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# THREE ESSAYS ON FIRM FINANCING AND GROWTH

TANG Wenxia

UNIL | Université de Lausanne

FACULTÉ DES HAUTES ÉTUDES COMMERCIALES

DÉPARTEMENT D'ÉCONOMIE

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THÈSE DE DOCTORAT

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pour l'obtention du grade de

Doctorat en économie

par

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Lausanne, le 19.09.2024

Professeure Marianne Schmid Mast, Doyenne

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# Three Essays on Firm Financing and Growth

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# Contents

# Introduction

This thesis includes three chapters that focus on international macroeconomics and development economics. These essays study the channels through which firms overcome financing constraints to achieve growth. Access to credit is critical for the survival and expansion of both large and small firms. However, credit is a limited resource, leading firms to seek various sources of external financing. The three chapters of this thesis explore different channels: (1) bank financing via short-term loans; (2) offshore borrowing through bond issuance in tax havens; and (3) trade credit, where firms extend credit to one another.

The first chapter examines short-term loan financing for small and medium-sized enterprises (SMEs) through the lens of a credit guarantee program in Morocco. My co-authors and I calibrate and estimate a quantitative model to evaluate the equilibrium growth and welfare gains of this loan guarantee program. We find that short-term finance promotes firm growth by enabling entrepreneurs to allocate their net worth more efficiently away from unproductive cash and towards productive capital. We also discover that these effects are persistent only if firms face intertemporal distortions in the form of exit risk or a tax on net worth. We empirically validate these findings by using data from the loan guarantee program designed to relax short-term external financial constraints in Morocco. The empirical results indicate that guaranteed firms expand their production scale and decrease their cash-to-asset ratio.

The second chapter investigates how large firms in China, particularly private companies in the real estate sector, circumvent domestic credit constraints by incorporating in tax havens. I analyze the spillover effects of a macroprudential policy aimed at tightening the domestic corporate bond market. By compiling data from various databases, I piece together the puzzle of Chinese offshore corporate behaviors in tax havens. My empirical estimation indicates that the macroprudential policy successfully reduces the bond issuances of non-state-owned enterprises (non-SOEs) from the domestic credit market. Further analysis reveals that this group of firms is more likely to issue bonds through shell companies in tax havens, compared to their state-owned counterparts after the new regulation. The effects are primarily driven by private firms in the real estate sector. Specifically, a 1% increase in private ownership corresponds to a 1% increase in bonds issued in tax havens for non-SOEs in the real estate industry after the regulatory changes.

The third chapter explores the role of trade credit, examining the impact of a new law aimed at improving payment delays for firms engaged in government procurement contracts in Morocco. Delayed payment among firms is a widespread phenomenon in the Moroccan business world. Payments usually take between 120 and 150 days, compared to an average of 70 days in France. It adversely affects the cash-flow dynamics of enterprises, especially SMEs. My coauthors and I study the impact of Act 49-15, a regulation designed to limit payment delays in government procurement contracts. Using a confidential database from the General Treasury under the Moroccan Ministry of Economy and Finance, we find that firms exposed to the reform reduce their trade credit post-reform relative to unexposed firms. Our analysis further indicates that the treatment effects on trade credit are mostly driven by large firms. This shows an unequal exposure to the new law across firm sizes. The findings highlight the size-dependent effects of policy interventions.

# Chapter 1

# Short-term Finance, Long-term Effects[1]

Kenza Benhima, Omar Chafik, Min Fang, Wenxia Tang

**Abstract**

We study the effect of short-term finance on firm growth and its aggregate implications in emerging economies. In theory, short-term finance promotes firm growth by enabling entrepreneurs to allocate their net worth more efficiently away from unproductive cash and towards productive capital. Importantly, these effects are persistent only if firms face intertemporal distortions in the form of exit risk or a tax on net worth. The quantitative model fitted to Moroccan data replicates qualitatively and quantitatively the observational impacts of a loan guarantee program (LGP) designed to relax short-term financial constraints. Fitting the model to the data also reveals that intertemporal distortions are large and that the costs of participating in the LGP are high. This implies that there are potentially large gains from increasing the guaranteed ratio and decreasing the participation costs. These two policies generate substantial growth and welfare gains, with the former generating relatively higher growth and the latter motivating relatively more participation.

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## 1.1 Introduction

The lack of access to external credit for Small and Medium Enterprises (SMEs) is one of the major development bottlenecks in MENAP (Middle East, North Africa, Afghanistan, and Pakistan) countries. For instance, Blancher et al. (2019) find that closing the SME financial inclusion gap in those countries would increase annual growth by up to 1%. The first purpose of this paper is to study the ability of loan guarantee programs (LGP) targeted at SMEs to improve their access to credit, the extent to which it actually promotes development, and through which channels, using a model calibrated to Moroccan firm-level data on an ongoing LGP.

The second purpose of this paper is to understand the effect of short-term external finance on firm growth and the aggregate economy in emerging countries. Financial frictions hinder the ability of firms to use inputs efficiently, affect firm growth, and, therefore, lower economic development, especially in emerging economies. Extensive literature addresses how the scarcity of long-term external finance leads to under-leveraged small and young enterprises and hinders economic development.[2] However, little attention has been paid to the scarcity of short-term external finance in emerging countries. This paper fills that gap.

The role of short-term external finance is different from long-term external finance. Long-term external finance promotes firm growth since it directly enlarges entrepreneurs' total asset scale, given their net worth. Short-term external finance, as we show in a simple dynamic model, promotes firm growth by enabling a more efficient allocation of the entrepreneurs' existing net worth. Since entrepreneurs in emerging economies tend to hoard a substantial amount of cash to meet their working capital needs, short-term external finance promotes firm growth by allowing entrepreneurs to allocate their net worth more efficiently away from unproductive cash and towards productive capital. Importantly, these effects are persistent only if firms face intertemporal distortions in the form of exit risk or a tax on net worth. These distortions, by discouraging net worth accumulation, prevent firms from outgrowing their financial constraints through savings and self-financing. Therefore, in the presence of intertemporal distortions, better access to external finance has both short-term and long-term impacts.

To evaluate the equilibrium growth and welfare gains of expanding credit guarantee programs, we calibrate and estimate a quantitative model using a loan guarantee program (LGP) designed to relax short-term external financial constraints in Morocco. Our dataset combines Moroccan firm-level data from Orbis with the national-level loan guarantee data from Tamwilcom.[3]

[2]See Cooley and Quadrini (2001), Albuquerque and Hopenhayn (2004), Clementi and Hopenhayn (2006), DeMarzo and Fishman (2007), Buera, Kaboski, and Shin (2011a), and Arellano, Bai, and Zhang (2012), among others.

[3]Tamwilcom is a public financial institution under the supervision of the Central Bank of Morocco, Bank Al-Maghrib. Therefore, the national loan guarantee data covers every firm that has ever been guaranteed in Morocco.

The model includes, for quantitative purposes, a uniformly distributed LGP participation cost and a size-dependent collateral constraint. The participation cost prevents small firms from entering the program as only firms with a high enough growth potential self-select into the program. In contrast, the large firms do not self-select into the program since they do not need the guarantee. These assumptions are important quantitatively to fit the endogenous selection of firms into the program and the observed hump-shaped participation rate in firm size.

There are two main takeaways from the model's calibration. First, a high participation cost, equivalent to one-third of the average net worth, is needed to fit the low LGP participation rate of 3.8%. This implies that only large constrained, or small fortunate firms can access the guarantee. This large number reflects a lack of access of small firms to the guarantee program due to geographical barriers or to the fact that many small firms do not have a relationship with the banking system. Making access to the LGP easier is, therefore, one channel through which the program can be expanded. Second, intertemporal distortions are large. These large distortions imply that credit guarantees have a strong impact on firms both in the short and long run.

We empirically validate the findings from our model by using our Moroccan data. The empirical results show that, relative to their peers, (i) guaranteed firms expand their production scale homogeneously by increasing sales, capital input, and labor input by about 10% relative to their matched peers and (ii) they decrease their cash-to-asset ratio. Importantly, the results are persistent, which is consistent with the presence of intertemporal distortions. We do not interpret these results as causal, as we cannot rule out self-selection: firms that are more productive can endogeneously self-select into the program. However, our quantitative model, which accounts for self-selection, is able to replicate the empirically estimated impact of the guarantee: the model-simulated data yields estimation results that are similar to the empirical model.

We then examine two policies aimed at expanding the LGP: a higher guaranteed ratio and a lower participation cost. We show that the gains from enlarging the loan guarantee programs by reducing both frictions are substantial. Increasing the guaranteed ratio from 60% to 80%, as in Indonesia, more than doubles the participation rate, decreases the cash ratio by 1.2 percentage points, and achieves an output growth of 0.09% and a welfare gain of 0.25%. Decreasing participation costs achieves less extra growth and welfare, but increases the participation rate further without substantially increasing the guaranteed portfolio. Interestingly, increasing the guaranteed ratio also has a positive effect on participation, as it increases the gains from participating in the program. This is especially true for small firms, which are more likely to be constrained while productive.

This paper relates to the large theoretical literature on firms' financial frictions and their aggregate implications, such as Cooley and Quadrini (2001), Albuquerque and Hopenhayn (2004), Quadrini (2004), Clementi and Hopenhayn (2006), DeMarzo and Fishman (2007), Huynh and Petrunia (2010), Arellano, Bai, and Zhang (2012), Moll (2014), Midrigan and Xu (2014), Gopinath et al. (2017), Jo and Senga (2019), Buera, Kaboski, and Shin (2021) and others. We contribute to this literature first by focusing on short-term finance. We show that short-term external finance promotes firm growth by allowing entrepreneurs to allocate their net worth more efficiently towards productive capital stock and away from unproductive cash holdings. Second, we contribute to a better understanding of the aggregate impact of external finance. The literature has shown that the gains from access to external finance are elusive. In particular, when productivity shocks are persistent, firms typically grow out of their collateral constraints through savings and self-financing (Moll, 2014; Buera, Kaboski, and Shin, 2021). We show that intertemporal distortions, that is, distortions in the consumption/saving choices of entrepreneurs, are an overlooked and crucial factor that amplifies the aggregate effect of external finance by hindering firms' self-financing abilities. This intertemporal friction interacts with the external finance frictions to determine the long-run scale of firms. We estimate these distortions and show that they are important in Morocco. We also show, by performing counterfactual exercises, that they contribute substantially to the aggregate effect of the LGPs.[4]

In our model, credit guarantees are modeled explicitly. Firms face a credit constraint that is micro-founded in the spirit of Hart and Moore (1994) and Kiyotaki and Moore (1997). The renegotiation-proof debt contract generates a debt limit equal to the creditors' outside value. This outside option depends on the liquidation value of the firm. Credit guarantees affect the firms' credit constraints by increasing the creditors' outside value. We assume that the firms' participation constraint is always satisfied, which we check ex-post in our quantitative analysis. In other studies, the credit policy is typically modeled as a technology that reduces or annihilates the enforceability constraint of the loan (Jo and Senga, 2019; Buera, Kaboski, and Shin, 2011b, 2021). We acknowledge here that credit guarantees do not improve the firms' credit access through better enforceability of the loan but rather by affecting the banks' incentives to provide loans. Another strand of the literature has shown that when markets are prone to adverse selection (Stiglitz and Weiss, 1981), credit guarantees improve aggregate outcomes (Smith and Stutzer, 1989; Gale, 1990; Philippon and Skreta, 2012). Other papers, some of which were motivated by the Covid-19 crisis, focus on emergency credit guarantee programs that are put in place to limit firms' defaults and debt overhang (Philippon and Schnabl, 2013; Elenev, Landvoigt, and Van Nieuwerburgh, 2022; Glode and Opp, 2023; Martin, Mayordomo, and Vanasco, 2023).

This paper also contributes to the empirical literature on credit guarantee schemes. A credit

[4]Intertemporal distortions are often in the background of theoretical papers studying firms' financial frictions in the form of an exit probability or a tax, but their role is not made explicit or quantified. See for instance Arellano, Bai, and Zhang (2012), Jo and Senga (2019), Cooley and Quadrini (2001).

guarantee scheme is one of the most common policy tools to facilitate SMEs' access to finance. Gudger (1998) and Green (2003) provide an overview of credit guarantee programs' typology, design, implementation, and general evaluation worldwide. Beck, Klapper, and Mendoza (2010) survey 76 partial credit guarantee schemes across 46 developed and developing countries. Saadani, Arvai, and Rocha (2011) focus on the Middle East and North Africa (MENA) and review credit guarantee programs in 10 countries in the MENA region. Some empirical contributions study the impact of guarantee programs using microdata, including Paravisini (2008), Oh et al. (2009), Lelarge, Sraer, and Thesmar (2010), Bach (2013), Banerjee and Duflo (2014), Brown and Earle (2017), Mullins and Toro (2018), Bhue, Prabhala, and Tantri (2019), Wilcox and Yasuda (2019), Barrot et al. (2019) and Bachas, Kim, and Yannelis (2021). Our paper contributes to the literature by presenting new empirical findings on the usage of cash and the profile of growth post-guarantee. Our evidence is not causal but is used to validate our model.

The rest of the paper is organized as follows. Section 1.2 lays down the full model. Section 1.3 presents the mechanism and implications of short-term financial constraints in a simple special case. Section 1.4 calibrates the full model to Moroccan firm-level data and a Moroccan loan guarantee program. Section 1.5 documents some empirical findings and validates the quantitative analysis. Finally, Section 1.6 performs policy experiments.

## 1.2 The Model

We consider an economy with heterogeneous entrepreneurs facing collateral and working capital constraints, intertemporal distortions, and access to short-term loan guarantees with fixed participation costs. Time is discrete. There is a unit mass of entrepreneurs indexed by $i \in [0, 1]$. Each entrepreneur owns a firm that is subject to idiosyncratic productivity shock. We do not distinguish entrepreneurs and firms. Firms decide how much investment to undertake, how much labor to hire, how much debt to issue, how much cash to hold, and how many dividends to pay. They face three frictions: first, external financial friction in the form of a collateral constraint and a working capital constraint; second, intertemporal distortions in the form of an exogenous exit risk and a tax on net worth; third, imperfect selection into the guarantee program in the form of a stochastic participation cost. A capital good producer and a unit mass of other households complete this general equilibrium model.

### 1.2.1 Technology and Production

**Technology** Each firm $i$ produces with capital $k_{i,t}$, and labor $l_{i,t}$ using the production function $y_{i,t} = z_{i,t}F(k_{i,t}, l_{i,t}) + (1 - \delta)q_t k_{it}$, where $z_{i,t}$ is the firm's idiosyncratic stochastic productivity, which follows an exogenous Markov process $\log(z_{i,t}) = \rho_z \log(z_{it-1}) + \sigma_z \varepsilon_{i,t}$, with $\varepsilon_{i,t}$ a standard normal i.i.d. process, $\delta$ is the capital depreciation rate and $q_t$ is the price of capital. The input combination $F(k_{i,t}, l_{i,t})$ features decreasing returns to scale.

**Working Capital Constraint** At the beginning of each period, before the realization of their productivity shocks, firms pay in advance for their working capital: they are required to pay their current period wage bill $w_t l_{i,t}$ before production. They can finance this working capital through both internal and external funds: they can use their cash holdings $c_{i,t}$ or short-term borrowing $b_{i,t} \leq \bar{b}_{i,t}$. Therefore, the working capital constraint is $w_t l_{i,t} \leq c_{i,t} + \bar{b}_{i,t}$.

### 1.2.2 Credit Market

Firms can access competitive financial intermediaries, who receive deposits from all households (including cash from entrepreneurs) and provide short-term loans to firms. The competitive financial intermediaries make zero net profit with the same market deposit interest rate and loan interest rate of $r_t$. Contracts are imperfectly enforceable, as entrepreneurs can walk away without completing production. This gives rise to a short-term borrowing constraint.

The short-term borrowing $b_{i,t} \leq \bar{b}_{i,t}$ of the firm $i$ is subject to a collateral constraint. Since entrepreneurs can easily transfer their liquid assets (cash holdings), financial intermediaries only consider their illiquid assets (fixed capital) as collateral. As in Hart and Moore (1994) and Kiyotaki and Moore (1997), we assume that the firm's creditors impose a repudiation-proof debt limit on the firm. This debt limit is equal to the liquidation value of the firm, which is the outside option of the creditors.[5] Since the liquidation of the firm's fixed capital would incur fixed costs and proportional re-structuring costs, the collateral constraint could, therefore, be nonlinear in the firm's capital stock (Gopinath et al., 2017). We write the liquidation value as $\Theta(k_{i,t})$, which will be specified later. Without guarantees, the debt limit $\bar{b}_{i,t}$ is equal to the liquidation value $\Theta(k_{i,t})$.

We are assuming here that the firm participation constraint is always satisfied and firms do not default in equilibrium. In the event of default, the firm does not pay back its debt but loses its capital stock. For default to be ruled out, it is enough that the depreciated value of capital $(1 - \delta)k_{it}$ is higher than the total value of debt repayments. We assume this condition is satisfied

[5] As in Kiyotaki and Moore (1997), we consider a situation where firms can negotiate the debt down to the liquidation value of the firm's capital because the value of the project is zero without the cooperation of the firm owner.

throughout, but check the ex-post to ensure it is indeed the case in the quantitative analysis.

### 1.2.3 Loan Guarantee

We now introduce the loan guarantee programs. For the sake of fitting the Moroccan credit market and LGP, we make the following assumptions. First, a certain fraction of credit is non-bank third-party credit, which the government can hardly guarantee.[6] Therefore, we assume that there are two types of intermediaries: banks, which can benefit from the LGP, and non-bank creditors, which cannot. We denote bank loans by $b_{it}^{b}$ and the non-bank loans by $b_{it}^{nb}$. In case of liquidation, banks are entitled to a fraction $s \in [0, 1]$ of the firm's collateral in case of default, while non-bank creditors are entitled to a fraction $1 - s$. This fraction $s$ represents the relative bargaining power of bank and non-bank creditors in appropriating the liquidation value once the firm has defaulted. This means that the firm could renegotiate its formal debt $b_{it}^{b}$ down to $s\Theta(k_{it})$ with the banks and its informal debt $b_{it}^{nb}$ down to $(1 - s)\Theta(k_{it})$ with its informal creditors. Banks and non-bank creditors thus each impose a separate renegotiation-proof debt limit: $b_{it}^{b} \leq s\Theta(k_{it})$ and $b_{it}^{nb} \leq (1 - s)\Theta(k_{it})$. This implies that in practice, a proportion $s$ of a firm's loans comes from formal banks, and a proportion $(1 - s)$ comes from non-bank creditors.

Second, the LGP does not restrict large firms from applying.[7] With a fixed commission fee, unconstrained large firms would self-select not to participate.

Third, a firm's selection into LGP does not perfectly reflect its profitability. Therefore, we assume that firms participating in the LGP incur a uniformly distributed random fixed participation cost $\xi \in [0, \bar{\xi}]$, which is paid in units of labor. This fixed cost accounts for all the explicit and implicit barriers to accessing the guarantees, which can be pecuniary, geographical, or cognitive. In practice, this cost can reflect an unequal access to the guarantee. Moreover, as we will see, because of this cost, only firms that are profitable enough will select into the guarantee. The model can then account for endogenous self-selection.[8]

Finally, upon successfully getting the guaranteed loan, a fraction $x$ of the bank loans $b_{i,t}^{b}$ is guaranteed, and the firm pays a commission fee $\mu$ on top of the interest rate for the guaranteed part of the loan to the government. If the guaranteed firm defaults, the government would repay the guaranteed proportion $x$ of the loan to the banks.

[6] For instance, a large fraction of credit in Morocco is composed of trade credit.

[7] In fact, the Moroccan LGP does impose a limit on firm size, but this limit is not binding in practice as very few large firms apply to guarantee. In alternative specifications, we impose a cap on firm size in the LGP. The quantitative results are almost equivalent to the baseline model.

[8] A random fixed cost setup is widely used in the lumpy investment literature, i.e., Khan and Thomas (2008), Fang (2021), and Fang (2023). It is also introduced in Chen, Deng, and Fang (2022) for patent collateral participation.

**The LGP multiplier** If the firm obtain the guarantee, then the bank's outside option increases by the amount $xb^b_{i,t}$, which means that the renegotiation-proof debt limit on bank loans becomes $b^b_{i,t} \leq s\Theta(k_{i,t}) + xb^b_{i,t}$. This generates a financial multiplier effect: for each unit of additional bank loan, the firm increases its liquidation value by $x$. The effective constraint, which takes into account this multiplier effect, is then: $b^b_{i,t} \leq s\chi\Theta(k_{i,t})$, where $\chi = \frac{1}{1-x}$ is the LGP multiplier. The LGP multiplier is greater than one and can be very large. For instance, if the government guarantees 60% of the bank loan, then $\chi = \frac{100\%}{100\%-60\%} = 2.5$.

**Effective constraint on total borrowing** Let $F = \{A, N\}$ indicate whether a firm decides to pay the fixed participation cost and participate in the LGP. When $F = N$, the firm does not pay the participation cost and can only borrow up to its original collateral constraint. When $F = A$, the firm pays the participation cost and relaxes its borrowing constraint. In that case, the firm can borrow up to $\chi$ times more in the form of bank loans. Therefore, the effective constraint on total borrowing $b_{i,t} = b^b_{i,t} + b^{nb}_{i,t}$ that the firm faces depends on $F$:

$$b_{i,t} \leq \begin{cases} (1 + (\chi - 1)s)\Theta(k_{i,t}) & \text{if } F = A \\ \Theta(k_{i,t}) & \text{if } F = N \end{cases}$$

### 1.2.4 Recursive Problem for Firms

The individual state variables of a firm are its idiosyncratic productivity $z_{i,t}$ and its beginning-of-period net worth $n_{i,t-1}$. Firm decisions are divided into three sub-periods. In the first sub-period, the firm learns about the participation cost for the next period and decides whether to participate in the LGP. In the second sub-period, the firm makes production and working capital decisions. In the third sub-period, it makes consumption and saving decisions.

**Production Decisions** It is useful first to consider the production decisions, which happen in the second sub-period, before the participation decisions in the first sub-period. In the second sub-period, after having drawn its participation cost $\xi^i_t$ and decided whether to participate ($F^i_t = A$) or not ($F^i_t = N$) in the LGP, the firm maximizes its total net revenue given its productivity, its beginning-of-period net worth, and its LGP participation decision $F^i_t$. The firm decides how much capital $q_t k_{i,t}$ to invest, how much cash $c_{i,t}$ to hold, and how much labor $w_t l_{i,t}$ to hire. Given the working capital and collateral constraints, the firm maximizes its net revenue

$$\pi^*(z_{i,t}, n_{i,t-1}, F_{i,t}) = \max_{k^i_t, c^i_t, l^i_t, b^i_t} \left\{ z_{i,t}F(k_{i,t}, l_{i,t}) - w_t l_{i,t} + (1-\delta)q_t k_{i,t} + (1+r_t)c_{i,t} - r_t b_{i,t} - \mu\tilde{b}_{i,t} \right\} \tag{1.1}$$

subject to the constraints

$$n_{i,t-1} = q_t k_{i,t} + c_{i,t} \tag{1.2}$$

$$w_t l_{i,t} \leq c_{i,t} + \mathbb{1}_{F_{i,t}=A} \cdot (1 + (\chi - 1)s)\Theta(k_{i,t}) + (1 - \mathbb{1}_{F_{i,t}=A}) \cdot \Theta(k_{i,t}) \tag{1.3}$$

$$b_{i,t} = w_t l_{i,t} - c_{i,t} \tag{1.4}$$

$$\tilde{b}_{i,t} \equiv \mathbb{1}_{F_{i,t}=A} \cdot \chi s \Theta(k_{i,t}) \tag{1.5}$$

where $\tilde{b}_{i,t}$ is the guaranteed proportion of loans that pays a commission fee $\mu$.

**Participation Decision** Now that we understand how net revenues are affected by participation in the LGP, we can characterize the participation decision $F_{i,t}$. In the first sub-period, the chooses to participate ($F_{i,t} = A$) whenever $\xi_t^i < \xi^*(z_{i,t}, n_{i,t-1})$, where $\xi^*(z_{i,t}, n_{i,t-1})$ is a threshold participation cost:

$$\xi^*(z_{i,t}, n_{i,t-1}) = \frac{\pi^*(z_{i,t}, n_{i,t-1}, A) - \pi^*(z_{i,t}, n_{i,t-1}, N)}{w_t} \tag{1.6}$$

A firm with state $(z_{i,t}, n_{i,t-1})$ which draws a fixed cost higher than $\xi^*(z_{i,t}, n_{i,t-1})$ will not participate ($F_{i,t} = N$) in the loan guarantee program. Otherwise, it pays the fixed cost and joins the program.

**Intertemporal Decisions and Intertemporal Distortions** Finally, in the third sub-period, the entrepreneur makes saving and consumption decisions. The entrepreneur faces two intertemporal distortions that are relevant to their consumption/saving choices. The first is due to the high exit risk that firms face in emerging countries. We assume an exogenous survival rate $\epsilon \leq 1$. Exiting firms are replaced with the same measure of entrants with an initial low net worth $\underline{n}_0$.[9] The second is the "erosion" of the firm's net worth, which we represent through a tax on net worth $\tau \geq 0$. This erosion could come from effective taxes, but also red tape, corruption, and expropriation risk and captures the entrepreneur's potential losses on their net worth in developing countries. Both the exit risk and net worth erosion work as intertemporal distortions. We will use our Moroccan firm-level data to discipline these distortions and quantify their economic effects.

The entrepreneur makes consumption/saving decisions to maximize her value function $v(z_{i,t}, n_{i,t-1}, F_{i,t})$ given her end-of-period total net revenue $\pi^*(z_{i,t}, n_{i,t-1}, F_{i,t})$ and her participation cost $\xi_t^i$. We write the entrepreneur's optimization recursively:

$$v(z_{i,t}, n_{i,t-1}, F_{i,t}, \xi_t^i) = \max_{d_{i,t}, n_t^i} \left\{ \frac{d_{i,t}^{1-\eta}}{1-\eta} + \beta \epsilon E_t[\tilde{v}(z_{i,t+1}, n_{i,t})] \right\} \tag{1.7}$$

where $\beta$ is the stochastic discount factor of the entrepreneur.

[9] The initial net worth $\underline{n}_0$ of all new entrants equals the post-restructuring net worth of exiting firms. We assume that financial institutions owned by all households conduct the restructuring, so such restructuring costs return to households' total income.

The net worth follows the accumulation rule:

$$n_{i,t} = (1-\tau)\{\pi^*(z_{i,t}, n_{i,t-1}, F_{i,t}) - d_{i,t} - w_t \xi_{i,t}\} \tag{1.8}$$

Finally, $\tilde{v}(z_{i,t+1}, n_{i,t})$ is the ex-ante value of the firm before drawing the new participation cost in period $t+1$. Therefore, $\tilde{v}(z_{i,t+1}, n_{i,t}) \equiv \frac{\xi^*(z_{i,t+1}, n_{i,t})}{\bar{\xi}} v(z_{i,t+1}, n_{i,t}, A) + (1 - \frac{\xi^*(z_{i,t+1}, n_{i,t})}{\bar{\xi}}) v(z_{i,t+1}, n_{i,t}, N)$, because $\xi^*(z_{i,t}, n_{i,t})$ is bounded between 0 and $\bar{\xi}$.

### 1.2.5 Other Households and the Capital Good Producer

The general equilibrium model is completed by introducing a unit mass of identical households that consume and supply labor and the capital goods producer who supplies investment goods.

**Other Households** There is a unit measure continuum of identical non-entrepreneur households with preferences over consumption $C_t$ and labor supply $L_t$ whose expected utility is:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left( \frac{C_t^{1-\eta}}{1-\eta} - \theta \frac{L_t^{1+\omega}}{1+\omega} \right)$$

subject to the budget constraint $C_t + \frac{1}{1+r_t} B_t \leq B_{t-1} + w_t L_t + \Pi_t$, where $\beta$ is the discount factor of households, $\theta$ is the disutility of working, $r_t$ is the interest rate, $B_t$ is a one-period bond, and $w_t$ is the wage. $\Pi_t$ summarizes all the profits from financial institutions and all the government gains or losses in the loan guarantee program that are transferred to other households. Households choose consumption, labor, and bonds, which supply two Euler equations that determine both the wage and the real interest rate:

$$w_t = -\frac{U_l(C_t, L_t)}{U_c(C_t, L_t)} = \theta L_t^{\omega} C_t^{\eta} \tag{1.9}$$

$$\frac{1}{1+r_t} = \beta \frac{U_c(C_{t+1}, L_{t+1})}{U_c(C_t, L_t)} = \beta \Big( \frac{C_t}{C_{t+1}} \Big)^{\eta} \tag{1.10}$$

**Capital Good Producer** There is a representative capital good producer who produces new aggregate capital using the technology $\Phi(I_t/K_t)K_t$, where $I_t$ are units of the final good used to produce capital, $K_t = \int k_{jt} dj$ is the aggregate capital stock at the beginning of the period, $\Phi(I_t/K_t) = \frac{\delta/\phi}{1-1/\phi} \left( \frac{I_t}{K_t} \right)^{1-1/\phi} - \frac{\delta}{\phi - 1}$, and $\delta$ is the steady-state investment rate. Profit maximization pins down the relative price of capital as $q_t = \frac{1}{\Phi'(I_t/K_t)} = \frac{I_t/K_t}{\delta}^{1/\phi}$.

### 1.2.6 Equilibrium Definition

We now characterize and define the equilibrium of the model. We focus on the stationary equilibrium, given current government policies.

**Definition 1 (Stationary Equilibrium)** *A stationary equilibrium for this economy is defined by a set of policy functions* $\{v(z,n,F,\xi),\ \xi^*(z,n),\ k(z,n,F),\ c(z,n,F),\ l(z,n,F),\ \pi(z,n,F),\ d(z,n,F,\xi)\}$*, a set of quantities* $\{C, L, Y, K\}$*, a set of prices* $\{w, r, q\}$*, and a distribution* $\mu'(z,n,F,\xi)$ *that solves the firms' problem, capital good producer's problem, household's problem, and market clearing of labor and final goods such that:*

*(i). [Firm Optimization] Taking the aggregate prices* $\{w, r, q\}$ *as given,* $\{v(z,n,F,\xi),\ \xi^*(z,n),\ k(z,n,F),\ c(z,n,F),\ \pi(z,n,F),\ d(z,n,F,\xi)\}$ *solve the firms' static participation and production choices and the dynamic consumption/saving choice.*

*(ii). [Household and Capital Good Producer Optimization] Taking the aggregate prices* $\{w, r, q\}$ *as given,* $C$ *and* $L$ *solve the household's utility maximization problem and* $I$ *and* $K$ *solve the capital producer's maximization problem.*

*(iii). [Market Clearing] Given the aggregate prices* $\{w, r, q\}$ *as given, the labor market clears* $L = \int l(z,n,F)d\mu'(z,n,F,\xi)$*, and the final goods market clears* $Y = C + D + I + \Delta$*, where* $D = \int d(z,n,F,\xi)d\mu'(z,n,F,\xi)$ *is the sum of entrepreneurs' dividend and* $\Delta = \int \tau n(z,n,F,\xi)d\mu'(z,n,F,\xi)$ *is the sum of the net worth erosion across all firms.*

*(iv). The quantities* $\{C, L, Y, K\}$*, prices* $\{w, r, q\}$*, and distribution* $\mu'(z,n,F,\xi)$ *are constant.*

## 1.3 The Mechanism and Predictions

Before we turn to the quantitative and empirical analyses, we consider a simple, special case of the model in partial equilibrium to illustrate the main mechanisms that could qualitatively guide our empirical analysis and discuss their aggregate implications.

The mechanism generates three predictions. First, we show how, in the joint presence of collateral and working capital constraints, getting access to the loan guarantee program alleviates cash needs, allocates resources to productive capital, and achieves higher sales in the short run. Second, we show that while the loan guarantee always benefits firms in the short run, the guarantee affects the firm's long-run scale only in the presence of intertemporal distortions. Third, selection into the program is endogenous: when the participation cost prevents the participation of small and unproductive firms, it also lowers the participation of large firms due to sufficient self-financing. As a result, we should observe a hump-shaped participation rate in the loan guarantee program over firm size.

### 1.3.1 A Special Case

We consider a special case of the full model defined as follows:

**Definition 2 (A Special Case)** *In the special case, we make the following assumptions:*

*(i). The technology is Leontief, and productivity is constant:*

$$y_t = Z[min(k_t, a^{-1} l_t)]^{\alpha} \tag{1.11}$$

*where $Z$ is the firm's constant total productivity, $a^{-1}$ measures the relative labor productivity of the firm, and $\alpha$ is the curvature of the production function.*

*(ii). The collateral constraint is linear: $\Theta(k_t) = \theta k$. Note that in a conventional calibration, labor share relative to capital share is $a \approx 2$, and $\theta < 1$, so we assume that $a > \theta$.*

*(iii). The loan guarantee does not incur fees: $\mu = 0$, the participation cost is zero: $\xi = 0$, and all credits are formal loans: $s = 1$.*

*(iv). The wage, the price of capital, and the interest rate are constant as the economy is in a partial equilibrium: $w_t = q_t = 1$ and $1 + r_t = 1/\beta$.*

In what follows, we discuss the effects of short-term finance in the form of LGPs on the firm's optimal choices of capital and cash and on firm growth. The LGP helps the firm scale up its pledgeable share of capital ($\theta$) from $\theta_{low}$ ($F = N$) to $\theta_{high}$ ($F = A$) where the additional share ($\theta_{high} - \theta_{low}$) is guaranteed by the government. We compare the trajectories of a firm in two worlds, one in which it obtains access to the guarantee and one in which it does not.

Define $\psi_t = Z[min(k_t, a^{-1} l_t)]^{\alpha} - q_t k_t - w_t l_t + (1 - \delta) k_t$ as the equilibrium profits. With the Leontieff assumption, $l_t = a k_t$ always holds. Then, since $w_t = q_t = 1$, $\psi_t = \psi(k_t) = Z k_t^{\alpha} - (a + \delta) k_t$. We consolidate the firm's problem, represented by Equation (1.1) and Equation (1.7). The following maximization can summarize the firm's choice:

$$v(n_{t-1}) = \max_{k_t, c_t, d_t, n_t} \left\{ \frac{d_t^{1-\eta}}{1-\eta} + \beta \epsilon v(n_t) \right\} \tag{1.12}$$

subject to the constraints

$$(1 - \tau)[\psi(k_t) + (1 + r_t) c_t] - d_t - n_t \geq 0 \tag{1.13}$$

$$n_{t-1} \geq k_t + c_t \tag{1.14}$$

$$c_t + \theta k_t \geq a k_t \tag{1.15}$$

$$c_t \geq 0 \tag{1.16}$$

### 1.3.2 Short-run Growth and Resource Reallocation from Cash to Capital

In the first step, we analyze how scaling up the firm's pledgeable share from $\theta_{low}$ to $\theta_{high}$ affects the static production and financing choices of the firm's decisions given its net worth $n_t$.

Denote by $\eta_t$ the shadow price of the budget constraint (1.13), and $\gamma_t \eta_t (1-\tau)$, $\lambda_t \eta_t (1-\tau)$ and $\zeta_t \eta_t (1-\tau)$ the shadow prices of, respectively, the net worth allocation constraint (1.14), the working capital constraint (1.15) and the non-negative cash constraint (1.16). The shadow prices are normalized by $[\eta_t(1-\tau)]^{-1}$ for convenience. From the first-order conditions of the objective equation (1.12) concerning capital and cash holdings, we can derive the relationship between the marginal benefit of capital ($MBK$) and marginal benefit of cash holding ($MBC$) through the three shadow prices. The full derivation is in Appendix F. Below we describe the relationship between $MBK$ and $MBC$.

$$\gamma_t = MBK_t = MBC_t + \zeta_t \tag{1.17}$$

where the marginal benefit of capital ($MBK$) and marginal benefit of cash holding ($MBC$) are

$$\begin{aligned} MBK_t &= \underbrace{\psi'(k_t)}_{\text{Physical Return of Assets}} + \underbrace{\lambda_t(\theta - a)}_{\text{Shadow Return of Finance}} \\ MBC_t &= \underbrace{1 + r_t}_{\text{Physical Return of Assets}} + \underbrace{\lambda_t}_{\text{Shadow Return of Finance}} \end{aligned}$$

with $\psi'(k_t) = Z\alpha k_t^{\alpha-1} - (a+\delta)$. Both capital and cash have a physical return on assets and a shadow return on finance. These physical and shadow returns may differ. First, capital has a large physical return $\psi'(k_t)$ for a sufficiently small firm, while cash has a low physical return. Second, capital has a negative shadow return of finance ($\lambda_t(\theta - a)$).[10] Increasing the capital stock increases the demand for labor and, hence, the need for working capital, thus increasing the tightness of the collateral constraint. However, cash provides a positive shadow return of finance ($\lambda_t > 0$) because increasing cash reduces the need for external working capital funds, thus relaxing the tightness of collateral constraint.

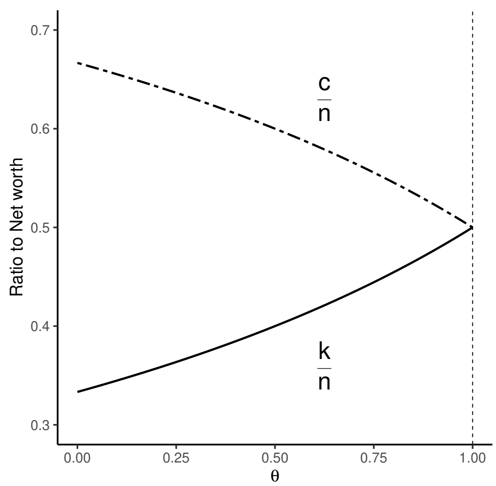
The optimal choice of a constrained firm whose demand for cash is positive ($\zeta_t = 0$), that is, a

[10]Since $a$ measures the input share of labor relative to capital, which is usually assumed to be around 2, without loss of generality, $a > \theta$ always holds. As a result, $(\lambda_t(\theta - a)) < 0$ always holds.

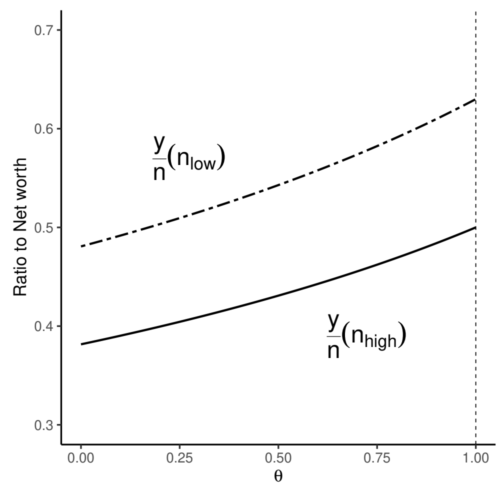
sufficiently small firm, would be to build cash holdings up to achieve a shadow benefit of relaxing the collateral constraint such that $MBK_t = MBC_t$. This yields an optimal shadow return of finance from cash $\lambda_t^* = \frac{\psi'(k_t)-r_t}{1+a-\theta}$. This shadow return is proportional to the return of capital $\psi'(k_t)$, which is higher for smaller firms, as $\psi(\cdot)$ decreases in $k$, and $k$ is constrained by $n$. Combining the binding collateral and working capital constraints ($ak_t = c_t + \theta k_t$) and the budget constraint ($c_t + k_t = n_t$), the constrained firm's choices of capital and cash are proportional to net worth:

$$k_t^* = k(\theta, n_t) = \frac{1}{1+a-\theta} n_t, \quad c_t^* = c(\theta, n_t) = \frac{a-\theta}{1+a-\theta} n_t. \tag{1.18}$$

Figure 1.1: Relationship between Optimal Choices and $\theta$




(a) Capital and Cash




(b) Production (or Sales)

Note: This plot shows the entrepreneur's optimal capital, cash, and production choices as a function of $\theta$. The numerical calibration of the parameters corresponds to an annual frequency: the annual depreciation in the capital is set to $\delta = 0.1$, returns to scale are set to $\alpha = 2/3$, the labor to capital share is set to $a = 2$, and $n_{low} = 1$ and $n_{high} = 2$ stand for small and large entrepreneurs.

Using the production function (1.11) along with the optimal capital and cash equations (1.18), we can determine how the capital, cash, and output choices are affected by $\theta$, which is summarized in Proposition 1 below. Figure 1.1 illustrates these properties in Proposition 1 visually.

**Proposition 1** *Suppose that the economy is described by the special case laid down in Definition 2. LGPs that increase the firm's collateral $\theta$ increase the shadow benefit of relaxing the collateral constraint and, therefore,*

*(i). increase the firm's optimal choice of capital* $\left(\frac{\partial k}{\partial \theta} > 0\right)$.

