Universidad del País Vasco/Euskal Herriko Unibertsitatea

Departamento de Fundamentos del Análisis Económico II

2006-2007

The impact of firm-level contracting on wage levels and inequality: Spain 1995-2002

A Master Thesis

Submitted for the Degree

MASTER IN ECONOMICS: EMPIRICAL APPLICATIONS AND POLICY

by

Ainara González San Román

Supervisor: Sara de la Rica

Bilbao, Spain

THE IMPACT OF FIRM-LEVEL CONTRACTING ON WAGE LEVELS AND INEQUALITY: SPAIN 1995-2002

by
Sara de la Rica
(University of the Basque Country and IZA)
Ainara González San Román
(University of the Basque Country)

Abstract

This paper examines the variations over time in collective bargaining in Spain and their effect on wages by comparing the first two waves of a large matched employer-employee data set (Structure of Earnings Survey (EES-95 and EES-2002)). Using both wage determination models and counterfactual distributions we analize the changes over time in the wage premium associated with firm-level contracting from 1995 to 2002, as well as its impact on wage inequality. For men the firm contract wage premium decreased only slightly over this seven-year period but we find a larger drop in the average firm contract wage premium for women. In regard to wage inequality, for men firm contracts are found to reduce the variance of wages only in 1995 whereas for women the opposite effect is observed in both periods. In addition to controls for individual, job and co-worker characteristics, we make use of the Spanish Structural Business Statistics (SBS) survey to add financial information at industry level. With this additional information, we improve the previous estimator of the propensity score and test the robustness of our findings. Finally, we find no contribution of the observed centralization in collective agreements to the slight increase in wage dispersion in Spain over this seven-year period.

1 Introduction

Several recent papers have analized the effects of unions on the wage structure by comparing wages in unionized and non-unionized labor markets. Most of the empirical evidence refers to Anglo-Saxon countries[1] because there the non-unionized sector represents a suitable counterfactual. However, most European countries are not characterized by "closed shop" labor markets and collective bargaining agreements extend rather to unionized and non-unionized workers. Furthermore in some European countries, such as Spain, Italy and the Netherlands, collective bargaining takes place simultaneously at firm level and sectoral level (on regional and national scales). Firm-specific contracts override the prevailing sectoral agreements and provide an opportunity for workers to improve the labor conditions attained at more centralized levels. Thus, the relevant question in these European countries is not how unionized and non-unionized labor markets compare, but rather how more and less centralized collective bargaining systems compare.

In this setting, there are many macroeconometric studies that analyze the effect of wage contracting centralization on the economic performance of a country, though no clear conclusion has been reached. For instance, Bruno and Sachs (1985) and Jimeno (1992) support the hypothesis of a linear relationship between centralization and macro results while Calmfors and Driffil (1988) defend the inverted-U shape with centralized and decentralized economies both performing better than their intermediate competitors. However, there is much less microeconometric evidence on how the level of contracting affects the structure of wages, due in part to the lack of suitable data. Several authors have found a statistically significant premium associated with firm-level contracting. Dell'Aringa and Lucifora (1994) find a positive effect of this type of contracting on wages for a sample of manufacturing plants in Italy. Card and de la Rica (2006) find that firm-level contracting is associated in Spain with a 5-10% wage premium. On the other hand, Hartog, Leuven and Teulings (2002) find no premium for the Netherlands.

[1] See for instance, Lewis (1986) who concludes that the average union effect on wages is positive in the US. See Card, Lemieux and Riddell (2002) for a study of the effect of unions on the wage structure of Canada, the UK and the US. Hirsch (2003) also finds higher premiums after correcting for estimation biases.

This paper provides microeconomic evidence on the variation over time of the firm-specific wage premium in Spain from 1995 to 2002, and its impact on wage inequality. We make use of two waves of a detailed linked employer-employee data set. In addition, a new data set with financial information on firms is used for 2002 to control as flexibly as possible for differences in the performance of firms (aggregated at industry level). To our knowledge, there is no microeconomic evidence on the dynamics of the firm-specific wage premium for Spain or for any other country with a similar institutional setting. Our results suggest that there is a clear tendency towards centralization in the collective bargaining process in Spain over this seven-year period, that the firm-level contract wage premium undergoes a substantial decrease, particularly for women, and finally that the "centralization" observed in the collective bargaining process has not resulted in any significant change in wage inequality.

The rest of the paper is organized as follows. Section 2 provides a description of the major institutional aspects of the Spanish labor market in recent years. Section 3 briefly describes the data used in the empirical analysis, presents some descriptives and carries out a semiparametric procedure to analyze the impact of firm-level contracting on the distribution of wages as a whole. Section 4 describes the model of earnings used to carry out the analysis. Section 5 presents the empirical findings from the estimation of the different regression models, with sub-sections on changes in the premium by types of worker, and compares OLS estimates with different propensity score estimates. Section 6 analyzes how the centralization in collective agreements affects wage dispersion. Section 7 concludes the study with a summary of our findings.

2 Institutional Framework

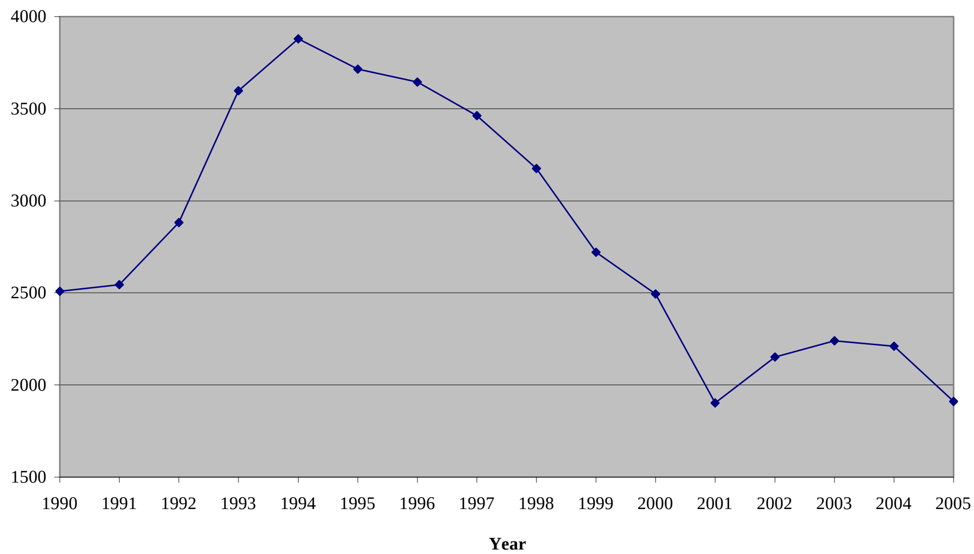
During the Franco dictatorship the Spanish labor market was centralized and highly regulated. Collective bargaining was already established through the old labor laws (*Ordenanzas Laborales*). Legal trade unions and employers' organizations negotiated contracts but the State intervened directly to determine the outcomes of the negotiation (Milner and Metcalf 1994). The post-Franco Spanish collective bargaining system is based on principles resulting from the 1980 Workers' Statute (*Estatuto de los Trabajadores*), which was the result of a consensus between unions and employers' organizations[2]. A system for the election of workers' representatives was established in order to form workers' councils and negotiate issues such as productivity, wages and working hours. In this context, the terms of the agreements reached were and still are legally binding on all workers within the scope of the agreement. Thus, bargaining coverage is very high in Spain, especially relative to the low union density. In 1995 union density was 19% but 81% of the work-force was covered. Both rates have decreased over time; in 2002 union density was 14.9% and the agreements reached covered 72% of workers.

From a macroeconomic point of view, the slow economic growth of Spain since 1990 demonstrates the limits and the negative effects of the political measures on employment taken during the eighties. The worldwide recession that began in the early nineties hit Spain around 1993. Unemployment increased considerably to 25% of working-age population (See Figure 1). Inflation also rose to the highest level of the decade. At that time, it became necessary to encourage employment and promote collective bargaining, requiring the attribution to the unions of a genuine function as a regulator of labor relations. The reform of 1994 repealed the old labor laws, encouraging the use of contractual instruments such as firm-level contracts in order to decentralize bargaining (Toharia and Malo 1997). In both 1997 and 2001 social actors (*employers" and union representatives*) signed the Interconfederal Agreements on Collective Bargaining (*Acuerdo Interconfederal de la Negociación Colectiva*), which were focused on increasing coordination in the collective bargaining mechanism.

[2] See Bentolila and Jimeno (2002) for a description of the evolution of the legislation on collective bargaining in Spain since the eighties and Valdés dal Ré (2006) for a review of the 25 years of the Workers' Statute.

As from 1996 the economy started to recover, with the best period of expansion being 2000, when 5.0% growth was achieved. Employment started to increase and the inflation rate was stabilized[3]. After that a decreasing trend set in for the annual growth rate until 2002, when growth (2.7%) was basically the same as in 1995. Therefore, 1995 and 2002 are characterized by the same GDP growth rate but not by the same expectations; in 2002 the trend and the future prospects were clearly better than 7 years earlier. Obviously, the economic climate affects wage results from bargaining and the behaviour of unions and employers may therefore be expected to differ under such different situations.

