. The value at $\gamma_I = \bar{\gamma}_B$ (red dashed line) is of particular interest. By equalising the risk aversion of insurance companies and banks, according to the model any residual difference in β between the two sectors is attributable to the preference for Treasuries. According to this measure, convenience yields can explain about 99% of the difference in β, so they play a dominant role in explaining the observed cross-sector heterogeneity of reactions to excess returns.

However, it is important to underscore how this striking result relies critically on two aspects of the modelling approach. First, investors' utility is assumed to depend on two structural parameters only, so there is no place for other features such as regulation, differences in potential convenience yield preferences for euro area government bonds (Jiang et al., 2020), or heterogeneity in home bias across sectors. Second, as noted above, the steepness of the marginal utility of holding Treasuries implied by the estimated σ produces an outsized effect despite small differences in ψ.

Figure 1.7. Percentage change in $\beta_B - \beta_I$ as a function of γ_I



$\beta_I(\gamma_I)$ is calculated using the means of parameters γ_B, ψ_B and ψ_I drawn from the simulated distribution, letting γ_I vary, and using the calibrated values displayed in Table 1.8 for other parameters. $\beta_I(\bar{\gamma_I})$ is calcualted using the same parameters as $\beta_I(\gamma_I)$, but using the mean of γ_I from the simulated distribution, defined as $\bar{\gamma}_I$. $\beta_B(\bar{\gamma}_B)$ is calculated using the means of parameters γ_B, γ_I, ψ_B, and ψ_I drawn from the simulated distribution, and using the calibrated values displayed in Table 1.8 for other parameters.

Risk premium and convenience yield over time

The previous two experiments demostrated that preferences for Treasuries have a quantitatively strong effect on portfolio choice. However, preferences for US Treasuries affects equilibrium excess returns as well. Equation 1.5 shows that equilibrium excess returns can be decomposed in two components: the risk premium $RP > 0$, and the convenience yield $\phi < 0$.

After estimating structural parameters γ_k, ψ_k and σ, I can calculate the model-implied excess returns, broken down into the risk premium and convenience yield components, to gauge how they compare to their empirical counterparts displayed in Table 1.1. Figure 1.8 plots the model-implied risk premium $RP(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t$ (blue line) and the total excess return $ER(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t$ (red line) at the means for structural parameters, over the sample period from 2015 Q1 to 2022 Q1. The difference between the two lines (blue shaded area) then represents the absolute value of the convenience yield component $|\phi(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t|$. [14]

The order of magnitude of excess returns is a calibration target for in the calculation of structural parameters, so it is close to the empirical estimate by construction. However,contrary to its empirical counterpart, model-implied excess returns are always negative, implying that the convenience yield component dominates.Furthermore, it is less volatile, with a minimum value of about -1.75 percentage points compared to -5.21 percentage points for the empirical proxy.

The risk premium component is notably much less volatile than total excess returns, hovering at about 10 basis points throughout the sample. Once again, the discrepancy is attributable to the large value of σ, which magnifies movements of $b_{US,k,t}$, the source of time series variation, in the convenience yield term.

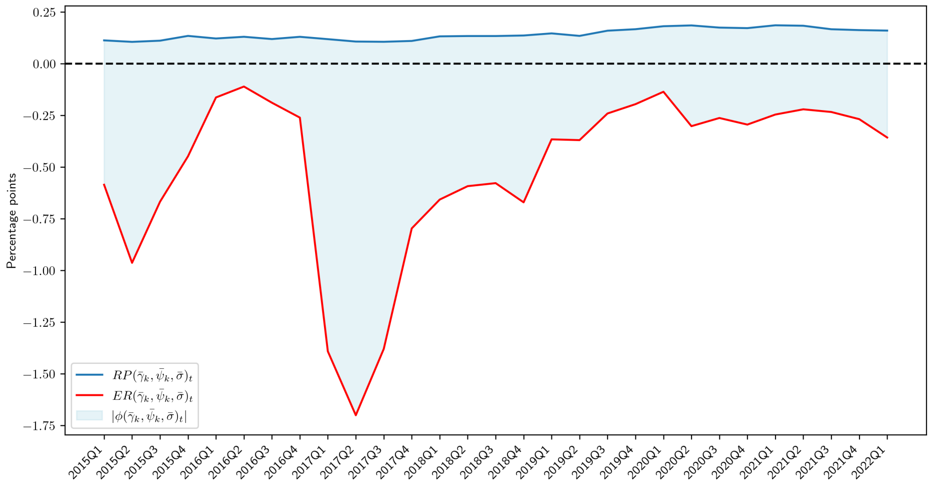
Overall, the model implies that convenience yields are quantitatively much larger than risk premia, driving excess returns consistently into negative territory and explaining much of their time-series variation. Interestingly, these properties match the behaviour of the empirical proxies for excess returns and convenience yields, which both have negative means and display a similar distribution.

Effect of convenience yield on equilibrium excess returns

Finally, I perform a counterfactual experiment on the model-implied excess returns, asking how their mean varies as a function of the Treasury preference parameter ψ_k for each sector.

[14]Note that the time-series variation in the figure is driven entirely by changes in the US portfolio share of the two sectors, which is the only term that is allowed to vary over time as I treat initial wealth, variances and covariances as fixed parameters. In order to purge excess returns from trends, for the purposes of this figure I scale US Treasury holdings $b_{US,k,t}$ by the total outstanding amount of US Treasuries.

Figure 1.8. Time-series decomposition of excess returns in risk premium and convenience yield

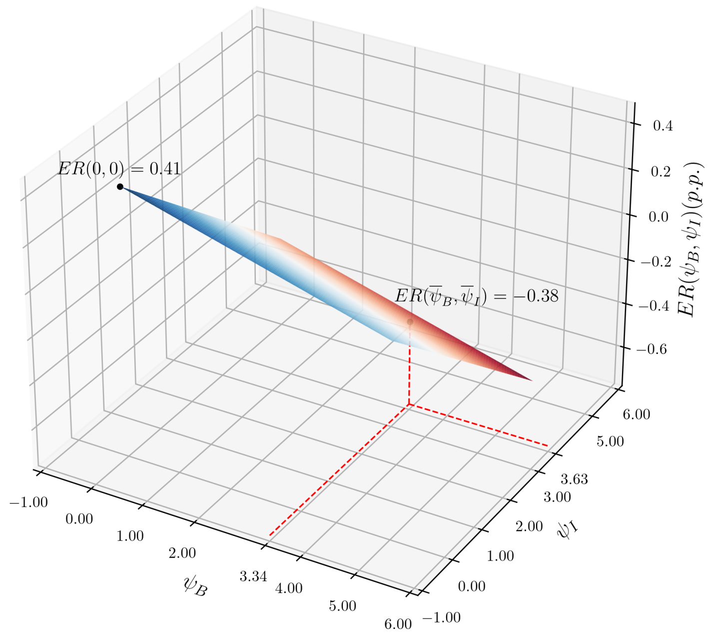


Excess returns ($ER(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t$), the risk premium ($RP(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t$) and the absolute value of convenience yields ($|\phi(\bar{\gamma}_k, \bar{\psi}_k, \bar{\sigma})_t|$) are calculated from equation 1.5 using the means of parameters γ_k, ψ_k and σ drawn from the simulated distribution, and using the calibrated values displayed in Table 1.8 for other parameters, except for $\bar{s}$. It is replaced by $s_{US,k,t}$ for each quarter-sector observation.

Figure 1.9 plots on the z axis the excess return $ER(\psi_B, \psi_I)$ as a function of varying levels of ψ_B (x axis) and ψ_I (y axis), leaving γ_k and σ at their mean values. Excess returns are steeply decreasing in both ψ_B and ψ_I, as the higher importance of US Treasuries in investors' utility functions implies that they require lower monetary returns. Therefore, any loss of confidence by foreign investors in the special value of US Treasuries, even partial, would result in a fast erosion of the US government's funding advantage.

In the extreme case of a complete loss of special status, represented by $\psi_k = 0 \, \forall \, k$, US Treasuries would have to pay a positive excess return of 41 basis points on euro area government bonds, a large jump of 79 basis points from their value of -38 basis points at parameter means. The stark penalty implied by the model suggest that the sustainability of US public finances relies on the special role of US Treasuries as a global asset, as observed in previous studies (Jiang et al., 2019; Bonam, 2020; Choi et al., 2024).

Figure 1.9. Exces returns as a function of ψk



Excess returns $ER(\psi_B, \psi_I)$ (z axis) are calculated from equation 1.5 using the means of parameters γ_k and σ drawn from the simulated distribution, letting ψ_B (x axis) and ψI (y axis) vary, and using the calibrated values displayed in Table 1.8 for other parameters.

1.5 Conclusion

This paper shows that the convenience yield that foreign investors derive from US Treasuries plays a key role in explaining the observed differences in demand sensitivity across sectors. Differences in risk aversion are also substantial, but alone they cannot reconcile the joint observation of positive US Treasury portfolio shares for ICPFs; negative excess returns of Treasuries with respect to eurozone government bonds; and a positive correlation between Treasury excess returns and the income of insurers.

Thanks to the estimation of structral parameters in investors' preferences, the relative importance of convenience yields and risk aversion can be quantified. Both a high risk aversion and convenience yields imply a lower reaction of portfolio shares to excess returns, so a structural approach is necessary to disentangle their impact.

The model implies that, absent convenience yields, the demand of banks would be 3 times more sensitive to the mean and variance of excess returns, while insurance companies would be 9 times more sensitive. As a result, the absorption of additional US government debt by foreign investors would be much more fickle and volatile.

The sizeable impact of convenience yields on the yield sensitivity of portfolios is matched by an equally large effect on equilibrium rates. The decomposition in the model shows that the convenience yield term accounts for the vast majority of the volatility in excess returns, and that it is large enough to turn excess returns consistently negative. Furthermore, the model-implied excess return is steeply increasing in the Treasury preference parameter of both sectors, and it would jump from -38 basis points to 41 basis points in the absence of convenience yields.

The policy implications are twofold. First, the sustainability of persistent US government deficits heavily relies on yield-insensitive foreign investors to absorb additional sovereign debt at a low rate. Second, the high sensitivity of returns to the convenience yield component highlights the risks of US Treasuries losing their special status, leading to a potentially large increase in the borrowing cost for the US government.

While the quantitative conclusions of this paper are by costruction limited to the context of European banks and insurance companies, they nevertheless offer insights for further research on the nuances of the foreign demand for Treasuries by different sectors in a global perspective.

Appendix

1.A Data sources

Table 1.6. Data sources

Data	**Source**
Government bond holdings of eurozone banks and ICPFs	ECB Securities Holdings Statistics
Income of eurozone banks	ECB Consolidated Banking database
Income of eurozone insurers	EIOPA Insurance Statistics
Government debt purchases and holdings under PSPP	ECB
Government bond indices and yields	Refinitiv Eikon
Spot and forward exchange rates	Refinitiv Eikon
EUR/USD cross-currency basis swap	Refinitiv Eikon
CDS rates	Refinitiv Eikon
Amount of government debt outstanding	Bank for International Settlements
	and Federal Reserve Economic Data (FRED)
Real GDP growth	OECD
CPI inflation	ECB and IMF
Debt/GDP ratio	Eurostat and FRED

1.B Mean-variance preferences with convenience yields

Consider the problem of an investor allocating their initial wealth $W_{0,k}$ between domestic government bonds $b_{j,k}$ with riskless return R_j, and US Treasuries $b_{US,k}$ with return R_{US}, which is risky because of fluctuations in the exchange rate. The investor derives utility from their final wealth W_k, and from holding US Treasuries. The utility function is

$$U(W_k, b_{US,k}) = -e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})}$$

This utility function preserves the desirable properties of standard exponential utility, namely it is increasing and concave in W_k, and it displays constant absolute risk aversion with risk aversion coefficient γ_k.

Furthermore, by taking the first and second derivatives with respect to $b_{US,k}$,

$$U'(b_{US,k}) = \gamma_k \psi_k b_{US,k}^{-\sigma} e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})} > 0$$
$$U''(b_{US,k}) = -\gamma_k \psi_k b_{US,k}^{-2\sigma} \left(\gamma_k \psi_k + \sigma b_{US,k}^{\sigma-1}\right) e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})} < 0.$$

Therefore, due to the CES specification the marginal utility of holding Treasuries is declining, so investors require a higher monetary return to absorb more Treasuries in equilibrium. This mechanism is widely used in the literature to link the outstanding amount of US government debt with the equilibrium convenience yield (Krishnamurthy and Vissing-Jorgensen (2012) and Engel and Wu (2018), among others).

The investor maximises their expected utility subject to their budget constraint, expressed in terms of $b_{j,k}$ as the outside risk-free asset.

$$\max_{b_{US,k}} \mathbb{E}\left[-e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})}\right]$$
$$\text{s.t.} W_k = \quad RW_0 + (R_{US} - R_j) b_{US,k} + Y_k,$$

Assume that $R_{US} \sim N(\mu_{US}, \sigma_{US}^2)$ and $Y_k \sim N(\mu_Y, \sigma_Y^2)$, so that $W_k \sim (\mu_W, \sigma_W^2)$, with $\mu_W = RW_{0,k} + (\mu_{US} - R)B_{US} + \mu_Y$ and $\sigma_{W_k}^2 = B_{US}^2 \sigma_{US}^2 + \sigma_Y^2 + 2\sigma_{US,Y}$.

The objective function can be re-written as a generalisation of mean-variance preferences. First, write expected utility in terms of the density function of W_k,

$$\mathbb{E}\left[-e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})}\right] = \int_{-\infty}^{\infty} -e^{-\gamma_k (W_k + \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma})} \frac{1}{\sigma\sqrt{2\pi}} e^{-\frac{W_k - \mu_W}{2\sigma_{W_k}^2}} dW_k$$
$$= e^{-\gamma_k \psi_k \frac{b_{US,k}^{1-\sigma}}{1-\sigma}} \int_{-\infty}^{\infty} -e^{-\gamma_k W_k} \frac{1}{\sigma\sqrt{2\pi}} e^{-\frac{W_k - \mu_W}{2\sigma_{W_k}^2}} dW_k$$

Now, following the same steps as the derivation of standard mean-variance preferences by collecting the terms under the integral that depend on W_k,

$$\begin{aligned}
\mathbb{E}\left[-e^{-\gamma_k(W_k+\psi_k\frac{b_{US,k}^{1-\sigma}}{1-\sigma})}\right] &= e^{-\gamma_k\psi_k\frac{b_{US,k}^{1-\sigma}}{1-\sigma}}\int_{-\infty}^{\infty} -e^{-\gamma_k\left(\mu_W-\frac{\gamma_k}{2}\sigma_{W_k}^2\right)}\frac{1}{\sigma\sqrt{2\pi}}e^{-\frac{\left(W_k-\mu_W+\gamma_k\sigma_{W_k}^2\right)^2}{2\sigma_{W_k}^2}}dW_k \\
&= e^{-\gamma_k\left(\mu_W-\frac{\gamma_k}{2}\sigma_{W_k}^2+\psi_k\frac{b_{US,k}^{1-\sigma}}{1-\sigma}\right)}\underbrace{\int_{-\infty}^{\infty}\frac{1}{\sigma\sqrt{2\pi}}e^{-\frac{\left(W_k-\mu_W+\gamma_k\sigma_{W_k}^2\right)^2}{2\sigma_{W_k}^2}}dW_k}_{=1} \\
&= e^{-\gamma_k\left(\mu_W-\frac{\gamma_k}{2}\sigma_{W_k}^2+\psi_k\frac{b_{US,k}^{1-\sigma}}{1-\sigma}\right)}.
\end{aligned}$$

It follows that

$$\max_{b_{US,k}}\mathbb{E}\left[-e^{-\gamma_k(W_k+\psi ln(B^{US}))}\right] = \max_{b_{US,k}}\mu_W-\frac{\gamma_k}{2}\sigma_{W_k}^2+\psi_k\frac{b_{US,k}^{1-\sigma}}{1-\sigma}.$$

Therefore, maximising expected utility with an exponential utility function in wealth and US Treasuries reduces to standard mean-variance preferences with an additive term for US Treasury holdings, which is increasing and concave due to the CES specification for Treasuries in the utility function.

1.C Logarithmic preferences for Treasuries

Consider the preferences for investors introduced in Section 1.2 where $\sigma \to 1$, so that term for US Treasuries in investors' utility is logartihmic. In this case, an analytical nonlinear solution for the optimal portfolio share exists. The problem of sector k investor is

$$\begin{aligned}
&\max_{s_{US,k}}\mathbb{E}[W_{0,k}]-0.5\gamma_k\mathbb{V}[W_k]+\psi_k\ln(b_{US,k}) \\
\text{s.t.}\quad &W_k = W_{0,k}(R_j+(R_{US}-R_j)s_{US,k})+Y_k \\
&s_{j,k}+s_{US,k}=1,
\end{aligned}$$

The following sub-sections derive propositions on the optimal share and on its sensitivity to the mean and variance excess returns, showing how the results in Section 1.2.1 extend to a nonlinear setting.

1.C.1 Proposition 2: optimal portfolio share

Proposition 2 (Optimal portfolio share). (i) *The optimal portfolio share is*

$$s_{US,k} = \frac{\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]} + \frac{\sqrt{(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k])^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j]}}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]}. \tag{1.14}$$

(ii) $s_{US,k} \in \mathbb{R}^+$ *if* $\gamma_k > 0$, $\psi_k > 0$, *and* $\mathbb{V}[R_{US} - R_j] > 0$.

Proof. Pproposition 2 (i): substitute the constraints in the objective function to re-cast the problem with $s_{US,k}$ as a choice variable. Take the first-order condition for $s_{US,k}$ to obtain the following quadratic equation:

$$\gamma_k \mathbb{V}[R_{US} - R](s_{US,k})^2 - (\mathbb{E}[R_{US} - R] - \gamma_k Cov[R_{US} - R, Y_k])\, s_{US,k} - \psi = 0.$$

The two solutions are

$$s_{US,k,1} = \frac{\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]} - \frac{\sqrt{(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k])^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j]}}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]}$$

and

$$s_{US,k,2} = \frac{\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]} + \frac{\sqrt{(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k])^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j]}}{2\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]}.$$

To select a solution, consider the conditions for $s_{US,k,1}, s_{US,k,2} > 0$, for for $\gamma_k > 0$, $\psi_k > 0$, and $\mathbb{V}[R_{US} - R_j] > 0$.

$$\begin{aligned} s_{US,k,1} > 0 \iff & \quad \mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k] \\ & > \sqrt{(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k])^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j]} \\ \iff & \quad 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j] < 0 \end{aligned}$$

There is no solution for $\gamma_k > 0$, $\psi_k > 0$, and $\mathbb{V}[R_{US} - R_j] > 0$. Therefore, I choose the solution $s_{US,k} = s_{US,k,2}$, and derive the conditions for $s_{US,k,2} \in \mathbb{R}^+$ in the next proof.

Pproposition 2 (ii): start with the conditions for $s_{US,k} \in \mathbb{R}$:

$$s_{US,k} \in \mathbb{R} \iff \left(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]\right)^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R] > 0.$$

It is immediate to see that it always holds for $\gamma_k > 0$, $\psi_k > 0$, and $\mathbb{V}[R_{US} - R_j] > 0$. Now consider the condition for $s_{US,k} > 0$.

$$\begin{aligned} s_{US,k} > 0 \iff & \mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k] \\ & < \sqrt{\left(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]\right)^2 + 4\gamma_k \psi_k \mathbb{V}[R_{US} - R_j]} \\ \iff & 4\gamma_k \psi \mathbb{V}[R_{US} - R_j] > 0 \end{aligned}$$

The condition also always holds for $\gamma_k > 0$, $\psi_k > 0$, and $\mathbb{V}[R_{US} - R_j] > 0$.

$\square$

Note that, for $\gamma_k > 0$, $\psi_k > 0$, and $\mathbb{V}[R_{US} - R_j] > 0$, $s_{US,k} > s_{US,k}|_{\psi=0}$. In the logarithmic case, investors choose a higher portfolio share than they would absent convenience yields.

Furthermore,

$$\lim_{\psi_k \to 0} s_{US,k} = \frac{1}{\gamma_k W_{0,k} \mathbb{V}[R_{US} - R_j]} \left(\mathbb{E}[R_{US} - R_j] - \gamma_k Cov[R_{US} - R_j, Y_k]\right).$$

Therefore, the optimal share collapses back to the standard case as the weight on Treasury preferences vanishes.

Pproposition 2 (ii) states the conditions under which the insurers' problem admits a real, positive solution for $s_{US,k}$. Note that there are no requirements on the risk-return profile of US Treasuries. Therefore, $s_{US,k} > 0$ even for $\mathbb{E}[R_{US} - R] < 0$ and $Cov[R_{US} - R, Y_k] > 0$ simultaneously. Due to the convenience yield of Treasuries, investors choose to hold a positive amount even if they offer neither an extra return, nor good insurance for income risk.

1.C.2 Proposition 3: sensitivity to excess returns

Proposition 3 (Sensitivity to the mean of excess returns). (i)

$$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} = \frac{1}{2}\frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]} \left(1-\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k]}{\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}\right) \tag{1.15}$$

(ii) *With* $\gamma_k > 0$*,* $\psi_k > 0$*, and* $\mathbb{V}[R_{US}-R_j] > 0$*,*
$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} \in \left(0, \frac{1}{2}\frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]}\right)$ *for* $\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k] > 0$*.*
$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} \in \left(\frac{1}{2}\frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]}, \frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]}\right)$ *for* $\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k] < 0$

Proof. Pproposition 3 (i): it follows immediately from differentiating $s_{US,k}$ with respect to $\mathbb{E}[R_{US}-R_j]$. □

Pproposition 3 (ii):

Proof.

$$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} > 0 \iff$$
$$\mathbb{E}[R_{US_j}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k] <$$
$$\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}$$

This condition holds for $\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US},Y_k] > 0$, with $\gamma_k > 0$ and $\mathbb{V}[R_{US}-R_j] > 0$.

$$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} < \frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]} \iff$$

$$\frac{1}{2}\frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]}\left(1-\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k]}{\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}\right)$$
$$< \frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]} \iff$$

$$\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k]$$
$$< \sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k])^2+4\gamma_k\psi\mathbb{V}[R_{US}-R_j]}$$
$$\iff 4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j] > 0.$$

This condition always holds for $\gamma_k > 0$ and $\mathbb{V}[R_{US}-R_j] > 0$.

$$\frac{\partial s_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} > \frac{1}{2}\frac{1}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]} \iff$$

$$-\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k]}{\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}} > 0 \iff$$

$$\mathbb{E}[R_{US}-R_j]-\gamma_k Cov[R_{US}-R_j,Y_k] < 0.$$

$\square$

1.C.3 Proposition 4: sensitivity to variance of excess returns

Proposition 4 (Sensitivity to the variance of excess returns). (i)

$$\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]} = \frac{1}{2W_{0,k}\mathbb{V}[R_{US}-R_j]^2}$$
$$\Bigg(- \frac{\sqrt{\left(\mathbb{E}[R_{US}-R_j] - \gamma_k Cov[R_{US}-R_j, Y_k]\right)^2 + 4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}{\gamma_k}$$
$$+ \frac{2\psi_k\mathbb{V}[R_{US}-R_j]}{\sqrt{\left(\mathbb{E}[R_{US}-R_j] - \gamma_k Cov[R_{US}-R_j, Y_k]\right)^2 + 4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}$$
$$+ Cov[R_{US}-R_j, Y_k] - \frac{\mathbb{E}[R_{US}-R_j]}{\gamma_k}\Bigg) \tag{1.16}$$

(ii) $\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]} < 0$ *for* $\gamma_k > 0$, $\psi_k > 0$, *and* $\mathbb{V}[R_{US}-R_j] > 0$

(iii) *With* $\gamma_k > 0$, $\psi_k > 0$, *and* $\mathbb{V}[R_{US}-R_j] > 0$,
$\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]}|_{\psi_k=0} < \frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]}|_{\psi_k>0}$ *for* $\mathbb{E}[R_{US}-R_j] - \gamma_k Cov[R_{US}-R_j, Y_k] > 0$

Proof. Pproposition 4 (i): it follows immediately from differentiating $s_{US,k}$ with respect to $\mathbb{V}[R_{US}-R_j]$. □

Proof. Pproposition 4 (ii):

$$\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]} < 0 \iff$$

$$-\frac{\sqrt{\left(\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]\right)^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}{\gamma_k}$$
$$+\frac{2\psi_k\mathbb{V}[R_{US}-R_j]}{\sqrt{\left(\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]\right)^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}$$
$$+\text{Cov}[R_{US}-R_j,Y_k]-\frac{\mathbb{E}[R_{US}-R_j]}{\gamma_k} < 0 \iff$$

$$-\left(\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]\right)^2-2\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]-$$
$$\left(\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]\right)$$
$$\sqrt{\left(\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]\right)^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]} < 0 \iff$$

$$4\left(\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]\right)^2 > 0$$

This condition always holds.

□

Proof. Pproposition 4 (iii):
For $\psi_k = 0$,

$$\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]}|_{\psi_k=0} = -\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k \text{Cov}[R_{US}-R_j,Y_k]}{\gamma_k W_{0,k}\mathbb{V}[R_{US}-R_j]^2}.$$

Then,

$$\frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]}|_{\psi_k=0} < \frac{\partial s_{US,k}}{\partial \mathbb{V}[R_{US}-R_j]}|_{\psi_k=0} \iff$$

$$2\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k]}{\gamma_k} <$$
$$+\frac{\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}{\gamma_k}$$
$$-\frac{2\psi_k\mathbb{V}[R_{US}-R_j]}{\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]}}$$
$$-\frac{\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k]}{\gamma_k} \iff$$

$$(\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k])$$
$$\sqrt{(\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k])^2+4\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]} <$$
$$(\mathbb{E}[R_{US}-R_j]-\gamma_k \mathrm{Cov}[R_{US}-R_j,Y_k])^2+2\gamma_k\psi_k\mathbb{V}[R_{US}-R_j] \iff$$

$$4\left(\gamma_k\psi_k\mathbb{V}[R_{US}-R_j]\right)^2 > 0$$

This condition always holds. □

Proposition ii shows that, even in the nonlinear solution, the derivative of the Treasury portfolio share to the variance of excess returns is negative. Furthermore, proposition iii confirms that the presence of convenience yields makes the reaction to market volatility more muted, that is less negative, compared to the standard mean-variance preferences case.

1.D Proofs

1.D.1 Proposition 1: reaction of excess returns to euro area debt supply

Proof. Proposition 1:

Differentiating equation 1.5 with respect to B_j, taking into account that $b_{US,k}$ is a function of B_j through excess returns only,

$$\frac{\partial \mathbb{E}\left[R_{US}-R_j\right]}{\partial B_j} = -\frac{\mathbb{V}[R_{US}-R_j]}{\sum_k \tau_k} + \frac{1}{\sum_k \tau_k}\frac{\partial b_{j,O}}{\partial B_j}$$
$$+\sigma\frac{\sum_k \tau_k \psi_k b_{US,k}^{-\sigma-1}\left(\frac{\partial b_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]}\frac{\partial \mathbb{E}[R_{US}-R_j]}{\partial B_j}\right)}{\sum_k \tau_k}$$

To find $\partial b_{US,k}/\partial \mathbb{E}[R_{US}-R_j]$, differentiate equation 1.1 with respect $\mathbb{E}[R_{US}-R_j]$, applying the implicit function theorem:

$$1-\gamma_k \mathbb{V}[R_{US}-R_j]\frac{\partial b_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} - \sigma\psi_k b_{US,k}^{-\sigma-1}\frac{\partial b_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} = 0 \iff$$

$$\frac{\partial b_{US,k}}{\partial \mathbb{E}[R_{US}-R_j]} = \frac{1}{\mathbb{V}[R_{US}-R_j]/\tau_k + \sigma\psi_k b_{US,k}^{-\sigma-1}}.$$

Then, substitute $\partial b_{US,k}/\partial \mathbb{E}[R_{US}-R_j]$ back into the expression for $\partial \mathbb{E}\left[R_{US}-R_j\right]/\partial B_j$

$$\frac{\partial \mathbb{E}\left[R_{US}-R_j\right]}{\partial B_j} = -\frac{\mathbb{V}[R_{US}-R_j]}{\sum_k \tau_k} + \frac{1}{\sum_k \tau_k}\frac{\partial b_{j,O}}{\partial B_j}$$
$$+\sigma\frac{\sum_k \frac{\tau_k \psi_k b_{US,k}^{-\sigma-1}}{\mathbb{V}[R_{US}-R_j]/\tau_k+\sigma\psi_k b_{US,k}^{-\sigma-1}}}{\sum_k \tau_k}\frac{\partial \mathbb{E}\left[R_{US}-R_j\right]}{\partial B_j} \iff$$

$$\frac{\partial \mathbb{E}\left[R_{US}-R_j\right]}{\partial B_j} = \frac{\mathbb{V}[R_{US}-R_j]\left(\frac{\partial b_{j,O}}{\partial B_j}-1\right)}{\sum_k \tau_k\left(1-\frac{\sigma\psi_k b_{US,k}^{-\sigma-1}}{\mathbb{V}[R_{US}-R_j]/\tau_k+\sigma\psi_k b_{US,k}^{-\sigma-1}}\right)}$$

To prove that $\partial \mathbb{E}\left[R_{US}-R_j\right]/\partial B_j < 0$, note that $\partial b_{j,O}/\partial B_j \in [0,1]$, so

$$\frac{\partial b_{j,O}}{\partial B_j} - 1 < 0$$

and

$$\frac{\partial \mathbb{E}\left[R_{US}-R_j\right]}{\partial B_j} < 0 \iff \sum_k \tau_k\left(1-\frac{\sigma\psi_k b_{US,k}^{-\sigma-1}}{\mathbb{V}[R_{US}-R_j]/\tau_k+\sigma\psi_k b_{US,k}^{-\sigma-1}}\right) > 0$$

A sufficient condition to satisfy this equaiton is

$$\tau_k \left(1 - \frac{\sigma\psi_k b_{US,k}^{-\sigma-1}}{\mathbb{V}[R_{US} - R_j]/\tau_k + \sigma\psi_k b_{US,k}^{-\sigma-1}}\right) > 0 \,\forall\, k \iff$$

$$\sigma\psi_k b_{US,k}^{-\sigma-1} < \mathbb{V}[R_{US} - R_j]/\tau_k + \sigma\psi_k b_{US,k}^{-\sigma-1} \iff$$

$$\mathbb{V}[R_{US} - R_j]/\tau_k > 0.$$

This condition always holds for $\tau_k > 0$ and $\mathbb{V}[R_{US} - R_j] > 0$. □

1.E Details on the recovery of structural parameters

1.E.1 Calibration

To solve for the structural parameters, I first proxy $\mathbb{V}[R_{US} - R_j]$ and $Cov[R_{US} - R_j, Y_k]$ with their empirical counterparts $\hat{\varsigma}^2_{ER}$ and $\hat{\varsigma}_{ER,Y_k}$. Likewise, initial wealth $W_{0,k}$ is measured as the total holdings of US and country j government bonds, averaged over the country-quarter distribution, $\bar{W}_{0,k}$. Note that the unit of measure of debt holdings is not specified in the theoretical model, so I calibrate it to hundreds of billions of euros so that the order of magnitude of model-implied excess returns matches the average of the empirical proxy of excess returns, $er_{j,t}$, as reported in Table 1.1.

I also convert the coefficient $-\hat{\pi}$ from the empirical model to account for the minus sign with respect to the theoretical model, and for the standardisation of PSPP purchases in the empirical model, such that the parameter $\tilde{\pi}$ used in the solution for structural parameters is

$$\tilde{\pi} = -\hat{\pi}\frac{\bar{B}}{\hat{\sigma}_{PSPP}}$$

where $\bar{B}$ is average outstanding government debt for euro area countries and $\hat{\varsigma}_{PSPP}$ is the sample standard deviation of PSPP purchases.

Finally, I estimate $\partial b_{j,O}/\partial B_j$ via an absorption regression. I start by breaking down the outstanding amount of government debt for country j in quarter t into holdings by four sectors: banks, insurance companies, the ECB, and other investors. The amount held by other investors is defined residually, while the amount held by the ECB includes only PSPP holdings. It is crucial to account for holdings by the ECB for this decomposition to map correctly to the model, as PSPP purchases are used to identify the parameter π and so should be excluded from the amount held by oher investors.

I then estimate the following regression separately for each sector on the same country-quarter panel used in the regressions in Section 1.3.4.

$$b_{j,k,t} = \zeta_j + \zeta_t + \theta_k B_{j,t} + \upsilon_{j,k,t}$$

where ζ_j and ζ_t are country and quarter fixed effects, added to account for common macroeconomic conditions and country-specific idiosyncracies. As shown in Table 1.7, coefficients θ_k of the absorption regression add up to 1 across sectors because sectoral holdings $b_{j,k,t}$ sum up to the total amount outstanding $B_{j,t}$ in each quarter. I use the estimated coefficient for other investors, from the model with time and country fixed effects, $\hat{\theta}_O$, as a proxy for $\partial b_{j,O}/\partial B_j$.

Table 1.7. Absorption regression

	(1)	(2)	(3)	(4)
$b_{j,B,t}$	0.19***	0.19***	0.02	0.04**
	(0.01)	(0.01)	(0.02)	(0.02)
$b_{j,I,t}$	0.22***	0.22***	0.16***	0.18***
	(0.01)	(0.01)	(0.02)	(0.01)
$b_{j,PSPP,t}$	0.19***	0.18***	0.63***	0.57***
	(0.01)	(0.01)	(0.05)	(0.04)
$b_{j,O,t}$	0.41***	0.41***	0.19***	0.20***
	(0.01)	(0.01)	(0.05)	(0.04)
Time fixed effects	No	Yes	No	Yes
Country fixed effects	No	No	Yes	Yes

Coefficients from regression model $b_{j,k,t} = \zeta_j + \zeta_t + \theta_k B_{j,t} + \upsilon_{j,k,t}$ estimated via OLS. Each row reports the estimates of coefficient θ_k for a different sector k. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 1.8 displays the values of the calibrated parameters.

1.E.2 Solution algorithm

After replacing all estimates and calibrated parameters, I am left with the following system of five equations in five unknowns: γ_k and ψ_k for $k = \{B, I\}$, and σ.

Table 1.8. Calibrated parameters

Parameter	Unit of measure	Value
$\hat{\varsigma}^2_{ER}$	N.A.	0.81
$\hat{\varsigma}_{ER_j,Y_B}$	N.A.	-0.04
$\hat{\varsigma}_{ER_j,Y_I}$	N.A.	0.03
$\bar{W}_{0,B}$	€ 100 bn.	2.56
$\bar{W}_{0,I}$	€ 100 bn.	1.63
$\bar{B}_j$	€ 100 bn.	5.16
$\hat{\varsigma}_{PSPP}$	€ 100 bn.	1.54

$$\hat{\alpha}_k = \bar{s} - \gamma_k \frac{\bar{W}_{0,k}\bar{s}\left(\hat{\varsigma}^2_{ER} - \bar{v}\right) + \hat{\varsigma}_{ER_j,Y_k} - \bar{c}}{\gamma_k \bar{W}_{0,k}\bar{v} + \sigma\psi_k \bar{W}_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}} \quad \text{for } k = \{B, I\}, \tag{1.17}$$

$$\hat{\beta}_k = \frac{1}{\gamma_k \bar{W}_{0,k}\bar{v} + \sigma\psi_k \bar{W}_{0,k}^{-\sigma}\bar{s}^{-\sigma-1}} \quad \text{for } k = \{B, I\}, \tag{1.18}$$

$$\hat{\pi} = \frac{\bar{v}\left(\hat{\theta}_O - 1\right)}{\sum_k \frac{1}{\gamma_k}\left(1 - \frac{\sigma \bar{b}_{US,k}^{-\sigma-1}\psi_k}{\gamma_k\bar{v} + \sigma\psi_k \bar{b}_{US,k}^{-\sigma-1}}\right)}. \tag{1.19}$$

To solve it, I exploit the structure of the system whereby equations 1.17 and 1.18, derived from the portfolio choice of each investor sector, depend only on the risk aversion and Treasury preference parameter for that sector, γ_k and ψ_k, and on the common Treasury preference curvature parameter σ. On the other hand, equation 1.19, derived from equilibrium excess returns, depends on σ and on the preference parameters of all investors. I can then use the following iterative algorithm:

1. Pick a starting value σ_0.
2. Solve equations 1.17 and 1.18 for γ_k and ψ_k separately for each sector letting $\sigma = \sigma_0$, obtaining solutions $\gamma_{k,0}$ and $\psi_{k,0}$ $\forall\, k$.
3. Solve equation 1.19 for σ letting $\gamma_k = \gamma_{k,0}$ and $\psi_k = \psi_{k,0}$ $\forall\, k$, obtaining solution σ_1.
4. Let $\sigma = \sigma_1$ in step 1 and iterate until convergence for γ_k, ψ_k and σ.

After solving for the structural parameters, I simulate their distribution to obtain empirical confidence intervals. I start by drawing 100,000 times from the joint asymptotic normal distribution of $\boldsymbol{\lambda} \coloneqq (\hat{\alpha}_B, \hat{\beta}_B, \hat{\alpha}_I, \hat{\beta}_I, \hat{\pi}, \hat{\theta}_O)$, assuming independent coefficients across regressions, but accounting for the correlation of within portfolio equations. Then, for each joint draw I solve for the structural parameters as outlined above. Then, I calculate the mean, 5^{th} and 95^{th} percentile for γ_k, ψ_k and σ over all values of $\boldsymbol{\lambda}$ that admit a positive solution of equations 1.17, 1.18 and 1.19. I then use the 5^{th} and 95^{th} percentiles of the distribution of solutions as bounds for the simulated confidence intervals.

Chapter 2

Treasury supply, relative convenience yields and exchange rates

The fiscal sustainability of US public debt depends crucially on the con- venience yield, the premium that investors pay to hold US Treasury. The- oretically, government debt in equilibrium is negatively associated with the convenience yield, which is also linked to the exchange rate through interest parity. However, the existing literature offers only correlational evidence, disregarding the active choice of debt issuance by the government. Using a simple open-economy model with optimal debt supply and liquidity prefer- ence for Treasuries, we show that outward shifts in debt supply reduce the convenience yield through dollar depreciation. Conversely, changes in liquid- ity preference generate positive comovements between debt supply, currency appreciation, and convenience yields. As a result, estimation strategies, like OLS that fail, to disentangle Treasury supply and demand shocks result in an understatement of the yield elasticity of Treasury demand, and of the impact of Treasury supply shocks on exchange rates. We confirm the predicions of our model via local projections using an instrument based on Treasury fu- tures price changes following auction announcements. An unexpected rise in US Treasury supply lowers the convenience yield and depreciates the dollar against G10 currencies by up to three times more than previously esimated.

This chapter is co-authored with Maxime Phillot. Both authors provided equal contributions.

2.1 Introduction

In times of historically high government debt, when the US government relies heavily on debt financing to support massive spending in the wake of the COVID-19 pandemic, understanding the simultaneous impact of debt issuances on exchange rates and the convenience yield *vis-à-vis* foreign government bonds takes on paramount importance.

The issuance of US Treasury debt—the global safe asset—is not a mere domestic policy decision; its effects being felt globally, it is a key element of global financial stability. Changes in Treasury yields impact the perceived risk-free rate, triggering adjustments in financial asset valuations worldwide, as highghlited in the Global Financial Cycle literature (e.g., Miranda-Agrippino and Rey, 2020; Miranda-Agrippino and Rey, 2022). Simultaneously, fluctuations in exchange rates due to Treasury supply shocks have implications for global trade and investment flows, affecting the competitive position of US exports and the cost of dollar-denominated debt for emerging economies.

From a policy standpoint, these dynamics are even more critical. As elucidated in Van Nieuwerburgh et al. (2021), the fiscal sustainability of the growing US public debt relies crucially on the willingess of foreign investors to pay a premium for US Treasuries. Therefore, assessing fiscal capacity in the US requires appropriately estimating the yield sensitivity of investor's demand, in particular in response to an increase in public debt. Likewise, the ability of the US Treasury to fulfill its mandate—to issue debt at the lowest cost to the government—also hinges on it capacity to anticipe market reactions accurately.

Yet, studies on the relationship between Treasury supply, convenience yields and exchange rates have focused almost exclusively on the demand side. Empirical analyses routinely treat the observable outstanding amount of US Treasuries as an exogenous variable (e.g., Du et al., 2018), while theories linking convenience yields and exchange rates tend to take the supply of debt either as fixed, or changing automatically in response to a postulated tax rule (e.g. Valchev, 2020).

However, governments can and do take into account price incentives when deciding the quantity of debt to issue. Within the bounds imposed by their intertemporal budget constraint, lower interest rates can induce tilting government funding towards debt, especially if Ricardian equivalence does not hold. In fact, the low-interest rate environment of the 2010s sparked a lively academic debate on the incentives for government to issue debt to finance expenditure and roll-over previous borrowing, and on the sustainability thereof[1].

The convenience yield itself is a determinant of the attractiveness of debt issuance: Choi et al. (2022) show that if foreigners derive a non-monetary payoff from holding US Treasuries, the difference between the yield of Treasuries and that of corporate bonds represents the marginal benefit in the optimality condition for debt in the government's Ramsey problem. They also provide evidence that the US acts as a monopolist, exploiting market power and hence restricting the global safe asset supply. While we do not explore strategic behaviour in this paper, this line of argument does lend additional support to the idea that the debt supply decision is not divorced from price considerations.

In this paper, we aim to construct a unified theoretical framework that can effectively analyze the dynamics between Treasury supply, the convenience yield, and exchange rates. We then propose to empirically test the key implications of this framework. This dual approach allows us not only to deepen our understanding of the impact of Treasury debt issuance, but also to validate these insights against empirical evidence.

A simple, deterministic two-country general equilibrium model illustrates the interplay between supply and demand of US Treasuries and frames our empirical analysis. US Treasuries offer a non-monetary liquidity payoff that generates an endogenous and time-varying convenience yield relative to foreign bonds. The convenience yield adjusts to changes in debt supply and liquidity preferences in equilibrium through the nominal exchange rate. The first testable implication of our model is that an increase in debt supply causes a drop in the convenience yield and an immediate depreciation of the US dollar, followed by an appreciation.

We also analyse the Ramsey-optimal choice of government bond issuance, which depends on the marginal cost of issuing debt and on its marginal benefit, which equals the convenience yield. As a result, the debt supply schedule is upward-sloping in the convenience yield, which introduces a positive correlation between debt supply, convenience yields and exchange rates in response to higher liquidity preference. The second testable implication of our model is that any regression of convenience yields and exchange rates on measures of Treasury supply will suffer from endogeneity, as it cannot be determined *a priori* whether the observed price-quantity pairs are the results of shifts in the demand or the supply curves. As a consequence, the estimated coefficient on the quantity of debt will be attenuated

[1] See for example Blanchard (2019), Valchev (2020)Jiang et al. (2019), Reis (2021), Mehrotra and Sergeyev (2021), Brunnermeier et al. (2022b) Mankiw (2022)

towards zero.

We address this endogeneity issue with a high-frequency instrumental variable *à la* Phillot (2024). The intuition behind the instrument is the following: if the futures market prices in all relevant information, as per the efficient market hypothesis (Malkiel and Fama, 1970), changes in Treasury futures prices in a tight window around auction announcements are caused by an unexpected changes in the supply of Treasuries. The strategy is complementary to that of Gorodnichenko and Ray (2017), who use instead futures price changes around the auction itself to tease out Treasury demand shocks.

Firstly, we examine the immediate impact of unexpected changes in the outstanding amount of US Treasuries on the daily US dollar exchange rate and the convenience yield relative to other G10 currencies between February 2001 and January 2020 via a 2SLS procedure. First-stage statistics validate our set of instruments both in terms of relevance and overidentifying restrictions. Second-stage results indicate that a median-sized US Treasury supply shock translates into a same-day decrease of the US relative convenience yield of about 0.65 basis points and a depreciation of the US dollar of about 19 basis points. Notably, the effects otherwise reported by OLS are much smaller and not statistically significant.

Secondly, we investigate the evolution and persistence of these effects. A local-projection instrumental-variable model shows that the drop in the convenience yield reaches up to 3 basis points and is persistent, i.e. statistically significant, over a 12-week horizon. On the other hand, the US dollar depreciation documented in the daily exercise lasts for a week and reverses into a statistically significant appreciation, reaching more than 50 basis points four weeks after the impact before vanishing four weeks later.

Finally, to illustrate how our instrumental variable approach solves the endogeneity problem that emerges naturally in our theoretical framework, we replicate the panel-data analysis of the relationship between the outstanding amount of US Treasuries and the convenience yield in Du et al. (2018). Our IV approach documents that a one percentage-point increase in US debt-to-GDP causes a 2.90 basis points decrease in the 5-year US relative convenience yield. Importantly, this coefficient is two times as large as its OLS equivalent.

Together, these empirical findings corroborate both of the testable implications of our model. First, truly unexpected increases in the outstanding amount of US Treasury debt do seem to cause an immediate US dollar depreciation followed by an appreciation, as well as a decline in the US convenience yield, relative to a panel of G10 currencies. Second, the downward bias OLS estimates exhibit in all three approaches is consistent with the presence of a positive correlation between Treasury quantity and convenience yield introduced by demand shocks along an upwards-sloping supply curve, which is left unaccounted for in the absence of a clean identification strategy for Treasury supply shocks.

Related Literature

A long-standing literature analyses the theoretical foundations of the observed premium, or convenience yield, of US Treasuries with respect to various comparable assets (Longstaff, 2004; Krishnamurthy and Vissing-Jorgensen, 2012; Nagel, 2016), motivating a downward-sloping demand curve for US Treasuries. We follow their approach by modeling the convenience yield with an additional term in households' utility function that depends on Treasury holdings. More recently, a series of papers provides theoretical frameworks linking convenience yields and exchange rate dynamics through demand-side effects (Engel and Wu, 2018; Jiang et al., 2023a; Kekre and Lenel, 2021; Jiang et al., 2021). Our studies contributes to this strand of the literature by highlighting the role of an upward-sloping supply curve of US Treasuries.

There is also a related literature on the optimal supply of government debt, modelling the benefit to households via a variety of mechanisms such as collateral constraints and liquidity (Aiyagari and McGrattan, 1998; Woodford, 1990; Angeletos et al., 2016). We contribute by building a model in which the benefit is motivated by a different channel: issuing debt frees up resources, previously tied up in taxation, to invest in foreign bonds, which pay a higher yield due to the liquidity payoff of Treasuries enjoyed by foreign households.

The papers closest to our theoretical model are Valchev (2020) and Choi et al. (2022). The former shows that time-varying convenience yields arise in a simple endowment economy with bonds in the utility function, and that monetary-fiscal policy interactions generate non-monotonic dynamics in the exchange rate. The demand-side of our theoretical model is similar, but we restrict US household to hold only foreign bonds. The most significant difference arises from the government debt supply side. Valchev (2020) imposes a linear rule for taxes, which then implies a given amount of bonds through the budget constraint. On the other hand, we solve for the bond supply curve deriving from the Ramsey problem of the government. Furthermore, the empirical section of the paper uses raw outstanding amounts of US Treasuries in regressions of convenience yields and exchange rates, subjecting the results to the threat of debt supply endogeneity.

Similarly to our paper, Choi et al. (2022) use a model in which optimal choice of government debt issuance results in an upward-sloping supply curve of US Treasuries. The marginal benefit for the government in their setup is however the Treasury premium with respect to dollar-denominated corporate bonds, instead of foreign government bonds as in our model. Furthermore, we study the dynamics of exchange rates, while Choi et al. (2022) focus only on the real implications of under-provision of safe assets in a regime of monopolistic supply.

Other papers have investigated empirically the interplay between Treasury supply, relative convenience yields and exchanges rates.[2] Du et al. (2018) propose a measure of relative convenience yields based on Treasury yield covered interest rate (CIP) deviations and find that it decreases when government bond supply increases. Engel and Wu (2018) contend that relative convenience yields are significantly correlated with G10 currency fluctuations. Krishnamurthy and Lustig (2019) find that safe dollar asset supply and demand affect the dollar exchange rate, bond yields, and other aspects of the global financial system.

Our paper builds upon this set of empirical studies by invoking a cleaner identification of Treasury supply, borrowed from Phillot (2024). The latter relates to the well-established literature that aims at identifying macroeconomic "random causes" (Slutsky, 1937), i.e., drivers of business cycle fluctuations (see Ramey, 2016, for a review of the literature on structural shock identification). Phillot (2024) proposes a so-called high-frequency identification strategy of US Treasury supply shocks, exploiting the design of US Treasury auctions. Much like the literature that identifies monetary policy shocks (Kuttner, 2001; Gürkaynak et al., 2005, among others), he interprets changes in US Treasury futures prices around announcements by the US Treasury as surprises about the supply of US debt securities.

We build upon Phillot (2024), whose focus is exclusively on US *domestic* financial outcomes and implement similar local projections (Jordà et al., 2020, 2015; Jordà, 2005) by considering exchange rates and convenience yields as dependent variables to explore the effects of US Treasury supply shocks on *global* macro-financial outcomes. By investigating transmission mechanisms between US Treasury supply shocks and global financial markets, and introducing a theoretical framework to understand these connections, we go beyond a mere addition of exchange rates

[2]Note that our paper does not study the relationship between fiscal policy and exchange rates (see, e.g., Monacelli and Perotti, 2010; Ravn et al., 2012; Alberola-Ila et al., 2021). Rather, we evaluate solely shocks to the funding composition of US debt and consider the nominal exchange rate as opposed to the real exchange rate.

and convenience yields to his approach.

