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Wage Determination in the Spanish Industry

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Abstract

In this paper we estimate wage equations for the Spanish industry using time series data on 85 industrial sectors, which allows us to distinguish between aggregate and sector specific effects in wage determination. Industry wages respond mainly to economy wide labour market conditions and to a much lesser extent to sector specific productivity gains. The size of the insider effect has not remained stable through the sample period. The estimated equations show a strong transitory effect of unemployment on wages, which is in accordance with the non- stationarity of the Spanish unemployment rate. This hysteresis effect seems well accounted for by the sharp rise in the proportion of long term unemployment.

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1 Introduction.

The process of wage determination has been extensively studied in Spain in an aggregate time series framework. Three main results seem firmly established. The labour market is not competitive; real wages respond to excess labour supply but not to the extent as to clear the market. Wages are, on the other hand, heavily indexed with respect to aggregate productivity. Finally, unemployment and productivity cannot account for the total variation in real wages; other variables must be considered even in the long run. These variables account for the observed non stationarity in the NAIRU whether exogenous (push factors) or endogenous (lagged unemployment or *hysteresis*).

Aggregate wage equations leave some interesting questions unanswered. In this paper we focus on two of these, using a richer data set and following Nickell and Wadhwani (1990) approach to discriminate among competing theories of wage formation. First, we analyze in more detail the influence of productivity on wages; the observed link among these variables in aggregated wage equations is consistent with alternative models of the labour market. However, the extent to which workers capture the productivity gains at the firm (or sector) level is of great importance to interpret the NAIRU and to design suitable microeconomic policies to reduce it.

Finally, we turn our attention to the determinants of the NAIRU and the reasons behind its non stationary behaviour. As in many other countries the estimated NAIRU in Spain follows closely the evolution of current unemployment and displays great inertia. This effect is usually captured in aggregate studies by heavily trended variables (such as taxes); in this paper we investigate the ability of labour market variables to (endogenously) account for non stationarity trying to discriminate among insider power-based and other explanations of *hysteresis*. In Section 2 we set up the theoretical framework discussing the impact of productivity on employment; the empirical results are discussed in Sections 3 and 4. The paper ends with a summary of the most relevant findings and their policy implications.

2 Productivity and Employment.

Productivity growth is good news for employment. Strictly speaking a positive productivity shock will reduce employment only in demand rationed firms. On the other hand, firms operating on their notional demand schedules will rise labour demand, and hence employment, following an increase in productivity. Other goods market features will influence that response: monopoly power, elasticity of demand, etc. However, to properly account for the impact of positive technological shocks on employment we must also look to the supply side. Nickell and Wadwhani (1990) have proposed the following equation which nests the most important non competitive labour supply models

$$w_i = \mu\{p_i+a_i-(1-\alpha)n^*\} + (1-\mu)(W-\delta U+Z) \qquad (1)$$

all variables in logs except U.
w_i, p_i and a_i are the firm nominal wage, price and labour productivity and capital letters refer to economy wide variables such as the aggregate wage (W) and the unemployment rate (U). Equation (1) also incorporates nominal homogeneity and includes a vector of variables **Z** which contains other exogenous determinants of wages such as mismatch, replacement ratio, etc.. It can be formally derived from an insider-outsider bargaining model in which firms keep the right to choose the optimal level of employment. Employers find costly to replace incumbent workers (due to hiring and firing costs). This in turn increases the employed bargaining power (Lindbeck and Snower (1987)), who enter in wage negotiations with some defined employment objective (n^*), which we shall assume alike across sectors. The influence of outside market conditions comes from the possibility of employed losing their jobs in unexpected bad times.

The parameter μ turns out to be of great importance since it determines the ability of employed workers to capture the firm specific productivity gains. When μ is low, the nominal norm for wages at the firm level (w_i) is the aggregate wage (W), and w_i is not isolated from the general labour market conditions. However, this does not mean that the labour market is competitive. In fact, equation (1) with μ set to 0 is compatible with non competitive explanations, such as efficiency wages and union models. The macro and microeconomic implications of the value of μ can be better understood in a labour supply and demand framework.

