

# Do Marginal Products Differ from User Costs? Micro-Level Evidence from Italian Firms

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## Abstract

Using micro-data on firm-specific borrowing costs and wages, we demonstrate that distortions in firms' employment and investment policies can be empirically measured using firm-level gaps between marginal revenue products and user costs (MRP-cost gaps). We estimate MRP-cost gaps for 4 million firm-year observations in Italy between 1997 and 2013, showing that the variation in these measures is closely related to the extent of credit market frictions and to the degree of labor market rigidities individual firms face. Using the estimated MRP-cost gaps, we propose a reallocation algorithm that helps us assess the scope of capital and labor misallocation in Italy, and its impact on aggregate output and total factor productivity (TFP). We calculate that, holding constant the aggregate capital and labor endowments in the economy, the Italian corporate sector could produce between 3% to 4% more output by reallocating resources from over-endowed producers toward higher-value users. The output losses from misallocation are larger during episodes of macro-financial instability, in non-manufacturing industries, and in geographical regions with less developed socio-economic institutions.

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The Appendix of the paper is available at [www.simonelenzu.com](http://www.simonelenzu.com) or [here](#)

# 1 Introduction

Research in economics and finance has long been interested in measuring the extent and implications of resource misallocation.<sup>1</sup> An intuitive way to conceptualize misallocation is to think of frictions and regulations as implicit taxes that generate wedges in the first-order conditions characterizing firms’ optimal investment and employment policies (Chari et al. 2007; Restuccia and Rogerson 2008). This approach captures the idea that producers may face differential relative costs when they try to acquire capital and labor inputs in the market, either because they are charged different prices or because they face quantity constraints (shadow prices). When differential costs do not reflect heterogeneity in fundamentals or risk, they cause some producers to be either too large or too small relative to their “socially efficient” size. This misallocation squanders scarce resources, reducing aggregate total factor productivity (TFP), and ultimately impairing economic growth (Banerjee and Duflo 2005).

Despite the large interest in this topic, data limitations have prevented researchers from directly measuring deviations from optimal capital and labor policies, mostly due of the inability to gather micro-data on the user costs of capital and labor paid by individual producers. To overcome these empirical constraints, the literature has produced appealing indirect measures of misallocation (e.g., Hsieh and Klenow 2009) that, however, rely on specific assumptions about firms’ demand and production technologies and therefore might over- or understate the extent of misallocation when these assumptions are violated (Asker et al. 2014; Haltiwanger et al. 2017).

In this paper, we shed light on the distribution of the firm-level gap between marginal revenue products of capital and labor and their user costs (MRP-cost gaps), provide evidence of the relation of such gaps to market frictions and regulations, and ultimately use them to quantify the impact of resource misallocation on aggregate TPF and output. We assemble a comprehensive bank-firm-employee matched panel database that contains micro-level information on firm-specific wages, borrowing costs, balance-sheet data, and bank credit for the sample of non-financial corporations active in Italy between 1997 and 2013. We link accounting variables from the census of corporations to the archives of the National Credit Register and to employer-employee records obtained from the social security administration. The coverage, granularity, and richness of our data puts us in the unique position of observing the distribution of the user cost of both capital and labor, and allows us to estimate the distribution of marginal revenue products of primary inputs (Gandhi et al. 2017b; De Loecker and Warzynski 2012).

To gain intuition on the economic content of MRP-cost gaps, let us consider a neoclassical environment with homogeneous producers and no risk. When input policies are fully unconstrained, firms accumulate assets and hire labor up to the point where their marginal revenue products are equal to their user costs. We show that this intuition can be generalized to a more realistic framework with heterogeneous producers, where capital structure matters and default risk is endogenous. In particular, when debt is the marginal source of financing and borrowing rates are pre-determined and rigid, the gap between the marginal revenue product of capital and its user cost (the sum of the interest rate and the depreciation rate on the capital stock) is positively related to the shadow cost of capital that is generated by binding credit constraints (Stiglitz and Weiss 1981, 1992). Similarly, when wages are rigid, the gap between the marginal revenue product of labor (MRPL) and the wage is proportional to the implicit cost of labor that firms face, such as the ones generated by regulatory interventions in labor markets (Petrin and Sivasadan 2013).

Our research has three primary empirical results. First, we characterize the distributions of MRP-cost

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<sup>1</sup>See Hsieh and Klenow (2009), Restuccia and Rogerson (2013), Hopenhayn (2014), and the literature cited.

gaps of capital and labor. According to our metric, the central percentiles of the distributions are occupied by firms whose capital and labor endowment appears to be relatively undistorted. The median gaps of capital and labor are 3.5 % and 6 thousand euros for capital and labor, respectively. We calculate that, to close the gaps of the median firm, investing an amount of capital worth 1% of firm assets and hire 3% extra workers would be sufficient.

Yet the distributions of MRP-cost gaps are dispersed and highly right-skewed. The average capital and labor gaps are 37% and 9 thousand euros, respectively; the 90-10 percentile differences are almost 3 times larger. Based on our estimates, 25% of the firm-year observations should have invested to acquire 6% or more capital, and expand their labor force by 15% or more. On the contrary, another 25% of firms should have sold 1% or more of their assets, and over 10% of observations should have reduced their labor demand by 1% or more. These findings are indicative of suboptimal investment and employment policies, and suggest that output gains might be attainable through a reallocation of resources.

Second, we show that the variation in firm-level MRP-cost gaps is related to the extent of financial frictions and to the degree of labor market rigidities individual firms face. On the capital side, we analyze the impact of asymmetric information in credit markets and the effect of bankruptcy costs, and we study the response of the MRP-cost to idiosyncratic shocks to credit supply. On the labor size, we use MRP-cost gaps of labor to analyze the impact of a labor market regulation that imposes severance payments that vary as a function of firm size.

Economic theory suggests that repeated interactions with financial intermediaries allow firms to overcome possible asymmetric information frictions, and gradually accumulate a capital endowment more consistent with profit maximization (Diamond 1991). In line with these theoretical predictions, we find a monotonic negative relation between MRP-cost gaps for capital and the length of the lending relationships of a firm with its current lenders. We estimate that the amount of investment needed to close the gap is worth 25% of the installed capital for firms with newly established lending relationships; this amount reduces to 10% after three years, and to 6% after 10 years of continuous bank-firm interactions. Importantly, the benefits associated with tighter bank-firm relationships are entirely concentrated among those borrowers that operate with an insufficient capital endowment, and they are stronger for highly productive firms. Both findings are consistent with the predictions of economic theory, according to which, *ceteris paribus*, the shadow cost of capital is higher for more productive capital-constrained firms.

Then we show that the variation in MRP-cost gaps is related to the costs of bankruptcy procedures. To do so, we use cross-province variation in the length of bankruptcy litigations across Italian provinces. We find MRP-cost gaps of observationally similar firms are significantly lower in jurisdictions with shorter bankruptcy litigations. This statistical relationship holds even if we restrict our focus to cross-province variation within the same industry-year-macro region (North-Center-South) into account for other socio-economic differences across geographical regions of Italy (Guiso et al. 2004a; Guiso et al. 2004b).

Next, we analyze how the MRP-cost gap of capital responds to changes in the supply of credit. A major challenge in answering this question is to separately identify time-varying credit-supply shocks from simultaneous firm-borrowing demand shocks. To unravel the demand and supply channels, we construct firm-time-specific credit supply shifters. We adopt a shift-share approach that, by leveraging on the granularity of the bank-firm matched records from the Credit Registry, allows us to disentangle nationwide changes in credit supply of individual financial institutions from the idiosyncratic credit demand for their borrowers (Greenstone et al. 2015; Amiti and Weinstein 2016).<sup>2</sup> Consistent with the theoretical prediction that variation

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<sup>2</sup>A contemporaneous work, Manaresi and Pierri (2017), uses firm-bank matched records from the Italian Credit Registry to

in gaps across firms captures heterogeneous shadow costs of capital, we find that, all else being equal, the exposure to a positive supply shock reduces the MRP-cost gaps, whereas a negative credit-supply shock increases them. We find that the response of MRP-cost gaps to credit-supply shocks is substantial for capital-constrained firms, especially the response to positive supply shocks, and that the sensitivity of gaps to credit shocks is particularly strong for more productive firms. On the contrary, the MRP-cost gaps of firms with zero or negative MRP-cost gaps (i.e., those that operate with a capital endowment close to or above target) show a small or no response to credit-supply shocks. That is, by and large, this group of firms respond to an expansion in the credit supply by rolling over their debt, rather than by undertaking new investments, and does not appear to be affected by a credit contractions.

Finally, we analyze the relation between labor gaps and labor market regulations. During our sample period, the provisions of the Italian Workers Statute imposed large severance payments to firms employing more than 15 employees, but significantly smaller payments to firms with 15 or fewer employees (Garibaldi and Violante 2005; Schivardi and Torrini 2008).<sup>3</sup> Size-dependent firing costs are an adjustment cost that generates variation in marginal revenue products and, if not undone by properly designed wage contracts (Lazear 1990), generates misallocation. We find that as the 15-employees threshold is approached, the average gap between the marginal revenue product of labor and wages increases. Higher labor costs also affect firms above this threshold, inducing them to operate with a smaller labor force than the one they might have chosen in the absence of the size-dependent regulation. From a dynamic point of view, we test the local response of firms labor demand to productivity shocks. We estimate that a 1% positive increase in firm-level productivity increases the rate of under-employment by 5 percentage points for firms at the threshold (15 employees), relative to firms immediately below the threshold (14 employees). These results are consistent with the hypothesis that the government-mandated severance payments curb economic growth by discouraging firms from increasing their size despite the growth opportunities that might be available.

Importantly, our research highlights that, for both capital and labor, the dispersion of MRP-cost gaps and the relation between gaps and market frictions is entirely driven by variation in marginal revenue products. Borrowing costs and wages, by contrast, display a limited cross-sectional variation. This finding suggests market prices are not the instruments that allocate resources across credit and labor market participants.

Other phenomena and frictions (i.e., different from credit and labor market frictions) are likely to contribute to the size and dispersion of the gap between inputs' marginal revenue products and user costs, such as economic uncertainty and real adjustment costs (Asker et al. 2014; Foster et al. 2016). Addressing these concerns, we show that the relation between MRP-cost gaps and credit and labor market frictions is robust to controlling for age, size, credit rating, firm-level productivity and profitability measures, and hold true if we restrict the analysis to within-industry-year-province variation or to within-firm variation. We also evaluate the robustness of our estimates of the marginal revenue products with respect to alternative production function estimations.

The third set of results in this paper cast light on the aggregate implications of resource misallocation in Italy. We use MRP-cost gaps to estimate how idiosyncratic distortions in input policies translate into aggregate output and TFP losses, and to document how gains from reallocation evolved over time and how they differ across sectors and different geographical regions. We calculate that, in any given year, aggregate TFP and output of the Italian corporate sector could be 3%–4% higher following a reallocation of production factors, by taking resources away from firms that over-utilize them, and redistributing these

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construct firm-level credit supply shifters for a subsample of large Italian corporations, to study the impact of credit availability on firm-level productivity.

<sup>3</sup>This size-dependent provision of the Italian Worker Statute (Article 18) was reformed in 2012 and finally abolished in 2014.

resources to the most productive producers who are lacking them. The majority of allocative inefficiencies take place within narrowly defined industries (roughly two-thirds), and roughly two-thirds of within-industry misallocation takes place within the same geographical regions. Also, we find that gains from reallocation are one-third higher during periods that are characterized by financial instability – the financial crisis (2008-2009) and following the burst of the sovereign debt crisis (2010-2013) – compared to those estimated during the 1997-2004 period. Examining sectoral heterogeneity, we find that the scope of misallocation is more severe outside of manufacturing industries (i.e., services and construction). This finding is important because data constraints has forced most of the existing literature in this topic to focus on manufacturing industries.<sup>4</sup> Our analysis suggests that this might lead researchers to underestimate the extent of resource misallocation. Finally, we examine the spacial variation in misallocation in Italy. Previous research has documented a large disparity in terms of quality of markets and institutions between Italy’s Southern regions and the rest of the country (Putnam et al. 1994). Accordingly, we document larger output and TFP losses directly imputable to misallocation in the Southern regions when compared to Northern and Central regions of the country.

Our research contributes to a broad scope of literature interested in studying the impact of market frictions and regulations on firms’ real activity. The empirical measures produced and analyzed in this paper (MRP-cost gaps) are linked to theory and, because they vary both between- and within-firms, they allow us to shed light on the differential impact of market distortions across heterogeneous types of firms. We see MRP-cost gaps as a particularly appealing empirical tool for researchers seeking to identify firms that are more likely to be financially constrained and for those interested in measuring the real effects of financial frictions, both of which are key topics in corporate finance and applied macroeconomics. In these respects, the value added of our approach is particularly relevant when studying investment policies of privately owned firms. For these firms, traditional measures such as Tobin’s Q (Hayashi 1982; Abel and Eberly 1994) or indexes of financial constraints (Kaplan and Zingales 1997; Whited and Wu 2006) are not computable because there no information is available on the market value of their assets and liabilities. By contrast, the estimation of MRP-cost gaps requires standard product variables and information on firm-specific user costs, both of which are observable for private firms, and are becoming accessible to researchers as more administrative databases are being disclosed.<sup>5</sup>

This paper directly speaks to the literature that studies the impact of a suboptimal allocation of resources on aggregate TFP and output (Banerjee and Dufo 2005; Restuccia and Rogerson 2008; Hsieh and Klenow 2009; Petrin et al. 2011; Gilchrist et al. 2013; Bartelsman et al. 2013).<sup>6</sup> Our contribution is twofold. First, to the best of our knowledge, our paper is the first to try to characterize the distribution of deviations from firms’ first-order conditions using detailed micro-data on borrowing costs and wages, thereby showing how these deviations relate to specific frictions and regulations in factor markets, and how to aggregate them to provide macro assessments. Secondly, substantial empirical evidence documents declines in aggregate TFP and output during economic downturns and, in particular, following episodes of financial instability (Calvo et al. 2006; Jermann and Quadrini 2012). An open question is whether a change in the scope of resource misallocation, on top of (or instead of) technology shocks, contributes to explaining the co-integration of

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<sup>4</sup>Relatively few papers have addressed misallocation in the service sector. Those empirical studies that do, also find that the scope of misallocation appears to be larger in services sectors than than in manufacturing (Busso et al. 2013; De Vries 2014; Dias et al. 2016).

<sup>5</sup>The ongoing effort of several national data providers to collect information on firm-level borrowing costs (e.g., the AnaCredit project by the ECB, or the CompNet Network) suggests we should expect databases similar to ours to become soon available in other countries. We hope our work can provide guidelines for future research interested in measuring policy distortions combining information on production, financing, and factor prices.

<sup>6</sup>See Restuccia and Rogerson (2013), Hopenhayn (2014), and Restuccia and Rogerson (2017) for a review.

business-cycle fluctuations and aggregate TFP. Our results speak to this question by showing that, indeed, boom and burst cycles in credit markets can affect aggregate TFP due to a deterioration in the efficiency of capital allocation (Gopinath et al. 2016; Oberfield 2013; Sandleris and Wright, 2014; Schivardi et al. 2017).<sup>7</sup>

Relatedly, our work also bridges the misallocation literature and the literature that studies the real effects of changes in the supply of credit by financial institutions. Previous works have analyzed the real effects of credit-supply shocks on firms' input accumulation and revenues (e.g., Khwaja and Mian 2008; Chodorow-Reich 2014; Banerjee and Duflo 2014; Cingano et al. 2016; Bottero et al. 2017) and, more recently, their impact on firm-level productivity (Manaresi and Pierri 2017; Duval et al. 2017). By combining our measure of policy distortions with quasi-experimental variation individual firms face in the supply of credit, this paper casts light on the distributional effects of changes in financial intermediaries' lending policies, and on their aggregate implications. Our analysis also sheds light on the relative importance of the price channel versus the quantity channel in the transmission of credit market frictions to the real economy. We document a substantial rigidity of loan prices, and provide evidence that credit limits (i.e., quantity rationing) are the most salient feature of business loan contracts.<sup>8</sup>

Finally, the analysis of the effects of size-dependent labor market regulations connects this paper to a strand of empirical works in labor economics (Schivardi and Torrini 2008; Hijzen et al. 2013; Bertrand et al. 2015) and applied macroeconomics (Guner et al. 2008; Garicano et al. 2016) that evaluates the micro- and macroeconomic impact of labor market regulations on firm policies.<sup>9</sup> Our approach parallels and extends the one in Petrin and Sivadasan (2013). Given the widespread presence of size-dependent labor market regulations across countries, and the evidence on the comparability of labor demand functions around the world (Hamermesh 1996, Heckman et al. 2006), lessons about the impact of the Italian employment protection regulation are likely applicable to other countries and to similar types of government interventions in labor markets.

The paper is organized as follows. Section 2 describes the data and the institutional features of the Italian credit and labor market that are relevant for our analysis. Section 3 presents the theory underpinning the MRP-cost gaps and illustrates their relationship to market frictions. Section 4 estimates the gaps and characterizes their empirical distribution. Section 5 explores the relationship between MRP-cost gaps and credit and labor market frictions. Section 6 presents firm-level counterfactuals useful to quantify the magnitude of firm-policy distortions. We examine the aggregate implications of resource misallocation in section 7. Section 8 concludes.

## 2 Data and Institutional Context

We assemble a comprehensive employee-employer-bank matched database that contains micro-level information on firm-specific wages, borrowing costs, balance-sheet data, and bank credit for the lion's share of

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<sup>7</sup>Recent papers examine the aggregate costs generated by firm-specific collateral constraints (Chaney et al. 2017) and sub-optimal capital structure (Whited and Zhao 2017).

<sup>8</sup>Economic theorists have emphasized that prices in credit markets are not the instrument used to allocate resources (Stiglitz and Weiss 1981; Stiglitz 1992). A body of evidence in the consumer lending market has corroborated these predictions (Ausubel 1991; Adams et al. 2009; Einav et al. 2012; Agarwal et al. 2017). Yet besides a few noteworthy exceptions, empirical evidence for firms remains scant, mostly due to the lack of longitudinal micro-level data that provide information on the borrowing costs paid by individual on firms (Hannan and Berger 1991; Petersen and Rajan 1994; Crawford et al. 2016).

<sup>9</sup>The analysis of the economic effects of firing costs in Hopenhayn and Rogerson (1993) is one of the earliest studies of misallocation due to regulation. See Cooper and Willis (2009) for a study the aggregate implications of different forms of establishment-level labor adjustment costs.

non-financial incorporated firms that were active, in Italy, between 1997 and 2013. We assemble our data by merging and harmonizing different administrative and proprietary sources.

We collected detailed information on yearly balance sheets, income statements, and registry variables from Cerved Group S.p.A. (Cerved database).<sup>10</sup> We merge the firm-level dataset with the archives of the national Credit Registry (CR) administered by the Bank of Italy, and to matched employer-employee records from the Italian National Social Security Institute (INPS). The CR provides us with information on firms' credit market participation, debt exposure, and corresponding borrowing cost (interest rates) for each bank-firm credit relationship. The Social Security records allow us to observe wages and a detailed snapshot of firms' workforce composition. We complement these data with information on industry-specific price deflators, industry-specific depreciation rates of fixed assets, and socioeconomic indicators measured at the province level, all of which are collected from the publicly available archives of the Italian National Statistical Institute (ISTAT).<sup>11</sup> From the archives of the Italian Ministry of Justice, we collect information on the average length of bankruptcy litigations in court.

Our final dataset includes 3.9 million firm-year observations, 6.5 thousand firms, and 13.3 million credit relationships. It amounts to circa 90% of the value added produced by the corporate sector in the selected industries, and over 70% of the total value added produced by the whole Italian corporate sector.<sup>12</sup> To the best of our knowledge, ours is the first longitudinal dataset that provides information on both production and financing, as well as firm-specific wages and borrowing costs for the large majority of the corporate sector of a country. Table 1 reports the summary statistics of the main variables used in our analysis. Appendix A provides a detailed description of each variable and of the steps followed to clean the database. By all means, our sample is composed predominantly of small and medium enterprises, matching the size and industry distribution of Italian firms.<sup>13</sup> Almost all the companies in the Cerved sample are unlisted, making our dataset particularly suited for the purpose of this study, because market failures are expected to have a greater impact on small and young producers (Gertler and Gilchrist 1994; Petersen and Rajan 1994; Chodorow-Reich 2014).<sup>14</sup>

**Credit market relations** – Over 80% of the firm-year observations report access to some form of bank credit ( $BORROWER=1$ ); only 9% of firms never engaged in any type of credit market transaction at some point between 1997 and 2013. Bank debt is worth on average 43% of firm total assets (54%, if we consider only firms with outstanding debt obligations).

Exploiting the panel dimension of the CR database, we gauge information on the number and length of active credit relationships between firms and individual credit institutions. On average, firms have four active credit relations with financial intermediaries ( $NUMBER\ RELATIONS_{it}$ ).<sup>15</sup> The variable  $LENGTH\ RE-$

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<sup>10</sup>Our database includes only incorporated businesses (limited liability companies), but not sole proprietorship and other non-incorporated firms. The unit of observation is a firm-year, no plant-level information is available. Compared to other publicly available datasets (such as Orbis and Amadeus by Bureau van Dijk Electronic Publishing; see Kalemli-Ozcan et al. (2015)), our database has the advantage of having no selection bias, no issues with merging different vintages, and a substantially richer set of balance sheet, income statement, and registry variables.

<sup>11</sup>Data available at <https://www.istat.it/en/>.

<sup>12</sup>We drop the following industries: Agriculture, Mining and quarrying, Utilities, Public administration and National defense, Education, Health services, Activities of membership organizations, Activities of households as employers, and Activities of extraterritorial organizations and bodies to avoid dealing with firms with complete or partial government ownership, or heavily subsidized by the government; Financial and insurance activities and Real estate activities because firms operating in these industries are themselves credit providers. See Appendix A for further details.

<sup>13</sup>The median firm in our dataset collects 825 thousand euros per year in revenues, has a book value of fixed assets worth 706 thousand euros, and only 6 employees (average number of employees through the year). The macro-industry composition mirrors the one of the Italian economy: 29% of the observations refer to firms operating in manufacturing (23% of the firms); 54% to firms operating in the service sector (61% of the firms); 17% to firms in construction industry (16% of the firms).

<sup>14</sup>In our final sample, only 224 firms are publicly listed.

<sup>15</sup>Multi-bank relations are a wide-spread phenomenon in business lending, including the United States (Detragiache et al.

$\text{LENGTH RELATION}_{it}^{wmean} = \sum_{b \in \mathcal{N}_{it}^B} \frac{\text{Credit}_{ibt}}{\text{Credit}_{it}} \cdot \text{LENGTH RELATION}_{ibt}$  measures the (weighted) average length of active relations, where  $\text{LENGTH RELATION}_{ibt}$  measures the number of years of continuous relationship between firm  $i$  and its lender  $b$ , and  $\mathcal{N}_{it}^B$  is the set of all its active lenders at time  $t$ , and  $\frac{\text{Credit}_{ibt}}{\text{Credit}_{it}}$  is the share of total credit provided by each lender. We also construct a second proxy that measures the length of the relationship with the most important lender in terms of outstanding credit ( $\text{LENGTH RELATION}_{it}^{lead}$ ) and, for completeness, we compute the unweighted average length of relations ( $\text{LENGTH RELATION}_{it}^{mean}$ ). The three relationship variables are, by construction, bounded between 0 (no credit relations) and 16 years (the span of our sample). A comparison of the three measures offers important insights into the nature of firm-bank interactions. The data highlight that credit relationships, once established, tend to be quite stable. The average relationship lasts over 5.4 years, about one-third of the span of our sample. Moreover, Table 1 shows that  $\text{LENGTH RELATION}_{it}^{mean} < \text{LENGTH RELATION}_{it}^{wmean} < \text{LENGTH RELATION}_{it}^{lead}$ . This finding indicates that, while engaging in multiple relations, not all of them are equally important or equally long-lasting. This evidence is in line with the empirical findings reported in Petersen and Rajan (1994) for small firms in the United States, and corroborates the theoretical predictions that banks gradually expand their credit supply as they develop a tighter relationship with their borrowers (Diamond 1991).

For each firm-year observation, we have information on their CREDIT SCORE measured by Altman Z-score (Altman 1968; Altman et al. 1994). This credit-rating metric is widely used by Italian financial intermediaries in their assessment of firms' creditworthiness (see Albareto et al. 2011). It ranges from 1 to 9, with lower numbers (1–4) indicating high solvency and low risk, and higher numbers (7–9) indicating troubled economic conditions and high default risk. Return on Assets (ROA), ASSETS TURNOVER (Revenues/Assets), and CASH FLOWS/ASSETS are measures of profitability, also commonly used in banks' credit assessments.

Finally, we construct an empirical proxy of the deadweight costs incurred in case of bankruptcy. Using data from the Italian Ministry of Justice, we collect information on the average length of bankruptcy trials. For every Italian province, we calculate the average length of cases concluded in years 2005–2007 ( $\text{LENGTH BANKRUPTCY}$ ). The length of the bankruptcy litigations increases the deadweight loss in case of bankruptcy, because the lender is more exposed to borrowers' moral hazard behavior, and the market value of firms' assets typically decays during the period of automatic stay.<sup>16</sup> The data show it takes, on average, almost nine years to resolve a bankruptcy dispute through Italian courts. The standard deviation is two years, with judgments taking "as little as" three years to become final in some provinces, but 13 years in others. We return to sources of variation in this variable in section 5.<sup>17</sup>

**Labor market relations** – Two institutional features of the Italian labor market are important for our paper. The first is the wage-setting mechanism. In Italy, wages are predominantly determined by a two-tier bargaining structure: (1) the first-level bargaining is collective and takes place at the national-sectoral level. It determines the general terms and conditions of employment for different occupations and basic minimum-wage guarantees (*minimi tabellari*); (2) bargaining at the second-level takes place at the regional level or at firm level, and it allows firms and workers to supplement national contracts. Second-level bargaining is optional, and, importantly, it is restricted to upward wage adjustments with respect to the minimum

2000).

<sup>16</sup>According to the World Bank's "Doing Business" report, Italy ranks 160th out of 185 countries in the enforcing contracts indicator. The poor performance of the Italian legal system largely due to extremely long judicial proceedings. The same report highlights that, in Italy, it takes on average 1,210 days to resolve a commercial dispute through the courts, which is about four times the number of days needed in the United States and three times the number of days needed in the UK and in Germany.

<sup>17</sup>See Appendix A for further information about this variable and its geographical variation.



wage guarantees set by the first-level negotiations.<sup>18</sup><sup>19</sup> Several papers have documented that second-level bargaining is rarely used, and only by medium-large firms. For example, D’Amuri et al. (2015) finds that less than 20% of firms with more than 20 employees use secondary bargaining. This suggests that, although *de jure* wages could adjust upwardly via firm-level bargaining, *de facto* they are anchored to the occupational wage rate periodically set at the national level.

The second institutional feature is the stringent employment protection regulation and its size-dependent nature.<sup>20</sup> Under the Italian employment protection legislation in place during our sample period, individual and collective dismissals of workers with open-end contracts are only allowed on a “just cause” basis. When workers appeal to the court against dismissal, and judges rule the dismissal unfair, firms must provide compensation in the form of severance payments that vary according to firm size. For firms above 15 employees, the firing costs are substantial. Under Art.18 of the Italian Worker’s Statute (Law 300/1970), such firms are obliged to reinstate the unfairly dismissed worker, unless the worker opts for a severance payment of at least 15 months of salary. Moreover, employers also have to compensate unfairly laid-off workers for the forgone wages in the time elapsing between the firm’s dismissal and the final sentence. This process can take up to five years due to the inefficiency of the Italian legal system. Thus, a firm larger 15 employee faces severe expected firing costs when it attempts to scale down its workforce (Garibaldi and Violante 2005; Schivardi and Torrini 2008).<sup>21</sup> For firms with 15 or fewer employees, Article 18 does not apply, and their expected firing costs in case of unfair dismissals are substantially lower: they must compensate unfairly dismissed workers with a severance payment that varies between 2.5 and 6 months of salary or, as an alternative to the severance payment, firms can opt for reinstating the worker.

### 3 A Theory of Gaps

Let us consider a neoclassical environment with homogeneous producers and no risk. When input policies are fully unconstrained, firms accumulate assets and hire labor up to the point where their marginal revenue products are equal to their user costs. In this section, we show this intuition can be generalized to a more realistic framework with heterogeneous producers, where capital structure matters and default risk is endogenous. Appendix C provides a full description of the model.