In a replication of Du et al. (2018) from a Swiss perspective, Benhima and Phillot (2023) report that the OLS supply price elasticity of Swiss relative convenience yields is underestimated by a factor of three relative to an equivalent instrumental approach based on Swiss auction data. Our estimates of this bias in the United States confirm their findings.

The remainder of this paper is organized as follows. Section 2.2 builds a theoretical framework to illustrate the interplay between supply and demand of US Treasuries and provide testable implications. Section 2.3 investigates this relationship empirically and revisits past evidence using an identification technique based on high-frequency changes in Treasury futures prices surrounding US Treasury announcements. Section 2.4 concludes.

2.2 Theoretical Model

2.2.1 The Setup

The model features two countries: the US, indexed H and the rest of the world (henceforth RoW), indexed F. The environment is deterministic and time is discrete and infinite. Consumers in either economy are endowed with real amounts of an undifferentiated good, with price P_t (P_t^*) in the US (RoW) currency.[3] The law of one price holds with $P_t = S_t P_t^*$ where S_t is the US dollar price of one unit of RoW currency (the US dollar depreciates when S_t increases).

Consumers choose consumption and investment in real government bonds. The US representative household can purchase only foreign bonds, while the RoW household can purchase both foreign and US bonds. This assumption is not meant to represent the actual set of assets available to US investors, but rather a snapshot of the external assets and liabilities position of the country as a whole. Thanks to the status of Treasuries as safe assets, the US can invest at a high yield while borrowing at lower rates, as highlighted by Gourinchas et al. (2010). In our model, this "exorbitant privilege" stems from foreign households deriving a non-monetary payoff from holding US Treasuries. Following Krishnamurthy and Vissing-Jorgensen (2012), we model the non-monetary payoff as an additional term in RoW households' utility function, which is meant to capture special liquidity or safety char-

[3] Hereafter, we denote with superscripts "$*$" variables pertaining to RoW.

acteristics. [4]

The US government finances a fixed amount of spending with a mix of lump-sum taxes levied on US households, and bonds purchased by foreign households. Following Choi et al. (2022), the government solves a Ramsey problem with a convex debt issuance cost to choose the optimal amount of debt.

RoW Households

The problem of the RoW household is

$$\max_{C_t^*, B_{H,t}^*, B_{F,t}^*} \sum_{s=0}^{\infty} \beta^s [U(C_{t+s}^*) + V(B_{H,t+s}^*)]$$

$$\text{s.t.} \quad C_t^* + B_{F,t}^* + B_{H,t}^* = \left(\frac{S_t}{S_{t-1}}\right)^{-1} \frac{(1+i_{t-1})}{\Pi_t^*} B_{H,t-1}^* + \frac{(1+i_{t-1}^*)}{\Pi_t^*} B_{F,t-1}^* + Y^*,$$

where C_t^* is consumption, $B_{H,t}^*$ and $B_{F,t}^*$ are real holdings of US and RoW bonds, Y^* is the RoW endowment, $1+i_t$ and $1+i_t^*$ are the US and RoW nominal interest rates, and S_t is the nominal exchange rate in terms of dollars per foreign currency, so that an increase of S_t is a dollar depreciation. $\Pi_t^* = P_t^*/P_{t-1}^*$ is gross inflation. $U(C_t)$ and $V(B_{H,t+s}^*)$ are increasing, concave functions representing the utility of consumption and the non-monetary payoff of US Treasuries. The Euler equations for foreign and domestic bonds, respectively, are

$$U'(C_t^*) = \beta \frac{1+i_t}{\Pi_{t+1}^*} \frac{S_t}{S_{t+1}} U'(C_{t+1}^*) + V'\left(B_{H,t}^*\right),$$

$$U'(C_t^*) = \beta \frac{1+i_t^*}{\Pi_{t+1}^*} U'(C_{t+1}^*).$$

Combining these equations, we obtain a modified uncovered interest parity (UIP) condition

$$\underbrace{\frac{1+i_t^*}{\Pi_{t+1}^*} - \frac{S_t}{S_{t+1}} \frac{1+i_t}{\Pi_{t+1}^*}}_{\equiv \phi_t} = \frac{1}{\beta U'(C_{t+1}^*)} V'\left(B_{H,t}^*\right).$$

The left-hand side of the equation is the conventional UIP condition, which is different from zero because of the liquidity benefit provided by US Treasuries, reflected by $V'\left(B_{H,t}^*\right)$ on the right-hand side. We define the wedge in UIP as ϕ_t and refer to it as convenience yield henceforth. Note that since the model features no risk, the UIP deviation is equivalent to the CIP deviation which we use as a proxy for the convenience yield in the empirical analysis.

[4]This approach is isomorphic to imposing a cash-in-advance constraint (Feenstra (1986)) or transaction costs (Valchev (2020)).

US Government & Fiscal Policy

The US government's budget constraint in real terms is

$$B_t^G + T_t = \bar{G} + \frac{B_{t-1}^G}{\Pi_t}(1+i_{t-1}) + \chi(B_{t-1}^G),$$

where B_t^G is the amount of government debt issued in period t, and $\bar{G}$ is a fixed amount of government spending. In line with the empirical analysis, we focus solely on changes in the *composition* of funding of government spending, rather than changes in spending itself. T_t are lump-sum taxes, which adjust in response to changes in B_t^G to satisfy the budget constraint. $\chi(B_{t-1}^G)$ is a cost function that is increasing and convex in the real amount of debt B_{t-1}^G.

The choice of the cost function $\chi(B_{t-1}^G)$ in the model may appear ad-hoc at first glance. However, it can be justified as a convenient representation of distortionary costs associated with financing debt repayments through taxation. This approach is common in the literature, and is used for example in Choi et al. (2022) and Gorton and Ordonez (2022). Another interpretation of this cost function is that it captures the costs associated with expanding the US government's balance sheet. This route is taken by Hall and Reis (2015) and Greenwood et al. (2016), wherein issuing debt incurs expenses related to interest rate risk and asset purchases. Then, $\chi(B_{t-1}^G)$ in our model can be thought of as a catch-all term for any frictions associated with increasing sovereign debt that are not explicitly modelled.

The government chooses the optimal amount of debt issued B_t^G to maximise the US household's utility, subject to the government's budget constraint and to the household problem's optimality conditions, taking interest rates as given. ,[5]

We show in Appendix 2.A that this Ramsey problem can be formulated as

$$\max_{C_t, A_t, B_t^G} \sum_{s=0}^{\infty} \beta^s U(C_{t+s})$$
$$\text{s.t.} \quad C_t + A_t + \bar{G} = \frac{1+i_{t-1}^*}{\Pi_t}\frac{S_t}{S_{t-1}}A_{t-1} + \phi_{t-1}B_{t-1}^G - \chi(B_{t-1}^G) + Y,$$

where we define $A_t \equiv B_{F,t} - B_t^G$ as net foreign assets. The optimality conditions

[5]Note that, unlike Choi et al. (2022), we assume that the US government is a price taker and does not exploit its monopolistic power to extract a rent from US bond holders. In other words, it does not internalize the effect of B_t^G on ϕ_t. If it did, an under-provision of government bonds would occur, but the mechanisms highlighted in this paper would still hold.

of this problem are

$$U'(C_t) = \beta \frac{1+i_t^*}{\Pi_{t+1}} \frac{S_{t+1}}{S_t} U'(C_{t+1}),$$
$$\phi_t = \frac{\partial \chi(B_t^G)}{\partial B_t^G}.$$

The first condition is the Euler equation for US households. The second one is a static optimality condition stating that the convenience yield ϕ_t equals the marginal cost of issuing a unit of real debt $\frac{\partial \chi(B_t^G)}{\partial B_t^G}$. By issuing a unit of debt, the government commits to pay the US interest rate i_t, but at the same time reduces by one unit the lump sum taxes levied on households. Since US households can invest only in foreign bonds, the marginal opportunity cost of taxation is equal to the RoW interest rate i_t^*. Thus, the difference between the interest rate on RoW and US bonds expressed in the same currency, i.e. the convenience yield, is the relevant marginal benefit for the choice of issuing debt.

Monetary policy

We examine the equilibrium under a fully flexible exchange rate policy with fixed nominal interest rates, so that the convenience yield can adjust through the exchange rate only, thus bringing the core mechanism of the paper into stark relief. In the US, we fix the nominal interest to a constant, arbitrary level

$$i_t = i \quad \forall t.$$

The nominal interest rate does not affect the real rate, which is fixed at $\frac{1-\beta}{\beta} - \frac{V'(B^G)}{\beta U'(C^*)}$ due to constant endowments. Thus, all adjustments in the equilibrium US real interest rate occur through US inflation only. The modified UIP condition on the other hand adjusts through the nominal exchange rate, which is linked to inflation by virtue of the LOP.

In the RoW, the real interest rate is fixed at $\frac{1-\beta}{\beta}$ due to constant endowments. Therefore, by letting the nominal interest rate be

$$i_t^* = \frac{1-\beta}{\beta} \quad \forall t,$$

$\Pi_t^* = 1$ obtains in equlibrium. A consequence is that $\Pi_t = \frac{S_t}{S_t - 1}$, as can be seen by combining the LOP for periods t and $t+1$.

2.2.2 Equilibrium

An equilibrium in this economy is characterized by an allocation $\{C_t, C_t^*, B_{H,t}^*, B_t^G, B_{F,t}, B_{F,t}^*\}$ and prices $\{\Pi_t, \Pi_t^*, S_{t+1}/S_t, i_t, i_t^*\}$ such that

1. Given prices, the allocation satisfies the optimality condition and budget constraint of the RoW household.
2. Given prices, the allocation satisfies the optimality condition and budget constraint of the US government's Ramsey problem.
3. Markets for goods, US government bonds and RoW government bonds clear:
   - Goods: $C_t + C_t^* = Y + Y^* - \bar{G}$,
   - US government bonds: $B_{H,t}^* = B_t^G$,
   - RoW government bonds: $B_{F,t} + B_{F,t}^* = \bar{B}_F$,

where $\bar{B}_F$ is the supply of foreign bonds. Note that in this equilibrium US and RoW consumption C and C^* are constant. Real interest rates are constant because endowments are fixed, so it follows from the US and RoW Euler equation for RoW bonds that $C_t = C_{t+1} = C$, and $C_t^* = C_{t+1}^* = C^*$.

Therefore, the only variables that are not constant over time in equilibrium are US Treasury supply B_t^G, US inflation and the exchange rate, which in turn causes the convenience yield ϕ_t to vary. The rationale for these modelling choices is to provide an environment that is as simple as possible, while maintaining the core mechanisms of liquidity preference for US Treasuries and debt supply choice. As real variables are fixed, with the exception of B_t^G, the model is meant to capture a within-quarter environment. Each period can be conceptualised as a week, in keeping with the timing of the local projections in section 2.3. Note that US Treasury auctions are held at a frequency of several per month, which justifies our assumption of B_t^G varying within a quarter. The equilibrium dynamics are then described by two equations: the modified UIP condition derived from the RoW household's optimisation, and the bond supply schedule implied by the government's Ramsey problem.

We now turn to analyse the effects of shifts in the debt supply curve and in the household's liquidity preference on the convenience yield and exchange rate.[6] In

[6]Note that we consider permanent changes to parameters that result in changes to both *steady-state* values and *equilibrium* values. For all variables that are fixed at the steay state in the equilibrium, such as consumption, the two coincide.

order to obtain an analytical solution of the model and a graphical representation of the equilibrium, we assume the following convenient functional forms

$$U(x) = log(x), \qquad V(x) = \Psi log(x), \qquad \chi(x) = x^2/2 + \nu x,$$

where $\Psi > 0$ is a parameter regulating the relative weight of bonds within the utility function that can be interpreted as liquidity preference, and $\nu > 0$ is a parameter regulating the constant component of the marginal cost of debt issuing. Conceptually, ν plays the same role as the identified Treasury supply shock does in our empirical analysis, with an increase in ν corresponding to a negative supply shock. However, the theoretical model is deterministic, so ν is to be interpreted as a shifter for the supply curve rather than a shock. Note that the results of the model do not hinge on these specific functional forms, but only on $U(x)$ and $V(x)$ being separable, increasing and concave, and on $\chi(x)$ being increasing and convex.

With these functional forms, the US bond supply schedule and the modified UIP condition yield two equilibrium equations

$$\phi_t = B_t^G + \nu, \tag{2.1a}$$

$$\phi_t = \frac{C*}{\beta} \frac{\Psi}{B_t^G}. \tag{2.1b}$$

We can then solve for the equilibrium values of ϕ_t and B_t^G as a function of parameters

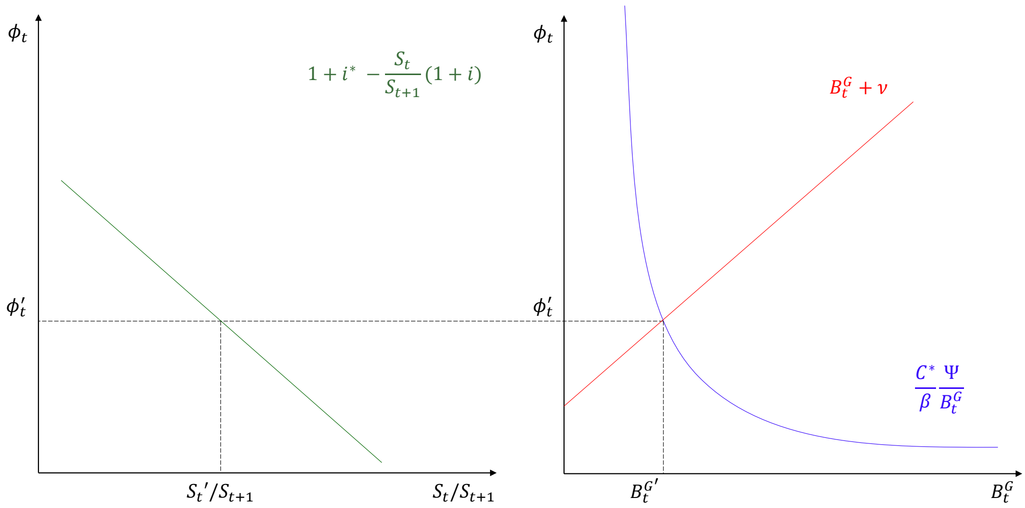
$$\phi_t = \frac{1}{2}\left(\nu + \sqrt{\nu^2 + 4\frac{C^*\Psi}{\beta}}\right), \qquad B_t^G = \frac{1}{2}\left(\sqrt{\nu^2 + 4\frac{C^*\Psi}{\beta}} - \nu\right).$$

The two latter equations tell us that the convenience yield ϕ_t and the optimal level of US debt B_t^G depend in equilibrium on the two parameters of interest, namely the marginal cost of debt issuance ν, and the preference for liquidity Ψ.
Figure 2.1 illustrates the equilibrium. The right panel shows the level of the convenience yield that clears the market for US government bonds. The red line is the supply of US debt as per Equation 2.1a, and the blue curve is the demand thereof by RoW households characterized by Equation 2.1b. The left panel portrays, for given constant levels of US and RoW nominal interest rates, the simultaneous depreciation required for the modified UIP condition to hold (green line) at the market-clearing level of the convenience yield.
Next, we show both graphically and analytically how this equilibrium is affected by changes in ν and Ψ.

Changes in the Marginal Cost of Debt

Figure 2.2 depicts the effects of a drop in ν on the equilibrium values of US government debt, the convenience yield and exchange rate. It corresponds to a

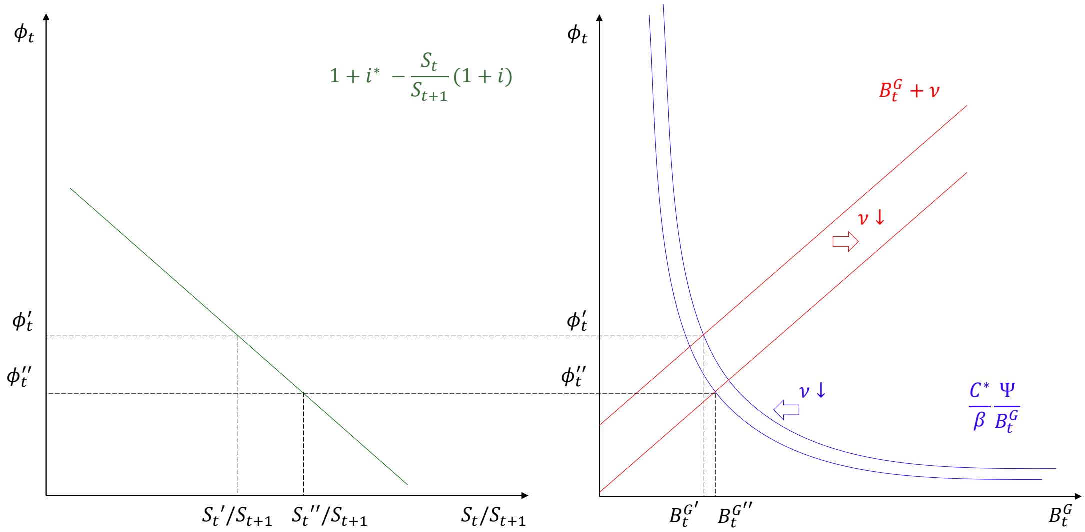
Figure 2.1. Equilibrium Debt, Convenience Yield and Exchange Rates



reduction in the fixed component of marginal cost of debt issuance, engendering an outward shift of the debt supply curve. As a result, the US government chooses to increase the supply of Treasuries *ceteris paribus*. Since the marginal liquidity value that foreign household derive from US Treasuries is decreasing in the amount held, they will require a higher monetary return to absorb the now higher supply. We define this as "debt supply effect".

In addition, a decrease in ν leads to an inward shift of the Treasury demand curve through an increase in the marginal utility of RoW consumption, which makes holding US debt less attractive at any level of ϕ_t. The decrease in C'^* originates through the goods market clearing from a contemporaneous increase in domestic consumption. In turn, the latter is due to subsitution from saving to consumption on the part of US households. As the convenience yield decreases following the first-order expansion in B_t^G, the economy-wide "carry" return from issuing Treasuries and investing in foreign bonds becomes less attractive, and so the US household invests less in foreign bonds and consumes more. We define this mechanism as the "marginal utility effect".

Figure 2.2. Marginal Cost of Debt Change



The responses of B_t^G and ϕ_t to a change in ν can be expressed formally as

$$\frac{\partial B_t^G}{\partial \nu} = \frac{1}{4}\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(2\nu + 4\frac{\Psi}{\beta}\frac{\partial C^*}{\partial \nu}\right) - \frac{1}{2},$$
$$\frac{\partial \phi_t}{\partial \nu} = \frac{1}{4}\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(2\nu + 4\frac{\Psi}{\beta}\frac{\partial C^*}{\partial \nu}\right) + \frac{1}{2}.$$

Note that the steady-state level of RoW consumtpion C^* depends on both ν and Ψ, as shown in Appendix 2.B. We can see immediately that $\frac{\partial \phi_t}{\partial \nu} > 0$ for $\frac{\partial C^*}{\partial \nu} > 0$, $\nu > 0$ $\beta > 0$ and $\Psi > 0$. Therefore, a decrease of ν results in a lower ϕ_t. We can confirm this graphically, as the debt supply curve shifts out after a drop in ν, while the debt demand curve shifts inward.

On the contrary, the sign of $\frac{\partial B_t^G}{\partial \nu}$ depends on the relative strengths of the debt supply and marginal utility effects. The latter is mediated by the $\frac{\partial C^*}{\partial \nu}$ term in $\frac{\partial B_t^G}{\partial \nu}$. The condition for $\frac{\partial B_t^G}{\partial \nu} < 0$ requires that the marginal utility effect is sufficiently weak, represented by an upper bound on $|\frac{\partial C^*}{\partial \nu}|$. If that were not the case, the US government would react to the weaker demand for Treasuries by choosing a lower B_t^G than in the original equilibrium. The derivatives $\frac{\partial \phi_t}{\partial \nu}$, $\frac{\partial B_t^G}{\partial \nu}$ and their signs are analysed formally in Appendix 2.D.

As a result of a drop in ν, the economy then moves to a new equilibrium characterized by a lower ϕ_t and, provided that the "debt supply effect" dominates, a higher B_t^G. Interest rates being fixed, the increase in returns from Treasuries will be achieved by an immediate depreciation and a later appreciation of the US dollar, as shown in the left-hand side plot. This theoretical mechanism translates into one simple testable implication which we state next.

Proposition 5. *An outward shift in US debt supply reduces the convenience yield through an immediate depreciation and a later appreciation of the US dollar.*

These dynamics are consistent with the correlational evidence in the literature, and with the results of our empirical analysis outlined in Sections 2.3.3 and 2.3.4. We contribute by showing that a positive shock to the supply of US Treasuries leads to a reduction of the US convenience yield and an immediate depreciation followed by an appreciation of the US dollar, relative to a panel of G10 currencies, with larger magnitudes than previously estimated.

Changes in Liquidity Preference

An increase in the liquidity preference of RoW households, i.e. an outward demand curve shift, also results in a higher equilibrium amount of US Treasuries, but it is instead associated with a dollar *appreciation* and a *higher* convenience yield.

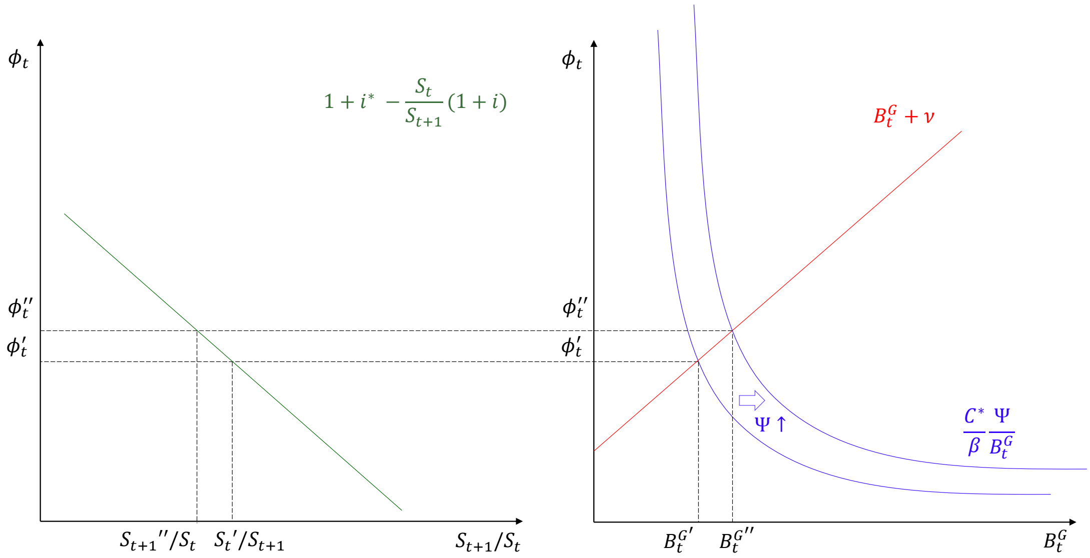
The responses of B_t^G and ϕ_t to a change in Ψ can be expressed formally as

$$\frac{\partial B_t^G}{\partial \Psi} = \frac{\partial \phi_t}{\partial \Psi} = \frac{1}{4}\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(\frac{4\Psi}{\beta}\frac{\partial C*}{\partial \Psi} + \frac{4C^*}{\beta}\right).$$

The two derivatives are the same, and they are both positive for $\Psi > 1$. This condition ensures that $|\frac{\partial C*}{\partial \Psi}|$ is low enough. In other words, the increase in the marginal utility of RoW consumption following an increase in Ψ should not be so strong as to reverse the first-order outward shift in US debt demand. In turn, this requires that the parameter Ψ that regulates the marginal liquidity benfit of Treasuries be high enough. We derive this condition formally in Appendix 2.D.
Figure 2.3 shows that a higher Ψ implies a higher marginal liquidity benefit for a given amount of US Treasuries held. Therefore, RoW households will accept a lower monetary payoff, which causes ϕ_t to increase. The left-hand plot shows that the higher convenience yield is achieved through a contemporaneous dollar appreciation.

The higher convenience yield will then incentivize the US government to issue more debt, due to the higher marginal benefit. If the Treasury supply curve is

Figure 2.3. Liquidity Preference Shift



upward-sloping in the convenience yield, changes in the liquidity preference for US Treasuries introduce a positive correlation between convenience yields, exchange rates and the equilibrium amount of debt. This theoretical mechanism can be summarized in a simple testable implication which we state next.

Proposition 6. *Changes in liquidity preference generate positive comovements between US debt supply, US dollar appreciation, and the convenience yield, introducing a bias towards zero in the coefficient of OLS regressions of the convenience yield on Treasury amounts.*

In other words, any empirical analysis that cannot distinguish whether the observed variability in outstanding Treasury amounts, convenience yields and exchange rates is due to changes debt supply or liquidity preference will incur in issues of endogeneity, which we address with an instrumental variable *à la* Phillot (2024) by isolating Treasury supply shocks.

Consistent with this type of endogeneity, we report in Sections 2.3.3, 2.3.4 and 2.3.5 a large downward bias of OLS estimates of the impact of Treasury supply shocks on the US convenience yield and exchange rate, relative to a panel of G10 currencies.

2.3 Empirical Evidence

Next, in an attempt to corroborate Implications 5 and 6 derived in the previous section, we present empirical evidence concerning the link between Treasury supply, convenience yields, and exchange rates. Our empirical contribution is to characterize how well-identified shocks to Treasury supply impact Treasury premia via currency fluctuations.

Although Du et al. (2018) discuss the long-term link between convenience yields and government bond supply, they do not discuss exchange rate fluctuations. Engel and Wu (2018), on the other hand, relate convenience yields to monthly currency swings, but they employ debt-to-GDP ratios as instruments for Treasury liquidity services. Our theoretical results support the view that government debt does not constitute a viable instrument, for fluctuations in the observed outstanding amount of Treasuries may very well depend on the liquidity services they offer. By invoking a cleaner identification of Treasury supply shocks, this paper aims at improving upon the existing literature.

2.3.1 Identification

Problem

It is challenging to empirically measure how changes in the supply of US Treasuries affect global financial markets and macroeconomic outcomes, because linear regressions of exchange rates or relative convenience yields onto US government debt inexorably face endogeneity issues. To see it, consider a flight-to-liquidity episode that leads to an increase in the US relative convenience yield and an appreciation of the US dollar. In other words, there are temporary demand-driven forces that make US debt relatively cheap to finance. Suppose further that the US government is more likely to issue Treasuries during those times when its debt trades at a relatively high convenience yield.[7] Then, a linear regression of that yield onto debt fails to disentangle supply- from demand-driven factors, reverse causality emerges and estimates are biased downwards. In turn, the link between convenience yields and exchange rates through the UIP transmits this issue to exchange rates too.

[7] In our theoretical framework, government debt issuance relaxes households' budget constraint via less taxes, allowing them to purchase more foreign bonds. The government pays the US interest rate on its borrowing, while households earn the foreign interest rate, converted into dollars, on their RoW bond investment.

To cope with this, we implement an identification strategy of US Treasury supply shocks *à la* Phillot (2024) using Treasury auction data. By interpreting changes in US Treasury futures prices around announcements by the US Treasury as surprises about the supply of US debt securities, we are able to recover shocks between 1998 and 2020. Following is a thorough description of the identification strategy.

Strategy

The US government finances its debt by issuing Treasuries, whose yield is determined via public auctions. Concurrently, several futures contract on US Treasuries—securities with a settlement price that the buyer agrees to accept delivery of on the settlement date—are being traded on the CBOT since 1977.

According to Phillot (2024), the design of US Treasury auctions offers an ideal set up for shock identification because the details about maturities and volumes of the issued securities are announced several days in advance and come with a report published on the same day. Under the efficient market hypothesis, intraday price variations of US Treasury futures around these announcements reflect surprises.

More formally, let $P_t^{TS,k}$ be the price of a k-year US Treasury and let $F_t^{TS,k}$ be its associated futures price for $k = 2, 5, 10, 30$. Phillot (2024) supposes that, at announcement time t,

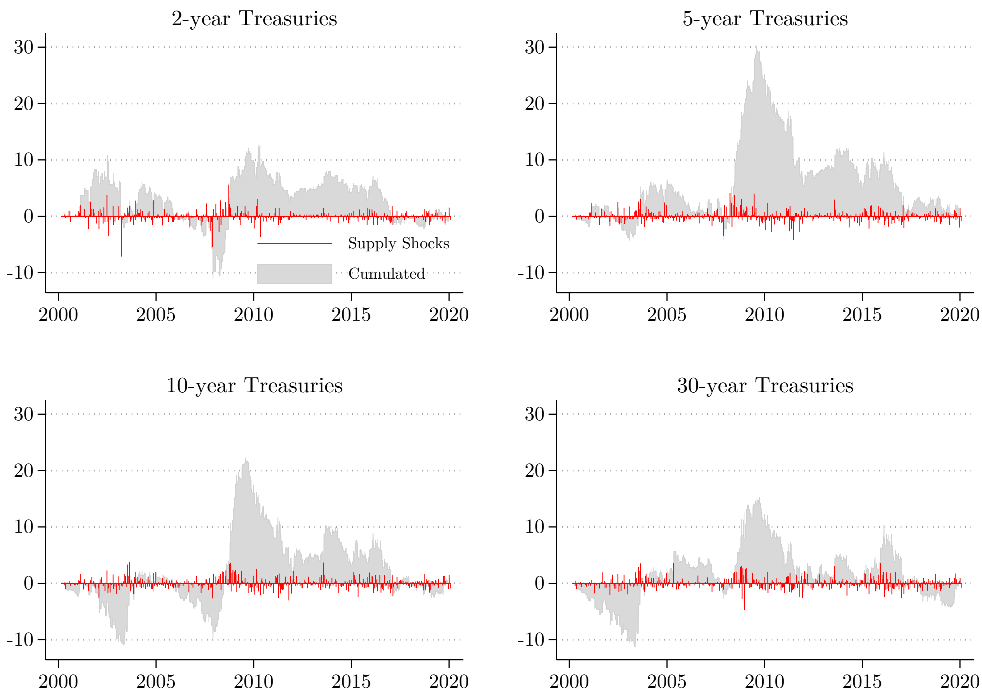
$$F_{t^+}^{TS,k} - F_{t^-}^{TS,k} = -\sigma^k \xi_t^k + u_t^k.$$

In this setting, futures price variations between t^- and t^+ have two drivers: US Treasury supply shocks ξ_t^k, scaled by Treasury demand price elasticity $-\sigma^k$, and a residual u_t. The latter consists of changes in Treasury futures prices that are orthogonal to Treasury supply shocks, such as the release of other macroeconomic news.

There is a trade off in picking the length of the time window (t^-, t^+). As argued by Nakamura and Steinsson (2018), a longer time window allows for capturing more detailed dynamics, yet it comes at the cost of potential confounding factors and noise contamination. In other words, shorter windows restrict the effects of u_t but generate instruments of little statistical power.

We chose a 15-minute window following the announcement so as to minimize the influence of cofounding factors while preserving a satisfying level of relevance for the instrumental variable exercise detailed below. Thus, our four Treasury supply shocks series $\{-\hat{\xi}_t^k\}_{k=2,5,10,30}$ are simply the 15-minute Treasury futures (inverse)

Figure 2.4. US Treasury Supply Shocks, 1998–2020



Source: Own calculations based on Phillot (2024).

returns following the announcements.[8]

2.3.2 Data

We now turn to describing the data used in our empirical approach.

US Treasury Supply Shocks

Announcements about US Treasury auctions are summarized in reports available on `TreasuryDirect.com`. Intraday data on US Treasury futures prices are provided by CQG. As mentioned above, we consider a 15-minute window following report official releases to compute the shocks.

[8]Phillot (2024), on the other hand, employs a 30-minute window. Unfortunately, using a 30-minute window results in weak instruments across most specifications, unlike the use of a 15-minute window. In other words, prolonging the window by 15 minutes introduces too much noise. Still, the conclusions outlined in the paper are less robust yet consistent under the 30-minute window. See Appendix 2.F for a discussion.

Figure 2.4 displays the series of shocks stemming from the identification strategy outlined above. The solid red spikes show the Treasury supply shocks $\hat{\xi}_t^k$ and the shaded areas show their running sums. For comparability, the shock series have been z-normalized. As a result, they have a zero mean and sum to zero.

Our measure of changes in Treasury supply at the daily frequency, which we instrument using the series of shocks depicted in Figure 2.4, is the net cash operation that reads on the US Treasury auction announcement reports. It corresponds to the sum of the dollar amounts of 2-, 5-, 10- and 30-year soon-to-be-auctioned Treasury securities, minus the dollar amounts of soon-to-mature 2-, 5-, 10- and 30-year Treasury securities.

Relative Convenience Yields

As outlined in Section 2.2, the relative convenience yield is a premium investors are willing to forego on their holdings of one country's Treasuries for the liquidity services they provide, relative other countries' Treasuries. In our theoretical framework, the relative convenience yield is a wedge in the UIP condition for government bonds. In the absence of uncertainty, UIP deviations are equivalent to CIP deviations.

Du et al. (2018) (based upon Du and Schreger, 2016) propose a measure of relative convenience yields based on CIP deviations, which they define as the yield difference between one country's government bond and US Treasuries, once cash-flows are hedged into that country's currency. In particular, letting $\iota_{t,j}^k - \iota_{t,US}^k$ be the time-t k-year own-currency government bond yield differential between country j and the US, and $\rho_{t,j}^k$ be the logarithm of the time-t k-year market-implied forward premium to hedge currency j against the US dollar, they define CIP deviations $\Phi_{t,j}^k$ as

$$\Phi_{t,j}^k = \iota_{t,j}^k - \iota_{t,US}^k - \rho_{t,j}^k.$$

Moreover, they argue that CIP deviations between country j and the US are mainly driven by their relative convenience yield $\phi_{t,j}^k$, their relative default risk $\kappa_{t,US}^k$, and risk-free CIP deviations $\tau_{t,j}^k$ caused by financial frictions

$$\Phi_{t,j}^k \approx \phi_{t,j}^k - \kappa_{t,US}^k + \tau_{t,j}^k,$$

such that relative convenience yields are well approximated by CIP deviations on government bonds, once relative default risk and CIP deviations on risk-free rates are taken into account. In our empirical exercise, we use their measure of convenience yields, which is available at daily frequencies for all G10 currencies

(Australia, Canada, Denmark, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, the United Kingdom, and the United States) and for bond maturities ranging from 3 months to 10 years.[9]

Using $\Phi^k_{t,j}$ as a proxy for $\phi^k_{t,j}$ assumes frictionless foreign exchange swap markets and default-free government bonds.[10] As a result, it abstracts from CIP deviations in FX markets and relative credit spreads. The former can be proxied using CIP deviations on observed risk-free rate proxies and the latter using credit default swaps (CDS) on sovereign credit, both of which can be found on Refinitiv Datastream. Unfortunately, CDS are not available before the Global Financial Crisis. As a result, we only control for market frictions in what follows and discuss the robustness of our results to controlling for sovereign credit risk in Appendix 2.E.

Finally, unless stated otherwise, CIP deviations are averaged along their maturity dimension (k). The panel structure at the day-currency-level (t, j) on the other hand is exploited.

Financial & Macro Variables

The financial variables used in the daily regression and the weekly local projections along with convenience yields are exchange rates and a set of controls. The latter are central banks policy rate differentials, MSCI stock market indices, WTI crude oil futures and gold futures prices, as well as the VIX.

Data on central banks policy rates come from the IMF ("International Financial Statistics" dataset), those on the VIX from the website of the Federal Reserve Economic Data (FRED), and the rest from Refinitiv Datastream. This daily sample covers the period between 2001 and 2020.

The macroeconomic variables deployed in the replication exercise of the quarterly panel-data analysis from Du et al. (2018) are debt-to-GDP ratios net of central banks' holdings, central banks policy rates, the VIX and real GDPs. Their sources are respectively the IMF ("Sovereign Debt Investor Base for Advanced Economies" and "International Financial Statistics" datasets), Refinitiv Datastream, FRED, and the OECD ("Quarterly National Accounts"). This quarterly sample covers the period between 2004 and 2020.

[9]The data is available at `https://sites.google.com/view/jschreger/CIP`.

[10]According to Du et al. (2018), this assumption is sound for developed economies.

2.3.3 Daily Regression

The first piece of empirical evidence we produce in this paper is a characterization of the impact of unexpected changes in the observable outstanding amount of US Treasuries on the US dollar exchange rate and the US convenience yield *vis-à-vis* other G10 currencies on a daily basis.

The goal is threefold. First, inspecting the first-stage statistics of the 2SLS procedure informs us on our instrument's performance. Second, once our instrument is deemed valid, the daily regressions provide a clear picture on the extent to which our instrument solves endogeneity issues associated with OLS. Third, the second-stage results do not only uncover the immediate effects of Treasury supply shocks, but also guide the subsequent weekly estimations in terms of the variables worth controlling for.

Methodology

We estimate two separate baseline pooled regressions on daily changes in Treasury supply ΔB_t^G of daily changes in the US convenience yield relative to country j, $\Delta\phi_{t,j}$, and daily log-changes in the US dollar exchange rate *vis-à-vis* currency j, $\Delta \log(S_{t,j})$

$$\Delta\phi_{t,j} = \beta_0 + \beta_1 \Delta B_t^G + \beta_2 \Delta \log(S_{t,j}) + v_{t,j}, \tag{2.2}$$

$$\Delta \log(S_{t,j}) = \gamma_0 + \gamma_1 \Delta B_t^G + \gamma_2 \Delta\phi_{t,j} + w_{t,j}. \tag{2.3}$$

The coefficients of interest in Equation 2.2 and 2.3 are β_1 and γ_1, for they respectively measure the contemporaneous effect in basis points of an increase in Treasury supply on relative convenience yields and exchange rates.[11]
As argued before, the OLS estimation of Equation 2.2 and 2.3 suffer from endogeneity. Phillot (2024) argues that the four series of Treasury supply shocks $\{-\eta_t^k\}_{k=2,5,10,30}$ are valid instruments for ΔB_t^G. The results shown in this paper therefore come from the 2SLS estimation of these two Equations.

Albeit parsimonious, these two models will arguably fail to account for a substantial part of the variability in convenience yields and exchange rates movements. In particular, one might be worried about the importance of cross-currency heterogeneity, macroeconomic low-frequency factors and other relevant financial market

[11] Although our results are robust to excluding them, including $S_{t,j}$ and $\phi_{t,j}$ in the two regressions is important if one thinks that Treasury supply, convenience yields and exchange rates admit an infinite vector MA representation whereby structural shocks to one variable affects the others contemporaneously.

outcomes.

As a result, in what follows, we supplement Equation 2.2 and 2.3 with country fixed effects, year fixed effects as well as set of controls. These are (daily changes of) US and country j stock market price indices, policy rate differentials between the US and country j, the VIX, gold and oil prices, endogenous changes in expectations about future monetary policy (measured as 15-minute changes in Fed funds futures prices around Treasury announcements), risk-free CIP deviations.[12] We also include dummies that take on the value one on days when an auction is either open for bidding, or held.

The inclusion of central bank policy rate differentials is meant to reflect the relative stance of monetary policy and capture exchange rate fluctuations pertaining to excess currency returns, as predicted by the UIP which our theoretical framework features. On the other hand, stock market indices and the VIX, beyond being indicators of investor sentiment and uncertainty, should capture the previously documented short-term interdependency of stock prices and exchange rates (Nieh and Lee, 2001). Finally, adding commodity prices should not only proxy inflation expectations but also explain some degree of exchange rate variations for so-called commodity currencies, i.e., the Australian, the Canadian, and the New-Zealand dollar (Chen and Rogoff, 2003).

It is noteworthy that these two regressions are equivalent to a zero-horizon instrumental variable local projection applied to panel data. Following Montiel Olea and Plagborg-Møller (2021), we augment Equation 2.2 and 2.3 with the two lags of all the financial variables (including the dependent variables) and compute standard errors that are robust to the presence of arbitrary heteroskedasticity.[13]

Results

Table 2.1 displays the results stemming from the estimation of Equation 2.2 and 2.3. The elements of columns 1 to 3 pertain to convenience yields, and those of columns 4 to 6 pertain to exchange rates. Columns 1 and 4 contain the OLS estimates, while columns 2 and 5 contain the IV estimates of the baseline model described by Equation 2.2 and 2.3. Columns 3 and 6 show the IV estimates when

[12]Risk-free CIP deviations control for frictions on the swap markets, one of two potential drivers of relative convenience yields according to Du et al. (2018). The other one, relative credit risk, is accounted for by Table 2.3 in Appendix 2.E.

[13]Table 2.4 from Appendix 2.E reports the estimates that are robust to both arbitrary heteroskedasticity and arbitrary autocorrelation. The statistical significance of our results is unchanged.

Table 2.1. On-Impact Effects of Treasury Supply Shocks

	Convenience Yield			Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	IV	OLS	IV	IV
ΔB_t^G (abs. median)	-0.0404	-0.635*	-0.668**	2.232***	30.54***	18.85***
	(0.0392)	(0.371)	(0.317)	(0.792)	(8.455)	(6.351)
$\Delta \log(S_t)$ (bp.)	-0.0319***	-0.0309***	-0.0248***			
	(0.00612)	(0.00615)	(0.00605)			
$\Delta\phi_t$ (bp.)				-4.840***	-4.694***	-5.005***
				(0.936)	(0.939)	(1.253)
Observations	49810	49810	48460	49810	49810	48460
Effective F-stat		13.3	17.3		13.3	17.3
$\hookrightarrow$ Critical Value		12.8	12.5		12.8	12.5
Hansen J-stat p-val.		0.31	0.24		0.71	0.18
Lags	No	No	Yes	No	No	Yes
Fixed Effects	No	No	Yes	No	No	Yes
Controls	No	No	Yes	No	No	Yes

Robust standard errors in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

the above-mentioned controls are included.
Looking at the bottom of Table 2.1, one finds a set of important statistics. First, our sample, by covering the period between February 2001 and January 2020 for 10 countries, amounts to nearly 50'000 business days. Second, our set of instruments appears relevant, as all the computed robust F-statistics (Olea and Pflueger, 2013) are above critical values.[14] In all cases, the Hansen J-statistic reflects a p-value exceeding conventional confidence levels, preventing us from rejecting the joint null hypothesis that the instruments are valid.

The first row of coefficients are to be interpreted as the contemporaneous effect of a change in Treasury supply on convenience yields and exchange rates. To recover interpretability, Treasury supply changes are scaled to be the size of an absolute median increase, i.e., \$13 billions.

Upon examination of the convenience yield, we note that OLS regression is unable to establish a statistically significant reduction in the US convenience yield following an increase in Treasury supply. On the contrary, the IV coefficients provided in column (2) indicate that, on impact, unexpected median-sized Treasury

[14]The critical values are computed under the null hypothesis that the Nagar (1959) bias is greater than 10% of the benchmark with a 95% confidence level.

supply increases result in a decline of approximately 0.65 basis points in the US convenience yield, relative to G10 currencies. This latter observation suggests that our model is susceptible to omitted variable bias in the absence of an instrument, consistent with Implication 6. The inclusion of additional controls in column (3) does not affect these findings.

Secondly, concerning exchange rates, our analysis shows that Treasury supply shocks lead to a significant 2.2 basis point depreciation of the US dollar on impact, as indicated by OLS. However, this effect is substantially underestimated by OLS, as our IV approach yields a much higher figure of 30.5 basis points. After adding the controls, we estimate that a median-sized surprise increase in Treasury supply results in an immediate depreciation of the US dollar by approximately 18.9 basis points.

Taken together, these results align with Implications 5, at least qualitatively speaking. In terms of their magnitude, on the other hand, these on-impact effects are fairly small. Indeed, the standard deviation of changes in convenience yields and exchange rates is 5.6 and 68.7 basis points respectively. Hence, their immediate response to debt supply shocks lies within the range of a tenth to a third of a standard deviation.

Notwithstanding, the weekly analysis below—computed over a time-period that speaks more adequately to our model—documents effects with a magnitude of greater economic significance as well as a non-negligible degree of persistence.

2.3.4 Weekly Local Projections

The second piece of empirical evidence we present in this paper is a general picture of the dynamics of exchange rates and relative convenience yields a few weeks following the shocks characterized above. We are particularly interested in knowing how the effects of Treasury supply shocks outlined above evolve within a quarter, for consistency with the timing of the theoretical model, and whether they are persistent.

Methodology

We estimate the cumulative impulse response functions (IRFs) of relative convenience yields and exchange rates to changes in Treasury supply via the following local projections

$$\phi_{t,j} - \phi_{t-h-1,j} = \beta_{0,h} + \beta_{1,h}\Delta B^G_{t-h} + \beta_{2,h}\Delta\log(S_{t-h,j}) + v_{t,j,h}, \tag{2.4}$$

$$\log(S_{t,j}) - \log(S_{t-h-1,j}) = \gamma_{0,h} + \gamma_1\Delta B^G_{t-h} + \gamma_{2,h}\Delta\phi_{t-h,j} + w_{t,j,h}, \tag{2.5}$$

for $h = 0, ..., 12$.[15] Because we are interested about the dynamic causal effects of US Treasury supply shocks in a time window that speaks to our theoretical framework, we lower our data frequency to the weekly level and compute these IRFs over a horizon of 12 weeks.[16] We take averages for all the financial outcomes (convenience yields, exchange rates, volatility, stock and commodity prices, policy rate differentials, FX swap market frictions), and sums for the auction-related variables (shifts in expectations about monetary policy, auction and bidding dummies).

As argued before, because the OLS estimation of Equation 2.4 and 2.5 suffers from endogeneity, we use the four series of Treasury supply shocks $\{-\eta^k_t\}_{k=2,5,10,30}$, summed over each week, as instruments for ΔB^G_t. The results shown below therefore come from their 2SLS estimation.

We add the same controls as in the daily regressions described earlier. Again, the confidence intervals that we report are based on a lag-augmented model with standard errors that are robust to heteroskedasticity (Montiel Olea and Plagborg-Møller, 2021).

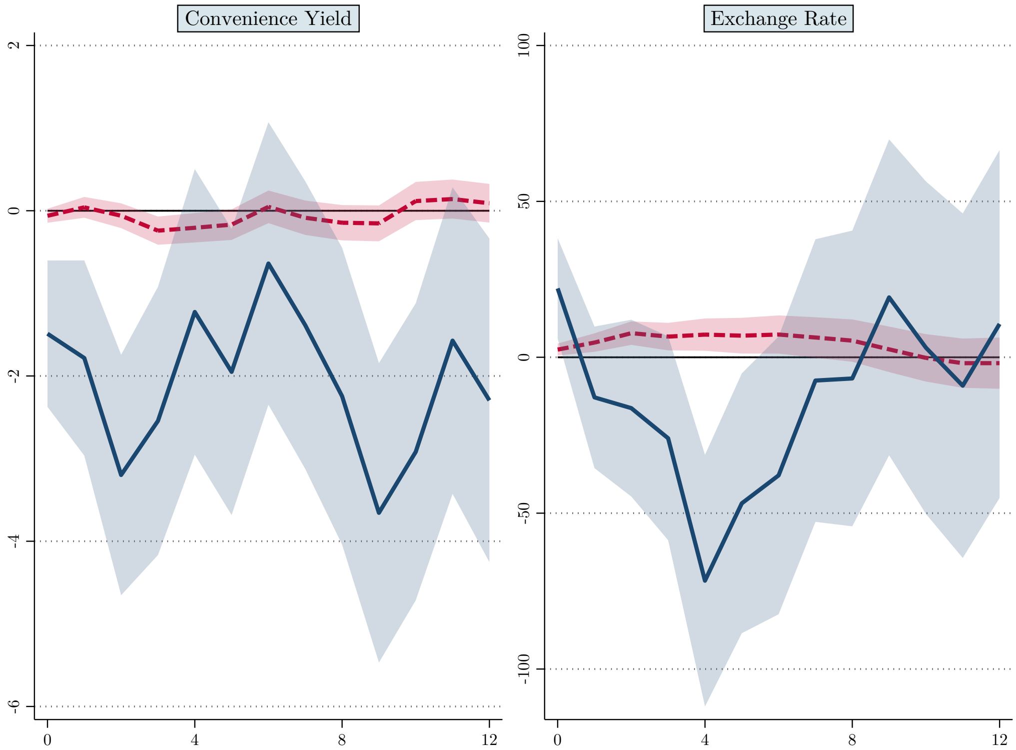
Results

Figure 2.5 displays the IRFs to Treasury supply shocks of convenience yields and exchange rates over a 12-week period (in basis points). The solid blue lines in each subplot displays the cumulative IRF to a 13-billion-dollar positive change in US Treasury supply of the above-mentioned variable stemming from the 2SLS estimation of Equation 2.4 and 2.5. The dashed red lines are the estimates of the same object via OLS. Blue and red shaded areas are the respective 90% confidence intervals. The horizontal axis represents weeks from impact. Variables can be interpreted in levels as the lines depict cumulative changes, as per Equation 2.4 and 2.5.
Concentrating on the IV outcomes for convenience yields, the initial observation is that the daily regression analysis results remain valid. Specifically, our instrumented approach indicates that, upon impact, the US convenience yield relative to

[15] Although presented with slight differences from the specification used in Phillot (2024), these local projections are equivalent as they generate cumulative IRFs to Treasury supply shocks.

[16] The results we obtain with weekly data are similar, both qualitatively and quantitatively, to the ones we obtain using daily data over 60 business days. The advantage of weekly IRFs over daily ones lies in their smoothness. See Figure 2.6 in Appendix 2.E.

Figure 2.5. IRFs to Treasury Supply Shocks of Convenience Yields and Exchange Rates



Notes: The solid blue lines in each subplot displays the IRF to a 13-billion-dollar positive change in US Treasury supply of the above-mentioned variable stemming from the 2SLS estimation of Equation 2.4 and 2.5. The dashed red lines are the estimates of the same object via OLS. Blue and red shaded areas are the respective 90% confidence intervals. The horizontal axis represents weeks from impact. Variables are in basis points.

