Rent sharing in Portuguese Banking*

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December 4, 2009

Abstract

Using a fixed effects estimator and a dynamic panel data system-GMM estimator on a sample of 77 banks, covering the period 1988-2005, this paper estimates how wages in the Portuguese banking sector depend on the employers' ability to pay. The results indicate that wages are strongly positively correlated with rents even after controlling for firm and workforce characteristics. A conservative Lester's range of wages due to rent sharing is around 56% of the mean wage of the Portuguese banking sector, a number that is considerably larger than in previous studies.

Keywords: Rent sharing; Portuguese banking industry; Dynamic panel data.
Jel classification: J31, J45, L33.

*We are grateful to the editor and three anonymous referees, Odd Rune Straume, and participants at seminars in Minho and Porto Universities, for their valuable comments on an earlier version of the paper. Corresponding author: Miguel Portela, Universidade do Minho, Departamento de Economia, Gualtar, 4710-057 Braga, Portugal. Tel.: +351 253601910; fax: +351 253601380 E-mail: mangelo@eeg.uminho.pt.

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1 Introduction

Several studies have examined how wages vary systematically by firm or industry profitability. Abundant direct evidence, compatible with a wide range of competing explanations, has shown that a persistent wage-profit correlation exists, both at national and sectoral levels, in quite different economies. Arai (2003) and Johansen *et al.* (2001), for example, look at the relation between pay and profits in Sweden and Norway, respectively, while Estevão and Tevlin (2003), Dobbelaere (2004) and Grosfeld and Nivet (1999) focus on such correlation in manufacturing industries, in the US, Bulgaria and Poland, respectively. More recently, Budd *et al.* (2005) extend this domestic focus of the literature by assessing the existence of international rent sharing between European multinational parent companies and their affiliates. Güertzgen (2009) and Rusinek and Rycx (2008), in turn, make use of variation in the domestic institutions of the labour market to examine how different wage-setting institutions affect the magnitude of profit sharing. In particular, Güertzgen (2009) finds that in Germany, wages respond stronger to firm quasi-rents in non-unionised sectors with firm-specific wage contracts than in unionised sectors with industry-wide wage agreements. Along the same lines, Rusinek and Rycx (2008) find that, in Belgium, rent sharing is higher in industries with more decentralized wage setting.

This paper contributes to this current debate on the linkages between wages, firm-specific profitability and collective bargaining by focusing on one industry: banking in Portugal. This industry provides an interesting case to test different rent-sharing hypotheses. First, banking is a heavily unionised industry. In contrast with nationwide figures, the average union density rate in the industry has been rising since the mid-80s and covers all the workforce (Cerdeira, 1997). Moreover, the industry is covered by one sectoral wage agreement. The oldest trade unions in the country and a group of banks, regardless of their ownership, meet each year to negotiate a single collective wage agreement in the industry. Thus, in contrast with previous work, we have close to ideal conditions for analysing how profit sharing varies across firms with different ownership types. These wage-setting conditions are particularly appealing in our context, as the industry experienced, in the 80s and 90s, a successful deregulatory and privatisation reform (OECD, 1999) which changed significantly its size and ownership structure. Furthermore, these reforms were accompanied by a simultaneous rise in the wage premium and firm profits (Monteiro, 2009). Despite firms being subject to industry-level wage agreements, there is also clear evidence of firm-specific wage flexibility. The wage cushion – the difference between the wage defined by the collective agreement and the actual wage paid – in banking has increased since the early 90s (Aperta *et al.*, 1994) and is the second highest (in 1999) among sixteen industries in Portugal (Cardoso and Portugal, 2005).[1]

We also introduce a new methodology to this literature. To our knowledge, this is the first study to provide evidence of rent sharing allowing for wage inertia and adopting a system of Generalized Method of Moments (GMM) to estimate the dynamic wage-profit equation.[2] One of the most serious problems in this literature is how to deal with the endogeneity of firm performance and measurement error in the wage-profit equation, when instruments are not available and/or not applicable. We try to overpass these difficulties by using alternative measures of profits, as in Estevão and Tevlin (2003), Dobbelaere (2004) and Martins (2009), and by implementing a system–GMM. The system-GMM uses internal instruments based on lags of the instrumented variables for the first-differenced and level equations. This access to multiple instruments is particularly welcome in our context, since previous work has shown that failure to tackle properly these problems leads to an underestimation of the rent-sharing effect.

Finally, the present study also benefits from using rich firm level data.[3] We rely on aggregate data collected for the whole industry by *Associação Portuguesa de Bancos*. Beyond the standard financial information, these unexplored data offer several firm characteristics and portray the workforce in many different attributes. Therefore, in contrast with typical studies based on aggregate data, we are able to control for differences in firms' human capital characteristics, which can induce spurious positive correlations between firm-level wages and profits. These data also cover an extended period, eighteen years, from 1988 through 2005, which makes them well suited for testing our hypothesis. Indeed, a (short-run) positive correlation between profits and pay is consistent with both competitive and non-competitive theories of wage determination. Although the concept of short or long run is naturally controversial, past empirical evidence on the correlation between profits and wages frequently relies on cross-sectional analysis or on relatively short panels.[4]

The remainder of the paper is structured into four sections. Section 2 offers a brief overview of related theoretical and empirical literature. Regulatory reforms and wage bar-

[1] Bastos *et al.* (2009) also find evidence for a the existence of a considerable wage cushion in most Portuguese industries.

[2] Recently, Guertzen (2009) also implements this system-GMM to measure the level of rent-sharing. However, her analysis should be interpreted with some caution, as it fails to pass the overidentification tests.

[3] Matched worker-firm level data is clearly superior but not available for our analysis.

[4] There are other studies, comparable to ours in terms of length of time, but at *industry* level. For example, Estevão and Tevlin (2003) and Blanchflower *et al.* (1996) use data aggregated at industry-level for 37 and 22 years in the US, while Johansen (1996,1999) uses similar data for 25 years in Norway.

gaining developments that took place in Portuguese banking since the mid-80s are presented in Section 3. Section 4 addresses data and model specifications. Section 5 presents and discusses the empirical results. Section 6 offers some concluding remarks.

