Financial Development, Labor Markets, and Aggregate Productivity: Evidence from Brazil

Job Market Paper

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Abstract

We estimate the effect of an increase in the availability of bank credit on the employment and the earnings of high- and low-skilled workers. To do so, we consider a bankruptcy reform that increased the legal protections of secured creditors, which led to an expansion of bank credit to Brazilian firms. We use detailed administrative data and an empirical strategy that compares changes in outcomes for financially constrained firms, which were affected by the bankruptcy reform, with unconstrained firms, which were largely unaffected by the reform. Following the bankruptcy reform and subsequent expansion in credit, constrained firms increased employment, especially of high-skilled workers. We also observe an increase in earnings, with gains concentrated on skilled workers and on workers who were employed at constrained firms prior to the reform. To rationalize these findings, we design a model in which heterogeneous producers face constraints in their ability to borrow and have production functions featuring capital-skill complementarity. Using this framework, we estimate that the reallocation of resources induced by the bankruptcy reform accounts for 36 percent of the observed increase in aggregate productivity in Brazil during the 2000s.

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Financial constraints are a pervasive characteristic of low- and middle-income economies. In Brazil, for instance, 45 percent of firms identify access to finance as a major constraint (2009 World Bank Enterprise Survey). Extensive literature has found that financial frictions affect economic development not only by slowing down capital accumulation but also by causing capital to be misallocated across producers.[1] Because there is complementarity between capital and labor, these findings suggest a role for the cost and availability of external finance in determining labor market outcomes. Moreover, financial frictions may also impact the skill composition of a firm's workforce, as well as the returns to skill, as skilled and unskilled labor potentially differ in how complementary they are to capital.

This paper sheds light on the effect of increased access to bank credit on the employment and earnings of high- and low-skilled workers. To conduct our analysis, we assemble a comprehensive firm-level panel of formally registered Brazilian firms by merging matched employer-employee data with credit registry data and with data on real outcomes such as assets, investment, and output. Our identification strategy makes use of a 2005 reform to the legislation governing bankruptcy proceedings in Brazil, which significantly strengthened the rights of secured creditors and led to an increase in the borrowing capacity of firms. We also develop a framework that allows us to quantify the effect of the credit expansion caused by the 2005 bankruptcy reform on aggregate productivity, taking into account its impact on capital and on both skilled and unskilled labor.

We start by documenting that the 2005 bankruptcy reform significantly loosened credit constraints. In order to identify financially constrained firms, we estimate the firm-level wedge between the marginal revenue product of capital and its user cost. We sort firms into quartiles according to this measure, with firms in the fourth quartile being the most financially constrained and firms in the first quartile being the least constrained. We find that the reform led to a sizable and significant increase in firms' access to bank credit, with the most-constrained firms experiencing a 7 percent increase in bank credit following the reform relative to the least-constrained firms.

Our research has three main empirical results. First, we find that constrained firms increase their employment of both skilled and unskilled workers, but especially of skilled workers, when access to credit is expanded. In particular, we find that firms in the fourth quartile of our measure of financial constraints experience an 8.5 percent increase in the ratio of skilled to unskilled workers and a 6.3 percent increase in total employment relative to firms in the first quartile. The increase in total employment is driven by both an increase in hiring and in worker retention, and the expansion of constrained firms is partly driven by the poaching of employees previously working in unconstrained firms. Following the loosening of credit constraints, constrained firms see a 3.9 percent increase in

[1] Early work includes King and Levine (1993), Jayaratne and Strahan (1996), and Rajan and Zingales (1998). For studies on the impact of financial frictions on capital accumulation, see also Levine and Zervos (1998) and Rioja and Valev (2004). See, for instance, Bertrand, Schoar, and Thesmar (2007), Buera, Kaboski, and Shin (2011), Moll (2014), Cong et al. (2017), Catherine et al. (2017), and Bai et al. (2018) for studies on the impact of financial frictions on the allocation of capital.

the share of their employees with a previous employment spell in an unconstrained firm. This result provides direct evidence that the loosening of financial constraints causes labor to be reallocated from unconstrained firms to constrained firms—which are, on average, younger, smaller, and more productive.

Second, we further find that workers' earnings rise at constrained firms following the bankruptcy reform and subsequent credit expansion. These gains are concentrated on skilled workers, leading to an increase in the within-firm return to skill. In particular, we find that firms in the fourth quartile of our measure of financial constraints see a 10 percent increase in the skill premium relative to firms in the first quartile. This effect is sizable, accounting for almost 60 percent of the difference between the skill premium paid by firms in the first and in the fourth quartiles of our measure of financial constraints in the pre-reform period. We also find that the increase in earnings is stronger for incumbent workers, i.e. for workers that were at the firm prior to the reform.

What can explain the observed increase in the relative utilization of skill and in the return to skill following a credit expansion? One possibility is that capital and skilled labor are relative complements.[2] If that is the case, an increase in capital accumulation will cause skilled labor to become more productive relative to unskilled labor. This in turn will lead to an increase in the employment of skilled workers relative to unskilled workers and/or an increase in the skill premium. In order to shed light on the mechanism behind our results, we leverage the fact that we can link employer-employee data with firm balance sheet information to investigate how firms adjust their investment decisions.

Our third empirical finding is that firms that were more financially constrained increased their level of investment following the expansion in credit. Moreover, we provide direct evidence in favor of the capital-skill complementarity hypothesis by exploiting variation in the the degree of capital-skill complementarity at the industry level. We find that constrained firms in high-complementarity industries increase their utilization of skilled labor and their within-firm return to skill by more than constrained firms in low-complementarity industries.[3]

To rationalize these findings, we design a model in which heterogeneous producers face constraints in their ability to borrow (Moll 2014) and technology is such that skilled labor is more complementary to capital than unskilled labor. Our model features workers who have idiosyncratic tastes for

[2]Another possibility is that financial frictions directly impact employment decisions due to a mismatch between payments to labor and the generation of cash flows or the fact that labor has a fixed-cost component including, for instance, hiring and firing costs (Benmelech, Bergman, and Seru 2015; Bai, Carvalho, and Phillips (2018); Caggese, Cunat, and Metzger (2018)). Since skilled labor requires larger wages than unskilled labor and is arguably associated with higher recruiting costs, this has the potential to explain our results. While we do not rule out this channel, we provide evidence that our results are at least partly driven by capital-skill complementarity.

[3] We use estimates of the parameters of a nested CES production function to compute the elasticities of substitution between unskilled labor and capital and between skilled labor and capital, for each 2-digit industry. Our industry-level measure of capital-skill complementarity is the ratio between the two elasticities—a high ratio implies that, in relative terms, skilled labor is a considerably more complementary to capital than unskilled labor.

different workplaces, allowing firms to set wages with some degree of market power (Card et al. 2018; Kline et al. 2017). Moreover, production functions have a nested CES form and feature capital-skill complementarity as in Krusell et al. (2000). In the presence of complementarity between capital and skilled labor, an increase in capital will cause the productivity of skilled workers to increase by more than the productivity of unskilled workers. These theoretical predictions are in line with the observed increase in the employment of skilled workers and in the skill premium given that we also observe an increase in investment.

We use this framework to quantify the effect of the credit expansion caused by the 2005 bankruptcy reform on aggregate productivity. To do so, we estimate production function parameters for each two-digit industry. Our estimation procedure involves first estimating a second-order approximation of the production function as in De Loecker and Warzynski (2012). In a second step, we use these reduced-form estimates to recover the structural parameters of our nested CES production function using a minimum distance estimation procedure. Once the model is parametrized, we then compute the change in aggregate productivity implied by both the model and the data following Petrin and Levinsohn (2012). We compute the model-implied change in aggregate productivity as a response to a relaxation in credit constraints calibrated to match the 2005 bankruptcy reform. This exercise implies a 2 percent increase in aggregate productivity, which accounts for 36 percent of the observed increase in productivity in Brazil during the 2000s.

We confirm the robustness of our empirical findings to flexibly controlling for industry- and region-specific trends, alleviating concerns that our results are biased by differential firm growth across product categories or localities. We obtain similar estimates when we control for firm-level exposure to the business cycle, reducing concerns that our results are driven by the fact that constrained firms are more productive and, as such, may be expected to grow faster during economic booms. Moreover, our results are qualitatively similar when we consider other common proxies of financial constraints, such as firm size or age.

Overall, our results add to a growing body of evidence supporting the existence of an important link between financial frictions and labor markets. Moreover, our findings suggest that by introducing distortions in the allocation of capital, financial frictions lead to distortions in the allocation of skill. We thus provide new evidence on the specific channels through which financial development can improve the allocation of production factors and hence increase aggregate productivity.

Our work contributes to recent literature on the impact of financial frictions on long-term labor market outcomes. For instance, Benmelech, Bergman, and Seru (2015), Bai et al. (2018), and Caggese et al. (2018) find that financial frictions impact firm-level employment decisions, with consequences for the allocation of labor across producers as well aggregate unemployment rates. We add to this literature by providing evidence that access to external finance impacts the types of workers a firm employs, in terms of both skill and previous experience, as well as the within-firm returns to skill. Moreover, we provide direct evidence that the shift in skill composition and the rise in the skill premium triggered by increased access to credit are at least partly driven by complementarities between capital and skill.

The present work is also connected to a rich body of literature in macroeconomics and finance that studies the impact of financial frictions on the allocation of capital across producers (Bertrand, Schoar, and Thesmar 2007; Buera, Kaboski, and Shin 2011; Moll 2014; Cong et al. 2017; Catherine et al. 2017; Bai et al. 2018). We contribute to this literature in three ways. First, we provide causal, micro-level evidence of the effect of financial constraints on the allocation of both capital and labor in the context of a middle-income country. Second, we find that financial frictions affect not only investment and total employment, but also the types of workers that a firm employs. Finally, we develop a framework that allows us to estimate the aggregate effect of increased access to external finance, taking into account the reallocation of capital and of both skilled and unskilled labor.

This paper also relates to the literature that estimates the effect of transient negative credit supply shocks on employment (Peek and Rosengren 2000, Chodorow-Reich 2014, Greenstone et al. 2015, Duygan-Bump, Levkov, and Montoriol-Garriga 2015, Benmelech, Bergman, and Seru 2015, Bottero, Lenzu, and Mezzanotti 2017, Huber 2018, Benmelech, Frydman, and Papanikolaou 2018). We complement this literature by analyzing the effect of a credit expansion on a wide range of long-term labor market outcomes in a middle-income country. In particular, we shed light on the characteristics of employees whose hiring and firing is impacted by a firm's access to credit. We also provide new evidence on the impact of access to bank credit on workers' earnings, demonstrating that earnings gains are concentrated on skilled workers and on incumbents.

Finally, this paper is also related to previous work analyzing the 2005 Brazilian bankruptcy reform. Ponticelli and Alencar (2016) find that firms in municipalities with less-congested courts experienced a larger increase in credit, investment, and output following the reform. Although we do not exploit variation in the enforcement of the legislation and instead focus on measures of financial frictions, our findings are consistent with these results. We add to this work by investigating the effect of the credit expansion triggered by the reform on the skill composition of firms and on the within-firm returns to skill, and by estimating the impact of the reform on aggregate productivity.

The remainder of this paper is structured as follows. Section 1 describes the data and the institutional features of the Brazilian bankruptcy reform. Section 2 develops the conceptual framework that guides our empirical work. Section 3 describes our empirical strategy. Section 4 reports our main results and evaluates their robustness. Section 5 quantifies the effect on aggregate productivity. Section 6 concludes.

1 Institutional Setting and Data

1.1 The 2005 Bankruptcy Reform in Brazil

Our empirical strategy uses the 2005 Brazilian bankruptcy reform as a source of exogenous variation in the availability of credit to firms. In this section, we describe the key features of the reform and discuss how these changes resulted in increased access to corporate credit. For a thorough discussion of the changes implemented by the new bankruptcy legislation, see Araujo and Funchal (2005).

The bankruptcy legislation that came into effect in Brazil in 2005 was the most consequential reform to the country's insolvency procedures since 1945, when the previous insolvency statute was enacted. The pre-2005 legislation was considered punitive to creditors and was criticized for contributing to Brazilian interest rate spreads ranking among the highest in the world.[4] The main issues with the existing legislation were: (i) the bankruptcy priority rule, which prioritized both labor claims and tax claims before of creditors, and (ii) what is generally referred to "successor liability" (Araujo, Ferreira, and Funchal 2012). Successor liability meant that tax claims, labor claims, and all other liabilities were transfered to the buyer of an asset sold in liquidation which, according to anecdotal accounts, led to depressed market value of the pool of bankruptcy assets. These issues resulted in an estimated rate of recovery in the event of insolvency of about 0.2 percent in 2004, which is extremely low even in comparison with other Latin American countries (World Bank Doing Business database).

Efforts to reform Brazilian bankruptcy laws started in 1993, with the goals of making legislation more creditor-friendly and increasing the recovery rate of creditors. The reform was seen as a crucial step toward reducing bank spreads and increasing the volume of private credit to corporations.[5] After several amendments, the reform package was approved by the House of Representatives in October 2003 and by the Senate in December 2004. The approved bill was signed into law in February 2005 and became effective 120 days later.

In this paper, we focus on two key aspects of the new legislation that introduced changes to the liquidation procedure. First, secured creditors were given priority over tax claims in the bankruptcy priority rule. Second, tax claims, labor claims, and other liabilities were no longer transferred to the buyer of an asset sold in liquidation. Figure 1 shows the expected recovery rate estimated by the World Bank from 2004 to 2013. According to these estimates, the recovery rate increased sharply from about 0.2 in 2004 to 12.1 cents on the dollar in 2007, in line with what we would expect given the nature of the changes introduced by the new legislation.

[4] For instance, according to Paiva Muniz and Palhares Basilio (2005), "The inefficiency of [the prior] Brazilian insolvency rules ha[d] severe negative impacts on the economy, to the extent that they adversely affect[ed] the spread in the interest rates charged by financial institutions, which are among the highest in the world."

[5] For instance, this argument was made by the then Minister of Finance Antonio Palocci in his inauguration speech, in January 2003 (http://www1.folha.uol.com.br/folha/dinheiro/ult91u61397.shtml)

As a consequence of higher rates of recovery, we expect an increase in the availability of credit. In Panel A of Figure 2, we show that private credit expanded rapidly following the reform, from under 30 percent of GDP in 2004 to close to 63 percent in 2013. In all likelihood, this aggregate trend is partially attributable to a credit boom that was felt throughout the continent. But while Brazil was not the only Latin American country to experience a private credit expansion during the 2000s, Brazil's expansion seems to have surpassed those of other countries. We illustrate this in Panel B of Figure 2 by using data from Latin America to construct a synthetic control for the evolution of credit in Brazil, as in Abadie et al. (2010) and Abadie et al. (2015). This exercise suggests that at least some of this expansion is due to conditions that were specific to Brazil's credit market as opposed to trends common to all Latin American countries.

1.2 Data Sources

Our analysis combines credit registry data from the Central Bank of Brazil, matched employer-employee data from the Brazilian Ministry of Labor and Employment, and data on real outcomes from the Brazilian Statistics and Geography Institute (*Instituto Brasileiro de Geografia e Estatística*, or IBGE).

Credit registry data from 2003 onwards are available at quarterly frequency. This data contain time-invariant identifiers for each loan, bank, and firm, allowing us to track any corporate loan above 5,000 BRL granted by a financial institution operating in Brazil. This information is reported by banks to the Central Bank of Brazil and is of high quality because loan amounts reported to the credit registry must match banks' quarterly accounting figures. We keep all private firms in manufacturing and extracting sectors that were present in our sample prior to the 2005 bankruptcy reform and collapse data from 2003 to 2010 to the firm-quarter-year level.