*(ii). decrease the firm's optimal choice of cash holdings* $\left(\frac{\partial c}{\partial \theta} < 0\right)$.

*(iii). increase the firm's optimal sales* $\left(\frac{\partial y}{\partial \theta} > 0\right)$.

***Proof.*** *(i), (ii), and (iii) are derived directly from Equation* (1.18).

### 1.3.3 Long-run Scales of the Firms with Intertemporal Distortions

In the second step, we analyze how increasing the firm's collateral ability from $\theta_{low}$ to $\theta_{high}$ affects the long-run scale of the firm. We now consider the firm's intertemporal choices and derive the first-order conditions of the firm's objective (1.12) with respect to dividends, future net worth, and cash. The full derivation is in Appendix F. The derived Euler equation is as follows:

$$1 = \beta\epsilon(1-\tau)\left(\frac{d_{t+1}}{d_t}\right)^{-\eta}((1+r_{t+1})+\lambda_{t+1}) \tag{1.19}$$

where $\epsilon(1-\tau) < 1$ is the total intertemporal distortion resulting from both the exit risk and the net worth erosion and distorts the entrepreneur's net worth accumulation. If the firm survives long enough, the shadow price of the constraint admits a long-term value that we denote by $\lambda^{LT}$. It is defined by the Euler equation (1.19) where $d_t = d_{t+1} = d^{LT}$:

$$\lambda^{LT} = \frac{1}{\beta\epsilon(1-\tau)} - (1+r^{LT}) \tag{1.20}$$

where $\beta(1+r^{LT}) = 1$ if the economy is in a steady state.

As a result, the long-term shadow price of cash is positive ($\lambda^{LT} > 0$) in the long run only if $\epsilon(1-\tau) < 1$, that is, if entrepreneurs face intertemporal distortions. Therefore, a key implication is that the firm's long-run scale is affected by financial constraints only in the presence of intertemporal distortions. Otherwise, the financial constraints are irrelevant ($\lambda^{LT} = 0$) in the long run, so financial constraints only affect the speed at which the firm converges to that long-term scale. The following proposition summarizes how the financial constraints affect the firm's long-run scale.

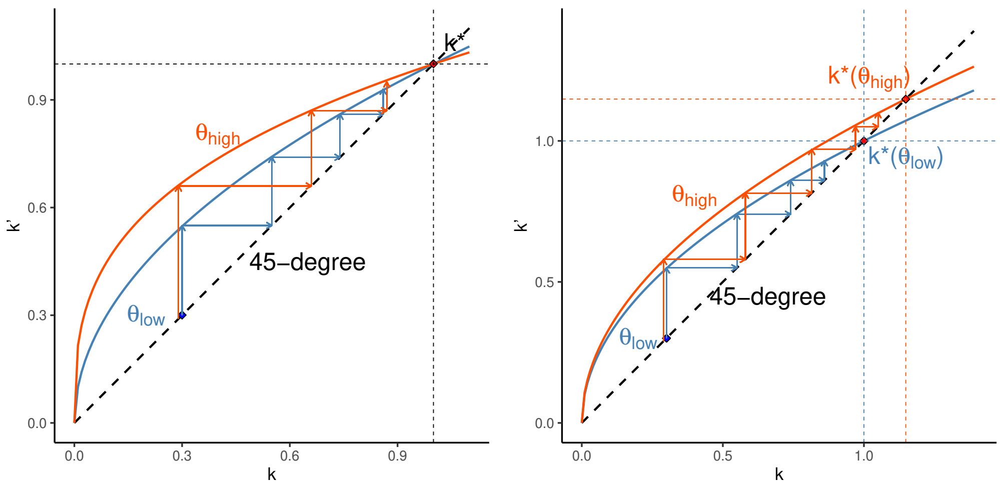
**Proposition 2** *Suppose that the economy is described by the special case defined in Definition 2.*

*(i) The financial constraints remain relevant in the long run (i.e.,* $\lambda^{LT} > 0$*) if and only if the intertemporal distortions are non-negligible:* $\epsilon(1-\tau) < 1$*. In that case*

$$\lambda^{LT} = \frac{1}{\beta}\left(\frac{1}{\epsilon(1-\tau)} - 1\right) > 0 \tag{1.21}$$

*(ii) In the long run, the gap between long-term capital* $k^{LT}$ *and the undistorted capital stock* $k^{opt}$

Figure 1.2: DYNAMICS OF FIRM'S LONG-TERM GROWTH




(a) Without Intertemporal Distortions $\epsilon(1-\tau)=1$

(b) With Intertemporal Distortions $\epsilon(1-\tau)<1$

Note: Given the erosion conditions, this plot shows the entrepreneur's growth dynamics. The numerical calibration of the parameters is conventional to an annual model: $\delta = 0.1$ stands for annual depreciation in the capital, $\alpha = 2/3$ stands for decreasing return to scale, $a = 2$ stands for labor share in production, $n_{low} = 1$ and $n_{high} = 2$ stand for smaller and larger entrepreneurs.

*is affected by the long-term shadow value of the financial constraint $\lambda^{LT}$ following*

$$\psi'(k^{LT}) - \psi'(k^*) = (1 + a - \theta)\lambda^{LT} > 0 \tag{1.22}$$

*where $\lambda^{LT}$ is defined in equation* (1.21)*, and $\psi'(k^*) = 1/\beta - (1-\delta)$.*

***Proof.*** *See Appendix F.*

The first point of Proposition 2 restates the above discussion. The second point describes in more detail how the distortion in long-term capital accumulation depends on the interaction between the intertemporal distortions (through $\lambda^{LT}$) and the collateral constraint (through $\theta$). The expression (1.22) shows clearly that the intertemporal distortions and the collateral constraint reinforce their respective impact on capital accumulation in the long run. An increase in $\theta$ increases the stock of capital for firms that are below their long-run scale compared to a firm that does not benefit from an increased $\theta$. This increase is temporary without intertemporal distortions as the long-run scale is unchanged. In the presence of intertemporal distortions, a firm that benefits from an increased $\theta$ converges to a larger scale.

### 1.3.4 Endogenous Selection into the Guarantee Program

In the final step, we analyze how increasing the firm's collateral from $\theta_{low}$ to $\theta_{high}$ and the participation cost affects the participation of firms in the guarantee program and discuss the potential aggregate implications.

We still consider the same special case above, except that the participation cost is now strictly positive $\xi > 0$. In this case, we examine which type of firm self-selects into the loan guarantee program. A firm decides to ask for the guarantee only if $\psi(k(\theta_{high}, n_t)) - \psi(k(\theta_{low}, n_t)) > \xi$, which we can approximate as follows:

$$\psi'(k(\theta_{low}, n_t)) \cdot n_t > \frac{\xi}{\Delta\theta} \tag{1.23}$$

where we used $k(\theta_{high}, n_t) - k(\theta_{low}, n_t) \simeq n_t \Delta\theta$ with $\Delta\theta \equiv \theta_{high} - \theta_{low}$. Importantly, considering the participation cost $\xi$ and the guarantee increment $\Delta\theta$, essentially two factors decide whether a firm joins the loan guarantee program: marginal productivity $\psi'(\cdot)$ and net worth $n_t$.

**Proposition 3** *Suppose that the economy is described by the special case defined in Definition 2, but now the participation cost is positive: $\xi > 0$.*

*(i) Given net worth $n_t$, more productive firms (with high marginal productivity $\psi'(\cdot)$) would be more likely to self-select into the program.*

*(ii) Given marginal productivity $\psi'(\cdot)$, firms with a high net worth $n_t$ would be more likely to self-select into the program.*

*(iii) Neither very large firms ($\psi'(\cdot) \to 0$) nor very small firms ($n \to 0$) will self-select into the program. The participation rate would be hump-shaped in firm size.*

***Proof.*** *(i), (ii), and (iii) are derived directly from Equation* (1.23).

The first and second points of Proposition 3 indicates that the self-selection into the guarantee program benefits the aggregate as resources are allocated to firms with a higher growth potential or with a higher total output. The third point shows that the limited guarantee resources are not allocated to firms which cannot afford it or firms which do not need it at all.

However, there is still a group of firms that are inefficiently constrained. These high-potential firms with high $\psi'(\cdot)$ but a low net worth $n_t$ may have low incentives to ask for a guarantee. In that case, the firm is limited in its capacity to increase its scale, which limits the incentives to pay the participation cost $\xi$. As a result, some small but productive firms may still not ask for a guarantee if the participation cost is high. Lower participation costs $\xi$ or an increase in the guaranteed ratio $\Delta\theta$ could potentially motivate these firms to participate.

### 1.3.5 Remarks on the Mechanisms and Predictions

This simple special case suggests that lowering participation costs or increasing the guaranteed ratio could push productive but small firms to participate in the LGPs and achieve a higher scale. We will test these policy suggestions in the fully parameterized quantitative model in Section 1.4 below. This simple model also provides three predictions on how LGPs should affect firm static production and financing choices in the short run and in the long run. First, firms with LGPs lower their unproductive cash holdings and increase their productive capital stock. Second, firms with LGPs achieve a persistent increase in their production scale if they face non-negligible intertemporal distortions. Third, the participation rate in the LGPs is hump-shaped over firm size because small firms and firms with lower marginal productivity self-select out of the loan guarantee program. We will confront these predictions to the data in Section 1.5 below.

## 1.4 Quantitative Analysis

We now assess quantitatively how short-term finance shapes firm financing and growth. We parameterize the model to our Moroccan firm-level data and LGP program, using information on guaranteed and non-guaranteed firms. The key parameters that capture financial frictions are set to match some cross-sectional and dynamic patterns observed in the data. We then find that the model can quantitatively account for the observed growth effect of LGPs in Morocco.

### 1.4.1 Institutional Background and Data

Our analysis merges loan-level data from Tamwilcom, a Moroccan government institution that provides SME-related loan guarantees, with firm-level data from Orbis.

**Collateral Requirements and Loan Guarantees in Morocco** Collateral requirements for loans are exceptionally high in Morocco. Approximately 84% of the loans in Morocco require collateral, as reported by World Bank (2013). Most collaterals are asset-based. To reduce potential inefficiency caused by such high collateral requirements, Tamwilcom, as a public financial institution under the supervision of the Central Bank of Morocco, *Bank Al-Maghrib*, cooperates with four leading banks which jointly cover an extensive credit network to provide loan guarantee programs to SMEs.[11] Firms that apply for bank loans at these four leading banks and which do not

[11] Tamwilcom (formerly *Caisse Centrale de Garanties*) has a long history as a credit institution dating back to 1949. Since its reform in 2012, Tamwilcom has focused on SME-related loan guarantees (Tamwilcom, 2013-2018). Our study focuses on the post-reform period from 2012 to 2018.

have insufficient collateral but still eligible for guarantees are transferred to Tamwilcom for further assessment. Once approved, the bank grants credit to qualified borrowers, and Tamwilcom underwrites a share of the loan.

Firms in Morocco are severely financially constrained. They have almost no access to long-term formal bank loans ($<$ 1% of total corporate credit) and very limited access to short-term formal bank loans ($\approx$ 20% of total corporate credit). To fulfill their liquidity needs, they rely heavily on non-bank loans ($\approx$ 80% of total corporate credit), such as trade credit. Meanwhile, they also stockpile cash holdings, equivalent to about 42% of total corporate credit and more than 200% of bank credit[12]. To alleviate such severe financial frictions, the Moroccan government started to provide extensive LGPs for short-term loans to SMEs in 2012. This loan guarantee program is the ideal setup to evaluate the impact of short-term external finance.

**Products of the Loan Guarantee Program** Among the range of products Tamwilcom offers, we focus on two main products catering to the firm's working capital needs: Damane Exploitation and Damane Express. Damane Exploitation targets medium-sized firms requesting a short-term loan of up to 18 months. Access to Damane Exploitation is subject to a firm size threshold, but this size threshold does not appear to be binding in the data.[13] The loan size varies substantially, ranging from 180 million dirhams to as small as 1 million dirhams. Tamwilcom guarantees 60% of the loan and requires a commission fee of 0.5% of the loan amount. Damane Express is a product that targets working capital loans and is designed explicitly for small firms.[14] It deals with loans below 1 million dirhams and provides a guarantee coverage of up to 70%. The commission fee is 0.5% for loans up to 12 months and 1.5% for those beyond 12 months.

Since both programs are designed to alleviate credit constraints of firms ranging from small to medium and jointly cover almost all SMEs in Morocco, we will pool both programs together as one in our analysis. In practice, firms self-select into different programs based on the size of their liquidity needs, which is implied by the smooth sales distribution for firms guaranteed under both products, as shown in Figure 1.11 in Appendix C.

**Tamwilcom Loan-level Data** The Tamwilcom loan-level database is a unique confidential database. It covers loans guaranteed by Tamwilcom through the short-term loan guarantee programs and provides information on 43,195 loans associated with 23,017 firms guaranteed by Tamwilcom from 2009 to 2019. The database includes firm identifiers (name, national ID, address, creation date) and loan characteristics (loan approval date, maturity, loan amount, guarantee amount, commission, and maturity). The total number of guaranteed loans amounts to 87

[12] All calculations are based on firm-level data in Orbis.

[13] The size threshold is 175 million dirhams ($\approx$ 18 million US dollars). Approximately 92% of firms in the program have sales of less than 100 million dirhams. Damane Exploitation was renamed Damane Atassyir in 2019.

[14] Damane Express is also associated with a much-simplified process and a fast approval period of 48 hours.

billion dirhams, which constitutes about 3.2% of the total short-term loans to SMEs in Morocco.[15]

**Orbis Firm-level Data** The Orbis firm-level database is a commercial database by Bureau van Dijk (BvD) that is widely used in economics research (Kalemli-Ozcan et al., 2015). For Morocco, BvD collects firm-level balance sheet data from Morocco's business register, the Office of Industrial and Commercial Property (OMPIC), and standardizes it to its global format. OMPIC maintains the central trade register of all firms under the supervision of the Moroccan government. It is the most comprehensive source of firm data in Morocco. Consequently, the data's coverage and quality of Orbis are well-suited for our analysis. Orbis covers firms throughout the period of our loan-level data from Tamwilcom.

We pair the Tamwilcom guarantee dataset with the Orbis balance sheet dataset to construct our final data sample. Details for the pairing procedure are in Appendix A. The targeted moments are computed using the sample of guaranteed and non-guaranteed firms.

### 1.4.2 Quantitative Specifications

We specify the model's functional forms in the quantitative analysis. First, we assume that the production function is the conventional Cobb-Douglas form:

$$F(k, l) = k^{\alpha} l^{\nu}, \quad \alpha + \nu < 1$$

Second, we follow Gopinath et al. (2017) to model the size-dependent collateral constraint:

$$b_{i,t} \leq \bar{b}_{i,t} \equiv \Theta(k_{i,t}) = \theta_0 k_{i,t} + \theta_1 \Psi(k_{i,t}) = \left[\theta_0 + \theta_1 \frac{\Psi(k_{i,t})}{k_{i,t}}\right] k_{i,t} \tag{1.24}$$

where $\Psi(k) = exp(\gamma k) - 1$ is an increasing and convex function of capital and $\theta_0$ and $\theta_1$ are parameters characterizing the borrowing constraint. In this micro-foundation, the $\Psi(\cdot)$ function denotes an increasing and convex cost firms incur from disrupting their productive capacity. In contrast to Gopinath et al. (2017), we introduce the elasticity $\gamma$ to change the convexity of the size-dependent component of the collateral constraint to provide additional freedom to match the moments in the Moroccan firm-level data.

[15] We exclude canceled guarantees and only consider the first guarantee in case of renewal.

### 1.4.3 Parameterization

We group parameters into two categories. The first category includes preference and technology parameters that are difficult to identify using our data. We fix these parameters using values that are standard in existing work. The second category includes parameters that determine the process for productivity, financial frictions, and LGPs. We pin down these parameters by requiring that the model fits the salient features of the Moroccan data.

Table 1.1: Fixed Parameters

| Parameter | Description | Value |
|---|---|---|
| ***Firms*** | | |
| $\alpha$ | Capital coefficient | 0.21 |
| $\nu$ | Labor coefficient | 0.64 |
| $\delta$ | Capital depreciation | 0.10 |
| $\phi$ | Capital adjustment cost | 4.0 |
| ***Households*** | | |
| $\beta$ | Discount factor | 0.96 |
| $\eta$ | Elasticity of intertemporal substitution | 1 |
| $\theta$ | Leisure preference | 2 |
| $\omega$ | Inverse Frisch | 0.5 |

Table 1.2: Fitted Parameters

| Parameter | Description | Value |
|---|---|---|
| ***Output Dynamics*** | | |
| $\rho_z$ | Persistence of TFP shock | 0.90 |
| $\sigma_z$ | Volatility of TFP shock | 0.06 |
| $\underline{n}_0$ | Net worth of entrants | 0.08 |
| $\epsilon$ | Survival rate | 0.91 |
| $\tau$ | Net worth erosion | 0.02 |
| ***Financial Frictions*** | | |
| $s$ | Share of formal bank loans | 0.20 |
| $\theta_0$ | Collateral constraint (size-irrelevant) | 0.01 |
| $\theta_1$ | Collateral constraint (size-dependent) | 0.26 |
| $\gamma$ | Collateral constraint (size-dependent) | 1.35 |
| ***Loan guarantee program*** | | |
| $\mu$ | Guaranteed loan commission fee | 0.005 |
| $\chi$ | Multiplier of LGP on loans | 2.5 |
| $\bar{\xi}$ | Upper bound of LGP fixed cost | 0.35 |

**Fixed Parameters** Table 1.1 lists the parameters that are calibrated from the literature. The frequency of the model is a year, so we set the discount factor $\beta = 0.96$ to match an annual interest rate of 4%. We assume log utility, which implies a unit elasticity of intertemporal substitution ($\eta = 1$). We set the Frisch elasticity of labor supply to 2, within the range of macro elasticities

Table 1.3: Target Moments

| Moments | Data | Model |
|---|---|---|
| ***Output Dynamics*** | | |
| 1-year autocorrelation of output | 0.89 | 0.89 |
| 3-year autocorrelation of output | 0.69 | 0.71 |
| 5-year autocorrelation of output | 0.53 | 0.56 |
| Size ratio of entrant relative to average | 17% | 16.4% |
| Annual exit rate of firms | 9.0% | 9.0% |
| ***Financial Frictions*** | | |
| Mean debt/asset ratio (non-guaranteed) | 51% | 39% |
| Mean debt/asset ratio (guaranteed) | 64% | 64% |
| Mean cash/asset ratio (non-guaranteed) | 22% | 21% |
| Mean cash/asset ratio (guaranteed) | 9% | 6% |
| Guaranteed loan/current liability ratio | 22% | 22% |
| ***Loan guarantee program*** | | |
| Guaranteed loan commission fee | 0.5% | 0.5% |
| Percentage of loan guaranteed | 60% | 60% |
| Percentage of firms participating LGP | 3.4% | 3.4% |

Note: This table reports the moments from both the Orbis firm-level database and the Tamwilcom loan-level database. Moments of *productivity* and *entry/exit* are from all the Moroccan firms in the Orbis firm-level database. The output level of a firm is measured by its sales. The size ratio of the entrant relative to the average is calculated using total assets. The exit rate is calculated for Moroccan firms from 2006 to 2017. Moments of *financial frictions* are calculated from both the sample of the Orbis firm-level database and the sample of the Tamwilcom guarantee-level database. The debt/asset ratio only includes current liability because, for SMEs in Morocco, the long-term debt is less than 1% of their total credit.

identified by Chetty et al. (2011), which implies an inverse Frisch $\omega = 0.5$. We then set leisure preference $\theta = 2$ to match the fact that households spend a third of their time working. On the firm side, we set the capital coefficient $\alpha = 0.21$ and the labor coefficient $\nu = 0.64$ to match a labor share of two-thirds and return to scale of 85%. Capital depreciates at a rate of $\delta = 0.10$ annually, and the capital adjustment cost is set to $\phi = 4.0$, which generates an average aggregate nonresidential fixed investment rate as in Bachmann, Caballero, and Engel (2013).

**Fitted Parameters** The second category of parameters, listed in Table 1.2, are jointly pinned down by the requirement that the model accounts for the firm-level facts in Morocco to match the moments in Table 1.3. First, the four parameters related to output dynamics: persistence of TFP shock $\rho_z$, volatility of TFP shock $\sigma_z$, the net worth of entrants $\underline{n}_0$, and survival rate $\epsilon$ jointly match the three moments of productivity persistence, the relative size ratio of entrants to an average firm, and the annual exit rate of firms in the data. Second, the loan guarantee parameters: commission fee $\mu$, multiplier of LGP on loans $\chi$, and share of guaranteed loans $s$ explicitly match the three corresponding moments of the commission fee, percentage of loan guaranteed, and guaranteed loan to current liability ratio in the data.

Then, we parameterize the other three financial friction parameters: the size-irrelevant collateral parameter $\theta_0$, the size-dependent collateral parameter $\theta_1$, and the net worth erosion parameter $\tau$ to jointly match the five moments of LGP participating rate, cash asset ratios, and debt asset ratios. The collateral constraint parameters and net worth erosion then jointly pin down the cash asset ratios and debt asset ratios of both the guaranteed and non-guaranteed firms. These four moments reflect that guaranteed firms have, on average, a 13 percentage point lower cash ratio and a 13 percentage point higher debt ratio.

The fitted exit rate $(1-\epsilon)$ and net worth erosion $\tau$ yield a high intertemporal distortion $(1-\epsilon)+\tau = 0.11$. Note that the exit rate $(1-\epsilon)$ is probably overestimated, as part of the firms that exit the dataset, in fact, fail to update their registration in the trade registry. However, as the average cash and debt ratios are determined mostly by the saving behavior of firms, the total level of intertemporal distortions is strongly disciplined by these variables. Therefore, the overestimation of the exit rate is compensated by the underestimation of the net worth erosion.[16]

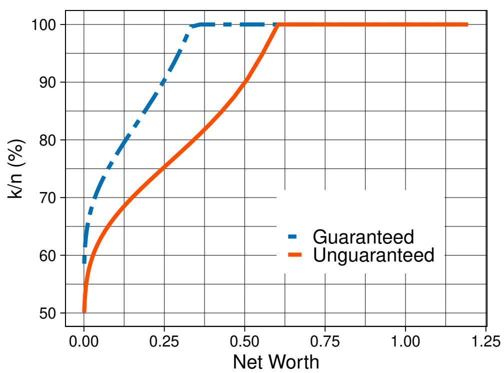
Finally, the upper bound of the LGP participation cost $\bar{\xi}$ mostly uniquely pins down the percentage of firms participating in LGP of 3.8%. The value of that upper bound implies an average cost of 0.175. This represents one-third of the average net worth. This number is large, implying that the barriers to participation remain important so that only large constrained firms or small lucky firms would have the chance to self-select into the program. This large number reflects an imperfect penetration of the guarantee program among small firms due to geographical barriers or the fact that many small firms do not have a relationship with the banking system. In 2023, many firms still did not have checking or deposit accounts (27% against 11% in the rest of the world, according to the World Bank Enterprise Survey) and did not interact with the banking system, which could increase the informational barriers. Many firms also cannot provide reliable financial information (only 24% of firms produce an annual financial statement reviewed by external auditors, compared to 46% in the rest of the world, according to the same source), which prevents them from submitting relevant information to banks. Administrative and cognitive costs related to the application procedure are also far from negligible. As an illustration, after 2018, when the guarantee allocation procedure was automated, the number of applications increased dramatically. Policies that lower the fixed cost by simplifying the application procedure, advertising the LGP, or diversifying the application channels are, therefore, one potential important margin to expand guarantee programs, which we discuss in the policy section.

[16] We later check that our parametrization is consistent with the estimated long-term effect of the guarantees. Besides, as we show below, a higher exit rate reduces the aggregate impact of the guarantee for a given total intertemporal distortion. Fitting $\epsilon$ in this way thus generates conservative parameters.

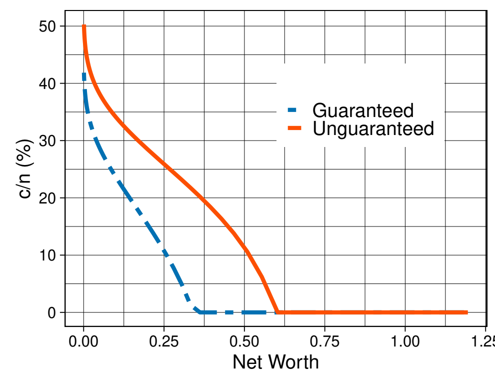
### 1.4.4 Model Implications

With the calibrated model, we show how the loan guarantee program affects the firm's static choice between capital and cash, the participation rate in the loan guarantee program, and the firm's long-term scale. We find that the quantitative model reproduces the mechanisms from the special case in Section 1.3.

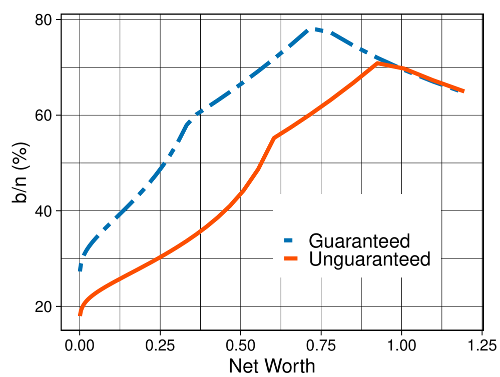
Figure 1.3: Reallocation of Resources from Cash to Capital



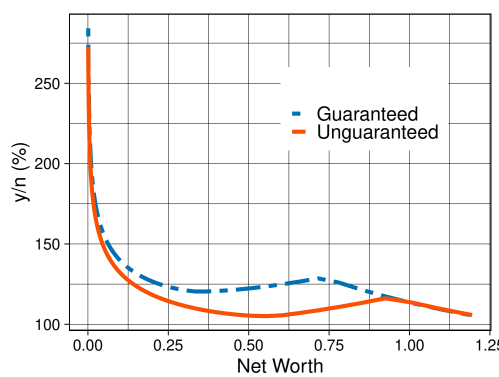

(a) Capital Policy




(b) Cash Policy




(c) Debt Policy



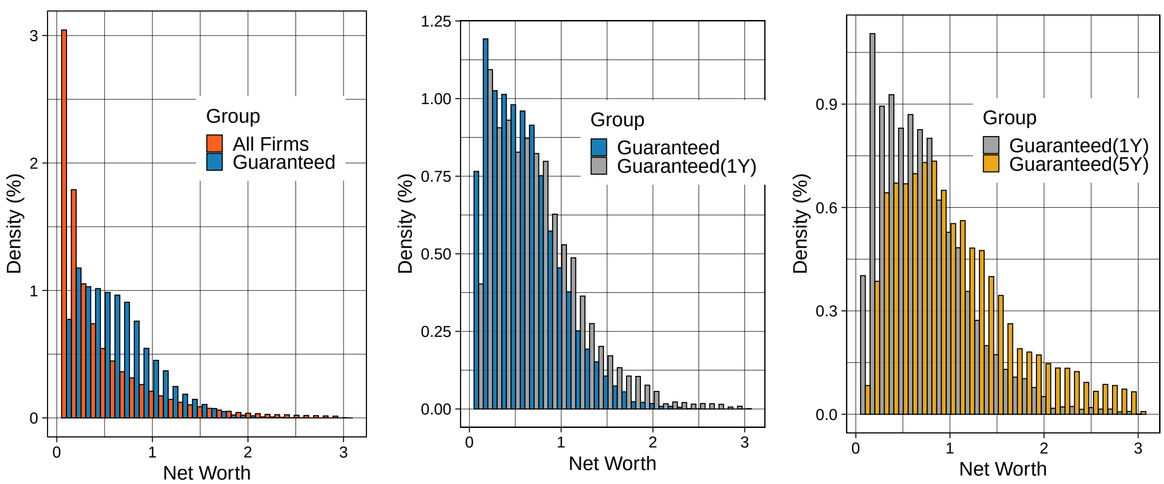

(d) Realized Output

Notes: This figure shows the capital, cash, debt, and output policies as a function of net worth for firms with median productivity. The blue line stands for guaranteed firms, and the red line stands for unguaranteed firms. Net worth is truncated at 1.25 because the measure of firms larger than 1.25 is tiny, and the decision rules are monotone in net worth beyond 1.25.

**Short-term Growth and Resource Reallocation** We first show the effects of short-term finance on the reallocation of resources from cash to capital. Figure 1.3 plots the optimal capital, cash, and debt policies and the realized output of guaranteed and unguaranteed firms. First, we compare the optimal policies among the dimension of net worth, focusing on unguaranteed firms. Due to their limited access to short-term external finance, smaller firms borrow less, hoard more cash, and accumulate less capital. When firms grow, their short-term financial constraints are relaxed, so they start borrowing more and lowering their cash holdings. Finally, if firms grow further, they become unconstrained.

Second, in Figure 1.3, we compare the policies and realized output of guaranteed firms versus unguaranteed firms. Generally, guaranteed firms with access to the guarantee accumulate more productive capital, lower their unproductive cash holdings, borrow more external debt, and produce more output per net worth. The changes are most significant for median-sized firms. The policies converge when firms become large and unconstrained. These effects on the reallocation of resources from cash to capital and its consequences are consistent with our empirical findings in Section 3.4 and our analytical findings in Section 1.3.2.

Figure 1.4: Distribution of Firm Long-term Scale



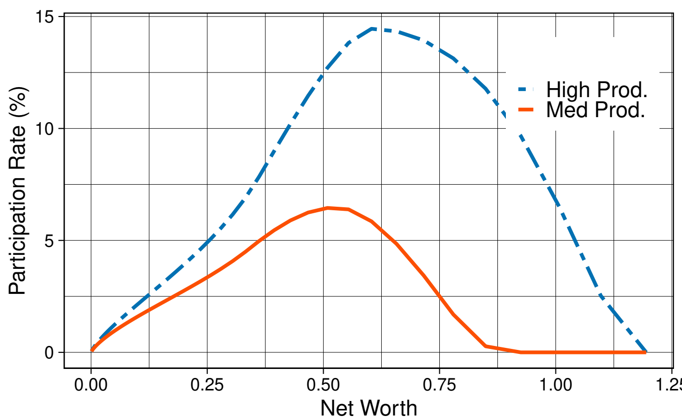

(a) LGP Self-Selection (b) One Year Growth (c) Five Year Growth

Notes: This figure shows the net worth distributions of all firms, of guaranteed firms upon their self-selection into the loan guarantee program, of guaranteed firms after one-year growth without exiting, and of guaranteed firms after five years, conditional on not exiting. It helps to distinguish the selection effect and growth effect. The mean net worth of the distributions are 0.51, 0.72, 0.85, and 1.27, respectively.

**Firms Long-Term Scale** We then show the effects of short-term finance on the long-term scale of firms. Figure 1.4 plots the distributions of all firms and guaranteed firms alone. More specifically, the distribution of guaranteed firms is shown in three stages. The first stage is the distribution of firms that self-select into the loan guarantee program before accessing the additional credit line. Given the random fixed participation cost, medium firms are more likely to be able to pay the fixed cost and enter the program, which is consistent with our data in Section 3.4. The difference between the distributions of *Guaranteed* and *All Firms* shows the selection effect of the loan guarantee program.

The second stage is the distribution of these guaranteed firms in the first stage after one year of growth without exiting. The additional credit line from the loan guarantee program helps them accumulate more net worth, so the distribution shifts to the right. Finally, we show the third stage, the distribution of the guaranteed firms after five years of guarantee without exiting. These firms grow significantly larger with a distribution shifting to the right. These effects on the long-term scale of firms are consistent with our empirical findings in Section 3.4 and our analytical findings in Section 1.3.3.

Figure 1.5: Participation In the Guarantee Program




Notes: This figure shows the proportion of firms participating in the LGP as a function of net worth for two different productivity levels. The blue line stands for the higher-productivity firms, and the red line stands for the lower-productivity firms (in this plot, the median productivity). Net worth is truncated at 1.25 because the measure of firms above 1.25 is tiny, and the decision rules are monotone beyond 1.25.

**Participation In the Guarantee Program** Figure 1.5 shows the participation probability for firms with two different productivity levels as a function of net worth. First, the participation rate is hump-shaped over the net worth dimension. The smallest firms hardly participate in the LGP programs since they can hardly support the fixed cost of participation, while the largest firms also hardly participate because they are much less financially constrained. As a result, the median-sized firms are the most engaged in the LGP. Second, high-productivity firms have a higher participation rate than low-productivity peers with the same net worth since these firms gain more from participating in the LGP and are more able to pay for the participation costs with their higher profits. These effects on the participation rate in the LGP are consistent with our empirical findings in Section 1.5.2 and our analytical findings in Section 1.3.4.

## 1.5 Model Validation with Firm-level Data

In this section, we empirically validate the findings illustrated in our quantitative model using Moroccan firm-level data. Specifically, we document empirical evidence on a guaranteed firm's enlarged production scale and reduced cash holdings.

Our empirical strategy combines nearest-neighbor matching with a difference-in-difference (DID) approach. The matching procedure is to find statistical twins for a guaranteed firm based on a series of time-varying and observable variables. Details on the matching procedure are in Appendix A and the summary statistics of the firm-level data are in Appendix C. The DID method controls for unobservable group-specific time effects, where the "group" refers to the guaranteed firm and its matched unguaranteed firms.

It is important to note that we do not interpret these results as causal evidence. Because we cannot fully control for time-varying unobservable drivers of selection into the program (like the firm productivity), the differences in outcome variables between guaranteed and matched firms cannot be fully interpreted as causal evidence. Our results are instead compared with the model, which accounts for selection since only firms that are productive enough enter the program (see Proposition 3). This will constitute our model validation.

### 1.5.1 Estimation Results on Short-term Growth and Long-Term Scale

We first try to validate the short-term growth and resource reallocation, as shown in Figure 1.3 and firms' long-term scale, as shown in Figure 1.4 through regressions. Our regression follows Brown and Earle (2017)'s matched DID setup:

$$\Delta Y_{igs} = \delta D_{it} + \lambda_{gs} + \epsilon_{igs}, \tag{1.25}$$

where $i$ indexes the firm, $g$ is the group (the guaranteed firm and its matched unguaranteed firms), $t$ is the guarantee year, and $s = t+1, t+2, t+3$ refers to three post-guarantee years. The dependent variable $\Delta Y_{igs}$ is the change in the selected outcome variable in the post-guarantee period compared to the year before obtaining the guarantee. It has the form $\Delta Y_{igs} = Y_{igs} - Y_{igt-1}$, where year $t-1$ is considered as the base year and $s = t+1, t+2, t+3$ refers to three post-guarantee years. We cannot estimate $t+5$ as in the model due to short data length. All $Y$ variables are in logs so that the dependent variable can be read as a growth rate. $D_{it}$ is a dummy variable indicating whether firm $i$ has been granted a guarantee in year $t$. $\lambda_{gs}$ are the group-year fixed effects, which control for the group-specific trend. We also include city-year fixed effects to control local credit demand and financial conditions.[17] Standard errors are clustered at the city level.

We first examine whether the guarantee is associated with an expansion in production scale, measured in sales growth, total asset growth, cost of employee growth, and fixed asset growth. Columns (1) to (3) of Table 1.4 report the estimation results for sales growth. Firm sales growth under a Tamwilcom guarantee increases by 13.5% in the first year, compared to the pre-guarantee period, relative to non-guaranteed firms. The impact is close to 12.5% in the third year after obtaining the guaranteed loan. This large and significant effect on sales indicates that the guarantee goes together with a firm's expansion in production. Columns (4) to (6) of Table 1.4 report the significant and positive association of the guarantee with total asset growth. This shows that the firm simultaneously increases its net worth. All in all, access to the loan guarantee coincides with a persistent increase in firm scale.

Table 1.4: Effects on Firm Sales and Total Assets

| | Sales Growth | | | Total Asset Growth | | |
|---|---|---|---|---|---|---|
| | (1) t+1 | (2) t+2 | (3) t+3 | (4) t+1 | (5) t+2 | (6) t+3 |
| Guaranteed | 0.135$^{***}$ | 0.101$^{***}$ | 0.125$^{***}$ | 0.092$^{***}$ | 0.092$^{***}$ | 0.166$^{***}$ |
| | (0.010) | (0.021) | (0.020) | (0.011) | (0.021) | (0.014) |
| $N$ | 18503 | 10610 | 5585 | 18959 | 11018 | 5952 |
| adj. $R^2$ | 0.332 | 0.315 | 0.357 | 0.299 | 0.264 | 0.268 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of ("Guaranteed") treatment from the DID regression (1.25). Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variable "Sales Growth" is the log difference between sales in year $t+1$, $t+2$ or $t+3$, and sales in year $t-1$. The dependent variable "Total Assets Growth" is the log difference between total assets in year $t+1$, $t+2$ or $t+3$, and total assets in year $t-1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the city level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

We then examine how guaranteed firms change their production inputs in Table 1.5. We use the variable "cost of employee growth" to detect changes in a firm's hiring since we do not have good coverage for the number of employees in the Orbis database. As Table 1.5 shows, labor costs increase by 11.5% in the year following the granting of a guarantee relative to non-guaranteed firms and remain 10.6% and 10.5% higher in the two following years. Along with the increase

[17] Firm-level fixed effects are not included since our dependent variable has differenced out any individual fixed effects relevant to the outcome. We do not include fixed effects for sector, year, and size because we use group fixed effects. A group consists of one guaranteed firm and five matched unguaranteed firms from the same sector, year, and size category. Therefore, the group fixed effects account for fixed effects across these three dimensions. Further details are provided in in Appendix A.

Table 1.5: Effects on Firm Production Inputs

| | Cost of Employee Growth | | | Fixed Asset Growth | | |
|---|---|---|---|---|---|---|
| | (1) t+1 | (2) t+2 | (3) t+3 | (4) t+1 | (5) t+2 | (6) t+3 |
| Guaranteed | 0.115*** | 0.106*** | 0.105*** | 0.116** | 0.230*** | 0.241** |
| | (0.015) | (0.023) | (0.025) | (0.037) | (0.062) | (0.079) |
| $N$ | 17852 | 10422 | 5416 | 18344 | 10624 | 5760 |
| adj. $R^2$ | 0.252 | 0.223 | 0.239 | 0.183 | 0.174 | 0.236 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of ("Guaranteed") treatment from the DID regression (1.25). Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variable "Cost of Employee Growth" is the log difference between labor costs in year $t+1$, $t+2$ or $t+3$, and labor costs in year $t-1$. The dependent variable "Fixed Asset Growth" is the log difference between fixed tangible assets in year $t+1$, $t+2$ or $t+3$, and fixed tangible assets in year $t-1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the city level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

in the wage bill, guaranteed firms also experience increases in fixed tangible assets after the treatment, according to Table 1.5. This variable is a good proxy for investment in productive assets (Amamou, Gereben, and Wolski, 2020). It shows that guaranteed firms allocate more resources to long-term productive assets, consistent with an expansion in productive capacities.

Table 1.6: Effects on Firm Balance Sheet

| | Current Liability Growth | | | Cash Growth | | |
|---|---|---|---|---|---|---|
| | (1) t+1 | (2) t+2 | (3) t+3 | (4) t+1 | (5) t+2 | (6) t+3 |
| Guaranteed | 0.131*** | 0.122*** | 0.167*** | -0.061 | -0.210*** | 0.088 |
| | (0.014) | (0.015) | (0.025) | (0.069) | (0.054) | (0.070) |
| $N$ | 19299 | 11171 | 5969 | 18761 | 10683 | 5814 |
| adj. $R^2$ | 0.252 | 0.276 | 0.243 | 0.321 | 0.304 | 0.291 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of ("Guaranteed") treatment from the DID regression (1.25). Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variable "Current Liability Growth" is the log difference between current liabilities in year $t+1$, $t+2$ or $t+3$, and current liabilities in year $t-1$. The dependent variable, "Cash Growth," is the log difference between cash and cash equivalents in years $t+1$, $t+2$ or $t+3$, and cash and cash equivalents in years $t-1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the city level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

We then explore changes in the firm's balance sheet, including current liabilities and cash, summarized in Table 1.6. There is a persistent 12-17% increase in current liabilities in guaranteed firms relative to control firms. This arises naturally from the buildup of current liabilities from newly granted working capital loans. Conversely, we observe stagnation and decline (though only significant for the period $t + 2$) in cash for guaranteed firms. The decrease in cash holdings is significantly negative in $t + 2$ and not significant in the other two years. Combined with the guaranteed firms' sharp increase in total assets documented in Table 1.4, this result implies that cash holdings decrease relative to the production scale. To further validate our mechanism, we test for cash ratio changes in Table 1.7. The results are consistent with our predictions.

