Bargaining power and local labour market influences on wage determination

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Abstract

This paper uses a unique panel of data at the level of the bargaining group to examine aspects of `right-to-manage' models of wage determination. Empirical measures of firms' and unions' bargaining power are identified and found to be important influences on wage setting. The role of union characteristics in wage determination is examined; results confirm their importance and illuminate previous survey findings. Features of the local labour market are shown to affect bargained wages over and above the influence of aggregate factors.

Keywords: Wage determination, bargaining power, panel data.

JEL classification: J30, J51, R23.

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1. Introduction

Models in which wages are determined as the outcome of bargaining between ¯rm and union have become widespread over the last decade, following the pioneering work of Nickell and Andrews (1983) and Layard and Nickell (1986). This paper makes use of a unique dataset at bargaining unit level, covering hundreds of settlement groups in a major UK industry over more than a decade, to make one of the ¯rst assessments of the in°uence of bargaining power and alternative options. In addition, the paper presents new ¯ndings relating to the role of local labour markets in wage determination.

Because the level of aggregation of data in this paper matches that of the theory, our results enable us to assess whether the success of bargaining models at a more aggregate level stems from their accurately capturing the behaviour of micro units. Many interesting theoretical concepts arising from bargaining models cannot be examined at the aggregate level because at that level they are unobservable or have no meaning (the `representative' ¯rm's outside option and union members' alternative wage, for example). Micro datasets rich enough to allow even a subset of these concepts to be tested are rare. Those that exist often have serious °aws in terms of data measurement in addition to the constraints of data availability.

Two other UK studies have recently reported results at a similar level of disaggregation. Nickell et al. (1994) include one equation estimated on basic wage data from the Confederation of British Industry (CBI) Pay Databank, matched to performance data from company accounts.[1] Hildreth and Oswald (forthcoming) report results of estimating bargaining and `modi¯ed competitive' models for 58 establishments operating in a variety of industries, mainly based in the West Midlands, during 1980 to 1986.[2] Our work is unique in providing insights into what factors in°uence the distribution of rents between ¯rm and workers and assessing the particular in°uence of local labour markets. Other micro-level studies have used subjective reports of rela-

[1] The dependent variable in equation 6, Table 2, p. 25 of Nickell et al. (1994) is the \weekly rate for given hours and given skill group" paid to certain workers in 78 settlement groups from various industries in the UK during 1981 to 1986.

[2] Hildreth and Oswald's dependent variable | the establishment wage bill divided by employment | is potentially a®ected by problems discussed in Section 2 in relation to similar measures at company level. We also adopt a di®erent approach from Hildreth and Oswald to potential endogeneity of regressors, aided by a larger sample.

tive company performance from the CBI Pay Databank (Gregory et al. 1986, 1987), or the British Workplace Industrial Relations Surveys (WIRS) in a cross-sectional examination of wage determination in the UK (Blanchflower et al. 1990). Subjectivity presents its own problems, while it is difficult to determine causality and evade endogeneity problems when working with cross-sections.

Our results indicate that a major determinant of firms' bargaining power is their inventory. In accordance with strategic considerations, higher stocks seem to enable the firm to better survive a stoppage, increasing the firm's negotiating strength. We also find support for the hypothesis that company bargaining power is reduced by greater capital intensity (which might raise interest and overhead costs during a stoppage). Although alternative interpretations of these results are possible, they are consistent with a bargaining model of wage determination. Higher outside wages, representing the reward that might be obtained if negotiations broke down, enable the union to push for higher wages in the current agreement. Workers' bargaining power also depends on features specific to the main union involved in bargaining (such as the size of its membership, its reputation, financial strength or information). We examine the impact of multi-unionism on union bargaining power: results suggest that more unions bargaining together increase negotiating strength, rather than result in coordination problems. This could explain why multi-unionism has persisted despite the general decline in unionisation. We also find that greater union density within the bargaining unit raises pay.

We find that wages are affected by characteristics of the local labour market over and above the influence of aggregate labour market factors. Our results also indicate that improvements in company performance are associated with higher basic pay. We estimate that a 10% increase in profitability raises wages by 4%.

The structure of this paper is as follows. The data used are described in Section 2 and issues arising in econometric estimation in Section 3. Section 4 explains the basis for the modelling of bargaining power. Section 5 discusses our results, comparing them with previous findings, and Section 6 concludes. The theoretical framework for the analysis is set out in an Appendix.

2. The data

The dataset is a large unbalanced and holed panel, containing a reasonably large number of cross-sectional units on which there are relatively few time-series observations spanning 1981 to 1989.[3] The data feature four levels of aggregation: bargaining group, company, region and whole economy. Summary statistics can be found in Table 4 in the Appendix.

The units of study are bargaining groups that cover unskilled manual workers in the UK chemical and allied industries. A bargaining group might be cover these workers in a plant — for example, Albright and Wilson's Marchon Works in Whitehaven, Cumbria (producer of detergent chemicals) — or might span a division — for example, Fisons Pharmaceuticals — or might cover a whole company, such as Yorkshire Chemicals. Bargaining unit level data are drawn from annual surveys conducted by a major union in the industry — the General, Municipal, Boilermakers and Allied Trades Union — and from the Reports and Bulletins of two UK pay research bodies, Incomes Data Services and Industrial Relations Services.

