BANCO DE ESPAÑA

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Juan J. Dolado and Samuel Bentolila (*)

(*) Banco de España and CEMFI, respectively.
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ABSTRACT

This paper analyzes the wage formation process in Spain, taking into account the effect of the institutional change caused by the creation of low-firing-cost, fixed-term labor contracts. An insider-outsider model which assumes that the interests of workers under these contracts are disregarded in wage bargains is set up and tested. The estimation of a wage equation for a sample of large Spanish manufacturing firms yields an estimate of the insider weight and confirms that an increase in the proportion of fixed-term employees increases insiders' wages.

1. Introduction

The unemployment rate in Spain was the highest in the OECD in the 1980s, and it still hovers around 17% nowadays. Recent estimates put the non-accelerating inflation rate of unemployment (NAIRU) around 15% (De Lamo and Dolado, 1991). This persistence reveals the structural nature of the problem, so that labor supply issues in general, and wages in particular, must play a key role in any attempted explanation of the Spanish unemployment experience.

An important result from the previous two Chelwood Gate conference papers on Spain[1] was that large wage hikes were the main factor behind the rise in unemployment in the 1970s, but that in the first half of the 1980s the blame should fall much more on contractionary aggregate demand policies, since real wage growth was quite moderate.

From 1985 to 1990 the unemployment rate fell by 5 percentage points, as employment grew by 3% per year. Apart from a strong surge in aggregate demand, this good performance was helped by low real wage growth, at least in the first few years. This can be better understood with reference to the recent evolution of the wage drift, i.e. the difference between the rates of growth of wage rates agreed in collective bargains and of average earnings. In the 1980s wage drift used to be around 2 percentage points but since 1987 it has been abnormally low (see Figure 1), while at the same time, bargained wage rates have been abnormally high in real terms

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1 Dolado, Malo de Molina, and Zabalza (1986) and Andrés *et al.* (1990).

Figure 1: Bargained and average wages

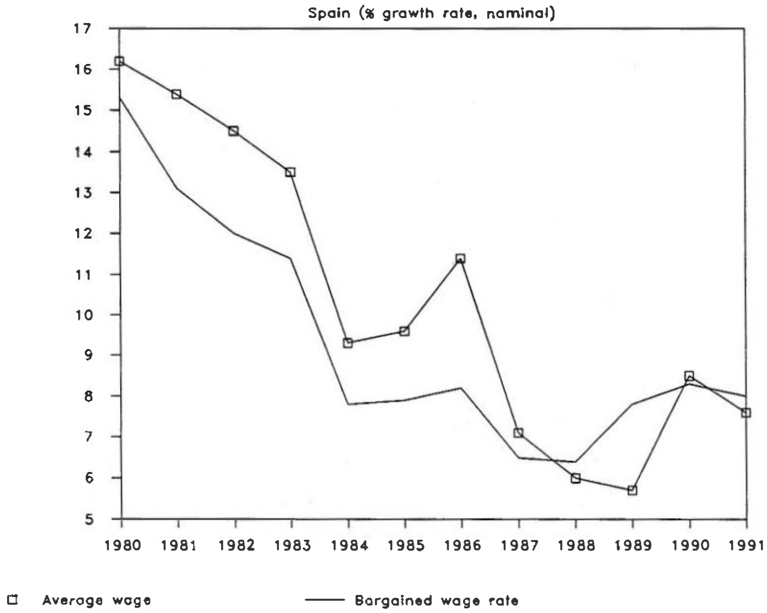
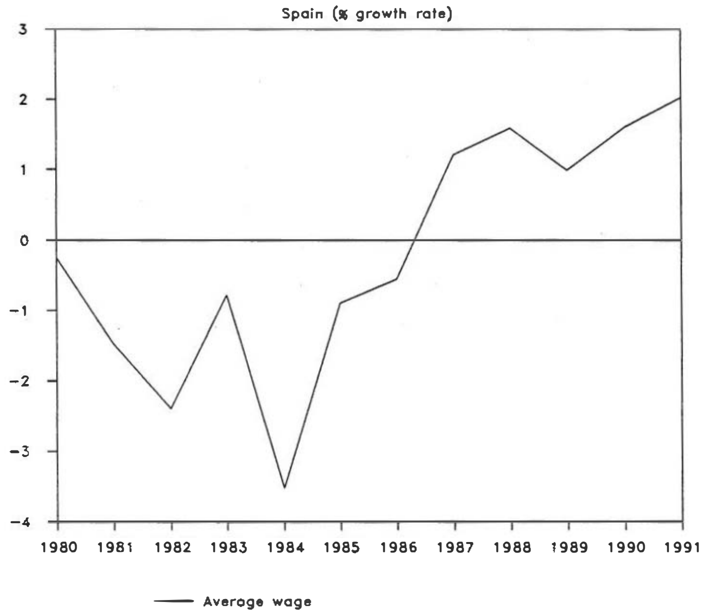


Figure 2: Real bargained wage rates



(see Figure 2).[2]

Understanding these facts is crucial for predicting whether future wage growth will be compatible with further reductions in unemployment. In this respect, any rigorous analysis must take into account the institutional change occurred in 1984, when the Spanish government created new fixed-term labor contracts, which can be signed for short periods and bear low firing costs. These contracts have been used very widely and now comprise around a third of all employees. This change has strongly affected most key variables in the Spanish labor market, including wages. For example, the existing evidence reveals that permanent workers appear to enjoy higher wages than temporary workers (Jimeno and Toharia, 1991a, and Alba, 1991).

In this paper we focus on the relationship between fixed-term contracts and wage growth, from a microeconomic perspective. Our model has both theoretical and empirical implications, so that, besides helping to explain the Spanish experience, it might be relevant for other countries. Let us first sketch the theoretical framework, and then come back to its application to the case of Spain.

Wage setting in Europe is widely seen today as dominated by *insider* employees, that is, incumbent workers whose jobs are protected either by substantial labor turnover costs or by the

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2 Bargained wages refer to industry and services, while average earnings are for the whole economy. Source: García Perea (1992). Similar patterns are found in both the data in Albarracín and Artola (1990), which correct official data for lack of updating of employment composition, and in unpublished, more homogeneous data referred to in Banco de España (1992, p. 18).

possession of specific skills which are necessary in the production process.[3] Insiders are assumed to maximize the rents extracted from the wage bargain with employers, so that their wages depend both on their bargaining power and the firm's performance, the so-called inside factors, and on general conditions affecting their expected alternative income, the so-called outside factors. The relative importance of each kind of factor in each firm is a key determinant of wages.

Also, insiders are defined in contrast to unemployed *outsiders*, who do not participate in the wage bargain. Within the insider-outsider framework it is common (see Layard *et al.*, 1991, chapter 2) to equate insiders with last period employees of the firm. These are assumed to take into account their own interests solely when bargaining. This assumption is a special case of a more general membership rule in which some of the outsiders (like recently dismissed employees) would be considered in the insiders' objective function. In this setup, unemployment persistence can arise both through the operation of these dynamic rules, i.e. *insider or membership hysteresis* (Blanchard and Summers, 1986), and through the influence of unemployment dynamics on the wage bargain, i.e. *outsider hysteresis* (Layard and Nickell, 1987).

We believe that the insider-outsider framework is applicable to the Spanish wage-setting process, but we submit that the insider group may be even smaller than the number of employees in the previous period or, more generally, that not all employees may be equally weighted in the insiders' objective function. This

[3] Lindbeck and Snower (1988b) contains the series of papers in which they started the *insider-outsider* approach.

hypothesis is not new and has already been considered, for example, in the median voter model with layoffs by seniority (see Oswald, 1987). What we do here is to specialize such framework to the case of a two-tier labor market, which currently prevails in Spain.