Let us assume for simplicity that firms are price takers in the goods market, and that their technological constraint is represented by the following production function:

$$Y_i = A_iN_i^{\alpha} \qquad (2)$$

so that the labour demand looks like (3):

$$n_i = (1-\alpha)^{-1}\{a_i-(w_i-p)\} \qquad (3)$$

Let us first assume that firms are homogeneous, in that case upon aggregation we may write down:

$$n = (1-\alpha)^{-1}\{a-(w-p)\} \qquad (4)$$

$$w = p+a-(1-\alpha)n^* - (1-\mu)\mu^{-1}(\delta U-Z) \qquad (5)$$

and using

$$U \approx \log.L - \log.N = l^s - n \qquad (6)$$

we get

$$n = [\mu(1-\alpha) + \delta(1-\mu)]^{-1} [\mu(1-\alpha)n^* + (1-\mu)(\delta l^s - Z)] \qquad (7)$$

Equation (7) gives us a more realistic picture of the impact of productivity on equilibrium unemployment. The NAIRU is independent of technological shocks regardless of the value of μ. In fact, this parameter is not even identified in aggregate wage equations like (5). The close to unit elasticities that are usually found in many econometric studies may well be a pure aggregation effect and carry no information whatsoever about the economic relationship they meant to measure[1].

Despite its apparent unimportance there are reasons why the magnitude of μ is relevant. Although we shall not deal explicitly with this issue here, there may be different components of productivity that are unequally perceived by price and wage setters, in that case the invariance of employment to productivity growth no longer holds[2]. Most important, the value of μ determines the nature of equilibrium unemployment. Setting $\mu=0$ in (7) we get $U=\delta^{-1}Z$, so that unemployment is explained by aggregate push factors[3]; on the other hand, for μ close to 1 aggregate employment is proportional to n with little response to excess labour supply. The policy implications of this extreme cases have little in common: policies aimed to reduce the impact of Z factors in the former versus less market regulation in the later.

These two reasons become still more important if we allow for persistence effects in (7). Under some behavioural assumptions we can write the following dynamic expression[4]

$$n = [\mu(1-\alpha)+\delta(1-\mu)]^{-1}[\mu(1-\alpha)n_{-1}+(1-\mu)(\delta l^s-U_{-1})] \qquad (7')$$

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[1] Andrés et al. (1988) find an elasticity of 0,83. Unconstrained estimation yields a value slightly above 1.

[2] If we split a_i in a permanent and a transitory component

$$a_i = a_i^P + a_i^T$$

and, if we assume further that a term like $(a_i^P+\beta a_i^T)$ enters equation (1), it may be shown upon aggregation that a^T affects equilibrium unemployment if and only if μ is not zero.

[3] The same would obtain if $n^* = l^s$, regardless of the value of μ.

[4] If we assume for instance $n^*=n_{-1}$ and that U_{-1} belongs to Z.

In this case even transitory productivity misalignaments can produce permanent real effects. According to (7') we have again two alternative explanations for the observed non stationarity in equilibrium unemployment in Spain (Andrés and García (1990)). Whether μ is 1 or 0, unemployment shows great inertia: If the insider power is very low, ***hysteresis*** is caused by a deterioration in the stabilizing role of unemployment (due to long term unemployment, reduced search effort, etc.); however, if employed workers manage to capture any productivity gain, a selfish employment target (n^*) may bring the NAIRU into a non stationary path.

Finally, if we allow for differences in productivity across sectors, the size of μ is a key determinant of the wage and employment structure. Along the process of industrial change in Spain employment grew much faster or fell by less in high productivity growth sectors. Jaumandreu (1986) has shown that the employment reallocation across industries accounts for a significant share of aggregate productivity growth. This reallocation is not compatible with any labour market structure; in fact, sizable employment gains in the most technologically dynamic industries cannot be achieved if employed workers exploit in full their insider power; in other words massive employment reallocation is *prima facie* evidence in favour of a small μ parameter.