Bargaining unit data have been matched with data from company accounts (taken from Extel Financial Services data tape and Kompass Company Directories).[4] Using group or other 'large' profit centre accounting variables is potentially no less misleading than using small-centre (such as divisional, or even site or plant) data. For example, it may be difficult to monitor the profit which is due to small-scale operations. In addition, small profit centres are constituent parts of a large organisation, and the large company will often be able, and might be motivated, to distort the profit or balance sheet picture at the lower level by inter-group transfers, purchases, credit, etc. Pay negotiators might have good reason to refer to profit centres at significantly higher levels of aggregation than the actual bargaining unit. Daniel's (1976) survey of UK establishments revealed that union negotiators are frequently given little or no information about financial performance by management, and so have to rely on

[3] The shortness of individual groups' time series is almost entirely due to missing bargaining group level data. The fact that bargaining group level data are derived from surveys leaves open the possibility of bias from this source, but Smith (1994) examines this in detail and concludes on the basis of out-of-sample information that bias is unlikely to be substantial.

[4] The companies are mainly, but not exclusively, publicly quoted.

publicly available data. The 1990 Workplace Industrial Relations Survey (WIRS) also records that when asked about in°uences on the most recent pay settlement, many managers in private sector establishments cited \considerations about the employers' ability to pay. The majority of these responses either mentioned the commercial and ¯nancial performance of the establishment itself (or the company to which it belonged) or else, more generally, the employers' ability to pay" (Millward et al. 1992, p. 239, emphasis added).[5]

A major problem with data used in previous micro-level studies is the wage variable itself. Most researchers have resorted to a wage bill-based measure | generally the wage bill divided by employment | which is necessitated by the use of company accounts data alone. Measuring remuneration as the wage bill per employee has several drawbacks. First, if employment is used also as a regressor, measurement error bias might result. Second, the wage bill aggregates skill groups, so changes in the skill composition of employment could distort results. Third, a substantial body of evidence suggests that the basic wage, and not the total pay bill, is the focus of bargaining. Fourth, the wage bill includes all sorts of payments: overtime payments might generate spurious correlation between the wage variable and cyclical demand; the interpretation of a signi¯cant relationship between earnings and company performance is confused by the inclusion in the wage variable of payments directly related to company performance (pro¯t sharing, productivity and other bonuses).

To avoid these problems, it is necessary to resort to a `pure' wage variable: one which is directly observed for a given (type of) worker, and one which excludes performance-related elements. These factors argue in favour of looking to data on `basic' rates of pay. This paper is one of the ¯rst to make use of such data for a given occupation at the level of individual settlement groups.

The dependent variable in all equations is the natural logarithm of the annual[6] basic wage paid to the lowest grade of unskilled manual worker who is not a cleaner or a canteen worker, de°ated by the UK retail price index excluding mortgage interest

[5] It is relevant to the results in this paper that, for unionised private sector establishments, performance was the second most cited in°uence on pay settlements after the cost of living.

[6] The raw data refer to annual, weekly or hourly basic wages; weekly wages were annualised by multiplying by 52, and hourly wages were annualised by multiplying by standard weekly hours times 52.

payments. This grade rate is often described as `general worker' or `labourer'. Since the job is the same over time and across establishments, this wage variable covers workers with similar skill bundles, thus controlling to a large extent for di®erences in human capital (education, age, experience).

On the face of it, our results relate to the determination of wages of unionised unskilled production and general workers in the UK chemicals industry. However, the results can be regarded as applying to unionised chemicals production and general workers in general. The wage of the unskilled worker, which is usually the lowest basic wage in the production and general workers' bargaining group,[7] ubiquitously forms the focal point of wage negotiations for that category of workers in unionised ¯rms in the UK.[8] Furthermore, it is consideration of factors applying to all workers in the bargaining unit that will in°uence the outcome of negotiations, as all workers in that group are a®ected by the result: each settlement applies across all workers in the bargaining unit. The wage of unskilled production workers is taken to be representative of production and general workers' bargaining group as a whole.

The dataset is unbalanced: the number of time periods varies across individuals. Although twelve years are covered by the sample, there are only a few observations on most groups. The low number of time periods in such small-T panels limits the number of aggregate variables (those with no cross-sectional variation), since the degrees of freedom in the time dimension are restricted. Including too many aggregate variables could lead to multicollinearity and badly-determined estimates. Earlier work using company-level data has used outside wages, unemployment and vacancies at national level because company accounts do not usually include information about the location of production. The bargaining unit level of the data in this study allows the regional dimension of these variables to be exploited.

[7] Occasional exceptions arise where canteen or cleaning workers are included in the bargaining group. Canteen and cleaning workers were deliberately excluded from our analysis to obtain uniformity of grade across groups.

[8] Evidence for this is widespread: for example, the wage of lowest-grade manual labourers who are not canteen or cleaning workers is used as the rate representative of the settlement for the whole bargaining group by commercial pay research groups, including Incomes Data Services and Industrial Relations Services.