**Economic environment** – Consider a firm run to maximize the present discounted value of cash flows to risk-neutral shareholders in an environment where firms are heterogeneous with respect to the realization of firm-specific revenue productivity ( $\omega_{it}$ , TFPR). Every period, the manager observes the realization of productivity, and then he decides whether (i) to repay its outstanding debt or (ii) default and exit. From a firm’s standpoint, a default on bank debt is the optimal decision when the realization of  $\omega$  is below an endogenously determined threshold level  $\bar{\omega}$  (Hennessy and Whited 2007).<sup>22</sup> If the firm is worth more as an

<sup>18</sup>The general terms and conditions of employment contracts and minimum-wage guarantees agreed upon in the first-level bargaining are renegotiated, for different occupations, every four and two years, respectively. The amendments to the national contracts renegotiated in the second-level bargaining are valid for four years.

<sup>19</sup>Only in well-delimited cases of firm’s restructuring or crisis, second-level deals can (temporarily) cut wages below the nationally set sectoral minimum. Still, although legally possible, evidence of firm-level agreement envisaging a decrease in the wage below these minima is scant (see D’Amuri et al. 2015).

<sup>20</sup>According to the OECD index of strictness of employment protection regulation, Italy ranks fifth among the OECD countries. Size-dependent regulations in labor markets are common in both developed and developing countries (see Guner et al. 2008). Gourio and Roys (2014) and Garicano et al. (2016) analyze the effect of size-dependent regulation in France; Braguinsky et al. (2011) in Portugal; Abidoye et al. (2009) in Sri Lanka; Martin et al. (2017) in India; in the United States under the US Affordable Care Act, penalties are levied against firms with more than 50 full-time employees that do not offer health care insurance to their employees.

<sup>21</sup>Article 18 was substantially reformed in 2012 and finally abolished in 2014.

<sup>22</sup>Revenue productivity is a combination of technical Hicks neutral productivity and consumer demand (Foster et al., 2008).

ongoing concern, the manager repays its obligations, and he chooses new factor demands (capital  $K_{it+1}$ , labor  $L_{it}$ , and intermediate inputs  $M_{it}$ ) and how to finance these purchases (bank debt  $B_{it}$ , internally generated cash flows, or capital injection from shareholders). In case of default, creditors acquire ownership and control of the firm. They produce during the current period, and liquidate the firm at the end of the period. We assume liquidation is costly, as a fraction  $X \geq 0$  of firm assets are lost during the bankruptcy process.

**Firm policies and MRP-cost gaps** – We heuristically characterize firms’ investment policies using the augmented Euler equation of capital and the first-order condition for labor.

We assume new capital injections from shareholders are costly and restrict our attention to cases in which debt is the marginal source of financing for incremental investment.<sup>23</sup> We consider a credit market where lenders offer loan contracts that consist of a single interest rate for each group of observationally similar firms ( $r_{it+1} = \bar{r}_{t+1}$ ), and deal with diversity by rationing those firms within the group that have a loan demand exceeding the loan offer (Stiglitz and Weiss 1981).<sup>24</sup> For example, as in Kiyotaki and Moore (1997), banks might tie credit supply to firms’ net worth  $B_{it+1} \leq \lambda_{it}K_{it+1}$ ,  $\lambda_{it} \geq 0$ . Banks might also pass through higher deadweight costs of bankruptcy in the form of more severe credit constraints (lower  $\lambda$ ). As a result, a firm might prefer to pay a higher interest rate in order to obtain a larger loan, but charging higher interest rates would conflict with the purpose of the bank and its classification scheme. We return to this point below.

In this environment, the investment optimality condition is characterized by the following equation<sup>25</sup>

$$\begin{aligned} \rho \int_{\bar{\omega}}^{\infty} [MRP_{it+1}^K - (\bar{r}_{t+1} + \delta)] d\Phi(\omega_{it+1}|\omega_{it}) &= \psi_2^K(K_{it}, K_{it+1}) + \rho \int_{\bar{\omega}}^{\infty} \psi_1^K(K_{it+1}, K_{it+2}) d\Phi(\omega_{it+1}|\omega_{it}) + \\ &\quad \chi_{it}(1 - \lambda_{it}) \\ &\quad \equiv \tau_{it}^K. \end{aligned} \tag{1}$$

where  $\Phi(\omega_{it+1}|\omega_{it})$  denotes the conditional density function of TFPR. The left-hand side represents the difference between the marginal revenue product of capital and the user cost of capital ( $r_{it+1} + \delta$ ). On the right-hand side, the first line denotes real adjustment costs of capital (Cooper and Haltiwanger 2006).  $\psi^K(K_{it}, K_{it+1})$  is an adjustment cost function of capital, and  $\psi_j^K(\cdot)$  the derivative with respect to its  $j$ th argument. The existence and impact of these costs on investment policies might be related to firms’ lifecycle (e.g., age and size) or product market conditions (Asker et al. 2014). The second term -  $\chi_{it}(1 - \lambda_{it})$  - is the *shadow cost of capital* firms face. In the presence of binding credit constraints, the gap between the marginal revenue product and the user cost of capital is an increasing function of the multiplier attached to the borrowing constant ( $\chi_{it} \geq 0$ ) and of the tightness of the constraint ( $1 - \lambda$ ). We group the terms on

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We assume TFPR evolves stochastically following a first-order Markov process. Because today’s investment becomes productive tomorrow ( $K_{it+1} = (1 - \delta)K_{it} + I_{it}$ ), uncertainty about the realization of  $\omega_{it+1}$  generates idiosyncratic investment risk, which makes capital and debt imperfect substitutes in the firms’ problem and generates endogenous default risk.

<sup>23</sup>This assumption is largely consistent with what we find in our data, in which over 99% of firms are not listed in the stock market, and 80% of the firm-year observations borrow from financial institutions to finance their operations. Of the remaining 20% of the observations, 80% finances capital expenditure with some combination of self-financing and trade credit, and less than 5% uses only capital from shareholders, either in the form of debt from shareholders or in-kind contributions. A large literature addresses the evidence of, reason for, and consequence of the limitation of equity financing (e.g., Greenwald et al., 1984; Myers and Majluf 1984; Stiglitz 1992).

<sup>24</sup>The interest rate  $\bar{r}_{it+1}$  and the tightness of the borrowing constraint  $\lambda_{it}$  are set jointly to maximize bank profits when lending to firms similar to firm  $i$ . Pooling observationally similar borrowers, banks set the interest rate based on the expected probability of default for firms similar to firm  $i$ , and may cope with risk imposing a borrowing constraint that links credit supply to firms’ net worth. For every group of similar borrowers, banks can choose multiple lending contracts, defined by the pair  $(\bar{r}, \lambda)$ . For example, a competitive lender might follow a two-step optimization process. As a first step, interest rates are chosen to maximize expected profits from borrowers similar to type  $i$ . Then  $\lambda_{it}$  is chosen to satisfy the zero profit condition, irrespective of firm-specific productivity, which is unobservable to the bank. A similar two-step optimization can be followed by a monopolistic competitive lender that faces a downward-sloping residual demand for its financial services.

<sup>25</sup>Gopinath et al. (2016) derive a similar expression in a model with no default risk.

the right-hand side and denote them by  $\tau_{it}^K$ . Abstracting from the impact of adjustment costs, MRP-cost gaps are positive for credit-constrained firms. Their magnitude is proportional to the degree of credit market frictions (e.g., asymmetric information frictions and bankruptcy costs) individual producers face.

Similarly, we express the first-order condition that characterizes optimal employment policies isolating the difference between the Marginal Product of Labor and its user cost ( $w_{it}$ ) from a residual quantity  $\tau_{it}^L$

$$\begin{aligned} MRP_{it}^L - w_{it} &= \psi_2^L(L_{it-1}, L_{it}) + \rho \int_{\bar{\omega}}^{\infty} \psi_1^L(L_{it}, L_{it+1}) d\Phi(\omega_{it+1} | \omega_{it}) \\ &\equiv \tau_{it}^L. \end{aligned} \quad (2)$$

Intuitively, when labor is flexibly hired on the spot market after the realization of productivity, firms choose labor demand equalizing the marginal revenue product of labor to the wage rate. The presence of labor adjustment costs ( $\psi^L(L_{it-1}, L_{it})$ ) invalidates this neoclassical prediction (Cooper and Willis 2009).<sup>26</sup> For our purposes, the incidence of adjustment costs that vary as a function of firm size is particularly relevant. The following adjustment cost function models the size-dependent provisions of Article 18 of the Italian Worker Statute:

$$\psi^L(L_{it-1}, L_{it}) = \begin{cases} \frac{c^L}{2} \frac{(\Delta L_{it})^2}{L_{it-1}} & \text{if } L_{it-1} < \bar{L} \\ (\mathbf{1}_{[\Delta L_{it} < 0]} f_i^L) \Delta L_{it} + \frac{c^L}{2} \frac{(\Delta L_{it})^2}{L_{it-1}} & \text{otherwise,} \end{cases}$$

where  $f_i^L$  is a size-dependent, government-mandated severance payment that firms with a workforce larger than  $\bar{L}$  (=15, under Article 18) have to pay to laid-off workers. Because firing costs are only born by companies whose employment is above  $\bar{L}$ , MRP-cost gaps for labor are expected to display a discontinuous behavior around the threshold.

**Discussion** – The characterization of firm policies in terms of MRP-cost gaps is convenient. From an empirical point of view, realized MRP-cost gaps are measurable quantities, once estimates of marginal revenue products and information of user costs are available. Thus, they can be used to cast light on the distribution of the unobservable residuals  $\tau_{it}^K$  and  $\tau_{it}^L$ , and to test the incidence of specific types of frictions and regulations that affect firm policies. On the capital-side, the gap  $\tau_{it}^K$  is a particularly valuable empirical tool for investigating the efficiency of investment policies for privately owned firms. For them, traditional measures, such as Tobin’s Q or indexes of financial constraints (e.g., Kaplan and Zingales (1997) or Whited and Wu (2006)), are not computable because no information is available about the market value of firm’s assets and liabilities.<sup>27</sup>

It is important to emphasize that the role plaid by price adjustments, or lack of adjustment thereof, in interpreting the sign and magnitude of MRP-cost gaps. As shown in equations (1) and (2), when borrowing costs and wages do not vary to accommodate factors demands of heterogeneous producers, MRP-cost gaps capture the pass-through of credit and labor market frictions to firm policies via distorted accumulation of capital and labor. The interpretation of  $\tau^K$  and  $\tau^L$  changes when prices are the instruments that allocate resources in capital and labor markets.

In Appendix C, we consider a credit contract by which banks do not constraint their credit supply but adjust the interest rate as a function of firm characteristics (bank leverage, capital endowment, productivity)

<sup>26</sup>González and Miles-Touya (2012) and Petrin and Sivadasan (2013) construct similar value of the marginal product minus wage gaps to study the impact of changes in the employment protection regulation in Chile and Spain.

<sup>27</sup>Unable to observe user costs of capital and/or labor, previous literature has frequently relied on both time-specific effects and firm-specific effects in the empirical specifications to control for the variation in these terms. See Bond and Meghir (1994) or Whited and Wu (2006). A noteworthy exception is Philippon (2009), who proposes an implementation of the q-theory of investment using variation in bond prices.

and in response to credit market frictions (bankruptcy costs):  $r_{it+1} = r(K_{it+1}, B_{it+1}, \omega_{it}, X)$ . Under this credit contract, anything that affects the firm-specific likelihood of default, cost of credit provision, or loss given default affects individual firms' investment decisions and the allocation of credit through an adjustment of the cost of credit. In this case, the term  $\chi_{it}(1 - \lambda_{it})$  is replaced by the term  $\left(\frac{\partial r_{it+1}}{\partial K_{it+1}} + \frac{\partial \tilde{r}_{it+1}}{\partial B_{it+1}}\right)$ , and positive gaps would no longer signal constrained access to credit.

Similarly, the characteristics of the wage contract affect the interpretation of the labor gap. Since the seminal work of Lazear (1990), it is well known that, in the absence of contractual and market frictions, the transfer  $f_t^L$  can be neutralized by an appropriately designed wage contract: the firm reduces the entry wage of the worker by an amount equal to the expected present value of the future transfer, so as to leave the expected cumulative wage bill arising from the employment relationship unchanged. On the contrary, when wages are inflexible, firms resort to quantity adjustments that are then reflected in the distribution of  $\tau^L$ .

The literature suggests several explanations for why the price terms in credit and employment contracts might be rigid. Prominent examples are asymmetric information frictions (Stiglitz and Weiss, 1981, 1992; Campbell III and Kamlani 1997), imperfect competition (Petersen and Rajan 1995; Ashenfelter et al. 2010), and government interventions that prevent or limit price discrimination, forcing sellers/buyers of credit or labor services to charge/demand the same price in types of transactions that are intrinsically different (Calmfors and Horn 1986; Benmelech and Moskowitz 2010; Banerjee and Duflo 2014; Hurst et al. 2016). In section 5, we document a relative stickiness of interest rates and wages, and provide evidence that is consistent with credit limits and workforce adjustments as the primary margin of adjustment in response to credit and labor market frictions.

Finally, note that besides credit and labor markets frictions, other phenomena contribute to the size and dispersion of realized MRP-cost gaps. Equations (1) and (2) highlight that economic uncertainty and real adjustment costs naturally drive a wedge between realized marginal revenue products and user costs. Also, market power and imperfect competition, heavy taxation, the bureaucratic costs of doing business, tariffs and subsidies, and frictions in the market of corporate ownership and control also drive a wedge between user costs and marginal revenue products of production factors (see the review in Restuccia and Rogerson (2013)). We design empirical tests that allow us to disentangle the effect of these alternative phenomena from the extent of credit and labor market frictions that individual firms face. For credit market frictions, we focus on the relevance of asymmetric information, bankruptcy costs, and idiosyncratic shocks to credit supply. For labor market frictions, we study the static and dynamic effects of the provisions of Article 18 of the Italian Workers' Statute.

## 4 The Distribution of MRP-cost gaps in the Micro Data

In this Section, we describe the empirical procedure that allows us to produce measurable counterparts of the MRP-cost gaps in equations (1) and (2). A unique feature of our database is the availability of information on *both* firm-specific wages and interest rates, collected from highly reliable administrative sources, for the lion-share of the corporate sector of a country. This feature gives us a significant edge in obtaining measurable proxies of the distribution of firms' user costs of capital and labor. We estimate realized marginal revenue products of capital and labor following the literature on production function and markup estimation (Gandhi et al. 2017b; De Loecker and Warzynski 2012). Thus, the empirical counterparts of MRP-cost gaps are

$$\hat{\tau}_{it}^K \equiv \rho(1 - \hat{P}\{Exit_{it+1}|X_{it}\}) \cdot \left[\widehat{MRP}_{it+1}^K - (r_{it+1} + \delta_s)\right] \quad (3a)$$

$$\hat{\tau}_{it}^L \equiv \widehat{MRP}_{it}^L - w_{it}. \quad (3b)$$

The discount factor is set to  $\rho = 0.95$ , a standard assumption in the literature (Gopinath et al. 2016). In order to approximate the conditional expectation in Equation (1), we evaluate the expectation of the marginal revenue product minus user costs gap at their realizations, adjusting the latter by multiplying them by the expected probability of exit ( $P\{Exit_{it+1}\}$ ). This procedure naturally introduces an expectational error that is going to generate variation in the estimated MRP-cost gaps (Asker et al. 2014). The probability adjustment accounts for the fact that, because today’s investments become productive with a lag, expected returns are lower for firms with higher exit probability.

We estimate expected probabilities of exit via a probit model. The left-hand side variable is an indicator function equal one when we observe the firm exiting in year  $t + 1$ .<sup>28</sup> The explanatory variables include the set of state variables of the firm problem as well as a set of firm-specific and macro-financial variables that allow us to better capture firms’ expectations (see Appendix G for details). Our estimates of the unconditional probability of exit is 7.3% on average, matching the unconditional exit rate in our sample. In line with the guidelines of economic theory, the estimated exit probability is decreasing in firm’s size, age, productivity, and credit rating. It is higher for more leveraged firms and for those producers that defaulted on their debt obligations.

In the remaining of this Section, we describe our proxies of the user costs of capital and labor, illustrate the estimation procedure of marginal revenue products, and finally present the estimates of the MRP-cost gaps.

## 4.1 User costs of Capital and Labor

### 4.1.1 The User Cost of Capital

We construct firm-time varying user costs of capital as the sum of borrowing costs and depreciation rates of fixed assets ( $r_{it+1} + \delta$ ). Industry-specific depreciation rates ( $\delta$ ) are collected from the Italian Statistical Agency (National Accounting Tables). To measure the borrowing rate  $r_{it+1}$ , we use the Average Percentage Rates (APR) on firm-bank matched loans from the Credit Registry (Taxia database).<sup>29</sup> While alternative credit products are available to firms, bank loans represent around 3/4 of total bank debt and they are the typical credit product used to finance expenditures in fixed assets.<sup>30</sup> We calculate the firm-year-level APR as follows. When multiple banks are lending to the firm, we compute the weighted average APR with weights

<sup>28</sup>The dummy variable  $EXIT_{it+1}$  takes value one in year  $t$  when a firm does report any balance sheet and income statements from year  $t + 1$  onward. It also takes value equal one for firms that report in  $t + 1$  zero amounts of two of the three production inputs: capital stock, wage bill expenses, or purchases of intermediate inputs.

<sup>29</sup>The Taxia database covers the large majority of the financial intermediaries operating under the supervision of the Bank of Italy. Until 2003, the subgroup of banks in Taxia was composed by around 90 banks, accounting for more than 80% of total bank lending. Starting from 2004, the pool of banks in Taxia sample has been expanded to 103 national banks and 10 branches and subsidiaries of foreign banks. Banks in the Taxia sample must report information on the APR charged to every borrower if the total amount of credit granted plus guarantees provided by the borrower exceeds 75,000 euro. Taxia allows us to distinguish between two types of loans: term loans and loans backed by account receivables. We use the APR on term loans as a baseline rate, and the APR on credit backed by receivables when the interest rate is not available (25% of the cases). 73% of firms with an outstanding balance of term loans also have outstanding credit backed by receivables. In terms of observable characteristics, firms that use only one or the other credit product are similar. Two APR are similar in level (on average: 5.6% term loans, 6.2% credit BAR), highly correlated with each other (raw correlation 39%), and they correlate in the same way with firm characteristics.

<sup>30</sup>Appendix B.1 shows that changes in bank loans can explain a larger share of the variation in investment rates and that the elasticity of investment with respect to changes in loans is three times as large as the elasticity with respect to changes in credit line draws.

equal to the fraction of total loans granted by each institution. When a firm has only one outstanding loan from a single bank, no aggregation is needed.<sup>31</sup>

Firms that do not actively engage in credit market transactions (20% of our sample) pose an empirical challenge because we have no information on borrowing rates for them. These observations are of interest since they allow us to investigate the relationship between credit market participation and firm policies. Thus, we would like to construct a plausible estimate of their user cost. There is ample empirical evidence, corroborated by the analysis of Section 5, that banks set their rates based on a limited number of observable characteristics (Jaffee and Modigliani 1969; Crawford et al., 2016). Moreover, it is well established that financing of small and medium firms - the lion share in our data - is tied to their local credit markets as proximity between borrowers and lenders facilitates information acquisition (Petersen and Rajan, 2002; Degryse and Ongena 2005). We use firm characteristics and geographical location to infer the interest rate that non-borrowers could have been charged had they engaged in credit market transactions. Within each year and local credit market - defined by the perimeter of Italian provinces -, we estimate loan pricing regressions in the relationship level database. The set of predictors includes industry, age, assets, credit score, assets turnover, ROA, and whether the firm has any credit in default during that year of the previous ones.<sup>32</sup> These variables are selected to meet two criteria. On the one end, they represent a parsimonious choice that ensures the existence of a common support between the group of borrowers and non-borrowers for every year-market combination. On the other hand, they are observable indicators commonly used by banks to assess firms' riskiness and creditworthiness. The Altman Z-score is a widely used metric used by Italian banks to assess firms credit risk (Albareto et al. 2011) and the Cerved database is the source of firms' balance sheet information used by banks to collect balance sheet data on current and perspective borrowers. Moreover, our data on firms total debt exposure is the same information that banks can obtain when they send a query to the CR.<sup>33</sup> The pricing regression is estimated on the subsample of newly established relations ( $\text{LENGTH RELATION}_{it} \leq 1 \text{ year}$ ).<sup>34</sup> Appendix B.2 provides a detailed description of this procedure and provides a number of robustness tests of our estimates.<sup>35</sup>

Firms that engage in credit market transactions, but for which we are unable to observe the interest rate, represent a second empirical challenge for the construction of the user cost of capital. These observations refer to firms that only use credit lines; to those firms borrowing from lenders are not part of the group of banks in the Taxia database; or to firms that borrow small amounts that are not reported in the CR.<sup>36</sup> For

<sup>31</sup>That is, we calculate the value-weighted average APR for each firm-year as  $r_{it+1} = \sum_b w_{ibt} r_{ibt+1}$ , where  $w_{ibt} = \text{Loans}_{ibt} / \sum_b \text{Loans}_{ibt}$ . When we observe multiple APRs for the same firm-bank pair, we calculate the weighted average using as weights the share of interest expense imputable to each loan. See Appendix A for details.

<sup>32</sup>Italian provinces are the natural candidates for the definition of local credit markets for small business lending (see Guiso et al. 2012). They constitute administrative units comparable to US counties. The the Bank of Italy uses the administrative boundaries of provinces as a proxy of local credit markets for regulatory and supervisory purposes.

<sup>33</sup>Through the Central Credit Registry financial companies supervised by the Bank of Italy exchange information about the global risk position (total outstanding banks credit and total credits in default) of their customers and of those of other institutions. After it receives information on the loans granted by the participating intermediaries to individual customers, the Bank of Italy aggregates the data for each borrower and calculates their total debt exposure vis-a-vis the financial system, and possible amounts past due or in default.

<sup>34</sup>The focus on new relationships is important because non-borrowers would be new customers for the bank in case they approach them. Moreover, for new relationships, we do not have to account for the dynamics of firm-bank relationships, and the acquisition of soft information and lower monitoring costs that repeated interactions bring about.

<sup>35</sup>Matching-on-observables raises concerns related to unobserved heterogeneity - as soft information might be available to the bank but not to the econometrician -, and to possible selection issues, since only transactions for which borrowing/lending is economical for both firms and banks are observed. We discuss these issues in Appendix B.2.

<sup>36</sup>The debt amounts from the Credit Registry are recorded at a monthly frequency. To harmonize them with firms' annual balance sheets, we calculate the average credit exposure of a firm across all lenders in each fiscal year. Intermediaries report to the Credit Registry any relationship with a client whose total amount of credit granted plus guarantees provided by the borrower exceeds 30,000 euro (75,000 euros before 2008). See Appendix A for details.

these observations, the missing price problem is less severe because, beside firm-specific characteristics and geographical location, we can augment the pricing regressions with information about total bank leverage, the length of each individual credit relation, the total number of lending relations, and dummies that identify lenders (see Appendix B.2).

Table 2 (panel a) presents summary statistics describing the distribution of user costs of capital and its components. We present them for the whole sample, and splitting observations into borrowers with outstanding loans (BORROWERS-LOANS), borrowers with no loans (BORROWERS-NOLOANS, i.e. firms with no outstanding loans but positive draws from credit lines), and non-borrowers (NON-BORROWERS). For observations belonging to the first subsample, the interest rate is observed; for the last two groups, we report the estimated interest rate. Consider first the subsample of borrowers with loans. Over our sample period, their user cost of capital was on average 16.4%. One-third of it is imputable to the borrowing cost (5.5%), and two thirds to depreciation rates (10.8%). On average, the borrowing costs inferred for credit lines-only borrowers and for non-borrowers are respectively 40 and 130 basis points higher than the ones observed for borrowers, reflecting the compositional differences among the observations that form the three sub-samples. In Appendix A we show that, compared to firms with outstanding bank loans, producers that do not engage in credit market transactions and those who only used credit lines are younger and smaller; over-represented in Southern regions of Italy, and in industries with lower tangible to intangible assets ratio (such as services).<sup>37</sup> Credit lines-only firms also tend to have shorter lending relationships with their lenders when compared to companies that utilize bank loans. Not by chance, all these variables are commonly regarded as proxies of credit constraints.<sup>38</sup> Consistent with this, the empirical analysis of Section 5.1 finds that firms that do not engage in credit market transactions and credit lines-only borrowers tend to have a higher marginal revenue product of capital than borrowers with outstanding term loans.

#### 4.1.2 The User Cost of Labor

Employer-employee records from the Italian National Security Institute provide us with detailed information on workforce compensation. We use the average annual wage as a proxy for the user cost of labor  $w_{it}$ . We calculate it considering the annualized compensations of all fixed-term contract workers (white collars, blue collars, middle managers, and full-time interns) hired by the firm throughout the year.<sup>39</sup> Table 2, panel a shows that the average nominal wage is about 19 thousand euros per year, the median is one thousand euros lower.

One may worry that the average wage may differ significantly from the wage paid to higher an extra worker. To address this, we construct an alternative proxy of the user cost of labor using individual workers' wage records from the matched employer-employee panel database. In particular, we calculate the average annualized wage paid by firms to *newly hired* workers in each industry-province-year triplet.<sup>40</sup> The advantage of this measure is that it can be thought of as the cost that a firm would incur when hiring an additional worker in the same industry and labor market. The drawback of this measure is that, by averaging across

<sup>37</sup>See Appendix A for a comparison of borrowers and non-borrowers based on observable characteristics.

<sup>38</sup>Because credit lines are a more expensive type of credit and they can be revoked at lenders' discretion, firms should rarely turn to credit lines to finance capital expenditures in fixed assets, unless bank loans are constrained or denied by credit institutions.

<sup>39</sup>The firm-level records are aggregated by the Italian National Security Institute and provided to us at a monthly frequency. For each firm-year observation, we first calculate the average monthly wage (simple average) and then we annualize it. While not perfect, this procedure is better than using the annualized end-of-year wage (month of December) because end-of-year compensations are more likely to be susceptible to *una tantum* adjustments.

<sup>40</sup>The employer-employee matched database follows the employment history of a random sample of 20% of every cohort of workers. In our dataset, the subsample of firm-year observations that (i) hires new workers and (ii) for which we have information on at least one wage rate of the newly hired workers from the employer-employee database is 48%.

companies, it washes away any firm-level link between wages and the marginal product of labor. We find that the average wage paid to new workers exceeds the average wages by approximately four thousand euros (18% of the average wage). As we discuss below, our main empirical findings are ultimately unaffected by using this alternative proxy of user costs of labor.

## 4.2 Identification and Estimation of Marginal Revenue Products

Let us consider a firm facing a downward sloping demand curve, and posit a general characterization of production technologies:

$$Q_{it} = e^{\omega_{it} + \epsilon_{it}} F(K_{it}, L_{it}, M_{it}; \gamma),$$

where  $\gamma$  is a vector of structural parameters to be estimated.  $\omega_{it}$  is firm-level productivity, observed by the firm at the moment of its production decisions, and  $\epsilon_{it}$  is a production shock taking place after input decisions have been made. Without loss of generality, we can decompose the marginal revenue product of an input  $X = \{K, L, M\}$  into the Value of the Marginal Product ( $VMP_{it}^X$ ) and the inverse-markup ( $\mu_{it}^{-1}$ )

$$MRP_{it}^X \equiv \frac{\partial (P_{it}(Q_{it})Q_{it})}{\partial X_{it}} = \underbrace{P_{it} \frac{\partial Q_{it}}{\partial X_{it}}}_{VMP_{it}^X} \underbrace{\left(1 + \frac{Q_{it}}{P_{it}} \frac{\partial P_{it}}{\partial Q_{it}}\right)}_{\mu_{it}^{-1}} = \theta_{it}^X \frac{PQ_{it}}{X_{it}} \frac{1}{\mu_{it}}. \quad (4)$$

The last equation decomposes the physical Marginal Product Value into output elasticity ( $\theta^X$ ) and average product ( $\frac{PQ_{it}}{X_{it}}$ ) using the definition of output elasticity. We estimate marginal revenue products taking Equation (4) to the data.

We measure average products of capital ( $PQ_{it}/K_{it}$ ) and labor ( $PQ_{it}/L_{it}$ ) directly in the data.  $PQ_{it}$  is total sales. The ideal empirical measures of capital ( $K_{it}$ ) and labor ( $L_{it}$ ) shall capture the flow of services provided by these inputs. Toward this end, we re-construct the sequence of capital from investments in fixed assets (both tangibles and intangibles) following the Perpetual Inventory Method (Becker and Haltiwanger 2006) and measure labor services in units of effective labor (annual wage bill over average annual wage). The Perpetual Inventory Method provides us with a better proxy of capital services than the book value of physical assets.<sup>41</sup> With respect to other measures - such as the number of workers -, by measuring labor services in effective labor unites we can better accounts for differences in the quality of firms' workforce (Fox and Smeets 2011).<sup>42</sup>

**Output elasticities** – We estimate output elasticities via production function estimation following the structural approach proposed in Gandhi et al. (2017b).<sup>43</sup> This approach identifies the parameters of the production function addressing the simultaneity bias that derives from the correlation between input choices and unobserved (to the econometrician) productivity (Marschak and Andrews 1944), and it solves the non-identification problem that affects the estimates of output elasticity with respect to flexible inputs.<sup>44</sup>

<sup>41</sup>See Appendix A.6 for details on the construction of the capital sequence using the Perpetual Inventory Method (PIM).