The second tier in that market consists of fixed-term (or *temporary*) employees, who are subject to a high job turnover and thus have an unstable attachment to the firm. This leads us to believe that they do not participate in the wage bargain on equal footing with permanent employees. In fact, we make the extreme assumption that temporary employees' interests are completely disregarded in wage bargains. To formalize this idea we write down a modified model of insider wage setting, in which permanent employees are the insiders and temporary workers are outsiders. In this model, the impact of the proportion of temporary employment on wages is threefold: (a) a *buffer effect* (arising from the knowledge by insider permanent workers that temporary employees would be the first ones to be dismissed[4]), (b) a *composition effect* (stemming from the potential difference between permanent and transitory workers' wages), and (c) a *bargaining effect* (arising from the potential influence of the fixed-term proportion on insiders' bargaining power).

Two wide sets of implications of the model are empirically studied. The first is the assessment of the importance of inside

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[4] This may seem counterfactual, since over the quarters 1990:3-1991:4 (i.e. outside our sample period), permanent employment fell by a yearly rate of 2.5% while temporary employment grew by 8.4%. These are, however, net figures. Lacking data on gross flows, we can only point to the huge number of fixed-term contracts per net job change observed (see section 2) as evidence of a very large turnover of temporary employees.

factors in wage setting in Spain. This has already been done at the sectoral level (by Andrés and García, 1991) but not at the firm level,[5] which is the best suited to the model. We estimate a wage equation using panel data on a set of 1,167 private, manufacturing, and mainly large firms which reported to the Bank of Spain's Balance Sheet Survey (*Central de Balances del Banco de España*), for the period 1985-88. Our estimate of the so-called *insider weight* is small, around 11%, which is similar to the estimates for some other countries and also for Spain at the sectoral level. We also find some sectoral variation in this coefficient. As to hysteresis effects, we obtain strong evidence of membership but not of outsider hysteresis.

The second set of issues we address is the impact of the proportion of fixed-term employment on wage growth. We test, and are not able to reject, the hypothesis that temporary workers can be considered as outsiders, which confirms the presence of a *buffer effect*. We also separately identify the *composition* and the *bargaining* effects mentioned above, finding that the latter is positive.

The rest of the paper is structured as follows. In Section 2 we briefly describe the structure of wage setting in Spain, the nature and evolution of fixed-term contracts and the existing evidence on their impact on key variables in the Spanish labor market. Section 3 contains a model of insider wage setting with two types of workers. Section 4 presents the estimation of the model and Section 5 provides our conclusions.

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[5] The only exception is the recent work by Draper (1992).

2. Wage setting and fixed-term contracts in Spain

The current wage bargaining system is quite new. Although collective bargaining existed before, unions were not legally recognized until 1977. There are two main unions, which comprise around 70% of all representatives bargaining with employers. Union affiliation is low, around 10%, but this is a poor measure of union power because most industry-wide agreements apply to all firms within the sector. Over our sample period, 1983-88, 4,000 agreements per year were officially registered. Those corresponding to a single firm are concentrated in large firms and comprise about 20% of workers covered by collective bargains. Agreements are signed yearly. The object of the bargain is normally wages and hours of work; there are often clauses on union representation but hardly ever on employment levels. (See Jimeno and Toharia, 1991b, chapter 3, for details.)

It appears, therefore, that the *right-to-manage* model we use below -in which unions and firms bargain about wages and let the firm choose employment- is a good enough approximation to reality, at least for large firms, which are a majority in our database, and especially after 1986, when nationwide wage agreements ceased to exist.

The other aspect we focus on is the change in the institutional setting brought about by fixed-term contracts. These contracts can be used in any type of activity (temporary in nature or not), they can be signed for short periods (previously 6 months, 1 year since April 1992) and can be renewed up to three years, have low firing costs,[6] and, contrary to permanent contracts, their

[6] Some types bear no firing costs at all. Other types imply 12 days

extinction does not require administrative approval and cannot be appealed to labor courts.

The introduction of these contracts shortly preceded the boom experienced over the period 1986-90 -in which GDP grew by 5% on average- and they were extensively used. Over those 5 years total employment increased by 3% per year, the number of employees by 4.8%, and 98% of all contracts registered at employment offices were fixed-term. As a result these contracts now cover one third of all employees, and 38% in the private sector. The proportion of this type of contract varies across workers and sectors. They are more prevalent for women than for men (34% vs. 28%), for young workers (78% for those aged 16-19), for less educated workers (36% for illiterates, 17% for college graduates), in agriculture and construction (51% and 54%) and in Southern regions (e.g. 45% in Murcia). (In 1990:2, when the average temporary employment ratio was 30%.)[7]

Such a change in the structure of labor contracts should affect most key labor market variables. For example, the lower firing costs entailed by these contracts should induce an increase in the cyclical variability of employment. This is what Bentolila and Saint-Paul (1992) have found in a sample of large manufacturing firms (almost the same one we use here), also estimating an increase in employment of 1.5 percentage points over three years (in which employment actually fell in the sample) due to the

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of wages per year of service versus 20 days for permanent employees or 45 days if the dismissal is ruled unfair.

[7] See Segura *et al.* (1991) or Jimeno and Toharia (1991b), chapter 4, for more details.

introduction of fixed-term contracts.

The second important effect of fixed-term contracts has been an increase in labor turnover. In 1991 there were 5 million contracts signed, while in net terms 25 thousand dependent jobs were being destroyed. The higher turnover may affect productivity, for several reasons discussed in the next section. For a sample of large firms in 1989, Alba (1991) finds a negative or null effect of the proportion of temporary employees on labor productivity, but a positive effect of the flow of temporary contracts in newly hired workers. Jimeno and Toharia (1991a) find a negative effect of the proportion of fixed-term employment on sectoral labor productivity growth in 1988, but not in 1989. However, Hernando and Vallés (1992) do not find any influence of temporary employment on total factor productivity during 1986-89. These findings point towards a negative correlation between the capital-labor ratio and the proportion of temporary employment,[8] which is quite sensible.

Another consequence of high labor turnover has been a disequilibrium in public finances. Since workers had the right to collect unemployment benefits after having worked for just 6 months, the revenues of the unemployment compensation system fell short of benefit payments as soon as the economy slowed down (second half of 1990). In response, the government raised (in April 1992) the minimum period of previous work to be eligible for unemployment benefits to 1 year and lowered the amount of unemployment benefits, causing some labor unrest.

8 This is straightforward from the definition of total factor productivity.

It has also been found that workers on fixed-term contracts tend to receive lower wages than those on permanent contracts.[9] Jimeno and Toharia (1991a) estimate the wage premium in favor of the latter, after a wide set of observable characteristics is controlled for, at about 11%. This is consistent with the previous remarks about the evolution of the wage drift. Nevertheless, the lower employment security enjoyed by temporary employees implies that they should be compensated through higher, not lower, wages. The estimated negative premium for temporary employees might be due to unobservable characteristics (this is Alba's, 1991, reading of the data), but we believe it is more likely due to their low power in wage bargains. Again, their unstable attachment to the firm makes them much less likely to be part of the workers' bargaining unit. In the model presented below we assume that these workers' interests are not represented at all in wage bargains.[10] We think this is plausible in manufacturing, but much less in the construction sector or even in services, where the proportion of fixed-term employment is so high that such employees must surely count in wage bargains.

[9] Discrimination based on contract type is illegal in Spain, but apparently employers are relatively free to choose the occupation in which they classify new workers, so that by underclassifying them they can actually pay them less.

[10] This should not be thought of as arising from the law: regardless of his/her contract type, any worker with six-months seniority can be elected as a workers' representative and any worker with one-month seniority can vote.