Figure 1. Total Unemployment (millions of people)



For our analysis, in addition to a knowledge of the macroeconomic situation in Spain we also need to analyze the dynamics of wage contracts from 1995 to 2002. Does the labor market in Spain tend towards a decentralized wage setting or a centralized one? The level of firm collective bargaining measured as the number of workers affected by firm-level contracting as a percentage of the total number of affected workers was stable in the eighties (Milner and Metcalf 1994) but declined slightly at the beginning of the nineties.

[3]In spite of the economic expansion in recent years Spain has not undergone a significant increase in workers' purchasing power, due in part to political measures to encourage employment but at the same time restrain wages.

Table 1 presents the number of firm-level collective agreements and the number of affected workers, plus the number of affected workers by firm-level contracting as a percentage of the total. Although the number of firm agreements increased in 2002, the percentage of workers covered by firm-specific contracts decreased from fourteen to ten percent. Collective bargaining in Spain is at an intermediate level of centralization in comparison with other developed countries, but we find a tendency towards a centralized wage setting scheme over the last ten years.

Table 1. Evolution of Collective Bargaining in Spain 1995-2002

	Total Agreements		**Firm Collective Agreements**		
	Collective Agreements (thousands)	Affected Workers (thousands)	Collective Agreements (thousands)	Affected Workers (thousands)	Affected Workers (%)
1995	4.827	7.605,1	3.461	1.043,7	13,72
1996	5.028	8.128,2	3.661	1.061,5	13,06
1997	5.040	8.365,1	3.669	998,3	11,93
1998	5.091	8.750,6	3.690	1.021,5	11,67
1999	5.110	9.008,1	3.704	1.063,7	11,81
2000	5.252	9.230,4	3.849	1.083,3	11,74
2001	5.421	9.496	4.021	1.039,5	10,95
2002	5.462	9.696,5	4.086	1.025,9	10,58

Source: Boletín de Estadísticas Laborales

As Disney, Gosling and Machin (1996) point out, an employer´s decision to accept a firm-specific contract is comparable to the voluntary union recognition process in the United Kingdom. The decline in the percentage of workers covered by firm-specific agreements could be compared with the decrease in the unionization rate in the Anglo-Saxon countries. For instance, Dinardo, Fortin and Lemieux (1996) find de-unionization in the US to be an important factor in explaining the rise in wage inequality during the eighties. Related to this study, Gosling and Lemieux (2001) argue that the steeper decline in unionization in the UK explains why wage inequality increased faster than in the US. In section 6 we analyze how the decrease in the number of workers covered by firm collective agreements affects wage inequality.

3 Data and Descriptives

The primary source of data for our analysis lies in the Spanish Structure of Earnings Surveys (SES) of 1995 and 2002, large-sample surveys of wages, job information and worker characteristics. These datasets consist of random samples of workers from establishments with 10 or more employees in the private sector, which accounted for approximately 70 percent of the working population in Spain in both 1995 and 2002 in the manufacturing, construction, trade and service industries[4]. They are sampled in two stages. In the first stage, establishments are selected randomly and stratified by region and size. In the second stage, a sample of workers from each of the selected establishments is also randomly selected[5]. Moreover, in order to add financial information on the establishments we use the Spanish Structural Business Statistics (SBS) survey and match SES and SBS at the level of the 2-digit NACE industry because there is no information available at establishment level. We add this information only for 2002 due to the lack of financial data for 1995.

We focus on the subsample of full-time workers to ensure that our findings are comparable to those of other studies in the relevant literature. This gives a sample of 130,170 workers and 14,347 establishments for 1995 and a sample of 129,378 workers and 14,778 establishments for 2002. There is almost the same number of total observations in both years but the percentage of firm-specific contracts is much lower in 2002, as mentioned before (dropping from 23% to 10%). Since the structure of pay in Spain is similar under regional and national contracts, as Card and de la Rica (2006) show using the EES-95, we group the two together and concentrate on the difference between firm-specific and sectoral contracts. Table 2 presents weighted[6] descriptive statistics for the overall sample and for workers covered by the two different levels of contracts for each year.

[4] In 2002 there are also establishments in the fields of education, health and social work, but we exclude these from the sample for comparability reasons.

[5] An average of 5 workers are interviewed in establishments with 10-20 employees, 7 workers in establishments with 21-50 employees, 12 workers in establishments with 51-100 employees, 20 workers in establishments with 100-200 employees, and 25 workers in establishments with more than 200 employees.

[6] In our analysis we use sampling weights for each worker that reflect the relative probabilities of sample selection for different establishments.

Table 2. Workers' Characteristics by Type of Contract 1995-2002

	All		*Firm-Specific*		*Other Contract*	
Independent variable	**1995**	**2002**	**1995**	**2002**	**1995**	**2002**
Fraction Male	0.77	0.69	0.83	0.76	0.76	0.68
Age Distribution						
Under 30	0.25	0.29	0.15	0.21	0.27	0.30
30-44	0.44	0.44	0.43	0.42	0.45	0.45
45-55	0.23	0.20	0.33	0.28	0.20	0.19
Over 55	0.08	0.06	0.09	0.09	0.07	0.06
Education Distribution						
Primary	0.34	0.30	0.35	0.21	0.34	0.31
Secondary	0.55	0.57	0.51	0.61	0.57	0.57
University	0.10	0.13	0.14	0.18	0.09	0.13
Fraction Temporary Contracts	0.27	0.27	0.10	0.11	0.32	0.29
Mean Tenure (years)	10.68	7.18	15.36	13.24	9.30	6.54
Establishment Size Distribution						
11-20	0.20	0.44	0.05	0.08	0.24	0.48
21-50	0.26	0.15	0.12	0.08	0.30	0.16
51-100	0.17	0.09	0.15	0.06	0.17	0.09
101-200	0.15	0.03	0.21	0.02	0.13	0.03
Over 200	0.22	0.28	0.46	0.75	0.16	0.23
Industry Distribution						
Manufacturing	0.65	0.48	0.84	0.59	0.59	0.47
Construction	0.07	0.10	0.01	0.01	0.08	0.11
Trade	0.09	0.12	0.04	0.10	0.11	0.12
Hotels	0.05	0.07	0.01	0.02	0.06	0.08
Transportation	0.04	0.06	0.05	0.17	0.03	0.05
Financial Services	0.06	0.06	0.03	0.04	0.07	0.06
Other Services	0.04	0.11	0.02	0.07	0.05	0.11
Product Market Orientation						
Local-Regional	0.86	0.88	0.77	0.75	0.88	0.89
International	0.14	0.12	0.23	0.25	0.12	0.11
Occupational Distribution						
Managers and Technicians	0.15	0.19	0.21	0.31	0.13	0.18
Clerical Workers	0.14	0.12	0.12	0.10	0.15	0.12
Service Workers	0.07	0.11	0.02	0.08	0.08	0.11
Qualified Manual Workers	0.48	0.42	0.53	0.42	0.47	0.42
Non-qualified Manual Workers	0.12	0.14	0.08	0.05	0.13	0.15
Number of Observations	130,170	129,377	29,599	12,512	100,571	116,866

Notes: Samples are weighted and include all full time workers with valid information on key variables in EES-95 and EES-02 for each year respectively.

If we look at the overall sample we can see that in 2002 there is higher proportion of women and young workers. That may be why we find a much lower average number of years of tenure in 2002 than in 1995. We also observe that workers are more highly educated in 2002. While the establishment size distribution was equally distributed in 1995, seven years later we find an extreme size distribution with either very small or very large establishments. There are no significant differences between the two years in other characteristics, except for the lower incidence of manufacturing firms and the higher proportion of trade and services in 2002.

If we compare workers covered by the two different types of contract we can see that employees at establishments with firm-specific contracts are more highly educated, older, have a higher average job tenure and are more likely to be employed on a permanent basis. In addition, men are more likely to be covered by firm-specific contracts than women. Establishments with firm-specific contracts are larger and there is a higher proportion of manufacturing firms and plants with an international market orientation for their products. This is so for both 1995 and 2002. If we compare the characteristics of firms covered by firm-specific contracts in the two time periods, we can observe that in 1995 around half the employees covered by firm-level contracts were working at establishments with 200 or more workers, but by 2002 this figure had increased to two thirds. It can also be observed that in 2002 there was a higher proportion of firms working in transport services offering firm-specific contracts. Finally, in 2002, the average skill level of workers covered by firm-specific contracts was higher, as the percentage of managers and technicians under firm-level contracting was clearly higher than in 1995.

The dependent variable for this analysis is the logarithm of the gross wage expressed in euros per hour. Gross wage is defined as the sum of the gross base wage and gross wage complements. Base wages are determined from the corresponding contract by occupation and grade within a firm, while wage complements are defined as the set of payments above the base wage, and include factors such as seniority as well as discretionary suplements such as incentives for productivity, attendance and punctuality awarded to individual employees. Our measure does not include bonuses for nightshifts[7] or extraordinary payments. Table 3 presents the mean and the standard deviation of log hourly wages for 1995 and 2002.

[7] Generally, young men are more likely to have this kind of complement, so by including them in the hourly wage we are making a distintion between groups of workers.