G10 currencies experiences a substantial decrease of around 1 basis point. Furthermore, this decline is persistent throughout a quarter, and in some cases, reaches up to 3 basis points. Ordinary least squares (OLS) estimations, despite indicating a significant decrease four weeks after the shock, do not account for the magnitude and persistence of the effects revealed by our instrumental approach.

Concerning exchange rates, our IV local projections show that, similar to our daily regression, Treasury supply shocks lead to an immediate 25 basis points depreciation of the US dollar (USD). Consistently with the modified UIP condition in our theorretical model, the effects of these shocks reverse over time and result in a statistically significant appreciation of the USD, reaching more than 50 basis points four weeks after the impact, before vanishing four weeks later. Meanwhile, OLS estimates show a moderate but statistically significant depreciation of the USD over the same period.

Overall, our empirical findings not only highlight the presence of endogeneity in OLS estimates of IRFs, as per Implication 6, but they also align well with the dynamics prescribed by Implication 5.

2.3.5 Quarterly Panel-Data Regression

The third piece of empirical evidence we present in this paper is a replication of the panel-data quarterly analysis of the relationship between the outstanding amount of US Treasuries and the convenience yield in Du et al. (2018).
In an empirical study, Benhima and Phillot (2023) previously report that the OLS supply price elasticity of Swiss relative convenience yields is underestimated by a factor of three relative to an equivalent instrumental approach based on Swiss auction data. Here, we address whether the size of this bias is similar for the United States.

Methodology

As in Du et al. (2018), we regress relative convenience yields at quarterly frequencies onto a set of macroeconomic and financial variables, for the panel of G10 countries between 2004 and 2020.[17]

In particular, we model the US dollar 5-year maturity relative convenience yield as

$$\phi^{5Y}_{t,j} = \beta_0 + \beta_1 \log\left(\tfrac{\text{Debt}}{\text{GDP}}\right)_{US,t} + \beta_2 \log\left(\tfrac{\text{Debt}}{\text{GDP}}\right)_{j,t} + \beta_3' X_{j,t} + \varepsilon_{j,t}, \qquad (2.6)$$

where $\log\left(\tfrac{\text{Debt}}{\text{GDP}}\right)_{US,t}$ and $\log\left(\tfrac{\text{Debt}}{\text{GDP}}\right)_{j,t}$ are the log of the debt net of central banks's holdings as a ratio of GDP at time t for the US and country j respectively, $X_{j,t}$ is a set of controls. Du et al. (2018) consider for $X_{j,t}$ three variables other than currency fixed effects: The US policy rate, country j policy rate and the VIX.

Earlier, we argued that OLS estimates of β_1 from Equation 2.6 are likely biased downwards due to endogeneity.[18] To cope with this, we instrument the US debt-to-GDP ratio with the first principal component of US Treasury futures returns around announcements, cumulated over quarters and first-differenced. We resort

[17] Note that one departure from their specification lies in the data coverage, as their sample ranges between 2000 and 2016.

[18] Arguably, β_2 estimates face a similar bias. We unfortunately do not have auction data for the other countries in the sample and therefore interpret our estimates of β_2 with caution.

to principal component analysis because it performs better as an instrument for quarterly US debt-to-GDP ratio compared to using the set of four shock series as four separate instruments.

Unfortunately, our instrument's relevance erodes when we restrict ourselves to the above set of controls. Hence, we complement $X_{j,t}$ with additional controls, whose choice is guided by Bacchetta et al. (2022), namely the log of US real GDP, the log of country j real GDP and a quadratic time trend. In order to make our specification as similar as possible with the previously reported daily and weekly evidence, we additionally account for risk-free CIP variations.

Finally, because the error term $\varepsilon_{j,t}$ might very well be autocorrelated and could suffer from heteroskedasticity, we compute HAC robust standard errors (8 lags). Note that unlike Du et al. (2018), we do not resort to cluster-robust standard errors at the country level, since the consistency thereof calls for an infinite number of clusters and that we have only 10 of them (Angrist and Pischke, 2009). Arguably, the ample availability of observations along time dimension on the other hand does not invalidate the use of kernel-robust standard errors.

Results

Table 2.2 displays the estimates of Equation 2.6. Odd-numbered columns present OLS estimates, even-numbered ones present IV estimates. Columns 1 and 2 excludes all types of controls, columns 3 and 4 includes the set of controls used in Du et al. (2018) plus country fixed-effects, while columns 4 and 6 adds the controls suggested by Bacchetta et al. (2022).
Looking at the bottom of Table 2.2 reveals that our instrument is weak when we do not control for anything (column 2), or only for the VIX and policy rates (column 4). Indeed, the effective F-stats fail to exceed the critical values (95% confidence of a 20% worst-case bias). The resulting coefficients must therefore be interpreted with caution. The instrument's relevance is restored once we control for log-real GDPs and a quadratic trend (column 6), as the effective F-stat is well above its critical value.

In general, both OLS and IV associate government debt with relative convenience yield in a negative fashion. The reported OLS estimates predict that a one pp. increase in US debt-to-GDP weakens significantly the 5-year US relative convenience yield by between 0.49 and 1.50 basis points depending upon the inclusion of control variables. In particular, the effect is largest under the full set of controls, in itself reinforcing the view that statistical models linking Treasury supply to relative convenience yields are prone to committed variable biases.

Table 2.2. Du et al. (2018) Revisited

	5-Year Convenience Yield					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	OLS	IV	OLS	IV
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{US}}$ (pp.)	-0.49***	-3.32***	-0.81***	-1.72	-1.50***	-2.90***
	(0.10)	(1.07)	(0.17)	(1.65)	(0.36)	(1.10)
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{j}}$ (pp.)	0.04	0.86**	0.02	-0.05	0.11*	0.12*
	(0.07)	(0.41)	(0.08)	(0.13)	(0.06)	(0.07)
Observations	640	640	640	640	640	640
Effective F-stat		9.1		2.1		19.2
↪ Critical Value		15.1		15.1		15.1
Controls:						
↪ Du et al. (2018)	No	No	Yes	Yes	Yes	Yes
↪ Bacchetta et al. (2022)	No	No	No	No	Yes	Yes

HAC robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Our IV approach, on the other hand, report effects that have the same sign, yet with a higher magnitude. Indeed, a one percentage-point increase in US debt-to-GDP causes a decrease in the 5-year US relative convenience yield ranging between 1.72 and 3.32 basis points. Although the specification without control does well in delivering our central result that endogeneity issues prevent OLS to consistently estimate this coefficient, its IV estimate should be interpreted with caution, as it suffers from weak instruments. Adding the restricted set of controls (column 5) does not solve this issue. It is upon the inclusion of the full set of controls (column 6) that our instrument's relevance is restored. The corresponding decrease of roughly 3 basis points in the US relative convenience yield for each percentage-point increase in the US debt as a ratio of GDP is not only highly statistically significant, it is also twice as large a the OLS estimate. The latter observation is consistent with the OLS coefficient being muddled by the positive correlation between Treasury quantity and convenience yield introduced by demand shocks along an upwardly sloping supply curve, which is unaccounted for in the absence of a clean identification strategy for Treasury supply shocks.

2.4 Conclusions

In this paper, we develop a comprehensive theoretical framework to examine the interplay among Treasury supply, the convenience yield, and exchange rates. We conduct an empirical examination of the principal implications of this theoretical framework. This dual approach enables us to not only enhance our comprehension of the consequences of Treasury debt issuance for two variables of high relevance for market participants and policymakers alike, but also to validate these insights against empirical evidence.

We construct a simple two-country model to uncover the dynamics of supply and demand for US Treasuries. Our model describes how US Treasuries, as a result of a liquidity payoff derived by foreign investors, generate an endogenous and time-varying convenience yield relative to foreign bonds characterized by a wedge in the UIP condition. In turn, because US government debt issuance, for a given spending, relaxes households' budget constraint via less taxes, it allows them to purchase more foreign bonds and pocket the convenience yield—the marginal debt issuance benefit. As a result, the debt supply schedule is upward-sloping in the convenience yield because the government solves a Ramsey problem with a convex debt issuance cost. The framework reveals how the US convenience yield and exchange rate respond to changes in the marginal cost of debt issuance and shifts in liquidity preference.

An essential part of our theoretical discussion is the unveiling of two testable implications. Firstly, an increase in debt supply instigates a reduction in the convenience yield and an immediate depreciation of the US dollar followed by an appreciation. Secondly, as the US government faces incentives to issue debt in times when the US convenience yield is high, regression analyses of convenience yields and exchange rates onto US Treasury supply measures may suffer from endogeneity.

In a set of empirical exercises, we address these testable implications using an instrumental variable approach, thus tackling the endogeneity issue. In particular, we exploit the US Treasury auction design to elicit US Treasury supply shocks measuring intraday US Treasury futures price changes around announcements by the US Treasury. Because they reflect surprises to the supply of US Treasury securities, these futures prices changes qualify as valid instruments.

Crucially, the empirical results corroborate our model's testable implications. We demonstrate how unexpected increases in US Treasury supply lead to immediate depreciation of the US dollar followed by an appreciation, as well as a decline in the US convenience yield, relative to other G10 currencies between 2001 and 2020. Our findings also shed light on the downward bias exhibited by OLS estimates, demonstrating the presence of a positive correlation between US debt supply, the US relative convenience yield, and US dollar appreciation introduced by demand shocks, consistent with an upward-sloping US debt supply schedule.

The implications of these findings are twofold. They can inform policy discussions, given the significance to the global financial system of the US dollar and the liquidity and safety attributes of US Treasury securities, whereby the US enjoys a currency hegemony and an exorbitant privilege.[19] Additionally, they contribute to the literature, as mischaracterizations of government debt management strategy could arise from biased estimates of the impact of Treasury supply on the convenience yield, with important repercussions on the US debt sustainability and fiscal capacity.[20]

Our paper sets the stage for exciting further research. First, incorporating Treasury demand shocks into the empirical analysis, which could offer empirical verification of the conjecture that the US government solves a Ramsey-like optimal debt issuance problem. Second, introducing production under uncertainty, non-tradable goods, and active monetary policy regimes may refine results and help quantitatively test our predictions. Last but not least, developing an accurate dichotomy of Treasury supply shocks from changes in government spending versus changes in the composition of funding could elucidate the "exchange rate appreciation puzzle", by revealing conditions under which the convenience yield drop following a rise in debt supply can reconcile the observed exchange depreciation in response to an expansionary fiscal shock.

[19] See Gourinchas et al. (2019) for a recent discussion.
[20] See, e.g., Mian et al. (2021) and Jiang et al. (2022) for recent contributions.

Appendix

2.A US government Ramsey problem

The problem of the US household is

$$\max_{C_t, B_{F,t}} \sum_{s=0}^{\infty} \beta^s U(C_{t+s})$$
$$\text{s.t.} \quad C_t + B_{F,t} + T_t = \frac{1+i^*_{t-1}}{\Pi_t}\frac{S_t}{S_{t-1}} B_{F,t-1} + Y.$$

The resulting Euler equation for RoW bonds is

$$U'(C_t) = \beta \frac{1+i^*_t}{\Pi_{t+1}} \frac{S_{t+1}}{S_t} U'(C_{t+1}).$$

By substituting the T_t in the household's budget constraint using the government's budget constraint, the Ramsey problem of the US government is

$$\max_{C_t, B_{F,t}, B^G_t} \sum_{s=0}^{\infty} \beta^s U(C_{t+s})$$
$$\text{s.t.} \quad C_t + B_{F,t} + \bar{G} + \frac{B^G_{t-1}}{\Pi_t}(1+i_{t-1}) + \chi(B^G_{t-1}) - B^G_t = \frac{1+i^*_{t-1}}{\Pi_t}\frac{S_t}{S_{t-1}} B_{F,t} + Y$$
$$U'(C_t) = \beta \frac{1+i^*_t}{\Pi_{t+1}} \frac{S_{t+1}}{S_t} U'(C_{t+1})$$

Now consider a modified problem in which only the first constraint is present. The combined FOCs for C_t and $B_{F,t}$ yield the RoW bonds Euler equation, which appears as the second constraint in the original problem. Therefore, the Euler equation constraint is redundant, and we can write the Ramsey problem as in section 2.2.1 by removing the Euler equation and re-writing the budget constraint as a function of net foreign assets $A_t \equiv B_{F,t} - B^G_t$.

2.B Steady State

We solve the model assuming that domestic and foreign interest rates remain fixed at their zero-inflation steady-state values. In this appendix, we provide the solution for such steady state under the functional form assumptions made in section 2.2.2.

Given the zero-inflation assumption, at the steady state $\Pi = 1$, $\Pi^* = 1$. Therefore, given LOP and the assumption of $P_t^* = 1$, without loss of generality as P_t^* is indetermiante, it follows that $S = 1$.

The domestic and ROW households' Euler equations for foreign bonds at the steady state imply

$$i^* = \frac{1-\beta}{\beta}.$$

The ROW household's Euler equation for domestic bonds at the steady state implies

$$i = \frac{1-\beta}{\beta} - \frac{\Psi}{\beta B^G} C^*.$$

From the goods market clearing condition, we can express C^* as a function of C and exogenous variables

$$C^* = Y + Y^* - \bar{G} - C.$$

Substituting i and C^* in the ROW household's budget constraint evaluated at the steady state and exploiting the foreign government bond market clearing condition $B_{F,t} + B^*_{F,t} = \bar{B}_F$, we obtain

$$(1 + \frac{\Psi}{\beta})C + \frac{1-\beta}{\beta} B^G + \frac{1-\beta}{\beta} B_F^* - (1 + \frac{\Psi}{\beta})(\beta Y - \bar{G}) - \frac{\Psi}{\beta} Y^* = 0. \tag{2.7}$$

By substituting T from the domestic government's budget constraint into the domestic household's budget constraint at the steady state, and exploiting the market clearing condition for the foreign bond, we can also express it as an equation in B^G, B_F^* and C

$$\frac{\beta + \Psi}{\beta} C + \frac{\beta + \Psi}{\beta} \bar{G} + \frac{1-\beta}{\beta}(B^G - \bar{B}_F + B_F^*) - \frac{\Psi}{\beta} Y^* - \frac{\beta + \Psi}{\beta} Y + \frac{B^{G2}}{2} + \nu B^G = 0 \tag{2.8}$$

Furthermore, the first-order condition of the domestic government at the steady state reads

$$\frac{\Psi}{\beta B^G}(Y + Y^* - \bar{G} - C) = B^G + \nu. \tag{2.9}$$

Therefore, we can solve for B^G, B^*_F and C, through equations 2.7, 2.8 and 2.9.

This process results in a quadratic equation in B^G:

$$\frac{B^{G2}}{2} + B_G\nu + \left(1 - \frac{1}{\beta}\right) B_F = 0$$

It has two real solutions for $\beta \in (0,1)$:

$$B^G_1 = \frac{\sqrt{\beta\left(2\bar{B}_F\left(1-\beta\right) + \beta\nu^2\right)} - \beta\nu}{\beta},$$
$$B^G_2 = \frac{\sqrt{\beta\left(2\bar{B}_F\left(1-\beta\right) + \beta\nu^2\right)} + \beta\nu}{\beta}.$$

Note that for $\beta \in (0,1)$, $B^G_2 > 0$. Consistently with our approach for the dynamic equilibrium, we want restrict our attention to a unique positive value of steady-state US government debt. Therefore, we impose $B^G_1 < 0$, which holds for $\bar{B}_F \in [\frac{\beta\nu^2}{2(\beta-1)}, 0)$, and choose $B^G = B_{G,2} > 0$. A positive steady-state supply of Treasuries then requires a negative $\bar{B}_F$, implying that the RoW government is a net creditor in the steady state.

We can then solve for C and B^*_F as a function of B^G.

$$B^*_F = \frac{\beta}{\beta-1}\left(Y^* - \frac{\beta(1+\Psi)}{\Psi}B^{G2} + \frac{\Psi(1-\beta) - \beta^2\nu(1+\Psi)}{\Psi\beta}B^G\right)$$

$$C = Y + Y^* - \bar{G} - \frac{\beta}{\Psi}B^{G2} - \frac{\beta\nu}{\Psi}B^G$$

By substituting B^G, B^*_F and C into the ROW Euler equation, goods market clearing and ROW bond market clearing conditions, respectively, we find i, C^* and B_F.

2.C Equilibrium Existence and Uniqueness

Combining the two equilibrium equations, we are left with a quadratic equation that expresses ϕ_t as a function of parameters

$$\phi_t^2 - \nu\phi_t - \frac{C^*\Psi}{\beta} = 0$$

This equation has two solutions:

$$\phi_{1,t} = \frac{\nu - \sqrt{\nu^2 + 4\frac{C^*\Psi}{\beta}}}{2}, \qquad \phi_{2,t} = \frac{\nu + \sqrt{\nu^2 + 4\frac{C^*\Psi}{\beta}}}{2}.$$

First, note that $\nu^2 + 4\frac{C^*\Psi}{\beta} > 0 \, \forall \Psi > 0, \beta > 0, C^* > 0$. Therefore, $\phi_{1,t}$ and $\phi_{2,t} \in \mathbb{R}$ everywhere in the region of interest of the parameter space.

The model then features two equilibria, characterized by $\phi_{1,t}$ and $\phi_{2,t}$. For $\nu > 0, \Psi > 0, \beta > 0, C^* > 0$, we have $\phi_{2,t} > 0$, so an equilibrium with a positive convenience yield exists. Consider the sign of $\phi_{1,t}$. The condition for $\phi_{1,t} > 0$ is

$$\nu - \sqrt{\nu^2 + 4\frac{C^*\Psi}{\beta}} > 0 \iff \frac{C^*\Psi}{\beta} < 0.$$

This condition never holds for $\Psi > 0, \beta > 0, C^* > 0$. Therefore, $\phi_{1,t} < 0$ in the region of interest of the parameter space. By 2.1b, $\phi_{1,t} < 0 \implies B^G_{1,t} < 0$, so the convenience yield can only be negative in equilibrium if the US government is a net creditor. Our proxy for the convenience yield, that is observed CIP deviations for US government bonds, can take positive or negative values depending on currencies, but we discard the negative convenience yield equilibrium in our model because of the counterfactual implication of negative equilibrium levels of US government debt.[21]

We are then left with the equilibrium characterized by $\phi_{2,t}$ since $\nu^2 + 4\frac{C^*\Psi}{\beta} > 0$, $\phi_{2,t} > 0$ and, by 2.1b, $B^G_{1,t} > 0$. Therefore, what is presented in the main body of the paper is the unique equilibrium characterized by positive values of the convenience yield and the US debt level.

[21]See Sushko et al. (2016) for a discussion on the determinants of CIP deviation signs.

2.D Derivations for Comparative Statics

In this appendix we derive the formal conditions for the signs of $\frac{\partial \phi_t}{\partial \nu}$, and $\frac{\partial \phi_t}{\partial \Psi}$.

2.D.1 Steady-State Derivatives

We start by calculating useful derivatives of steady-state variables and determining their sign.

First, consider the derivative of US government debt supply $B_{G,t}$ with respect to ν at the steady state

$$\frac{\partial B^G}{\partial \nu} = 1 + \frac{\beta \nu}{\sqrt{\beta\left(\beta\nu^2 + 2\bar{B}_F(1-\beta)\right)}}. \tag{2.10}$$

The derivative of $B_{G,t}$ with respect to ν at the steady state.

Note that $\frac{\partial B^G}{\partial \nu} > 0$ for $\nu > 0$ and $\beta \in (0, 1)$.
Let us then turn to domestic consumption. Note that by goods market clearing, $C^* = Y + Y^* - \bar{G} - C$, so $\frac{\partial C^*}{\partial \nu} = -\frac{\partial C}{\partial \nu}$ and $\frac{\partial C^*}{\partial \Psi} = -\frac{\partial C}{\partial \Psi}$.
We have

$$\frac{\partial C}{\partial \Psi} = \frac{\beta B^G (B^G + \nu)}{\Psi^2},$$

so $\frac{\partial C}{\partial \Psi} > 0$ and $\frac{\partial C^*}{\partial \Psi} < 0$ for $\beta \in (0, 1)$, $\nu > 0$ and $\Psi > 0$.
We also have

$$\frac{\partial C}{\partial \nu} = -\frac{\beta}{\Psi}\frac{\partial B^G}{\partial \nu}(2B^G + \nu),$$

so $\frac{\partial C}{\partial \nu} < 0$ and $\frac{\partial C^*}{\partial \nu} > 0$ for $\beta \in (0, 1)$, $\Psi > 0$ and $\frac{\partial B^G}{\partial \nu} > 0$.

2.D.2 Comparative Statics

We will now use the results derived above to establish signs for the derivatives presented in the comparative statics in Sections 2.2.2 and 2.2.2.

To ensure that a decrease in ν results in a higher equilibrium level of US debt, we set

$$
\begin{aligned}
&\frac{\partial B_t^G}{\partial \nu} < 0 \\
\iff &\frac{1}{4}\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(2\nu + 4\frac{\Psi}{\beta}\frac{\partial C^*}{\partial \nu}\right) - \frac{1}{2} < 0 \\
\iff &\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(2\nu + 4\frac{\Psi}{\beta}\frac{\partial C^*}{\partial \nu}\right) < 2 \\
\iff &\frac{\partial C^*}{\partial \nu} < \frac{\beta}{2\Psi}\left(\left(\nu^2 + \frac{4\Psi}{\beta}C^*\right) - \nu\right).
\end{aligned}
$$

For $\frac{\partial C^*}{\partial \nu} > 0$, this condition requires that RoW consumption does not react too strongly to the increase in ν, so that the marginal utility effect does not dominate the debt supply effect.

To ensure that the increase in the marginal liquidity preference of RoW households results in a higher convenience yield and equilibrium level of US debt, we set

$$
\begin{aligned}
&\frac{\partial \phi_t}{\partial \Psi} > 0 \iff \frac{\partial B_{G,t}}{\partial \Psi} > 0 \\
\Psi \iff &\frac{1}{4}\left(\nu^2 + 4\frac{C^*\Psi}{\beta}\right)^{-1/2}\left(\frac{4\Psi}{\beta}\frac{\partial C^*}{\partial \Psi} + \frac{4C^*}{\beta}\right) > 0 \\
\iff &\frac{4\Psi}{\beta}\frac{\partial C^*}{\partial \Psi} + \frac{4C^*}{\beta} > 0 \\
\iff &\frac{4\Psi}{\beta}\frac{-\beta B^G(B^G + \nu)}{\Psi^2} + \frac{4C^*}{\beta} > 0 \\
\iff &\frac{\beta B^G(B^G + \nu)}{\Psi^2} < \frac{\beta B^G(B^G + \nu)}{\Psi} \\
\iff &\Psi > 1,
\end{aligned}
$$

where we substitute in C^* as a function of B^G and we use $\frac{\partial C}{\partial \Psi} = -\frac{\partial C^*}{\partial \Psi}$. This condition requires that the is strong enough that the outward shift in the debt demand curve engendered by a higher Ψ is not undone by the concurrent increase in the marginal utility of consumption due to $\frac{\partial C^*}{\partial \Psi} < 0$.

Table 2.3. Relative Credit Risk

	Convenience Yield			Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	IV	OLS	IV	IV
ΔB_t^G (abs. median)	-0.0404	-0.635*	-0.703**	2.232***	30.54***	18.45***
	(0.0392)	(0.371)	(0.318)	(0.792)	(8.455)	(6.249)
$\Delta \log(S_t)$ (bp.)	-0.0319***	-0.0309***	-0.0241***			
	(0.00612)	(0.00615)	(0.00605)			
$\Delta\phi_t$ (bp.)				-4.840***	-4.694***	-4.831***
				(0.936)	(0.939)	(1.249)
Observations	49810	49810	48460	49810	49810	48460
Effective F-stat		13.3	17.6		13.3	17.6
$\hookrightarrow$ Critical Value		12.8	12.6		12.8	12.6
Hansen J-stat p-val.		0.31	0.24		0.71	0.20
Lags	No	No	Yes	No	No	Yes
Fixed Effects	No	No	Yes	No	No	Yes
Controls	No	No	Yes	No	No	Yes

Robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

2.E Robustness

2.E.1 Daily Analysis

Relative Credit Risk

One limitation of our convenience yield metric is that it presupposes frictionless foreign currency swap markets and default-free government bonds. As a consequence, it abstracts from risk-free CIP deviations and relative credit spreads. Thus, we use CIP deviations on observed risk-free rate proxies in the study to approximate the former. Du et al. (2018) advise employing sovereign CDS for the latter. But, owing to the shortage of CDS data, we do not include it in our benchmark findings.

Table 2.3 re-estimates the figures in Table 2.1, by controlling for prices of CDS of both the US and country j. In order to avoid dropping half of the observations compared to our benchmark regressions, we set CDS prices to zero for all countries in the period prior to the Global Financial Crisis. Although this might seem a strong assumption, it is guided by the documented fact that prior to the crisis, credit spread differentials were a negligible a driver of our measure of relative convenience yields (Du et al., 2018).

Table 2.4. Heteroskedasticity and Autocorrelation-Robust Standard Errors

	Convenience Yield			Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	IV	OLS	IV	IV
ΔB_t^G (abs. median)	-0.0404	-0.635*	-0.668**	2.232***	30.54***	18.85***
	(0.0392)	(0.366)	(0.315)	(0.784)	(8.345)	(6.397)
$\Delta \log(S_t)$ (bp.)	-0.0319***	-0.0309***	-0.0248***			
	(0.00673)	(0.00677)	(0.00655)			
$\Delta\phi_t$ (bp.)				-4.840***	-4.694***	-5.005***
				(1.029)	(1.039)	(1.366)
Observations	49810	49810	48460	49810	49810	48460
Effective F-stat		13.1	17.1		13.1	17.1
$\hookrightarrow$ Critical Value		12.6	12.3		12.6	12.3
Hansen J-stat p-val.		0.31	0.24		0.70	0.19
Lags	No	No	Yes	No	No	Yes
Fixed Effects	No	No	Yes	No	No	Yes
Controls	No	No	Yes	No	No	Yes

HAC robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

By construction, only columns (3) and (6) of Table 2.3 contain alternative estimates. If anything, controlling for relative credit risk amplifies the reaction of relative convenience yields to Treasury supply shocks. Indeed, the estimated coefficient decreases from 0.668 to 0.703 basis points, the latter coefficient remaining highly statistically significant.

The immediate US dollar depreciation obtained from the benchmark results on the other hand, though it is conserved, is somewhat smaller. An unexpected median-sized increase in US Treasury supply, controlling for CDS prices, leads to a statistically significant depreciation of 18.45 basis points (18.85 basis points in the benchmark).

Autocorrelation-Robust Standard Errors In the benchmark daily regressions whose estimates are displayed in Table 2.1, we compute standard errors that are robust to arbitrary heteroskedasticity, as suggested by Montiel Olea and Plagborg-Møller (2021). Nonetheless, computing heteroskedasticity and autocorrelation consistent (HAC) standard errors instead is a reasonable alternative.

Thus, Table 2.4 informs on the significance of the same coefficients under HAC standard errors. The resulting minor variations in the estimated standard errors leaves all the conclusions drawn in the paper unchanged.

CIP Deviations on 2-, 5- and 10-Year Treasuries Our measure of US Treasury supply is the net operation that the US Treasury plans on achieving with an upcoming auction. Since our instrument is based on Treasury futures prices, and that these futures only exist for the 2-, 5-, 10- and 30-year bonds, this measure of supply considers the auctions of securities with one of these four maturities exclusively. Yet, in benchmark regressions, we average relative convenience yields along the maturity dimension, although the latter ranges from 3 month to 10 years.

What does that mean for our estimates $\hat{\beta}_1$? In theory, our benchmark estimates (in absolute terms) ought to be taken as lower bounds. This is because they fail to account for the variability in CIP deviations at maturities outside our supply measure that are, in truth, associated with Treasury supply shocks.

One natural robustness check therefore consists in considering CIP deviation on government bonds at maturities that match our supply measure. In particular, we re-estimate the coefficients from Table 2.1 using the average relative convenience yield for 2-year, 5-year and 10-year bonds. Table 2.5 displays the results.

As expected, the magnitude of the effect of Treasury supply shocks on relative convenience yields is larger. Our instrumental variable approach without control now associates a sudden median-sized increase in Treasury supply with an immediate decrease of the convenience yield of about 1 basis point, with very high statistical significance. Adding lags and fixed effects, and controlling for other important factors leads to an estimate of about 0.9 basis point.

Note that the coefficients for exchange rates are mostly unaffected.

2.E.2 Weekly Analysis

Long-Term Daily Local Projections As a robustness check, we replicate Figure 2.5 but using the daily sample instead of weekly averages thereof. The results are shown on Figure 2.6.

Qualitatively speaking, all the conclusions drawn in the paper are intact. Quantitatively speaking, although the overall magnitude is preserved, daily local projections produce IRFs that are somewhat more volatile. As a consequence, the estimated impact of Treasury supply shocks on relative convenience yield temporarily returns to a level that is statistically indiscernible from zero between 20 and 40 business days.

Table 2.5. CIP Deviations on 2-, 5- and 10-Year Treasuries Only

	Convenience Yield			Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	IV	OLS	IV	IV
ΔB_t^G (abs. median)	-0.0495	-0.987***	-0.897***	2.191***	29.90***	17.81***
	(0.0330)	(0.370)	(0.290)	(0.790)	(8.404)	(6.370)
$\Delta \log(S_t)$ (bp.)	-0.0277***	-0.0262***	-0.0280***			
	(0.00542)	(0.00547)	(0.00516)			
$\Delta\phi_t$ (bp.)				-9.200***	-8.829***	-10.10***
				(1.831)	(1.818)	(1.871)
Observations	49714	49714	48390	49714	49714	48390
Effective F-stat		13.2	17.2		13.2	17.2
$\hookrightarrow$ Critical Value		12.8	12.5		12.8	12.5
Hansen J-stat p-val.		0.00	0.00		0.69	0.21
Lags	No	No	Yes	No	No	Yes
Fixed Effects	No	No	Yes	No	No	Yes
Controls	No	No	Yes	No	No	Yes

Robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Exchange rates on the other hand seem to display an almost identical response to Treasury supply changes, both qualitatively and quantitatively.

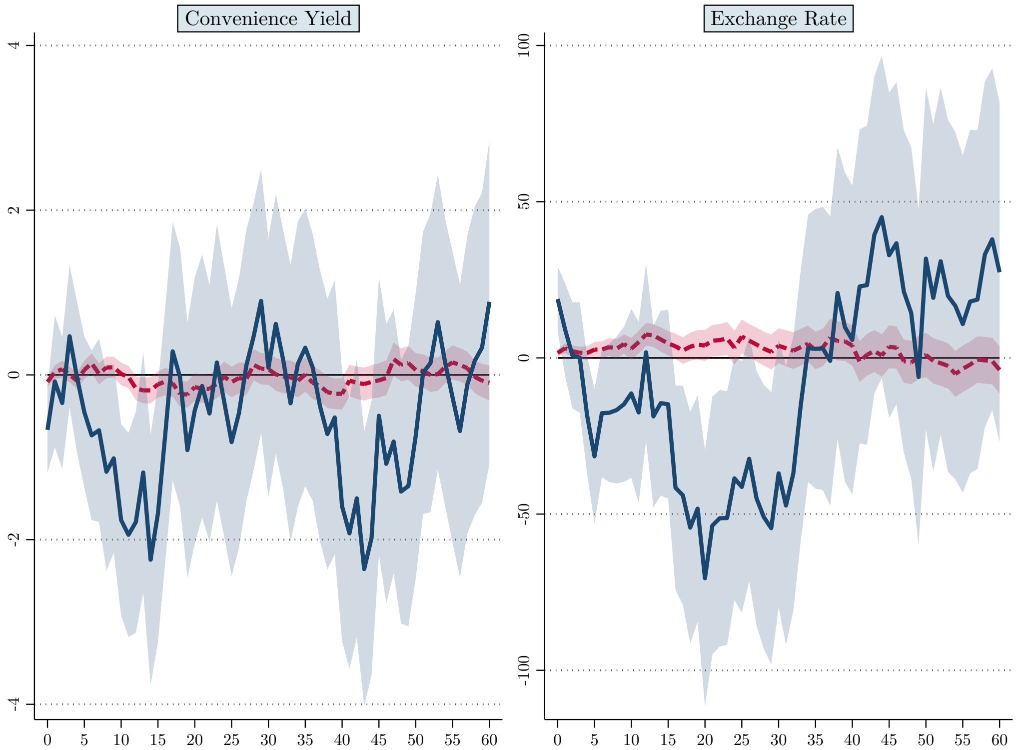
2.E.3 Quarterly Analysis

CIP Deviations on 2-, 5- and 10-Year Treasuries Recall that our instrument is the quarterly difference of the first principal component of within-quarter cumulated changes in Treasury futures prices occurring around 2-, 5-, 10- and 30-year US Treasury auctions. As a result, it is of interest to reassess the relationship between debt increases and the relative convenience yield, when the latter is measured using maturities that match the instrument.

Table 2.6 displays the re-estimation of the coefficients from Table 2.2 using the average relative convenience yield for 2-year, 5-year and 10-year bonds. As can be expected, the conclusions drawn from this refinement of our dependent variable are univocally identical to the ones exposed in the main body of this paper.

If anything, the effects reported in the specification with all the available controls are even larger than in the benchmark, i.e., the one with the strongest instrument. In particular, a 1 percentage-point increase in the US debt-to-GDP ratio is

Figure 2.6. IRFs to Treasury Supply Shocks of Convenience Yields and Exchange Rates



Notes: The solid blue lines in each subplot displays the IRF to a 13-billion-dollar positive change in US Treasury supply of the above-mentioned variable stemming from the 2SLS estimation of Equation 2.4 and 2.5. The dashed red lines are the estimates of the same object via OLS. Blue and red shaded areas are the respective 90% confidence intervals. The horizontal axis represents business days from impact. Variables are in basis points.

associated with a significant decrease in the relative convenience yield of 3 basis points.

Relative Credit Risk

As earlier, a natural robustness check consists in controlling for relative credit risk, for the latter is a potential driving force of the CIP deviations on Treasury securities, especially in the period after the Global Financial Crisis.

To this end, Table 2.7 re-estimates the figures in Table 2.2, by controlling for prices of CDS of both the US and country j. As before, we set CDS prices to zero for all countries in the period prior to the Global Financial Crisis.

Controlling for relative credit risk erodes the instrument's effective F-statistic in

Table 2.6. Du et al. (2018) Revisited, 2-, 5- and 10-year Maturities

	5-Year Convenience Yield					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	OLS	IV	OLS	IV
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{US}}$ (pp.)	-0.43***	-2.77***	-0.80***	-2.76	-1.59***	-3.00**
	(0.09)	(0.95)	(0.16)	(1.88)	(0.37)	(1.20)
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{j}}$ (pp.)	0.03	0.70**	-0.03	-0.17	0.04	0.06
	(0.07)	(0.36)	(0.07)	(0.15)	(0.08)	(0.08)
Observations	640	640	640	640	640	640
Effective F-stat		9.1		2.1		19.2
↪ Critical Value		15.1		15.1		15.1
Controls:						
↪ Du et al. (2018)	No	No	Yes	Yes	Yes	Yes
↪ Bacchetta et al. (2022)	No	No	No	No	Yes	Yes

HAC robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

the first stage. It comes as no surprise that once such an important driver of debt supply is controlled for, the power of a daily instrument aggregated at the quarter level is lowered. In any case, the shown coefficients have to be interpreted with caution as they suffer from weak instruments.

The second stage reveals results that are qualitatively similar to the specification used in the paper. Qualitatively speaking, the effects reported are somewhat smaller and of lower statistical significance. Overall, this robustness check confirms the findings highlighted in this paper.

Table 2.7. Du et al. (2018) Revisited, Relative Credit Risk

	5-Year Convenience Yield					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	OLS	IV	OLS	IV
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{US}}$ (pp.)	-0.49***	-3.32***	-0.81***	-1.72	-0.98***	-2.46*
	(0.10)	(1.07)	(0.17)	(1.65)	(0.35)	(1.41)
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{j}}$ (pp.)	0.04	0.86**	0.02	-0.05	0.10	0.11
	(0.07)	(0.41)	(0.08)	(0.13)	(0.08)	(0.08)
Observations	640	640	640	640	640	640
Effective F-stat		9.1		2.1		11.8
↪ Critical Value		15.1		15.1		15.1
Controls:						
↪ Du et al. (2018)	No	No	Yes	Yes	Yes	Yes
↪ Bacchetta et al. (2022)	No	No	No	No	Yes	Yes

HAC robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

2.F Alternative Instrument Window

In this appendix, we present an alternative analysis using the 30-minute shock window approach employed by Phillot (2024). While our main analysis in this paper utilized a tighter 15-minute shock window, it is natural to inquire about the differences that would arise if we were to adopt the same instrument as Phillot (2024).

We provide this appendix to offer a comparative perspective on the results obtained from both approaches. Specifically, we comment on the outcomes of the three baseline approaches employed: (i) the on-impact daily analysis, (ii) the weekly local projections over 12 weeks, and (iii) the quarterly replication of Du et al. (2018).

First, it is important to note that the instrument's F-statistics indicate weak instruments in the context of the daily and weekly analysis. Therefore, the results should be interpreted with caution and are presented here for completeness.

Nakamura and Steinsson (2018) argue that while a longer time window allows for capturing more detailed dynamics, it also introduces potential confounding factors and noise contamination. In our analysis, the choice of a 15-minute window strikes a balance between capturing relevant dynamics and minimizing these potential issues. Indeed, it appears that the inclusion of 15 additional minutes in the instrument computation introduces so much noise that it poses a challenge for instrument relevance.

Furthermore, it's worth considering that the analysis in Phillot (2024) involves distinct financial time series compared to our study, which could play a role in the divergent instrument performance we observe. While Phillot (2024) concentrates on domestic financial outcomes, our research centers on global macro-financial variables. Notably, our variables exhibit variation not only over time but also across countries, introducing a panel dimension not present in his paper. As a result, our adoption of the 15-minute window appears to offer a more targeted empirical approach for investigating convenience yields and exchange rates.

Despite these limitations, the analysis using the 30-minute window reveals a negative relationship between the US relative convenience yield and US Treasury supply both on impact (daily) and over the course of 12 weeks, whereby it takes over a month and a half for that effect to become significant. Exchange rates, on the other hand, fail to exhibit a statistically significant relationship with Treasury supply. Over the course of 12 weeks, an immediate currency change cannot be significantly estimated, but a future US dollar appreciation can. As such, it appears that our testable implication 1 can be confirmed by the data with a limited degree of robustness.

Second, for the quarterly analysis, the conclusions remain qualitatively similar, whereby a positive change in debt-to-GDP induces a significant decrease in the US relative convenience yield against a panel of G10 currencies. The relevance of the instrument, when it is cumulated quarterly, seems to perform extremely well (in terms of F-statistcs) while controlling for the same variables as Du et al. (2018). On the other hand, under our preferred set of instruments (Bacchetta et al., 2022), the instrument is weak and the result have to be interpreted with caution. That being said, the estimated coefficient under the latter specification of 8.89 is almost sixfold compared to its OLS equivalent, further demonstrating our testable implication 2.

Nonetheless, it is crucial to recognize the limitations of the presented empirical results. Due to the presence of weak instruments, which reduce the precision and robustness of the estimated coefficients, the findings should be interpreted with caution. Weak instruments not only introduce bias into the parameter estimates, but also produce larger standard errors, thereby lowering the reliability of the reported effects.

Next, we describe the detailed results for the daily, weekly and quarterly analyses under the alternative instrument.

2.F.1 Daily Analysis

We estimate the same-day effect of changes in US Treasury supply on the convenience yields and exchange rates of the US relative to G10 currencies. The OLS specification is identical to the one used in the main body of the paper, while the IV specification employs the US Treasury futures returns over a 30-minute window around Treasury announcements, as in Phillot (2024).

As can be seen from Table 2.8, the instrument fails in all specifications to achieve the corresponding critical value. Hence, we face weak instruments and cannot interpret the coefficients as causal effects with a high enough degree of certainty.

Regarding relative convenience yields (columns 1 and 3), we still find a negative relationship with Treasury supply throughout. The only coefficient that is statistically significant in this respect is the one from IV that accounts for the lags, fixed effects, and the set of controls (column 3). Albeit suffering from weak instruments, the estimated coefficient implies that an absolute median increase in Treasury sup-

Table 2.8. On-Impact Effects of Treasury Supply Shocks, 30-Minute-Window Instrument

	Convenience Yield			Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	IV	OLS	IV	IV
ΔB_t^G (abs. median)	-0.0404	-0.454	-0.999*	2.232***	-0.406	-5.745
	(0.0392)	(0.507)	(0.545)	(0.792)	(8.473)	(8.963)
$\Delta \log(S_t)$ (bp.)	-0.0319***	-0.0312***	-0.0244***			
	(0.00612)	(0.00619)	(0.00611)			
$\Delta\phi_t$ (bp.)				-4.840***	-4.854***	-5.376***
				(0.936)	(0.937)	(1.247)
Observations	49810	49810	48460	49810	49810	48460
Effective F-stat		11.1	9.4		11.1	9.4
$\hookrightarrow$ Critical Value		12.2	12.3		12.2	12.3
Hansen J-stat p-val.		0.12	0.45		0.55	0.15
Lags	No	No	Yes	No	No	Yes
Fixed Effects	No	No	Yes	No	No	Yes
Controls	No	No	Yes	No	No	Yes

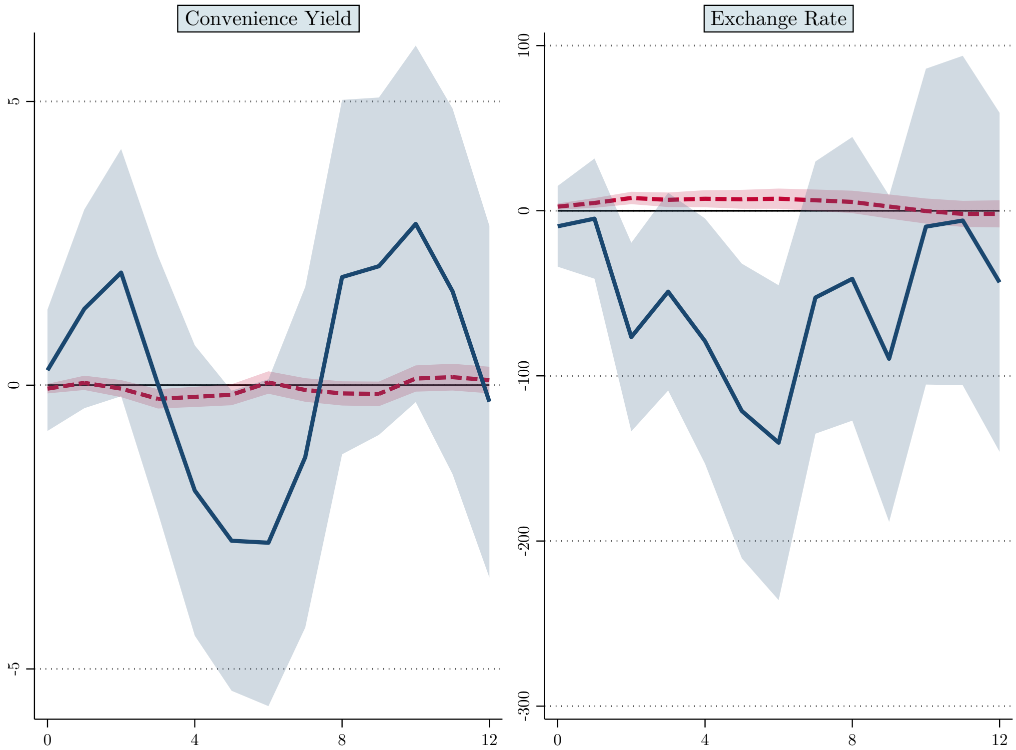
Robust standard errors in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

ply leads to a 1 basis-point decrease of the US convenience yield relative to the other currencies in the sample. This effect is significant at the 90% level. This result is consistent with testable implication 1, further supporting the theoretical framework of the paper.

As for exchange rates (columns 4-6), not only do we suffer from weak instruments, but we also find statistically insignificant relationships once Treasury supply is instrumented. Even the coefficients are negative, going against our testable implication 1.

Finally, regarding testable implication 2, our analysis suggests that it can only be supported to the extent that the significant coefficient from column 3 exceeds the statistically insignificant coefficient from column 1. However, the results on exchange rates do not provide additional support for this implication. Therefore, while the findings on relative convenience yields align with testable implication 2, the lack of significant relationships in exchange rates limits the overall support for this testable implication.

Figure 2.7. IRFs to Treasury Supply Shocks of Convenience Yields and Exchange Rates, 30-Minute-Window Instrument



Notes: The solid blue lines in each subplot displays the IRF to a 13-billion-dollar positive change in US Treasury supply of the above-mentioned variable stemming from the 2SLS estimation of Equation 2.4 and 2.5. The dashed red lines are the estimates of the same object via OLS. Blue and red shaded areas are the respective 90% confidence intervals. The horizontal axis represents weeks from impact. Variables are in basis points.

2.F.2 Weekly Analysis

We apply the same methodology as in the main body of the paper, estimating local projections of the dynamic causal effect of Treasury supply shocks on relative convenience yields and the US dollar relative to G10 currencies, but at a weekly frequency. The only difference is that we use the alternative 30-minute window for measuring the instrument, as in Phillot (2024).

First, similar to the daily analysis, the instrument does not achieve relevance in this exercise. The F-statistic of 7.29 falls below the critical value of 10.27, raising doubts about the validity of the estimates. Despite this limitation, Figure 2.7 presents the IRFs following a $13 billion increase in US Treasury supply, mirroring the approach in the main body of the paper.

As observed in Figure 2.7 , there is no significant immediate impact on the US dollar exchange rate and the US convenience yield following the shock. However, after approximately one and a half months, the US relative convenience yield experiences a significant decrease of about 2.5 basis points, which quickly diminishes. Treasury supply shocks seem to cause the US dollar to appreciate by over 100 basis points after one and a half months, with the effect dissipating within a quarter.

The weekly analysis, conducted using this alternative instrument, provides partial confirmation of testable implication 1, aligning with the daily findings outlined above. Treasury supply tends to lead to a future decrease in US convenience yields and a future appreciation of the US dollar, but does not immediately impact these variables. On the other hand, the decoupling of the effect magnitude with the use of this instrument appears to provide supportive evidence for testable implication 2.

However, it is important to note that the weakness of the instrument in this analysis raises concerns about the validity of the approach. As a result, we favor the 15-minute window used in the main body of the paper as a more robust approach for measuring these relationships.

2.F.3 Quarterly Analysis

In quarterly analysis, we replicate the exercise conducted by Du et al. (2018), which is presented in the main body of the paper. However, we introduce a 30-minute-window instrument for this replication. The results of this replication exercise can be found in Table 2.9.

Columns 1 and 4 in Table 2.9 correspond to the same specifications as those in Table 2.2, as they use identical parameters. On the other hand, columns 2, 3, 5, and 6 incorporate the alternative instrument. It is worth noting that the instrument, represented by the first principal component derived from the four series of daily instruments (cumulated over quarters), performs exceptionally well in both the specification without controls (column 2) and the one with the same controls as used by Du et al. (2018) in their paper (column 3). The F-statistics in these cases greatly exceed their respective critical values, indicating sufficient instrument relevance. However, the same cannot be said for the specification that includes the controls proposed by Bacchetta et al. (2022) (column 6), raising doubts about the validity of the estimates.

Table 2.9. Du et al. (2018) Revisited, 30-Minute-Window Instrument

	5-Year Convenience Yield					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	OLS	IV	OLS	IV
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{US}}$ (pp.)	-0.49***	-1.58***	-0.81***	-1.34***	-1.50***	-8.89***
	(0.10)	(0.30)	(0.17)	(0.32)	(0.36)	(3.12)
$\log(\frac{\text{Debt}}{\text{GDP}})_{\text{j}}$ (pp.)	0.04	0.36**	0.02	-0.02	0.11*	0.17*
	(0.07)	(0.14)	(0.08)	(0.08)	(0.06)	(0.09)
Observations	640	640	640	640	640	640
Effective F-stat		52.5		47.0		7.3
↪ Critical Value		15.1		15.1		15.1
Controls:						
↪ Du et al. (2018)	No	No	Yes	Yes	Yes	Yes
↪ Bacchetta et al. (2022)	No	No	No	No	Yes	Yes

HAC robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

The estimated impact of a one percentage-point increase in US debt-to-GDP on the US convenience yield varies between -1.6 and -8.9 basis points, depending on whether and which controls are included in the model.

Under this alternative instrument, the conclusions outlined in the paper remain consistent. Most notably, instrumenting debt-to-GDP ratios using Treasury futures prices changes around announcements by the US Treasury resolves an endogeneity issue, as suggested by testable implication 2. However, it is important to exercise caution when interpreting the figure of 8.8 basis points due to the potential problem of weak instruments.

Chapter 3

Mutual funds and safe government bonds: do returns matter?

This paper investigates the sensitivity of the demand for safe government debt to currency unhedged and hedged excess returns in a sample of US mutual funds. We find evidence of active rebalancing towards government bonds that offer relatively higher returns on an unhedged basis, in particular euro-denominated securities. The size of the effect is large, leading to a change in portfolio share by around one percentage point on average in response to a change by one percentage point in the currency-specific excess return. Interestingly, mutual funds rebalance their portfolio towards currencies, such as the Japanese yen, that display large deviations in the covered interest parity and offer higher returns than US Treasuries on an hedged basis. Finally, when global financial risk is on the rise, US mutual fund managers repatriate their investments towards US government debt securities, mainly at the expenses of euro-denominated ones. Our results imply that deviations in pricing conditions like uncovered and covered interest parity for sovereign bonds affect capital flows from the United States towards other major currency areas.