3. The model

According to the insider-outsider theory, wages are the outcome of a bargaining process whereby firms and their insiders share the economic rent from insider employment. In this context, the insiders' wages will be higher, *ceteris paribus*, the more the firm stands to lose from a breakdown in wage negotiations. To illustrate this theory we draw on Nickell and Wadhwani (1990), slightly amended to highlight the theme that permanent workers are the insiders. In Lindbeck and Snower (1988a) new workers are called *entrants* and gain the insider status after one period in the firm. Although our model is similar to theirs, strictly speaking we do not follow this path, since the probability of becoming a permanent employee appears to be very low in Spain, between 10% (Segura *et al.*, 1991) and 15% (Alba, 1991).

Following the convention in this literature, the model has two parts. The first part describes the behavior of a representative imperfectly competitive firm which sets price, production, and employment levels given wages and expected demand. The second part deals with wage determination in the context of the insider-outsider theory with two types of labor. These decisions are made in two stages. First, wages are set according to the *right-to-manage* model (see Nickell and Andrews, 1983) by which the firm and workers bargain over wages, taking the expected effect of this decision on prices, production, and employment into account. Then the firm sets the previous variables taking wages are given.[11]

[11] Our model is static. Since the power of insiders stems essentially from adjustment costs, we should have a dynamic model. This is an extension we leave for further research.

3.1 The Firm

Let there be a representative monopolistic firm, producing a homogeneous product with a Cobb-Douglas technology in two types of labor:

$$Y = \tilde{A}\, N_p^{\alpha}\, N_T^{\gamma}\ ; \qquad \alpha + \gamma < 1 \tag{1}$$

where Y is output, N_p is permanent employment, N_T is temporary employment, and $\tilde{A}$ is a technical progress coefficient. Note that other factors of production could be included in $\tilde{A}$.

We can think of many reasons why the intensities of permanent and temporary labor may differ (i.e., why α differs from γ). Three are general in nature: temporary workers may be less productive than permanent ones because the firm invests less in the formers' firm-specific human capital, given their shorter expected tenure; also temporary workers may devote either less effort than permanent ones -in an efficiency wage context- as a result of the pay structure established by the firm (as in Rebitzer, 1987, and Saint-Paul, 1991), or more effort, if they want to show themselves as good workers in order to boost their chances of being hired on a permanent basis (as in Jimeno and Toharia, 1992). The other two reasons are specific to the Spanish experience: over 1985-89 most newly employed workers have been less educated and less experienced than the average employee (Jimeno and Toharia, 1991b, chapter 4).

Except for one, all of these reasons point towards γ being lower than α. The approach we follow in this paper avoids having to formalize these ideas explicitly. We just allow for potentially different intensities for each type of labor, without assuming anything about their relative magnitudes.

Using the marginal rate of substitution (MRS) between both types of labor, (1) can be rewritten as

$$Y = A \, (W_p/W_T)^{\gamma} \, N_p^{\alpha+\gamma} \tag{2}$$

where $A = \tilde{A}(\gamma/\alpha)^{\gamma}$, and W_p and W_T are the wages for both types of workers. The first wage is subject to bargaining, while the second is taken as exogenous, possibly corresponding to the outsiders' reservation wage.

The firm faces an isoelastic demand curve given by

$$Y = \varepsilon \, P^{-\eta} \, Y_d \; ; \qquad \eta > 1 \tag{3}$$

where P is the price of output, Y_d is a demand index and ε is a random variable whose value is revealed only after wage bargaining has taken place. The firm maximizes profits, Π, which are equal to

$$\Pi = P \, Y - W_p \, N_p \, (1 + \gamma/\alpha) \tag{4}$$

where use has been made of the following relationship implied by the MRS between N_p and N_T:

$$W_p \, N_p + W_T \, N_T = W_p \, N_p \, (1 + \gamma/\alpha)$$

From the first order condition of the maximization of (4) subject to (2) and (3), permanent employment is given by:

$$N_p = \tilde{\varepsilon} \left[\frac{W_p}{\alpha \, k \, Y_d^{1/\eta} \, A^k (W_p/W_T)^{\gamma k}} \right]^{-\frac{1}{1-(\alpha+\gamma)k}} \tag{5}$$

where k $(= 1-1/\eta)$ is the degree of market competition and $\tilde{\varepsilon}=\varepsilon^{1/\eta[1-(\alpha+\gamma)k]}$. Under the assumption, without loss of generality, that $E(\tilde{\varepsilon}) = 1$, expected permanent employment N_p^e satisfies

$$N_p = \tilde{\varepsilon}\, N_p^e \tag{6}$$

where, making use of (3), N_p^e can also be expressed in terms of the expected real wage

$$N_p^e = \left[\frac{W_p^{1-\gamma}\, W_T^{\gamma}}{\alpha\, k\, A\, P^e} \right]^{-\frac{1}{1-(\alpha+\gamma)k}} \tag{7}$$

Substituting (7) into (4), it is then easy to show that ex-ante maximum profit (Π^e) is given by

$$\Pi^e = \left[\frac{1 - k\ (\alpha+\gamma)}{\alpha\ k} \right] W_p\ N_p^e \tag{8}$$

So, with the Cobb-Douglas production function, the profit share in terms of the wage bill of permanent workers is independent of W_T and is decreasing in both labor intensities, α and γ, and competitiveness, k.

3.2 The Union

We assume that insiders are workers with a permanent job, and are represented in the bargaining by an organization (for example, a works council) which pursues their interests rather than those of outsiders, as represented by temporary workers. In particular, the bargaining body is assumed to maximize the expected income of the median permanent voter, taking W_T as exogenous. We will see below that, under the technological assumptions made here, the way in which W_T is determined is not relevant for the empirical formulation of the model. The objective, V_p, is thus given by

$$V_p = S_p W_p + (1 - S_p) W_p^a$$

where S_p is the survival probability (of being employed next period) and W_p^a is the expected income of a laid-off worker given by

$$W_p^a = (1 - \varphi(u)) \bar{W}^e + \varphi(u) B = \bar{W}^e [1 - \varphi(u)(1 - b)]; \ \varphi'>0 \qquad (9)$$

where $\varphi(u)$ is the probability of unemployment, $\bar{W}^e$ is the expected outside wage, B is the unemployment benefit, and b ($\equiv B/\bar{W}^e$) is the expected benefit replacement ratio.

If we denote the target employment of the union in the bargaining by M, then the survival probability can be written as

$$S_p = \text{prob } [N_p \geq M] + E(N_p/M \mid N_p<M) \text{ Prob } [N_p < M]$$

$$= \text{prob } [\tilde{\varepsilon} \geq M/N_p^e] + (N_p^e/M) \ E(\tilde{\varepsilon} < M/N_p^e) \text{ prob } [\tilde{\varepsilon} < M/N_p^e]$$

$$= S_{p1} \ (M/N_p^e) \ = \ S_{p2} \ (W_p) \qquad (10)$$

where S'_{p1}, S'_{p2} < 0. Thus, the probability of survival is smaller the larger the insiders' target relative to expected employment and, therefore, the larger is the wage. In what follows we will denote the absolute values of the elasticities of S_p with respect to M/N_p^e and W_p as $|e_{MN}|$ and $|e_{SW}|$, respectively.

3.3 The Nash Bargain

The bargaining outcome is assumed to be based on the standard Nash model with *status quo* points given by $\bar{\Pi} = 0$ and $\bar{V}_p = W_p^a$. So the Nash objective to be maximized is

$$\Omega = \left[S_p \; (W_p - W_p^a) \right]^{\beta} \; \Pi^e \tag{11}$$

where β measures workers' relative bargaining power. β will itself be a function of a set of variables, among which we can find those related to the firm's likelihood of avoiding bankruptcy arising from wage demands or its financial health (e.g. real profits per head (bn), liquidity variables (l), etc.).