<sup>42</sup>Using total wage bill as a measure of labor inputs delivers estimates very similar to the ones obtained using effective labor. Results are available upon request.

<sup>43</sup>We provide the details of the estimation routine in Appendix D and refer to Gandhi et al. (2017b) for a more detailed exposition and its underlying assumptions. We thank the authors of Gandhi et al. (2017b) for sharing their code, and to David Rives for his practical advice.

<sup>44</sup>Gandhi et al. (2017b) shows that the standard proxy-variable approach applied to gross output production functions does not identify the elasticities of flexible inputs, unless the production function takes specific functional forms (e.g. the Leontief case discussed in Akerberg et al. (2015)) or external sources of variation in firms' demand for flexible inputs (e.g. Doraszelski and Jaumandreu (2013)).



We specify a Translog functional form for production technologies  $f$ , parametrized by the vector  $\gamma$ . For the purpose of approximating the full distribution of marginal revenue products, the flexibility of Translog represents a significant advantage over more standard (but less flexible) functional forms such as Cobb-Douglas or CES. Translog does not impose any restriction on the elasticity of substitution of different inputs. Moreover, it allows us to recover a distributions of firm-time specific elasticities that are a function of industry-specific structural parameters  $\gamma$ s and of the input-mix utilized by each firm:  $\theta_{it}^X = \theta^X(k_{it}, l_{it}, m_{it}; \gamma)$   $X = \{K, L, M\}$ .<sup>45</sup>

The production function estimation is performed separately for every four-digits industry (NACE, rev.2 industry classification system). This allows the structural technology parameters  $\gamma$ s to vary by narrowly defined industries (467 in total) that encompass both the manufacturing and non-manufacturing sectors of the economy. We use deflated revenues in place of physical output, and deflate capital and intermediate inputs (measured as total expenditures in raw materials, services, and energy consumption) by the corresponding industry-year price deflators.<sup>46</sup> Finally, we need to take a stand on the vector of instruments that identify  $\theta^K$  and  $\theta^L$  in the estimation routine. We assume capital is quasi-fixed and predetermined.<sup>47</sup> Thus, in principle,  $k_{it}$  does not require an instrument. Nevertheless, we use (lagged) firm-specific borrowing costs to construct an additional moment condition that strengthens the identification of the elasticity  $\theta^K$ , which typically suffers from attenuation bias due to the difficulty to measurement capital services (Collard-Wexler and De Loecker 2016).<sup>48</sup> Given the institutional features of the Italian labor market, we consider labor a flexible input (chosed in period  $t$  after observing  $\omega_{it}$ ) but dynamic (subject to adjustment costs). Thus, we rely on  $l_{it-1}$  as an instrument for  $l_{it}$  and address the endogeneity problem due to correlation with unobserved productivity.

Once estimates of the structural parameters  $\gamma$  are available, we infer the realization of firm-level revenue productivity (TFPR, Foster et al. (2008)) as

$$(\omega_{it} + \epsilon_{it}) = q_{it} - f(k_{it}, l_{it}, m_{it}; \gamma),$$

where lower case letters denote the log of variables. With a slight abuse of notation, we denote  $(\omega_{it} + \epsilon_{it})$  with  $\omega_{it}$ .

**Markups** – To estimate markups we follow the production side approach pioneered by the seminal work of Hall (1988) and recently revisited by De Loecker and Warzynski (2012). The identification rests on the theoretical intuition that, conditional on the state variables of the problem, the first-order conditions of the cost minimization problem for inputs that are flexible and static provides an expression relating revenue cost shares and output elasticities to markups:

$$\hat{\mu}_{it} = \hat{\theta}_{it}^M \left( \frac{P_{it} Q_{it}}{\exp(\hat{\epsilon}_{it})} \frac{1}{P_{it}^M M_{it}} \right),$$

where  $P_{it} Q_{it} / P_{it}^M M_{it}$  is the inverse of the expenditure share on intermediate inputs in revenues (directly observed in the data) and  $\hat{\theta}_{it}^M$  is the output elasticity with respect to intermediate inputs (obtained via production function estimation). We follow De Loecker and Warzynski (2012) and correct expenditures

<sup>45</sup>Consider the following log-version of production functions:  $q_{it} = f(k_{it}, l_{it}, m_{it}; \gamma) + \omega_{it} + \epsilon_{it}$ . Under Translog, the expression for output elasticities of any input  $X = \{K, L, M\}$  is  $\theta_{it}^X = \gamma_X + 2\gamma_{XX} x_{it} + \sum_{x' \neq X} \gamma_{xx'} x'_{it}$ . See Appendix D for details.

<sup>46</sup>While unsatisfactory, this is the predominant approach in the Industrial Organization literature since most of the available firm- and plant-level database, including ours, do not separately report prices and physical quantities of inputs (with the exception or labor, in our case) and/or output. We use industry-specific investment deflators for capital, and industry-specific value added deflators for intermediate inputs. These data are freely available on the website of the Italian National Statistical Agency (<http://dati.istat.it/?lang=en>).

<sup>47</sup>This is a standard assumption, consistent with the capital accumulation equation:  $K_{it} = I_{it-1} + (1 - \delta)K_{it-1}$ .

<sup>48</sup>See Appendix Appendix D for more details, and Doraszelski and Jaumandreu (2013) for a discussion of how information factor prices can be used to identify production functions.

shares using the residuals of a regression of a polynomial function of deflated inputs on deflated revenues. This adjustment helps to net out variation in output not correlated with changes in input utilization (such as the one due to demand, inputs prices, or productivity).<sup>49</sup> The flexibility of the Translog functional form adopted in the production function estimation also helps to address this issue.

Table 3 displays our estimates of elasticities, returns to scale, markups, and productivity. Block-bootstrapped standard errors are reported in parenthesis (Horowitz 2001). The deflated revenues of the average firm responds by 4%, 29% and 67% to a one-percent increase in capital, labor and intermediate inputs, respectively, which implies average local returns to scale close to unity. These parameters are precisely estimated and in line with the ones found in the literature.<sup>50</sup> Importantly, our estimates highlight substantial heterogeneity in the parameters characterizing production technologies, both within and across industries.<sup>51</sup> The interquartile range spans between 57% and 79% for intermediate inputs, and 2%–6% and 18%–38% for capital and labor, respectively. In term of markups, our estimates suggest that, on average, firms price 2% above their marginal cost of production. The right skewness of the distribution drives the dispersion of markups. Firms located at the 75th and 90th percentile of the distribution price 5% and 15% above marginal cost, respectively.

In the Appendix of the paper, we present a number of sanity and robustness checks on our estimates. Appendix D shows that the estimates output elasticities are consistent with the ones obtained using a cost-share approach (Hall et al., 1986). We also discuss the robustness of our estimates with respect to alternative functional forms of production technologies and estimation routines. In Appendix E we conduct a series of robustness checks of our estimates of markups. We find a strong positive correlation between markups and firm’s profitability (either EBITDA over total assets or ROA), and with product market concentration measured by the Herfindahl concentration index. Our estimates of firm-level markups also display a strong and positive correlation with productivity (in both levels and changes), which is an empirical relationship documented by previous literature (see De Loecker and Warzynski 2012).

**Marginal revenue products** – Combining average products, output elasticities, and markups, we construct estimates of realized marginal revenue products of  $K$  and  $L$  (Equation 4). Table 2 (panel b) reports descriptive statistics of their distribution. Over the 1997-2013 period, the median firm in our dataset has a marginal product of capital of 21%, while that of labor is slightly lower than 25 thousand euros.<sup>52</sup> We point out that the estimated marginal revenue product of capital is 1.5 times higher for non-borrowers and 0.5 times higher for those borrowers that access only to credit lines, which suggests that constrained access to credit markets might prevent some firms from harvesting profitable investment opportunities. We will return to the distinction between the three groups of firms in Section 5.

Finally, two remarks are in order. First, we treated deflated sales as a measure of physical quantity when estimating output elasticities. Therefore our estimates are potentially subject to the omitted price variable bias discussed in Klette et al. (1996), and our estimates of productivity are a proxy for revenue productivity (TFPR). Not controlling for firm-specific output prices would be particularly problematic if estimating physical productivity (TFPQ) was the ultimate goal of this paper (Foster et al. 2008). It is less of

<sup>49</sup>See Appendix E for more details and De Loecker et al. (2016) for a discussion and application of this methodology.

<sup>50</sup>See for example De Loecker (2011), Akerberg et al. (2007), Petrin and Sivadasan (2013), and Gandhi et al. (2017b) for estimates referring to manufacturing industries.

<sup>51</sup>Appendix D provides a graphical comparison of output elasticities across firms of different age and size. We find a significant decline of  $\theta^K$  with firm size and age, while  $\theta^L$  increases as firms grow older but decrease with firm size.

<sup>52</sup>In Appendix F, we also investigate the sources of dispersion of  $MRPs$ . Two findings are worth mentioning. First, marginal revenue products are more dispersed outside manufacturing. Second, the bulk of the dispersion in  $MRPs$  is found within industries rather than between industries. The within-industry dispersion exceeds the between industry dispersion by a factor of two for  $MRP^K$  and a factor of 1.4 for  $MRP^L$ .

a concern for our analysis because TFPR is the relevant productivity measure to test the theory underlying the MRP-cost gaps (see Section 3).<sup>53</sup> Second, we must also recognize that our data does not allow to distinguish between single and multi-product firms. If firms operate across multiple industries or produce differentiated goods, our estimates might be biased because the estimation routine implicitly assumes a single production function and a single consumer’s demand curve faced by each firm (see Bernard et al. 2010 and De Loecker 2011). We cannot identify which companies operate across industries because our data reports only the primary industry code of each observation. However, because large firms are more likely to expand their activity across industries, the small size of the producers in our data suggests that multi-product firms are unlikely to be the majority of our sample.

### 4.3 The Variability of User Costs and Marginal Returns, and the Empirical Distribution of MRP-cost Gaps

**Dispersion in MRP and User Costs** – Before presenting the empirical distribution of the MRP-cost gaps, it is instructive to analyze the joint distribution of user costs and marginal revenue products of capital and labor. In Figure 1, we parse the data according to the percentile of the distribution of user costs of capital (panel a) and labor (panel b). For each percentile, the x-axis reports the median value of the user cost. The y-axis reports the median value and interquartile range of the MRP for the group of firm-year observations belonging to each percentile of the distributions of user costs. Two observations are in order.

First, the central percentiles of the distribution of *MRPs* map onto the corresponding moments of the distributions of the user costs. The correlation between the median (mean) value of MRP and the median (mean) value of user costs within each percentile of the distribution of user cost is 98% (95%) percent for capital and 98% (96%) for labor, with p-values lower than 1%.<sup>54</sup> This finding suggests that user costs are an economically meaningful benchmark for the realized marginal revenue products of capital and labor of individual producers, as profit maximization predicts.

Secondly, the large dispersion of marginal revenue products is in stark contrast with the fairly symmetric and compact distribution of user costs. This is particularly evident in the case of capital. For example, the variation in realized  $MRP^K$  within each percentile of the distribution of the  $r + \delta$  is greater than the unconditional variation of  $r + \delta$ . A similar observation applies to the dispersion in  $MRP^L$  and wages.

**Distribution of MRP-cost Gaps** – We combine the estimate of MRP and the observed user costs, to produce empirical counterparts of MRP-cost gaps (Equations (3a) and (3b)). To limit the impact of outliers, we winterize the 2.5 tails of the distribution of  $\tau_{it}^K$  and  $\tau_{it}^L$ . Table 2, panel c reports summary statistics of our estimates. Figure 2 displays their full distribution.

According to our metric, the central percentiles of the distributions are occupied by firms whose capital and labor endowment appears to be relatively undistorted. The median gaps of capital and labor are 3.5% and 6 thousand euros for capital and labor, respectively.

Yet, the distributions of MRP-cost gaps are dispersed and highly right-skewed, reflecting the right-skewness of the corresponding distributions of marginal revenue products. In fact, the average capital and labor gaps are 37% and 9 thousand euros, respectively; the 90-10 percentile differences are almost 3 times larger.

<sup>53</sup>It must be kept in mind, however, that the inability to control for heterogeneous prices may also generate a downward bias in the estimates of output elasticities (De Loecker 2011) and, thus, of our estimates of marginal revenue products.

<sup>54</sup>The correlation between marginal revenue products and user costs is economically and statistically significant also at the firm-level (6% capital and 37% for labor, p-values lower than 1%).

**Correlation with observable characteristics** – The large dispersion in marginal revenue products could be entirely driven by measurement error or production function misspecification. Alternatively, as discussed in section 3, market frictions might distort the quantity of capital and labor employed by firms, and generate a dispersion of marginal revenue products above and beyond the variation observed in user costs. The correlation between MRP-gaps with firms’ observable characteristics provides preliminary evidence in this direction.

We regress gaps on life cycle variables (firm age and size), credit score, measures of productivity and profitability (TFPR and ROA), and proxies of internal and external financing (cash-over-assets and bank leverage, respectively).<sup>55</sup> We focus on within-year and within-industry variation by controlling for year and industry fixed effects. Table 4 reports the regression results: panel a for capital and panel b for labor. Because  $\tau^K$  and  $\tau^L$  have different variability, coefficients are expressed as Z-scores to facilitate their comparison across the two panels.

MRP-cost gaps of capital monotonically decrease with firm age and size. In contrast, labor gaps are higher for larger and older firms. The availability of financing, either internally generated liquidity or bank debt, is negatively correlated with  $\tau^K$ . Like capital gaps, labor gaps are lower for firms with high cash flows but, unlike capital gaps, they increase with bank leverage.

MRP-cost gaps are low for firms with poor credit scores. For capital, this relation is non-linear, and it is driven by a combination of lower marginal revenue products and higher interest rates charged by banks. For labor, the negative correlation is entirely driven by the variation in marginal products of labor, while wages display little sensitivity and, if anything, they tend to be lower for firms with poor credit scores.

The relation between both capital and labor gaps and productivity and ROA is positive and economically relevant, suggesting that higher MRP-cost gaps might capture unexpressed growth potentials.

One interpretation of these patterns is that firms tend to substitute labor inputs for capital inputs as they grow older and bigger. On the one hand, labor is more costly for larger firms than for smaller ones due to the size-dependent provisions of the Italian employment protection regulation (see section 2). On the other hand, access to external finance is more expensive and possibly constrained in early stages of firms’ life cycle (Gertler and Gilchrist 1994). Alternatively, it is possible that younger and smaller firms might hold on to partially irreversible capital investments because they face a more volatile demand (Foster et al. 2016). Larger capital gaps for young and small firms might also be the results of a form of non-classical error in the measurement capital that decreases with firm size and age. In the next section, we construct empirical tests that allow us to investigate to what extent the sign and magnitude of MRP-cost gaps reflect the degree of financial constraints and labor market rigidities faced by individual firms.

## 5 MRP-cost Gaps and Market Frictions

This section presents empirical evidence of the relationship between MRP-cost gaps and credit and labor market frictions. On the capital side, we analyze the impact of asymmetric information in credit markets, investigate the relationship to bankruptcy costs, test the response of MRP-cost gaps to credit-supply shocks, and study the dynamic of the capital gap as firms transition into the credit market. On the labor side, we

<sup>55</sup>Age groups are defined as follows: young if age  $\leq 5$ , medium if age  $\in (10]$ , old if age  $> 10$ . Assets groups are defined based on the terciles of the distribution of assets (average assets across firms in each tercile are 190 thousand, 760 thousand, and 8.8 million Euros, respectively). Credit score groups are defined as follows: safe firms are those with a credit score ranging from "Excellent" to "Solvent" (credit score from 1 to 4); a second group includes firms classified as "Vulnerable" and "Very vulnerable" (credit score from 5 and 6); Risky are firms with credit score ranging from "Risky" to "Very very risky" (credit score from 7 and 9).

show the relationship between MRP-cost gaps of labor and labor market frictions by analyzing the impact of the size-dependent severance payment requirements on firms’ employment policies.

## 5.1 Credit Market Distortions

### 5.1.1 Information Frictions

Theory suggests repeated interactions with financial intermediaries allow firms to overcome possible asymmetric-information frictions, and gradually accumulate a capital endowment more consistent with profit maximization (Botsch and Vanasco, 2015). Enduring bank-firm relations typically translate into a reduction in the expected costs of credit provision for lenders, because, conditional on past experience with the borrower, the lender now expects loans to be less risky (Diamond 1991; Petersen and Rajan 1994). Moreover, besides the effect on the probability of default, monitoring and screening costs related to information acquisition are generally lower for existing customers, because information obtained at one date may also be used to assess risk at a later date. The discussion in section 3 highlights that lenders could respond to a decline in the expected cost of credit provision by adjusting the price term of the loan contract or by relaxing credit limits that might be in place. We provide empirical evidence in favor of the latter, and show that firm-level MRP-cost gaps for capital can be used to study the impact of asymmetric information on capital accumulation by firms.

**Price versus quantity adjustments** – We begin by analyzing the relationship between probability of default and the duration of lending relationships. We focus on the subsample of observations that engage in credit markets transactions and for which we have information on borrowing rates (see section 4.1).

We define the dummy variable  $\text{DEFAULT}_{t+1}$  that takes the value of 1 in year  $t$  when we observe in year  $t + 1$  any credit in default, or any debt restructured, or in the process of being restructured.<sup>56</sup> Then, we estimate the following linear model:

$$\text{Default}_{it+1} = \beta_1 \cdot \text{Length Relation}_{it}^{wmean} + \Gamma X_{it} + \iota_{spt} + \epsilon_{it}. \quad (5)$$

To claim that longer lending relationships are less likely to culminate in default events, we must control for the underlying local credit market conditions, as well as loan- and firm-specific characteristics that are related to the strength of consumers’ demand and might affect firm profitability and credit risk. Thus, the empirical model includes year-by-province-by-industry dummies and a vector of firm-specific characteristics ( $X_{it}$ ) that includes firm-level productivity ( $\omega_{it}$ ), assets turnover, ROA, cash flows over assets, current bank leverage (bank debt/assets), nine dummy variables corresponding to each value taken by the Altman Z-score, and decile dummies for firm age and size.<sup>57</sup> As discussed in section 4.1, these variables are a set of observable indicators commonly used by banks to assess firms’ riskiness and creditworthiness. We also control for the number of active credit relations to account for heterogeneity in the intensity of credit market participation. Conforming with the prediction of economic theory, we find a negative correlation between default and length of lending (Table 5, column (1)).

Next, we investigate if, and to what extent, the reduction in credit is passed through a reduction of the interest rates or, rather, through a relaxation of existing credit-supply constraints. We estimate the

<sup>56</sup>This definition is similar to the one adopted by Panetta et al. (2009) and Crawford et al. (2016). The unconditional probability of default is 2.6% among firms in the regression sample.

<sup>57</sup>In the baseline regressions, we use 2-digits industries for the construction of year-by-province-by-industry dummies. This choice allows us to control for fairly granular industry heterogeneity while avoiding a reduction in the sample size due to singleton observations once we interact industry, year, and provinces. This choice does not affect our results. In fact, using more (4-digits industries) or less restrictive (macro industries) definition of industries, coefficients remain remarkably stable.

regression model in equation (5) using borrowing rates and  $MRP^K$  as a left-hand-side variable. Despite the incidence of productivity on default rates, the data show a relative insensitivity of borrowing rates to the duration of lending relations. Conditional on bank leverage and other observable characteristics, one extra year of lending relationships reduces interest rates by 2 basis points (Table 5, column (2)). Instead, the length of lending relationships is strongly and negatively associated with  $MRP^K$  (column (3)). Comparing two observationally similar firms that differ by one year in terms of length of lending relationships, the firm with the shorter relationship displays a marginal revenue product of capital 138 basis points higher.

The relationship between interest rates and  $MRP^K$  with the length of lending relationships is consistent with the predictions of theories of credit rationing (Stiglitz and Weiss 1981; Stiglitz and Weiss 1992). Lacking complete information about their clients, lenders are reluctant to adjust the price of credit, because such adjustment affects both the composition of the borrowing pool and their borrowing behavior. Credit limits - rather than credit prices - adjust as bank-firm relations unfold and more information is acquired (Hoshi et al. 1990a 1990b, 1991; Petersen and Rajan 1994; Schenone 2009), and the  $MRP^K$  drops as profitable investments are undertaken.<sup>58</sup>

Another explanation is lack of competition in credit markets. If information about a firm’s creditworthiness is difficult to acquire and not easily transferable, relationship lending gives current lenders monopoly power over other intermediaries, which allows them to extract rents from highly productive firms they manage to “lock-in” (Berger and Hannan 1989; Petersen and Rajan 1995).<sup>59</sup> In Appendix H.1 we test to what extent the relation between the length of lending relationships, interest rates, and marginal revenue products is a function of the degree of competition of credit markets (Berger and Hannan 1989; Petersen and Rajan 1995). Interest rates are higher and the correlation between interest rates and the length of lending relations is less negative in more concentrated markets. Both predictions are in line with the imperfect competition hypothesis. Interestingly, the correlation between Marginal Products of capital and the length of lending relationships is also less negative in more concentrated markets. However, the effect of duration on marginal revenue products of capital swamps the variation in interest rates regardless of the degree of credit market competition.

Before examining the relation between MRP-cost gaps and the length of lending relationships, we highlight an additional piece of empirical evidence in line with the asymmetric information hypothesis that comes from the relation between productivity, defaults, and interest rates. In frictionless credit markets, theory predicts a negative correlation between firm-specific productivity and the cost of debt. In the model of Appendix C we show that, under an efficient risk classification system and frictionless credit markets, high-productivity firms are safer customers from a bank’s perspective because, *ceteris paribus*, they are less likely to default on their debt obligations. Column (1) shows that, conforming with these theoretical predictions, more productive firms are indeed less likely to default on their credit obligations. *Ceteris paribus*, one interquartile range difference in TFPR (0.41 in the subsample of borrowers with loans) is associated with a reduction of the observed probability of default of 1.2 percentage points. This effect is economically significant, considering that the unconditional probability of default is 2.6% among firms in the regression sample.

Yet despite the incidence of productivity on default rates, the data provide weak support for the proposition that interest rates vary with firm-level productivity. In fact, we find a positive correlation between

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<sup>58</sup>The stickiness of interest rates and the importance of credit limits as the primary margin of adjustment of credit contracts has been also shown in the market for credit cards (see Agarwal et al. (2017) and references cited). Other types of non-price adjustments of the terms of credit contracts have been documented in other consumer credit markets, for example, the downpayment requirements for subprime auto loans in Adams et al. (2009) and Einav et al. (2012).

<sup>59</sup>See Sharpe (1990), Rajan (1992), and Hauswald and Marquez (2006) for a theoretical treatment on the link between credit market competition, information acquisition incentives, and credit-supply.

productivity and borrowing costs. This effect is statistically significant but economically negligible: all else being equal, one interquartile range difference in TFPR  $\omega$  is associated with a 3.6 basis points increase in the observed interest rate (less than 2% of a standard deviation in borrowing rates). These results suggest productivity may not belong to the variables in banks’ pricing kernel, possibly because it is unobservable to banks, and that the positive coefficient is a reflection of more productive firms’ greater willingness to pay.<sup>60</sup> Imperfect competition (“lock-in” hypothesis) might also explain the relation between the two variables. However, as we show in Appendix H.1, we find no economically significant response of borrowing rates to productivity, irrespective of the degree of credit market concentration.<sup>61</sup>

**Length of lending relationships and MRP-cost gaps** – Given the sluggish response of interest rates and much larger sensitivity of MRPK, we expect to find a strong relation between MRP-cost gaps of capital and the variable  $\text{LENGTH RELATION}_{it}^{wmean}$ . Figure 3 (panel a) shows that indeed this is the case. With respect to the year in which relationships are established, the gap  $\tau^K$  is 2 times (3 times) lower after 3 years (6 years) of continuous interactions. This descriptive analysis suggests a remarkably strong association between the two variables, yet other confounding factors might explain this pattern. We turn to regression analysis to try to isolate the effect of a relaxation of information frictions from alternative explanations. We estimate the following regression model:

$$\tau^K_{it} = \beta_1 \cdot \omega_{it} + \beta_2 \cdot \text{Length Relation}_{it}^{wmean} + \Gamma X_{it} + \iota_{spt} + \epsilon_{it}. \quad (6)$$

The vector  $X_{it}$  includes TFPR, ROA, cash-flow-to-assets ratio, assets turnover, leverage, credit score dummies, the number of active credit relationships, and a battery of age and size fixed effects (decile dummies). By controlling for TFPR and profitability measures, and by restricting our analysis to variation within industry-year-province bins ( $\iota_{spt}$ ), we tackle the concern that the dispersion in the realized MRP-cost gaps is driven by idiosyncratic variation in investment opportunities, industry-specific demand shocks (Asker et al. 2014), or time-varying risk premia.<sup>62</sup> The flexible controls for age and size are also crucial. As Foster et al. (2016) point out, young and small firms face a more volatile demand that might discourage them from undertaking partially irreversible investments, regardless of the cost and availability of external financing.

Regression results are reported in Table 5, Panel a. Column (1) shows that MRP-cost gaps strongly correlate with the average length of lending relationships between a bank and its lenders. Net of the variation explained by firm characteristics and local market dynamics, longer lending relationships allow firms to gradually implement more efficient investment policies. Ceteris paribus, one additional year of continuous borrower-lenders interactions is associated with a reduction in the *absolute value* of MRP-cost gaps of capital by about 98 basis points. We estimate model (6), replacing the continuous variable  $\text{LENGTH RELATION}_{it}^{wmean}$  with a set of dummy variables. Figure 3, Panel b, plots the regression coefficients associated with each dummy variable ( $\text{LENGTH RELATION}_{it}^{wmean} \leq 1/2$  is the baseline category, which is omitted in the regression). It shows that the monotonic relation between the two variables holds across the entire distribution of LENGTH

<sup>60</sup>An econometric explanation of this result would be that our estimates of productivity have no empirical content, due to measurement and/or misspecification errors. Prima facie, this explanation seems implausible. Our estimates of TFPR are highly correlated with credit-default outcomes and, as we show in Appendix D, both investment rates and changes in labor demand are closely related to productivity dynamics, as theory would predict.

<sup>61</sup>A possible explanation for this result is that competition among credit suppliers works to eliminate systematic misclassifications due to imperfect information. For example, if a firm is - by mistake - classified as excessively risky or not creditworthy by one lender, competitive lenders may offer a lower interest rate to attract that customer.

<sup>62</sup>In the baseline regressions, we use 2-digit industries for the construction of year-by-province-by-industry dummies. This choice does not affect our results. In fact, using a more (4-digit industries) or less restrictive (macro industries) definition of industries, the coefficients remain remarkably stable. We also experimented with replacing the vector of contemporaneous controls with its lagged counterpart. Results are unchanged. Results are available upon request.

RELATION $_{it}^{wmean}$ . Examining the control variables, we find  $\tau^K$  is positively related to measures of profitability (productivity, ROA, and assets turnover). This finding is consistent with the interpretation that a positive gap between the MRP-cost gap  $\tau^K$  signals potential investment opportunities. The analysis of the coefficients of age and size - not reported in the regression table - shows that, as expected, gaps are smaller for older and larger firms. The relation between MRP-cost gaps and credit scores is non-monotonic: everything else being equal,  $\tau^K$  increases as we move from firms with high credit rating to firms with intermediate ratings (Altman Z-score from 1 to 5), and then  $\tau^K$  sharply drops once we consider firms with the lowest credit ratings (Altman Z-score from 6 to 9). The negative sign of the coefficient associated with the number of active credit relations is also in line with our interpretation, because a larger pool of lenders provides firms with a greater set of financing options. Gaps are also negatively related to the availability of internal and external finance, as the coefficients of cash flows and leverage indicate.<sup>63</sup>

**Heterogeneous effects** – The average effect, however, masks substantial heterogeneity across producers. In column (2), we interact the length of the lending relationships with a dummy variable that indicates whether the firm was operating below its target capital endowment in period  $t - 1$  ( $UNDERCAPITALIZED_{it-1} = \mathbf{1}\{\tau_{it-1}^K > 0\}$ ), as well as a the full set of interactions of the dummy  $UNDERCAPITALIZED_{it-1}$  with the vector of controls and fixed effects in the regression model (6).<sup>64</sup> Consistent with MRP-cost gaps being proportional to the shadow cost of capital, we find that the economic benefits of longer credit relations are almost entirely concentrated among under-capitalized producers, helping them overcome potential information frictions that constrained the availability of bank finance. Figure 3, Panel c, shows this point clearly by showing the effect of longer relationships for  $UNDERCAPITALIZED_{it-1} = \{0, 1\}$ , across the distribution of LENGTH RELATION $_{it}^{wmean}$ . We find a negligible impact of longer lending relationships on the MRP-cost gap of firms that operate with a capital endowment that, according to our measure, exceeds the one more consistent with unconstrained profit maximization. Despite its small magnitude, the negative sign of the coefficient on LENGTH RELATION $_{it}^{wmean}$  suggests longer relationships might actually allow some overcapitalized firms to maintain, or even increase their capital endowment.