3. The choice of estimator

The equations to be estimated take the following general form:

$$w_{it} = \alpha_i + \beta' X_{it} + \gamma' Z_{jt} + \phi' R_{Rt} + \varepsilon_{it}. \tag{3.1}$$

X_{it} is a matrix of bargaining unit level variables, Z_{jt} is a matrix of company-level variables, and R_{Rt} is a matrix of regional variables that may be replaced or augmented with aggregate variables.

Because we use annual data it is possible that, within a period, the pay settlement occurs before the accounts reporting date. We lag company-level accounting variables to eliminate this possibility, as we want accounts-related data to be known at the time wages are set. We can allow for potential remaining endogeneity in regressors derived from company accounts by instrumenting these with their own previous two lags.

Panel data enable the researcher to take account of heterogeneity across individuals. Our wage equation allows the intercept term to differ across individuals, while slope coefficients are constrained to be the same. Individual effects are presumed to take one of two possible forms. First, individual effects α_i might be fixed and certain, in which case they can be captured by incorporating N dummy variables, one for each individual i, as in the 'fixed effects' model, which is estimated using least squares.[9] This least squares dummy variables (LSDV) estimator fully controls for between-group variation and implies that the coefficients measure what are essentially time series effects, averaged over bargaining groups. Cases where there is only one observation to a bargaining group would be dummied out by LSDV and are therefore dropped. Alternatively, individuals in the sample might be considered randomly drawn from some underlying population, in which case their specific characteristics α_i are better modelled by allowing them to be random (the 'random effects' or 'error components' model) which is estimated using feasible generalised least squares (FGLS). Hsaio (1986) has argued that use of the LSDV estimator implies that inferences should be restricted to the sample alone, whereas the FGLS estimator allows inferences to be

[9] In practice, this model is estimated by OLS after the data have been transformed by differencing from individual means or by first differencing. Both transformations have the advantage over OLS with N dummy variables of reducing the size of the matrix to be inverted.

drawn regarding the underlying population because of its random element. Statistical methods can also help decide between competing estimators.

We therefore conducted preliminary estimation of a basic specification by different methods.[10] In addition to comparing LSDV and FGLS estimators, we also investigated the relative merits of ordinary least squares (OLS), two stage least squares (instrumental variables), and two stage least squares with fixed effects. The sample was restricted to bargaining groups for which there were at least four observations ($T_i \geq 4$), two of which were used to create lags used for instrumenting.[11] It turns out that the choice of estimator is important, as the coefficient estimates differed across equations in their sizes, significance and, occasionally, signs. Fortunately, we are able to make some assessment of which estimator is preferred. The Breusch-Pagan Lagrange multiplier statistic (Breusch and Pagan, 1979) rejects the null hypothesis that there are no individual (random) effects, indicating that the FGLS estimator is to be preferred to OLS (statistic $\chi^2_1 = 383.37$, p-value < 0.001). Inspection reveals quite substantial differences between the coefficients estimated by LSDV and FGLS, which might well indicate that the individual effects are correlated with the regressors, in which case the FGLS estimates would be biased. This is supported by a Hausman test (Hausman, 1978); the null hypothesis that the FGLS is inconsistent whereas the LSDV is consistent and efficient cannot be rejected (statistic $\chi^2_8 = 23.99$, p-value $= 0.002$). Our interpretation will therefore focus on specifications including individual fixed effects. As remarked, we will also use instrumental variables to control for endogeneity.

4. Modelling bargaining power

The wage equation that forms the basis of our empirical analysis can be derived as a log-linear approximation to the first order condition for the Nash bargaining problem when a profit maximising firm faces a union that cares about the wage alone (as in

[10] For brevity, full results are not reported here but are available from the author on request.

[11] It would have been possible to use all observations ($T_i \geq 1$) for the OLS and FGLS regressions, and all cases where there were at least two observations ($T_i \geq 2$) for the LSDV estimation. The use of the same sample for all estimators is intended to provide a level playing field for comparisons. Results were similar when the maximum possible number of observations was used.

the ‘right-to-manage’ model). The wage depends on the union's outside option $\underline{\Omega}$, the firm's outside option (per employee) ($\underline{\Pi}/N$), profit per employee (Π/N), and the relative bargaining power of firm and union, Θ:[12]

$$W = W\left(\underline{\Omega}, \frac{\underline{\Pi}}{N}, \frac{\Pi}{N}, \Theta\right). \tag{4.1}$$

The interpretation of $\underline{\Omega}$ and $\underline{\Pi}$, the fall-back options of union and firm, and their relative bargaining power Θ, is crucial for the formulation of empirical models of wage determination. It is possible to represent the bargaining power of the two parties by their ‘inside’ options — the cost/benefit of a delay to agreement.[13] The inside option is non-zero when at least one of the bargaining parties has an exogenous source of income or incurs exogenous costs during the bargaining process. Inside options could well affect the outcome of negotiations by affecting the parties' bargaining power through their relative time preferences. If one party has greater utility due to income from other sources while a bargain is being struck, that party will be more willing to bear some period of disagreement and might therefore be able to bargain for a higher share of the pie.