Relaxing the homogeneity assumption, the i_{th} firm employment level (setting **Z** to 0) is given by (8)

$$n_i = (1-\alpha)^{-1}[(1-\mu)a_i+\mu(1-\alpha)n^*-(1-\mu)(w-p)+\delta(1-\mu)U] \tag{8}$$

and upon aggregation we get

$$n = [\mu(1-\alpha)+\delta(1-\mu)]^{-1}\Big[\mu(1-\alpha)n^*+(1-\mu)\delta l^s+\mu(1-\alpha)\lg\Sigma\lambda_i\Big] \tag{9}$$

where:

$$n = \lg\Sigma N_i \tag{10}$$

$$A = \prod_i (A)_i^{\theta_i} \tag{11}$$

$$\lambda_i = \left[\frac{A_i}{A}\right]^{(\frac{1-\mu}{1-\alpha})} \tag{12}$$

$$\theta_i = \left[\frac{N_i}{N}\right] \tag{13}$$

When firms are not all alike, λ_i is different from one and aggregate employment is no longer independent of the productive structure unless μ equals one. Whenever incumbent workers do not enjoy full monopoly power, aggregate employment also depends on industry specific productivity shocks. The intuition of this result can be seen in (14) (15). From (1) and (8) we get:

$$(w_i - w_j) = \mu(a_i - a_j) \tag{14}$$

$$(n_i - n_j) = (1-\mu)(1-\alpha)^{-1}(a_i - a_j) \tag{15}$$

If μ is 0, firm wages are fully indexed to aggregate wages and across firms variations in productivity lead to huge differences in employment. Technical progress favours a rapid reallocation towards the most dynamic firms. On the opposite side, if μ equals 1, wage structure fully reflects productivity differentials, while employment (growth) converges. In the former case productivity gains in one firm are not fully passed into wages and hence lead to an increase in employment; firms with low productivity growth suffer a substantial wage pressure with increasing unit labour costs. In the latter, unit labour costs remain unchanged across firms and so does the employment structure. There is some evidence in favour of a μ value well below one. Andrés and García (1991) find that the elasticity of relative wage differentials with respect to industry specific productivity is significant but fairly low (below 0.12). Similarly, Pérez and Doménech (1990) estimate a wage equation for the financial sector of the Spanish economy using cross section data obtaining an even smaller elasticity (below 0.09).

3 Productivity and Wages.

In this section we estimate a suitably modified version of equation (1) to evaluate the relative influence of specific and economy wide determinants of wages in Spain. We move from the firm to the industry level taking in account institutional as well as data availability considerations. Wage negotiations in Spain take place mainly at industry level. On the other hand, our basic data set come from the Industrial Survey and contains information relative to 89 industrial sectors for the 1978-1986 period. Aggregate variables are mainly drawn from the National Accounts and the Labour Force Survey. Our data set suffers two major limitations; there is no available information at this stage about financial issues, and the cross section dimension is relatively small[5]. However, working at the industry level we avoid problems of sample representation. The time dimension of data is also particularly relevant, the Spanish industrial sector has gone trough a process of massive jobs destruction both regulated, until 1982, and deregulated thereafter; by 1986 this downward trend is reversed and employment has kept growing very fast in recent years.

The empirical model is represented in (16)

$$w_i = \beta w_{i-1} + (1-\beta)(\mu(p_i+y_i-n_i+z_i) + (1-\mu)(W-\delta U+Z)) \qquad (16)$$

where we allow for slow adjustment in nominal wages and in which a_i has been substituted out by observed labour productivity $(p_i+y_i-n_i)$. Equation (16) also includes some variables that may affect the bargaining power of the insiders (z_i) as well as outside opportunities (Z). Vector z_i includes average firm size (s_i), lagged profits per employee (π_i), and lagged real wages to approximate an error correction mechanism. We have also corrected for average hours (h_i-n_i); the exclusion of this variable leads to a significant downwards bias in the estimate of μ. This can be explained by the strong negative correlation among average hours worked and labour productivity (per man) which seems surprising on a time series perspective but can be reflecting huge technological differences among sectors. On the other hand, is not clear whether hourly or per man variables are relevant in wage negotiations. To account for different speeds in the downwards trend in average hours we have chosen to work in *per capita* terms and to include the average hours correction. Strictly speaking a similar correction should be applied to aggregate wages, nevertheless given the degrees of freedom constraint we have chosen not to include additional aggregate variables and let this effect to be incorporated in the constant term in the model in first differences. This caveat does not apply when we use time dummies to capture aggregate influences; these equations give then the most reliable estimate of μ.

Aggregate variables include wages (both economy wide and labour cost in the industrial sector), unemployment (aggregate rate, industrial sector rate and a less aggregate rate u_j[6]), long term unemployment (proportion

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5 After eliminating four industries for the lack of information about prices we were left with a sample size of 85.