Table 1.7: Effects on Firm Cash Ratio

| | Cash Ratio Growth | | |
|---|---|---|---|
| | (1) t+1 | (2) t+2 | (3) t+3 |
| Guaranteed | -0.004$^{+}$ | -0.010$^{*}$ | -0.008$^{*}$ |
| | (0.002) | (0.004) | (0.004) |
| $N$ | 18766 | 10716 | 5818 |
| adj. $R^2$ | 0.215 | 0.161 | 0.126 |
| Group × Year FE | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes |

Note: This table reports the coefficients of ("Guaranteed") treatment from the DID regression (1.25). Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variable, "Cash Ratio Growth," is the difference between cash ratio in years $t + 1$, $t + 2$ or $t + 3$, and cash and cash ratio in years $t - 1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the city level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

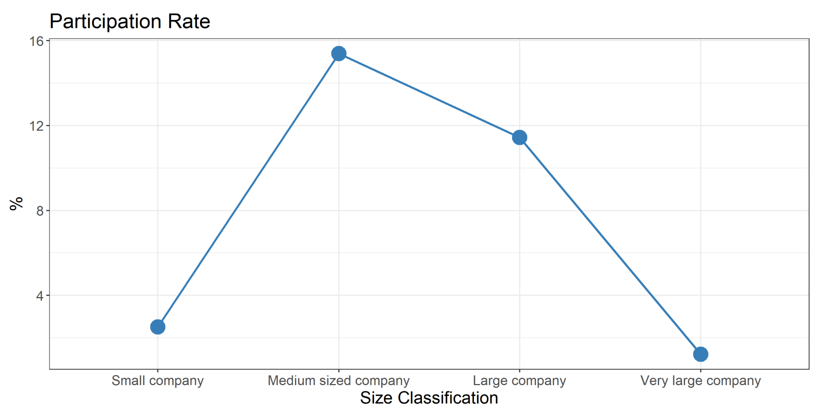
**Robustness Checks** We also perform a series of robustness checks, which are reported in Appendix E. We check for (1) changing the number of pre-treatment years in matching, (2) correcting the data attrition bias, (3) emphasizing cash in matching, and (4) including propensity score in Mahalanobis distance. The estimation results remain mostly unchanged.

### 1.5.2 Participation in the Loan Guarantee Program

We finally examine the propensity of firms to participate in the loan guarantee program. Figure 1.6 represents the percentage of Orbis firms that have been identified as benefiting from a guarantee, which proxies for the participation rate. The participation rate is hump-shaped, as both small and very large companies have a low rate compared to medium and large firms. This hump-shaped distribution is maintained when looking at the participation rate by total asset quantile bins. More details are in Appendix D.

It is important to note that these results are biased by the size-dependent probability of Tamwilcom firms being paired with Orbis firms, as this probability increases in firm size. In particular, the probability of a small firm being paired may be underestimated. However, the probability of a small firm being paired is 1.7 times lower than that of a medium firm (30% versus 52%). This cannot explain the strong difference in the estimated participation rate. Moreover, this concern is valid only if we have a lot of false negatives in our dataset, that is, if many firms identified as not having a guarantee do, in fact, have a guarantee. While this risk cannot be ruled out altogether, it is minimized by our thorough pairing procedure described in Appendix A. In fact, the most likely reason for a failed pairing is that a guaranteed firm is not in the Orbis dataset. In that case, the participation rates by firm size will not be significantly underestimated.

Figure 1.6: Participation Rate by Size




Notes: This figure shows the participation rate of our Orbis firm sample in Morocco by size. The size classification defines very large companies as those with an operating revenue larger than 100 million euros, total assets larger than 200 million euros, and more than 1000 employees. Large companies have an operating revenue larger than 10 million euros, total assets larger than 20 million euros, and more than 150 employees. Medium-sized companies have an operating revenue larger than 1 million euros, total assets larger than 2 million euros, and more than 15 employees. All the remaining companies are defined as small. The participation rate is calculated as the ratio of guaranteed firms to the total number of firms existing in Orbis.

### 1.5.3 Quantitative Model Validation with Regressions

Our empirical estimates from Tables 1.4 to 1.7 cannot be interpreted causally because unobservable time-dependent variables, like productivity, may drive selection. We, however, can use these estimates as “untargeted moments” to validate our model. To do so, we simulate a panel of five million firm-year observations and conduct the same empirical procedure of nearest neighbor matching and DID. We refer to these results as "Matched-DID". We also show how "Naive OLS" regressions would overestimate the effects of short-term loan guarantees since relatively larger and higher growth SME firms are more likely to self-select into loan guarantee programs.

Table 1.8: Quantitative Model Validation with Empirical Analysis

| Effects of Credit Guarantee | Matched-DID (Data) | | | Matched-DID (Model) | | | Naive OLS (Model) | | |
|---|---|---|---|---|---|---|---|---|---|
| | t+1 | t+2 | t+3 | t+1 | t+2 | t+3 | t+1 | t+2 | t+3 |
| $\Delta log(Sales)$ | 0.135*** | 0.101*** | 0.125*** | 0.187*** | 0.163*** | 0.150*** | 0.524*** | 0.511*** | 0.461*** |
| | (0.010) | (0.021) | (0.020) | (0.013) | (0.016) | (0.018) | (0.007) | (0.010) | (0.012) |
| $\Delta log(TotalAsset)$ | 0.092*** | 0.092*** | 0.166*** | 0.143*** | 0.147*** | 0.139*** | 0.257*** | 0.238*** | 0.215*** |
| | (0.011) | (0.021) | (0.014) | (0.010) | (0.013) | (0.014) | (0.006) | (0.008) | (0.009) |
| $\Delta log(CurrentLiability)$ | 0.131*** | 0.122*** | 0.167*** | 0.191*** | 0.196*** | 0.187*** | 0.391*** | 0.371*** | 0.341*** |
| | (0.014) | (0.015) | (0.025) | (0.012) | (0.016) | (0.019) | (0.008) | (0.011) | (0.013) |
| $\Delta log(Cash)$ | -0.061 | -0.210*** | 0.088 | -0.451*** | -0.445*** | -0.632*** | -1.772*** | -1.787*** | -1.545*** |
| | (0.069) | (0.054) | (0.070) | (0.051) | (0.063) | (0.077) | (0.049) | (0.061) | (0.065) |

Note: "Matched-DID (Data)" is taken from our empirical analysis above, and "Matched-DID (Model)" is obtained by applying exactly the same method to our model simulated data. "Naive OLS (Model)" runs the following OLS regression: $\Delta Y_{is} = \delta D_{it} + \gamma_z' Z_{is-1} + \gamma_i + \gamma_s + \epsilon_{is}$, where $i$ is the firm, $t$ is the guarantee year, and $s$ is the year ahead of the guarantee year. The dependent variable $\Delta Y_{it}$ is the same as in the Matched-DiD. Similarly, $D_{it}$ is a dummy variable indicating whether firm $i$ has been granted a guarantee in year $t$, and $gamma_j$ and $\gamma_t$ are the firm and year fixed effects, respectively. $Z_{js-1}$ is the group of control variables that are used in the matching process for the regression (1.25). Significance level: $^+$ $p < 0.10$, $^*$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$. Robust standard errors are in parentheses.

We show the results in Table 1.8. In the table, "Matched-DID (Data)" is taken from our empirical analysis above, and "Matched-DID (Model)" is obtained by applying exactly the same method to our model-simulated data. "Naive OLS (Model)" runs the following OLS regression: $\Delta Y_{is} = \delta D_{it} + \gamma_z' Z_{is-1} + \gamma_i + \gamma_s + \epsilon_{is}$, where $i$ is the firm, $t$ is the guarantee year, and $s$ is the year ahead of the guarantee year. The dependent variable $\Delta Y_{it}$ is the same as in the Matched-DiD. Similarly, $D_{it}$ is a dummy variable indicating whether firm $i$ has been granted a guarantee in year $t$, and $\gamma_j$ and $\gamma_t$ are the firm and year fixed effects, respectively. $Z_{js-1}$ is the group of control variables that are used in the matching process for the regression (1.25).

Our model matches the empirical analysis results well both in the magnitude and significance levels for sales, total assets, and current liabilities. Although the model over-predicts the effects on cash growth relative to our empirical findings, the directions are mainly consistent. In contrast, the effects of credit guarantee are highly overestimated in the naive OLS regression.

## 1.6 Policy Counterfactuals and Aggregate Implications

We finally conduct two groups of policy counterfactuals to demonstrate the aggregate implications of liquidity constraints. More specifically, we show how expanding the short-term loan guarantee program could further reduce liquidity constraints and promote firm growth in Morocco. The loan guarantee program can be expanded mainly along two dimensions, as our analysis suggested in Section 1.3.4: the fixed participation cost and the guaranteed ratio. We show below in this Section how much firm growth we can achieve by further relaxing these restrictions in four counterfactuals. We focus on both the aggregate output expansion and the coverage and "inclusiveness" of the loan guarantee program.

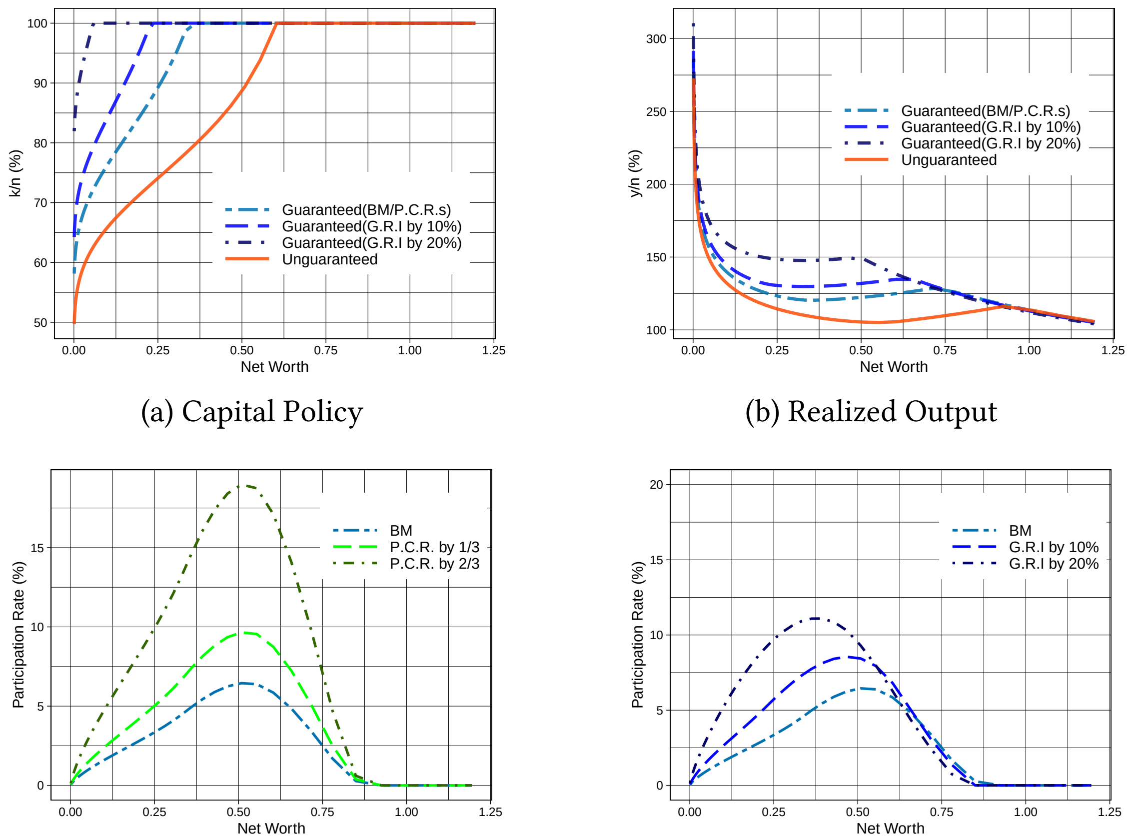
### 1.6.1 Policy Counterfactuals

We show four counterfactuals with realistic alternative participation costs and guaranteed ratios that stand for LGP expansions.

**Participation Cost Reduction (PCR)** First, in the original benchmark, the upper bound of the participation cost ($\bar{\xi}_{bm} = 0.35$) gives us a participation rate of 3.4%. More importantly, the fixed cost is an overhead cost that is relatively expensive for smaller entrepreneurs, as our analytical analysis suggested in Section 1.3.4. We explore two counterfactuals below ($PCR{\downarrow}_{by\,\frac{1}{3}}$ and $PCR{\downarrow}_{by\,\frac{2}{3}}$), where we cut the upper bound of the participation cost to two-thirds ($\bar{\xi}_{PCR\downarrow_{\text{by }\frac{1}{3}}} = 0.233$) and to one-third ($\bar{\xi}_{PCR\downarrow_{\text{by }\frac{1}{3}}} = 0.117$). These could be understood in the real world as the government requiring fewer financial documents, simplifying the evaluation procedures, directly assisting the application, and subsidizing applications to the LGP.

**Guaranteed Ratio Increment (GRI)** Second, in the original benchmark, the guaranteed ratio is 60%, which gives a bank loan multiplier of $\chi_{bm} = \frac{100\%}{100\%-60\%} = 2.5$. This is lower than guaranteed ratios in many other countries. For instance, in Kazakhstan, it is equal to 70%; in India, it is equal to 75%; in Indonesia and Japan, it is equal to 80% according to Yoshino and Taghizadeh-Hesary (2019). We explore two counterfactuals below ($GRI{\uparrow}_{by\,10\%}$ and $GRI{\uparrow}_{by\,20\%}$), where the guaranteed ratio goes up to 70% and 80%. Correspondingly, the new multipliers are $\chi_{GRI\uparrow_{\text{by }10\%}} = \frac{100\%}{100\%-70\%} = 3.33$ and $\chi_{GRI\uparrow_{\text{by }20\%}} = \frac{100\%}{100\%-80\%} = 5.00$. Participating firms have thus more relaxed financial constraints. Moreover, the increased guaranteed ratio could potentially incentivize more firms to participate.

Figure 1.7: Effects on Firm Financing, Output, and Participation




(a) Capital Policy

(b) Realized Output

(c) Participation Rate (Scale I)

(d) Participation Rate (Scale II)

Notes: This figure shows the capital and output policy, as well as the participation rate, as a function of net worth for firms with the median productivity level. In plot (a) of the capital policy, the k/n ratio will essentially decrease as the net worth grows over a certain scale. However, we do not show these patterns since we focus on small and median-sized firms.

### 1.6.2 Firm-level and Aggregate Implications

**Implications for Firm Decisions** Figure 1.7 shows the effects of LGP expansions on firm decisions. We mainly focus on how the expansions of LGP affect the choices of capital, cash, output, and participation in the LGP as a function of firm size. First, regarding capital and cash choices, the PCR has no effect conditional on participation. In contrast, the GRI would benefit smaller guaranteed firms more, as shown in plot (a). Second, the same patterns also hold in terms of output, as shown in plot (b).

Third, both types of policies affect the participation rate. Plot (c) shows the changes in the participation rate of the PCR. Lowering the fixed participation cost would increase the participation rate across all firm sizes. Median-sized firms benefit the most and increase their participation even more. Plot (d), on the other hand, shows the changes in participation rate following a GRI. Since an increased guaranteed ratio changes the capital policy and the output, which benefits the smaller firms more, the participation rate increments are skewed towards smaller firms. These distributional results show that although all counterfactuals benefit firm financing in general, the effects are quite different across firms of different sizes.

Table 1.9: Aggregate Implication of LGP Expansion

| **Model Outcomes (%)** | **Benchmark** | $PCR\downarrow_{by\ \frac{1}{3}}$ | $PCR\downarrow_{by\ \frac{2}{3}}$ | $GRI\uparrow_{by\ 10\%}$ | $GRI\uparrow_{by\ 20\%}$ |
|---|---|---|---|---|---|
| ***Penal A: Firm Financing*** | | | | | |
| LGP participation rate | 3.4 | 5.1 | 10.1 | 5.0 | 7.6 |
| Guaranteed credit/total credit | 1.3 | 2.0 | 3.8 | 3.1 | 8.0 |
| Mean cash/asset ratio (guaranteed) | 6.0 | 6.0 | 6.0 | 2.9 | 0.1 |
| Mean cash/asset ratio (all firms) | 20.0 | 19.8 | 19.3 | 19.6 | 18.8 |
| Mean debt/asset ratio (guaranteed) | 64.3 | 64.2 | 64.0 | 72.1 | 85.8 |
| Mean debt/asset ratio (all firms) | 40.0 | 40.3 | 41.3 | 40.8 | 42.7 |
| ***Penal B: Aggregate Outcomes*** | | | | | |
| Δ Total Credit | n.a. | 0.28 | 1.25 | 0.60 | 1.63 |
| Δ Aggregate TFP | n.a. | 0.04 | 0.14 | 0.08 | 0.25 |
| Δ Total Output | n.a. | 0.01 | 0.10 | 0.05 | 0.09 |
| Δ Total Consumption | n.a. | 0.06 | 0.29 | 0.15 | 0.47 |
| Δ Welfare | n.a. | 0.04 | 0.15 | 0.08 | 0.25 |

Note: This table reports the counterfactual results. The results are reported in two groups: (1) firm financing, which shows how the counterfactual changes the financing patterns of firms in the model, and (2) economic outcomes, which show how the counterfactual changes the aggregate economic conditions. The Benchmark stands for our benchmark calibration of both the fixed participation cost and the guaranteed ratio ($\bar{\xi}_{bm} = 0.26$, $\chi_{bm} = 2.5$). Each of the four counterfactuals changes only one parameter, reflecting participation cost reduction or guaranteed ratio increment.

**Aggregate Implications** Table 1.9 shows the effects of LGP expansions on aggregate financing, economic growth, and welfare. All four counterfactuals improve aggregate TFP, output, and consumption. Participation cost reductions significantly change firm financing patterns. The participation rate increases from 3.4% to 5.1% and 10.1%, and the ratio of guaranteed credit in the economy increases from 1.3% to 2.0% and 3.8%, respectively. The fixed participation cost changes do not substantially affect the guaranteed firm cash and debt ratios. However, more firms are guaranteed, so the average credit ratio increases, and the average cash ratio decreases. Given that most firms that benefit from these policies are small and medium firms, the changes in aggregate outcomes are not substantial at first glance. Nonetheless, we find that reducing the fixed participation cost increases aggregate TFP, total output, consumption, and welfare. Considering the small changes in total credit, the gains are substantial.

Increasing the guaranteed ratio also significantly changes firm financing patterns. The participation rate increases from 3.4% to 5.0% and 7.6%, and the ratio of guaranteed credit in the economy increases from 1.3% to 3.1% and 8.0%, respectively. Contrary to the PCR counterfactuals, guaranteed firms substantially decrease their cash ratio and increase their debt ratio. This reduces the average cash ratio and significantly increases the debt ratio. Finally, we also find that increasing the guaranteed ratio significantly increases aggregate TFP, total output, consumption, and welfare. As for PCR, considering the small changes in total credit, the gains are substantial.

It is worth noticing that two counterfactuals $PCR\downarrow_{by\,\frac{1}{3}}$ and $GRI\uparrow_{by\,10\%}$ almost increase participation by the same amount, but they have different aggregate outcomes. Because $GRI\uparrow_{by\,10\%}$ affects credit access over the whole distribution of firms, and not only self-selection into the program, total credit increases more, as well as output and welfare. In that sense, a guaranteed ratio increment kills two birds with one stone: it can both increase participation and total output. However, participation cost reduction is more "cost-effective" if the objective is to increase participation without significantly increasing the guaranteed portfolio.

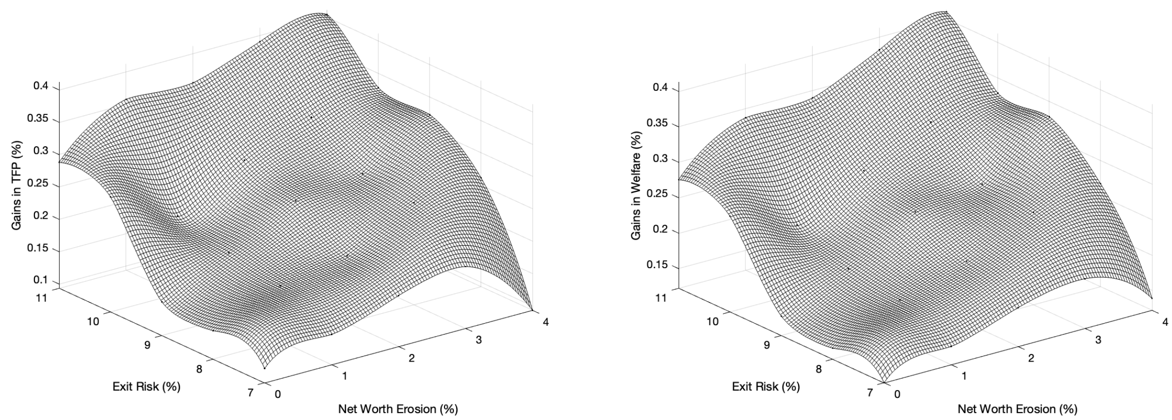
Unfortunately, we cannot directly discuss the exact cost-benefit analysis of such loan guarantee program expansions since we do not know exactly the direct and indirect costs of such expansions in the data, that is, how much direct organization costs the government pays to reduce the participation cost or how much indirect costs the government pays to increase the guaranteed ratio. We leave these to future work due to data limitations.

### 1.6.3 The Essential Role of Intertemporal Distortions

Finally, we examine quantitatively how intertemporal distortions determine the aggregate impact of relaxing liquidity constraints. In Section 1.3.3, we show in Proposition 2 that the long-term growth depends on the interaction between the intertemporal distortions and the collateral constraint. Without intertemporal distortions, the effect of short-term LGPs is temporary as the long-run scale is unchanged. Below, we show how the aggregate long-term economic impacts of a guaranteed ratio increment, demonstrated with the counterfactual of guaranteed ratio increment by 20% above, change under variations of both intertemporal distortions.

First, both intertemporal distortions determine the effects of the LGP expansion. Figure 1.8 shows how the intertemporal distortions affect the aggregate impact of a higher guarantee ratio. We consider alternative net worth erosion $\tau$ and exit risk $(1-\epsilon)$ around the benchmark specification ($\tau = 2\%, 1-\epsilon = 9\%$). The impact on TFP and welfare is lower if either intertemporal distortion is reduced. This arises from a smaller scale effect, as discussed earlier. In other counterfactuals, with higher net worth erosion and exit risk, the effect of the policy on output, TFP, and welfare is higher than in the benchmark because of the more severe intertemporal distortions.

Figure 1.8: The Role of Intertemporal Distortions




(a) Effects on Productivity

(a) Effects on Welfare

Notes: This figure shows how the intertemporal distortions affect the aggregate impact of a higher guarantee ratio. We consider alternative net worth erosion $\tau$ and exit risk $1-\epsilon$ around the benchmark specification ($\tau = 2\%, 1-\epsilon = 9\%$). The impact on TFP and welfare is more than halved if both intertemporal distortions are reduced by 2%. This arises from a smaller-scale effect, as discussed earlier. In other counterfactuals, with higher net worth erosion and exit risk, the effect of the policy on output, TFP, and welfare is higher than in the benchmark because of the more severe intertemporal distortions.

Table 1.10: Effect under alternative intertemporal distortions

| **Model Outcomes** | **Benchmark** ($\tau = 0.02, \epsilon = 0.91$) | **Lower** $\tau$ ($\tau = 0, \epsilon = 0.91$) | **Higher** $\epsilon$ ($\tau = 0.02, \epsilon = 0.93$) |
|---|---|---|---|
| Changes in Aggregate TFP (%) | 0.25 | 0.14 | 0.17 |
| Changes in Total Welfare (%) | 0.25 | 0.15 | 0.18 |

Note: This table reports the effects of policy $\mathrm{GRI}\uparrow_{\text{by }20\%}$ with different assumptions on $\tau$ and $\epsilon$ to illustrate the differences between the roles of net worth erosion and exit risk.

Second, net worth erosion matters more than exit risk. Table 1.10 shows the aggregate impact of a 20% higher guarantee ratio under two alternative assumptions on $\tau$ and $\epsilon$ from Figure 1.10. We first consider a lower $\tau$ ($\tau = 0$ as opposed to $\tau = 0.02$ in the benchmark) while $\epsilon$ is unchanged. The impact on output and welfare dropped by about 45% (0.25% to 0.14%). This arises from a smaller scale effect, as discussed above. In another counterfactual, we consider a higher $\epsilon$ ($\epsilon = 0.93$ as opposed to $\epsilon = 0.91$ in the benchmark) while $\tau$ remains unchanged. The effect of the policy on TFP and welfare dropped by about 30% (0.25% to 0.18%). Even though the lower $\tau$ and higher $\epsilon$ counterfactuals generate the same reduction in intertemporal distortion $(1-\tau)\epsilon$, the effects of the policy are stronger in the high $\epsilon$ counterfactual than in the low $\tau$ counterfactual.

With a higher $\epsilon$, firms exit less frequently, which gives them time to reach their long-run scale. This reinforces the long-term scale channel as opposed to the case with a lower $\tau$. However, the effect of the policy remains lower than in the benchmark because of the milder intertemporal distortions. LGP effects are thus particularly strong when intertemporal distortions come from net worth erosion as opposed to when they come from firm exit risk.

## 1.7 Conclusion

In this paper, we study the effect of short-term finance and how relaxing liquidity constraints affects firm growth and the aggregate economy.

We first build a heterogeneous firm model in which firms face collateral and working capital constraints and show the theoretical predictions. In the model, constrained firms preserve many resources in unproductive cash instead of productive capital to finance short-run working capital. A loan guarantee program mitigates credit constraints by inducing firms to reduce their cash holdings and expand their production scale. Additionally, a loan guarantee program generates a permanent increase in production scale due to intertemporal distortions.

We then take the model to a unique firm-level dataset of a credit guarantee program in Morocco. The model matches Moroccan firm-level moments well and replicates the patterns of our empirical findings: we show that firms with guaranteed loans expand their production scale homogeneously and persistently with increased labor and capital inputs and decreased cash ratios, which is consistent both qualitatively and quantitatively with the model. We then conduct counterfactual analyses to relax the severity of the short-term financial constraints. The gains from relaxing the severity of the short-term financial constraints by expanding the loan guarantee programs are substantial in terms of firm growth and welfare.

# Empirical Appendix

## A Sample Construction and Statistics

**Sample Cleaning** We clean the Orbis firm-level data following the standard procedure as in Kalemli-Ozcan et al. (2015): (i) the data series deflated by the Moroccan national GDP deflator from the World Bank (2007 base); (ii) the entire series of company data is dropped if total assets, sales, tangible fixed asset are negative in any year; (iii) values of zero are dropped for all financial variables; (iv) the series are winsorized by year at 1% following Amamou, Gereben, and Wolski (2020); (v) as a final step, firms that are in the finance and insurance, public administration and utilities sectors are excluded since firms are not eligible for Tamwilcom guarantees.

**Merge Databases** We pair the Tamwilcom guarantee dataset and the Orbis balance-sheet dataset in four rounds. In the first round, considering that the firm's registered ID with the chamber of commerce is not unique across regions, a unique combination of two variables of national ID and date of firm creation is applied to conduct the first pairing round. This yields good pairing results owing to the good coverage of both variables. In the second round, we use the firm's national ID and name as a unique combination. As a first step, redundant elements in firm names are trimmed away, such as STE, SARL, and Société. With more compact firm names, the Levenshtein distance between the names of two firms is calculated to locate the closest match. A string distance of up to two generally indicates a good match. The third pairing round relies on combining the firm name and address. Paired results from this step only yield a small number of matches. The final round is based on the firm's name and the date of firm creation; the pairing rate is also low.

Table 1.11: Summary Statistics of Tamwilcom-Guaranteed Firms: Whole Sample vs. Merged Sample

| Statistics | Guaranteed Amount | | Guaranteed Loan | | Sales | |
|---|---|---|---|---|---|---|
| Sample | Whole | Merged | Whole | Merged | Whole | Merged |
| Mean | 545 | 663 | 967 | 1,162 | 14,610 | 15,949 |
| Std | 1,336 | 1,467 | 3,401 | 3,598 | 28,120 | 28,314 |
| Min | 2 | 4 | 3 | 5 | 3 | 3 |
| 25% | 35 | 42 | 50 | 60 | 775 | 1,148 |
| Median | 105 | 140 | 150 | 200 | 3,219 | 4,462 |
| 75% | 400 | 560 | 550 | 800 | 14,176 | 17,039 |
| Max | 10,000 | 10,000 | 190,000 | 190,000 | 163,235 | 163,235 |

Notes: This table reports summary statistics of three variables (sales, guaranteed loan, and guaranteed amount) from the Tamwilcom sample and the merged sample between the Tamwilcom database and Orbis. All variables are in thousands of Moroccan Dirhams.

**Sample Statistics** After linking the Tamwilcom database to Orbis, we can identify 11,344 out of 23,017 guaranteed firms in the Orbis database, implying a rate of the successful pairing of 49.3%. Further participation of identified and unidentified guaranteed firms in Orbis shows that the two groups have similar characteristics. As shown in Table 1.11, loan amount, guaranteed amount, and sales reported by Tamwilcom are comparable for the two groups. One expected difference is that firm size is slightly higher for the subset of guaranteed firms that have been merged with Orbis data. This is because small firms usually report less complete information, making it less likely to be identified in Orbis.

Although we could successfully pair about half of the firms in our confidential loan-level data, a substantial portion of the successfully paired firms do not have key financial variables. Unfortunately, in our final matched sample, only 2.2% of the Tamwilcom-guaranteed firms are included.[18] The rate is admittedly low but is consistent with other studies using Orbis data, especially considering Morocco as an African country. Even studies focusing on the EU and the US, which have much better data coverage, also suffer from low pairing rates. For instance, Gereben et al. (2019) report a rate of only 3.6% using data from Central and Eastern Europe. Fortunately, the sample size is still sufficiently large for our analysis. Our main concern about the pairing rate is the attrition of small firms in the final sample. To correct this bias, we follow Amamou, Gereben, and Wolski (2020) and use the technique of inverse probability weight (henceforth IPW) to recover the shares of firms of different sizes in the original treated population as a robustness check.

**Potential Concerns** We have two main concerns about our sample construction. The first is that some unidentified treated firms could be mistaken as untreated control firms and are matched with other treated firms later in the procedure. This would bias the estimation downward. However, this concern is marginally relevant due to the low treatment rate. Suppose the total number of firms in Orbis is taken as a representation of the whole business world in Morocco. In that case, there are approximately 1.58 million firms, of which only 23,017 have been treated. The resulting treatment rate is only 1.5%, indicating a very small possibility of a treated firm being matched with another unpaired treated firm.

The second is survivor bias. It mainly results from the fact that only businesses that actively report their balance sheet to the local trade register's office for the last five years are maintained in Orbis' online version. To reduce this bias, we complement the main online version with Orbis historical vintages, which have records of firms exiting the market.

[18] Only 7505 firms have sales data for the year in which they are granted the guarantee. The number drops even further when a panel of at least three consecutive years is required for the matching process later.

## B Matching Procedure

**Mahalanobis Distance Matching** The matching, implemented under the assumption of "selection on observables," consists of finding statistical twins (control firms) for a guaranteed firm based on a series of time-varying and observable variables relevant to selection into the program. We use the Mahalanobis distance matching (MDM) method to construct a control sample in which a treated firm is matched with the five nearest "neighbors."[19]

**Variable Choices** Following Caliendo and Kopeinig (2008)'s recommendation, we choose the matching variables based on the existing literature and the institutional setting. As a result, total assets, sales, current liabilities, cash and cash equivalents, and firm age are used to measure the statistical distance between observations. Total assets and sales are selected as matching criteria since they are essential to balance sheet items, which reflect the firm's size and overall performance. Current liabilities, namely short-term debt, shed light on the firm's ability to rely on bank credit and the amount of existing indebtedness and risks associated with external credit. Cash and cash equivalents contain short-term investments and funds that can be used to pay current invoices and represent the firm's liquidity situation. The financial variables are log-transformed.

**Matching Choices** The matching is based on the firm's three-year history before receiving the credit guarantee. Firms with insufficient data coverage are inevitably excluded from the matching procedure. We apply exact matching on the sector, year, and firm size classification. We further divide the firms into 20 quantiles based on their sales and impose exact matching on their quantile bin. The purpose is to maximize the similarities between matched firms while maintaining a decent sample size. In robustness tests, we extend the three-year pre-treatment period to four and five. Our results are robust for both tests.

We conduct a new round of matching for each outcome variable in Section 3.4. In each round, we impose two requirements to ensure sample quality. First, we restrict the sample used for matching only to firms with data points for that outcome variable in that year. If a firm's data is missing for this variable in that period, this observation is dropped automatically. This ensures we match firms with data for the specified outcome variable to be tested in the regression. Second, we drop out outliers for that outcome variable before conducting matching. An observation is considered an outlier if that log change is huge (we use ten as the cutoff). A guaranteed firm is matched with a maximum of five closest control firms based on their Mahalanobis distances. Matched observations of treated firms are assigned a weight of one, whereas those of control firms are allocated a weight based on their distance from the corresponding treated firm.

[19]The Mahalanobis distance is a matrix that measures the multivariate proximity between two observations based on selected variables.

**Caliper** A caliper is implemented to ensure the common support assumption. A caliper refers to the maximum distance allowed between a treated firm and its controls. Any control firm beyond this caliper is dropped. This is to ensure that all control firms in the final sample are similar enough to the treated firm that it is matched with. The choice of the caliper is derived from the 0.9-quantile of the distribution of distances between observations in nearest neighbor pairwise matching with replacement, multiplied by 1.5. The choice is based on Jann (2017), Huber, Lechner, and Wunsch (2013), Huber, Lechner, and Steinmayr (2015) after considering the variance-bias trade-off: choosing a large caliper includes more control observations, thus decreasing variance; however, the bias increases if a non-comparable and distant control is included.

**Weighing** The analysis unit is firm-year, based on a similar procedure in Brown and Earle (2017). Matched observations of treated firms are assigned a weight of one, whereas those of control firms are allocated a weight based on their distance from the corresponding treated firm. We first calculate the kernel weight of each matched control observation based on its distance from the treated firm, using the Epanechnikov kernel function with the same bandwidth used in the matching. Subsequently, the weight of each control observation is rescaled as the share of its kernel weight in the sum of kernel weights of all controls matched with the same treated firm. This weight rescaling intends to up-weight those control firms close to treated firms and down-weight those that are far away. For treated firms, only the firm-year observation of the guarantee receipt year is kept. This is to avoid the situation where a treated observation is matched with another observation from a treated firm in a year in which it does not receive a guarantee. For control firms, multiple firm-year observations that belong to the same firm are maintained in the pool of potential controls for matching, provided that the firm's data covers a three-year history of selected financial variables. The matching is performed with a replacement, which implies that one firm-year observation of an untreated firm can be selected more than once.

**Matching Outcomes** All matched samples have similar characteristics. For ease of discussion on matching outcomes, we hereby use the matched sample where the outcome of interest is sales growth between year $t + 1$ and year $t - 1$ as a representative sample. We obtain a final matched sample of 506 guaranteed firms and 1937 control firms, among which 60% have been matched only once, and 26% are used twice. The maximum number of times a control firm has been matched is eight. There are only eight firms in this situation. Since most untreated firms are matched only once, we expect the estimation results to be similar to a matching procedure without replacement. This is confirmed later by a robustness check.

**Balancedness Tests** Figure 1.9 represents the standardized mean difference (SMD) and variance ratios between the treated and control groups in the raw and matched sample.[20] The SMD

[20] See Table 1.12 for the statistics represented in Figure 1.9.

Figure 1.9: Standardized Mean Difference and Variance Ratio in Raw and Matched Sample


Notes: This figure is a visualization of Table 1.12. The standardized mean differences ("StdDif") and variance ratios ("Ratio") of the raw sample and matched sample are reported by the Stata *kmatch* package as in Jann (2017). All financial variables are log-transformed.

measures the mean difference of a given variable between two groups, normalized by the standard deviation of that variable. Variance ratio refers to the ratio between the variances of a variable across two groups. A value of zero for the SMD and a value of one for the variance ratio indicate a good balance in the sample. As shown in the Figure, the matching procedure substantially improves the overall balancedness for most variables, except for cash. Guaranteed firms have a lower cash holding level on average compared to their matched control firms, which also appears in Figure 1.10.

As a second balancedness test suggested by Caliendo and Kopeinig (2008), we evaluate the probability of obtaining a guarantee through a logit model based on the variables used in the matching. Ideally, a drop in $R^2$ indicates a good balance in the sample. We observe that the pseudo $R^2$ of the logit model falls from 0.11 with the raw sample to 0.01 with the matched sample. This confirms the loss of the predictive power of the selected variables after matching. It confirms that the matching procedure has eliminated differences in the pre-treatment observable characteristics between the two groups and that the treatment status is "randomized" in the matched sample conditional on the selected variables.

**Trend Inspection** Figure 1.10 illustrates the weighted average of the log value of the four vari-

Table 1.12: Standardized Mean Difference and Variance Ratio: Raw and Matched Sample

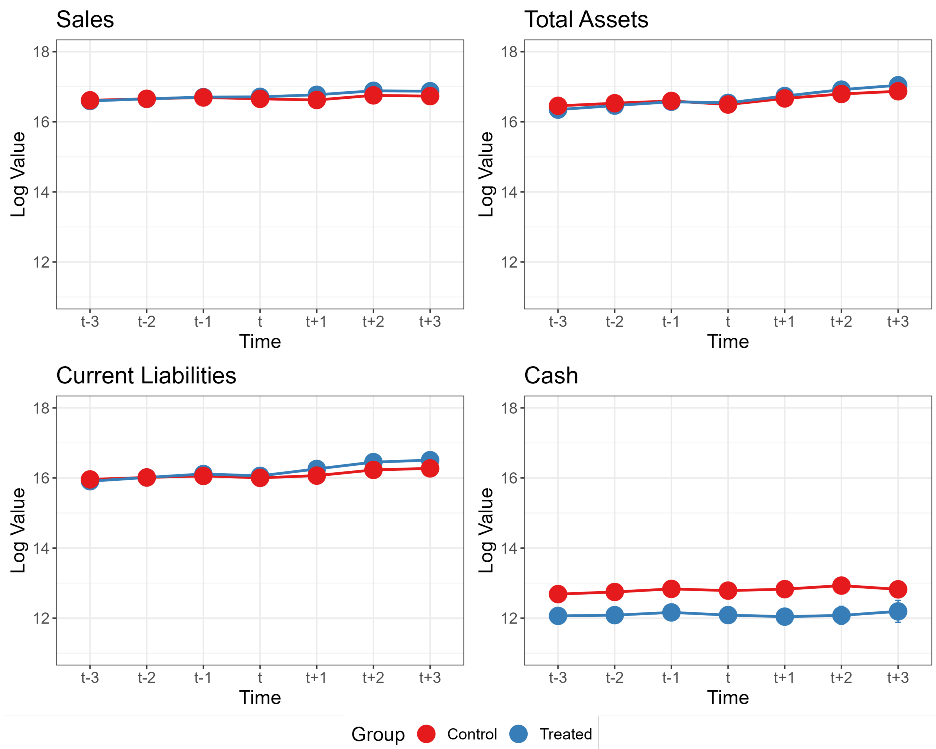
| Sample | Raw | | | Matched | | |
|---|---|---|---|---|---|---|
| Mean | Treated | Untreated | StdDif | Treated | Untreated | StdDif |
| ln(total assets)$_{t-1}$ | 15.60 | 14.01 | 0.88 | 16.61 | 16.62 | -0.01 |
| ln(sales)$_{t-1}$ | 15.62 | 11.23 | 1.06 | 16.71 | 16.70 | 0.003 |
| ln(current liabilities)$_{t-1}$ | 15.09 | 13.42 | 0.71 | 16.16 | 16.08 | 0.03 |
| ln(cash)$_{t-1}$ | 11.35 | 10.76 | 0.25 | 12.21 | 12.90 | -0.30 |
| ln(total assets)$_{t-2}$ | 15.40 | 13.94 | 0.79 | 16.50 | 16.55 | -0.03 |
| ln(sales)$_{t-2}$ | 15.13 | 11.19 | 0.89 | 16.66 | 16.66 | -0.001 |
| ln(current liabilities)$_{t-2}$ | 14.88 | 13.28 | 0.64 | 16.06 | 16.04 | 0.01 |
| ln(cash)$_{t-2}$ | 11.35 | 10.77 | 0.26 | 12.13 | 12.81 | -0.30 |
| ln(total assets)$_{t-3}$ | 15.12 | 13.82 | 0.67 | 16.39 | 16.48 | -0.05 |
| ln(sales)$_{t-3}$ | 14.39 | 10.78 | 0.72 | 16.60 | 16.61 | -0.004 |
| ln(current liabilities)$_{t-3}$ | 14.48 | 12.96 | 0.50 | 15.97 | 15.99 | -0.01 |
| ln(cash)$_{t-3}$ | 11.31 | 10.78 | 0.24 | 12.13 | 12.74 | -0.28 |
| ln(age) | 5.18 | 5.15 | 0.07 | 5.43 | 5.45 | -0.06 |
| Variances | Treated | Untreated | Ratio | Treated | Untreated | Ratio |
| ln(total assets)$_{t-1}$ | 2.17 | 4.35 | 0.50 | 1.65 | 1.64 | 1.01 |
| ln(sales)$_{t-1}$ | 2.26 | 32.33 | 0.07 | 1.35 | 1.34 | 1.01 |
| ln(current liabilities)$_{t-1}$ | 3.32 | 7.66 | 0.43 | 1.96 | 1.91 | 1.03 |
| ln(cash)$_{t-1}$ | 5.14 | 5.56 | 0.92 | 4.57 | 3.72 | 1.23 |
| ln(total assets)$_{t-2}$ | 2.46 | 4.34 | 0.57 | 1.73 | 1.71 | 1.01 |
| ln(sales)$_{t-2}$ | 7.32 | 31.99 | 0.23 | 1.37 | 1.38 | 0.99 |
| ln(current liabilities)$_{t-2}$ | 3.72 | 8.60 | 0.43 | 2.01 | 1.95 | 1.03 |
| ln(cash)$_{t-2}$ | 4.95 | 5.27 | 0.94 | 4.55 | 3.81 | 1.20 |
| ln(total assets)$_{t-3}$ | 3.09 | 4.47 | 0.69 | 1.79 | 1.77 | 1.01 |
| ln(sales)$_{t-3}$ | 15.34 | 34.52 | 0.44 | 1.37 | 1.41 | 0.98 |
| ln(current liabilities)$_{t-3}$ | 6.67 | 11.31 | 0.59 | 2.08 | 1.94 | 1.07 |
| ln(cash)$_{t-3}$ | 4.78 | 5.00 | 0.96 | 4.51 | 3.78 | 1.19 |
| ln(age) | 0.23 | 0.20 | 1.13 | 0.22 | 0.18 | 1.27 |

Notes: This table reports the standardized mean differences (“StdDif") and variance ratios (“Ratio") of the raw sample and the matched sample, reported by Stata *kmatch* package (see Jann (2017)). All variables are log-transformed.

ables used in the matching procedure. It confirms the parallel pre-treatment trend between the treated and control firms and provides preliminary evidence on the dynamic impact of working capital loan guarantees on a firm's growth. As shown in Figure 1.10, guaranteed firms experience growth in sales, total assets, current liabilities, and a decline in cash. This will be confirmed later in the regressions. Overall, standard balancedness tests indicate a good balance in the sample.