Linked employer-employee data come from the *Relação Anual de Informações* (RAIS), a mandatory survey filled out annually by all tax-registered firms in Brazil. Incomplete or late information results in severe penalties, which leads to a high degree of compliance and essentially complete coverage of all employees in the Brazilian formal sector. The RAIS contains a time-invariant identifier for each worker as well as time-invariant firm identifiers. This allows us to link all workers to the firm that employs them and to follow a given worker over time. We observe data on average gross monthly earnings and average number of hours worked, as well as worker characteristics such as education, occupation, race, age, and gender. Using information on employees' start and termination dates, we convert this panel to quarterly frequency by assigning a worker-firm observation to a quarter if the worker was employed at the firm for at least one month during that quarter. We restrict our attention to full-time workers between the ages of 18 and 65 and use data from 2000 to 2010.

Finally, firm-level data on real outcomes come from the *Pesquisa Industrial Anual da Empresa* (PIA), which is based on annual surveys filled out by firms in the manufacturing and mining sector. The surveys are mandatory for all firms with 30 or more employees or above a certain threshold of revenues (300,000 USD in 2012), and there are fines for non-compliance. The PIA dataset also includes a random sample of firms with 5 to 29 employees, referred to as "sampling stratum" (*estrato amostrado*). We restrict our analysis to the universe of larger firms, which are sampled with a probability of one, because we are unable to follow firms over time in the sampling stratum or observe information such as the municipality where the firm is located or the firm's 4-digit industry classification. These data contain information on operational and non-operational costs, revenues, assets, and investments, as well as time invariant firm identifiers. We restrict our attention to private firms present in our sample prior to the 2005 bankruptcy reform and use data from 2000 to 2010.

Using firm identifiers, we merge the RAIS dataset with both the credit registry data and the PIA data, after collapsing the RAIS to the firm level. Unfortunately, we are unable to merge the PIA dataset with credit registry data as neither dataset is allowed to leave its institution of origin. Table 1 provides descriptive statistics of the final datasets, with credit registry data and matched employer-employee data at the firm-quarter-year level, and data on real outcomes at the firm-year level. We report statistics for employment and real outcomes from the PIA-RAIS dataset and for credit outcomes from the SCR-RAIS dataset. The average firm has 216 thousand BRL in bank debt and pays an annualized interest rate of 26 percent. The average firm in the PIA-RAIS dataset employs approximately 74 workers and has been in operation for 10 years.

2 Conceptual Framework

In this section, we introduce a model in which heterogeneous producers face constraints in their ability to borrow and technology is such that skilled labor is more complementary to capital than unskilled labor. The model features workers with idiosyncratic tastes for different workplaces, which allows firms to set wages with some degree of market power. Moreover, production functions have a nested CES form and feature capital-skill complementarity as in Krusell et al. (2000).

The main goal of the model is to shed light on how we should expect the 2005 bankruptcy reform to affect the employment and earnings of high- and low-skilled workers. We start by describing the model and our estimation procedure. Next, we discuss the effect of loosening credit constraints on firms' employment and investment decisions in the context of our model.

2.1 Model

2.1.1 Preferences and Technology

The model has two periods, $t = 0, 1$. There is a continuum of entrepreneurs indexed by their productivity Z and their initial wealth A. Productivity and initial wealth are distributed uniformly and independently across entrepreneurs.

Each entrepreneur i owns a private firm which uses K_i units of capital, S_i units of skilled labor, and N_i units of unskilled labor at $t = 0$ to produce Q_i units of the final good at $t = 1$ according to the following production technology

$$Q_i = F(Z_i, K_i, S_i, N_i) = Z_i \left(\nu N_i^{\sigma} + (1-\nu)(\tau K_i^{\rho} + (1-\tau)S_i^{\rho})^{\frac{\sigma}{\rho}}\right)^{\frac{1}{\sigma}}, \tag{1}$$

where $\nu,\ \sigma,\ \tau,\ \rho,\ \in\ (0,1)$. This production function is a version of the technology in Krusell et al. (2000) without capital differentiation and, as in Krusell et al. (2000), there is capital-skill complementarity as long as $\sigma > \rho$.

Firms are monopolistically competitive, and each firm faces an isoelastic demand curve with a common elasticity of demand $\varepsilon > 1$.

2.1.2 Labor supply

There is a unit mass of workers, which can be skilled (s) or unskilled (n). The supply of each type of labor to firm i is given by

$$S_i(w_{is}) = S(w_{is} - b_{is})^{\beta}$$
$$N_i(w_{in}) = N(w_{in} - b_{in})^{\beta},$$

where w_{is} and w_{in} denote the wage rates posted by firm i, b_{is} and b_{in} denotes the reservation wages below which no worker is willing to work at firm i, and S and N denote the mass of skilled and unskilled workers, respectively.

Differences in reservation wages are meant to capture firm-specific amenities, such as commuting time and workplace culture, as well as differences in the value attributed to leisure by skilled and unskilled workers. This differentiation endows entrepreneurs with monopsony power to set wages, as in Card et al. (2018) and Kline et al. (2017). We assume also that workers cannot save, so they live, in effect, hand-to-mouth.

2.1.3 Financial Markets

The only asset in this economy is productive capital. A perfectly-competitive financial intermediary collects deposits and rents out capital to entrepreneurs. The return on deposited assets is r and the break-even condition of the intermediary implies that the rental price of capital is $r + \delta$, where

δ is the rate at which capital depreciates.

The key friction is this market is limited enforcement. In period $t = 1$, an entrepreneur can steal a fraction $1 - \eta$ of rented capital K_i. As punishment, the entrepreneur would loose her wealth. The intermediary will then allow the entrepreneur to rent capital as long as the entrepreneur's incentive compatibility constraint is satisfied. This requires that

$$R_i K_i - (1 + r)(K_i - A_i) \geq R_i K_i - \eta K_i,$$

where R_i denotes the gross return to capital investment of entrepreneur i. This implies that an entrepreneur faces a collateral constraint given by

$$K_i \leq \lambda(r, \eta) A_i, \tag{2}$$

where

$$\lambda(r, \eta) \equiv \frac{1 + r}{1 + r - \eta}, \tag{3}$$

While simple, this formulation yields a tractable model of capital market imperfections that cause initial wealth to limit investment. Moreover, by varying η (and consequently λ), we are able to outline all degrees of capital-market efficiency. This formulation of a capital rental market in which entrepreneurs face collateral constraints is similar to that of Buera et al. (2013) and Moll (2014), and captures the intuition that the amount of capital available to an entrepreneur is limited by his personal assets (Kiyotaki and Moore 1997)

2.1.4 Firm Optimization

Each entrepreneur faces the following profit maximization problem, which will determine her factor demands

$$\begin{aligned} \max_{P_i, Q_i, K_i, w_{is}, w_{in}} \quad & P_i(Q_i) Q_i(Z_i, K_i, S_i, N_i) - w_{is} S_i(w_{is}) - w_{in} N_i(w_{in}) - (r + \delta) K_i \\ \text{s.t.} \quad & K_i \leq \lambda(r, \eta) A_i \quad [\chi_i], \end{aligned}$$

The first-order condition with respect to wage rates for an active entrepreneur (i.e., an entrepreneur with production Q_i greater than zero) implies the following wage rule for a worker of skill $j \in \{s, n\}$

$$w_{ij} = (1-\varphi)b_{ij} + \varphi \left(1 - \frac{1}{\varepsilon}\right) P_i \frac{\partial Q_i}{\partial S_i}, \tag{4}$$

where $\left(1 - \frac{1}{\varepsilon}\right) P_i \frac{\partial Q_i}{\partial S_i}$ is the marginal revenue product of a worker of skill $j \in \{s, n\}$ working at firm i and $\varphi = \frac{\beta}{1+\beta}$.

The first-order condition with respect to capital for an active entrepreneur is

$$\left(1 - \frac{1}{\varepsilon}\right) P_i \frac{\partial Q_i}{\partial K_i} = r + \delta + \chi_i, \tag{5}$$

where $\left(1 - \frac{1}{\varepsilon}\right) P_i \frac{\partial Q_i}{\partial K_i}$ is the marginal revenue productivity of capital, $r + \delta$ is the user cost of capital, χ_i is the Lagrange multiplier associated with the collateral constraint. We can see that the Lagrange multiplier introduces a wedge in the first-order condition of constrained firms (for whom $\chi_i > 0$), causing the marginal product of capital to exceed the user cost.

2.2 Estimation

The key determinants of the response of employment and earnings of high- and low-skilled workers to a loosening of credit constraints in this model are the set of production function parameters: the parameters governing income shares (μ and τ) and the parameters governing the elasticities of substitution between unskilled labor, capital, and skilled labor (σ and ρ). Accordingly, in order to use this framework to conduct a realistic assessment of the impact of credit constraints on the skill composition and the skill premium, we use our firm-level PIA-RAIS sample to estimate these parameters.

Our estimation procedure consists of two steps. In the first step, we estimate an approximation of the production function in Equation (1). Letting lower case variables represent logged upper case variables, a second-order approximation yields

$$q_i = \gamma_s s_i + \gamma_n n_i + \gamma_k k_i + \sum_{x \in \{s,n,k\}} \gamma_{xx} x_i^2 + \sum_{w \neq x} \sum_{x \in \{s,n,k\}} \gamma_{xw} x_i w_i + z_i, \tag{6}$$

where

$$\gamma_k = (1-\nu)\tau \tag{7}$$
$$\gamma_n = \nu \tag{8}$$
$$\gamma_s = (-1+\nu)(-1+\tau) \tag{9}$$
$$\gamma_{kk} = -\frac{((-1+\nu)^2(-1+\sigma)\tau^2)}{2} + \frac{(-1+\nu)\tau(-\rho+\rho\tau-\sigma\tau)}{2} \tag{10}$$
$$\gamma_{nn} = \frac{\nu^2+\nu\sigma-\nu^2\sigma}{2} \tag{11}$$
$$\gamma_{ss} = \frac{(-1+\nu)^2(-1+1/\sigma)(-\sigma+\sigma\tau)^2}{2\sigma} - \frac{(-1+\nu)(\sigma^2+\rho\sigma\tau-2\sigma^2\tau-\rho\sigma\tau^2+\sigma^2\tau^2)}{2\sigma} \tag{12}$$
$$\gamma_{kn} = (-1+\nu)\nu(-1+\sigma)\tau \tag{13}$$
$$\gamma_{ks} = (-1+\nu)\tau(-1+\nu+\rho-\nu\sigma+\tau-\nu\tau-\rho\tau+\nu\sigma\tau) \tag{14}$$
$$\gamma_{sn} = (1-\nu)\nu(-1+\sigma)(-1+\tau)), \tag{15}$$

Our preferred method follows De Loecker and Warzynski (2012) and estimates Equation (6) separately for each 2-digit industry relying on proxy methods developed by Olley and Pakes (1996), Levinsohn and Petrin (2003), and Ackerberg et al. (2015) in order to control for unobserved productivity shocks, which are potentially correlated with input choices. More specifically, we proxy for productivity using the demand for materials and follow Ackerberg et al. (2015) in estimating all production function parameters using second-stage moments. From this step, we obtain estimates of the coefficients in Equation (6) as well as estimates of markups. We discuss the details of this production function estimation procedure and of how we obtain estimates of markups in Appendix A.

In a second step, we use reduced-form estimates of the coefficients in Equation (6) to recover the structural parameters of Equation (1) using a minimum distance estimation procedure. Let $\theta = \{\mu,\ \tau,\ \sigma,\ \rho\}$ represent the vector of structural parameters and $\gamma = h(\theta)$ represent the nonlinear system of Equations (7)–(15). With an estimate $\hat{\gamma}$ of the coefficients in Equation (6) obtained in the first step of our estimation procedure, we compute an efficient minimum distance estimator of the vector of structural parameters θ by solving

$$\min_{\theta\in\Theta}\{\hat{\gamma}-h(\theta)\}'\hat{\Xi}^{-1}\{\hat{\gamma}-h(\theta)\}, \tag{16}$$

where $\hat{\Xi}$ is the variance-covariance matrix of the reduced-form coefficients obtained in the first step of our estimation procedure.

This two-step estimation procedure produces estimates of production function parameters for each 2-digit industry, which we summarize in Panel A of Table 9, and our average estimates are in line with previous findings in the literature (Krusell et al. 2000). Regarding the parameters governing the elasticities of substitution between inputs (σ and ρ), we find that $\sigma > \rho$ for all 2-digit industries, suggesting that all industries display some degree of capital-skill complementarity. This result is consistent with previous work that finds evidence of capital-skill complementarity for all industries in manufacturing using US data (Larrain 2015).

Panel B of Table 9 displays model parameters. We set production function parameters equal to the average across all 2-digit industry parameter estimates, and the elasticity of demand ε is set to match the average estimated markup. We set the interest rate r and the depreciation rate δ equal to the average borrowing and depreciation rates in the PIA-RAIS sample, respectively. Finally, we set the elasticity of labor supply $\beta_S = \beta_N = \beta$ to match an estimated pass-through of mean log value added per worker to log earnings of 0.16. In our model, this pass-through is equal to $\varphi\left(1 - \frac{1}{\varepsilon}\right)$, where $\varphi = \frac{\beta}{1+\beta}$. Given a value for the elasticity of demand of $\varepsilon = 5.26$, this implies $\beta \approx 0.25$.

2.3 The Effect of Loosening Credit Constraints

In Section 1.1, we argued that the 2005 bankruptcy reform increased the recovery rate of creditors. Through the lens of our model, this can be interpreted as an increase in the recovery rate η. From Equation (3), this implies an increase in the maximum leverage rate λ, i.e., a relaxation of the credit constraint modeled in Equation (2). For that reason, it will be useful to consider the implications of an increase in λ in the context of our model.

Intuitively, a constrained entrepreneur sees a direct increase in her ability to rent capital as a result of looser credit constraints. An increase in capital, in turn, raises the marginal productivity of labor, leading to an increase in the optimal level of total labor. Thus, when credit constraints are relaxed, constrained entrepreneurs increase their levels of capital and labor.

Next, let us consider the effect of loosening credit constraints on the ratio of skilled to unskilled workers (S_i/N_i) and on the skill premium (w_{is}/w_{in}). Under the assumption of capital-skill complementarity, the marginal productivity of skilled labor rises by more that the marginal productivity of unskilled labor when capital utilization increases. This causes the skill composition to shift toward skilled workers, and the ratio of skilled to unskilled workers to increase. Moreover, recall from Equation (4) that wages for a skill type at a given firm are a convex combination of the reservation wage and the marginal productivity of that skill type. Hence, under capital-skill complementarity, the skill premium also rises when capital utilization increases, as skilled workers become relatively more productive.

We formalize this intuition regarding the effect of loosening credit constraints on firms' real decision in our model by conducting the following numerical exercise. After solving the model, we simulate a large number of entrepreneurs and, for each level of the the maximum leverage rate λ, compute the average difference in the amount of capital held by constrained and unconstrained entrepreneurs. We define constrained entrepreneurs as those entrepreneurs for whom the collateral constraint holds with equality when $\lambda = 1$, i.e. when there is no borrowing. Results of this simulation are qualitatively identical if we define constrained entrepreneurs as those entrepreneurs for whom the collateral constraint holds with equality for other values of λ. We repeat this exercise for overall employment, for the ratio of skilled to unskilled workers, and for the skill premium, and show results in Figure 3.

