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Essays on Labor, Product, and Credit Market Imperfections

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Abstract

Market frictions or market imperfections are diverse, broadly present in most markets, and affect most transactions in the economy. These market failures may prevent buyers and sellers from trading, even if they agree on a price. This means that the central assumption of perfectly competitive markets that markets clear fails to hold, and some buyers and sellers remain unmatched.

Since the 1970s, a growing literature has emerged addressing the importance of market frictions in most markets of the economy, namely in the labor, product, and credit markets. Information asymmetries, transaction costs, heterogeneous preferences, and coordination failure are examples of sources of market imperfections. In labor markets, these frictions imply that firms possess some market power over their employees and that a one cent wage cut does not lead all workers to leave the firm. In product markets, a key ingredient for the sluggish price adjustment is coordination failure among firms. Firms respond incompletely to an aggregate shock because other firms have not yet responded. In turn, asymmetric information and costly contract enforcement provide the foundations of credit market frictions, and are used to explain credit rationing as a market equilibrium.

In Chapter 1 we use matched employer-employee data and firm balance sheet data to investigate the importance of firm productivity and firm labor market power in explaining firm heterogeneity in wage formation. We use a linear regression model with one interacted high dimensional fixed effect to estimate 5-digit sector-specific elasticity of output with respect to input factors directly from the production function. This allows us to derive firm specific price-cost mark-up and elasticity of labor supply. The results show that firms possess a considerable degree of product and labor market power. Furthermore, we find evidence that a firm's monopsony power negatively affects the earnings of its workers, and firm's total factor productivity is closely associated with higher earnings, *ceteris paribus*. We also find that firms use monopsony power for wage differentiation between male and female workers.

Chapter 2 describes price setting behavior using a very rich dataset of producer prices collected for Portuguese firms. The Industrial Producer Prices

Index dataset is comprised of monthly transaction prices collected for products defined at a detailed level. We proceed with the analysis in two steps. First, we estimate a hazard function model for the probability of a price change with high dimensional fixed effects to extensively account for product- and firm-specific time-invariant heterogeneity, splitting price changes between price decreases and price increases. Second, we estimate a peer-effects model to document how market competition affects firms' price setting rules. The results suggest that the likelihood of price adjustment depends on both idiosyncratic and sectoral conditions. Furthermore, when we fully account for heterogeneity, duration dependence is estimated to be positive in the case of both a price increase and price decrease. The results of the peer-effects model suggest that firms timidly respond to their competitors' price setting behavior.

Chapter 3 examines the importance of credit demand and credit supply-related factors in explaining the evolution of credit granted to Portuguese small and medium-sized enterprises (SMEs). The results suggest that the interest rate is a strong driver of SMEs' demand for bank loans, as well as their internal financing capacity. On the other hand, credit supply mostly depends on firms' ability to generate cash-flows and reimburse their debt, and on the amount of collateral. The model was estimated for the period between 2010 and 2012. The results suggest that a considerable fraction of Portuguese SMEs were affected by credit rationing in this period.

Keywords: labor market frictions, wage setting policy, high dimensional fixed effects, gelbach decomposition, price rigidity, coordination failure, duration model, high dimensional fixed effects, peer effects, credit market frictions, bank lending, financial crisis, disequilibrium model, SMEs

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Contents

Introduction

Market frictions or market imperfections are diverse, broadly present in most markets, and affect most transactions in the economy. These market failures may prevent buyers and sellers from trading, even if they agree on a price. This means that the central assumption of perfectly competitive markets that markets clear fails to hold, and some buyers and sellers remain unmatched in the market.

Until the 1970s, economists while recognizing the importance of market frictions in real-world transactions did not incorporate market frictions in formal economic models. Since then, a growing literature has emerged addressing the importance of market frictions in most markets of the economy, namely in the labor, product, and credit markets. Imperfect information about available jobs, moving and learning costs, firm specific human capital, reputation costs, exploitation of rents, and worker heterogeneous preferences are examples of sources of frictions in the labor market. These market imperfections may generate upward sloping labor supply curves to a particular firm. This is in line with the "new monopsony" literature popularized by Manning (2003), in which firms possess some degree of market power and are able to mark down their wages below the marginal revenue product. This means that the assumption of a single market wage that would cause all employees to instantaneously leave the firms after a one cent wage cut seems unrealistic. Moreover, wages of workers with similar characteristics and working for similar firms show considerable dispersion.

In business cycle models, frictions are used to explain the non-neutrality of monetary shocks and the amplifying effects of demand shocks on output. A central hypothesis for these findings has been the sluggish price adjustment to aggregate conditions. In this new class of macroeconomic models, an important source of market incompleteness is coordination failure that leads firms to adjust their prices incompletely because other firms have not responded yet to aggregate shocks. The combination of sticky prices and coordination failure in product markets may amplify and generate long-lasting effects of demand shocks on output.

In credit markets, the seminal work of Stiglitz & Weiss (1981) show that in the presence of credit market frictions, banks may decide not to grant credit to borrowers with a high probability of default, even if they are willing to pay the respective risk premium. Furthermore, information asymmetries, moral hazard, and adverse selection problems may be so severe for some firms that credit rationing may arise as an equilibrium.

The importance of frictions in driving market outcomes is crucial to understand these markets and promoted a large number of empirical contributions. The increasing availability of very detailed micro data allows researchers to study most of these relations. It is the combination of economic theory and empirical testing that renders economics fruitful. Economic theory promotes new research topics and leads to the construction of new databases, while empirical results can challenge economic theory, motivating new theories. This microeconomic evidence is of great importance for drawing macroeconomic implications, making it of interest for policymakers.

This thesis contributes to the literature on labor, product, and credit market frictions.

The first chapter was written in co-authorship with Pedro Portugal and studies the importance of labor market frictions and total factor productivity to explain earnings of individuals. We depart from the central assumption of perfectly competitive labor markets that markets clear, and show that firms possess a considerable degree of labor market power, which is in contradiction with the general assumption made by economists until the 1950s that labor power is given and not augmentable. We use the high dimensional fixed effects estimation procedure proposed by Portugal & Guimarães (2010) to estimate an empirical distribution of the firm-specific product and labor markets imperfections parameters as directly estimated from the production function, following the theoretical framework developed by Dobbelaere & Mairesse (2011). We proceed by estimating what must be one of the most common regression equations in microeconomics, a Mincer wage equation augmented with the estimated labor market imperfections parameters and firm-specific total factor productivity. Furthermore, we apply the Gelbach's methodology (Gelbach (2016)) to decompose the impact of the estimated firm's wage setting power and firm's total factor productivity within the firm, worker, and job title dimensions. The results suggest that the firm elasticity of labor supply is positively and significantly related to wages, meaning that firms with more monopsonistic power pay lower wages, *ceteris paribus*. We also show that firm's total factor productivity helps considerably to explain heterogeneity in wage formation.

The second chapter describes price setting behavior using a very rich micro dataset of producer prices collected for Portuguese firms. Macroeconomists have been studying for a long time the large effects of demand shocks on output and, in particular, the non-neutrality of the monetary policy on real output. A growing empirical literature has emerged in the last decades documenting the large effects of demand shocks on output (Blanchard & Perotti (2002), Ramey (2009), and Nakamura & Steinsson (2014)) and the non-neutrality of monetary shocks (Friedman & Schwartz (2008), Christiano *et al.* (1999), Romer & Romer (2004)). A central hypothesis for explaining these findings has been the sluggish price adjustment to aggregate conditions. We proceed with the analysis in two steps. First, we estimate a hazard function model for the probability of a price change with high dimensional fixed effects to extensively account for product- and firm-specific time-invariant heterogeneity, splitting price changes between price decreases and price increases. Second, we estimate a peer-effects model to document how market competition affects a firm's price setting rules. The results suggest that the likelihood of price adjustment depends on both idiosyncratic and sectoral conditions. Furthermore, when we fully account for heterogeneity, duration dependence is estimated to be positive in the case of both a price increase and price decrease. The results of the peer-effects model suggest that firms timidly respond to their competitors' price setting behavior.

The recent global financial crisis and the subsequent European sovereign debt crisis led to a surge of studies exploiting the effects of financial market frictions on real economic activity. One of the strands of this literature focuses the bank lending channel, specifically whether shocks to the financial position of a bank affect lending supply. In Portugal, credit granted to firms tightened considerably since the onset of the economic and financial crisis. Portuguese firms were affected by the significant contraction of the economy and worse economic prospects, and banks were severely affected by internacional financing restrictions and stronger capital requirements. In this context, the reduction in lending was a result of increased restrictiveness in credit standards and conditions applied to loans as well as of decreased demand by firms. The seminal work of Stiglitz & Weiss (1981) presents a theoretical framework for credit rationing as arising from adverse selection and moral hazard problems. In fact, these market failures can be so severe that creditworthy small and medium-sized enterprises (SMEs) may not have access to credit mainly due to their reduced capacity to provide collateral (Berger & Udell (2006) and Beck *et al.* (2010)). The third and fourth chapters of this thesis are concerned with credit rationing for Portuguese SMEs since the onset of the economic and financial crises.

The third chapter of this thesis, written in co-authorship with Luísa Farinha, is an evaluation of the relative contribution of demand and supply to explain credit developments in Portugal since the onset of the economic and financial crises. We consider a disequilibrium model based on microeconomic data to identify credit restrictions for Portuguese SMEs, as these are expected to rely more on bank loans and are, therefore, more vulnerable to credit rationing. The model assumes that the credit market may be in disequilibrium, which means that the observed interest rate does not ensure that credit demand is equal to credit supply. The results suggest that approximately 15 percent of Portuguese SMEs were partially rationed while 32 percent of SMEs with no bank loans were fully rationed during the period between 2010 and 2012. These results are of great importance for policy makers due to the economic implications of the shortage of credit (see for example Bentolila *et al.* (2015)).

Chapter 1

Labor market imperfections and the firm's wage setting policy

1.1 Introduction

A central feature of perfectly competitive markets is that markets clear, meaning that all workers with similar quality should be paid the same market clearing wage. The assumption of a single market wage that would cause all employees to instantaneously leave the firm after a one cent wage cut seems unrealistic. Recent empirical evidence suggests the presence of considerable wage dispersion among workers with similar characteristics and among similar firms. Torres *et al.* (2012) use a longitudinal matched employer-employee dataset for Portugal to estimate a wage equation with three high-dimensional fixed effects and decompose the variation in real hourly wages into three different components related to worker, firm, and job title heterogeneity. The authors find that worker permanent heterogeneity accounts for about 36 percent of wage variation, firm permanent effects account for almost 29 percent, and job title effects are less important, although still explaining almost 10 percent of wage variation.

The firm effects estimated in wage regressions can be thought of as arising from distortions in the labor markets (Abowd *et al.* (1999b) and Goux & Maurin (1999)). Search frictions in the labor market such as imperfect information on alternative available jobs (Burdett & Mortensen (1998) and Shimer (2005)), moving and learning costs (Boal & Ransom (1997)), firm specific human capital, reputation costs, exploitation of rents, and worker heterogenous preferences namely over nonwage job characteristics (Stevens (1994) and Bhaskar *et al.* (2002)) are sources of labor market power, and help to explain why firms have market power and why the labor supply curve faced by an individual firm is not perfectly elastic. These search frictions in the labor market may generate upward sloping labor supply curves to a particular firm. In a standard wage setting model this means that firms possess some power to mark down their wages below the marginal revenue product. This is in line with the "new monopsony" literature popularized by Manning (2003), in which employers gain some market power derived from search frictions when setting wages. Monopsony is not understood in the traditional sense of a unique employer in the labor market, but instead as synonymous with imperfect competition, monopsonistic competition, upward sloping labor supply curve to the firm, or finite labor supply elasticity. A particular firm may face an upward labor supply curve even if there is no concentration on the demand side of the market.

Recent empirical literature provides robust evidence consistent with the existence of monopsony power and upward-sloping labor supply curves to individual firms. There is considerable heterogeneity in the estimated labor supply elasticity. Ransom & Oaxaca (2010), Hirsch *et al.* (2010), and Weber (2013) estimate the labor supply elasticity to range between 1 to 10. Dobbelaere & Mairesse (2011) estimate a production function for 38 French manufacturing industries and derive product and labor market imperfection parameters as a wedge between the factor elasticities and their corresponding shares in revenue. Then, the authors classify each industry into six different regimes according to the type of competition in the product and the labor market (perfect vs. imperfect competition in the product market and efficient bargaining vs. right to manage or perfect competition vs. monopsonistic competition in the labor market). Their analysis of the within-regime firm heterogeneity through the Swamy methodology suggests considerable dispersion in the estimated price-cost mark-up and rent-sharing or labor supply elasticity parameters. Depew & Sørensen (2013) use employee records from Ford Motors in Michigan and A. M.

Byers in Pennsylvania and find that the workers' labor supply elasticity to a firm is counter-cyclical so that monopsony power is pro-cyclical. The estimates of the labor supply elasticity to the firm are typically between 4 during expansions and 1.6 during recessions.

The primary contribution of this study is twofold: first, we estimate a measure of labor supply elasticity to the firm directly from the production function and at a very granular level (by estimating a standard production function using the one-iterative high dimensional estimation procedure and considering the 5-digit sector variable as interaction variable), which allows us to account for the heterogeneity across and within labor markets in the analysis; second, and to the best of our knowledge, this is the first study that disentangles the importance of firm's wage setting power and firm's total factor productivity to explain the firm's wage setting policies.

In this study we use employer-employee matched data and firm balance sheet data to obtain an empirical distribution of the firm-specific product and labor market imperfection parameters as directly estimated from the production function, following the theoretical framework developed by Dobbelaere (2004) and Dobbelaere & Mairesse (2011) and using the high dimensional fixed effects estimation procedure proposed by Portugal & Guimarães (2010). We proceed by estimating the impact of monopsony power on the firm's wage setting policy by plugging the estimated labor supply elasticity and the firm total factor productivity in a Mincer wage equation. We also estimate the importance of rent-sharing to explain wage formation within the efficient bargaining setting. Furthermore, we use the Gelbach's methodology (Gelbach (2016)) to decompose the impact of the estimated labor supply elasticity on wages within the firm, worker, and job title dimensions.

In fact, little empirical literature can be found on the effects of monopsonistic competition on earnings of individuals. Weber (2013) estimates firm-level labor supply elasticities for the U.S. labor market through an extension of the dynamic model of labor supply proposed by Manning (2003) and examines the effects of monopsonistic competition on the earnings distribution. The author provides evidence of substantial heterogeneity in the market power possessed by firms and shows a positive relationship between the firm's labor supply elasticity and the wages of its workers. The author estimates that the impact of a one unit increase in a firm's labor supply elasticity is associated with an increase in earnings that ranges from 5 to 16 percent.

We find strong evidence that the firm elasticity of labor supply is positively and significantly related to wages, meaning that firms with more monopsonistic power pay on average lower wages, *ceteris paribus*. We also find that the elasticity of labor supply to the firm affects wages differently according to the gender of workers, which reveals the importance of considering the firm's labor market power when studying the wage pay gap between women and men. In turn, and surprisingly, our results suggest that firms with higher relative extent of rent-sharing pay lower wages. We also show that firm productivity helps considerably to explain heterogeneity in wage formation in both labor market settings.

The paper proceeds as follows. In the next section we briefly present the theoretical framework. This is followed by a discussion of the data used in the empirical analysis and the estimation procedure. Sections 1.4 and 1.5 report the results on the estimation of the labor and product market imperfection parameters and firm's total factor productivity, respectively. Section 1.6 presents the wage regressions. Section 1.7 discusses the results and Section 1.8 concludes.

1.2 Theoretical framework

We closely follow Dobbelaere & Mairesse (2011) to jointly estimate product and labor market imperfections as a wedge between factor elasticities for labor and materials in the production function and their corresponding shares in revenue. This approach extends the framework of Hall (1988) abstaining from the assumption of perfect competition in the labor market and builds on the estimation of the firm price-cost mark-up and rent-sharing parameters directly from the production function. The analysis relies crucially on the assumption that output elasticities of labor and materials are equal to their revenue shares when prices equal the marginal cost of production.

We consider a production function $\mathcal{Q}_{ft} = \Theta_{ft}F(N_{ft}, M_{ft}, K_{ft})$, where $\mathcal{Q}_{ft}$ represents physical output of firm f in period t, and F(.) is a function of labor N_{ft}, materials M_{ft}, and capital K_{ft}. The term $\Theta_{ft} = A\exp(\eta_f + \nu_t + u_{ft})$ is the Hicksian neutral shift of firm f in period t, η_f is an unobserved firm-specific time-invariant effect, ν_t is a set of time effects, and u_{ft} is a firm-year idiosyncratic disturbance term with the conventional properties.

Taking natural logarithms on both sides of the production function and denoting q_{ft}, l_{ft}, m_{ft}, k_{ft}, and θ_{ft} the logarithm of $\mathcal{Q}_{ft}$, N_{ft}, M_{ft}, K_{ft}, and Θ_{ft}, respectively, results in a linear production function

$$q_{ft} = (\varepsilon_N)_{ft}n_{ft} + (\varepsilon_M)_{ft}m_{ft} + (\varepsilon_K)_{ft}k_{ft} + \theta_{ft}, \tag{1.1}$$

where ε_J (J=N, M, K) is the factor-cost elasticity of output with respect to input factor J.

1.2.1 Perfect competition in the product and labor market

In perfectly competitive labor and product markets, in which firms are price-takers in the product and input factor markets, short-run profit maximization implies that:

$$(\varepsilon_N)_{ft} = \frac{w_{ft}N_{ft}}{P_{ft}\mathcal{Q}_{ft}} \equiv (\alpha_N)_{ft}$$
$$(\varepsilon_M)_{ft} = \frac{j_{ft}M_{ft}}{P_{ft}\mathcal{Q}_{ft}} \equiv (\alpha_M)_{ft},$$

where w_{ft} and j_{ft} represent labor and material factor prices, respectively, and P_{ft} is the price of output. Therefore, $(\alpha_N)_{ft}$ and $(\alpha_M)_{ft}$ are the firm shares of labor and material costs in total revenue, respectively.