Table 3. Mean Log Wages by Type of Contract 1995-2002

	Mean Log Wage	*Standard Deviation*	*Standardized Mean Log Wage*	*Standardized Standard Deviation*
Group		**1995**		
All Workers				
Overall	1.68	0.51		
Firm Contract	1.95	0.47	1.84	0.50
Other Contract	1.61	0.50	1.73	0.50
Men				
Overall	1.74	0.51		
Firm Contract	1.99	0.46	1.89	0.49
Other Contract	1.67	0.50	1.78	0.50
Women				
Overall	1.47	0.45		
Firm Contract	1.76	0.49	1.64	0.49
Other Contract	1.42	0.43	1.52	0.44
Group		**2002**		
All Workers				
Overall	1.70	0.50		
Firm Contract	1.99	0.52	1.95	0.52
Other Contract	1.66	0.48	1.85	0.54
Men				
Overall	1.77	0.51		
Firm Contract	2.07	0.48	2.03	0.49
Other Contract	1.73	0.50	1.93	0.54
Women				
Overall	1.53	0.45		
Firm Contract	1.74	0.55	1.72	0.52
Other Contract	1.51	0.43	1.66	0.48

Notes: For 1995, samples are 130.170 for all workers,100.533 for men and 29.637 for women. For 2002, samples are 129.377 for all workers, 89.320 for men and 40.056 for women.

The first two columns show the unadjusted sample moments, while columns (3) and (4) show adjusted moments obtained using the re-weighting technique of DiNardo, Fortin and Lemieux (1996) based on separate probit models for the probability of employment in each of the two sectors. This procedure adjusts the sample for differences in observed characteristics in such a way that the weighted distribution of skill characteristics is the same in each sector as in the overall sample[8]. In order to measure wages in real terms,

[8] The idea behind this semiparametric procedure is to allocate a lower weight to individuals who are overrepresented in the subsample of workers covered by non firm-specific contracts and viceversa.

2002 wages are deflated to their 1995 value.

The information in the table above gives us some indication of how mean wages and inequality varied from one contracting level to another over time. Starting with mean wages, we observe that unadjusted mean wages are systematically higher in 2002, except for women covered by firm-specific contracts. In both periods mean log hourly wages are higher under firm-level contracts. Furthermore, the raw firm-specific wage premium does not change substantially for men, but it decreases by around 30% for women over the seven-year period. Looking at the standardized moments, we find that estimated wages under firm-level contracts are 11-12% higher for both men and women in 1995 and 6-10% higher in 2002. These results suggest that on average workers with firm-specific contracts have higher skills.

Various issues emerge in regard to wage dispersion: First, looking at all workers in 1995, it can be seen that there is a lower variance associated with firm-specific contracts than with other contracts, similar to the findings of Plasman, Rusinek and Rycx (2006) for Spain with the same database (EES-95). However, this result is reversed in 2002. If we dissagregate the sample of workers by gender, it can be seen that the change in the dynamics of wage dispersion is completely driven by the substantial increase in the dispersion of wages of those women covered by firm-specific contracts. This result is similar to the findings of Card (2001) for US women.

To give a better picture of the relationship between contract type and wages for different types of workers, Figures 2 and 3 plot mean log wages for men and women for workers in different age-education cells[9] for the two periods. In doing this, we restrict individual heterogeneity in observed skills (age and education) of workers covered by different types of contracts. If mean wages were the same in the two sectors for workers with similar observed skills, the points would lie along the reference 45 degree line. Examination of the graphs shows that wages are higher under firm-level contracts than under sectoral contracting in both periods for both men and women (as we saw in Table 3). Furthermore, wage dispersion seems to have increased considerably for both genders over the seven-year period, and the increase is higher for more highly-skilled workers, particularly for women. This suggests the existence of a more heterogeneous group of women in the 2002 sample.

[9]Specifically, workers are divided into 56 cells using eight age categories and seven education ranges.

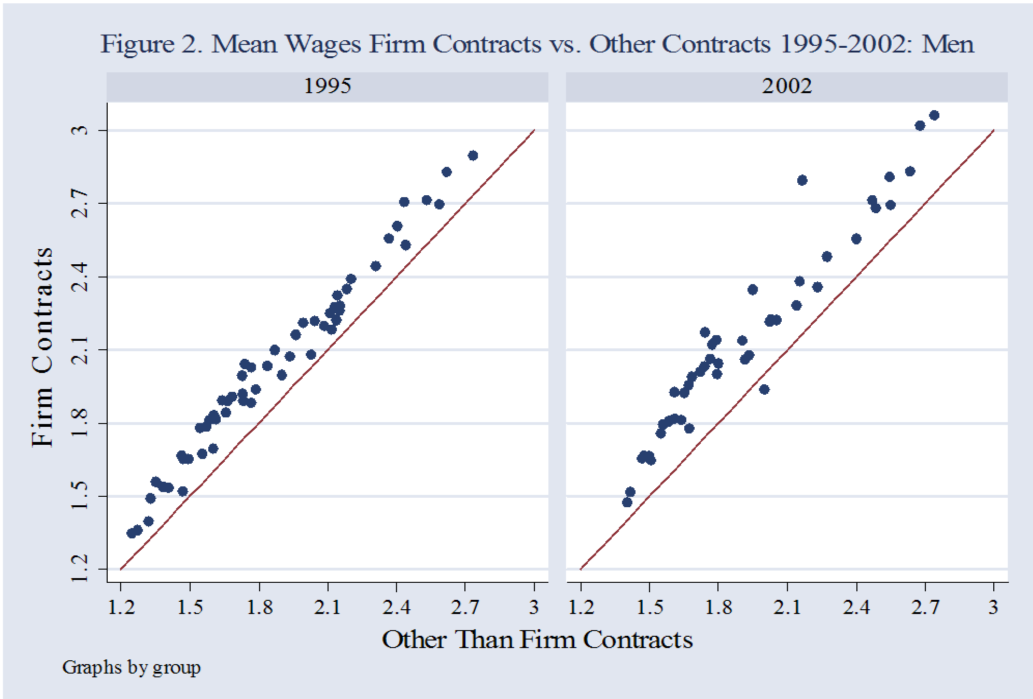


Figure 2. Mean Wages Firm Contracts vs. Other Contracts 1995-2002: Men
1995
2002
Firm Contracts
1.2
1.5
1.8
2.1
2.4
2.7
3
Other Than Firm Contracts
Graphs by group

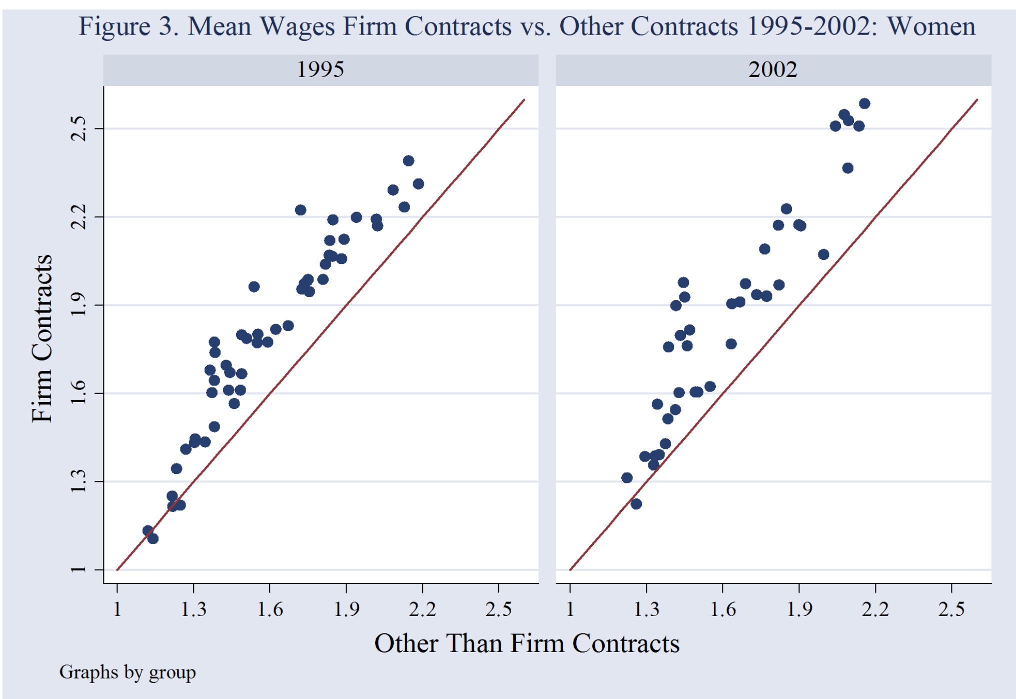


Figure 3. Mean Wages Firm Contracts vs. Other Contracts 1995-2002: Women
1995
2002
Firm Contracts
1
1.3
1.6
1.9
2.2
2.5
Other Than Firm Contracts
Graphs by group

Finally, to offer a broader view of the dynamics of wage dispersion over the seven year period by type of contract, Figures 4 and 5 plot the whole distribution of wages for both periods by gender[10]. The two figures plot both the actual and counterfactual distributions. The latter is the distribution that would prevail in the absence of firm-level contracts. It is obtained using a semiparametric method to estimate the counterfactual distribution of wages. The procedure was developed by Dinardo, Fortin and Lemieux (1996) and consists of a simple reweighting of the distribution of wages of workers covered by non firm-level contracts (See Appendix for further details). To carry out the analysis, we focus on a general method for describing changes in the whole distribution of wages such as kernel density estimates.

From the graphs it can be seen that firm-level contracts tend to move the distribution of wages rightwards for both men and women. This effect is more pronounced in 1995, which is consistent with the higher average wage premium associated with firm-level contracts in that year. In fact, the actual and counterfactual distributions of wages for 2002 are almost identical. While a clear positive relationship between firm-specific contracts and wage levels can be extracted from the figures, no clear relationship can be found between firm-specific contracts and wage dispersion.