As for the level of financial variables before treatment, they are similar, except for cash. Guaranteed firms have a lower level of cash holding on average compared to their matched control firms. This is likely linked to the firm's short-term credit demand. Firms that apply for a guarantee have insufficient cash to cover their liquidity needs. To address this issue, we conduct two robustness tests of a matching procedure focused only on cash-related variables. We show that the results are consistent. This is detailed in E.

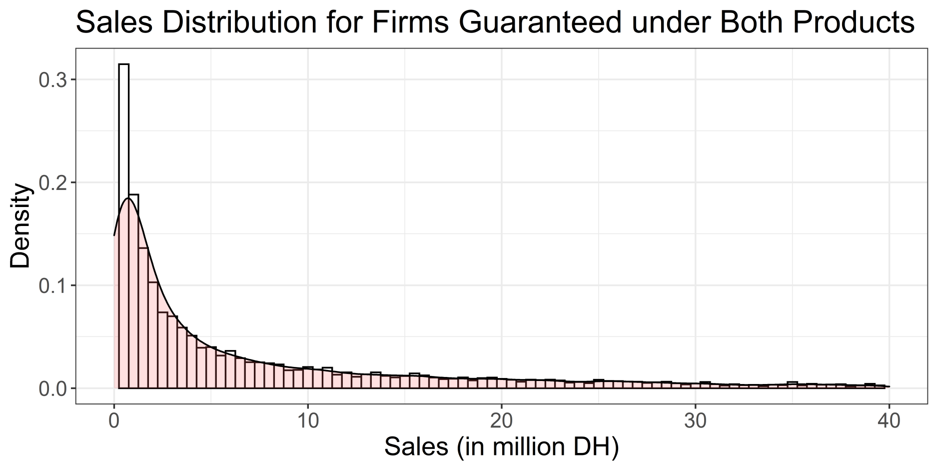
Figure 1.10: Trend Inspection of Four Financial Variables Used in Matching




Notes: This figure depicts the weighted average of the log values of sales, total assets, current liabilities, and cash in years $t-3$ to $t+3$ of treated and control firms in the final matched sample. Confidence intervals are at the 95% level.

## C Supplementary on Summary Statistics

Figure 1.11: Sales Distribution of Guaranteed Firms




Notes: This figure presents the sales distribution (density) of firms guaranteed under Damane Exploitation and Damane Express. The sales number is from the Tamwilcom database.

Table 1.13: Summary Statistics: Guaranteed Firms VS. Non-Guaranteed Firms

| Statistic | N | Mean | St. Dev. | Pctl(25) | Median | Pctl(75) |
|---|---|---|---|---|---|---|
| Guaranteed firms (treated sample) | | | | | | |
| Sales Growth | 3,178 | 0.152 | 1.836 | −0.108 | 0.073 | 0.290 |
| Total Assets Growth | 3,184 | 0.177 | 1.728 | −0.015 | 0.148 | 0.367 |
| Costs of Employees Growth | 3,091 | 0.204 | 0.597 | −0.017 | 0.130 | 0.331 |
| Fixed Assets Growth | 3,184 | 0.099 | 2.308 | −0.323 | −0.053 | 0.340 |
| Current Liabilities Growth | 3,184 | 0.189 | 1.537 | −0.091 | 0.123 | 0.388 |
| Cash Growth | 3,124 | 0.058 | 1.984 | −0.942 | 0.066 | 1.038 |
| Non-guaranteed firms (control sample) | | | | | | |
| Sales Growth | 15,921 | 0.086 | 2.214 | −0.126 | 0.034 | 0.224 |
| Total Assets Growth | 15,932 | 0.095 | 1.804 | −0.048 | 0.087 | 0.265 |
| Costs of Employees Growth | 15,338 | 0.138 | 0.573 | −0.038 | 0.089 | 0.258 |
| Fixed Assets Growth | 15,932 | −0.043 | 2.374 | −0.377 | −0.096 | 0.182 |
| Current Liabilities Growth | 15,951 | 0.094 | 1.558 | −0.137 | 0.045 | 0.290 |
| Cash Growth | 15,674 | 0.104 | 1.608 | −0.624 | 0.076 | 0.860 |

Notes: The summary statistics are based on the matched sample of treated firms and control firms. The growth rate of financial variables is the first difference between logged variables.

Table 1.14: Distribution Statistics of Firm Characteristics

| Assets Quantile | Assets | Sales/Assets | Debt/Assets | Cash/Assets |
|---|---|---|---|---|
| Orbis Sample | | | | |
| 1 | 77 | 1.21 | 0.24 | 0.59 |
| 2 | 338 | 1.24 | 0.53 | 0.25 |
| 3 | 1, 102 | 1.02 | 0.61 | 0.15 |
| 4 | 3, 509 | 0.89 | 0.65 | 0.10 |
| 5 | 41, 965 | 0.68 | 0.65 | 0.06 |
| Guaranteed Sample | | | | |
| 1 | 770 | 1.69 | 0.56 | 0.18 |
| 2 | 2, 462 | 1.31 | 0.62 | 0.08 |
| 3 | 5, 422 | 1.21 | 0.65 | 0.06 |
| 4 | 12, 866 | 1.12 | 0.66 | 0.05 |
| 5 | 54, 923 | 0.91 | 0.68 | 0.03 |
| Non-Guaranteed Sample (Control) | | | | |
| 1 | 2, 897 | 2.02 | 0.60 | 0.12 |
| 2 | 8, 704 | 1.54 | 0.64 | 0.08 |
| 3 | 16, 971 | 1.32 | 0.64 | 0.06 |
| 4 | 33, 254 | 1.08 | 0.63 | 0.06 |
| 5 | 111, 372 | 0.72 | 0.60 | 0.04 |
| Non-Guaranteed Sample (Whole) | | | | |
| 1 | 73 | 1.20 | 0.23 | 0.60 |
| 2 | 309 | 1.25 | 0.52 | 0.26 |
| 3 | 996 | 1.01 | 0.61 | 0.16 |
| 4 | 3, 166 | 0.87 | 0.65 | 0.11 |
| 5 | 40, 929 | 0.65 | 0.65 | 0.06 |

Notes: This table reports the means of indicated financial variables and ratios based on five quantile groups of total assets. The unit of total assets is a thousand. Observations with ratios of current liabilities/total assets and cash/total assets greater than one or less than 0 are dropped. The Orbis sample comprises all firms in Morocco. The guaranteed sample refers to the whole sample of guaranteed firms. The non-guaranteed sample (whole) comprises all firms that do not possess a credit guarantee. The non-guaranteed sample (control) refers to those non-guaranteed firms selected during the empirical analysis's matching process.

## D Supplementary on Participation Rate

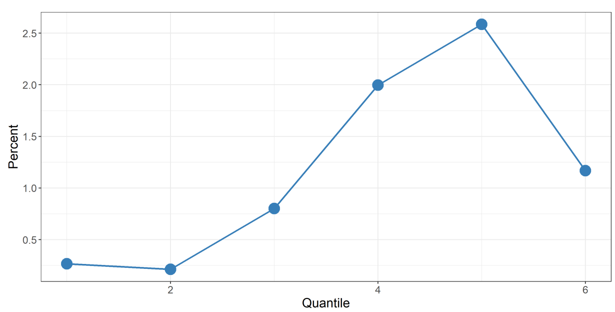
Table 1.15: PARTICIPATION RATE BY SIZE

| Size | Participation Rate (%) |
|---|---|
| Small company | 2.51 |
| Medium-sized company | 15.40 |
| Large company | 11.44 |
| Very large company | 1.22 |

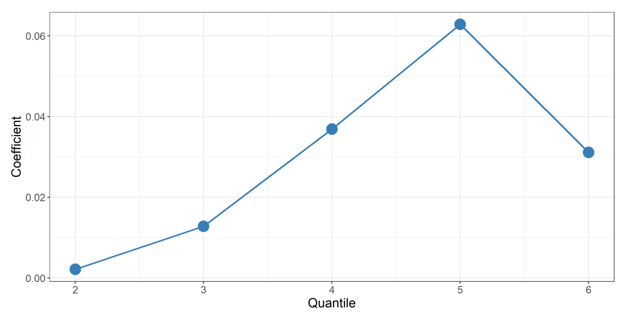
Table 1.16: PROBABILITY OF SUCCESSFUL PAIRING BETWEEN TAMWILCOM AND ORBIS BY SIZE

| Size | Probability (%) |
|---|---|
| Small company | 30.37 |
| Medium-sized company | 52.54 |
| Large company | 74.41 |
| Very large company | 79.03 |

Figure 1.12: DISTRIBUTION OF PARTICIPATION RATE




(a) Distribution by Total Assets Quantile Bin




(b) Coefficient of Quantile Regression

Notes: The cut points of total assets quantile bins are set at 0.1, 0.5, 0.75, 0.9, 0.99. In Figure 1.12 (a), the participation rate is calculated as the ratio of the number of guaranteed firms to the total number of firms in each bin. In Figure 1.12 (b), the coefficient is from the following regression: $\text{Participation}_{it} = \sum_{q=1}^{6} \beta_q$ Quantile $\text{Bin}_q + \delta_j + \delta_t$, where $i$ indexes individual firms, $j$ indexes sector, and $t$ indexes year. $\text{Participation}_{it}$ is a dummy variable of one if the firm is guaranteed and zero otherwise. Quantile $\text{Bin}_q$ is a dummy variable of one of the firm's total assets located in quantile $q$ ($q \in [1, 6]$), and zero otherwise. $\delta_j$ and $\delta_t$ refer to sector and year fixed effects. Observations in quantile one are dropped automatically by Stata due to collinearity. Coefficients of $\beta_q$ are reported in Figure 1.12 (b).

# E Robustness Checks

### Number of Pre-treatment Years in Matching

The first robustness test corresponds to concerns regarding the number of pre-treatment years used for matching. Existing literature suggests that we should rely on at least three years' pre-treatment performance for matching, which is our main estimation. In this robustness check, we extend the number of years to four and five. Table 1.17 reports the estimated results when we match four years' data. As a result of the stricter matching requirement, the number of treated firms that have at least one matched control firm drops to 345. Most results in year $t+1$ remain robust and significant on a similar level, consistent with our baseline results. When we increase the number of years used for matching to five, we only have 213 guaranteed firms that enter the final sample. The estimated results for the year $t+1$ in Table 1.18 are mostly significant except for the coefficient on cash. This is also in line with the baseline.

Table 1.17: Results of Year $t+1$ from Matching on Four Pre-Treatment Years' Data

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.130*** | 0.094** | 0.129*** | 0.090 | 0.090*** | 0.146* |
| | (0.029) | (0.029) | (0.031) | (0.106) | (0.025) | (0.069) |
| $N$ | 13432 | 13723 | 13952 | 13531 | 12636 | 13460 |
| adj. $R^2$ | 0.216 | 0.236 | 0.213 | 0.338 | 0.228 | 0.213 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment ("Guaranteed") from DID regression in the robustness test, where we match on four pre-treatment years' data. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t-1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

Table 1.18: Results of Year $t + 1$ from Matching on Five Pre-Treatment Years' Data

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.169*** | 0.137*** | 0.130*** | 0.015 | 0.081** | 0.273*** |
| | (0.039) | (0.032) | (0.036) | (0.133) | (0.029) | (0.075) |
| $N$ | 8664 | 8805 | 8902 | 8752 | 8343 | 8641 |
| adj. $R^2$ | 0.265 | 0.236 | 0.133 | 0.361 | 0.223 | 0.229 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we match five pre-treatment years' data. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t + 1$ from year $t - 1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

### Data Attrition Bias

The second robustness test is to correct the bias from the data attrition issue. The main concern arises from the loss of observations of small firms during matching. Considering that small firms often report minimal financial data, it could lead to their exclusion in the matching process due to missing data points. In order to correct this bias, we use inverse probability weighting (ipw) (Amamou, Gereben, and Wolski, 2020) to increase the weight of underrepresented SMEs and decrease the weight of those who are over-represented. As a first step, we calculate the number of small, medium, and large firms in the sample of Tamwilcom-guaranteed firms that can be merged with Orbis. As discussed earlier, this sample shares similar statistical properties with the sample of all Tamwilcom-guaranteed firms. The reason for choosing this merged sample rather than the full sample is that we can use the size information provided by Orbis. We assume that information on firm size composition in this merged sample can reflect that of the full sample. As a second step, we count the number of firms of different sizes in the processed sample after matching and divide the number of small, medium, and large firms in the processed sample by the number in the original sample before matching. The inverse of the proportion is then used as a weight to re-scale the representation of different-sized firms in the final sample. As Table 1.19 shows, estimation results are similar to the main ones, with the exception of total and fixed assets.

Table 1.19: Results of Year $t + 1$ with Inverse Probability Weight

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.129** | 0.039 | 0.109** | -0.224 | 0.131** | 0.003 |
| | (0.046) | (0.037) | (0.041) | (0.176) | (0.040) | (0.103) |
| $N$ | 17199 | 17344 | 17520 | 17017 | 16571 | 17117 |
| adj. $R^2$ | 0.201 | 0.222 | 0.193 | 0.323 | 0.319 | 0.201 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we use the technique of inverse probability weight to correct data attrition bias. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t-1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

### Emphasizing Cash in Matching

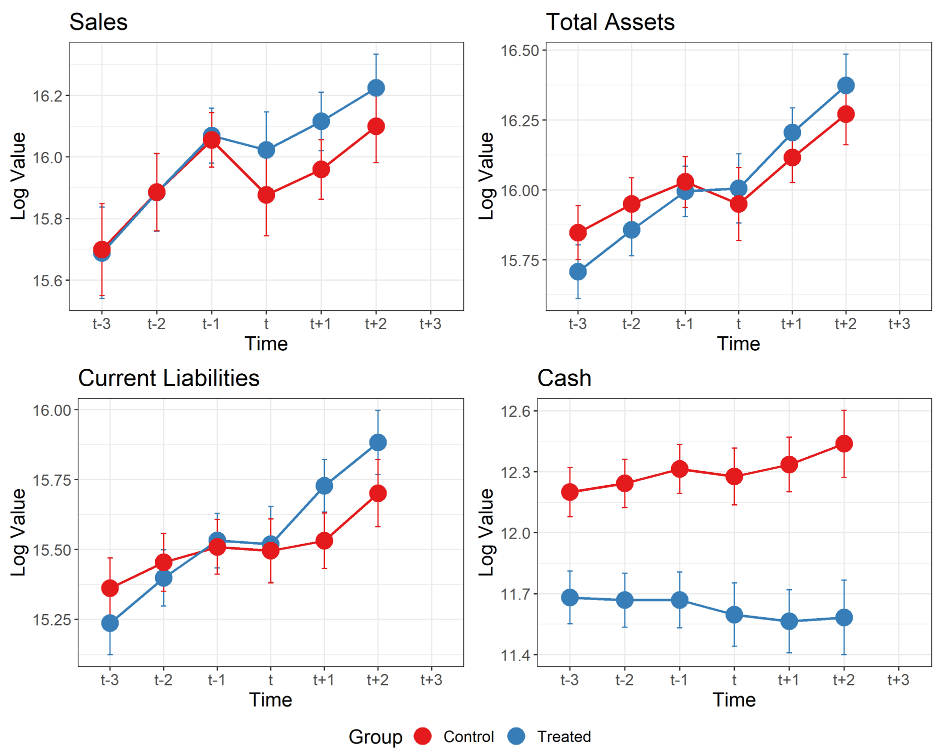
The third set of checks intends to test the robustness of the main results when we emphasize matching on cash-related variables to reduce the difference in the cash level of treated and control firms after matching in Figure 1.10. As a first test, we use one-to-one nearest-neighbor matching to ensure that only the closest control firm is selected. This is to see if the difference in the gap comes from any chosen control firm that is not similar enough to its matched neighbor. As shown in Figure 1.13, the gap remains large and is very similar to the five-to-one nearest neighbor matching. In view of this, we rule out the possibility that remote control firms contribute to the difference in cash. As a further test, we only include logged cash and the ratio of cash to total assets in the matching process. This setup “forces" a good matching result on cash by not including other variables so that the measurement of Mahalanobis distance is only based on cash-related variables. In addition, we divide the variable of logged cash into 20 quantile intervals and apply exact matching on the interval. Figure 1.14 shows that this procedure manages to substantially improve the matching performance on cash. Furthermore, total assets are balanced as well due to the incorporation of the ratio of cash to total assets. However, we observe a gap in sales. To reduce this gap, we modified the setup to match on cash ratio and logged sales. As Figure 1.15 indicates, the good balancedness in cash, total assets, and current liabilities are preserved while the difference in sales is decreased. Estimation results for both matchings are reported in Table 1.20 and Table 1.21. They are consistent with the main results.

Table 1.20: Results of Year $t + 1$ from Matching on Logged Cash and Cash Ratio

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.217** | 0.162** | 0.203*** | -0.067 | 0.088** | 0.256* |
| | (0.080) | (0.052) | (0.050) | (0.114) | (0.032) | (0.114) |
| $N$ | 6109 | 6435 | 6604 | 6144 | 4963 | 6209 |
| adj. $R^2$ | 0.233 | 0.231 | 0.101 | 0.325 | 0.359 | 0.215 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

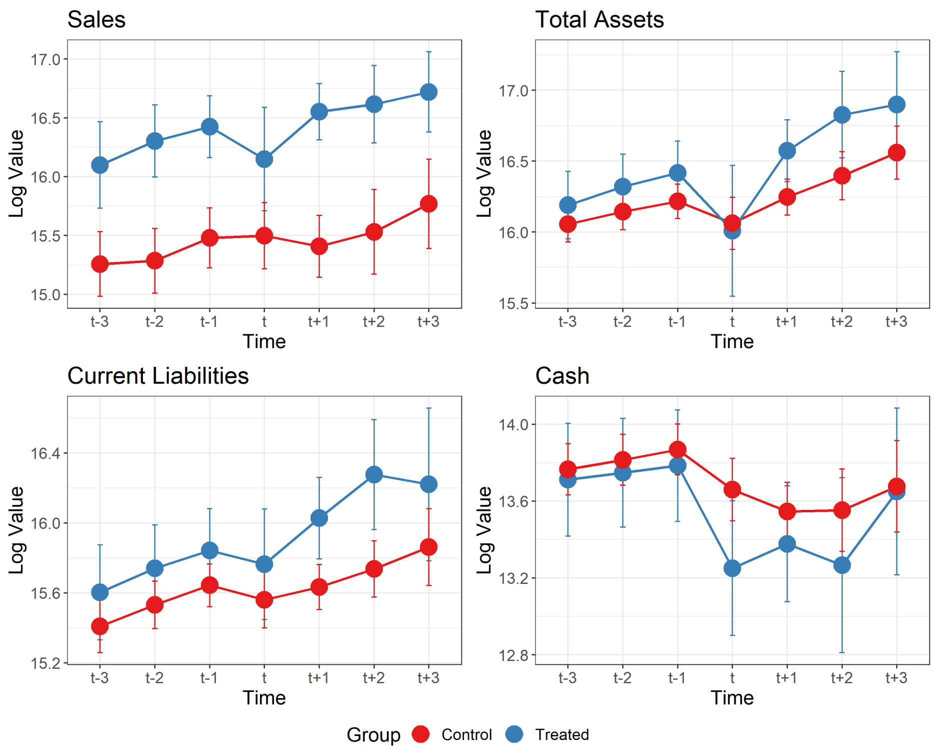
Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we only include logged cash and the ratio of cash to total assets from three pre-treatment years for matching. In addition, we divide the log cash variable into 20 quantile intervals and apply exact matching on the interval. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t + 1$ from year $t - 1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

Figure 1.13: Robustness: Trend Inspection from Matching with One Nearest Neighbor




Notes: This figure depicts the weighted average of the log values of sales, total assets, current liabilities, and cash in year $t-3$ to $t+2$ of treated and control firms from the robustness test, where we match only one nearest control firm for a treated firm. The confidence interval is at 95% level.

Figure 1.14: Robustness: Trend Inspection from Matching on Log Cash and Cash Ratio




Notes: This figure depicts the weighted average of the log values of sales, total assets, current liabilities, and cash in year $t-3$ to $t+2$ of treated and control firms from the robustness test, where we only include logged cash and the ratio of cash to total assets from three pre-treatment years for matching. In this robustness test, we also divide the log cash variable into 20 quantile intervals and apply exact matching on this interval. The confidence interval is at 95% level.

Figure 1.15: Robustness: Trend Inspection from Matching on Log Sales and Cash Ratio


Notes: This figure depicts the log values of sales, total assets, current liabilities, and cash in year $t-3$ to $t+2$ of both treated and control firms from the robustness test, where we only include logged sales and the ratio of cash to total assets from three pre-treatment years for matching. In this robustness test, we also divide the log cash variable into 20 quantile intervals and apply exact matching on this interval. The confidence interval is at 95% level.

Table 1.21: Results of Year $t + 1$ from Matching on Logged Sales and Cash Ratio

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.202** | 0.136** | 0.205*** | -0.091 | 0.096** | 0.243* |
| | (0.071) | (0.052) | (0.050) | (0.115) | (0.031) | (0.104) |
| $N$ | 6478 | 6750 | 6873 | 6496 | 5335 | 6595 |
| adj. $R^2$ | 0.324 | 0.215 | 0.137 | 0.297 | 0.364 | 0.196 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment ("Guaranteed") from DID regression in the robustness test, where we only include logged sales and the ratio of cash to total assets from three pre-treatment years for matching. In addition, we divide the log cash variable into 20 quantile intervals and apply exact matching on the interval. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t + 1$ from year $t - 1$. "Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

### Propensity Score in Mahalanobis Distance

In the fourth set of checks, we include propensity score as one variable in calculating Mahalanobis distance. We exploit the predictive power of a logit model, where the dependent variable is a dummy of one if a firm is guaranteed in a certain year, and independent variables are the same as those selected for calculating Mahalanobis distance. Table 1.22 reports the estimation results, which are similar to our main results. We conduct another robustness test where we increase the number of nearest neighbors matched with guaranteed firms to ten. We find that the results are not sensitive to the number of controls chosen for the treated firm, as shown in Table 1.23. We also apply the matching procedure without replacement and confirm that estimation results remain similar, as shown in Table 1.24.

Table 1.22: Results of Year $t+1$ with Propensity Score in Multivariate Matching

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.139*** | 0.093*** | 0.143*** | -0.054 | 0.107*** | 0.113+ |
| | (0.025) | (0.024) | (0.027) | (0.091) | (0.021) | (0.061) |
| $N$ | 18268 | 18464 | 18841 | 18141 | 17418 | 17976 |
| adj. $R^2$ | 0.190 | 0.204 | 0.213 | 0.313 | 0.241 | 0.199 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we include propensity score as one variable in calculating Mahalanobis distance. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variable in the logit model is a dummy of one if a firm is guaranteed in a certain year, and the independent variables are the same ones selected for calculating Mahalanobis distance in the main setup. Outcome variables are the log difference of six main variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t-1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

Table 1.23: Results of Year $t + 1$ from Matching on 10 Nearest Neighbors

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.136*** | 0.084*** | 0.131*** | 0.135 | 0.098*** | 0.163* |
| | (0.023) | (0.022) | (0.025) | (0.101) | (0.023) | (0.064) |
| $N$ | 23583 | 24054 | 24569 | 23644 | 22796 | 23410 |
| adj. $R^2$ | 0.253 | 0.253 | 0.249 | 0.348 | 0.278 | 0.246 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we match up to 10 nearest control firms for a treated firm. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t - 1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

Table 1.24: Results of Year $t + 1$ from Matching without Replacement

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | 0.147*** | 0.105*** | 0.146*** | -0.198* | 0.086** | 0.109+ |
| | (0.030) | (0.024) | (0.028) | (0.084) | (0.028) | (0.063) |
| $N$ | 16165 | 16631 | 16682 | 16131 | 14681 | 16226 |
| adj. $R^2$ | 0.297 | 0.207 | 0.261 | 0.328 | 0.267 | 0.234 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we apply the matching procedure without replacement. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t - 1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

### Placebo Test Assuming Treatment Occurred Three Years Earlier

We conduct a falsification test assuming the treatment occurred three years earlier than it actually took place. Table 1.25 reports the estimation results. As it shows, most coefficients are not significant, which confirms the robustness of the main results.

Table 1.25: RESULTS OF YEAR $t + 1$ FROM PLACEBO TEST

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Sales | Total Assets | Current Liabilities | Cash | Costs of Employees | Fixed Assets |
| Guaranteed | -0.013 | 0.098 | -0.006 | -0.022 | 0.104** | 0.091 |
| | (0.117) | (0.089) | (0.057) | (0.100) | (0.039) | (0.126) |
| $N$ | 17833 | 17859 | 17874 | 17575 | 17219 | 17859 |
| adj. $R^2$ | 0.069 | 0.037 | 0.032 | 0.102 | 0.083 | 0.029 |
| Group × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| City × Year FE | Yes | Yes | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients of treatment (“Guaranteed") from DID regression in the robustness test, where we assume the treatment occurred three years earlier. Each outcome variable in each year is based on a different matched sample where we drop firms without data for that outcome variable in that year before matching and excluding outliers. The dependent variables are the log difference of six main outcome variables (sales, total assets, labor costs, fixed assets, cash, and current liabilities) in year $t+1$ from year $t-1$. “Guaranteed" indicates that a firm receives a Tamwilcom guarantee in year $t$. Group-year and city-year fixed effects are included. Standard errors are clustered at the group-year level. Significance level: $^{+}$ $p < 0.10$, $^{*}$ $p < 0.05$, $^{**}$ $p < 0.01$, $^{***}$ $p < 0.001$.

# Theoretical Appendix

## F Proofs in the Analytical Model

### The entrepreneur's program in the special case

The first-order conditions of the Lagrangian problem associated with objective (1.12) are the following:

$$/d_t \ : d_t^{-\eta} - \eta_t = 0 \tag{1.26}$$

$$/n_t \ : \beta\epsilon v'(n_t) - \eta_t = 0 \tag{1.27}$$

$$/k_t \ : -\gamma_t + [\psi'(k_t) + 1 - \delta] + \lambda_t(\theta - a) = 0 \tag{1.28}$$

$$/c_t \ : -\gamma_t + (1 + r_t) + \lambda_t + \zeta_t = 0 \tag{1.29}$$

where $\eta_t$ is the shadow price of the budget constraint (1.13), and $\gamma_t$, $\lambda_t$ and $\zeta_t$ the shadow prices of, respectively, the net worth allocation constraint (1.14), the working capital constraint (1.15) and the non-negative cash constraint (1.16), normalized by $[\eta_t(1-\tau)]^{-1}$. The envelope theorem yields

$$v'(n_{t-1}) - \gamma_t\eta_t(1-\tau) = 0 \tag{1.30}$$

We use the first-order conditions to derive the equations (1.17) and (1.19) in the paper. FOC equation (1.28) and equation (1.29) yield the relationship between MBK and MBC equation (1.17) immediately. Considering a sufficiently small SME which needs positive cash holdings ($\zeta = 0$), FOC equation (1.27), combined with equation (1.30) evaluated in $t+1$, together with equation (1.29) evaluated in $t+1$, and finally replace both $\eta_t$ and $\eta_{t+1}$ using equation (1.26), yield equation (1.19).

### Proof of Proposition 2

Point (i) derives immediately from the long-term Euler equation (1.21) and from the stationarity of household consumption that implies $\beta(1 + r_{t+1}) = 1$.

To establish the point (ii), we use the expression for $\lambda_t^*$, where we replace $\lambda_t^*$ with its long-term value $\lambda^{LT}$. We obtain an implicit definition of the long-run capital stock $k^{LT}$:

$$\lambda^{LT} = \frac{\psi'(k^{LT}) + 1 - \delta - (1 + r_t)}{1 + a - \theta}$$

we then replace $1 + r_t$ with $1/\beta$. Then, we define the optimal capital stock as $k^{opt}$. Noting that, $k^{opt}$ is determined by $\psi'(k^{opt}) + 1 - \delta = (1 + r_t)$, we replace $1 - \delta - (1 + r_t)$ with $-\psi'(k^{opt})$. This yields point (ii).

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# Chapter 2

# Macroprudential Policy and Spillovers: Evidence from Chinese Corporate Credit in Tax Havens

Wenxia Tang [1]

**Abstract**

This article examines the spillover effects of a macroprudential policy in the Chinese bond market. My analysis reveals that the tightening of domestic credit regulations in 2018 generated unintended spillovers in tax haven countries through the international borrowing activities of large Chinese firms. These spillover effects are predominantly linked to non-state-owned enterprises (non-SOEs) in the real estate sector, which have been crowded out from the domestic credit market. The findings indicate that a 1% increase in private ownership corresponds to a 1% increase in bonds issued in tax havens for non-SOEs in the real estate sector.

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## 2.1 Introduction

Recent evidence points to the growing importance of tax havens as conduits for Chinese firms to access the international capital market through foreign subsidiaries. Coppola et al. (2021) find that the amount of investments flowing from the US to China in corporate bonds, after adjusted reallocation from tax havens to its true destination, increases from $3 to $48 billion in 2017. While often overlooked in mainstream research, this amount of substantial capital involved in offshore financing activities by emerging market firms raises crucial questions for policymaking. A fundamental question is to understand the motivations that drive these firms to engage in offshore bond offerings in tax havens. Existing qualitative literature (Buckley et al., 2015) points out that the use of tax haven vehicles is not mainly for the purpose of tax avoidance for Chinese firms. In fact, a tax law back in 2008 discouraged enterprises from incorporating offshore and round-tripping investments by specifying that firms whose "de facto management body" is located in China are subject to Chinese taxation laws. My research offers the first empirical evidence on this issue by leveraging a macroprudential policy change in China and studying its link with the offshore financing behaviors of domestically credit-constrained firms.

I analyze the impact of a macroprudential policy on regulating wealth management products (WMPs) in China in 2018. This policy was intended to tighten the WMP market, reduce associated financial risks, and stabilize the credit market. WMPs were the main trading channel of corporate bonds and represented 64% of their investor base in 2016 (Miao, 2019). By targeting WMPs, the policy adversely affected demand for corporate bonds. The reduction in corporate bond financing through domestic demand channel provides a unique opportunity to estimate if offshore corporate bond issuances in tax havens, a funding channel beyond regulation, are associated with domestic credit tightening. This points to the possibility that the main motivation for Chinese firms to engage in activities in tax havens comes from overcoming financing constraints in the domestic credit market. Incorporation in tax havens grants access to the capital market of advanced economies (Buckley et al., 2015). This corresponds to the growing body of literature documenting that macroprudential regulation in a domestic credit market can generate unintended spillovers through international borrowing of large firms (Forbes, 2020)[2]. This paper intends to investigate if China's macroprudential regulation on domestic credit growth provokes spillovers of corporate debt in tax haven countries.

Although this policy does not single out non-state-owned enterprises (non-SOEs) in particular, the market witnesses a substantial decline in demand for their bonds. In contrast, bonds

[2] Forbes (2020) differentiates spillovers from leakages of a macroprudential policy. He defines spillovers as cross-border exposures, compared to leakages which refer to credit shifting to other domestic entities not subject to regulations. This paper focuses on cross-border spillovers rather than leakages

of state-owned enterprises (SOEs) remain relatively unaffected due to the implicit government guarantee. This sudden market exclusion of non-SOEs serves as the key treatment measure in my subsequent difference-in-differences (DID) analysis. Moreover, an additional layer of complexity introduced to the consequences of the policy is its impact on firms in the real estate sector, in particular the non-state-owned ones. These firms, characterized by high levels of indebtedness, have been subject to strict regulations in China since the early 2010s. This macroprudential policy further crowds them out in the domestic credit market and pushes them to seek alternative sources of financing that are beyond the supervision of Chinese regulatory agencies. This aspect of the policy is explored in a difference-in-difference-in-differences (DDD) setup.

I compile data from five commercially accessible databases to piece together the puzzle of Chinese corporate behaviors in tax havens. Bond-level data is sourced from Capital IQ and Refinitive Eikon, which is screened following Coppola et al. (2021) to identify bonds issued by Chinese firms incorporated in tax havens. Additionally, firm-level data is gathered from China Stock Market and Accounting Research (CSMAR), RESSET, and Worldscope. I combine bond-level data with firm-level data to obtain a sample from 2010 to 2020. Distinct from Coppola et al. (2021), who estimate the total amount of capital flowing through tax havens to other destinations, my focus is on investigating firm-level motivation for offshore bond offerings. While my sample does not comprehensively cover all Chinese corporate activities in tax havens, it effectively serves the purpose of my analysis.

I begin my analysis with DID estimates of the effectiveness of the macroprudential reform. The findings indicate that the regulations targeting WMPs successfully reduce the borrowing activities of non-SOEs from the domestic credit market. As a result, credit-constrained non-SOEs turn to banks for additional loans to cover the financing gap on their balance sheets. Nevertheless, the extent of bank financing is not able to adequately address the funding gap. This indicates a compelling incentive for non-SOEs to explore alternative channels of financing. I don't identify a more pronounced impact specifically linked to the borrowing behaviors of non-SOEs in the real estate sector compared to other industries. However, it is revealed that non-SOEs in the real estate sector manage to sustain the accumulation of corporate net income, whereas non-SOEs in other sectors experience diminished retained earnings as they use internal funds to offset the external funding gap. This suggests that these firms in the real estate sector have alternative channels of funding to support corporate debt rollover and investments in new projects.

I continue to examine the spillovers of corporate borrowing in tax havens through bond issuances following the implementation of the macroprudential policy in 2018. My results confirm that non-SOEs when compared to their SOE counterparts, are more likely to offer bonds through shell companies in tax havens after the new regulation. The effects are predominantly driven by private firms within the real estate sector. My analysis indicates that a 1% increase in private ownership corresponds to a 1% increase in bonds issued in tax havens for non-SOEs in the real estate sector after the regulatory changes. This substantiates the existence of spillover effects of Chinese corporate activities in tax havens following the credit tightening measures implemented in 2018.

The primary threat to my identification strategy is the potential endogeneity of the regulation's timing with changes in firm-level outcomes. To address this concern, I estimate a specification with leads and lags to verify that outcome variables for SOEs and non-SOEs do not trend differently prior to the implementation of the new regulation. In addition, existing literature has confirmed that this policy event stands as the key factor contributing to the surge in SOE premiums and the crowd-out of non-SOEs (Geng and Pan, 2022). This mitigates the concerns that the observed spillover effects associated with non-SOEs could be attributed to other factors during the same period.

The article is organized as follows. Section 2 presents a detailed institutional background on the new macroprudential regulation. Section 3 discusses related literature. Section 4 describes the data. Section 5 provides the empirical analysis. Section 6 concludes.

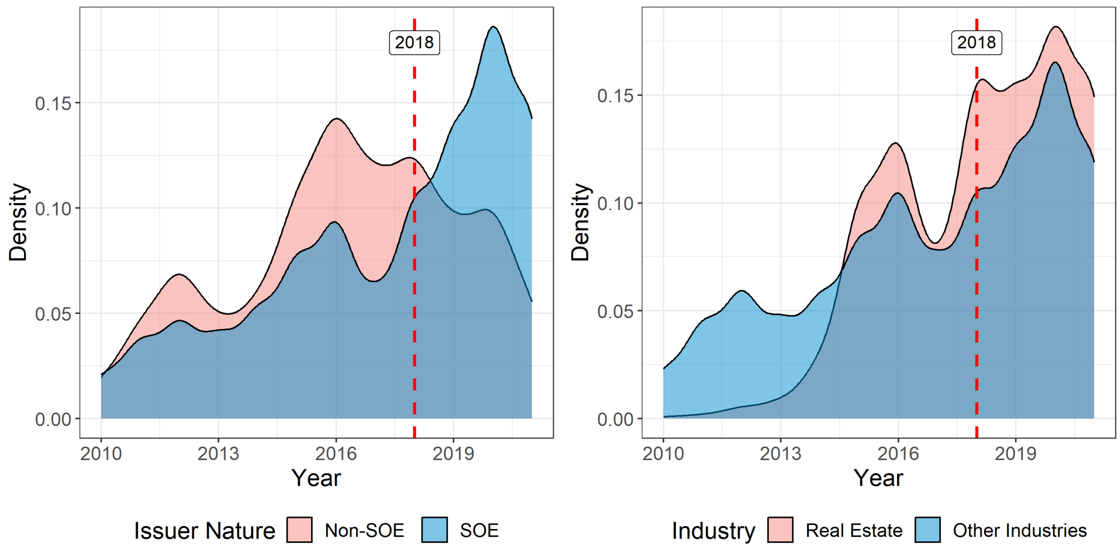
## 2.2 Policy Background

Chinese corporate bond market has expanded rapidly during the past decade. Domestic debt securities issued by non-financial companies have grown from 0.5 trillion USD in 2010 to 5.1 trillion in the first quarter of 2022, making it the second largest corporate bond market next to the US (BIS Debt Securities Statistics). It is worth noting that banks still dominate China's financing system. In fact, 70% of firm funding is sourced through bank loans. It is significantly larger than bond financing, which takes up only 10% (Zhang and Wu, 2019). However, the rapid expansion suggests future growth potential and highlights the importance of evaluating the implications of the expanding corporate credit market.

Excessive credit expansion is often correlated with financial instability (Jordà, Schularick, and Taylor, 2011). This particularly applies to the Chinese corporate bond market, which is largely fueled by shadow banking activity in recent years. Ehlers, Kong, and Zhu (2018) highlight the close financial linkage between the bond market and the shadow banking sector. They estimate that approximately 25% of corporate bonds were purchased by the proceeds of WMPs in 2016, which constituted the largest shadow banking component.

WMPs are issued by commercial banks as alternative saving instruments with higher invest-

Figure 2.1: Chinese Onshore Bond Issuance Trend




Notes: This figure depicts the density of bond issuances in the domestic credit market since 2010. The left figure shows the evolution of bond issuances by SOEs and non-SOEs. The right figure demonstrates the trends by firms in real estate sector versus other sectors.

ment returns than traditional bank deposits. It is used to attract funding from retail investors and wealthy individuals. They are considered a safe investment by households since they are sold at the bank counter (Ouyang and Wang, 2022). However, WMPs are not explicitly guaranteed by banks, neither are they recorded on banks' balance sheets. Unlike deposits, WMP is not subject to banking regulations on interest rate ceilings and capital reserve requirements (Ehlers, Kong, and Zhu, 2018). WMPs were loosely regulated before 2018. They could take up various product forms issued by different entities regulated by different authorities under different sets of rules before 2018. This left room for regulatory arbitrage and risk-taking behaviors (Miao, 2019). The most notable risk is maturity mismatch. Most WMPs have maturities of three to six months, whereas the underlying investments in corporate bonds have maturity terms of two to four years (Ehlers, Kong, and Zhu, 2018).