Assuming that the elasticity of scale is known (λ), the elasticity of capital can be written as:

$$(\varepsilon_K)_{ft} = \lambda_{ft} - (\alpha_N)_{ft} - (\alpha_M)_{ft}. \tag{1.2}$$

Then, combining equation (1.2) with equation (1.1) yields:

$$q_{ft} - k_{ft} = (\alpha_N)_{ft}[n_{ft} - k_{ft}] + (\alpha_M)_{ft}[m_{ft} - k_{ft}] + [\lambda_{ft} - 1]k_{ft} + \theta_{ft}. \tag{1.3}$$

1.2.2 Imperfect competition in the product market

Perfectly competitive labor market

In turn, if firms act as price-setters in the product market but price-takers in the input factor markets, profit maximization leads to:

$$(\varepsilon_N)_{ft} = \mu_{ft}(\alpha_N)_{ft} \tag{1.4}$$

$$(\varepsilon_M)_{ft} = \mu_{ft}(\alpha_M)_{ft}, \tag{1.5}$$

where $\mu_{ft} = \frac{P_{ft}}{(C_Q)_{ft}} > 1$ refers to the mark-up of price (P) over marginal cost (C_Q). In this setting, the capital-output elasticity can be written as:

$$(\varepsilon_K)_{ft} = \lambda_{ft} - \mu_{ft}(\alpha_N)_{ft} - \mu_{ft}(\alpha_M)_{ft}, \tag{1.6}$$

and equation (1.3) can be rewritten as:

$$q_{ft} - k_{ft} = \mu_{ft}\big[(\alpha_N)_{ft}[n_{ft} - k_{ft}] + (\alpha_M)_{ft}[m_{ft} - k_{ft}]\big] + [\lambda_{ft} - 1]k_{ft} + \theta_{ft}. \tag{1.7}$$

Therefore, the mark-up of price over marginal cost can be estimated using the previous equation[1].

[1] In the right-to-manage bargaining framework, the firm can bargain with risk-neutral workers over wages but retains the right to set employment afterwards unilaterally. Since the firm uniquely sets the amount of labor and material inputs to contract, this is equivalent to perfect competition in the labor market and equation (1.7) still provides an estimate of the mark-up μ_{ft}.

Efficient bargaining

In this setting, risk-neutral workers and the firm bargain over wages and employment. Workers maximize $U(w_{ft}, Nft) = N_{ft}(w_{ft} - \bar{w}_{ft})$ where $\bar{w}_{ft} < w_{ft}$ is the reservation wage. Given that capital is assumed to be quasi-fixed, the firm wants to maximize short-run profits $\Pi_{ft} = R_{ft} - w_{ft}N_{ft} - j_{ft}M_{ft}$, where $R_{ft} = P_{ft}\mathcal{Q}_{ft}$ is the firm total revenue. The outcome of the bargaining is the generalized solution to the following maximization problem:

$$\max_{w_{ft}, N_{ft}, M_{ft}} \{N_{ft}(w_{ft} - \bar{w}_{ft})\}^{\phi_{ft}}\{R_{ft} - w_{ft}N_{ft} - j_{ft}M_{ft}\}^{1-\phi_{ft}}. \tag{1.8}$$

The first-order condition for material input is given by equation (1.5) because material input is unilaterally determined by the firm. Setting the relative extent of rent-sharing equal to $\gamma_{ft} = \phi_{ft}/(1-\phi_{ft})$ and denoting the marginal revenue product of labor as $(R_N)_{ft}$, the first-order conditions with respect to wage and labor are, respectively,

$$w_{ft} = \bar{w}_{ft} + \gamma_{ft}\left[\frac{R_{ft} - w_{ft}N_{ft} - j_{ft}M_{ft}}{N_{ft}}\right] \tag{1.9}$$

$$w_{ft} = (R_N)_{ft} + \phi_{ft}\left[\frac{R_{ft} - (R_N)_{ft}N_{ft} - j_{ft}M_{ft}}{N_{ft}}\right], \tag{1.10}$$

which yield the following contract curve:

$$(R_N)_{ft} = \bar{w}_{ft}. \tag{1.11}$$

Given that in equilibrium $\mu_{ft} = \frac{P_{ft}}{(R_Q)_{ft}}$, where $(R_Q)_{ft}$ is the marginal revenue, the marginal revenue of labor can be expressed as the product of marginal revenue and marginal product of labor, $(\mathcal{Q}_N)_{ft}$:

$$(R_N)_{ft} = (R_Q)_{ft}(Q_N)_{ft} = (R_Q)_{ft}(\varepsilon_N)_{ft}\frac{\mathcal{Q}_{ft}}{N_{ft}} = \frac{P_{ft}(\mathcal{Q}_N)_{ft}}{\mu_{ft}}. \tag{1.12}$$

Combining equation (1.11) with equation (1.12) yields:

$$(\varepsilon_N)_{ft} = \mu_{ft}\left(\frac{\bar{w}_{ft}N_{ft}}{P_{ft}\mathcal{Q}_{ft}}\right) = \mu_{ft}(\bar{\alpha}_N)_{ft}, \tag{1.13}$$

which is equivalent to:

$$(\varepsilon_N)_{ft} = \mu_{ft}(\alpha_N)_{ft} - \mu_{ft}\gamma_{ft}[1 - (\alpha_N)_{ft} - (\alpha_M)_{ft}]. \tag{1.14}$$

The previous equation shows that employment does not directly depend on the bargained wage. The elasticity of capital is given by:

$$(\varepsilon_K)_{it} = \lambda - \mu_{it}(\alpha_N)_{it} + \mu_{it}\gamma_{it}[1 - (\alpha_N)_{it} - (\alpha_M)_{it}] - \mu_{it}(\alpha_M)_{it}, \tag{1.15}$$

and the corresponding modified production function can be expressed as:

$$q_{ft} - k_{ft} = (\varepsilon_N)_{ft}[n_{ft} - k_{ft}] + (\varepsilon_M)_{ft}[m_{ft} - k_{ft}] + [\lambda_{ft} - 1]k_{ft} + \theta_{ft}. \tag{1.16}$$

This equation allows the identification of the mark-up of price over marginal cost as well as the labor market imperfections parameter as measured by the extent of rent-sharing parameter.

The authors derive a joint market imperfections parameter (ψ) through the comparison between the factor elasticities as directly estimated from the production function and the factor shares for labor and materials. The sign and significance of this parameter characterize the type of competition in the product and labor markets:

$$\psi_{ft} \equiv \frac{(\varepsilon_M)_{ft}}{(\alpha_M)_{ft}} - \frac{(\varepsilon_N)_{ft}}{(\alpha_N)_{ft}}. \tag{1.17}$$

In the efficient bargaining setting:

$$\psi_{ft} \equiv \frac{(\varepsilon_M)_{ft}}{(\alpha_M)_{ft}} - \frac{(\varepsilon_N)_{ft}}{(\alpha_N)_{ft}} = \mu_{ft}\gamma_{ft}\left[\frac{1 - (\alpha_N)_{ft} - (\alpha_M)_{ft}}{(\alpha_N)_{ft}}\right]. \tag{1.18}$$

If ψ is positive, then an efficient bargaining model prevails and we can derive estimates for the price-cost mark-up and the (absolute and relative) extent of rent-sharing parameters. In this case the worker obtains a wage higher than her marginal revenue and, therefore, the ratio between the output elasticity of labor and the share of labor costs in revenue becomes smaller than the respective ratio for materials.

1.2.3 Monopsony

In this study we analyze the importance of firm labor market power in explaining the wage setting policy followed by firms, and we therefore focus our analysis on the monopsony regime. The theoretical framework for the monopsony model can be described as follows. Consider a firm that operates under imperfect competition in the product market and faces a labor supply $N_{ft}(w_{ft})$. N_{ft}, the labor supply curve of the individual firm, is an increasing function of the wage, w_{ft}. The short-run profit function for the monopsonist firm taking the labor supply as given is:

$$\max_{N_{ft},M_{ft}} \Pi(w_{ft}, N_{ft}, M_{ft}) = R_{ft}(N_{ft}, M_{ft}) - w_{ft}(N_{ft})N_{ft} - j_{ft}M_{ft}. \qquad (1.19)$$

The first-order condition with respect to the material input leads to $(R_M)_{ft} = j_{ft}$ with the marginal revenue of materials $(R_M)_{ft}$ equal to the price of materials j_{ft} (equation (1.5)). The first-order condition with respect to the labor input is given by:

$$w_{ft} = \beta_{ft}(R_N)_{ft}, \beta_{ft} \equiv \frac{(\varepsilon_w)_{ft}}{1 + (\varepsilon_w)_{ft}}, \qquad (1.20)$$

where $(\varepsilon_w)_{ft} \in \mathcal{R}_+$ is the wage elasticity of the labor supply. The firm's degree of monopsony power can be measured by $\frac{(R_N)_{ft}}{w_{ft}}$, and therefore the more inelastic the labor supply, the larger the gap between the marginal revenue of labor and the wage. This means that the monopsony power depends negatively on the elasticity of the labor supply.

Equation (1.20) can be rewritten as:

$$(\varepsilon_N)_{ft} = \mu_{ft}(\alpha_N)_{ft}\left(1 + \frac{1}{(\varepsilon_w)_{ft}}\right). \tag{1.21}$$

Then, under monopsony, estimating equation (1.16) yields an estimate of the mark-up of price over marginal cost and of the labor supply elasticity to the firm. Moreover, the parameter of joint market imperfections is given by:

$$\psi_{ft} \equiv \frac{(\varepsilon_M)_{ft}}{(\alpha_M)_{ft}} - \frac{(\varepsilon_N)_{ft}}{(\alpha_N)_{ft}} = -\mu_{ft}\frac{1}{(\varepsilon_w)_{ft}}. \tag{1.22}$$

In this setting we expect ψ to be negative and labor market frictions to generate upward-sloping labor supply curves to individual firms giving some degree of market power to employers. In the monopsony setting, the marginal employee receives a wage that is less than her marginal revenue.

1.3 Data description

In the first part of this study we use the Portuguese dataset Simplified Corporate Information - IES (*Informação Empresarial Simplificada*) - which covers the population of virtually all Portuguese nonfinancial corporations[2]. Data are compiled and disseminated by Statistics Portugal (*Instituto Nacional de Estatística (INE)*) and consists of a new system to collect firm mandatory annual economic, financial, and accounting information for a single moment and a single entity.

Firms report detailed balance sheet information as well as information on several important variables, namely employment and transactions of goods and services by geographical area. Even though data on IES started being collected in 2006, there was a report collecting data in 2005 that was also taken into consideration in the analysis. We obtain an unbalanced panel of nonfinancial Portuguese firms spanning eight years. We restrict our sample to manufacturing firms with at least six years of

[2]The sampling method consists of non-financial corporations covering all sectors of activity defined in the Portuguese Classification of Economic Activities with the following exceptions: financial intermediation, general government, private households with employed persons, international organizations, and other non-resident institutions.

observations for identification purposes. We consider only observations with nonzero sales and capital, employing at least one worker, and observations with factor shares of labor or materials inside the unit interval. Also, we consider 1 and 99 percentiles as cutoff levels for output and input growth rates. We use sales as the measure of output (Q), labor is the average number of employees (N), capital is the net book value of fixed assets (K), and material is intermediate consumption (M).

The main descriptive statistics of the variables included in the analysis are reported in Table 1.1. The shares of labor and materials in output are obtained by dividing the firm total labor cost and intermediate consumption, respectively, by the firm production as measured by firm sales.

Table 1.1: Main summary statistics

	2006-2012				
	Mean	St. Dev.	Q_1	Q_2	Q_3
Δ_q: Output growth	0.015	0.261	-0.115	0.014	0.143
Δ_n: Labor input growth	-0.004	0.221	-0.065	0.000	0.061
Δ_m: Materials input growth	0.004	0.329	-0.152	0.010	0.165
Δ_k: Capital input growth	-0.045	0.417	-0.220	-0.078	0.051
α_n: Share of labor costs	0.307	0.172	0.184	0.280	0.393
α_m: Share of materials	0.584	0.182	0.484	0.606	0.714
$1-\alpha_n-\alpha_m$: Share of capital	0.109	0.082	0.051	0.090	0.145
Solow Residual (SR)	0.015	0.159	-0.057	0.009	0.079

Notes: The sampling period goes from 2006 to 2012. The number of observations is 127,869. The variables Δ_q, Δ_n, Δ_m, and Δ_k represent the annual growth rates of output, labor, materials, and capital, respectively, in the sampling period. The variables α_n, α_m, and $\alpha_k = 1-\alpha_n-\alpha_m$ are the shares of labor, materials, and capital averaged over adjacent periods. The Solow residual is calculated as follows: SR$=\Delta_q-\alpha_n\Delta_n-\alpha_m\Delta_m-(1-\alpha_n-\alpha_m)\Delta_k$. Q_1 and Q_3 correspond to the first and third quartiles and Q_2 corresponds to the median.

Then, in the second part of this study we merge the estimated firm labor supply elasticity and firm total factor productivity with a matched employer-employee-job title dataset known as *Quadros de Pessoal* (Personnel Records). This dataset was created by the Portuguese Ministry of Employment and is an annual mandatory employment survey addressed to establishments employing at least one wage earner.

Data are available from 1986 to 2012 for each wage earner, with the exception of workers of the Public Administration sector and domestic servants.

Detailed data are available on the establishment (location, economic activity, and employment), the firm (location, economic activity, employment, sales, year of formation, and legal framework), and for each and every of its workers (gender, age, education, occupation, earnings - base wage, seniority-related earnings, other regular and irregular benefits, overtime pay, normal and overtime hours, and tenure)[3].

To estimate a Mincerian wage equation we considered a subset of these variables and some restrictions were imposed in the dataset. We restricted the analysis to full-time workers who were aged between 18 and 65 years old, and who earn a nominal wage of at least 80 percent of the mandatory minimum wage. Also, we excluded from the analysis workers from the agriculture and fishery sectors. Finally, we dropped around two percent of the observations that did not belong to the largest connected set. The dependent variable considered in the estimation is the natural logarithm of the real hourly wage.

1.4 Product and labor market imperfections parameter

The baseline model formulated to derive the product and labor market imperfection parameters is presented in equation (1.16). We directly estimate from the production function the labor and material output elasticities to derive the joint imperfections parameter as the difference between the output elasticity-revenue share ratio for labor and materials. The sign and significance of this parameter will determine which regime applies. Dobbelaere & Mairesse (2011) use the Swamy (1970) methodology and document considerable within-regime firm differences in the estimated product and labor markets imperfection parameters.

We believe production function estimates differ across firms due to firms' idiosyncratic heterogeneity and heterogeneity in the product and labor markets they operate in. Hence, in this study we resort to an empirical methodology that allows

[3]For a more detailed description of the dataset *Quadros de Pessoal* see Torres *et al.* (2012), for example.

us to derive a distribution of the labor and product markets imperfection parameters through the estimation of 5-digit sector-specific factor elasticities. The estimation uses the high dimensional fixed effects procedure developed by Portugal & Guimarães (2010) to compute the elasticity of output with respect to labor, materials, and capital through the estimation of a linear regression model with one interacted high dimensional fixed effect. The high dimensional fixed effect considered in the analysis is the 5-digit classification of economic activities. This level of disaggregation of the economic activity leads us to believe that we are close to the firm definition. In this way we are able to draw a distribution of the 5-digit sector estimated output elasticities of labor, materials and capital as directly estimated from the production function. Then, we can obtain an estimate for the firm-specific joint market imperfections parameter and derive firm-specific estimates of the price-cost mark-up and labor market imperfection parameters. The baseline empirical specification to be estimated considers constant returns to scale[4] ($\lambda_{ft} = 1$) and is given by:

$$q_{ft} - k_{ft} = (\varepsilon_N)_s[n_{ft} - k_{ft}] + (\varepsilon_M)_s[m_{ft} - k_{ft}] + \theta_{ft}. \tag{1.23}$$

Hence, the output elasticity for capital is given by $(\varepsilon_K)_s = 1 - (\varepsilon_N)_s - (\varepsilon_M)_s$.

The distribution of the estimates for the elasticities of labor, materials, and capital with respect to output obtained through the estimation of the production function presented in equation (1.23) are shown in Figure 1.1. These figures show considerable dispersion in the estimated output elasticities.

These firm-level estimates are then considered to calculate a distribution of the joint market imperfections parameter[5]. The results are depicted in Figure 1.2. This figure shows that many firms in the sample are characterized by an efficient bargaining model ($\widehat{\psi}_f > 0$), while several others are classified as a monopsony ($\widehat{\psi}_f < 0$).

[4]The estimation results should be robust to this assumption since the first-order conditions do not depend on the elasticity of scale parameter.

[5]$\widehat{\psi}_f = \frac{(\widehat{\varepsilon}_M)_s}{(\bar{\alpha}_M)_f} - \frac{(\widehat{\varepsilon}_N)_s}{(\bar{\alpha}_N)_f}$, where $(\widehat{\varepsilon}_M)_s$ and $(\widehat{\varepsilon}_N)_s$ are the 5-digit sector-specific output elasticities of materials and labor, respectively, estimated from the production function, and $(\bar{\alpha}_N)_f$ and $(\bar{\alpha}_M)_f$ are the firm time-averaged shares of labor costs and intermediate consumption in total revenue.