This chapter is co-authored with Maurizio Michael Habib. Both authors provided equal contributions.

3.1 Introduction

Since the global financial crisis in 2008, a scarcity of safe assets emerged, leading to a dramatic decline in the yields of government debt issued by major advanced economies with a reserve currency status and a growing interest in the characteristics of safe assets (Caballero et al., 2017). Safe assets command a premium for liquidity (Krishnamurthy and Vissing-Jorgensen, 2012), are information insensitive (Gorton, 2017) and have a negative market beta, appreciating when global risk aversion is on the rise, i.e they are like a "good friend": valuable and liquid when one needs them (Brunnermeier et al., 2022a). Most of the empirical research on this topic focuses on what is considered the world's safe asset, US Treasury debt, while there are only few studies that try to offer a global perspective to this debate (Du et al., 2018; Habib and Stracca, 2015; Habib et al., 2020). Understanding the drivers of demand for safe assets becomes even more important as geopolitical risk is on the rise following the war in Ukraine, sanctions on Russia by advanced economies, and tensions in the Middle East, which might cause long-term consequences for the international monetary system (Brunnermeier et al., 2022).

In this paper, we tackle this issue from a specific angle, studying the portfolio of *safe* government debt securities – debt issued by sovereigns with the highest credit rating, double A or higher according to Standard & Poor's – held by US mutual funds, which play a key role in intermediating savings from the world's largest economy to the rest of the world.[1] Specifically, we study whether asset managers of US mutual funds shift the allocation of these portfolios towards currencies that offer higher yields ("search for yield"). We also ask whether the demand for high-rated government bonds by mutual funds is affected by global financial conditions, as typical of safe haven assets.("search for safety"). By studying the sensitivity of the portfolio shares of safe government securities to returns differentials for a major class of investors such as mutual funds, we help to shed light on the broader nature of demand for safe assets.

To a very large extent, government debt of major advanced economies is issued in domestic currency and therefore the US dollar return of the portfolio of US mutual funds will be influenced by exchange rate movements of these currencies against the US dollar. Portfolio managers may decide to accept the currency risk or hedge

[1]Total assets managed by the fund industry in the United States rose from little over USD 5 trillion at the turn of the century to more than USD 16 trillion at the end of 2022 (see Financial Accounts of the United States, Table L.122, Federal Reserve Board), and are a growing source of financing in the sovereign debt markets of advanced economies (Fang et al., 2023).

it. As a novelty compared to the existing literature, we use a granular fund-level panel dataset to investigate the reaction of the portfolio shares of sovereign "safe" assets to both currency *unhedged* excess returns – i.e the total return differential between US dollar debt and that issued in another currency – and currency *hedged* excess returns, where asset managers use derivatives to neutralise the impact of fluctuations in the exchange rate of the US dollar against other currencies in the portfolio. Indeed, around 90% of US fixed income funds with an international focus use currency forwards to manage their foreign exchange exposure (Sialm and Zhu, 2022). To neutralise this currency risk, managers have to sell the foreign currency in the forward market against the US dollar.

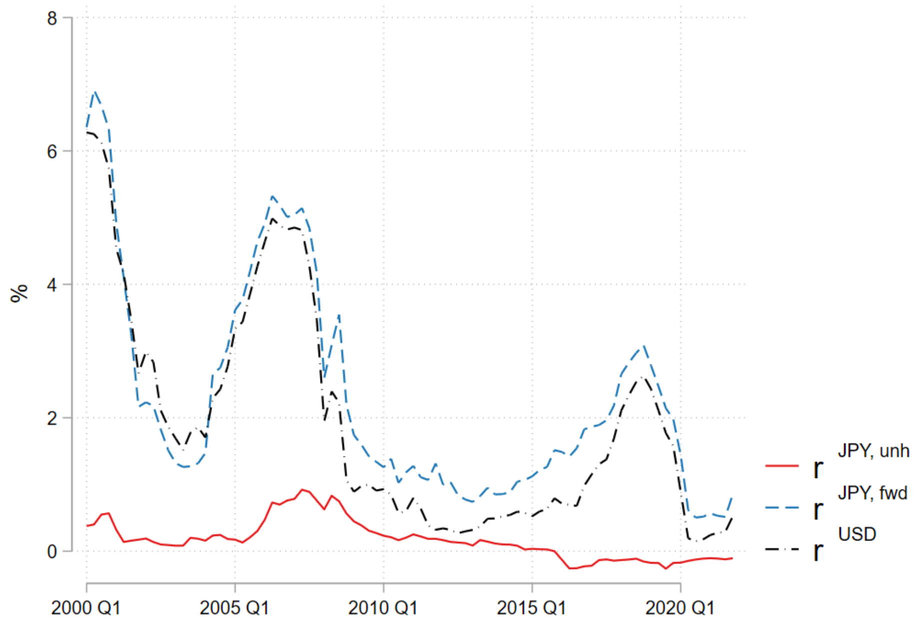
In theory, the Covered Interest Parity (CIP) should ensure that the cost of this currency hedging operation match the forward premium or discount, i.e the difference between the spot exchange rate and the forward one. This way, the dollar return equals the return from investing in another currency that is hedged into dollars.[2] In practice, there may be large deviations from CIP, which create a wedge between the dollar return and the foreign currency *hedged* return.[3] As a result, the portfolio manager of a US dollar-based fund that invests in government debt securities issued by sovereigns with similar credit risk is confronted with a choice between three potential returns: (i) the return on US dollar debt; (ii) the currency *unhedged* return from investing in debt issued by another country and denominated in a foreign currency, including an exchange rate fluctuation between the current spot rate and future spot rate; (iii) a *hedged* return in US dollar terms from investing in the foreign currency debt and covering the exchange rate risk with derivatives. Figure 3.1 shows concretely the return opportunities of a US investor that must decide between investing in US (dollar) debt or Japanese (yen) debt securities. On an unhedged basis, US Treasuries (black dashed line) yield higher returns than Japanese government bonds (red solid line) so that an investment in Japanese yen offers a lower return unless the investor expects that the yen will appreciate over the investment horizon. The comparison between the re-

[2]For instance, assuming similar credit and liquidity risk and the same maturity between two debt securities issued by the United States and another sovereign, if the interest rate on the foreign currency - say debt issued by Japan in Japanese yen - is lower than the interest rate on US dollar debt, then the US dollar should be priced at discount against the Japanese yen in the forward market with respect to the spot market.

[3]CIP deviations have been attributed to either the unique safety and liquidity of US Treasuries (Krishnamurthy and Vissing-Jorgensen (2012), Du et al. (2018)) or frictions in FX markets (Borio et al. (2016)), and their implications are usually understood in terms of affecting the cost of USD hedging for non-dollar investors. From the perspective of dollar-based US investors, CIP deviations can instead represent an opportunity for higher returns on foreign government bonds on a hedged basis, in line with the speculative motive of hedging described by Anderson and Danthine (1981).

turn from an investment in US Treasuries (black dashed line) and one in Japanese debt, currency hedged (blue dashed line) is particularly interesting. Under CIP, the black and blue lines should coincide. Indeed, the two lines tend to comove. However, we can observe large deviations in our sample: on average more than 50 basis points and, in some instances, even up to 100 basis points in favour of an investment into Japanese debt, currency hedged. Most importantly, since the global financial crisis in 2008, the sign of this deviation is consistently positive so that this *risk-free* excess return is *predictable*. To our knowledge, so far no one investigated whether US professional investors, such as fund managers, tried to exploit these opportunities to boost the overall return of their dollar portfolios, and how these strategies differ across currencies.

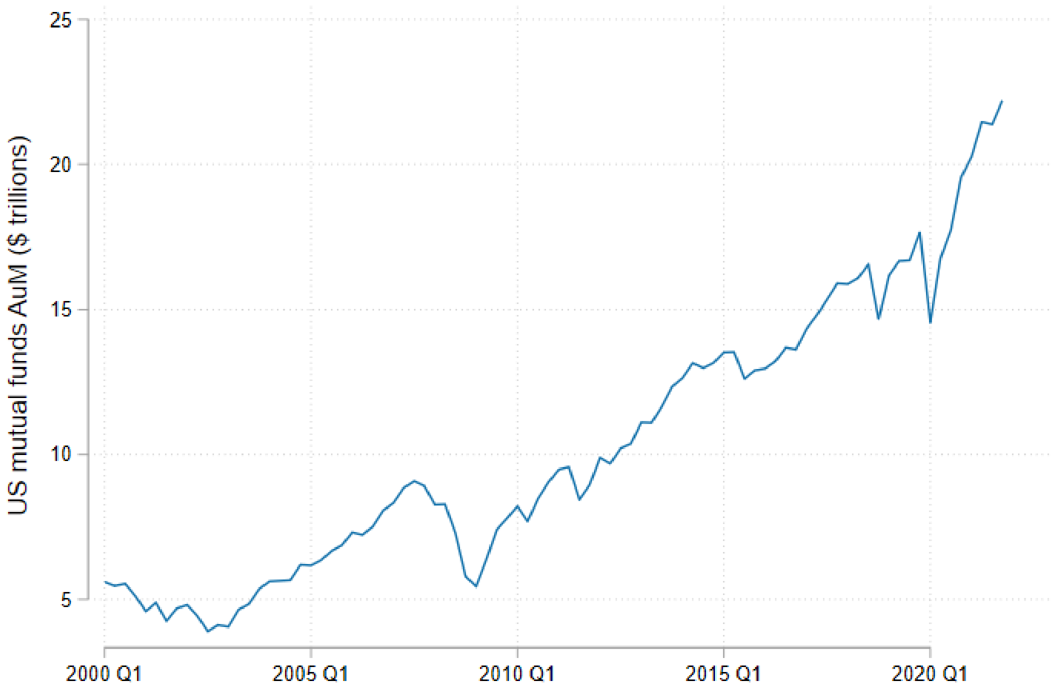
Figure 3.1. Hedged and unhedged returns from investing in Japanese debt against the return from US Treasuries



Average yields on US Treasuries in dollars, r_t^{US} black line . Average yield of Japanese government bonds in yen, $r_t^{j,unh}$ red line. Average yield of Japanese government bonds hedged into dollars, $r_t^{j,fwd}$ blue line. All yields are averaged over the 3 month, 1 year, 2 year and 5 year maturities. Hedged returns are calculated based on forward contract of the same maturity as the corresponding government bond. Source: Refinitiv Eikon.

Non-bank financial intermediation more in general also grew in importance relative to banks since the 2008 financial crisis, raising questions for financial stability as discussed for example in Aramonte et al. (2021). Mutual funds then play a systemic and increasingly large role in intermediating capital from savers in the

Figure 3.2. Mutual funds industry growth



Assets under management of all mutual funds resident in the United States. Source: Federal Reserve Finanial Accounts of the United States, Table F.122

world's largest economy. Therefore, it is crucial to understand how they manage their portfolio, and especially international investments due to the implications for capital flows.

To assess the sensitivity of the portfolio of US mutual funds to differentials in currency *unhedged* and *hedged* returns, this study uses panel regressions of the shares of currencies of major advanced economies with an elevated credit rating: the US dollar, the euro [4] the Japanese yen, the pound sterling, the Swiss franc, and the Australian dollar. We estimate a model for each currency, including fund and time fixed-effects to isolate idiosyncratic variation in fund-level, currency-specific excess returns. The objective of our analysis is the behaviour of fund managers and their decision to change the currency allocation of their portfolio in response to returns. Therefore, we net out valuation effects that are driven by changes in bond prices and exchange rates and focus on the active rebalancing of their currency portfolio,

[4]For the euro area, we include only government debt issued by Austria, Belgium, France, Germany and the Netherlands, which maintained a credit rating from S&P of AA or higher throughout our sample period.

which tends to dominate the overall variation in the currency shares.

Our main findings are the following. We find evidence of active rebalancing towards the government bonds that offer relatively higher returns on an *unhedged* basis, in particular euro-denominated securities. The size of the effect is large, leading to a change in the portfolio share of the US dollar, the euro, the pound sterling, the Swiss franc or the Australian dollar by around one percentage point in response to a change by one percentage point (one standard deviation approximately) in currency-specific excess returns. The ensuing impact on capital flows is also sizeable. For instance, an increase by one percentage point in the excess return of euro area economies issuing highly-rated debt triggers capital flows from the United States towards these euro area economies in the order of magnitude of $300 million. This accounts for around 2 percentage points of total quarterly foreign flows into highly-rated euro area government debt securities on a quarterly basis, according to the balance of payments. There is also evidence of active rebalancing into the Japanese yen and the Canadian dollar when they offer relatively higher returns on an *hedged* basis. The evidence for the Japanese yen is particularly intriguing as this currency typically offers the highest hedged excess return among the currencies in our portfolio, with the lowest volatility (around 40 basis points). A one standard deviation change in Japanese yen hedged excess returns is estimated to trigger a reallocation by around 150 basis points in its portfolio share. As regards the other currencies, the lower sensitivity of portfolio shares to hedged returns is consistent with recent findings of relatively exchange-rate inelastic demand for forwards for the investment fund sector as a whole (Wallen (2022), Bräuer and Hau (2022)). Therefore, our results suggest that the sign, magnitude and persistence of CIP deviations on government bonds do affect the portfolio choice of mutual funds, which in turn drive large capital flows from the United States.

This study offers additional insights that are relevant for the theoretical and empirical literature on the portfolio choice of private institutional investors. First, we find evidence of strong frictions, since the coefficients associated with the lagged currency shares in our regressions are positive and statistically significant. Second, there is some tentative evidence of *currency momentum* for debt issued by the euro area and Japan, as the response of currency shares to past currency movements is positive. Third, we ask whether global financial turbulence, proxied by the VIX index measuring investors' risk aversion, is associated with a reallocation within our portfolio of safe assets. We find that when global risk is on the rise, US mutual fund investors repatriate their investments towards US government debt securities, mainly at the expenses of euro-denominated ones. Finally, we study whether the low interest environment of the past decade had an impact in the currency allocation. When the policy rate of the Federal Reserve is low, US mutual funds rebalance their sovereign portfolio shares more intensely in response to excess returns for the euro, consistently with the mechanism of searching for yield abroad in a low interest rate environment. Interestingly, times of tight monetary policy in the US are instead associated with *higher* portfolio shares for the yen and the Swiss franc, as well as a lower sensitivity to their excess returns. This finding is consistent with "search for safety" behaviour, but only to the extent that high US policy rates contain information on global financial stress that is distinct from investors' risk aversion as measured by the VIX, as suggested by Habib and Venditti (2019) .

Our paper relates to three main strands of literature. The first one is the analysis of the demand for safe assets. Krishnamurthy and Vissing-Jorgensen (2012) argue that the premium commanded by US Treasuries over other assets with similar credit risk is a reflection of demand by investors that value their safety and liquidity even at the expense of lower returns. Jiang et al. (2021) posit that that foreigners have a special demand for U.S. dollar, buying U.S. Treasuries when they are expensive. Jiang et al. (2022) provide further empirical support by documenting low rates of returns on Treasuries for foreign investors in particular. Tabova and Warnock (2022) partly challenges this view by providing evidence of elastic demand for Treasuries from foreign private investors, who actually achieve high returns on their Treasury portfolio. In a similar vein, Fang et al. (2023) and Eren et al. (2023) find that, on aggregate, mutual funds display a particularly yield-elastic demand for advanced-economy sovereign bonds.[5] We contribute to this debate by studying the demand for safe assets issued by several countries, not only US Treasuries, and from the perspective of US rather than foreign investors. Our analysis highlights a strong reaction of mutual funds' demand to the excess returns of safe government debt securities. This elasticity is different across investment currencies and also appears to depend on the opportunities provided by deviations in the CIP. These deviations allow US-based investors to obtain higher returns on a currency-hedged basis when they invest invest in non-US dollar currencies.

The second strand of related literature examines the role of mutual funds and the search for yield in driving capital flows across countries. As regards mutual funds, these studies include the analysis of both flows in and out of funds, and the portfolio choice of managers. A notable contribution in this area is Raddatz and Schmukler (2012), which finds that both injections into/redemptions

[5] Faia et al. (2022) reports high sensitivity of euro area mutual funds to the returns of corporate bonds as well.

from funds and changes in country shares by asset managers respond to country returns. Other papers that zoom in specifically on portfolio choice include Falkenstein (1996) and Camanho et al. (2022), but they focus on US *equity* funds.[6] The latter fits most closely with our analysis in that it studies portfolio rebalancing in response to foreign excess returns, taking exchange rate movements into account. Several papers use mutual funds data too, but are concerned mainly fund flows rather than the portfolio choice within each fund. Some examples are Kroencke et al. (2015), Banegas et al. (2022), Fratzscher et al. (2016), Hau and Lai (2016), Fratzscher et al. (2018), and Bubeck et al. (2018). The findings suggest that both conventional and unconventional monetary policies result in substantial mutual fund flows, with a stronger effect for outflows from corporate bond funds and inflows into equity funds. Other studies investigate the search for yield and the resulting capital flows, not necessarily driven by mutual funds. Frankel and Engel (1984) and Bohn and Tesar (1996) are among the first papers to test empirically the asset demand implications of portfolio choice models and to uncover evidence of search for yield behaviour. More recent papers focus on risk-taking, including Ammer et al. (2016) and Ammer et al. (2019). They use confidential data on US bond holdings by foreign investors to show that low interest rate in investors' residence countries lead to inflows into US equities and corporate bonds, especially in the high-yield, high-risk segment. Ahmed et al. (2023) considers the relationship between foreign excess returns, currency hedging and search for yield, but focusing instead on the investment of euro area investment funds in US corporate bonds. They find that institutional investors hedge their foreign exchange risk exposure, which compresses their excess returns and leads them to shift their portfolio towards riskier corporate bonds when US monetary policy tightens.

We extend the literature on mutual funds and capital flows across a number of dimensions. First, we provide a systematic analysis of the determinants of US mutual funds' demand for *safe* government bonds, within a large sample in terms of total Assets under Management (AuM) and coverage of the bond fund universe. In particular, we zoom in on a portfolio of safe government debt securities and calculate *fund-specific* excess returns that closely reflect the idiosyncratic incentive for each fund to invest in a given country/currency. Similarly to other studies, we find that currency excess returns are an important driver of active reallocation on the fund managers' part. We also complement the literature by highlighting how injections and redemptions by investors are not the sole key determinant of capital flows originating from mutual funds. Instead, our results show that the

[6]This feature is common to much of the literature, mainly due to better data availability. Papers that do use data on bond funds, such as Raddatz and Schmukler (2012),Raddatz et al. (2017) and Cenedese and Elard (2021), typically rely on rather limited samples.

response of managers' portfolio choices to international investment opportunities plays a significant role too.

Finally, this paper is also related to studies of portfolio frictions. Giglio et al. (2021) establishes that equity portfolios of US Vanguard funds move sluggishly in response to expected returns, consistently with investor inattention and trading frequency frictions. Similarly, Bacchetta et al. (2020) shows that US equity funds respond significantly to expected returns, but portfolio frictions lead to a weaker and more gradual response. We confirm in the context of bond funds the importance of both the lagged share and passive changes due to valuation effects in explaining portfolio shares, mirroring the findings of Bacchetta et al. (2020).

The remainder of the paper is structured as follows. Section 3.2 provides background and descriptive analysis on the mutual funds portfolio data. We present the results of our econometric models of portfolio shares in Section 3.3. Finally, Section 3.4 concludes.

3.2 Data

3.2.1 Description of the dataset

We use a commercial dataset, Refinitiv Lipper, which provides detailed information regarding the portfolios of US-domiciled mutual funds. We restrict the scope of the analysis to fixed-income funds and exclude mixed-funds which may have an incentive to substitute equity for bonds in response to common shocks that affect all bond yields. Our initial sample includes 880 funds with an active management style, which we observe quarterly from 2010 Q1 to 2021 Q4.[7] Our object of interest is the share of highly-rated government debt in portfolios of advanced economy government debt with high credit ratings. The share $s^j_{i,t}$ of country (currency) j bonds held by fund i in quarter t is calculated as the ratio of the market value of country (currency) j government bonds, summed over all maturities; to the market value of government bonds of all countries of interest, summed over all maturities.

We select countries that maintained an S&P credit rating of AA or above rating throughout our sample period: Australia, Canada, highly-rated Euro Area countries, Japan, Switzerland, the United Kingdom and the United States.[8] The

[7]Our panel is unbalanced, since funds enter and drop out of the sample. In appendix 3.E.4, we perform robustness checks on both descriptive and econometric analysis, to ensure that our results are not affected by sample composition.

[8]Consistently with our focus on safe government bonds, we include in the Euro Area share

government bonds held in this portfolio are almost exclusively denominated in domestic currency – the country of issuance and currency denomination therefore overlap and we will be using "country" and "currency" interchangeably when mentioning portfolio shares.

The rationale for choosing these countries in particular is twofold. First, we want to study safe assets, looking at a portfolio of major international currencies with a low degree of credit risk. These are also the major currencies that account for the bulk of foreign exchange reserves in the IMF's COFER dataset and are particularly valued for their safety by official investors with a conservative investment mandate. Second, we include both currencies for which CIP deviations offer on average an extra return from the point of view of a US dollar investor (EUR, JPY), and those for which the CIP deviation is usually negative (AUD). The aim is to investigate whether the reaction to hedged excess returns changes with the sign of the CIP deviations. We remove from the sample all fund-quarter observations for which there are no holdings of sovereign bonds for any of these countries.

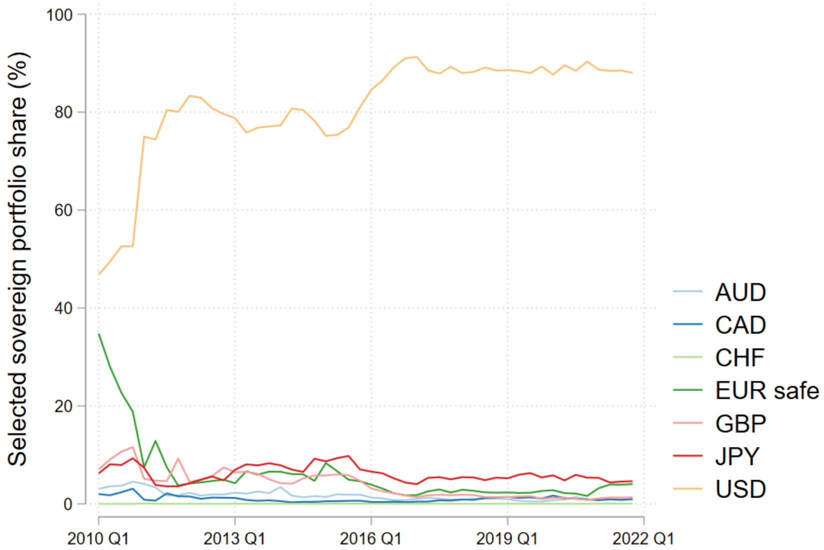
There is a strong *home bias* in the portfolio of US-based funds that must be addressed before starting our empirical analysis, since we are interested in the decisions of fund managers with a diversified international portfolio.[9] Our initial sample includes more than 600 funds that invest only in US Treasuries, with an average AuM of \$1.6 billion. As a result, the aggregate portfolio share s_t^j is strongly biased towards US Treasuries, accounting for more than 80% of the total portfolio (see Figure 3.3). The share of debt securities issued by highly-rated euro area economies, Japan and the United Kingdom ranges between 5% and 10%. To account for this home bias, we exclude all funds that have an average portfolio share greater than 95% in any of the countries of interest.[10] The sample thus restricted includes 186 funds. Figure 3.4 plots aggregate portfolio shares in the restricted sample excluding funds with a country focus. The US share drops to a level much closer to that of other countries, indicating a sub-sample of internationally-oriented funds. As a robustness check, in appendix 3.E.5, we show that the results of the analysis including only funds whose average portfolio is close to an International CAPM benchmark are similar to the results using this portfolio.

only bonds issued by countries which maintained an S&P credit rating of AA or above throughout our sample period. Namely, they are Austria, Belgium, France, Germany and the Netherlands.

[9] Home bias is a common feature of international portfolios, see for example Hau and Rey (2008), Coeurdacier and Rey (2013), Coeurdacier and Gourinchas (2016)

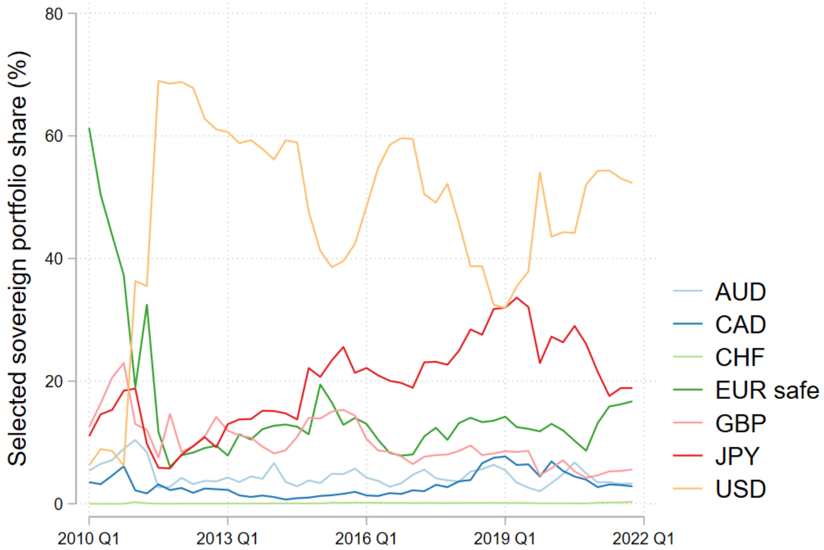
[10] In addition, we also exclude funds that never invest in country j when studying the allocation towards currency j. Note that the evolution of flows into funds with a geographical focus on a single country can provide useful information on the choice of retail investors, but this is outside the scope of this study.

Figure 3.3. Aggregate portfolio shares: full sample



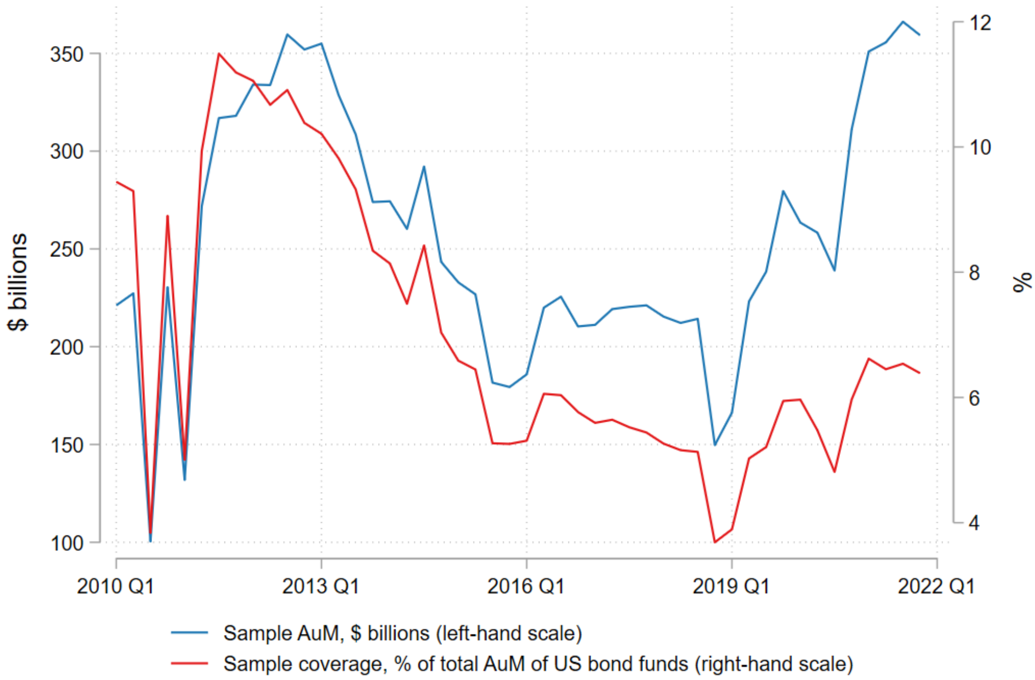
The chart reports the aggregate currency shares in a portfolio of selected sovereign issuers with an elevated rating standard (AA or more). The aggregate share s_t^j of country (currency) j bonds in quarter t is calculated as the ratio of the market value of country (currency) j government bonds, summed over all maturities and all funds; to the market value of government bonds of all countries of interest, summed over all maturities and all funds. Source: Refinitiv Lipper.

Figure 3.4. Aggreagate portfolio shares: sample excluding funds with a country focus



The chart reports the aggregate currency shares in a portfolio of selected sovereign issuers with an elevated rating standard (AA or more), excluding funds that have an average portfolio share greater than 95% in any of the countries of interest. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. The aggregate share s_t^j of country (currency) j bonds in quarter t is calculated as the ratio of the market value of country (currency) j government bonds, summed over all maturities and all funds; to the market value of government bonds of all countries of interest, summed over all maturities and all funds. Source: Refinitiv Lipper.

Figure 3.5. Sample excluding funds with a country focus: coverage



Sample coverage calculated as total AuM of all funds in the sample excluding funds that have an average portfolio share greater than 95% in any of the countries of interest, divided by total AuM of all bond mutual funds domiciled in the US. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Source: Refinitiv Lipper and Federal Reserve Finanial Accounts of the United States, Table F.122

Figure 3.5 shows that the total AuM of funds in our sample excluding funds with a country focus varies over time, ranging from $100 billion in 2010 Q3, to $366 billion in 2021 Q3. This corresponds to a coverage of about 4 to 11% of the AuM held by all US-based mutual funds investing in the fixed-income market. The size and coverage of our sample are larger than those of previous studies using portfolio-level data on fixed-income mutual funds. For example, Cenedese and Elard (2021) use EPFR data covering a maximum of 75 bond funds with aggregate $106 billion AuM, while Raddatz and Schmukler (2012) report total AuM ranging from $10 to $100 billion out of 121 bond funds.
Sovereign bonds constitute a minority of assets in the sample, with the average fund holding $482 million worth of sovereign bonds, which is roughly one fifth of total assets (see Table 3.1). The remaining assets consist primarily of corporate bonds, cash, and derivatives. The debt issued by the highly-rated countries that constitute our portfolios account for more than a half (approximately 57%)

of holdings of sovereign debt of the average fund. Panel B of Table 3.1 reports portfolio shares in the sample. The US share across time and funds averages around 43%, while highly-rated euro area issuers and Japan have an average share 16%. These currency shares are not very far from a theoretical benchmark from the international-CAPM, where the US debt would represent 50% of the market capitalisation of our portfolio of currencies, the highly-rated euro-denominated debt 13% and Japanese debt 23%. Therefore, our selection procedure results in a subsample of geographically well-diversified funds displaying substantial time-series and cross-sectional variation in country shares, and whose average portfolio broadly reflects relative sizes in the sovereign bond market.

Table 3.1. Sample excluding funds with a country focus: summary statistics

	N	Mean	SD	Min	P5	P95	Max
A. Fund characteristics							
Assets under management ($Mil.)	4,823	2,544	10,528	0.10	11.20	8,470	179,914
Total sovereign holdings ($Mil.)	4,906	482	1,669	0.03	2.51	2,295	46,539
Selected sovereign holdings ($Mil.)	4,906	277	1,126	0.01	0.77	1,277	29,880
Reporting quarters	4,906	41	11	1	13	48	48
B. Selected sovereign portfolio shares							
United States	4,906	0.43	0.44	0	0	1.00	1
Euro Area safe	4,906	0.16	0.24	0	0	0.69	1
Australia	4,906	0.09	0.22	0	0	0.64	1
Canada	4,906	0.04	0.13	0	0	0.23	1
Japan	4,906	0.16	0.23	0	0	0.61	1
Switzerland	4,906	0.00	0.02	0	0	0.01	0.56
United Kingdom	4,906	0.12	0.20	0	0	0.52	1

Statistics are calculated on the sample that excludes funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Highly-rated sovereign holdings include government bonds of all maturities issued by Australia, Canada, euro area highly-rated issuers (Austria, Belgium, Germany, France, the Netherlands), Japan, Switzerland, United Kingdom and United States. Shares refer to the portfolio that comprises all debt securities issued by the selected highly-rated sovereigns. Source: Refinitiv Lipper.

Comparing the summary statistics for fund characteristics and portfolio shares with those pertaining to the whole sample in Table 3.10 in Appendix 3.B reveals that the selection procedure did not significantly alter the size and aggregate exposure to sovereign holdings of funds in the sample. At the same time, the portfolio of the average fund matches much more closely that of a global investor.

3.2.2 Disentangling active and passive reallocation in portfolio shares

The currency portfolio shares, our main variable of interest, is affected not only by the active reallocation by fund managers, but potentially also by valuation effects from exchange rate and bond price movements. It is necessary to disentangle these two components to study whether the active reallocation may be influenced by return differentials. To do that, the change in fund i's currency j portfolio share $\Delta s_{i,t}^{j}$ can be decomposed as

$$\Delta s_{i,t}^{j} = \Delta s_{i,t}^{j,A} + \Delta s_{i,t}^{j,P,R} + \Delta s_{i,t}^{j,P,XR}. \tag{3.1}$$

Here, $\Delta s_{i,t}^{j,P,R}$ is the passive change in portfolio share due to changes in the issue-currency market value of country j bonds; $s_{i,t}^{j,P,XR}$ is the passive change in portfolio share due to the appreciation or depreciation of the issuance currency of country j bonds *vis à vis* the dollar; and $\Delta s_{i,t}^{j,A}$ is the change in portfolio shares due to active rebalancing on the fund manager's part. This decomposition follows closely the method used in Curcuru et al. (2011) and Bubeck et al. (2018), among others.

The passive reallocation due to bond returns in the currency of issuance is

$$\Delta s_{i,t}^{j,P,R} = s_{i,t-1}^{j} \left(\frac{R_t^j}{\bar{R}_{i,t}} - 1 \right), \tag{3.2}$$

where R_t^j is the growth in the total return index of country j's government bonds between quarters $t-1$ and t, including both changes in prices and the reinvestment of coupon payments and interests, averaged across all maturities; and $\bar{R}_{i,t} = \sum^k s_{i,t-1}^{k} R_t^j$ is the overall performance of the portfolio, more precisely the weighted average of the performance of the total return indices of government bonds in fund i's portfolio. Intuitively, if the return of securities issued by country j is higher with respect to the rest of the portfolio, the share of country j will automatically increase even if the fund manager did not perform any active portfolio reallocation.

Similarly, the passive reallocation due to exchange rate effects is

$$\Delta s_{i,t}^{j,P,XR} = s_{i,t-1}^{j} \left(\frac{A_t^j}{\bar{A}_{i,t}} - 1 \right), \tag{3.3}$$

where $A_t^j = \frac{E_t^{USD/j} - E_{t-1}^{USD/j}}{E_{t-1}^{USD/j}}$ is the percentage appreciation of currency j with respect to the dollar between quarters $t-1$ and t, with $E_t^{USD/j}$ denoting the spot exchange rate in terms of dollars per unit of currency j ; and $\bar{A}_{i,t} = \sum^k s_{i,t-1}^{k} A_t^j$

is the average appreciation of fund i's portfolio with respect to the dollar between quarters $t-1$ and t [11]. Once again, passive reallocation measures are defined *relative* to other countries in the portfolio. For instance, assuming that the US dollar depreciates across the board against all the other currencies in the portfolio. Not only the dollar share will decrease, but also the share of currency j will increase passively due to the currency valuation effects if its appreciation with respect to the dollar is stronger than the appreciation of other currencies in the portfolio.

Figure 3.6 in the appendix shows the decomposition of changes in aggregate portfolio shares into active and passive reallocation components of the changes in currency shares in the aggregate portfolio, indicating that active reallocation dominates for all currencies.

Table 3.2. Fund-level active reallocation

	N	Mean	SD	Skewness	Kurtosis	Min	P1	P5	P95	P99	Max
$\Delta s_{i,t}^{AUD,ACT}$	3,442	-0.26	11.31	-0.52	32.29	-100	-43.50	-9.87	9.78	37.37	100
$\Delta s_{i,t}^{CAD,ACT}$	3,333	0.04	8.13	1.63	56.97	-99	-26.73	-6.05	5.91	25.63	100
$\Delta s_{i,t}^{CHF,ACT}$	692	0.00	1.00	-5.84	122.49	-16	-2.20	-0.65	0.70	2.90	8
$\Delta s_{i,t}^{EUR_S,ACT}$	3,668	-0.51	15.43	0.06	21.71	-100	-59.59	-17.90	14.16	59.64	100
$\Delta s_{i,t}^{JPY,ACT}$	3,425	0.23	12.43	0.20	26.83	-100	-42.95	-13.05	13.69	42.98	100
$\Delta s_{i,t}^{GBP,ACT}$	3,442	0.17	13.15	0.33	25.24	-100	-46.04	-14.07	15.05	47.71	100
$\Delta s_{i,t}^{USD,ACT}$	4,359	0.29	17.39	0.05	18.60	-100	-71.93	-17.14	20.55	73.59	100

All variables in percentage points. Summary statistics are calculated over the distribution of fund-quarter observations in the sample excluding funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries.

Table 3.2 reports summary statistics for the active reallocation across currencies at the fund-level. We may note that, even though the average change in the share is close to zero, reflecting inertia in the portfolio shares, large changes in the portfolio share by more than 10 percentage points quarter-on quarter (see standard deviation) are not infrequent for most currencies. Importantly, we identify the presence of several outliers with active rebalancing in the portfolio share by 100% in both directions. [12] Most likely, these reflect misreporting by the fund manager or the data provider (e.g when there are two consecutive large changes with the opposite sign) or, if not a reporting issue, they represent radical changes in the strategy that are most likely unrelated to excess returns. In order to prevent these

[11]The very near totality of government bond in our sample are denominated in the issuer country's currency.

[12]Appendix 3.B provides the same set of summary statistics for passive reallocation.

outliers from affecting our results, we augment our regressions with a dummy that takes a value of 1 for fund-quarters observations in the top or bottom 1% by active rebalancing $\Delta s^{j,A}_{i,t}$.[13]

3.3 Econometric analysis

3.3.1 Baseline specification

Our variable of interest is the currency share in the portfolio of highly-rated government debt securities. Rearranging the terms of equation 3.1, we may see that the currency share at time t, $s^j_{i,t}$, is a function of its value in the previous period, the active reallocation by fund managers and the passive reallocation due to price effects and exchange rate effects that have been described in Section 3.2.2:

$$s^j_{i,t} = s^j_{i,t-1} + \Delta s^{j,A}_{i,t} + \Delta s^{j,P,R}_{i,t} + \Delta s^{j,P,XR}_{i,t}. \tag{3.4}$$

Therefore, controlling for the past currency share, $s^j_{i,t-1}$, and for the valuation effects, $\Delta s^{j,P,R}_{i,t} + \Delta s^{j,P,XR}_{i,t}$, we can attribute the residual variation in the currency share exclusively to the active reallocation component, $\Delta s^{j,A}_{i,t}$, which, in turn, we aim to explain with the excess currency return, $r^{ex,j}_{i,t}$. Concretely, our empirical approach consists of regressing fund i's portfolio share in currency j on its own lag, fund-level excess returns, and passive reallocation components.[14] We estimate a separate model for each currency j to examine variations in investor behaviour that may arise from factors specific to the destination country, such as its status as a safe haven and the direction of dollar CIP deviations.[15] The baseline regression equation is

$$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}, \tag{3.5}$$

where α^j_i and γ^j_t are fund and quarter fixed effects; $\mathbb{1}^{Out,j}_{i,t}$ is a dummy that takes value 1 if observation i, t is below the 1st percentile or above the 99th percentile

[13] In Appendix 3.E.3, we show that the vast majority of our results survives a more stringent definition of both country-focused funds, selected as $\bar{s}^j_i \geq 0.90$; and outliers, selected as observations in the top or bottom 2.5% by $\Delta s^{j,A}_{i,t}$

[14] Note that the Nickell (1981) bias due to the inclusion of the lagged share is small in our setting because of the relatively large T. It is bounded at around 3% for the average fund run of 30 quarters, while it drops to circa 2% if we consider the full 40-quarter run of our sample.

[15] In Appendix 3.H we present results from estimating the currency-specific models jointly by exploiting the restriction $\sum_j s^j_{i,t} = 1$

by $\Delta s_{i,t}^{j,A}$, and 0 otherwise; $r_{i,t}^{ex,j} = r_t^j - \sum_{k \neq j} f_{i,t-1}^k r_t^k$ is currency j' s excess return with respect to the fund's sovereign portfolio; $f_{i,t-1}^k$ is the lagged share of currency k in the portfolio that excludes currency j; and $r_t^j = r_t^{j,unh}$ or $= r_t^{j,fwd}$ is the unhedged or hedged return of country j's government bonds. The other variables are defined in Section 3.2.2. We estimate the model separately for each currency, j.

Before discussing our main explanatory variable of interest, the fund-specific excess return of country j government bonds $r_{i,t}^{ex,j}$, it is worth noting that this model allows us to test two additional hypothesis regarding the drivers of the current currency portfolio share: (i) *portfolio frictions* and (ii) *valuation effects*. First, the inclusion of the lagged share, $s_{i,t-1}^j$, allows us to examine the share autocorrelation, a marker of delayed portfolio adjustments. In line with Bacchetta et al. (2020), we interpret the parameter β_1^j as a gauge of *portfolio frictions*. A positive β_1^j implies a positive correlation between past and current shares, given currency-specific excess returns, indicating slow portfolio adjustment. Such sluggishness can be attributed to several factors such as delayed reaction to new information (highlighted in Bohn and Tesar (1996) and Froot et al. (2001)), transaction costs, or targeted currency-share levels. Second, the coefficients on the passive reallocation components, β_3^j and β_4^j, offer insights into whether fund managers proactively alter portfolio shares in response to *valuation effects*. A coefficient of β_3^j or β_4^j equal to 0 denotes that any changes in share due to valuation effects are completely offset by active reallocation of the opposite sign. Conversely, coefficients between 0 and 1 imply partial offsetting. A coefficient equal to 1 implies full pass-through of valuation effects to share changes, so that the current share is only a function of the past level and valuation effects. Lastly, a coefficient of β_3 or β_4 greater than 1 suggests that the active reallocation goes in the same direction of valuation effects, an indication of a *currency momentum* strategy. For example, if fund managers increase their investments in Japanese bonds after witnessing the relative appreciation of the Japanese yen against the dollar between the previous and current quarters, we would observe $\beta_4^{JPY} > 1$.

3.3.2 Unhedged and hedged excess returns

The main explanatory variable of interest is the fund-specific excess return of country j government bonds $r_{i,t}^{ex,j}$. It measures the attractiveness of investing in currency j relative to other currencies in the fund's portfolio. We use the lagged shares to calculate excess returns, both as a benchmark observable to fund managers at the point of decision, and to allay concerns of multicollinearity with current shares. We construct excess returns on a fund-specific basis, so that we can exploit fund and time fixed effects to single out idiosyncratic variation in the investment opportunity available to funds.

Fund fixed effects control for fund-specific characteristics, such as management style or geographical focus, that might explain a significant portion of the variation in share levels. Time fixed effects control for any global and country j-level variables that might affect both demand for sovereign bonds and their yield. Shocks to aggregate demand by global investors constitute a prime example, as they have been shown to have an important impact on bond pricing[16]. Importantly, time fixed effects also absorb all time-varying factors that affect currency-specific demand for sovereign bonds. For example, changes in investors' risk appetite can result in generalised flight-to-quality behaviour towards US Treasuries, introducing a positive correlation between returns and portfolio shares in our regressions.

If we are ready to assume further that fund-specific demand shocks do not affect the price of government bonds or exchange rates, the residual within-fund variation in fund-specific excess returns after controlling for global and currency-specific aggregate demand shocks identifies the average sensitivity of portfolio shares to excess returns.

We calculate excess returns on both a hedged and an unhedged basis, which correspond to a different definition of r_t^j. Our portfolio includes countries that in a traditional carry trade play the role of investment country, like Australia, as well as funding country, like Japan. Furthermore, average country interest rates are negatively correlated with the sign of CIP deviations with respect to the dollar, as shown in Borio et al. (2016). Then, from the perspective of a US investor, the incentives to invest in a given currency might point in opposite directions on an unhedged or hedged basis. We aim to capture these competing forces by analysing responses to unhedged and hedged returns separately.

The unhedged return $r_t^{j,unh}$ is simply the yield of country j's government bond in the domestic currency, averaged over the 3 month, 1 year, 2 year and 5 year maturities. We assume that expectations of the future exchange rate are equal to their current value, as it has been observed since Meese and Rogoff (1983) that exchange rates behave very closely to a random walk, making their current value a reasonably accurate forecast for future rates. [17]

[16] see for example Faia et al. (2022) and Schmidt and Yeşin (2022)

[17] Note that in this paper we adopt the perspective of the fund manager investing at time t. Therefore, we need a proxy for the manager's expectations of exchange rates at $t+1$ to assess the predictability of excess returns, rather than using exchange rates at $t+1$ as in the Fama (1984) regressions used by most of the macro-focused literature. Instead, in keeping with the random

We calculate hedged returns as $r_t^{j,fwd} = r_t^{j,unh}\frac{F_t^{USD/j}}{E_t^{USD/j}}$, where $F_t^{USD/j}$ is the forward exchange rate in terms of dollars per unit of currency j averaged over the 3 month, 1 year, 2 year and 5 year maturities. We choose $F_t^{USD/j}$ because forwards are the preferred exchange rate hedging instruments of US mutual funds. The evidence in Sialm and Zhu (2022) shows that 90% of international bond funds use forwards to manage their foreign exchange exposure. Hedged excess return could be interpreted as a weighted average deviation from CIP for country j's government bonds with respect to other countries in the portfolio. The hypothesis we want to test is whether mutual funds modify their country shares based on the resulting cross-currency differences in hedged returns. [18]

Summary statistics for unhedged and hedged returns are displayed in the left-hand side of Table 3.3. A clear pattern of cross-currency heterogeneity emerges from the averages: high-interest rate currencies (e.g the Australian dollar) offer positive unhedged excess returns but negative hedged excess returns, while the opposite holds for low-interest-rate currencies (e.g. the Japanese yen). There is a clear mapping to the negative cross-currency correlation between interest rate levels and signs of dollar CIP deviation observed in Borio et al. (2016). Funds then face markedly different incentives to invest in any given currency on a unhedged or hedged basis. Note also that the negative unhedged and positive hedged excess returns of Switzerland and Japan are consistent with different signs of deviations from UIP and CIP *vis à vis* the dollar for safe haven currencies, documented in Bacchetta et al. (2023).

walk view of exchange rates, we use the exchange rate at t as the fund manager's forecast for $t + 1$. This approach could be refined by using data on exchange rate expectations by market participants.

[18]It is important to point out that we are not suggesting that US mutual funds engage in CIP arbitrage. In an intermediary-based FX pricing framework à la Gabaix and Maggiori (2015), they are best conceptualised as originators of hedging flows from different countries into dollars, who take as given the forward rates offered by international dealer banks. In turn, FX market rates are possibly driven by the banks's leverage constraints and the heterogeneous nature of aggregate demand across currencies. As shown by Rime et al. (2022), CIP arbitrage necessitates short-selling and possibly expanding the balance sheet with leverage. Therefore, "true" CIP deviations that generate arbitrage profits must take into account both the institution-specific funding costs and balance sheet constraints of FX intermediaries, as well as transaction costs. US mutual funds rarely engage in short selling, as shown in An et al. (2021). Furthermore, they face rather stringent leverage regulations, and when leverage does arise it typically does so in the form of index funds synthetically increasing their exposure to the reference index through derivatives, as shown in Boguth and Simutin (2018). Therefore, bond mutual funds are not well-equipped to act as CIP arbitrageurs.

Table 3.3. Fund-level unhedged and hedged returns

	Summary statistics							Autocorrelations			
	N	Mean	SD	Skewness	Kurtosis	Min	Max	AC(1)	AC(2)	AC(3)	AC(4)
A. Unhedged excess returns (percentage points)											
$r_{i,t}^{ex,AUD,unh}$	4,906	1.47	1.33	0.46	2.68	-0.91	5.13	0.90***	0.84***	0.77***	0.69**
$r_{i,t}^{ex,USD,unh}$	4,906	0.56	1.07	-0.91	6.36	-4.38	3.02	0.92***	0.85***	0.76***	0.67**
$r_{i,t}^{ex,CAD,unh}$	4,906	0.47	0.76	-1.04	7.39	-3.51	2.43	0.48***	0.31	0.20	0.17
$r_{i,t}^{ex,GBP,unh}$	4,906	-0.02	0.77	-1.39	6.24	-4.01	1.31	0.39***	0.24	0.16	0.17
$r_{i,t}^{ex,JPY,unh}$	4,906	-0.68	0.79	-1.71	6.76	-4.92	0.56	0.73***	0.55***	0.34	0.16
$r_{i,t}^{ex,EUR_s,unh}$	4,906	-0.82	0.89	-0.45	3.86	-4.15	1.74	0.79***	0.74***	0.66**	0.70**
$r_{i,t}^{ex,CHF,unh}$	4,906	-1.02	0.82	-1.08	4.33	-4.56	0.79	0.76***	0.69***	0.58**	0.54*
B. Hedged excess returns (percentage points)											
$r_{i,t}^{ex,JPY,fwd}$	4,906	0.56	0.40	3.05	14.93	-0.01	3.15	0.44***	0.29*	0.24	0.17
$r_{i,t}^{ex,CHF,fwd}$	4,906	0.39	0.44	2.38	10.54	-0.32	2.95	0.58***	0.48**	0.40*	0.26
$r_{i,t}^{ex,EUR,fwd}$	4,906	0.22	0.44	2.48	11.32	-0.49	2.79	0.36**	0.22	0.14	0.16
$r_{i,t}^{ex,GBP,fwd}$	4,906	0.02	0.44	2.78	13.78	-1.01	2.68	0.09	0.08	-0.03	0.04
$r_{i,t}^{ex,CAD,fwd}$	4,906	0.01	0.42	3.06	15.59	-0.70	2.66	0.15	0.05	0.01	0.18
$r_{i,t}^{ex,AUD,fwd}$	4,906	-0.17	0.55	1.59	8.14	-1.49	2.66	0.19	0.10	0.20	0.24

Summary statistics are calculated over the distribution of fund-quarter observations in the sample excluding funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Autocorrelations up to 4 quarters, AC(q), are calculated over the sample of cross-sectional average excess returns for each currency $\bar{r}_t^j = \frac{1}{I}\sum_{i=1}^{I} r_{i,t}^{ex,j}$. Standard errors for autocorrelations are calculated using the Bartlett (1946) formula. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

To assess the predictability of excess returns, the right-hand side of Table 3.3 displays autocorrelations for cross-sectional average excess returns for each currency $\bar{r}_t^j = \frac{1}{I}\sum_{i=1}^{I} r_{i,t}^{ex,j}$. Unhedged returns display a positive, significant, and large autocorrelation up to four quarters ahead for most currencies. On the other hand, hedged returns display much smaller autocorrelation coefficients, significant only up to a maximum of two quarters, and only for the Japanese yen, Swiss franc and euro. Therefore, current unhedged returns provide a better signal for forward-looking investors than their hedged counterpart, and so we would expect an overall stronger reaction of portfolio rebalancing to excess returns on an unhedged basis. We discuss the predictability of excess returns in more detail in Appendix 3.C.