It is sometimes claimed that in a static model of wage determination “it is natural to assume that, in the event of a dispute with the workers, the inside option of the employer is to make zero profits” (Manning 1991, p. 327). This is the assumption made by Layard et al. (1991, p. 101); they define the firm's objective to be operating profit and infer that Π in expression (4.1) represents excess income when there is no strike and in the case of a strike the firm gets nothing. This might not tell the whole story: in the case of a strike, a firm may well have fixed costs even when no production is taking place, in which case the firm's inside option will be less than zero profits. The fixed costs might be in the form of overheads (heating, lighting, storage, support staff), hiring charges on leased equipment or interest payments on debt. The greater the firm's liquid assets in relation to these fixed costs, the easier it will be for the firm to continue to meet its financial commitments during a stoppage: liquid assets might raise bargaining power and reduce wages. Nickell and Wadhwani (1990) argue instead

[12] For further details see the Appendix.

[13] This ‘inside’ option can be distinguished from the ‘outside’ option, which is the cost/benefit of quitting the current relationship and, for example, finding an alternative bargaining partner.

that ¯rm liquidity can be used to represent one aspect of insider power, in which case the e®ect on wages is reversed.

Since a ¯rm must continue to pay rental on its capital equipment during a strike, the ¯rm's inside option should vary negatively with capital intensity (see Currie and McConnell 1992). Furthermore, the inclusion of the ratio of capital stock to employment has been commonly supposed to \control for the potential e®ects [on wages] of capital accumulation and technical progress" (Denny and Machin 1991). The capital-labour ratio might capture the trend of labour productivity; (some proportion of) productivity gains may accrue to workers in the form of higher wages. Unambiguously, a higher capital-labour ratio should raise the wage.[14]

Binmore et al. (1986) suggest that the ¯rm's inside option can be regarded as the \income from temporary arrangements that keep the business running" (p. 177). This is assumed to vary positively with the ¯rm's holdings of inventories, since these can be used to bu®er the ¯rm against the e®ects of a strike (for a theoretical model incorporating these e®ects in a Nash bargaining framework, see Clark 1991 and 1993; see also Currie and McConnell 1992; Doiron 1992).[15] Loss of production due to stoppages associated with settlement delays might cause potential customers of that ¯rm to turn to alternative suppliers unless the ¯rm has su±cient stocks to maintain supplies, which would impose costs on the ¯rm due to lower future revenues and pro¯ts if the probability that the customers would not return was less than unity.[16] The higher are stocks, the longer the ¯rm can last before the e®ects of lost production are felt in terms of reduced sales and pro¯ts. Larger inventory should increase ¯rms' bargaining power and reduce wages. It is clear from equation (4.1) that all factors representing

[14] The capital intensity of the ¯rm could a®ect the speci¯city of workers (i.e. their relative value to the ¯rm) and hence in°uence the bargaining power of the union. The higher the capital stock per employee, the more responsible each employee might be for the production performance of the ¯rm. Again, a higher capital-labour ratio would tend to raise the bargained wage. But this interpretation is improbable in the case of the data used here, since they refer to unskilled production workers whose ¯rm-speci¯city is likely to be low.

[15] The accumulation of inventory increases the demand for labour in earlier periods. Clark (1991; 1993) shows that despite this it might bene¯t the ¯rm to undertake strategic inventory accumulation, in order to reap the bene¯t of lower wage rates in later periods.

[16] Lower future revenues are equivalent to a smaller future pie, which can reduce the union's future payo®s as well as the ¯rm's. Thus there may be negative elements within the union's inside option. Dalmazzo's (1992) static model incorporates this possibility by assuming that the pie diminishes in size over time.

the ¯rm's outside option should be de°ated by employment. This also makes sense empirically, as a control for ¯rm size, since larger ¯rms are likely to have larger liquid assets and inventories.

What does the union receive in terms of `inside option' if there is a delay in reaching an agreement? Binmore et al. (1986) mention several possibilities: if the delay is associated with a strike, union members might receive income from union strike funds or from temporary employment elsewhere (if they can get a job).[17] The income of the rest of the worker's household could form an inside option; this income might rise if the striking worker's spouse is motivated or enabled to obtain temporary employment as a result of the worker being at home because of the strike. Thus the union's inside option will be related to wages elsewhere in the economy. Various measures of the alternative wage have been included in previous empirical work. The sample average wage for each year can be used as a measure of the `inside option' wage (see, for example, Currie and McConnell 1992). During a delay, workers (or their spouses) might be unwilling to look for temporary jobs outside their local area, in which case the average wage prevailing in the worker's region might be more in°uential.

Variables are incorporated to capture workers' expectations about whether, in the case of a bargaining delay, they or their households will bene¯t from the alternative wage, which depends on the probability of getting the `alternative' job. This probability might be inversely related to the rate of unemployment, suggesting a negative relationship between unemployment and pay. The chance of ¯nding alternative employment (the disutility of losing the current job) might also be related to the ratio of vacancies to unemployment: the greater the number of jobs o®ered in relation to the numbers searching for jobs, the greater a striking worker's chance of ¯nding temporary employment. We would therefore expect a higher vacancy/unemployment ratio to be associated with higher union bargaining power and higher wages.