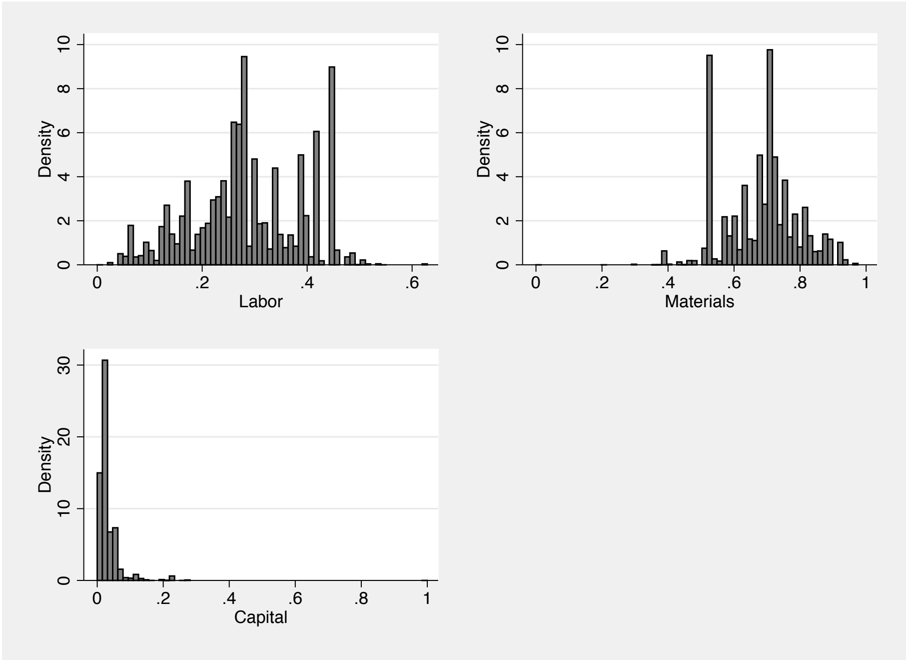

Figure 1.1: Distribution of 5-digit sector estimated elasticities of labor, materials, and capital with respect to output weighted by the firm average number of employees.

In the first case, workers are assumed to receive a wage that exceeds their marginal revenue, while in the second case workers receive a wage that is less than their marginal revenue[6].

Once the regime is identified we can compute the product and labor market imperfections parameters as measured by the firm price-cost mark-up and rent-sharing or monopsony power, respectively. The empirical distribution of the estimated firm price-cost mark-up[7] is shown in Figure 1.3, and suggests that a great number of firms operate in an imperfect competitive product market. Therefore, the estimates suggest that firms possess a considerable degree of market power in the product market.

In this study we focus the analysis on firms that possess some degree of monopsony power ($\widehat{\psi}_f < 0$) and explore the distribution of the estimated labor supply elasticity to a particular firm. The results for the firm labor supply elasticity $\widehat{\beta}_f$[8]

[6]The case $\widehat{\psi}_f = 0$ corresponds to the right-to-manage model (see Dobbelaere & Mairesse (2011) for details).

[7]$\widehat{\mu}_f = \frac{(\widehat{\varepsilon}_M)_s}{(\bar{\alpha}_M)_f}$.

[8]$\widehat{\beta}_f = \frac{(\bar{\alpha}_N)_f}{(\bar{\alpha}_M)_f}\frac{(\widehat{\varepsilon}_M)_s}{(\widehat{\varepsilon}_N)_s}$.

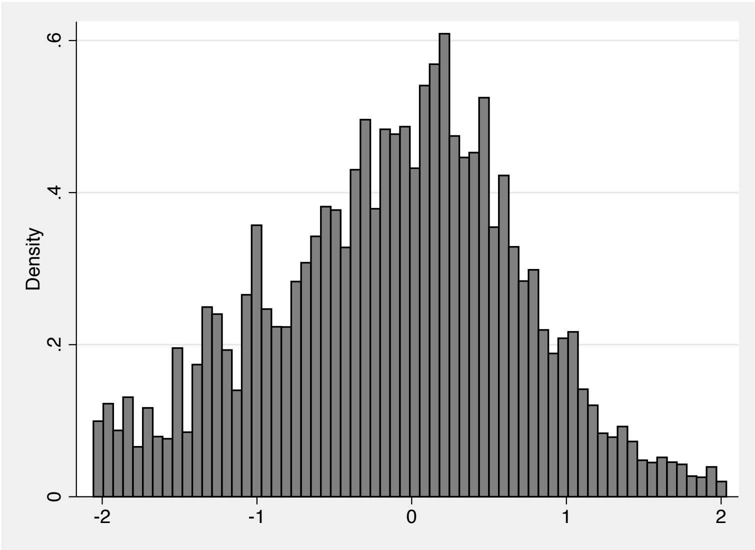

Figure 1.2: Distribution of firm estimated joint market imperfections parameter $(\widehat{\psi}_f)$weighted by the firm average number of employees.

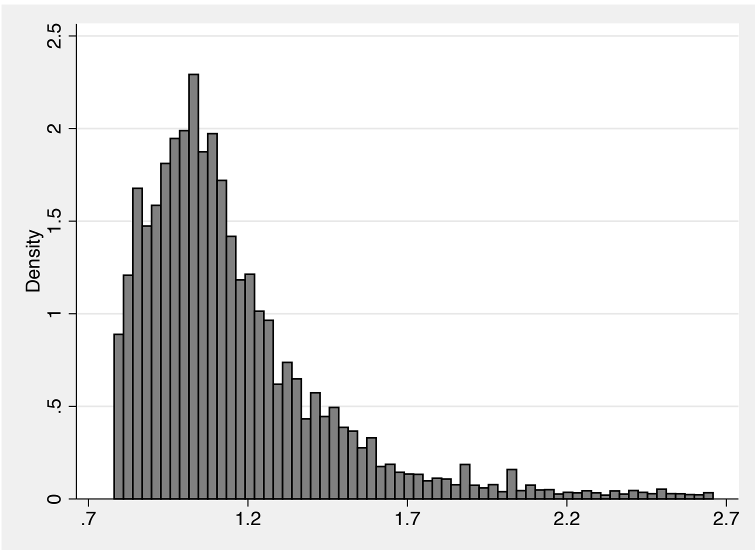

Figure 1.3: Distribution of firm estimated price-cost mark-up $(\widehat{\mu}_f)$ weighted by the firm average number of employees.

and $(\widehat{\varepsilon}_w)_f = \frac{\widehat{\beta}_f}{1+\widehat{\beta}_f}$ are shown in Figure 1.4. We find evidence for imperfect competition in the labor market with considerable dispersion in the estimated labor supply elasticity across firms even within the same labor market. The empirical distributions of the absolute and relative extent of rent-sharing ($\widehat{\phi}_f$ and $\widehat{\gamma}_f$, respectively) calculated for firms in the efficient bargaining setting ($\widehat{\psi} < 0$) are depicted in Figure 1.5. The results show considerable dispersion in the extent of rent-sharing within the efficient bargaining setting. The main descriptive statistics of the estimated parameters are reported in Table 1.2.

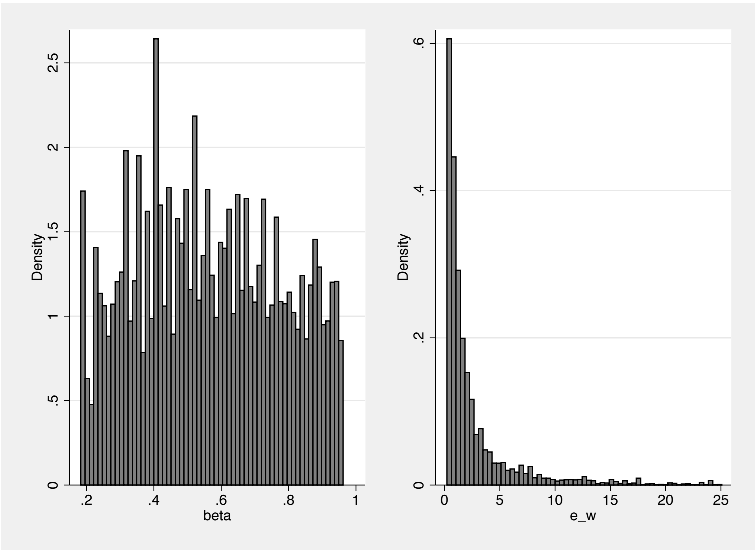

Figure 1.4: Distribution of firm estimated elasticity of labor supply ($\widehat{\beta}_f$ and $(\widehat{\varepsilon}_w)_f$) weighted by the firm average number of employees.

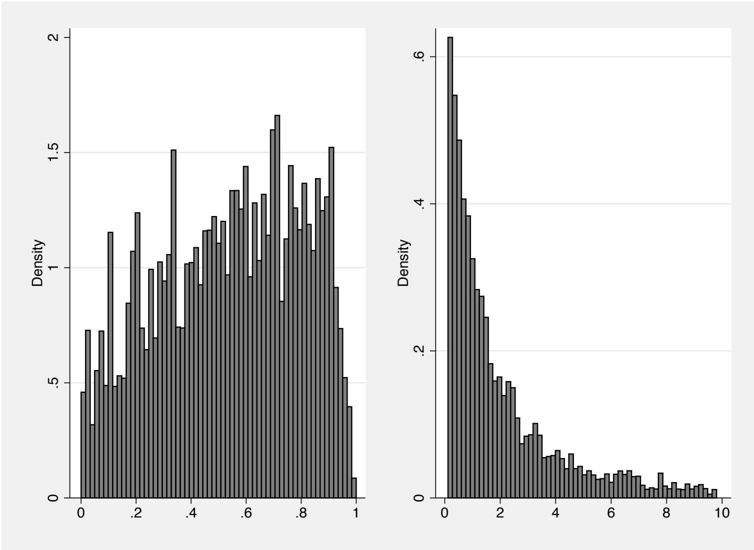

Figure 1.5: Distribution of firm estimated absolute ($\widehat{\phi}_f$) and relative ($\widehat{\gamma}_f$) extent of rent-sharing weighted by the firm average number of employees.

Table 1.2: Main summary statistics: Labor and product markets imperfections parameters

	2006-2012				
	Mean	St.Dev.	Q_1	Q_2	Q_3
Joint markets imperfections parameter ($\widehat{\psi}$)	0.091	0.769	-0.357	0.157	0.597
Price-cost mark-up ($\widehat{\mu}$)	1.226	0.330	1.002	1.146	1.359
Wage elasticity of labor supply ($\widehat{\varepsilon}_w$)	3.271	4.228	0.767	1.624	3.824
Relative extent of rent-sharing ($\widehat{\gamma}$)	2.166	2.057	0.666	1.433	2.981

Notes: The sampling period goes from 2006 to 2012. Q_1 and Q_3 correspond to the first and third quartiles and Q_2 corresponds to the median. The descriptive statistics of the joint markets imperfections parameter and the price-cost mark-up are based on 109,812 observations. The efficient bargaining and the monopsony parameters are based on 63,539 and 46,620 observations, respectively.

These results show that some firms possess a considerable degree of product and labor market power and confirm that the hypothesis of perfectly competitive product and labor markets is not suitable to characterize these markets.

1.5 Total factor productivity

The total factor productivity (TFP) is estimated through the following equation:

$$\mathcal{Q}_{ft} = \varepsilon_N N_{ft} + \varepsilon_M M_{ft} + \varepsilon_K K_{ft} + \eta_f + \nu_{st} + u_{ft}, \tag{1.24}$$

where η_f accounts for time-invariant observed and unobserved firm heterogeneity, ν_{st} is a 5-digit sector s specific time trend that allows us to control for sector-specific productivity shocks, and u_{it} is a residual component. Therefore, firm-level TFP is given by $\Theta_{ft} = A\exp(\eta_f + \nu_{st} + u_{ft})$. The results of the high dimensional fixed effects estimation (see Portugal & Guimarães (2010) for details on the estimation procedure) are reported in Table 1.3[9]. Figure 1.6 depicts the distribution of firm TFP weighted by the number of employees. Our results are in line with earlier literature showing considerable variation in the productivity of Portuguese firms, with a large number of firms being relatively low productive and a small number of firms being more productive.

Table 1.3: Estimation results: Total factor productivity

	Coef.	Std. Error
Materials	0.6222	0.0004
Labor	0.2229	0.0003
Capital	0.0184	0.0002
Observations	127,869	
R^2	0.996	

Notes: The sampling period goes from 2006 to 2012. The dependent variable is the natural logarithm of sales. Materials refers to firm's intermediate consumption, labor is measured by the average number of employees, and capital is the net book value of the tangible assets. Linear regression estimation with two high dimensional fixed effects: firm fixed effects and time fixed effects interacted with 5-digit sector dummies.

[9]We also estimate this model using the two semi-parametric approaches proposed by Olley & Pakes (1996) and Levinsohn & Petrin (1999). The first uses the firm's investment decision to proxy for the unobserved time-varying productivity shock to account for the problem of simultaneity, and considers survival probabilities to address the problem of selectivity. The second is similar but uses the intermediate inputs to proxy for unobservable variables. The results are very similar to those obtained through the high dimensional fixed effects estimation with firm fixed effects and year fixed effects interacted with 5-digit sector dummies.

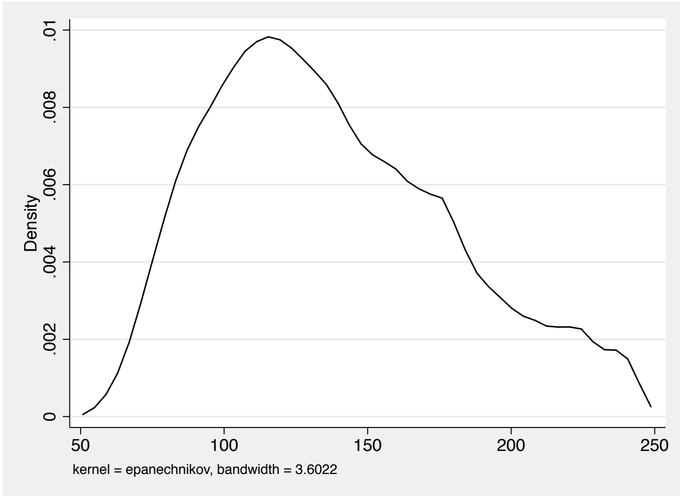

Figure 1.6: Kernel density estimate of firm TFP weighted by the firm average number of employees.

1.6 Wage regressions

Empirical evidence reported in recent decades suggests the presence of considerable variability in wages (Abowd *et al.* (1999b), Abowd *et al.* (1999a) and Cardoso *et al.* (2016)). Researchers have estimated wage regressions incorporating both worker effects and firm effects with the goal of disentangling the effects of worker decisions and firm wage policies in wage formation. More recently, Torres *et al.* (2012) use Portuguese longitudinal matched employer-employee data to estimate a wage regression and add job title heterogeneity as an important third dimension of wage formation. In fact, the characteristics of some tasks (namely, the risk of fatal or serious accidents, the workplace conditions in which the tasks are performed, and the specific training or skills that some tasks require) contribute to wage differentiation. The wage decomposition shows that in Portugal, firm, worker, and job title time-invariant heterogeneity accounts for a significant fraction of the total wage variation. The authors estimate that worker permanent heterogeneity is the primary source of wage variation, accounting for approximately 36 per cent, followed by firm permanent effects, which account for almost 29 per cent of the total wage variation. The job title permanent heterogeneity plays a less significant but non-negligible role, explaining close to 10 per cent of wage variation.

The presence of frictions in the labor market may explain why wages vary across labor markets and even across firms within a given labor market. In this section we explore the importance of the firm-specific degree of monopsony power and firm total factor productivity in explaining wage formation. Furthermore, we study the contribution of rent-sharing for wage formation. We pick the estimates of the labor supply elasticity and relative extent of rent-sharing as obtained in Section 1.4, and combine these with firm-level productivity as calculated in Section 1.5.

Next we present the methodology applied in this study to understand the importance of monopsony power, relative extent of rent-sharing, and total factor productivity in explaining the firm wage setting policy. First, we follow Torres *et al.* (2012) and estimate a standard Mincerian wage equation with the inclusion of three high dimensional fixed effects to account for firm, worker, and job-title time-invariant observed and unobserved heterogeneity:

$$\ln w_{ifjt} = \mathbf{X}_{ift}\beta + \phi_i + \gamma_f + \omega_j + \tau_t + \epsilon_{ifjt}, \tag{1.25}$$

where the dependent variable $\ln w_{ifjt}$ is the natural logarithm of the real hourly wage of worker i ($i = 1, ..., N$) working at firm f ($f = 1, ..., F$) with the job title j ($j = 1, ..., J$) in year t ($t = 1, ..., T$). The vector $\mathbf{X}_{ift}$ is a row-vector of k observed characteristics of the worker i and firm f (and includes the quadratic terms on age and tenure within the firm and the worker qualifications). The term ϕ_i is a worker fixed effect, γ_f is a firm fixed effect, ω_j is a job fixed effect, and τ_t is a set of year dummies. The disturbance term ϵ_{ifjt} has the conventional properties.

This equation is estimated using the *Quadros de Pessoal* matched employer-employee data for the period between 1986 and 2012 and applying the iterative algorithm developed by Portugal & Guimarães (2010), which produces the exact solution of the least squares estimation of equations with three high dimensional fixed effects. From the estimation of this equation we obtain the estimated firm $\widehat{\gamma}_f$, worker $\widehat{\phi}_i$, and job title $\widehat{\omega}_j$ fixed effects, which represent firm, worker, and job title time-invariant observed and unobserved heterogeneity, respectively.

Then, we combine the estimates of the labor supply elasticities and total factor productivity computed in the previous sections with the *Quadros de Pessoal* dataset[10]. We proceed by applying the Gelbach's exact decomposition (Gelbach (2016)) to quantify the importance of the firm monopsony power and total factor productivity to explain total wage variation[11].