2 Related literature

This study is related with two strands of the empirical labour literature. The most relevant is whether individual wages are influenced by employers' ability to pay. Given that the neoclassical model of perfect competition is unable to explain the existence of persistent wage differentials among comparable workers, several theories, offering a wide range of competing explanations, predict a positive relationship between firms' performance and wages. For instance, the efficiency-wages hypothesis argues that workers may be paid above the market clearing wage in order to raise their productivity. Higher wages increase the cost of job loss, and thus create incentives to prevent shirking, minimize turnover and attract more able job-seekers. This is particularly relevant for high-skilled and more educated workers, who are more likely to accumulate firm-specific human capital and hence require more expensive training costs for replacement workers. Search theories also emphasize a similar argument: given the between-firm competition for labour, firms may prefer to pay a higher wage in order to reduce labor search costs and attract more able job-seekers. The union bargaining models (right-to-manage and efficient bargaining) also predict a positive pay-performance link which depends on the relative bargaining strength of the unions and firms. Stronger unions allow workers to appropriate a higher share of firms' surplus. The insider-outsider theory suggests that insiders (incumbent workers) due to their stronger bargaining power relative to outsiders (unemployed) are able, in the presence of demand shocks, to obtain a higher pay rather than more jobs for the outsiders. In this context, larger (rather than smaller) firms will provide stronger bargaining position for their insiders. According to the implicit contract theory, wages also move together with profits as wages are set to provide "insurance" against random demand shocks.

Despite the diversity of channels for explaining the positive link between pay and ability to pay, in all of these models the wage-setting can be seen as a form of rent sharing, where pay is determined by a mixture of external and internal forces; typically, the outside temporary wage in the labour market (external) and the level of profits per employee (internal). The validation of these models' prediction has been done for a wide range of countries[5] at national level, but more frequently at sectoral level. In all studies, wages measured at different aggregation levels, either individual-worker, firm, firm-union bargaining unit or industry level are linked with profits per worker measured at firm or industry level. In this respect, our study is closely related to Hildreth and Oswald (1997), who use UK data at firm level for both key variables but contains limited information about other firm and workforce attributes. It is also related to Grosfeld and Nivet (1999) and Dobbeleare (2004), who use similar, but richer data, for Poland and Bulgaria, respectively. Our data are even richer and available for a much larger time span.

There is also significant variation in the empirical specification, but, in general, it corresponds to an expanded form of the wage equilibrium equation. Frequently, other potential sources of rent extraction, such as unionisation level or worker attributes, size or ownership of the firm and industry affiliation, are included in the model as extra controls. Despite all these differences in the specifications, six systematic findings emerge in the literature. First, a positive and significant link from profits to wages is found, regardless of the inclusion of other controls and how wages and profits are measured. The corresponding estimated wage-profit elasticity lies between 0.006 (Christofides and Oswald, 1992) and 0.086 (Rusinek and Rycx, 2008). Second, studies that try to account for the simultaneous determination of wages and profits (inherent in the economic models and in the accounting relation between the two measures) by using external instrumental variables have found profit-effects at least ten times larger than those that do not (see, e.g., Abowd and Lemieux, 1993, Van Reenen, 1996, Estevão and Tevlin, 2003, and Arai and Heyman, 2008). Third, wage reactions to profits also tend to be asymmetric. Wages appear to be more responsive to rises than falls in profits, implying some downward wage rigidity (see, e.g., Grosfeld and Nivet, 1999, Martins, 2009, and Arai and Heyman, 2008). Fourth, profit-effects also tend to be larger in larger firms, in accordance with the insider-outsider model, and in state-owned firms and to be alike in unionised and non-unionised sectors (see, e.g., Hildreth and Oswald, 1997, Estevão and Tevlin, 2003, and Dobbeleare, 2004). Fifth, wage responses to profitability also appear to be more pronounced in industries with more decentralised wage setting agreements (see, e.g., Güertzgen, 2009, and Rusinek and Rycx, 2008). Finally, in conformity with the efficiency-wages hypothesis, rents seem to be unequally shared across different worker groups. In

[5] These include, for example, USA [Estevão and Tevlin (2003), Blanchflower *et al.*(1996)], UK [Nickell and Wadhwani (1990), Hildreth and Oswald (1997)] Canada [Abowd and Lemieux (1993), Christofides and Oswald (1992)], Norway [Johansen et al. (2001)], Sweden [Arai (2003) and Arai and Heyman (2008)], Poland [Grosfeld and Nivet (1999)], Bulgaria [Dobbelaere (2004)] and Ghana [Teal (1996)].

particular, men, the high-skilled, and more educated and more experienced workers appear to benefit more from profit-sharing mechanisms.

Evidence on the sustained link from profitability to pay rejects the validity of the competitive model as an adequate representation of the labour market. Another strand of the literature, which dates from Hendricks (1977), recognizes that labour markets are far from being competitive due to product market regulation, and uses (de)regulatory reforms to test how wages or, more rarely, profits change. Although this literature does not focus on measuring the *level* of rent sharing in each specific industry analysed, it presents empirical specifications linking wages to firm and/or product market conditions before and after reforms, similar to the ones typically estimated in the rent-sharing literature. For example, Crémieux (1996) estimates the effect of the Airline Deregulation Act in the USA on the wages of pilots, flight attendants and mechanics using, apart from deregulatory variables and a time trend for controlling for "natural" earnings trend, six different firm characteristics, including an instrumented profits variable. The main message of this literature is that each regulatory experience is unique. The effects of deregulation on earnings or profits depend on a number of direct and indirect channels between labour and product markets, which make generalizations from one industry to another a dangerous exercise (for a survey of multiple industries, see Peoples, 1998). We connect these two literatures by evaluating the level of rent sharing in a single particular industry, during and after the liberalisation period and considering the type of industrial labour relations in the industry.[6]

3 Reforms and wage-setting in Portuguese banking

The Portuguese banking industry has successively undergone tremendous transformations during the last two decades. Prior to 1984, the industry was almost exclusively composed of a small number of public firms which were overstaffed and inefficient, reflecting an activity severely limited by state control (OECD, 1999). Like in many OECD countries, credit and interest rate ceilings and other capital controls governed daily banking operations. Furthermore, borrowing on public debt was compulsory, which created an additional source of credit misallocation. Entry barriers, either to new or already installed banks through branch expansion, also contributed to the lack of competition and development of the sector.

In 1984, the reversal of the regulated financial system started. The first three legal actions

[6] Assessing the impact of liberalisation on the rent-sharing level is beyond the scope of this study given the limited availability of data for the pre-reform period.