As we would expect, the difference between the amount of capital held by constrained and unconstrained entrepreneurs rises as constraints are loosened because constrained entrepreneurs are able to increase their level of investment. The increase in capital leads to an increase in the marginal productivity of labor, causing constrained entrepreneurs to also increase employment relative to unconstrained entrepreneurs. Moreover, under capital-skill complementarity, the productivity of skilled workers rises by more than the productivity of unskilled workers when the level of capital increases. This leads to an increase in both the average ratio of skilled workers and the average skill premium for constrained entrepreneurs relative to unconstrained entrepreneurs.

3 Research Design

Our identification strategy uses the 2005 Brazilian bankruptcy reform as exogenous variation in the recovery rate of lenders and, consequently, in the availability of credit to firms. In order to identify the causal effect of increased access to credit, we exploit cross-sectional variation in financial constraints prior to the reform. In this section, we first discuss our empirical strategy and then describe how we measure financial constraints in the data.

3.1 Empirical Strategy

We estimate the effect of increased access to credit using the 2005 Brazilian bankruptcy reform as a quasi-natural experiment and employing a difference-in-differences research design, in which we compare outcomes for financially constrained firms (the “treatment” group) and unconstrained firms (the “control” group), before and after the reform. The framework of Section 2 implies that an increase in the recovery rate of creditors should lead to looser credit constraints, directly increasing the borrowing capacity of financially constrained firms. Hence, financially constrained firms should be differentially exposed to the increase in recovery rates triggered by the bankruptcy reform.

The role of the control group is to provide a counter factual of what would have happened to firms' outcomes if this legislation had not been implemented. Accordingly, the identifying assumption is that, in the absence of the bankruptcy reform, outcomes for constrained and unconstrained firms would have maintained parallel trends. Our main approach to assess the validity of this assumption is to examine outcomes for firms in the treatment and control groups in the pre-reform period. As we discuss in Section 4, our estimates show that outcomes for the treatment and control groups move in close parallel prior to the reform. We take these results as evidence that our control group establishes an accurate counter factual for what would have happened to the treatment group in the absence of the reform.

Our baseline specification consists of a difference-in-differences specification of the form

$$\log(Y_{it}) = \alpha + \beta_2 Reform_t \times Q2_i + \beta_3 Reform_t \times Q3_i + \beta_4 Reform_t \times Q4_i + \kappa_i + \theta_t + \epsilon_{it}, \quad (17)$$

where Y_{it} is the outcome of interest for firm i at time t; $Reform_t$ is a dummy that equals 0 prior to the reform and 1 after the reform, Qk_i is an indicator for firm i being in the k^{th} quartile of our measure of financial constraints in the pre-reform period; κ_i is a vector of firm fixed effects; θ_t is a vector of time fixed effects.

Our coefficients of interest β_k, with $k \in \{2, 3, 4\}$, represent the average within-firm change in our outcome variables for firms in k^{th} quartile of our measure of financial constraints relative to firms in the first quartile, following the 2005 bankruptcy reform. We opt for this specification as our baseline due to its ability to capture nonlinearities in the treatment effect in a transparent way. Our results are qualitatively identical if we sort firms in two groups according to their position relative to the median or if we allow for a continuous treatment effect.

In order to provide evidence in favor of the parallel trends assumption discussed above, we also estimate equations of the following form

$$\log(Y_{it}) = \alpha + \sum_{\tau \in \mathcal{T}} \beta_{2\tau} I(\tau) \times Q2_i + \sum_{\tau \in \mathcal{T}} \beta_{3\tau} I(\tau) \times Q3_i + \sum_{\tau \in \mathcal{T}} \beta_{4\tau} I(\tau) \times Q4_i + \kappa_i + \theta_t + \epsilon_{it}, \quad (18)$$

where $I(\tau)$ is a dummy equal to one exactly τ years after (or before if τ is negative) the reform.

3.2 Measuring Financial Constraints

A large body of work in corporate finance studies how frictions in the process of raising external funding can generate financial constraints for firms. As financial constraints are hard to directly observe, many possible measures of constraints have been proposed, and there is considerable disagreement regarding the merits of each approach.

In this paper, we follow an emerging branch of literature that directly estimates gaps between the marginal revenue product of capital (MRPK) and the user cost of capital. Intuitively, in the absence of frictions, a firm will keep adding units of a given input until the marginal product equals the marginal cost of this input. By estimating wedges between marginal product and marginal cost, we can construct a flexible measure of deviations from what we would expect in the absence of frictions. In practice, MRPK-cost wedges have been found to be closely related to the severity of credit-market frictions (Lenzu and Manaresi 2018). Moreover, MRPK-cost wedges can be constructed for private firms, unlike measures that rely on market value (Kaplan and Zingales 1997, Whited and Wu 2006). This feature is of particular importance in the context of our study given the relatively small number of listed firms in Brazil. In 2005, the year the bankruptcy reform was enacted, Brazil had 342 listed firms while the US had 5,145.

In order to estimate MRPK-cost wedges, we first need an estimate of marginal revenue products. Let us consider, as in the model of Section 2, a firm i that uses K_i units of capital, S_i units of skilled labor, and N_i units of unskilled labor to produce Q_i units of the final good according to a production technology $Q_i = F(Z_i, K_i, S_i, N_i)$, and that faces a downward-sloping demand curve with elasticity parameter ε.

Recall from Equation (5) that, under these assumptions, we can write the marginal revenue product of capital as

$$MPRK_{it} = \frac{\partial P_{it}(Q_{it})Q_{it}}{\partial K_i} = \left(1 - \frac{1}{\varepsilon}\right) P_i \frac{\partial Q_i}{\partial K_i}, \tag{19}$$

Define θ_i^k as the output elasticity of capital $\frac{\partial Q_i}{\partial K_i}\frac{K_i}{Q_i}$. We can rewrite Equation (19) as

$$MPRK_{it} = \theta_i^k \frac{P_{it}Q_{it}}{K_{it}} \left(1 - \frac{1}{\varepsilon}\right), \tag{20}$$

While we can directly observe the average revenue product of capital ($P_{it}Q_{it}/K_{it}$), an estimate of the marginal revenue product of capital requires estimates of output elasticities and markups, which we obtain from the production function estimation procedure described in Section 2.2 and further detailed in Appendix A. With estimates of markups and of the output elasticity of capital, we compute MRPK-cost wedges as

$$\text{MRPK-cost wedge}_{it} = \hat{\theta}_{it}^k \frac{P_{it}Q_{it}}{K_{it}} \frac{1}{\hat{\mu}_{it}} - R_{it}, \tag{21}$$

where $\hat{\theta}^k$ is the estimated output elasticity of capital, $P_{it}Q_{it}$ is deflated revenues, K_{jt} is the deflated book value of fixed assets, and $\hat{\mu}_{jt}$ is estimated markup. We proxy for the user cost of capital R_{jt} using the sum of the average borrowing rate at the 4-digit industry level and the firm-level depreciation rate of fixed assets. We use average borrowing rates at the industry level as we are unable to merge the PIA dataset with credit registry data at the firm level. Average rates are computed as weighted averages, with weights equal to the share of each firm's borrowing relative to the industry total loan amount. We measure depreciation rates at the firm level as total depreciation divided by the book value of fixed assets. We construct our baseline measure of pre-reform MRPK-cost wedges by computing firm-level averages of estimated wedges between 2000 and 2004—the years preceding the bankruptcy reform.

Table 2 reports results of this estimation. In Panel A, we report summary statistics for the user cost of capital, the marginal product of capital, and the MRPK-cost wedge. The average firm in our sample has an MRPK-cost wedge of 20 percentage points, and there is considerable dispersion. In Panel B, we display estimates of average output elasticities and markups, with block-bootstrapped standard errors in parenthesis. Our estimates are precise and in line with previous findings in the literature (De Loecker 2011; Petrin and Sivadasan 2013; Ackerberg et al. 2015; Lenzu and Manaresi 2018).

In Table 3, we report averages of key variables by quartile of MRPK-cost wedge. Firms with high MRPK-cost wedges appear to have characteristics that one would associate with financial constraints. For instance, firms with high MRPK-cost wedges are younger, smaller (both in terms of assets and in terms of number of employees), have lower capital-to-output ratios, and have lower industry-level debt-to-assets ratios. We report debt-to-assets ratio at the industry level because we are unable to merge the PIA dataset with credit registry data at the firm level. Instead, we merge loan amounts at the 4-digit industry level with the PIA dataset and compute debt-to-assets at the 4-digit industry level. These firms also pay lower wages, have a lower share of skilled workers, and have a lower skill premium. We further discuss these differences and their potential to shed light on the mechanism behind our results in Section 4.

In Section 4, we also consider the robustness of our results to alternative measures of financial constraints. Our results are qualitatively identical if we sort firms according to common proxies such as size and age, which have been found to be particularly successful in predicting financial constraints (Hadlock and Pierce 2010). Our results are also robust to sorting firms according to the ratio between estimated total factor productivity (TFP) and the book value of fixed assets, which is a predictor of financial constraints in models with heterogeneous producers and collateral constraints, such as the framework in Section 2.

4 Results

4.1 Evidence of Credit Expansion

We start by providing evidence that the 2005 bankruptcy reform led to an increase in credit for financially constrained firms relative to unconstrained firms. Establishing this result is crucial to our identification strategy because we use the bankruptcy reform as a quasi-natural experiment that led to increased credit availability to financially constrained firms. We do so by estimating Equation (17) with log bank credit as the dependent variable, using our firm-level SCR-RAIS sample. We define total bank credit as the sum of all existing loans granted by any financial institution to a given firm at a point in time. As we discuss in Section 1.2, we cannot merge credit registry data with the PIA dataset on real outcomes. For that reason, we exploit industry-level variation in our measure of MRPK-cost wedges for dependent variables relating to credit availability and interact the post-reform dummy (post-2005Q1) with an indicator for a firm's being in a 4-digit industry that is in the fourth quartile of our measure of financial constraints in the pre-reform period. This specification estimates how access to bank credit changed after the reform for firms in the fourth quartile of our financial constraint measure, relative to firms in the bottom quartiles. All specifications include firm fixed effects to account for any level differences between these groups of firms, as well as time fixed effects to flexibly control for any time trends common to all firms. Standard errors are clustered at the industry level throughout.

Column 1 of Table 4 shows estimation results for Equation (17). We see that bank credit expanded to constrained firms by approximately 7 percent relative to unconstrained firms following the reform. This result suggests that the 2005 bankruptcy reform led to increased access to credit for firms that were more financially constrained prior to the reform. Moreover, the timing of this effect is entirely consistent with the reform. We provide evidence for this by estimating a by-quarter version of Equation (17). This specification replaces the $Reform$ dummy with a dummy for each quarter-year, hence separately estimating the difference in bank debt for constrained and unconstrained firms at each time period. The omitted period is 2004Q3, two quarters prior to the reform, so effects can be interpreted relative to this period.

Figure 4 plots the coefficients of this regression model, along with 95 percent confidence intervals. We estimate a sizable and significant increase in bank debt for firms in constrained relative to unconstrained industries starting in the first quarter of 2005, which grows steadily over the next few years. Importantly, our coefficient estimates are close to zero and statistically insignificant in the period preceding the reform. This implies that our estimated treatment effect is consistent with the timing of the reform at the quarterly level and, in particular, we find no evidence of pre-existing trends.

Column 2 of Table 4 two shows results are robust to adding municipality$\times$time fixed effects in order to flexibly control for any differential local trends. The specification in the third and final column of Table 4 includes two-digit-industry$\times$time fixed effects to allow for differing trends by industry, and we find similar results as in our baseline specification. This implies that we observe a similar increase in credit availability to financially constrained firms following the reform even when comparing firms producing similar products or firms in narrowly defined geographical regions. We take this as evidence that our results are not driven by differing local- or industry-specific trends.

4.2 Effect on Skill Composition and on Skill Premium

In this section, we present and discuss our key empirical findings. We show that constrained firms experience a shift in the skill composition of their workers and in the return to skill following the reform. In Table 5, we present estimates of Equation (17) for outcomes relating to the employment of skilled workers, using our firm-level PIA-RAIS sample. This equation compares outcomes for firms in each of the top three quartiles in our measure of financial constraints with those of firms in the bottom quartile, before and after the reform. All models include firm fixed effects to account for any level differences between these groups of firms, as well as time fixed effects to flexibly control for any time trends common to all firms. Standard errors are clustered at the industry level.

In column 1, we document an increase in the ratio of skilled to unskilled employment, where workers are considered skilled if they possess at least some post-secondary education and unskilled otherwise. Firms in the fourth quartile of MRPK-cost wedges experienced an 8.5 percentage point increase in the ratio of skilled to unskilled workers after the reform relative to firms in the first quartile. While we document in subsequent sections that overall employment also increases as a consequence of increased borrowing capacity, this results speaks to the characteristics of the workers that a firm employs and to how these characteristics change when financial constraints are relaxed.

These estimates suggest that, following the increase in credit availability, financially constrained firms increased their employment of skilled workers, relative to the employment of unskilled workers. Moreover, these results are economically significant: an 8.5 percentage point increase in the ratio of skilled to unskilled workers implies that firms in the fourth quartile of the MRPK-cost wedge distribution close approximately 15 percent of the gap in skill composition between them and firms in the first quartile.

We also observe an increase in employment in occupations that are traditionally performed by skilled workers. Consistent with classifications based on the International Standard Classification of Occupations, we classify legislators, senior officials and managers as "managers" and science, engineering, health, teaching, and business professionals and associate professionals and technicians as "professionals." In columns 2 and 3 of Table 5, we see that the share of managers and professionals also increases for constrained relative to unconstrained firms following the reform. Firms in the fourth quartile of MRPK-cost wedges saw an increase of 4.5 percent in the share of managers and of 7.6 percent in the share of professionals and technicians after the reform relative to firms in the first quartile. This is further evidence that constrained firms were better positioned to hire and retain skilled workers as a consequence of increased access to credit.

Column 4 of Table 5 documents a rise in the skill premium following the reform, with firms in the fourth quartile of our measure of financial constraints experiencing a 10 percent increase relative to firms in the first quartile. This result suggests that increased access to credit leads to a higher within-firm return to skill. This effect is not only highly statistically significant, but it is also economically meaningful. Firms in the fourth quartile of the MRPK-cost wedge distribution pay a premium to skill about 17 percent lower than the skill premium paid by firms in the first quartile. Hence, our estimated change in the skill premium of firms in the fourth quartile of our constraint measure following the reform is approximately 60 percent of the gap in skill premium between them and firms in the first quartile.

Taken together, these findings suggest that access to credit allows financially constrained firms to hire and retain relatively more skilled workers and that these firms do so by making their return to skill more comparable to that of unconstrained firms. This is of particular importance given that the ability to hire and retain skilled labor has been show to meaningfully impact firm-level productivity (Bloom et al. 2013; Bhattacharya et al. 2013). Our findings corroborate this existing body of work as we also find, as we discuss in a subsequent section, that constrained firms see an increase in measures of efficiency such as value added per worker.

Finally, in Panel A of Figure 5, we show coefficient estimates and 95 percent confidence intervals for Equation (18), which separately estimates the difference in the ratio of skilled to unskilled for constrained and unconstrained firms at each time period, by replacing the $Reform$ dummy with a dummy for each quarter-year. The omitted period in these specifications is 2004Q3, two quarters prior to the reform, so effects can be interpreted relative to this period.