The base specification is given by:

$$\ln w_{ift} = \alpha_0 + \alpha_1 \widehat{LMI}_{ft} + \alpha_2 \widehat{\Theta}_{ft} + \mathbf{X}_{ift}\xi + \tau_t + \vartheta_{it}, \tag{1.26}$$

where the dependent variable is the natural logarithm of the real hourly wage ($\ln w_{it}$) and the explanatory variables are the estimated labor market imperfection ($\widehat{LMI}_{ft}$) parameter, which is the firm's monopsony power ($1/(\widehat{\varepsilon}_w)_f$) in the case of monopsonist firms and the relative extent of rent-sharing ($\widehat{\gamma}_{ft}$) in the case of firms in the efficient bargaining setting, the estimated firm total factor productivity ($\widehat{\Theta}_{ft}$), and a vector of explanatory variables $\mathbf{X}_{ift}$ (quadratic terms on age and tenure, worker qualifications, and worker gender). The term τ_t is a vector of year dummies and ϑ_{it} is a disturbance term with the conventional properties.

In turn, the full specification is given by the following equation:

$$\ln w_{ifjt} = a_0 + a_1 \times \widehat{LMI}_{ft} + a_2 \widehat{\Theta}_{ft} + \mathbf{X}_{ift}\beta + \widehat{\phi}_i + \widehat{\gamma}_f + \widehat{\omega}_j + \tau_t + \epsilon_{ifjt}, \tag{1.27}$$

where $\widehat{\phi}_i$, $\widehat{\gamma}_f$, and $\widehat{\omega}_j$ are the estimated worker, firm, and job-title time-invariant heterogeneity.

[10] We end with a four-dimensional panel data (firm, worker, job-title, and year dimensions) spanning from 2006 to 2012, since the labor market imperfection parameters and the total factor productivity are estimated using the *IES* information, which is available only for this period.

[11] A more detailed presentation of the Gelbach's decomposition can be found in the Appendix.

1.7 Estimation results

1.7.1 Monopsony

The estimation results of equation (1.26) are reported in columns (1) and (2) of Table 1.4. According to these estimates, the elasticity of labor supply is positively and significantly associated with wages. Since the elasticity of labor supply is inversely related to monopsony power, this means that firms with more market power manage to pay lower wages to their workers. The estimates presented in column (2) suggest that a one standard deviation increase in firm's monopsony power leads earnings of workers to decrease by approximately 2 percent, *ceteris paribus*. Also, we find that more productive firms pay on average higher wages, holding everything else constant. It is interesting to notice that the elasticity of labor supply to the firm and the firm total factor productivity alone explain a considerable fraction of the variation in the earnings of its workers (approximately 39 percent).

To shed further light on the importance of market power and productivity to explain the wage policy of the firm as measured by the firm time-invariant heterogeneity, we proceed with the Gelbach's decomposition. The estimation results of the full model presented in equation (1.27) are reported in column (3) of Table 1.4. Then, in columns (1), (2), and (3) of Table 1.5 we find the decomposition of the wage differential given by the difference between the estimates of the base and full models reported in columns (2) and (3) of Table 1.4, respectively. The results suggest that, in fact, monopsony power is mostly related to the firm permanent heterogeneity. This reveals the role of monopsony power and firm's total factor productivity to explain heterogeneity in wage formation, even after controlling for detailed firm, worker, and job title characteristics. This is in line with the suggestion of Goux & Maurin (1999) and Abowd *et al.* (1999b) that the presence of firm effects in wage regressions, after controlling for person and industry characteristics, is strongly suggestive of market power. The coefficients of firm's monopsony power and firm's productivity converge to zero in the case of the worker and job title permanent

Table 1.4: Monopsony power and total factor productivity

	(1) $\ln w_{ifjt}$	(2) $\ln w_{ifjt}$	(3) $\ln w_{ifjt}$
Monopsony power $(1/(\widehat{\varepsilon}_w)_f)$	-0.0234***	-0.0108***	-0.0002**
	(0.0003)	(0.0002)	(0.0001)
Total factor productivity $(\widehat{\theta}_{ft})$	0.5411***	0.4075***	0.0039***
	(0.0010)	(0.0010)	(0.0008)
Age		0.0290***	0.0199***
		(0.0002)	(0.0001)
Age^2		-0.0002***	-0.0002***
		(0.0000)	(0.0000)
Tenure		0.0158***	0.0081***
		(0.0001)	(0.0001)
Tenure^2		-0.0003***	-0.0002***
		(0.0000)	(0.0000)
Gender		-0.2607***	
		(0.0007)	
Constant	-2.4622***	-2.4366***	0.1232***
	(0.0058)	(0.0069)	(0.0047)
Fixed effects	No	No	Yes
Observations	1,022,391	1,022,391	1,022,391
Adjusted R^2	0.391	0.546	0.889

Notes: The dependent variable is the natural logarithm of real hourly wage ($\ln w_{ifjt}$). The sampling period goes from 2006 to 2012. All specifications include year dummies and dummy variables for education levels. Ordinary Least Squares (OLS) estimation with bootstrap standard errors (using 1000 draws) in parentheses. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

heterogeneity[12].

Table 1.5: Gelbach decomposition

	(1) $\widehat{\gamma}_f$	(2) $\widehat{\phi}_i$	(3) $\widehat{\omega}_j$
Monopsony power $(1/(\widehat{\varepsilon}_w)_f)$	-0.0112***	0.0015***	-0.0009***
	(0.0001)	(0.0001)	(0.0001)
Total factor productivity $(\widehat{\theta}_{ft})$	0.3886***	0.0173***	-0.0023***
	(0.0004)	(0.0006)	(0.0003)
Observations	1,022,391	1,022,391	1,022,391
Adjusted R^2	0.593	0.376	0.359

Notes: The dependent variable is the firm, worker, and job title time-invariant heterogeneity in columns (1), (2), and (3), respectively. The sampling period goes from 2006 to 2012. All specifications include year dummies and dummy variables for education levels. Ordinary Least Squares (OLS) estimation with bootstrap standard errors (using 1000 draws) in parentheses. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

1.7.2 Efficient bargaining

The contributions of the relative extent of rent-sharing parameter and total factor productivity to explain wage formation are presented in Table 1.6. The estimates related to the role of total factor productivity in explaining wages are consistent with those obtained in the previous section and show that total factor productivity is a crucial determinant of wage heterogeneity.

Based on the standard collective bargaining literature, we would expect a positive correlation between the relative extent of rent-sharing parameter and wages. However, the estimation results suggest a negative impact of the relative extent of rent-sharing on wages and therefore workers with a larger share of the rents fail to obtain extra income. Instead, a larger share of the rents is estimated to depress the wages paid by the firm.

While not intuitive, this result is similar to the results presented in Dobbelaere & Mairesse (2011). The authors estimate the correlation between the relative extent

[12]These results are robust to using different criteria to characterize firms according to the labor market setting, namely using a statistical significance criterion.

Table 1.6: Efficient bargaining and total factor productivity

	(1) $\ln w_{ifjt}$	(2) $\ln w_{ifjt}$	(3) $\ln w_{ifjt}$
Relative extent of rent-sharing ($\widehat{\gamma}_{ft}$)	-0.0279***	-0.0130***	-0.0007***
	(0.0002)	(0.0002)	(0.0001)
Total factor productivity ($\widehat{\theta}_{ft}$)	0.5343***	0.4460***	0.0097***
	(0.0013)	(0.0011)	(0.0008)
Age		0.0265***	0.0192***
		(0.0003)	(0.0001)
Age^2		-0.0002***	-0.0002***
		(0.0000)	(0.0000)
Tenure		0.0162***	0.0077***
		(0.0001)	(0.0001)
Tenure^2		-0.0003***	-0.0002***
		(0.0000)	(0.0000)
Gender		-0.2787***	
		(0.0008)	
Constant	-2.3529***	-2.4728***	0.1180***
	(0.0069)	(0.0076)	(0.0047)
Fixed effects	No	No	Yes
Observations	838,563	838,563	838,563
Adjusted R^2	0.382	0.542	0.884

Notes: The dependent variable is the natural logarithm of real hourly wage ($\ln w_{ifjt}$). The sampling period goes from 2006 to 2012. All specifications include year dummies and dummy variables for education levels. Ordinary Least Squares (OLS) estimation with bootstrap standard errors (using 1000 draws) in parentheses. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

of rent-sharing and firm size, capital intensity, among other firm variables, and also find a negative correlation between these variables.

1.7.3 The gender pay gap

Monopsonistic competition may help to explain one of the stylized empirical results in the labor economics literature, which is the gender pay gap. Ransom & Oaxaca (2010) and Hirsch *et al.* (2010) investigate women's and men's labor supply to the firm separately using a dynamic model of monopsony and find that women have lower elasticities than men. The reasons for this result may be different preferences over nonwage job characteristics (namely, hours of work and job location) and a higher degree of worker immobility. Monopsonist employers may take advantage of this lower female elasticity of labor supply to the firm and pay lower wages to women, *ceteris paribus*. Hirsch *et al.* (2010) suggest that this result implies that at least one-third of the gender pay gap might be wage discrimination by monopsonist employers[13].

These differences in the labor supply elasticity between women and men suggest that it is likely that the marginal impact of increasing the elasticity of labor supply at the firm level may differ considerably across these two groups. Since in our model monopsony power is inversely related to the labor supply elasticity at the firm level, this means that the ability of monopsonist firms to mark down wages is greater in the case of female workers. In fact, in our sample the average estimated labor supply elasticity for female and male workers is approximately 1.95 and 2.316, respectively. This means that firms hiring a large fraction of male workers have on average less monopsony power.

Table 1.7 presents the results of the estimation of equation (1.26) separately for male and female workers. These results make it clear that the marginal impact of increasing the labor supply elasticity to a particular firm is much lower for female workers and that there are considerable differences in how market power on the

[13] This explanation aligns with the Robinsonian monopsony model of wage discrimination (Robinson (1933)).

firms' side affects workers' wages.

Table 1.7: Monopsony power and total factor productivity by gender

		By gender	
	All	Male	Female
Monopsony power ($1/(\widehat{\varepsilon}_w)_f$)	-0.0108***	-0.0152***	-0.0058***
	(0.0002)	(0.0003)	(0.0003)
Total factor productivity ($\widehat{\theta}_{ft}$)	0.4075***	0.4190***	0.3834***
	(0.0010)	(0.0013)	(0.0014)
Age	0.0290***	0.0349***	0.0201***
	(0.0003)	(0.0003)	(0.0004)
Age^2	-0.0002***	-0.0003***	-0.0001***
	(0.0000)	(0.0000)	(0.0000)
Tenure	0.0158***	0.0174***	0.0127***
	(0.0001)	(0.0002)	(0.0002)
Tenure^2	-0.0003***	-0.0003***	-0.0002***
	(0.0000)	(0.0000)	(0.0000)
Gender	-0.2607***		
	(0.0007)		
Constant	-2.4366***	-2.9231***	-2.5561***
	(0.0068)	(0.0093)	(0.0106)
Fixed effects	No	No	No
Observations	1,022,391	616,780	405,611
Adjusted R^2	0.546	0.501	0.513

Notes: The dependent variable is the natural logarithm of real hourly wage ($\ln w_{ifjt}$). The sampling period goes from 2006 to 2012. All specifications include year dummies and dummy variables for education levels. Ordinary Least Squares (OLS) estimation with bootstrap standard errors (using 1000 draws) in parentheses. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

1.8 Conclusions

A central feature of perfectly competitive markets is that markets clear, meaning that all workers with similar quality should be paid the same market clearing wage. Recent empirical evidence suggests the presence of considerable wage dispersion among workers with similar characteristics and among similar firms. A potential explanation for the presence of firm effects in wage regressions after accounting for detailed firm, worker, and job title heterogeneity rely on the presence of considerable frictions in the labor market, namely asymmetric information, worker immobility, and heterogenous preferences, which may constitute sources of market power for employers.

In the new monopsony literature, search frictions imply that firms may face an upward labor supply curve even if operating in a labor market with many competing firms.

In this study we use matched employer-employee data and firm balance sheet data to investigate the importance of firm total factor productivity and firm labor market power in explaining firm heterogeneity in wage formation. We use a linear regression model with one interacted high dimensional fixed effect to estimate 5-digit sector-specific elasticity of output with respect to input factors directly from the production function. This allows us to derive firm-specific price-cost mark-up and firm-specific elasticity of labor supply. The results suggest that many Portuguese firms are classified as monopsonist, and that there exists a broad range of firm market power among monopsonist firms. The hypothesis that the elasticity of labor supply is finite has major implications for theoretical models of labor economics.

We proceed by investigating the impact of the elasticity of labor supply to a particular firm and firm total factor productivity on individuals' earnings. Furthermore, we use the Gelbach's exact decomposition to understand how firm's monopsony power is associated with the firm's wage setting policy. The results suggest that a one standard deviation increase in the labor supply elasticity increases wages by approximately 1.51 percent, *ceteris paribus*. This means that monopsony power affects negatively the wages of workers. Also, we find evidence that the elasticity of labor supply is mainly correlated with the firm effects as hypothesized in the labor economics literature. This suggests that firm market power is a key ingredient to explain heterogeneity in wage formation.

Last, we analyze if there are any gender differences on the impact of the labor supply elasticity on earnings. The results reveal that the marginal impact of increasing the labor supply elasticity to a particular firm is much lower for female workers and that there are considerable differences in how market power on the firms' side affects workers' wages. This finding is intimately related with the gender pay gap, and suggests that we should consider firms' market power when analyzing wage differentials arising from gender differences.

Appendix

Gelbach decomposition

In this section we closely follow Cardoso *et al.* (2016) to present the methodological details related to the Gelbach decomposition proposed by Gelbach (2016). The linear wage equation estimated is given by:

$$\ln w_{ifjt} = \mathbf{X}_{ift}\beta + \phi_i + \gamma_f + \omega_j + \tau_t + \epsilon_{ifjt}, \tag{1.28}$$

where $\ln w_{ifjt}$ is the natural logarithm of the real hourly wage of individual i (i=1,...,N) working at firm f (f=1,...,F) holding a job title j (j=1,...,J) at year t (t=1,...,T). The vector X_{ift} contains k observed time-varying characteristics of individual i (quadratic terms on age and yearly seniority within the firm). The terms ϕ_i, γ_f, and ω_j represent the individual, firm, and job-title fixed effects, respectively, and measure observed and unobserved individual, firm, and job-title time-invariant heterogeneity. The term τ_t is a set of year dummies.

Consider the basic regression of the natural logarithm of hourly wages on the set of explanatory variables defined above and time dummies. This can be expressed in matrix notation as:

$$\mathbf{Y} = \mathbf{X}\mathbf{b} + \epsilon. \tag{1.29}$$

Then, following Gelbach (2016) and the omitted variable bias formula we can write the difference between the coefficients of the basic specification defined in equation (1.29) and those of the full specification presented in equation (1.28) as:

$$\widehat{\mathbf{b}} - \widehat{\beta} = \mathbf{P_X}\mathbf{D}_i\widehat{\phi} + \mathbf{P_X}\mathbf{D}_f\widehat{\gamma} + \mathbf{P_X}\mathbf{D}_j\widehat{\omega}, \tag{1.30}$$

where $\mathbf{P_X} = (\mathbf{X}'\mathbf{X})^{-1}\mathbf{X}'$ and $\mathbf{D}_i\widehat{\phi}$, $\mathbf{D}_f\widehat{\gamma}$, and $\mathbf{D}_j\widehat{\omega}$ are column vectors containing the estimates of the fixed effects for the worker, firm, and job title, respectively. This means that $\mathbf{P_X}\mathbf{D}_i\widehat{\phi}$ is the coefficient of the regression of the worker fixed effects on the set of variables $\mathbf{X}$ in the base model. A similar interpretation applies to the two remaining terms in the right-hand side of equation (1.30). Then, we can rewrite the previous equation more succinctly as:

$$\widehat{\mathbf{b}} - \widehat{\beta} = \widehat{\delta}_\phi + \widehat{\delta}_\gamma + \widehat{\delta}_\omega. \tag{1.31}$$

Then, the change in the coefficient of interest is partitioned into the role of the different additional covariates, and the conditional contribution of the worker, firm, and job title fixed effects to explain the firm labor market imperfection parameter can be identified.

Chapter 2

Price setting and market competition

2.1 Introduction

Macroeconomists have been studying for a long time the large effects of demand shocks on output and, in particular, the non-neutrality of monetary policy on real output. A growing empirical literature has emerged in the last decades documenting the large effects of demand shocks on output (Blanchard & Perotti (2002), Ramey (2009), and Nakamura & Steinsson (2014)) and the non-neutrality of monetary shocks (Friedman & Schwartz (2008), Christiano *et al.* (1999), and Romer & Romer (2004)). A central hypothesis for explaining these findings has been the sluggish price adjustment to aggregate conditions. A large class of macroeconomic models has been conceived in which firms, acting in a dynamic and stochastic environment, face barriers to adjusting their prices in response to aggregate shocks. From our standpoint, it is possible to summarize the theoretical literature along three lines of research, according to how frictions to price-setting are formalized and, consequently, to their predictions regarding inflation-output dynamics[1].

[1]At the same time, some of the most critical contributions to our understanding of nominal rigidity were introduced through static partial-equilibrium models. Two of the most notable examples are the seminal studies of Akerlof & Yellen (1985) and Ball & Romer (1990). The former suggests that slightly non-optimizing behavior - "near rationality" - from price-setters could induce sizable real effects from monetary disturbances, while the latter shows that explicit barriers

Until recently, the literature was divided between models assuming that firms would adjust their prices at an exogenously set timing and models that feature the timing of price changes as endogenous to the firm. The first strand of literature, known as time-dependent models, evolved from assuming that price changes follow a deterministic pattern (Fischer (1977), Taylor (1980)) to postulating that opportunities for a firm to reset its price come randomly, as suggested by Calvo (1983). Another strand of the literature, denoted as state-dependent models, formalizes the idea that firms stand ready to change prices whenever it is worth doing so with the introduction of a real cost to change a price, the menu cost. In these models, firms with a large difference between the current and its desired price are the most likely to change prices in response to a given shock (see, for example, Golosov & Lucas (2007) and Dotsey *et al.* (1999)). Time-dependent models tend to yield larger and more persistent real effects from monetary disturbances than state-dependent models because the set of firms that adjust prices are exogenously determined in the former.