Another variable affecting β is the ratio of temporary to total employment in the firm (ϕ). This is the *bargaining effect* of ϕ, which comprises two sub-effects with opposite signs. On the one hand, there is the so called *harassment effect* (see Lindbeck and Snower, 1988a), by which insiders may have a stronger bargaining position with the firm by threatening to use uncooperative behavior against fixed-term contract employees, in case that their wage claims are not accepted by the firm. Under the assumption that the smaller the number of insiders (the higher the proportion ϕ) the more essential they are in the process of production, it is likely that by withdrawing cooperation or damaging personal relations with temporary workers they will inflict a higher damage on the firm's position in case of disagreement, so that the firm will prefer to accept their wage claims rather than keep the permanent workers' positions vacant or replace them by new workers with lower productivity.[12]

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[12] Another potential factor by which the proportion of temporary employment might affect insiders' power positively is the accumulation of strike funds over time, as the number of permanent employees shrinks. But strike funds seem to be largely absent from the Spanish system, so we deem this channel to be irrelevant.

On the other hand, there is the *discipline effect*, by which the higher the number of workers under fixed-term contracts the lower will be the workers' relative power, since the firm can easily lay them off. Temporary workers should be wary of engaging in strikes led by permanent employees and this will tend to reduce β.

In summary, β can be written as:

$$\beta = \beta(bn, \quad 1, \quad \phi) \tag{12}$$

where $\beta'_i>0$ $(i = 1,2)$, and the sign of β'_3 will depend on the relative strength of the two aforementioned effects.

Differentiating Ω in logs with respect to W_p yields

$$\frac{W_p}{W_p - W_p^a} - |e_{SW}| + \frac{1}{\beta} \frac{\partial \Pi^e}{\partial W_p} \frac{W_p}{\Pi^e} = 0 \tag{13}$$

From the envelope theorem it is easy to show that

$$\frac{\partial \Pi^e}{\partial W_p} \frac{W_p}{\Pi^e} = \frac{- \alpha k}{1-(\alpha+\gamma)k}$$

and since $|e_{SW}| = |e_{MN}| \, |\varepsilon_{NW}|$, (13) can be expressed in terms of the mark up of wages over alternative income as follows:

$$\frac{W_p - W_p^a}{W_p} = \frac{1-k(\alpha+\gamma)}{|e_{MN}|(1-\gamma k) + k\alpha/\beta} \tag{14}$$

The comparative statics of (14) imply that the mark-up of the wage of permanent workers with respect to the outside wage is higher the greater the union power (β) and the firm's demand (Y_d) (the latter through its effect on $|e_{MN}|$) and the lower the product market degree of competition (k), the targeted number of insiders (M) and the intensity of both types of labor (α and γ) (the latter because more rent is available to extract in the bargaining).

3.4 Empirical formulation

On the basis of the dependence of $|e_{MN}|$ on M/N_p^e and the expression for expected labor demand given in (7), it is possible to log-linearize (14) so as to explicitly derive the impact of all the relevant variables on firm-level wages. Taking the lower-case letters, a, m, w, p, and n to represent logs, we have

$$w_p = \text{constant} + \lambda \left[a + p^e + \gamma\, (w_p - w_T) - (1-(\alpha+\gamma))\, m \right] +$$

$$+ (1-\lambda) \left[\bar{w}^e - c_1\, u + c_2\, b \right] + c_3\, \beta + c_4\, k \qquad (15)$$

where $0 \le \lambda \le 1$, given the homogeneity of W_p in $\bar{W}^e$ and M/N_p^e (see Jackman Layard and Nickell, 1992, p. 183). Thus λ represents the *insider weight*. Since a is not observable, we can substitute it out using the production function given in (2), implying that the first bracketed term in (15) can be rewritten as

$$\left[p^e + y^e - (\alpha + \gamma)\, n_p^e - (1-(\alpha+\gamma))\ \ m \right] \qquad (16)$$

If we now assume that $m = n_{p,-1}$, i.e. the targeted level of employment is last period's permanent employment, (16) can be expressed as

$$\left[p^e + y^e - n^e + (n-n_p)^e + (1-(\gamma+\alpha))\ \ \Delta n_p \right] \qquad (16')$$

Denoting the proportion of temporary workers in total employment by ϕ ($\equiv N_T/N$), we have that $n_p - n \simeq \log(1-\phi) \simeq -\phi$ if such a proportion is small, and (16) now reads

$$\left[(p + y - n)^e + \phi^e + (1-(\alpha+\gamma))\ \Delta n_p^e \right] \tag{16"}$$

Note that the membership hysteresis term Δn_p and the proportion ϕ appear as a consequence of the assumption that permanent workers are the insiders, and also that the way in which W_T is determined is irrelevant, given the MRS condition derived from the Cobb-Douglas production function.

Finally, it is important to notice that, due to lack of information, our wage data correspond to the average labor cost per employee in each firm and not to the wage earned by permanent workers. This is not a problem, since the aggregate average wage is given by

$$W = \frac{W_T\ N_T + W_P\ N_P}{N} = \frac{W_P\ N_P\ (1+\gamma/\alpha)}{N}$$

which in logs can be approximated by

$$w = w_p - \phi + \text{constant} \tag{17}$$

Thus given (16"), (17), and a linear approximation of (12), we are able to write the log-linear approximation of (14) in terms of the average wage, w, as

$$w = \text{constant} + \lambda \left[(p + y - n)^e + (1-(\alpha+\gamma))\ \Delta n_p{}^e \right] +$$

$$(1-\lambda)\left[\bar{w}^e - c_1 u + c_2 b\right] + c_{31}bn + c_{32}l + (c_{33}-(1-\lambda))\,\phi^e + c_4 k \quad (18)$$

This equation has the standard property that bargained wages are a convex combination of inside factors relating to the firm's ability to pay $((p+y-n)^e,\ k)$ and the size and strength of the union $(\Delta n_p,\ \beta)$; and outside factors affecting the firm's ability to control its workforce $(\bar{w},\ u,\ b)$.

The distinctive feature of equation (18) relative to other empirical specifications in this class of wage determination models (cf. Nickell and Wadhwani, 1990) is the role played by Δn_p and ϕ. The first variable just arises from the definition of insiders, i.e. $M = N_{p,-1}$. If it were the case that the bargaining unit also included a proportion (τ) of the temporary workers in their employment target for period t, i.e. $M = N_{p,-1} + \tau N_{T,-1}$, then, assuming that the product $\tau\phi_{-1}$ is small, we would have, in logs, $m \simeq n_{p,-1} + \tau\tilde{\phi}_{-1}$, where $\tilde{\phi}_{-1} = \phi_{-1}(1-\phi_{-1})^{-1}$. Thus the new equation would be identical to (18), except that $\tilde{\phi}_{-1}$ should appear significantly, with a negative sign, in equation (18), once the current value of ϕ has been controlled for. Since the coefficient of τ is identified, we could estimate the proportion of temporary workers in the insiders' target. An estimate for τ lower than unity would imply the presence of the *buffer effect*.

The second variable, ϕ, appears in two ways. First, through having total labor productivity instead of that of permanent workers, as in (16'). This is the *composition effect* which stems from the fact that wages are not the same for both types of workers. Second, through its strategic effect on β, the *bargaining effect*. Note from (17) and (18) that the (semi)elasticity of w_p with respect to ϕ is given by $(c_{33}+\lambda)$, where the sign of c_{33} will depend upon the relative strength of the *harassment effect* (positive) to the *discipline effect* (negative). Note that if both effects' sizes did exactly compensate each other c_{33} would be zero. If this second effect were negligible, then the coefficient of ϕ in the equation for w would be identical, with opposite sign, to the coefficient of $\bar{w}$, a testable implication of the model.