We find that the timing of the effect is consistent with the reform at the quarterly level. The estimates indicate that, for firms in the fourth quartile of our financial constraint measure, the skilled-to-unskilled ratio for rises by 2.6, 3.1, 5.2, and 7.6 percent, in the first, second, third, and fourth quarters of 2005, respectively. Our points estimates continue to rise modestly, stabilizing around the end of 2006. Importantly, our estimates are close to zero and statistically insignificant prior to the reform, showing no evidence of pre-existing trends. We show analogous results with the skill premium as the outcome variable in Panel B of Figure 5 and again find no evidence of pre-trends.

4.3 Effect on Earnings

Next, we present evidence that the relaxation of credit constraints as a consequence of the 2005 bankruptcy reform led to increased earnings. In Table 6, we show estimates of Equation (17) with outcome variables relating to earnings using our firm-level PIA-RAIS sample. All models include firm fixed effects, which account for any level differences between these groups of firms, as well as time fixed effects to flexibly control for any time trends common to all firms. Standard errors are clustered at the industry level. This specification compares outcomes for firms in each of the top three quartiles in our measure of financial constraints with those of firms in the bottom quartile, before and after the reform. We start by documenting, in column 1 of Table 6, the total wage bill increase at constrained firms relative to unconstrained firms following the reform, which can be driven by an increase in the number of employees, an increase in average earnings, or both.

In column 2 of Table 6, we show that constrained firms indeed experienced an increase in earnings following the reform. Firms in the fourth quartile of MRPK-cost wedges saw an increase of 2.2 percent in average earnings following the reform relative to firms in the first quartile. The effect is, on average, stronger if we restrict our attention to workers who were matched with the firm prior to the 2005 bankruptcy reform, which we refer to as incumbents. In column 3 of Table 6, we show that constrained firms in the fourth quartile of MRPK-cost wedges experienced an increase of 4.2 percent in average earnings of incumbents following the reform relative to firms in the first quartile. The estimated rise in earnings is precisely estimated and economically significant—a 4.2 percent increase corresponds to about approximately 20 percent of the difference in average earnings between firms in the fourth quartile of the MRPK-cost wedge distribution and firms in the first quartile.

Finally, in column 4 of Table 6, we show that the increase in earnings is, on average, stronger still for incumbents who remained at the firm following the reform, which we refer to as "stayers." A "stayer" at time t is a worker that was matched with the firm prior to the reform and is still with the firm at time t. We are cautious in our interpretation of this estimate as the determination of which workers remain employed at a firm is endogenous. However, this result could be taken as evidence that the increase in earnings experienced by workers at constrained firms has an important firm-specific component and that workers are unable to fully carry this premium with them as they switch jobs.

In Panel C of Figure 5, we show coefficient estimates and 95 percent confidence intervals for Equation (18) with average earnings as the outcome variable. These specifications separately estimate the difference in the ratio of skilled to unskilled for constrained and unconstrained firms at each time period, by replacing the $Reform$ dummy with a dummy for each quarter-year. The omitted period in these specifications is 2004Q3, two quarters prior to the reform, so effects can be interpreted relative to this period. Again, we find no evidence of pre-trends, and the timing of the effect is consistent with the reform at the quarterly level. The estimates indicate that, for firms in the fourth quartile of our financial constraint measure, the skilled-to-unskilled ratio for rises by 2.5, 4.1, 6.0, and 7.2 percent, in the first, second, third, and fourth quarters of 2005, respectively.

4.4 Analyzing the Mechanism: Effect on Total Employment and Investment

To help shed light on the mechanism behind the effect of a relaxation of credit constraints on the skill composition and on earnings, we investigate the response of firms' employment and investment decisions to the reform. In Table 7, we present estimates of Equation (17) for employment outcomes using our firm-level PIA-RAIS sample. As before, these specifications compare outcomes for firms in each of the top three quartiles in our measure of financial constraints with those of firms in the bottom quartile, before and after the reform. All specifications include both firm fixed effects, to account for any level differences between these groups of firms, and time fixed effects, to control for any time trends common to all firms. Standard errors are clustered at the industry level.

Column 1 of Table 7 shows that constrained firms increased overall employment following the reform, driven by both an increase in hiring and a decrease in terminations (as shown in columns 2 and 3 of Table 7). We define employment as the the total number of full-time workers at a firm in a given moment in time, while hiring is defined as the number of workers hired in the last 4 quarters. We define "terminations" as all dissolutions of a worker-firm match in the last 4 quarters, including both instances in which a worker quit and instances in which a worker was fired. We find that firms in the fourth quartile of MRPK-cost wedges saw a 6.3 percent increase in overall employment after the reform—resulting from an 61.1 percent increase in hiring and a 52.1 percent decrease in terminations—relative to firms in the first quartile.[6]

These results suggest that increased access to credit causes firms to employ more workers, and that this adjustment occurs both through increased hiring and through higher worker retention. Increased worker retention could be a consequence both of higher earnings, which potentially make the firm more appealing to workers, or of firms' being better equipped to weather temporary shocks without laying off workers so as to economize on firing, hiring, and training costs. The latter, a phenomenon known as "labor hoarding," has been found to be negatively impacted by financial constraints (Giroud and Mueller 2017). Our results are consistent with these findings, although we cannot distinguish between this channel and the potential effect of higher earnings on worker retention.

We also ask whether increased availability of credit allows credit-constrained firms to "poach" workers from unconstrained firms. In column 4 of Table 7, we show estimates of Equation (17) with the share of workers with prior experience in a firm with below-median MRPK-cost wedge as the outcome variable. We find that firms in the fourth quartile of MRPK-cost wedges saw a 3.9 percent increase in the share of workers with previous experience at an unconstrained firm after the reform relative to firms in the first quartile. This result lends support to the theory that the availability of credit impacts not only the number of workers, but also the type of workers that a firm is able to hire. It also provides direct evidence that the bankruptcy reform and subsequent credit expansion cause labor to be reallocated from less-financially-constrained firms to more-financially-constrained firms.

Next, we investigate the effect of looser credit constraints on investment decisions. In column 1 of Table 8, we show thar investment, measured as total capital expenditures scaled by lagged total assets, goes up by 6 percent for firms in the fourth quartile of MRPK-cost wedges following the reform relative to firms in the first quartile. In columns 2 and 3, we document that constrained firms also experience an increase in output and in value added per worker following the reform.

[6] Note that we measure employment as a stock, while new hires and terminations are flow variables. This explains why percentage changes in new hires and terminations are one order of magnitude larger than percentage changes in employment.

In Panels (d) and (e) of Figure 5, we show coefficient estimates and 95 percent confidence intervals for Equation (18) with employment and investment over assets as the outcome variable, respectively. These specifications separately estimate the difference in outcomes for constrained and unconstrained firms at each time period by replacing the $Reform$ dummy with a dummy for each quarter-year. The omitted period in these specifications is 2004Q3, two quarters prior to the reform, so effects can be interpreted relative to this period. As before, the timing of the effect is consistent with the introduction of the bankruptcy reform at the quarterly level, and we find no evidence of pre-existing trends.

Overall, these results suggest that financially constrained firms increase their levels of capital and employment as a response to increased access to credit. Along with the shift in the skill composition toward better-educated workers and the increase in the skill premium, these results are consistent with a production function featuring complementarities between capital and skilled labor.

We also provide direct evidence in support of the capital-skill complementarity hypotheses by showing that financially constrained firms in industries with high capital-skill complementary see a larger increase in the ratio of skilled-to-unskilled workers and in the skill premium following the bankruptcy reform relative to financially constrained firms in industries with low capital-skill complementary. To do so, we use the parameter estimates obtained from the estimation procedure described in Section 2.2 to compute the elasticity of substitution between capital and unskilled labor ($\varepsilon_{nk} = \frac{1}{1-\sigma}$) and between capital and skilled labor ($\varepsilon_{sk} = \frac{1}{1-\rho}$) for each 2-digit industry.

As we discuss in Section 2.2, we find that capital and skilled labor are relative complements in all industries, meaning that our estimates imply that $\sigma > \rho$ (and hence $\varepsilon_{nk} > \varepsilon_{sk}$). In order to sort industries according to the degree of capital-skill complementarity, we use the ratio between the two elasticities of substitution ($\frac{\varepsilon_{nk}}{\varepsilon_{sk}}$) as a measure. Industries with high $\frac{\varepsilon_{nk}}{\varepsilon_{sk}}$ are such that the elasticity of substitution between unskilled labor and capital is much higher than the elasticity of substitution between skilled labor and capital, meaning that capital is much more complementary to skilled labor than to unskilled labor.[7]

We sort firms into high and low complementarity according to this measure, and estimate the following equation

$$\log(Y_{ijt}) = \alpha + \beta Reform_t \times Constrained_i + \beta Reform_t \times Constrained_i \times HighCSC_i + \kappa_i + \theta_t + \epsilon_{ijt} \tag{22}$$

[7]The highest complementarity industries are Chemicals and Newspapers, Periodicals, and Book Publishers, while the lowest complementarity industries are Wood and Machinery and Equipment. This result is consistent with previous findings suggesting that Chemicals are among the manufacturing industries with the strongest complementarity, and that Wood and Machinery and Equipment have the weakest complementarity (Larrain 2015).

where Y_{ijt} is an outcome of firm i in industry j at time t; $Reform_t$ is a dummy that equals 0 prior to the reform and 1 after the reform, $Constrained_i$ is a dummy for firm i's having an above-median MRPK-cost wedge prior to the reform; $HighCSC_i$ is a dummy for firm i's being in an industry j above the median in our measure of capital-skill complementarity; κ_i is a vector of firm fixed effects; and θ_t is a vector of quarter-year fixed effects.

We show estimation results in Appendix Table B2. We can see from columns 1 and 2 that financially constrained firms in high-complementarity industries see the ratio of skilled-to-unskilled workers and the skill premium rise by an additional 6.0 and 7.1 percent following the bankruptcy reform relative to firms in low-complementarity industries. In column 3-4, we show that financially constrained firms in high-complementarity industries also see a larger increase in the employment and earnings following the reform relative to financially constrained firms in low-complementarity industries.

Finally, we document in column 5 that there is no significant difference between the investment response of financially constrained firms in high- and low-complementarity industries following the reform. This suggests that, given a similar increase in capital accumulation, constrained firms in high-complementarity industries increase their utilization of skilled labor and their return to skill by more than constrained firms in low-complementarity. These findings thus lend support to the theory that capital-skill complementarity is the mechanism behind the effect of a relaxation of credit constraints on the skill composition and on earnings in our setting.

4.5 Controlling for Industry- and Region-Specific Trends

One potential challenge to our identification strategy is that firms we categorize as financially constrained might systematically differ from unconstrained firms in terms of sectors or regions in which they operate, and that these sectors or regions might have experienced differential growth from the first quarter of 2005 onwards. This would be problematic as it would be consistent with the lack of pre-existing trends documented in Section 4, but would imply that something other than access to credit is the driving force behind our results.

In order to alleviate these concerns, we assess the robustness of our results to flexibly controlling for industry- and region-specific trends through the inclusion of fixed effects. In Appendix Table B3, we show that our results are robust to including either 4-digit-industry$\times$time or municipality$\times$time fixed effects to the specification in Equation (17). Columns 1, 3, 5, 7, and 9 include 4-digit-industry$\times$time fixed effects, while columns 2, 4, 6, 8, and 10 include municipality$\times$time fixed effects.

Let us first focus our attention to odd-numbered columns, which control for industry-specific trends. This specification compares outcomes for firms in each of the top three quartiles in our measure of financial constraints with those for the group of firms in the first quartile that are in the same 4-digit industry, before and after the reform. Estimates from this specification are qualitatively identical to our baseline results, thus providing strong evidence against the possibility that our results are driven by industry-specific trends.

Next we turn our attention to even-numbered columns of Appendix Table B3, which report estimates of Equation (17) including municipality×time fixed effects. This specification compares outcomes for firms in each of the top three quartiles in our measure of financial constraints with those for the group of firms in the first quartile that are in the municipality, before and after the reform. Currently, Brazil has over 5,500 municipalities, with the average municipality population close to 34,300. Hence, this specification compares constrained and unconstrained firms in narrowly defined regions, before and the reform. Estimates are again qualitatively identical to our baseline results, which we interpret as strong evidence against the theory that our results are driven by differences in local economic conditions.

4.6 Controlling for Exposure to the Business Cycle

It is also plausible that financially constrained firms might be differentially exposed to the business cycle. This would present a challenge to our identification strategy as it would imply that our estimates would be driven by a differential response to economic growth as opposed to increased access to credit. In order to determine whether this mechanism plays a role in explaining our results, we compute firm-specific GDP betas as

$$\text{GDP beta}_{it} = \frac{Cov(\Delta y_{it}, \Delta y_t)}{Var(\Delta y_t)}, \tag{23}$$

where Δy_{it} is log real revenue growth and Δy_{it} GDP growth.

We then use this proxy for firm-level exposure to the business cycle to control for systematic differences along this dimension between constrained and unconstrained firms. Specifically, we compute deciles of the distribution of GDP betas and estimate Equation (17), including beta-decile×time fixed effects.

This specification exploits variation across firms in the same decile of the distribution of GDP betas, which should at least partly control for firms' exposure to the business cycle. We show results of this exercise in Appendix Table B4 and confirm that our baseline estimates are relatively unchanged by the inclusion of these additional controls. We take this as evidence that differential exposure to the business cycle is not the driving force of our results.

4.7 Placebo Checks

Another form of assessing the plausibility of alternative explanations, such as the ones described above, is to conduct placebo checks. We do so by splitting our data into blocks of eight consecutive quarters and estimating the following specification for each block

$$\log(Y_{it}) = \alpha + \sum_{\tau\in\mathcal{T}} \beta_{2\tau} I(\tau) \times Q2_i + \sum_{\tau\in\mathcal{T}} \beta_{3\tau} I(\tau) \times Q3_i + \sum_{\tau\in\mathcal{T}} \beta_{4\tau} I(\tau) \times Q4_i + \kappa_i + \theta_t + \epsilon_{it}, \quad (24)$$

where Y_{it} is the outcome of interest for firm i at time t and Qk_i is an indicator for firm i being in the kth quartile of our measure of financial constraints in the pre-reform period, and $I(\tau)$ is a dummy equal to one exactly τ years after (or before if τ is negative) the placebo reform period, with $\tau \in \{-4, -3, \ldots, 3, 4\}$. We split quarterly data from 2000 until 2014 into seven blocks that do not contain the 2005 bankruptcy reform and one block that does contain the reform (from 2004Q1 until 2005Q4).

We show results of this exercise in Figure B1, in which we plot coefficients $\beta_{4\tau}$, with $\tau \in \{-4, -3, \ldots, 3, 4\}$, for each block of time. In each panel of Figure B1, we plot coefficient estimates for each block of data that does not contain the 2005 reform in dashed lines, and the estimates for the period ranging from 2004Q1 until 2005Q4, which contains the reform period, in a solid line. In each panel, we can see that placebo estimates fall far below estimates from the true reform period, lying outside the 95 percent confidence interval in the quarters following the reform.

Results from this exercise lend support to the robustness checks reported in previous sections, and help alleviate additional concerns such as the possibility that our results could be driven by life-cycle effects. We show that firms in the fourth quartiles of our measure of financial constraints are, in general, smaller and younger. Hence, because we sort firms at the beginning of our sample period and hold this classification fixed over time, one could argue that we should expect these firms to grow differentially due to their differing positions in their life cycle. This would be a violation of the parallel trends assumption and would present a challenge to our identification strategy.