Another testable implication of the theory of gaps is the relation to firm-level productivity. As discussed in section 3, theory suggests the MRP-cost  $\tau^K$  is proportional to the multiplier attached to the borrowing constraint ( $\chi_{it}$ ). The shadow cost of capital  $\chi_{it}$  is increasing with the firm’s productivity because, ceteris paribus, more productive firms are capable of transforming one extra unit of capital into more revenues. Thus, if variation in  $\tau^K$  truly reflects heterogeneous shadow costs due to binding financial constraints, the benefits of bank-firm interactions should be larger for more productive firms that appear to be undercapitalized. These theoretical predictions find strong empirical support. We augment the model with the interaction between TFPR ( $\omega_{it}$ ) and the length of lending relationships (column (3)), and the triple interaction with  $UNDERCAPITALIZED_{it-1}$  (column (4)). To facilitate the interpretation of estimates, we de-mean  $\omega_{it}$ , so that the coefficient associated with LENGTH RELATION $_{it}^{wmean}$  represents the average response of  $\tau^K$  to one additional year of firm-bank interactions for a firm located at the mean of the distribution of TFPR. In column (4), the same coefficient refers to an overcapitalized firm located at the mean of the distribution of TFPR. We find a stronger correlation between gaps and the length of lending relationships for more productive firms. In particular, the sign and magnitude of the coefficient associated with the triple interaction (LENGTH RELATION $_{it}^{wmean}$  x  $UNDERCAPITALIZED_{it-1}$  x TFPR $_{it}$ ) shows that the benefits of relationship lending accrue, for the most part, to the subsample of the most productive firms that operate with too little capital. Figure

<sup>63</sup>We observe a statistically significant, positive relation between the length of lending relations and bank leverage. On average, one more year of lending relationships is associated with a 5% increase in bank leverage (p-value lower than 1%).

<sup>64</sup>The full regression table is available upon request.



3 (Panel d) provides a visual representation of the heterogeneous effects of longer lending relationships along the productivity spectrum.<sup>65</sup>

**Robustness** – We augment the regression model with firm fixed effects, and study the impact of a relaxation of borrower-lender information frictions over the firm’ life cycle (Table 5, Panel b). By doing so, we strengthen the identification of the coefficient of interest, because we now control for time-invariant unobservable firm characteristics and we also better address measurement error problems. The within-firm estimates largely confirm the results of the between-firm regressions. Also, Appendix H shows results are qualitatively similar if we measure the degree of information frictions using the unweighted average length of relations ( $\text{LENGTH RELATION}_{it}^{mean}$ ), or using only the length of the relation with the main lender ( $\text{LENGTH RELATION}_{it}^{lead}$ ).

### 5.1.2 Bankruptcy costs

Next, we investigate the relationship of MRP-cost gaps and bankruptcy costs. Inefficient bankruptcy procedures have an unambiguous, detrimental effect on firm activity. On the one hand, higher bankruptcy costs might affect investments because interest rates rise, which reduces the credit demand. On the one hand, when the cost of credit is inflexible or only partially adjusts, bankruptcy costs affect investment through a reduction of the availability of external finance (lower  $\lambda$ , through the lens of our theoretical model), which, as we previously discuss, would raise the marginal revenue product of capital of credit constrained firms.

We test these alternative hypotheses using the length of bankruptcy litigations in court as an empirical proxy of the deadweight cost of bankruptcy. The length of the bankruptcy procedures increases the deadweight loss in case of bankruptcy for several reasons.<sup>66</sup> First, long trials increase legal expenses, and for disputed loans, interest income is forgone when collateral does not cover judicial costs. Second, during the trial, the creditor is exposed to the danger of asset substitution by the debtor. Third, even in the absence of moral hazard behaviors, the market value of firm assets typically decays during the period of automatic stay. The average length of judicial proceedings across different court jurisdictions displays significant geographical variation (see Figure A.1, Panel a, in the Appendix). The between-province standard deviation is two years, with judgments taking “as little as” three years to become final in some provinces, but as much as 13 years in others.

We augment the regression model (6) with the length of bankruptcy cases ( $\text{LENGTH BANKRUPTCY}$ ) and investigate its covariance with borrowing costs,  $MRP^K$ , and the gap  $\tau^K$ . Because the bankruptcy variable is fixed over time, we cannot include industry-year-province fixed-effects, which we replace with industry-year fixed-effects plus a rich set of province-level controls (population, GDP, unemployment rate, active firms per resident, firm exit rate, Herfindahl-Hirschman Index (HHI) of credit market concentration, and number of active credit institutions).

Results show that borrowing costs are only marginally affected by heterogeneous bankruptcy costs, whereas  $MRP^K$  responds markedly (Columns (1) and (2) of Table 6). Consistent with bankruptcy costs generating more severe credit constraints rather than higher borrowing rates, we find that, on average, one extra year of legal controversies increase the Marginal Product of Capital by 42 basis points but increases

<sup>65</sup>In the graph, the variable  $\text{HIGH-TFPR}_{it}$  that takes value of 1 for observations whose productivity is above the median, and zero otherwise.

<sup>66</sup>The seminal work of La Porta et al. (1997; 1998) highlights that law and its enforcement by the judiciary are essential to credit markets. Djankov et al. (2008) provide empirical evidence of the impact of costs associated with inefficient debt enforcement procedures. Qian and Strahan (2007) and Bae and Goyal (2009) show how legal differences shape the ownership and terms of bank contracts. A body of empirical works investigate the connections between legal institutions and firm size (e.g., Laeven and Woodruff 2007).

the interest rate by 0.9 basis points. These results are in line with Jappelli et al. (2005), who finds that judicial efficiency in Italy correlates positively with the volume of lending and negatively with proxies for credit constraints. Given the small sensitivity of borrowing rates and significant response of  $MRP^K$ , it follows that inefficiencies in the legal system translate into larger MRP-cost gaps. Comparing similar firms that operate in the same industry-year, we find that one extra year of bankruptcy litigations in court translates into an average increase of 37 basis points in  $\tau_{it}^K$ .

We worry that the coefficient associated with the length of bankruptcy litigations might be simply picking up the stark difference in the quality of institutions or in the level of human and social capital between the northern and the southern regions of the country (Putnam et al. 1994; Guiso et al. 2004b; Guiso et al. 2004b). In fact, bankruptcy litigations are, on average, two years longer in the South than in the rest of the country. In column (10), we focus on within macro-region variation to try to disentangle the effect of bankruptcy costs from the North-South effect.<sup>67</sup> As expected, year-by-industry-by-macro region fixed effects reduce the correlation of bankruptcy length with both interest rates and marginal revenue products of capital, but the differential effect of LENGTH BANKRUPTCY on the two variables becomes even stronger. The correlation coefficient between bankruptcy costs and  $\tau^K$  also shrinks. However, the relationship between the two variables remains statistically relevant.

### 5.1.3 Credit availability

The analysis presented so far uses the length of lending relationships and bankruptcy costs as proxies of the supply of credit available to individual firms, providing indirect evidence of the impact of credit availability on the size and dispersion of MRP-cost gaps. We now estimate the direct effect of credit supply on MRP-cost gaps of capital. We start from the spurious correlation between changes in credit amounts granted and changes in  $\tau^K$ . We estimate the following linear model:

$$\Delta\tau_{it}^K = \beta_1 \cdot g(\text{Credit}_{it}) + \Gamma X_{it-1} + \iota_{spt} + \epsilon_{it}, \quad (7)$$

where  $g(\text{Credit}_{it}) = (\text{Credit}_{it} - \text{Credit}_{it-1})/0.5 \cdot (\text{Credit}_{it} + \text{Credit}_{it-1})$  is the symmetric growth rate of bank credit (Davis et al. 1996). With respect to  $\Delta \ln(\text{Credit}_{it})$ , this growth rate has the advantage of being defined also for firms that stop borrowing ( $\text{Credit}_{it} = 0$ ) and, being bounded between -2 and +2, it is more robust to the presence of outliers. The vector  $X_{it-1}$  is the same set of firm-level controls of model (6), but measured in period  $t - 1$ ;  $\iota_{spt}$  are industry-province-year fixed effects.<sup>68</sup> Results are reported in column (1) of Table 7. All else being equal, comparing two firms that face a one-standard-deviation difference in the firm-level growth rate of credit (37%), the MRP-cost gap  $\tau^K$  of the firm with the greater growth rate of credit drops by 1.5 percentage points ( $-3.99 \cdot 0.37$ ).<sup>69</sup> Although suggestive, this correlation results from the simultaneous effect of firms' demand and banks' supply of credit. Thus, we cannot infer from OLS coefficients whether the variation in MRP-cost gaps reflects changes in the shadow cost of capital generated by constrained supply of credit, or rather heterogenous investment opportunities. To disentangle the demand and supply channels, we construct firm-year-specific credit-supply shifters adopting an empirical design that

<sup>67</sup>Significant variation exists in the length of bankruptcy cases within macro-regions (Figure A.1, Panel b): the standard deviation between macro-regions is 1.11 years, and the average standard deviation within-region is 1.93 years.

<sup>68</sup>The controls are assets turnover, ROA, cash-flow-to-assets ratio, leverage, credit score dummies, and the number of active credit relationships, length of lending relationships, age, age square, the natural logarithm of assets, and credit score dummies. We replicate our analysis using  $\Delta \ln(\text{Credit}_{it})$  as a left-hand side variable, and find estimates that are qualitatively and quantitatively very similar to the ones in Table 7. Regressions are available upon request.

<sup>69</sup>In the estimation sample (borrowers in  $t - 1$  with information on APR on loans), the average change in firm-level credit is 4%, its standard deviation is 37%.

is a variant of the shift-share approach of Amiti and Weinstein (2016).<sup>70</sup> Specifically, using the bank-firm matched records of the CR, we decompose the yearly growth rate of credit at the relationship-level into a supply factor and demand factor using the following linear model:

$$g(\text{Credit}_{ibt}) = \mathbf{b}_{bt} + \mathbf{i}_{it} + e_{ibt}, \quad (8)$$

The left-hand side variable is the growth rate of credit from bank  $b$  to firm  $i$ , between year  $t$  and  $t - 1$ . The vectors  $\mathbf{b}_{bt}$  and  $\mathbf{i}_{it}$  are bank-year and firm-year fixed effects. The regression is run via weighted least squares with weights equal to  $\text{Credit}_{ibt-1}$ . The coefficients of interest are the estimated bank-year fixed effects  $\mathbf{b}_{bt}$ . They capture the nationwide growth rate of credit of individual financial institutions, net of the overall change in lending that can be explained by firms' idiosyncratic demand, which is absorbed by the firm-year fixed effects  $\mathbf{i}_{it}$ . The identification of both  $\mathbf{b}_{bt}$  and  $\mathbf{i}_{it}$  is guaranteed by the presence of multiple banks lending to the same firm at the same time (and banks lending to multiple firms). As discussed in section 2, the widespread presence of multi-bank firms ensures model (8) has enough power to accurately estimate the fixed effects of interest.<sup>71</sup>

Using the estimated bank-year fixed effects, we construct a firm-year-specific credit-supply shifter as

$$\text{Credit Shifter}_{it} = \sum s_{it-1} \hat{\mathbf{b}}_{bt},$$

where  $s_{ibt-1} = \frac{\text{Credit}_{ibt-1}}{\sum_b \text{Credit}_{ibt-1}}$  is the share of bank  $b$  on firm  $i$  total credit in period  $t - 1$ . Appendix I shows the distribution of  $\text{Credit Shifter}_{it}$ : the average and standard deviation are -10% and 13%, respectively.

In general, several factors create variation in the predicted credit-supply shifters, such as bank-specific events that affect the cost and availability of external financing of individual banks (Khwaja and Mian, 2008; Chodorow-Reich 2014; Cingano et al. 2016), shocks that weaken or strengthen bank balance sheets (Peek and Rosengren 2000; Gan 2007; Paravisini et al. 2014; Bottero et al. 2017), banks heterogeneous response to monetary policy (Jiménez and Ongena 2012), or merger and acquisition events that might temporarily slow down or freeze the provision of credit (Sapienza 2002). We take an agnostic view on what drives the change in credit for every individual bank. Below, we provide evidence of the information content of our estimated shifters showing their relation to banks' lending patterns during the recent European sovereign debt crisis.

We find an economically large and statistically strong relation between predicted supply shifters and growth rates of credit at the firm-level (Figure 4). Column (2) of Table 7 (Panel a) examines this relationship more formally. The regression includes the battery of firm-level controls and fixed effects of model (7). Comparing two observationally similar firms that are exposed to a one-standard-deviation difference in the credit-supply shock, the firm facing the larger (positive) shock increases its bank debt by 2.7% more (0.21\*13). Importantly, the coefficient of the correlation is remarkably stable and significant if we constrain the regression model to using only within-firm variation (column (3)). As we discuss below, these findings leave us confident that the predicted credit supply-shifters are not correlated with time-invariant borrowers' characteristics,

<sup>70</sup>Greenstone et al. (2015) propose a similar shift-share approach using more aggregated data. In Appendix I, we construct alternative credit supply shifters following their approach and find that our baseline results are very similar in terms of economic magnitude and statistical significance.

<sup>71</sup>We run the regression on the full sample of firms that appear in the CR in order to maximize the representativeness of the sample and improve the precision of the estimates of the fixed effects. We exclude from the estimation sample the stock of loans in default. We deal with mergers and acquisitions, treating the acquired and acquiring bank as a single entity over the pair of years in which mergers and acquisitions take place (Bernanke et al. 1991). In the Italian Credit Registry, almost 70 percent of the firms borrow from multiple lenders during the same year, with an average number of 3.3 active lending relationships per year.

including persistently high credit demand or investment opportunities.<sup>72</sup>

**Reduced form estimates** – Next, we use the predicted lending shocks to test the effect of changes in credit supply on  $\tau_{it}^K$ . We find that gaps shrink in response to a supply-driven change in the availability of bank finance (Table column (4)). Ceteris paribus, one standard deviation difference in credit-supply shock is associated with a 0.24 percentage points reduction of the capital gaps (-1.88\*0.13). The statistical relationship between the two variables holds true even if exploit only within-firm variation (column (5)).

Once again, the average effect masks the differential impact across firms that, before being exposed to the credit-supply shock, were operating with either an excessive or an insufficient capital endowment (column (6)). All else being equal, a one-standard-deviation difference in the credit-supply leads to a reduction of  $\tau^K$  that is 11 times larger for undercapitalized ( $\mathbf{1}\{\tau_{it-1}^K > 0\} = 1$ ) firms than for firms with zero or negative MRP-cost gaps ( $\mathbf{1}\{\tau_{it-1}^K > 0\} = 0$ , i.e., those that operate with a capital endowment close to or above target). The heterogeneous effect along the productivity margin is also economically relevant. We normalize the variable  $\omega_{it-1}$  to have zero mean, and interact it with the credit shifter and with the dummy variable that flags firms with positive gaps in period  $t-1$ . Consistent with the analysis of the relation between capital gaps and the length of lending relationships, column (7) and (8) suggest the shadow cost of capital of more productive firms experiences a more pronounced drop compared to the drop in the shadow cost of capital of the less productive firms when hit by an equally large credit-supply shock, and that this effect is entirely driven by those producers our metric classifies as credit constrained.<sup>73</sup>

Finally, we study whether the response of MRP-cost gaps to positive changes in credit supply differs from the response to negative changes in credit supply, and whether this difference is across with positive or negative gaps (column (9) and (10)). We find that change in gaps is substantial for capital-constrained firms, and especially in response to positive supply shocks. On the contrary, the MRP-cost gaps of firms with zero or negative MRP-cost gaps show an economically small response to negative credit shocks and no statistically significant response to positive ones. That is, by and large, this group of firms respond to an expansion in the credit supply by rolling over their debt, rather than by undertaking new investments, and does not appear to be affected by a credit contractions.

**Identifying assumptions and robustness** – The interpretation of the central result that variation in gaps is related to variation in credit constraints crucially depends on our ability to effectively disentangle credit supply movements from simultaneous changes in firms’ idiosyncratic credit demand. One way to investigate the validity of our identification strategy is to study how the correlation coefficient between  $g(\text{Credit}_{it})$  and  $\text{CREDIT SHIFTER}_{it}$  and the model-explained variance change as we vary the set of firm-level controls and fixed effects in the regression model (Altonji et al. 2008; Oster 2014). We do so in Appendix I. If both the  $R^2$  of the regression and the magnitude of the coefficient fluctuate substantially while changing the model specification, we would conclude that the correlation between firm-level changes in credit and our proxy of credit supply shifts might be the result of a spurious correlation of the two variables with local market conditions and firm characteristics. In contrast to this, we find that while the (adjusted)  $R^2$  gradually increases as we augment the model with a larger set of fixed effects and controls (year fixed, province-by-industry-by-year

<sup>72</sup>Although a formal test of the equality of the coefficients in column (2) and (3) rejects the null hypothesis with statistical confidence levels, the change in magnitude is small in economic terms and small if compared to the change in the explained variance the inclusion of fixed effects brings about (adjusted  $R^2$  increases by 11 percentage points). The increase in the coefficient from column (2) to column (3) does not appear to be driven by the reduction in sample size due to singleton observations in the firm-fixed effect regression. If we replicate model (3) on the estimation sample of model (4) we obtain a correlation coefficient of 0.208 (standard error 0.003). One interpretation of the larger correlation coefficient in column (3) is that unobserved time-invariant factors that affect credit growth at the firm-level are negatively correlated with credit-supply movements.

<sup>73</sup>In Column (8), the coefficient associated to Credit Shifter $_{it}$  measures the effect of a shifter of 100 percent on a firm with  $\tau_{it-1}^K \leq 0$  and lagged TFPR equal to the average of  $\omega_{it-1}$  in the estimation sample.

fixed effects, firm-level controls, and firm fixed effects), the correlation coefficient between  $g(Credit_{it})$  and  $CREDIT\ SHIFTER_{it}$  remains remarkably stable across specifications.

Another important assumption needed to identify credit supply movements via model (8) is the absence of endogenous sorting between firms and banks. We provide two pieces of evidence that suggest that assortative matching is unlikely to be the sole driver of our results. First, if a systematic assortative matching is in place because good banks specialize in lending to fast-growing industries or local markets, we would see the coefficient of both the first-stage regression and of the reduced form regression change markedly as we control for more or less coarse industry fixed effects. Results provided in Appendix I show the coefficient of interest is remarkably robust to model specification. Second, if bank-firm matching is persistent - as the duration of lending relationships seem to suggest -, firm fixed effects would control for it and possibly wipe out the statistical relation between credit supply and gaps. As previously discussed, including firm fixed effects in the regression does not erode the economic and statistical significance of the regression coefficients.<sup>74</sup> In Appendix I, we also experiment with alternative econometric models used to disentangle simultaneous movements in supply and demand for credit at the firm level (Greenstone et al., 2015), and find estimates that are comparable in sign and magnitude to the ones in Table 7.

The firm-year fixed effects estimated by model (8) convey useful information about the demand for credit of individual firms (Bottero et al. 2017). Thus, we use the estimated  $\mathbf{i}_{it}$  to test two propositions. First, from a theoretical point of view, larger gaps should be a reflection of profitable investment opportunities not undertaken by firms. Thus, we expect to see a positive association between  $\tau_{it-1}$  and the estimated firm-year fixed effects for year  $t$ . The data provide strong support in favor of this prediction, suggesting the larger gaps are associated with a greater demand for credit. Unconditionally, a one-standard-deviation increase in  $\hat{\mathbf{i}}_{it}$  (0.47) is associated with a 6.6 percentage points larger  $\tau_{it-1}$ . Second, we use  $\hat{\mathbf{l}}_{it}$  to verify that the estimated credit-supply shocks are in fact orthogonal with respect to firms' idiosyncratic credit demand. Including the estimated firm-year fixed effects  $\hat{\mathbf{l}}_{it}$  as a control, we fail to reject the hypothesis that the coefficient of the baseline regression of Table 7 (Panel b) is statically unchanged. Results of these tests are reported in Appendix I.

Our interpretation of the negative relation between gaps and credit availability is that constrained access to external finance prevents some firms from undertaking profit-maximizing investments. An alternative but related explanation is that gaps respond to credit-supply shocks because the latter affect firm-level productivity, which in turn affects the realization of  $MRP^K$  (Manaresi and Pierri 2017). In Appendix I, we show our results continue to hold controlling for the *simultaneous* change in productivity, which suggests that our estimates are driven by an efficient adjustment in the quantity of capital utilized by firms, rather than a possible increase in firm-level productivity.<sup>75</sup>

**Economic magnitudes** – To gauge an understanding of the magnitude of the effect on credit availability on gaps, we instrument the change in firm-level credit supply ( $g(Credit_{it})$ ) with our predicted lending shifter. Before commenting on the estimation results of the 2SLS model, we must emphasize the limitations of this approach. First, the exclusion restriction is problematic, because the estimated effect of an expansion/contraction of credit supply also encompasses the effect of other outcomes that impact  $\tau_{it}^K$ .<sup>76</sup> Second,

<sup>74</sup>Another assumption needed in the Amiti and Weinstein (2016) methodology is the absence of spillover effects across supplies of different banks. This assumption seems to hold in the data: indeed, estimating the shifters controlling for these spillovers (i.e., by iterating several times estimates of bank-fixed effects, including past estimates of other banks into the decomposition) yields very similar credit-supply shocks (Manaresi and Pierri 2017).

<sup>75</sup>We fail to reject the hypothesis that the coefficient in column (1) of Table 7 (Panel b) equals the coefficient of the same regression that also includes  $\Delta\omega_{it}$  as a control.

<sup>76</sup>Shifts in credit supply affect both the quantity of credit as well as other terms of the credit contract, such as interest rates,

the IV is going to give us a “local average treatment effect (LATE, Angrist 1990), that is, the effect of an additional unit of credit on the MRP-cost gaps firms for which credit actually changed (the “compliers”).<sup>77</sup> With these caveats in mind, we present the estimates of the 2SLS model in Table 7, panel b. The economic magnitude of changes in credit availability on MRP-cost gaps appears to be substantial. On average, comparing two observationally similar firms whose change in credit supply is one standard deviation apart, we find the gap of the firm experiencing the larger credit expansion is reduced by 2.9 percentage points more with respect to the gap of the other firm (column (1)), and by 1.7 percentage points if we control for firm fixed effects (column (2)). Importantly, a supply-driven credit expansion worth one standard deviation of  $g(Credit_{it})$  leads to a reduction of 3.6 percentage points in the  $\tau^K$  of those firms that, before experiencing the shock, appear to be over-capitalized (column (3)). By contrast, an expansion of credit supply widens the distance between the marginal revenue product of capital and the user cost of those overcapitalized firms. The magnitude of this effect, however, is nine times smaller than the one estimated for undercapitalized producers. Columns (4) and (5) confirms the shadow cost of capital of more productive firms drops more pronouncedly compared to the one of the less productive firms when hit by an equally large credit-supply shock, and that this effect is entirely driven by those producers that appear to be credit constrained.<sup>78</sup> For example, consider two undercapitalized firms ( $\tau_{it-1} > 0$ ), located at the 75th and 25th percentiles of the TFPR distribution in  $t - 1$ . All else being equal, a one-standard-deviation increase in firm-level credit supply reduces the gap of the latter by 1.9 percentage points more (-4.4 vs -2.5). Comparing the most productive of the two firms to another producer with average productivity and  $\tau_{it-1} \leq 0$ , the differential effect of a one-standard-deviation increase in credit supply is a larger reduction in  $\Delta\tau_{it}^K$  of 3.7 percentage points for the former firm (-4.4 vs -0.7).

**Quasi-experimental evidence from the European sovereign crisis** – In Appendix I, we provide an additional piece of evidence that suggests part of the observed dispersion of  $\tau_{it}^K$  can be explained by binding credit constraints that generate heterogeneous shadow costs of capital across firms. Building on the work of Bottero et al. (2017) (BLM henceforth), we study the relation between firms’ exposure to the recent European Sovereign crisis and the gap  $\tau^K$ .<sup>79</sup> Following BLM, we construct a measure of banks’ exposure to the sovereign crisis, exploiting variation in firms’ exposure to banks with differential holdings of government bonds issued by distressed sovereigns, and construct a firm-level credit-supply shifter as

$$\text{Sovereign Shock}_{iPRE} = \sum_b s_{iPRE} \text{Sovereigns}_{bPRE},$$

where  $s_{iPRE} = \frac{\text{Credit}_{ibPRE}}{\sum_b \text{Credit}_{ibPRE}}$  is the share of bank  $b$  on firm  $i$  total credit, measured before the Greek bailout;  $\text{Sovereigns}_{iPRE}$  is the exposure of bank  $b$  to Italian government bonds at the end of 2010:Q1 scaled by risk-weighted assets, which is a bank-specific measure of financial institutions’ exposure to the sovereign shock. We find a strong, negative correlation between the variable  $\text{Sovereign Shock}_{iPRE}$  and the estimated

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maturity, covenants – all of which can independently affect the availability of credit supply and the MRP-cost gap. Importantly, although these concerns implicate the consistency of the second stage coefficients, they do not invalidate the reduced form estimates reported in the rest of this section.

<sup>77</sup>For example, a contraction in credit supply ( $g(Credit_{it}) < 0$ ) is going to affect those borrowers whose loans are maturing during the year. Loans maturing in subsequent years are likely not responding to negative credit supply shifts in period  $t$ , unless failure or delays in debt repayments or covenant violations allow lenders to renegotiate the term of those contracts (Chodorow-Reich and Falato 2017).

<sup>78</sup>As in columns (7) and (8) of Panel a, we normalize the variable  $\omega_{it-1}$  to have zero mean and interact it with the credit shifter and with the dummy variable that flags firms with positive gaps in period  $t - 1$  ( $\text{UNDERCAPITALIZED}_{it-1}$ ).

<sup>79</sup>BLM shows the sovereign default and subsequent bailout of Greece in spring 2010 led banks more exposed to sovereign securities issued by Southern European countries (including Italian bonds) to sharply reduce their credit supply in response to a reassessment of the riskiness to their portfolios.

credit-supply shifter between 2009 and 2010 (Credit Shifter<sub>*i*2010</sub>). The raw correlation is 16%. In terms of magnitude, we find a 9% reduction of Credit Shifter<sub>*i*2010</sub> (0.6 of a standard-deviation) associated to a one-standard-deviation increase in the variable Sovereigns<sub>*i*PRE</sub>. Both correlations are significant with a p-value below 1%, and provide direct evidence of the link between the credit-supply shifters constructed using the shift-share approach in model (8) and events that affect the credit provision of individual banks.

Using the sovereign shifter, we investigate how the contraction of credit by more exposed banks affected the MRP-costs gap of their borrowers. Comparing two similar firms that, at the onset of the sovereign crisis, are one standard deviation apart in terms of lenders' exposure to the distressed sovereigns, we observe an increase of 0.5 percentage points in  $\tau^K$  the year following the burst of the crisis, and a cumulative effect of 1 percentage point over a four-year period. To the extent that the change in the gap captures a change in the shadow cost of capital of affected firms, it suggests the real effects of the sovereign crisis are long-lasting.

#### 5.1.4 Access to credit

Up to now, we have restricted the analysis of investment policy distortions and credit market frictions to producers who actively engage in credit market transactions. The summary statistics reported in Table 2, however, shows that the gap between realized marginal revenue products of capital and user costs is, on average, three times as large for firms that do not engage in credit market transactions.<sup>80</sup> Interestingly, we also observe a significant difference between the MRP-cost gaps of firms that report outstanding loan obligations and the estimated MRP-cost gaps of firms that utilize only revokable credit lines. Credit lines are often the first type of credit granted by banks in order to test borrowers' creditworthiness. The high interest rates, relatively low amounts, and revokable nature, however, make this type of credit product an inappropriate and expensive source of financing for capital expenditures in fixed assets (see Appendix B.1).