(law 11/83 of 16$^{\text{th}}$ August, decree-law 406/83 of 19$^{\text{th}}$ November and decree-law 51/84 of 11$^{\text{th}}$ February) opened the financial intermediation to the private sector. At the same time, some of the deposit and borrowing interest rates were liberalised. This process involved a cautious sequencing of step-by-step measures which dismantled most of the regulatory instruments that directly affected the behaviour of firms. In 1992, the complete removal of price and entry barriers was accomplished with the lifting of the remaining capital controls and barriers to branch expansion.[7]

In the second phase, between 1989 and 1997, the full ownership of eleven out of twelve public banks was transferred to the private sector. This privatisation program (law 84/88 from 20$^{\text{th}}$ July and decree-law 11/90 from 5$^{\text{th}}$ April), following two Constitutional Amendments, shared the common goals of the worldwide privatisation processes: independence and improvement of public banks' performance and further enhancement of banking competition.

During the same period, conglomeration and technological innovations also reshaped the industry. The conglomeration process – involving the formation of groups (bancassurance) – took place mainly in the mid 90s, while the consolidation process – starting in 1998 with the merger of three recently privatised banks – is still ongoing. The widespread use of new technologies, such as the automated teller machines and the electronic fund transfer at the point of sale, also contributed to a reduction in time and costs associated with financial transactions.

How did regulatory reforms affect the wage bargaining in the industry? The developments described previously conditioned the type of industrial labour relations prevailing in banking, but did not affect the bargaining process itself. Covering three different geographical areas, the oldest trade unions in the mainland represent all employees, regardless the ownership of the bank. These trade unions and a group of banks, public and now private (domestic or foreign), meet each year to negotiate the collective bargaining agreement. This collective agreement, the most detailed and extensive in Portugal, regulates the employment conditions, the remuneration and the duration of work. It also delimits the starting wage level and the compulsory wage progressions for each of its 18 levels of the 4 groups defined to cover all the banking workforce.[8] In practice, the bargained wage works as a wage floor and firms are free to set wages above the negotiated benchmark according to their specific conditions. The differential between the wage defined by the collective wage agreement and the

[7] OECD (1999) offers a detailed and chronological description of all reform measures.

[8] There are two exceptions to this industry wage agreement. In 2001 and 2004, the two largest groups in the financial intermediation sector, BCP and CGD, signed individual firm-level wage agreements.

actual wage paid, also known as the “wage cushion”, has increased since the early nineties (Aperta *et al.*, 1994). More recently, Cardoso and Portugal (2005) estimate a gap of 57% between the actual wage and the contractual wage for the financial intermediation industry in 1999, the second largest among sixteen industries in Portugal.

Beyond this broad scope of the collective agreement, the union attachment in the industry is one of the strongest in the economy. Between the periods 1974-78 and 1991-95, average union density increased from 71% to 106% (Cerdeira, 1997).[9] Despite this increased union density, banking unions did not contest the new market environment to any significant extent. The resistance was limited, not coordinated, mostly being made through internal speeches and pamphlets which were rarely reported in the national press. The total number of strikes was limited as well: five strikes occurred in 1986, 1988 and 1989, each involving less than half of the total workforce (MSST, 1987-2001).

4 Data and model specification

This study uses data released by *Associação Portuguesa de Bancos* in the *Boletim Informativo.* This consists of annually collected firm-level data from company accounts of all banks in Portugal. Beyond standard financial and operational information, the dataset reports other firm characteristics such as ownership, firm size measured either in terms of employment or number of branches per bank, location of the branches, capital-labour ratio and age, and portrays the workforce in terms of schooling, tenure, age, type of activity and occupation in each bank. Table 1 displays some statistics that summarise the main changes occurring in Portuguese banking, in terms of size and workforce, in three points in time: 1988, 1998 and 2005.[10] The years 1988 and 2005 correspond to the first and last year, respectively, for which information is available. The year 1988 portrays the sector during the deregulation reform, one year before the implementation of the privatisation program. For symmetry, we choose the year 1998, which corresponds to one year immediately after the privatisation program is concluded in the industry.

The key variables, log wages and profitability, are computed following the usual practice in the literature (Hildreth and Oswald, 1997). Thus, the log average wage is constructed as the logarithm of the ratio between the reported wage bill and total employment. For

[9] The union density rate also includes retired employees, making a density rate in excess of 100% possible.

[10] Table A.1 in Appendix contains the summary of all variables across ownership groups over the period 1988-2005.

Table 1: Industry and labour market in Portuguese banking, 1988-2005

Variable	1988	1998	2005
Number of banks	26	48	42
Firm size (10^3 employees)	2.227	1.234	1.080
Number of branches	1467	4513	4949
Geographical Herfindahl index	.329	.423	.405
Ownership (%)			
Public	46.2	4.2	2.4
Private domestic	23.1	56.3	54.7
Foreign	30.7	39.5	42.9
Firm age	41.7	23.9	20.3
Real capital per worker (prices=2005)	35.7	83.2	57.2
Profits per worker	49.3	89.5	120.2
Real profits per worker (prices=2005)	113.2	109.8	120.2
Value added per worker	64.2	140.7	188.3
Real value added per worker (prices=2005)	147.3	172.7	188.3
Log average wage	2.57	3.81	4.04
Log average real wage (prices=2005)	3.40	4.01	4.04
Schooling (%)			
Primary	n.a.	14.5	6.2
High school	n.a.	42.3	38.9
University degree	n.a.	43.2	54.9
Tenure, in years (%)			
$[0, 6)$	38.0	49.0	45.3
$[6,\ 11)$	23.0	24.1	25.9
$[11, -)$	39.0	26.9	28.7
Age, in years (%)			
$[-, 35)$	47.6	53.7	49.5
$[35, 55)$	49.6	42.8	47.2
$[55, -)$	2.8	3.5	3.3
Commercial activity (%)	44.8	45.1	43.5
Occupation (%)			
1	20.1	23.5	25.9
2	14.1	31.1	40.3
3	55.8	42.9	32.6
4 (lowest)	10.0	2.5	1.2

Source: Own computations based on *Boletim Informativo*, *Associação Portuguesa de Bancos*.

profitability, we use two variables that are relevant for wage determination according to union bargaining models: value added per worker and profits per worker. Valued added per worker is obtained by taking sales revenues less operational costs and dividing by total employment. Profit per employee is obtained by taking value added per worker minus our constructed wage variable. More precisely, our profits measure corresponds to the item *Resultado Bruto de Exploração* in the income sheet and corresponds closer to the economic concept of surplus available to share between workers and firms. The variable value added per worker has the advantage of circumventing the downward bias induced by the accounting relationship between wages and profits. The capital-labour ratio is obtained by dividing real fixed assets by total employment. All monetary measures are expressed in 10^3 Euros.