6 The rate u_j is the unemployment rate corresponding to each of the 11

of unemployed searching over twelve (U12) or twenty four months (U24)), mismatch (MM) and replacement ratio (RR). Finally we have also included a transitory nominal variable (lagged unanticipated inflation) to capture any different response to unexpected nominal shocks; this variable also approximates cyclical aggregate shocks given the trendy nature of all other components of the Z vector.

In Table 1, column (1) we present the basic model including eleven industry dummies and with nominal homogeneity imposed. All the equations have been estimated using the Generalized Method of Moments estimator of Arellano and Bond (1988)[7]. There is neither evidence of mispecification in the choice of instruments nor of first order autocorrelation in the undifferenced model (or second order autocorrelation in the differenced equation). Nominal wages adjust very quickly and both hours correction and plant size have the correct sign and are highly significant. Unemployment (in logs) exerts the expected negative influence on wage growth, this effect is nonetheless non permanent since the exclusion of U_{-1} worsens the model properties. The mismatch coefficient is also positive and significant.

The point estimate of μ is around 0.10. Equations in columns (2) and (3) give a similar value of μ. In column (2) nominal homogeneity is relaxed and tested; point estimates of coefficients of nominal variables add up to 1,2 so that homogeneity cannot be rejected at least at the 10 per cent significance level. In equation (3) aggregate variables are replaced by time dummies improving the overall fit but with no substantial change in the remaining coefficients. The model has also been estimated excluding industry dummies yielding a point estimate μ=0.15. Nevertheless the exclusion of this dummies is strongly rejected by the data; there are significant differences in the rate of growth of wages across industries that are not fully explained by the model, if they are not taken into account their effect is partially captured by a trendy productivity variable producing an upwards bias in the corresponding elasticity. We can safely conclude that 0.15 is an upper limit to the value of μ.

The insider power itself may change both over time and across sectors. The Spanish economy has gone during the sample period through a process of deep institutional and political changes. Structural break cannot be

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industrial branches in the Labour Force Survey. For each of the 85 industries we have considered the rate to the branch it belongs to.

7 The GMM estimator enlarges the instrument set providing sizable efficiency gains. Alternative instrument sets have been tried including additional lags of endogenous regressors. Degrees of freedom limitations prevent the use of all orthogonality restrictions in the data set; nevertheless point estimates are not substantially altered ($\hat{\mu} \approx 0.09$).

confused in this case with misspecification. The empirical model comes from a fully specified theoretical framework and the well behaved statistics do not suggest any relevant omitted variable; on the other hand, as we shall see below, unrestricted estimation do not lead to big changes in coefficients. We have carried out a careful search for structural breaks with alternative sample splits. The results in column (4) consider a break in 1984; this seems the most satisfactory specification on empirical grounds and is also justified by the changes in the legal environment of industrial relations. Until 1984 most labour contracts in Spain were permanent with expensive and slow dismissals procedures; by that year the government launched a new set of legal measures to facilitate flexible contracting; in recent years an extraordinary proportion of new contracts is made on a temporary basis[8]. This may have led to a much weaker position of employed in wage negotiations and hence to a lower μ. In fact, this is what equation in column (1) shows. There is a sensible downwards change in μ falling from 0.11-0.15 in the first subsample to 0.0-0.05 in the second one. The Wald test of structural break is highly significant, therefore whatever weight we attach to industry specific variables in wage determination this seems to have been falling over time. The full sample coefficients simply averages out the two subsample estimates.

Equations in Table 1 all include aggregate wage and unemployment to approximate outside opportunities for workers in the i_{th} sector. However, these could be referred to other aggregates, two of which are considered in equations in Table 2. Industrial sector wages can be used instead of aggregate ones as they give close point estimates for μ; however, this way was not pursued further because of some mispecification signs as well as some meaningless parameter values. Industrial sector unemployment rates also capture properly the relevant excess labour supply as equation in column (1) shows. Despite a significant rise in the residual sum of squares and some signs of second order autocorrelation in the differenced model, neither μ nor other relevant parameters change substantially (μ is around 0.07).