However, under this theory, we would expect that estimates reported in Figure B1 for time periods that do not contain the 2005 bankruptcy reform would be similar to estimates obtained from the true reform period. Our findings are in direct opposition to this prediction, which we take as evidence that our results are not driven by life-cycle effects.

4.8 Alternative Definitions of Credit Constraints

Next we assess the robustness of our results to exploiting alternative measures of financial constraints. Hadlock and Pierce (2010) assess the informativeness of several measures of financial frictions and find that size and age are the most successful predictors of financial constraints. Specifically, financially constrained firms are, on average, smaller and younger than unconstrained firms. Based on these finding, we estimate the following specification

$$\log(Y_{it}) = \alpha + \beta Reform_t \times Q1_i + \kappa_i + \theta_t + \epsilon_{it}, \tag{25}$$

where Y_{it} is an outcome of firm i at time t; $Reform_t$ is a dummy that equals 0 prior to the reform and 1 after the reform, $Q1_i$ is either (i) a dummy for firm i's being in the first quartile of firm size prior to the reform or (ii) a dummy for firm i's being in the first quartile of the firm age distribution prior to the reform; κ_i is a vector of firm fixed effects; and θ_t is a vector of quarter-year fixed effects. We measure firm size as the number of employees because this measure is available in both the SCR-RAIS dataset and the PIA-RAIS dataset.

Panels A and B of Appendix Table B5 report the results of this exercise and we are reassured to find that our results are qualitatively identical if we sort firms along these two well-known predictors of financial constraints. Our results are also robust to sorting firms according to the ratio between estimated total factor productivity (TFP) and the book value of fixed assets, which is a predictor of financial constraints in models with heterogeneous producers and collateral constraints, such as the framework in Section 2. Specifically, financially constrained firms are more productive and have lower levels of assets than unconstrained firms, leading them to have a higher TFP-to-assets ratio.

We conduct this analysis by estimating the following specification

$$\log(Y_{it}) = \alpha + \beta Reform_t \times Q4_i + \kappa_i + \theta_t + \epsilon_{ijt}, \tag{26}$$

where Y_{it} is an outcome of firm i at time t; $Reform_t$ is a dummy that equals 0 prior to the reform and 1 after the reform, $Q4_i$ is a dummy for firm i's being in the fourth quartile of TFP-to-assets ratio; κ_i is a vector of firm fixed effects; and θ_{jt} is a vector of industry-time fixed effects.[8]

Panel C of Appendix Table B5 presents results with the TFP-to-assets ratio as the source of cross-sectional variation. We again find that results are qualitatively identical to our baseline estimates.

4.9 Alternative Sources of Variation

In this section, we assess the robustness of our results by exploiting particular features of the reform. Although the estimates presented in this section do not necessarily speak to the effect of access to credit on financially constrained firms, they should alleviate concerns that our results may be driven by some event contemporaneous to the reform that differentially impacted firms we classify as financially constrained.

[8] As in our main analysis, we are unable to exploit firm-level variation in TFP-to-assets ratio when estimating the effect on credit outcomes because we are unable to merge credit registry data with the PIA dataset. Thus, when the outcome variable is bank debt, we interact the $Reform_t$ dummy with a $Q4_{ij}$ dummy for firm i's being in a 4-digit industry j in the fourth quartile of TFP-to-assets ratio.

We focus on two key changes introduced by this legislation: (1) secured creditors were given priority over tax claims in the bankruptcy priority rule, and (2) tax claims, labor claims, and other liabilities were no longer transferred to the buyer of an asset sold in liquidation. The first change should lead to a finding that recovery rates went up by more for firms with a high tax burden. The second change should lead to a finding that recovery rates went up more for firms with more tangible assets, which can be sold in the event of liquidation.

As we do not observe these key dimensions at the firm level in our dataset, we turn to data from Economatica for publicly traded Brazilian firms. We proxy for a firm's tax burden by using the average ratio of tax expenses to earnings before income and taxes prior to 2005, and the average share of assets that are tangible prior to 2005 as a firm-level measure of tangibility. We conduct this analysis by estimating the following specification

$$\log(Y_{it}) = \alpha + \beta Reform_t \times Q4_i + \kappa_i + \theta_t + \epsilon_{it}, \tag{27}$$

where Y_{it} is either total debt or investment of firm i at time t; $Reform_t$ is a dummy that equals 0 prior to the reform and 1 after the reform, $Q4_i$ is either (i) a dummy for firm i's being in the fourth quartile of tax burden prior to the reform or (ii) a dummy for firm i's being in the fourth quartile of the share of tangible assets prior to the reform; κ_i is a vector of firm fixed effects; and θ_t is a vector of quarter-year fixed effects. We measure debt as total debt position minus trade credit. This measure will not only include bank debt, but also other debt instruments. Our measure of investment is given by capital expenditures, which we scale by total assets.

Results of this exercise are presented in Appendix Table B6. As we would expect as an outcome of the reform, debt increases more for firms with a high tax burden and asset tangibility. We find that firms in the fourth quartile of pre-reform tax ratio experienced a 10 percent increase in total debt and a 7 percent increase in investment following the reform, relative to firms in the first 3 quartiles. Similarly, firms in the fourth quartile of pre-reform share of tangible assets saw an 8 percent increase in total debt and a 5 percent increase in investment following the reform. These results are further evidence that the 2005 bankruptcy reform led to increased access to credit, which in turn resulted and higher rates of investment.

5 Quantifying the Effect on Aggregate Productivity

In this section, we use the model developed in Section 2 to quantify the effect of the credit expansion caused by the 2005 bankruptcy reform on aggregate productivity and compare the magnitude of this effect to the observed change in aggregate productivity during this period. We start by measuring the observed growth in aggregate productivity following Petrin and Levinsohn (2012). Next, we compute the model-implied change in aggregate productivity as a response to a relaxation in credit constraints calibrated to match the 2005 bankruptcy reform. We do so by applying the same growth accounting procedure to model-generated data.

5.1 Observed Aggregate Productivity Growth

As in Petrin and Levinsohn (2012), we define aggregate productivity growth (APG) as the change in aggregate final demand minus the change in aggregate expenditures in primary inputs. Following the framework of Section 2, we consider primary inputs to be capital (K), skilled labor (S), and unskilled labor (N). Letting Y_i represent the amount firm i's output that ultimately goes to final demand, we have that

$$APG = \sum_i P_i dY_i - \sum_i \sum_{X \in K,S,N} W_{ix} dX_i, \tag{28}$$

We follow Petrin and Levinsohn (2012) and decompose aggregate productivity growth into two components, the change in firm-level technical efficiency (Z) and the reallocation of inputs across firms, by totally differentiating output

$$APG = \sum_i D_i d \log Z_i + \sum_i D_i \sum_{X \in K,S,N,M} (\theta_i^x - s_i^x) d \log X_i, \tag{29}$$

where M represents intermediate materials, $D_i = \frac{P_i Q_i}{\sum_i VA_i}$ is the Domar weight of firm i, θ_i^x is the output elasticity of input X, and $s_i^x = \frac{W_{ix} X_i}{P_i Q_i}$ is the revenue share of input X, with $X \in \{K, S, N, M\}$.

The second term in this decomposition measures productivity gains from the reallocation of inputs across firms—the magnitude of which depends on how distorted the initial allocation is relative to a frictionless benchmark. As we can infer from the first-order condition in Equation (5), profit maximization implies that the output elasticity of an input will equal the input's revenue share ($\theta_i^x = s_i^x$) in the absence of distortions such as market power and financial constraints. In practice, however, we would expect such distortions to be present, implying that the reallocation term in Equation (29) is different than zero and $APG \neq \sum_i D_i d \log Z_i$.

In order to take Equation (29) to the data, we rely on the following discrete-time approximation, obtained by use of Tornquist-Divisia approximations

$$APG_t \approx \sum_i \bar{D}_{it} \Delta \log Z_{it} + \sum_i \bar{D}_{it} \sum_{X \in K,S,N,M} (\theta_{it}^x - s_{it}^x) \Delta \log X_{it}, \tag{30}$$

where bars denote averages across periods $t-1$ and t, and Δ is a first-difference operator from period $t-1$ to period t.

We divide our PIA-RAIS sample into two periods: 2000 to 2004 (before the bankruptcy reform) and 2005 to 2010 (after the bankruptcy reform). We take time-series averages of all relevant variables within these periods, so that variable X_{i1} is the average of X_{it} between 2000 and 2004, and X_{i2} is the average of X_{it} between 2005 and 2010. We measure value added as the difference between deflated revenue and deflated intermediate inputs, skilled labor as the number of workers with at least some college education, unskilled labor as the number of workers with no college education, capital as the deflated book value of fixed assets, and materials as deflated intermediate inputs. Our measures of firm-level productivity and output elasticities come from the production function estimation procedure described in Section 2.2 and further detailed in Appendix A.

Using Equation (30), we estimate aggregate productivity growth of 5.5 percent between the pre- and post-reform periods. We find that all of the increase in productivity comes from the reallocation of inputs, with the change in technical efficiency contributing a negative 2.9 percent to aggregate productivity growth. This result is consistent with previous findings of declining technical efficiency in manufacturing and extraction sectors in Brazil during this period (De Negri and Cavalcante 2014; Miguez and Moraes 2014).

5.2 Model-Implied Aggregate Productivity Growth

Next, we compute the model-implied aggregate productivity growth as a result of the bankruptcy reform. To do so, we must first map the 2005 bankruptcy reform to our theoretical framework. As we discuss in Section 2, a natural way to interpret the reform through the lens of our model is as an increase in the recovery rate η. We can then use Equation (3) to translate an increase in the recovery rate to an increase in the maximum leverage rate λ—a relaxation of the credit constraint modeled in Equation (2). Guided by the rise in the estimated recovery rate shown in Figure 1, we interpret the 2005 bankruptcy reform as an unanticipated increase in the recovery rate η from 0.002 to 0.15, which implies an increase in the maximum leverage rate λ from 1.002 to 1.135, given the parameters in Table 9.

Recall from Section 2 that each firm has two state variables: productivity and initial wealth. For each firm, we take productivity to be the time-series average of estimated productivity in the post-reform period and initial wealth to be the time-series average of the book value of fixed assets in the pre-reform period. Using policy functions implied by maximum leverage rate λ of 1.002, we compute optimal levels of output, capital, and skilled and unskilled labor for each firm, which we take as that firm's outcome in a counter-factual scenario in which there was no reform. To obtain the change in output and input usages attributable to the bankruptcy reform, we use the same values for each firm's state variables, but consider a maximum leverage rate λ of 1.135.

We then use Equation (30) to aggregate firm-level changes, with $\Delta \log X_i$ being the change in the usage of (logged) input X by firm i when we raise the maximum leverage rate. Note that the model abstracts from intermediate input usage, which means that we set $M = 0$ for all firms and periods and equate value added with $P_i Q_i$. This implies that weights become $D_i = \frac{P_i Q_i}{\sum_i P_i Q_i}$ and sum to one, unlike the Domar weights in Equation (30), which sum to more than one.

This exercise implies a 2 percent increase in aggregate productivity, accounting for 36 percent of observed aggregate productivity growth. Note that technical efficiency is constant by construction in this comparative statics, and hence all of the productivity increase comes from the reallocation of inputs across firms. This exercise suggests that the reallocation of inputs triggered by the 2005 bankruptcy reform led to sizable aggregate productivity gains.

6 Conclusion

In this paper, we investigate the effect of increased access to bank credit on the employment and earnings of high- and low-skilled workers. Our comprehensive dataset provides information not only on bank lending, output, investment, employment, and earnings, but also on detailed characteristics of workers—such as education, occupation, and previous experience. Our identification strategy exploits a considerable reform to bankruptcy legislation undertaken by Brazil in 2005 that strengthened creditor rights and led to an increase in the borrowing capacity of financially constrained firms relative to unconstrained firms. We show that the credit expansion resulting from the reform led constrained firms to increase employment, especially of high-skilled workers. We also observe an increase in earnings, especially for skilled workers and for workers who were matched with constrained firms prior to the reform. We also quantify the aggregate productivity gains from the bankruptcy reform and subsequent credit expansion.

Taken together, our results suggest that increased access to bank credit impacts not only investment and total employment, but also the type of worker a firm employs, in terms of both skill and previous experience. We establish a credible causal link between access to credit and a firm's utilization of skilled labor, providing new evidence on the specific channels through which financial development can improve the allocation of production factors and increase aggregate productivity. In addition, we show that financial development, in the form of increased access to bank credit, impacts within-firm earnings inequality through its effect on the return to skill.

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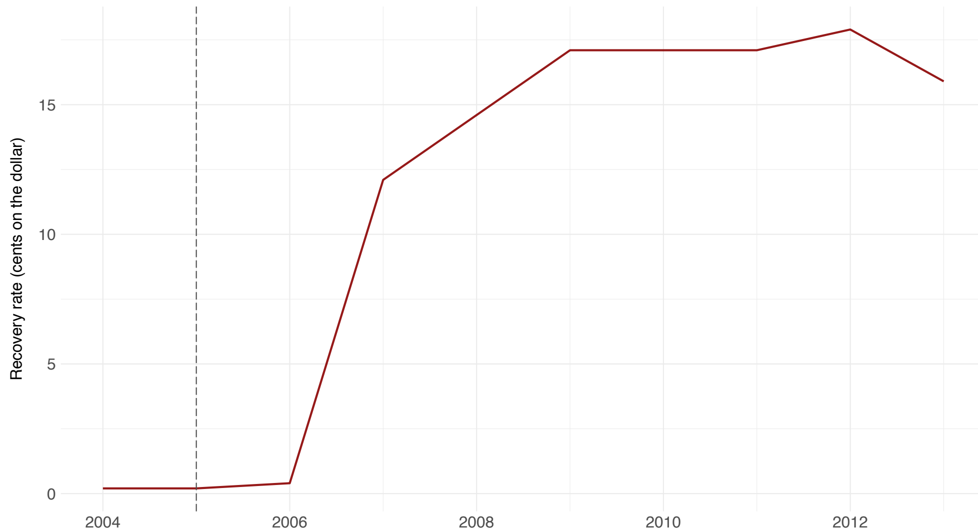
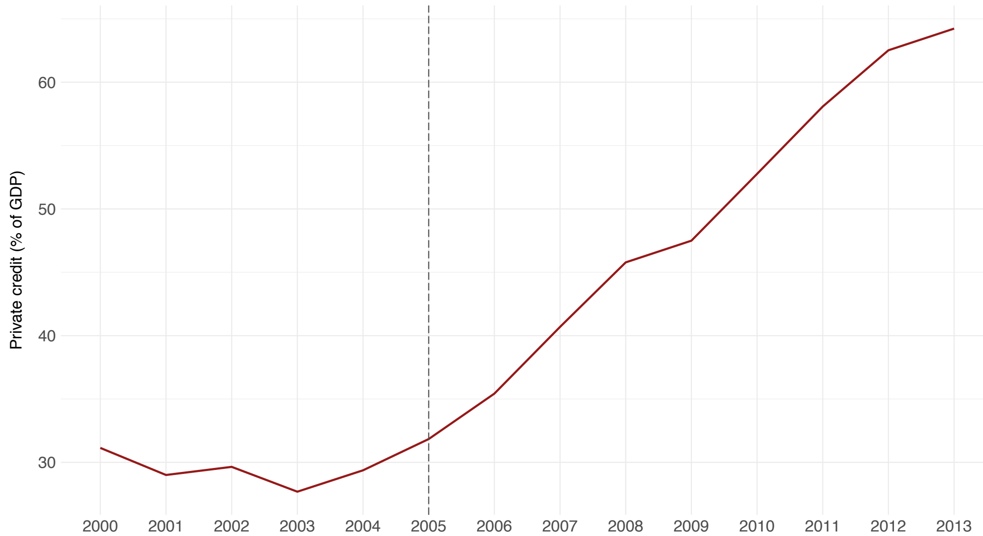
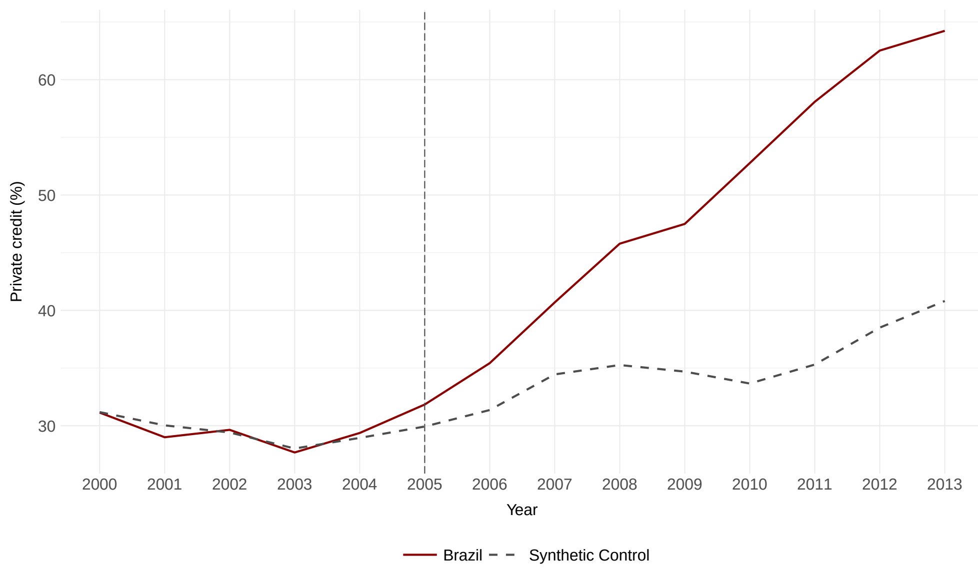