We also examine whether increases in the proportion of the unemployed who have been without jobs for more than one year a®ect wages. Following the work of Layard and Nickell (1986) it is commonly supposed that, for a given unemployment rate, an increase in this proportion will have a positive e®ect on the bargained wage and

[17] No strike data are available at bargaining unit or company level, so we have no measure of how often the costs associated with this type of bargaining delay are incurred.

the expected value of employees' inside options. The long-term unemployed might be `outsiders' incapable of seriously contending for `insider' jobs, perhaps because they search less hard, or perhaps because employers regard the fact they have been unemployed so long as a bad signal.

The disaggregated level of the data allows us to include of bargaining unit level variables that capture trade union power more directly. Bargaining power might be systematically related to the identity of the (main) trade union: for example, trade unions with higher total memberships might have greater bargaining power as a result of better finance, information, or reputation. The number of unions involved in bargaining could also affect the outcome — given the possibility of coordination problems, a higher number of unions bargaining together might not achieve wage increases as high as a single union would. Alternatively, unions might gain strength through a `collective voice' effect, in which case the expected sign would be reversed.

Union bargaining power, and hence wages, are likely to be positively related to union density. The greater the proportion of the workforce who are members of the union, the more damage the union could inflict on the firm by a strike, and therefore the less willing the firm might be to pursue a bargaining strategy that could induce a stoppage. Union density affects the bargaining parties' relative bargaining power through (its negative effect on) the firm's inside option.

A summary of predictions concerning the signs of coefficients on the explanatory variables considered is presented, together with our findings, in Table 3 (page 24).

5. Results and interpretation

The discussion of empirical results first focuses on the role of bargaining power, then turns to the importance of regional factors, and finally assesses the influence of company performance in wage determination.

5.1. Bargaining power

5.1.1. Firms' bargaining power

In accordance with strategic considerations which inform bargaining models of wage determination, the level of the basic wage is affected by the level of inventory per employee, $(\widehat{stk} - n)_{jt-1}$ (see, for example, column [1]).[18] The higher the level of stocks, the greater the ¯rm's ability to survive a temporary stoppage and the greater its bargaining power, with the consequence that the ¯rm is more able to resist workers' wage demands.

This `strategic' argument might be thought to apply most strongly to stocks of ¯nished goods, as opposed to work in progress and raw materials, which are not saleable without further work that could not be carried out if there were a strike. Company accounts data allow a decomposition to be made of total stocks into ¯nished goods, work in progress and raw materials (see column [2]). It turns out that the only signi¯cant negative e®ect on wages comes from stocks of raw materials, $(\widehat{raw} - n)_{jt-1}$. Firms with stocks of raw materials per employee of one standard deviation above the average ($7,974, compared to an average of $4,025) pay real wages that are 10% lower (the equivalent of $11.65 a week for the average-wage ¯rm).[19] The importance of raw materials could indicate ¯rms beginning to stockpile in advance of wage negotiations (recall that inventories are lagged).

But the empirical relationship between inventories and wages could re°ect a more traditional, cyclical, e®ect. An unanticipated positive demand shock leading to a low level of stocks will tend to induce ¯rms to raise output in later periods to restore stocks to their original levels. Labour demand, and wages, would rise (the rise in wages could re°ect workers' ability to grab some of the higher rents resulting from the increased product demand). A high level of inventories might indicate a lower-than-expected level of demand for the ¯rm's product, which might induce workers to accept a lower pay settlement.

Despite compelling theoretical arguments discussed in Section 4, we ¯nd that

[18] Variables with hats are instrumented with their own two previous lags. Lower case letters denote variables in natural logarithms. j subscripts denote company level, and i bargaining unit level, variables. Company level variables are expressed in real terms and are de°ated by company employment. Until 1982 UK accounting regulations required companies to disclose only the number of UK employees. After this date, group totals are reported. In this paper, company employment is `Domestic employment' where available and `Total employment' otherwise. The lack of measurement error was con¯rmed by re-estimation of equations using data covering 1982{ and 1983{1989 (results not reported here but available on request).

[19] Wald tests conducted to check for equality across the stocks coe±cients con¯rm what is obvious from inspection of column [6]: that (because they are so ill-determined) there is no signi¯cant di®erence between the coe±cients on ¯nished goods and work in progress, but that the coe±cient on raw materials is signi¯cantly di®erent from each of these.