Although suggestive, we should be careful ascribing the difference between borrowers and non-borrowers to financial frictions. Credit market participation naturally co-varies with other phenomena affecting firm policies over their life cycle, and borrowers differ from non-borrowers in terms of age, size, and industry affiliation. Young and small firms may voluntarily restrain themselves from undertaking (partially) irreversible investments even if debt financing is available, especially when they operate in industries characterized by high demand uncertainty (Foster et al. 2016; Bloom 2009; Asker et al., 2014). We turn to regression analysis to try to disentangle the impact of credit market participation from these confounding effects. We run the following non-parametric difference-in-differences regression and analyze, within firm, the evolution of  $\tau^K$  around the year in which firms enter the credit market:

$$y_{it} = \sum_{j=-3}^7 \beta_j \mathbf{1}\{(t - t_0) = j\}_{it} + \Gamma X_{it} + \iota_t + \iota_i + \epsilon_{it}, \quad (9)$$

where  $t_0$  represents the year of the change in status: BORROWER<sub>*i* $t_0-1$</sub>  = 0 (no outstanding bank debt) and BORROWER<sub>*i* $t_0$</sub>  = 1 (positive outstanding bank debt). We control for time trends with year fixed effects ( $\iota_t$ ); firm fixed effects ( $\iota_i$ ) allow us to exploit only within-firm variation. The vector of time-varying controls  $X_{it}$  includes a second-order polynomial in age, the natural logarithm of lagged assets, and lagged credit score. We allow the error term  $\epsilon_{it}$  to display serial correlation at firm level. Figure 5 (Panel a) displays the estimates of the coefficients.  $\hat{\beta}_j$  captures the average change in  $\tau_{it}^K$  from year  $j = -1$  (the baseline category in Model

<sup>80</sup>We do not observe the borrowing costs for firms that do not engage in credit market transactions. We follow the procedure described in section 4.1 and Appendix B.2 to construct an estimate of the interest rate they might have charged had they been able/willing to borrow.

(9)) to year  $j \neq -1$ . Comparing the level of the gap the year before credit market entry to the level of the year of entry and to one observed the following year, we estimate an average drop of 5 and 10 percentage points in  $\tau^K$ , respectively. These stylized facts are revealing. They highlight that credit market participation (extensive margin) matters as much as, or even more than, the intensity of credit market interactions (Jeong and Townsend 2007; Midrigan and Xu 2014). Despite its importance, the extensive margin has received limited attention in the empirical literature, typically because of the lack of micro-level records that allow researchers to follow firms in their transition into credit markets.

**Robustness** – A concern with the comparison of borrowers and non-borrowers is related to an incorrect estimation of the missing prices. We worry that the larger gaps observed for non-borrowers might be driven by a systematic underestimate of their user cost of capital. A simple back-of-the-envelope exercise suggests that imprecise estimates of  $r$  are unlikely to explain the large differences between borrowers and non-borrowers, as the interest rate of non-borrowers should be 22 *percentage points* higher than our estimate in order to equalize their average gap to the average gap of borrowers. In Appendix B.2, we provide other formal tests that suggest that an incorrect assessment of the potential borrowing rate that non-borrowers face is unlikely to be the driver of the estimates of model (9). First, we look at “crossover firms.” That is, we identify those firms that are borrowers in year  $t$  but were not borrowers in  $t - 1$ . For these observations, the difference between the observed interest rate in period  $t$  and the imputed interest rate in  $t - 1$  is only 28 basis points on average (median 0.18). We also perform an out-of-sample test, excluding a random sample of 10% of the firms for which we observe the interest rates, and implement our imputation procedure using the remaining 90% of the observations. For the subsample of excluded observations, the difference between the imputed and observed rate is, on average, economically negligible (-0.1 percentage points) and not significantly different from zero. Thus, consistent with previous evidence on the rigidity of interest rates, the change in  $\tau^K$  as firms transition into the credit market is by and large driven by a drop in the  $MRP^K$  that signals firms use bank credit to harvest profitable investment opportunities.

Another concern with the previous calculations is related to discounting. The macro-finance literature has long recognized the presence of heterogeneous and state-contingent discount factors (Cochrane 2011). We abstracted from this issue assuming shareholders are risk-neutral, and they discount cash flows at a constant rate  $\rho$ . This suggests that significant differences in the level of risk-aversion across entrepreneurs, and differences in the prices of risk of the same entrepreneur over time, could explain the larger gaps of non-borrowers. A proper account of the heterogeneous and stochastic nature of firms’ discount factors is beyond the scope of the paper. However, in the spirit of the previous calculations, we calculate that non-borrowers shall have a discount factor twice as large as borrowers, in order to equalize the gap between these two group of firms. That is, if one dollar tomorrow is worth 95 cents today for borrowers (i.e.,  $\rho^{Borrower} = 0.95$ , as we assume in the paper), the same dollar should be worth 45 cents to non-borrowers in order to equalize the average gap between these two groups. This level of a one-period discount factor for firms is a particularly high, especially because we are already “discounting” the realized MRP-cost gap of non-borrowers more than the one of borrowers through a lower estimated probability of exit ( $P\{Exit_{it+1}\}$ ) (see equation (3a)).<sup>81</sup>

### 5.1.5 Extension to the Full Sample of Borrowers and Alternative Interest Rates

**Full sample of borrowers** – Analyzing the relation between credit market frictions and credit-supply shocks, we restricted our attention to the subsample of borrowers for which we observe the information on the APR on term loans. In Appendix I, we replicate all the analysis on the sample of borrowers for which

<sup>81</sup>Our estimates suggest an expected probability of exit of 14% for non-borrowers and of 9% for borrowers.



we have information on the identity of the lender. As discussed in Section 4, this subsample include firms that only borrow drawing from credit lines, firms whose lenders are not in the TAXIA sample. Results are confirmed in both economic magnitude and statistical significance.

**Alternative interest rates** – Throughout the paper, we have used the APR on bank loans as a measure of borrowing costs. We have argued and shown empirically that investments in fixed assets display higher sensitivity to loans than to other types of credit products. How would the results of our analysis change if the APR on credit lines is used to construct  $\tau^K$ ? Appendix L shows the MRP-cost gaps  $\tau_{it}^K$  and the percentage deviations  $(K_{it}^* - K_{it})/K_{it}$  are lower if we use the APR on credit lines, reflecting the higher level of  $r^{CredLines}$  with respect to  $r^{Loans}$ .<sup>82</sup> However, all the statistical relationships discussed in this section are found to be highly robust to changes in the reference interest rate.

## 5.2 Labor Market Distortions

We now move to the analysis of labor policies and labor market frictions. We focus, in particular, on the firing restrictions imposed by Article 18 of the Italian Workers’ Statute. As discussed in section 2, contractual frictions prevent wage adjustments from undoing the government-mandated severance payment imposed on firms larger than 15 employees. In this situation, we expect the size-dependent provisions of Article 18 to have an effect on employment policies and on the allocation of labor across producers.

We begin our analysis showing the impact of the employment protection provision on firms’ propensity to grow and on the size distribution of firms around the 15-employees threshold. Figure 6 Panel a (left), displays the probability of increasing the labor force ( $EMPLOYEES_{it} > EMPLOYEES_{it-1}$ , left axis) and the probability of labor-force inertia ( $EMPLOYEES_{it} = EMPLOYEES_{it-1}$ , right axes) as a function the number of employees of the firm. The probability of workforce inertia jumps by 0.5 percentage points at the threshold.<sup>83</sup> This finding suggests the provisions of Article 18 generate a significant rigidity in the labor market. Because adjustment has an option value, firing costs also affect hiring decisions (Hopenhayn and Rogerson 1993): in anticipation of a possible reversal of consumer demand, firms will hire less than they would have in a frictionless environment, in order to avoid incurring high firing costs when downscaling is needed. Consistent with this theoretical prediction, we find a sharp drop in the probability of hiring new workers right-after the 15-employees threshold, inverting an upward tendency observed to the left of the threshold. Figure 6 panel a (right) shows that Article 18 has implications for the size distribution. The fraction of firms by number of employees generally decays with size following a power law, but the rate of decay changes markedly past the 15-employees threshold (Hijzen et al. 2013).

**Price versus quantity adjustments** – In the absence of wage rigidities, theory predicts that government-mandated severance payment would be neutralized by appropriately designed wage contracts. Firing costs would be transferred to workers in the form of reduced wages, with no effects on firms’ employment policies (Lazear 1990). This is not what we find in the data.

Figure 6, Panel b (left), plots average yearly wages and average Marginal Revenues Products of labor by size class. To account for time, industry, and local labor trends, we use firm-level deviations from the year-industry-province mean; then we aggregate the data, taking the average of firm wages within each size cell. We normalize both series to zero at size 15, to make the patterns more easily comparable. Wages increase smoothly with firm size; however, we do not find a noticeable reduction in the wage rate at or after

<sup>82</sup>On average, the nominal APR on credit lines is 12.5%.

<sup>83</sup>These results are in line with previous empirical studies that analyzed the impact of Article 18 on firms employment policies (Garibaldi et al. 2004, Schivardi and Torrini 2008).

the 15 employees threshold. By contrast, the figure highlights a significant response of the marginal revenue product around the threshold, suggesting that - due to the rigidity of wages - the labor market regulation affects firms' labor demand.

The rigidity of wages is consistent with the institutional features of the Italian labor market. Evidence of wage rigidity can be gathered from nationally representative firm-level surveys. According to the 2010 Bank of Italy survey on industrial and service firms, only 20% of firms engage in some form of firm-employees wage negotiations. On average, the nationally negotiated minimum wage contracts account for 80% of the total wage, whereas the residual 20% is set at the firm level (Adamopoulou et al. 2016). Heckman et al. (2006) reports similar estimates for the 1990s. Firm-level wage bargaining, in the sporadic instances when it takes place, is limited to upward adjustments, because nationally-set minimum wages are legally binding (Devicienti et al. 2007). Data from the 2010 Structure of Earning Survey administered by the National Statistical Institute show the firm-level component of workers' compensation is mostly composed of *una tantum* bonuses, rather than adjustments in the hourly or daily wage rate.

Our data corroborate these findings. We run a linear regression that includes a set of year-by-province-by-industry dummies, and dummies for age and size deciles. This simple model explains 50.3% of the dispersion in average annual wages.<sup>84</sup> Firm-level (log) TFPR and ROA - two measures of productivity and profitability - are able to explain only an additional 1% of the variance in wages.

Because wages are rigid, labor demand is expected to react. Indeed, Figure 6 panel b (left) highlights a significant response of the marginal revenue product around the threshold, suggesting that - due to the rigidity of wages - the labor market regulation affects firms labor demand.

**Static effects** - Figure 6, Panel b (right), studies the impact of the size-dependent EPL on the distribution of the labor gap  $\tau^L$ . The average labor gap increases as we approach the regulatory threshold, possibly reflecting forgone growth opportunities in order to avoid the cost being subject to the labor irreversibility. Mirroring the distribution of  $MRP^L$ , we find a significant kink right at the threshold: gaps become flat immediately past the threshold, as labor adjustment takes place. Above 20 employees, gaps start growing again, albeit at a slower pace.

**Dynamic effects** - Next, we study the dynamic effects of the provisions of Article 18 on the static distribution of gaps. We study how firms of different size respond to realized changes in TFPR productivity ( $\Delta\omega$ ).<sup>85</sup> More formally, for all firms below 30 employees, we estimate the model:

$$y_{it} = \sum_{j=1}^{30} \beta_j D(j)_{it} + \gamma \Delta\omega_{i,t} + \sum_{j=1}^{30} \delta_j D(j)_{it} \Delta\omega_{i,t} + X_{it} \theta + \epsilon_{it}, \quad (10)$$

where  $y_{it}$  is  $\tau_{it}^L$ ,  $MRP_{it}^L$ ,  $w_{it}$ , and  $(L_{it}^* - L_{it})/L_{it}$ , alternatively. The variable  $\Delta\omega_{i,t}$  is the change in TFPR between year  $t - 1$  and  $t$ .  $D(j)_{it}$  are a set of dummies equal to the value of 1 if  $\text{EMPLOYEES}_{it-1}$  equals  $j$ . The vector of controls includes lagged TFPR, a quadratic in lagged age, and (alternatively) industry-by-year-by-province fixed effects or year and firm fixed effects. The vector  $\delta_k$  provides an estimate of the association between changes in productivity and the dependent variables for different size categories. We are interested, in particular, in firms around the 15-employees threshold. To ease the discussion, here we focus on the comparison of  $\delta_{15}$  with  $\delta_{14}$  and  $\delta_{16}$ , that is, firms at the threshold vis-a-vis those below and immediately

<sup>84</sup>In a similar institutional context, Doraszelski and Jaumandreu (2015) shows that a large share of the observed wage dispersion across Spanish manufacturing firms is explained by geographic and temporal differences in labor markets.

<sup>85</sup>See Garicano et al. (2016) for a detailed analysis of the dynamic implications of size-dependent labor market regulation using structural methods.

above it.<sup>86</sup> Results are provided in Table 8. The first column of Panel a shows that changes in TFP for firms at the threshold induce larger gaps for firms right below the threshold relative to smaller or larger firms. In column 2, we focus on the effect of positive innovations in TFPR ( $\Delta\omega_{it}^+$ ), restricting the estimation sample to firm-year observations for which we observe a  $\Delta\omega_{it} > 0$ . The lower response of firms at the threshold relative to firms before the threshold become larger in magnitude and statistically significant. Columns 3 and 4 confirm that results hold true also when controlling for firm time-invariant, unobserved heterogeneity using firm fixed effects. This result is consistent with the hypothesis that firms with 15 employees do not increase their size in response to a positive TFPR shocks in order to avoid becoming subject to Article 18 provisions. Panels b and c of the table provide additional evidence in favor of this conjecture: they show that the difference in MRP-cost gaps is entirely driven by labor-quantity adjustments (change in  $MRP^L$ ), with no significant differences in wages. Finally, Panel d shows the effect of TFPR shocks on employment policy distortions ( $(L_{it}^* - L_{it})/L_{it}$ ). A one-percentage-point positive increase in TFPR is found to increase under-employment by 5 percentage points for firms of size 15, relative to those of size 14.

### 5.2.1 Robustness checks

**Average versus marginal wages** – One might worry that average annual wages are poor proxies of the marginal ones. To gauge an understanding on the impact of an imperfect measurement of the user cost of labor, we use wages of newly hired workers as an alternative proxy of the user cost of labor (see section 4.1). We find that wages of new workers are higher than the wages paid to previously employed workers (on average, 15% higher), an effect likely due to a combination of nominal rigidities in the inter-temporal adjustment of wages of infra-marginal workers, and to the different skill composition of newly hired employees. This difference, however, has a limited impact on MRPL-wage gaps and percentage deviations. In Appendix L, we show that, consistent with rigid wages and labor-quantity adjustments, the level of  $\tau^L$  and  $(L_{it}^* - L_{it})/L_{it}$  is lower, but using wages of new hires as a proxy of  $w_{it}$  does not affect our findings with respect to the static and dynamic implications of the size-dependent EPL.

**Informal labor and misreporting** – Considering the incidence of the informal labor on the Italian economy, one might argue that these patterns could be explained by firms misreporting their labor force, in order to avoid being subject to the provisions of Article 18. Anecdotal evidence suggests the problem of hidden labor is more severe in the informal sector of the economy. All firms in our sample, instead, are incorporated entities, which are subject to a closer scrutiny by government officials and unions. In Appendix H, we provide two pieces of evidence suggesting that the patterns in Figure 6 reflect true distortions rather than misreporting. First, we show that even in highly unionized industries (e.g., manufacturing), we find patterns very similar to industries with low unionization levels (e.g., services). Second, if capital and labor are partially substitutable in production, theory suggests that firms should respond to an increase in the cost of labor demanding more capital. Consistent with a stronger substitution effect due to changes in the relative cost of inputs, the capital gap  $\tau^K$  spikes down to the left of the threshold and then up to the right. A similar pattern is observed for the distribution of  $(K_{it} - K_{it}^*)/K_{it}$ .<sup>87</sup> The ability to effectively substitute labor services for capital services is a function of the relative intensity of these two inputs in production. Thus, we expect employment policy distortions to be higher in labor-intensive industries than in capital-intensive

<sup>86</sup>In Appendix J, we provide the entire distribution of the estimates  $\delta_j$ , showing that results are not affected by narrowing our discussion to these three categories.

<sup>87</sup>Autor et al. (2007) provide similar evidence of capital deepening in response to variation in EPL in the United States. Cingano et al. (2010) study the capital-labor substitution when EPL interact with financial frictions. Rodano et al. (2016) study the impact of Article 18 on capital accumulation within a general equilibrium framework.

ones. Using the industry ratio of  $\theta^L/\theta^K$  as a measure of labor intensity (relative to capital) of the production technology, we find that  $\tau^L$  responds more strongly to the provisions of Article 18 in more labor-intensive industries.

**Overtime hours** – Another concern is related to the possibility that firms might respond to the size-dependent EPL by adjusting the number of hours they ask their employees to work, or by incentivizing overtime. The data in our possess do not contain information on these variables. However, the Italian institutional framework and previous empirical studies leave us confident that these factors are not a major force driving our results. First, firms are limited in their ability to increase hours, because the Italian labor law allows a maximum 40 hours per week and 8 hours per day. Moreover, Adamopoulou et al. (2016) shows that overtime pay accounted for a relatively low portion (around 4%) of monthly earnings in the industrial sector. They also find that overtime hours (as a fraction of total hours) are uncorrelated with the degree of wage rigidity at the firm level.

## 6 Firm-level Counterfactuals and Policy Distortions

Do the estimated MRP-cost gaps imply economically relevant investment and employment distortions? How much labor and capital do firms need to acquire/dismiss in order to close them? To answer these questions, we define the following counterfactual input demands:

$$\begin{aligned} K_{it}^* &:= MRP^K(K_{it}^*) = r_{t+1} + \delta \\ L_{it}^* &:= MRP^L(L_{it}^*) = w_{it}. \end{aligned} \tag{11}$$

We refer to these counterfactual quantities as *target labor force* and *target capital endowment* ( $L_{it}^*$  and  $K_{it}^*$ ). Figure 7 provides a graphical representation of the relation between gaps and target endowments. Under the assumptions that the user costs each firm faces do not change for moderate adjustments of its input demand, we calculate the *deviations from targets* as

$$\frac{L_{it}^* - L_{it}}{L_{it}} = -\tau_{it}^L \cdot \left( \frac{\partial MRP^L}{\partial L} \Big|_{L \cong L_{it}; MRP^L \cong MRP_{it}^L} \right)^{-1} \tag{12a}$$

$$\frac{K_{it}^* - K_{it}}{K_{it}} = -\frac{\tau_{it-1}^K}{\rho} \cdot \left( \frac{\partial MRP^K}{\partial K} \Big|_{K \cong K_{it}; MRP^K \cong MRP_{it}^K} \right)^{-1}, \tag{12b}$$

where the terms in parentheses are the inverse of the slope of marginal revenue product schedules evaluated in a neighborhood of the observed input demands and realized *MRPs*.  $(L_{it}^* - L_{it})/L_{it}$  and  $(K_{it}^* - K_{it})/K_{it}$  have an intuitive interpretation. They help us read MRP-cost gaps in terms of how many extra workers a firm should hire (or fire), and how much capital expenditures should change, in order to close the gap between realized marginal revenue products and observed user costs.

Being able to construct firm-level counterfactual input demands is important for two reasons. First, evaluating the extent of misallocation necessarily requires computing counterfactual levels of output. As we discuss in the next section,  $L^*$  and  $K^*$  are key ingredients in our attempt to do so. Second, thinking about policy distortions in terms of percentage deviations is also important if one wants to compare the magnitude of the capital and labor adjustments needed to close firm-specific gaps. To clarify this point, consider a case where  $\tau_{1t}^K > \tau_{2t}^K > 0$ . Based on the ordinality of MRP-cost gaps, one might be tempted to conclude firm 1 investment policies are more distorted than firm 2 policies, because the distance between firm 1 capital stock

from its optimal endowment is larger than the distance of firm 2. This logic has a caveat. What matters is not only the size of the gap but also the rate at which one additional unit of capital closes the gap. For example, consider small- and large-size firms with a similar  $\tau^K > 0$ . Despite the similar gaps, the implied investment policy distortion is expected to be larger for the latter because, as the data suggest, the slope of the  $MRP^K$  schedule in the  $K - MRP^K$  plane is much steeper for small firms than for large firms. Similar considerations apply to labor gaps.<sup>88</sup>

We estimate slopes of marginal revenue products via local linear regressions. For each input and each macro-industry (1-digit code), we sort observations into 100 cells defined by deciles of the distribution of  $L$  and  $MRP^L$  ( $K$  and  $MRP^K$ ). Within each cell, we run a linear model in first difference:  $\Delta L_{it} = \beta^L \Delta MRP_{it}^L$  and  $\Delta K_{it} = \beta^K \Delta MRP_{it}^K$ .<sup>89</sup> Regressions are run separately for each macro-industry (1-digit code) to account for heterogeneous adjustments due to different technologies of production. Table A.15 in Appendix F reports the summary statistics of the distribution of the estimated slopes. On average, a change in capital of 1,000 Euros reduces the marginal revenue product of capital by 0.13%. A positive 1-unit-change of effective labor reduces the marginal revenue product of labor by 8,000 Euros.

Table 2, Panel d, reports the summary statistics of deviations from targets  $(L_{it}^* - L_{it})/L_{it}$ ,  $(K_{it}^* - K_{it})/K_{it}$ . We multiply them by 100 to express them as percentages. Figure 2 (Panel b) provides a graphical representation of their distribution. On average, to close their gaps, firms should increase capital expenditure by an amount worth 16% of their assets in place, and they should expand their (effective) labor force by 11% more. Mirroring the distribution of MRP-cost gaps, these numbers are driven by the right-tail of the distributions. In fact, central percentiles are occupied by firms whose capital and labor endowment appears to be relatively undistorted, according to our metric. To close the gaps of the median firm, it would be sufficient to invest an amount of capital worth 1% of firm assets and hire 3% more units of effective labor.

Importantly, we observe both positive and negative deviations at the tails of the distributions of  $(L_{it}^* - L_{it})/L_{it}$  and  $(K_{it}^* - K_{it})/K_{it}$ . Based on our estimates, 25% of the firms-year observations should have invested to acquire 6% more physical capital and expand their labor force 15% more. On the contrary, another 25% of firms should sell 1% or more of their fixed assets and over 10% of observations should reduce their quality-adjusted labor demand by 1% or more.

## 6.1 Investment Policy Distortions

To gauge an understanding of the relation between credit market frictions and policy distortions, we replicate the analysis of Section 5.1 using the percentage deviations from target capital  $(K_{it}^* - K_{it})/K_{it}$  as an outcome variable.

**Information frictions and relationship lending** – Figure 8 shows the amount of investment needed to close the gap is worth 25% of installed capital for firms with newly established relations, but it reduces to 10% after three years, and to 6% after 10 years of continuous bank-firm interactions. We use regression analysis to disentangle the effect of relationship lending from other correlated variables. The empirical specifications mirror the ones adopted to test the MRP-cost gap  $\tau_{it}^K$  (equation 6), but using percentage deviations  $((K_{it}^* - K_{it})/K_{it})$  as the dependent variable. Results are presented in Table 9, Panel a (continuous variable), and

<sup>88</sup>The estimated slopes confirm this intuition. Small firms tend to have higher Marginal Products, but significantly steeper MRP slopes than larger firms: large (small) MRP-cost gaps might translate to relatively small (large) distortions depending on the slope of the MRP-cost schedule.

<sup>89</sup>First differencing allows us to exploit only within-firm variation, and smooth out the impact of outliers. We also experimented with other specifications in levels with fixed effects. This specification produces estimates of  $\hat{\beta}$ s in the same ballpark of the first-difference estimator, with some more extreme values.

Figure 8, Panels b–d (discrete variable). On average, one additional year of continuous interactions leads to a reduction in the investment gap of almost 0.5 percentage points for those firms that operate below scale. The costs of asymmetric information in credit markets are particularly high for highly productive firms, especially for those under scale.<sup>90</sup>

Figure 9 shows how the benefits of relationship lending vary across firms in different stages of their life cycle as a function of the strength of their credit market interactions. If scale economies exist in information production, and the monitoring cost per dollar falls with the size of the loan, the cost of asymmetric information falls with borrower size and transparency. Confirming a robust empirical finding in the literature, asymmetric information frictions appear to be particularly taxing for small and young firms (Gertler and Gilchrist 1994; Chodorow-Reich, 2014; Bottero et al. 2017).<sup>91</sup>

**Bankruptcy costs** – The association between investment policy distortions of capital and bankruptcy costs is also confirmed. All else being equal, our estimates suggest that longer bankruptcy litigations depress investment, because the average percentage deviations from the target capital are higher in geographical areas where the judiciary system is more inefficient (Table 9, Panel a).

**Credit availability** – Table 9 (Panel b) studies how  $(K_{it}^* - K_{it})/K_{it}$  responds to changes in the availability of credit. In line with the negative relation found for  $\tau_{it}^K$ , the distance between capital and target capital decreases as more credit becomes available (reduced form regressions). Using the 2SLS regressions to quantify the relation between credit availability and gaps, we find that, everything else being equal, an increase of one standard deviation in credit (driven by an expansion of credit supply) is associated, on average, with a 0.45 percentage points drop in  $(K_{it}^* - K_{it})/K_{it}$  ( $-1.20 \cdot 0.37$ ) for firms that were undercapitalized the period before they experienced the credit shock.<sup>92</sup> This effect is entirely driven by the most productive firms that operate below their optimal capital endowment. By contrast, overcapitalized firms seem to respond to an expansion of credit supply mostly rolling over their debt, rather than undertaking new investment.

**Access to credit** – Finally, we assess the impact of credit market participation on investment policies. We calculate that the amount of investment non-borrowers need to close their capital gaps is 3.5 times larger than the one borrowers need. We use the non-parametric, within-firm difference-in-differences model in equation (9) to study the transition into credit markets. We estimate a significantly higher deviation from optimal capital endowment before a firm enters into a credit market transaction (Figure 5, Panel b). Then, we estimate that, on average,  $(K_{it}^* - K_{it})/K_{it}$  drops by 3 and 5 percentage points the year of entry and to one of the following year, respectively.<sup>93</sup>

## 6.2 Employment Policy Distortions

We conclude this section studying the local impact of the size-dependent employment protection law on percentage deviations from target labor  $(L_{it}^* - L_{it})/L_{it}$ .

**Static effects** – Figure 6, Panel c plots the average percentage deviations as a function of firm size, highlighting a structural change at around 15-employees. Whereas employment policy distortions monotonically

<sup>90</sup>The inclusion of firm fixed effects and control for simultaneous productivity does not affect the results.

<sup>91</sup>Age groups are defined as follows: young if  $\text{age} \leq 5$ , medium if  $\text{age} \in (10]$ , old if  $\text{age} > 10$ . Asset groups are defined based on the terciles of the distribution of assets; average assets across firms in each tercile are 190,000, 760,000, and 8.8 million Euros.

<sup>92</sup>As in Table 7, we normalize the variable  $\omega_{it-1}$  to have zero mean in the estimation sample. The IV regressions in Table 9 (Panel b) are subject to the same caveats discussed above.

<sup>93</sup>Comparing Figure 5 Panel a to Panel b, we find a pre-entry upward trend in the MPR-cost gap  $\tau^K$  that is not found in the dynamic of the percentage deviation  $(K_{it}^* - K_{it})/K_{it}$ . This trend is due to the fact that, for small and young producers, a large  $\tau^K$  translates into a moderate increase in  $(K_{it}^* - K_{it})/K_{it}$ , because the slope of the  $MRP_{it}^K$  schedule is steeper for these firms.

decrease in size for firms with less than 15 employees, possibly reflecting the relaxation of other frictions as firms become older and bigger. After the 15-employees threshold, however, the relation between distortions and size becomes positive. This finding suggests the increase in labor adjustment costs due to the labor market regulation offsets the reduction of other components of the shadow cost of labor that decrease as firms become older and larger (e.g., reduced adjustment costs of production, less economic uncertainty, or a relaxation of financial frictions).

**Dynamic effects** – Next, we estimate the empirical model in equation (10) to study the response of  $(L_{it}^* - L_{it})/L_{it}$  for firms to the left and to the right of the regulatory threshold. Results are reported in Table 8, Panel d. Confirming the results of section 5.2, we find that, in response to a positive TFPR shock,  $(L_{it}^* - L_{it})/L_{it}$  increases more for firms below the threshold (14 employees) than for firms at the threshold (15 employees). A 1-percentage-point positive increase in TFPR is found to increase under-employment by 5 percentage points for firms of size 15, relative to those of size 14. This result, which is robust to the inclusion of firm fixed effects, suggest firms with 15 employees are less likely to increase their size in response to a positive TFPR shock in order to avoid becoming subject to the provisions of Article 18 of the Italian Workers Statute.

## 7 Aggregate Implications: Misallocation, TFP, and Lost Output

Resource misallocation has received much recent attention as an important explanatory factor of the disparity in aggregate productivity and economic growth across and within countries (Banerjee and Duflo 2005; Restuccia and Rogerson 2008; Hsieh and Klenow 2009; Hopenhayn 2014). In the case of Italy, the broad consensus is that the country’s spectacular failure to sustain aggregate productivity growth and contemporaneous economic stagnation of the last 20 years can be attributed in large part to malfunctioning financial, labor, and product markets that jeopardized the efficiency of the allocation of resources across their different uses (Bugamelli and Lotti 2017).<sup>94</sup>

In this section, we use our micro-estimates to perform aggregate counterfactual exercises that bear on the question of how idiosyncratic distortions in investment and employment policies translate into a loss in aggregate output and TFP, and how the gains from reallocation of resources evolve over time and in different sectors of the Italian economy.