Figure 1: Expected recovery rate of secured creditors (cents on the dollar)

This Figure shows the expected recovery rate for secured creditors in Brazil. Data comes from World Bank's Doing Business database.



(a) Brazil



(b) Comparison with Latin America

Figure 2: Private Credit as Percentage of GDP (%)

This Figure shows private credit as percentage of GDP. In panel A, we plot private credit as percentage of GDP in Brazil. In panel B, we superimpose the evolution of a synthetic control for private credit in Brazil, constructed using data from Latin America and following the methodology of Abadie et al. (2010) and Abadie et al. (2015), and the evolution of private credit for a subsample of countries used in the construction on the synthetic control. Data on private credit comes from the IMF.

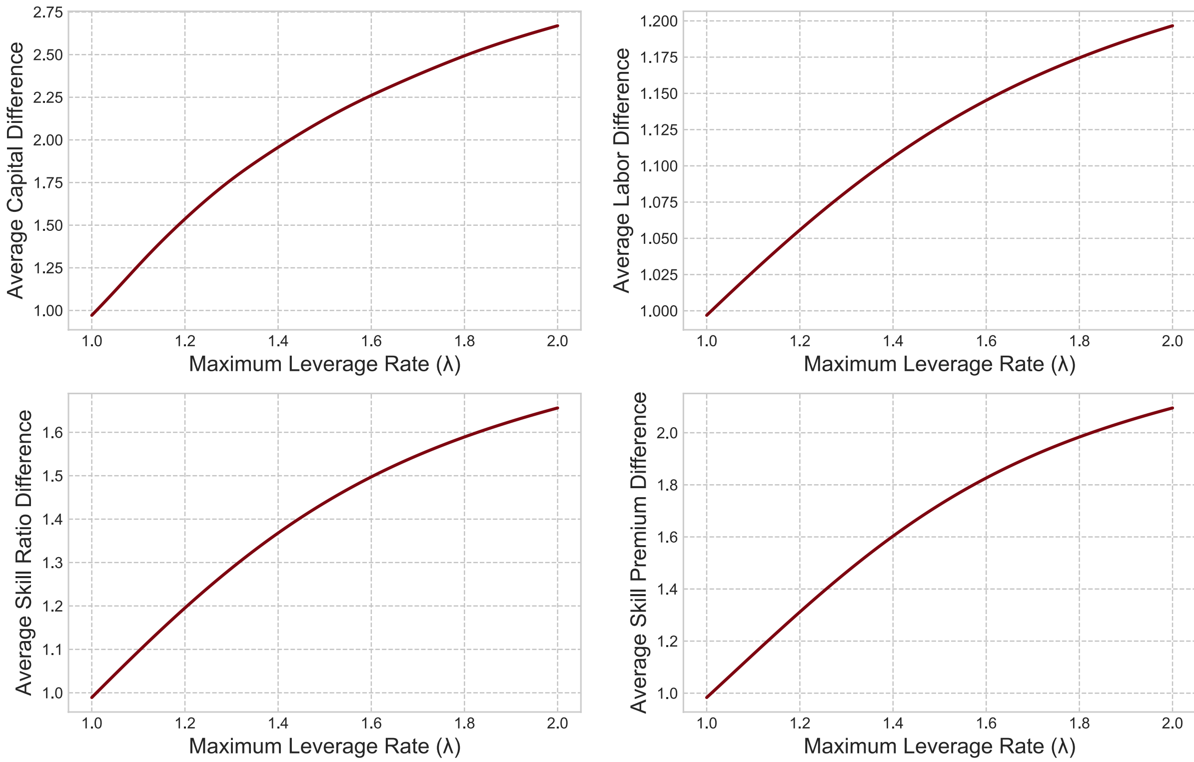


Figure 3: The Effect of Loosening Credit Constraints

This Figure shows the average difference between capital, total, employment, the ratio of skilled workers, and the skill premium for constrained and unconstrained firms, as a function of the maximum leverage λ. Results are from a simulation of 10,000 entrepreneurs. For each variable, we plot the difference between the the average value for constrained entrepreneurs minus the average value for unconstrained entrepreneurs, normalizing this difference to 1 when $\lambda = 1$. We define constrained entrepreneurs as those entrepreneurs for whom the collateral constraint holds with equality when $\lambda = 1$, i.e. when there is no borrowing.

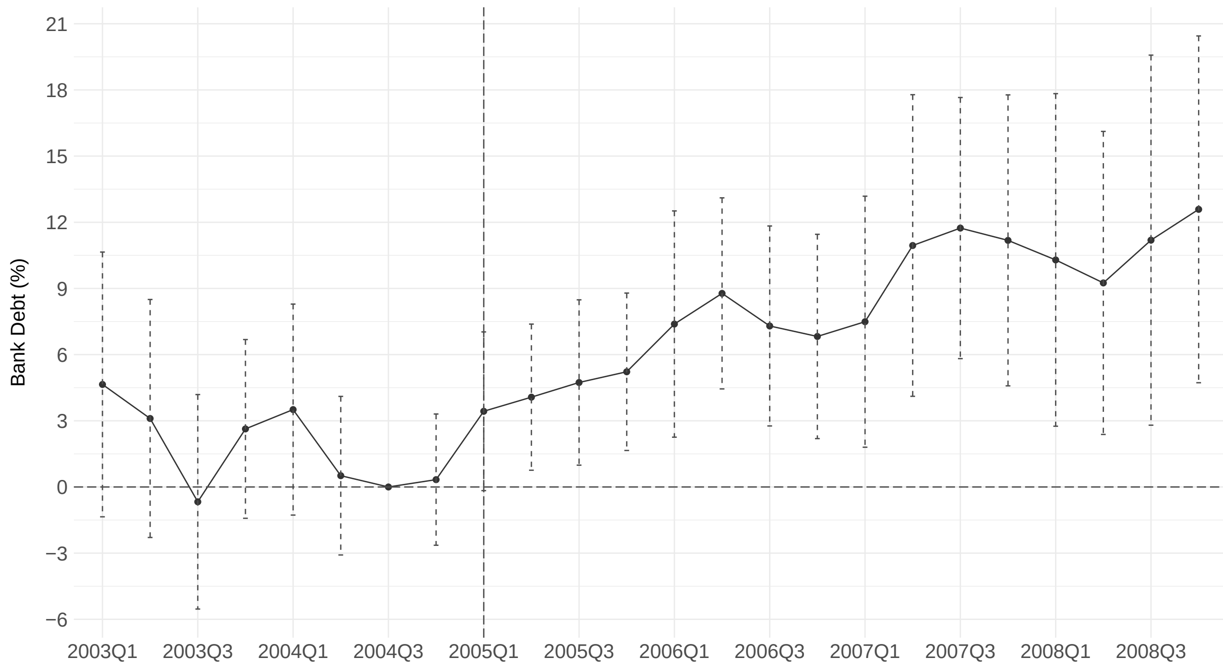
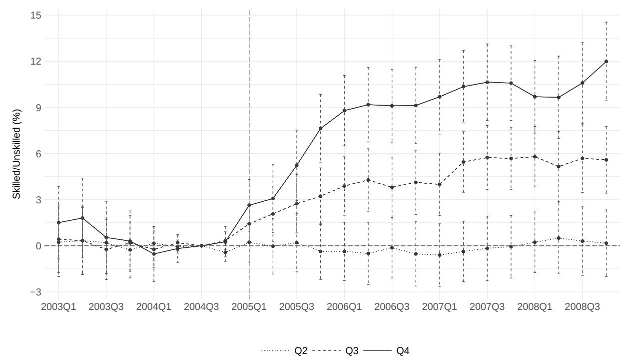
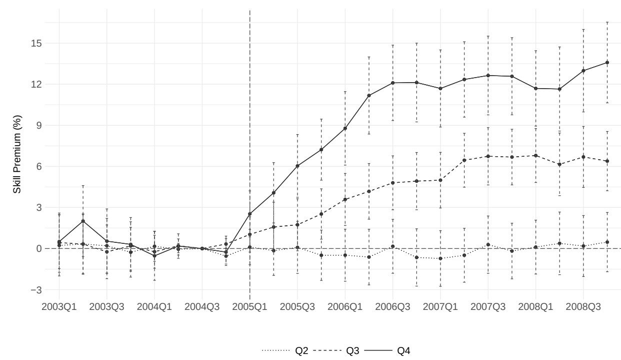


Figure 4: Timing of Effect on Bank Debt

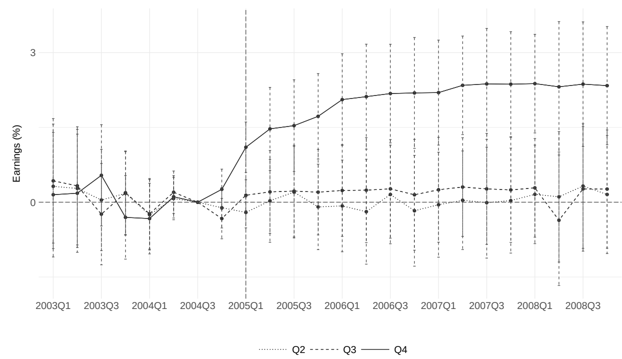
This Figure shows the timing of the effect of the 2005 bankruptcy reform on bank credit. We plot coefficient estimates from specification (**??**) along with 95 percent confidence intervals are plotted. Observation is at the firm-quarter level and standard errors are clustered at the industry level.



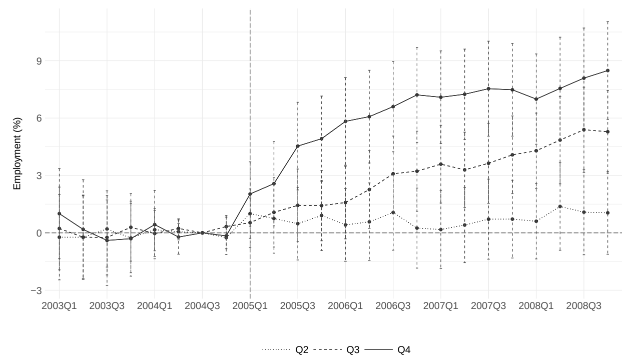
(a) Skill Ratio



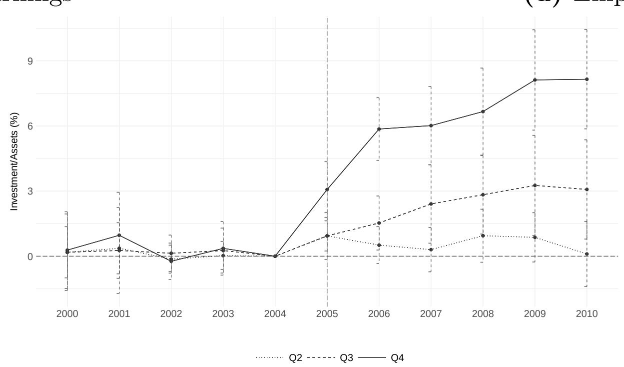
(b) Skill Premium



(c) Earnings



(d) Employment



(e) Investment/Assets

Figure 5: Timing of Effect on Skill Composition, Earnings, Employment, and Investment

This Figure shows the timing of the effect of the 2005 bankruptcy reform on the ratio of skilled to unskilled workers (panel A), on the skill premium (panel B), on earnings (panel C), on employment (panel D), and on investment over assets (panel E). We plot coefficient estimates from specification (18) along with 95 percent confidence intervals are plotted. Observation is at the firm-quarter level and standard errors are clustered at the industry level.

Table 1: Summary Statistics

	Mean	Med.	St. Dev.	N
Total bank debt	216.74	40.84	4,410.01	2,907,501
Number of loans	9.82	6.00	10.29	2,907,501
Interest rate	25.94	20.07	20.33	2,147,499
Firm age	9.58	7.10	8.55	944,656
Number of workers	73.61	30.00	137.55	944,656
Skilled-to-unskilled ratio	0.14	0.04	0.43	944,656
Share managers	0.03	0.00	0.05	944,656
Share professionals	0.07	0.01	0.16	944,656
Average monthly earnings	0.88	0.62	0.65	944,656
Average skill premium	2.23	1.96	1.45	944,656
Investment	906.78	10.85	3,249.09	227,920
Investment/assets	0.05	0.02	0.09	227,920
Assets	12,009.94	2,169.54	22,054.36	227,920
Output	14,351.56	3,709.67	23,697.11	227,920
Capital/output	0.64	0.55	0.57	227,920
Value added	4,040.93	1,101.92	6,483.41	227,920
Value added per worker	71.12	37.88	209.18	227,920
Industry debt/assets	0.10	0.07	0.04	227,920

Notes: This Table shows descriptive statistics for firms in our sample, with credit registry data and matched employer-employee data at the firm-quarter-year level, and data on real outcomes at the firm-year level. We restrict our attention to private firms present in our sample prior to the 2005 bankruptcy reform. We report statistics for employment and real outcomes from the PIA-RAIS dataset and for credit outcomes from the SCR-RAIS dataset. Monetary values are in thousands of 2003 BRL.