In the last decade, renewed attention to Lucas' (1972) imperfect information model spurred a new strand of literature building on the idea that monetary non-neutrality arises from imperfect information about the state of the economy. The model developed by Mankiw & Reis (2002) posits that firms, while being able to set prices every period at no cost, adjust their planned price paths only when they receive new information on the general price level, which happens at a Calvo-like random frequency. The idea that there is too much information for agents to absorb was later refined by the rational inattention model of Maćkowiak & Wiederholt (2009), which specifies that agents have a fixed endowment of information-processing capacity and gives a formal treatment to the agent's problem of allocating its limited attention to different sources of shocks.

Nakamura & Steinsson (2013) note that one important ingredient for sluggish price adjustment in most models is coordination failure that leads to incomplete

to price adjustment are not enough for monetary disturbances to produce real effects, rather it is also required that firms face strategic incentives to not fully and immediately adjust to any shock (i.e., there must also exist real rigidity).

price adjustment, even if prices change. Coordination failures among price setters arise due to two essential factors. First, the timing of price changes is staggered, either because it is exogenous - time-dependent models - or because the timing of price changes is largely determined by idiosyncratic shocks that are orthogonal across firms - modern state-dependent models; second, due to the presence of strategic complementarity in price setting, which may result from intermediate inputs, demand structures in which the elasticity of demand is an increasing function of a firm's relative price (Kimball (1995)), or heterogeneous factor markets (Gertler & Leahy (2008)). Firms respond incompletely to an aggregate shock because other firms have not yet responded. The combination of sticky prices and coordination failure can generate large and long-lasting effects of demand shocks on output due to long-lasting sluggishness of the aggregate price level. Business cycle models combining these features are often referred to as New Keynesian models.

In the macroeconometric literature, some of the first notable studies attempted to test the New Keynesian Phillips Curve through structural estimation of DSGE models, generally finding that this relationship accounts for inflation dynamics in a satisfactory way (Galí & Gertler (1999), Galí *et al.* (2001), Smets & Wouters (2003), and Christiano *et al.* (2005)). More recent studies based on dynamic factor models have focused on the response of sectoral price indices to both macroeconomic and sectoral shocks, and conclude that they respond sluggishly to the former but relatively quickly to the latter (Boivin *et al.* (2009), Maćkowiak *et al.* (2009), Kaufmann & Lein (2013)), which is evidence in favor of the rational inattention model of Maćkowiak & Wiederholt (2009) or multi-setor menu cost models as developed by Nakamura & Steinsson (2010).

Alternatively to assessing the macroeconomic implications of the different sticky price models, and given the importance of the assumptions made about price adjustment, some authors have followed the seminal work of Bils & Klenow (2004) and studied the dynamics of prices at micro level. The increasing availability of detailed data such as that underlying the national Consumer Price Indices (CPI) or scanned data from retailers promoted this area of research[2]. Klenow & Kryvtsov (2008) and

[2] See, among other studies, Dhyne *et al.* (2006) and Vermeulen *et al.* (2007) for the euro area;

Nakamura & Steinsson (2008) use CPI data from the United States (U.S.) and extensively analyze the behavior of individual prices over time, namely the frequency and size of price changes, both concluding that the data favor state- over time-dependent models (mostly because the latter are inconsistent with a frequency of price change that comoves with inflation and the fact that the size of a price change is independent of price duration). Further evidence in favor of menu-cost models was provided by Eichenbaum *et al.* (2011) using scanner data from a large U.S. retailer. The authors observe that prices seem to be adjusted so that the mark-up is not distant from a given reference value[3].

In this study we rely on a rich micro dataset of producer prices collected at a very granular level and analyze price setting behavior of manufacturing firms in Portugal[4]. The first part of this study characterizes the duration of price spells and the probability of a price change. We proceed by first estimating unconditional survival functions for price adjustments by inflation regime. Then, we estimate a duration model that extensively accounts for unobserved heterogeneity to examine the importance of idiosyncratic and sectoral characteristics in explaining the probability of a price change. To the best of our knowledge, we are the first to estimate a duration model for price adjustment with high dimensional fixed effects, which allows us to extensively account for time-invariant (observed and unobserved) product- and firm-specific heterogeneity. In both cases we distinguish between positive and negative price changes. In fact, the shape of the hazard function of price changes conveys important information regarding the pricing rules of firms. If the conditional probability of a price change depends positively on the elapsed duration, the hazard function is upward sloping. If, conversely, the conditional probability of

and Dias *et al.* (2007) for Portugal.

[3]Klenow & Malin (2010) present a review of the micro price studies and summarize ten stylized facts: prices change at least once a year, with temporary price discounts and product turnover often playing an important role; reference prices are stickier and more persistent than regular prices; the frequency of price changes differs widely across goods, with more cyclical goods exhibiting greater price flexibility; the timing of price changes is little synchronized across sellers; the hazard (and size) of price changes does not increase with the age of the price; the cross-sectional distribution of price changes is thick-tailed, but many small price changes also occur; and finally, strong linkages exist between price changes and wage changes.

[4]One important feature of producer price data is the high correlation between the frequency of price change for manufacturing firms and consumer prices (Nakamura & Steinsson (2008)).

a price change decreases with the elapsed duration, the hazard function is downward sloping. This means that the Calvo model predicts that the hazard function of price change is flat because the probability of price adjustment is constant and does not depend on idiosyncratic shocks, and that the Taylor model predicts a zero hazard except at a single age, when the hazard equals one. In contrast, menu cost models can give rise to a wide variety of hazard functions, depending on the relative importance of inflation and idiosyncratic shocks. In fact, permanent shocks tend to produce upward-sloping hazards while transitory shocks tend to flatten or even produce downward-sloping hazards (see Nakamura & Steinsson (2008) for a discussion).

In the second part we analyze how the degree of market competition influences price setting behavior of manufacturing firms. In particular, we identify firms' competitors in the product market as those producing the same product in a given time period, considering a very detailed definition of products. To this aim we consider a peer-effects model to study the importance of competitors' price setting behavior in explaining a firm's price change.

The results suggest that competitive pressure increases the likelihood of a price increase and decreases the likelihood of a price decrease. Moreover, the probability of a price increase or price decrease is estimated to comove with inflation, at both product and sectoral level. A striking result is that when we account for time-invariant (observed and unobserved) heterogeneity, duration dependence is estimated to be positive in the case of both a price increase and a price decrease, which suggests that prices that remained longer unchanged are more likely to change in the next short period. Also, when we account for product and firm heterogeneity, prices are estimated to be more likely to increase and less likely to decrease if prices decreased the last time they changed. Overall, these results suggest that price setting depends on both idiosyncratic and sectoral shocks, and that price adjustment is considerably heterogeneous across products and firms.

Furthermore, we estimate that firms' price setting behavior responds to the price setting behavior of their competitors, even though the small magnitude of the estimated peer effect suggests that firms respond incompletely and sluggishly to their competitors' price setting decisions. This result aligns with the presence of coordination failures in price adjustment among price setters.

The paper is organized as follows: Section 3.5 presents the data and describes the main facts of price duration in Portugal. Section 2.3 presents the survival function for price adjustment. Section 2.4 presents the empirical models considered in the estimation of the hazard function of price adjustment and discusses the results. Section 2.5 presents the empirical methodology to estimate how market competition affects firms' price setting behavior and discusses the results. Section 2.6 concludes.

2.2 The data

In this study we use the Industrial Production Prices Index dataset (IPPI) collected by the Portuguese Statistics Institute (*Instituto Nacional de Estatística - INE*). These data are a monthly compilation of transaction prices obtained from companies whose principal or secondary activity is mining and quarrying, manufacturing, and electricity, gas and water supply. The sample covers a large fraction of the Portuguese manufacturing industry, representing approximately 90 percent of the production value of manufacturing firms[5]. The data are available for the period between January 1995 and September 2002. We restrict our analysis to manufacturing firms and to the period ending in December 2001 to avoid the escudo-euro contamination.

The unit of observation is a specific product defined at a very granular level and collected at the three-digit NACE level, produced in a given establishment. Items are classified using the Prodcom nomenclature at the 12-digit level. The sample is comprised of about 722 classes of products. The 14,590 items in the sample correspond to 2,279 different firms. The surveyed firms are asked on average about the price of five products. The sample is not balanced because of entries and exits of firms in the sampled population and due to the product sampling method. The

[5]Firms that provide only manufacturing services are excluded from the sample.

average number of units interviewed is 1,966 and approximately 10,379 prices are collected monthly.

The data are obtained through a monthly survey sent to firms located in Portugal and conducted in the first week of each month. Firms are asked about the transaction prices[6] of a predetermined set of products with day 15 of each month as the reference date. This allows us to construct a dataset with multiple completed price spells for the same product.

The dataset is organized at the firm-product-month level. For each firm-product combination in the sample we can observe multiple spells, with each spell corresponding to a price change. This raises an important issue because there are some observations corresponding to spells for which the first observation does not follow a price change, meaning that these observations are left-censored[7]. Because we are estimating the probability of price changes since the last price change, we follow the standard practice in the duration literature of dropping these observations. We ended with a sample comprising approximately 2,278 firms and 6,228 monthly collected prices.

The frequency of price changes is depicted in Figure 2.1. The distribution is skewed to the right, with the number of price spells concentrated at small values. The median price spell is equal to 5 and very few products register more than nine price spells over the sampling period. The average price spell is equal to 5.4. This distribution of price adjustments may be *prima facie* evidence of some nominal price rigidity in price changes.

In our sample, approximately 30 percent of price changes are price decreases and 70 percent are price increases[8]. This pattern of price changes is consistent with the hypothesis that idiosyncratic factors play an important role in explaining price dynamics since inflation in Portugal remained steadily positive over the sampling

[6]Prices are not deducted of any discounts or subsidies and taxes are not added.

[7]If not properly accounted for, this may introduce an initial-conditions problem leading the estimation of duration to be downward biased (Heckman & Singer (1986)). This means that, as in the case of the stock sampling problem well-known in the duration literature (Lancaster (1979)), longer spells are more likely to be left-censored.

[8]Nakamura & Steinsson (2008) report that price declines are very common, accounting for about 40 percent of producer price changes in the Bureau of Labor Statistics (BLS).

period.

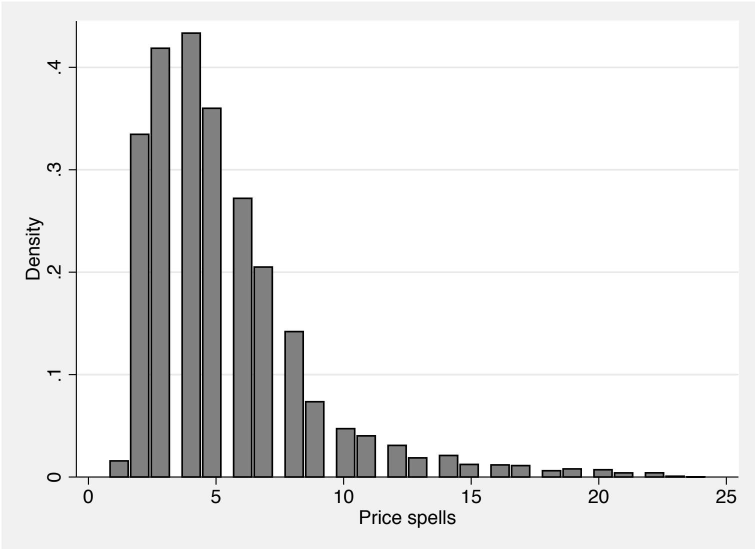

Figure 2.1: Distribution of the frequency of monthly price changes.

2.3 The survival function of price adjustment

Figures 2.2, 2.3, and 2.4 present the survival functions for price spells by inflation regime, distinguishing between positive and negative price changes. We consider three inflation regimes calculated as the average price change at the product level: negative inflation, moderate inflation, and high inflation. The threshold for high inflation is equal to 0.001, which roughly corresponds to the median inflation at the product level.

Figure 2.2 depicts the survival function for price variation (without distinguishing between price increases and price decreases) according to the inflation regime. At first glance, price behavior is considerably heterogeneous by inflation regime. Prices of products in the negative inflation regime exhibit considerably lower survival rates in the initial periods than the prices of products in the moderate and high inflation regimes. This means that prices of products are more likely to change in the case of negative inflation. The survival functions of price adjustment in the moderate and high inflation regimes reveal similar patterns.

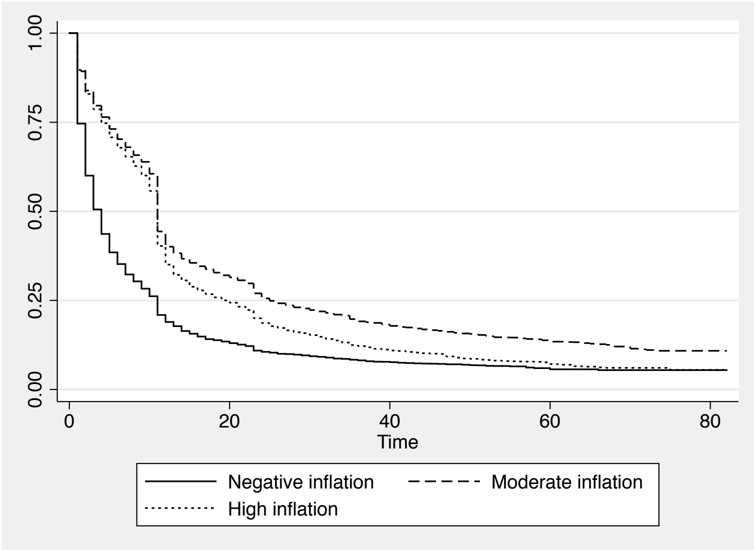

Figure 2.2: Kaplan-Meier survival function by inflation regime.
Notes: Inflation regimes are defined at the product level. For detailed data definitions see Section 3.5.

Figures 2.3 and 2.4 show the survival rates for positive and negative price changes by inflation regime, respectively. These figures reveal that the patterns of price changes are clearly different when we split price changes into price increases and price decreases. The survival function for positive price changes indicates that the likelihood of a price increase is considerably higher in periods of negative inflation up to twelve months. From this moment on, the survival rates are lower for prices in the high inflation regime, suggesting that the likelihood of price adjustment is higher for prices in this regime.

In turn, Figure 2.4 shows the survival function for negative price changes and suggests that the likelihood of a price decrease is considerably lower in the negative inflation regime. Furthermore, prices in the high and moderate inflation regimes exhibit high survival rates, which suggests that the probability of a price decrease in these regimes is quite low.

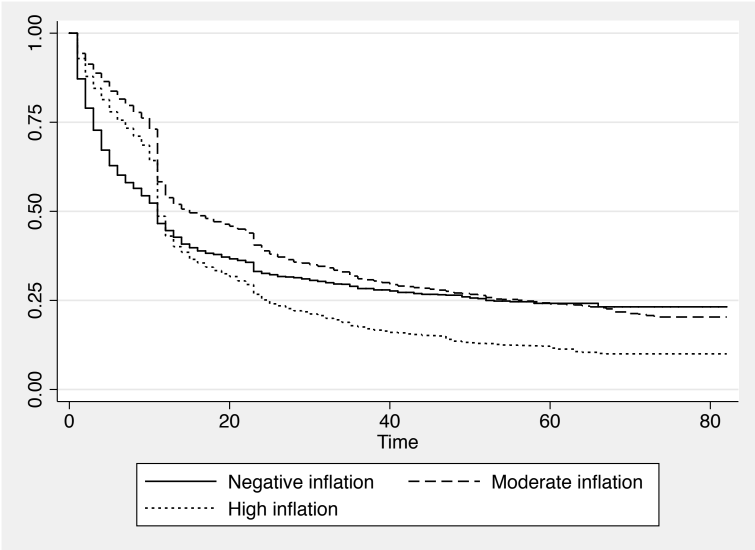

Figure 2.3: Kaplan-Meier survival function for positive price changes by inflation regime.
Notes: Survival functions calculated for products that registered a price increase the last time they changed. Inflation regimes are defined at the product level. For detailed data definitions see Section 3.5.

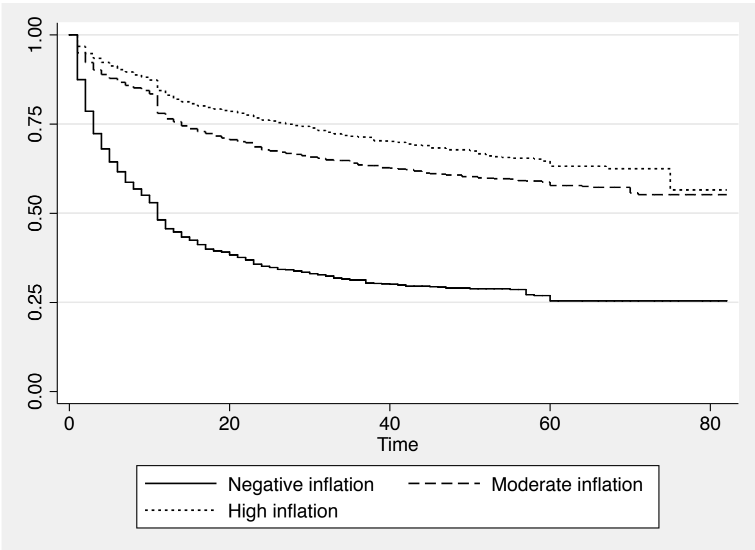

Figure 2.4: Kaplan-Meier survival function for negative price changes by inflation regime.
Notes: Survival functions calculated for products that registered a price decrease the last time they changed. Inflation regimes are defined at the product level. For detailed data definitions see Section 3.5.