We begin by showing that, at the firm level, output could be higher if firms were able to adjust their input-demand mix to the one that closes the MRP-cost gaps at the observed market prices. For every firm-year observation, we reconstruct a counterfactual level of output ( $Y_{it}^*$ ) that it could have produced employing  $K_{it}^*$  and  $L_{it}^*$  (equation 11), and contrast it with a comparable measure of output ( $Y_{it}$ ) that uses the observed input demands  $K_{it}$  and  $L_{it}$ :

$$Y_{it}^* = e^{a_{it}} \cdot (K_{it}^*)^{\gamma_s} (L_{it}^*)^{1-\gamma_s} \quad (13)$$

$$Y_{it} = e^{a_{it}} \cdot (K_{it})^{\gamma_s} (L_{it})^{1-\gamma_s} , \quad (14)$$

where  $a_{it} = va_{it} - \gamma_s k_{it} - (1 - \gamma_s)l_{it}$  is the estimated firm-level valued added productivity using value added cost shares of each 4-digit industry. The value-added production function above is obviously an over-

<sup>94</sup>Data from the Italian National Statistical Institute show the Italian year-on-year TFP growth was, on average, 0.05% in the decade 1997-2007, and of -0.3% between 2008 and 2013. Data available at <https://dati.istat.it> (access October 2017). Bugamelli and Lotti (2017) highlights the importance of frictions in labor, capital, and output markets in preventing Italian firms from effectively responding to the competitive pressures of increasingly globalized markets, and benefit from the opportunities offered by technological innovation and EU integration.

simplification of firms' production function. However, it serves our purpose. Indeed, any misspecification of firms' production process is held constant in (13) and (14).<sup>95</sup> Also, note that productivity and factor elasticities are held constant. Thus, the difference between  $Y_{it}^*$  and  $Y_{it}$  only arises from a different input mix.

We calculate the implied deviations from target output as  $(Y_{it}^* - Y_{it})/Y_{it}$ , and present their distribution in Table 2 and Figure 2 (Panel c). On average, output could be 12% higher if inputs were chosen to equalize marginal revenue products to the observed user costs. One might expect  $Y_{it} < Y_{it}^*$  for the large majority of the observations if firm policies were somehow constrained, and if the adjustments  $K_{it} \rightarrow K_{it}^*$  and  $L_{it} \rightarrow L_{it}^*$  move firms close to their production possibility frontier. Consistent with this prediction, we find  $Y_{it} < Y_{it}^*$  in 85% of the cases. Importantly, this positive output gap does not simply result from more inputs. To see this, note that target inputs demands (either  $K_{it}^*$  or  $L_{it}^*$ , or both) are lower than firms actual input demands for 40% of the observations. That is, a significant fraction of firms in the economy could produce *more output employing fewer resources*, by simply utilizing a more efficient input-mix.

At an aggregate level, we calculate that the output produced by the Italian corporate sector as a whole would grow by 8% to 9% if firms could adjust their input demands so to close estimated MRP-cost gaps at the observed user costs.<sup>96</sup> These aggregate effects result from the interplay of two forces. Mechanically, output grows because - on aggregate - 9% more labor and 1% more capital is needed in the economy in order to fully close all gaps.<sup>97</sup> The second force generating the output gain is resource reallocation. Holding constant the aggregate endowments of capital and labor, output would grow as resources freed by negative-gap producers reach positive-gap producers.<sup>98</sup>

**Reallocation algorithm** – To isolate the output gains that can be *directly* imputed to aggregate gains in TFP driven by a more efficient allocation of the resources, we define the following counterfactual outcomes:

$$\begin{aligned} Y_{it}^{**} &= e^{a_{it}} \cdot (K_{it}^{**})^{\gamma_s} (L_{it}^{**})^{1-\gamma_s} \\ \text{s.t.} \quad L_t^{**} &= \sum_i m_{it}^L L_{it}^* = L_t \\ K_t^{**} &= \sum_i m_{it}^K K_{it}^* = K_t, \end{aligned} \tag{15}$$

where  $m_{it}^L \geq 0$  and  $m_{it}^K \geq 0$  are reallocation weights that meet the following criteria:

1.  $m_{it}^X \geq 0$  when  $\tau_{it}^X > 0$ .
2. when  $a_{jt} \geq a_{it}$ : (i)  $m_{jt}^X \leq m_{it}^X$  if  $(\tau_{jt}^X \leq 0 \ \& \ \tau_{it}^X \geq 0)$ ; (ii)  $m_{jt}^X \geq m_{it}^X$  otherwise.

The constraints in (15) ensure that the reallocation takes place with no change in the aggregate capital and labor endowment of the economy. The reallocation weights require resources to move in a welfare-enhancing direction: from negative MRP-cost gap producers toward positive gap firms (criterion 1), and following a productivity rank (criterion 2).<sup>99</sup> Appendix K.1 provides a detailed explanation of a reallocation algorithm that satisfies these criteria. In short, we group firms into positive and negative MRP-gaps. Then, we reallocate

<sup>95</sup>Value added specifications, and the corresponding estimates of productivity and elasticities are problematic from both a theoretical and an empirical point of view (see Gandhi et al. 2017a). In this context, however, we are interested in reconstructing counterfactual output levels obtained using the target labor and capital ( $K_{it}^*$  and  $L_{it}^*$ ), and to contrast this quantity to a comparable output level obtained using the actual input demands observed in the data ( $K_{it}$  and  $L_{it}$ ).

<sup>96</sup>For every year  $t$ , we calculate the loss in aggregate output aggregating firm-specific deviations from target output  $\frac{Y_t^* - Y_t}{Y_t} = \frac{\sum_i Y_{it}^* - \sum_i Y_{it}}{\sum_i Y_{it}}$ . See Appendix K.2.

<sup>97</sup>Appendix K.2 analyzes the difference between aggregate input demands in the economy ( $K_t = \sum_i K_{it}$  and  $L_t = \sum_i L_{it}$ ) and the aggregate endowments that would fully close the gaps at the observed prices ( $K_t^* = \sum_i K_{it}^*$  and  $L_t^* = \sum_i L_{it}^*$ ).

<sup>98</sup>For the United States, Chaney et al. (2017) estimate the aggregate losses generated by binding collateral constraints at the firm-level. They find that the aggregate endowment channel matter twice as much as the misallocation channel.

<sup>99</sup>Another way to restate criterion 2 is  $\frac{\partial P\{m_{it}^X > 0\}}{\partial \omega_{it}} \geq 0$  if  $\tau_{it}^X \geq 0$ ;  $\frac{\partial P\{m_{it}^X < 0\}}{\partial \omega_{it}} \leq 0$  if  $\tau_{it}^X < 0$ .



resources away from the lowest-TFP firm belonging to the negative-gap group, and toward the highest-TFP firms in the positive-gap group. The reallocation stops when the aggregate constraint binds ( $X_t^{**} = X_t$ ).<sup>100</sup>

To study the scope of resource reallocation, we contrast  $Y_{it}^{**}$  to  $Y_{it}$  and construct our measure of aggregate output and TFP gains from reallocation as:

$$\frac{Y_t^{**} - Y_t}{Y_t} = \frac{TFP_t^{**} - TFP_t}{TFP_t},$$

where  $Y_t^{**} = \sum_i Y_{it}^{**}$  and  $Y_t = \sum_i Y_{it}$ . Because resources are reallocated with no change in the aggregate endowment, the gap between  $Y_t^{**}$  and  $Y_t$  measures output gains as well as aggregate TFP gains from reallocation.

Table 10 presents the output (TFP) gains that accrue from reallocation of resources from over-endowed to under-endowed producers. The solid line in Figure 10 provides a graphical visualization of its time-series. Averaging across years, we find that aggregate output and TFP could be 3% to 4% higher if approximately 1% of capital and labor could be re-allocated toward high-value use producers, without changing the aggregate endowments. The dashed line in Figure 10 shows that reallocation alone can explain between 35% and 45% of the output gain that would accrue if producers could fully close the estimated capital and labor gaps.<sup>101</sup>

A number of remarks are in order. First, because productivity and factor elasticities are held constant, the difference between  $Y_{it}^{**}$  and  $Y_{it}$  only arises from a different input mix. Although firm-level productivity growth is of high interest *per se*, it is not the focus of this paper. By shutting down this channel, we focus on gains from reallocation only. Second, gains from reallocation are a function of the size of the absolute gap of firm-level distortions in capital and labor. Thus, for reallocation to have significant aggregate effects, producers with both positive and negative MRP-cost gaps must be present in the economy. Figure 2 shows that, indeed, both types of producers are present in our database. Finally, our reallocation algorithm moves resources away from over-endowed and low-productivity producers toward under-endowed and high-productivity producers. This implies that, in our calculations, what matters is *both* the relative size of investment and employment distortions (i.e., the distance between  $K_{it}^{**}$  and  $K_{it}$  and between  $L_{it}^{**}$  and  $L_{it}$ ) and the correlation between firm-level distortions and firm-level productivity (Restuccia and Rogerson 2008). In the data, the correlation between distortions and productivity is positive and highly significant: within narrowly defined industries (4 digits) and years, a one-standard-deviation increase in TFPR ( $\omega_{it}$ ) is associated with a 0.07- and 0.16-standard-deviations increase in  $\frac{(K_{it}^{**}-K_{it})}{K_{it}}$  and  $\frac{(L_{it}^{**}-L_{it})}{L_{it}}$ , respectively. The last column in Table 10 shows that by shutting down the productivity pecking order (criterion 2 of the reallocation weights), and reallocating the resources of over-endowed firms toward randomly selected under-endowed firms, the gains from reallocation would be lower by a factor of 2.

**Business cycle fluctuations** – The times-series evolution of  $(Y_t^{**} - Y_t)/Y_t$  offers interesting insights into the importance of resource misallocation in different phases of the business cycle, and in particular during periods of credit expansion versus periods of credit crunch. We find that financial crisis periods are characterized by an increase in resource misallocation. Compared to the 1997–2004 period, gains from reallocation are 1/3 higher after the transmission of the global financial crisis to Italy (2008–2009) and the burst of the European Sovereign debt crisis that followed (2010–2013). In fact, reallocation can explain 35%

<sup>100</sup>Evidently, it must be the case that  $X_{it}^{**} < X_{it}^*$  for some firms in the economy when  $X_t^* > X_t$ , and  $X_{it}^{**} > X_{it}^*$  for some firms in the economy when  $X_t^* < X_t$ . When  $X_t^* < X_t$  (not the case in our dataset), we first set  $X_{it}^{**} = X_{it}^*$  for all firms, and then we reallocate the difference  $X_t - X_t^*$  across firms in a way that is proportional to the relative productivity  $a_{it}$ .

<sup>101</sup>In Appendix L we calculate the gains from reallocation using MRP-cost gaps constructed with alternative proxies for the user costs of capital and labor ( $r^{Cred Lines}$  and  $w^{New Hires}$ ) and find slightly higher gains from reallocation.

of the output (TFP) gains during the 1997–2004 period, but 40%–45% of the gains during financial and sovereign crisis periods.

A substantial body of empirical evidence documents a decline in TFP during episodes of financial crisis (Calvo et al. 2006).<sup>102</sup> Our findings show the strong co-integration of business-cycle fluctuations and TFP observed during episodes of financial instability might be explained, at least in part, by a deterioration of the efficiency of resource allocation (Ziebarth 2012; Oberfeld 2013; Sandleris and Wright 2014). In a contemporaneous empirical study, Manaresi and Pierri (2017) and Dörr et al. (2017) document a loss in average *firm-level productivity* of Italian corporations during the financial and sovereign crisis, and relate it to credit market frictions. Our results complement theirs, as we focus on the *reallocation channel* rather than the productivity channel.<sup>103</sup> Our findings are also in line with Schivardi et al. (2017) who document evidence of credit misallocation in Italy during the financial crisis.

Figure 10 also provides interesting insights into the long-term trends in misallocation in Italy. The potential output and TFP loss due to a detrimental allocation of capital and labor across industries are relatively flat in the late 90s and early 2000s. Then misallocation starts to increase from the second half of the 2000s, reaching its peak during the financial crisis. Below, we compare this pattern with the time-series patterns obtained using alternative measures of misallocation proposed by previous literature.

**Sectoral and spatial components** – To better understand in which parts of the economy misallocation is generating larger output losses, we disentangle the contribution of unconstrained reallocation (between-industries and within-industries) from the contribution of within-industry reallocation, and the one of a reallocation that is constrained to take place only within the same industry and same macro-region.<sup>104</sup> Comparing the solid and dotted line in Figure 11, we find that the majority of allocative inefficiencies take place within narrowly defined 4-digit industries (roughly two-thirds). The remaining third of the welfare gain could be achieved mobilizing resources across industries (Bartelsman et al. 2013). Finally, roughly two-thirds of the within-industry gains can be achieved if we constraint reallocation to take place within the same geographical macro-regions (dashed line in Figure 11).

Given these findings, a natural question concerns whether the scope for reallocation is similar in all industries and local markets, or rather is driven by some specific sectors or geographical regions. Figure 11 (Panel b) displays the output gains from within-industry reallocation for three different macro-sectors of the economy: manufacturing, services, and construction. Our estimates indicate output could grow about 2.5%–3% by improving the allocation of capital and labor among produces in manufacturing firms, 3.5%–4% in the services industry, and over 4% in the construction industry. These are novel findings. Indeed, mostly due to the lack of comprehensive data, the large majority of the empirical studies that assess the costs of misallocation focus on manufacturing industries. Considering the growing importance of non-manufacturing industries in both developed and developing economies, our analysis suggests that by extrapolating the evi-

<sup>102</sup>Calvo et al. (2006) analyzes 22 severe crises in emerging markets and finds that output and TFP typically decline by 10% and 9.5%, respectively. Examining the Chilean economy, Oberfeld (2013) performs an analysis similar to that of Hsieh and Klenow (2009) that documents a significant change in misallocation and consequent loss in TFP during the Chilean crisis of 1982. Sandleris and Wright (2014) and Ziebarth (2012) find a simultaneous decline of TFP and allocative efficiencies studying the Argentine Crisis of 2001 and several U.S. industries during the Great Depression.

<sup>103</sup>Manaresi and Pierri (2017) also use CR data and firm-level balance sheet data from Cerved, but they focuses on a subsample of larger firms. Although we adopt a different technique for production function estimation, the time-series evolution of TFPR is fully consistent across papers. For the subsample of our data that overlaps with theirs, the correlation between our measure of TFPR and theirs exceeds 85%.

<sup>104</sup>The reallocation algorithm adopted to construct the three aggregate measures is the same. The difference is in the level at which the aggregate constraints are required to hold. In the first case, we hold constant the aggregate amount of resources for every year ( $X_t^{**} = X_t$ ). In the second case we do so for every 4 digits industry-year pair ( $X_{st}^{**} = X_{st} \forall$  industry  $s$ ), in the third case for every 4 digits industry-macro-region-year pair ( $X_{srt}^{**} = X_{srt} \forall$  industry  $s$  and  $\forall$  macro-region  $r$ ).

dence on the manufacturing sector to the whole economy, researchers might be underestimating the potential welfare losses resulting from market frictions and regulations.

Figure 11 (Panel c) plots output gains from reallocation for different Italian macro-regions. We previously pointed out that southern and northern regions of the country are characterized by significantly different socio-economic outcomes, which can be traced to different historical backgrounds, quality of institutions, culture, and stock of social and human capital (Putnam et al. 1994). For example, we highlighted the difference in terms of efficiency of bankruptcy courts between northern and southern regions of the country. In line with these observations, we find the extent of resource misallocation is significantly higher in south than in the north or center of Italy. Our estimates indicate that, without changing the amount of capital and labor in use, a better allocation of resources might rise output and productivity in Southern regions by 4.5%–5%. The scope for reallocation is lower as we move north (3.5%–4.5% in the center, and over 2.5–3.5 percent in the north), although, interestingly, the post-2004 rise in misallocation seems to be driven by these regions, whereas no upward trend is observed for southern regions.

A number of remarks are due. First, we must note that the reallocation of capital and labor across producers is expected to have an impact on the level and distribution of factor prices, even when the aggregate amount of capital and labor in the economy does not change. Our analysis does not consider these important general equilibrium effects. In general, the direction and magnitude of these effects depend on the characteristics of the firms from and to which resources are mobilized. Second, our sample includes only incorporated firms. Thus, there is a question related to the generalization of the previous results to the whole Italian economy. Due to their smaller size, greater opacity, and lack of managerial capital, the non-corporate sector may exhibit investment and employment policies more distorted than the ones we found for the limited liability firms in our database (Midrigan and Xu 2014). Thus, our calculations might understate the scope of misallocation in the whole economy. Finally, we shall emphasize that our exercise takes the set of producers as given. Thus, it does not account for a particular form of misallocation that has to do with the choice of which producers end up operating (selection effect).<sup>105</sup>

**Alternative measures of misallocation** – We conclude this section with a comparison of our estimates of allocative efficiency gains to alternative measures adopted in the literature. Figure A.15, Panel a, in Appendix K.3 presents the time-series evolution of the OP-covariance term (Olley and Pakes, 1996), that is, the correlation between firm-level productivity ( $\omega_{it}$ ) and local market share ( $\text{Revenues}_{ispt}/\sum_i \text{Revenues}_{ispt}$ ,  $s$ =industry and  $p$ =province). Although this indicator of allocative efficiency displays pretty large variability over time, we obtain a significant steady decline of it starting in 2005, which is consistent with the gradual increase in gains from reallocation shown in Figure 10. Panel b of Figure A.15 displays the within-industry standard deviation of  $\ln(MRP^K)$  and  $\ln(MRP^L)$ . Starting from the seminal work of Hsieh and Klenow (2009), the dispersion of marginal revenue products has been largely used in cross-country analyses because, under specific model assumptions, it is proportional to the dispersion in TFPR which, in turn, is inversely proportional to the efficiency of within-industry resources allocation. Consistent with our measures, we find that dispersion in both Marginal Products is increasing over time, particularly for capital and especially during the financial and sovereign crisis.<sup>106</sup> Panel c of Figure A.15 plots inferred allocative efficiency for

<sup>105</sup>Jeong and Townsend (2007) and Midrigan and Xu (2014) highlight the importance of financial friction on the extensive margin.

<sup>106</sup>An influential study by Gopinath et al. (2016) (GKKV, henceforth) looks at the dispersion of log MRP to investigate the extent of misallocation in Europe, arguing that the productivity slowdown in Spain, Italy, and other Southern European countries may have been driven by the credit expansion that followed the establishment of the European Monetary Union. For Italy, the data source for the accounting variables used by GKKV is ultimately the Cerved database, in its release by Bureau Van Dijk. Thus, it is reassuring that the trends in log- $MRP^K$  and log- $MRP^L$  dispersion we obtain are very similar to those

Italy, calculated by taking to the data the model in Bils et al. (2017) (BKR). Consistent with our results, the BKR procedure highlights a downward trend in the allocative efficiency starting in 2005. Finally, in Panel d, we plot the average absolute deviation from target output ( $\sum_i \frac{1}{N_t} |(Y_{it}^* - Y_{it})/Y_{it}|$ , where  $N_t$  is the number of observations in year  $t$ ) across the different years of our sample. This measure, inspired by the measure of allocative efficiency developed by Amil Petrin and coauthors (Petrin et al. 2011; Petrin and Levinsohn 2012; Petrin and Sivadasan 2013), correlates strongly with the economic downturn of early 2000s, and it confirms the aggregate dynamics obtained with our reallocation algorithm starting in 2005.<sup>107</sup>

## 8 Concluding Remarks and Future Research

In this paper, we combine information on firm-specific borrowing costs and wages with estimates of the marginal returns to produce empirical measures of deviations in firms' first-order conditions. The analysis of the distribution and variation of MRP-cost gaps provides valuable insights into the impact of credit and labor market frictions on firm policies, and into the aggregate implications of resource misallocation in terms of output and aggregate productivity loss. Our approach is capable of guiding researchers toward the primitive frictions affecting firm policies and can be used to test the effects of policy interventions that economists want to confront directly with the data. Because they require no information on the firm market value of assets or liabilities, MRP-cost gaps are a particularly valuable tool to study investment and employment policies of privately owned firms, for which standard empirical measures are unavailable.

The natural extension of our analysis is the estimation of a fully-fledged dynamic structural model that, by using MRP-cost gaps as additional moments to target in the estimation, can further cast light and disentangle the impact of other market frictions such as taxes, moral hazard behavior, or uncover the implicit cost of equity financing for private firms. With this respect, a key modeling issue is to incorporate the effect of frictions and regulations on endogenous decisions that affect future realizations of firm-level productivity.

The procedure that calculates the aggregate output and TFP gains from a more efficient allocation of resources is transparent and intuitive, but does not consider the general equilibrium effects on the level and distribution of interest rates and wages that may result from reallocation. In the future, it would be interesting, and important, to incorporate these general equilibrium spillovers, without neglecting the heterogeneity across producers and its micro-foundation (Buera and Moll 2015).

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of GKKV, with a particularly remarked upward trend for capital. We shall emphasize, however, that the time series of the dispersion of log-MRP<sup>L</sup> differs from the one in GKKV: downward in their paper, but upward in our data. This is likely due to the fact that we measure labor services in effective units, while the authors use the total wage bill.

<sup>107</sup>In a sequence of papers, Amil Petrin and coauthors (Petrin et al. 2011; Petrin and Levinsohn 2012; Petrin and Sivadasan 2013) have proposed measures of allocative efficiency based on the aggregation of firm-level gaps between a firm Value of the Marginal Products and factor prices (VMP-cost gaps). They theoretically demonstrate that the integration of VMP-cost gaps provides us with a direct estimate of Aggregate Productivity Growth (APG) in the spirit of Hulten (1978) and Basu and Fernald (2002).

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Table 1: Summary statistics

This table reports the summary statistics of the main variables used in the paper. A description of the variables is provided in Section 2 and Appendix A.

	Mean	Std	p10	p25	p50	p75	p90
REVENUES	3598	11230	142	320	825	2314	6661
TOTAL ASSETS	3246	10321	121	269	706	2027	5986
AGE	12	11	2	4	9	18	28
EMPLOYEES	17	39	1	3	6	14	33
ASSETS TURNOVER	1.47	1.42	0.48	0.82	1.25	1.80	2.54
ROA	0.03	0.18	-0.09	0.01	0.04	0.09	0.16
CASH FLOWS / ASSETS	0.04	0.32	-0.04	0.02	0.05	0.10	0.17
BANK LEVERAGE	0.44	0.45	0.00	0.06	0.35	0.65	0.98
LENGTH RELATIONS <sup>mean</sup>	3.6	2.6	0.9	1.6	3.0	4.9	7.1
LENGTH RELATIONS <sup>wmean</sup>	3.8	2.8	0.9	1.7	3.0	5.3	7.9
LENGTH RELATIONS <sup>lead</sup>	4.1	3.6	0.8	1.3	3.0	5.8	9.5
NUMBER RELATIONS	4.1	3.5	1.0	2.0	3.0	5.0	8.0
CREDIT RATING	5.4	5.7	2.0	4.0	5.0	7.0	8.0
LENGTH BANKRUPTCY CASE	8.2	2.0	6.1	6.3	7.9	9.6	10.8
BORROWER	0.80						
MANUFACTURING	0.31						
SERVICES	0.53						
CONSTRUCTION	0.17						
OBSERVATIONS	3933209						
FIRMS	650489						

Table 2: marginal revenue products, user costs, MRP-cost gaps, and percentage deviations

Panel a reports summary statistics describing the distribution of the marginal revenue product of capital ( $MRP_{it}^K$ ) and labor ( $MRP_{it}^L$ ).  $MRP_{it}^K$  is expressed in percent;  $MRP_{it}^L$  is expressed in thousands of Euros. Panel b reports summary statistics describing the distribution of the annual percentage rate (APR) on bank loans ( $r_{it+1}$ ), depreciation rate on capital ( $\delta_s$ ), the sum of the two (user cost of capital,  $r_{it+1} + \delta_s$ ), and wage (user cost of labor,  $w_{it}$ ). Interest rates, depreciation rates and user costs of capital are expressed in percentages; wages are expressed in thousands of Euros. Panel c presents the descriptive statistics of the distribution of MRP-cost gaps  $\tau_{it}^K$  and  $\tau_{it}^L$ . Capital gaps are expressed in percentages; labor gaps are expressed in thousands of Euros. Panel d reports summary statistics of the distribution of percentage deviations from target input demands ( $\frac{L_{it}^* - L_{it}}{L_{it}}$ ,  $\frac{K_{it}^* - K_{it}}{K_{it}}$ ) and implied deviation from target output  $\frac{Y_{it}^* - Y_{it}}{Y_{it}}$ , which are expressed in percentages. Summary statistics are reported for the full sample and, for capital-related variables, also splitting the sample into borrowers with active loans (BORROWER-LOANS = 1), borrowers with credit lines only (BORROWER-NOLOANS = 1), and non-borrowers (BORROWER = 0). Block-bootstrapped standard errors are in parenthesis.

	<i>Whole Sample</i>				<i>Borrowers-Loans</i>				<i>Borrowers-NoLoans</i>				<i>Non-Borrowers</i>			
	MEAN	MEDIAN	10 <sup>th</sup>	90 <sup>th</sup>	MEAN	MEDIAN	10 <sup>th</sup>	90 <sup>th</sup>	MEAN	MEDIAN	10 <sup>th</sup>	90 <sup>th</sup>	MEAN	MEDIAN	10 <sup>th</sup>	90 <sup>th</sup>
<b>Panel a: User Costs</b>																
$r + \delta$	16.4	16.7	12.1	19.9	16.4	16.4	12.6	20.0	15.9	16.3	11.1	19.3	16.8	17.3	12.0	19.3
$r$	5.9	5.8	3.5	8.4	5.5	5.3	3.0	8.3	5.9	5.8	4.1	8.0	6.8	6.6	4.8	8.7
$\delta$	10.5	11.4	5.6	12.5	10.8	11.4	7.5	12.5	10.0	10.1	5.5	11.9	10.1	10.3	5.6	11.9
$w$	18.8	17.8	10.7	27.7												
<b>Panel b: Marginal Revenue Products</b>																
$MRP^K$	54.6	19.6	2.8	113.4	44.6	17.7	2.7	180.1	62.0	19.7	2.6	131.4	80.0	27.0	3.4	180.1
	(0.11)	(0.04)	(0.01)	(0.28)												
$MRP^L$	28.0	24.6	9.8	49.1												
	(0.02)	(0.02)	(0.01)	(0.05)												
<b>Panel c: MRP-Cost Gaps</b>																
$\tau^K$	36.4	3.2	-12.4	91.6	26.9	1.4	-12.5	71.1	43.8	4.0	-12.4	109.0	60.1	9.8	-12.0	154.4
	(0.10)	(0.03)	(0.01)	(0.26)												
$\tau^L$	9.3	6.1	-6.4	28.3												
	(0.02)	(0.01)	(0.01)	(0.04)												
<b>Panel d: Percentage Deviations</b>																
$(K^* - K)/K$	15.5	0.4	-1.9	16.6	9.2	0.2	-1.9	12.0	17.3	0.5	-1.8	19.5	32.4	1.2	-1.7	42.1
	(0.04)	(0.00)	(0.00)	(0.08)												
$(L^* - L)/L$	10.7	3.0	-7.2	35.4												
	(0.06)	(0.01)	(0.01)	(0.11)												
$(Y^* - Y)/Y$	12.0	4.3	-4.0	35.6	9.4	4.0	-3.6	26.1	12.1	4.5	-5.4	37.7	17.8	5.1	-4.6	59.5
	(0.00)	(0.00)	(0.00)	(0.00)												



Table 3: Revenue elasticities, returns to scale, markups, and elasticities

This table displays the estimates of firm-level production function parameters, returns to scale, markups, and revenue productivity. We report average, interquartile range, and block-bootstrapped standard errors of the mean (in parenthesis). The first block reports the statistics across all firm-years. The second and third block split the sample into manufacturing and non-manufacturing firms, respectively. In each block, the first four rows of table show the estimates of output elasticities with respect to capital ( $\theta_{it}^K$ ), labor ( $\theta_{it}^L$ ), intermediate inputs ( $\theta_{it}^M$ ). The fourth row reports the estimated returns to scale ( $RS_{it} = \sum_X \theta_{it}^X$ ,  $X = \{K, L, M\}$ ). The fifth and sixth rows report the summary statistics of the estimated markups ( $\mu_{it}$ ) and revenue productivity (TFPR,  $\omega_{it}$ ), respectively.