The u_j unemployment rate fares rather worse. The equation including current and lagged u_j is presented in column (2), and shows a significant drop in $\hat{\mu}$; nevertheless, there are good reasons to reject this model too. The fit worsens substantially, and, most important, when aggregate influences are captured by time dummies u_j coefficients become wrongly signed. In other words, u_j seems a bad proxy for the influence of labour slack on wages, and it only works because its time series behaviour is close to that of aggregate unemployment; however,

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8 Bentolila and Saint-Paul (1991) find some evidence of this effect by estimating labour demand equations using data on Spanish firms.

once this effect is controlled for, the cross section variation of u_j adds nothing to the explanation of wages dispersion. When across industries mobility is not too low narrow unemployment definition do not approximate properly the outside opportunities for employed workers.

The high elasticity of wages with respect to productivity gains found in many aggregate equations (Dolado et al. (1986), Andrés et al. (1991), Andrés and García (1990)) are consistent with a fairly low ability of employed workers to capture productivity gains right where they are generated. That elasticity comes just from aggregation and has little explanatory power. Industry wages in Spain respond more to relative pay considerations and to aggregate labour market conditions than to industry specific factors than is commonly asserted. This process of wage determination facilitates employment reallocation from less dynamic to more dynamic industries leading to a more homogeneous wage structure.

The estimated μ value in Spain lies within the range of values reported by Nickell and Wadhwani (1990) for the UK and other countries. Low values (around (0.05)) characterize highly centralized wage formation (Sweden, etc.), whereas other countries, such as Japan and USA with more decentralized labour markets present much higher values (0.30). This is also consistent with the commonly held view of the Spanish labour market somewhere in between the two extremes in the degree of decentralized bargaining.

4 Hysteresis in Unemployment.

The structure of the wage equation is a center piece of the supply side of a macroeconomic system. Jointly with the labour demand it determines the equilibrium unemployment. Aggregate wage equations are of little help to discriminate among alternative causes of the observed stationarity in the NAIRU in Spain. In some causes the importance of exogenous supply shocks is stressed (Dolado et al. (1986), Andrés et al. (1991)), whereas the estimation of Phillips Curves leads to an endogenous interpretation of *hysteresis* (Coe (1988), Andrés and García (1990)), where the unemployment rate exerts strong transitory and small permanent downwards influence on wage inflation.

There are two basic approaches to explain unemployment persistence (Layard and Bean (1989)). When insider power is high, wage and employment targets of incumbent workers determine the path of unemployment. On the other hand, transitory shocks may have permanent effects if long term unemployed devote little effort to search or are discriminated in job offers by firms. These two hypothesis can be evaluated in our empirical framework. Results in Tables 1 and 2 display a very low permanent effect of unemployment on wages; any attempt to exclude lagged unemployment rates from the equations were unsuccessful on statistical grounds. In this section we try to account for that effect in a more structural way, bringing into the model two additional variables. A positive coefficient in employment growth (Δn_l) is consistent with an insider-based explanation of persistence, whereas the alternative explanation would be consistent with a positive coefficient for the long term unemployment rate (U12).

In Table 3 we observe a positive and significant coefficient of the long term unemployment rate even in presence of lagged unemployment. In column (1) neither the estimated model nor the μ estimate (0.10) are substantially changed. Although non significant at the 5 per cent level, U12 is a close substitute for U_{-1} which now becomes smaller. The long term rate is strongly significant in an equation with the industrial sector unemployment rate, as in column (2), it also improves the overall fit (standard error) and other statistics (autocorrelation is now lower than in equation (1) in Table(2)). In columns (3) and (4) we find the expected coefficients in U and U12 with opposite signs in well behaved wage equations. Neither μ nor the coefficient on unemployment change and the long term unemployment rate fully captures the impact of lagged unemployment.

The influence of Δn_l is analyzed in Table 4. Unlike the analysis in Table 3 the results are not robust to alternative specifications. Despite a remarkable stability of μ we find a wide range of values for the coefficients of employment growth. In columns (1) and (2) $(\Delta n_l)_{l-1}$ displays a weak positive effect, whereas equations with different aggregate variables give either positive, negative or non significant coefficients on Δn_l. According to the model in (6) one would expect Δn_l and $(p_l+y_l-n_l)$ having a similar dynamics. As we saw in Tables 1 and 2, the effect of industry productivity works with some delay and it has significantly changed over the sample period. We have tried a similar specification search for Δn_l as we did with μ above, allowing for richer dynamics an structural break from 1984 onwards. In column (3) and (4) we present the basic model with time dummies and in both cases the break in μ show up clearly as in Table 1. The behaviour of Δn_l is remarkably similar; hence the influence of insider power in unemployment persistence is significant in the first period and vanishes in recent years. The influence of Δn_l moves according to what happened to μ as it could not be otherwise.