This descriptive evidence seems to contrast with other evidence on the effect of the level of collective bargaining on wage dispersion for some European countries. Plasman, Rusinek and Rycx (2006) and Dell'Aringa and Lucifora (1994) conclude that firm-specific contracts contribute significantly to a reduction in wage dispersion in France and Italy, respectively. By contrast, other authors such as Teulings and Hartog (1998) or Blau and Kahn (2002), show that the more centralized wage setting, the more compressed the wage distribution is. If we compare analysis of this type with the effect of unions on the distribution of wages for the Anglo-Saxon countries, most of them find that unionization tends to compress the distribution of wages[11].

[10] Previous studies for the Anglo-Saxon countries, such as Dinardo, Fortin and Lemieux (1996) have shown that the effect of unions is best captured by modelling the whole distribution of wages.

[11] Metcalf (1982) and Freeman (1993) show that unions tend to reduce wage inequality among UK and US males respectively. Lemieux (1997) finds no significant effect of unionization on wage inequality among US women. Other studies by Dinardo, Fortin and Lemieux (1996) and Gosling and Lemieux (2001) find that de-unionization contributes to the increase in wage inequality in both the US and the UK.

Figure 4. Effect of Firm Contracts on the Distribution of Wages 1995

Figure 5. Effect of Firm Contracts on the Distribution of Wages 2002

Next, we look at wage estimation models to quantify the effect of firm-specific contracts on wages and their dynamics over this seven year period. In the next section we describe the econometric model used here to explain the wage determination process and test the descriptive facts previously observed.

4 The model

To quantify collective bargaining wage premiums, we consider a wage equation for worker i at establishment j of the following form,

$$w_{ij} = \beta X_i + \delta_j + \mu_i + \varepsilon_{ij} \tag{1}$$

where w_{ij} is the log hourly wage of individual i, X_i is a set of observed characteristics of workers (such as age, education, job tenure and contract status) associated with a vector of coefficients β, δ_j represents the wage premium earned by workers due to their belonging to establishment j, μ_i represents the unobserved characteristics of individual i, such as ability. Finally, ε_{ij} is the error term, which is assumed to follow a normal distribution. To model the effect of firm-level contracting it is assumed that the wage premium earned by workers at establishment j, that is, δ_j, depends on four factors: a dummy variable for the presence of a firm-specific contract at the workplace (F_j), the average characteristics of the workforce at the establishment (X_j), a vector of observed chacacteristics of firms (Z_j) such as size and market orientation of products, and the unobserved component of firms (v_j), e.g. their profitability.

These assumptions lead to a model for individual wages of the following form,

$$w_{ij} = \beta X_i + \alpha F_j + \eta X_j + \gamma Z_j + \mu_i + v_j + \varepsilon_{ij} \tag{2}$$

The main focus of this paper is on estimating the impact of firm-level contracting on wages, α. Is OLS the most appropriate way to estimate that coefficient in this context? If μ_i and v_j are uncorrelated with F_j, conditional on the observed worker and firm characteristics, α can be consistently estimated by OLS applied to equation (2). But generally, as pointed out by Lewis(1986), such estimates suffer from upward bias resulting from the omission of unobserved variables correlated with contract status. If this is the case, the estimation of α by OLS will include the true effect as well as the bias factors arising from both worker and firm unobserved heterogeneity.

Therefore, the relevant question when estimating the firm-level contract effect is how to eliminate these biases. This would be easier if we had longitudinal data[12], but alternative solutions can be found that exploit the advantages of our data, such as the availability

[12] See Blanchflower (1999) for an estimation of the fixed effects of the union premium on wages.

of information on the characteristics of co-workers (meaning workers in the same firm with the same occupation). If we assume that workers with higher unobserved ability tend to have co-workers with higher average skill levels, some of the effects of unobserved ability can be eliminated by controlling for average characteristics of co-workers.

In order to deal with the unobserved heterogeneity of firms, we control for the full set of observed firm-level characteristics that determine the level of bargaining[13]. As shown by Imbens (2004), if contract status is ignorable conditional on the observed control variables, then conditioning on the probability that establishment j has a firm-level contract will eliminate any bias in the estimation of α arising from the correlation of contract status with firm characteristics. To implement this idea, a probit model for the probability of a firm-level contract is fitted, the predicted probability is estimated and then included as a polynomial function in the individual earnings equation. Using this predicted probability all the relevant information is condensed into the one dimensional propensity score.

5 Empirical estimation of the Firm-level contract premium

5.1 Standard regression models

Tables 4 and 5 present a series of individual regression models for men and women respectively for the two years. The first model includes only a dummy variable which takes value one if the worker is covered by a firm-specific contract. Column (2) incorporates observed worker and firm characteristics such as age, education, contract status, market orientation, firm size and public ownership status of the firm. Finally, the third model also includes the average characteristics of co-workers at the same establishment in the same one-digit occupational group[14].

[13] For 1995 the same information used by Card and De la Rica (2006) is included, but for 2002 we also add financial information at the industry level.

[14] Columns (1), (2) and (3) for 1995 were reported in Card and De la Rica (2006), but we show them here for the sake of comparability. Models which include mean co-worker characteristics in narrower occupational subgroups were fitted by Card and de la Rica (2006) for the EES-95 and no significant differences were found.

Table 4. Log Wage Regressions for MEN

	1995			**2002**		
Independent Variable	*(1)*	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*
Firm Contract	0.317	0.082	0.075	0.335	0.072	0.069
	(0.013)	(0.012)	(0.009)	(0.022)	(0.018)	(0.016)
Worker´s Characteristics						
Education		0.026	0.024		0.023	0.021
		(0.001)	(0.001)		(0.001)	(0.001)
Age		0.011	0.010		0.011	0.010
		(0.0002)	(0.0002)		(0.0002)	(0.0002)
Temporary Contract		-0.199	-0.184		-0.106	-0.102
		(0.006)	(0.006)		(0.007)	(0.006)
Firm´s Characteristics						
International Market		0.025	0.022		0.076	0.076
		(0.013)	(0.010)		(0.013)	(0.011)
Publicly Owned		0.054	0.054		0.039	0.034
		(0.055)	(0.035)		(0.030)	(0.024)
20-50 Workers		0.070	0.070		0.086	0.085
		(0.009)	(0.008)		(0.009)	(0.008)
51-100 Workers		0.134	0.134		0.102	0.102
		(0.012)	(0.009)		(0.015)	(0.013)
101-200 Workers		0.149	0.151		0.149	0.151
		(0.013)	(0.010)		(0.022)	(0.019)
Over 200 Workers		0.219	0.220		0.156	0.156
		(0.013)	(0.011)		(0.017)	(0.014)
Average Characteristics of Coworkers						
Education			0.004			0.003
			(0.002)			(0.002)
Age			0.003			0.003
			(0.0009)			(0.001)
Proportion under 30			-0.088			-0.035
			(0.017)			(0.019)
Proportion over 50			-0.111			-0.052
			(0.020)			(0.024)
Proportion Female			-0.045			0.008
			(0.016)			0.021
Intercept	1.672	1.244	0.743	1.735	1.372	0.809
	(0.007)	(0.022)	(0.039)	(0.007)	(0.023)	(0.046)
R-Squared	0.068	0.506	0.509	0.045	0.474	0.477

Notes: Sampling weights used for estimation. For 1995, samples are 100.533 for men. For 2002, samples are 89.320 for men. Standard errors are calculated with clustering by firms. All models except (1) also include controls for occupation, industry and region.

Table 5. Log Wage Regressions for WOMEN

	1995			2002		
Independent Variable	*(1)*	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*
Firm Contract	0.343	0.126	0.108	0.238	0.043	0.026
	(0.024)	(0.015)	(0.012)	(0.048)	(0.030)	(0.026)
Worker´s Characteristics						
Education		0.022	0.017		0.024	0.016
		(0.001)	(0.001)		(0.002)	(0.003)
Age		0.010	0.009		0.010	0.007
		(0.0004)	(0.0004)		(0.0005)	(0.0004)
Temporary Contract		-0.173	-0.157		-0.109	-0.105
		(0.011)	(0.012)		(0.007)	(0.007)
Firm´s Characteristics						
International Market		0.033	0.026		0.124	0.113
		(0.016)	(0.015)		(0.020)	(0.019)
Publicly Owned		0.124	0.137		0.012	-0.0097
		(0.041)	(0.033)		(0.043)	(0.039)
20-50 Workers		0.029	0.025		0.071	0.061
		(0.011)	(0.011)		(0.011)	(0.009)
51-100 Workers		0.081	0.067		0.064	0.052
		(0.014)	(0.013)		(0.014)	(0.013)
101-200 Workers		0.120	0.096		0.129	0.110
		(0.017)	(0.015)		(0.019)	(0.018)
Over 200 Workers		0.177	0.149		0.131	0.113
		(0.014)	(0.013)		(0.018)	(0.014)
Average Characteristics of Coworkers						
Education			0.008			0.012
			(0.002)			(0.003)
Age			0.005			0.005
			(0.001)			(0.002)
Proportion under 30			-0.036			-0.060
			(0.021)			(0.024)
Proportion over 50			-0.153			-0.104
			(0.031)			(0.036)
Proportion Female			-0.156			-0.182
			(0.015)			(0.022)
Intercept	1.419	1.11	0.684	1.51	1.182	0.806
	(0.007)	(0.034)	(0.060)	(0.008)	(0.045)	(0.079)
R-Squared	0.074	0.494	0.506	0.022	0.432	0.447

Notes: Sampling weights used for estimation. For 1995, samples are 29.637 for women. For 2002, samples are 40.056 for women. Standard errors are calculated with clustering by firms. All models except (1) also include controls for occupation, industry and region.