	ALL INDUSTRIES		MANUFACTURING		NON MANUFACTURING	
	MEAN	75-25	MEAN	75-25	MEAN	75-25
$\theta^K$	0.04 (0.6·10 <sup>-4</sup> )	0.04	0.05 (0.6·10 <sup>-4</sup> )	0.04	0.04 (0.6·10 <sup>-4</sup> )	0.04
$\theta^L$	0.29 (2.2·10 <sup>-4</sup> )	0.21	0.30 (2.2·10 <sup>-4</sup> )	0.16	0.29 (2.2·10 <sup>-4</sup> )	0.23
$\theta^M$	0.67 (1.9·10 <sup>-4</sup> )	0.22	0.67 (1.9·10 <sup>-4</sup> )	0.16	0.68 (1.9·10 <sup>-4</sup> )	0.24
$RS$	1.01 (2.2·10 <sup>-4</sup> )	0.05	1.02 (2.2·10 <sup>-4</sup> )	0.05	1.01 (2.2·10 <sup>-4</sup> )	0.06
$\mu$	1.02 (0.6·10 <sup>-4</sup> )	0.16	1.01 (0.6·10 <sup>-4</sup> )	0.15	1.02 (0.6·10 <sup>-4</sup> )	0.16
$\omega$	2.52 (14.5·10 <sup>-4</sup> )	0.71	2.64 (14.5·10 <sup>-4</sup> )	0.36	2.46 (14.5·10 <sup>-4</sup> )	0.90

Table 4: MRP-Cost gaps and firm's characteristics

This table reports the correlation between MRP-cost gaps and firm characteristics. Panel a focuses on the MRP-cost gap of capital ( $\tau^K$ ) and panel b on the MRP-cost gap of labor ( $\tau^L$ ). We regress gaps on life cycle variables (firm age and size), credit score, measures of productivity and profitability (TFPR and ROA), and proxies of internal and external financing (cash-over-assets and bank leverage, respectively). Age groups are defined as follows: young if age  $\leq 5$ , medium if age  $\in (10]$ , old if age  $> 10$ . Assets groups are defined based on the terciles of the distribution of assets (average assets across firms in each tercile are 190 thousand, 760 thousand, and 8.8 million Euros). Credit score groups are defined as follows: safe firms are those with a credit score ranging from "Excellent" to "Solvent" (credit score from 1 to 4); a second group includes firms classified as "Vulnerable" and "Very vulnerable" (credit score from 5 and 6); Risky are firms with credit score ranging from "Risky" to "Very very risky" (credit score from 7 and 9). We focus on within-year and within-industry variation by controlling for year and industry fixed effects. All variables in the regressions are standardized so that coefficients are expressed as Z-scores. Standard errors (in parenthesis) are clustered at the firm level.

**Panel a: MRP-Cost gap of capital ( $\tau_{it}^K$ )**

AGE		CREDIT SCORE		TFPR
YOUNG	OMITTED CATEGORY	SAFE	OMITTED CATEGORY	0.411 (0.004)***
MEDIUM	-0.273 (0.002)***	VULNERABLE	0.019 (0.002)***	
OLD	-0.316 (0.002)***	RISKY	-0.048 (0.002)***	ROA
				0.151 (0.003)***
ASSETS		LEVERAGE		
SMALL	OMITTED CATEGORY	-0.066 (0.001)***		
MEDIUM	-0.077 (0.002)***			
LARGE	-0.116 (0.003)***	CASH / ASSETS		
		-0.193 (0.007)***		
YEAR FIXED EFFECTS		Y		
INDUSTRY FIXED EFFECTS		Y		
ADJ. $R^2$		0.106		
$R^2$ YEAR AND INDUSTRY FE ONLY		0.049		
OBSERVATIONS		3511678		

**Panel b: MRP-Cost gap of labor ( $\tau_{it}^L$ )**

AGE		CREDIT SCORE		TFPR
YOUNG	OMITTED CATEGORY	SAFE	OMITTED CATEGORY	0.798 (0.003)***
MEDIUM	0.082 (0.002)***	VULNERABLE	-0.011 (0.002)***	
OLD	0.059 (0.002)***	RISKY	-0.079 (0.002)***	ROA
				0.152 (0.001)***
ASSETS		LEVERAGE		
SMALL	OMITTED CATEGORY	0.0552 (0.001)***		
MEDIUM	0.009 (0.002)***			
LARGE	0.382 (0.003)***	CASH / ASSETS		
		-0.007 (0.001)***		
YEAR FIXED EFFECTS		Y		
INDUSTRY FIXED EFFECTS		Y		
ADJ. $R^2$		0.292		
$R^2$ YEAR AND INDUSTRY FE ONLY		0.128		
OBSERVATIONS		3863961		

Table 5: Information frictions and relationship lending

This table explores the relation between length of lending relationships ( $LENGTH\ RELATION_{it}^{wmean}$ ) and borrowing rates ( $r_{t+1}$ ), marginal revenue products of Capital ( $MRP_{it+1}^K$ ), and MRP-cost gaps of capital ( $\tau_{it}^K$ ). In Panel a, firm-level controls include: age and size dummies (deciles), credit score dummies (9 values), assets turnover, ROA, cash flows over assets, leverage, and the number of active credit relationships. These regressions include year by province by industry (2-digits industry codes) fixed effects. In Panel b we augment the specification with firm fixed effects, and replace age and size deciles with a second order polynomial in age and lag of log assets. Columns (5)–(7) in Panel a and Columns (2)–(4) in Panel b, also include the interactions of all variables and fixed effects with  $UNDERCAPITALIZED_{t-1}$  ( $=1$  if  $\tau_{it-1}^K > 0$ ),  $TFPR_t$  (mean-zero  $\omega_{it}$ ), and  $UNDERCAPITALIZED_{t-1} \times TFPR_t$ . Regressions are run on the sub-sample of borrowers with outstanding loans for which we observe the APR on loans. Standard errors (in parenthesis) are clustered at the firm level.

Panel a: Between Firm Regressions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$Default_{it+1}$	$r_{t+1}$	$MRP_{it+1}^K$	$ \tau_{it}^K $		$\tau_{it}^K$	
LENGTH RELATION $^{wmean}_t$	-0.001 (0.000)***	-0.020 (0.001)***	-1.138 (0.060)***	-0.936 (0.052)***	-0.029 (0.011)***	-2.273 (0.048)***	-0.019 (0.012)*
LENGTH RELATION $^{wmean}_t$ X UNDERCAPITALIZED $_{t-1}$					-0.851 (0.092)***		-0.744 (0.091)***
LENGTH RELATION $^{wmean}_t$ X TFPR $_t$						-1.215 (0.130)***	0.097 (0.029)***
LENGTH RELATION $^{wmean}_t$ X UNDERCAPITALIZED $_{t-1}$ X TFPR $_t$							-2.962 (0.199)***
TFPR $_t$	-0.027 (0.002)***	0.036 (0.009)***	36.998 (0.860)***	30.355 (0.757)***			
ASSETS TURNOVER $_t$	-0.004 (0.001)***	0.044 (0.003)***	23.495 (0.371)***	16.924 (0.306)***			
ROA $_t$	-0.141 (0.003)***	1.288 (0.038)***	114.577 (4.629)***	85.482 (3.660)***			
CASH FLOWS $_t$ / ASSETS $_t$	-0.008 (0.002)***	-0.933 (0.051)***	-131.025 (7.382)***	-102.490 (5.781)***			
BANK LEVERAGE $_t$	-0.004 (0.000)***	-0.565 (0.005)***	-8.774 (0.418)***	-6.843 (0.362)***			
NUMBER RELATIONS $_t$	0.004 (0.000)***	0.027 (0.001)***	-0.496 (0.067)***	-0.610 (0.059)***			
FIRM CONTROLS	-	-	-	-	Y	Y	Y
AGE, SIZE, CREDIT SCORE FE	Y	Y	Y	Y	Y	Y	Y
INDUSTRY X YEAR X PROVINCE FE	Y	Y	Y	Y	Y	Y	Y
YEAR FE	N	N	N	N	N	Y	N
FIRM FE	N	N	N	N	N	N	N
ADJ. R2	0.146	0.486	0.114	0.093	0.167	0.070	0.167
OBS.	1887314	1887314	1887314	1887314	1633484	1633697	1633484

Panel b: Within Firm Regressions

	(1)	(2)	(3)	(4)
	$ \tau_{it}^K $		$\tau_{it}^K$	
LENGTH RELATION $^{wmean}_t$	-1.489 (0.053)***	-0.473 (0.033)***	-0.476 (0.046)***	-0.462 (0.034)***
LENGTH RELATION $^{wmean}_t$ X UNDERCAPITALIZED $_{t-1}$		-0.509 (0.061)***		-0.462 (0.061)***
LENGTH RELATION $^{wmean}_t$ X TFPR $_{t-1}$			-0.685 (0.111)***	0.031 (0.051)
LENGTH RELATION $^{wmean}_t$ X UNDERCAPITALIZED $_{t-1}$ X TFPR $_t$				-1.568 (0.132)***
FIRM CONTROLS	Y	Y	Y	Y
AGE, SIZE, CREDIT SCORE FE	Y	Y	Y	Y
INDUSTRY X YEAR X PROVINCE FE	Y	Y	Y	Y
YEAR FE	N	N	Y	N
FIRM FE	N	N	N	N
ADJ. R2	0.575	0.638	0.624	0.639
OBS.	1846643	1614149	1614591	1614149

Table 6: **Bankruptcy costs**

This table explores the relation between bankruptcy costs (LENGTH BANKRUPTCY) and borrowing rates ( $r_{t+1}$ ), marginal revenue products of Capital ( $MRP_{it+1}^K$ ), and MRP-cost gaps of capital ( $\tau_{it}^K$ ). Firm-level controls include: age and size dummies (deciles), credit score dummies (9 values), assets turnover, ROA, cash flows over assets, leverage, weighted average length of lending relationships, and the number of active credit relationships. Province-level controls, measured in 2007, include: population, GDP, unemployment rate, active firms per resident, firm exit rate, Herfindahl-Hirschman Index of credit market concentration, and the number of active credit institutions. Columns (1)–(3) include year by industry fixed effects. Columns (8) and (9) include year by industry by macro-region (North, Center, South of Italy) fixed effects. Regressions are run on the sub-sample of borrowers with outstanding loans for which we observe the APR on loans. Standard errors (in parenthesis) are clustered at the firm level.

	(1)	(2)	(3)	(4)	(5)	(6)
	$r_{t+1}$	$MRP_{t+1}^K$	$\tau_t^K$	$r_{t+1}$	$MRP_{t+1}^K$	$\tau_t^K$
LENGTH BANKRUPTCY	0.013 (0.001)***	0.423 (0.086)***	0.363 (0.077)***	0.003 (0.001)**	0.208 (0.088)**	0.181 (0.079)**
FIRM CONTROLS	Y	Y	Y	Y	Y	Y
PROVINCE CONTROLS	Y	Y	Y	Y	Y	Y
INDUSTRY X PROVINCE FE	Y	Y	Y	N	N	N
INDUSTRY X YEAR X M. REGION FE	N	N	N	Y	Y	Y
ADJ. R2	0.468	0.113	0.110	0.471	0.114	0.111
OBS.	1822631	1822631	1822631	1822618	1822618	1822618

Table 7: Credit supply shifters

This table investigates the relationship between firm-specific credit-supply shocks and MRP-cost gap ( $\tau_{it}^K$ ). In Panel a, Column (1) reports the OLS coefficient; Column (2) and Column (3) the first stage regression, where the percentage change in credit ( $g(\text{CREDIT}_{it})$ ) is projected onto the credit supply shifter ( $\text{CREDIT\_SHIFTER}_{it}$ ); Column (4) and Column (5) reports the reduced form regressions, where we project the gap onto the credit shifter. In Columns (6)-(8) regressions also include the interactions of all variables with  $\text{UNDERCAPITALIZED}_{it-1}$  ( $=1$  if  $\tau_{it-1}^K > 0$ ),  $\text{TFPR}_{it-1}$  (mean-zero  $\omega_{it-1}$ ), and  $\text{UNDERCAPITALIZED}_{it-1} \times \text{TFPR}_{it-1}$ . Columns (9) and (10) split the variation in the credit supply shifters in a positive changes ( $\text{MAX}\{0, \text{CREDIT\_SHIFTER}_{it}\}$ ) and negative changes ( $\text{MAX}\{0, -\text{CREDIT\_SHIFTER}_{it}\}$ ). In Column (10) we include the interactions of all variables with  $\text{UNDERCAPITALIZED}_{it-1}$ . Panel b reports the Instrumental Variables (IV) regressions, where we instrument the percentage change in credit supply ( $g(\text{CREDIT\_SHIFTER}_{it})$ ) using the credit supply shifter. All regressions include 2-digit industry by year by province fixed effects and the following set of lagged controls: productivity (TFPR), the weighted average of the length of lending relationships, a second order polynomial in age, log assets, credit score dummies (9 values), assets turnover, ROA, cash flows over assets, leverage, and number of active credit relationships. In Columns (3)-(5) regressions also include the interactions of all variables with  $\text{UNDERCAPITALIZED}_{it-1}$ ,  $\text{TFPR}_{it-1}$ , and  $\text{UNDERCAPITALIZED}_{it-1} \times \text{TFPR}_{it-1}$ . In Columns (3) and (5) in Panel a, and in Column (2) in Panel b, we include firm fixed effects. Regressions are run on the sub-sample of borrowers with outstanding loans for which we observe the APR on loans. Standard errors (in parenthesis) are clustered at the firm level.

Panel a: OLS, first stage, and reduced form regressions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	
	OLS	First Stage		Reduced Form						Reduced Form	
	$\Delta\tau_t^K$	$g(\text{CREDIT}_t)$		$\Delta\tau_t^K$						Positive VS	Negative Shocks
										$\Delta\tau_t^K$	$\Delta\tau_t^K$
$g(\text{CREDIT}_t)$	-4.173 (0.062)***										
CREDIT SHIFTER <sub>t</sub>		0.209 (0.003)***	0.246 (0.003)***	-1.801 (0.149)***	-1.512 (0.160)***	-0.145 (0.054)	-2.373 (0.152)***	-0.145 (0.059)***			
CREDIT SHIFTER <sub>t</sub> x UNDERCAPITALIZED <sub>t-1</sub>						-2.131 (0.261)***		-2.019 (0.259)***			
CREDIT SHIFTER <sub>t</sub> x TFPR <sub>t-1</sub>							-1.625 (0.361)***	-0.008 (0.137)			
CREDIT SHIFTER <sub>t</sub> x UNDERCAPITALIZED <sub>t-1</sub> x TFPR <sub>t-1</sub>								-1.501 (0.614)**			
MAX{0,CREDIT SHIFTER <sub>t</sub> }									-3.101 (0.423)***	-0.074 (0.159)	
MAX{0,-CREDIT SHIFTER <sub>t</sub> }									1.319 (0.201)***	0.183 (0.074)**	
MAX{0,CREDIT SHIFTER <sub>t</sub> } x UNDERCAPITALIZED <sub>t-1</sub>										-4.293 (0.722)***	
MAX{0,-CREDIT SHIFTER <sub>t</sub> } x UNDERCAPITALIZED <sub>t-1</sub>										1.345 (0.359)***	
FIRM CONTROLS	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	
INDUSTRY x YEAR x PROVINCE FE	Y	Y	N	Y	N	Y	Y	Y	N	Y	
YEAR FE	N	N	Y	N	Y	N	N	N	Y	N	
FIRM FE	N	N	Y	N	Y	N	N	N	Y	N	
ADJ. R2	0.059	0.189	0.307	0.055	0.126	0.096	0.065	0.096	0.055	0.095	
OBS.	1610438	1610438	1578340	1610438	1578340	1595056	1610438	1595056	1610438	1595056	

Table 7 (cont'ed): **Credit supply shifters****Panel b: IV regressions**

	(1)	(2)	(3)	(4)	(5)
			$\Delta \tau_t^K$		
$g(\text{CREDIT}_t)$	-8.581	-4.675	-0.722	-13.580	-0.733
	(0.713)***	(0.649)***	(0.272)***	(0.877)***	(0.305)**
$g(\text{CREDIT}_t)$ X UNDERCAPITALIZED $_{t-1}$			-10.015		-9.630
			(1.234)***		(1.220)***
$g(\text{CREDIT}_t)$ X TFPR $_{t-1}$				-9.115	-0.093
				(1.661)***	(0.699)
$g(\text{CREDIT}_t)$ X UNDERCAPITALIZED $_{t-1}$ X TFPR $_{t-1}$					-8.974
					(2.787)***
FIRM CONTROLS	Y	Y	Y	Y	Y
INDUSTRY X YEAR X PROVINCE FE	Y	N	Y	Y	Y
YEAR FE	N	Y	N	N	N
FIRM FE	N	Y	N	N	N
OBS.	1610438	1578340	1595056	1610438	1595056

Table 8: Labor market regulations and response to TFP shocks

This table investigates the response of labor gaps ( $\tau_{it}^L$ ) and marginal revenue product of labor ( $MRP_{it}^L$ ), wages ( $w_{it}$ ), and percentage deviations from target labor ( $\frac{L_{it}-L_{it}^*}{L_{it}^*}$ ) to changes in firm-level productivity (TFPR,  $\omega_{it}$ ). The regression model is specified in equation (10). The vector of controls includes lagged TFPR, a quadratic in age, and (alternatively) industry by year by province fixed effects or year and firm fixed effects. Standard errors (in parenthesis) are clustered at the firm level.

	(1)	(2)	(3)	(4)
	CHANGES IN $\omega$	POSITIVE CHANGES IN $\omega$	CHANGES IN $\omega$	POSITIVE CHANGES IN $\omega$
<b>Panel a: Dep. Var. <math>\tau_{it}^L</math></b>				
$\delta_{15} - \delta_{14}$	1.468 (1.326)	5.013 (2.226)**	.607 (.992)	5.346 (2.226)**
$\delta_{15} - \delta_{16}$	1.222 (1.367)	5.156 (2.633)*	1.13 (1.337)	4.634 (2.285)**
<b>Panel b: Dep. Var. <math>MRP_{it}^L</math></b>				
$\delta_{15} - \delta_{14}$	1.774 (1.865)	5.022 (2.659)*	.738 (1.031)	5.28 (2.275)**
$\delta_{15} - \delta_{16}$	1.606 (1.634)	5.419 (2.744)**	1.403 (1.445)	4.96 (2.404)**
<b>Panel c: Dep. Var. <math>w_t</math></b>				
$\delta_{15} - \delta_{14}$	-.163 (.509)	-.403 (1.204)	-.13 (.277)	-.472 (.396)
$\delta_{15} - \delta_{16}$	-.145 (.682)	.047 (1.146)	.011 (.241)	.167 (.408)
<b>Panel d: Dep. Var. <math>(L_{it}^* - L_t)/L_t</math></b>				
$\delta_{15} - \delta_{14}$	.285 (1.327)	5.141 (1.931)**	-.448 (1.121)	3.191 (2.177)**
$\delta_{15} - \delta_{16}$	1.13 (1.271)	1.589 (3.92)	1.137 (1.051)	.715 (2.168)
FIRM CONTROLS	Y	Y	Y	Y
INDUSRY X YEAR X PROVINCE FE	Y	Y	N	N
YEAR	N	N	Y	Y
FIRM	N	N	Y	Y
OBS	1,759,883	727,419	1,794,323	724,687

Table 9: Percentage deviations from target capital and credit market frictions

This table investigates the relationship between percentage deviations from target capital ( $\frac{K_{it}-K_{it}^*}{K_{it}}$ ) and proxies of credit market frictions (Panel a) and credit-supply shocks (Panel b). In Panel a, the main regressors are: the weighted average length of lending relationships (LENGTH RELATION $_{it}$ <sup>wmean</sup>) and the length of bankruptcy litigations in court (LENGTH BANKRUPTCY). The set of firm-level controls includes age and size dummies (deciles), credit score dummies (9 values), assets turnover, ROA, cash flows over assets, leverage, and the number of active credit relationships. These regressions include year by province by industry (2-digits industry codes) fixed effects. In Columns (2)–(4) regressions also include the interactions of all variables and fixed effects with UNDERCAPITALIZED $_{it-1}$  (=1 if  $\tau_{it-1}^K > 0$ ), TFPR $_{it}$  (mean-zero  $\omega_{it}$ ), and UNDERCAPITALIZED $_{it-1}$  x TFPR $_{it}$ . In Columns (5) and (6) we include the following set of province-level controls, measured in 2007: population, GDP, unemployment rate, active firms per resident, firm exit rate, Herfindahl-Hirschman Index of credit market concentration, and the number of active credit institutions. In Column (5) we include year by industry fixed effects.; in Columns (6) we include year by industry by macro-region (North, Center, South of Italy) fixed effects. In Panel b, Column (1) reports the OLS coefficient; Columns (2)–(4) report the reduced form regressions, where we project the percentage deviation  $\frac{K_{it}-K_{it}^*}{K_{it}}$  onto the credit shifter (CREDIT SHIFTER $_{it}$ ); Columns (5)–(7) report the Instrumental Variables (IV) regressions, where we instrument the percentage change in credit supply using the credit supply shifter. All regressions include 2-digit industry by year by province fixed effects and the following set of lagged controls: productivity (TFPR), the weighted average of the length of lending relationships, a second order polynomial in age, log assets, credit score dummies (9 values), assets turnover, ROA, cash flows over assets, leverage, and number of active credit relations. In Columns (3)–(4) and (6)–(7) regressions also include the interactions of all variables and fixed effects with UNDERCAPITALIZED $_{it-1}$ , TFPR $_{it-1}$  (mean-zero  $\omega_{it-1}$ ), and UNDERCAPITALIZED $_{it-1}$  x TFPR $_{it-1}$ . Regressions are run on the sub-sample of borrowers with outstanding loans for which we observe the APR on loans. Standard errors (in parenthesis) are clustered at the firm level.

Panel a: Asymmetric information and bankruptcy costs

	(1)	(2)	(3)	(4)	(5)	(6)
	$ (K_{it}^* - K_{it})/K_{it} $		$(K_{it}^* - K_{it})/K_{it}$		$(K_{it}^* - K_{it})/K_{it}$	
LENGTH RELATION $_{t-1}$ <sup>wmean</sup>	-0.317 (0.029)***	-0.020 (0.014)**	-0.775 (0.026)***	-0.023 (0.014)	-0.171 (0.026)***	-0.178 (0.026)***
LENGTH RELATION $_{t-1}$ <sup>wmean</sup> x UNDERCAPITALIZED $_{t-1}$		(0.049)***		(0.049)***		
LENGTH RELATION $_{t-1}$ <sup>wmean</sup> x TFPR $_t$			-0.355 (0.064)***	0.049 (0.030)		
LENGTH RELATION $_{t-1}$ <sup>wmean</sup> x UNDERCAPITALIZED $_{t-1}$ x TFPR $_t$				-0.781 (0.104)***		
LENGTH BANKRUPTCY					0.103 (0.041)**	0.032 (0.042)**
FIRM CONTROLS	Y	Y	Y	Y	Y	Y
PROVINCE CONTROLS	N	N	N	N	Y	Y
AGE AND SIZE AND CRED. SCORE FE	Y	Y	Y	Y	Y	Y
INDUSTRY X YEAR X PROVINCE FE	Y	Y	Y	Y	N	N
INDUSTRY X PROVINCE FE	N	N	N	N	Y	N
INDUSTRY X YEAR X M. REGION FE	N	N	N	N	N	Y
ADJ. R2	0.025	0.032	0.032	0.032	0.030	0.030
OBS.	1887314	1633484	1633697	1633484	1822631	1822618

Panel b: credit-supply shocks

Dep Var: $\Delta((K_{it}^* - K_{it})/K_{it})$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	Reduced Form				IV	
$g(\text{CREDIT}_t)$	-0.566 (0.016)***				-1.262 (0.188)***	-0.426 (0.076)***	-0.409 (0.084)***
CREDIT SHIFTER $_t$		-0.265 (0.039)***	-0.086 (0.015)**	-0.082 (0.016)***			
CREDIT SHIFTER $_t$ x UNDERCAPITALIZED $_{t-1}$			-0.168 (0.073)**	-0.151 (0.072)**			
CREDIT SHIFTER $_t$ x TFPR $_{t-1}$				0.034 (0.039)			
CREDIT SHIFTER $_t$ x UNDERCAPITALIZED $_{t-1}$ x TFPR $_{t-1}$				-0.299 (0.172)*			
$g(\text{CREDIT}_t)$ x UNDERCAPITALIZED $_{t-1}$						-0.779 (0.345)**	-0.733 (0.342)**
$g(\text{CREDIT}_t)$ x TFPR $_{t-1}$							0.145 (0.196)
$g(\text{CREDIT}_t)$ x UNDERCAPITALIZED $_{t-1}$ x TFPR $_{t-1}$							-1.594 (0.781)*
FIRM CONTROLS	Y	Y	Y	Y	Y	Y	Y
INDUSTRY X YEAR X PROVINCE FE	Y	Y	Y	Y	Y	Y	Y
ADJ. R2	0.021	0.019	0.038	0.038			
OBS.	1611783	1611783	1596403	1596403	1611783	1596403	1596403



Table 10: **Aggregate implications: output and TFP gains from resource reallocation**

This table presents the gains in aggregate output and TFP that accrue from resource reallocation. The reallocation algorithm is described in Section 7. Column (1) and (2) show the percentage of capital and labor reallocated; Column (3) the implied output and productivity gains. Column (4) shows the reallocation gains when resources are reallocated without following a priority rule based on productivity.

	(1)	(2)	(3)	(4)
	CAPITAL REALLOCATED	LABOR REALLOCATED	OUTPUT (TFP) GAIN	OUTPUT (TFP) GAIN
<b>Panel a: Reallocation Within &amp; Between Industries</b>				
1998 - 2001	0.76 %	0.92 %	3.05 %	1.17 %
2002 - 2007	0.67 %	1.11 %	3.25 %	1.19 %
2008 - 2013	0.71 %	0.71 %	3.83 %	1.55 %
AVERAGE	0.71 %	1.14 %	3.38 %	1.30 %
<b>Panel b: Reallocation Within Industries</b>				
1998 - 2001	0.68 %	0.91 %	2.39 %	
2002 - 2007	0.60 %	1.10 %	2.45 %	
2008 - 2013	0.64 %	0.64 %	3.08 %	
AVERAGE	0.64 %	1.13 %	2.64 %	
<b>Panel c: Reallocation Within Industries &amp; Macro Regions</b>				
1998 - 2001	0.67 %	0.91 %	1.45 %	
2002 - 2007	0.60 %	1.10 %	1.58 %	
2008 - 2013	0.64 %	0.64 %	2.24 %	
AVERAGE	0.64 %	1.13 %	1.76 %	

Figure 1: Joint distribution and dispersion of MRP and user costs

This figure investigates the joint distribution of marginal revenue products and User Costs, and their dispersion. We parse the data according to the percentile of the distribution of user costs of capital (panel a) and labor (panel b). The x-axis reports the median value of the user cost and the y-axis reports the median value and interquartile range of the MRP for the group of observations belonging to the same percentile of the distribution of user costs.