To stabilize the financial market, the authorities introduced new regulations to reign in the growth of WMPs in 2018, notably "The Guiding Opinions on Regulating the Asset Management Business of Financial Institutions". The draft of this regulation started to circulate in the market in November 2017 and the official regulations were announced in April 2018. This regulation intends to tighten the WMP market, reduce liquidity mismatches, and raise requirements for WMP issuers[3]. The tightened condition has severely shrunk the investor base and financing

[3]See Miao (2019) for detailed discussions on the regulation.

channels of corporate bonds, especially for non-SOEs. SOEs are barely affected owing to the implicit guarantee by the government. However, demand for non-SOEs bonds is significantly reduced, which resembles a market run. Tightened credit conditions have triggered investors' concerns over default and rollover risks in the corporate bond market for non-SOEs (Geng and Pan, 2022). Figure 2.1 shows the market reactions to bonds issued by SOEs and non-SOEs after the announcement of the 2018 regulation. As the left figure demonstrates, the number of newly issued bonds by non-SOEs started to decline due to reduced market demand in 2018 whereas those of SOEs continued to rise with little influence from the regulation change. This is consistent with existing literature documenting this market event (Geng and Pan, 2022). The right figure illustrates the changing trends between firms in the real estate sector and those in other sectors. As it shows, the policy announcement puts the sharp rise of the real estate sector to a transient decline and subsequently a much slower growing pace. In contrast, the number of bonds issued in other sectors demonstrates a continuous upward trajectory throughout the observed period. In Figure 2.6 in Appendix B, the four categories of firms (non-SOEs non-real-estate, non-SOEs real estate, SOEs non-real-estate, and SOEs real estate) are compared, showing trends consistent with those in 2.1. Examining their alternative funding sources of bank loans, as shown Figure 2.7 shows, there are no significant changes in the amount of borrowing from banks post-reform.

The onshore corporate bond market is closely linked to firms' offshore behaviors. This macroprudential policy in 2018 carries important implications for firms' offshore bond financing, in particular for funding channels that are beyond regulation. This applies especially to firms that are crowded out from the domestic credit market, namely non-SOEs. They start to look for other financing sources to raise funding and roll over their current corporate debt. This will be explored in the empirical analysis.

## 2.3 Related Literature

This paper contributes to the small group of literature on the rising economic importance of tax havens in intermediating international capital flows. The recent seminal work by Coppola et al. (2021) redraws the map of capital flowing through tax havens based on the true economic destination of investments. Their research reveals the underestimated scale of bilateral investment from advanced economies to emerging market countries. It corresponds to the evolving role of offshore financial centers documented by a series of studies (Lane and Milesi-Ferretti, 2007, 2011a,b, 2018). Their role as banking centers subsides after the global crisis, whereas we observe a surge in their FDI positions. This is closely linked to the increasingly complicated corporate structures of multinationals and their intra-firm cross-border balance sheet operations.

This paper relates to the strand of literature focused on the international aspect of domestic macroprudential policies, which have become common policy instruments to limit systemic financial risks after the global financial crisis. There is subsequently a growing body of research on the effectiveness of different macroprudential tools and their unintended leakages. The strand of literature that is most relevant to this paper is those focused on the cross-border spillovers of macroprudential policy. Forbes (2020) provides a comprehensive review of various international aspects of macroprudential policies. Ahnert et al. (2021) focuses on domestic macroprudential foreign exchange (FX) regulation and its cross-border spillovers from the corporate sector. They find that FX regulations significantly reduce bank FX borrowing; however, firms increase their FX corporate bond issuance. Economists at the Bank for International Settlements (BIS) call for international policy coordination in view of the spillovers and spillbacks of macroprudential policy (Agénor and da Silva, 2018, 2019).

This paper is also closely related to the large body of literature on the surge of offshore bond issuance by emerging market economies (EMEs) in the international market after the global financial crisis. Shin (2013) describes it as the "second phase of global liquidity", where the main stage is the emerging market bond market open to international investors. It features a substantial retrenchment from cross-border banking to international bond issuance (Lane and Milesi-Ferretti, 2018). A strand of literature focuses on the financial motive of non-financial EME firms that borrow internationally under a low US interest rate environment. Bruno and Shin (2017) note that non-US firms issue dollar-denominated bonds mostly to exploit favorable dollar carry trade based on evidence from 47 countries. Caballero, Panizza, and Powell (2016) emphasize the interaction between carry trade activities and capital account restrictions in 18 emerging market economies. Their evidence suggests that non-financial firms issue bonds through offshore affiliates and bring the proceeds of issuance into the home country via an inter-company loan to escape capital controls. Rodrigues-Bastos, Kamil, and Sutton (2015) document this bond issuance uptrend in five large Latin American economies and term it "Bon(d)anza". They corroborate Caballero, Panizza, and Powell (2016)'s finding on regulatory arbitrage through bond issuance by offshore vehicles. Huang, Panizza, and Portes (2018) focus on China and find that risky firms are more likely to do inter-firm lending in the face of prudential regulations on capital inflows.

This paper is broadly related to the large body of literature on Chinese SOEs and credit misallocation. The most closely related work is by Geng and Pan (2022). They study the extent of SOE premiums in the Chinese credit market due to perceived government support. Their analysis is based on the same policy event in 2018 used in this paper to study the domestic credit market segmentation between SOEs and non-SOEs.

## 2.4 Data

The main analysis of this paper is based on five commercially available databases. The pieces of the puzzle on Chinese corporate behaviors in tax havens are brought together through varying sources of information. The main dataset combines corporate bond-level data from the S&P Capital IQ platform (CIQ) and Refinitiv Eikon (Eikon) with firm-level data from China Stock Market & Accounting Research (CSMAR), RESSET, and Worldscope.

### A Bond-level Data (Offshore)

The sources for offshore bond data are CIQ and Eikon. The two sources cover partially overlapping but largely different sets of bond issuances. By combining the two, I am able to obtain a relatively comprehensive picture of Chinese bond issuance in tax haven countries.

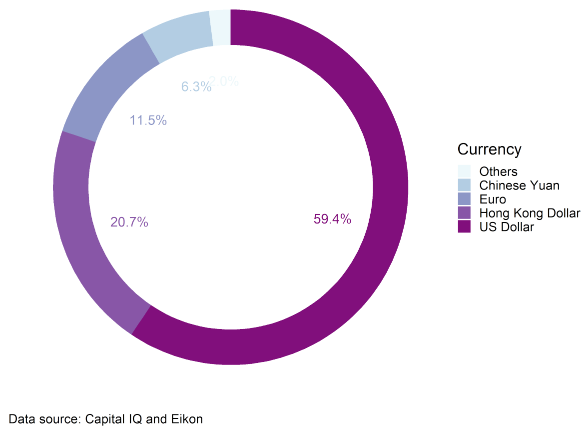
To retrieve the set of bonds issued by Chinese companies incorporated in tax havens, several steps are followed. As a first step, I gather all bonds issued by all firms incorporated in tax havens from CIQ and Eikon between 2010 and 2020. To consolidate the two sets of bond data from these two sources, I merge them by eliminating the duplicated bond-level entries.[4]. In total, CIQ lists 8842 bonds issued by firms in tax havens between 2010 and 2020, while Eikon reports 5041. After the merging procedure, the combined sample consists of 13701 distinct bond issuances, which indicates that only 182 bonds are repeatedly reported across the two data sources. This shows that the two databases rely on largely different sources to collect bond data.

As a second step, I focus on identifying bonds issued by shell companies of Chinese firms. Several criteria are applied to establish the company's connection with China including Hong Kong: (i) the headquarter of the bond issuer is located in China or Hong Kong; (ii) the name of the bond issuer or bond issuer's ultimate parent contains "China" or "Hong Kong"; (iii) the native language of the issuer company name or parent company name is Chinese; (iv) the website of the issuer or issuer's parent company ends with ".cn" or ".hk", which are the country codes of internet domains of Chinese entities. This is a round of coarse screening to identify Chinese companies, which could potentially include firms that are not Chinese but share those criteria (for example Singaporean or Japanese firms).

At a later stage of merging bond-level with firm-level data, this set of bond data will be screened to a finer extent by verifying if it can be matched to a publicly listed company in China.

[4]Overlap in bond data is identified by cross-referencing bond offering dates, maturity dates, and common firm identifiers (CIQ identifier, company name, and website) of the issuers. Bonds sharing identical characteristics across these variables are considered duplicates.

Figure 2.2: Currency Composition of Offshore Bonds Issued by Chinese Firms




Notes: This figure depicts the currency composition of bonds issued by Chinese firms incorporated in tax haven countries between 2010-2020. There are 2338 bonds with currency information out of a total number of 3723.

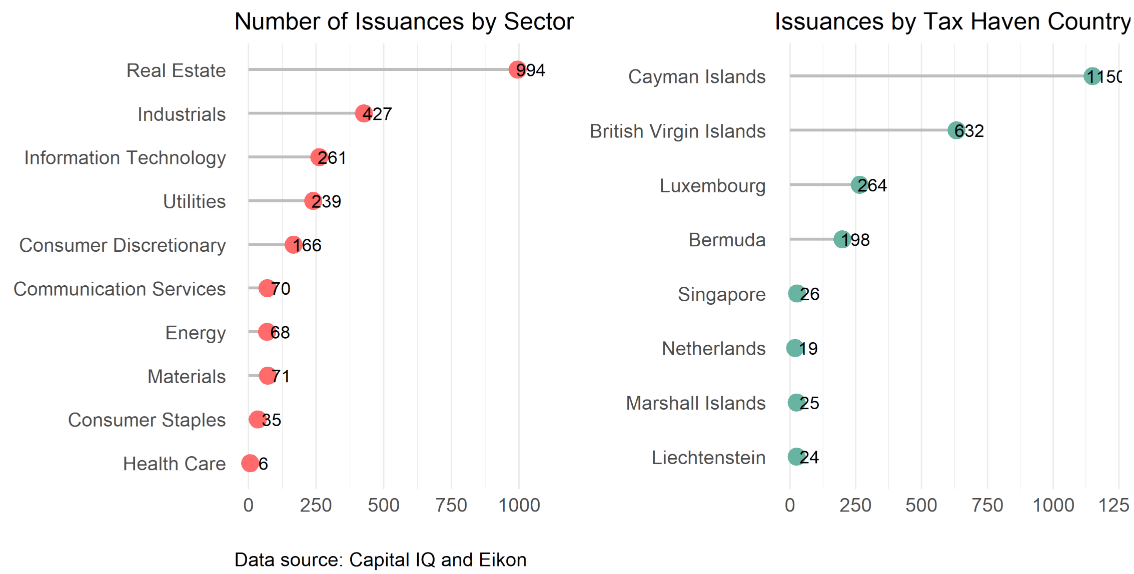
If a bond can be paired up with a parent company in China, I assume that it is issued offshore by a Chinese company. Otherwise, it is dropped out of the final sample. It is important to acknowledge the potential presence of false negatives in cases where an offshore firm is Chinese but remains unidentified in the screening process. This suggests that the estimated regression coefficient aligns with the lower boundary of the actual effects.

The final bond data consists of 3723 bonds issued by Chinese firms incorporated in tax havens for the period 2010-2020, issued by 592 firms. Figure 2.2 shows the currency composition of all the bonds. Nearly 60% of the bonds are issued in US dollars and about 20% are issued in HK dollars. Both are international currencies that are widely accepted in mainland China.

Figure 2.3 demonstrates the industry composition and number of issuances by originating countries. Nearly 42% of firms are in the real estate industry, which highlights that real-estate firms tend to use tax havens as a channel to absorb international capital and round-trip it back to China. This corresponds to the reality that real-estate firms in mainland China are highly indebted and strictly regulated due to the government's fears of the housing bubble. Consequently, they escape domestic regulations to raise international capital offshore. In terms of country composition, Cayman Island is the most popular destination owing to its large financial sector and its close ties with the US capital market. British Virgin Island is the second most popular destination due to its historic link with Hong Kong.

The dataset covers basic information about bonds (coupon rate, offering date, maturity date, amount, currency, etc) and their issuers (residence country, parent company, headquarter, date of

Figure 2.3: Industry and Country Compositions of Offshore Bonds Issued by Chinese Firms




Notes: The left figure demonstrates the sector composition of bonds issued by Chinese firms incorporated in tax haven countries. The right figure shows the composition of the originating countries. There are 2340 bonds with this information out of 3723.

incorporation, etc). While it is true that CIQ and Eikon do not cover all bonds issued between 2010 and 2020, this coverage limitation does not impede the explorations of firm-level motivations for offshore bond issuance. Thus, the representative sample from CIQ and Eikon serves this purpose.

## B Firm-level Data

The firm-level balance sheet data is sourced from CSMAR and RESSET from 2010 to 2020. The focus of the study is on publicly listed Chinese companies in the Shanghai, Shenzhen, and Hong Kong Stock Exchanges. This choice is based on the assumption that listed firms represent the primary group capable of establishing shell companies and raising funding through this channel. CSMAR covers all the listed companies in mainland China in the Shanghai and Shenzhen Stock Exchange, while RESSET provides information on Hong Kong-based listed companies. I include only non-financial corporations.

The final firm sample consists of 5677 firms. Summary statistics are reported in Table 2.1. Panel A reports relevant firm-level financial variables from the balance sheet. The key variables of bank borrowing and cash inflow from the market indicate two major sources of funding for publicly listed firms. The other financial variables included in the summary statistics are control variables. The selection of control variables follows Yuan, Ouyang, and Zhang (2022) and covers variables that affect corporate borrowing structure. It includes the log of total assets, profitability (the ratio of net profits over total assets), tangibility (the ratio of fixed assets over total assets), and liability (the ratio of total liability over total assets). Financial variables are winsorized at 1% and 99% level by year. The final sample is restricted to firms with reported borrowing activities from the credit market.

Table 2.1: Summary Statistics

| Panel A: Publicly Listed Company Financial Variables (CSMAR and RESSET) | | | | | | |
|---|---|---|---|---|---|---|
| Statistic | N | Mean | St. Dev. | Pctl(25) | Median | Pctl(75) |
| $(\text{Cash Inflow from Market/Total Assets})_{it}$ | 31,812 | 0.08 | 0.78 | 0.00 | 0.004 | 0.08 |
| $(\text{Bank Borrowing/Total Assets})_{it}$ | 29,576 | 0.18 | 0.43 | 0.03 | 0.13 | 0.26 |
| $\ln(\text{Total Assets})_{it}$ | 31,812 | 21.99 | 1.62 | 20.94 | 21.85 | 22.94 |
| $(\text{Net Profit/Total Assets})_{it}$ | 31,072 | 0.03 | 0.32 | 0.01 | 0.04 | 0.07 |
| $(\text{Total Liability/Total Assets})_{it}$ | 31,811 | 0.46 | 0.39 | 0.26 | 0.43 | 0.61 |
| $(\text{Fixed Assets/Total Assets})_{it}$ | 31,798 | 0.20 | 0.17 | 0.06 | 0.16 | 0.29 |
| **Panel B: Dummy Variable** | | | | | | |
| Statistic | N | 1 | 0 | | | |
| State-owned Enterprise (SOE) | 31,812 | 10,233 | 21,579 | | | |
| Issuance in Tax Haven | 31,812 | 384 | 31,428 | | | |

Note: This table reports the summary statistics of variables used in regressions.

Another important aspect of firm-level data is the nature of a firm, specifically if it is state-owned or privately owned. The two key variables of defining the SOE nature of a firm are selected based on Geng and Pan (2022). The first one is an SOE dummy variable furnished by CSMAR and RESSET, indicating if the ultimate control of a firm belongs to the government. This is a piece of information that a publicly listed company has to disclose in its annual financial report. A second measure gauging the SOE nature is the share of ownership that belongs to the government. This is based on the top ten shareholders' information (ownership share and owner's nature) of a publicly listed company, which is also required for disclosure. The top ten shareholders, even though not covering the full picture of the ownership nature, provide a representative proxy to compare the relative difference in shares of a firm owned by the state or private entities. In the firm sample, the average holding percentage for the top ten shareholders is 59%, similar to what is reported in Geng and Pan (2022), which uses a different data source (Wind Financial Information Database) and reports 61.2%. Based on the provided ownership information, I calculate the ratio of private ownership share to the total reported share to rescale it.

## C Merging Offshore Bond-level Data with Firm-level Data

The final sample is constructed by merging bond-level data with firm balance sheet data. Several company identifiers are used to match bond entries with its issuing firms.[5] Most bonds issued offshore can be paired with their parent firms based in China within this step using firm identifiers. For those in the sample that cannot be paired with an issuing firm, I resort to manual merging. The main source of information for manual inspection is the website associated with the bond provided by CIQ and Eikon. I glean information on parent-child firm relationships, and mergers and acquisitions from company websites to verify if it can be identified with any firm in CSMAR and RESSET. This manual checking step also complements the previous coarse screening to exclude bonds issued by non-Chinese firms. In the end, 22 bonds are excluded from the sample. They are not issued by Chinese companies, or issued by companies that are not publicly listed, have been delisted, or are publicly listed outside China.

These stages of merging and filtering result in a final merged sample of 6946 firms (104278 firm-year observations) where 155 firms have issued bonds in tax havens in different years (384 firm-year observations). This represents approximately 3% of the entire firm sample.

## D Tax Haven Countries

The list of tax haven countries follows Coppola et al. (2021). Hong Kong is excluded, thus not being considered a tax haven. The main assumption of doing so is that publicly listed Chinese firms in Hong Kong are also affected by the new regulation, therefore not being considered as a destination of macroprudential spillover. This is based on the fact that many Hong Kong firms conduct their main business activities in mainland China and share close links with the credit market and local banks. When the new regulation on WMPs was implemented in 2018, Hong Kong firms were impacted in the same way as firms in mainland China.

# 2.5 Results

## A Effectiveness of macroprudential regulation (onshore)

I begin the analysis by examining if the regulation on WMPs has effectively reduced the firm's domestic borrowing through bond issuances and if it aligns with its objective. The empirical

[5] Company identifiers used are Legal Entity Identifier (LEI), CUSIP, Stock Ticker Symbol, SEDOL, ISIN, and company website.

strategy is a DID setup that exploits the policy experiment in 2018. I intend to use this policy change as a source of shock to the availability of domestic credit for SOEs and non-SOEs. I estimate the following equation for each firm $i$ in year $t$:

$$\mathrm{Y}_{it} = \beta_0 + \beta_1 \mathrm{NSOE}_i \times \mathrm{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \tag{2.1}$$

where $\mathrm{Y}_{it}$ represents a set of outcome variables related to a firm $i$'s borrowing or financing through different channels in year $t$: bank borrowing, cash inflow from the market, total borrowing, and retained earnings. The first two variables represent flow concepts from a firm's cash flow statement, indicating funding sourced from banks and the domestic credit market. Bank borrowing refers to the loans obtained from banks. Cash inflow from the market represents the funding raised through the issuance of bonds and stocks in the domestic credit market. Even though this variable covers both bonds and equities, I use it to measure the changes in bond issuance before and after the new regulation in 2018. This choice is based on two reasons. First, it is due to data limitation that firms in the dataset do not report borrowing from bonds and equity separately. Second, there are no major reforms in the stock market that have coincided with the new regulation in 2018. Hence, substantial changes in cash inflow from the market are likely the result of variations in bond issuance. Total borrowing is the sum of market and bank borrowing, indicating a firm's external financing from banks and credit market. Retained earnings represent the profits held back by the company rather than distributed as dividends and can therefore be considered as an internal source of financing for investment activities.

$\mathrm{NSOE}_i$ is a dummy variable of one if the firm is private (non-state owned) and $\mathrm{Post}_t$ is a dummy variable of one if it is the year of 2018 or afterward. The regression controls for firm and year fixed effects ($\delta_i$ and $\delta_t$). Control variables are denoted by $X_{it}$, which includes firm-level characteristics that are related to corporate borrowing structure, i.e. firm size (the log of total assets), profitability (net profits/total assets), tangibility (fixed assets/total assets), liability (total liability/total assets). The standard errors are clustered at a city level given that local bank branches and local credit markets are important sources of financing for firms.

Table 2.12 summarizes results on a firm's cash inflow from market and bank borrowing in the domestic credit market. Both outcome variables are log-transformed. The sample is restricted to firms that engaged in both bank borrowing and bond/equity issuance between 2010 and 2020. Firms that have not participated in either activity during this period are excluded.

The first column provides raw results on cash inflow from market without control variables. The coefficient is significantly negative, confirming that the regulation has taken effect for non-SOEs. Compared to their state-owned peers, they have experienced unfavorable market reactions. Consequently, they have reduced their borrowing from the credit market after the regulation.

Table 2.2: Onshore Effects on Market and Bank Borrowing

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Cash Inflow from Market) | | ln(Bank Borrowing) | |
| $NSOE_i \times Post_t$ | -0.802* | -1.143*** | 1.205*** | 0.793*** |
| | (0.315) | (0.307) | (0.340) | (0.220) |
| $\ln(\text{Total Assets})_{it}$ | | 2.420*** | | 1.783*** |
| | | (0.340) | | (0.119) |
| $(\text{Net Profits/Total Assets})_{it}$ | | -0.662+ | | -0.188 |
| | | (0.384) | | (0.389) |
| $(\text{Fixed Assets/Total Assets})_{it}$ | | -10.678*** | | 2.966*** |
| | | (1.461) | | (0.334) |
| $(\text{Total liability/Total Assets})_{it}$ | | -3.323*** | | 2.334*** |
| | | (0.846) | | (0.629) |
| $N$ | 26673 | 26191 | 26673 | 26191 |
| adj. $R^2$ | 0.338 | 0.379 | 0.519 | 0.547 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Outcome variables are the log of cash inflow from market and bank borrowing for firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

When control variables are added in the second column, the significance of the coefficient further increases, underlining the importance of controlling for corporate borrowing structure to obtain a more precise estimation. This result underlines the effectiveness of the regulation, in particular for non-SOEs. It is consistent with anecdotal evidence and previous literature on SOE premium and market discrimination against non-SOEs (Geng and Pan, 2022; Bai et al., 2020; Dollar and Wei, 2007).

Results on domestic bank borrowing are summarized in Columns 3 and 4. Column 3 shows estimation results without control variables, which indicate an increase in bank loans in the post-reform period on a firm level for non-SOEs. Results in Column 4 are with control variables and confirm what we observe in Column 3. The estimation implies that private firms, faced with funding cuts in the credit market, resort back to bank lending to finance their activities.

Table 2.21 summarizes results on a firm's total borrowing and retained earnings. The first two columns present results on the total borrowing, with and without control variables. The coefficient in the first column is not significant. However, when we impose control variables in the second column, the coefficient becomes significantly negative, indicating an 11% drop in a non-SOE's total borrowing compared to its SOE peers. This attests to the effectiveness of the macroprudential policy and shows that non-SOEs have taken on less debt since the reform. To-

Table 2.3: Onshore Effects on Total Borrowing and Retained Earning

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Total Borrowing) | | ln(Retained Earnings) | |
| $NSOE_i \times Post_t$ | 0.093 | -0.115*** | -2.280*** | -2.485*** |
| | (0.089) | (0.033) | (0.515) | (0.382) |
| ln(Total Assets)$_{it}$ | | 1.205*** | | 3.279*** |
| | | (0.071) | | (0.205) |
| (Net Profits/Total Assets)$_{it}$ | | 0.044 | | 2.423** |
| | | (0.057) | | (0.886) |
| (Fixed Assets/Total Assets)$_{it}$ | | -0.919*** | | -2.302** |
| | | (0.178) | | (0.851) |
| (Total liability/Total Assets)$_{it}$ | | 0.603*** | | -5.055*** |
| | | (0.146) | | (0.959) |
| $N$ | 26673 | 26191 | 26537 | 26084 |
| adj. $R^2$ | 0.713 | 0.777 | 0.634 | 0.665 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Outcome variables are the log of total borrowing and retained earnings by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

gether with results from Table 2.12, it reveals that they are not able to fully recover the reduction in cash inflow from the credit market despite an increase in bank loans. There is a potential funding gap on their balance sheet, suggesting crowd-out effects from the domestic market and the possibility of looking for alternative financing sources.

Results in Columns 3 and 4 present changes in retained earnings, a firm's internal channel of adjustment after macroprudential policies. Both coefficients are significantly negative, confirming the deteriorated financial position of non-SOEs.

## B Effectiveness of macroprudential regulation for real estate sector (onshore)

To look at heterogeneous effects on a sector level and focus on firms in real estate, I extend the DID setup to a DDD specification by interacting the Post and NSOE dummies with a sector dummy of Real Estate, which indicates if a firm $i$ is in the real estate sector. The regression I run

Table 2.4: Onshore Effects on Market and Bank Borrowing of Firms in Real Estate Sector

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Cash Inflow from Market) | | ln(Bank Borrowing) | |
| $\text{NSOE}_\text{i} \times \text{Post}_\text{t}$ | -0.877** | -1.118*** | 1.286*** | 0.868*** |
| | (0.291) | (0.301) | (0.333) | (0.213) |
| $\text{Post}_\text{t} \times \text{Real Estate}_\text{i}$ | 0.659 | 0.467 | -0.183 | -0.428** |
| | (0.698) | (0.851) | (0.207) | (0.164) |
| $\text{NSOE}_\text{i} \times \text{Post}_\text{t} \times \text{Real Estate}_\text{i}$ | 0.676 | -0.275 | -0.767* | -0.721** |
| | (0.536) | (0.600) | (0.322) | (0.261) |
| $\ln(\text{Total Assets})_\text{it}$ | | 2.416*** | | 1.795*** |
| | | (0.336) | | (0.118) |
| $(\text{Net Profits/Total Assets})_\text{it}$ | | -0.663+ | | -0.185 |
| | | (0.383) | | (0.389) |
| $(\text{Fixed Assets/Total Assets})_\text{it}$ | | -10.688*** | | 2.996*** |
| | | (1.476) | | (0.338) |
| $(\text{Total liability/Total Assets})_\text{it}$ | | -3.321*** | | 2.314*** |
| | | (0.842) | | (0.626) |
| $N$ | 26673 | 26191 | 26673 | 26191 |
| adj. $R^2$ | 0.338 | 0.379 | 0.519 | 0.548 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.2). Outcome variables are the log of cash inflow from market and bank borrowing by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

is:

$$\text{Y}_{it} = \beta_0 + \beta_1 \text{NSOE}_i \times \text{Post}_t \times \text{Real Estate}_i + \beta_2 \text{NSOE}_i \times \text{Real Estate}_i + \beta_3 \text{Post}_t \times \text{Real Estate}_i + \beta_4 \text{NSOE}_i \times \text{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \quad (2.2)$$

where the set of outcome variables, independent variables, and control variables are the same as in Regression (2.10). This setup corresponds to the exploratory analysis in Figure 2.3 where real estate firms account for the majority of entities issuing fixed-income securities through tax havens.

Table 2.14 illustrates the outcomes of corporate borrowing within the real estate sector through the domestic credit market and banking institutions. The principal coefficients related to the triple interaction in Columns 1 and 2 are insignificant. This suggests that non-SOEs in the real estate sector did not further cut their borrowing from the credit market compared to SOEs in other sectors post-reform. Considering the already stringent borrowing conditions proceeding 2018 for

Table 2.5: Onshore Effects on Total Borrowing and Retained Earnings of Firms in Real Estate Sector

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Total Borrowing) | | ln(Retained Earnings) | |
| $NSOE_i \times Post_t$ | 0.093 | -0.105** | -2.586*** | -2.700*** |
| | (0.097) | (0.037) | (0.617) | (0.464) |
| $Post_t \times Real\ Estate_i$ | 0.165 | 0.030 | 0.108 | -0.184 |
| | (0.108) | (0.058) | (0.671) | (0.620) |
| $NSOE_i \times Post_t \times Real\ Estate_i$ | -0.007 | -0.102 | 2.969*** | 2.152** |
| | (0.115) | (0.072) | (0.740) | (0.677) |
| $\ln(\text{Total Assets})_{it}$ | | 1.206*** | | 3.260*** |
| | | (0.071) | | (0.209) |
| $(\text{Net Profits/Total Assets})_{it}$ | | 0.044 | | 2.419** |
| | | (0.057) | | (0.887) |
| $(\text{Fixed Assets/Total Assets})_{it}$ | | -0.918*** | | -2.344** |
| | | (0.178) | | (0.849) |
| $(\text{Total liability/Total Assets})_{it}$ | | 0.602*** | | -5.020*** |
| | | (0.145) | | (0.959) |
| $N$ | 26673 | 26191 | 26537 | 26084 |
| adj. $R^2$ | 0.713 | 0.777 | 0.634 | 0.666 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.2). Outcome variables are the log of total borrowing and retained earnings for firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

firms in the real estate sector, the macroprudential policy did not invoke any particular market effects targeting non-SOEs in the real estate sector.

As shown in Columns 3 and 4, banks demonstrate an even more heightened aversion towards real estate firms' loan requests after the policy announcement. The triple interaction coefficient for bank borrowing is significantly negative, while it registers a positive significance for overall non-SOEs. This indicates that banks display an increased reluctance to lend to the real estate sector post-policy announcement.

The results of total borrowing and retained earnings are presented in Table 2.15. The lack of significance in the triple interaction term for total borrowing mirrors the findings observed in cash inflow from market in Table 2.14. This implies that the regulation did not generate any significantly different borrowing constraints for non-SOEs within the real estate sector compared to SOE firms in other industries.

The key coefficients on retained earnings for the triple interaction in Columns 3 and 4 are significantly positive, suggesting that non-SOEs in the real estate sector were able to accumulate their net income, even as other non-SOEs had to reinvest this internal funding to offset diminished borrowing from other sources. This outcome may be attributed to the potential capital raised by subsidiaries located outside of China.

## C Spillovers of macroprudential regulation in tax havens (offshore)

I now examine if the implementation of the new regulation on WMPs generates spillovers of corporate bond issuance in tax havens for non-SOEs post-reform. I evaluate the spillover effects through the following equation:

$$\mathrm{TH}_{it} = \beta_0 + \beta_1 \mathrm{NSOE}_i \times \mathrm{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \tag{2.3}$$

where $\mathrm{TH}_{it}$ represents a dummy variable of one if a firm $i$ issues bonds in tax havens in year $t$. The independent variables remain the same as in Regression 2.10. The only change is the addition of two more control variables: the log of cash inflow from market from year $t$ and year $t-1$. They reflect the onshore borrowing activities for each firm and, therefore are essential to be controlled to obtain a relatively precise estimation of a firm's offshore borrowing behavior.

Table 2.16 summarizes the main results. The first column provides the estimation results for the full firm sample without any control variables. The estimate is positive but not significant. This may represent an underestimate of the spillover effect considering that a number of firms in the full sample do not report borrowing from the domestic credit market and the new regulation would not affect their bond-issuing behavior. In the second column, I restrict the sample to firms with activities related to cash inflow from the market only. As expected, the point estimate increases to 0.7%, statistically significant at 5% level. It indicates that a private firm, compared to a state-owned one, is 0.7% more likely to finance bonds through shell companies in tax havens after the 2018 new macroprudential regulation. In the third column, I add control variables. The coefficient increases to 0.9%, significant at 5% level, confirming the unintended spillover effects.

## D Spillovers from real estate sector (offshore)

To examine the spillover effects from non-SOEs in the real estate sector, I extend the DID setup to a DDD specification:

$$\begin{aligned}\mathrm{TH}_{it} = \beta_0 &+ \beta_1 \mathrm{NSOE}_i \times \mathrm{Post}_t \times \text{Real Estate}_i + \beta_2 \mathrm{NSOE}_i \times \text{Real Estate}_i + \beta_3 \mathrm{Post}_t \times \text{Real Estate}_i \\ &+ \beta_4 \mathrm{NSOE}_i \times \mathrm{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}. \end{aligned} \tag{2.4}$$

Table 2.6: Spillovers of Macroprudential Regulation in Tax Havens

| | Issuance in Tax Haven$_{it}$ | | |
|---|---|---|---|
| | Full sample | Restricted | Restricted |
| | (1) | (2) | (3) |
| NSOE$_i$×Post$_t$ | 0.002 | 0.007* | 0.009* |
| | (0.002) | (0.003) | (0.004) |
| ln(Total Assets)$_{it}$ | | | 0.011*** |
| | | | (0.002) |
| (Net Profits/Total Assets)$_{it}$ | | | 0.002 |
| | | | (0.003) |
| (Fixed Assets/Total Assets)$_{it}$ | | | 0.016+ |
| | | | (0.010) |
| (Total liability/Total Assets)$_{it}$ | | | 0.002 |
| | | | (0.003) |
| ln(Cash Inflow from Market)$_{it}$ | | | -0.000 |
| | | | (0.000) |
| ln(Cash Inflow from Market)$_{it-1}$ | | | -0.000** |
| | | | (0.000) |
| *N* | 43890 | 30979 | 23097 |
| adj. $R^2$ | 0.319 | 0.314 | 0.335 |
| Firm and Year FE | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.3). The outcome variable is a dummy variable of one if a firm $i$ issues at least one bond in tax haven countries in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

The results are provided in Table 2.17. Column 1 summarizes the results without any control variables. $\beta_1$ evaluates the mean difference in the likelihood of issuing bonds in tax havens between an SOE in non-real estate sectors and a non-SOE in the real estate sector before and after the new macroprudential regulation. The coefficient is 9.7% and statistically significant at 1% level. Compared to the results from the previous specification Equation (2.3), the emphasis on private firms in the real estate sector brings about a significant upward shift in coefficient scale on a stricter significance level. $\beta_3$ looks at the mean difference in spillover effects between firms in real estate and non-real estate sectors. It is positively significant and confirms that spillovers are more likely associated with firms in the real estate sector despite the private or public nature of the firm. In addition, what is interesting to observe is that the spillover effects for non-SOEs after the reform ($\beta_4$) now become negatively significant after the strong effects from the real estate sector are purged out. It underscores a potential double “crowd-out" effect, where non-SOEs in the non-real estate sector are pushed away both in the domestic and offshore credit market.

In Column 2 where control variables are added, the key coefficient $\beta_4$ associated with the triple interaction item increases to 18.7%, doubling the previous coefficient scale. This result is consistent with Table 2.16 where the coefficient scale increases after the control of firm-level characteristics. In the third column, I cluster standard errors at a city level given that local bank branches are important sources of financing. The point estimate remains similar to Column 2, and the standard errors are larger, as expected. However, the coefficient still stays significant at 1% level, underlining the robustness of this result. This estimation result points to the fact that the spillovers of this macroprudential policy are mostly driven by non-SOEs in the real estate sector, whose domestic financing sources are severely shrunk after the new regulation of WMPs.

## E Leads and lags

I estimate a specification with leads and lags to measure pre-reform and post-reform trends:

$$\mathrm{TH}_{it} = \beta_0 + \sum_{k=2011}^{2020} \beta_k \mathrm{NSOE}_i \times \mathrm{Year}_k + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \tag{2.5}$$

where $\mathrm{Year}_k$ is a year indicator variable. The reference year is 2017, the year before the reform. The specification includes six lags and three leads, corresponding to the studied period of 2011-2020. Figure 3.3 plots the $\beta_k$ coefficients and its 95% confidence intervals. As it shows, there is no obvious trend before the new regulation in 2018. The spillovers slowly take effect in the year after the announcement of the new regulation and become more significant in the second year as non-SOEs are more likely to issue bonds in tax havens post-reform.

Table 2.7: Spillovers from Firms in Real Estate Sector

| | (1) | (2) | (3) |
|---|---|---|---|
| | Issuance in Tax $\text{Haven}_{it}$ | | |
| $\text{NSOE}_i \times \text{Post}_t$ | -0.006** | -0.011** | -0.011* |
| | (0.002) | (0.004) | (0.005) |
| $\text{Post}_t \times \text{Real Estate}_i$ | 0.033*** | 0.055*** | 0.055 |
| | (0.006) | (0.010) | (0.035) |
| $\text{NSOE}_i \times \text{Post}_t \times \text{Real Estate}_i$ | 0.097*** | 0.187*** | 0.187*** |
| | (0.007) | (0.012) | (0.022) |
| $\ln(\text{Total Assets})_{it}$ | | 0.008*** | 0.008 |
| | | (0.002) | (0.005) |
| $(\text{Net Profits/Total Assets})_{it}$ | | 0.002 | 0.002** |
| | | (0.003) | (0.001) |
| $(\text{Fixed Assets/Total Assets})_{it}$ | | 0.013 | 0.013 |
| | | (0.009) | (0.017) |
| $(\text{Total liability/Total Assets})_{it}$ | | 0.004 | 0.004+ |
| | | (0.003) | (0.003) |
| $\ln(\text{Cash Inflow from Market})_{it}$ | | -0.000 | -0.000** |
| | | (0.000) | (0.000) |
| $\ln(\text{Cash Inflow from Market})_{it-1}$ | | -0.000** | -0.000+ |
| | | (0.000) | (0.000) |
| $N$ | 43890 | 23097 | 23097 |
| adj. $R^2$ | 0.339 | 0.371 | 0.371 |
| Firm and Year FE | Yes | Yes | Yes |
| SE clustered | Yes | No | Yes |

Note: This table reports the coefficients from the DID regression (2.4). The outcome variable is a dummy variable of one if a firm $i$ issues at least one bond in tax haven countries in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Figure 2.4: Leads and Lags of Spillover Effects

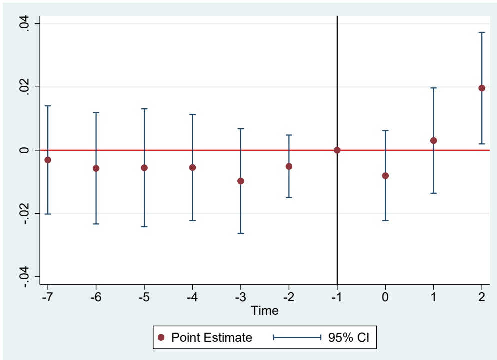
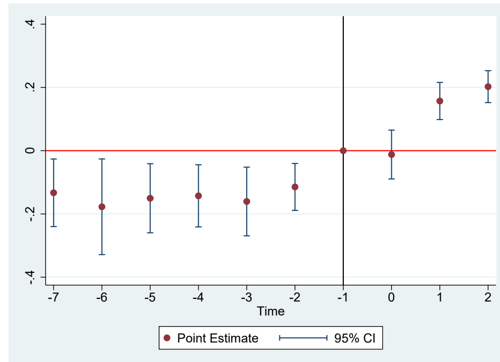



Figure 2.5: Leads and Lags of Spillover Effects of Firms in Real Estate Sector




Note: These two figures plot the coefficients of $\beta_k$ in Regressions (3.5) and (2.6). The reference year is 2017, the year before the implementation of the macroprudential policy.

A similar coefficient plot featuring non-SOEs in the real estate sector is provided as Figure 2.5. The specification is as follows:

$$\mathrm{TH}_{it} = \beta_0 + \sum_{k=2011}^{2020} \beta_k^1 \mathrm{NSOE}_i \times \text{Real Estate}_i \times \mathrm{Year}_k + \sum_{k=2011}^{2020} \beta_k^2 \mathrm{NSOE}_i \times \mathrm{Year}_k + \sum_{k=2011}^{2020} \beta_k^3 \text{Real Estate}_i \times \mathrm{Year}_k + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it} \quad (2.6)$$

Aligned with the DDD regression estimations, the post-reform trend exhibits increased bond issuances among non-SOEs within the real estate sector. Notably, the confidence intervals have narrowed indicating a higher level of statistical significance. Furthermore, it is noteworthy that the pre-reform coefficients are negative, suggesting that non-SOEs in the real estate sector were less likely to issue offshore bonds in tax havens prior to the implementation of the new regulation. This highlights the post-reform spillover effects for non-SOEs within the real estate sector as identified in the previous estimations.