	[1]	[2]	[3]	[4]	[5]	[6]
$(\widehat{\pi} - n)_{jt-1}$	0.0408 [3.05]	0.0557 [2.73]	0.0369 [2.35]	0.0434 [4.13]	0.0504 [3.75]	0.0522 [1.38]
$(\widehat{k} - n)_{jt}$	0.0605 [2.31]	0.0294 [0.97]	0.0579 [2.16]	0.0301 [1.54]	0.0183 [0.86]	−0.0207 [−0.19]
$(\widehat{stk} - n)_{jt-1}$	−0.107 [−2.92]		−0.115 [−2.84]	−0.0517 [−1.58]	−0.123 [−3.79]	−0.0336 [−0.27]
$(\widehat{liq} - n)_{jt-1}$			0.0119 [0.46]			
$(\widehat{fin} - n)_{jt-1}$		0.0076 [0.18]				
$(\widehat{wip} - n)_{jt-1}$		0.0114 [0.45]				
$(\widehat{raw} - n)_{jt-1}$		−0.160 [−2.94]				
UR_{Rt}	0.317 [1.17]	−0.129 [−0.41]	0.290 [1.05]	0.0180 [0.07]	0.0837 [0.27]	−0.100 [−0.15]
$(V/U)_{Rt}$	0.186 [1.23]	0.154 [1.02]	0.180 [1.19]	0.112 [0.75]	0.112 [0.70]	0.275 [0.90]
$UR52_t$	−0.212 [−1.74]	0.0211 [0.16]	−0.210 [−1.73]	−0.152 [−1.39]	−0.0275 [−0.20]	−0.0112 [−0.04]
w_{Rt}	0.342 [5.18]	0.203 [2.66]	0.342 [5.17]	0.241 [4.14]	0.203 [3.12]	0.165 [1.02]
$wedge_{Rt}$	−0.772 [−2.97]	−1.194 [−4.15]	−0.749 [−2.83]	0.118 [0.39]	−0.375 [−1.27]	0.250 [0.26]
$TU_{i} \times t$				0.0009 [3.54]		
$TUNO_{i} \times t$					0.0029 [4.48]	
$DENS_{i} \times t$						0.0234 [1.95]
$\bar{R}^2$	0.904	0.908	0.904	0.944	0.946	0.749
F test	37.90 (119;349)	33.61 (80;184)	37.55 (120;348)	69.84 (78;243)	71.91 (59;180)	110.64 (49;116)
Log likelihood	780.6	514.9	781.0	623.0	463.9	234.2
Wald (wedge)	0.776 [0.381]	0.454 [0.501]				
Autocor. coef.	-0.032	-0.095	-0.033	0.173	0.175	-0.095
Observations	469	265	469	322	240	166
Bargaining units	112	71	112	70	51	41

Table 1: The influence of bargaining power on wage determination

Notes: Dependent variable is natural logarithm of lowest real basic wage in bargaining group including production and general workers. All equations estimated by instrumental variables with individual fixed effects. Sample 1981-89. $\widehat{}$ over mnemonic indicates variable instrumented with two previous own lags. Upper case letters refer to real-valued variables, lower case to natural logarithms. t-statistics in square brackets. F test: for joint significance of regressors and individual dummies. Wald (wedge): test of hypothesis that coefficient on $wedge_{Rt}$ is unity (χ^2_1). For mnemonic definitions see Table 3.

higher liquid assets $(\widehat{liq - n})_{jt-1}$ do not seem to increase ¯rms' ability to resist wage claims by reducing the costs of a temporary stoppage.[20] Nor do they reliably indicate a greater `ability to pay', although the positive coe±cient suggests that this is more likely to be the nature of their role (see column [2]).[21] Insigni¯cance of liquid assets could then re°ect collinearity with pro¯ts.

Workers in more capital-intensive ¯rms receive higher wages, as predicted (see the generally positive coe±cient on $(\widehat{k - n})_{jt}$ in Tables 1 and 2). This is consistent with hypotheses discussed in Section 4: that capital intensity lowers ¯rms' bargaining power by raising (rental or interest) costs during a temporary stoppage; or that this ratio captures productivity increases which workers with bargaining power are able to grab in the form of higher wages.

5.1.2. Union bargaining power

Trade union bargaining power, modelled by elements of workers' `inside' options, signi¯cantly a®ects the bargained wage. Our results indicate that the wage elsewhere is positively related to the bargained wage (the coe±cient on the regional wage w_{Rt} is estimated at between 0.2 and 0.3). Among other things, this con¯rms that wage setting has an important role in the dissemination of shocks across ¯rms.

The coe±cient on the regional tax and price wedge $wedge_{Rt}$ is quite unstable. It is sometimes indistinguishable from unity (see the Wald test reported in columns [1] and [2], for example), which suggests that employees bear all the costs of higher employment-related taxes on the employer and lose all the bene¯ts of lower direct taxes, in the form of lower basic wages. This might be expected, given that workers care about the real consumption wage (the dependent variable), whereas the bargaining framework actually determines the real production wage.[22]

The performance of other labour market variables is relatively disappointing. The regional vacancy/unemployment ratio $(V/U)_{Rt}$ is insigni¯cant, although its positive

[20] Liquid assets face the same potential endogeneity problem as pro¯ts, in that, for a given level of company performance, higher wages will be associated with lower liquid assets. We control for this potential endogeneity by lagging and instrumenting, as described in earlier.

[21] In their study of US contract wages, Currie and McConnell (1992) also report an insigni¯cant (but large) coe±cient on liquid assets per employee.

[22] See, for example, Barrell (1994), pp. 229-30.

coefficient is consistent with the hypothesis that this variable indicates the ease of finding work. Wages also appear to be unaffected by the rate of regional unemployment, UR_{Rt}.[23] A higher proportion of long-term unemployment within the total ($UR52_t$) also appears to have little effect on real pay, contrary to the widely-held `insider-outsider' view. One possible explanation for a negative coefficient on long-term unemployment could be that a higher long-term unemployment rate might indicate to workers worsening economic prospects, thereby inducing them to accept lower wage rises. Our results compare with those of Christofides and Oswald (1992) for Canadian settlement groups, where there is a negative coefficient on the proportion of long-term unemployed (their estimate is also not well determined).