Table 1 indicates that the abandonment of regulatory restrictions in 1984 led to a proliferation of new firms and branches, which fueled the demand for labour, especially for skilled employees. Indeed, the number of banks (and branches) operating in the industry increased markedly from 26 to 48 (1467 to 4949) between 1988 and 1998, though the number of banks is marginally squeezed during the merger wave after 1998. Total employment and the number of branches in the market also did not grow at the same pace due to the downsizing process that took place during the privatisation program and the merger wave. Our geographical Herfindahl concentration index, measured at branch level, also rose over the period, implying an increasingly widespread presence of the banks in all districts of Portugal.[11]

Reforms also changed significantly the ownership and average age of Portuguese banks. Nowadays, private firms, either domestic or foreign, dominate the market and firms' average age more than halved compared to its level in 1988. Nevertheless, significant discrepancies still emerge in terms of firm size and age, among the three ownership categories. In 2005, the publicly owned bank is the largest and the oldest firm in the sector, employing approximately 20% of the total workforce. Privately owned banks, despite being much smaller, are on average three times larger and twice as old as their foreign counterparts.

Payment in the industry, measured at firm level, experienced a strong rise of 64% between 1988 and 2005. The workforce became more educated, younger (less than 35 years) and less experienced (less than 6 years of tenure). Again, this is more pronounced for all privately owned firms, but in particular for foreign firms, which have more than half of the total workforce in these categories. Banking employees are also working mainly in intermediate/middle occupations and in other activities than the commercial one (public employees are the only exception). Finally, profitability, either measured by valued added per worker or profits per worker, also increased over the period, although for the latter the rise was modest and observed only after 1998. Within these aggregate figures, there has been a continuous

[11] This measure is calculated per year t and bank i by summing up the square of the share of the number of branches of bank i in all districts.

decline in profitability in foreign firms, whose market share increased in the sample period. However, profitability has increased significantly in domestically owned firms, both public and private, implying that reforms are associated with improvements in firm performance.

For empirical purposes, we explore the structure of our data – an unbalanced panel data covering 77 banks in the period 1988-2005 – by specifying two panel data models. We start with the estimation of a static model given by:

$$w_{it} = \alpha_1 + \beta.R_{it} + \alpha_2.Z_{it} + v_i + \varepsilon_{it}, \tag{1}$$

where subscripts i and t index bank i and time t; w is real average log wage, R refers to our profitability or rents variable by worker and Z is a set of other regressors that vary by bank and time. In our specifications, Z includes two dichotomous variables that identify three ownership categories (public, private and foreign), time-varying measures of firm size (thousands of employees and number of branches) and firm age, real capital-labour ratio, geographic dispersion of branches given by the geographical Herfindahl index and five attributes of the workforce: share of workers in three different educational categories, share of workers in three seniority groups, share of workers in three age groups, share of workers in commercial and in other activities and share of workers in four occupational categories. The remaining unobserved time-invariant bank heterogeneity that affects wages is captured by the unobservable bank specific effect v. Following Hildreth and Oswald (1997), Z also includes a full set of time dummies effects to control for any external "outsider" variable considered in the theoretical models, such as unemployment, or to control for any unobserved heterogeneity common to all banks. ε represents an error term. We also estimate some versions of (1) where the coefficient β is specified to depend on the ownership of the bank:

$$\beta = \beta_0 + \beta_1.P_{it} + \beta_2.F_{it},$$

where P and F are dummies variables that identify public and foreign ownership, respectively.

Equation (1) is estimated using a standard within-groups (fixed effects) estimator. The estimates of the parameters of interest (β or β_0, β_1 and β_2) may be biased for different reasons. The first concern relates with the measure of profitability. Using accounting profits, which decrease when wages increase, leads to a direct downward bias since wages appear on both sides of the equation (1). An alternative measure to circumvent this problem is to use value added per worker which is not affected directly by changes in wages. Nevertheless, both measures may suffer from measurement error. If the measurement error is random, then the effect is the well-known *attenuation bias* and the rent-sharing effect will be downward biased. If instead, the measurement error is nonrandom, then the direction of bias is not known a priori (Margolis and Salvanes, 2001, discuss the case when firms spread losses in order to reduce income tax liabilities).

Apart from this bias related to the choice and quality of variables used, another downward bias occurs in some economic models. For example, in the right-to-manage model, where firms set the level of employment subsequent to wage bargaining, the rent-sharing parameter depends, as shown by Estevão and Tevlin (2003), on the labour demand elasticity, the value added employment elasticity and the unions' bargaining power. Thus, it works as a lower bound for the parameter representing the bargaining power of workers, which corresponds to the rent-sharing coefficient in an efficient bargaining framework.

In order to investigate whether these potential issues affect our results and to accommodate for the sluggish adjustment of wages, we also estimate a dynamic version of wage equation (1) by including a lagged dependent variable as a regressor,

$$w_{it} = \alpha_1 + \delta . w_{it-1} + \beta . R_{it} + \alpha_2 . Z_{it} + v_i + \varepsilon_{it}. \tag{2}$$

Given the presence of bank-specific effects, the estimation of equation (2) by OLS, fixed or random effects, is inconsistent.[12] One possible solution is to take first differences to eliminate bank-fixed effect and apply the Arellano and Bond (1991) first-differenced generalized method of moments (GMM). This method allows us to control for the bias introduced by omitted time-invariant variables, and the inconsistency resulting from correlation between the transformed lagged dependent variable and the transformed residual. However, if "the lagged levels of the series are only weakly correlated with subsequent first-differences, so that the instruments available for the first-differenced equations are weak" (Bond *et al.*, 2001, p. 6), this solution has poor finite sample properties on bias and precision. Another possibility, which we follow in this study, consists of estimating the system-GMM discussed by Blundell and Bond (1998). The system estimation combines a set of moment conditions for the first-differenced equations with a set of moment conditions implied for level equations. Relative to the original GMM, the estimation of the system-GMM is preferable when the dependent

[12] Nickell (1981) shows the inconsistency of the fixed effects estimator when applied to a dynamic model with fixed effects, for $N \longrightarrow \infty$ and fixed T.

and/or independent variables are persistent. Estimation of equation (2) by system-GMM implies a smaller sample size as compared to the estimation of equation (1) by a fixed effects procedure.