From column (1), it can be seen that over the seven-year period the unadjusted firm-specific wage premium remains stable for men (around 32%) but decreases by almost 30% for women (from 34% to 24%). Adding together individual and firm characteristics (column 2) increases the explanatory power of the model considerably and reduces the coefficients, suggesting that a striking proportion of the correlation between firm-level contracts and wages is due to systematic sorting of workers across firms. For men the adjusted premium is still positive and significant in both periods, but for women the average firm contract wage premium decreases from 0.12 in 1995 to zero in 2002. Adding in co-worker characteristics (column 3) decreases the premium slightly for men but the main features remain.

Several additional features arise from Tables 4 and 5. Belonging to a firm with an international product orientation increases wages particularly in 2002. The penalty for holding a fixed-term contract decreases by 5-10% over this seven-year period for both men and women. This could be one of the results of the reforms that took place in Spain during the nineties, which tried to reduce the amount of this kind of contracting but also to improve the conditions of those workers who had a fixed-term contract[15]. Co-worker average characteristics have a significant, positive effect on individual wages. However, the proportion of female co-workers affects wages differently depending on gender and period. For men, working in women's jobs decreased wages in 1995 by around 4 percent, but this penalty had dissappeared by 2002. However, for women, working in women's jobs poses a big, increasing penalty (15 percent in 1995 and 18 percent in 2002).

5.2 Adding the Propensity Score

Our earlier considerations (in Section 3) suggested that the estimated wage effects might be biased due to a non-random selection of firms with unobservable characteristics in the different contracting regimes. Thus, in this paper one of the principal goals is to control for unobserved heterogeneity arising from the side of the firm. Since we have no financial information at establishment level which could give us an idea about the rents of firms in each reference year, we will try to find an alternative solution for this source of heterogeneity. As in Card and de la Rica, following Imbens (2004) this problem can

[15] See de la Rica (2004) for a detailed analysis of the wage gap between workers with indefinite and fixed-term contracts.

be solved in two steps. First, we estimate the probability of a firm offering a firm-level contract using a discrete choice model. This estimated probability for each establishment is asigned to all individuals working at the same firm. In the second step, we introduce this variable as an additional regressor in the individual earnings equation.

Table 6 shows the results of adding a third order polynomial function of the estimated propensity score to the more general wage determination model covered in the previous subsection. Although not reported, due to the similarity of the coefficients, estimations also include all covariates included in model (3) of Tables 4-5, that is, individual observable skills, job characteristics, average skills of co-workers, and indicators for region, occupation and industry.

Table 6. Log Wage Regressions. p-score Added as a Regressor

	1995		**2002**	
Description	*Men*	*Women*	*Men*	*Women*
Firm Contract	0.054	0.067	0.046	0.007
	(0.010)	(0.013)	(0.016)	(0.022)
Propensity Score	0.663	0.019	0.967	0.901
	(0.114)	(0.146)	(0.223)	(0.372)
Squared Propensity Score	-1.269	0.597	-0.736	0.720
	(0.316)	(0.413)	(0.625)	(0.986)
Cubed Propensity Score	0.847	-0.277	-0.019	-1.591
	(0.242)	(0.288)	(0.548)	(0.969)
R-Squared	0.512	0.514	0.480	0.461

Notes: Sampling weights used for estimation. For 1995, samples are 130.170 for all workers, 100.533 for men and 29.637 for women. For 2002, samples are 129.377 for all workers, 89.320 for men and 40.056 for women. Standard errors are calculated clustering by firms.

If we compare Tables 4-5 and 6, adding the estimated probability of a firm-level contract decreases the premium associated to this kind of contracting for both men and women, due in part to the fact that establishments offering firm-level contracts tend to have higher propensities. In 1995 the premium is still lower for men than for women (5-7% respectively). In 2002 the coefficient is lower but significant for men (around 5%) and insignificant for women. The propensity score is statistically significant in all cases. Summarizing, from 1995 to 2002 the premium decreases on average for men and drops to zero for women.

5.3 Analysis by Skill Group

In this section we try to determine whether the effect associated with firm-level contracting is different for different skill groups and whether it depends on the period under study. This issue is even more relevant for the 2002 group of females, because we find that on average there is no significant effect of firm-level contracting on women's wages. To adress these questions we first divide the wage distributions for men and women into percentiles (the 20th, 40th, 60th and 80th) for each year. Ordered probit models (See Appendix Table A1) are then estimated separately by gender and period to predict the probability of a given person being in any one of the five quintiles. These predicted probabilities are used as weights and five separate models for each gender and year are estimated.

Table 7 summarizes the estimated firm-level contract premiums from four different wage determination models for 1995 and 2002. These different models correspond to the specifications in Tables 4-5 and 6, with model (4) being the most accurate specification, adding in co-worker characteristics and a cubic in the estimated propensity score. Several points common to 1995 and 2002 can be adressed from Table 7. For men, the estimated effect of working under firm-level contracts is roughly constant across wage groups when we exclude any other observed covariates. But in specifications (2)-(4) we find a tendency for a higher premium among higher paid workers. The results for women are similar to those for men, although among women there is an even stronger tendency for the firm-level contracting premium to increase across the wage distribution.

A comparison of the two years, looking at the more general specification, shows that premiums in 2002 are in general below those obtained 7 years earlier. For men, the premium ranges from 5% to 9% between the lowest and highest wage quintiles in 1995, whereas in 2002 it goes from 5% to 8%. The premium for the lowest-paid women is almost zero in both periods, but in the top quintile of the wage distribution is nearly 10% in 1995 and 6% in 2002. In addition, while all premiums are statistically significant for the 1995 sample, regardless of gender, after 7 years premium for women are only found to be significant at the upper end of the distribution (quintiles 4 and 5), i.e. for the highest paid females.

Table 7. Estimation of Firm-Contract Impact by Wage Quintile

	Quintile 1	Quintile 2	Quintile 3	Quintile 4	Quintile 5
Model			**1995**		
Men					
Model (1)	0.253	0.267	0.257	0.241	0.222
	(0.009)	(0.007)	(0.007)	(0.008)	(0.010)
Model (2)	0.075	0.095	0.103	0.105	0.102
	(0.008)	(0.007)	(0.007)	(0.007)	(0.009)
Model (3)	0.064	0.085	0.094	0.098	0.098
	(0.008)	(0.007)	(0.007)	(0.007)	(0.009)
Model (4)	0.051	0.072	0.082	0.088	0.088
	(0.009)	(0.007)	(0.007)	(0.008)	(0.011)
Women					
Model (1)	0.168	0.242	0.276	0.294	0.288
	(0.019)	(0.015)	(0.013)	(0.012)	(0.015)
Model (2)	0.056	0.101	0.123	0.134	0.131
	(0.014)	(0.012)	(0.011)	(0.010)	(0.013)
Model (3)	0.037	0.077	0.098	0.112	0.113
	(0.013)	(0.011)	(0.011)	(0.010)	(0.013)
Model (4)	0.003	0.044	0.068	0.086	0.097
	(0.013)	(0.012)	(0.011)	(0.011)	(0.014)
			2002		
Men					
Model (1)	0.244	0.293	0.309	0.310	0.291
	(0.016)	(0.013)	(0.011)	(0.011)	(0.014)
Model (2)	0.077	0.097	0.101	0.098	0.084
	(0.014)	(0.011)	(0.010)	(0.009)	(0.011)
Model (3)	0.068	0.088	0.094	0.094	0.084
	(0.014)	(0.011)	(0.009)	(0.009)	(0.011)
Model (4)	0.051	0.071	0.078	0.081	0.080
	(0.015)	(0.012)	(0.010)	(0.009)	(0.011)
Women					
Model (1)	0.122	0.172	0.219	0.274	0.316
	(0.020)	(0.020)	(0.021)	(0.022)	(0.024)
Model (2)	0.037	0.056	0.070	0.083	0.095
	(0.018)	(0.017)	(0.016)	(0.016)	(0.020)
Model (3)	0.022	0.039	0.052	0.063	0.075
	(0.017)	(0.016)	(0.015)	(0.015)	(0.019)
Model (4)	0.002	0.011	0.023	0.037	0.055
	(0.017)	(0.016)	(0.015)	(0.015)	(0.018)

Notes: Models (1)-(3) corresponds to specifications in Tables 4-5. Model (4) corresponds to specifications in Table 6. Sampling weights used for estimation. Standard errors are calculated with clustering by firms.

This is consistent with the preliminary analysis carried out at the beginning of the paper, where we found the highest wage dispersion in the 2002 sample of women, and higher premiums for older and more highly educated women.

Summarizing, two important conclusions can be drawn. The premium found for firm-specific contracting is higher for more highly paid workers in both periods. This result contrasts with the union wage premium in the United States and United Kingdom, which is generally found to be lower for highly paid workers. In addition, firm-level contracting matters for all skill groups in 1995, but in 2002 it is not significant for less skilled women. In the next section we extend this analysis using the new source of financial data available for 2002.