Columns [4], [5] and [6] investigate the influence on wage setting of observable features of trade unions. We are interested in the bargaining power of unions which represent lowest-grade manual workers (or, more precisely, those unions which are involved in bargaining on behalf of the group of workers that includes lowest-grade manual workers). So it is at the level of this bargaining group that we measure the features of unions.

Column [4] shows the results of adding TU, a dummy variable denoting the identity of the main trade union recognised for bargaining purposes.[24] This is intended to allow for differences in bargaining power across unions that are not picked up in general

[23] The regional unemployment rate generally enters positively. A majority of previous studies has found a negative coefficient, and a `pressure of demand' interpretation has generally been invoked (see Blackaby and Manning, 1990; Christofides and Oswald, 1992; Blanchflower and Oswald, 1994). But there are precedents for positive coefficient (see, for example, Nickell and Kong's (1992) results for the chemicals industry; Beckerman and Jenkinson, 1990; Sanfey, 1992; Forslund, 1994). A positive coefficient on (lagged) unemployment could reflect hysteresis (Nickell and Kong 1992; Beckerman and Jenkinson 1990). An alternative interpretation is that the inclusion of the vacancy/unemployment ratio has already captured the difficulty and costs of search, which are often presumed to lie behind the hypothesised negative unemployment effect on wages. The regional unemployment rate is then free to capture `compensating variations', namely the wage premium that must be paid to induce workers to accept jobs which are identical to other jobs in all respects except that they are located in an area where unemployment is higher (Roback, 1982; Harris and Todaro, 1970). High local unemployment could induce wage premia for a number of reasons. If the risk of layoff is related to the level of local unemployment or the size of the pool of labour looking for work in the area, the worker will need to be compensated for the increased risk of working in a high-unemployment location. High unemployment might also be associated with undesirable `social' factors | high crime being an obvious example.

[24] TU is included interacted with time dummies, so that general unobserved fixed effects can also be included, but results are not changed when TU is included alone, nor when time effects are included separately.

regional and aggregate labour market conditions. TU effectively controls for trade unions' differing abilities to translate these elements of inside options into strategic advantage. The identity of the main trade union might also pick up: differences in density; differences in job type and product not already controlled for (note that UK unions tend to be organised along `craft', or occupational, lines); and differences in the contents of unions' utility functions (for example, some unions might care relatively more about employment vis à vis wage increases, and some might prefer lower intra-group differentials in return for a lower wage level). The identity of trade unions is found to have a significant effect on wages. Its inclusion tends to dominate the effect of what we have interpreted as outside variables, local labour market effects, and other measures of union bargaining power.

Column [5] examines whether the wage outcome is affected by the number of trade unions that bargain together on behalf of the bargaining group of workers that includes lowest-grade manual labourers. TUNO is the number of trade unions in the bargaining group. As reported in Table 4, we observe as many as nine unions bargaining together; all groups are represented by at least one union, as our sample is restricted to unionised groups. A group of unions bargaining jointly is sometimes given a name, such as the `Joint Consultative Committee'.[25]

Previous research (Machin et al. 1991) using 1984 UK WIRS2 data has found little difference in the wages paid in establishments where workers were represented by a large number of trade unions (single-table-multiple-union bargaining) and those where there was a single union.[26] Results here suggest differently. A larger number of unions

[25] It is possible that a larger number of unions is associated with greater diversity of occupations, as bargaining groups can contain different numbers of job category. (Because we operate at the level of the bargaining group, for which we take the wage of the lowest paid manual worker to be representative, all that we require is that the bargaining group contain lowest grade manual workers). This suggests two possibilities. First, that a larger number of unions could be associated with a higher average skill level in the bargaining group, which could imbue the bargaining group with greater bargaining power (for asset-specificity type reasons, for example). Second, that a larger number of unions could be associated with the size of the establishment or company. Size is generally acknowledged to be positively related to pay. We investigate the second issue further later. Our ability to investigate the first issue is limited by lack of data.

[26] According to Machin et al., what matters is the structure of bargains, namely whether there are one or many agreements within the establishment. The data here concern one bargaining group, and we have little information about whether the bargaining group covers all or only some workers in an establishment.

is associated with a higher wage level. This indicates that coordination problems (which could result from a higher number of trade unions and decrease the union side's e®ectiveness) do not arise, and indeed that unions draw greater bargaining strength from negotiating jointly. This is a striking result which could shed light on one of the ¯ndings of the 1990 WIRS | that \active inter-union competition [for membership] was not prevalent where multi-unionism was already established" (Millward et al. 1992, p. 85): only 10% of manual representatives in private manufacturing reported its occurrence. It would not be in their interests to do so, given that any reduction in the number of unions in a bargaining group seems to result in lower wages. Our results also illuminate the comment by Gregg and Yates (1991), based on a survey of companies, that given a general picture of unionism in decline, the lack of evidence of moves away from multi-unionism \implies multi-unionism is harder to remove" (p. 11).