There is yet another piece of empirical evidence drawn from aggregate wage equations that deserves further analysis. The *hysteresis* effect may not be structural. The *level effect* and the *variation effect* can be alternative representations for the impact of excess labour supply on wage inflation (Gordon (1988)). Andrés and García (1990) cannot reject this hypothesis on estimated Phillips Curves for the Spanish economy during the period 1964-1988; the *level effect* dominates while jobs destruction was still moderate (up to 1978) and the *variation effect* has been more important when the unemployment rate speeds up during the last decade. Given the extraordinary trendy nature of the unemployment rate in Spain the *variation effect* could merely reflect the impact of some exogenous shocks (industrial reorganization, falling capital formation, etc.). If that is the case the changing nature of these shocks could lead to structural instability in the *hysteresis* effect.

Again we have carried out a careful search for different breaking points in unemployment coefficients (U, U_{-1}, U12). Results in columns (1) and (2) in Table 5 suffer a severe degrees of freedom constraint and must be taken cautiously, however, according to Wald and t statistics, they give no signs of significant changes in estimated coefficients. In column (2) we find a significant change in U and U12 coefficients from 1984 onwards that, nevertheless, leaves unaltered their sign and relative size. A similar analysis for the u_j rate confirms its poor information contents: the *level effect* remains significant only for the first period, whereas the *variation effect* vanishes altogether.

The results in this section point towards an explanation of unemployment persistence based on the progressive deterioration of labour market conditions reflected in the sharp increase of long term joblessness. Conservative employment objectives by employed workers may to some extent help to explain persistence in the first subsample. The *variation effect* appears nonetheless fairly robust and seems to be present over the whole period. We must bear in mind that 1978-1986 were years of massive job losses in the Industry in Spain and of even faster growth of the long term unemployment rate, we cannot tell whether this result is truly structural and carries over high employment growth periods.

5 Summary and Conclusions.

Sector wages in Spain are mainly driven by relative pay considerations with a significant influence of aggregate labour market conditions. The impact of productivity on wages is low and decreasing, whereas the aggregate unemployment rate seems the best indicator of outside employment opportunities. These features are consistent both with union bargaining and efficiency wage models, and provide a more balanced view than is usually held about the Spanish labour market. This mechanism of wage formation favours the most dynamic industries in which productivity gains increase employment, while the less productive ones suffer from increasing pressures in their unit labour costs. The small permanent effect of unemployment on wages is nevertheless and unpleasant feature for it suggests that the labour market stabilization role has been severely damaged. This may be one of the explanations of the inflationary pressures in recent years when rapid economic recovery led to a fall in unemployment rates for the first time since 1970.

The impulse to employment reallocation enhanced by a low insider power can also be potentially dangerous on the eve of 1993 economic integration in Europe. To the extent that industry specific factors are of little relevance on wage formation, joining a broader labour market with lower unemployment rates and higher real earnings may put an additional pressure on wage inflation in most industrial sectors. As long these enjoy lower productivity levels than the corresponding industries in other European countries, this will contribute to a quick erosion of the relative unit labour costs advantage of the Spanish economy vis a vis the European Community.

We have not considered across industries differences in insider power. A positive correlation of μ_1 with productivity growth is another strand of influence of productivity differentials on aggregate employment. If low productivity industries is mainly influenced by alternative wages, whereas productivity gains in most dynamic industries are fully passed into wages the process of labour reallocation may be interrupted; job losses in less productive industries would not be matched by new jobs in the most productive ones. Industrial reorganization in Spain was one of the causes of employment fall (Bentolila and Blanchard, 1990) and would be consistent with an unequal distribution of insider power. A natural extension of the work in this paper is to asses whether these differences in insider power exist and to what extent they are related to the technological structure across sectors.

Table 1.

Wage Equations. Industrial Survey. 85 Industries. 1978-1986.
Dependent Variable w_i.