Following Bowsher (2002), the choice of our instruments is parsimonious. We use w_{it-2} and w_{it-3} as instruments for the first-differenced equation and Δw_{it-1} as instruments for the equations in level. Furthermore, the explanatory variables, R_{it}, $R_{it}.F_{it}$ and $R_{it}.P_{it}$ are treated as endogenous.[13] In the estimation, we use two to three lags of their levels as instruments for the first-differenced equations and lagged first-differences as instruments for the level equations. With the exception of time dummies, the remaining explanatory variables are also instrumented as endogenous variables. Therefore, in the first-differenced equations, we use their levels lagged one to three periods and contemporaneous first-differences as instruments for the equations in levels. We treat time dummies as exogenous variables. The number of instruments increases significantly with the number of periods, which leads to overfitting and tends to cause biased standard errors. In order to limit the number of instruments, we have not applied each moment condition underlying the system-GMM procedure to each time period and lag available. Instead, we apply a single moment condition for each period and regressor. Finally, the results we report are for the two-step GMM estimation procedure, following the correction proposed by Windmeijer (2005). The reported statistics are robust to heteroskedasticity and serial correlation in the errors.

Since we suspect the errors are non-spherical, we report the Hansen consistent test instead of the Sargan statistic. When the estimation is overidentified, the Hansen statistic allows for the test of the validity of the instruments used. This test follows a χ^2 with degrees of freedom equal to the number of instruments over the number of regressors. In order to test the validity of the extra instruments used in the system-GMM, we also report the difference in the Hansen test. In such a test, we compare the Hansen statistic between the two estimation procedures: first-differenced and system-GMM. Under the null hypothesis of valid instruments, the resulting statistic follows a χ^2 distribution, with degrees of freedom equal to the number of extra instruments associated with the level equations. In line with the usual procedure in the literature, we also report tests for serial correlation in the residuals of the dynamic equation. These tests are applied to differenced residuals, and provide another check on the validity of the instruments used in the estimation.

[13] Given the unavailability of external instruments, our instrumenting strategy is not directly driven from the rent-sharing theoretical models, as in Estevão and Tevlin (2003) or Van Reenen (1996).

5 Empirical results

Table 2 reports empirical results for different versions of the static model (1).[14] Columns 1 to 5 show results when profitability is measured by profits per worker while columns 6 to 10 repeat the previous five specifications using instead value added per worker. Comparing the two set of estimates, we note that although all rent-sharing estimates are similar in magnitude, using value added, rather than profits per worker, yields marginally larger effects, thus confirming the downward bias inherent to the profits per worker variable.[15]

The point estimate of rents per employee on log wages is statistically significant and around 0.002 (for both profitability measures) when only time effects are used as controls. Adding firm attributes (columns 2 and 7) improves the quality of the model specification and reduces marginally the estimates of our parameter of interest.[16] Including further workforce attributes (columns 3 and 8) reduces our sample size and the coefficients of interest.[17] These, despite being small, imply very large elasticities of wages with respect to profits and value added per worker: between 0.137 and 0.254, respectively. As discussed previously, these figures are well above the range of elasticities usually found in the literature and above those found by Martins (2009) for Portugal using linked employer-employee data (-0.031 and 0.078). Moreover, these figures are much higher than the range from 0.01 to 0.086 found by Arai (2003), Güertzgen (2009) and Rusinek and Rycx (2008), under similar wage setting mechanisms. Our rent-sharing estimates might be upward biased as we do not control for variation in hours worked. We explore this issue in detail below.

The remaining columns in Table 2 explore whether rent sharing varies according to the bank ownership structure. Models 4 and 9 include interactions between ownership and rents while models 5 and 10 add several workforce regressors beyond the referred interactions. These regressors are jointly significant at 10%, making our preferred specifications the most expanded ones in columns 5 and 10.[18] In contrast with Estevão and Tevlin (2003), the inclusion of the workforce regressors also changes the estimates of the coefficients of interest. Following earlier findings, the estimate of the rent-sharing parameter is higher in state-

[14] The variable foreign ownership is not included in the results as foreign banks do not change ownership over the sample period.

[15] Therefore, rents per worker seem to be a better measure for this type of analysis.

[16] The corresponding likelihood ratio tests are $2(-134.0839 + 148.9285) = 29.6892$ and $2(-77.2704 + 87.1218) = 19.703$, with a critical value of $\chi^2_{(6)} = 12.59$.

[17] These regressors are jointly significant in both models with F-statistics of 1.94 and 2.03 and p-values of 0.053 and 0.042, respectively,

[18] The joint significance test has a F-statistic of 1.80(1.93), with a p-value of 0.076(0.054) for profits (value added) per worker.

owned firms in both models (0.0024) implying an elasticity of wages with respect to rents of 0.174 and 0.307, when using profits and value added per worker, respectively. This range of elasticities is above the one found for domestic firms ($0.098 - 0.220$), though similar to that found for foreign firms ($0.182 - 0.304$). These large elasticities found for foreign firms reflect larger figures for both profits and value added per worker.

In terms of pay level, public banks pay 22 to 24 per cent less than their private counterparts when workforce characteristics are controlled for.[19] Firm age has a significant positive effect on log wages of 0.029 or 0.021, implying that more visible banks pay higher wages. Capital per worker also has a positive effect of 0.001, similar in magnitude to that reported in Sweden or USA (Arai, 2003), but which is statistically significant when rents are measured by value added per worker. Finally, variables controlling for the bank size, either in terms of level of employment or number of branches, and the dispersion of branches across the Portuguese territory, appear to be insignificant.

Table 3 verifies the robustness of our evidence and interpretation of Table 2 and explores how rent sharing varies across some groups. For simplicity, we show results using only value added per worker. The first concern is about not controlling for the number of hours worked, a variable that is not available in our database. If profitability increases in response to increases in product demand, which are accommodated by an increase in demand for labour hours, we may observe a positive correlation between profitability and annual wages per worker that is still compatible with competitive wage determination. We explore this issue by splitting the period of analysis in two sub-periods. Columns 1 and 2 use specification 9 from Table 2, applied to the period before and after 2000, respectively. The period before 2000 corresponds to a "boom" period characterised by high GDP growth rates after Portugal joined the EU in 1986. The period after 2000 is characterised by stagnation/recession in the Portuguese economy. We believe that the high-growth period is related with a stronger product demand, and thus higher level of working hours, implying larger rent-sharing effects when hours worked are not controlled for.

[19] These figures are obtained using the exponential function: $e^{-.2442(-.2736)} - 1 \cong -0.217(-.239)$.