3.3.3 Search for yield without hedging currency risk

Table 3.4 reports results from the baseline model using unhedged excess returns as the explanatory variable for the active reallocation by fund managers. We find that the coefficient associated with the unhedged excess return is positive and statistically significant for several currencies (AUD, CHF, GBP, EUR and USD), indicating a broad-based search for yield behaviour among safe government bonds. The size of the coefficients is also economically relevant. For instance, an increase by one percentage point in the excess return of euro area safe government bonds, slightly more than one standard deviation (see Table 3.3), leads to an increase by

0.84 percentage points in the portfolio share of the euro for the average fund. Such portfolio adjustment triggers capital flows from the United States towards the euro area economies issuing highly-rated debt securities in the order of magnitude of \$300 million, amounting to around 2 percent of total foreign flows into highly-rated euro area government debt securities on a quarterly basis, according to balance of payments data. The size of the impact of excess unhedged returns on the currency share of the US dollar, the Australian dollar, and the pound sterling is of a similar magnitude, even though slightly less significant from a statistical point of view. Note that funds do not rebalance towards the Japanese yen, in response to unhedged excess returns, and the rebalancing towards the Swiss franc is statistically significant but small. Note that these two currencies offered the highest positive *hedged* excess returns, on average, in our sample period (see Table 3.3). In the next subsection, we shall investigate whether the portfolio shares of these two currencies are more sensitive to their returns on a hedged rather than unhedged basis.

The results in Table 3.4 show strong evidence of portfolio frictions and a rather limited impact of valuation effects on currency shares. First, the coefficient on the lagged currency share is large, ranging between 0.63 for the euro and 0.77 for the Australian dollar, and always highly statistically significant. This suggests that portfolio frictions are important as funds tend to benchmark portfolio shares and seldom deviate from these benchmarks.[19] Second, in general, the coefficients associated with valuation effects are not statistically different from zero, suggesting that fund managers tend to offset valuation effects and do not let the currency shares change passively as a result of movements in bond prices and exchange rates. Nevertheless, fluctuations in the exchange rate of the Japanese yen and the euro seem to have an impact. For the Japanese yen, the coefficient associated with the exchange rate valuation effects is close to 1, implying a nearly full pass-through of valuation effects to changes in the share for this currency. For the euro, the coefficient associated with the exchange rate valuation effects is greater than one, suggesting the presence of a currency momentum strategy, whereby fund managers actively increase (or decrease) their exposure to euro-denominated government bonds whenever the euro appreciates (or depreciates) in the previous quarter.

3.3.4 Search for yield hedging currency risk

Table 3.5 shows the results of the model where the unhedged excess return has been replaced by the hedged excess return. This model describes the portfolio

[19] In appendix 3.F we show that adding up to four lags to the model does not reveal any evidence of unit root behaviour, confirmed by the results of unit root tests.

Table 3.4. Baseline unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^{j}_{i,t-1}$	0.77***	0.74***	0.71***	0.70***	0.71***	0.63***	0.74***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)	(0.03)
$r^{ex,j,unh}_{i,t}$	1.03*	0.46	0.10*	0.94*	0.46	0.84***	1.86**
	(0.61)	(0.28)	(0.06)	(0.53)	(0.63)	(0.27)	(0.90)
$\Delta s^{j,P,R}_{i,t}$	-0.25	-0.17	0.93	-1.69	0.53	-0.27	-2.81
	(4.38)	(5.61)	(4.20)	(2.77)	(2.22)	(1.75)	(1.77)
$\Delta s^{j,P,XR}_{i,t}$	0.26	0.86	2.76	-0.14	0.89**	1.53***	0.50
	(1.11)	(1.45)	(1.83)	(1.11)	(0.40)	(0.48)	(1.16)
Within R^2	0.68	0.62	0.55	0.53	0.57	0.48	0.61
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668	4,359
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	41	42	43	41	42	41	40

Coefficients from regression model
$s^{j}_{i,t} = \alpha^{j}_{i} + \gamma^{j}_{t} + \eta^{j}\mathbb{1}^{Out,j}_{i,t} + \beta_1 s^{j}_{i,t-1} + \beta^{j}_{2} r^{ex,j,unh}_{i,t} + \beta^{j}_{3}\Delta s^{j,P,R}_{i,t} + \beta^{j}_{4}\Delta s^{j,P,XR}_{i,t} + \varepsilon^{j}_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^{j}_{i} \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

choice of those fund managers who do not assume the currency risk and are thus confronted with returns from non-US dollar currencies that include the cost of hedging exchange rate fluctuations. In this case, evidently, a model for the currency share of the US dollar cannot be estimated since there is no currency risk to be hedged. The coefficient associated with the hedged excess return is not always statistically significant and not always positive. However, quite interestingly, the regressions for the Canadian dollar and, importantly, for the Japanese yen show that fund managers active rebalance in response to hedged excess returns. The portfolio adjustment is large. A one standard deviation change in Japanese yen excess returns, around 40 basis points, hedged into US dollars, is estimated to trigger a reallocation by around 150 basis points in its portfolio share. Intriguingly, Japan is one of the few countries displaying predictability in hedged returns, and at the same time dollar CIP deviations *vis à vis* are the largest among currencies in the advanced economy sovereign bond portfolio for the majority of our sample period, as shown in Figure 3.1 and Table 3.3. Therefore, there is some suggestive evidence that US mutual funds exploit their favourable position as suppliers of dollars in the USD-JPY hedging market to seek higher returns.[20]

The relatively low sensitivity of portfolio shares to hedged excess returns, compared with unhedged returns, might be attributed to the low predictability of hedged excess returns for most currencies. These results are in line with the findings of low elasticity of hedging demand to exchange rates in Bräuer and Hau (2022). Our approach differs because our measure of hedged returns includes the forward premium, a more precise measure of the price of hedging than spot exchange rates. In addition, we do not observe holdings of FX forward contracts directly. However, hedged returns are relevant for fund managers only to the extent that positions are actually hedged, so portfolio shares being relatively unresponsive to excess hedged returns is consistent with either low hedge rates, or insensitivity of hedging demand to forward premia.

3.3.5 Search for the safest asset

In this section, we investigate whether US mutual funds increase their exposure to domestic (i.e. US) or foreign safe government bonds in periods of financial stress, searching for the safest among safe assets from the point of view of an investor based in the United States. The previous section demonstrated how US mutual funds actively react to currency-specific opportunities for excess returns. However, we are analysing safe assets that offer desirable properties above and beyond monetary returns. One may wonder whether within the segment of high-rated government bonds some might be considered more desirable by investors because of their safe-asset properties. These safe havens provide a form of insurance that is particularly valuable in times of financial stress, when they experience large purchases by foreign investors in flight to quality episodes (Longstaff (2004), Beber et al. (2009), Habib and Stracca (2015)).

We test the flight to safety hypothesis by adding to the baseline model the VIX, which measures the expected volatility in the US stock market and is generally considered a good proxy of global risk appetite. In this model, flight to safety toward country j would be captured by a positive correlation between the VIX and active reallocation into country/currency j. This specification does not allow

[20] The model for the United Kingdom indicates an active *decrease* in portfolio shares in response to hedged excess returns, even though the statistical significance of this result is not very strong. Note also that predictability of excess returns is very poor for the United Kingdom on both an unhedged and a hedged basis, as shown in Figures 3.7e and 3.8e. Therefore, current excess returns for this currency do not appear to offer a consistent signal for the future.

Table 3.5. Baseline hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.73***	0.71***	0.71***	0.70***	0.63***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-1.64	3.58***	-0.42	-3.60*	3.92**	1.94
	(1.98)	(1.00)	(0.38)	(1.95)	(1.73)	(1.58)
$\Delta s^{j,P,R}_{i,t}$	-0.39	0.02	0.89	-1.89	0.45	-0.31
	(4.30)	(5.47)	(4.17)	(2.86)	(2.20)	(1.74)
$\Delta s^{j,P,XR}_{i,t}$	0.33	0.74	2.77	-0.16	0.93**	1.44***
	(1.05)	(1.44)	(1.82)	(1.08)	(0.42)	(0.53)
Within R^2	0.68	0.62	0.56	0.53	0.57	0.48
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

us to use time fixed effects, as they would absorb the level of the VIX. Therefore, we replace them with a set of destination-country-specific macroeconomic variables to control for factors that affect demand for government bonds in each currency j. We estimate the following model for both hedged and unhedged returns:

$$s^j_{i,t} = \alpha^j_i + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j}_{i,t} + \beta^j_3 VIX_t + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \boldsymbol{\beta^j_6 W^j_t} + \varepsilon^j_{i,t}, \tag{3.6}$$

where VIX_t is the average value of the VIX index in quarter t in standard deviation units, and $\boldsymbol{W^j_t}$ is a vector of country-level controls that includes inflation, which affects the real payoff of country j's government bonds; and the Citigroup Economic Surprise Index, which accounts for macroeconomic shocks that may influence the incentives to invest in country/currency j across all funds. We then interpret $\beta^j_3 > 0$ as consistent with flight-to-quality behaviour toward country j.

A higher VIX index is associated with high risk aversion and retrenchment of the global financial cycle, and active reallocation in such times would indicate that fund managers treat the government debt of that country as a safe haven.

Table 3.6 reports the results including unhedged returns in the regressions.[21] Across most currencies, high risk aversion in financial markets shows no correlation with portfolio shares, even for safe-haven currencies like the Swiss franc and the Japanese yen. In a nutshell, investors might move money in and out of safe government debt (which we do not control here), but it does not seem that they alter their portfolio of safe securities. However, there is some tentative evidence that US fund managers actively reallocate towards domestic government bonds, at the expense of euro area sovereign debt, when global financial risk is on the rise. A one standard deviation increase in the VIX index is associated with an active reallocation of 59 basis points towards US bonds, and of 69 basis points away from euro area bonds. The retrenchment towards domestic government securities is consistent with both a flight-to-quality argument, as US Treasuries are the global safe asset *par excellence*, and with heightened home bias in uncertain times.

3.3.6 The role of central bank policy rates

Much of the literature on search for yield frames this behaviour in the low-interest-rate environment prevailing in the aftermath of the 2007 global financial crisis. As central bankers pushed the policy rate toward zero and engaged in unconventional policy, investors looked for higher returns abroad or in riskier assets. This line of argument has been applied to mutual funds too. Cenedese and Elard (2021) show that unconventional monetary policy operations by the ECB, BoE, BoJ and Fed triggered reallocation of mutual fund portfolio shares away from countries conducting Unconventional Monetary Policies (UMP) towards other advanced economies. Similarly, Kaufmann (2021) documents that accommodative monetary policy shocks by the Fed result in flows into euro area corporate bond funds.

It is then natural to ask whether the search for yield behaviour detected in the safe government bond portfolio of US mutual funds is stronger when the domestic policy rate is low. We test this hypothesis by augmenting our baseline regression model with the US policy rate and its interaction with excess returns. Much like

[21] For the sake of brevity, we report results for hedged returns in Appendix 3.I.2. They confirm the message that the VIX is not a significant driver of active portfolio reallocation for US fund managers.

Table 3.6. Search for safety unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^{j}_{i,t-1}$	0.78***	0.74***	0.71***	0.71***	0.73***	0.64***	0.75***
	(0.03)	(0.03)	(0.10)	(0.04)	(0.04)	(0.03)	(0.03)
$r^{ex,j,unh}_{i,t}$	0.72***	0.48	0.00	0.81	-0.06	1.01***	0.71
	(0.26)	(0.29)	(0.04)	(0.54)	(0.44)	(0.25)	(0.49)
VIX_t	23.12	10.83	-3.03	-10.28	-38.62	-69.17*	59.27*
	(22.77)	(18.51)	(2.57)	(27.21)	(37.03)	(36.21)	(33.02)
$\Delta s^{j,P,R}_{i,t}$	0.67	-1.06	1.03	0.46	-0.09	0.20	-1.48
	(4.47)	(5.46)	(3.76)	(2.45)	(1.61)	(1.54)	(1.70)
$\Delta s^{j,P,XR}_{i,t}$	0.13	0.99	2.66	-0.19	0.82**	0.86*	0.55
	(1.04)	(1.29)	(1.64)	(0.86)	(0.35)	(0.50)	(1.22)
Within R^2	0.68	0.61	0.53	0.51	0.56	0.47	0.59
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668	4,359
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	32	33	32	33	33	33	32

Coefficients from regression model $s^{j}_{i,t} = \alpha^{j}_{i} + \eta^{j}\mathbb{1}^{Out,j}_{i,t} + \beta^{j}_{1}s^{j}_{i,t-1} + \beta^{j}_{2}r^{ex,j,unh}_{i,t} + \beta^{j}_{3}VIX_t + \beta^{j}_{4}\Delta s^{j,P,R}_{i,t} + \beta^{j}_{5}\Delta s^{j,P,XR}_{i,t} + \boldsymbol{\beta^{j}_{6}W^{j}_{t}} + \varepsilon^{j}_{i,t}$. Each column reports results for a different currency j. $\boldsymbol{W_t}$ includes year-on-year inflation for country j in quarter t in percentage points, the Citi Economics Surprise Index for country j in quarter t in standard deviation units, and the VIX in quarter t in standard deviation units. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^{j}_{i} \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

the regressions in Section 3.3.5, we cannot use time fixed effects as they would absorb the level of the US policy rate. Therefore, we estimate fund fixed effects model augmented with country-specific macroeconomic variables and the VIX as a factor that accounts for the global financial cycle, with the aim of controlling for demand shocks. In Appendix 3.I.1 we present the results of models containing only the interaction between the policy rate and excess returns. They allow the use of times and funds fixed effects with the same identification strategy as the main models in Section 3.3. We estimate the following model for both hedged and unhedged returns:

$$s_{i,t}^{j} = \alpha_i^j + \eta^j \mathbb{1}_{i,t}^{Out,j} + \beta_1^j s_{i,t-1}^j + \beta_2^j r_{i,t}^{ex,j} + \beta_3^j cb_t^{US} + \beta_4^j cb_t^{US} \times r_{i,t}^{ex,j} + \beta_5^j \Delta s_{i,t}^{j,P,R} + \beta_6^j \Delta s_{i,t}^{j,P,XR} + \boldsymbol{\beta_7^j W_t} + \varepsilon_{i,t}^j, \quad (3.7)$$

where cb_t^{US} is the average mid-point for the Federal Reserve target rate in quarter t, and $\boldsymbol{W_t}$ is a vector of country-level and global controls. The country-level controls are inflation and the Citigroup Economic Surprise Index. As in Section 3.3.5, we use the VIX as a global factor to capture swings in risk appetite that drive co-movement in global bond prices. We then interpret $\beta_4 < 0$ as evidence of stronger rebalancing into currency j in response to excess returns when the Fed policy rate is low.

Table 3.7 displays the results for the unhedged returns models. There is evidence of higher responsiveness to excess returns for euro area government bonds at times of low US interest rates, albeit the size of the effect is small. A one percentage point decrease in the Federal Funds target rate is associated with a one basis point higher reallocation into the euro area for a given increase in euro area excess returns.[22] This result extends the findings in Kaufmann (2021), which studies investment fund flows into equity and corporate debt, to government bonds as well. Furthermore, low US interest rates are associated with a higher euro area portfolio share conditional on excess returns, which is also consistent with searching for yield abroad.

Strikingly, the results are reversed for the Japanese yen and Swiss franc, the two foreign safe haven currencies in the portfolio. A contractionary Fed policy rate is correlated with a *high* portfolio share, and the interaction coefficient is positive, albeit again very small. High US policy rates lead to a worsening of global financial conditions accompanied by a drop in risk appetite, as argued in Miranda-Agrippino and Rey (2020). In turn, the swings in the global financial cycle caused by both monetary policy and risk aversion shocks are an important driver of international portfolio flows. The fact that US mutual funds actively *increase* their exposure to Swiss and Japanese government bonds when US monetary policy is contractionary, combined with the relative insensitivity to excess returns for these two currencies, suggests flight to quality.

[22]Note that we are interested in the *level* of the interest rate as an indicator of the overall policy stance of the Federal Reserve, rather than monetary policy shocks. Therefore, we do not instrument cb_t^{US}. As a result, we do not interpret coefficients β_3 and β_4 causally. Although monetary policy decisions are plausibly exogenous to mutual funds portfolio choice, the domestic macroeconomic conditions that drive the former might not be.

3.4 Conclusions

One of the main features of safe assets, such as the government debt securities issued by economies with elevated credit ratings, is the relatively low elasticity of their demand with respect to yields. In this paper, we contribute to the mounting evidence that this is not always the case. Cross-currency yield differentials in the sovereign bond market of high-rating issuers can affect the relative appeal of currencies for an important class of investors such as US mutual funds, shaping the overall demand for global safe assets.

Specifically, US-based fund managers actively rebalance towards government bonds offering higher returns than the portfolio-weighted average return on an *unhedged* basis, i.e. without hedging the currency risk. The size of the effect is significant. A change by one percentage point in their unhedged excess return, approximately one standard deviation, leads to a change in their portfolio shares by around 100 basis points, on average, across several currencies. This portfolio adjustment has also important implications for capital flows. For instance, an increase in the excess return of euro area government debt securities by one percentage point would trigger capital flows from the United States towards the euro area economies issuing highly-rated debt securities in the order of magnitude of $300 million, amounting to around 2 percent of total foreign flows into highly-rated euro area government debt securities on a quarterly basis, according to balance of payments data. Importantly, there are significant differences in the reaction to excess returns on an unhedged or hedged basis. Currencies such as the Japanese yen, that offer lower returns on an unhedged basis, seem capable of increasing their pull for US-based institutions by offering relatively higher returns on a currency hedged basis. These results reveal that US mutual funds do exploit the advantage conferred by their role of liquidity providers in the market for forward dollars, where mismatches in hedging flows combined with balance sheet frictions of intermediaries open up CIP deviations. In this respect, the discrepancies between the reactions to unhedged and hedged returns appear important and merit further research.

There are additional results that are important for portfolio choice theory and the demand for safe assets. We find evidence of strong frictions in portfolio adjustment, with lagged shares displaying positive and significant coefficients in the portfolio regressions, probably due to fund-specific targets and benchmarking. With a few exceptions, valuation effects do not seem to influence the currency shares, as fund managers offset through active rebalancing the impact of changes in bond prices and exchange rates on the various currency shares. When global financial risk is on the rise, there is evidence that US mutual fund investors repatriate their investments towards US government debt securities, mainly at the expenses of euro-denominated ones. Finally, in periods of low US policy rates, the sensitivity of sovereign portfolio shares to excess returns is more elevated than in other periods, indicating that the mechanism of searching for yield abroad in a low-interest-rate environment highlighted in previous studies is not unique to corporate bonds and equities. However, when US policy rates are high, they increase their exposure to safe-haven currencies like the Japanese yen and the Swiss franc. Overall, our results have significant implications for capital flows from the United States towards other major currency areas, as well as for the impact of the failures of arbitrage conditions, such as the CIP, on the incentives of institutional investors.

Table 3.7. Policy rate unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.78***	0.74***	0.71***	0.71***	0.72***	0.64***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)
$r^{ex,j,unh}_{i,t}$	0.57*	0.50	-0.14**	1.33*	0.31	1.37***
	(0.33)	(0.44)	(0.07)	(0.78)	(0.50)	(0.37)
cb^{US}_t	-0.42	0.28	0.13*	-0.40	2.41***	-1.22**
	(0.40)	(0.25)	(0.08)	(0.36)	(0.80)	(0.52)
$cb^{US}_t \times r^{ex,j,unh}_{i,t}$	0.00	-0.00	0.00***	-0.01	0.01*	-0.01***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta s^{j,P,R}_{i,t}$	0.67	-1.08	1.09	0.82	0.51	0.20
	(4.43)	(5.46)	(3.79)	(2.57)	(1.66)	(1.52)
$\Delta s^{j,P,XR}_{i,t}$	0.12	1.03	2.66	-0.24	0.60*	0.82
	(1.02)	(1.31)	(1.64)	(0.84)	(0.33)	(0.50)
Within R^2	0.68	0.61	0.53	0.51	0.56	0.47
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 cb^{US}_t + \beta^j_4 cb^{US}_t \times r^{ex,j,unh}_{i,t} + \beta^j_5 \Delta s^{j,P,R}_{i,t} + \beta^j_6 \Delta s^{j,P,XR}_{i,t} + \boldsymbol{\beta^j_7 W_t} + \varepsilon^j_{i,t}$.
Each column reports results for a different currency j. $\boldsymbol{W_t}$ includes year-on-year inflation for country j in quarter t in percentage points, the Citi Economics Surprise Index for country j in quarter t in standard deviation units, and the VIX in quarter t in standard deviation units. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.8. Policy rate hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.73***	0.70***	0.72***	0.71***	0.64***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	0.96	2.95***	0.08	-4.42**	0.36	-1.14
	(0.61)	(1.04)	(0.17)	(1.85)	(2.31)	(1.73)
cb^{US}_t	-0.84**	0.36*	0.11*	-0.28	0.53	-0.82
	(0.37)	(0.20)	(0.06)	(0.41)	(0.91)	(0.51)
$cb^{US}_t \times r^{ex,j,fwd}_{i,t}$	-0.01	0.01	-0.01***	0.01	0.02	0.03**
	(0.01)	(0.01)	(0.00)	(0.01)	(0.03)	(0.01)
$\Delta s^{j,P,R}_{i,t}$	0.59	-0.69	0.94	0.38	0.33	0.20
	(4.21)	(5.28)	(3.76)	(2.61)	(1.65)	(1.49)
$\Delta s^{j,P,XR}_{i,t}$	0.14	0.90	2.65	-0.22	0.59*	0.83*
	(0.98)	(1.31)	(1.62)	(0.84)	(0.32)	(0.48)
Within R^2	0.68	0.61	0.53	0.51	0.56	0.47
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 cb^{US}_t + \beta^j_4 cb^{US}_t \times r^{ex,j,fwd}_{i,t} + \beta^j_5 \Delta s^{j,P,R}_{i,t} + \beta^j_6 \Delta s^{j,P,XR}_{i,t} + \boldsymbol{\beta^j_7 W_t} + \varepsilon^j_{i,t}$.
Each column reports results for a different currency j. $\boldsymbol{W_t}$ includes year-on-year inflation for country j in quarter t in percentage points, the Citi Economics Surprise Index for country j in quarter t in standard deviation units, and the VIX in quarter t in standard deviation units. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Appendix

3.A Data sources

Table 3.9. Data sources

Data	**Source**
Assets under management, US mutual fund industry	Federal Reserve Financial Accounts of the United States
Fund-level data on US fixed income funds	Refinitiv Lipper for Invsetment Managament and Lipper Global Data feed
Government bond yields	Refinitiv Eikon
Government bond indices	Refinitiv Eikon
Spot and forward exchange rates	Refinitiv Eikon
Amount of government debt outstanding	Bank for International Settlements
CPI inflation	Federal Reserve Economic Data (FRED)
Citigroup Economic Surprise Index	Haver Analytics
Federal Reserve policy rate	Bank for International Settlements

3.B Further descriptive statistics

Table 3.10. Whole sample summary statistics

	N	Mean	SD	Min	P5	P95	Max
A. Fund characteristics							
Assets under management ($Mil.)	15,438	2,382	8,656	0.08	12.84	8,977	179,914
Total sovereign holdings ($Mil.)	15,700	422	1,754	0.00	1.03	1,822	46,539
Selected sovereign holdings ($Mil.)	15,700	318	1,499	0.00	0.62	1,273	43,175
Reporting quarters	15,700	34	14	1	8	48	48
B. Selected sovereign portfolio shares							
United States	15,700	0.82	0.37	0	0	1.00	1
Euro Area safe	15,700	0.05	0.16	0	0	0.37	1
Australia	15,700	0.03	0.13	0	0	0.12	1
Canada	15,700	0.01	0.08	0	0	0.07	1
Japan	15,700	0.05	0.15	0	0	0.44	1
Switzerland	15,700	0.00	0.01	0	0	0.00	0.56
United Kingdom	15,700	0.04	0.12	0	0	0.24	1

Statistics are calculated on the whole sample of funds. Highly-rated sovereign holdings include government bonds of all maturities issued by Australia, Canada, euro area highly-rated issuers (Austria, Belgium, Germany, France, the Netherlands), Japan, Switzerland, United Kingdom and United States. Shares refer to the portfolio that comprises all debt securities issue by the selected highly-rated sovereigns. Source: Refinitiv Lipper.

Table 3.11. Fund-level price passive reallocation

	N	Mean	SD	Min	P5	P25	P50	P75	P95	Max
$\Delta s_{i,t}^{AUD,P,R}$	3,442	-0.00	0.12	-2.01	-0.11	0.00	0.00	0.00	0.13	1.08
$\Delta s_{i,t}^{CAD,P,R}$	3,333	0.00	0.10	-2.27	-0.08	-0.00	0.00	0.00	0.10	0.84
$\Delta s_{i,t}^{CHF,P,R}$	692	-0.00	0.04	-0.71	-0.03	-0.00	0.00	0.00	0.01	0.15
$\Delta s_{i,t}^{EUR_S,P,R}$	3,668	-0.02	0.17	-2.06	-0.29	-0.04	0.00	0.01	0.22	0.89
$\Delta s_{i,t}^{JPY,P,R}$	3,425	-0.00	0.21	-1.87	-0.36	-0.02	0.00	0.05	0.34	1.12
$\Delta s_{i,t}^{GBP,P,R}$	3,442	0.01	0.18	-1.03	-0.28	-0.05	0.00	0.03	0.32	1.17
$\Delta s_{i,t}^{USD,P,R}$	4,359	0.02	0.25	-1.17	-0.31	-0.00	0.00	0.00	0.41	2.27

Fund-level passive reallocation due to issue-currency bond returns. All variables in percentage points. Summary statistics are calculated over the distribution of fund-quarter observations in the sample excluding funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries.

Table 3.12. Fund-level exchange rate passive reallocation

	N	Mean	SD	Min	P5	P25	P50	P75	P95	Max
$\Delta s_{i,t}^{AUD,P,XR}$	3,442	-0.01	0.37	-3.21	-0.43	0.00	0.00	0.00	0.42	2.93
$\Delta s_{i,t}^{CAD,P,XR}$	3,333	0.00	0.23	-2.79	-0.26	0.00	0.00	0.01	0.29	2.36
$\Delta s_{i,t}^{CHF,P,XR}$	692	0.01	0.13	-0.67	-0.05	0.00	0.00	0.00	0.08	2.79
$\Delta s_{i,t}^{EUR_S,P,XR}$	3,668	-0.01	0.52	-3.82	-0.78	-0.08	0.00	0.07	0.84	3.23
$\Delta s_{i,t}^{JPY,P,XR}$	3,425	-0.04	0.76	-3.23	-1.34	-0.16	0.00	0.02	1.20	3.82
$\Delta s_{i,t}^{GBP,P,XR}$	3,442	0.02	0.45	-2.86	-0.69	-0.04	0.00	0.12	0.81	2.07
$\Delta s_{i,t}^{USD,P,XR}$	4,359	0.02	0.36	-2.73	-0.44	0.00	0.00	0.00	0.54	3.13

Fund-level passive reallocation due to exchange rate effects. All variables in percentage points. Summary statistics are calculated over the distribution of fund-quarter observations in the sample excluding funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries.

3.C Predictability of excess returns

Current excess returns provide insights on the relative yield that an investor can expect from government bonds in the same quarter. However, if mutual fund managers are forward-looking, current excess returns are more informative for their portfolio choice to the extent that they can forecast *future* excess returns. To test for predictability, we compute autocorrelations of cross-sectional average excess returns for each currency $\bar{r}_t^j = \frac{1}{I}\sum_{i=1}^{I} r_{i,t}^{ex,j}$. Unhedged returns, pictured in Figure 3.7 display a positive and high autocorrelation, significant at the 95% level for up to four quarters for all currencies. Therefore, a fund that observes a positive excess return from currency j in quarter t can expect a positive extra return for up to one year thereafter, a useful signal for rebalancing toward that currency. This is in line with a long-standing literature documenting the predictability in currency excess returns (e.g. Fama (1984)) and the profitability of carry trade strategies that go long high-interest rate currencies (e.g. Lustig and Verdelhan (2007)). In addition, our findings show forecastability of excess returns not only with respect to individual currencies, but even when weighted by fund-specific portfolio shares, so that individual investors do have incentives to exploit this margin for higher returns.

The picture is strikingly different for hedged returns, shown in Figure 3.8. Autocorrelations are only significantly different from zero at a one-quarter horizon for the euro and yen, and at a two-quarter horizon for Swiss franc. Even when autocorrelations are significant, they are much lower than for unhedged returns, never exceeding 0.6.

Overall, wedges for the UIP condition are larger than for CIP, as already evident in Table 3.3.[23] The marked difference in predictability between these deviations suggests that funds can glean more information from contemporaneous unhedged returns, and so possibly have more incentives to respond to them compared to hedged returns.

3.D Fixed versus random effects

The baseline model in Section 3.3 uses fund fixed effects, but a random effects estimator would be more efficent if fund-specific unobservables α_i were uncorrelated with the vector of explanatory variables $\boldsymbol{X}_{i,t}$. In this appendix, we test the null hypothesis $Cov[\alpha_i, \boldsymbol{X}_{i,t}] = 0$ via a Hausman (1978) test.

Table 3.13. Hausman (1978) test for random vs. fixed effects

	AUD	CAD	CHF	GBP	JPY	EUR	USD
P-value unhedged returns	0.00	0.00	0.00	0.00	0.00	0.00	0.00
P-value hedged returns	0.00	0.00	0.00	0.00	0.00	0.00	

P-values of Hausman (1978) test, obtained from comparing estimates from the fixed effects and random effects estimation of the baseline model $s^j_{i,t} = \alpha_i + \gamma_t + \beta_1 s^j_{i,t-1} + \beta_2 r^{ex,j}_{i,t} + \beta_3 \Delta s^{j,P,R}_{i,t} + \beta_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon_{i,t}$, with $r^{ex,j}_{i,t} = r^{j,unh}_{i,t}$ or $r^{j,fwd}_{i,t}$.

Table 3.13 reports the p-values of the test for the baseline unhedged and hedged returns models for each currency. The null hypothesis is rejected at the 1% significance level for all currencies and both models, suggesting that fund-specific unobservables are indeed correlated with explanatory variables. Therefore, the choice of a fixed-effects estimator in the baseline model is appropriate.

[23]This is a robust result in the literature. In a recent example, Bacchetta et al. (2023) report dollar CIP deviations with respect to the Japanese yen and the Swiss franc of less than 1%, while UIP deviations reach up to 10%.

3.E Robustness checks

3.E.1 Pooled OLS

In this appendix, we present results from running the baseline regressions of Section 3.3 with Pooled Ordinary Least Squares. Formally, the estimating equation is

$$s_{i,t}^{j} = \alpha + \eta^{j}\mathbb{1}_{i,t}^{Out,j} + \beta_1^j s_{i,t-1}^{j} + \beta_2^j r_{i,t}^{ex,j} + \beta_3^j \Delta s_{i,t}^{j,P,R} + \beta_4^j \Delta s_{i,t}^{j,P,XR} + \varepsilon_{i,t}^{j}. \tag{3.8}$$

Table 3.14. Pooled OLS unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s_{i,t-1}^{j}$	0.86***	0.81***	0.93***	0.80***	0.87***	0.76***	0.92***
	(0.02)	(0.02)	(0.05)	(0.03)	(0.03)	(0.02)	(0.01)
$r_{i,t}^{ex,j,unh}$	-0.01	0.40***	-0.03	0.87***	0.54	1.25***	-0.45
	(0.26)	(0.14)	(0.05)	(0.25)	(0.34)	(0.34)	(0.44)
$\Delta s_{i,t}^{j,P,R}$	0.05	-0.63	0.28	0.05	1.55	1.11	-1.13
	(5.67)	(5.41)	(3.69)	(2.43)	(1.72)	(1.55)	(1.56)
$\Delta s_{i,t}^{j,P,XR}$	0.62	0.72	2.76	-0.33	0.99**	1.25**	1.10
	(1.20)	(1.40)	(1.74)	(0.82)	(0.44)	(0.61)	(1.35)
R^2	0.78	0.70	0.97	0.66	0.76	0.64	0.85
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668	4,359
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	41	42	43	41	42	41	40

Coefficients from regression model $s_{i,t}^{j} = \alpha + \eta^{j}\mathbb{1}_{i,t}^{Out,j} + \beta_1^j s_{i,t-1}^{j} + \beta_2^j r_{i,t}^{ex,j,unh} + \beta_3^j \Delta s_{i,t}^{j,P,R} + \beta_4^j \Delta s_{i,t}^{j,P,XR} + \varepsilon_{i,t}^{j}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

These models serve as a benchmark for our identification strategy. By comparing the results in Tables 3.14 and 3.15 with the time and fund fixed effects models in Section 3.3, we can observe that the coefficients on excess returns for some currency, like the US dollar in the unhedged return model, are attenuated towards zero and/or lose significance. This is consistent with fixed effects removing the bias towards zero due to demand shocks.

Table 3.15. Pooled OLS hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.85***	0.80***	0.93***	0.81***	0.87***	0.78***
	(0.02)	(0.02)	(0.05)	(0.03)	(0.03)	(0.02)
$r^{ex,j,fwd}_{i,t}$	0.49	3.28***	0.05	-3.69***	-1.29	-2.23
	(0.70)	(0.92)	(0.18)	(1.25)	(1.40)	(1.75)
$\Delta s^{j,P,R}_{i,t}$	0.04	-0.03	0.27	-0.29	1.41	1.14
	(5.60)	(5.13)	(3.68)	(2.52)	(1.81)	(1.53)
$\Delta s^{j,P,XR}_{i,t}$	0.57	0.60	2.76	-0.33	1.00**	1.24**
	(1.16)	(1.39)	(1.74)	(0.81)	(0.45)	(0.57)
R^2	0.78	0.70	0.97	0.66	0.76	0.63
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model $s^j_{i,t} = \alpha + \beta^j_1 s^j_{i,t-1} + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

3.E.2 Whole sample

In this sub-section, we present results for the baseline unhedged and hedged returns regressions estimated for the whole sample, without applying the selection procedure outlined in Section 3.2.1.

The complete sample consists of all 880 funds, including all currency-specific funds whose portfolio shares change barely or not at all over time. While the presence of such funds is liable to bias the analysis substantially , the results presented here are nonetheless useful as a benchmark for the effects of our sample selection strategy.

Table 3.16 reports results for unhedged returns, and Table 3.17 for hedged returns. In both cases, the magnitude and significance of coefficients is rather different from the findings of the models estimated on the restricted sample that excludes funds with a country focus and presented in Section 3.3. The smaller and less significant coefficients on unhedged excess returns for most currencies in the whole sample regressions suggest that the presence of funds focused on single countries does introduce some bias towards zero. On the other hand, results from hedged return models are broadly consistent with the baseline. One exception is the coefficient for the euro, which turns weakly significant in the whole sample.

Table 3.16. Whole sample unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.78***	0.73***	0.65***	0.70***	0.71***	0.63***	0.73***
	(0.03)	(0.03)	(0.13)	(0.04)	(0.05)	(0.03)	(0.03)
$r^{ex,j,unh}_{i,t}$	0.07	0.30*	-0.00	0.33	1.02**	0.24	1.71*
	(0.25)	(0.15)	(0.00)	(0.27)	(0.42)	(0.20)	(0.94)
$\Delta s^{j,P,R}_{i,t}$	-0.52	-0.54	1.23	-0.73	2.11	1.36	-1.14
	(4.32)	(4.68)	(3.95)	(2.44)	(1.35)	(1.69)	(1.41)
$\Delta s^{j,P,XR}_{i,t}$	0.20	0.64	2.65	-0.24	0.95**	1.14***	0.85
	(1.00)	(1.14)	(1.61)	(0.87)	(0.38)	(0.38)	(1.12)
Within R^2	0.67	0.59	0.54	0.51	0.55	0.46	0.58
N (fund-quarter)	13,611	13,611	13,611	13,611	13,611	13,611	13,611
N (funds)	880	880	880	880	880	880	880
Avg. nr. quarters	34	34	34	34	34	34	34

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

3.E.3 Alternative sample and outlier selection

Our baseline sample excludes funds with a country focus, defined as having an average portfolio share of at least 95% for any currency. Furthermore, we define as outliers fund-quarter observations in the top and bottom 1% by active portfolio share reallocation. We include a dummy for these observations to absorb their impact on our estimates.

Table 3.17. Whole sample hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.72***	0.65***	0.71***	0.71***	0.63***
	(0.03)	(0.03)	(0.13)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-1.06	2.97***	0.01	-3.36**	3.19**	2.16*
	(1.34)	(0.77)	(0.02)	(1.54)	(1.29)	(1.16)
$\Delta s^{j,P,R}_{i,t}$	-0.49	-0.44	1.23	-0.96	1.95	1.30
	(4.32)	(4.55)	(3.95)	(2.50)	(1.36)	(1.70)
$\Delta s^{j,P,XR}_{i,t}$	0.22	0.47	2.65	-0.15	1.04**	1.08***
	(0.98)	(1.10)	(1.61)	(0.84)	(0.42)	(0.38)
Within R^2	0.67	0.59	0.54	0.52	0.55	0.46
N (fund-quarter)	13,611	13,611	13,611	13,611	13,611	13,611
N (funds)	880	880	880	880	880	880
Avg. nr. quarters	34	34	34	34	34	34

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \eta^j \mathbb{1}^{Out,j}_{i,t} + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

While we believe these thresholds capture country-focused funds and outliers satisfactorily, they are inevitably discretionary. Therefore, in this section we test the sensitivity of our results to more stringent crieria for both thresholds. We now define country-focused funds as having an average portfolio share of 90% or higher for any currency. We also tag as outliers fund-quarter observations in the top and bottom 2.5% by active portfolio share reallocation.

Tables 3.18 and 3.19 display the results of unhedged and hedged return models using this alternative selection procedure. The estimates are almost identical in terms of both size and statistical significance. One exception is the elasticity to US dollar excess returns, which loses significance with respect to the baseline model. We are then reassured that our results do not depend on the exact thresholds for sample and outlier selection, and can survive more exacting criteria for both.

Table 3.18. Alternative sample selection unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.77***	0.75***	0.68***	0.69***	0.72***	0.63***	0.74***
	(0.03)	(0.03)	(0.13)	(0.04)	(0.04)	(0.03)	(0.03)
$r^{ex,j,unh}_{i,t}$	1.20*	0.44	0.09	1.01*	0.30	1.07***	1.55
	(0.65)	(0.29)	(0.05)	(0.56)	(0.67)	(0.29)	(0.95)
$\Delta s^{j,P,R}_{i,t}$	0.15	-0.89	1.01	-1.93	0.02	-0.74	-3.16
	(4.79)	(5.67)	(4.29)	(2.76)	(2.40)	(1.77)	(2.22)
$\Delta s^{j,P,XR}_{i,t}$	0.10	0.44	2.73	0.07	0.87**	1.34***	0.03
	(1.09)	(1.27)	(1.80)	(1.13)	(0.39)	(0.46)	(1.26)
Within R^2	0.68	0.62	0.56	0.54	0.59	0.48	0.62
N (fund-quarter)	3,264	3,112	692	3,329	3,202	3,448	3,996
N (funds)	168	168	168	168	168	168	168
Avg. nr. quarters	40	40	40	40	40	40	40

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.90$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries.Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

3.E.4 Correcting for survivorship bias

Our dataset comprises funds that drop out of the sample due to liquidation or merging.At the same time, new funds enter the sample during the period we analyse, either because they are newly created or because they start reporting to Lipper. Furthermore, some funds abandon the sample temporarily. For example, the fund "John Hancock Government Income Fund" first joins in the sample in 2012 Q1, and it is observed until 2013 Q4, when it drops out to re-appear in 2016 Q1. The consequences are twofold: the panel is unbalanced, and the sample only includes funds that are active at any given quarter, potentially leading to survivorship bias.

Survivorship bias affects preeminently models of the performance of funds (Elton et al. (1996)), so in principle it should not be particularly severe for our study given that we examine only portfolio shares.

Table 3.19. Alternative sample selection hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.74***	0.68***	0.70***	0.71***	0.63***
	(0.03)	(0.03)	(0.12)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-1.25	3.74***	-0.29	-3.19	3.66**	1.80
	(1.99)	(1.06)	(0.41)	(1.93)	(1.70)	(1.66)
$\Delta s^{j,P,R}_{i,t}$	-0.03	-0.74	0.97	-2.10	-0.05	-0.77
	(4.72)	(5.51)	(4.27)	(2.87)	(2.37)	(1.77)
$\Delta s^{j,P,XR}_{i,t}$	0.17	0.30	2.74	0.04	0.90**	1.25**
	(1.05)	(1.25)	(1.79)	(1.10)	(0.41)	(0.50)
Within R^2	0.68	0.63	0.56	0.54	0.59	0.48
N (fund-quarter)	3,264	3,112	692	3,329	3,202	3,448
N (funds)	168	168	168	168	168	168
Avg. nr. quarters	40	40	40	40	40	40

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Nevertheless, to alleviate any concerns of results being affected by changes in sample composition over time, we repeat our descriptive and econometric analysis using only funds in the restricted sample, excluding funds with a country focus, that are present in the sample for at least five consecutive years. There are 112 funds that meet this condition out of 186 in this sample that we use in the baseline models. Ideally, we would only include funds that are present throughout the sample period, but only 32 funds would meet such a strict requirement, preventing us from relying on large-sample asymptotics in our panel regressions.

We first compare the size and portfolio shares of the sample excluding funds with a country focus with long-permanence funds to ascertain whether the latter differ in any meaningful way. Table 3.20 shows that the funds in the long-permanence sample are broadly similar to those in the baseline sample, in terms of both size

Table 3.20. Long-permanence sample summary statistics

	N	Mean	SD	Min	P5	P95	Max
A. Fund characteristics							
Assets under management ($Mil.)	3,826	2,905	11,670	0.10	14.40	11,937	179,914
Total sovereign holdings ($Mil.)	3,888	523	1,345	0.10	3.38	2,458	22,698
Selected sovereign holdings ($Mil.)	3,888	300	952	0.02	0.89	1,406	21,116
Reporting quarters	3,888	44	6	21	29	48	48
B. Selected sovereign portfolio shares							
United States	3,888	0.43	0.44	0	0	1.00	1
Euro Area safe	3,888	0.15	0.22	0	0	0.61	1
Australia	3,888	0.08	0.21	0	0	0.60	1
Canada	3,888	0.05	0.13	0	0	0.24	1
Japan	3,888	0.16	0.23	0	0	0.61	1
Switzerland	3,888	0.00	0.01	0	0	0.01	0.08
United Kingdom	3,888	0.12	0.20	0	0	0.54	1

Highly-rated sovereign holdings include government bonds of all maturities issued by Australia, Canada, euro area highly-rated issuers (Austria, Belgium, Germany, France, the Netherlands), Japan, Switzerland, United Kingdom and United States. Shares refer to the portfolio that comprises all debt securities issue by the selected highly-rated sovereigns. The sample includes only funds that are observed for at least five consecutive years and excludes all funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Source: Refinitiv Lipper.

and sovereign portfolio shares. As expected, the only large difference is the higher average of 44 reporting quarters, edging closer to the maximum of 48. These descriptive statistics already suggest that survivorship bias is unlikely to skew our results, but we seek to confirm it by running the baseline models on the long-permanence sample.

Tables 3.21 and 3.22 report results for the baseline regressions run on the sample of long-permanence funds. Overall, the results are similar, with the exception of the loss of significance in the excess return coefficient for JPY in the hedged return model, and the presence of very high coefficients on valuation effects for Switzerland, probably due to the drastic reduction in the number of funds which makes large-I asymptotics unreliable.

3.E.5 Global funds

The sample selection strategy used in the main body of the paper aims at isolating funds with an international investment horizon, and it produces aggregate portfolio shares that are consistent with this goal. However, it is based on the arbitrary threshold of a 95% average portfolio share. To test whether our core results survive a more systematic sample selection strategy, in this appendix we offer an alternative procedure that identifies global funds based on the distance from an

Table 3.21. Long permanence funds unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.79***	0.78***	0.56***	0.72***	0.75***	0.62***	0.75***
	(0.03)	(0.04)	(0.15)	(0.04)	(0.04)	(0.03)	(0.02)
$r^{ex,j,unh}_{i,t}$	0.58	0.58**	0.14**	1.14**	0.11	0.81**	1.68
	(0.58)	(0.24)	(0.06)	(0.53)	(0.52)	(0.31)	(1.03)
$\Delta s^{j,P,R}_{i,t}$	-0.71	-0.39	-6.89	-1.03	0.12	-1.21	-3.28
	(5.16)	(6.74)	(9.46)	(2.99)	(1.81)	(2.01)	(2.03)
$\Delta s^{j,P,XR}_{i,t}$	0.11	1.05	0.59	0.73	0.64	1.79***	0.37
	(0.99)	(1.57)	(2.30)	(0.93)	(0.43)	(0.47)	(1.10)
Within R^2	0.70	0.67	0.51	0.57	0.62	0.48	0.62
N (fund-quarter)	2,852	2,870	561	2,890	2,931	3,098	3,583
N (funds)	84	79	13	80	82	87	111
Avg. nr. quarters	44	45	46	44	44	45	44

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j, as well as funds that have less than five consecutive years of observations. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

International Capital Asset Pricing Model (ICAPM henceforth) benchmark.

The canonical ICAPM model predicts that a mean-variance investor would choose shares in a portfolio of non currency-hedged international bonds equal to the relative market capitalization of each asset (Solnik (1974)). We calculate the ICAPM portfolio weights as the relative market capitalization of each country in the selected sovereign portfolio. Formally,

$$w^{j,CAPM}_t = \frac{B^j_t}{\sum_k B^k_t} \tag{3.9}$$

Where B^j_t is the amount of country j's central government debt outstanding in quarter t.

For each fund, we then calculate the Euclidean distance $d_{i,t}$ between the vector of

Table 3.22. Long permanence funds hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.80***	0.77***	0.56***	0.74***	0.75***	0.63***
	(0.03)	(0.04)	(0.15)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-0.39	3.02***	-0.37	-1.92	1.45	-0.80
	(1.73)	(0.99)	(0.37)	(2.35)	(2.50)	(1.80)
$\Delta s^{j,P,R}_{i,t}$	-0.75	-0.27	-7.07	-1.05	0.10	-1.30
	(5.14)	(6.60)	(9.46)	(2.99)	(1.80)	(2.00)
$\Delta s^{j,P,XR}_{i,t}$	0.14	0.93	0.61	0.68	0.65	1.82***
	(0.97)	(1.54)	(2.29)	(0.90)	(0.43)	(0.48)
Within R^2	0.70	0.67	0.51	0.57	0.62	0.48
N (fund-quarter)	2,852	2,870	561	2,890	2,931	3,098
N (funds)	84	79	13	80	82	87
Avg. nr. quarters	44	45	46	44	44	45

Coefficients from regression model $s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j, as well as funds that have less than five consecutive years of observations. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

average portfolio shares and the vector of relative market capitalizations as

$$d_{i,t} = \sqrt{\sum_j \left(s^j_{i,t} - w^{j,CAPM}_t\right)^2} \tag{3.10}$$

For our empirical analysis, we keep the funds that are in the bottom 50% in the distribution of time-series average Euclidean distance $\bar{d}_i$. This criterion is meant to capture the set of funds that are closest to the theoretical benchmark of a fully-diversified ICAPM portfolio.

Figure 3.9 compares the theoretical ICAPM wieghts, on the left-hand side, with the aggregate portfolio shares in the low $\bar{d}_i$ sub-sample, on the right-hand side. The aggregate shares are volatile but close to the ICAPM benchmark, save for a home bias, especially in the latest parts of the sample. Overall, we are satisfied that this procedure yields a sub-sample of funds with a global investment horizon.