Moreover, even though prices do not seem to adjust continuously, we observe a spike at twelve months in the three figures, which suggests that prices adjust at least once a year. Empirical evidence for the U.S. producer prices documents that the estimated hazard function is downward sloping for the first few months, then mostly flat except for a spike in the hazard at twelve months (Nakamura & Steinsson (2008)).

2.4 Empirical methodology

A key concept in duration analysis is the hazard function, which represents the instantaneous probability of a price change at time t, $t = 1, 2, ..., k$, conditional on survival to time t. The shape of the hazard function is meaningful to characterize price adjustment because it is upward sloping if prices that remain unchanged longer are the most likely to change, downward sloping if prices that changed recently are more likely to change again, and flat if the probability of price adjustment is constant over time.

Let time be divided into k intervals of time $[T_0, T_1)$, $[T_1, T_2)$ ... $[T_{k-1}, \infty)$[9]. We observe discrete time $T \in 1, ..., k$ and $T = t$ denotes a transition within the interval $[T_{t-1}, T_t)$. Hence, the hazard function can be defined as:

$$h(t) = \Pr(T = t | T \geqslant t), t = 1, 2, ..., k - 1, \tag{2.1}$$

and the survival function, which gives the probability of surviving in the same state up to time t is given by:

$$S(t) = \Pr(T > t) = \prod_{j=1}^{t} [1 - h(j)]. \tag{2.2}$$

We estimate the hazard function for producer prices considering a discrete time proportional hazards model. The extension of the proportional hazards model to discrete time starts from the conditional survival function $S(t|\mathbf{x}_i)$ given by:

$$S(t|\mathbf{x}_i) = S_0(t)^{\exp(\mathbf{x}_i' \beta)}, \tag{2.3}$$

[9]The empirical formulation closely follows Addison & Portugal (2008).

where $S_0(t)$ is the baseline survival function, the vector $\mathbf{x}_i$ is a set of k observed covariates of firm i, and β is a vector of coefficients. Given the relationship between the hazard and the survival functions for discrete time, the conditional hazard function for the complementary log-log model can be written as:

$$h(t|\mathbf{x}_i) = 1 - [1 - h_0(t)]^{\exp(\mathbf{x}_i'\beta)}, \tag{2.4}$$

which can be factored into a non-parametric baseline hazard function $h_0(t)$ that is a function of t alone, and a function $\exp(\mathbf{x}_i'\beta)$ that depends on $\mathbf{x}$ alone. We estimate the model by maximum likelihood.

The vector of explanatory variables $\mathbf{x}_i$ is comprised of a measure of competitive pressure calculated as the (logarithm of the) number of firms producing the same product in the same time period (log *number of competitors*), the (logarithm of the) magnitude of the last price change (log *magnitude last price change*), and an indicator variable for the case in which the last price change was a price decrease (*Last change price decrease*). We also calculate two measures of inflation defined as the logarithm of the average price change at the product level (*Product inflation*$_{t-1}$) and at the sectoral level (*Inflation*$_{t-1}$). These two variables are introduced in the model estimation lagged by one period[10]. Duration dependence is accounted for in the logarithm of price duration (log *duration*).

Empirical evidence suggests that there is considerable seasonal synchronization on price changes, meaning that a fraction of firms is likely to adjust their prices at regular intervals[11]. This evidence in favor of price rigidity as arising from seasonal

[10] Klenow & Malin (2010) document that the main determinants of the frequency of price change mentioned in the literature are the level and volatility of inflation, the frequency and magnitude of cost and demand shocks, the structure and degree of market competition, and the price collecting methods of statistical agencies (namely on whether temporary sales and product turnover are reported). Fabiani *et al.* (2005) report that the main four reasons why Eurozone firms refrain from changing prices are implicit and explicit contracts with customers, cost-based pricing (i.e. staggered input prices), and coordination failure.

[11] Nakamura & Steinsson (2008) document that producer prices in the U.S. are substantially more likely to change in the first quarter of the year and Alvarez *et al.* (2006) find the same pattern of price change for consumer prices in the euro area. Dias *et al.* (2007) use consumer prices to study the duration of price spells and find that seasonal dummies are statistically significant, suggesting that some price changes follow a regular schedule. Dias *et al.* (2004) suggest that this pattern of price change can be interpreted as an indication of time-dependency on price setting but also as a

behavior in price setting has important implications for monetary policy, because shocks in the first part of the year may yield greater effects than shocks later on. To account for seasonality we include a set of monthly dummy variables in the estimation. Moreover, we include a linear time trend in the analysis (*Time*).

2.4.1 Results

In this section we analyze the importance of idiosyncratic and sectoral characteristics in explaining the probability of a price change. The analysis of the determinants of price duration without accounting for product or firm unobserved heterogeneity is presented in Table 2.1. The estimates are reported for overall prices[12], price increases, and price decreases. The results suggest that market competition as measured by the (logarithm of the) number of competitors negatively affects the likelihood of a price change in the case of overall prices and price decreases, and contributes positively to the likelihood of a price change in the case of a price increase. This result implies that competitive pressure increases the likelihood of a price increase and decreases the likelihood of a price decrease. The likelihood of a price increase or price decrease is higher for prices that decreased the last time they changed, but the coefficient is much higher in the case of a price decrease. Furthermore, even though of small magnitude, the (logarithm of the) last price change is negatively correlated with the likelihood of a price change in all cases and, therefore, the higher the magnitude of the last price change the lower the probability of changing prices, once again.

In this model aggregate inflation and inflation calculated at the product level are not statistically significant when we analyze overall price variation. However, according to the results reported in column (2), price increases are predicted to comove with inflation, at either the sectoral or product level. Moreover, the results reported in column (3) suggest that price decreases also comove with inflation, meaning that prices are more likely to decrease when inflation decreases. This result

result of the high concentration of changes in production costs in the beginning of the year.

[12] We use the term "overall prices" to denote price variation without distinguishing between price increases and price decreases.

is especially strong in the case of sectoral inflation. These results partly contrast with the findings in Nakamura & Steinsson (2008) for the U.S., which suggest that the frequency of price increases covaries with inflation while the frequency of price decreases is largely unresponsive to inflation. Nevertheless, the results are consistent with a menu cost model influenced by idiosyncratic shocks in both studies.

Table 2.1: Determinants of price duration

	Price variation (1)	Price Increase (2)	Price Decrease (3)
log number of competitors	-0.0228***	0.0165***	-0.0917***
	(0.0048)	(0.0057)	(0.0087)
Last change price decrease	0.3393***	0.0980***	0.7524***
	(0.0124)	(0.0157)	(0.0212)
log magnitude last price change	-0.0040***	-0.0027***	-0.0126***
	(0.0005)	(0.0005)	(0.0015)
Product inflation$_{t-1}$	0.0027	0.0903***	-0.0907***
	(0.0052)	(0.0057)	(0.0061)
Inflation$_{t-1}$	-0.0236	0.1262**	-0.4851***
	(0.0484)	(0.0573)	(0.0906)
log duration	-0.3481***	-0.2266***	-0.5818***
	(0.0053)	(0.0064)	(0.0099)
Time	-0.0002	0.0003	-0.0019***
	(0.0003)	(0.0003)	(0.0005)
constant	-0.9088***	-1.5180***	-1.9459***
	(0.0247)	(0.0296)	(0.0462)
Observations	495,789	495,789	495,789

Notes: The sampling period goes from March 1995 to December 2001. Complementary log-log estimates and standard errors in parentheses. All specifications include monthly dummies. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

The estimated coefficient for duration dependence is negative, suggesting a downward hazard function of price adjustment, which implies that the probability of a price change is negatively associated with price duration, or in other words, that prices that have remained unchanged for longer periods are the ones that are less likely to make a transition. The general finding in the empirical literature of price adjustment is that hazard rates are mostly flat, except for a spike at twelve months. Nevertheless, the fact that we do not account for product- or firm-specific effects in the estimation results reported in Table 2.1 may bias the estimated pattern of duration dependence.

2.4.2 Models with frailty

A major empirical issue in the estimation of hazard rates of price changes is how to account for cross-product heterogeneity in the level of the hazard function. If not properly accounted for, cross-product heterogeneity can lead to downward-sloping hazard functions, even if the true hazard function for any individual product is upward sloping or flat (see, for example, Lancaster (1979) and Heckman & Singer (1986)). Klenow & Malin (2010) show that the frequency of price adjustment depends on the firms' technology with firms that are more labor-intensive tending to change prices less frequently (in comparison to firms with higher shares of energy and non-energy intermediate inputs)[13].

To deal with this issue we follow two estimation strategies. First, we estimate the model proposed by Lancaster (1979), which allows for multiplicative unobserved heterogeneity in the level of the hazard function. Under the proportional hazards model and assuming the exponential mean with a multiplicative error as functional form, the complementary log-log (cloglog) model specification can be written as follows:

$$\text{cloglog}\left[p(t, \mathbf{x}_i|\beta, \nu_i)\right] = D(t) + \mathbf{x}_i^{'}\beta + u_i \tag{2.5}$$

where ν_i denotes an unobserved heterogeneity term for product i, u_i is the logarithm of ν_i distributed with zero mean, and $D(t)$ characterizes the baseline hazard function. Then, prices of products with above-average values of ν change relatively quickly, meaning that, other things remaining equal, the corresponding hazard rate is higher and, consequently, their survival times are shorter. The opposite holds true for prices of products with below-average values of ν. We estimate a random intercepts

[13] Dias *et al.* (2004) use Portuguese data on consumer and producer prices and document that the frequency of price adjustment differs significantly across sectors but also across firms and stores. In particular, the frequency of price adjustment in the food sector is considerably higher than in the services sector.

model assuming that the unobserved heterogeneity term is normally distributed.

The estimation of this model requires expressions for survival and density functions that do not condition on the unobserved effects. This is because the random term ν for each individual is unobserved. One solution is to integrate out the unobserved effect (see Jenkins (2005)), by specifying a distribution for the random effect ν characterized in terms of parameters. Then, the survival function is written in terms of the latter.

A second way to deal with product and firm unobserved heterogeneity is to estimate a discrete time proportional hazards model with high dimensional fixed effects. We consider product and firm high dimensional fixed effects in the estimation of the probability of a price change in order to extensively account for (observed and unobserved) time-invariant heterogeneity at the product and firm level. The fact that we have multiple spells with completed durations for each product in the dataset allows us to account for product-level frailties in addition to firm-level frailties. The estimation procedure uses a re-iterated weighted least squares algorithm.

Results

The estimation results of the random intercepts model are presented in Table 2.2. The estimated coefficient for duration dependence is now estimated to be positive in columns (1) and (2). This means that the hazard rate is estimated to be upward sloping in the case of overall prices and price increases, which is consistent with prices that remained unchanged for longer periods of time being more likely to change. This result holds true in the model estimated with fixed effects and highlights the importance of unobserved heterogeneity in the estimation of the probability of a price change. In fact, an interesting result is that duration dependence is estimated to be negative in the fixed effects model if we do not account for firm permanent heterogeneity. Also, it is worth noting that the estimated coefficient for duration dependence is increasing in the number of fixed effects considered in the estimation.

In the case of negative price changes, the estimated coefficient for duration dependence in the random intercepts model and in the fixed effects model remains negative, which suggests that the distribution of negative price changes has a negative slope and, therefore, the longer the prices have remained unchanged the lower the probability of a price decrease in the next short period. The estimated coefficient for duration dependence is much smaller in the case of the fixed effects model. These results suggest that the frequency of price change is correlated with the age of prices, which is evidence in favor of state-dependent models in which shocks accumulate over time, and contrasts with the documented lack of relationship between the frequency and timing of price changes (see Klenow & Malin (2010)).

Table 2.2: Determinants of price duration: random effects model

	Price variation (1)	Price Increase (2)	Price Decrease (3)
log number of competitors	-0.0218**	0.0271***	-0.1587***
	(0.0110)	(0.0099)	(0.0164)
Last change price decrease	0.1234***	0.2204***	0.1112***
	(0.0156)	(0.0190)	(0.0256)
log magnitude last price change	-0.0040***	-0.0032***	-0.0092***
	(0.0006)	(0.0006)	(0.0016)
Product inflation$_{t-1}$	0.0229***	0.0891***	-0.0632***
	(0.0043)	(0.0055)	(0.0063)
Inflation$_{t-1}$	0.1035**	0.2356***	-0.4204***
	(0.0496)	(0.0581)	(0.0927)
log duration	0.1126***	0.1041***	-0.2283***
	(0.0077)	(0.0093)	(0.0128)
Time	-0.0053***	-0.0035***	-0.0051***
	(0.0003)	(0.0004)	(0.0005)
constant	-1.8608***	-2.3339***	-2.7798***
	(0.0451)	(0.0451)	(0.0701)
Observations	495,789	495,789	495,789

Notes: The sampling period goes from March 1995 to December 2001. Complementary log-log with random effects estimates and standard errors in parentheses. The likelihood ratio test suggests that frailty is statistically significant in the three models. All specifications include monthly dummies. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

According to the results reported in Table 2.3, the likelihood of price increases and price decreases comove with inflation at the product and sectoral level. Sectoral inflation is not statistically significant in the case of price variation without distinguishing between price increases and price decreases. Furthermore, the likelihood of a price decrease is now estimated to depend negatively on prices having decreased the last time they changed, which seems to be a more intuitive result than otherwise.

Table 2.3: Determinants of price duration: two high dimensional fixed effects model

	Price variation (1)	Price Increase (2)	Price Decrease (3)
log number of competitors	0.3614***	0.5862***	-0.1572*
	(0.0560)	(0.0684)	(0.0921)
Last change price decrease	0.0460***	0.1484***	-0.1094***
	(0.0166)	(0.0215)	(0.0243)
log magnitude last price change	-0.0021**	-0.0022**	-0.0005
	(0.0008)	(0.0009)	(0.0014)
Product inflation$_{t-1}$	0.0171***	0.0678***	-0.0503***
	(0.0050)	(0.0062)	(0.0059)
Inflation$_{t-1}$	0.0782	0.2386***	-0.4705***
	(0.0512)	(0.0587)	(0.0950)
log duration	0.1764***	0.2906***	-0.0590***
	(0.0074)	(0.0090)	(0.0128)
Time	-0.0052***	-0.0051***	-0.0051***
	(0.0003)	(0.0004)	(0.0005)
Observations	469,365	457,334	353,481
Product fixed effects	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes

Notes: The sampling period goes from March 1995 to December 2001. Complementary log-log estimation with two high dimensional fixed effects. Standard errors in parentheses. All specifications include monthly dummies. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

A notable result is the greater magnitude of the estimated coefficient for product competition in the high dimensional fixed effects model for price increases when compared to the random intercepts model. Moreover, the (logarithm of the) number of a firm's competitors is now estimated to positively affect the likelihood of a price change in the case of overall price variation. These results seem to be partly driven by product permanent heterogeneity since they are estimated to be considerably smaller when we do not consider the product fixed effect in the estimation.

Table 2.4 presents fixed effects estimates for the determinants of price duration considering the interaction between firm and product fixed effects. In the previous case, product fixed effects were introduced in the model estimation to account for product-specific heterogeneity while firm fixed effects captured the price strategy of the firm. In this case, the interaction term between product and firm fixed effects accounts for the heterogeneity in the price policy of the firm for each product produced by the firm. The results are mostly robust to this model specification. Nevertheless, the estimates suggest that duration dependence in the case of a price decrease is estimated to be positive, meaning that the likelihood of a price decrease increases with the elapsed price duration.

Table 2.4: Determinants of price duration: high dimensional fixed effects model

	Price variation (1)	Price Increase (2)	Price Decrease (3)
log number of competitors	0.4529***	0.6684***	-0.1636
	(0.0620)	(0.0759)	(0.1038)
Last change price decrease	0.0030	0.2945***	-0.3484***
	(0.0195)	(0.0264)	(0.0268)
log magnitude last price change	-0.0037***	-0.0043***	0.0008
	(0.0010)	(0.0012)	(0.0020)
Product inflation$_{t-1}$	0.0258***	0.0914***	-0.0660***
	(0.0050)	(0.0067)	(0.0065)
Inflation$_{t-1}$	0.0950*	0.2424***	-0.4889***
	(0.0521)	(0.0599)	(0.0964)
log duration	0.3911***	0.5094***	0.1217***
	(0.0091)	(0.0112)	(0.0153)
Time	-0.0083***	-0.0077***	-0.0078***
	(0.0004)	(0.0004)	(0.0006)
Observations	425,751	399,650	208,793
Product×Firm fixed effects	Yes	Yes	Yes

Notes: The sampling period goes from March 1995 to December 2001. Complementary log-log estimation with high dimensional fixed effects. Standard errors in parentheses. All specifications include monthly dummies. ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

Overall, these results suggest that price setting of manufacturing firms depends on idiosyncratic and sectoral shocks, and that price adjustment is considerably heterogeneous across products and firms. Furthermore, our results show that when product- and firm-specific heterogeneity is properly accounted for in the estimation, the hazard rate of price adjustment is predicted to be upward-sloping in the case of both positive and negative price changes.

2.5 Peer effects

In this section we analyze the presence of coordination failures in price setting. In particular, we estimate how a firm price setting behavior responds to its competitors' price setting decisions. In fact, the estimation results reported in Tables 2.3 and 2.4 suggest that competitive pressure at the product level is an important determinant of the duration of price spells for Portuguese manufacturing firms, but little or nothing is known on how these firms react to their competitors' price setting behavior.

We consider a peer-effects model that allows us to estimate how firm prices adjust to competitors' price changes.