4. Data, estimation and results

4.1 Empirical specification

Having obtained an empirical log-linear specification of the basic wage equation in the previous section, we also allow for the lagged dependent variable to enter, for various potentially relevant reasons: the existence of long-term contracts, the concern of the bargaining unit with both wage changes and wage levels, and the long-run nature of the relationship between the wage bills of permanent and temporary employees, which is based upon the MRS (which justifies equation (4)). Thus, the equation we wish to estimate is:

$$w_{it} = a_i + a_1 w_{it-1} + a_2 (p+y-n)_{it} + a_3 \Delta n_{pit} + a_4 \bar{w}_t + a_5 u_t + a_6 b_t$$

$$+ a_7 k_{jt} + a_8 bn_{it} + a_9 l_{it} + a_{10} \phi_{it} + a_{11} \tilde{\phi}_{it-1} + \varepsilon_{it} \quad (19)$$

where $i = 1,\ldots,N$ and $t = 1,\ldots,T$.

So the (i,t) subscript denotes the i-th firm in period t; α_i denotes a firm-specific fixed effect which controls for all proportional differences in skill levels, production technologies, etc. that remain fixed over time; j refers to industry; and we assume that ε_{it} is distributed n.i.d.$(0, \sigma_\varepsilon^2)$.

The hypotheses of interest (H_0) are the following: *homogeneity* in inside and outside factors (h); a full *buffer effect*, i.e. exclusion of temporary workers in the bargaining unit's employment target (be); and no *bargaining effect*, i.e. no influence of the temporary employment ratio on the relative power of the bargaining unit (nb). They can be represented as follows:

$$H_0^h: a_1 + a_2 + a_4 = 1; \qquad H_0^{be}: a_{11} = 0; \qquad H_0^{nb}: a_4 + a_{10} = 0$$

The definitions of the actual variables used in estimating equation (19) are:[13] w_{it} is the average labor cost per employee; $(p+y-n)_{it}$ is the value of sales per employee; Δn_{pit} is the rate of change of permanent employment; $\bar{w}_t$ is the alternative wage, as measured by average labor cost in industry and services; u_t is the national unemployment rate (in order to capture composition effects we also included the proportion of long-term unemployed in some of the regressions); b_t is the replacement ratio; k_{jt} is the market concentration index (the percentage of sales of either the five (c5) or ten (c10) largest firms in each sector)[14]; and bn_{it} is the level of after-tax profits per employee.

Lastly, l_{it} represents liquidity and financial variables. We used a variety of proxies, following similar studies (e.g. Wadhwani, 1987, Nickell and Wadhwani, 1989 and 1990, or Blanchflower, Oswald and Garrett, 1990). Specifically, we tried: (a) rd_{it}= interest bill-debt ratio; (b) dc_{it}= debt-capital ratio

[13] Throughout n stands for employment. We intend to estimate alternatively with hours of work in the next version of this paper.

[14] Note that, to save notation, here k denotes the inverse of the degree of competition in the market.

(indicator of insolvency); (c) mld_{it}= percentage of medium- and long-term debt; (d) bd_{it}= percentage of debt with banks; and (e) al_{it}= *quick assets* ratio (current assets less stocks/current liabilities).

4.2 The data

Our data come from the accounts of 1,167 manufacturing, non-energy, private firms over the period 1983-88. The source is the database of balance sheet records kept at the Bank of Spain. Not all firms in the original sample could be used; both the details of the selection process and the precise definitions of the variables are in the Appendix. Our final sample represents around 13.5% of total manufacturing employment in Spain in that period.

Table 1 presents descriptive statistics on some of these variables. These are firms: (a) which are mainly large, (b) whose employment fell more than manufacturing employment, (c) whose sales grew more than average; (d) with less temporary workers than average (in 1988 the manufacturing average temporary ratio was 18.3%, vs. 9% in our sample), and (e) paying higher wages than average. These differences mean that, at the very least, we should be wary of extrapolating our results to the whole manufacturing sector.

4.3 Estimation method

Equation (19) is estimated using the generalized method of moments (GMM) technique due to Arellano and Bond (1991),[15] which is

[15] With their DPD program (Arellano and Bond, 1988).

Table 1: Descriptive statistics[1]

	Mean	Std. dev.	Max.	Min.	Avg. value 1983	Avg. ann. growth (%) Sample	Manuf.
(a) Firm-level data:							
Levels:							
Total employment	314	1004	21718	2	326	-0.9	-0.2
Permanent employment	300	997	21694	1	317	-1.9	-
Temporary employment	15	43	914	0	9	22.0	-
Labor cost p/empl.[2]	21	8	119	3	16	9.6	9.0
Sales[3]	41	164	4619	10	28	14.0	12.4
Ratios (%):							
Temporary employment	6.6	12.9	94.1	0.0	4.6	4.7	
Concentration index[4]	19.5	12.2	99.2	5.5	19.7	0.5	
Med./long term debt	14.0	19.0	99.8	0.0	6.2	52.4	

(b) Economy-wide data:

	Mean	Growth rate	1983 value
Labor cost[5]	17.0	9.0	14.0
Long-term unempl.[6]	62.5	4.5	57.3
Unemployment rate	20.0	2.2	17.3
Replacement ratio	54.0	-3.0	54.7

Notes:

[1] 1983-88 data unless otherwise noted. Monetary figures at 1991 peseta-dollar exchange rates.

[2] In thousand dollars.

[3] In million dollars.

[4] Percentage of sales corresponding to largest 5 firms in the sector.

[5] Average labor cost in manufacturing and services in thousand dollars.

[6] Unemployed for more than a year as percentage of total unemployment.

a more efficient extension of the traditional instrumental variables (IV) method suggested by Anderson and Hsiao (1982). Fixed effects, α_i, are eliminated by transforming the corresponding equation into first difference form. Thus, under the assumption that ε_{it} is white noise, the resulting error now has an MA(1) structure and is correlated with the lagged dependent variable, requiring instruments (to be dated at t-2) for those variables which are not assumed to be strictly exogenous. Of course, the validity of the instruments depends on the previous assumption that the errors in the equation in levels are serially uncorrelated, which can be tested by means of the m_2 statistic for the absence of second-order serial correlation in the differenced equation. Arellano and Bond (1991) show that the limiting distribution of m_2 is a standardized normal variate. We also report the Sargan statistic for the validity of the (over)identifying restrictions, S_{IV}. Its limiting distribution is chi-square with (m-k) degrees of freedom, where m is the number of instruments and k that of regressors in the equation.

In the estimation, the variables w_{it-1}, $(p+n-y)_{it}$, Δn_{pit}, and ϕ_{it} are treated as endogenous, whereas the remaining variables with firm or sectoral variation are lagged one period and treated as strictly exogenous. An extended set of instruments is used, consisting of all lags of w_i, n_i, $(p+y-n)_i$, and ϕ_i from t-2 back, plus a constant, the exogenous variables in the equation, and rd_i, dc_i, mld_i and al_i lagged once, making a total of fifty-two instruments. A reduced instrument set consisting of thirty-five instruments, selected from the most significant regressors in the reduced form equations for the instrumented variables, is also used, to check on the robustness of the results with respect to the choice of instruments.

4.4 Main Empirical results

Due to lag requirements only the last four years of the sample can be used for estimation. This makes for a total of 4668 observations. Our preferred version of equation (19) is given in Table 2, column 1. Several points are worth noting.

First, the long-run homogeneity in the inside and outside factors has been imposed after the corresponding null hypothesis (H_0^h) was tested for and not rejected (t-ratio = 0.36). The estimated insider weight is about 0.11, a value which is in line with that estimated by Nickell and Wadhwani (1990) for the UK (0.11) and in between those pertaining to countries where bargaining is rather decentralized (Brunello and Wadhwani, 1989, report a value of 0.33 for Japan) or very centralized (Holmlund and Zetterberg, 1989, estimate a value of 0.04 for Sweden). More interestingly, our estimate is almost identical to the one obtained by Andrés and García (1991), using equivalent sectoral regressions across 89 industry groups of the Spanish Industrial Survey during 1978-1986.