Table 2: MRPK-Wedges and Elasticities

Panel A: MRPK-wedge summary statistics

	Mean	St. Dev.	10^{th}	50^{th}	90^{th}
$r+\delta$	31.24	9.99	20.28	29.58	44.57
MRPK	51.98	58.00	3.96	19.94	89.57
MRPK-cost wedge	19.93	45.59	-29.12	-9.87	53.63

Panel B: Estimated elasticities and markups

	θ^K	θ^S	θ^N	θ^M	μ
	0.09	0.14	0.10	0.68	1.23
	(0.003)	(0.005)	(0.003)	(0.008)	(0.006)

Notes: This Table reports summary statistics for the MRPK-cost wedge, as well as estimates of production function parameters and markups. Panel A displays summary statistics for the marginal revenue product of capital (MRPK), the user cost of capital ($r+\delta$), and the wedge between the marginal revenue product of capital and its user cost. The first four columns of panel B report the average (standard errors of the mean in parenthesis) output elasticity with respect to capital (θ^K), skilled labor (θ^S), unskilled labor (θ^N), and materials (θ^M), while the last column reports the average (standard errors of the mean in parenthesis) markup (μ).

Table 3: Summary Statistics by quartile of MRPK-cost wedge

	Most constrained			Least constrained	
	Q4	Q3	Q2	Q1	(Q4 - Q1)
Firm age	7.00	9.50	10.39	11.43	-4.42***
Number of workers	52.38	79.70	80.10	82.27	-29.89***
Assets	7,626.50	12,303.62	13,804.50	14,303.34	-6,676.84***
Investment/assets	0.06	0.05	0.05	0.05	0.01***
Capital/output	0.45	0.65	0.80	0.83	-0.38***
Industry debt/assets	0.08	0.10	0.10	0.11	-0.03***
Skilled-to-unskilled ratio	0.10	0.14	0.16	0.16	-0.06***
Average monthly earnings	0.78	0.91	0.92	0.93	-0.15***
Average skill premium	1.98	2.29	2.34	2.33	-0.35***

Notes: This Table shows descriptive statistics for firms in our sample, sorted by quartile of average MRPK-cost wedge prior to the 2005 bankruptcy reform. Monetary values are in thousands of 2003 BRL. The last column reports the difference between the fourth and the first quartile. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 4: Effect of Bankruptcy Reform on Bank Debt

Dependent Variable:	Bank Debt		
	(1)	(2)	(3)
Reform×Q4	**0.069**	**0.070**	**0.055**
	(0.02)***	(0.02)***	(0.02)***
Firm FE	x	x	x
Quarter-Year FE	x		
Municipality×Quarter-Year FE		x	
Industry×Quarter-Year FE			x
Observations	2,907,501	2,885,953	2,907,501

Notes: All columns report estimates of the linear regression model specified in equation (17). Total bank debt is the sum of all outstanding bank loans for a given firm, in a given quarter. Standard errors, clustered at the 4-digit industry level, are reported in parentheses. The bottom rows specify the fixed effects included in each column. Industry refers to 2-digit industry fixed effects. Observations at the firm-quarter-year level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: Effect of Bankruptcy Reform on Skill Composition and Skill Premium

Dependent Variable:	Skilled/Unskilled	Managers/Total	Professionals/Total	Skill Premium
	(1)	(2)	(3)	(4)
Reform×Q4	**0.085*****	**0.049*****	**0.076*****	**0.100*****
	(0.009)	(0.008)	(0.010)	(0.015)
Reform×Q3	**0.039*****	**0.007**	**0.046*****	**0.045*****
	(0.011)	(0.008)	(0.010)	(0.014)
Reform×Q2	**-0.011**	**0.002**	**0.004**	**-0.002**
	(0.009)	(0.008)	(0.010)	(0.013)
Firm FE	x	x	x	x
Quarter-Year FE	x	x	x	x
Observations	524,496	384,490	520,076	524,496

Notes: All columns report estimates of the linear regression model specified in Equation (17). Skilled/Unskilled is the ratio of skilled to unskilled workers, with a worker being categorized as skilled if possessing at least some post-secondary education. Managers/Total is the ratio of managers (according to ISCO) to total employment. Professionals/Total is the ratio of professionals and technicians (according to ISCO) to total employment. Skill Premium is the ratio of average earnings of skilled and unskilled workers. Standard errors, clustered at the 4-digit industry level, are reported in parentheses. The bottom rows specify the fixed effects included in each column. Observations at the firm-quarter-year level. Differences in the number of observations are due to the fact that our baseline specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Effect of Bankruptcy Reform on Earnings

Dependent Variable:	Wage Bill	Earnings	Incumbent Earnings	Stayer Earnings
	(1)	(2)	(3)	(4)
Reform×Q4	**0.086*****	**0.022*****	**0.042*****	**0.059*****
	(0.014)	(0.005)	(0.008)	(0.008)
Reform×Q3	**0.051*****	**0.002**	**0.019****	**0.022****
	(0.011)	(0.004)	(0.008)	(0.011)
Reform×Q2	**-0.003**	**-0.002**	**0.001**	**-0.001**
	(0.013)	(0.004)	(0.008)	(0.008)
Firm FE	x	x	x	x
Quarter-Year FE	x	x	x	x
Observations	911,680	911,680	911,680	786,520

Notes: All columns report estimates of the linear regression model specified in Equation (17). Wage Bill is the the sum of all wage payments to employees. Earnings are the average earnings paid by the firm. Incumbent Earnings are the average earnings of workers who were matched with the firm prior to the reform. Stayer Earnings are the average earnings of workers who were matched with the firm prior to the reform, and are still with the firm at a given period in time. Standard errors, clustered at the 4-digit industry level, are reported in parentheses. The bottom rows specify the fixed effects included in each column. Observations at the firm-quarter-year level. Differences in the number of observations are due to the fact that our baseline specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 7: Effect of Bankruptcy Reform on Employment Outcomes

Dependent Variable:	Employment	New Hires	Terminations	Share Poached
	(1)	(2)	(3)	(4)
Reform×Q4	**0.063*****	**0.611*****	**-0.521****	**0.039*****
	(0.013)	(0.201)	(0.265)	(0.010)
Reform×Q3	**0.032*****	**0.380*****	**-0.143**	**0.019****
	(0.015)	(0.139)	(0.094)	(0.009)
Reform×Q2	**0.009**	**0.166**	**0.051**	**0.010**
	(0.008)	(0.120)	(0.090)	(0.008)
Firm FE	x	x	x	x
Quarter-Year FE	x	x	x	x
Observations	911,680	813,812	882,236	607,136

Notes: All columns report estimates of the linear regression model specified in Equation (17). Employment is the total number of employees of a given firm. New Hires is the number of workers hired by a firm in the last 4 quarters. Terminations is the number of workers whose employment was terminated in the last 4 quarters, including both workers who quit and workers who are fired. Share Poached is the fraction of workers with a previous employment spell in a firm bellow the median in measure of financial constraints. Standard errors, clustered at the 4-digit industry level, are reported in parentheses. The bottom rows specify the fixed effects included in each column. Observations at the firm-quarter-year level. Differences in the number of observations are due to the fact that our baseline specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 8: Effect of Bankruptcy Reform on Investment and Output

Dependent Variable:	Investment/Assets	Output	Value Added/Worker
	(1)	(2)	(3)
Reform×Q4	**0.060*****	**0.100*****	**0.103*****
	(0.007)	(0.012)	(0.014)
Reform×Q3	**0.021****	**0.038*****	**0.023***
	(0.010)	(0.013)	(0.012)
Reform×Q2	**0.005**	**-0.014**	**-0.018**
	(0.004)	(0.012)	(0.011)
Firm FE	x	x	x
Year FE	x	x	x
Observations	227,920	227,920	227,920

Notes: All columns report estimates of the linear regression model specified in Equation (17). Investment/Assets is total capital expenditures divided by lagged assets. Output is defined at net revenues. Value Added/Worker is defined as value added (net revenues minus intermediate inputs) divided by the number of workers. Standard errors, clustered at the 4-digit industry level, are reported in parentheses. The bottom rows specify the fixed effects included in each column. Observations at the firm-quarter-year level. Differences in the number of observations are due to the fact that our baseline specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 9: Parameters

Panel A: Parameter estimates

	Mean	Median	Min	Max
μ	0.14	0.10	0.05	0.41
τ	0.78	0.81	0.52	0.96
σ	0.68	0.56	-0.10	0.98
ρ	-0.23	-0.31	-0.43	0.04

Panel B: Model parameters

	Value	Description
r	0.26	Average borrowing rate
δ	0.09	Average depreciation rate
ϵ	5.26	Match average markup
β	0.25	Match rent-sharing elasticity
μ	0.14	Average estimate
τ	0.78	Average estimate
σ	0.68	Average estimate
ρ	-0.23	Average estimate

Notes: This table reports model parameters. In panel A, we summarize results from the two-step estimation of production function parameters for each 2-digit industry described in Section 2.2. We report all model parameters and a short description in panel *B*.

A Production Function Estimation

A.1 Baseline Estimation Procedure

Our production function estimation procedure closely follows De Loecker and Warzynski (2012). Consider the following production function

$$q_{it} = f(s_{it}, n_{it}, k_{it}; \gamma) + \omega_{it} + \varepsilon_{it} \tag{31}$$

where q_{it} is logged value added, s_{it} is logged skilled labor, n_{it} is logged unskilled labor, k_{it} is logged capital, γ collects all coefficients, and ω_{it} is logged physical productivity (TFPQ). Our baseline specification relies on a translog functional form for $f()$, which is equivalent to approximating $f()$ by a second-order polynomial in which all inputs, inputs squared, and interaction terms between all inputs are included (in log form). We consider a translog production function of the form

$$q_{it} = \gamma_s s_{it} + \gamma_n n_{it} + \gamma_k k_{it} + \sum_{x\in\{s,n,k\}} \gamma_{xx} x_i^2 + \sum_{w\neq x} \sum_{x\in\{s,n,k\}} \gamma_{xw} x_{it} w_{it} + \omega_{it} + \varepsilon_{it} \tag{32}$$

In order to consistently estimate production function coefficients, we need to control for unobserved productivity shocks, since those are potentially correlated with input choices. We deal with this issue by relying on proxy methods developed by Olley and Pakes (1996) and Levinsohn and Petrin (2003), use material demand

$$m_{it} = m_t(k_{it}, \omega_{it}, s_{it}, n_{it}) \tag{33}$$

to proxy for productivity by inverting $m_t()$. We hence assume that the demand for materials is strictly monotone in ω_{it}.

We follow Ackerberg et al. (2015) and estimate all relevant coefficients using second-stage moments, instead of attempting to identify labor coefficients in the first stage as in Levinsohn and Petrin (2003).[9] In the first stage, we estimate

$$q_{it} = \phi(s_{it}, n_{it}, k_{it}, m_{it}) + \varepsilon_{it} \tag{34}$$

and obtain an estimate of expected output ($\hat{\phi}$) and an estimate of ε_{it}. In the second stage, we rely on the assumed law of motion for productivity

[9] See Ackerberg et al. (2015) and Wooldridge (2009) for a discussion of the issues with this approach.

$$\omega_{it} = g_t(\omega_{it}) + \xi_{it} \tag{35}$$

For a given set of parameters γ, we can compute $\omega_{it}(\gamma) = \hat{\phi} - \gamma_s s_{it} - \gamma_n n_{it} - \gamma_k k_{it} - \sum_{x \in \{s,n,k\}} \gamma_{xx} x_{it} - \sum_{z \neq x} \sum_{x \in \{s,n,k\}} x_{it} w_{it}$. We can then regress $\omega_{it}(\gamma)$ on its lag and recover the innovation to productivity (conditional on the set of parameters γ) $\xi_{it}(\gamma)$. We then estimate the production function parameters using GMM and moment conditions of the form

$$\mathbb{E}[\xi_{it}(\gamma) z^j] = 0 \; j \;\in\; \{s, n, k\}$$
$$\mathbb{E}[\xi_{it}(\gamma) z^j z^h] = 0 \; j, h \;\in\; \{s, n, k\}$$

where z^j, $j \;\in\; \{s, n, k\}$, is an instrument for skilled labor, unskilled labor, capital, or materials. We assume capital is decided one period ahead and is thus not correlated with the innovation in productivity. Under that assumption, we can use capital as its own instrument. We use lagged skilled and unskilled labor as instruments for skilled and unskilled labor, respectively. In order for these instrument to be valid, we require that skilled and unskilled wages be correlated over time, an assumption that is supported by our data.

We measure value added as the difference between deflated net revenue and deflated intermediate inputs, and measure materials as the deflated value of intermediate inputs. We measure skilled labor as the number of workers with at least some college education and unskilled labor as the number of workers with no college education. Finally, we measure capital as the deflated book value of fixed assets.

This procedure produces estimates of the coefficients γ in Equation (32). In addition to using these estimates in our two-step procedure to recover structural production function parameters, we use these coefficients and the observed input utilization for each firm to produce estimates of output elasticities and of markups. The estimated output elasticity of capital (K), for instance, is given by

$$\hat{\theta}^k_{it} = \hat{\gamma}_k + 2\hat{\gamma}_{kk} k_{it} + \hat{\gamma}_{sk} s_{it} + \hat{\gamma}_{nk} n_{it} \tag{36}$$

With estimates of output elasticities, we estimate markups following the approach of De Loecker and Warzynski (2012), who use the first-order condition of an input free from adjustment costs to derive an following expression for the markup. As in De Loecker and Warzynski (2012), we correct for the fact that we do not observe the correct expenditure share since we observe $P_{it} Q_{it} \exp(\varepsilon_{it})$ and not $P_{it} Q_{it}$. We do so by utilizing our estimate of ε_{it} from the first stage and computing markups as

$$\hat{\mu}_{it} = \hat{\theta}^n_{it}\left(\frac{P_{it}Q_{it}/\exp(\hat{\varepsilon}_{it})}{W^n_{it}N_{it}}\right) \tag{37}$$

This method of estimating markups assumes that unskilled labor (N) is a static input into production. If unskilled labor is instead a dynamic input due, for instance, to hiring and firing costs, this procedure would still produce consistent estimates of the coefficients in Equation (32), but not of markups. In this case, the wedge between a firm's output elasticity of unskilled labor and its expenditure share will capture an additional component reflecting adjustment costs.

To address this potential concern, we consider a gross output production function and compute markups using the output elasticity of materials and its expenditure share, and obtain similar results. The gross output version of Equation (32) is given by

$$y_{it} = \gamma_s s_{it} + \gamma_n n_{it} + \gamma_k k_{it} + \gamma_m m_{it} + \sum_{x\in\{s,n,k,m\}} \gamma_{xx} x^2_{it} + \sum_{w\neq x}\sum_{x\in\{s,n,k,m\}} \gamma_{xw} x_{it} w_{it} + \omega_{it} + \varepsilon_{it} \tag{38}$$

where y_i is measured as deflated revenue and m_{it} as the deflated value of intermediate inputs. The first stage remains identical, and we estimate

$$y_{it} = \phi(s_{it}, n_{it}, k_{it}, m_{it}) + \varepsilon_{it} \tag{39}$$

to obtain an estimate of expected output ($\hat{\phi}$) and an estimate of ε_{it}. The second stage is also similar, with the exception that, for a given set of parameters γ, we compute productivity $\omega_{it}(\gamma)$ as $\omega_{it}(\gamma) = \hat{\phi} - \gamma_s s_{it} - \gamma_n n_{it} - \gamma_k k_{it} - \gamma_m m_{it} - \sum_{x\in\{s,n,k,m\}} \gamma_{xx} x_{it} - \sum_{z\neq x}\sum_{x\in\{s,n,k,m\}} x_{it} w_{it}$. After regressing $\omega_{it}(\gamma)$ on its lag and recovering the innovation to productivity $\xi_{it}(\gamma)$, we form an analogous set of moment conditions including materials. We instrument for materials using lagged materials, relying on the assumption that material input prices are serially correlated.