Table 3.23. Global funds unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.67***	0.58***	0.66***	0.69***	0.72***	0.60***	0.68***
	(0.05)	(0.06)	(0.08)	(0.05)	(0.05)	(0.04)	(0.04)
$r^{ex,j,unh}_{i,t}$	-0.34	0.20	0.02***	-0.75**	1.75***	-0.07	0.64
	(0.32)	(0.24)	(0.01)	(0.37)	(0.43)	(0.34)	(0.88)
$\Delta s^{j,P,R}_{i,t}$	3.99	-1.71	-0.85	-1.86	2.98	-2.55	-1.62
	(4.14)	(4.43)	(2.70)	(1.64)	(1.94)	(1.96)	(1.36)
$\Delta s^{j,P,XR}_{i,t}$	-1.16	0.49	-0.18	-0.33	1.03**	1.49***	1.06
	(1.20)	(1.32)	(1.11)	(0.69)	(0.39)	(0.42)	(0.91)
Within R^2	0.58	0.44	0.58	0.56	0.60	0.48	0.51
N (fund-quarter)	4,985	4,985	4,985	4,985	4,985	4,985	4,985
N (funds)	443	443	443	443	443	443	443
Avg. nr. quarters	32	32	32	32	32	32	32

Coefficients from regression model $s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model only includes funds in the bottom 50% of the average ICAPM Euclidean distance $\bar{d}_i$ distribution. Each model excludes funds for which the average portfolio share of currency j is equal to 0. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Tables 3.23 and 3.24 present the results of running the baseline time and fund fixed effects models on the sub-sample of global funds identified through ICAPM distance. Interestingly, the results for are rather different from the baseline. Only the Swiss franc and Japanese yen display a positive and significant coefficient in the unhedged returns model, with Japan the only currency with a positive reaction to excess returns on both an unhedged and a hedged basis.

3.F Autocorrelation of portfolio shares

The autocorrelation of portfolio shares is a strikingly robust result of our analysis, with displaying highly significant coefficients on $s^j_{i,t-1}$ across all currencies

Table 3.24. Global funds hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.70***	0.58***	0.66***	0.69***	0.72***	0.60***
	(0.05)	(0.06)	(0.08)	(0.05)	(0.05)	(0.05)
$r^{ex,j,fwd}_{i,t}$	-7.54**	0.28	0.03	-0.98	4.38**	-1.59
	(3.56)	(0.85)	(0.05)	(1.58)	(1.76)	(1.67)
$\Delta s^{j,P,R}_{i,t}$	3.86	-1.72	-0.86	-2.07	2.65	-2.54
	(4.14)	(4.43)	(2.72)	(1.56)	(1.86)	(1.95)
$\Delta s^{j,P,XR}_{i,t}$	-0.93	0.48	-0.18	-0.23	1.16***	1.54***
	(1.12)	(1.34)	(1.12)	(0.72)	(0.43)	(0.43)
Within R^2	0.59	0.44	0.58	0.56	0.60	0.48
N (fund-quarter)	4,985	4,985	4,985	4,985	4,985	4,985
N (funds)	443	443	443	443	443	443
Avg. nr. quarters	32	32	32	32	32	32

Coefficients from regression model $s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model only includes funds in the bottom 50% of the average ICAPM Euclidean distance $\bar{d}_i$ distribution. Each model excludes funds for which the average portfolio share of currency j is equal to 0. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

and specifications. While the estimated coefficients are always below one, it is important to investigate the possibility of unit roots in portfolio shares that would jeopardise inference in our models. In this appendix, we perform unit root tests and augment our baseline models with longer lags of $s^j_{i,t}$ to explore the higher-order autocorrelation of portfolio shares.

We perform a Fisher-type panel unit root test that allows for unbalanced panels with gaps. The test is based on performing Dickey and Fuller (1979) unit-root tests on each panel and combining the resulting p-values to test the null hypothesis that all panels exhibit a unit root.

Table 3.25 reports the p-values for the Fisher test performed on each currency's

Table 3.25. Portfolio shares unit root tests

	AUD	CAD	CHF	GBP	JPY	EUR	USD
P-value Choi (2001) statistic	0.00	0.00	0.00	0.00	0.00	0.00	0.00

P-values from Fisher-type test using the Choi (2001) modified inverse χ^2 transformation that is consistent for $I \to \infty$. The individual Dickey and Fuller (1979) regressions for each panel i are
$\Delta \tilde{s}^j_{i,t} = \alpha^j + \beta^j \tilde{s}^j_{i,t-1} + \xi^j \delta \tilde{s}^j_{i,t-1} + \varepsilon^j_{i,t}$, where $\tilde{s}^j_{i,t} = s^j_{i,t} - \bar{s}^j_t$

portfolio share, showing that we can reject at the 1% level of significance the null hypothesis that all panels display a unit root.

We then augment the baseline regressions for unhedged and hedged returns with four lags of the portfolio share, allowing us to investigate autocorrelation up to one year ahead. Tables 3.26 and 3.27 present the results of these regressions, broadly confirming the results of the baseline models for what concerns excess returns and valuation effects for hedged returns. The addition of further portfolio share lags does not affect the size or significance of the coefficient on $s^j_{i,t-1}$, and none of the further lag coeffficients are statistically significant, save for weak significance on for the Japanese Yen and the Euro. On the other hand, they seem to add noise to the model, turning most coefficients on excess returns statistically insignificant. Overall, these results point at a very weak autocorrelation beyond the first quarter, implying that portfolio stickiness is short-lived and portfolio shares are rather flexible.

3.G Lagged excess returns

Figures 3.7 shows evidence of autocorrelation of fund-specific unhedged excess returns for all currencies for up to two quarters, and 3.8 displays predictability at a one-quarter horizon for the Euro, the Japanese Yen and the Swiss Franc. Therefore, it is natural to ask whether funds use lagged values as a predictor of future excess returns that drives active portfolio reallocation.

To answer this question, we run the baseline models augmening the unhedged and hedged excess returns models with their first lag $r^{ex,j}_{i,t-1}$. Formally, the estimating equation is

$$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j}_{i,t} + \beta^j_3 r^{ex,j}_{i,t-1} + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}, \tag{3.11}$$

Tables 3.28 and 3.29 report the results, showing that lagged excess returns are not associated with active reallocation on an unhedged or hedged basis for any currency. These results are consistent with the weaker predictability of excess returns at longer horizons.

3.H Joint estimation

The dependent variables in our econometric analysis are the shares in a portfolio of sovereign bonds, $s^j_{i,t}$. In the main body of the paper, we run portfolio regressions separately on each individual currency share, but the shares sum up to 1 across currencies by construction. In this appendix, we exploit the cross-equation restriction $\sum_j s^j_{i,t} = 1$ to discipline the fitted values. We implement the summing-up constraint by estimating the equations jointly, following Bubeck et al. (2018).

3.H.1 Methodology

Let $\boldsymbol{X}^j_{i,t} = (\alpha_i, \gamma_t, \mathbb{1}^{Out,j}_{i,t}, s^j_{i,t-1}, r^{ex,j}_{i,t}, \Delta s^{j,P,R}_{i,t}, \Delta s^{j,P,XR}_{i,t})$ the $K = 7$-dimensional vector fund- and time-specific intercepts and explanatory variables in the baseline regression. We can rewrite the model for each currency j compactly by stacking the observations over the T quarters and I funds:

$$\boldsymbol{s}_j = \boldsymbol{\beta}'_j \boldsymbol{X}_j + \boldsymbol{\varepsilon}_j \tag{3.12}$$

Where $\boldsymbol{s}_j = (s^j_{1,1}, \dots s^j_{I,T})$ is the vector of currency j portfolio shares, $\boldsymbol{\beta}^j = (1, 1, \eta^j_1, \beta^j_1, \beta^j_2, \beta^j_3, \beta^j_4)$ is the vector of coefficients and constants, $\boldsymbol{X}_j = (\boldsymbol{X}^j_{1,1}, \dots \boldsymbol{X}^j_{I,T})$ is the vector of explanatory variables in the baseline portfolio regression, and $\boldsymbol{\varepsilon}_j = (\varepsilon^j_{1,1}, \dots \varepsilon^j_{I,T})$ is the vector of residuals, all stacked over the time and fund dimensions.

The $J = 6$ portfolio regressions can be estimated jointly as $\boldsymbol{s} = \boldsymbol{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon}$, without a summing-up constraint, by stacking them using a block diagonal matrix $\boldsymbol{X}$ of dimension $JL \times LK$, where L is the number of fund-quarter observations. [24]

$$\begin{bmatrix} \boldsymbol{s}_1 \\ \boldsymbol{s}_2 \\ \vdots \\ \boldsymbol{s}_J \end{bmatrix} = \begin{bmatrix} \boldsymbol{X}_1 & \boldsymbol{0}_L & \dots & \boldsymbol{0}_L \\ \boldsymbol{0}_L & \boldsymbol{X}_2 & \dots & \boldsymbol{0}_L \\ \vdots & \vdots & \ddots & \vdots \\ \boldsymbol{0}_L & \boldsymbol{0}_L & \dots & \boldsymbol{X}_J \end{bmatrix} = \begin{bmatrix} \boldsymbol{\beta}_1 \\ \boldsymbol{\beta}_2 \\ \vdots \\ \boldsymbol{\beta}_J \end{bmatrix} + \begin{bmatrix} \boldsymbol{\varepsilon}_1 \\ \boldsymbol{\varepsilon}_2 \\ \vdots \\ \boldsymbol{\varepsilon}_J \end{bmatrix} \tag{3.13}$$

[24]Note that $L \neq IT$ because the panel is unbalanced.

To introduce the summing-up restriction $\sum_j s^j_{i,t} = 1$, we start by rewriting fund i's portfolio share of currency j in quarter t as

$$s^j_{i,t} = 1 - \sum_{m \neq j} s^m_{i,t} = \boldsymbol{\iota}^{'}_{j,i,t}\boldsymbol{s}, \tag{3.14}$$

where $\boldsymbol{\iota}_{j,i,t}$ is a JL-sized vector containing $(L-1)J+J$ zeros and $J-1$ ones, which extracts the entries of $\boldsymbol{s}$ corresponding to the i,t-th observation for currency j.

We can then re-write the vector s_j of all observations for currency j:

$$\boldsymbol{s}_j = \mathbf{1} - \boldsymbol{\iota}^{'}_j\boldsymbol{s} = \mathbf{1} - \boldsymbol{\iota}^{'}_j(\boldsymbol{X\beta}) + \boldsymbol{\varepsilon}, \tag{3.15}$$

where $\boldsymbol{\iota}^{'}_j$ is a $L \times JL$ matrix that stacks vectors $\boldsymbol{\iota}^{'}_{j,i,t}$ across the fund and time dimensions. We then substitute $\boldsymbol{s} = \boldsymbol{X\beta} + \boldsymbol{\varepsilon}$ from the joint estimation system in the second step.

The resulting equation is used as constraint by appending it to the joint estimation system:

$$\begin{bmatrix} \boldsymbol{s} \\ \boldsymbol{s}_j \end{bmatrix} = \begin{bmatrix} \boldsymbol{X\beta} + \boldsymbol{\varepsilon} \\ 1 - \boldsymbol{\iota}^{'}_j(\boldsymbol{X\beta} + \boldsymbol{\varepsilon}) \end{bmatrix}. \tag{3.16}$$

By rearranging the system, we can write it as

$$\underbrace{\begin{bmatrix} \boldsymbol{s} \\ \boldsymbol{s}_j - \mathbf{1}_{L\times 1} \end{bmatrix}}_{\equiv \tilde{\boldsymbol{s}}} = \underbrace{\begin{bmatrix} \boldsymbol{I}_{JL} \\ -\boldsymbol{\iota}^{'}_j \end{bmatrix} \boldsymbol{X}}_{\equiv \tilde{\boldsymbol{X}}} \boldsymbol{\beta} + \underbrace{\begin{bmatrix} \varepsilon \\ -\boldsymbol{\iota}^{'}_j\varepsilon \end{bmatrix}}_{\equiv \tilde{\boldsymbol{\varepsilon}}}. \tag{3.17}$$

The coefficients $\boldsymbol{\beta}$ can then be estimated simply by running an OLS regression with the same standard errors as the baseline model on the transformed system $\tilde{\boldsymbol{s}} = \tilde{\boldsymbol{X}}\tilde{\boldsymbol{\beta}} + \tilde{\boldsymbol{\varepsilon}}$. Therefore, the summing-up restriction is imposed by adding a $(L+1)$-th group of observations that contains transformed values of the dependent variables for one currency j, which are excluded from the other rows. [25]

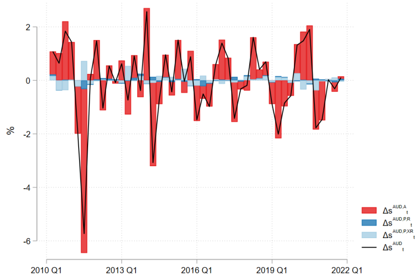
3.H.2 Results

Tables 3.30 and 3.31 display the results of the joint estimation of the baseline model, run separately for unhedged and hedged returns. Unhedged returns results are broadly comparable with those of the baseline model, with higher elasticities

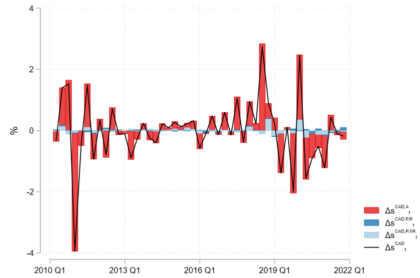
[25] The estimation procedure is invariant to which currency j is "excluded".

for the euro and US dollar. The hedged return models show instead highly significant and economically large coefficients on $r_{i,t}^{ex,j,unh}$ for the Australian dollar, the Canadian dollar and the euro.

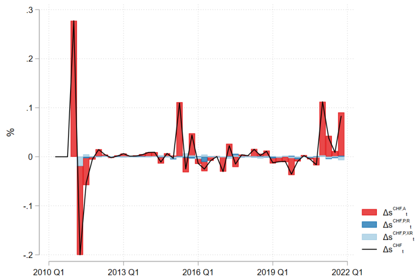
Figure 3.6. Decomposition of aggregate active and passive reallocation



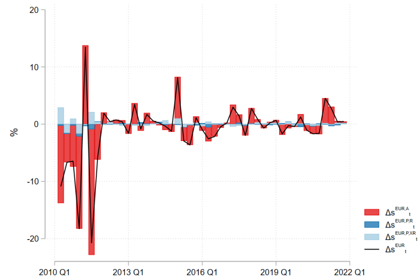
(a) Australia



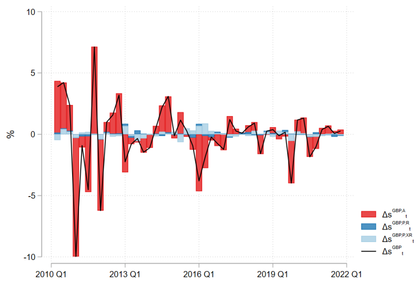
(b) Canada



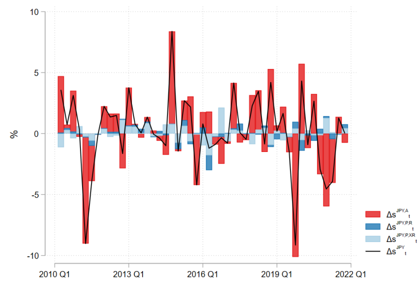
(c) Switzerland



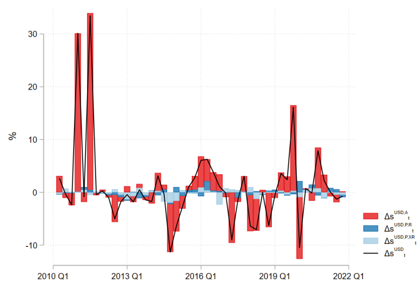
(d) Euro Area



(e) United Kingdom



(f) Japan



(g) United States

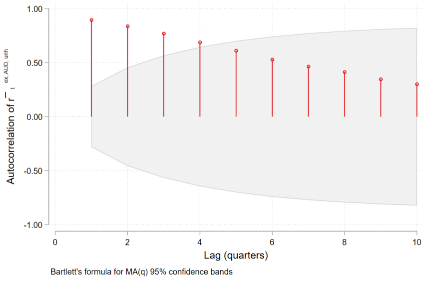
Portfolio shares are calculated on the basis of aggregate amounts held by all funds in the sample excluding funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Δs_t is the change in aggregate country j portfolio share;

$\Delta s_t^{P,R} = s_{t-1}^j \left(\frac{A_t^j}{A_t} - 1 \right)$ is the passive reallocation due to bond returns in the currency of issuance;

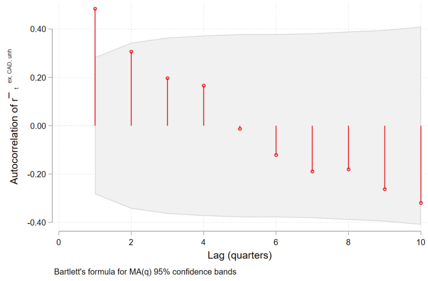
$\Delta s_t^{P,XR} = s_{t-1}^j \left(\frac{A_t^j}{A_t} - 1 \right)$ is the passive reallocation due to exchange rate effects;

$\Delta s_t^A = \Delta s_t - \Delta s_t^{P,R} - \Delta s_t^{P,XR}$ is the active reallocation.

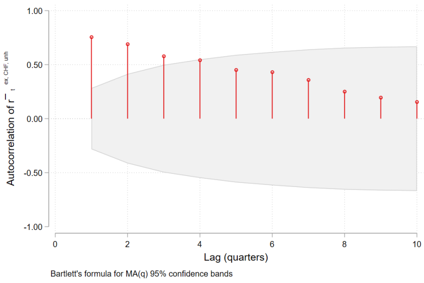
Figure 3.7. Autocorrelation of fund-level unhedged excess returns $r_{i,t}^{ex,j,uhn}$.



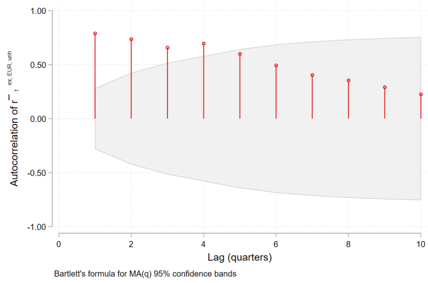
(a) AUD



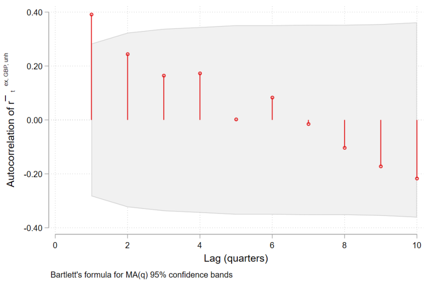
(b) CAD



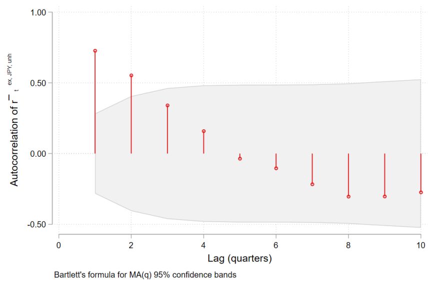
(c) CHF



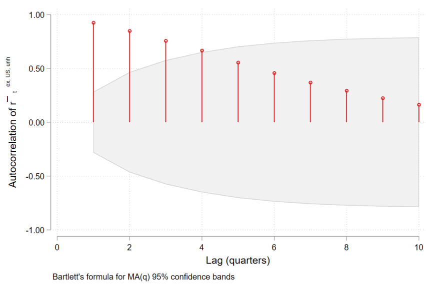
(d) EUR



(e) GBP



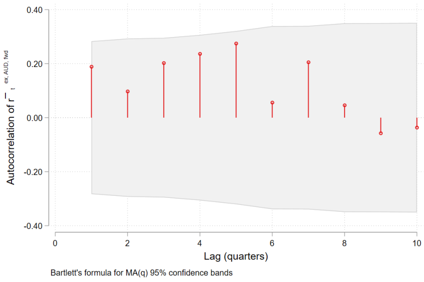
(f) JPY



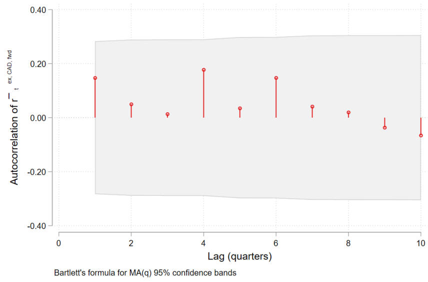
(g) USD

Quarterly autocorrelations calculated on the basis of average fund-level excess returns in the sample that excludes funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Confidence bands are calculated at the 95% level using Bartlett (1946)'s formula.

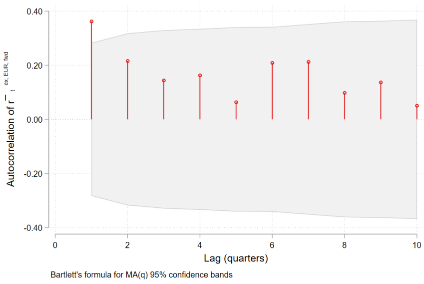
Figure 3.8. Autocorrelation of fund-level hedged excess returns $r_{i,t}^{ex,j,fwd}$.



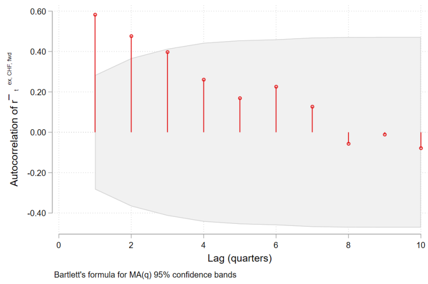
(a) AUD



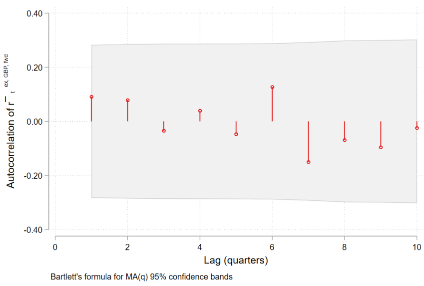
(b) CAD



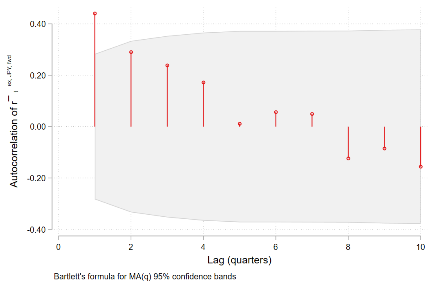
(c) EUR



(d) CHF



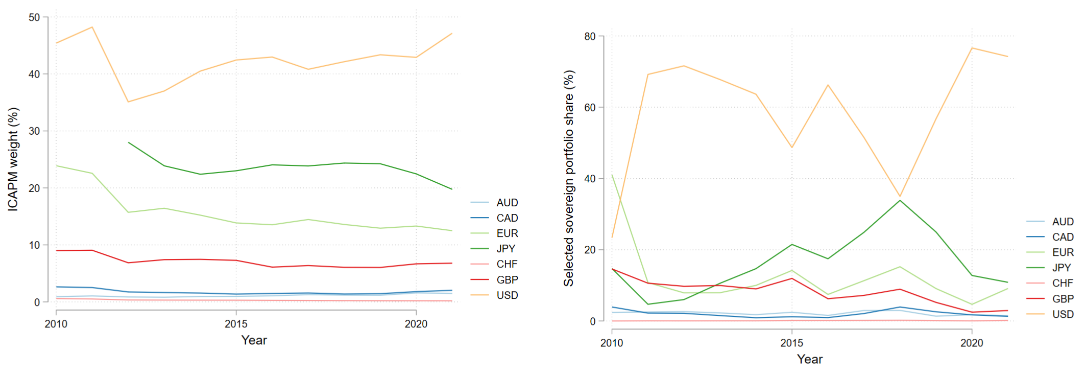
(e) GBP



(f) JPY

Quarterly autocorrelations calculated on the basis of average fund-level excess returns in the sample that excludes funds with $\bar{s}_i^j \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Confidence bands are calculated at the 95% level using Bartlett (1946)'s formula.

Figure 3.9. ICAPM and global fund sample share comparison



The left-hand side chart shows ICAPM weights for selected sovereigns, calculated as the relative market capitalization of outstanding central government debt. The right-hand side chart shows aggregate shares in the portfolio of selected sovereigns for funds in the bottom 50% of the the $\bar{d}_i$ distribution. Source: BIS government bond statistics (Bogdanova et al. (2021)) and Refinitiv Lipper.

Table 3.26. Lagged portfolio share unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.62***	0.71***	0.73***	0.67***	0.60***	0.56***	0.66***
	(0.05)	(0.07)	(0.19)	(0.05)	(0.08)	(0.05)	(0.05)
$s^j_{i,t-2}$	0.06	0.02	0.01	0.04	0.12*	0.09	-0.01
	(0.06)	(0.05)	(0.12)	(0.06)	(0.06)	(0.06)	(0.05)
$s^j_{i,t-3}$	-0.01	-0.01	0.01	-0.00	-0.01	-0.01	0.06
	(0.07)	(0.07)	(0.06)	(0.05)	(0.05)	(0.04)	(0.04)
$s^j_{i,t-4}$	0.02	0.01	-0.08	0.04	0.01	-0.04	0.01
	(0.04)	(0.05)	(0.10)	(0.04)	(0.04)	(0.03)	(0.02)
$r^{ex,j,unh}_{i,t}$	0.68	0.18	0.17***	0.61	0.53	0.24	0.96
	(0.63)	(0.23)	(0.04)	(0.50)	(0.64)	(0.44)	(0.90)
$\Delta s^{j,P,R}_{i,t}$	0.19	1.87	-3.45	-1.18	0.44	0.34	-1.55
	(3.54)	(4.82)	(3.26)	(3.10)	(1.95)	(1.91)	(1.94)
$\Delta s^{j,P,XR}_{i,t}$	-0.47	0.60	7.07**	-0.37	0.71	1.27**	-0.20
	(1.09)	(1.74)	(3.18)	(0.95)	(0.43)	(0.57)	(1.24)
Within R^2	0.57	0.58	0.57	0.54	0.57	0.43	0.53
N (fund-quarter)	2,651	2,654	564	2,717	2,744	2,882	3,296
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	41	42	43	41	42	41	40

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \sum_{p=1}^4 \beta^j_1 s^j_{i,t-p} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.27. Lagged portfolio share hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.64***	0.70***	0.74***	0.68***	0.60***	0.57***
	(0.05)	(0.06)	(0.19)	(0.05)	(0.07)	(0.05)
$s^j_{i,t-2}$	0.06	0.02	0.01	0.04	0.11*	0.09*
	(0.06)	(0.05)	(0.12)	(0.06)	(0.06)	(0.06)
$s^j_{i,t-3}$	-0.01	-0.02	0.01	-0.00	-0.02	-0.01
	(0.07)	(0.07)	(0.06)	(0.05)	(0.05)	(0.04)
$s^j_{i,t-4}$	0.02	0.01	-0.08	0.04	0.02	-0.04
	(0.04)	(0.04)	(0.10)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-3.54	3.50**	-0.23	-0.59	5.70***	-2.43
	(2.61)	(1.32)	(0.48)	(2.38)	(1.51)	(2.05)
$\Delta s^{j,P,R}_{i,t}$	0.15	2.00	-3.49	-1.21	0.28	0.24
	(3.47)	(4.69)	(3.26)	(3.15)	(1.92)	(1.89)
$\Delta s^{j,P,XR}_{i,t}$	-0.40	0.49	7.07**	-0.41	0.76*	1.32**
	(1.02)	(1.73)	(3.19)	(0.93)	(0.45)	(0.58)
Within R^2	0.57	0.58	0.57	0.54	0.57	0.43
N (fund-quarter)	2,651	2,654	564	2,717	2,744	2,882
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model $s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \sum_{p=1}^4 \beta^j_1 s^j_{i,t-p} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 \Delta s^{j,P,R}_{i,t} + \beta^j_4 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.28. Lagged unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.77***	0.75***	0.71***	0.70***	0.71***	0.63***	0.74***
	(0.03)	(0.03)	(0.12)	(0.04)	(0.04)	(0.03)	(0.03)
$r^{ex,j,unh}_{i,t}$	1.22	0.79	-0.03	1.08	0.91	1.11*	2.01*
	(0.95)	(0.83)	(0.12)	(0.67)	(1.04)	(0.61)	(1.17)
$r^{ex,j,unh}_{i,t-1}$	-0.24	-0.45	0.20	-0.19	-0.60	-0.36	-0.22
	(0.87)	(0.89)	(0.13)	(0.60)	(0.70)	(0.62)	(0.80)
$\Delta s^{j,P,R}_{i,t}$	-0.23	-0.06	0.88	-1.67	0.55	-0.28	-2.84
	(4.36)	(5.66)	(4.17)	(2.78)	(2.19)	(1.75)	(1.80)
$\Delta s^{j,P,XR}_{i,t}$	0.26	0.89	2.83	-0.15	0.89**	1.53***	0.48
	(1.11)	(1.43)	(1.84)	(1.11)	(0.40)	(0.48)	(1.14)
Within R^2	0.68	0.62	0.56	0.53	0.57	0.48	0.61
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668	4,359
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	41	42	43	41	42	41	40

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 r^{ex,j,unh}_{i,t-1} + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.29. Lagged hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.73***	0.71***	0.71***	0.70***	0.63***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-1.03	3.85***	-0.44	-3.70*	4.17**	1.73
	(2.08)	(1.06)	(0.39)	(2.07)	(1.88)	(1.52)
$r^{ex,j,fwd}_{i,t-1}$	-0.84	-0.45	0.03	0.19	-0.55	0.42
	(0.87)	(0.42)	(0.10)	(0.67)	(0.59)	(1.20)
$\Delta s^{j,P,R}_{i,t}$	-0.35	-0.00	0.89	-1.90	0.48	-0.27
	(4.31)	(5.45)	(4.17)	(2.86)	(2.18)	(1.75)
$\Delta s^{j,P,XR}_{i,t}$	0.36	0.73	2.78	-0.17	0.94**	1.46***
	(1.06)	(1.45)	(1.82)	(1.08)	(0.42)	(0.53)
Within R^2	0.68	0.62	0.56	0.53	0.57	0.48
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \eta^j \mathbb{1}^{Out,j}_{i,t} + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 r^{ex,j,unh}_{i,t-1} + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.30. Unhedged returns joint estimation

	AUD	CAD	CHF	GBP	JPY	EUR	USD
$s^j_{i,t-1}$	0.62***	0.58***	0.85***	0.59***	0.57***	0.51***	0.47***
	(0.03)	(0.02)	(0.18)	(0.03)	(0.04)	(0.03)	(0.02)
$r^{ex,j,unh}_{i,t}$	-0.03	0.84*	-2.27**	0.72	0.50	1.29***	2.47**
	(0.25)	(0.48)	(1.05)	(0.64)	(0.62)	(0.46)	(1.04)
$\Delta s^{j,P,R}_{i,t}$	-2.58	-1.60	-2.40	-1.64	1.11	0.18	-0.06
	(3.86)	(3.48)	(5.40)	(2.42)	(1.76)	(1.44)	(1.67)
$\Delta s^{j,P,XR}_{i,t}$	-0.21	0.52	2.70	-0.00	0.83*	0.76	0.20
	(0.90)	(1.10)	(1.92)	(0.56)	(0.43)	(0.49)	(0.79)
Within R^2	0.52	0.52	0.52	0.52	0.52	0.52	0.52
N (fund-quarter)	3,778	3,606	739	3,744	3,700	3,999	4,830
N (funds)	129	113	22	122	117	131	177
Avg. nr. quarters	41	41	41	41	41	41	41
N (joint estimation) 26720							

Coefficients from joint regression model $\tilde{\boldsymbol{s}} = \tilde{\boldsymbol{X}}\tilde{\boldsymbol{\beta}} + \tilde{\boldsymbol{\varepsilon}}$ using unhedged returns $r^{ex,j,unh}_{i,t}$. Each column reports results for a different currency j, but all coefficients are estimated jointly. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.31. Hedged returns joint estimation

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.58***	0.58***	0.96***	0.61***	0.60***	0.51***
	(0.02)	(0.02)	(0.22)	(0.03)	(0.05)	(0.02)
$r^{ex,j,fwd}_{i,t}$	1.79***	4.02***	2.91	-0.14	-0.93	2.63**
	(0.60)	(1.23)	(2.02)	(1.29)	(1.51)	(1.23)
$\Delta s^{j,P,R}_{i,t}$	-1.89	-1.66	-3.30	-1.55	1.31	0.39
	(3.98)	(2.66)	(5.86)	(2.34)	(1.78)	(1.48)
$\Delta s^{j,P,XR}_{i,t}$	-0.44	0.38	2.93	0.08	0.91*	0.75
	(0.88)	(0.98)	(1.92)	(0.53)	(0.46)	(0.51)
Within R^2	0.52	0.52	0.52	0.52	0.52	0.52
N (fund-quarter)	3,778	3,606	739	3,744	3,700	3,999
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	33	33	33	33	33	33
N (joint estimation) 26720						

Coefficients from joint regression model $\tilde{\boldsymbol{s}} = \tilde{\boldsymbol{X}}\tilde{\boldsymbol{\beta}} + \tilde{\boldsymbol{\varepsilon}}$ using hedged returns $r^{ex,j,fwd}_{i,t}$. Each column reports results for a different currency j, but all coefficients are estimated jointly. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

3.I Further results

3.I.1 Policy rates models with time and fund fixed effects

Table 3.32. Policy rate time and fund FE unhedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.77***	0.75***	0.71***	0.70***	0.71***	0.63***
	(0.03)	(0.03)	(0.12)	(0.04)	(0.04)	(0.03)
$r^{ex,j,unh}_{i,t}$	0.89	0.59	-0.08	1.71**	-0.16	1.23***
	(1.82)	(0.39)	(0.14)	(0.75)	(0.58)	(0.35)
$cb^{US}_t \times r^{ex,j,unh}_{i,t}$	0.00	-0.00	0.00*	-0.01**	0.01*	-0.01*
	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta s^{j,P,R}_{i,t}$	-0.26	-0.18	0.91	-1.59	0.64	-0.22
	(4.35)	(5.62)	(4.20)	(2.81)	(2.22)	(1.74)
$\Delta s^{j,P,XR}_{i,t}$	0.27	0.85	2.78	-0.16	0.86**	1.50***
	(1.13)	(1.46)	(1.82)	(1.10)	(0.41)	(0.49)
Within R^2	0.68	0.62	0.56	0.53	0.57	0.48
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,unh}_{i,t} + \beta^j_3 cb^{US}_t * r^{ex,j,unh}_{i,t} + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

3.I.2 Search for safety hedged returns

Table 3.33. Policy rate time and fund FE hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.73***	0.69***	0.71***	0.70***	0.63***
	(0.03)	(0.03)	(0.12)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-0.88	2.57**	0.19	-5.89***	1.11	-1.80
	(2.98)	(1.12)	(0.41)	(1.90)	(3.49)	(1.70)
$cb^{US}_t \times r^{ex,j,fwd}_{i,t}$	-0.00	0.01	-0.01***	0.02	0.02	0.03**
	(0.01)	(0.01)	(0.00)	(0.02)	(0.03)	(0.01)
$\Delta s^{j,P,R}_{i,t}$	-0.37	-0.05	0.86	-2.02	0.43	-0.31
	(4.30)	(5.53)	(4.22)	(2.93)	(2.21)	(1.73)
$\Delta s^{j,P,XR}_{i,t}$	0.33	0.74	2.77	-0.18	0.90**	1.54***
	(1.06)	(1.44)	(1.81)	(1.07)	(0.41)	(0.51)
Within R^2	0.68	0.62	0.56	0.53	0.57	0.48
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	41	42	43	41	42	41

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \gamma^j_t + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 cb^{US}_t * r^{ex,j,unh}_{i,t} + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.34. Search for safety hedged returns regressions

	AUD	CAD	CHF	GBP	JPY	EUR
$s^j_{i,t-1}$	0.79***	0.73***	0.71***	0.72***	0.73***	0.64***
	(0.03)	(0.03)	(0.11)	(0.04)	(0.04)	(0.03)
$r^{ex,j,fwd}_{i,t}$	-0.16	3.88***	-0.16	-3.26*	2.21*	1.87
	(0.74)	(0.88)	(0.16)	(1.69)	(1.25)	(1.30)
VIX_t	24.48	6.10	-2.81	-13.58	-44.02	-52.02
	(21.23)	(14.81)	(2.92)	(28.42)	(39.11)	(34.85)
$\Delta s^{j,P,R}_{i,t}$	0.55	-0.58	1.03	0.12	-0.09	0.31
	(4.34)	(5.20)	(3.75)	(2.47)	(1.57)	(1.52)
$\Delta s^{j,P,XR}_{i,t}$	0.34	0.82	2.66	-0.20	0.77**	0.93*
	(1.02)	(1.29)	(1.63)	(0.86)	(0.32)	(0.50)
Within R^2	0.67	0.61	0.53	0.51	0.56	0.47
N (fund-quarter)	3,442	3,333	692	3,442	3,425	3,668
N (funds)	129	113	22	122	117	131
Avg. nr. quarters	32	33	32	33	33	33

Coefficients from regression model
$s^j_{i,t} = \alpha^j_i + \beta^j_1 s^j_{i,t-1} + \beta^j_2 r^{ex,j,fwd}_{i,t} + \beta^j_3 VIX_t + \beta^j_4 \Delta s^{j,P,R}_{i,t} + \beta^j_5 \Delta s^{j,P,XR}_{i,t} + \boldsymbol{\beta^j_6 W^j_t} + \varepsilon^j_{i,t}$. Each column reports results for a different currency j. $\boldsymbol{W_t}$ includes year-on-year inflation for country j in quarter t in percentage points, the Citi Economics Surprise Index for country j in quarter t in standard deviation units, and the VIX in quarter t in standard deviation units. Each model excludes funds for which the average portfolio share of currency j is equal to 0 and funds with $\bar{s}^j_i \geq 0.95$ for any j. The sample also excludes fund-quarter observations with no holdings of sovereign debt issued by the selected countries. Driscoll and Kraay (1998) standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Chapter 4

Spillovers of LSAPs through US Treasuries on foreign balance sheets

This paper introduces and quantifies a new financial channel for the international transmission of large scale asset purchase programmes (LSAP) by the Federal Reserve. European banks hold a significant amount of U.S. Treasuries on their balance sheet, and quantiative easing (QE) (or tightening) policies in the USA affect the value of their holdings via two opposing channels. It raises the price of long-term U.S. Treasuries, resulting in a capital gain for banks (price channel*). On the other hand, it depreciates the U.S. dollar, resulting in a capital loss absent perfect hedging (*exchange rate channel*). The relative size of the two effects is not obvious* ex ante*, and we measure it, as well as the ramifications for credit provision, using granular data on European banks. The exchange rate channel dominates, and banks actively rebalance away from U.S. Treasuries in response to QE. In addition, bank net worth drops and lending contracts. QE by the Federal Reserve can then have* negative *spillovers to the real economy in Europe through banks' exposure to U.S. Treasuries. The overall reaction of net worth and credit is more muted for large banks, suggesting that they can better cushion the impact of the shock.*

This chapter is co-authored with Marius Koechlin and Andreas Tischbirek. All authors provided equal contributions.

4.1 Introduction

In the early months of 2023, the dramatic failure of Silicon Valley Bank precipitated unrealized losses on their US Treasury portfolio, showed that the world's safe asset of choice is not devoid of risk. The recent round of conventional policy tightening brought into sharp relief the severity of interest risk embedded in government bonds, and the ongoing process of reversing quantitative easing (QE) raises questions on the implications of valuation effects on government bonds for the financial sector and credit conditions, both in the USA and abroad. Such concerns have been raised for banks in Europe too, and they did not escape the attention of policymakers. The vulnerability to interest rate hikes demonstrated by the American banking sector prompted the European Banking Authority (EBA) to conduct an ad-hoc exercise on bond holdings by European banks. They concluded that unrealised losses on debt securities held at amortised cost amounted to €75 billions as of February 2023, on a steadily increasing trend since December 2021 (European Banking Authority, 2023).

European banks are also substantially exposed to US Treasuries, and they face an additional source of risk in the form of exchange rate fluctuations, provided that they do not perfectly hedge. By both raising of the price of US Treasuries in US dollars and causing a US dollar depreciation, QE policies by the Federal Reserve (Fed) affect the value of US government bonds held by banks in Europe in two opposite directions. Through the lens of a financial accelerator framework in which banks are leverage-constrained (Bernanke et al., 1999), the resulting capital gains or losses cause bank lending to expand or contract. The relative size of these two effects is an open question, to which this paper provides an empirical answer.

Studies on the spillover of the Fed' large scale asset purchases (LSAP) abound, but they chiefly point to a *positive* effect on foreign economies, be it via a reduction of foreign yields (Neely, 2010; Bauer and Neely, 2014) and term premia (Alpanda and Kabaca, 2020), an increase in equity prices (Chen et al., 2012), or a boost to local aggregate demand (Kolasa and Wesołowski, 2020). In this paper, we instead highlight the potential for QE policies to have *negative* spillovers to the real economy through the financial sector if the exchange rate effect dominates.

First, we provide evidence via local projections that QE shocks cause a drop in the yield of long-term US Treasury and a depreciation of the US dollar through a fall in the spread between long-term US and German government bonds. These results show that the price and exchange rate channels do indeed operate in the expected directions, thereby confirming the results of previous literature.

Then, we estimate the effect of QE shocks on the balance sheet of European banks. We exploit bank-level data provided by the EBA, allowing us to observe the exposure of European banks to US Treasuries at a granular level of detail. We find that QE reduces the value of US Treasury holdings on European banks' balance sheets, consistent with the exchange rate channel dominating. A rebalancing away from US Treasuries is also apparent. We then trace out the ramifications to banks' net worth, finding that a QE shock raises net worth and lending through the price effect, and lowers them through the exchange rate effect. Our estimates of the net balance-sheet effect of QE show that net worth as well as credit decrease, conditional on macroeconomic and financial controls that account for other possible channels of QE spillovers. This finding indicates that the exchange rate effect on US Treasury dominates. The effect on the real economy is large, with a one basis point QE shock by the Fed resulting in an overall aggregate decline of €22 billions through balance sheet effects.

Overall, our results shed new light on the spillovers of the Fed's QE to credit conditions in Europe through the exposure of leverage-constrained banks to US Treasuries. Contrary to the positive spillovers highlighted by previous literature, our results demonstrate that the US dollar depreciation of US Treasuries on banks' balance sheets in response to QE leads to a contraction in lending. On the flipside, they suggest that the ongoing QT policies might instead have a *positive* effect on the European economy through a US dollar appreciation.

Related literature

Our paper sits in the broad literature of studies that investigate empirically the effects of unconventional monetary policy by the Fed. Some foundational contributions include Gagnon et al. (2011), which uses an event-study approach and concludes that QE resulted in substantial reductions in yield for a broad range of securities through lower risk premia; and Krishnamurthy and Vissing-Jorgensen (2011), which identifies the effects of QE through a broad range of channels such as signaling, safe asset demand and inflation depending on the specific type of operation. More references can be found in Bhattarai and Neely (2016), a survey of empirical literature on QE. Many of the papers in this line of inquiry focus on disentangling empirically the effects of the portfolio balance and signaling channels on interest rates. According to the former, unconventional policy lowers long-term yield through a commitment by the central bank to keep short rates lower than optimal in the future (Eggertsson and Woodford, 2003). On the other hand, the latter operates directly on US Treasury rates by altering their supply, which will in turn affect the equilibrium rate in a framework of imperfect subsititutability between assets (Vayanos and Vila, 2021). The mechanism we examine is consistent with a portfolio balance argument, in which the effect on US Treasury rates is potentially offset by a US dollar depreciation.

The transmission channel we introduce operates directly through the balance sheet of banks, which other studies also identify as crucial mediators for the effects of QE. In a closed-economy context, several papers examine the effects of unconventional policy in financial accelerator models with leverage-constrained banks. QE can operate either directly through the prices of assets held by banks (Karadi and Nakov, 2021), or indirectly by easing financial constraints in times of crises (Gertler and Karadi, 2011). A strand of this literature also adopt an open-economy approach, studying the effects of the Fed's QE through foreign financial institutions. For example, Morais et al. (2019) finds evidence of credit expansion in Mexico following unconventional monetary easing in the QE through reaching for yield in higher-risk assets. Some papers do examine the effects of exchange rate fluctuations on banks' balance sheets in a financial accelerator framework, by either remaining agnostic on their source (Longaric, 2022), or by modelling them as resulting from US monetary policy (Aoki et al., 2016; Akinci and Queraltó, 2019). However, they do so in the context of emerging markets where banks issue US dollar liabilities in excess of their US dollar assets, hence making them vulnerable to US dollar *appreciations*. Instead, we focus on the role of US dollar-denominated *assets*, which is more relevant for advanced economies as their banking sectors are generally less reliant on dollar liabilities.

In a narrower sense, the papers most closely related to our own, study the international spillovers of QE by the Fed. Early research in this area focused on the effects of QE on foreign financial markets, with evidence of lower international bond yields (Neely, 2010) and term premia (Bauer and Neely, 2014), and portfolio rebalancing towards non-US assets (Fratzscher et al., 2018). Concerning real spillovers, (Chen et al., 2016) finds expansionary effects on advanced and emerging economies alike through lower corporate spreads. In a model with nominal and real rigidities and imperfect subsititutability between domestic and foreign bonds, Alpanda and Kabaca (2020) shows that QE in the USA has an expansionary effect on the foreign economy through lower term premia and rebalancing towards foreign long-term bonds. Note that in this framework, unlike in our paper, the partial effect of a QE-induced US dollar depreciation on the foreign economy is *expansionary* due to trade balance effects. Kolasa and Wesołowski (2020) highlights instead the role of asset market segmentation for the international transmission of QE. In a small-open economy model where only some households can access US bonds, QE by the Fed results in capital inflows for the foreign economy, accompanied by higher demand but lower output.

Finally, our study also contributes to the strand of literature focusing on the effects of unconventional monetary policy on exchange rates, and the implications thereof. Dedola et al. (2013) shows that QE by the Fed results in a large and persistent US dollar depreciation through a standard uncovered interest parity (UIP) channel. Likewise, Gourinchas et al. (2022) and Jiang et al. (2024) highlight that QE affects the exchange rate by altering the supply of US Treasuries in a framework of imperfect subsitititutability motivated by liquidity preference or preferred habitat. We contribute to this literature by tracing out the effects of US dollar depreciation to the real economy in Europe through banks' balance sheets.

The rest of the paper is structured as follows. Section 4.2 describes the EBA Transparency Exercise and Stress Test data and reports summary statistics on the exposure of European banks to US Treasuries. Section 4.3 presents evidence on the effect of the Fed's QE shocks on term spreads, international spreads and exchange rates through a local projection approach. Section 4.4 lays out a simple framework to explain the effects of the Fed's QE on the balance sheet of leverage constrained banks through their holdings of US Treasuries. Section 4.5 investigates the effects of the Fed's QE shocks on the balance sheet, analysing the reaction of their Treasury holdings, net worth and credit provision through the lens of the price, portfolio rebalancing and exchange rate channels. Section 4.6 concludes.

4.2 Data

4.2.1 The EBA Transparency Exercise and Stress Test datasets

We use bank-level data from the European Banking Authority's (EBA) Transparency Exercise and Stress Test databases. They are based on regulatory reports filed by the largest banks domiciled in the European Economic Area, which includes the European Union plus Iceland, Liechtenstein, and Norway.

The reports are semi-annual, available from December 2010 to June 2022 for a total of 20 semesters.[1] For the purpose of this study, we restrict our attention to banks operating in the euro area to simplify exchange rate effects to the EUR/USD currency pair only.

[1]The Transparency Exercise was first conducted in 2013, and did not take place in 2014. We supplement the data with information from the 2011 and 2014 Stress Tests, which contain data for December 2011 and December 2013. Therefore, we miss observations for the first semesters of 2012 and 2014.

This restriction leaves us with 152 out of 189 banks, not all of which are observed every semester. The resulting dataset is an unbalanced panel with 1671 observations.[2] Appendix 4.B.3 details the number of banks by size quartile and country, showing that banks from Germany, Italy, Spain and France are the most represented.

We can observe the exposure of individual banks to US government debt, broken down by maturity buckets and accounting portfolio.[3] The database also includes information on other balance sheet items such as *Total Assets, Tier 1 Capital ratios* and *Credit*. The variable *Total Assets* is interpolated with data from CapitalIQ.

The QE shock series used in our empirical analysis is identified by Jarocinski (2021). Jarocinski adopts a high-frequency identification approach, focusing on the 30-minute window surrounding the Federal Open Market Committee (FOMC) meetings. This approach allows him to isolate monetary policy shocks affecting both long-term and short-term interest rates in the USA. Notably, Jarocinski takes into consideration the non-Gaussian nature of market reactions, employing a Student-t distribution. This unique approach enables him to identify these shocks without imposing additional economic restrictions. To facilitate our analysis, we aggregate these shocks at the quarterly level. The aggregation involves summing the shocks that occurred within each quarter. On average, there are approximately four to five such shocks within each quarterly period.

Summary statistics

Table 4.1 reports summary statistics of the bank-level variables that we include in our analysis. Notwithstanding the restriction to banks subject to EBA reporting, our sample spans a wide range of sizes, with total assets ranging from €280 million to more than €2 trillion. In total, the dataset covers approximately 67% of the

[2]Some variables such as total assets and credit are not reported in every semester. Furthermore, the Jarocinski (2021) QE shock series, which we use for our estimates, is available only until June 2019. Therefore, the number of observations in regressions is lower. More specifically, we lose 595 observations which leaves us with a total of 1076 observations.