Second, the long-run coefficient of the proportion of temporary workers in the firm is very significant and equal to -0.65, which is not too far in absolute value from the outsider weight, 0.89. However, the null hypothesis of equality of both coefficients (H_0^{nb}) is definitely rejected (t-ratio = 3.29), implying that there is a direct effect of such proportion, ϕ, on the relative power of insiders, β. Taking into account the estimated value of λ and the difference between the previous two values (c_{33} = 0.89-0.65), the long-run (semi)elasticity of w_p with respect to ϕ is 0.34, so that an increase of ϕ of one percentage point (e.g. from 6% to 7%) raises the growth rate of permanent

Table 2: Wage equation (19)
(Dependent variable: w_{it})

	(1) Preferred specification	(2) Time dummies	(3) Sectoral dummies	(4) Reduced IV set	(5) Altern. Insiders
w^{*}_{it-1}	0.267	0.245	0.244	0.250	0.273
	(8.14)	(5.81)	(8.52)	(8.16)	(7.31)
$(p+y-n)^{*}_{it}$	0.078	0.077	0.072	0.079	0.090
	(2.76)	(2.42)	(2.80)	(2.93)	(2.93)
Δn^{*}_{pit}	0.030	0.028	0.022	0.029	-0.021
	(2.90)	(2.98)	(3.17)	(3.15)	(2.76)
$\bar{w}_t$	0.655	-	0.684	0.671	0.637
	(-)	(-)	(-)	(-)	(-)
u_t	-0.928	-	-0.836	-0.985	-0.978
	(5.60)	(-)	(4.53)	(6.92)	(6.07)
b_t	0.108	-	0.106	0.112	0.119
	(3.82)	(-)	(3.50)	(4.07)	(4.37)
k_{jt-1}	0.011	0.015	0.020	0.032	0.017
	(1.47)	(1.60)	(2.67)	(2.92)	(1.19)
bn_{it-1}	0.093	0.110	0.116	0.088	0.092
	(3.25)	(3.62)	(3.68)	(3.25)	(3.20)
rd_{it-1}	-0.016	-0.021	-0.012	-0.018	-0.006
	(1.51)	(1.40)	(1.19)	(1.22)	(0.78)
mld_{it-1}	-0.034	-0.033	-0.034	-0.028	-0.018
	(2.58)	(2.56)	(2.60)	(2.31)	(1.49)
ϕ^{*}_{it}	-0.477	-0.511	-0.510	-0.505	-0.302
	(8.68)	(8.29)	(8.40)	(8.43)	(2.76)

Table 2 (continued)

	(1) Preferred specification	(2) Time dummies	(3) Sectoral dummies	(4) Reduced IV set	(5) Altern. Insiders
m_2	1.050	0.586	0.166	0.812	0.896
S_{IV}	39.13	37.17	40.36	27.43	39.66
	(42)	(41)	(41)	(25)	(42)
σ^2_ε	0.58	0.57	0.59	0.58	0.60

Notes:

- Number of observations in all regressions: 4668. Period: 1985-88.

- Heteroskedasticity consistent t-ratios in parentheses.

- A superscript (*) denotes instrumented variables.
  The "extended" and "reduced" IV sets are described in the main text.

- A constant term was initially estimated but was never significant.

- m_2 = test for second-order serial correlation in the differenced equation.

- S_{IV} = Sargan's test of overidentifying restrictions (degrees of freedom in parentheses).

- σ^2_ε = variance of the residuals (multiplied by 100).

workers' wages by about one-third of a percentage point.

Third, the null hypothesis that permanent workers do not care about temporary employees when setting their employment targets (H_0^{be}) is not rejected (t-ratio = 1.13), implying that our choice of the former as the insiders is not at odds with the data, i.e. that the *buffer effect* is present.

Fourth, both the market power measures and the firm-specific liquidity effects are by no means insignificant, though individually some of them (k and rd) show weak effects. We have kept them on the preferred equation on the grounds that they appear more significant in other specifications.

Fifth, the membership hysteresis effect,[16] captured by the variable Δn_p, is also well determined and shows a small positive coefficient, again in agreement with the theory. Given that its coefficient in (18) should be $\lambda(1-(\alpha+\gamma))$, the estimated share of both labor factors is 0.62, a rather sensible value.

Sixth, outsider effects are correctly signed and strong. In this respect it is important to note that due to the shortage of degrees of freedom (we can only use four cross-sections), a priori we expected badly determined effects for aggregate variables. This is not the case, however, for the unemployment rate, the replacement ratio, and the outsider wage.

We have anyway been prevented from introducing some other

[16] Note that the dynamic monopsony model would provide an alternative interpretation of the membership hysteresis effect, different from the one derived from the insider-outsider framework.

interesting variables with no inter-firm variation. In particular, we were interested in the proportion of long-term unemployment, as a way to capture an outsider hysteresis effect due to the composition of the unemployed (Nickell, 1987). To be able to include this variable, which shows up very significant in the industry study of Andrés and García (1991), we formed a composite variable with u and b, using the coefficients estimated in the preferred equation, and then introduced the proportion of unemployed for more than a year. The variable showed up positive, as expected, but not significant at all (coefficient = 0.36, t-ratio = 0.27). In fact, Spanish data on unemployment flows do not indicate that the long-term unemployed have had unsurmountable difficulties in finding jobs. In the case of men, those who have got a job come from unemployment rather than from out of the labor force. The *long-term unemployment trap* would predict a rise in the proportion of long-term unemployed when employment rises, whereas in Spain employment grew in 1985-88 at an average rate of 2.8%, and at the same time that proportion actually fell by 4 percentage points. Similarly, the inclusion in the equation of unemployment changes did not prove successful. In any case, the shortage of cross-sections makes our data set not the best suited for examining the influence of such variables and we will not make a strong case about them.

Columns 2 to 5 in Table 2 report the results of some modifications of the preferred equation. In column 2, the aggregate variables have been substituted by time dummies which capture all the aggregate common variation in the data. The results are very similar, which is not surprising since the aggregate variables (three) were almost playing the role of the dummies (four). In column 3, we introduce thirteen sectoral dummies in the differenced equation, which correspond to sectoral linear trends in the levels equation. A Wald test for their joint significance takes a value of $\chi^2(13) = 22.23$, marginally insignificant at the five percent level. The remaining variables do not change much except the concentration ratio, which doubles its value and becomes more significant. Conversely, some of the liquidity variables lose significance.

The relative invariance of the results to the introduction of sectoral dummies has an interesting interpretation. A possible criticism of the joint *bargaining-composition* interpretation given to the coefficient on the variable ϕ in equation (18), is that we are not explicitly controlling for skill composition of both permanent and temporary workers.[17] To the extent that this composition has changed across sectors and over time (the latter would be captured in the model by changes in the coefficients α and γ in the production function), some of the estimated coefficients could be biased. By including sectoral dummies intersected with time dummies in the differenced version of the model, we are allowing for a raw measure of those differences in pay across sectors which have varied over time. The fact that this set of dummies is not significant ($\chi^2(52)=63.2$) suggests that skill composition and the like have tended to be stable over the sample period. However, some of the genuine explanatory variables such as $(p+y-n)$, ϕ, and Δn_p could be affected by those shifts, an issue to which we will come back below.

In column 4, the reduced instrument set is used and the results are again very similar, except in the case of the

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[17] Blanchflower, Oswald and Garret (1990) estimate, within an insider model, separate wage setting equations for unskilled, semi-skilled and skilled workers, finding statistically different coefficients in each.

concentration ratio, which again increases in size and significance. Given that the results also remained fairly invariant to slight variations in this reduced IV set, we feel somewhat protected against the usual criticism about the lack of identification of wage equations containing *inside* variables (see Manning, 1992).