2.5.1 Empirical methodology

The estimation of peer effects building on individual behavior as a function of group averages can be challenging. Angrist (2014) documents the main sources of estimation bias in models of peer effects and notes the importance of *controlling ourselves* in *y on y-bar* regressions[14]. In this study we follow Arcidiacono *et al.* (2012) and assume that the importance of own characteristics is proportional to that of competitors' characteristics. Then, the model specification to estimate the peer effect of competitors' price setting behavior is given by:

$$y_{ift} = \rho y_{ift-1} + \mathbf{x}_{ft}\beta + \alpha_{if} + \delta\bar{\alpha}_{ift} + \mu_t + \varepsilon_{ift} \tag{2.6}$$

where y_{ift} is the logarithm of the price of product i produced by firm f at month t and y_{ift-1} is an autoregressive term; $\mathbf{x}_{ft}$ comprises the (logarithm of the) price of each firm competitors at each time period (log *competitors prices*), two indices of price variation calculated at the product and sectoral level lagged by one period (*Index product prices*$_{t-1}$ and *Index prices*$_{t-1}$, respectively), and the (logarithm of the) number of the firm's competitors; α_{if} is a time-invariant firm-product fixed

[14]Other sources of estimation bias in peer effects models are selection bias, measurement error, nonlinear effects, and the common shocks problem.

effect; $\bar{\alpha}_{ift}$ is the mean of the fixed effects of the competitors of firm f at time t; μ_t is a vector of time dummies; and ε_{ift} is a zero mean disturbance term capturing all other omitted factors. The competitors of firm f are defined as all firms that, in a given month, produce the same product as firm f and, therefore, exhibits some time-variability. The parameter δ is the one of main interest and measures the peer effect.

The term $\bar{\alpha}_{ift}$ measures the average price of a firm's competitors and is identified even in the presence of the product-firm fixed effect (α_{if}) because the composition of a firm's competitors changes over time.

The model specification formulated in (2.6) can be estimated by ordinary least squares even though its solution is nonlinear. This happens because the fixed effect for a given firm-product combination enters directly in some observations and indirectly in other observations via prices of competitors. Arcidiacono *et al.* (2012) identify the set of assumptions under which the least squares solution provides consistent estimates for the parameters of the model presented in equation (2.6). Then, the model can be estimated using the algorithm proposed by Portugal & Guimarães (2010) to estimate models with high dimensional fixed effects.

2.5.2 Results

The estimation results of the peer-effects model are reported in Table 2.5. These results suggest that a firm's prices depend positively on its competitors' prices, meaning that firms increase prices when the average price of their competitors also increase. According to the estimates, firm's prices are predicted to increase by about 0.0318 percent when competitors' prices increase by 1 percent, *ceteris paribus*[15]. The small magnitude of the estimated peer effect seems to be consistent with the incomplete price adjustment that arises from coordination failures across firms, although it is partly explained by the fact that the set of competitors is quite constant over time.

[15]Despite the high magnitude of the autoregressive term, the Im-Pesaran-Shin unit root test leads to the rejection of the hypothesis of panel non-stationarity.

Furthermore, the estimation results suggest that prices are predicted to respond positively to product and sectoral inflation. The long-run effects of these two variables are non-negligible even though the magnitude of the short-run effects seems to be small.

It is important to note that even though the estimation is performed considering the dependent and explanatory variables in levels, the fact that we are considering the interaction term between firm and product fixed effects in the estimation leads to a similar interpretation of a first-differences model from a time series perspective.

All in all, the estimated peer effect in producer price setting is positive and statistically significant even though of small magnitude. This tallies with the presence of coordination failures in price setting behavior across firms.

Table 2.5: Price adjustment: peer-effects model

	log prices (1)
log prices$_{t-1}$	0.9802***
	(0.0002)
log competitors prices	0.0318***
	(0.0026)
Index product prices$_{t-1}$	0.0104***
	(0.0004)
Index prices$_{t-1}$	0.0061***
	(0.0007)
log number of competitors	0.0004***
	(0.0000)
Observations	534,735
Product×Firm fixed effects	Yes

Notes: The sampling period goes from January 1995 to December 2001. Interacted high dimensional fixed effects estimation with standard errors in parentheses. *** stands for statistical significance at 1%.

2.6 Conclusions

A central hypothesis for the large effects of demand shocks and the non-neutrality of monetary policy on real output is the sluggish price adjustment to aggregate conditions.

The theoretical literature on price rigidity can be broadly summarized in three lines of research: time-dependent models, state-dependent models, and imperfect information models. Time-dependent models assume that firms adjust their prices at an exogenously set timing while state-dependent models presume that firms decide to change prices according to the state of the economy at each moment. Last, Lucas' (1972) imperfect information model builds on the idea of monetary non-neutrality as arising from imperfect information about the state of the economy. A key ingredient for sluggish price adjustment in most models is coordination failure, which leads firms to adjust incompletely to aggregate shocks because other firms have not yet responded.

Our study contributes to the existing empirical literature on the micro-foundations of aggregate price setting in two different ways. First, to the best of our knowledge, we are the first to estimate a duration model to estimate price transitions with high dimensional fixed effects, which allows us to control for (observed and unobserved) time-invariant heterogeneity at the product and firm level; and second, we provide direct evidence on the impact of market competition on firms' price setting rules by estimating peer effects in a price setting model.

We conduct the analysis in two steps: first, we estimate a duration model for price adjustment in which we extensively account for permanent product- and firm-specific heterogeneity; then we estimate a peer effects model to document how firms react to the price setting behavior of their competitors.

The results suggest that competitive pressure increases the likelihood of a price increase and decreases the likelihood of a price decrease. Moreover, the probability of a price increase or price decrease is estimated to comove with inflation, at both product and sectoral level. A notable result is that when we fully account for time-invariant (observed and unobserved) heterogeneity, duration dependence is estimated to be positive in the case of both a price increase and price decrease, which suggests that prices that have remained unchanged longer are more likely to change in the next short period. Also, when we account for product and firm heterogeneity, prices are estimated to be more likely to increase and less likely to decrease if prices decreased the last time they changed. Overall, these results suggest that price setting depends on both idiosyncratic and sectoral shocks, and that price adjustment is considerably heterogeneous across products and firms.

Furthermore, we estimate that firms' price setting behavior responds to the price setting behavior of their competitors, even though the small magnitude of the estimated peer effect suggests that firms respond incompletely and sluggishly to their competitors' price setting decisions. This result aligns with the presence of coordination failures in price adjustment among price setters.

Chapter 3

Credit Rationing for Portuguese SMEs

3.1 Introduction

In the years that preceded the recent crisis, the Portuguese non-financial corporations accumulated high levels of debt, with the ratio between the non-financial corporate sector debt and GDP registering values higher than 130 percent[1] in 2008. This evolution of credit was induced by favorable financing conditions and optimistic expectations of productivity growth that did not materialize, highlighting the need for an adjustment process. The global financial crisis and the subsequent sovereign debt crisis led to an adjustment process characterized by a significant contraction of the economic activity and worse prospects of economic agents. Furthermore, this adjustment involved a bank lending channel with the Portuguese banks being severely affected by international financing restrictions and stronger capital requirements. In this context, the adjustment process comprised a simultaneous contraction in the demand and supply of credit. According to the Bank Lending Survey (BLS), the observed credit evolution during the economic and financial crisis was a result of increased restrictiveness in credit standards and conditions applied on loans as well

[1]This value refers to the ratio between the non-financial corporate sector debt and GDP in consolidated terms, calculated according to ESA 1995.

as of decreased loan demand by firms. Nevertheless, it is a difficult task to rigorously evaluate the relative contribution of each dimension[2].

The availability of microeconomic datasets with a maximum level of granularity has been crucial for the development of new approaches to the identification of credit rationing of non-financial corporations. Antunes & Martinho (2012) estimated a model to explain the evolution of credit granted by Portuguese banks to non-financial corporations and to examine the eventual presence of credit restrictions, using data for the period between the first quarter of 1995 and the first quarter of 2012. Even though this analysis does not allow to unequivocally identify the relative contribution of credit demand and credit supply to explain the evolution of credit, the results suggest that the access to credit by Portuguese firms became more difficult after 2009 and that credit restrictions were particularly relevant for firms seeking credit for the first time. Farinha & Félix (2014) considered an econometric methodology similar to that in Rottmann & Wollmershäuser (2013) and estimated a two-step model to evaluate the importance of credit demand and credit supply factors in explaining the evolution of credit in the period between the first quarter of 1997 and the second quarter of 2013. The results suggest that after controlling for firm idiosyncratic characteristics, the evolution of credit is largely explained by banks liquidity and solvency conditions. The fact that banks react in a significant way to changes in their liquidity structure and to stricter minimum capital requirements, with banks having less access to funding and smaller capital buffers granting less credit to firms, suggests the presence of a bank lending channel through which banks can affect firms financing conditions (assuming that at least a fraction of firms does not have easy access to alternative sources of financing)[3]. Nevertheless, this analysis is only indicative of the presence of credit restrictions in the Portuguese economy.

The disequilibrium models, originally developed to evaluate the presence of credit rationing at the macroeconomic level (see, for example, Laffont & Garcia (1977)),

[2]An analysis of this topic using the BLS data can be found in the Box 1.2.1 "Decomposing credit growth on the basis of the Bank Lending Survey", published in the Financial Stability Report of Banco de Portugal, November 2013. The authors construct indices based on the qualitative answers of bank officials regarding the developments in the credit market and include these indices as explanatory variables in a model to explain credit growth.

[3]This result contributes to the discussion on the bank lending channel and bank capital channel.

have been used more recently in the empirical literature to identify credit restrictions based on microeconomic data. These models assume that the credit market may be in disequilibrium and therefore, the observed interest rate does not ensure that credit demand is equal to credit supply. Ogawa & Suzuki (2000), Atanasova & Wilson (2004), Shikimi (2005) and Carbo-Valverde *et al.* (2009) estimated disequilibrium models to identify the presence of credit rationing in the economy. More recently, Kremp & Sevestre (2013) considered this methodology to analyse the determinants of credit demand and credit supply for French small and medium-sized enterprises (SMEs) and found that, even though banks adopted tighter standards in credit approval, French firms were not significantly affected by credit rationing, even after 2008.

In this study we assess the importance of credit demand and credit supply-related factors in explaining the evolution of credit granted to Portuguese SMEs and evaluate whether this evolution stems from a more restrictive credit policy in the period between 2010 and 2012. We restrict our sample to SMEs since they are expected to rely more on bank loans and to have less access to external sources of financing, namely financing by non-resident banks. Therefore, it is expected that small and medium-sized enterprises (SMEs) are potentially the most affected by credit restrictions. For that purpose, we consider a methodology that allows to simultaneously estimate credit supply and credit demand functions, assuming that the credit market may be in disequilibrium, as modeled in Kremp & Sevestre (2013).

The results suggest that credit supply depends positively on the amount of assets available to be provided as collateral and negatively on firms' indebtedness. In turn, credit demand depends negatively on the interest rate and on the amount of firms' internal resources, and positively on their working capital needs. We estimate that approximately 15 percent of Portuguese SMEs with bank loans were partially rationed and 32 percent of firms with no bank loans were fully rationed.

The paper is organized as follows: in the next section we present the model and the variables considered in the analysis, and in Section 3.3 we discuss some econometric questions. In Section 3.4 we present the methodology to evaluate the intensive and extensive margins of credit rationing. The data and main descriptive statistics are presented in Section 3.5. In Section 3.6 we discuss the main results and Section 3.7 concludes.

3.2 The demand and supply of credit: model and variables

3.2.1 The disequilibrium model

A disequilibrium model comprehends essentially three equations: one for credit demand, one for credit supply, and one additional equation that links credit demand to credit supply. We closely follow Kremp & Sevestre (2013) to present the disequilibrium model in the case of the credit market:

- An equation for the demand of new loans, $NL_d^\star$: $NL_d^\star = f_d(X_d b_d; u_d)$;
- An equation for the supply of new loans, $NL_s^\star$: $NL_s^\star = f_s(X_s b_s; u_s)$;

  where X_d and X_s represent the vectors of explanatory variables that affect credit demand and credit supply, respectively, and b_d and b_s are the vectors of coefficients associated with those variables. The terms u_d and u_s with zero mean and variance σ_d^2 and σ_s^2, respectively, represent the unobservable factors that affect credit demand and credit supply, respectively, and it is assumed that they may be correlated with each other (ρ);
- And an equation that links the observed quantity of credit with the unobservable credit demand and supply: $NL = \min(NL_d^\star, NL_s^\star)$. This equation assumes that the observed quantity of credit is the minimum between credit demand and credit supply.

This system of equations is estimated using the maximum likelihood estimator (see Maddala & Nelson (1974)).

When the demand and supply of credit are strictly positive and the interest rate is observed, the contribution to the maximum likelihood function is given by:

$$I_{it} = f_d(NL_{it})(1 - F_s(NL_{it})) + f_s(NL_{it})(1 - F_d(NL_{it})), \tag{3.1}$$

where $f_d(NL_{it}) = \frac{1}{\sigma_d\sqrt{2\pi}} \exp\left[-\frac{1}{2\sigma_d^2}(NL_{it} - X_{d,it}b_d)^2 \right]$ is the density function of new loans if the demand is observed;

$F_d = \Phi\left(\frac{NL_{it} - X_{d,it}b_d - \rho(\sigma_d/\sigma_s)(NL_{it} - X_{s,it}b_s)}{\sigma_d^2\sqrt{1-\rho^2}} \right)$ is the corresponding cumulative distribution function which considers the possibility that credit demand and credit supply may be correlated;

$f_s(NL_{it}) = \frac{1}{\sigma_s\sqrt{2\pi}} \exp\left[-\frac{1}{2\sigma_s^2}(NL_{it} - X_{s,it}b_s)^2 \right]$ is the density function of new loans if the supply is observed;

$F_s = \Phi\left(\frac{NL_{it} - X_{s,it}b_s - \rho(\sigma_s/\sigma_d)(NL_{it} - X_{d,it}b_d)}{\sigma_s^2\sqrt{1-\rho^2}} \right)$ is the corresponding cumulative distribution function.

3.2.2 Variables

In this section we present the variables included in the model. We consider that the demand for new loans depends on the following economic factors:

Firm activity:

- Firm size - smaller firms are expected to have more difficulties in accessing credit from sources other than bank loans and, simultaneously, are the most affected by credit restrictions due to their financial structure and their higher exposure to information asymmetries. This paper voluntarily restricts the analysis to SMEs. We consider three size categories: very small firms, small firms, and medium-sized firms[4].

- Short-term financing needs - the ratio between the firm's working capital needs and total assets is included in the analysis as a measure of the firm's short-term financing needs.

[4] This classification follows the European classification of SMEs.

- Long-term financing needs - the investment, calculated as the ratio between the first differences of tangible assets and total assets, is considered as a measure of long-term financing needs.

Availability of substitute funds on the demand for bank loans (internal or external funds):

- Amount of internal resources - the pecking order hypothesis says that firms with higher capacity to finance their activity through internal resources are less likely to seek bank loans since the latter have a higher cost, in particular for smaller firms that are expected to be more exposed to information asymmetries; nevertheless, firms may decide to increase their debt if the advantages associated with debt exceed the costs (namely, firms may benefit from fiscal advantages); the considered measure for the amount of internal resources is the ratio between the firm's operational results (EBITDA)[5] and total assets.
- Other sources of external financing - the non-financial debt that, in the case of SMEs, is largely composed by loans granted by shareholders and accounts payable, is an alternative source of financing less costly than bank loans; therefore, we include in the demand function the difference between firm's total debt and debt to credit institutions, and the accounts payable, measured in terms of total assets.

Cost of bank credit:

- Interest rate - it is included in the demand function as a measure of the firm's financing costs and is calculated by dividing the interest expenses by the firm's total debt.

Time and sectoral dummy variables - time dummies are included in the demand equation to account for macroeconomic shocks that affect firms' credit demand and

[5]The operational result is considered as a proxy variable for the amount of internal resources. In fact, this variable can be interpreted as a measure of the potential amount of internal resources that the firm may decide to retain or distribute as dividends.

sectoral dummies are considered to capture sector systematic differences in credit demand.

In turn, the following determinants are considered in the equation of credit supply:

Firm-specific risk:

- Firm size - this variable reflects the credit risk associated with a given firm since firms' survival probability is lower for small firms and the risk to go bankrupt is higher for these firms when compared to larger firms.
- Age of the firm - the default risk is higher for younger firms (see, for example, Fougère *et al.* (2012)). The analysis includes an indicator variable equal to one for firms less than or five years old, and zero for older firms.
- Ratio between financial debt and operational results - measures the capacity of the firm to generate cash-flows to pay its debt and is also a measure of credit risk: it is evaluated by the ratio between the firm's total debt and operational results (EBITDA), lagged by one period.

Firm collateral:

- Collateral - the ability to provide collateral when negotiating a bank loan limits the losses to be faced by banks in the case of firm's bankruptcy; the indicator used in the analysis is the ratio between tangible assets and total assets.

We follow Ogawa & Suzuki (2000), Atanasova & Wilson (2004), Shikimi (2005), and Kremp & Sevestre (2013), and do not consider the interest rate in the supply function. In fact, it is considered that in a context of credit restrictions, banks first decide on the amount they are willing to lend and then bargain the interest rate with firms. This is founded on the theoretical model of Stiglitz & Weiss (1981) that predicts that in the presence of credit market frictions, banks may decide not to grant credit to borrowers with higher probability of default, even if they are willing to pay the respective risk premium. The decision of granting credit to a firm is determined by variables other than the interest rate, namely the firm idiosyncratic characteristics, as suggested in the credit channel literature.

Finally, we also consider time and sectoral dummy variables to account for common shocks in the supply of credit and sectoral idiosyncratic characteristics constant over time.