Table 2: Rent-sharing - results for static models

Rent measure	Profits per worker					Value added per worker				
Model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Rents per worker(RW)	.0015***	.0013***	.0012***	.0013***	.0009**	.0019***	.0017***	.0015***	.0017***	.0013***
	(.0003)	(.0003)	(.0003)	(.0004)	(.0004)	(.0002)	(.0002)	(.0003)	(.0004)	(.0003)
RW*Foreign				-.00001	.0006				-.00006	.0004
				(.0005)	(.0005)				(.0004)	(.0004)
RW*Public				-.0040***	.0015**				-.0013	.0011***
				(.0006)	(.0007)				(.0008)	(.0004)
Public		-.0585	-.1545	.1392	-.2442*		-.0515	-.1458	.0582	-.2736**
		(.1209)	(.1120)	(.0992)	(.1457)		(.1031)	(.1008)	(.1212)	(.1303)
Firm age		.0524***	.0265**	.0542***	.0291**		.0447***	.0190**	.0456***	.0210**
		(.0065)	(.0107)	(.0070)	(.0113)		(.0058)	(.0095)	(.0062)	(.0103)
Capital per worker		.0010*	.0007*	.0010*	.0007		.0011*	.0009*	.0011*	.0010*
		(.0005)	(.0004)	(.0005)	(.0004)		(.0006)	(.0005)	(.0006)	(.0005)
Branches		.00009	-.0002	.00009	-.0002		-.00009	-.0003	-.00004	-.0004
		(.0004)	(.0006)	(.0004)	(.0006)		(.0004)	(.0005)	(.0004)	(.0005)
Employees (10^3)		-.0175	-.0339	-.0310	-.0395		-.0048	.0012	-.0124	-.0006
		(.0491)	(.0669)	(.0444)	(.0671)		(.0436)	(.0644)	(.0422)	(.0652)
Herfindahl		.0352	.0079	.0294	.0041		.0479	.0121	.0490	.0038
		(.1064)	(.1139)	(.1030)	(.1175)		(.0908)	(.0981)	(.0911)	(.1004)
Worker characteristics	no	no	yes	no	yes	no	no	yes	no	yes
Observations	731	731	571	731	571	731	731	571	731	571
Banks	77	77	75	77	75	77	77	75	77	75
Adj. R^2	.3761	.3967	.1734	.4093	.1764	.4732	.4836	.262	.4845	.2637
LogLikelihood	-148.9285	-134.0839	-88.8039	-125.3278	-86.7342	-87.1218	-77.2704	-56.4695	-75.556	-54.7458

Notes: Significance levels: $*$: 10% $**$: 5% $***$: 1%. Robust standard errors in parentheses, clustered by bank. The dependent variable in each model is log average wage. All regressions are estimated by fixed effects and include time dummies. Worker characteristics are defined in Section 4.

The evidence shown in columns 1 and 2 is mixed. For domestic private firms, the rent-sharing estimate reduces significantly in the "stagnation" period (compared with the "boom" period) while it rises (from 0.0014 to 0.0022) for foreign firms and remains relatively stable (varying from 0.0009 to 0.0015) for public firms.[20] Thus, the results are only consistent with our hypothesis when we consider domestic private firms. For the remaining firms, the evidence is not compelling, suggesting that other factors, beyond number of hours worked, are playing an important role in the process of wage formation.

Columns 3 and 4 test if wages are equally sensitive to both increases and decreases in rents. Column 3 splits the sample according to whether the change in value added per worker is negative while column 4 refers to positive changes in rents. This partition is based on the sign of the change in rents, regardless of its level. In contrast with previous research, in particular Martins (2009), the coefficients are close in both cases and still very precisely estimated. This similar response of wages to different changes in profits might be partly explained by the existence of a high "wage cushion" in the industry. If firms pay a wage differential well above the wage defined by the collective agreement, then this "wage cushion" would allow firms to accommodate to different demand shocks. Notice, however, that public firms do not follow this general pattern.

We now inspect if firms with different workforce attributes differ in terms of rent-sharing behaviour. For this purpose, we compute for each year of our sample the median of schooling at the industry level and classify firms according to this threshold. Column 5 shows the results for firms with a level of schooling lower than the industry median while column 6 reports results regarding banks in the upper part of the schooling distribution. Column 6 does not present estimates for the public banks as these employ lower educated workers (see table A.1 in Appendix). Estimates from this partition suggest that, in contrast with previous research, rent sharing is higher in firms with a less educated workforce. This finding may explain why the banking wage premium in Portugal is higher for less skilled workers (Monteiro, 2009).

Finally, columns 7 to 9 allow the process of wage formation to be fully determined according to ownership of the bank – private, public and foreign, respectively – instead of imposing equal returns to both time and firm attributes as in Table 2, column 8. Notice that, although we do not include worker attributes in the specifications, the rent-sharing estimates across different ownership types follows previous pattern.

[20] After 1998 the number of observations of public banks is limited, preventing the estimation of the public ownership coefficient.

Table 3: Rent-sharing - robustness check using value added

Model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	≤ 2000	> 2000	$\Delta RW \leq 0$	$\Delta RW > 0$	$Low - Edu$	$High - Edu$	$Private$	$Public$	$Foreign$
Rents per worker(RW)	.0025***	.0009***	.0018***	.0023***	.0034**	.0009***	.0015***	.0062***	.0018***
	(.0004)	(.0003)	(.0006)	(.0004)	(.0015)	(.0003)	(.0003)	(.0007)	(.0002)
RW*Foreign	-.0011**	.0013***	-.0006	-.0011**	-.0014	.0007*			
	(.0005)	(.0005)	(.0007)	(.0005)	(.0013)	(.0004)			
RW*Public	-.0016**	.0006	-.0041***	.0006	.0006				
	(.0008)	(.0016)	(.0012)	(.0009)	(.0005)				
Public	.1089		.0317	-.1297	-.2292**				
	(.1156)		(.1669)	(.1242)	(.1006)				
Firm age	.0555***	.0504**	.0524***	.0353***	.0262***	.0440***	.0399***	.0055	.0446***
	(.0070)	(.0243)	(.0125)	(.0073)	(.0076)	(.0107)	(.0076)	(.0152)	(.0093)
Capital per worker	.0007	.0024***	.0014*	.0006	-.0003	.0013**	.0021***	.0035	-.0003
	(.0005)	(.0006)	(.0007)	(.0007)	(.0015)	(.0007)	(.0006)	(.0026)	(.0003)
Branches	-.00005	.0031*	-.0006	.0004*	-.0003	-.0002	.0002	.00001	-.00004
	(.0002)	(.0016)	(.0006)	(.0002)	(.0006)	(.0006)	(.0006)	(.0007)	(.0003)
Employees (10^3)	-.0281	-.2805***	.0667	-.0277	.0616	-.4119**	-.0147	.1467**	-.3370
	(.0360)	(.0913)	(.0687)	(.0370)	(.0881)	(.1761)	(.0627)	(.0599)	(.3988)
Herfindahl	.1281	-.0806	-.0680	.2046**	.2169	.0475	-.0537	-.4500	.1296
	(.0821)	(.0782)	(.1682)	(.0987)	(.2290)	(.1235)	(.1161)	(1.5302)	(.1046)
Observations	504	227	264	388	288	283	384	76	271
Adj. R^2	.6677	.3413	.4706	.5269	.2401	.3357	.3914	.923	.4811
LogLikelihood	101.0577	12.0381	63.4391	7.2622	-17.2622	-.4874	-66.1455	91.7637	22.4089