[3]Banks can report their balance sheet holdings of government bonds as either available for trading or held to maturity. In the former case, they report them at market value, while in the latter they report them at amortised historical cost. The EBA datasets record information on both types of holdings separately, which we refer to in the text as *Book Value Exposure* and *Market Value Exposure*. It also reports information on *Non-derivative Exposure*, which is the sum of market and book value exposure, and *Derivative exposure*. The sum of derivative and non-derivative exposure is reported as *Total Exposure*.

eurozone's banking sector by assets as of June 2022.[4]

US Treasury holdings make up a relatively small percentage of banks' total assets on average, although some banks are heavily exposed. Holdings of US government bonds are small compared to total credit as well, on average less than €2 billion of exposure at market value versus €59.88 billion of credit, reflecting a sample of mostly commercial banks whose core business is lending. However, Figure 4.1 shows that US Treasuries add up to a non-negligible share of capital for many banks, with an average total exposure of 10.25 % of Tier 1 capital and several instances in which US Treasury holdings exceed the entire capital cushion. Therefore, fluctuations in the value of US Treasuries through prices and exchange rates resulting from unconventional monetary policy can significantly affect bank capital. In turn, these fluctuations in net worth can impact credit provision and the real economy through leverage constraints.

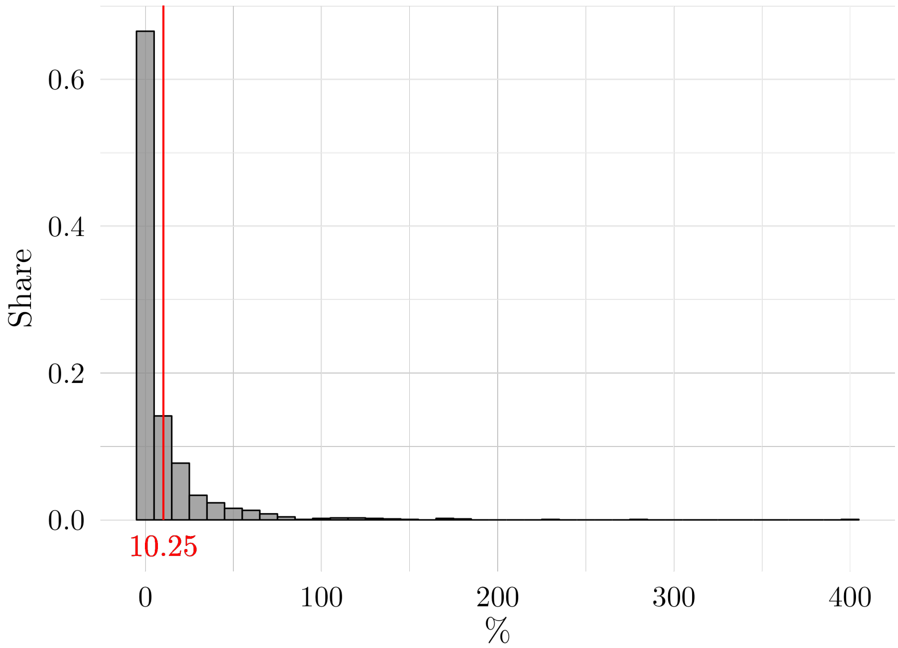
Table 4.1. Summary statistics

Variable	Obs.	Mean	SD	Min	Median	Max
Total assets	749	247.85	400.13	0.28	75.32	2171.39
Tier1 capital	1019	13.05	18.73	0.08	5.35	91
Tier1 leverage ratio (%)	749	6.95	5.96	1.15	5.81	76.65
Total U.S. Treas exp	1076	2.14	6.75	0	0	58.76
Market-val. U.S. Treas exp	598	1.87	6.79	-0.16	0	78.83
Book-val. U.S. Treas exp	598	0.7	2.73	-0.08	0	34.26
Total credit	1020	59.88	92.27	0.01	21.53	510.48

Notes: The table shows summary statistics for the estimation sample, which includes 152 European banks. All values except the *Tier1 leverage ratio* are in Euro billions. Negative amounts of market and book value exposure reflect short positions in US Treasuries, excluding positions on US Treasuries reported among financial assets held for trading that are netted out with cash short position.

[4] Aggregate total assets of eurozone-domiciled banks in the EBA Transparency Exercise dataset divided by total assets of Monetary and Financial Institutions domiciled in the eurozone as reported in the European Central Bank's Balance Sheet Items dataset.

Figure 4.1. US Treasury holdings as a fraction of Tier 1 capital



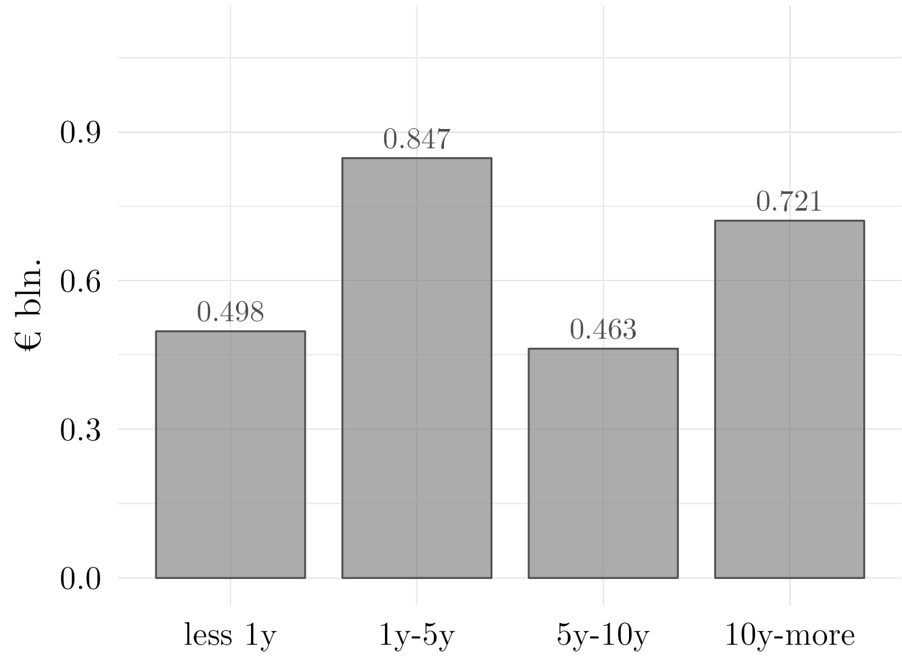
Notes: This figure shows the distribution of total exposure to US Treasuries as a percentage of the banks' Tier 1 capital.

4.2.2 Heterogeneity in US Treasury holdings by bank size

The profile of banks' exposure to US Treasuries is varied in terms of both their importance on the balance sheet and their term structure. The most relevant determinant is bank size (measured as total assets), so it is important to investigate the different sensitivity of large and small banks to QE shocks that affect the value of their US Treasury holdings.

European banks tend to hold US Treasuries mainly in the 1- to 5-year and 10-year or more maturity brackets, with average amounts of nearly €1 billion in either category (Figure 4.2).

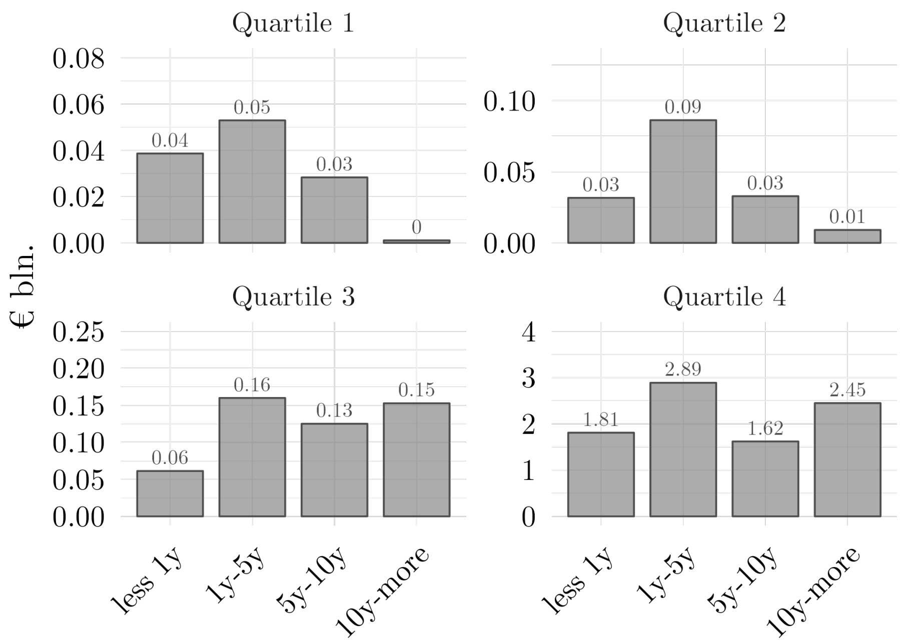
Figure 4.2. Term structure of US Treasury holdings



Notes: European banks holdings of US Treasuries split up in different maturity buckets. US Treasuries with a maturity of less than 1 year are called *Treasury Bills*, *Treasury Notes* have a maturity of 2-10 years, and *Treasury Bonds* of 10 or more years.

Drilling down to the difference in term structure across the bank size distribution, we can see how the pattern observed in the aggregate is driven by large banks. While banks in the first and second size quartile, on average, hardly hold any long-term US Treasuries, those in the upper two quartiles display a US Treasury portfolio that is heavily tilted towards middle and longer maturities (Figure 4.3).

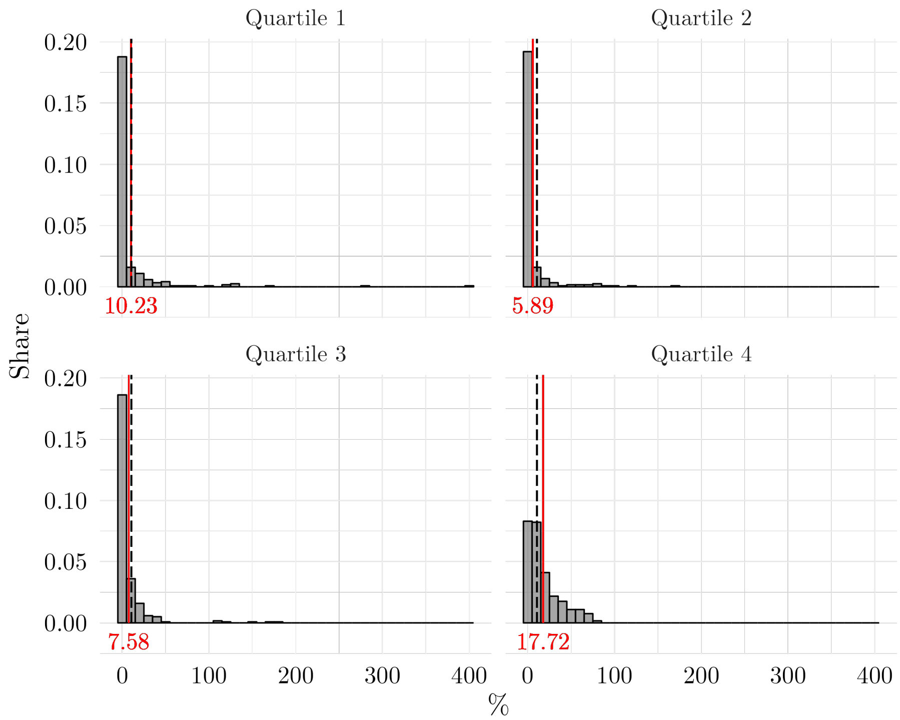
Figure 4.3. Term structure of US Treasury holdings by bank size



Notes: US Treasury holdings split up by maturity and bank size quartile (where bank size is measured by total assets). *Quartile 1* contains the smallest banks, while the biggest banks are reported in *Quartile 4*.

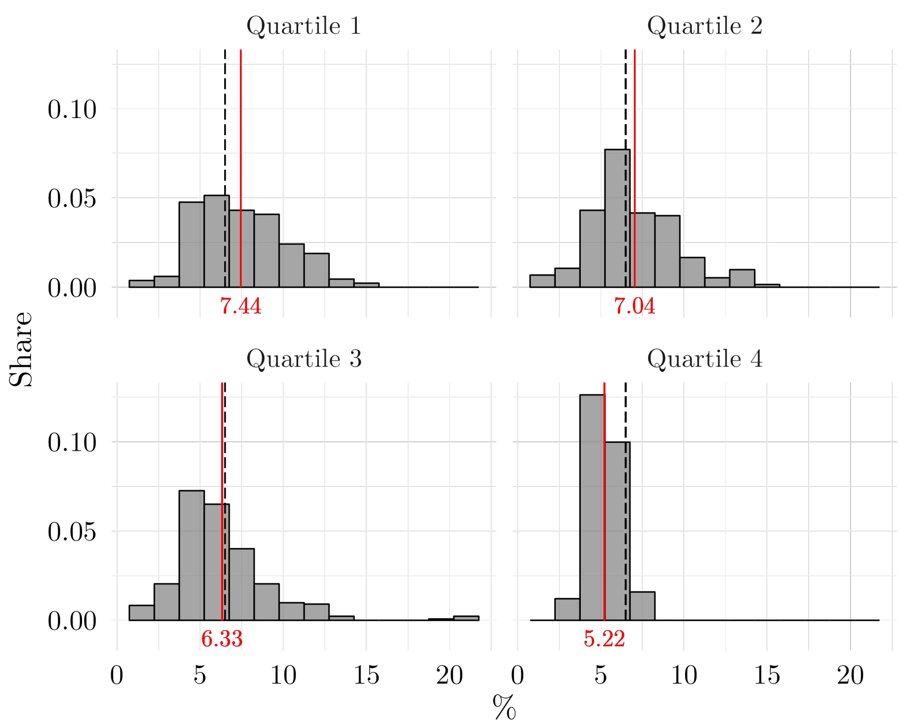
Likewise, larger banks are overall more exposed to US Treasuries as a fraction of net worth. Figure 4.4 shows that banks in the third and fourth quartile by size have an average exposure of 17.5% of their Tier 1 capital, more than 6 percentage points higher than the average of 11% for banks in the first quartile. Large banks also display an altogether less right-skewed distribution characterised by many fewer zeros. In addition, bigger banks also more leveraged, with an average ratio of Tier 1 capital to total asset of 5.23% in the fourth quartile, compared to 7.48% in the first quartile (Figure 4.5). Therefore, the higher US Treasury exposure and leverage might render lending by large banks *ceteris paribus* more sensitive to capital gains or losses incurred on their US Treasury portfolio. On the other hand, large banks might be able to soften the impact of valuation changes of US Treasuries through wider asset diversification, higher rates of foreign exchange and interest rate risk hedging, and better risk-bearing and risk-management expertise (Millon Cornett et al., 2009).

Figure 4.4. US Treasury holdings as fraction of Tier 1 capital by bank size



Notes: Total exposure to US Treasuries as a share of the banks' Tier 1 capital. The top-left panel (*Quartile 1*) shows the distribution of the smallest banks, while the bottom-right panel (*Quartile 4*) shows the distribution for the largest banks.

Figure 4.5. Leverage ratio by bank size



Notes: Tier 1 leverage ratio (defined as the Tier 1 capital divided by total assets) split up by bank size. The top-left panel (*Quartile 1*) shows the distribution of the smallest banks, while the bottom-right panel (*Quartile 4*) shows the distribution for the largest banks.

This heterogeneity suggests that the reaction of bank capital and credit to QE shocks is likely to vary significantly by bank size, and it is not obvious *a priori* whether large banks should react more or less strongly. We address this question in our econometric models by including bank size and interacting it with QE shocks.

4.3 Reaction of bond yields and exchange rates

The two main avenues through which QE affects the value of US Treasuries on European banks' balance sheets are their price in US dollars, and the EUR/USD exchange rate. The existing empirical literature on the effects of QE has focused mainly on the former, finding that LSAP programs by the Fed lead to a drop in the yield of targeted assets (D'Amico and King, 2013; Krishnamurthy and Vissing-Jorgensen, 2011) and a compression in the term premium (Gagnon et al., 2011; Li and Wei, 2013), in accordance with with a portfolio balance channel through imperfect substitution between maturities (Vayanos and Vila, 2021; Greenwood and Vayanos, 2014). Evidence on the exchange rate effect is less abundant, but the general consensus is that QE leads to a US dollar depreciation through a form of UIP, either standard (Dedola et al., 2021), or modified by preferred-habitat (Gourinchas et al., 2022) or convenience-yield (Jiang et al., 2024) mechanisms.

In this section, we provide evidence of the joint response of US Treasury yields, exchange rate and international spreads to the Fed's QE. The aim is to establish whether the QE shock series we use in the bank-level panel regressions has the expected effects on financial variables relevant for US Treasuries. Together with the EUR/USD exchange rate, we consider the 10-year US Treasury yield, the 3-month US Treasury yield, the spread between the two, and the long-term international spread between the yield of US Treasuries and German government bonds at the 3-month and 10-year maturities.[5]

The long-term US rate, here measured at the 10-year maturity, is directly targeted by QE, and which is also associated with a compressed term spread. The long-term spread between US and German bonds should then decrease, mediating a dollar depreciation on the exchange rate. However, the short-term US-German spread might move the exchange rate in the opposite direction if QE shifts the US yield curve downwards through signaling effects rather than flattening it Bauer and Rudebusch (2014).

We estimate monthly local projections of the variables listed above on the QE shock series identified by Jarocinski (2021) exploiting excess kurtosis in financial market responses. We choose this shock series as it reflects the state of the art in the identification of monetary policy surprises. Previous approaches distinguish between conventional and generic unconventional shocks (Gürkaynak et al., 2007), or more granularly between conventional, forward guidance, and QE shocks as in (Swanson, 2021). The Jarocinski (2021) can in addition distinguish QE from Delphic and Odyssean forward guidance.

The estimating equation, which follows the Jordà (2005) approach, is

$$y_{t+h} - y_{t-1} = \alpha^h + \beta^h \eta_t + \textstyle\sum_{i=1}^{p} \gamma_i^h \eta_{t-i} + \Gamma_i^h X_t + e_{t+h}, \tag{4.1}$$

where η_t is the QE shock in month t and β^h is the h-horizon component of the impulse response function.[6] In addition, lags of the shock series are included ($p = 7$)

[5]We use long- and short-term yields on the German government bonds because they are the safest in the eurozone and thus likely to be the relevant rate for UIP effects.

[6]The original shock series by Jarocinski (2021) has negative values for QE shocks. In order

and X_t is a vector of contemporaneous controls. We include the standard monetary policy shock, and the Delphic and Odyssean forward guidance surprises which were identified by Jarocinski (2021) in the regression, to account for possible residual contamination in the QE shock series.

Figure 4.6 depicts the impulse response functions to a one basis point QE shock. The 10-year yield on US Treasuries drops by one basis points on impact by construction, and then declines persistently. On the contrary, the 3-month yield does not display a significant reaction. Combined, these two effects result in a compression of the US term premium, which decreases by approximately 1 basis points on impact and then persistently declines for up to one year.

The 10-year spread between US and German government bonds drops by less than 1 basis point on impact and then declines further, while the 3-month counterpart does not display a statistically significant reaction at any horizon. Consistent with UIP, the EUR/USD exchange rate declines by around 0.25% on impact, which corresponds to a US dollar depreciation, and further decreases over the year.

The local projection results are overall consistent with the existing empirical evidence on the drop in US long-term rates and term premia through portfolio balance effects, rather than signaling effects. The estimated US dollar depreciation, combined with the reduction in the 10-year US - Germany spread matches the theoretical prediction of (modified) UIP as well as empirical evidence. The lack of response of the 3-month international spread corroborates the findings on the signaling effects for US rates, and suggests that the 10-year international spread is a more relevant mediator of the exchange rate effect of QE.

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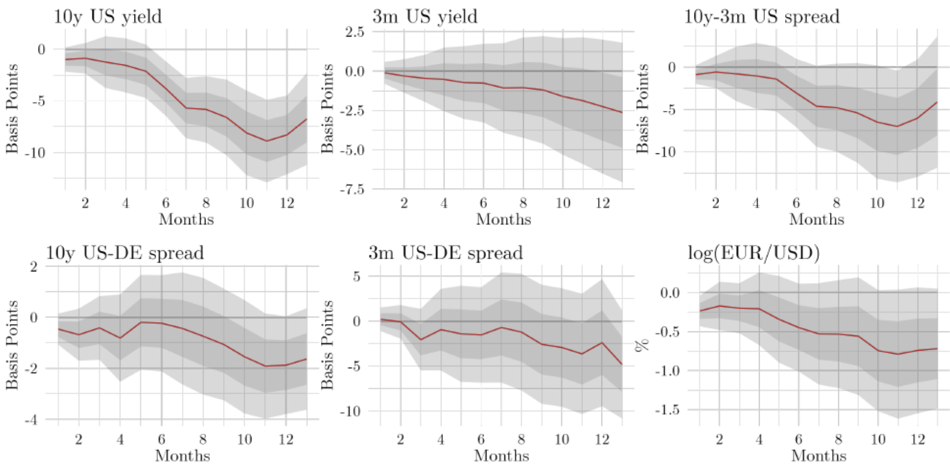
to interpret an increase in η_t as a QE shock, we therefore construct this variable by multiplying the original series by minus one. We adopt this approach because the original shock series is associated with an *increase* in the 10-year US government bond rate, thus representing a contractionary unconventional policy shock.

Figure 4.6. Impulse response of financial variables



Notes: The figure shows the estimates of the response of various financial variables to the QE shock by Jarocinski (2021). The estimates are obtained from the local projection regression outlined in Equation (4.1). Shown in gray are the 68% and 95% confidence bands, using the Newey-West standard errors which correct for heteroscedasticity and autocorrelation.

4.4 Conceptual framework

4.4.1 A simple bank balance sheet

In order to fix ideas and guide our empirical analysis, we build a simple conceptual framework of the channels through which the Fed's QE can affect the balance sheet of European banks. Consider a simplified bank balance sheet, represented in Table 4.2.

Table 4.2. Simplified bank balance sheet

Assets		Liabilities	
S_t:	Credit	D_t:	Deposits
$E_t^{EUR/USD} B_{S,t}^{US}$:	US Treasuries LT	N_t:	Net worth
$E_t^{EUR/USD} B_{L,t}^{US}$:	US Treasuries ST	$E_t^{EUR/USD} L_t$:	US dollar liabilities

Notes: This table shows a hypothetical balance sheet of a European bank. LT and ST stand for long- and short-term respectively.

On the asset side, the bank holds credit S_t denominated in euro, and dollar-denominated US Treasuries, either at long ($E_t^{EUR/USD} B_{L,t}^{US}$) or short ($E_t^{EUR/USD} B_{S,t}^{US}$) maturities. $E_t^{EUR/USD}$ is the EUR/USD exchange rate defined in euros per US dollar, such that an increase is a US dollar appreciation. On the liability side, banks fund themselves with euro-denominated deposits D_t, generic US dollar-denominated liabilities L_t, and net worth N_t. All quantities are in euro at market value.

Both assets and liabilities denominated in dollars are multiplied by $E_t^{EUR/USD}$ because we make the simplifying assumption of no hedging of exchange rate risk. The motivation is twofold. First, information on banks' hedging positions is notoriously opaque as they are not reported on balance sheets (Borio et al., 2017, 2022; Kloks et al., 2023). We cannot directly observe foreign exchange hedging in the EBA dataset, so we need to make an assumption on hedging behaviour. Second, the little evidence on foreign exchange risk exposure available in this dataset shows at the very least that banks tend to not hedge completely. The average exposure to foreign exchange risk across all currencies is €446 millions with peaks of upwards of €10 billions, on a similar order of magnitude as average US Treasury exposure at market value. In our empirical analysis, we will only be able to infer whether the estimated responses to QE are consistent with imperfect hedging.

4.4.2 Transmission through leverage constraints

In the financial accelerator framework, bank net worth is a crucial determinant of credit provision because banks operate under a leverage constraint stemming from an agency problem between bank owners and managers (Bernanke et al., 1999; Gertler and Kiyotaki, 2010; Gertler and Karadi, 2011).[7] As an optimality condition of the agency problem, the bank must maintain its leverage ratio ϕ_t under a threshold level such that

$$\phi_t \equiv \frac{S_t + E_t^{EUR/USD} B_{S,t}^{US} + E_t^{EUR/USD} B_{L,t}^{US}}{N_t} \leq \bar{\phi} \tag{4.2}$$

The maximum leverage $\bar{\phi}$ depends positively on the excess returns of asset prices, which in turn have a positive effect on net worth.

As shown by previous studies and confirmed by our own estimation in Section 4.3, a QE shock has a significant effect on the excess returns of long-term US Treasuries, in terms of both their dollar yield and the EUR/USD exchange rate.

[7]Note that the Basel III banking supervision rules impose a regulatory limit to leverage, so the sensitivity of credit to capital gains and losses is an underlying institutional feature and not strictly dependent on the microfoundations used in the literature.

Therefore, the banks' positions in US Treasuries expose them to two kinds of QE-induced valuation effects: *price* and *exchange rate*. QE by the Fed pushes up the dollar price of long-term US government bonds (price effect), relaxing the leverage constraint and allowing banks to lend more, as discussed in Karadi and Nakov (2021). At the same time, the exchange rate effect lowers the price of US Treasuries in euro through a contemporaneous dollar depreciation (exchange rate effect). The overall effect of the dollar depreciation induced by Fed QE is a function of the net balance sheet exposure to US dollar-denominated assets and liabilities $M_t \equiv E_t^{EUR/USD}(B_{S,t}^{US} + B_{L,t}^{US} - L_t)$. If $M_t > 0$ the bank is net long US dollars and a depreciation of the US dollar will lead to a fall in net worth. If $M_t < 0$, the bank is net short US dollars and a US dollar depreciation will have a positive effect on net worth instead. The EBA dataset does not allow us to observe the amount of US dollar-denominated liabilities, but we attempt to distinguish the exchange rate impact on US Treasuries and on the overall balance sheet with an instrumental variable approach.

The relative size of these two effects is *a priori* not obvious. It depends not only on the relative sensitivity of long-term US Treasury yields and exchange rates to QE, but also on the duration of the bank's US Treasury portfolio. For movements of the same size in long-term US Treasury prices and the EUR/USD exchange rate, banks with a longer-duration portfolio should be more sensitive to the price effect, while we can expect the exchange rate effect to dominate for banks with a shorter-duration portfolio.

In turn, the relative size of the exchange rate and price valuation effects has important implications for the transmission of Fed's QE to the real economy in Europe. If the former dominates, leverage-constrained banks reduce credit so that QE has a contractionary effect, while if the latter dominates the effect on credit will be expansionary.

In addition to valuation effects, portfolio rebalancing by banks in response to changes in relative asset returns can also influence the sensitivity to the Fed's QE. This mechanism has been documented empirically for European banks in response to unconventional monetary policy by the ECB (Koijen et al., 2017; Albertazzi et al., 2021), as well as for other investors in the context of QE programs by the Fed (Carpenter et al., 2015; Fratzscher et al., 2018; Goldstein et al., 2018) and Bank of England (Joyce et al., 2014). Since all US dollar-denominated assets have the same sensitivity to the exchange rate channel, QE reduces the returns of long-maturity US Treasuries compared to shorter maturities through the price effect. Provided that short- and long-term US Treasuries are not perfect substitutes, banks will have an incentive to tilt their portfolio towards short-term US Treasuries (and potentially other assets) due to the higher relative returns, while still demanding positive quantities of both assets in equilibrium.[8] Such portfolio rebalancing can then attenuate the sensitivity of credit to the price effect, hence making banks more vulnerable to the exchange rate effect.

In Section 4.5, we use panel regressions with bank-level EBA data to establish empirically whether price or exchange rate effects dominate. We also investigate the portfolio rebalancing channel by estimating the changes in exposure to both long- and short-term US Treasuries in response to QE. Most importantly, we trace out the effects of the Fed's QE to the real economy in Europe through banks' US Treasury exposure by estimating models of credit provision and net worth.

4.5 Effect of QE on banks' balance sheets

We adopt a panel regression approach to investigate the effect of QE shocks on the holdings of US Treasuries of European banks (Section 4.5.1) and its transmission to credit conditions in the real economy through leverage constraints (Section 4.5.2). The baseline model is the following:

$$y_{i,t}^{hy} = \alpha + \beta\eta_t^q + \gamma(\eta_t^q \cdot BS_{i,t}^{hy}) + \delta BS_{i,t}^{hy} + \Gamma_1 X_{t-1}^q + \Gamma_2 W_t^q + FE_i + \varepsilon_{i,t}. \qquad (4.3)$$

$y_{i,t}^{hy}$ is either holdings of US Treasuries, the leverage ratio or credit depending on the model, measured at the end of semester hy on bank i's balance sheet. η_t^q is the Jarocinski (2021) QE shock, summed up over the quarter q leading up to the end of semester hy. $BS_{i,t}^{hy}$ is the bank size, defined as total bank assets. X_{t-1}^q is a vector of one-quarter lagged control variables, which includes key macroeconomic variables such as the *harmonized unemployment rate*, *real gross domestic product (GDP) growth rate* and the year-on-year *CPI inflation rate*, and also the *CBOE Volatility Index (VIX)* (accessed through Fred). W_t are contemporaneous financial control variables, including the *3-months* and *10-years interest rate surprises* in the euro area (identified by Kearns et al. (2022)), the main *refinancing operations rate* of the ECB, and a *short-* and *long-term borrowing rate* in the eurozone.

The marginal effect of the shock η_t^q on the outcome variable in Equation (4.3), is

[8] Imperfect substitution can be modeled as preferred-habitat (Vayanos and Vila, 2021) different diversion rates in the bank's agency problem (Karadi and Nakov, 2021), or transaction costs in long-term bonds (Chen et al., 2012).

calculated as follows

$$\frac{\partial y}{\partial \eta} = \beta + \gamma\, BS \tag{4.4}$$

where the relevant standard deviation is the following

$$\widehat{\sigma}_{\frac{\partial y}{\partial \eta}} = \sqrt{Var(\hat{\beta}) + (BS)^2\, Var(\hat{\gamma}) + 2\, BS\, Cov(\hat{\beta}, \hat{\gamma})}. \tag{4.5}$$

QE can affect bank lending through a potentially large number of channels, including a reduction of yields (Neely, 2010; Bauer and Neely, 2014) and term premia (Alpanda and Kabaca, 2020) in Europe, an increase in equity prices (Chen et al., 2012), and a boost to local aggregate demand (Kolasa and Wesołowski, 2020). Therefore, we also control for GDP growth rate, inflation and unemployment, a parsimonious set of macroeconomic variables that account for the aggregate demand channel of QE transmission. In the period since 2011 the ECB pursued unconventional monetary policy concurrently with the Fed, so we control for 3-month and 10-year monetary policy surprises from Kearns et al. (2022). By controlling for the direct effect of the ECB's monetary policy on credit conditions in the eurozone, we partial out the documented spillovers of the Fed's QE through changes in yields abroad. Finally, we control for the VIX, a proxy for risk appetite in financial markets that correlates with US monetary policy, asset demand by global investors and exchange rates (Miranda-Agrippino and Rey, 2020; Bruno and Shin, 2015).

FE_i are bank fixed effects, which we add to account for unobserved time-invariant bank-level characteristics. However, since our main explanatory variable is common for all banks, fixed effects tend to absorb much of the bank-level variability in dependent variables and make estimates more noisy, so we also present results without fixed effects. For all regressions we use the inverse hyperbolic sine transformation for the variables in levels, which approximates the natural logarithm of that variable but preserves zero-valued observations.[9]

We add total assets BS_{hy} as an explanatory variable and keep balance sheet amounts in levels. We adopt this approach instead of rescaling amounts by total assets because it allows us to control for the correlation between the size of the bank and amounts of the variables of interest, while avoiding the impact of potential outliers on ratios. In addition, we showed in Section 4.2 that the structure of banks' US Treasury holdings varies by bank size, so we interact QE shocks with BS_{hy} throughout all specifications to investigate hetereogenous responses by bigger banks.

[9]The *Inverse Hyperbolic Sine Transformation* for a variable x is defined as $\tilde{x} = arcsinh(x) = \log(x + \sqrt{x^2+1})$.

4.5.1 Effects on US Treasuries holdings

As outlined in Section 4.4, QE affects the value of US Treasuries on European bank's balance sheet through both prices and exchange rates, which move in opposite directions. Furthermore, banks' exposure to US Treasuries can change in response to QE due to active portfolio rebalancing. In this section, we estimate these effects to investigate whether their signs are consistent with our conceptual framework. It is also important to gauge their relative size to understand the net effect on banks' net worth and lending.

Disentangling the channels of QE

The EBA requires banks to report the exposure to US government debt in both book-value and market-value portfolios. We can exploit this to disentangle the valuation and portfolio rebalancing effects. Exposures at book value are recorded at amortised historical cost, so they do not vary with the market price of US Treasuries. Therefore, by modelling the response of book value amounts in euros we can partial out the price effect, while book value amounts in US dollars are insensitive to exchange rates effects as well, providing us with a clean estimate of portfolio rebalancing effects. Note, however, that only government bonds that are held to maturity are eligible to be held at amortised costs. Estimated changes in book value exposures net of valuation effects can then only reflect purchases, or lack of purchases that would have happened under the counterfactual. Therefore, coefficients are likely to be biased upwards, providing a lower bound for portfolio rebalancing effects if they are negative, and an upper bound if they are positive.

We run models for market value exposure in *euros* to quantify the overall effect of the three channels and establish which one dominates. Likewise, models for market value exposure in *US dollars* allow us to measure the relative size of the price and portfolio rebalancing effects. As a robustness check, we estimate models of market-value exposures in US dollars controlling for the price index of US government bonds in quarter q, which controls for the price effect. Since US dollar amounts are not sensitive to the exchange rate effect, this approach provides an alternative estimate of the portfolio rebalancing effect.

Results

Tables 4.3, 4.4 and 4.5 report results of the estimation of Equation (4.3), where y_t equals the log holdings of US Treasuries by European banks at all maturities, either at market or book value, and in either euros or US dollars. Panels on the left-hand side contain models with fixed effects, while those on the right-hand side do not.

Table 4.3 shows the results for models of exposures in US dollars, which net out the exchange rate effect. Focusing on models without fixed effects, column (3) shows that a one basis point QE shock leads to a statistically significant decrease of 7 percentage points in book value exposure for the average bank. The interaction with bank size in column (4) reveals a more heterogeneous reaction, with banks at the 25^{th} percentile of the size distribution displaying a rebalancing of -3.52 percentage points, while the median bank reduces its US Treasury exposure by 4.6 percentage points and a bank at the 75^{th} percentile by 6.37 percentage points (Figure 4.8). Since book values partial out price effects as well, we can interpret this as evidence of a negative portfolio rebalancing effects for US Treasuries of *all* maturities.[10] Estimates for book value exposures broken down by maturity in Appendix 4.C show that rebalancing happens at short and especially at medium maturities, consistently with a rebalancig away from the upper segment of the yield curve due to lower yields. The market-value results in US dollars also display negative coefficients, although they are smaller for larger banks as shown in column (8). Since holdings measured in US dollars do not react to exchange rate movements, this suggests that rebalancing effects alone are enough to more than offset price effects, which should *raise* the market value exposure all else being equal.

Table 4.4 displays the results for models in US dollar that include US government bond price index as a control in market value regressions. The left-hand side panel is equivalent to that of Table 4.3 and is repeated for ease of comparison. The last row shows that the coefficients on the bond price index are positive and statistically significant, which reassures us that they are an appropriate control for the price effect of QE. The coefficients on QE shocks on the right-hand side panel then provide an alternative estimate of the portfolio rebalancing effect. They are negative and statistically significant at the 1% level for all models, and they are substantially larger than the book-value based ones, with an average rebalancing of -5.2 percentage points in response to a one basis point QE shock (column (7)). This is consistent with the interpretation of book-value based estimates as a lower bound for portfolio rebalancing effects. Differently from the estimate based on book value exposures, the effect appears slightly *smaller* for larger banks, although the interaction coefficient is small and not very statistically significant. Based on this coefficient, a bank at the 25^{th} percentile by size has a reaction of -7.3 percentage points, while one at the 75^{th} percentile reacts by -5.5 percentage points (Figure 4.8).

Finally, Table 4.5 shows the results for models in euros, which include the exchange

[10]We calculate size effects including the size of the non-interacted coefficient, despite its lack of statistical significance. The estimated portfolio rebalancing effect would be even larger if we treated the non-significant coefficient as a 0.

rate and portfolio rebalancing effect for book value, and all three channels in the case of market value. The book-value results confirm the negative impact of both the exchange rate and the portfolio rebalancing channels on US Treasury exposure. Based on column (4), the size of the joint exchange rate and portfolio rebalancing effects is of -8.18 percentage points for the median bank and -11.25 percentage points for a bank in the 75^{th} percentile by size (Figure 4.9, left-hand side). The coefficients are suitably larger in absolute value than those in Table 4.3, at least for large banks, as one additional channel is included when exposures are measured in euros. Regressions for market values (column (8)) also display larger coefficients in absolute values in euros compared to Table 4.3 and 4.4 along the whole bank size distribution. A one basis point QE shocks leads to a reduction in US Treasury exposure by 10 percentage points for the average bank, 12.4 percentage points for the median bank, and by 10.76 percentage points for a bank at the 75^{th} percentile (Figure 4.9, right-hand side).

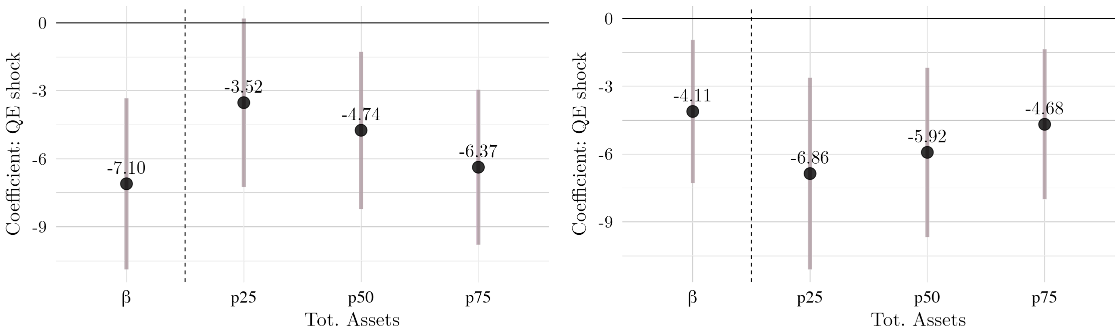
In conclusion, we find that QE reduces the value of US Treasury holdings on European banks' balance sheets. Disentangling the three theoretical channels show that valuation effects have an overall negative impact and that the portfolio rebalancing channel alone offsets the price channel. Jointly, these results imply that the exchange rate channel is stronger. The exposure to US Treasuries falls in response to QE even when shutting down both the price and the exchange rate channels, indicating a rebalancing away from US Treasuries. Larger banks rebalance their portfolio away from US Treasury particularly strongly, and especially so for shorter maturities. The market value results in euros, where all three channels are at play, show that the negative impact of the portfolio rebalancing and exchange rate channel dominate and result in a large drop in the value of US Treasuries on European banks' balance sheets. In turn, the associated capital losses open the door to a potentially sizeable decline in credit provision by banks. In the next section, we turn to estimating directly the effects of QE on bank leverage and lending, providing evidence for spillovers to the real economy through a financial accelerator mechanism.

Table 4.3. Reaction of US Treasury exposure in USD

	Book value				Market value			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE shock$_t^q$	-0.065^{***}	0.041	-0.071^{***}	0.029	-0.028	-0.078^{**}	-0.041^{**}	-0.118^{***}
	(0.020)	(0.036)	(0.019)	(0.034)	(0.017)	(0.034)	(0.016)	(0.040)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$		-0.016^{***}		-0.015^{***}		0.008^{*}		0.012^{**}
		(0.005)		(0.005)		(0.004)		(0.005)
log(Tot. assets)$_t^{hy}$	-0.032	-0.041	0.141^{***}	0.131^{***}	-0.045	-0.040	0.251^{***}	0.258^{***}
	(0.087)	(0.076)	(0.030)	(0.029)	(0.054)	(0.056)	(0.038)	(0.038)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	536	536	536	536	536	536	536	536
N Clusters	124	124	124	124	124	124	124	124

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.7. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.3 **(b)** Column (7) and (8) of Table 4.3

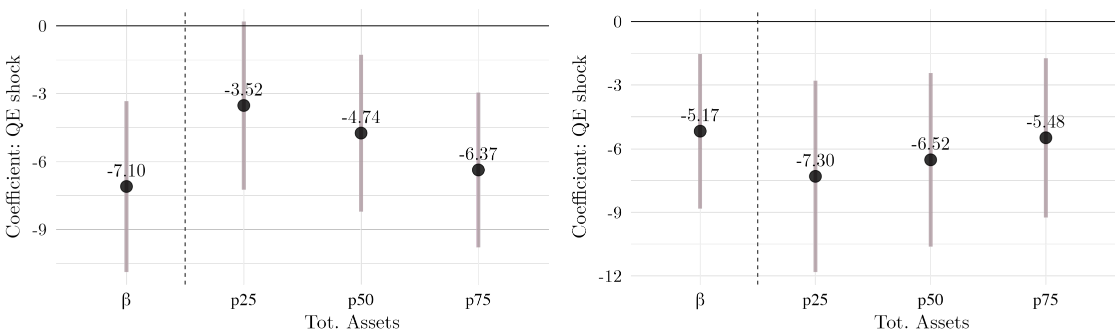
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.4. Reaction of US Treasury exposure in USD - including a price index

	Book value				Market value			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE shock$_t^q$	-0.065^{***}	0.041	-0.071^{***}	0.029	-0.067^{**}	-0.092^{**}	-0.052^{***}	-0.114^{***}
	(0.020)	(0.036)	(0.019)	(0.034)	(0.025)	(0.036)	(0.019)	(0.040)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$		-0.016^{***}		-0.015^{***}		0.004		0.010^{*}
		(0.005)		(0.005)		(0.004)		(0.005)
log(Tot. assets)$_t^{hy}$	-0.032	-0.041	0.141^{***}	0.131^{***}	-0.037	-0.035	0.253^{***}	0.259^{***}
	(0.087)	(0.076)	(0.030)	(0.029)	(0.055)	(0.056)	(0.038)	(0.038)
Bond Price Index$_t^{US,q}$					0.080^{***}	0.076^{***}	0.043^{**}	0.037^{**}
					(0.024)	(0.023)	(0.018)	(0.018)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	536	536	536	536	536	536	536	536
N Clusters	124	124	124	124	124	124	124	124

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.8. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.4

(b) Column (7) and (8) of Table 4.4

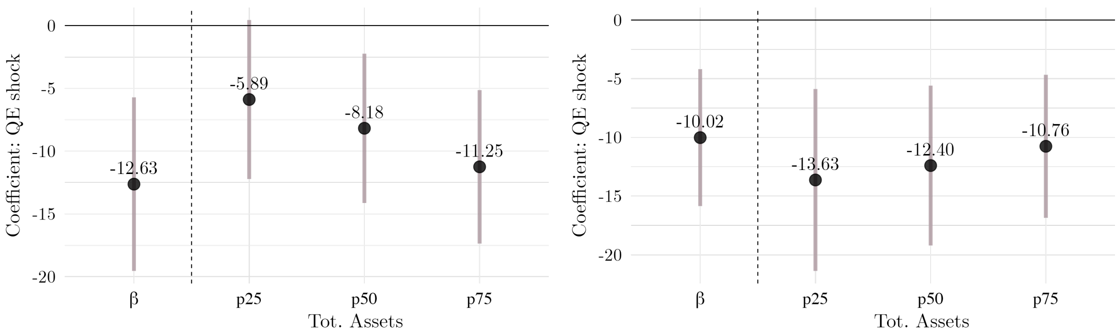
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.5. Reaction of US Treasury exposure in EUR

	Book value				Market value			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE shock$_t^q$	-0.111^{***}	0.083	-0.126^{***}	0.062	-0.072^{**}	-0.131^{**}	-0.100^{***}	-0.201^{**}
	(0.036)	(0.067)	(0.035)	(0.063)	(0.031)	(0.064)	(0.030)	(0.077)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$		-0.030^{***}		-0.029^{***}		0.009		0.015
		(0.010)		(0.011)		(0.008)		(0.011)
log(Tot. assets)$_t^{hy}$	-0.019	-0.035	0.238^{***}	0.219^{***}	-0.091	-0.086	0.455^{***}	0.465^{***}
	(0.161)	(0.143)	(0.056)	(0.052)	(0.094)	(0.095)	(0.084)	(0.082)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	536	536	536	536	536	536	536	536
N Clusters	124	124	124	124	124	124	124	124

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.9. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.5 **(b)** Column (7) and (8) of Table 4.5

Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

4.5.2 Effects on net worth and lending

For the following estimations we use an instrumental variable (IV) approach, where we use the QE shocks as the instrument for one of three financial variables θ_t^q. We aim to isolate the direct effect of QE on lending through US Treasuries by instrumenting the relevant spreads for price and exchange rate valuation effects. For the price effect, we use changes in the 10-year US Treasury yield, which is targeted by

QE and directly linked to the price of long-term US government bonds.

Quantifying the exchange rate effect on US Treasuries alone is more challenging, because a QE-induced US dollar depreciation results in equal losses on all US dollar-denominated assets on the bank's balance sheet, abstracting from hedging. We attempt to gauge the impact of exchange rate valuation effects on net worth and credit through US Treasuries by instrumenting changes in the international spread between US and German 10-year government bonds with QE shocks. QE causes the spread to shrink through a reduction in the yield of long-term US Treasuries. In turn, the interest rate differential between US and German government bonds is linked to the exchange rate through the UIP condition, with a lower US rate associated to a depreciation.[11] Given the tight theoretical link between the exchange rate and the spread between US Treasuries and German government bonds, we instrument the latter with the QE shock to to approximate the impact of the QE-induced US dollar depreciation on bank leverage and lending through US Treasuries alone. We then estimate the overall impact of exchange rate valuation effects through any US dollar denominated asset and liability on banks' balance sheets by instrumenting changes in the exchange rate directly rather than the US-DE spread.

We summarise here the reaction of the three instrumented variables θ_t^q to QE shocks, shown in in Figure 4.6. The US term spread drops ($\eta_t^q \uparrow \implies \Delta\text{US 10y-3m}_t^q < 0$), the international spread decreases ($\eta_t^q \uparrow \implies \Delta\text{US-DE 10y}_t^q < 0$), and the US dollar depreciates ($\eta_t^q \uparrow \implies \Delta\text{EUR/USD} < 0$). Note that a QE shock is associated with a *fall* in the US term spread and in the US-DE international spread. Therefore, for consistency with the rest of the paper, in this section we present coefficients for the *negative* of the US term spread, $-\Delta(\text{US 10y-3m})_t^q$, and for the international spread, $-\Delta(\text{US-DE 10y})_t^q$. Thus, we can interpret an increase in the fitted values of the relevant instrumented variable as a QE shock in all three cases.

For the IV estimation (Equation (4.6)), we replace the shock series η_t^q in Equation (4.3) with the fitted values $\widehat{\theta_t^q}$ from the first-stage regression of θ_t^q on η_t^q and controls, where θ_t^q = -(US 10y-3m)$_t^q$, -(US-DE 10y)$_t^q$ or Δlog(EUR/USD).[12]

[11] Note that even in models where the UIP condition does not hold exactly because of wedges introduced by financial frictions or segmented markets, a decrease in the interest rate is still associated with a depreciation *ceteris paribus*. This is the case in much of the recent literature on exchange rate determination in financial markets, such as Gabaix and Maggiori (2015), Itskhoki and Mukhin (2021), Jiang et al. (2024), and Jiang et al. (2021).

[12] Table 4.17 in Appendix 4.D shows the first-stage regression results, demonstrating how all three instruments are very strong with F-stats well over any thresholds suggested by the literature

$$y_{i,t}^{hy} = \alpha + \beta\widehat{\theta}_t^q + \gamma\widehat{\theta}_t^q \cdot BS_{i,t}^{hy} + \delta BS_{i,t}^{hy} + \Gamma_1 X_{t-1}^q + \Gamma_2 W_t^q + FE_i + \varepsilon_{i,t} \tag{4.6}$$

As the change in the 10-year US yield is directly targeted by Fed QE interventions, the price of long-term US Treasuries is positively affected. Therefore, we can expect a positive effect on net worth and credit through the price effect of US Treasuries. On the contrary, the US dollar depreciation associated with QE is expected to have a contrasting impact on banks, resulting in a drop in their net worth and credit.

To analyse the overall net effect of a QE shock on the banks' balance sheet, we use the QE shock directly as an explanatory variable. This allows us to gauge the net impact of all channels at play, after controlling for the effects of QE on macroeconomic conditions in Europe.

We expect a negative overall effect because Section 4.5.1 showed that the exchange rate effect dominates the price effect, engendering capital losses on US Treasuries. However, other effects of QE on the balance sheet of banks might offset valuation effects on the US Treasury portfolio. By allowing all balance-sheet avenues to affect bank net worth and credit in this specification, we can better understand whether the US Treasury channel is quantitatively important.

Leverage

Tables 4.6, 4.7, and 4.8 show the results for banks' net worth, proxied by their leverage ratio. The leverage ratio is used because it is consistent with the theoretical framework of a financial accelerator model where banks operate under a leverage constraint.[13] Therefore, any capital gains through their portfolio of sovereign holdings should result in an expansion in lending, while losses should lead to a contraction. In Appendix 4.H.1, we run a robustness check, where we proxy net worth as the regulatory Tier 1 capital ratio.

Table 4.6 shows that a QE induced decrease in the 10-year US yield, i.e. a higher US dollar price of US Treasuries, has indeed a positive effect on banks' net worth. On average the banks net worth increases by 13 percentage points (column (3)) for a one basis point decrease in US $10y_t^q$. The interaction with bank size reveals that banks at the 25^{th} percentile react more strongly (32 percentage points increase),

(Stock and Yogo, 2005).