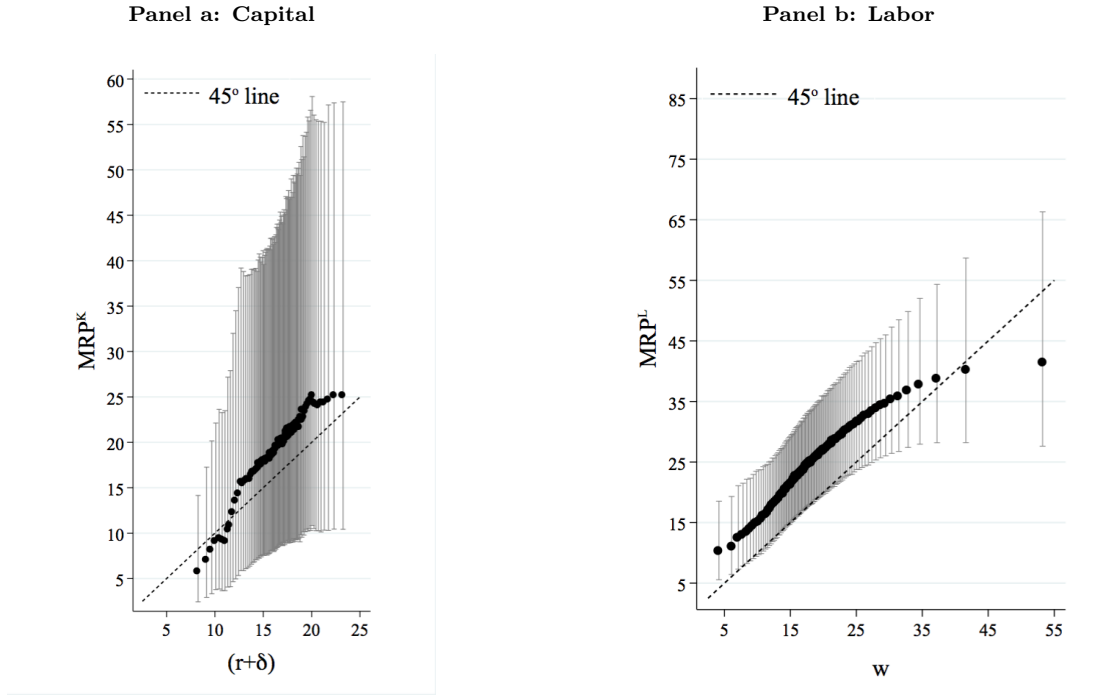
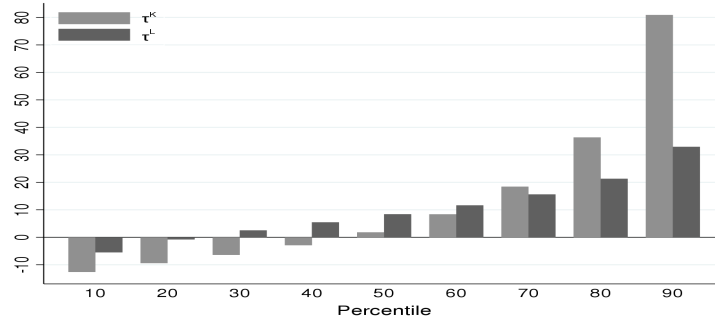


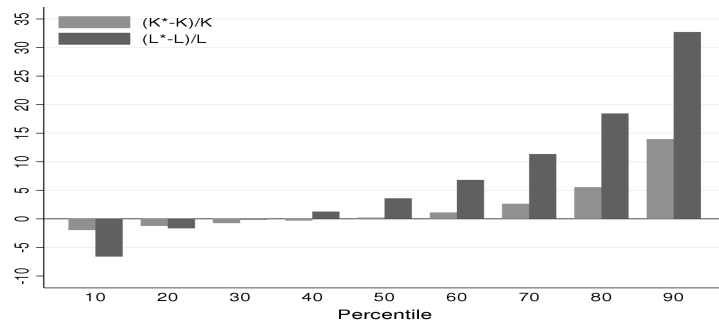
Figure 2: Distribution of percentage deviations from target capital, labor, and output

Panel a presents the distribution of MRP-cost gaps of labor  $\tau_{it}^L$  and capital  $\tau_{it}^K$ . Labor gaps are expressed in thousands of Euros; capital gaps in percentages. Panel b presents the distribution of percentage deviations from targets input demands  $\left(\frac{K_{it}^* - K_{it}}{K_{it}}\right)$  and  $\left(\frac{L_{it}^* - L_{it}}{L_{it}}\right)$  and percentage deviations from output  $\left(\frac{Y_{it}^* - Y_{it}}{Y_{it}}\right)$ , both of which are expressed in percentages. In panel a and b the distribution is asset-weighted. In panel c we present both the unweighted and the asset-weighted distribution.

Panel a: MRP-cost gaps



Panel b: Percentage deviations for target input demands



Panel c: Percentage deviations from target output

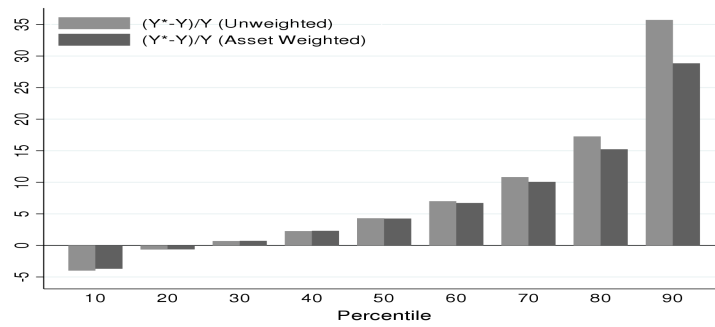


Figure 3: MRP-cost gaps of capital and length of lending relations

This figure displays the relation between MRP-cost gaps of capital ( $\tau_{it}^K$ ) and the length of lending relationships ( $\text{LENGTH RELATION}_{it}^{wmean}$ ). Panel a displays the raw correlation. Panels b, c, and d plot the regression coefficients associated with dummy variables indicating different length of lending relationships (omitted category:  $\text{LENGTH RELATION}_{it}^{wmean} \leq 0.5$  years). The regression model includes firm-level controls and industry by province by year fixed effects. In panel c, undercapitalized firms are those with  $\tau_{it-1}^K > 0$ . In panel d, high-TFPR firms are those with TFPR above the median of the distribution of TFPR ( $\omega_{it}$ ). All quantities on the y-axis are expressed in percentages. The length of relationships is expressed in years.

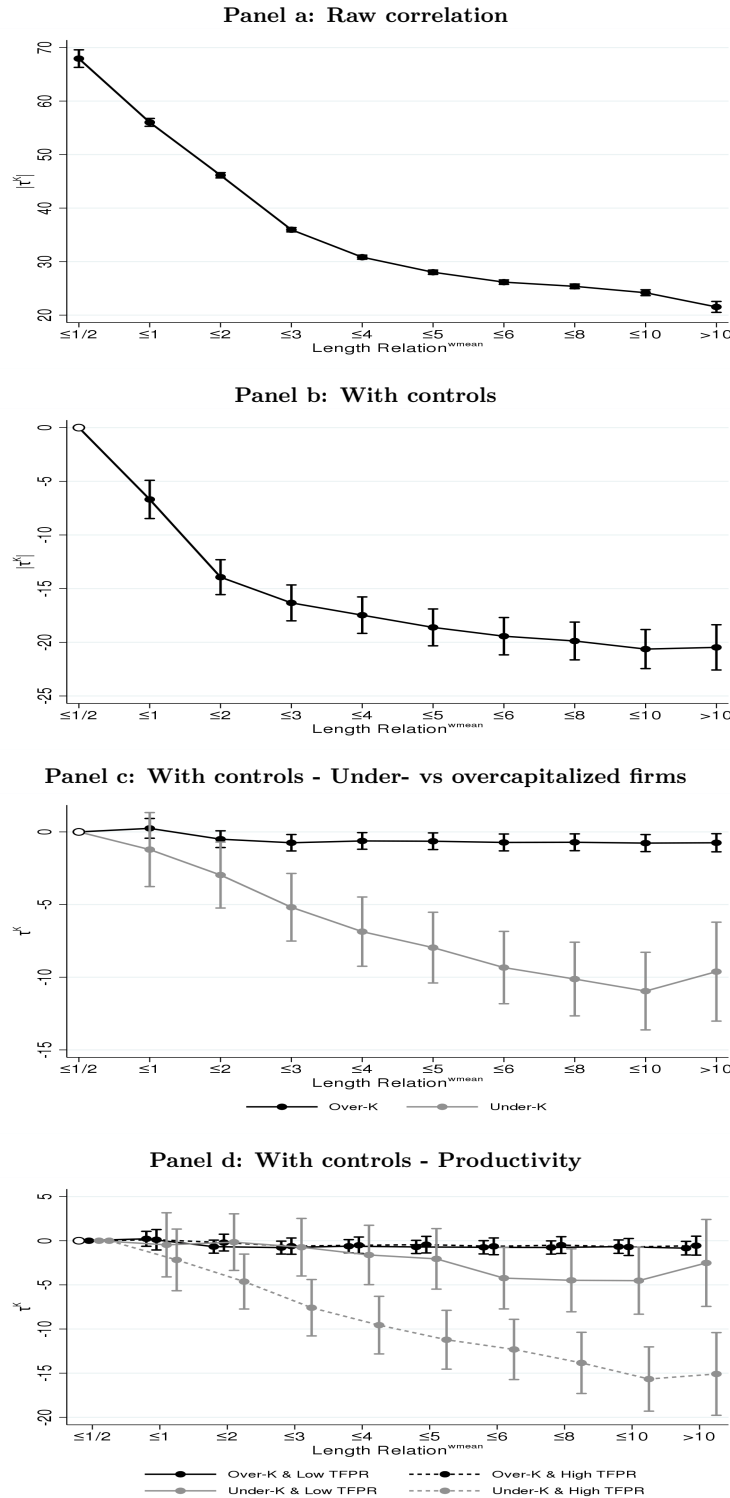


Figure 4: **Relation between growth rates of credit and credit supply shifters**

This figure shows the correlation between the growth rate of credit and the estimated credit supply shifters. A dot in the graph represents the average value of  $g(\text{CREDIT}_{it})$  (y-axis) and the average value of  $\text{CREDIT SHIFTER}_{it}$  (x-axis) across observations that belong to the same percentile of the distribution of  $\text{CREDIT SHIFTER}_{it}$ .

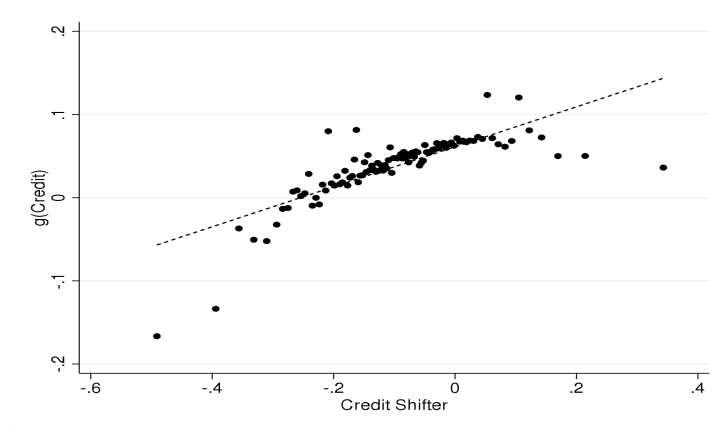
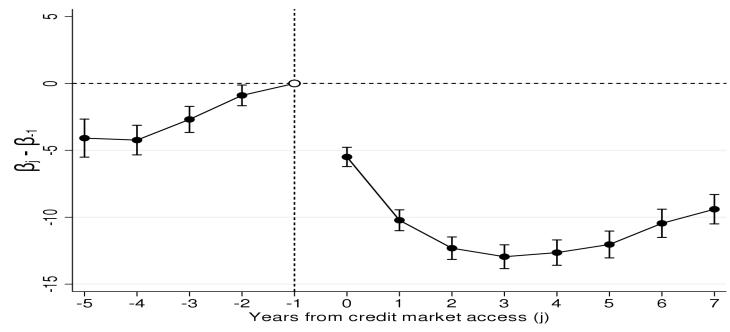


Figure 5: Access to credit markets

This figure displays the dynamic of MRP-cost gaps for capital ( $\tau_{it}^K$ ) (panel a) and percentage deviations from the target capital endowment ( $(K_{it}^* - K_{it})/K_{it}$ ) (panel b) before-to-after transition of a firm into the credit market. The regression model is described by equation (9). All quantities on the y-axis are expressed in percentage points.

Panel a: MRP-cost gaps



Panel b: Percentage deviations from target capital

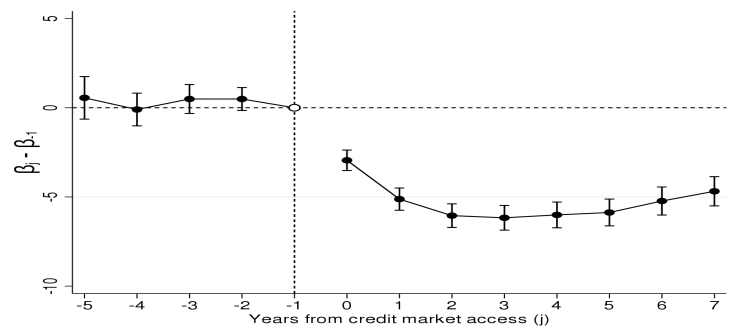
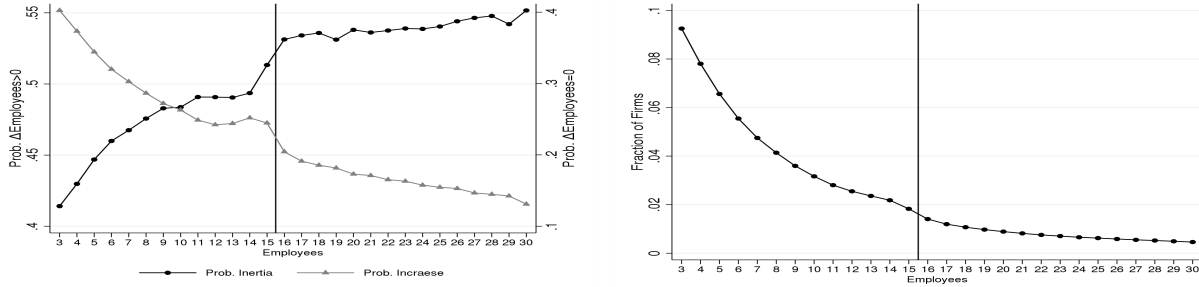


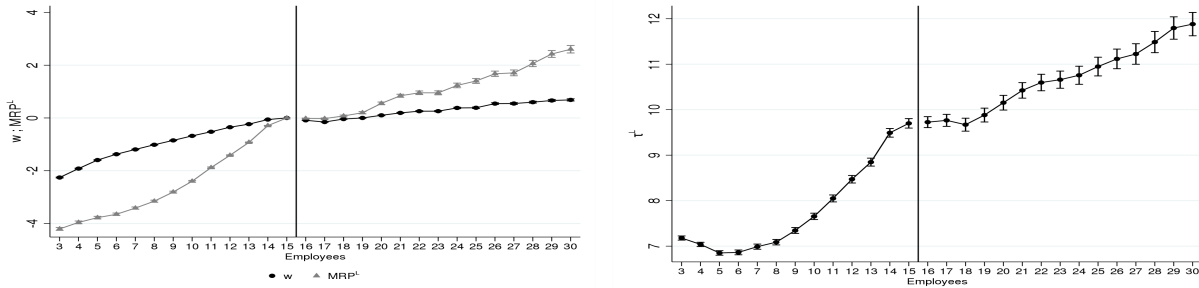
Figure 6: Labor market frictions

This figure studies the impact of size-dependent government-mandated severance payments on firms' employment policies. Panel a reports the probability of employment inertia and probability of upward adjustment across firms of different size (left) and the size distribution (right) as a function of firm size. The probability of inertia is the probability that  $EMPLOYEES_{it} = EMPLOYEES_{it-1}$ ; the probability of upward adjustment is  $EMPLOYEES_{it} > EMPLOYEES_{it-1}$ . Panel b displays the average marginal revenue product of labor ( $MRP_{it}^L$ ) and wages ( $w_{it}$ ) (left) and the average MRP-cost gap of labor  $\tau_{it}^L$  (right) as a function of firm size. Panel c displays the average percentage deviation from target employment  $(L_{it}^* - L_{it})/L_{it}$  as a function of firm size. marginal revenue products, wages, and MRP-cost gaps are expressed in thousands of Euros.

Panel a: Inertia and adjustment probability (left) and size distribution (right)



Panel b:  $MRP^L$  and wages (left) and MRP-cost gaps (right)



Panel c: Employment policy distortions

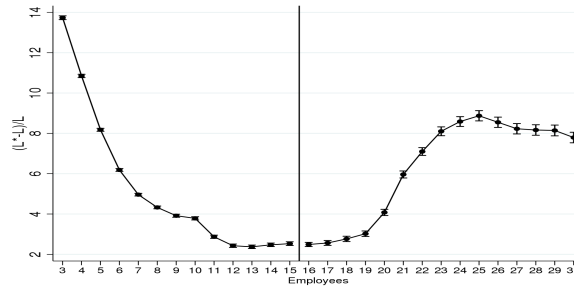


Figure 7: MRP-cost gaps and firm policies

This figure provides a graphical representation of gaps ( $\tau_{it}^K$  and  $\tau_{it}^L$ ) and their relation to target input demands ( $K_{it}^*$  and  $L_{it}^*$ ).

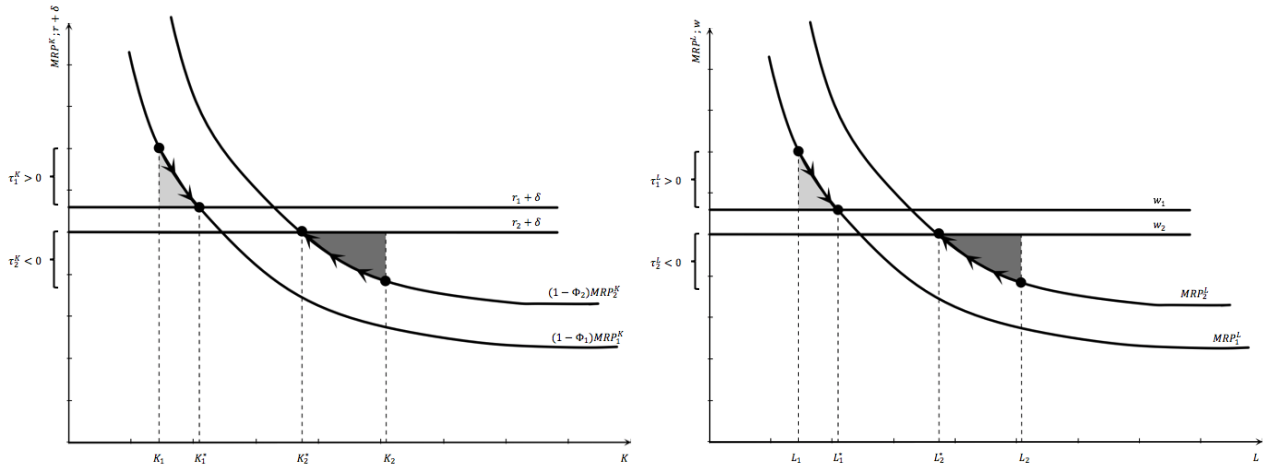




Figure 8: Percentage Deviations from target capital and length of lending relations

This figure displays the relation between percentage deviations from target capital ( $(K_{it}^* - K_{it})/K_{it}$ ) and the length of lending relationships ( $\text{LENGTH RELATION}_{it}^{\text{wmean}}$ ). Panel a displays the raw correlation. Panels b, c, and d plot the regression coefficients associated with dummy variables indicating a different length of lending relationships (omitted category:  $\text{LENGTH RELATION}_{it}^{\text{wmean}} \leq 0.5$  years). The regression model includes firm-level controls and industry by province by year fixed effects. In panel c, undercapitalized firms are those with  $\tau_{it-1}^K > 0$ . In panel d, high-TFPR firms are those with TFPR above the median of the distribution of TFPR ( $\omega_{it}$ ). All quantities on the y-axis are expressed in percentages. The length of relationships is expressed in years.

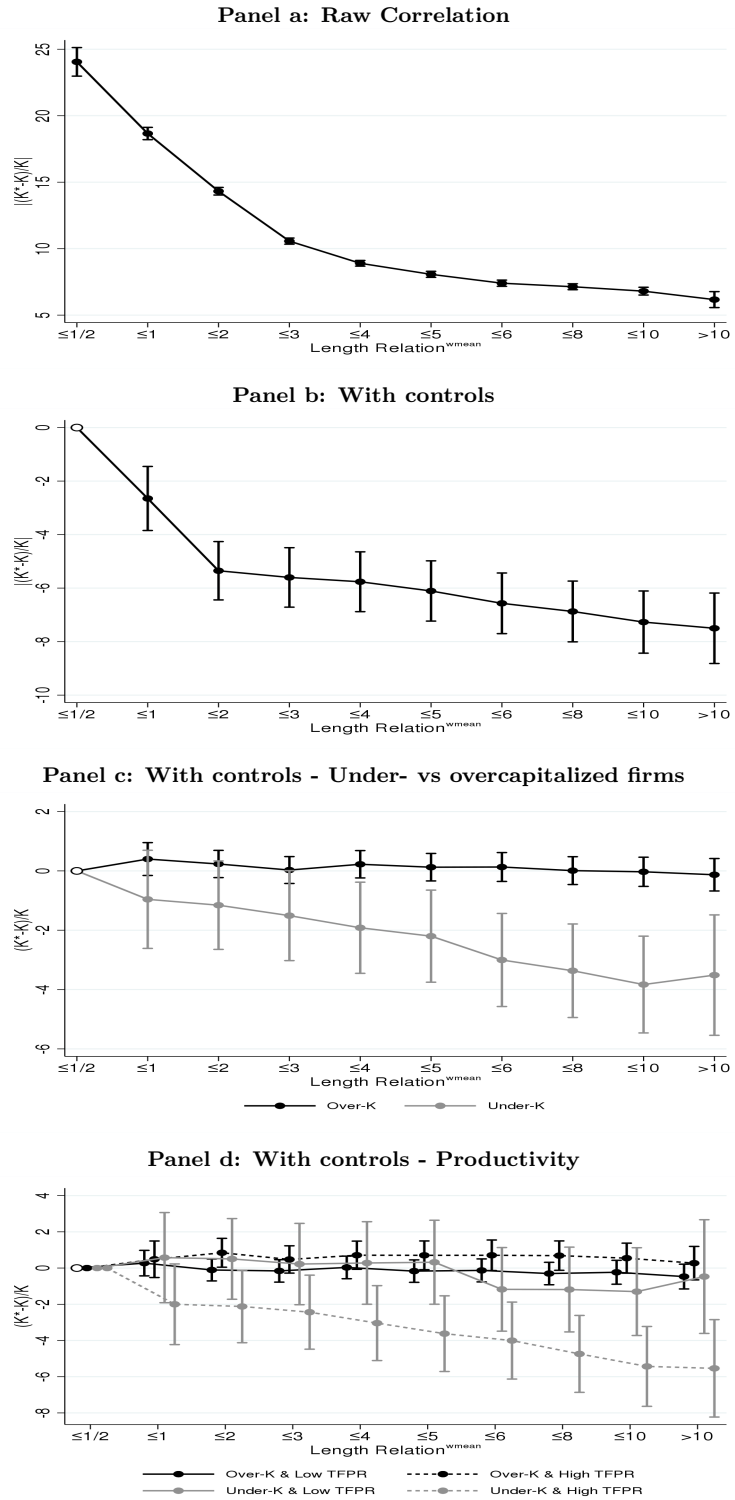
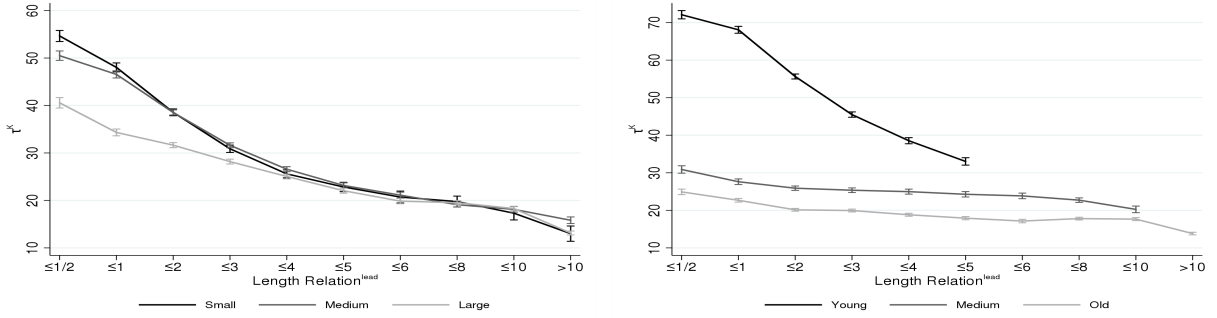


Figure 9: Age-size heterogeneity

This figure shows the average MRP-cost gap for capital ( $\tau_{it}^K$ ) (panel a), and average deviation from the target capital endowment ( $(K_{it}^* - K_{it})/K_{it}$ ) (panel b) across the distribution of  $\text{LENGTH RELATION}_{it}^{wmean}$  (x-axis) by age and size groups. All quantities on the y-axis are expressed in percentages. The length of relationships is expressed in years.

Panel a: MRP-cost gaps



Panel b: Percentage deviations from target capital

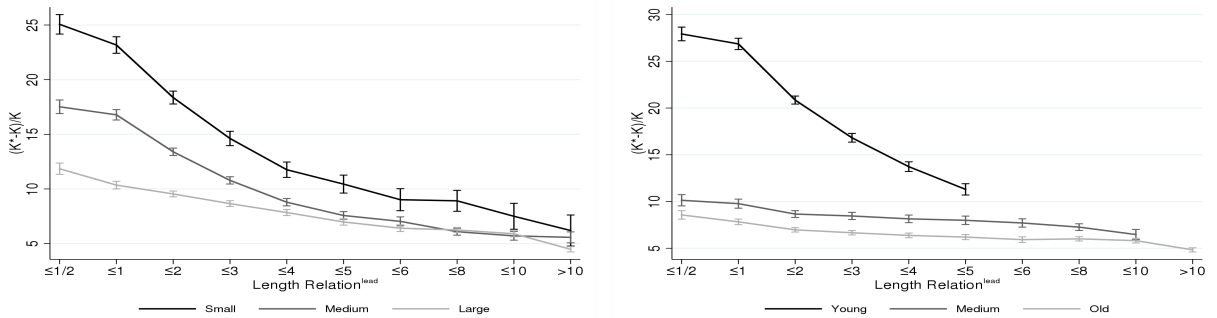


Figure 10: Aggregate output gain from resource reallocation

This figure presents the gains in aggregate output and TFP that accrue from resource reallocation (solid line) and the relative contribution of reallocation to total output gains achievable if all gaps were closed (dotted line). The relative contribution of misallocation is calculated as the ratio of the output gains from reallocation over the total gains that would accrue if all firms could adjust their capital endowment and workforce to completely close the observed MRP-cost gaps.

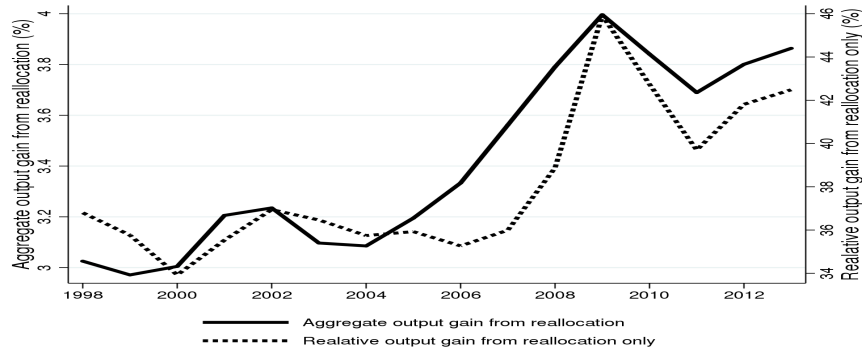
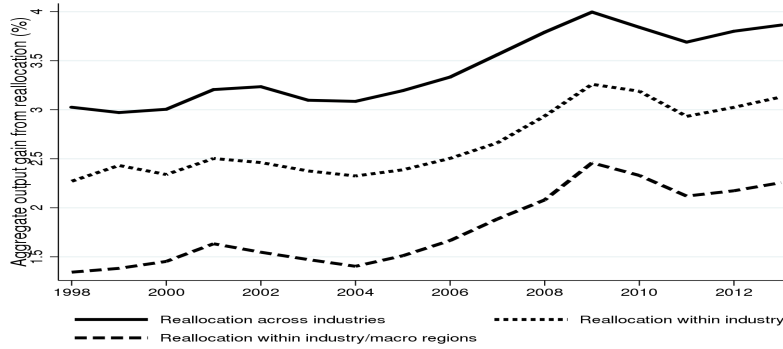


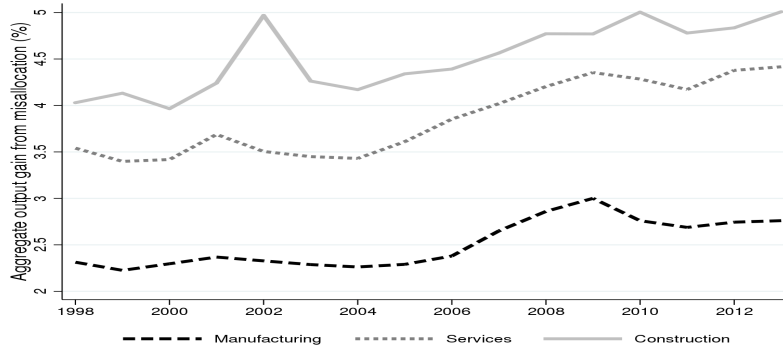
Figure 11: Aggregate Implications: Output and TFP Gains from Resource Allocation

This figure explores the extent of misallocation within and between industries, and across different geographical regions in Italy. Panel a presents the gains from reallocation allowing capital and labor reallocation both between and within industries (dotted line), only within 4-digits code industries (solid line), and only within 4-digits code industries/macro-regions (dashed line). Panel b presents the gains from reallocation (within 4-digits code industries) separately for each macro-industry (manufacturing, services, and construction). Panel c presents the gains from reallocation (within 4-digits code industries) separately for each macro-region (north, south, and center of Italy).

Panel a: Within versus between reallocations



Panel b: Sectoral heterogeneity



Panel c: Spacial heterogeneity