5.4 Adding in firm-performance information for 2002

5.4.1 Does the premium change when financial information is included?

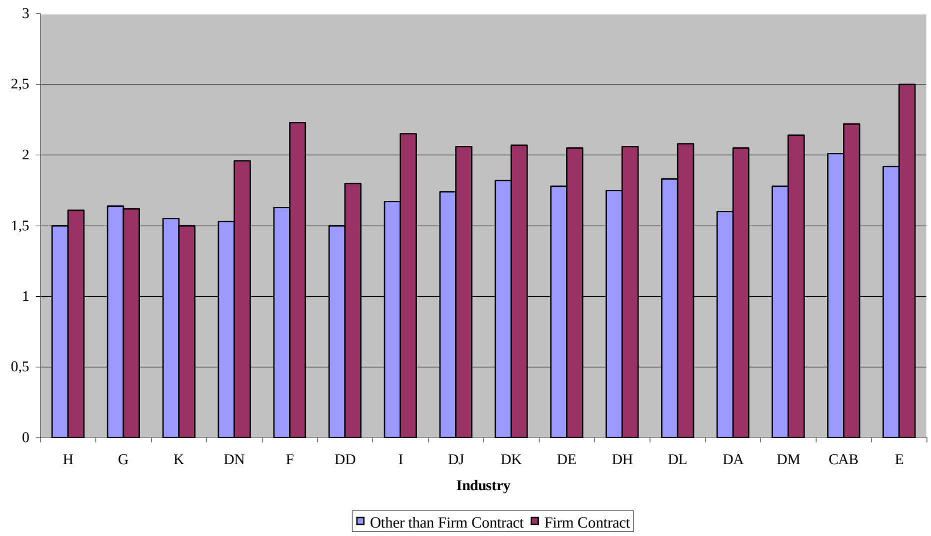
The wage premium could still be reflecting, at least partly, unobservable differences among firms. We have very little information about their economic performance, and these unobserved differences might well be correlated with the firm-specific wage premium. In order to take into account this possible bias, we use a new source of data, the Structural Statistics Business Survey, which allows us to introduce financial information on some industries, although this information is only available for 2002. Thus, in this section we test the sensitivity of the premium estimated to the new data and estimation methods available. We construct some measures of economic performance in line with the relevant literature and then choose the most suitable variables to include in the above estimation of the propensity score[16].

For a preliminary descriptive approach, Figure 6 shows the average log hourly wage by type of contract and economic performance of industry for 2002. Industries are sorted by volume of business (from lowest to highest). (See Appendix for a list of industries). It is revealed that mean wages fluctuate considerably from one industry to another.[17]

[16] Instead of including them in the regression directly, we add them to the propensity score because in that way we eliminate the bias arising from firms' unobserved heterogeneity. (Imbens 2004)

[17] The best paying industry is the electricity, gas and hot water supply sector. The results regarding high and low paying industries are similar to those obtained by Plasman et al. (2006) for Belgium.

Figure 6. Mean Log Wages by economic performance of the industry and type of contract



The figure suggests a positive correlation between workers' wages and firms' ability to pay. Dissagregation of Figure 6 by gender (not reported) reveals that this positive relationship is very similar for both men and women. In addition, as seen above, wages are greater on average when bargaining takes place at firm level, regardless of the industry.

Since the raw data suggest a positive relationship between wages and profits, we set out to check whether the premium associated with firm-level contracting varies when this financial information is included. For our purpose, the most useful variables from this survey are volume of business and the investment rate[18], per employee in both cases, in logarithms and measured in 1995 euros for the sake of comparability. Volume of business is defined as the amounts invoiced by a company during the reference year in return for the provision of services and sale of goods that are the object of its trading. Investment can be described as the transfers carried out during the reference year to obtain goods to be used in the long run in the company's activity.

These variables do not exactly measure the economic performance of industry, but they may be considered as good proxies for it, and from an econometric perspective the

[18] The variable used as a proxy of economic performance by some studies in Belgium and England is the investment rate. Other authors such as Abowd, Kramarz and Troske (2001) and Marsden (2005) use the per capita value added as an instrument for dealing with the problem of endogeneity of profits in the earnings equation.

use of proxies enables us to avoid the endogeneity problem between profits and wages, which are determined simultaneously in the wage equation. Other variables such as gross operating surplus and a productivity indicator are available, but they are highly correlated with investment and give the same estimate of the propensity score. Table 8 presents the results of the estimation when we include these variables together in the propensity score.

Table 8. Turnover and Investment added to the former p-score 2002

Description	*Men*	*Women*
Firm Contract	0.055	0.009
	(0.017)	(0.024)
Propensity Score	0.825	0.679
	(0.162)	(0.322)
Squared Propensity Score	-1.751	-0.181
	(0.473)	(0.844)
Cubed Propensity Score	1.190	-0.072
	(0.375)	(0.617)
R-Squared	0.482	0.447

Notes: Sampling weights used for estimation.Financial variables at the 2-digit industry level. Estimation also includes all covariates of model (3) of Tables 4-5 Standard errors are calculated with clustering by firms.

The first point to be mentioned is that both variables are significant and both increase the probability of having a firm-level contract when they are included in the probit model (See Appendix Table A2). As can be seen from the table above, the earnings regression coefficient associated with firm-level contracting is still significant for men, and it is significantly higher. For women, the premium was not significant before and including profitability of firms does not change that result. Therefore, the main conclusion of this robustness test is that the estimate of the propensity score has improved but the premium has not changed substantially.

The next question to be answered in this section is whether the premium is still higher for highly paid workers with the improvement in the estimation of the propensity score. We carry out the analysis by quintile group for the more general specification (model 4) substituting the new estimate of the p-score. Table 9 shows the results of this estimation.

Table 9. Estimation of Firm-Contract Impact by Wage Quintile for 2002

Model (4)	*Quintile 1*	*Quintile 2*	*Quintile 3*	*Quintile 4*	*Quintile 5*
Men	0.050	0.071	0.078	0.082	0.074
	(0.015)	(0.013)	(0.011)	(0.010)	(0.011)
Women	0.004	0.012	0.024	0.041	0.066
	(0.018)	(0.016)	(0.016)	(0.016)	(0.017)

Note: Sampling weights used for estimation. Specification (4) is estimated with the new propensity score. Standard errors are calculated with clustering by firms.

We find no significant differences when comparing Table 9 with the corresponding rows for 2002 in Table 7. Only the premium for women in the top quintile is significantly higher (up from 5.5% to 6.6%) when the estimate of the propensity score that includes financial information at the industry level is used.

5.4.2 Is there a rent component behind this premium?

There are several possible explanations for the positive wage premium associated with firm-level contracts. Workers covered by firm-level contracts could be required to work harder, raising the possibiltiy of the efficiency wage premium explanation (Akerlof 1982; Weiss 1990). On the other hand, there is a growing literature in favour of the rent-sharing hypothesis, showing that firms share rents with their employees. Some findings from the Anglo-Saxon countries suggest that profitable firms tend to pay higher wages to their workers. For instance, Blanchflower et al. (1996) using data for the US manufacturing sector, shows that an increase in the industry´s ablility to pay rises the long run level of wages. Abowd and Lemieux (1993) find the profit per employee elasticity of wages for the US to be around 0.3. Studies for European countries, such as the analysis by Goos and Konings (2001) also support a positive correlation, finding a profit per worker elasticity of wages of 0.1 for Belgium.

The availability of financial variables for 2002 may help us to test the hypothesis of rent-sharing at least for this period. The rent sharing hypothesis in our framework suggests that, through collective bargaining agreements, unions could be forcing large firms with positive financial results to share the rents obtained during the economic year with their employees. To test for this, we add the two possible measures of economic performance separately to the more general specification of the wage equation used in previous sections, in order to analize the profit per employee elasticity of wages. Interaction terms between profitability and contract status are also included to see whether the elasticity is greater under firm-level contracting. The results are shown in table 10.

Table 10. Log Wage Regressions adding profitability and the interactions 2002

Variable	*Men*				*Women*			
F. Contract	0.055	0.055	-0.347	0.003	0.009	0.009	-0.513	-0.158
	(0.017)	(0.017)	(0.128)	(0.038)	(0.024)	(0.024)	(0.145)	(0.068)
Financial var.								
B. Volume	0.054		0.042		0.034		0.023	
(per employee)	(0.013)		(0.013)		(0.016)		(0.016)	
Investment		0.045		0.041		0.027		0.024
(per employee)		(0.009)		(0.009)		(0.012)		(0.012)
Interactions								
F.Contract * BV			0.088				0.116	
			(0.026)				(0.030)	
F.Contract * Inv.				0.028				0.082
				(0.016)				(0.038)
Propensity	0.878	0.837	0.855	0.826	0.705	0.679	0.657	0.647
	(0.187)	(0.189)	(0.185)	(0.189)	(0.319)	(0.323)	(0.316)	(0.321)
Squared Prop.	-1.887	-1.759	-1.774	-1.698	-0.267	-0.178	-0.094	-0.045
	(0.463)	(0.472)	(0.457)	(0.474)	(0.832)	(0.844)	(0.820)	(0.845)
Cubed Prop.	1.307	1.187	1.112	1.101	0.011	-0.075	-0.303	-0.331
	(0.348)	(0.358)	(0.352)	(0.365)	(0.603)	(0.615)	(0.586)	(0.064)

Notes:Sampling weights used for estimation.Samples are 89.320 for men and 40.056 for women Models used the covariates of specification (4), and include controls for occupation and region. Financial variables at the 2-digit industry level.