[13] Note that here we use the regulatory definition of leverage ratio as Tier 1 capital/assets, such that an increase corresponds to a higher capital buffer, and thus to a *decrease* in leverage in its corporate finance definition.

while banks at the 75^{th} percentile have a slightly weaker reaction of 28 percentage points (Figure 4.10).

Table 4.7 displays estimates of the exchange rate effect through US Treasuries, proxied by the spread between US and German 10-year government bonds. A decrease in US-DE $10y_t^q$ appears to have a significant impact on net worth, with a one basis point change associated with a 5 percentage point fall in the leverage ratio on average (column (3)). Contrary to the findings for the 10-year US rate where small and large banks reacted in the same direction, larger banks seem to react in the opposite direction to exchange rate valuation effects on US Treasuries (column (4)). A bank at the 25^{th} percentile by size experiences an *decrease* in the leverage ratio by 1.2 percentage point upon a 1 basis point decrease in the international spread, while a bank at the 75^{th} percentile sees its leverage ratio rise by 2 percentage points (Figure 4.11). This might be due to a higher hedge ratio for larger banks, who are on the other hand more exposed to Treasuries than smaller banks in both absolute and relative terms. However, this is only a speculative hypothesis absent data on banks' hedging practices.

We can observe in Table 4.8 that the overall effect of the Fed's QE policy results in a serious hit to to the net worth of European banks. This decline is in line with the notion of capital losses on their US Treasury holdings, primarily attributable to the depreciation of the US dollar. A one basis point QE shock leads to a very large decrease of 32 percentage point in the leverage ratio (column (3)). Again as for the price effect, the overall impact of QE on net worth is attenuated for larger banks. A one basis point QE shock leads to a 56 percentage point reduction in the leverage ratio for a bank at the 25^{th} percentile by size, while it decreases by just 13 percentage points for a bank at the 75^{th} percentile (Figure 4.12).

In conclusion, the separate estimates of the price and exchange rate effects on banks' net worth are consistent with the prediction of a financial accelerator framework: the former being positive and the latter negative. Overall, the Federal Reserve's QE has a negative effect on European banks' net worth that is consistent with the valuation effect due to US dollar depreciation, operating through both US Treasuries and other US dollar-denominated assets on their balance sheets. Using the formalism introduced in Section 4.4, the estimates of the exchange rate effect, combined with those of the overall effects of QE on net worth are consistent with $M_t > 0$. Both the price effect on US Treasuries and the overall impact of a QE shock appear less intense for larger banks. On the other hand, the exchange rate valuation effect seems to go in the opposite direction. Appendix 4.F.1 shows that the reaction of net worth does not appear to depend on the bank's exposure to US

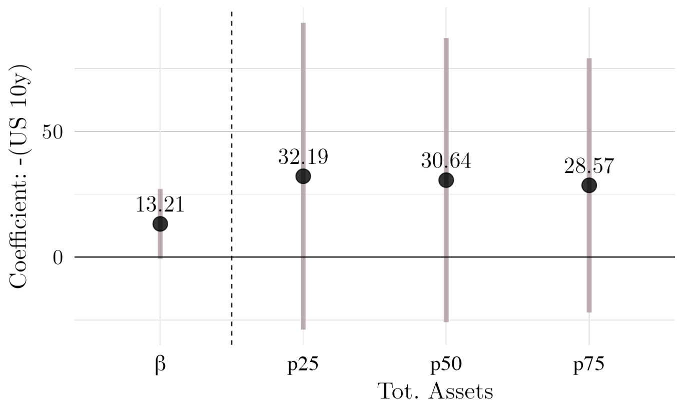
Treasuries in the previous semester.

Table 4.6. Reaction of net worth through US 10-year rate

	(1)	(2)	(3)	(4)
$-(\Delta \text{US10y})^q_t$	0.018	0.036	0.132^{*}	0.404^{*}
	(0.015)	(0.025)	(0.071)	(0.436)
$-(\Delta \text{US10y})^q_t \times \log(\text{Tot. assets})_{i,t}$		-0.006		-0.020^{*}
		(0.005)		(0.031)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	749	749	749
N Clusters	128	128	128	128

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.10. Column (3) and (4) of Table 4.6



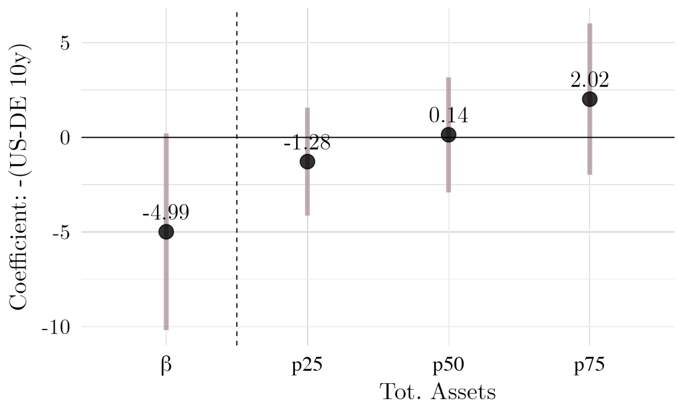
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.7. Reaction of net worth through US-DE 10-year spread

	(1)	(2)	(3)	(4)
-(ΔUS-DE 10y)$_t^q$	-0.012	0.161	-0.050^{*}	-0.087^{**}
	(0.010)	(0.114)	(0.027)	(0.039)
-(ΔUS-DE 10y)$_t^q$ $\times$ log(Tot. assets)$_{i,t}$		-0.027		0.018^{**}
		(0.020)		(0.008)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	749	749	749
N Clusters	128	128	128	128

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.11. Column (3) and (4) of Table 4.7



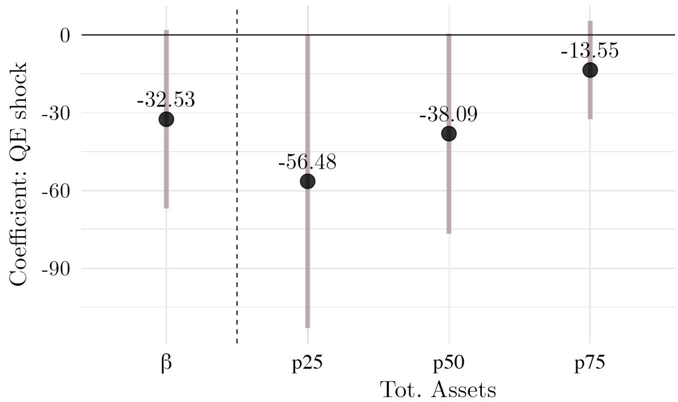
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.8. Overall effect of QE on net worth

	(1)	(2)	(3)	(4)
QE shock$_t^q$	-0.060	-0.376	-0.325^*	-1.533^*
	(0.051)	(0.276)	(0.176)	(0.798)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$		0.061		0.232^*
		(0.045)		(0.124)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	749	749	749
N Clusters	128	128	128	128

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.12. Column (3) and (4) of Table 4.8



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Lending

The final, and most important, dependent variable centers on credit provision by European banks. Our goal is to investigate the transmission of QE through US

Treasuries on European banks' balance sheets to the real economy. As discussed above, based on the theoretical framework of a financial accelerator model, we would expect that credit reacts in the same direction as net worth. Tables 4.9, 4.10, and 4.11 present the findings from our estimations of spread changes, which are instrumented by a QE shock.

In contrast to what we observed in Table 4.6, we show in Table 4.9 that upon a one basis point decrease in the 10-year US Treasury yield caused by QE, total credit decreases by 0.2 percentage points for the average bank (column (3)). This corresponds to €11 billions when multiplied by aggregate credit for all banks, averaged across all semesters. Whereas the response is positive for small banks (increase of 0.15 percentage points at the 25^{th} percentile in figure 4.13), it is negative for large banks (decrease of 0.2 percentage points at the 75^{th} percentile in figure 4.13). These results are consistent with larger banks displaying a disproportionately stronger rebalancing away from Treasuries in response to QE, more than offsetting the positive valuation effects and resulting in a contemporaneous hit to their net worth. Similarly, exposures to Treasuries for small banks might not be substantial enough to translate changes of their dollar price into changes in credit provision, so the weak and not very statistically significant negative effect we estimate might be also be due to rebalancing effects.

Table 4.10 shows the effect of QE shocks on credit through the exchange rate channel as mediated by the US-DE 10 year spread. The average bank cuts credit by 0.2 percentage points in response to a one basis point decrease in the US-DE 10-year spread caused by QE. The credit contraction adds up to €8.7 billion when multiplied by average aggregate credit. The effect seems to be stronger for big banks: banks at the 75^{th} percentile react with a 0.2 percentage point decrease of credit provision, whereas the median reaction is estimated to be at -0.12 percentage points. (Figure 4.14).

The total effect of QE, conditional on our macroeconomic and financi al controls, is shown in Table 4.11. Supporting the theoretical framework, the effect is negative, with an average reduction in credit of 1.2 percentage points in response to a one basis point QE shock. The effect is quantitatively significant, corresponding to an aggregate drop in credit by €21 billions. We find once again an attenuated impact on larger banks, with an contraction in credit by 2.9 percentage points at the 25^{th} percentile by size, and by 0.5 percentage points at the 75^{th} percentile (Figure 4.15).

The higher dollar prices of US Treasuries caused by a Fed QE shock leads to European banks extending more credit *ceteris paribus*. Taken separately, the estimates from Tables 4.9 and 4.10 would indicate that the average impact of the exchange rate channel through US Treasuries is smaller than that of the price channel. However, when estimating a model that uses QE shocks directly as an explanatory variable, the overall effect on total credit is negative. The finding is consistent with a large exchange rate effect, potentially due to the total balance sheet exposure to the USD/EUR rate, that more than offsets the positive effect of higher US Treasury prices. We attempt to corroborate this hypothesis by running the same regression as in Table 4.10, but instrumenting the EUR/USD exchange rate rather than the international spread. The results, shown in Section 4.E.2 show that the negative impact of QE on credit through the exchange rate at-large is indeed larger than the price effect, both on average and along the bank size distribution.

Overall, the results for credit are coherent with the drop in net worth documented in the previous section, as banks face a leverage constraint. Larger banks seem to display a muted response of credit to QE shock, as we found for net worth in the previous section. This heterogeneity seems to suggest that large banks are better equipped to cushion the effects of the Fed's QE, possibly through hedging.

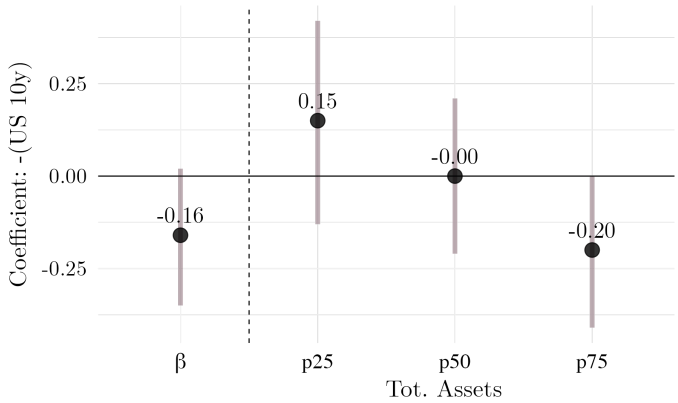
We carry out some extensions on the analysis of credit in the appendices. In Appendix 4.F.2 that lagged exposure to US Treasuries is not associated with a particularly strong or weak reaction of credit either. In Appendix 4.G, we estimate separate models for credit extended domestically and in the USA, finding that the overall response is driven by domestic credit.

Table 4.9. Reaction of credit through US 10-year rate

	(1)	(2)	(3)	(4)
-$(\Delta \text{US10y})^q_t$	-0.003^{***}	0.008^{***}	-0.002^{*}	0.009^{**}
	(0.001)	(0.003)	(0.001)	(0.004)
-$(\Delta \text{US10y})^q_t \times \log(\text{Tot. assets})_{i,t}$		-0.002^{***}		-0.002^{*}
		(0.001)		(0.001)
$\log(\text{Tot. assets})^{hy}_t$	0.563^{***}	0.138	0.970^{***}	0.510^{***}
	(0.098)	(0.099)	(0.035)	(0.168)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	946	946	946
N Clusters	129	129	129	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.13. Column (3) and (4) of Table 4.9



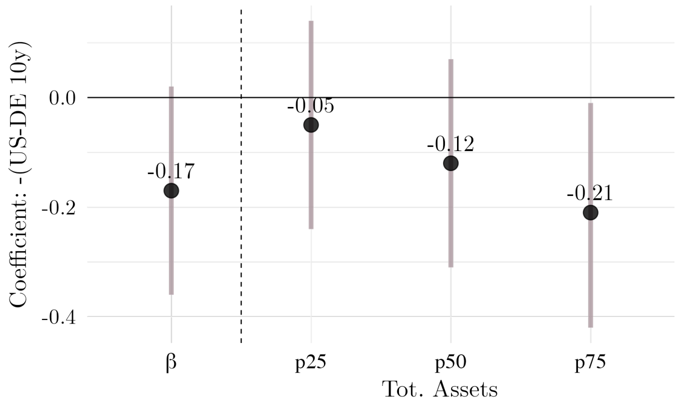
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.10. Reaction of credit through US-DE 10-year spread

	(1)	(2)	(3)	(4)
-(ΔUS-DE 10y)$_t^q$	-0.004^{***}	0.002	-0.002^{*}	0.003^{**}
	(0.001)	(0.001)	(0.001)	(0.002)
-(ΔUS-DE 10y)$_t^q$ × log(Tot. assets)$_{i,t}$		-0.001^{***}		-0.001^{***}
		(0.000)		(0.000)
log(Tot. assets)$_t^{hy}$	0.555^{***}	0.428^{***}	0.971^{***}	0.795^{***}
	(0.115)	(0.105)	(0.035)	(0.059)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	946	946	946
N Clusters	129	129	129	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.14. Column (3) and (4) of Table 4.10



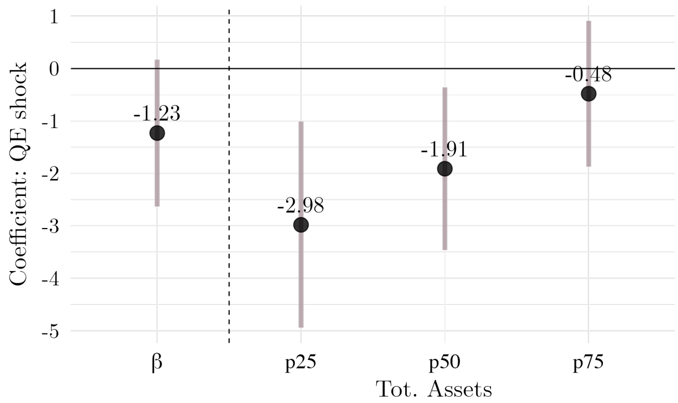
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

Table 4.11. Overall effect of QE on credit

	(1)	(2)	(3)	(4)
QE shock$_t^q$	-0.025^{***}	-0.087^{***}	-0.012^{*}	-0.086^{***}
	(0.005)	(0.018)	(0.007)	(0.027)
QE shock$_t^q$ $\times$ log(Tot. assets)$_{i,t}$		0.011^{***}		0.013^{***}
		(0.003)		(0.004)
log(Tot. assets)$_t^{hy}$	0.561^{***}	0.583^{***}	0.969^{***}	0.975^{***}
	(0.106)	(0.109)	(0.035)	(0.035)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	946	946	946
N Clusters	129	129	129	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3-months* and *10-years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.15. Column (3) and (4) of Table 4.11



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval, which is calculated based on Equation (4.5).

4.6 Conclusion

This paper puts forward a new mechanism through which unconventional monetary policy by the Fed can spill over to foreign economies. Differently from previous literature, we focus on the direct impact through holdings of the main targets of large-scale asset purchase programmes: US Treasuries.

We show that European banks'exposure to US Treasuries, while small in the context of their overall balance sheet, is on average a large enough fraction of their net worth to affect credit through valuation effects in a financial accelerator framework. Furthermore, exposures are concentrated in longer maturities that are more sensitive to price valuation effects.

The interplay of the price and exchange rate valuation effects on Treasuries triggered by QE is particularly interesting. The former are positive, implying a relaxing of leverage constraints and an expansion of credit. On the other hand, the dollar depreciation following QE produces a negative valuation effect, which should tighten constraints and lower credit provision. Our results suggest that the exchange rate effect dominates, carrying the implication, novel in the literature, that QE might have a *contractionary* effect through Treasury exposure. The presence of a negative transmission channel for quantitative easing raises the intriguing possibility that Treasury holdings may have a mitigating effect on the adverse spillovers of quantiative tightening, funcitoning in this respect as a hedging device.

We document heterogeneity in the reaction of Treasury holdings, net worth and credit to QE shocks along the bank size dimension. This facet is particularly important because large banks are simultaneously more exposed to Treasuries, and responsible for a larger share of credit provision. Therefore, their behaviour carries a disproportionately large weight on the spillovers of balance sheet effects to the real economy.

According to our results on Treasury exposures, large banks appear to both suffer larger negative valuation effects, consistent with their larger Treasury exposure, and rebalance their portfolios away from US Treasury more strongly than smaller banks. Likewise, their credit provision declines more intensely through valuation effects. On the other hand, the overall negative impact of QE shocks on credit is decreasing in bank size, suggesting that large banks nevertheless manage to cushion the blow, possibly through unobserved hedging strategies, or contemporaneous gains from other assets on their balance sheet. We observe the same relationship with bank size for the overall reaction of net worth too, consistent with the financial accelerator mechanism. However, net worth seems to actually increase for large banks in response to exchange rate valuation effects mediated by QE, again hinting at the possible involvement of hedging.

This studiy leaves some avenues open for future research. First, better data on banks' forex risk hedging might help in quantifying the exchange rate effect. Second, an investigation of European banks' overall exposure to dollar-denominated assets and liabilities could help single out the valuation effects on Treasuries, and quantify their importance relative to those on the rest of the balance sheet. Finally, the general-equilibrium implications of the Treasury channel could be studied in a open-economy financial accelerator model, possibly with heterogeneity in bank size to account for the asymmetric effects that we document.

Appendix

4.A Sources of data

Table 4.12. Data sources

Data	**Source**
Bank balance sheet data	EBA Transparency Exercise and Stress Test databases
Bank total assets	Capital IQ
Aggregate banking sector assets	ECB Balance Sheet Items database
QE shock series	Jarocinski (2021)
ECB interest rate shock series	Kearns et al. (2022)
Government bond indices and yields	Refinitiv Eikon
ECB main refinancing rate	ECB
Short- and long-term eurozone borrowing rates	ECB
EUR/USD exchange rates	Refinitiv Eikon
Unemployment rate	Datastream
CPI inflation	Datastream
Real GDP growth	Datastream
VIX	Federal Reserve Economic Data (FRED)

4.B Further descriptive analysis

4.B.1 Summary statistics for additional bank-level variables

Variable	Obs.	Mean	SD	Min	Median	Max
Tier1 capital / RWA (%)	1019	17.27	11.04	0.56	14.5	120.87
U.S.Treas exp ($\leq$ 1y)	720	0.4	1.58	0	0	18.17
U.S.Treas exp (2-10y)	776	1.1	3.47	0	0	33.44
U.S.Treas exp ($>$ 10y)	776	0.59	2.54	0	0	31.97
U.S. credit	548	7.34	13.49	0	1.93	66.46
Domestic credit	928	31.76	44.9	0	16.14	343.41

Notes: The table shows summary statistics for the estimation sample. *RWA* stands for risk-weighted assets. The abbreviation *US Treas exp* stands for US Treasuries exposure for the respective maturity buckets. All values except the *Tier1 capital / RWA* are in Euro billions

4.B.2 Descriptive statistics of financial variables

Variable	Obs.	Mean	SD	Min	Median	Max
US 10y yield	1076	2.35	0.39	1.69	2.27	3.03
US 3m yield	1076	0.84	0.93	0.01	0.26	2.35
US-DE 10y yield	1076	1.68	0.75	0.04	1.63	2.67
QE shock (Jarociński (2021))	1076	0.58	1.72	-3.05	0.27	3.95
log(EUR/USD)	1076	0.17	0.08	0.08	0.13	0.31

Notes: The table shows descriptive statistics for the financial market variables over the sample period. All variables are quarterly aggregates. The *QE shock* (Jarocinski, 2021) is in basis point shocks, the EUR/USD exchange rate is in log(), and all other values are in percentage points.

4.B.3 Banks by country

Table 4.13. Number of banks per country and quartile

Country	Q1	Q2	Q3	Q4	Total
Germany	3	12	6	6	27
Italy	5	5	7	2	19
Spain	1	2	7	4	14
France	2	3	2	6	13
Austria	6	3	0	0	9
Ireland	3	2	3	0	8
Netherlands	1	1	2	4	8
Belgium	3	2	1	1	7
Portugal	3	2	2	0	7
Luxembourg	5	0	1	0	6
Cyprus	4	0	0	0	4
Finland	2	1	0	1	4
Greece	0	1	3	0	4
Malta	4	0	0	0	4
Slovenia	4	0	0	0	4
Estonia	2	0	0	0	2
Latvia	2	0	0	0	2
Lithuania	1	0	0	0	1
Total	51	34	34	24	143

Notes: This table summarizes the number of banks in each quartile and for each country. The total number (143) does not add up to 152, since not all banks reported their total assets. In addition, the bank size can change over time, implying that banks can change their quartile.

Table 4.14. Number of observations per country and quartile

Country	Q1	Q2	Q3	Q4	Total
Germany	34	84	55	87	260
Italy	25	39	55	30	149
Spain	11	18	64	51	144
France	19	27	9	87	142
Netherlands	3	15	25	45	88
Austria	56	29	0	0	85
Belgium	38	25	2	15	80
Ireland	11	30	23	0	64
Portugal	15	28	12	0	55
Luxembourg	40	0	11	0	51
Greece	0	12	36	0	48
Finland	16	15	0	8	39
Cyprus	37	0	0	0	37
Malta	36	0	0	0	36
Slovenia	34	0	0	0	34
Estonia	13	0	0	0	13
Latvia	11	0	0	0	11
Lithuania	8	0	0	0	8
Total	407	322	292	323	1344

Notes: This table summarizes the number of observations in each quartile and each country. The total number of observations reflects the number of bank-quarter observations.

4.C Portfolio rebalancing across maturities

Table 4.15. Reaction of Treasury exposure across maturities - Book value

	Short		Medium		Long	
	(1)	(2)	(3)	(4)	(5)	(6)
QE shock$_t^q$	0.058	0.041	0.128**	0.118*	−0.042	−0.037
	(0.050)	(0.051)	(0.061)	(0.059)	(0.045)	(0.042)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$	−0.014	−0.013	−0.029***	−0.030***	−0.005	−0.007
	(0.008)	(0.009)	(0.010)	(0.010)	(0.005)	(0.006)
log(Tot. assets)$_t^{hy}$	0.065	0.059***	−0.020	0.137***	0.026	0.125***
	(0.083)	(0.021)	(0.096)	(0.037)	(0.107)	(0.038)
Bank FE	Yes	No	Yes	No	Yes	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	489	489	534	534	534	534
N Clusters	124	124	124	124	124	124

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Table 4.16. Reaction of Treasury exposure across maturities - Market value

	Short		Medium		Long	
	(1)	(2)	(3)	(4)	(5)	(6)
QE shock$_t^q$	−0.065*	−0.104**	−0.087	−0.133**	−0.104	−0.113
	(0.038)	(0.039)	(0.055)	(0.064)	(0.065)	(0.072)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$	0.004	0.008	0.008	0.011	0.005	0.005
	(0.005)	(0.005)	(0.007)	(0.008)	(0.008)	(0.010)
log(Tot. assets)$_t^{hy}$	−0.028	0.228***	−0.066	0.322***	−0.068	0.262***
	(0.091)	(0.062)	(0.077)	(0.067)	(0.061)	(0.066)
Bank FE	Yes	No	Yes	No	Yes	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	489	489	534	534	534	534
N Clusters	124	124	124	124	124	124

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

4.D First stage analysis for IV regressions

Table 4.17. First-stage regression results

	Net Worth			Total credit		
-(US 10y)	7.63			7.56		
	(0.35)			(0.34)		
-(US-DE 10y)		7.51			7.39	
		(0.24)			(0.24)	
log(EUR/USD)			-2			-2.01
			(0.05)			(0.05)
F-stat	2064.62	11580.53	778.71	1992.13	10306.82	779.3
Num. obs.	946	946	946	946	946	946

Notes: This table shows the first stage estimation of the Jarocinski (2021) QE shocks on the three instrumented variables that we use in the the regressions. The bottom row shows the first-stage F-statistic. The results differ for net worth and total credit, as the sample for which these variables are observed is not exactly the same.

4.E Effects on net worth and credit: overall impact of the exchange rate

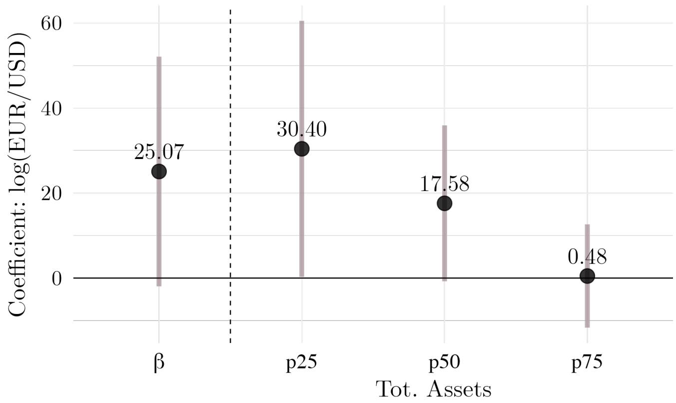
4.E.1 Net worth

Table 4.18. Reaction of net worth through EUR/USD exchange rate

	(1)	(2)	(3)	(4)
ΔEURUSD_t^q	0.042	0.590	0.251*	0.979*
	(0.036)	(0.494)	(0.138)	(0.510)
$\Delta\text{EURUSD}_t^q \times \log(\text{Tot. assets})_{i,t}$		−0.096		−0.161*
		(0.081)		(0.087)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	749	749	749
N Clusters	128	128	128	128

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.16. Column (3) and (4) of Table 4.18



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

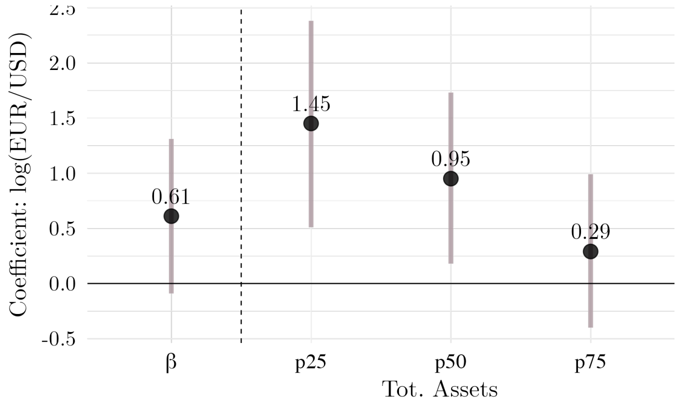
4.E.2 Credit

Table 4.19. Reaction of credit through EUR/USD exchange rate

	(1)	(2)	(3)	(4)
ΔEURUSD_t^q	0.012^{***}	0.045^{***}	0.006^{*}	0.041^{***}
	(0.002)	(0.009)	(0.004)	(0.011)
$\Delta\text{EURUSD}_t^q \times \log(\text{Tot. assets})_{i,t}$		-0.006^{***}		-0.006^{**}
		(0.001)		(0.002)
$\log(\text{Tot. assets})_t^{hy}$	0.566^{***}	0.691^{***}	0.968^{***}	1.059^{***}
	(0.104)	(0.124)	(0.035)	(0.045)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	946	946	946
N Clusters	129	129	129	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.17. Column (3) and (4) of Table 4.19



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

4.F Effects on net worth and credit: interaction with Treasury exposure

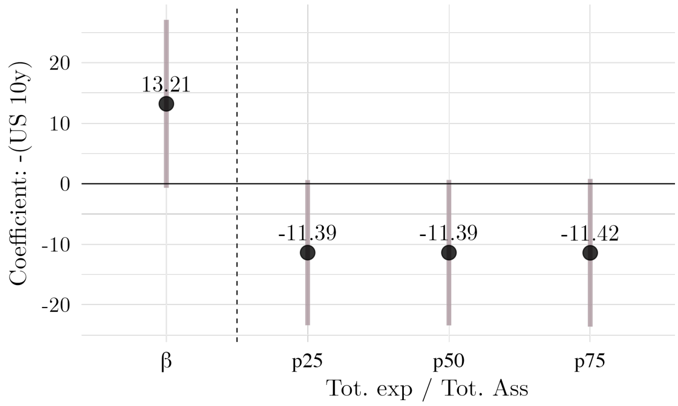
4.F.1 Net worth

Table 4.20. Reaction of net worth through US 10-year rate - Treasury exposure interaction

	(1)	(2)	(3)	(4)
$-(\Delta\text{US10y})_t^q$	0.018	-0.011	0.132^{*}	-0.114^{*}
	(0.015)	(0.010)	(0.071)	(0.061)
$-(\Delta\text{US10y})_t^q \times \text{Exp./assets}_{i,t-1}$		-0.001		-0.000
		(0.002)		(0.004)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	564	749	564
N Clusters	128	125	128	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.18. Column (3) and (4) of Table 4.20



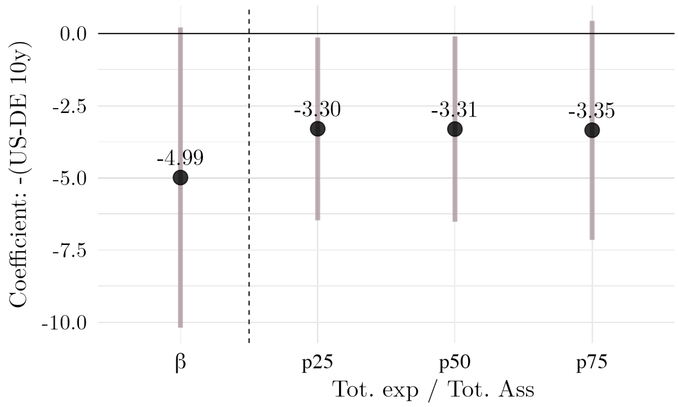
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.21. Reaction of net worth through US-DE 10-year spread - Treasury exposure interaction

	(1)	(2)	(3)	(4)
-(ΔUS-DE 10y)$_t^q$	-0.012	-0.001	-0.050^{*}	-0.033^{*}
	(0.030)	(0.013)	(0.02)	(0.016)
-(ΔUS-DE 10y)$_t^q$ $\times$ Exp./assets$_{i,t-1}$		-0.002		-0.001
		(0.009)		(0.017)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	564	749	564
N Clusters	128	125	128	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.19. Column (3) and (4) of Table 4.21



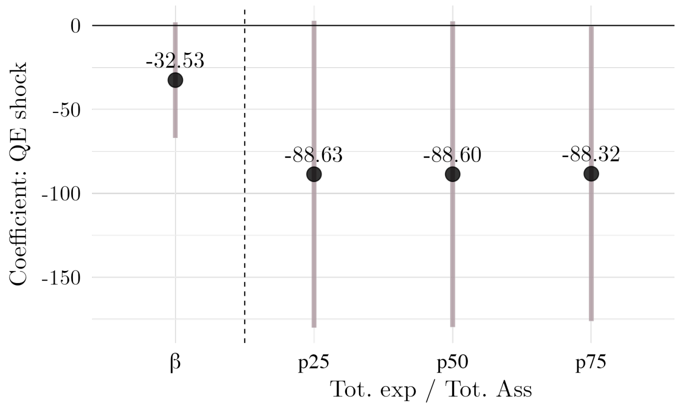
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.22. Overall effect of QE on net worth - Treasury exposure interaction

	(1)	(2)	(3)	(4)
QE shock^q_t	-0.060	-0.109	-0.325^*	-0.886^*
	(0.051)	(0.088)	(0.176)	(0.466)
QE shock^q_t × Exp./assets$_{i,t-1}$		0.011		0.006
		(0.035)		(0.058)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	749	564	749	564
N Clusters	128	125	128	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.20. Column (3) and (4) of Table 4.22



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

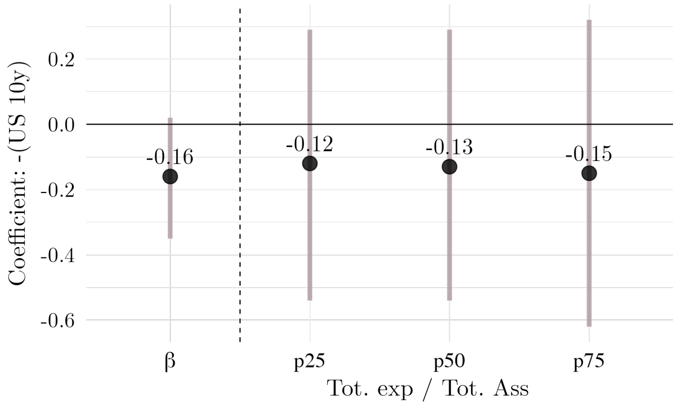
4.F.2 Credit

Table 4.23. Reaction of credit through US 10-year rate - Treasury exposure interaction

	(1)	(2)	(3)	(4)
-$(\Delta \text{US10y})^q_t$	-0.003^{***}	-0.002^{**}	-0.002^{*}	-0.001
	(0.001)	(0.001)	(0.001)	(0.002)
-$(\Delta \text{US10y})^q_t \times$ Exp./assets$_{i,t-1}$		0.000		-0.000
		(0.000)		(0.001)
log(Tot. assets)$^{hy}_t$	0.563^{***}	0.735^{***}	0.970^{***}	0.985^{***}
	(0.111)	(0.124)	(0.035)	(0.047)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	608	946	608
N Clusters	129	125	129	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.21. Column (3) and (4) of Table 4.23



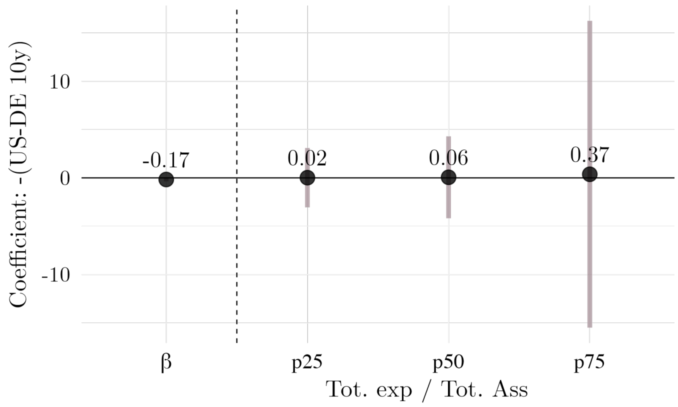
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.24. Reaction of credit through US-DE 10-year spread - Treasury exposure interaction

	(1)	(2)	(3)	(4)
-(ΔUS-DE 10y)$_t^q$	-0.004^{***}	0.000	-0.002^{*}	0.000
	(0.001)	(0.008)	(0.001)	(0.016)
-(ΔUS-DE 10y)$_t^q$ $\times$ Exp./assets$_{i,t-1}$		-0.001		0.007
		(0.008)		(0.129)
log(Tot. assets)$_t^{hy}$	0.555^{***}	0.951	0.971^{***}	1.114
	(0.115)	(1.237)	(0.035)	(2.239)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	608	946	608
N Clusters	129	125	129	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.22. Column (3) and (4) of Table 4.24



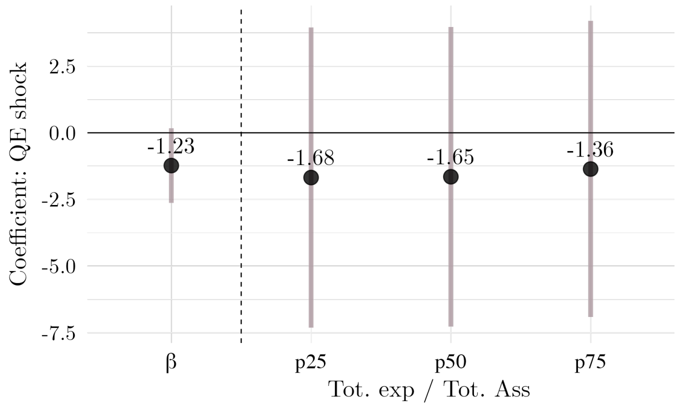
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.25. Overall effect of QE on credit - Treasury exposure interaction

	(1)	(2)	(3)	(4)
QE shock$_t^q$	-0.025^{***}	-0.027^{**}	-0.012^{*}	-0.017
	(0.005)	(0.011)	(0.007)	(0.029)
QE shock$_t^q$ $\times$ Exp./assets$_{i,t-1}$		-0.001		0.006
		(0.003)		(0.005)
log(Tot. assets)$_t^{hy}$	0.561^{***}	0.725^{***}	0.969^{***}	0.993^{***}
	(0.106)	(0.126)	(0.035)	(0.041)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	946	608	946	608
N Clusters	129	125	129	125

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level ($^{***}p < 0.01$; $^{**}p < 0.05$; $^{*}p < 0.1$).

Figure 4.23. Column (3) and (4) of Table 4.25



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

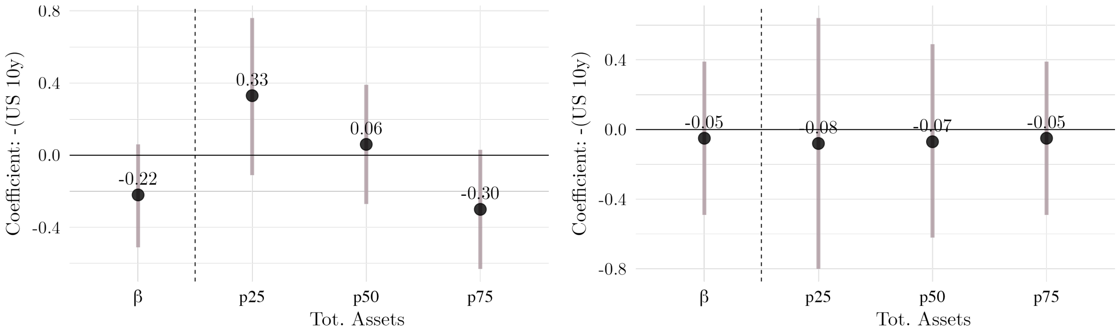
4.G Effects on domestic and US credit

Table 4.26. Reaction of domestic and US credit through US 10-year rate

	Domestic				US			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
-$(\Delta \text{US10y})_t^q$	-0.004^{***}	0.008^{*}	-0.002	0.017^{**}	0.001	0.007^{*}	-0.001	-0.001
	(0.001)	(0.004)	(0.001)	(0.007)	(0.001)	(0.004)	(0.002)	(0.009)
-$(\Delta \text{US10y})_t^q \times \log(\text{Tot. assets})_{i,t}$		-0.002^{*}		-0.003^{***}		-0.001		0.000
		(0.001)		(0.001)		(0.001)		(0.001)
$\log(\text{Tot. assets})_t^{hy}$	0.379^{***}	-0.067	0.914^{***}	0.099	0.402^{**}	0.176	0.695^{***}	0.724^{*}
	(0.118)	(0.202)	(0.048)	(0.273)	(0.141)	(0.183)	(0.075)	(0.359)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	854	854	854	854	515	515	515	515
N Clusters	123	123	123	123	69	69	69	69

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.24. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.26 **(b)** Column (7) and (8) of Table 4.26

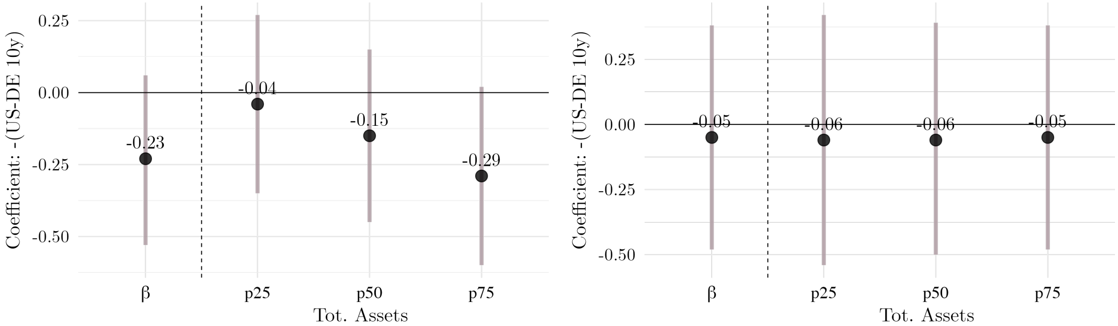
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.27. Reaction of domestic and US credit through US-DE 10-year spread

	Domestic				US			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
-(ΔUS-DE 10y)$_t^q$	-0.005^{***}	0.000	-0.002	0.005^{**}	0.002	0.004^{*}	-0.001	-0.001
	(0.001)	(0.002)	(0.002)	(0.003)	(0.001)	(0.002)	(0.002)	(0.004)
-(ΔUS-DE 10y)$_t^q$ $\times$ log(Tot. assets)$_{i,t}$		-0.001^{**}		-0.001^{***}		-0.000		0.000
		(0.000)		(0.000)		(0.000)		(0.001)
log(Tot. assets)$_t^{hy}$	0.366^{**}	0.259^{**}	0.915^{***}	0.652^{***}	0.403^{**}	0.344^{**}	0.697^{***}	0.707^{***}
	(0.130)	(0.122)	(0.048)	(0.081)	(0.138)	(0.146)	(0.075)	(0.141)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	854	854	854	854	515	515	515	515
N Clusters	123	123	123	123	69	69	69	69

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.25. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.27 **(b)** Column (7) and (8) of Table 4.27

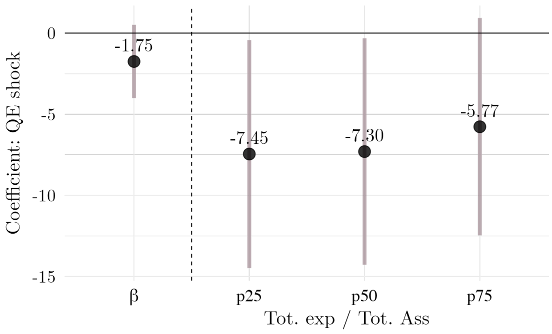
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.28. Overall effect of QE on domestic and US credit

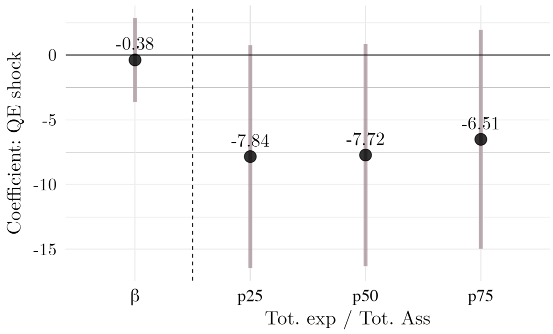
	Domestic				US			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE shock$_t^q$	-0.031^{***}	-0.053^{**}	-0.017	-0.075^{**}	0.010	-0.050^{***}	-0.004	-0.078^{*}
	(0.009)	(0.026)	(0.012)	(0.036)	(0.008)	(0.017)	(0.017)	(0.044)
QE shock$_t^q$ × Exp./assets$_{i,t-1}$		0.009		0.033^{*}		0.001		0.026
		(0.006)		(0.017)		(0.002)		(0.015)
log(Tot. assets)$_t^{hy}$	0.376^{***}	0.680^{***}	0.912^{***}	0.953^{***}	0.406^{**}	0.229	0.696^{***}	0.744^{***}
	(0.117)	(0.193)	(0.048)	(0.058)	(0.138)	(0.123)	(0.075)	(0.077)
Bank FE	Yes	Yes	No	No	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. obs.	854	546	854	546	515	342	515	342
N Clusters	123	107	123	107	69	65	69	65

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.26. Coefficient size over the distribution of bank size



(a) Column (3) and (4) of Table 4.28



(b) Column (7) and (8) of Table 4.28

Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

4.H Robustness checks

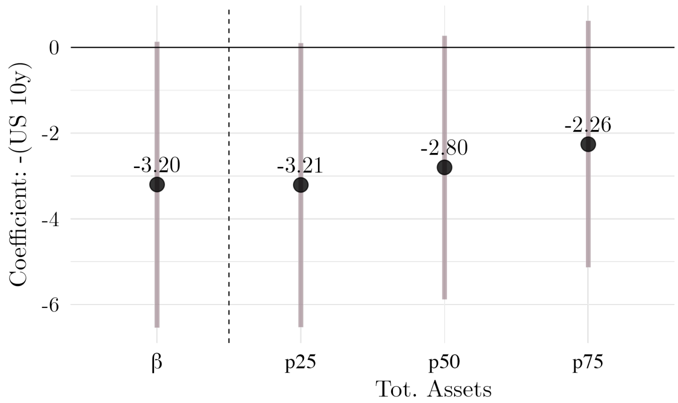
4.H.1 Tier 1 capital ratio as proxy for net worth

Table 4.29. Reaction of tier 1 capital ratio through US 10-year rate

	(1)	(2)	(3)	(4)
-$(\Delta \text{US10y})^q_t$	-0.016	-0.074	-0.032^{*}	-0.054^{*}
	(0.011)	(0.062)	(0.017)	(0.027)
-$(\Delta \text{US10y})^q_t \times \log(\text{Tot. assets})_{i,t}$		0.009		0.005
		(0.010)		(0.003)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	1019	946	1019	946
N Clusters	137	129	137	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.27. Column (3) and (4) of Table 4.29



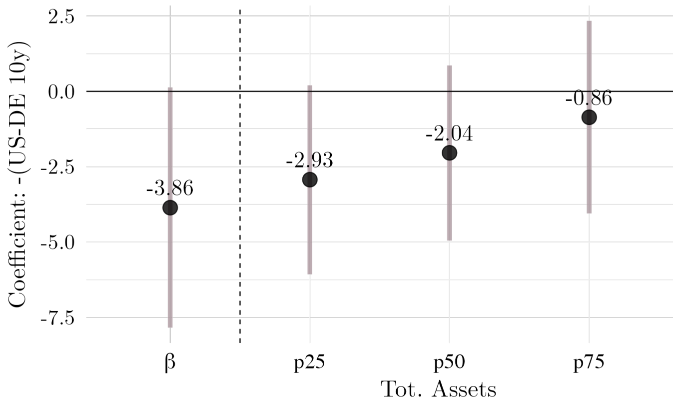
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.30. Reaction of tier 1 capital ratio through US-DE 10-year spread

	(1)	(2)	(3)	(4)
-(ΔUS-DE 10y)$_t^q$	-0.022	-0.045^{*}	-0.039^{*}	-0.076^{*}
	(0.015)	(0.022)	(0.020)	(0.038)
-(ΔUS-DE 10y)$_t^q$ × log(Tot. assets)$_{i,t}$		0.004		0.011
		(0.004)		(0.007)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	1019	946	1019	946
N Clusters	137	129	137	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.28. Column (3) and (4) of Table 4.30



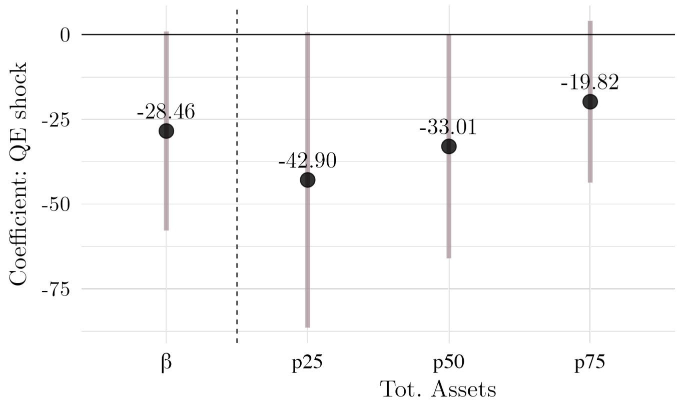
Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

Table 4.31. Overall effect of QE on tier 1 capital ratio

	(1)	(2)	(3)	(4)
QE shock$_t^q$	-0.130	0.199	-0.285^{*}	-0.950^{*}
	(0.090)	(0.287)	(0.150)	(0.561)
QE shock$_t^q$ × log(Tot. assets)$_{i,t}$		-0.057		0.125
		(0.045)		(0.086)
Bank FE	Yes	Yes	No	No
Controls	Yes	Yes	Yes	Yes
Num. obs.	1019	946	1019	946
N Clusters	137	129	137	129

Notes: The left hand side of the table shows the regression results including bank fixed effects, whereas the right hand side does not include bank fixed effects. The *Macro Control* variables are the one-period lagged *harmonized unemployment rate*, *CBOE Volatility Index*, *real GDP growth rate*, and *inflation rate*. In addition, the *3 months* and *10 years interest rate surprises* in the Euro area are included (identified by Kearns et al. (2022)). Standard errors are clustered at the bank level and shown in parentheses below the estimated coefficient. Asterisks indicate significance at the 1%, 5%, and 10% level (***$p < 0.01$; **$p < 0.05$; *$p < 0.1$).

Figure 4.29. Column (3) and (4) of Table 4.31



Notes: This figure shows the coefficient size of the regression. For each subfigure, the column on the left shows the coefficient of the regression which only includes the shock (β), and the three columns on the right show the coefficient for the regression that includes the interaction term with the bank size calculated at the 25^{th}, 50^{th}, and 75^{th} percentiles (of the bank size distribution). The gray bars show the 95% confidence interval.

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