Notes: Significance levels: $*$: 10% $**$: 5% $***$: 1%. Robust standard errors in parentheses, clustered by bank. The dependent variable in each model is log average wage. All regressions are estimated by fixed effects and include time dummies.

Table 4 presents our preferred specifications of Table 2 (columns 3, 5, 8 and 10) augmented by lagged wage in order to capture some dynamics of the wage formation. For the four specifications, the Hansen and the Difference Hansen (Diff-H-AB) overidentification tests do not reject the validity of instruments used in each regression. In particular, Difference Hansen tests indicate the validity of the extra instruments used in the system-GMM estimation when compared to the original Arellano and Bond (1991) proposal. The serial correlation tests based on the first-differenced residuals fail to reject the hypothesis of no second-order correlation. These results, read in conjunction with (the statistically significant) high persistence in wages (within the range of coefficients, $0.2-0.6$, found for this type of data), indicate that the system-GMM procedure is appropriate to estimate our dynamic panel model.

Table 4 provides further support for the existence of strong rent sharing in the Portuguese banking sector when value added per worker is employed, although the magnitude of the effect is smaller when compared to the one found earlier with the static model (Table 2, column 8). However, wages are invariant to profits per employee. The respective estimate is marginally lower than before but loses its statistical significance.

With this dynamic specification, the rent-sharing effect does not vary across firm ownership (columns 2 and 4). In particular, the coefficients for the interaction term between public and rents per employee are similar, or even larger, but are not statistically significant. A test on the joint significance of both interactions yields F-statistics of 0.57 and 0.46 with corresponding p-values of 0.556 and 0.635. The remaining estimates shown in Table 4 are mixed. Some results are in line with those found earlier with the static model, but present a lower magnitude (firms age, capital per worker and number of branches) and are statistically insignificant. The other estimates are not significant, though they do not follow our earlier evidence.

When seen in conjunction, our static and dynamic models establish a short-run elasticity of wages with respect to value added per capita that varies between 0.152 and 0.254. With respect to the long-run, our dynamic estimates imply an elasticity of 0.32.[21] These figures imply that doubling the per capita value added raises wages about 15 to 25 per cent in the short-run and 32 per cent in the long-run. Hence, our more conservative rent-sharing estimate implies an impact on wages due to a movement from the bottom to the top of the profits distribution of around 56%.[22] This Lester's (1952) "range of wages", despite risking

[21] This number is obtained by multiplying 0.152 by $(1-.5261)^{-1}$.

[22] Following Blanchflower *et. al* (1996), Lester's (1952) range of wages is obtained by multiplying the

to be upward biased as we cannot control for variations in average hours of work, is well above the 16% and 24% established for the UK (Hildreth and Oswald,1997) and the US (Blanchflower *et al.*, 1996), or the 12% − 24% found in Sweden, where the wage setting mechanisms are similar to our case.

Table 4: Rent-sharing - results for the dynamic model

Rent measure	Profits per worker		Value added per worker	
Model	(1)	(2)	(3)	(4)
LagWage	.6042***	.6164***	.5261***	.5287***
	(.0949)	(.0909)	(.1028)	(.1040)
Rents per worker(RW)	.0008	.0009	.0009*	.0011**
	(.0006)	(.0008)	(.0005)	(.0005)
RW*Foreign		-.0004		-.0003
		(.0012)		(.0010)
RW*Public		.0028		.0010
		(.0026)		(.0011)
Public	.3127	.0945	.2121	.0194
	(.4116)	(.5233)	(.4418)	(.3643)
Foreign	.0654	.0993	.2021	.2557
	(.4447)	(.4974)	(.3429)	(.3637)
Firm age	-.0010	-.0003	.0007	.0015
	(.0052)	(.0056)	(.0057)	(.0042)
Capital per worker	.0004	.0003	.0005	.0004
	(.0006)	(.0006)	(.0006)	(.0006)
Branches	-.0009	-.0007	-.0009	-.0009
	(.0011)	(.0010)	(.0010)	(.0010)
Employees (10^3)	.0786	.0743	.0987	.1011
	(.1403)	(.1222)	(.1271)	(.1136)
Herfindahl	-.3100	-.2261	-.2526	-.2410
	(.2057)	(.2452)	(.2226)	(.1966)
Observations	527	527	527	527
Banks	75	75	75	75
Hansen test	36.8762	37.7169	37.3114	38.3902
Hansen-*df*	37	41	37	41
Hansen-*pv*	.4748	.6173	.4548	.5872
Diff-H-AB	16.2408	14.3855	17.4132	16.3579
Diff-H-AB-*pv*	.5757	.8104	.4949	.6942
AR1	-3.0022	-3.0084	-2.9978	-3.1125
AR1-*pv*	.0027	.0026	.0027	.0019

Continued on next page...

elasticity of wages with respect to rents by 4*SD(rents)/Mean(rents). This measures the fraction of the wage inequality that is due to rents, assuming that the rent distribution has a width of 4 standard deviations (SD).