We find a positive profit per employee elasticity of wages of around 4-5% for men and 3-4% for women. Moreover, the positive and significant coefficient of the interaction terms (using both measures) reveals a greater elasticity in those establishments that offer firm-level contracts for both men and women. This result is consistent with the existence of a rent-sharing component in the firm contract wage premium in 2002[19]. Card and de la Rica (2006) support the rent-based explanation for the firm contract wage premium for 1995 for men, but not for women.

[19]Nevertheless, we advise some caution with this interpretation given that we cannot control perfectly for firm size (the firm size variable is in intervals only), and firm size is usually correlated with monitoring costs. If monitoring costs are not adequately controlled for, we might not be able to disregard the efficiency wages explanation.

6 The impact of "centralization" on wage inequality

The last issue that we want to address in this paper is whether centralization in the wage bargaining process has any effect on wage inequality. Table 1 reveals that in this seven-year period there has been a clear tendency towards centralization in the collective bargaining process, illustrated by the decline in the fraction of workers covered by firm-specific agreements. Furthermore, Section 3 reports a slight increase in the variance of wages for both men and women. These two facts lead us naturally to ask whether centralization in wage bargaining may account for some of the growth in wage dispersion observed in this seven-year period. Following Gosling and Lemieux (2001) we show the effect of firm contracts on wage inequality by performing a simple variance decomposition. One way of decomposing the variance of wages is the following,

$$Var(w) = \hat{F}Var(w \mid F_j = 1) + (1 - \hat{F})Var(w \mid F_j = 0) + \hat{F}(1 - \hat{F})\Delta^2 \tag{3}$$

where $\hat{F}$ is the fraction of firm-level contracts and Δ is the wage premium associated with firm-level contracting, that is, $E(w \mid F_j = 1) - E(w \mid F_j = 0)$.

Table 11 shows the different elements of this variance decomposition. The information in the 6th row gives us the counterfactual variance that would prevail if the fraction of firm agreements remained at the 1995 level, which is constructed using equation (3) but with $\hat{F}$ being the fraction of firm level contracts in 1995. In this way we can see the net contribution of "centralization" to the increase in the overall variance of wages.

Table 11. Variance decomposition of log wages 1995-2002

	Men			*Women*		
	1995	**2002**	**Change**	**1995**	**2002**	**Change**
1. Variance firm-contract	0.205	0.237	0.032	0.229	0.305	0.076
2. Variance other contract	0.247	0.247	0.000	0.165	0.185	0.020
3. Firm-contract wage differential	0.317	0.335	0.018	0.339	0.239	-0.100
4. Fraction with firm level contracts	0.228	0.116	-0.112	0.153	0.085	-0.068
5. Overall variance	0.255	0.257	0.002	0.189	0.199	0.010
6. Variance with 1995 firm contracts	0.255	0.264	0.009	0.189	0.211	0.022
7. "Centralization" effect			-0.007			-0.012
			(0%)			(0%)

Notes: For 1995, samples are 130.170 for all workers, 100.533 for men and 29.637 for women. For 2002, samples are 129.377 for all workers,89.320 for men and 40.056 for women. The "centralization" effect is computed as the difference between the change in rows 5 and 6.

From the table above a clear conclusion can be drawn. The decline in the number of workers covered by firm-specific agreements did not contribute to the slight increase in wage dispersion in Spain from 1995 to 2002. If anything, centralization decreased wage dispersion, although the effect is very small. This result contrasts with some findings previously quoted for the Anglo-Saxon countries.

7 Conclusions

This paper provides microeconomic evidence on the changes over time in the firm-specific wage premium in Spain from 1995 to 2002, and on its impact on wage inequality. We make use of two waves of a detailed linked employer-employee data set. In addition, a new dataset with financial information on firms is used for 2002 to control as flexibly as possible for differences in firm performance (aggregated at industry level).

Descriptive comparisons across workers suggest that workers covered by firm-level contracts are older, more highly educated and tend to be working in the largest establishments. Looking at the descriptives, wages under firm-specific agreements are systematically higher for both periods. Using different age-education cells, we see that the premium associated with firm contracts is a little higher for older and more highly educated women, particularly in 2002. Stylized facts also reveal a clear tendency towards centralization in the collective bargaining process in Spain over this seven-year period.

Our results conclude that for men the firm-level contract wage premium has slightly decreased over this seven-year period (from 5.4 percent to 4.6 percent). However, one of the most surprising results is the drop in the average firm-level contract wage premium for women: in 1995 the average premium for women was around 7 percent, but it had dropped to zero by 2002. When we look at the overall wage distribution, however, it can be seen that for both men and women in both periods, firm-level contracts tend to raise wages more for more highly paid workers. For female workers in 2002, the firm contract wage premium is only positive for women in the 4th and 5th quantiles of the wage distribution.

For 2002, availability of financial information on firms (at industry level) from the Structure of Business Survey allows us to test the robustness of our findings, once some unobserved firm heterogeneity is removed. Adding in this information the estimate of the propensity score improves but the premium does not change substantially.

We also address the empirical relationship between firm-level contracts and wage inequality. Wage dispersion in Spain increased slightly from 1995 to 2002. A semiparametric method is used to estimate the counterfactual distribution of wages, i. e. the distribution of wages that would prevail in the absence of firm-level contracts. Results suggest that for men, firm-level contracts tend to reduce the variance of wages in 1995, but the effect is not longer found seven years later. By contrast, for women the reverse effect is observed. Finally, the increased centralization in collective agreements has not been found to contribute to the slight increase in wage dispersion in Spain.

APPENDIX

Table A1
Models for the Probability of Being in Different Wage Quintiles

	Men		*Women*	
Variable	**1995**	**2002**	**1995**	**2002**
Age	0.038	0.034	0.041	0.027
	(0.0006)	(0.0006)	(0.001)	(0.001)
Education	0.101	0.099	0.092	0.091
	(0.002)	(0.002)	(0.004)	(0.003)
Temporary Contract	-0.839	-0.469	-0.721	-0.452
	(0.014)	(0.013)	(0.022)	(0.019)
Occupations				
Managers and Technicians	0.601	0.787	1.075	1.133
	(0.023)	(0.024)	(0.044)	(0.033)
Clerical Workers	0.061	0.069	0.368	0.261
	(0.023)	(0.027)	(0.032)	(0.028)
Service Workers	-0.409	-0.236	-0.016	0.243
	(0.029)	(0.028)	(0.037)	(0.029)
Qualified Manual Workers	-0.004	0.204	0.091	0.223
	(0.016)	(0.017)	(0.029)	(0.025)
Pseudo-R^2	0.150	0.124	0.148	0.116
Number of Observations	100.533	89.320	29.637	40.056

Notes: Models are ordered probit models with five ranges based on unconditional quintiles of gender specific wage distribution. Robust standard errors in parentheses.

Industries according to NACE rev.1

A: Agriculture, hunting and forestry.

B: Fishing.

C: Mining and quarrying.

D: Manufacturing

DA: Manufacture of food products, beverages and tobacco.

DD: Manufacture of wood and wood products.

DE: Manufacture of pulp, paper and paper products: publishing and printing.

DH: Manufacture of rubber and plastic products.

DJ: Manufacture of basic metals and fabricated metal products.

DK: Manufature of machinery and equipment, n.e.c.

DL: Manufacture of electrical and optical equipment.

DM: Manufacture of transport equipment.

DN: Manufacturing n.e.c.

E: Electricity, gas and water supply.

F: Construction.

G: Wholesale and retail trade; repair of motor vehicles and personal goods.

H: Hotels and Restaurants.

I: Transport, Storage and Communication.

K: Real Estate, Renting and Business Activities.

Table A2

Probit Model for the Probability of having a firm level contract

Variable	**1995**	**2002**	
Average characteristics of the workforce			
Age	0.162	0.065	0.121
	(0.042)	(0.060)	(0.063)
Squared Age	-0.001	0.001	0.002
	(0.0005)	(0.0007)	(0.0008)
Education	0.065	0.036	0.032
	(0.015)	(0.017)	(0.019)
Occupational distribution by firm			
Managers and Technicians	0.928	1.122	0.693
	(0.215)	(0.266)	(0.272)
Clerical Workers	0.487	0.322	0.933
	(0.222)	(0.315)	(0.273)
Service Workers	-0.428	0.833	0.910
	(0.279)	(0.285)	(0.269)
Qualified Manual Workers	0.521	0.843	0.503
	(0.153)	(0.237)	(0.233)
Product Market Orientation			
International market	0.151	0.252	0.038
	(0.078)	(0.122)	(0.108)
Establishment Size Distribution			
21-50	0.248	0.515	0.606
	(0.114)	(0.078)	(0.088)
51-100	0.912	0.484	0.615
	(0.132)	(0.096)	(0.104)
101-200	1.049	0.499	0.521
	(0.130)	(0.162)	(0.188)
Over 200	2.114	1.788	1.902
	(0.146)	(0.079)	(0.087)
Financial Variables at the Industry Level			
Business Volume per employee			0.344
			(0.085)
Investment per employee			0.541
			(0.075)
Pseudo R^2	0.389	0.336	0.425
Number of establishments	14.347	14.768	14.768

Notes: The individual explanatory variables are averaged over the characteristics of the workforce at each establishment. Financial variables at the 2-digit industry level. Robust standard errors in parentheses.