	(1)	(2)	(3)[b]	(4)
$(w_i)_{-1}$	0.06	0.06	0.06	0.08
	(2.8)	(2.7)	(2.6)	(2.6)
$(p_i+y_i-n_i)$				0.11
				(5.3)
$(p_i+y_i-n_i)_{-1}$	0.11	0.09	0.07	0.04
	(4.3)	(4.0)	(3.0)	(1.7)
$(p_i+y_i-n_i-W)^*$				-0.20
				(9.2)
$(p_i+y_i-n_i-W)^*_{-1}$				0.10
				(3.3)
(h_i-n_i)	0.40	0.49	0.48	0.45
	(6.3)	(8.0)	(7.4)	(6.3)
s_i	0.06	0.05	0.06	0.04
	(3.4)	(2.5)	(3.0)	(2.0)
U	-0.24	-0.29		-0.25
	(9.1)	(8.4)		(9.6)
U_{-1}	0.29	0.28		0.26
	(13.3)	(12.5)		(9.1)
W	0.94[c]	1.08		0.81[c]
		(10.1)		
W_{-1}	-0.11[c]			-0.04[c]
MM	0.13	0.14		0.11
	(6.9)	(6.2)		(4.6)
RSS	0.939	0.928	0.928	0.916
$\sigma^2 x10^3$	0.954	0.945	0.947	0.937
χ_1	31.6 (29)	30.3 (29)	33.0 (29)	19.22 (26)
A_1	-3.23	-3.05	-2.99	-2.87
A_2	-1.18	-1.11	-0.92	-0.97

U: Log of the aggregate unemployment rate.

Table 2
Wage Equations. Industrial Survey. 85 Industries. 1978-1986.
Dependent Variable w_i.

	(1)	(2)	(3)[b]
$(w_i)_{-1}$	0.09	0.10	0.09
	(3.5)	(4.2)	(3.4)
$(p_i+y_i-n_i)$		0.04	
		(2.7)	
$(p_i+y_i-n_i)_{-1}$	0.06		0.06
	(2.6)		(2.6)
(h_i-n_i)	0.78	0.54	0.47
	(10.9)	(5.2)	(7.3)
s_i		0.04	0.05
		(2.0)	(2.5)
U	-0.11	-0.004	0.003
	(5.8)	(3.5)	(1.7)
U_{-1}	0.18	0.004	-0.003
	(7.5)	(1.7)	(0.8)
W	0.91[c]	0.86[c]	
W_{-1}	-0.06[c]		
MM	0.15	0.08	
	(6.6)	(2.5)	
$\Delta^2 P_{-1}$	0.3	0.29	
	(2.5)	(3.2)	
RSS	0.991	1.010	0.937
$\sigma^2 x10^3$	1.008	1.028	0.960
χ_1	42.2 (29)	43.1 (27)	31.1 (27)
A_1	-2.77	-2.64	-3.11
A_2	-1.37	-1.66	-0.67

Column (1). U: Log of the industrial sector unemployment rate.
Columns (2) and (3) U: j_{th} industrial branch unemployment rate.

Table 3.

Wage Equations. Industrial Survey. 85 Industries. 1978–1986.
Dependent Variable w_i.

	(1)	(2)	(3)	(4)
$(w_i)_{-1}$	0.06	0.09	0.06	0.06
	(3.0)	(3.6)	(2.6)	(2.4)
$(p_i+y_i-n_i)_{-1}$	0.10	0.07	0.08	0.07
	(3.8)	(2.9)	(3.2)	(2.9)
(h_i-n_i)	0.42	0.58	0.50	0.64
	(6.5)	(11.3)	(6.5)	(7.5)
s_i	0.06		0.05	0.03
	(3.3)		(2.6)	(1.6)
U	-0.23	-0.18	-0.20	-0.17
	(8.8)	(8.1)	(8.5)	(7.8)
U_{-1}	0.21	-0.20		
	(3.9)	(3.4)		
U12	0.19	1.45	0.57	0.84
	(1.5)	(7.6)	(10.4)	(10.1)
W	0.94^c	0.91^c	0.94^c	0.94^c
W_{-1}	-0.10^c	-0.07^c	-0.08^c	-0.07^c
MM	0.14	0.17	0.18	0.18
	(6.7)	(8.0)	(8.4)	(8.3)
$\Delta^2 P_{-1}$		0.28	0.35	0.28
		(2.3)	(2.8)	(2.3)
RSS	0.935	0.936	0.936	0.944
$\sigma^2 x10^3$	0.952	0.954	0.953	0.961
χ_1	31.0 (29)	32.4 (29)	34.3 (29)	36.6 (29)
A_1	-3.18	-2.87	-2.82	-2.61
A_2	-1.07	-1.00	-1.06	-1.14

Columns (1) and (3). U: Log of the aggregate unemployment rate.
Columns (2) and (4). U: Log of the industrial sector unemployment rate.