## F Spillovers linked to firm's private ownership

Geng and Pan (2021) argue that not all SOEs are the same and the market has grown more sensitive to the percentage of government holdings in SOEs and non-SOEs. In the following specification, I replace the $\mathrm{NSOE}_i$ dummy with the percentage of equity owned by private entities in a company, as mentioned in Section 3.3. It provides a yearly representation of the share of private ownership in a company. Instead of treating all non-SOEs as one dummy group, this variable provides a way to continuously measure the spillovers associated with the percentage of private ownership. The outcome variable is the log of the total amount of offshore bond funding for firm $i$ in year $t$.

$$\ln(\text{Offshore Bond Amount})_{it} = \beta_0 + \beta_1 \text{Private Percent}_{it} \times \text{Post}_t \times \text{Real Estate}_i + \beta_2 \text{Post}_t \times \text{Real Estate}_i + \beta_3 \text{Private Percent}_{it} \times \text{Real Estate}_i + \beta_4 \text{Private Percent}_{it} \times \text{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}. \quad (2.7)$$

Table 2.18 provides the estimation results. Column 1 summarizes the results without any control variables. The point estimate of $\beta_1$ suggests that a 1% increase in private ownership for a firm in the real estate sector leads to a 0.67% increase in the amount of funding raised through bond issuance by shell companies in tax havens. In Column 2 where control variables are specified, the coefficient rises to 1% and is significant at 1% level. In Column 3 where standard errors are clustered on a city level, the point estimate stays the same and is statistically significant at a 10% level.

This result is also economically significant, indicating that a 1% rise in private ownership is associated with a 1% increase in bonds issued in tax havens by non-SOEs in the real estate sector. This highlights the substantial spillover effects attributed to the macroprudential policy from private enterprises within the real estate sector, compared to its SOE peers across other industries.

## G Real effects and risk implication

To examine the impact of offshore financing on a firm's balance sheet, I run the following regression:

$$\text{Y}_{it} = \beta_0 + \beta_1 \text{TH}_{it} \times \text{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \quad (2.8)$$

where $\text{Y}_{\text{it}}$ represents the log of fixed assets, sales, cash, and costs of employees. The other variables in the regression are as defined in previous estimations. The results are shown in Table 2.9. The coefficients $\beta_1$ for the log of fixed assets and sales are not significant, suggesting that financing through offshore sources does not significantly impact the firm's investment and revenue. However, the coefficients for the log of cash and wage are significantly positive. This indicates that firms incorporated in tax havens mostly round-trip funds from offshore to stock up cash reserves and expand employment.

I extend the previous regression setup to a triple interaction including a real estate sector

Table 2.8: Spillovers Linked to Private Ownership

| | (1) | (2) | (3) |
|---|---|---|---|
| | ln(Issuance Amount in Tax Haven)$_{it}$ | | |
| Private ownership$_{it}$ | -0.004 | -0.024 | -0.024 |
| | (0.026) | (0.047) | (0.068) |
| Post$_t$×Private Ownership$_{it}$ | -0.049** | -0.088** | -0.088 |
| | (0.015) | (0.032) | (0.054) |
| Real Estate$_i$×Private Ownership$_{it}$ | -0.375** | -0.550** | -0.550* |
| | (0.115) | (0.201) | (0.255) |
| Post$_t$×Real Estate$_i$ | -0.100** | -0.155* | -0.155 |
| | (0.037) | (0.068) | (0.144) |
| Post$_t$×Real Estate$_i$×Private Ownership$_{it}$ | 0.675*** | 1.033*** | 1.033+ |
| | (0.063) | (0.112) | (0.530) |
| ln(Total Assets)$_{it}$ | | -0.003 | -0.003 |
| | | (0.012) | (0.016) |
| (Net Profits/Total Assets)$_{it}$ | | -0.002 | -0.002 |
| | | (0.014) | (0.003) |
| (Fixed Assets/Total Assets)$_{it}$ | | -0.098 | -0.098+ |
| | | (0.062) | (0.056) |
| (Total liability/Total Assets)$_{it}$ | | -0.013 | -0.013 |
| | | (0.026) | (0.011) |
| ln(Cash Inflow from Market)$_{it}$ | | -0.001 | -0.001 |
| | | (0.001) | (0.001) |
| ln(Cash Inflow from Market)$_{it-1}$ | | -0.001 | -0.001 |
| | | (0.001) | (0.000) |
| $N$ | 19777 | 10923 | 10923 |
| adj. $R^2$ | 0.241 | 0.236 | 0.236 |
| Firm and Year FE | Yes | Yes | Yes |
| SE clustered | Yes | No | Yes |

Note: This table reports the coefficients from the DID regression (2.7). The outcome variable is the log of the total amount of offshore bond funding for firm $i$ in year $t$. The variable "Private ownership" refers to the share of private ownership of firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.9: Effects on Fixed Assets, Sales, Cash and Wage

| | (1) $\ln(\text{Fixed Assets})_{\text{it}}$ | (2) $\ln(\text{Sales})_{\text{it}}$ | (3) $\ln(\text{Cash})_{\text{it}}$ | (4) $\ln(\text{Wage})_{\text{it}}$ |
|---|---|---|---|---|
| Tax $\text{Haven}_{\text{it}}$ | 0.090*** | 0.056*** | 0.096*** | -0.090+ |
| | (0.006) | (0.014) | (0.022) | (0.046) |
| Tax $\text{Haven}_{\text{it}} \times \text{Post}_{\text{t}}$ | 0.007 | -0.003 | 0.155** | 0.128+ |
| | (0.012) | (0.038) | (0.055) | (0.076) |
| $\ln(\text{Total Assets})_{\text{it}}$ | 1.029*** | 0.865*** | 0.901*** | 0.726*** |
| | (0.017) | (0.046) | (0.015) | (0.023) |
| $(\text{Net Profits/Total Assets})_{\text{it}}$ | -0.023* | 0.061+ | 0.004 | -0.033 |
| | (0.011) | (0.031) | (0.013) | (0.045) |
| $(\text{Fixed Assets/Total Assets})_{\text{it}}$ | 5.917*** | 0.333*** | -1.871*** | 0.652*** |
| | (0.315) | (0.050) | (0.192) | (0.093) |
| $(\text{Total liability/Total Assets})_{\text{it}}$ | -0.004 | 0.021* | -0.028** | 0.033 |
| | (0.004) | (0.010) | (0.009) | (0.038) |
| $N$ | 42449 | 42439 | 42403 | 29663 |
| adj. $R^2$ | 0.952 | 0.934 | 0.869 | 0.985 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Outcome variables are the log of fixed assets, sales, cash, and wage by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

dummy as shown below:

$$\begin{aligned} \text{Y}_{it} = \beta_0 + \beta_1 \text{TH}_{it} \times \text{Post}_t \times \text{Real Estate}_i + \beta_2 \text{TH}_{it} \times \text{Real Estate}_i + \beta_3 \text{Post}_t \times \text{Real Estate}_i \\ + \beta_4 \text{TH}_{it} \times \text{Post}_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}. \quad (2.9) \end{aligned}$$

Table 2.10 reports the estimation results. Notably, the coefficient for fixed assets is positive and significant. It indicates that real estate firms, compared to firms in other sectors, invest their offshore funding into tangible properties. This finding underscores the spillover effects of real estate firms seeking alternative funding sources beyond regulation when domestic credit restrictions aim to reduce the housing bubbles and high leverage in the real estate sector. In addition, the results for sales and cash are not significant for real estate firms, while the coefficient for wage cost is significantly negative.

An important question to ask is whether firms issuing bonds offshore are more financially unstable and riskier, give that they circumvent regulations to obtain funding outside the authorities' supervision. To analyze the risk profiles of these offshore companies, I calculate the Altman's

Table 2.10: Effects on Fixed Assets, Sales, Cash and Wage

| | (1) $\ln(\text{Fixed Assets})_{it}$ | (2) $\ln(\text{Sales})_{it}$ | (3) $\ln(\text{Cash})_{it}$ | (4) $\ln(\text{Wage})_{it}$ |
|---|---|---|---|---|
| Tax $\text{Haven}_{it}$ | 0.145*** | 0.018* | 0.069 | -0.115** |
| | (0.021) | (0.008) | (0.044) | (0.042) |
| Tax $\text{Haven}_{it}\times\text{Post}_{t}$ | -0.067* | -0.021 | 0.044 | 0.145 |
| | (0.029) | (0.026) | (0.053) | (0.089) |
| Tax $\text{Haven}_{it}\times$Real $\text{Estate}_{i}$ | -0.117*** | 0.085*** | 0.084* | 0.265*** |
| | (0.022) | (0.023) | (0.042) | (0.042) |
| $\text{Post}_{t}\times$Real $\text{Estate}_{i}$ | 0.045 | -0.008 | 0.147* | 0.130* |
| | (0.030) | (0.025) | (0.059) | (0.059) |
| Tax $\text{Haven}_{it}\times\text{Post}_{t}\times$Real $\text{Estate}_{i}$ | 0.114** | 0.029 | 0.067 | -0.343+ |
| | (0.042) | (0.036) | (0.052) | (0.191) |
| $\ln(\text{Total Assets})_{it}$ | 1.028*** | 0.865*** | 0.900*** | 0.725*** |
| | (0.016) | (0.046) | (0.015) | (0.023) |
| (Net Profits/Total $\text{Assets})_{it}$ | -0.023* | 0.061+ | 0.004 | -0.032 |
| | (0.011) | (0.031) | (0.013) | (0.045) |
| (Fixed Assets/Total $\text{Assets})_{it}$ | 5.917*** | 0.333*** | -1.871*** | 0.649*** |
| | (0.316) | (0.050) | (0.194) | (0.093) |
| (Total liability/Total $\text{Assets})_{it}$ | -0.004 | 0.021* | -0.028** | 0.034 |
| | (0.004) | (0.010) | (0.009) | (0.038) |
| $N$ | 42449 | 42439 | 42403 | 29663 |
| adj. $R^2$ | 0.952 | 0.934 | 0.869 | 0.985 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Outcome variables are the log of fixed assets, sales, cash, and wage by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Z-score [6] to access the likelihood of insolvency, using this z-score as an outcome variable in regressions 2.8 and 2.9. A lower z-score indicates higher financial risk.

The estimation results are reported in Table 2.11. The coefficient in Column 1 with no control variables is positive and significant, but it becomes insignificant after controlling for a firm's corporate borrowing structure. This suggests that firms with offshore funding do not have a riskier financial profile compared to other firms. However, the result changes for firms in the real estate sector. The coefficient $\beta_1$ for the triple interaction term is negative and significant, suggesting that firms in the real estate sector demonstrate higher financial instability after issuing offshore bonds in tax havens.

This finding has important risk implications for firms in the real estate sector. While offshore financing provides these firms with alternative funding sources, they invest these funds into fixed assets and property development, alleviating the constraints of the tight regulatory environment in the domestic housing market. This offshore financing pattern increases the risk profile of the involved firms, increasing the financial instability in the credit market.

## 2.6 Conclusion

This article explores the spillover effects resulting from the implementation of a Chinese macroprudential policy aimed at regulating WMPs in 2018. My analysis reveals an observable impact of this policy on reducing the borrowing activities of non-SOEs within the domestic credit market. This provokes these non-SOEs to actively seek alternative avenues for financing beyond the regulatory constraints. I detect significant spillover effects of this policy, particularly linked to non-SOEs in the real estate sector, in tax haven countries.

My research explores the importance of tax havens as conduits for Chinese firms to raise international capital and contributes to the understanding of the implications of domestic macroprudential policy on a global scale. This is a crucial question for policymaking, particularly given China's status as one of the primary countries engaging in substantial offshore borrowing facilitated through tax havens. Despite the substantial amount of capital that flows through tax havens, there remains a gap in the existing body of research dedicated to this issue. My study seeks to fill this gap, contributing to our comprehension of the offshore behaviors of Chinese firms.

The consequences of these spillover effects hold substantial importance for the Chinese capital

[6] I use the Altman z-score model for emerging market firms. As in Altman (2005), EM Score = 6.56($X_1$) + 3.26($X_2$) + 6.72($X_3$) + 1.05($X_4$)+ 3.25, where $X_1$ = working capital/total assets, $X_2$ = retained earnings/total assets, $X_3$ = operating income/total assets, $X_4$ = book value of equity/total liabilities.

Table 2.11: Effects on Risks

| | (1) Z Score | (2) Z Score | (3) Z Score | (4) Z Score |
|---|---|---|---|---|
| Tax $\text{Haven}_{it}$ | 0.263 | 0.113 | 0.489* | 0.162 |
| | (0.201) | (0.196) | (0.231) | (0.312) |
| Tax $\text{Haven}_{it}\times\text{Post}_t$ | 0.649* | 0.311 | 0.588 | 0.633+ |
| | (0.251) | (0.355) | (0.370) | (0.331) |
| ln(Total $\text{Assets})_{it}$ | | 1.511*** | | 1.508*** |
| | | (0.326) | | (0.325) |
| (Net Profits/Total $\text{Assets})_{it}$ | | 2.898* | | 2.897* |
| | | (1.303) | | (1.303) |
| (Fixed Assets/Total $\text{Assets})_{it}$ | | -7.244*** | | -7.261*** |
| | | (1.263) | | (1.266) |
| (Total liability/Total $\text{Assets})_{it}$ | | -5.972*** | | -5.970*** |
| | | (0.820) | | (0.818) |
| Tax $\text{Haven}_{it}\times$Real $\text{Estate}_i$ | | | -0.567* | -0.086 |
| | | | (0.224) | (0.338) |
| $\text{Post}_t\times$Real $\text{Estate}_i$ | | | 2.340*** | 1.527*** |
| | | | (0.492) | (0.341) |
| Tax $\text{Haven}_{it}\times\text{Post}_t\times$Real $\text{Estate}_i$ | | | -1.747* | -2.092** |
| | | | (0.713) | (0.712) |
| $N$ | 35018 | 35011 | 35018 | 35011 |
| adj. $R^2$ | 0.511 | 0.797 | 0.511 | 0.797 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

market since the funds raised in tax havens flow back to the domestic credit market. This raises important yet unanswered questions. It is crucial to investigate if it undermines the efficacy of macroprudential policy, and increases the riskiness of involved firms. This calls for more research on related topics on the implications of international capital sourced from tax havens.

# Appendix

## A Regression Results on A Balanced Sample

The estimations of the main results are repeated on a balanced sample, dropping firms that have been de-listed or have newly entered the market. The results are consistent with the main ones.

Table 2.12: Onshore Effects on Market and Bank Borrowing

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Cash Inflow from Market) | | ln(Bank Borrowing) | |
| $NSOE_i \times Post_t$ | -0.619** | -0.773** | 0.965*** | 0.701*** |
| | (0.225) | (0.261) | (0.263) | (0.193) |
| ln(Total Assets)$_{it}$ | | 2.689*** | | 1.819*** |
| | | (0.356) | | (0.131) |
| (Net Profits/Total Assets)$_{it}$ | | -0.392 | | -0.212 |
| | | (0.310) | | (0.395) |
| (Fixed Assets/Total Assets)$_{it}$ | | -10.707*** | | 2.850*** |
| | | (1.133) | | (0.375) |
| (Total liability/Total Assets)$_{it}$ | | -3.040*** | | 2.063** |
| | | (0.844) | | (0.685) |
| $N$ | 19656 | 19391 | 19656 | 19391 |
| adj. $R^2$ | 0.353 | 0.396 | 0.527 | 0.556 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10) on a balanced sample. Outcome variables are the log of cash inflow from market and bank borrowing for firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.13: Onshore Effects on Total Borrowing and Retained Earning

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Total Borrowing) | | ln(Retained Earnings) | |
| $NSOE_i \times Post_t$ | 0.077 | -0.041 | -2.510*** | -2.563*** |
| | (0.109) | (0.045) | (0.542) | (0.355) |
| $\ln(\text{Total Assets})_{it}$ | | 1.218*** | | 3.779*** |
| | | (0.074) | | (0.189) |
| $(\text{Net Profits/Total Assets})_{it}$ | | 0.038 | | 2.128** |
| | | (0.073) | | (0.730) |
| $(\text{Fixed Assets/Total Assets})_{it}$ | | -0.641*** | | -2.416* |
| | | (0.159) | | (0.956) |
| $(\text{Total liability/Total Assets})_{it}$ | | 0.727*** | | -5.303*** |
| | | (0.168) | | (1.294) |
| $N$ | 19656 | 19391 | 19595 | 19342 |
| adj. $R^2$ | 0.729 | 0.795 | 0.654 | 0.687 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10) on a balanced sample. Outcome variables are the log of total borrowing and retained earnings by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.14: Onshore Effects on Market and Bank Borrowing of Firms in Real Estate Sector

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Cash Inflow from Market) | | ln(Bank Borrowing) | |
| $NSOE_i \times Post_t$ | -0.662* | -0.656* | 1.019*** | 0.783*** |
| | (0.267) | (0.284) | (0.287) | (0.208) |
| $Post_t \times Real\ Estate_i$ | 0.335 | 0.167 | -0.204 | -0.420* |
| | (0.770) | (0.893) | (0.170) | (0.164) |
| $NSOE_i \times Post_t \times Real\ Estate_i$ | 0.234 | -0.904 | -0.344 | -0.514 |
| | (0.664) | (0.642) | (0.393) | (0.326) |
| ln(Total Assets)$_{it}$ | | 2.698*** | | 1.832*** |
| | | (0.347) | | (0.131) |
| (Net Profits/Total Assets)$_{it}$ | | -0.391 | | -0.210 |
| | | (0.310) | | (0.393) |
| (Fixed Assets/Total Assets)$_{it}$ | | -10.690*** | | 2.880*** |
| | | (1.148) | | (0.377) |
| (Total liability/Total Assets)$_{it}$ | | -3.048*** | | 2.054** |
| | | (0.842) | | (0.685) |
| $N$ | 19656 | 19391 | 19656 | 19391 |
| adj. $R^2$ | 0.353 | 0.396 | 0.527 | 0.556 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.2) on a balanced sample. Outcome variables are the log of cash inflow from market and bank borrowing by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.15: Onshore Effects on Total Borrowing and Retained Earnings of Firms in Real Estate Sector

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ln(Total Borrowing) | | ln(Retained Earnings) | |
| $NSOE_i \times Post_t$ | 0.058 | -0.027 | -2.910*** | -2.819*** |
| | (0.133) | (0.054) | (0.619) | (0.413) |
| $Post_t \times Real\ Estate_i$ | 0.092 | -0.025 | -0.136 | -0.383 |
| | (0.103) | (0.062) | (0.721) | (0.615) |
| $NSOE_i \times Post_t \times Real\ Estate_i$ | 0.117 | -0.103 | 2.958** | 1.990* |
| | (0.176) | (0.119) | (0.922) | (0.793) |
| $\ln(\text{Total Assets})_{it}$ | | 1.220*** | | 3.759*** |
| | | (0.075) | | (0.188) |
| $(\text{Net Profits/Total Assets})_{it}$ | | 0.039 | | 2.126** |
| | | (0.073) | | (0.733) |
| $(\text{Fixed Assets/Total Assets})_{it}$ | | -0.638*** | | -2.452* |
| | | (0.158) | | (0.954) |
| $(\text{Total liability/Total Assets})_{it}$ | | 0.726*** | | -5.285*** |
| | | (0.168) | | (1.301) |
| $N$ | 19656 | 19391 | 19595 | 19342 |
| adj. $R^2$ | 0.729 | 0.795 | 0.655 | 0.687 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.2) on a balanced sample. Outcome variables are the log of total borrowing and retained earnings for firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.16: Spillovers of Macroprudential Regulation in Tax Havens

| | Issuance in Tax $\text{Haven}_{it}$ | | |
|---|---|---|---|
| | Full sample | Restricted | Restricted |
| | (1) | (2) | (3) |
| $\text{NSOE}_i \times \text{Post}_t$ | 0.006* | 0.016*** | 0.018*** |
| | (0.002) | (0.004) | (0.005) |
| ln(Total Assets)$_{it}$ | | | 0.013*** |
| | | | (0.002) |
| (Net Profits/Total Assets)$_{it}$ | | | 0.002 |
| | | | (0.003) |
| (Fixed Assets/Total Assets)$_{it}$ | | | 0.020+ |
| | | | (0.011) |
| (Total liability/Total Assets)$_{it}$ | | | 0.005 |
| | | | (0.003) |
| ln(Cash Inflow from Market) | | | -0.000+ |
| | | | (0.000) |
| L.ln(Cash Inflow from Market) | | | -0.000** |
| | | | (0.000) |
| $N$ | 29458 | 22127 | 17766 |
| adj. $R^2$ | 0.345 | 0.346 | 0.353 |
| Firm and Year FE | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.3) on a balanced sample. The outcome variable is a dummy variable of one if a firm $i$ issues at least one bond in tax haven countries in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.17: Spillovers from Firms in Real Estate Sector

| | (1) | (2) | (3) |
|---|---|---|---|
| | Issuance in Tax Haven$_{it}$ | | |
| NSOE$_i$×Post$_t$ | -0.005* | -0.008 | -0.008 |
| | (0.003) | (0.005) | (0.006) |
| Post$_t$×Real Estate$_i$ | 0.035*** | 0.057*** | 0.057 |
| | (0.007) | (0.011) | (0.037) |
| NSOE$_i$×Post$_t$×Real Estate$_i$ | 0.100*** | 0.179*** | 0.179*** |
| | (0.008) | (0.014) | (0.025) |
| ln(Total Assets)$_{it}$ | | 0.010*** | 0.010 |
| | | (0.002) | (0.006) |
| (Net Profits/Total Assets)$_{it}$ | | 0.002 | 0.002** |
| | | (0.003) | (0.001) |
| (Fixed Assets/Total Assets)$_{it}$ | | 0.017 | 0.017 |
| | | (0.011) | (0.022) |
| (Total liability/Total Assets)$_{it}$ | | 0.006+ | 0.006+ |
| | | (0.003) | (0.004) |
| ln(Cash Inflow from Market) | | -0.000 | -0.000** |
| | | (0.000) | (0.000) |
| L.ln(Cash Inflow from Market) | | -0.000** | -0.000+ |
| | | (0.000) | (0.000) |
| $N$ | 29458 | 17766 | 17766 |
| adj. $R^2$ | 0.363 | 0.384 | 0.384 |
| Firm and Year FE | Yes | Yes | Yes |
| SE clustered | Yes | No | Yes |

Note: This table reports the coefficients from the DID regression (2.4) on a balanced sample. The outcome variable is a dummy variable of one if a firm $i$ issues at least one bond in tax haven countries in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

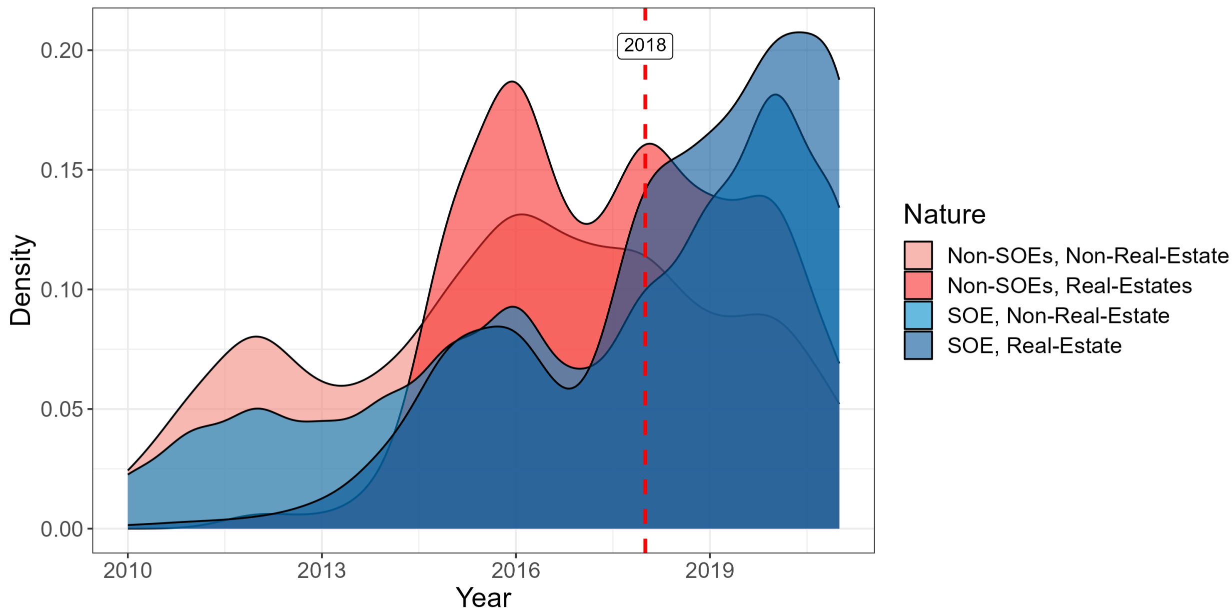
Table 2.18: Spillovers Linked to Private Ownership

| | (1) | (2) | (3) |
|---|---|---|---|
| | ln(Issuance Amount in Tax Haven)$_{it}$ | | |
| Private Ownership$_{it}$ | -0.004 | -0.024 | -0.024 |
| | (0.026) | (0.047) | (0.068) |
| Post$_t$×Private Ownership$_{it}$ | -0.049** | -0.088** | -0.088 |
| | (0.015) | (0.032) | (0.054) |
| Real Estate$_i$×Private Ownership$_{it}$ | -0.375** | -0.550** | -0.550* |
| | (0.115) | (0.201) | (0.255) |
| Post$_t$×Real Estate$_i$ | -0.100** | -0.155* | -0.155 |
| | (0.037) | (0.068) | (0.144) |
| Post$_t$×Real Estate$_i$×Private Ownership$_{it}$ | 0.675*** | 1.033*** | 1.033+ |
| | (0.063) | (0.112) | (0.530) |
| ln(Total Assets)$_{it}$ | | -0.003 | -0.003 |
| | | (0.012) | (0.016) |
| (Net Profits/Total Assets)$_{it}$ | | -0.002 | -0.002 |
| | | (0.014) | (0.003) |
| (Fixed Assets/Total Assets)$_{it}$ | | -0.098 | -0.098+ |
| | | (0.062) | (0.056) |
| (Total liability/Total Assets)$_{it}$ | | -0.013 | -0.013 |
| | | (0.026) | (0.011) |
| ln(Cash Inflow from Market) | | -0.001 | -0.001 |
| | | (0.001) | (0.001) |
| L.ln(Cash Inflow from Market) | | -0.001 | -0.001 |
| | | (0.001) | (0.000) |
| $N$ | 19777 | 10923 | 10923 |
| adj. $R^2$ | 0.241 | 0.236 | 0.236 |
| Firm and Year FE | Yes | Yes | Yes |
| SE clustered | Yes | No | Yes |

Note: This table reports the coefficients from the DID regression (2.7) on a balanced sample. The outcome variable is the log of the total amount of offshore bond funding for firm $i$ in year $t$. The variable "Private ownership" refers to the share of private ownership of firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

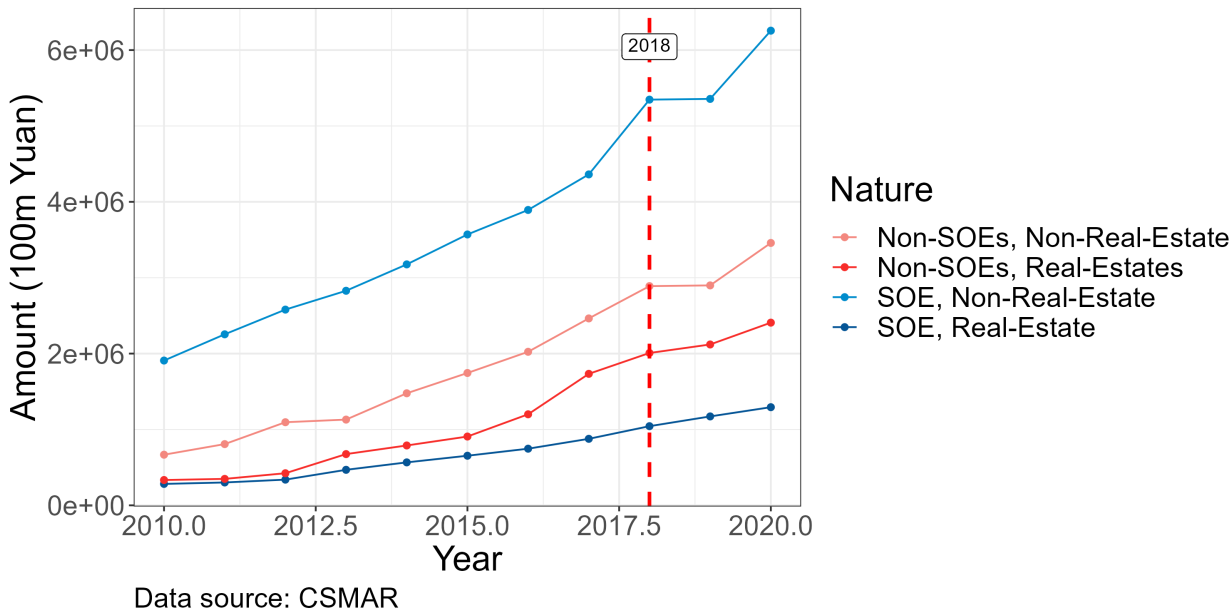
# B Onshore Trend of Bond Issuance and Bank Borrowing

Figure 2.6: Onshore Trend of Bond Issuance




Notes: This figure depicts the density of bond issuances in the domestic credit market since 2010 for four categories of firms: non-SOEs non-real-estate, non-SOEs real estate, SOEs non-real-estate, and SOEs real estate.

Figure 2.7: Onshore Trend of Bank Borrowing




Notes: This figure depicts the amount of total bank loans for four categories of firms: non-SOEs non-real-estate, non-SOEs real estate, SOEs non-real-estate, and SOEs real estate.

## C Spillovers Linked to Firm's Exposure to Domestic Bond Market

It is assumed that the degree of exposure to the domestic bond market affects the firm's offshore financing patterns. The hypothesis is that firms more dependent on the domestic bond market are more exposed to the macroprudential policy, making them more likely to issue bonds offshore. To test this hypothesis, I run the following regression:

$$\begin{aligned}\mathrm{T}_{it} = \beta_0 + \beta_1 \mathrm{NSOE}_i \times \mathrm{Post}_t \times \mathrm{Exposure}_{it} + \beta_2 \mathrm{NSOE}_i \times \mathrm{Exposure}_{it} \\ + \beta_3 \mathrm{Post}_t \times \mathrm{Exposure}_{it} + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it},\end{aligned} \tag{2.10}$$

where $\mathrm{Exposure}_{\mathrm{it}}$ is the ratio of the amount of domestic bonds to sales for firm $i$ in year $t$. The other variables are as defined in previous estimations. As shown in Table 2.19, the results are not significant. The insignificance also applies to two other measures of exposure: the ratio of the amount of domestic bonds to total liabilities and the log of the total amount of domestic bonds.

Table 2.19: Spillover Effects on Bond Issuance in Tax Havens

| | (1)<br>Tax Haven | (2)<br>Tax Haven | (3)<br>Tax Haven |
|---|---|---|---|
| $NSOE_i \times Post_t$ | 0.031** | 0.047** | 0.047 |
| | (0.011) | (0.015) | (0.034) |
| $Exposure_{it}$ | -0.000 | -0.000 | -0.000 |
| | (0.001) | (0.001) | (0.000) |
| $NSOE_i \times Exposure_{it}$ | 0.000 | 0.000 | 0.000 |
| | (0.001) | (0.001) | (0.000) |
| $Post_t \times Exposure_{it}$ | -0.002 | -0.003 | -0.003* |
| | (0.002) | (0.003) | (0.001) |
| $NSOE_i \times Post_t \times Exposure_{it}$ | 0.001 | 0.001 | 0.001 |
| | (0.002) | (0.003) | (0.001) |
| ln(Total Assets)$_{it}$ | | -0.021+ | -0.021 |
| | | (0.011) | (0.018) |
| (Net Profits/Total Assets)$_{it}$ | | 0.001 | 0.001 |
| | | (0.070) | (0.030) |
| (Fixed Assets/Total Assets)$_{it}$ | | -0.092+ | -0.092+ |
| | | (0.049) | (0.047) |
| (Total liability/Total Assets)$_{it}$ | | 0.018 | 0.018 |
| | | (0.043) | (0.026) |
| ln(Cash Inflow from Market) | | 0.000 | 0.000 |
| | | (0.001) | (0.001) |
| L.ln(Cash Inflow from Market) | | -0.000 | -0.000** |
| | | (0.000) | (0.000) |
| $N$ | 3176 | 2419 | 2419 |
| adj. $R^2$ | 0.346 | 0.281 | 0.281 |
| Firm and Year FE | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10). Firm and year fixed effects are included. Standard errors are clustered at the firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

## D Regression Results without Control Variables

The results of Tables 2.9 and 2.10 without control variables are reported in the two tables below. They are consistent with the main estimations.

Table 2.20: Effects on Fixed Assets, Sales, Cash and Wage

| | (1)<br>ln(Fixed Assets)$_{it}$ | (2)<br>ln(Sales)$_{it}$ | (3)<br>ln(Cash)$_{it}$ | (4)<br>ln(Wage)$_{it}$ |
|---|---|---|---|---|
| Tax Haven$_{it}$ | 0.320*** | 0.254*** | 0.318*** | -0.056 |
| | (0.056) | (0.042) | (0.023) | (0.113) |
| Tax Haven$_{it}$×Post$_t$ | 0.127* | 0.113* | 0.283*** | 0.309+ |
| | (0.052) | (0.055) | (0.064) | (0.177) |
| $N$ | 43858 | 42462 | 43840 | 29821 |
| adj. $R^2$ | 0.853 | 0.871 | 0.787 | 0.978 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10) without control variables. Outcome variables are the log of fixed assets, sales, cash, and wage by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2.21: Effects on Fixed Assets, Sales, Cash and Wage for Firms in Real Estate Sector

| | (1)<br>$\ln(\text{Fixed Assets})_{it}$ | (2)<br>$\ln(\text{Sales})_{it}$ | (3)<br>$\ln(\text{Cash})_{it}$ | (4)<br>$\ln(\text{Wage})_{it}$ |
|---|---|---|---|---|
| $\text{Tax Haven}_{it}$ | 0.423*** | 0.214*** | 0.287*** | -0.130 |
| | (0.118) | (0.059) | (0.029) | (0.114) |
| $\text{Tax Haven}_{it} \times \text{Post}_t$ | -0.035 | -0.018 | 0.054 | 0.179 |
| | (0.073) | (0.071) | (0.079) | (0.185) |
| $\text{Tax Haven}_{it} \times \text{Real Estate}_i$ | -0.222* | 0.103* | 0.107*** | 0.396*** |
| | (0.108) | (0.051) | (0.029) | (0.119) |
| $\text{Post}_t \times \text{Real Estate}_i$ | 0.104* | 0.079 | 0.254*** | 0.132+ |
| | (0.052) | (0.073) | (0.050) | (0.079) |
| $\text{Tax Haven}_{it} \times \text{Post}_t \times \text{Real Estate}_i$ | 0.246* | 0.165 | 0.188+ | 0.169 |
| | (0.098) | (0.100) | (0.098) | (0.332) |
| $N$ | 43858 | 42462 | 43840 | 29821 |
| adj. $R^2$ | 0.854 | 0.871 | 0.787 | 0.978 |
| Firm and Year FE | Yes | Yes | Yes | Yes |

Note: This table reports the coefficients from the DID regression (2.10) without control variables. Outcome variables are the log of fixed assets, sales, cash, and wage by firm $i$ in year $t$. Firm and year fixed effects are included. Standard errors are clustered at the city level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

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# Chapter 3

# Trade Credit and Financial Vulnerability: Evidence from Morocco

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**Abstract**

This paper explores the impact of a new law aimed at improving payment delays for firms engaged in procurement contracts with government entities in Morocco. Our analysis at the sector level indicates that the reform is effective at reducing the excessive amount of trade credit of firms in sectors that are more exposed to the new law. Our firm-level analysis shows that the effects are mostly driven by large firms, which experienced a significant reduction in trade credit post-reform compared to smaller firms.

## 3.1 Introduction

Trade credit is the most important source of external short-term finance in advanced economies (Petersen and Rajan, 1997). In most OECD countries, trade credit represents more than half of firms' short-term liabilities and more than one-third of their total debt (Altinoglu, 2021). It carries more significance for a developing economy such as Morocco, where the credit market is less developed. Firms, constrained by limited access to bank loans, heavily rely on trade credit for short-term financial needs. In Morocco, it accounts for nearly 100% of firms' current liabilities, underlying the importance of trade credit in a firm's day-to-day operations.

Delayed payment among firms is a widespread phenomenon in the Moroccan business world. Payments usually take between 120 and 150 days, compared to an average of 70 days in France (Allianz Trade, 2023). It adversely affects the cash-flow dynamics of enterprises, especially those small and medium-sized ones (SMEs). Heavy reliance on trade credit as short-term financing, together with delayed payments, severely tightens firms' cash flow constraints and limits firms' growth prospects. In fact, it is behind 40% of firm insolvencies (World Bank, 2019).

To address this issue, the government enacted Act 49-15 in 2016 to limit payment delays in procurement contracts by public entities, including local and central administration. The law mandates a maximum of 60 days between government invoice issuance and payment to involved enterprises. Exceeding the 60-day period results in penalties. Additionally, the government implemented mandatory electronic billing for invoices of public contracts above a certain threshold on a platform named *Gestion Intégrée de la Dépense* (GID) in 2019. The threshold was set at 5 million dirhams in 2019 and was reduced to 2 million dirhams in 2020. This digitizes the billing procedures, automatically detects any payment delays, and implements the penalty fee. These measures collectively target the reduction of payment delays by the government. Our paper analyzes the impact of this legislative reform, providing the first empirical evidence on the subject. Through our analysis, we aim at a better understanding of the importance of trade credit in the financial resilience of firms in a developing country and more generally its impact on firms' outcomes.

We use a confidential database from the General Treasury under the Moroccan Ministry of Economy and Finance (TGR, short for *Trésorerie Générale du Royaume*). It provides good-quality data on the invoices of government procurement contracts during 2010-2021. In addition, we use firm-level balance sheet data from Bank Al-Maghrib, the Central Bank of Morocco, and combine it with contract-level data from TGR. With these two databases, we are able to construct a sizeable sample to evaluate how the reform impacts the dynamics of trade credit of firms with public contracts.

We begin our empirical analysis by focusing on sector-level exposure to the new law. We rely on input-output tables to measure the degree of exposure of each sector to government spending. Specifically, we extract data on government expenditure across all sectors and divide it by the total production of each sector. This ratio serves as a measure of sector-level exposure to government spending, which is the primary focus of the reform on payment delays. We use a difference-in-differences (DID) approach to estimate changes in the trade credit of firms from sectors with different degrees of exposure to government spending. Our results indicate that firms in sectors that are more exposed to public administration are more likely to have lower amounts of trade credit in particular with their customers after the reform.

We continue our analysis by conducting firm-level estimation where the firms that are exposed to public contracts (treated firms) are compared to firms that are unexposed (control firms). We find that exposed firms reduce their trade credit post-reform relative to unexposed firms. We also see that these firms reduce their cash holding, increase their fixed assets and costs of employees, and have a reduced probability of exiting the market.

Our results further indicate that the treatment effects on trade credit are mostly driven by large firms, which experience a significant reduction in trade credit, in particular, compared to small firms. This indicates an unequal exposure to the new law across firm sizes. The impact is larger for bigger firms. One potential explanation is the implementation of mandatory billing on the GID platform for contracts exceeding 5 million dirhams associated with relatively large firms. The GID platform streamlines the billing process, identifying any payment delays once the invoice deadline has passed. This mechanism incentivizes the government to expedite payments and mitigate delays that firms encounter in receiving their dues.

To explore the hypothesis that large firms are more exposed to the new law, we exploit the threshold of 5 million dirhams and employ a regression discontinuity (RD) design. By comparing firms with TGR contracts just above and below the threshold, we can evaluate the treatment effect of the GID platform on large firms, which are capable of securing contracts around 5 million dirhams with the government. Our estimation shows that there are no significant changes in trade credit associated with those firms affected by the mandatory GID filing of invoices.

This paper is related to the large group of empirical literature on a firm's trade credit and its role in redistributing liquidity, such as Meltzer (1960); Schwartz (1974); Petersen and Rajan (1997); Fisman and Love (2003); Mateut, Bougheas, and Mizen (2006); Cuñat (2007); Love, Preve, and Sarria-Allende (2007); Shenoy and Williams (2017)[1]. We contribute to this literature by analyzing additional balance sheet effects. Similar to Benhima et al. (2023) regarding the impact of credit guarantee on Moroccan firms, we observe that exposed firms, following the reform, reduce their

[1]Please see a full discussion of these literature in the Appendix.

cash holding and reallocate it to other productive assets, such as capital and labor. This shift results in an expansion of sales and enhances the likelihood of survival for these firms. Our evidence on firm survival is aligned with the findings from Barrot (2016). Our paper is the first one to perform a causal analysis of the impact of trade credit on firm survival in an emerging economy.