3.3 Some econometric questions

3.3.1 New loans

Using firm's balance-sheet information to compute the amount of bank loans does not allow to immediately identify new loans. In this paper, we follow the same econometric strategy considered in Kremp & Sevestre (2013) and include in the analysis a set of lagged terms of the dependent variable, calculated within the firms' sector of activity and dimension category, imposing that the estimated coefficient is the same in both the credit demand and credit supply equations. This is equivalent to assume the relation $NL_{it} = L_{it} - (1 - \delta_g)L_{it-1}$, where δ_g is a reimbursement rate assumed to be constant over time and common to firms within a given sector of activity and dimension category.

3.3.2 Endogeneity and unobservable heterogeneity

Some of the variables included in the econometric model can raise endogeneity problems. More precisely, we assume that the amount of collateral in the supply equation, and investment, working capital needs, and the interest rate in the demand equation, may be endogenous variables[6]. In fact, the firm's amount of investment and collateral may depend on the amount of the loan it was granted; likewise the working capital needs and the accounts payable are likely to depend on the firm's

[6] The firm's amount of collateral, investment, working capital needs, and accounts payable are measured in terms of total assets.

activity levels that, in turn, may be affected by the firm's access to bank loans. On the other hand, the interest rate may also be endogenous because firms negotiate with banks the loan amount as well as the price. To limit endogeneity problems in the model estimation we follow the strategy used in Kremp & Sevestre (2013) and developed by Rivers & Vuong (1988), which consists in the introduction of the residuals of the estimation of each endogenous variable on a set of instruments (first differences of the explanatory variables lagged one and two periods) in the equations of credit demand and supply.

The firms' unobservable characteristics that may influence the demand and supply of credit, and that are potentially correlated with the explanatory variables may also raise an estimation issue. To account for the bias that omitted variables may induce in the estimation, we include the sample averages of the explanatory variables as well as the first observation of the dependent variable in the equations of demand and supply of new loans.

3.4 Credit rationing

The estimation of the equations for credit demand and supply allows us to calculate in what extent Portuguese SMEs faced credit restrictions in the sampling period. Based on these estimates, it is possible to identify which firms obtained a loan but in a lower amount than that they applied for - intensive margin - and in what extent firms that did not get any loan were fully rationed - extensive margin.

We follow Gersovitz (1980) and define the probability of credit rationing as the probability that the latent credit demand is higher that the supply of credit, conditional on the observed amount of loans:

$$\Pr(\text{Rationing}|NL_{it}) = \frac{f_s(NL_{it})(1 - F_d(NL_{it}))}{f_d(NL_{it})(1 - F_s(NL_{it})) + f_s(NL_{it})(1 - F_d(NL_{it}))}. \quad (3.2)$$

We consider that a firm was affected by credit restrictions if this probability is greater than 0.5[7].

[7]We considered different thresholds in a reasonable neighborhood of 0.5 and the results are

To assess the extensive margin of credit rationing it is necessary to estimate an interest rate for those firms with no bank loans. In the case of firms with no bank loans in a given year but that have a positive amount of bank loans in other years, the interest rate is calculated as follows:

$$\widehat{i}_{it} = i_{it-1} \times (1 + \Delta r_{jt}), \tag{3.3}$$

where $\widehat{i}_{it}$ is the estimated interest rate for firm i in year t, i_{it-1} is the last observation of the firm's interest rate, and Δr_{jt} is the average growth of the interest rate between $t-1$ and t, calculated within activity sector and size dimension.

In the case of firms with no bank loans in the sampling period, the estimated interest rate is the average interest rate calculated by size category and sector of activity.

3.5 Data

The variables included in the analysis were computed using the Portuguese dataset Simplified Corporate Information (Portuguese acronym for *Informação Empresarial Simplificada* - IES), which comprehends detailed balance-sheet information on the universe of Portuguese non-financial corporations. This dataset is also the data source for bank loans, and therefore it is not possible to decompose the total amount of bank loans by bank. Consequently, the specification we developed for credit supply does not take into consideration bank specific factors such as banks solvency and liquidity positions. The IES information is available for the period between 2005 and 2012. We consider a sample of Portuguese SMEs because these are expected to be the most exposed to credit restrictions. The activity sectors included in the analysis comprise the manufacturing industry, construction and real estate, trade, and the different segments of services.

The relevant period for the estimation starts in 2010 because we impose that firms stay at least for four consecutive years in the sample. This restriction is

qualitatively the same.

justified by the inclusion of the first differences and the first and second lags of the first differences in the estimation as instruments for the endogenous variables. We considered the observations above the 99th percentile or below the 1st percentile as outliers.

The sample includes 50,020 observations with positive bank loans and observed interest rate. The main descriptive statistics associated with the variables considered in the analysis are reported in Table 3.1. The sample is mostly constituted by SMEs that belong to the trade and manufacturing industry sectors. Moreover, the disinvestment of Portuguese SMEs is particularly evident over the sample period. It is also possible to find that Portuguese SMEs mostly rely on bank loans and accounts payable as financing sources of their activity. The firms' non-financial debt corresponds only to a small part of their total assets.

Table 3.1: Main descriptive statistics

	Mean	St. Dev.	Q1	Q2	Q3
Total assets	1,938,798	11,900,000	231,248.1	558,436.7	1,456,533
Loans/Assets	0.248	0.170	0.110	0.222	0.358
EBITDA/Assets	0.059	0.133	0.036	0.068	0.113
Interest rate	0.051	0.045	0.023	0.041	0.065
Investment/Assets	-0.014	0.078	-0.038	-0.014	0.001
Collateral/Assets	0.245	0.214	0.069	0.186	0.370
Non-financial debt/Assets	0.054	0.112	0.000	0.001	0.056
$(\text{Debt/EBITDA})_{t-1}$	2.857	7.994	0.846	2.487	5.098
Accounts payable/Assets	0.217	0.175	0.083	0.180	0.311
Working capital/Assets	0.299	0.257	0.117	0.293	0.480
Very small firms	0.098				
Small firms	0.432				
Medium firms	0.470				
Young SMEs	0.023				
Dummy Construction	0.147				
Dummy Trade	0.423				
Dummy Hotel and Restaurant	0.018				
Dummy Info. and Communication	0.020				
Dummy Services	0.145				
Dummy Manufacturing industry	0.247				

Notes: Number of observations: 52,020 (17,104 enterprises in 2010, 17,752 enterprises in 2011 e 17,164 enterprises in 2012). Q1 and Q3 refer to the first and third quartiles, and Q2 refers to the median.

3.6 Estimation results

3.6.1 Disequilibrium model

The main estimation results of the equations of demand and supply of new loans are presented in Table 3.2[8]. The coefficients reported in the first column were obtained assuming that credit demand and supply are independent (this corresponds to the $\rho = 0$ case, meaning that shocks in credit demand are not correlated with shocks in credit supply) and those reported in the third column were obtained assuming that credit demand and supply may be correlated (and, therefore, $\rho \neq 0$). The coefficients are mostly statistically significant and have the expected sign, with the exception of firm size in the credit supply equation.

Table 3.2: Estimation results

Supply equation	Disequilibrium model Demand and supply are independent		Disequilibrium model Demand and supply are correlated	
	Coeff.	St. Error	Coeff.	St. Error
Very small firms	0.0646***	0.0080	0.0346***	0.0051
Small firms	0.0384***	0.0081	0.0217***	0.0052
$(\text{Debt/EBITDA})_{t-1}$	-0.0017***	0.0005	-0.0013***	0.0003
Collateral/Assets	1.6622***	0.0846	1.0843***	0.0549
Young SMEs	-0.0259	0.0200	-0.0108	0.0130
Dummy 2010	-0.0880***	0.0080	-0.0556***	0.0054
Dummy 2011	0.0060	0.0090	-0.0045	0.0061
sigma_s	0.1699***	0.0030	0.1274***	0.0019

[8]The remaining estimated coefficients are reported in Tables A1 and A2 in the Appendix.

Demand equation	Disequilibrium model Demand and supply are independent		Disequilibrium model Demand and supply are correlated	
	Coeff.	St. Error	Coeff.	St. Error
Very small firms	0.0038	0.0025	0.0051*	0.0026
Small firms	-0.0012***	0.0019	-0.0017	0.0020
Non-financial debt/Assets	-0.5458***	0.0115	-0.5615***	0.0122
EBITDA/Assets	-0.2893***	0.0077	-0.2988***	0.0084
Interest rate	-0.7092***	0.0657	-0.7004***	0.0712
Accounts payable/Assets	-0.0401**	0.0161	-0.0468***	0.0174
Working capital/Assets	0.0635***	0.0098	0.0628***	0.0106
Investment/Assets	0.0099	0.0366	0.0211	0.0384
Dummy 2010	0.0661***	0.0019	0.0771***	0.0024
Dummy 2011	-0.0044***	0.0013	-0.0048***	0.0014
sigma_d	0.1009***	0.0004	0.0994***	0.0004
rho			0.4496***	0.0280
log-likelihood	-45,959.53		-46,085.26	
No. of observations	52,020		52,020	

Notes: ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

The results of the estimation of the credit supply equation in both models suggest that the credit ceiling is higher for firms of smaller dimension, controlling for the other firm characteristics. The estimated coefficient for the ratio between debt and operational results indicate that it is more difficult for indebted firms to access bank loans. In fact, firms with a stronger relative weight of debt have less capacity to generate cash-flows to reimburse the amount they have borrowed and have, therefore, a higher risk of default. The availability of collateral is particularly important in the banks' decision to grant credit to firms. Firms that are able to provide better guarantees when negotiating a bank loan are able to borrow significantly more from banks. In turn, the results of the credit demand function estimation suggest that credit demand strongly depends on the interest rate. The negative effect of the amount of internal resources, as measured by the firm's operational result, and the amount of external resources (loans granted by shareholders and accounts payable) is in line with the hypothesis that due to information asymmetries, firms prefer other financing sources than bank loans (pecking order hypothesis).

The coefficient associated with investment is not statistically significant in both specifications. On the other hand, working capital needs affect credit demand. These two results may be related to some delay between granting credit and investment but also suggest that firms have demanded credit mainly to finance their operational activities rather than for investment projects. Moreover, this result indicates that the slowdown in firms' investment during the financial and economic crisis in Portugal does not seem to be associated with credit restrictions, but with lower credit demand by firms, perhaps justified by the uncertainty about the future macroeconomic scenario.

Finally, in the model where credit demand and credit supply may be correlated the estimates suggest that smaller firms are those that most seek bank loans. This result is commonly found in the literature: smaller firms are not able to access alternative financing sources and tend to rely more on bank loans (see, for example, Gertler & Leahy (2008)).

3.6.2 Credit rationing

Table 3.3 presents the estimates associated with the probability that the latent credit demand is higher than the credit supply, leading to the presence of credit restrictions. The results for the intensive margin are reported in the first column and represent the percentage of firms that (with probability greater than 0.5) obtained a loan but in a lower amount than that they applied for, given the interest rate. The results of full rationing are reported in the second column - extensive margin. According to the results reported in the first column, approximately 15 percent of SMEs with a bank loan were partially rationed. The firms that are mostly affected by credit rationing are the younger ones, with approximately 26 percent, and the smaller ones, with approximately 19 percent. The most affected activity sectors are construction and trade (21 percent and 16 percent, respectively). The estimates for the extensive margin are reported in the second column of Table 3.3. The figures suggest that about 32 percent of Portuguese SMEs were affected by credit restrictions and were not granted a loan even though their credit latent demand was positive. Smaller and younger firms are the most affected, but the estimates for the remaining dimension categories indicate that those were also considerably affected, with approximately 23 percent of medium firms facing credit restrictions. All the activity sectors faced credit restrictions, with estimates between 31 and 41 percent.

Table 3.3: Estimates (probability of credit rationing)

	Partial (in % of firms with loans)	Full (in % of firms with no loan)
SMEs	15%	32%
Very small firms	19%	35%
Small firms	13%	27%
Medium firms	9%	23%
Young SMEs	26%	29%
Manufacturing	12%	32%
Construction	21%	31%
Trade	16%	32%
Hotel and restaurant	9%	41%
Information and Communication	11%	32%
Other services	14%	32%
2010	41%	59%
2011	3%	19%
2012	3%	22%

These findings contrast with the results presented by Kremp & Sevestre (2013) for French SMEs. The main reason for this finding is twofold: on one hand, the annual rate of change of total credit to Portuguese SMEs fell around 10 per cent by the end of 2008 to approximately -5 per cent in mid-2012[9]. Furthermore, the number of new firms with access to credit decreased by more than half in the period between 2010 and 2012, according to the Portuguese Credit Register dataset, which suggests that a considerable number of firms did not have access to credit in this period. In fact, Portuguese banks faced increased deleveraging pressures and stricter minimum capital requirements in the context of the Financial and Economic Assistance Program, contributing to more restrictive credit standards and conditions applied on loans. On the other hand, Portuguese SMEs seem to rely more on bank loans as source of financing than French SMEs (as measured by the ratio between bank loans and total assets, which is equal to 0.27 and 0.18 for the sample of Portuguese and French SMEs in 2010, respectively). The results presented for partial and full rationing of Portuguese SMEs are in line with the results on firms' access to bank loans of the ECB's Survey on Access to Finance of Enterprises (SAFE).

[9]See Financial Stability Report, May 2013, Banco de Portugal.

3.7 Conclusions

This study aims to identify the relative contribution of credit demand and supply factors in explaining the evolution of new loans granted to Portuguese SMEs. We voluntarily restrict the analysis to SMEs because we consider that larger firms have easier access to financing sources other than loans granted by resident banks and are, therefore, less exposed to credit restrictions. The methodology relies on the estimation of credit demand and credit supply functions simultaneously, assuming that the credit market may be in disequilibrium and, therefore, the interest rate does not lead to equilibrium in the demand and supply of credit. This methodology is particularly suitable since credit markets may be affected by information asymmetries between debtors and creditors. The results suggest that during the financial crisis Portuguese SMEs searched for bank loans mainly to finance their operational activity and not for investment. The smaller firms and those with smaller amounts of internal resources have higher demand for bank loans. In turn, firms with a higher capacity to pay their debt and with more collateral can borrow more from banks. The estimates relative to the intensive and extensive margins suggest that a significant fraction of Portuguese SMEs was affected by credit rationing. Furthermore, the results suggest that smaller and younger firms were the most credit constrained.

Appendix

Table A1 Estimation results

Other estimates	Disequilibrium model Demand and supply are independent		Disequilibrium model Demand and supply are correlated	
	Coeff.	St. Error	Coeff.	St. Error
Supply equation				
Average $(\text{Debt/EBITDA})_{t-1}$	0.0030***	0.0006	0.0024***	0.0004
Average Young SMEs	0.0369	0.0295	0.0362*	0.0197
Average Collateral/Assets	-0.2407***	0.0764	-0.3095***	0.0419
Instrument: Collateral. residuals	-0.4670***	0.0358	-0.2681***	0.0227
Initial value of y	-0.0749***	0.0134	-0.0359***	0.0089
Demand equation				
Average Non-financial debt/Assets	0.4557***	0.0131	0.4706***	0.0140
Average EBITDA/Assets	0.1587***	0.0091	0.1565***	0.0100
Average Interest rate	0.5580***	0.0219	0.6273***	0.0256
Average Accounts payable/Assets	0.2721***	0.0118	0.2872***	0.0128
Average Investment/Assets	-0.0097	0.0122	0.0045	0.0128
Average Working capital/Assets	-0.0514***	0.0080	-0.0549***	0.0087
Instrument: Interest rate. residuals	0.0088	0.0649	-0.0523	0.0701
Instrument: Accounts payable. residuals	-0.2668***	0.0127	-0.2748***	0.0139
Instrument: Investment. residuals	0.0463	0.0364	0.0258	0.0382
Instrument: Working capital. residuals	-0.0060	0.0070	0.0001	0.0076
Initial value of y	0.0312***	0.0034	0.0306***	0.0037

Notes: ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

Table A2 Estimation results

Other estimates	Disequilibrium model Demand and supply are independent		Disequilibrium model Demand and supply are correlated	
	Coeff.	St. Error	Coeff.	St. Error
Supply equation				
Dummy Construction	0.0894***	0.0088	0.0586***	0.0056
Dummy Trade	0.0667***	0.0074	0.0446***	0.0046
Dummy Hotel and restaurant	-0.0545*	0.0307	-0.0482***	0.0173
Dummy Information and communication	0.1793***	0.0403	0.0916***	0.0164
Dummy Services	0.0336***	0.0095	0.0293***	0.0061
Demand equation				
Dummy Construction	0.0131***	0.0030	0.0134***	0.0031
Dummy Trade	-0.0034	0.0023	-0.0053**	0.0025
Dummy Hotel and restaurant	-0.0401***	0.0067	0.0408***	0.0069
Dummy Information and communication	0.0232***	0.0067	0.0250***	0.0068
Dummy Services	0.0490***	0.0030	0.0482***	0.0032
Supply and demand equations				
Lagged loans, small, manufacturing	0.7775***	0.0079	0.7665***	0.0079
Lagged loans, large, manufacturing	0.8410***	0.0059	0.8347***	0.0059
Lagged loans, small, construction	0.7281***	0.0088	0.7157***	0.0087
Lagged loans, large, construction	0.7813***	0.0086	0.7720***	0.0085
Lagged loans, small, trade	0.7779***	0.0051	0.7687***	0.0051
Lagged loans, large, trade	0.8229***	0.0058	0.8190***	0.0058
Lagged loans, small, hotel	0.7276***	0.0224	0.7126***	0.0220
Lagged loans, large, hotel	0.7514***	0.0179	0.7449***	0.0177
Lagged loans, small, information	0.7538***	0.0197	0.7394***	0.0196
Lagged loans, large, information	0.8027***	0.0222	0.7994***	0.0222
Lagged loans, small, services	0.7237***	0.0069	0.7150***	0.0069
Lagged loans, large, services	0.7776***	0.0082	0.7749***	0.0082

Notes: ***, **, and * stand for statistical significance at 1%, 5%, and 10%, respectively.

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