Table 4.

Wage Equations. Industrial Survey. 85 Industries. 1978-1986.
Dependent Variable w_i.

	(1)	(2)	(3)	(4)[b]
$(w_i)_{-1}$	0.06	0.13	0.06	0.03
	(3.1)	(7.1)	(1.9)	(1.1)
$(p_i+y_i-n_i)$			0.13	0.13
			(4.8)	(4.8)
$(p_i+y_i-n_i)_{-1}$	0.10	0.08		0.03
	(4.3)	(3.1)		(1.9)
$(p_i+y_i-n_i)^*$				-0.22
				(8.8)
$(p_i+y_i-n_i-W)^*$			-0.22	
			(8.3)	
$(p_i+y_i-n_i-W)^*_{-1}$			0.08	
			(3.2)	
(h_i-n_i)	0.44	0.47	0.51	0.44
	(7.0)	(6.6)	(8.2)	(8.5)
s_i	0.05	0.06		
	(3.4)	(3.0)		
Δn_i		-0.05	0.13	0.14
		(1.7)	(3.2)	(2.7)
$(\Delta n_i)_{-1}$	0.02	0.04		
	(1.1)	(1.9)		
$(\Delta n_i)^*$			-0.14	-0.13
			(2.2)	(2.0)

U	-0.25	-0.23	-0.25	
	(10.8)	(10.3)	(9.5)	
U_{-1}	0.31		0.30	
	(14.0)		(9.8)	
U12		0.62		
		(7.5)		
W	0.94^{c}	0.87^{c}	0.94^{c}	
W_{-1}	-0.10^{c}	-0.08^{c}		
MM	0.12	0.14	0.16	
	(6.5)	(5.3)	(6.7)	
RSS	0.942	0.945	0.958	0.938
$\sigma^2 x10^3$	0.959	0.965	0.979	0.964
χ_1	37.7 (34)	40.1 (33)	21.6 (31)	23.9 (31)
A_1	-3.25	-3.42	-2.45	-2.15
A_2	-1.27	-1.28	-1.29	-0.56

U: Log of the aggregate unemployment rate.

Table 5.

Wage Equations. Industrial Survey. 85 Industries. 1978-1986.
Dependent Variable w_i.

	(1)	(2)	(3)	(4)
U	-0.32	-0.08	0.006	0.005
	(3.2)	(1.6)	(3.5)	(2.4)
U_{-1}	0.32		0.001	
	(4.2)		(0.9)	
U12		0.47		0.33
		(3.9)		(3.3)
U^*	0.07	-0.15	-0.009	-0.008
	(0.7)	(2.7)	(4.8)	(3.7)
U^*_{-1}	-0.03		0.003	
	(0.4)		(1.2)	
$U12^*$		0.92		0.77
		(4.5)		(4.9)
Ω	0.7 (2)	20.2 (2)	29.1 (2)	28.6 (2)

Columns (1) y (2). U: Log of the aggregate unemployment rate.
Columns (3) y (4). U:Log of the industrial sector unemployment rate.

Ω: Wald Test of the joint significance of the starred variables in the equation.

NOTE: Except for variables in the table, equations are equivalent to Table 1 column (4).

Notes to Tables 1-5.

RSS: Residual Sum of Squares.

σ^2: Estimated Variance.

χ_1: Test for validity of instruments.

A_1: First order autocorrelation in the differenced model, N(0,1)

A_2: Second order autocorrelation in the differenced model, N(0,1)

$X^* = X*D$, D is a dummy 0 in 1979-1983 and 1 otherwise.

b: Equation with time dummies.

c: Restricted coefficient.

All equations estimated in first differences. Industry specific variables are taken as endogenous. The instrument set includes the following lags, w_{i-2}, w_{i-3}, w_{i-4}, $(p+y-n)_{i-2}$, $(p+y-n)_{i-3}$, $(h-n)_{i-2}$, Δn_{i-2}. Estimation has been carried out using the DPD program of Arellano and Bond.

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