In our paper, we discover the size-dependent treatment effects of the new legislation. This is related to the strand of literature that emphasizes the heterogeneity of firm-level characteristics when analyzing trade credit, such as Garcia-Appendini and Montoriol-Garriga (2013) (cash level), Gonçalves, Schiozer, and Sheng (2018) (market power), Coricelli and Frigerio (2019) (firm size), Carbó-Valverde, Rodríguez-Fernández, and F. Udell (2016) (credit constraint).

The article is organized as follows. Section 2 lays out the institutional background. Section 3 describes the data. Section 4 provides the empirical analysis and Section 6 concludes.

## 3.2 Institutional Background

To tackle the payment delay issue, the Moroccan government passed a new law Act 49-15 in 2016 to improve the regulatory framework regarding payments between firms. Proposed measures are comprehensive, with a focus on reducing excessive delays in public contract payments. Specifically, the law requires that the maximum period between the issuance of invoices and the payment is 60 days for public procurement contracts from the state, local authorities, and other public administration entities. Any overrun of the 60-day period triggers penalties automatically.

To support the implementation of the new law, the government has introduced additional measures of digitizing the invoicing process for procurement transactions through GID. This system was set up in 2009 to facilitate government budget management. In 2019, the government made digital billing mandatory for public procurement contracts exceeding 5 million dirhams. In 2020, this threshold was reduced to 2 million dirhams. This measure intends to accelerate the invoicing process of public contracts.

In Morocco, there is a high engagement rate of firms in public contracts, as shown in Table A1 in the Appendix. There are no significant differences in the number of public contracts before and after the reform. The government's public orders remain consistent and do not have any changes due to the reform. Nearly 7% of all registered firms secured at least one public contract in 2018. The significance of this engagement rate is even more obvious when considering their contribution to the total value added. These firms collectively contribute to approximately 20% of the total value added across all sectors. In this regard, the reform aims at a substantial proportion of firms. It impacts not only those directly affected by delayed payments by the government but could also create a trickle-down effect on other firms' liquidity along the supply chain. The impact of the reform extends beyond the public establishments, exerting a significant influence on the entire Moroccan economy.

Preliminary evidence on the effects of the new law suggests that the reform has taken some effects. After the reform, the overall payment time is reported to be reduced. For the state, the number of payment days is reduced from 146 days in 2016 to 58 days in 2017 and further to 39 days in 2018 (Observatoire des Délais de Paiement, 2021). Similar reductions in payment time can be observed for contracts by local authorities. Additionally, there is an increase in penalty payments, which was 12.5 million dirhams in 2016 and jumped to 18.5 million dirhams in 2017, suggesting that the penalty measures take effect as well. Our empirical analysis is the first one to rigorously evaluate the impact of this reform.

## 3.3 Data

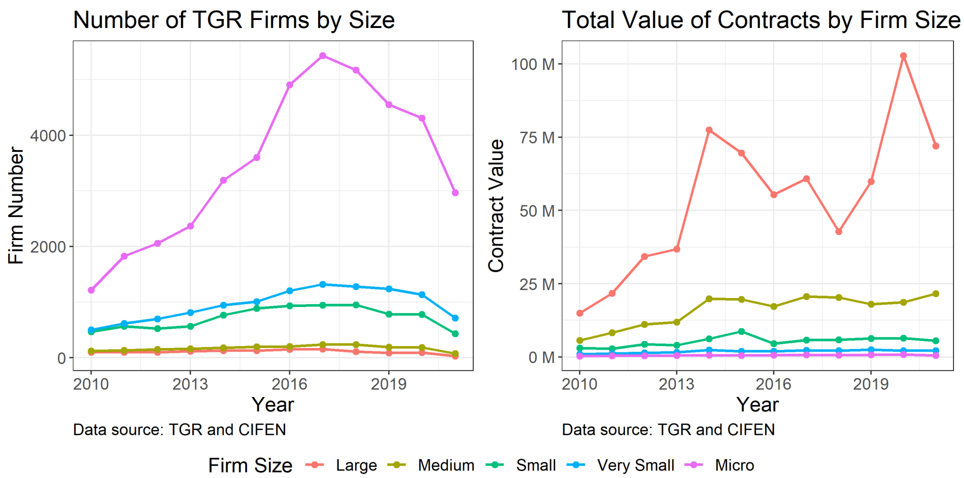
Our main empirical analysis combines contract-level data from TGR with firm-level balance sheet data from Bank Al-Maghrib (Central Bank of Morocco).

### A Contract-level data

Data of invoices for procurement contracts related to public establishments are sourced from TGR. For 2010-2021, the dataset covers 65116 public contracts conducted by more than 23812 firms. The data present detailed information related to public invoices, such as firm identifier, date of signing the contract, type of expense, and size of contract. Table A2 shows that contracts are concentrated in construction (34%), commerce (30%), and services industries (18%).

It is interesting to note that 64% of the TGR contracts are conducted by micro firms whereas large firms only account for 2% of the total number, as shown in Table A3. In contrast, the average size of contracts executed by micro firms is only 0.6 million dirhams, compared to an average size of 54 million dirhams by large firms. This unbalanced distribution is consistent over the years as demonstrated in Figure 3.1, where the proportion of micro firms dominates in terms of the number of contracts secured during 2010-2020. However, the monetary value of these contracts is minimal in contrast with those executed by large firms. This reveals significant heterogeneity across firm size and contract size.

Figure 3.1: Evolving Trend of TGR Firm Size and Contract Value




Notes: The left figure depicts the number of TGR firms by size over the years. The right figure shows the total value of contracts by firm size in the TGR database.

## B Firm-level data

The firm-level balance sheet data is sourced from Bank Al-Maghrib, which obtains data from the trade registry office *Office Marocain de la Propriete Industrielle et Commerciale* (OMPIC). New firms register with OMPIC and existing firms report their yearly balance sheets to OMPIC. We consider it as a comprehensive coverage of the entire universe of Moroccan firms.

Two variables are relevant to our main analysis of trade credit: accounts receivable and accounts payable. They gauge the amount of credit firms extend to or obtain from their customers or suppliers. In our main empirical analysis, we use the ratio of accounts receivable and accounts payable to sales.

Summary statistics of dependent and control variables are provided in Table 3.1. The average of accounts payable to sales is 64%, implying that the value of inputs purchased on credit from their suppliers is worth more than half of the sales of an average firm. The mean of accounts receivable to sales is 46%, showing that close to half of the sales are sold on credit to their customers. These statistics highlight the importance of trade credit in a firm's day-to-day operations.

## C Sample construction

We rely on the firm's national ID, a unique firm identifier in Morocco, to pair TGR contract-level data with OMPIC firm balance sheet data. We can identify 9330 TGR firms out of the total of 16037 in the OMPIC database. It represents a pairing rate of 58%, sufficient to construct a firm sample

Table 3.1: Summary Statistics

| Statistic | N | Mean | St. Dev. | Min | Max |
|---|---|---|---|---|---|
| Accounts Receivable/Sales | 580,085 | 0.46 | 0.42 | 0.00 | 2.00 |
| Accounts Payable/Sales | 580,085 | 0.64 | 0.50 | 0.00 | 2.00 |
| Cash/Total Assets | 576,813 | 0.19 | 0.26 | 0.00 | 1.00 |
| Fixed Assets/Total Assets | 579,279 | 0.15 | 0.22 | 0.00 | 1.00 |
| Current Liabilities/Total Assets | 498,710 | 0.55 | 0.29 | 0.00 | 1.00 |
| Costs of Employees/Total Assets | 569,029 | 0.19 | 0.28 | 0.00 | 2.00 |
| Sales/Total Assets | 580,085 | 1.64 | 1.80 | 0.0000 | 10.43 |

Notes: This table reports the summary statistics of key variables used in regressions.

for our empirical analysis. While there is a substantial number of firms in the TGR database that remain unidentifiable in the OMPIC database, it suggests the presence of false negatives within the final sample. Table A4 in the Appendix summarizes some key statistics of TGR and non-TGR firms. They are similar in terms of firm-level characteristics.

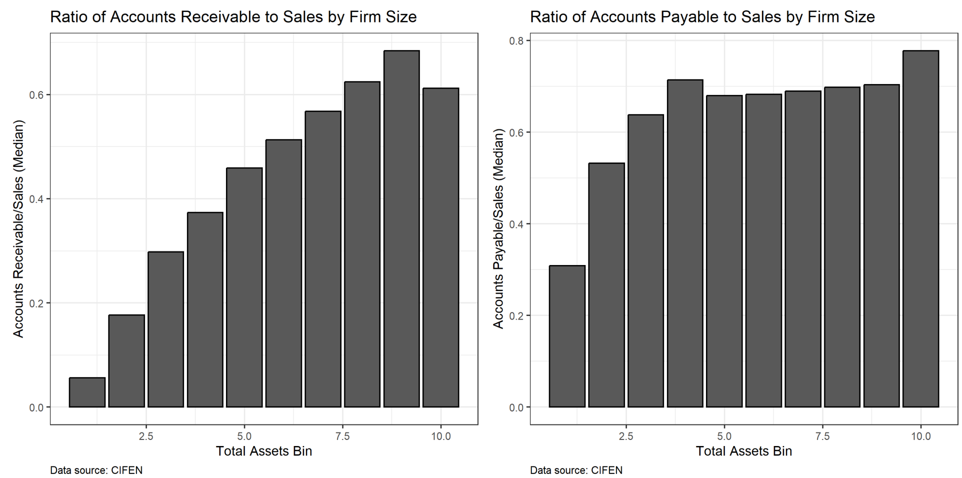
Large firms have high ratios of accounts receivable and accounts payable to sales. Figure 3.2 plots these two ratios across different sizes of firms proxied by different bins of total assets. The positive relation with the firm size is particularly pronounced for the ratio of accounts receivable to sales. This relation suggests that large firms are able to take on large amounts of short-term lending and borrowing with both suppliers and customers. In contrast, small firms face a different situation, characterized by a higher ratio of debt owed to suppliers and a lower ratio of credits extended to customers.

## 3.4 Empirical Analysis

### A Sector-level Heterogeneity

We begin our analysis by examining the effects of the new law, focusing on the sectoral exposure to the reform using the input-output table for intermediate goods. In Table A5, we list, for each sector, the percentage of goods and services procured by the public administrations, relative to each sector's production. This percentage serves as a measure of the exposure of industries to government spending. We use the average pre-reform exposure during 2007-2016 to mitigate annual fluctuations. The commerce and trade industry is the most exposed to government spending, with nearly a quarter of its output being consumed as intermediate goods by the public sector.

Figure 3.2: Distributions of Accounts Receivable and Payable to Sales Ratios by Firm Size




Notes: This plot provides the distribution of ratios of accounts receivable and payable to sales over 10 bins of total assets.

This is followed by hotels, restaurants, and manufacturing industries. Note that, because the sectoral composition of the total public final good consumption and of the total public investment is not known, this exposure abstracts from final consumption and investment goods.

We run the following regression to analyze the sectoral effects:

$$Y_{ijt} = \beta_0 + \beta_1 Exposure_j \times Post_t + \delta_j + \delta_t + \epsilon_{ijt} \tag{3.1}$$

The firms are indexed by $i$, and sectors and years are indexed by $j$ and $t$. This regression evaluates the impact on an average firm within one sector. The variable $Y_{ijt}$ denotes three outcome variables: the logged ratio of accounts receivable to sales, the logged ratio of accounts payable to sales, and the difference between the two logged ratios. These are three outcome variables measuring firm-level performance in trade credit, capturing the extent of credit they receive or extend to their suppliers or customers.

The independent variable $Exposure_j$ indicates the percentage of goods and services purchased by the public administration relative to the sector's total production. $Post_t$ is the time dummy variable indicating the post-reform period. We take fixed effects on sector and year levels, and standard errors are clustered at sector and year levels.

The results on accounts receivable and payable are summarized in Tables 3.2. A 1 percentage point increase in exposure for one sector is associated with a decrease of about 0.3% in ratios of accounts receivable to sales for firms in this sector. We do not observe significant effects for accounts payable. In Column 3, the coefficient associated with net trade credit is significantly negative, implying that the net credit that firms extend to their customers is in reduction. These results are statistically significant and economically important, suggesting that the reform has had some effects in reducing the amount of trade credit that firms accumulate from the public administration sector.

Table 3.2: Sector-level Effects

| | (1) $\ln(\frac{AR}{Sales})_{it}$ | (2) $\ln(\frac{AP}{Sales})_{it}$ | (3) $\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $\text{Exposure}_j \times \text{Post}_t$ | -0.003*** | 0.000 | -0.004*** |
| | (0.001) | (0.001) | (0.001) |
| ln(Total Assets)$_{it-1}$ | 0.145*** | 0.039*** | 0.106*** |
| | (0.007) | (0.006) | (0.007) |
| ln(Cash/Total Assets)$_{it-1}$ | -0.014*** | -0.040*** | 0.026*** |
| | (0.002) | (0.002) | (0.002) |
| ln(Fixed Assets/Total Assets)$_{it-1}$ | -0.012*** | 0.019*** | -0.031*** |
| | (0.003) | (0.002) | (0.003) |
| ln(Current Liabilities/Total Assets)$_{it-1}$ | -0.043*** | 0.032*** | -0.075*** |
| | (0.004) | (0.006) | (0.007) |
| ln(Sales/Total Assets)$_{it-1}$ | -0.015* | -0.053*** | 0.038*** |
| | (0.007) | (0.007) | (0.008) |
| $N$ | 229380 | 229380 | 229380 |
| adj. $R^2$ | 0.692 | 0.603 | 0.573 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.1). Outcome variables are the log of the ratios of accounts receivable to sales, the log of the ratio of accounts payables to sales, and the difference between the two logged ratios by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

## B Firm-level Heterogeneity

To study the effects of the reform on payment delays on a firm level, we adopt a standard DID approach.

### Treatment Dummy

Instead of using the actual treatment variable, which identifies firms with public contracts post-reform, we opt for a proxy treatment variable that takes the value of one if a firm has a contract before the reform. This is to avoid potential endogeneity issues, where financially constrained

Table 3.3: Prediction of Post-Reform Treatment based on Pre-Reform Situation

| | (1) | (2) |
|---|---|---|
| | $\text{TGR}_{post-2016,i}$ | $\text{TGR}_{post-2016,i}$ |
| $\text{TGR}_{pre-2016,i}$ | 0.467*** | 0.314*** |
| | (0.005) | (0.011) |
| 2nd Tercile | | 0.014*** |
| | | (0.002) |
| 3rd Tercile | | 0.009*** |
| | | (0.002) |
| 2nd Tercile×$\text{TGR}_{pre-2016,i}$ | | 0.164*** |
| | | (0.014) |
| 3rd Tercile×$\text{TGR}_{pre-2016,i}$ | | 0.211*** |
| | | (0.013) |
| $N$ | 126979 | 126979 |
| adj. $R^2$ | 0.193 | 0.200 |
| FE (Sector) | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.2). $TGR_{pre-2016,i}$ is a dummy variable of one if a firm $i$ had a public contract in or before 2016 and $TGR_{post-2016,i}$ is a dummy variable of one if a firm $i$ had a contract after 2016. Sector fixed effects are included. Robust standard errors are reported. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

firms, post-reform, may choose to engage in public contracts thanks to the reform whereas they could not do so pre-reform due to issues related to payment delays. By using a proxy treatment variable, we avoid this self-selection issue.

To test the validity of our treatment dummy, we conduct the following specification:

$$TGR_{post-2016,i} = \beta_0 + \beta_1 TGR_{pre-2016,i} + \delta_j + \epsilon_{ij}, \tag{3.2}$$

where $TGR_{pre-2016,i}$ is a dummy variable of one if a firm had a public contract in or before 2016 and $TGR_{post-2016,i}$ is a dummy variable of one if a firm had a contract after 2016. As we can see from Table 3.3 Column 1, a firm that had a public contract pre-reform is very likely to have a public contract post-reform. Thus, $TGR_{pre-2016,i}$ is a statistically robust predictor for $TGR_{post-2016,i}$ and thus can be used as our proxy treatment variable to avoid endogeneity issues.

### Firm-level DID Estimation

We examine the effects on accounts receivable and payable for TGR-related firms post-reform. We are interested in the following regression model:

$$Y_{it} = \beta_0 + \beta_1 TGR_{pre-2016,i} \times Post_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \quad (3.3)$$

where $Y_{ijt}$ denotes two set of outcome variables. The first set of variables is the same as in Regression 3.1, estimating the firm-level changes in trade credit. The second set of outcome variables estimates changes in other balance sheet items, which includes the log of cash, fixed assets, wage, and a dummy variable of one if the firm exits the market in the following year.[2]

Control variables included are the log of total assets, the ratio of cash to total assets, fixed assets to total assets, current liabilities to total assets, and sales to total assets. They are essential in accounting for variations in the assets and liabilities of a firm, given that accounts receivable and payable are two important elements on both sides of the balance sheet.

The regression results on the first set of outcome variables are summarized in Table 3.4. In the first two columns, we observe negative and statistically significant effects for accounts receivable and payable. This indicates that companies involved in public contracts demonstrate reductions in both variables at a similar scale. In the third column where we examine the difference between these two variables, the effects are near zero. This suggests that TGR-related firms borrow less from their suppliers while simultaneously extending less credit to their customers after the reform. Consequently, the net effect is close to zero.

Table 3.5 provides the estimation results related to a treated firm's changes in cash, fixed assets, costs of employees, sales, and probability of exiting the market. As Column 1 indicates, a TGR firm reduces cash holding post-reform. One of the reasons why a firm holds cash is to deal with unexpected payment delays from its customers. The improved situation has encouraged firms to reduce cash holding and allocate it to other productive assets. We observe increases in fixed assets and labor costs in Columns 2 and 3. This enhances the growth aspects for treated firms through sales expansion (Column 4) and a lower likelihood of exiting the market (Column 5).

[2] The variable "Exit" is a dummy of one in year $t$ if the firm disappears from OMPIC databases in year $t + 1$ and onward. We assume that if a firm stops reporting its balance sheet, it exits the market.

Table 3.4: Treatment Effects on Accounts Receivable and Payable

| | (1)<br>$\ln(\frac{AR}{Sales})_{it}$ | (2)<br>$\ln(\frac{AP}{Sales})_{it}$ | (3)<br>$\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $\text{TGR}_{pre-2016,i} \times \text{Post}_{t}$ | -0.029* | -0.029* | 0.000 |
| | (0.013) | (0.012) | (0.015) |
| $\ln(\text{Total Assets})_{\text{it-1}}$ | 0.141*** | 0.035*** | 0.106*** |
| | (0.007) | (0.006) | (0.008) |
| $\ln(\text{Cash/Total Assets})_{\text{it-1}}$ | -0.014*** | -0.040*** | 0.026*** |
| | (0.002) | (0.002) | (0.002) |
| $\ln(\text{Fixed Assets/Total Assets})_{\text{it-1}}$ | -0.011*** | 0.019*** | -0.030*** |
| | (0.003) | (0.003) | (0.003) |
| $\ln(\text{Current Liabilities/Total Assets})_{\text{it-1}}$ | -0.043*** | 0.030*** | -0.072*** |
| | (0.005) | (0.006) | (0.007) |
| $\ln(\text{Sales/Total Assets})_{\text{it-1}}$ | -0.013+ | -0.048*** | 0.035*** |
| | (0.007) | (0.007) | (0.008) |
| $N$ | 219040 | 219040 | 219040 |
| adj. $R^2$ | 0.695 | 0.607 | 0.575 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.3). Outcome variables are the log of the ratios of accounts receivable to sales, the log of the ratio of accounts payables to sales, and the difference between the two logged ratios by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 3.5: Treatment Effects on Sales, Cash, Fixed Assets, Wage and Exit

| | (1) $\ln(\text{Cash})_{\text{it}}$ | (2) $\ln(\text{Fixed Assets})_{\text{it}}$ | (3) $\ln(\text{Wage})_{\text{it}}$ | (4) $\ln(\text{Sales})_{\text{it}}$ | (5) $\text{Exit}_{\text{it}}$ |
|---|---|---|---|---|---|
| $\text{TGR}_{pre-2016,i} \times \text{Post}_{\text{t}}$ | -0.037+ | 0.052** | 0.029*** | 0.029*** | -0.021*** |
| | (0.019) | (0.018) | (0.008) | (0.008) | (0.003) |
| $\ln(\text{Total Assets})_{\text{it-1}}$ | -0.184*** | 0.424*** | 0.356*** | 0.345*** | 0.015*** |
| | (0.008) | (0.009) | (0.005) | (0.005) | (0.001) |
| $\ln(\text{Cash/Total Assets})_{\text{it-1}}$ | 0.050*** | 0.018*** | 0.002+ | -0.000 | -0.003*** |
| | (0.003) | (0.002) | (0.001) | (0.001) | (0.001) |
| $\ln(\text{Fixed Assets/Total Assets})_{\text{it-1}}$ | -0.031*** | 0.282*** | 0.022*** | -0.003* | 0.003*** |
| | (0.003) | (0.005) | (0.002) | (0.002) | (0.001) |
| $\ln(\text{Current Liabilities/Total Assets})_{\text{it-1}}$ | -0.037*** | -0.023*** | -0.006* | 0.026*** | 0.002** |
| | (0.004) | (0.004) | (0.002) | (0.002) | (0.001) |
| $\ln(\text{Sales/Total Assets})_{\text{it-1}}$ | -0.008 | 0.043*** | 0.200*** | 0.194*** | 0.007*** |
| | (0.007) | (0.008) | (0.005) | (0.005) | (0.001) |
| $N$ | 218510 | 181365 | 193236 | 219040 | 219040 |
| adj. $R^2$ | 0.644 | 0.895 | 0.934 | 0.945 | 0.655 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.3). Outcome variables are the log of cash, fixed assets, wage, sales, and probability of exiting the market by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

### Treatment Effects by Firm Size

We move on to analyze the heterogeneous effects for different-sized firms. We employ a triple interaction in the following regression:

$$\begin{aligned} Y_{it} = \beta_0 + \beta_1 TGR_{pre-2016,i} \times Post_t \times Size_i + \beta_2 TGR_{pre-2016,i} \times Post_t + \beta_3 Post_t \times Size_i \\ + \beta_4 TGR_{pre-2016,i} \times Size_i + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}. \quad (3.4) \end{aligned}$$

The variable $Size_i$ categorizes firms into three sizes: small (1st tercile), medium (2nd tercile), and large (3rd tercile). The control variables are the same as those specified in Regression (3.3). This regression model intends to measure heterogeneous treatment effects across different firm sizes, corresponding to our previous exploratory analysis that looks at variations in trade credit distribution based on firm size.

Table 3.6 summarizes the regression results. In Column 1, we observe that the treatment effects of accounts receivable for large firms (3rd tercile) are particularly pronounced compared to small firms (1st tercile). On average, large firms experience a significant reduction of 18.5% in the ratio of accounts receivable to sales post-reform, a statistically significant result at 1% level.

In the second column, we do not observe significant effects for large and medium firms related to accounts payable. The results on net trade credit in Column 3 are significantly negative for medium and large firms, suggesting that both experience reductions in net trade credit compared to small firms. The heterogeneous effects by firm size indicate that the treatment effects we observe in the previous estimation from Table 3.4 are mostly driven by large and medium firms.

This result is robust to different size measures. As Table A8 shows, when we use the log of total assets from year $t-1$ as a measure of size, the estimation results are consistent with our main analysis. As a second robustness test, we divide the variable of sales into 10 bins. Figure 3.9 in the Appendix corroborates our results and shows that it is the large firms that drive the treatment effects. Our results remain consistent when we also drop COVID years (2020 and 2021) and when we only use the ratios of AR/Sales and AP/Sales, instead of the log of these ratios.

Table 3.6: Heterogeneous Effects on Accounts Receivable and Payable

| | (1) $\ln(\frac{AR}{Sales})_{it}$ | (2) $\ln(\frac{AP}{Sales})_{it}$ | (3) $\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | 0.123* | -0.039 | 0.161* |
| | (0.055) | (0.048) | (0.064) |
| 2nd Tercile | -0.368*** | -0.313*** | -0.055* |
| | (0.020) | (0.017) | (0.023) |
| 3rd Tercile | -0.683*** | -0.641*** | -0.042 |
| | (0.023) | (0.021) | (0.026) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}$ | 0.065 | -0.028 | 0.093+ |
| | (0.052) | (0.044) | (0.057) |
| 3rd Tercile $\times\text{TGR}_{pre-2016,i}$ | 0.153** | -0.007 | 0.160* |
| | (0.059) | (0.051) | (0.063) |
| 2nd Tercile$\times\text{Post}_t$ | -0.044* | -0.094*** | 0.050* |
| | (0.021) | (0.018) | (0.025) |
| 3rd Tercile $\times\text{Post}_t$ | -0.147*** | -0.120*** | -0.027 |
| | (0.020) | (0.018) | (0.024) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | -0.088 | 0.035 | -0.123+ |
| | (0.058) | (0.052) | (0.067) |
| 3rd Tercile $\times\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | -0.185** | 0.019 | -0.203** |
| | (0.056) | (0.049) | (0.065) |
| ln(Total Assets)$_{it-1}$ | 0.192*** | 0.085*** | 0.107*** |
| | (0.007) | (0.006) | (0.008) |
| ln(Cash/Total Assets)$_{it-1}$ | -0.014*** | -0.040*** | 0.027*** |
| | (0.002) | (0.002) | (0.002) |
| ln(Fixed Assets/Total Assets)$_{it-1}$ | -0.011*** | 0.019*** | -0.030*** |
| | (0.003) | (0.002) | (0.003) |
| ln(Current Liabilities/Total Assets)$_{it-1}$ | -0.039*** | 0.032*** | -0.072*** |
| | (0.005) | (0.006) | (0.007) |
| ln(Sales/Total Assets)$_{it-1}$ | 0.015* | -0.022*** | 0.037*** |
| | (0.007) | (0.007) | (0.008) |
| $N$ | 219040 | 219040 | 219040 |
| adj. $R^2$ | 0.701 | 0.615 | 0.576 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.4). Outcome variables are the log of the ratios of accounts receivable to sales, the log of the ratio of accounts payables to sales, and the difference between the two logged ratios by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

### Leads and Lags

We estimate a specification with leads and lags to measure pre-reform and post-reform trends:

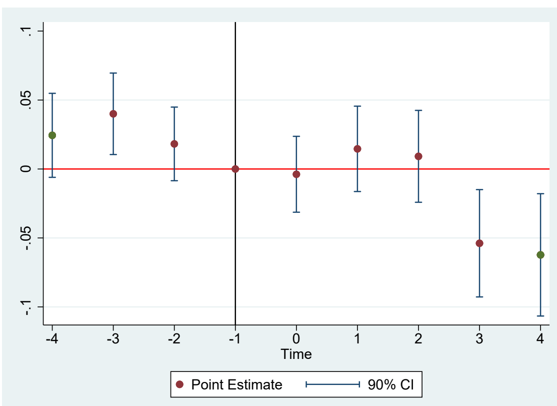
$$\text{Y}_{it} = \beta_0 + \sum_{k=2011}^{2020} \beta_k TGR_{pre-2016,i} \times \text{Year}_k + +\delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \tag{3.5}$$

where $\text{Year}_k$ is a year indicator variable. The reference year is 2015, the year before the reform. The specification includes three lags and five leads, corresponding to the studied period of 2011-2020. Figures 3.3 and 3.4 plots the $\beta_k$ coefficients and their 95% confidence intervals for all TGR-related firms as well as for different-sized firms.

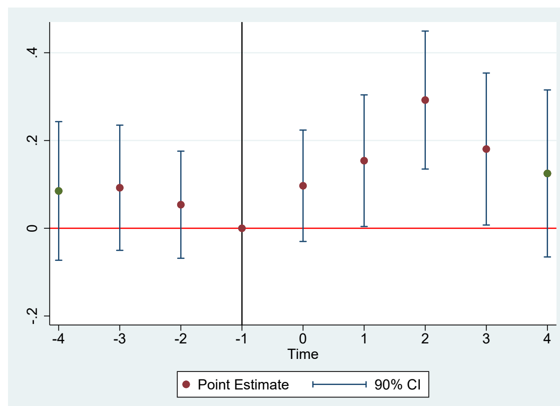
As Figure 3.3 shows, there is no obvious pre-reform trend for TGR-related firms. In addition, there is a slight reduction in AR/sale three years after the reform, aligning with our estimations from Table 3.4. It also corroborates our results in Table 3.6 where large firms demonstrate a reduction in trade credit right after the reform. However, the post-reform trend for medium and small firms is not obvious. Coefficient plots on accounts payable from Figure 3.4 also confirm our results from the previous estimation where we observe significant reductions in the ratio of accounts payable to sales on average but no heterogenous size effects.

Figure 3.3: Leads and lags of Treatment Effects on ln(AR/Sales)

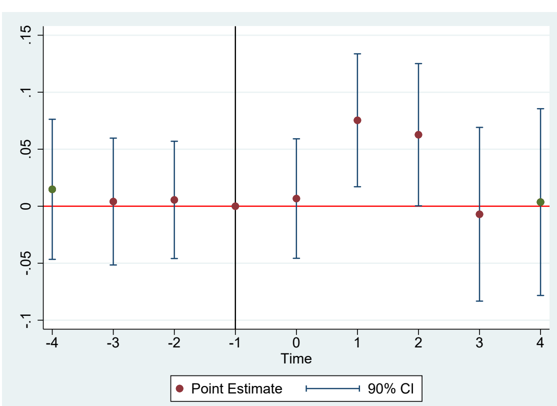
(a) All Firms



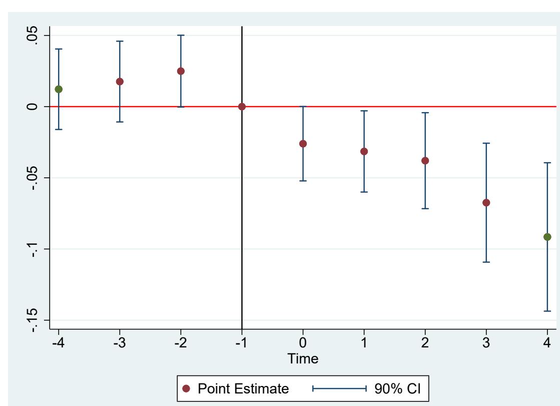

(b) 1st Tercile Firms (Small)




(c) 2nd Tercile Firms (Medium)



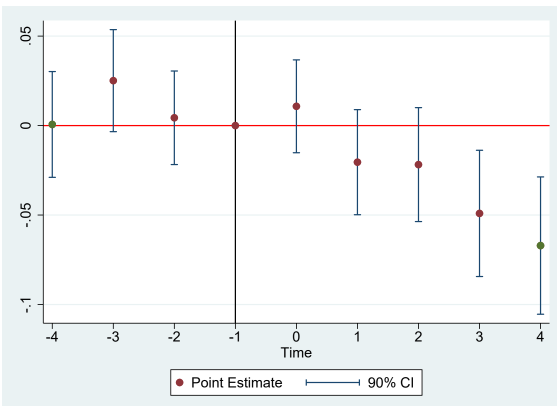

(d) 3rd Tercile Firms (Large)



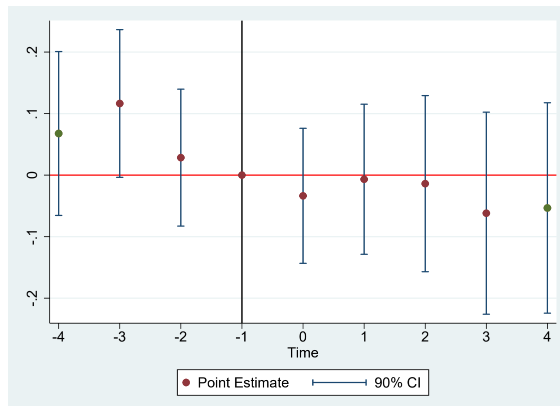

Notes: This figure plots the coefficients of Regression (3.5). The reference year is 2015, the year before the implementation of the new law.

Figure 3.4: Leads and Lags of Treatment Effects on ln(AP/Sales)

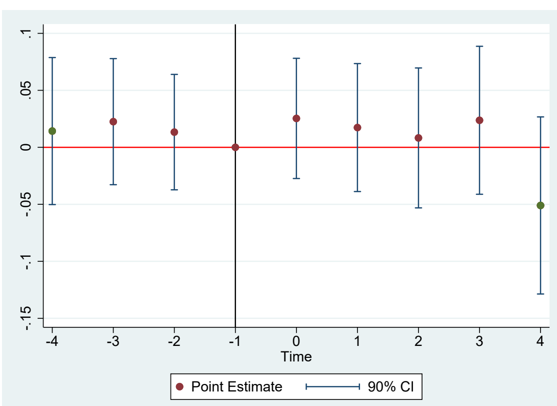
(a) All Firms



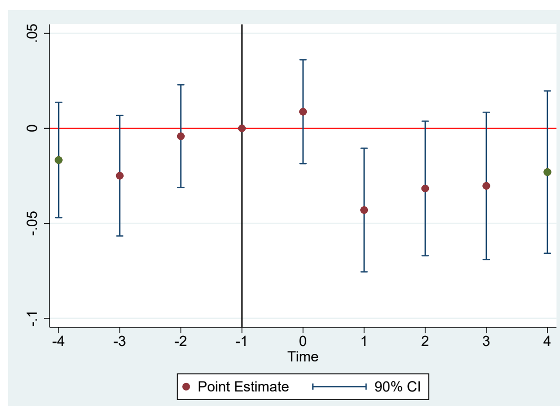

(b) 1st Tercile Firms (Small)




(c) 2nd Tercile Firms (Medium)




(d) 3rd Tercile Firms (Large)




Notes: This figure plots the coefficients of Regression (3.5). The reference year is 2015, the year before the implementation of the new law.

### Why Do Only Large Firms Benefit from the Reform?

Our firm-level results reveal that the treatment effects of the new law are more pronounced for large firms. One potential reason could be the implementation of compulsory invoicing on the GID platform for contracts over 5 million dirhams involving large firms. This mitigates payment delays by digitizing the billing procedures, automatically detecting any delays by the government, and implementing the penalty fee if needed.

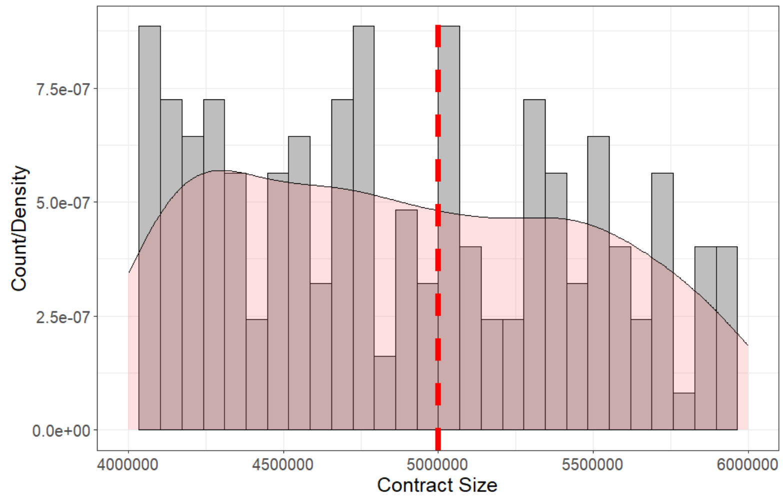
To test the hypothesis that big firms are more exposed to the reform, we employ an RD approach by leveraging the threshold for mandatory billing of procurement invoices exceeding 5 million dirhams on GID. Specifically, public contracts above 5 million dirhams were required to be electronically managed on GID in 2019 and this threshold was further reduced to 2 million dirhams in 2020. We assume that firms capable of securing public contracts larger than 5 million dirhams are relatively big firms. Using the RD regression, we evaluate the impact of the GID requirement on the trade credit dynamics of these large firms.

RD design allows us to evaluate the causal treatment effect by comparing trade credit changes associated with firms of TGR contracts just above and below the threshold. We anticipate that firms above the cutoff are more likely to reduce trade credit in comparison to firms below the cutoff thanks to the mandatory electronic billing process facilitated by the GID platform. If a specific impact of GID is identified, then it would mean that the fact that access to GID is dependent on the contract size may indeed have led to a differential effect of the reform across firm sizes.

Before discussing the RD results, we conduct two standard falsification tests. First, we examine if the density of TGR contracts is not discontinuous at the cutoff of 5 million dirhams. Specifically, we check if the number of contracts above the threshold is not significantly different from the number below. This is to ensure that the public contractor and the firm do not manipulate the contract size to avoid using GID, or on the contrary to be eligible for GID. As Figure 3.5 shows, there is no bunching below or above the threshold. We observe a peak at 5 million dirhams, which is a round-number effect similar to the peak observed at 4 million dirhams. Despite this peak, the density is smooth around the cutoff. The numbers of observations above and below the threshold are similar.

Second, we examine the control variables to detect any discontinuities at the threshold. This is to check if firms with similar TGR contract sizes at the cutoff are similar in pre-treatment observable characteristics. Figure 3.8 in the Appendix provides RD plots for each control variable, none of which demonstrates obvious discontinuity at the cutoff. We further test this statistically and conduct RD estimation for each control variable. As Table A6 shows, none of them shows any statistically significant jumps at the cutoff.

Figure 3.5: Histogram of TGR Contracts



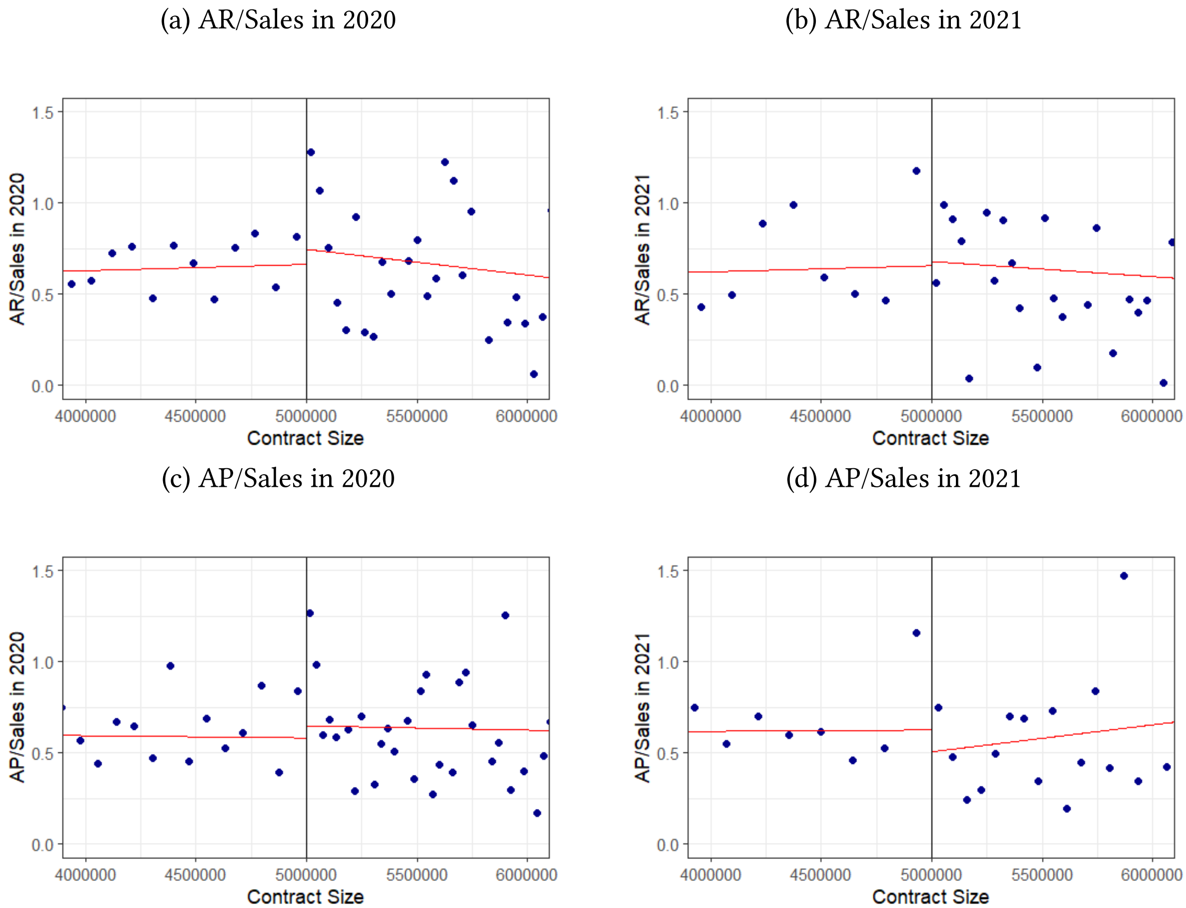

Notes: This figure presents the histogram and density plot of the variable of TGR contract size. The absolute numbers of observations below and above the threshold are 101 and 180.