Re-weighting procedure

Following the notation of the model presented in Section 4, the observed density of wages in the subsample of workers covered by other contracts is given by

$$f(w \mid F_j = 0) = \int f^{oc}(w \mid x) f(x \mid F_j = 0) dx \tag{4}$$

where $f^{oc}(w \mid x) = f(w \mid x, F_j = 0)$. The distribution that would prevail if all workers were paid under the wage structure of workers covered by other contracts would be,

$$f^{oc}(w) = \int f^{oc}(w \mid x) f(x) dx \tag{5}$$

As shown in full detail by Dinardo, Fortin and Lemieux (1996) this last equation can be written as follows,

$$f^{oc}(w) = \int \lambda(x) f^{oc}(w \mid x) f(x \mid F_j = 0) dx \tag{6}$$

A comparison of equations (4) and (6) indicates that the counterfactual density is simply the "reweighted" version of the actual density of wages in the subsample of workers covered by other contracts, where the sample weights are replaced by the weighting factor $\lambda(x)$, which is shown to be equal to $\lambda(x) = \Pr(F_j = 0) / \Pr(F_j = 0 \mid x)$. The numerator is just the proportion of workers covered by non firm-level contracts in the sample and the denominator is the probability of belonging to this subsample of workers conditional on the covariates, which is estimated using a probit model. The covariates used to estimate the probit model consist basically of a flexible functional form of human capital variables. Once the weighting factor is constructed each worker i is weighted by $\lambda_i = \lambda(x_i)$.

REFERENCES

Abowd, John, Francis Kramarz, D.N. Margolis and K.R. Troske. 2001. "The Relative Importance of Employer and Employee Effects on Compensation: A Comparison of France and the United States." *Journal of the Japanese and International Economies*, Vol. 15, pp. 419-36.

Akerlof, George. 1982 "Labor Contracts as Partial Gift Exchange." Quarterly Journal of Economics, Vol. 97, No. 4 (November), pp. 543-69.

Bentolila, Samuel, and Juan F. Jimeno. 2002. "La Reforma de la Negociación Colectiva en España". FEDEA Working Paper 2002-03. Madrid: FEDEA.

Blanchflower, David, Andrew Oswald, and Peter Sanfey. 1996. "Wages, Profits and Rent Sharing" Quarterly Journal of Economics, Vol. 111, No. 1 (February), pp. 227-51.

Blanchflower, David, and Alex Bryson. 2004. "Union Relative Wage Effects in the USA and the UK," *Industrial Relations Research Association*, 2004, pp. 133-140.

Bratsberg, B., and J. F. Ragan. 2002. "Changes in the Union Wage Premium by Industry – Data and Analysis.", *Industrial and Labor Relations Review*, 56 (1) October.

Bruno, Michael, and Jeffrey Sachs. 1985. *Economics of Worldwide Stagflation*. Cambridge, Mass.: Harvard University Press.

Calmfors, Lars and John Driffill. 1988. "Bargaining Structure, Corporatism and Macroeconomic Performance." *Economic Policy*, No. 6 (April), pp. 13-61.

Card, David. 1996. "The Effect of Unions on the Structure of Wages: A Longitudinal Analysis." *Econometrica*, Vol. 64, No. 4 (July), pp. 957-79.

Card, David. 2001. "The Effect of Unions on Wage Inequality in the U.S. Labor Market" i*ndustrial and Labor Relations RevieW*, Vol. 54 No. 2, pp. 296-315.

Card, David and Sara de la Rica. 2006. "Firm-level contracting and the structure of wages in Spain." *Industrial and Labor Relations Review*, Vol. 59, No. 4 (July), pp. 573-89.

De la Rica, Sara. 2004. "Wage gaps between workers with indefinite and fixed-term contracts: The impact of firm and occupational segregation." *Moneda y Crédito*, vol. 219, pp. 43-69.

Dell ´Aringa, Carlo, and Claudio Lucifora. 1994. "Collective Bargaining and Relative Earnings in Italy." *European Journal of Political Economy*, Vol. 10, pp.727-47.

Dell ´Aringa, Carlo, and Claudio Lucifora. 1994. "Wage Dispersion and Unionism: Do Unions Protect Low Pay?" *International Journal of Manpower*, Vol. 15, pp.150-70.

DiNardo, John E., Nicole Fortin, and Thomas Lemieux. 1996. "Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semi-Parametric Approach." *Econometrica*, Vol. 64, No. 5 (September), pp. 1001-44.

DiNardo, John, and Thomas Lemieux. 1997 "Changes in Wage Inequality in Canada and the United States: Do Institutions Explain the Difference?" *Industrial and Labor Relations Review*, Vol. 50, No. 5, pp. 629-51.

Disney, Richard, Amanda Gosling and Stephen Machin. 1996. "What Has Happened to Union Recognition in Britain?." *Economica*, Vol. 63, No. 249 (February), pages 1-18.

Freeman, Richard. 1993 "How much has De-unionisation contributed to the Rise in Male Earnings Inequality?" *Uneven Tides: Rising Inequality in America*. New York: Russell Sage Foundation: 133-63.

Goos, Maarten and Jozef Konings. 2001. "Does Rent-Sharing Exist in Belgium? An Empirical Analysis Using Firm Level Data" *Reflets et Perspectives de la vie économique,* Vol. XL No.1/2, pp. 65-79.

Hartog, Joop, Edwin Leuven, and Coen Teulings. 2002. "Wages and the Bargaining Regime in a Corporatist Setting: The Netherlands." *European Journal of Political Economy*, Vol. 18, pp. 317-31.

Hirsch, B. T., Edward J. Schumacher, "Private Sector Union Density and the Wage Premium: Past, Present, and Future," 2002. Journal of Labor Research, Vol. 22, No. 3, pp. 487-518.

Hirsch, B. T., D. A. Macpherson, and Edward J. Schumacher. 2002. "Measuring Union and Non-Union Wage Growth: Puzzles in Search of Solutions."

Hirsch, B. T. and Edward J. Schumacher. 2002. "Match Bias In Wage Gap Estimates Due To Earnings Imputation.", Mimeograph, Trinity University.

Imbens, Guido. 2004. "Nonparametric Estimation of Average Treatment Effects under Exogeneity: A Review." Review of Economics and Statistics, Vol. 86, No. 4 (November), pp. 4-29.

Jimeno, J. F. (1992): "Las implicaciones macroeconómicas de la negociación colectiva." 1992. *Moneda y Crédito*, 195, pp. 223-281.

Krueger, Alan B., and Lawrence H. Summers. 1988. "Efficiency Wages and the Inter-Industry Wage Structure." *Econometrica*, Vol. 56, No. 2, pp. 259-93.

Lemieux, Thomas, and John Abowd. 1993. "The Effects of Product Market Competition on Collective Bargaining Agreements: The Case of Foreign Competition in Canada." *Quarterly Journal of Economics* 108, November 1993, pp. 983-1014.

Lemieux, Thomas, and Amanda Gosling. 2001. "Labor Market Reforms and Wage Inequality in the United Kingdom and the United States." Chicago: University of Chicago Press for NBER, pp. 275-312.

Lemieux, Thomas. 2002. "Decomposing Changes in Wage Distributions: A Unified Approach." *Canadian Journal of Economics*, Vol. 35 No. 4, pp. 646-88.

Lewis, H. Gregg. 1986. *Union Relative Wage Effects: A Survey.* Chicago: University of Chicago Press.

Linneman, Peter D., Michael L. Wachter, and William H. Carter. 1990. "Evaluating the Evidence on Union Employment and Wages." *Industrial and Labor Relations Review*, Vol. 44, No. 1 (October), pp. 34-53.

Marsden, David and Richard Belfield. 2005. "Intra-establishment Pay Inequality and Performance Effects in Comparative Perspective." *Pay Inequality Economic Performance* papers.

Metcalf, David. 1982. "Unions and the dispersion of Earnings." *British Journal of Industrial Relations,* No. 20, pp. 170-185.

Milner, Simon, and David Metcalf. 1994. "Spanish Pay Setting Institutions and Performance Outcomes." Banco de España Documento de Trabajo No. 9420. Madrid: Banco de España.

Plasman, Robert, Michael Rusinek and François Rycx. 2006 "Wages and the Bargaining Regime under Multi-level Bargaining: Belgium, Denmark and Spain" *Iza discussion paper* No. 1990 (March)

Plasman, Robert, François Rycx and Ilan Tojerow. 2006 "Industry Wage Differentials, Unobserved Ability, and Rent-Sharing: Evidence from Matched Worker-Frim Data, 1995-2002" *Iza discussion paper* No.2387 (October)

Toharia C., Luis, and M. Angel Malo. 1997. "Economía y Derecho del Trabajo: las reformas laborales de 1994 y 1997." *Cuadernos económicos de ICE*, ISSN 0210-2633, No. 63, pp. 155-74.

Valdés dal Ré, Fernando. 2006. "Estatuto de los trabajadores. Edición conmemorativa del 25 aniversario." Ministerio de Trabajo y Asuntos Sociales.

Weiss, Andrew. 1990. *Efficiency Wages: Models of Unemployment, Layoffs, and Wage Dispersion.* Prenceton, N.J.: Princeton University Press.