Finally, in column 5, we substitute the rate of change of permanent employment, Δn_p, by the growth rate of total employment, Δn, in order to examine the robustness of our choice of the insider group, with respect to the more traditional view which associates lagged employment, n_{-1}, to the union's employment target in the bargaining process. Note that if this were the case, equation (16") would become:

$$\left[(p + y - n^e) + (\alpha+\gamma)\ \phi^e + (1-(\alpha+\gamma))\ \Delta n^e \right] \qquad (16"')$$

so that the coefficient of Δn would be positive. However, this is not what we find, since the estimated coefficient turns out to be negative and significant. A similar sign is obtained by Andrés and García (1991), though they present evidence that a twist in sign occurs when lagged employment changes are added to the equation. Unfortunately, we cannot test for this possibility due to the lack of cross-sections when estimating in differences. Nevertheless, we believe it is reasonable to interpret the above result as favorable to the hypothesis that permanent employees are the incumbent workers.

4.5 Further results

In this section we present further evidence on the robustness and the interpretation of the results in the previous section. We start with parameter stability tests carried out for the coefficients of productivity and changes in permanent employment. Given the small number of cross-sections available, we compare the estimated coefficients of both variables in the periods 1985-86 and 1987-88, by introducing in the preferred equation (column 1) both variables intersected with a dummy variable which takes a value of unity for the second period. In order to allow for long-run homogeneity in both periods, the first variable is introduced as $((p+y-n)_{it}-\bar{w}_t)$ so that its coefficient reflects the change in λ in equation (18) between both subsamples. The estimated coefficient for this variable is -0.03 (t-ratio = 1.12), whereas for Δn_p it is -0.01 (t-ratio = 0.76). Thus, the equation seems to be quite stable.

Given that in the second subperiod labor productivity in the industrial sector fell by 2 percentage points (see Hernando and Vallés, 1992), the reduction in the insider weight (λ), although insignificant, offers some weak evidence of asymmetries in the wage adjustment: i.e., λ is higher in good times and thus workers are happy to take a wage hike when productivity is high, but less happy to take a cut when productivity is low.

Similarly, we carry out a stability test for the coefficient on ϕ, yielding an estimated increase of 0.09 (t-ratio = 2.39), significant but small. Given that there are no serious signs of lack of stability in the insider weight, we interpret this significant shift as an increase in the coefficient c_{33}, i.e. a rise in the strength of the *harassment effect* relative to the *discipline effect*, which is compatible with the progressive resurgence of a positive wage drift as noted in the introductory section.

Thus, apart from some signs of a significant but not very sizable shift in the coefficient of the proportion of temporary workers, the equation seems fairly stable. This result is somewhat contradictory with that of Andrés and García (1991), who find strong parameter changes for productivity and employment growth after 1984, when fixed-term contracts we introduced. However, our sample covers a period where this type of contracts have always been used and, moreover, they assume that the employment target is lagged total employment.

Next, in order to analyze the degree of heterogeneity we may have missed by assuming a fixed-effects specification with common coefficients for all variables except for the constant term, we have included in the regression the interaction of the main insider variables (w_{i-1}, Δn_{pi}, $(p+y-n)_i$, ϕ_i) with sectoral dummy variables. The interacted terms for the first two variables are totally insignificant both individually and jointly ($\chi^2(24) = 18.36$), which implies both that the degree of inertia is rather homogeneous and that the insider hysteresis effect is rather robust across industries, as it happened to be across different samples. However, for the last two variables the interaction terms are only marginally insignificant ($\chi^2(12) = 18.56$ for $(p+y-n)$ and $\chi^2(12) = 20.72$ for ϕ) and the individual coefficients λ and c_{33}, reported in the first two columns of Table 3, show some degree of systematic variation across the 13 industry groups in our data set. The point estimates of the insider weight show a range of variation which goes from 0.7% (Sector 2: Mineral extraction) to 24% (Sector 4: Chemicals).

An interpretation of the latter differences, proposed by Draper (1992), classifies sectors into two types: laggard and dynamic. In the former (e.g. Construction materials, Food,

Table 3
The insider weight, the effect of temporary workers and productivity

Sector	λ	c_{33}	c_{33} ranking	strikes ranking	Productivity*	Product. ranking	λ ranking
1	0.200 (1.84)	0.570 (3.36)	2	1	9.5	2	2
2	0.007 (0.64)	0.402 (2.76)	8	8	4.1	7	13
3	0.008 (0.53)	0.422 (3.01)	-	-	4.2	6	12
4	0.239 (2.74)	0.510 (2.97)	5	9	9.6	1	1
5	0.166 (1.97)	0.545 (3.02)	4	2	5.0	4	6
6	0.173 (2.06)	0.298 (1.95)	9	7	7.3	3	4
7	0.117 (2.02)	0.852 (4.02)	1	4	3.2	10	10
8	0.138 (2.12)	-0.617 (3.76)	12	11	4.5	5	9
9	0.075 (1.67)	0.278 (2.00)	10	10	3.1	11	11
10	0.138 (2.35)	0.415 (3.66)	7	5	1.7	13	8
11	0.173 (2.62)	0.464 (3.27)	6	3	3.5	9	5
12	0.147 (2.33)	0.248 (2.72)	11	12	2.9	12	7
13	0.195 (2.26)	0.556 (3.86)	3	6	3.8	8	3

Rank correlation coefficients:

- Between c_{33} and strikes rankings: 0.80 (4.20)
- Between λ and productivity growth rankings: 0.50 (2.00)

Notes: * Average growth rate of real labor productivity in 1984-89.
Asymptotic t-ratios below parameter estimates and correlation coefficients.

Sources: For the strikes ranking, "Boletín de Estadísticas Laborales". For real labor productivity, Hernando and Vallés (1992).

Beverages, and Tobacco), sheltered from foreign competition, wages respond little to changes in the firm's productivity. In the latter (e.g. Chemicals, Automobiles, Other manufacturing), with state-of-the-art technology and open to foreign competition, wages respond to internal conditions. In Draper (1992), it was found in equivalent (more aggregated) industry regressions, that the insider weight varied across sectors and its variation was somewhat related to the degree of *openness* and *competition* across sectors. Although we have not been able to find a similar degree of systematic variation across our 13 industries, we are able to match both the top (sectors 4 and 6) and the bottom (sectors 2 and 9) of her classification.

Following up on this hypothesis, we provide in Table 3 the *real* labor productivity growth by sector (from Hernando and Vallés, 1992)). The rank correlation between λ and the average growth of real productivity over 1984-89 is 0.54 (t-ratio = 2.02). This may have a negative effect on employment growth. More specifically, if low productivity industries are mainly affected by the alternative wage (relative wage effect) whereas the productivity gains in the more dynamic industries feed through partially to wages, the process of employment reallocation can be curtailed, since the job destruction in the laggard industries would not match job creation in the more productive sectors.

We also analyze whether there exists any relation between the systematic variation found in the coefficient c_{33} across industries and some of its determinants as captured by the *harassment* and *discipline effects* described in section 3.3. In order to approximate them we have used an *industrial strife index*, measured by the average number of working days lost per employee in each of the 13 industries (sector 5 was excluded for lack of data) over the period 1986-89. The intuition behind the use of this index is that, although it should not have a clear correlation with the *harassment effect*, it should definitely be negatively correlated with the *discipline effect*. Since the latter affects c_{33} negatively, we should expect to find a positive correlation between the *industrial strife index* and c_{33}. For this purpose, we report in Table 3 Spearman's correlation between the rankings of both variables, which is surprisingly high (0.80) and significant (t-ratio = 4.20). Thus, the evidence in this respect turns out to be extremely favorable to our hypothesis, and may offer another explanation for the progressively wider use of fixed-term contracts, based on the potential increased discipline they provide.