A.2 Robustness to Alternative Procedures

As we describe in Section 3, we use our estimates of output elasticities to construct a measure of financial constraints given by the wedge between a firms' marginal revenue product of capital (MRPK) and its cost of capital. In this section, we consider the robustness of our output elasticity estimates and our MRPK-cost wedge measure to different production function estimation techniques. In particular, we compare our estimates with those obtained imposing a Cobb-Douglas functional form and using a control function approach with intermediate inputs as a proxy variable as in Levinsohn and Petrin (2003), a generalized method of moments estimation procedure as in Wooldridge (2009), and correcting for collinearity in the first stage as in Ackerberg et al. (2015).

We also compare our estimates to elasticities obtained by computing cost shares as in Bollard et al. (2013).

We measure output as deflated revenue and materials as the deflated value of intermediate inputs plus electricity. We measure skilled labor as the number of workers with at least some college education and unskilled labor as the number of workers with no college education. Finally, we measure capital as the deflated book value of fixed assets. Across control function estimation procedures, we use the demand for materials to proxy for productivity.

We present output elasticity estimates according to the different approaches in Table A1. The last row of Table A1 displays the correlation of our baseline measure of financial constraints with alternative measures using elasticities estimated through each method. More specifically, for each estimation procedure, we construct an alternative measure of MRPK-cost wedges at the 4-digit industry level as

$$\text{MRPK-cost wedge}_{jt} = \hat{\theta}_j^k \frac{P_{jt}Q_{jt}}{K_{jt}} \frac{1}{\hat{\mu}_{jt}} - R_{jt} \tag{40}$$

where $\hat{\mu}_{jt}$ is the estimated markup, $\hat{\theta}_j^k$ is the estimated output elasticity of capital, $P_{jt}Q_{jt}$ is deflated revenues, K_{jt} is the deflated book value of fixed assets, and R_{jt} is real financial expenditures plus depreciation, divided by the book value of fixed assets.

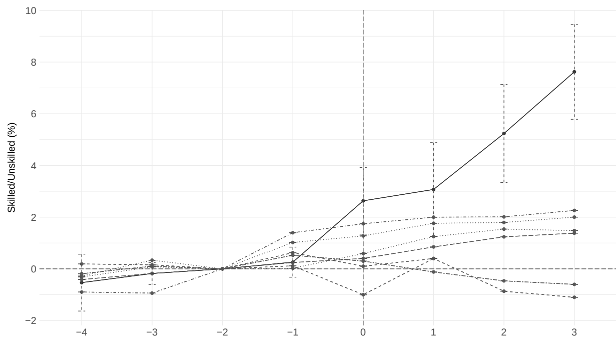
For each estimation procedure, we construct an alternative measure of pre-reform MRPK-cost wedges by computing averages at the 2-digit industry level of estimated wedges using elasticities estimated by each procedure, from 2000 to 2004. All alternative measures are highly and positively correlated with our baseline measure of financial constraints, with correlations ranging from 0.78 to 0.85.

Appendix Table A1: Production Function Estimation

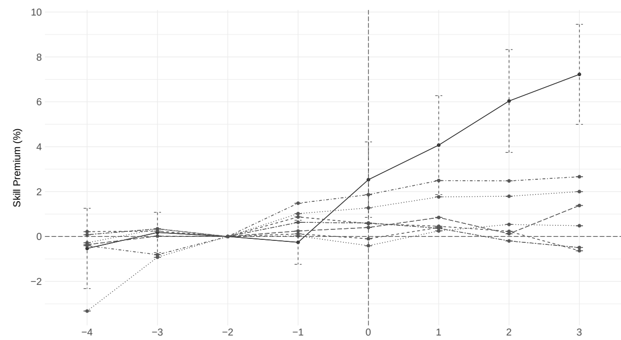
	Baseline	**LP**	**Wooldridge**	**ACF**	**Cost shares**
Capital	0.09	0.10	0.16	0.14	0.13
Skilled	0.14	0.15	0.11	0.16	0.03
Unskilled	0.10	0.10	0.04	0.06	0.06
Materials	0.68	0.63	0.55	0.65	0.77
Corr w/ MRPK-cost wedge$_i$	1	0.83	0.78	0.87	0.81

Notes: This Table shows averages of estimated output elasticities using a gross output translog specification as in De Loecker and Warzynski (2012), using a control function approach with intermediate inputs as a proxy variable as in Levinsohn and Petrin (2003), a generalized method of moments estimation procedure as in Wooldridge (2009), correcting for collinearity in the first stage as in Ackerberg et al. (2015), and by computing cost shares. The last row displays the correlation of our baseline measure of financial constraints with alternative measures using elasticities estimated through each method.

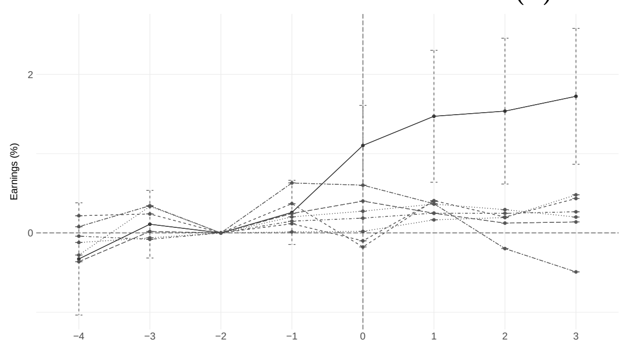
B Additional Results



(a) Skill Ratio



(b) Skill Premium



(c) Earnings

Appendix Figure B1: Placebo Checks

This Figure shows estimates of Equation (24). We split quarterly data from 2000 until 2014 into seven blocks that do not contain the 2005 bankruptcy reform, and one block from 2004Q1 until 2005Q4 that does contain the reform. We plot coefficient estimates of Equation (24) for each of the seven placebo blocks in dashed lines, and show estimates and 95 percent confidence bands of Equation (24) for the true reform period in solid lines. See text for details.

Appendix Table B2: Results by Degree of Capital-Skill Complementarity

Dependent Variable:	Skilled/Unskilled	Skill Premium	Earnings	Employment	Investment/Assets
	(1)	(2)	(3)	(4)	(5)
Reform×Constrained	**0.032*****	**0.037*****	**0.010*****	**0.038*****	**0.041*****
	(0.010)	(0.010)	(0.004)	(0.009)	(0.011)
Reform×Constrained× High CSC	**0.060*****	**0.071*****	**0.006****	**0.020*****	**-0.003**
	(0.010)	(0.011)	(0.003)	(0.006)	(0.003)
Firm FE	x	x	x	x	x
Time FE	x	x	x	x	x
Observations	524,496	524,496	911,680	911,680	227,920

Notes: All columns report estimates of the linear regression model specified in equation (22). The last rows specify which fixed effects are included in each column. Skilled/Unskilled is the ratio of skilled to unskilled workers, with a worker being categorized as skilled if possessing at least some post-secondary education. Skill Premium is the ratio of average earnings of skilled and unskilled workers. Earnings are the average earnings paid by the firm. Total employment is the total number of employees of a given firm. Investment/Assets is defined as capital expenditures scaled by lagged assets. Standard errors, clustered at the industry level, are reported in parentheses. Observations at the firm-quarter-year level in columns 1 to 5 and at the firm-year level in column 6, which explains the smaller number of observations in column 6. Additional differences in the number of observations are due to the fact that this specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Appendix Table B3: Robustness to Controlling for Industry- and Region-Specific Trends

Dependent Variable:	Skilled/Unskilled		Skill Premium		Earnings		Employment		Investment/Assets	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Reform×Q4	**0.054*****	**0.081*****	**0.068*****	**0.096*****	**0.013*****	**0.016*****	**0.035*****	**0.054*****	**0.032*****	**0.049*****
	(0.008)	(0.010)	(0.009)	(0.011)	(0.003)	(0.003)	(0.007)	(0.012)	(0.005)	(0.008)
Reform×Q3	**0.017*****	**0.048*****	**0.015*****	**0.030*****	**-0.004**	**0.009*****	**0.012****	**0.021****	**0.008**	**0.020****
	(0.005)	(0.010)	(0.005)	(0.011)	(0.003)	(0.003)	(0.004)	(0.010)	(0.005)	(0.009)
Reform×Q2	**0.008**	**0.005**	**-0.007**	**-0.017**	**-0.005**	**-0.004**	**0.004**	**0.003**	**-0.005**	**-0.002**
	(0.008)	(0.010)	(0.004)	(0.010)	(0.004)	(0.003)	(0.003)	(0.005)	(0.003)	(0.002)
Firm FE	x	x	x	x	x	x	x	x	x	x
Industry-Time FE	x		x		x		x		x	
Municipality-Time FE		x		x		x		x		x
Observations	524,496	524,496	524,496	524,496	911,680	911,680	911,680	911,680	227,920	227,920

Notes: All columns report estimates of the linear regression model specified in equation (17), including two-digit-industry×time or municipality×time fixed effects. The last rows specify which fixed effects are included in each column. Skilled/Unskilled is the ratio of skilled to unskilled workers, with a worker being categorized as skilled if possessing at least some post-secondary education. Skill Ratio is the ratio of average earnings of skilled and unskilled workers. Earnings are the average earnings paid by the firm. Total employment is the total number of employees of a given firm. Investment/Assets is defined as capital expenditures scaled by total assets. Standard errors, clustered at the industry level, are reported in parentheses. Observations at the firm-quarter-year level in columns 1 to 5 and at the firm-year level in column 6, which explains the smaller number of observations in column 6. Additional differences in the number of observations are due to the fact that this specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Appendix Table B4: Robustness to Controlling for Exposure to the Business Cycle

Dependent Variable:	Skilled/Unskilled	Skill Premium	Earnings	Employment	Investment/Assets
	(1)	(2)	(3)	(4)	(5)
Reform×Q4	**0.078*****	**0.106*****	**0.021*****	**0.057*****	**0.062*****
	(0.010)	(0.011)	(0.003)	(0.012)	(0.008)
Reform×Q3	**0.050*****	**0.036*****	**-0.007**	**0.029*****	**0.019****
	(0.010)	(0.011)	(0.004)	(0.015)	(0.009)
Reform×Q2	**0.010**	**-0.015**	**-0.006**	**0.004**	**-0.002**
	(0.010)	(0.010)	(0.003)	(0.005)	(0.003)
Firm FE	x	x	x	x	x
Beta-Time FE	x	x	x	x	x
Observations	524,496	524,496	911,680	911,680	227,920

Notes: All columns report estimates of the linear regression model specified in equation (17), including GDP-beta×time fixed effects. Skilled/Unskilled is the ratio of skilled to unskilled workers, with a worker being categorized as skilled if possessing at least some post-secondary education. Skill Ratio is the ratio of average earnings of skilled and unskilled workers. Earnings are the average earnings paid by the firm. Total employment is the total number of employees of a given firm. Investment/Assets is defined as capital expenditures scaled by total assets. Standard errors, clustered at the industry level, are reported in parentheses. Observations at the firm-quarter-year level in columns 1 to 5 and at the firm-year level in column 6, which explains the smaller number of observations in column 6. Additional differences in the number of observations are due to the fact that this specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Appendix Table B5: Robustness to Alternative Definitions of Financial Constraints

Panel A: Effect of reform according to firm size						
Dependent Variable:	Bank Debt	Skilled/Unskilled	Skill premium	Earnings	Employment	Investment/Assets
Reform×Small	**0.150*****	**0.094*****	**0.047*****	**0.025*****	**0.028****	**0.050*****
	(0.018)	(0.014)	(0.010)	(0.003)	(0.011)	(0.010)
Firm FE	x	x	x	x	x	x
Time FE	x	x	x	x	x	x
Observations	1,941,197	582,641	582,641	944,152	944,152	236,038

Panel B: Effect of reform according to firm age						
Dependent Variable:	Bank Debt	Skilled/Unskilled	Skill premium	Earnings	Employment	Investment/Assets
Reform×Young	**0.194*****	**0.072*****	**0.065*****	**0.021****	**0.043*****	**0.044*****
	(0.019)	(0.016)	(0.013)	(0.010)	(0.015)	(0.009)
Firm FE	x	x	x	x	x	x
Time FE	x	x	x	x	x	x
Observations	1,941,197	582,641	582,641	944,152	944,152	236,038

Panel C: Effect of reform according to TFPR over assets						
Dependent Variable:	Bank Debt	Skilled/Unskilled	Skill premium	Earnings	Employment	Investment/Assets
Reform×High TFPR/Assets	**0.040****	**0.052*****	**0.046*****	**0.018*****	**0.057*****	**0.035*****
	(0.018)	(0.012)	(0.010)	(0.005)	(0.014)	(0.008)
Firm FE	x	x	x	x	x	x
Time FE	x	x	x	x	x	x
Observations	1,941,197	524,496	524,496	911,680	911,680	227,920

Notes: All columns report estimates of the linear regression model specified in equation (**??**). Reform is a dummy that equals 0 prior to the reform and 1 after the reform. Small is a dummy for the firm being in the first quartile of firm size (in number of employees) to the reform. Young is a dummy for the firm being in the first quartile of firm age prior to the reform. High TFP/Assets is a dummy for the firm being in the fourth quartile of estimated logged revenue productivity (TFPR) scaled by logged assets prior to the reform. Total bank debt is the sum of all outstanding bank loans for a given firm, in a given quarter. Skilled/Unskilled is the ratio of skilled to unskilled workers, with a worker being categorized as skilled if possessing at least some post-secondary education. Skill Ratio is the ratio of average earnings of skilled and unskilled workers. Earnings are the average earnings paid by the firm. Total employment is the total number of employees of a given firm. Investment/Assets is defined as capital expenditures scaled by total assets. Standard errors, clustered at the industry level, are reported in parentheses. Observations at the firm-quarter-year level in columns 1 to 5 and at the firm-year level in column 6, which explains the smaller number of observations in column 6. Additional differences in the number of observations are due to the fact that this specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Appendix Table B6: Robustness to Alternative Sources of Variation

Dependent Variable:	Total Debt		Investment/Assets	
	(1)	(2)	(3)	(4)
Reform×High Tax Burden	0.10***		0.07**	
	(0.03)		(0.03)	
Reform×High Tangibility		0.08**		0.05***
		(0.04)		(0.02)
Firm FE	x	x	x	x
Quarter-Year FE	x	x	x	x
Observations	11,635	11,635	9,337	9,337

Notes: All columns report estimates of the linear regression model specified in equation (27), using data from Economatica. Reform is a dummy that equals 0 prior to the reform and 1 after the reform. High tax burden is a dummy for the firm being in the fourth quartile of tax burden prior to the reform, with tax burden defined as total tax expenses divided by earnings before interest and taxes. High tangibility is a dummy for the firm being in the fourth quartile of the share of tangible assets prior to the reform. Total Debt is defined as total debt liabilities minus trade credits. Investment/Assets is defined as capital expenditures scaled by total assets. Standard errors, clustered at the industry level, are reported in parentheses. Observations at the firm-quarter-year level. Differences in the number of observations are due to the fact that this specification uses a log transformation of dependent variables, and the number of zeros varies across variables. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.