... table 4 continued

Rent measure	Profits per worker		Value added per worker	
Model	(1)	(2)	(3)	(4)
AR2	.4369	.2532	.4153	.3202
AR2-*pv*	.6622	.8001	.6779	.7488

Notes: Significance levels: $*$: 10% $**$: 5% $***$: 1%. Robust standard errors in parentheses. The dependent variable is log average wage. The regressions are estimated by the system GMM procedure as discussed in Blundell and Bond (1998) and include time dummies and worker attributes. The estimation details, including the instrumens used, are described in Section 4. Hansen test is the robust test for overidentifying restrictions. Diff-H-AB is the difference in the Hansen test between the system GMM and the GMM estimation based on first-differences; *df* stands for degrees of freedom and *pv* stands for p-value of the test. AR(1) and AR(2) are tests for first- and second-order serial correlation in first-differenced residuals.

6 Conclusions

Despite the vast literature on rent sharing documenting whether wages vary systematically with firm or industry profitability, the examination of such a correlation accounting for wage bargaining has just begun. This study takes wage bargaining into consideration by measuring the degree of rent sharing in one single industry: the Portuguese banking. This industry provides a notable case study for several reasons. First, it is a heavily unionised sector covered by one single wage agreement. Therefore, we have almost ideal conditions to examine how rent sharing differs across key firm-specific variables, notably ownership type. The Portuguese banking has also experienced, in the 80s and 90s, a successful deregulatory reform which was accompanied by a simultaneous rise in the wage premium and firm profits. There is also evidence of bargaining over wages at local level in banking since the wage cushion – the differential between the wage defined by the collective agreement and the actual wage paid – has increased since the early 90s and is among the highest in Portugal.

We use a particularly rich firm-level data set for 77 banks for examining the level of rent sharing in the industry. Our principal result points to a positive and statistically significant relationship between rents per capita and wages in the Portuguese banking system. Indeed, our fixed effects estimates point to an elasticity of wages with respect to profits and value added per worker of 0.137 and 0.254, respectively. These estimates, possibly upward biased as we were not able to control for average hours worked per employee, are very large when compared with the usual range of estimates found in the rent-sharing literature or even with those found in Portugal by Martins (2009). Our results, in line with Güertzgen (2009) but contrary to Arai (2003), imply that wage-setting institutions may indeed matter for the magnitude of rent sharing. When we control for measurement error and endogeneity issues by estimating a dynamic model using a system-GMM estimator, our elasticity estimate with respect to value added per worker drops to 0.152. This is still a very high estimate when compared to previous evidence. Based on this elasticity, the cross-section variability of banking profits explains more than half of the cross-sectional variability in wages.

We also show some results that contrast with earlier evidence. First, our estimates of rent sharing across ownership are not conclusive. Although public and foreign banks show higher point rent-sharing estimates than domestic private banks, their coefficients are not statistically significant in the dynamic framework. Second, wage responses to changes in rents, either positive or negative, are nearly the same, thus rejecting the hypothesis of downward rigidity. We interpret this finding with the existence of a high "wage cushion" in the industry. Finally, our wage responses to changes in rents are stronger in firms whose work force presents average education levels below the median of the industry. This evidence might be consistent with the fact that the banking wage premium is larger for low skilled workers. Further research, ideally using linked employer-employee data, would be useful to confirm these findings.

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Appendix

Table A.1: Summary statistics

	Private			Public			Foreign			Total		
Variable	Mean	(Std. Dev.)	Obs.	Mean	(Std. Dev.)	Obs.	Mean	(Std. Dev.)	Obs	Mean	(Std. Dev.)	Obs.
Banks			43			13			32			77
Employees (10^3)	1.563	(1.995)	384	4.878	(3.292)	76	0.221	(0.545)	271	1.410	(2.259)	731
Branches	117.997	(172.968)	384	199.408	(191.82)	76	20.815	(66.575)	271	90.434	(156.690)	731
Herfindahl	0.294	(0.252)	384	0.143	(0.119)	76	0.552	(0.304)	271	0.374	(0.299)	731
Firm age	28.740	(43.394)	384	89.500	(42.457)	76	8.181	(6.533)	271	27.435	(41.575)	731
Capital per worker	103.437	(89.774)	384	67.914	(38.001)	76	93.160	(106.924)	271	95.934	(93.379)	731
Profits per worker	108.268	(126.807)	384	58.049	(54.087)	76	139.064	(153.283)	271	114.464	(134.177)	731
Value added per worker	166.717	(158.738)	384	94.690	(66.904)	76	194.417	(163.987)	271	169.497	(156.344)	731
LogWage	3.881	(0.617)	384	3.491	(0.433)	76	3.919	(0.433)	271	3.855	(0.551)	731
Schooling												
Primary	0.161	(0.171)	336	0.406	(0.130)	27	0.056	(0.091)	208	0.134	(0.165)	571
High school	0.424	(0.177)	336	0.398	(0.100)	27	0.425	(0.178)	208	0.423	(0.175)	571
University	0.415	(0.244)	336	0.196	(0.089)	27	0.518	(0.200)	208	0.442	(0.235)	571
Tenure, in years												
$[0-6[$	0.491	(0.297)	383	0.136	(0.088)	76	0.599	(0.290)	269	0.494	(0.309)	728
$[6-11[$	0.186	(0.134)	383	0.196	(0.109)	76	0.210	(0.163)	269	0.196	(0.143)	728
$[11-[$	0.323	(0.288)	383	0.668	(0.138)	76	0.191	(0.225)	269	0.310	(0.288)	728
Age, in years												
$[-,35[$	0.525	(0.207)	384	0.284	(0.107)	76	0.581	(0.179)	268	0.521	(0.207)	728
$[35,55[$	0.439	(0.182)	384	0.662	(0.100)	76	0.395	(0.166)	268	0.446	(0.186)	728
$[55,-[$	0.035	(0.036)	384	0.053	(0.028)	76	0.024	(0.033)	268	0.033	(0.035)	728

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... table A.1 continued

	Private			Public			Foreign			Total		
Variable	Mean	(Std. Dev.)	Obs.	Mean	(Std. Dev.)	Obs.	Mean	(Std. Dev.)	Obs	Mean	(Std. Dev.)	Obs.
Commercial activity	0.492	(0.260)	384	0.605	(0.127)	76	0.355	(0.222)	270	0.453	(0.249)	730
Occupation												
1	0.239	(0.112)	384	0.163	(0.025)	76	0.253	(0.094)	270	0.236	(0.103)	730
2	0.290	(0.173)	384	0.125	(0.055)	76	0.310	(0.183)	270	0.280	(0.177)	730
3	0.437	(0.190)	384	0.630	(0.063)	76	0.407	(0.193)	270	0.446	(0.193)	730
4 (lowest)	0.035	(0.034)	384	0.082	(0.031)	76	0.030	(0.050)	270	0.038	(0.043)	730