We begin our RD analysis with a graphical illustration of the RD design in Figure 3.6 for the cutoff of 5 million dirhams in 2019. The running variable is the contract size and the four outcome variables are the log of AR/Sales and the log of AP/Sales from the years 2020 and 2021. As the plots demonstrate, we do not detect obvious discontinuities at the threshold. To employ the RD approach, we run the following RD regression

$$Y_{it} = \beta_0 + \beta_1 GID_{i,2019} + \beta_2 TGR_{i,2019} + \beta_3 TGR_{i,2019} \times GID_{i,2019} + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it},$$

where $Y_{it}$ denotes the log of AR/Sales and the log of AP/Sales in 2020 and 2021. $GID_{i,2019}$ is a treatment dummy variable that is equal to one if the TGR contract size in 2019 is above 5 million dirhams and the firm is required to use the GID platform. $TGR_{i,2019}$ refers to the contract size. Coefficient $\beta_1$ is the local linear RD point estimation of treatment effects. Bandwidth is selected using the minimized mean squared error approach as specified in Cattaneo, Idrobo, and Titiunik (2019). A uniform kernel is employed, where all observations within the bandwidth are weighted equally.

Figure 3.6: Graphical Analysis of the RD Design for Cutoff at 5 Million Dirhams




Notes: These plots present the graphical RD analysis of the outcome variables of the log of AR/sales and the log of AP/sales in 2020 and 2021 when the threshold is set at 5 million dirhams. Each dot represents the mean in its corresponding bin of the outcome variable. The solid line is a linear regression of the outcome variable, fitted separately above and below the cutoff.

Table 3.7 provides the RD estimation results. As we can see, the coefficient $\beta_1$ is not significant for the four outcome variables, which corroborates our graphical analysis. This indicates that there are no significant changes in the trade credit level of large firms after the mandatory billing requirement of the GID platform. The results are also insignificant for the threshold of 2 million dirhams in 2020. Figure 3.7 shows that there are no jumps at the cutoff and Table 3.8 further confirms this statistically.

Table 3.7: RD Estimation for Cutoff at 5 Million Dirhams

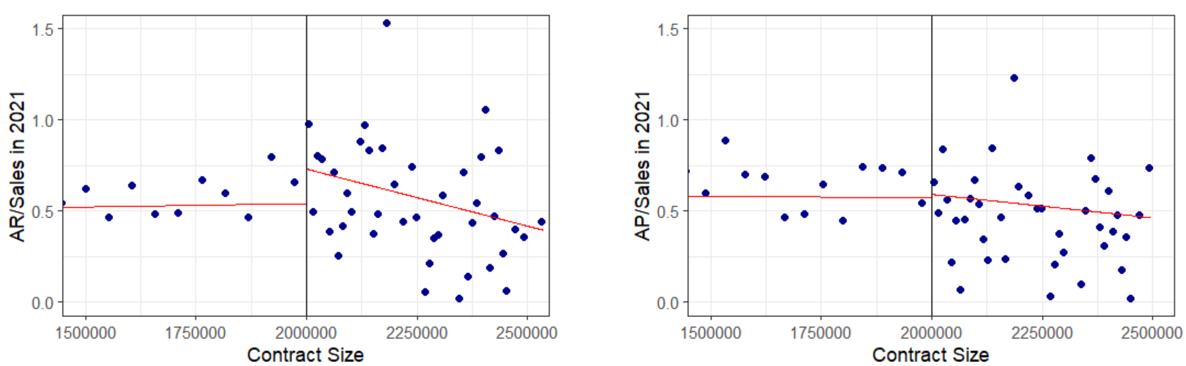
| | (1) $\ln(AR/Sales)_{2020}$ | (2) $\ln(AR/Sales)_{2021}$ | (3) $\ln(AP/Sales)_{2020}$ | (4) $\ln(AP/Sales)_{2021}$ |
|---|---|---|---|---|
| $GID_{i,2019}$ | 0.046 | -0.229 | 0.043 | 0.017 |
| | (0.107) | (0.160) | (0.109) | (0.156) |
| $N$ | 2682 | 1548 | 2682 | 1548 |
| Kernel | Uniform | Uniform | Uniform | Uniform |
| Bandwidth Selection Method | mserd | mserd | mserd | mserd |
| Bandwidth | 1635086 | 1400781 | 1677290 | 1514546 |
| Bias Bandwidth | 2953733 | 2505133 | 3476261 | 2753553 |

Notes: This table presents the RD estimation for the outcome variables of the log of AR/sales and the log of AP/sales in 2020 and 2021 when the threshold is set at 5 million dirhams.

Figure 3.7: Graphical Analysis of the RD Design for Cutoff at 2 Million Dirhams

(a) AR/Sales in 2021

(b) AP/Sales in 2021




Notes: These plots present the graphical RD analysis of the outcome variables of the log of AR/sales and the log of AP/sales in 2021 when the threshold is set at 2 million dirhams. Each dot represents the mean in its corresponding bin of the outcome variable. The solid line is a linear regression of the outcome variable, fitted separately above and below the cutoff.

Table 3.8: RD Estimation for Cutoff at 2 million dirhams

| | (1) $\ln(\text{AR/Sales})_{2021}$ | (2) $\ln(\text{AP/Sales})_{2021}$ |
|---|---|---|
| $\text{GID}_{i,2020}$ | 0.186 | 0.093 |
| | (0.105) | (0.121) |
| $N$ | 1773 | 1773 |
| Kernel | Uniform | Uniform |
| Bandwidth Selection Method | mserd | mserd |
| Bandwidth | 607565 | 513350 |
| Bias Bandwidth | 1128449 | 1019267 |

Notes: This table presents the RD estimation for the outcome variables of AR/sales and AP/sales in 2021 when the threshold is set at 2 million dirhams.

## 3.5 Conclusion

This paper assesses the impact of the new law Act 49-15 on payment delays of public contracts in Morocco. Our sector-level analysis shows that the reform has been effective at decreasing the amount of trade credit of firms in sectors exposed to government spending, the target of the reform. However, our firm-level analysis reveals that large firms benefit more from the reform and experience a significant decrease in trade credit, compared to small firms. In addition, the mandatory measure of the GID platform does not demonstrate any effect on reducing the trade credit of engaged firms. Our research highlights the fact that policy reforms can have size-dependent effects.

# Appendix

## Tables and Figures

Table A1: Engagement Rate by Year

| Year | Total Firm Number | TGR Contract Number | Ratio | Ratio of Value Added |
|---|---|---|---|---|
| 2010 | 50267 | 2411 | 4.80% | 25.45% |
| 2011 | 59559 | 3247 | 5.45% | 17.46% |
| 2012 | 65789 | 3532 | 5.37% | 13.63% |
| 2013 | 72051 | 4032 | 5.60% | 19.62% |
| 2014 | 75762 | 5212 | 6.88% | 22.05% |
| 2015 | 85338 | 5833 | 6.84% | 22.36% |
| 2016 | 107321 | 7411 | 6.91% | 18.61% |
| 2017 | 114146 | 8097 | 7.09% | 22.79% |
| 2018 | 113599 | 7752 | 6.82% | 20.46% |

Notes: This table shows the number of total firms, the number of TGR firms, the percentage of firms with at least one public contract to the total number of firms, and their contribution to the value added.

Table A2: Sector Composition of TGR Contracts

| Sector | Percent |
|---|---|
| Agriculture | 1.08% |
| Commerce | 30.10% |
| Construction | 33.67% |
| Education | 0.84% |
| Electricity and Water | 0.58% |
| Hotels and Restaurants | 3.73% |
| Mining Industry | 0.20% |
| Public Administration | 0.07% |
| Refining Industry | 6.23% |
| Services | 18.08% |
| Telecommunication | 3.55% |
| Transportation | 1.88% |

Notes: This table provides the sector distribution of TGR contracts.

Table A4: Summary Statistics of TGR and Non-TGR Firms

| Statistic | N | Mean | St. Dev. | Min | Max |
|---|---|---|---|---|---|
| TGR Firms | | | | | |
| Accounts Receivable/Sales | 53,932 | 0.49 | 0.41 | 0.00 | 2.00 |
| Accounts Payable/Sales | 53,932 | 0.65 | 0.46 | 0.00 | 2.00 |
| Cash/Total Assets | 53,926 | 0.18 | 0.25 | 0.00 | 1.00 |
| Fixed Assets/Total Assets | 53,924 | 0.12 | 0.17 | 0.00 | 1.00 |
| Current Liabilities/Total Assets | 47,710 | 0.59 | 0.26 | 0.00 | 1.00 |
| Costs of Employees/Total Assets | 53,489 | 0.17 | 0.22 | 0.00 | 2.00 |
| Sales/Total Assets | 53,932 | 1.61 | 1.70 | 0.0001 | 10.43 |
| Non-TGR firms | | | | | |
| Accounts Receivable/Sales | 493,760 | 0.45 | 0.42 | 0.00 | 2.00 |
| Accounts Payable/Sales | 493,760 | 0.64 | 0.50 | 0.00 | 2.00 |
| Cash/Total Assets | 493,691 | 0.19 | 0.26 | 0.00 | 1.00 |
| Fixed Assets/Total Assets | 493,707 | 0.16 | 0.23 | 0.00 | 1.00 |
| Current Liabilities/Total Assets | 423,714 | 0.55 | 0.29 | 0.00 | 1.00 |
| Costs of Employees/Total Assets | 485,385 | 0.19 | 0.29 | 0.00 | 2.00 |
| Sales/Total Assets | 493,760 | 1.64 | 1.80 | 0.0000 | 10.43 |

Notes: This table provides the summary statistics of key variables for TGR and non-TGR firms.

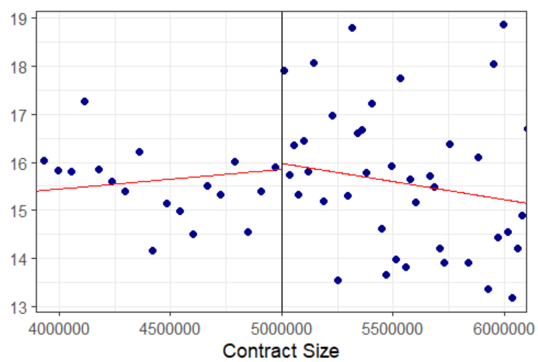
Table A3: TGR Contracts across Different Firm Sizes

| Firm Size | Percent | Average Contract Size (million dirhams) |
|---|---|---|
| Large | 2.0% | 53.75 |
| Medium | 3.2% | 16.68 |
| Small | 13.2% | 5.54 |
| Very Small | 17.6% | 2.00 |
| Micro | 63.9% | 0.63 |

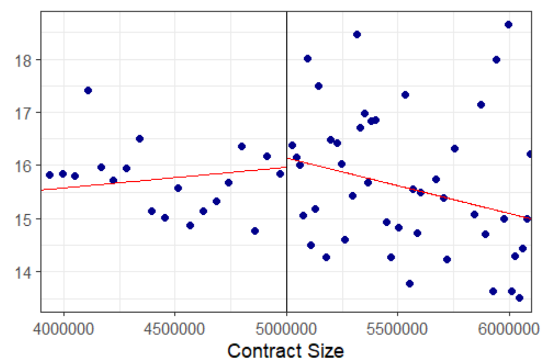
Notes: This table provides the distribution of TGR contracts by firm size and its corresponding average contract size.

Figure 3.8: RD Plots for Control Variables

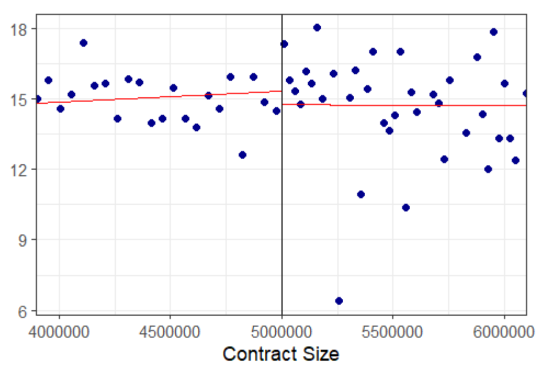
(a) Total Assets



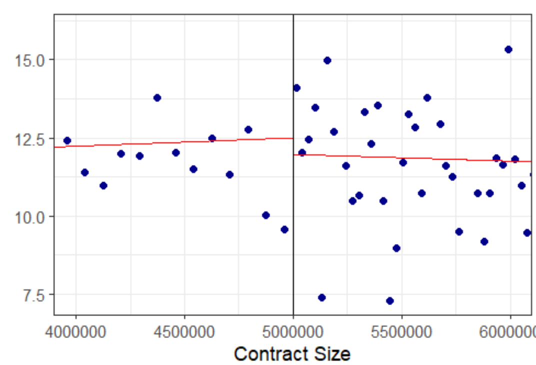

(b) Sales



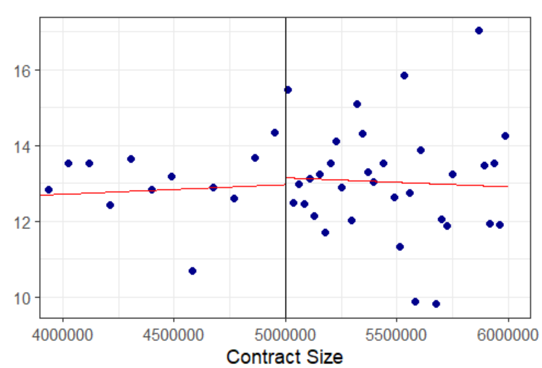

(c) Current Liabilities




(d) Cash




(e) Fixed Assets




Notes: This figure provides RD plots for control variables of total assets, sales, current liabilities, cash, and fixed assets.

Table A5: Sector Exposure to Public Administrations

| Industry | Exposure Percent |
|---|---|
| Agriculture, Forestry, and Fishing | 0.07% |
| Extraction Industry | 1.03% |
| Manufacturing Industries | 8.40% |
| Electricity and Water | 6.62% |
| Construction and Public Works | 0.66% |
| Commerce (Trade) | 23.86% |
| Hotels and Restaurants | 9.10% |
| Transportation | 2.10% |
| Postal and Telecommunications | 4.20% |
| Financial and Insurance Activities | 2.94% |
| Services to Businesses and Personal Services | 4.76% |
| Education, Health, and Social Action | 0.84% |

Notes: This table provides the percentage of goods and services from the listed sectors consumed by public administration relative to its total consumption.

Table A6: Balanced Covariate Tests

| Variable | Bandwidth | RD Estimator | Std. Err. | Number of Obs. |
|---|---|---|---|---|
| ln(Total Assets)$_{it-1}$ | 1315849 | 0.493 | 0.459 | 4650 |
| ln(Cash/Total Assets)$_{it-1}$ | 1791371 | 0.692 | 0.549 | 4601 |
| ln(Fixed Assets/Total Assets)$_{it-1}$ | 1001619 | 0.299 | 0.681 | 3493 |
| ln(Current Liabilities/Total Assets)$_{it-1}$ | 1766530 | -0.575 | 0.599 | 4650 |
| ln(Sales/Total Assets)$_{it-1}$ | 1148001 | 0.552 | 0.437 | 4650 |

Notes: This table provides the RD estimation of all control variables as part of the falsification tests. As it shows, none of the estimation results is significant.

Table A7: Heterogeneous Effects on Accounts Receivable and Payable

| | (1) $\ln(\text{Cash})_{it}$ | (2) $\ln(\text{Fixed Assets})_{it}$ | (3) $\ln(\text{Wage})_{it}$ | (4) $\text{Exit}_{it}$ |
|---|---|---|---|---|
| $\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | 0.005 | -0.005 | -0.055+ | -0.009 |
| | (0.053) | (0.066) | (0.032) | (0.010) |
| 2nd Tercile | 0.028 | 0.153*** | 0.249*** | -0.003 |
| | (0.021) | (0.023) | (0.013) | (0.004) |
| 3rd Tercile | 0.020 | 0.265*** | 0.501*** | -0.011* |
| | (0.027) | (0.026) | (0.016) | (0.005) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}$ | 0.120* | -0.114* | -0.117*** | -0.009 |
| | (0.054) | (0.057) | (0.029) | (0.008) |
| 3rd Tercile$\times\text{TGR}_{pre-2016,i}$ | 0.043 | -0.088 | -0.094** | -0.018+ |
| | (0.067) | (0.069) | (0.035) | (0.010) |
| 2nd Tercile$\times\text{Post}_t$ | 0.027 | 0.118*** | 0.005 | -0.023*** |
| | (0.022) | (0.024) | (0.012) | (0.004) |
| 3rd Tercile$\times\text{Post}_t$ | 0.108*** | 0.308*** | 0.038** | -0.033*** |
| | (0.021) | (0.023) | (0.012) | (0.004) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | -0.149* | 0.021 | 0.097** | -0.007 |
| | (0.061) | (0.070) | (0.033) | (0.011) |
| 3rd Tercile$\times\text{TGR}_{pre-2016,i}\times\text{Post}_t$ | -0.016 | 0.038 | 0.080* | -0.011 |
| | (0.059) | (0.068) | (0.033) | (0.010) |
| $\ln(\text{Total Assets})_{it-1}$ | -0.188*** | 0.398*** | 0.318*** | 0.017*** |
| | (0.008) | (0.009) | (0.005) | (0.001) |
| $\ln(\text{Cash/Total Assets})_{it-1}$ | 0.050*** | 0.017*** | 0.002 | -0.003*** |
| | (0.003) | (0.002) | (0.001) | (0.001) |
| $\ln(\text{Fixed Assets/Total Assets})_{it-1}$ | -0.031*** | 0.283*** | 0.022*** | 0.003*** |
| | (0.003) | (0.005) | (0.002) | (0.001) |
| $\ln(\text{Current Liabilities/Total Assets})_{it-1}$ | -0.038*** | -0.026*** | -0.008*** | 0.002** |
| | (0.004) | (0.004) | (0.002) | (0.001) |
| $\ln(\text{Sales/Total Assets})_{it-1}$ | -0.011 | 0.025** | 0.179*** | 0.008*** |
| | (0.007) | (0.008) | (0.004) | (0.001) |
| $N$ | 218510 | 181365 | 193236 | 219040 |
| adj. $R^2$ | 0.644 | 0.896 | 0.936 | 0.655 |

Standard errors in parentheses
+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: This table reports the coefficients from the regression (3.3). Outcome variables are the log of cash, fixed assets, wage, and probability of exiting the market by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

## Robustness Checks

Table A8: Different Size Measure: Total Assets from Past Year

| | (1) $\ln(\frac{AR}{Sales})_{it}$ | (2) $\ln(\frac{AP}{Sales})_{it}$ | (3) $\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $TGR_{pre-2016,i} \times Post_t$ | 0.389** | -0.385** | 0.774*** |
| | (0.139) | (0.123) | (0.158) |
| $\ln(\text{Total Assets})_{it-1}$ | 0.150*** | 0.043*** | 0.107*** |
| | (0.007) | (0.007) | (0.008) |
| $TGR_{pre-2016,i} \times \ln(\text{Total Assets})_{it-1}$ | 0.044* | 0.008 | 0.036+ |
| | (0.018) | (0.016) | (0.019) |
| $Post_t \times \ln(\text{Total Assets})_{it-1}$ | -0.048*** | -0.029*** | -0.018*** |
| | (0.004) | (0.003) | (0.004) |
| $TGR_{pre-2016,i} \times Post_t \times \ln(\text{Total Assets})_{it-1}$ | -0.027** | 0.024** | -0.050*** |
| | (0.009) | (0.008) | (0.010) |
| $\ln(\text{Cash/Total Assets})_{it-1}$ | -0.013*** | -0.040*** | 0.027*** |
| | (0.002) | (0.002) | (0.002) |
| $\ln(\text{Fixed Assets/Total Assets})_{it-1}$ | -0.011*** | 0.019*** | -0.030*** |
| | (0.003) | (0.003) | (0.003) |
| $\ln(\text{Current Liabilities/Total Assets})_{it-1}$ | -0.042*** | 0.030*** | -0.072*** |
| | (0.005) | (0.006) | (0.007) |
| $\ln(\text{Sales/Total Assets})_{it-1}$ | -0.017* | -0.050*** | 0.033*** |
| | (0.007) | (0.007) | (0.008) |
| $N$ | 219040 | 219040 | 219040 |
| adj. $R^2$ | 0.695 | 0.607 | 0.576 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: In this robustness test, we replace the size measure in our main analysis (size bin) with the total assets of a firm $i$ from year $t-1$. Outcome variables are the log of the ratios of accounts receivable to sales, accounts payables to sales, and the difference between accounts receivable and accounts payable to sales by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

## Exposure Measure

We assess firm-level exposure to the reform by measuring the ratio of contract value to sales. We estimate the following regression:

$$Y_{it} = \beta_0 + \beta_1 Exposure_{pre-2016,i} \times Post_t + \delta_i + \delta_t + \Gamma X_{it} + \epsilon_{it}, \tag{3.6}$$

Table A9: Estimation Results without Covid Years

| | (1) $\ln(\frac{AR}{Sales})_{it}$ | (2) $\ln(\frac{AP}{Sales})_{it}$ | (3) $\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $TGR_{pre-2016,i} \times Post_t$ | -0.019 | -0.029* | 0.010 |
| | (0.013) | (0.013) | (0.016) |
| $\ln(\text{Total Assets})_{it-1}$ | 0.115*** | 0.040*** | 0.075*** |
| | (0.008) | (0.008) | (0.009) |
| $\ln(\text{Cash/Total Assets})_{it-1}$ | -0.009*** | -0.036*** | 0.026*** |
| | (0.002) | (0.002) | (0.003) |
| $\ln(\text{Fixed Assets/Total Assets})_{it-1}$ | -0.006+ | 0.016*** | -0.022*** |
| | (0.003) | (0.003) | (0.004) |
| $\ln(\text{Current Liabilities/Total Assets})_{it-1}$ | -0.042*** | -0.019** | -0.023** |
| | (0.005) | (0.007) | (0.008) |
| $\ln(\text{Sales/Total Assets})_{it-1}$ | 0.018* | -0.013+ | 0.032*** |
| | (0.008) | (0.008) | (0.009) |
| $N$ | 170120 | 170120 | 170120 |
| adj. $R^2$ | 0.718 | 0.623 | 0.595 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: In this robustness test, we estimate our regression 3.3 without covid years. Outcome variables are the log of the ratios of accounts receivable to sales, accounts payables to sales, and the difference between accounts receivable and accounts payable to sales by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A10: Estimation Results without Covid Years (Firm Size)

| | (1) $\ln(\frac{AR}{Sales})_{it}$ | (2) $\ln(\frac{AP}{Sales})_{it}$ | (3) $\ln(\frac{AR}{Sales})_{it}$ - $\ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $TGR_{pre-2016,i} \times Post_t$ | 0.123* | -0.039 | 0.161* |
| | (0.055) | (0.048) | (0.064) |
| 2nd Tercile | -0.368*** | -0.313*** | -0.055* |
| | (0.020) | (0.017) | (0.023) |
| 3rd Tercile | -0.683*** | -0.641*** | -0.042 |
| | (0.023) | (0.021) | (0.026) |
| 2nd Tercile$\times TGR_{pre-2016,i}$ | 0.065 | -0.028 | 0.093+ |
| | (0.052) | (0.044) | (0.057) |
| 3rd Tercile$\times TGR_{pre-2016,i}$ | 0.153** | -0.007 | 0.160* |
| | (0.059) | (0.051) | (0.063) |
| 2nd Tercile$\times Post_t$ | -0.044* | -0.094*** | 0.050* |
| | (0.021) | (0.018) | (0.025) |
| 3rd Tercile$\times Post_t$ | -0.147*** | -0.120*** | -0.027 |
| | (0.020) | (0.018) | (0.024) |
| 2nd Tercile$\times TGR_{pre-2016,i} \times Post_t$ | -0.088 | 0.035 | -0.123+ |
| | (0.058) | (0.052) | (0.067) |
| 3rd Tercile$\times TGR_{pre-2016,i} \times Post_t$ | -0.185** | 0.019 | -0.203** |
| | (0.056) | (0.049) | (0.065) |
| ln(Total Assets)$_{it-1}$ | 0.192*** | 0.085*** | 0.107*** |
| | (0.007) | (0.006) | (0.008) |
| ln(Cash/Total Assets)$_{it-1}$ | -0.014*** | -0.040*** | 0.027*** |
| | (0.002) | (0.002) | (0.002) |
| ln(Fixed Assets/Total Assets)$_{it-1}$ | -0.011*** | 0.019*** | -0.030*** |
| | (0.003) | (0.002) | (0.003) |
| ln(Current Liabilities/Total Assets)$_{it-1}$ | -0.039*** | 0.032*** | -0.072*** |
| | (0.005) | (0.006) | (0.007) |
| ln(Sales/Total Assets)$_{it-1}$ | 0.015* | -0.022*** | 0.037*** |
| | (0.007) | (0.007) | (0.008) |
| $N$ | 170120 | 170120 | 170120 |
| adj. $R^2$ | 0.724 | 0.630 | 0.595 |
| FE (Firm, Sector × Year) | Yes | Yes | Yes |

Notes: In this robustness test, we estimate our regression 3.4 without covid years. Outcome variables are the log of the ratios of accounts receivable to sales, accounts payables to sales, and the difference between accounts receivable and accounts payable to sales by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A11: Estimation Results of Ratios

| | (1) $\frac{\text{AR}}{\text{Sales}}$ | (2) $\frac{\text{AP}}{\text{Sales}}$ |
|---|---|---|
| $\text{TGR}_{pre-2016,i}\times\text{Post}_\text{t}$ | -0.010* | -0.011* |
| | (0.005) | (0.005) |
| $\ln(\text{Total Assets})_{\text{it-1}}$ | 0.032*** | 0.016*** |
| | (0.002) | (0.002) |
| $\ln(\text{Cash/Total Assets})_{\text{it-1}}$ | -0.009*** | -0.016*** |
| | (0.001) | (0.001) |
| $\ln(\text{Fixed Assets/Total Assets})_{\text{it-1}}$ | -0.008*** | 0.004*** |
| | (0.001) | (0.001) |
| $\ln(\text{Current Liabilities/Total Assets})_{\text{it-1}}$ | -0.011*** | 0.026*** |
| | (0.001) | (0.001) |
| $\ln(\text{Sales/Total Assets})_{\text{it-1}}$ | -0.037*** | -0.042*** |
| | (0.002) | (0.002) |
| $N$ | 219040 | 219040 |
| adj. $R^2$ | 0.668 | 0.625 |
| FE (Firm, Sector × Year) | Yes | Yes |

Notes: This table reports the coefficients from the regression (3.3). Outcome variables are the ratios of accounts receivable to sales and, accounts payables to sales by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

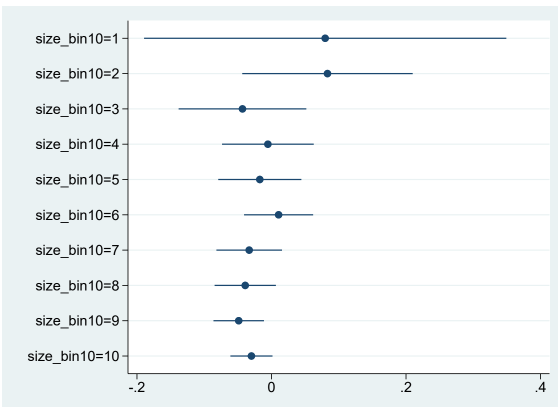
Table A12: Estimation Results with Ratios (Firm Size)

| | (1) $\frac{AR}{Sales}$ | (2) $\frac{AP}{Sales}$ |
|---|---|---|
| $\text{TGR}_{pre-2016,i}\times\text{Post}_\text{t}$ | 0.022 | -0.019 |
| | (0.014) | (0.016) |
| 2nd Tercile | -0.184*** | -0.213*** |
| | (0.006) | (0.006) |
| 3rd Tercile | -0.332*** | -0.406*** |
| | (0.007) | (0.008) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}$ | -0.006 | -0.040** |
| | (0.014) | (0.015) |
| 3rd Tercile$\times\text{TGR}_{pre-2016,i}$ | -0.026 | -0.063*** |
| | (0.016) | (0.018) |
| 2nd Tercile$\times\text{Post}_\text{t}$ | -0.002 | -0.025*** |
| | (0.006) | (0.006) |
| 3rd Tercile$\times\text{Post}_\text{t}$ | -0.048*** | -0.038*** |
| | (0.006) | (0.006) |
| 2nd Tercile$\times\text{TGR}_{pre-2016,i}\times\text{Post}_\text{t}$ | -0.007 | 0.023 |
| | (0.016) | (0.018) |
| 3rd Tercile$\times\text{TGR}_{pre-2016,i}\times\text{Post}_\text{t}$ | -0.036* | 0.013 |
| | (0.015) | (0.017) |
| $\ln(\text{Total Assets})_\text{it-1}$ | 0.057*** | 0.047*** |
| | (0.002) | (0.002) |
| $\ln(\text{Cash/Total Assets})_\text{it-1}$ | -0.009*** | -0.016*** |
| | (0.001) | (0.001) |
| $\ln(\text{Fixed Assets/Total Assets})_\text{it-1}$ | -0.008*** | 0.005*** |
| | (0.001) | (0.001) |
| $\ln(\text{Current Liabilities/Total Assets})_\text{it-1}$ | -0.009*** | 0.028*** |
| | (0.001) | (0.001) |
| $\ln(\text{Sales/Total Assets})_\text{it-1}$ | -0.023*** | -0.026*** |
| | (0.002) | (0.002) |
| $N$ | 219040 | 219040 |
| adj. $R^2$ | 0.686 | 0.647 |
| FE (Firm, Sector × Year) | Yes | Yes |

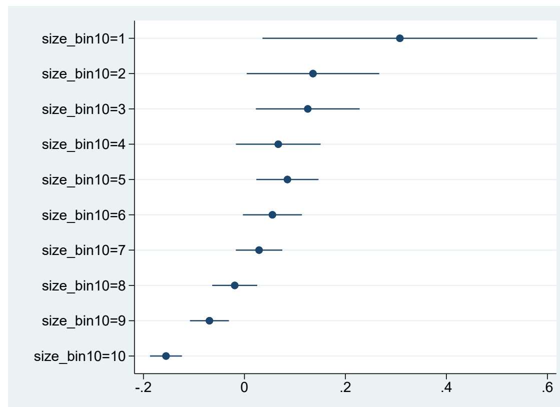
Notes: This table reports the coefficients from the regression (3.3). Outcome variables are the ratios of accounts receivable to sales and, accounts payables to sales by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Figure 3.9: Coefficient Plots with 10 Sales Bin

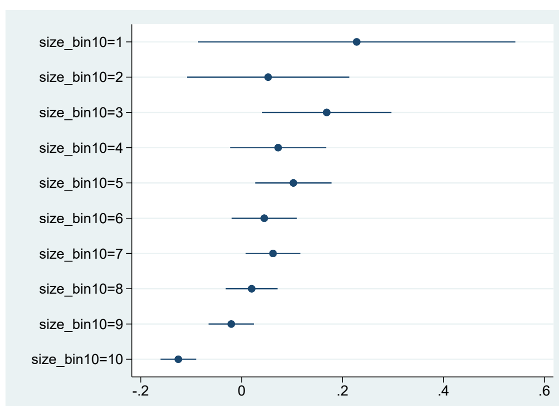
(a) Accounts Receivable




(b) Accounts Payable




(c) Net Trade Credit




Notes: This figure plots the coefficients of the $\beta_1$ from Regression (3.4) when we define size by dividing sales into 10 bins.

where $\text{Exposure}_{\text{pre-2016, i}}$ represents the average ratio of TGR contract value to sales in or before 2016 for firm $i$. The results are reported in Table A13. The coefficients are significantly negative for accounts receivable and payable in the first two columns, and not significant for net trade credit in Column 3. This result is consistent with our main estimation. However, the magnitude of the coefficient is very small. This suggests that while statistically significant, the economic effects in terms of exposure measures are relatively small.

Table A13: Treatment Effects on Accounts Receivable and Payable

| | (1)<br>$\ln(\frac{AR}{Sales})_{it}$ | (2)<br>$\ln(\frac{AP}{Sales})_{it}$ | (3)<br>$\ln(\frac{AR}{Sales})_{it} - \ln(\frac{AP}{Sales})_{it}$ |
|---|---|---|---|
| $\text{Exposure}_{\text{pre-2016,i}}$ | 1.97e-07 | 1.41e-07 | 5.65e-08 |
| | (6.38e-07) | (7.38e-07) | (1.63e-07) |
| $\text{Exposure}_{\text{pre-2016,i}} \times \text{Post}_t$ | -4.92e-07*** | -5.29e-07*** | 3.79e-08 |
| | (5.45e-08) | (7.27e-08) | (5.87e-08) |
| $\ln(\text{Total Assets})_{\text{it-1}}$ | 0.183*** | 0.059*** | 0.124*** |
| | (0.015) | (0.013) | (0.016) |
| $\ln(\text{Cash/Total Assets})_{\text{it-1}}$ | -0.013** | -0.037*** | 0.024*** |
| | (0.004) | (0.004) | (0.005) |
| $\ln(\text{Fixed Assets/Total Assets})_{\text{it-1}}$ | -0.012* | 0.020*** | -0.032*** |
| | (0.006) | (0.005) | (0.006) |
| $\ln(\text{Current Liabilities/Total Assets})_{\text{it-1}}$ | -0.030** | 0.023+ | -0.053*** |
| | (0.011) | (0.013) | (0.015) |
| $\ln(\text{Sales/Total Assets})_{\text{it-1}}$ | -0.005 | -0.039** | 0.035* |
| | (0.016) | (0.013) | (0.016) |
| $N$ | 47921 | 47921 | 47921 |
| adj. $R^2$ | 0.663 | 0.558 | 0.507 |

Standard errors in parentheses

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: This table reports the coefficients from the regression (3.6). Outcome variables are the log of the ratios of accounts receivable to sales, the log of the ratio of accounts payables to sales, and the difference between the two logged ratios by firm $i$ in sector $j$ in year $t$. Firm and sector-year fixed effects are included. Standard errors are clustered at the individual firm level. Significance level: + $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

## Literature Review

The early group of literature lays out broad discussions on the reasons why trade credit linkages are essential in the production network. Meltzer (1960) is among the first to empirically document the substitution effect between bank loans and trade credit. He argues that trade credit is a channel to redistribute liquidity from large firms, which are favored by credit rationing, to firms that are discriminated against by banks in a period of "tight money". Schwartz (1974) disciplines this finding by modeling trade credit as a part of firms' pricing policy. He summarizes that trade credit is passed from firms that have easy and cheap access to capital markets to their productive customers with less capital. Petersen and Rajan (1997) provide comprehensive empirical tests to explain why firms extend trade credit when banks could provide loans. They look at a sample of small US firms and emphasize suppliers' advantages in information acquisition, controlling the buyer, and salvaging the value from existing assets.

The relatively recent strand of literature specifically explores the role of trade credit in the business cycle and its macroeconomic implications. The empirical literature has identified two opposing forces of trade credit, as summarized by Coricelli and Frigerio (2019): (1) redistribution of liquidity (counter-cyclical) in normal times; (2) upstream transmission of shocks (pro-cyclical) in financial crisis.

Empirical evidence on the redistribution channel mainly emerged in the early 2000s before the major financial crisis broke out. It extends the findings of the previous group of literature and further explores the credit redistribution channel from bank lending to trade credit. Nilsen (2002) finds that, during monetary contraction, small firms increase trade credit to replace bank loans whereas large firms increase trade credit to substitute the financing from the capital market. Fisman and Love (2003) argue that trade credit is an important way of credit reallocation across industries. By exploring panel data of 37 industries and 44 countries, they confirm that industries with higher dependence on trade credit exhibit higher rates of growth. Mateut, Bougheas, and Mizen (2006) set up a model to illustrate how trade credit can smooth out the impact of a reduction in bank lending when monetary policy tightens. They validate their model by looking at UK manufacturing firms during 1990-1999. Cuñat (2007) model trade credit suppliers as liquidity providers and debt collectors in the presence of an efficient banking sector. Most recently, Shenoy and Williams (2017) exploit changes in interstate bank branching laws in the US as exogenous liquidity shocks. They implement DID and 2SLS estimation strategies to a manually linked sample of supplier-customer pairs of US public firms. They confirm that supplier firms with greater access to banking liquidity offer more trade credit to their customers.

Another group of empirical literature has emerged after the series of financial crises, namely the Asian financial crisis, the global financial crisis of 2007-2008, and the European debt crisis. They argue that the redistribution effect of trade credit is reduced to a limited extent during a financial crisis. Love, Preve, and Sarria-Allende (2007) provide empirical evidence on emerging market economies and document an overall reduced trend in trade credit supply after a bank credit crunch during a crisis. Love and Zaidi (2010) further expand this finding by examining the borrowing behaviors of small and medium enterprises (SMEs). Garcia-Appendini and Montoriol-Garriga (2013) emphasize the heterogeneity among suppliers' behaviors during a credit crunch and show that ex-ante cash-rich suppliers are able to extend more trade credit while cash-poor firms reduce the amount during the Great Recession in the US. Gonçalves, Schiozer, and Sheng (2018) focuses on the heterogeneity of suppliers' market power. They find that firms with high market power increased the supply of trade credit to avoid the loss of monopoly rents during the 2007-08 financial crisis in the US. On the contrary, Coricelli and Frigerio (2019) show that European SMEs with low bargaining power in fact extend more trade credit to their larger counterparts during the Great Recession. McGuinness, Hogan, and Powell (2018), using similar data on European SMEs, find that it is the cash-rich SMEs that extend more net trade credit, corroborating Garcia-Appendini and Montoriol-Garriga (2013)'s findings. They also show that trade credit, as an alternative source of external financing, has a significantly positive impact on SME survival. Carbó-Valverde, Rodríguez-Fernández, and F. Udell (2016) specify that the dependence of SMEs on trade credit is decided by its tightness of credit constraint.

One strand of recent empirical literature highlights the role of trade credit as a transmission channel for propagating financial shocks and transmitting financial distress upstream. The seminal work by Raddatz (2010) points out the output correlation on an industry level along the trade credit chain. Their results are based on a sample of 378 manufacturing industry pairs across 43 countries and show that the use of trade credit amplifies sectoral shocks. More recent studies confirm the role of trade credit in propagating financial shocks (Jorion and Zhang, 2009; Costello, 2020).

The recent availability of granular firm-to-firm payment data in different countries inspires an expanding group of literature furnishing rich empirical evidence on the role of trade credit in transmitting financial shocks. Boissay and Gropp (2013) use detailed microdata on payment defaults on a firm level in France and provide evidence on the pass-through of adverse liquidity shocks from credit-constrained firms to their suppliers. In a similar vein, Jacobson and Schedvin (2015) uses an exhaustive data set on Swedish corporate bankruptcies and associated trade credit claims between suppliers and customers. They show that trade credit is a quantitatively important mechanism to propagate corporate bankruptcy. Similar findings on the propagation of liquidity shocks can be found in Cortes, Silva, and Doornik (2019) for Brazil, Jiménez et al. (2020) for Spain, and Giannetti, Serrano-Velarde, and Tarantino (2021) for Italy.

The group of theoretical literature on the role of trade credit has been expanding rapidly, providing explanations of the mechanisms behind the empirical findings. Bigio and La'O (2020) is the first paper to include financial frictions in a production network. They highlight the transmission of sector-specific productivity shocks through the input-output network in the US during the Great Recession. The major limitation of their model is that the financial constraints is treated as exogenous. Their model has been later extended in a series of works, which endogenizes trade credit (Altinoglu, 2021; Miranda-Pinto and Zhang, 2022; Luo, 2020; Reischer, 2019; Shao, 2022).

Altinoglu (2021) imposes a constraint on the amount of trade credit a firm can obtain from its suppliers depending on the firm's cash flow. However, he only considers the amplification effect of financial shocks through trade credit channels, not its mitigation effect through substituting bank loans. Miranda-Pinto and Zhang (2022) discuss both effects and differentiate between two types of banking shocks. They find that when shocks are idiosyncratic, trade credit can act as a mitigation mechanism; when the shocks are strongly correlated, the amplification effect dominates. Luo (2020) endogenizes the supply of trade credit by adjusting its size and payment schedule in his model. His findings echo Miranda-Pinto and Zhang (2022) and emphasize that the types of shocks indeed matter. Trade credit can mitigate small shocks but can amplify big ones. Reischer (2019) includes endogenous adjustment of volume and cost of trade credit. She focuses on the amplification effect and quantifies the disruptions of production networks during the Great Recession. Shao (2022) explores the role of trade credit in a model with firm heterogeneity in financial constraints. He argues that if the shock impacts a small fraction of entrepreneurs, trade credit mitigates by redistributing liquidity. However, when the shock is on an aggregate level and impacts all entrepreneurs in an economy, trade credit amplifies the financial distortions.

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