Lastly we turn back to the issue of skill composition. As we mentioned above, the interpretation of the c_{33} coefficient might be mistaken because of the lack of control for skill levels. The argument is that since the skill composition affects the average wage, if the proportion of fixed-term contracts is correlated with skill levels and the latter variable is not included in the regression, the estimated coefficient in the latter would be biased. Therefore, we have performed an informal test of this hypothesis. We do not have information on the skill composition within firms but we do have a proxy for it by sector. Thus we have calculated a *skill index* equal to the proportion of workers in unskilled occupations by sector (the details are in the Appendix). We computed the ranking of the increase in the proportion of fixed-term employment and of the change in the skill index, both from quarter 1987:2 to 1991:2 (which does not correspond to our sample period, but there is no sectoral data on fixed-term contracts before 1987). The correlation is performed in differences instead of in levels, since fixed-term contracts are a new phenomenon. The rank correlation is 0.12 (t-ratio = 0.4), i.e. the necessary condition that there be correlation between skills and fixed-term contracts is not satisfied. So we do not worry further at this stage about this issue, although we think it merits future additional research.

5. Conclusions

In this paper we study wage setting in Spain, as a key factor in trying to explain this country's poor unemployment performance since the mid 1970s. In doing so we need to take into account the institutional change occurred in the mid 1980s, when fixed-term contracts with low firing costs were established.

We formalize the wage-setting process within the insider-outsider theory. We present a modified model in which permanent employees are the insiders and temporary employees are considered as outsiders, so that the latter's interests are not taken into account by the former in the wage bargain. Within the model, the proportion of temporary employment affects wages in three main ways. First, it may increase insiders' wage demands if they disregard temporary employees' interests, since the latter workers would be the first ones to be dismissed, i.e. a *buffer effect*. Second, it has a *composition effect*, stemming from the fact that wages may differ between the two types of workers. Third, it may affect insider's bargaining power. This *bargaining effect* is a composite one, since an increase in the proportion of temporary employment may raise such power because insiders may threaten to withdraw cooperation with temporary employees (the *harassment effect*) or it may lower it because temporary employees are more wary of engaging in strikes (the *discipline effect*).

We test our model using data on a panel of large manufacturing Spanish firms for the period 1985-88. We obtain two sets of results. On the one hand, we analyze the typical features of insider-outsider models: the homogeneity in inside and outside factors, predicted by the model, cannot be rejected, and a few liquidity and market power variables appear to be marginally significant as inside factors, while aggregate labor market variables come in strongly as significant outside factors. We take the significance of financial variables as another reason validating our approach of using firm-level data as opposed to sectoral data. Also, our estimate of the weight of firm-specific factors in wage formation (the so-called *insider weight*) is around 10%, a number close to those found for other countries with similar wage bargaining structures, and almost identical to the value estimated for aggregate wage setting in Spanish manufacturing by Andrés and García (1991). We also obtain strong evidence of membership but not of outsider hysteresis.

On the other hand, we test the features of our model which are more specific of the Spanish experience. First, it cannot be rejected that the interests of employees on fixed-term contracts are disregarded in the wage bargain (i.e., there is a *buffer effect*). We also find evidence of a *composition* and *bargaining effects*, the latter having a positive sign. This is probably the most striking of our results. We perform a sensitivity analysis on the estimated equation, finding that the results are relatively robust.

Lastly, we find some sectoral heterogeneity in the insider weight. This coefficient appears to be larger in sectors with larger productivity growth, whereas outside opportunities (and wages in particular) matter more in the case of sectors with lower productivity growth. As noted in the main text, this may hinder employment growth, since employment destruction in the latter sectors would not be matched with job creation in the former. This could be a structural explanation for the difficulties experienced in lowering the Spanish unemployment rate. There is also evidence of some heterogeneity in the effect of the proportion of temporary employment on wages, which seems to arise from its different dampening influence on wages across sectors.

We interpret these results as implying that the insider-outsider model is quite applicable to the set of Spanish firms we have studied. We infer that the flexibility enhancing measures introduced by the Spanish government have created a two-tier labor market in which temporary employees have little say on wage formation, at least in the manufacturing sector. The fact that permanent employees are protected by a *buffer* of temporary employees may lead to increase their wage demands. Specifically, we find that a 1 percentage point increase in the proportion of temporary employment raises the growth rate of permanent workers' wages by one-third of a percentage point. In order to gauge the overall impact of the introduction of temporary employment on Spanish wages, this effect has to be weighed against the lower wages earned by temporary workers and the disciplining effects they seem to have on insiders' bargaining power.

This application helps us understand the evolution of the wage drift we described in the introduction. On the one hand, the existing empirical evidence of a lower wage for temporary employees points towards the shrinking of the wage drift being the consequence of the *composition effect*: the rapid rise in the proportion of temporary employment lowered down computed average earnings growth. On the other hand, the net result of the *bargaining and buffer* effects of temporary employment on bargained wage rates may easily have been positive, thus explaining the latter's unusually large increases. Since the first effect has already disappeared (because of the stabilization of the temporary employment ratio), the other two may prevail in the future and, if their net effect is positive, the wage moderation experienced over most of the recent boom, 1985-1990, may be over, casting a shadow on the prospects for employment growth, and therefore for a decrease in unemployment, in the future.

We leave a set of issues for further research. On the theoretical side, we would like to build up a truly dynamic model arising from adjustment costs. On the empirical side, we want to try an alternative measure of the labor input, in terms of hours instead of employees, and of the wage, by using bargained wage rates instead of the average labor cost.

Appendix: Sample selection, definitions of variables and sectoral classification

A.1 Sample selection

We started with a sample of 2,445 public and private firms, to which we applied the following filters (details are in Bentolila and Saint-Paul, 1992): (a) Total and rigid employment, sales, net capital stock, cumulated depreciation and labor costs positive every year, (b) Change in permanent employment lower than 0,75 for firms above 50 workers, and (c) Change in labor productivity between -0,67 and 2.

Of the remaining 2,012 firms, 1,214 were private non-energy firms belonging to manufacturing (i.e. with more than 90% of their 1988 sales in manufacturing goods), of which 47 were deleted because their equity or borrowed funds were nonpositive.

A.2 Definition of variables

Employment: Average number of employees over the year. For fixed-term employees, the average number of workers is multiplied by their average number of weeks of work and divided by 52.
Labor cost per employee: Wages plus social security contributions paid by the firm plus other labor costs, divided by employment.
Productivity: Sales divided by employment.
Concentration ratio: Sales of 5 and 10 firms with largest market shares (Source: *Encuesta Industrial*, kindly provided by Cristina Mazón).
Profits per employee: After tax accounting profits divided by employment.
Interest rate: Financial costs divided by total debt. This is

multiplied by a dummy taking the value of zero if debt or financial costs are zero in the current year or if the calculated interest rate is larger than 50%.

A.3 Sector definitions and number of firms

No.	Sector	CNAE Sector[1]	No. of firms
1	Automobiles	36	38
2	Mineral extraction	21-23	29
3	Construction materials	24	90
4	Chemicals	25	174
5	Agricultural and industrial machinery	31-32	217
6	Office machinery, Electric materials, and Electronics	33-35	74
7	Ship-, train-, and plane-building	37-38	16
8	Precision and optical instruments	39	4
9	Food, Beverages, and Tobacco	41-42	180
10	Textiles, Leather, and Shoes	43-45	161
11	Wood, Pulp, and Paper	46-47	108
12	Rubber and Plastics	48	60
13	Other manufacturing	49	16

[1] CNAE is the Spanish National Classification of Economic Activities, at the 2 digit level.

A.4 Computation of the skill index

We grouped the 99 occupations in the Spanish National Classification of Occupations (*Clasificación Nacional de Ocupaciones*, from the Instituto Nacional de Estadística, revised in 1979) into 5 groups: (1) white collar skilled, (2) white collar middle skilled and unskilled, (3) sales people, (4) blue collar skilled and (5) blue collar unskilled. Our sectoral skill index is equal to the sum of employment in groups (2) and (5) divided by total employment. The assignment of the 99 groups into the 5 "skill" groups is available from the authors upon request.

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