As mentioned above, it is possible that the number of trade unions in a bargaining unit could be associated with the size of the unit, establishment or company. The 1990 WIRS showed that multi-unionism increases markedly with the size of the workplace (Millward et al. 1992, p. 81). If that were the case, the positive coe±cient on TUNO might re°ect the commonly-found `size e®ect' whereby wages are higher in larger organisations. But the signi¯cant positive e®ect on TUNO was robust to the inclusion of size as indicated by company employment.[27] And bargaining unit size is already controlled for to a large extent by the inclusion of individual ¯xed e®ects.

Finally, we turn to perhaps the most commonly-used measure of union bargaining power | union density (DENS). Unions with greater representation amongst the workforce appear able to extract higher wages (column [6]). Once again, density appears to dominate other measures of union bargaining power.[28]

5.2. Local labour markets

What is the e®ect of regional variables compared with their aggregate counterparts? From the point of view of estimation, regional variables might be preferred in bar-

[27] The coe±cient on company employment, lagged and instrumented, was negative but insigni¯cant. Full results are not reported here but are available from the author on request.

[28] The e®ects of trade union identity and number are robust whether included separately or together. The number of observations remaining when missing observations on TU, TUNO and DENS were simultaneously removed was considered too low to provide reliable results.

gaining unit level regressions, since their disaggregation might enable more of the variation to be explained. But it could be that regional e®ects are not important in determining bargaining unit level wages, in which case no signi¯cant di®erence would be expected between aggregate and regional formulations. The e®ect of replacing the regional unemployment rate, vacancy/unemployment ratio, wage and tax wedge with their aggregate counterparts is shown in column [7] of Table 2 (compare this with column [1] of Table 1). Inspection reveals that all coe±cients are less well determined in the aggregate formulation (although in combination with company-level regressors the aggregate variables seem to explain more of the variation in bargained wages: the $\bar{R}^2$ of the aggregate regression is higher). A comparison of the two formulations using non-nested tests proved inconclusive: the results (shown in column [7]) indicate that both regional and aggregate variables possess explanatory power for bargained wages.[29]

The additional e®ects of local labour market conditions can also be assessed by adding them to aggregate speci¯cations.[30] Including both sets of regressors could be meaningful if there is important co-movement which is captured in the aggregate labour market variables, in addition to regional di®erences. Column [8] includes both aggregate and regional variables: again, regional variation seems important, in that variables at that level of disaggregation are better-determined. The regional wage appears particularly in°uential, perhaps re°ecting low labour mobility.

A further means of testing the importance of regional e®ects is to include a regional dummy in the aggregate regression, as is reported in column [9]. As expected, regional e®ects are signi¯cant | the probability that the coe±cient on the regional dummy is zero is less than 1%. The structure of the chemicals industry means that we

[29] The tests used are J-tests (Davidson and MacKinnon, 1981): for dependent variable y and competing sets of regressors X and Z, y is regressed on X and ¯tted values obtained, then y is regressed on these ¯tted values and Z. If Z is the correct set of regressors, the coe±cient on the ¯tted values from the X-regression should be close to zero (a t-test is used to determine whether this is so). This procedure is then reversed; the set of regressors Z is preferred only if the results of the reverse procedure are consistent. The statistics reported in column [11] are t statistics (and their associated probabilities) for the ¯tted values from one formulation when included in the other, and hence provide an indication of additional variation explained. Unfortunately, non-nested tests often lack power to discriminate, as this case illustrates.

[30] As in Christo¯des and Oswald (1992), for example. Fewer observations are used for regression [8] because `national' bargaining units, whose `regional' regressors would be aggregate variables, are dropped. The same smaller sample was used for the J tests reported in column [7].

	[7]	[8]	[9]	[10]
(¼ $^{d}_{l}$ n)$_{jt¡1}$	0:0298 [2:22]	0:0311 [1:96]	0:0337 [2:48]	0:0443 [3:19]
(k $^{d}_{l}$ n)$_{jt}$	0:0486 [1:92]	0:0653 [1:85]	0:0474 [1:86]	0:0586 [2:77]
(stk^{d}¡ n)$_{jt¡1}$	¡ 0:0765 [¡ 2:22]	¡ 0:0880 [¡ 2:18]	¡ 0:0901 [¡ 2:57]	¡ 0:119 [¡ 4:12]
UR_t	0:478 [0:82]	0:359 [0:24]	0:377 [0:64]	
UR_{Rt}		¡ 0:121 [¡ 0:11]		0:410 [1:10]
$(V{=}U)_t$	0:149 [0:55]	¡ 0:295 [¡ 0:74]	0:202 [0:74]	
$(V{=}U)_{Rt}$		0:215 [1:07]		0:235 [1:26]
$UR52_t$	¡ 0:0617 [¡ 0:48]	¡ 0:136 [¡ 0:91]	¡ 0:0273 [¡ 0:21]	
w_t	0:424 [1:43]	0:475 [1:28]	0:399 [1:34]	
w_{Rt}		0:318 [3:29]		0:338 [4:88]
$wedge_t$	¡ 0:853 [¡ 1:21]	¡ 1:337 [¡ 1:53]	¡ 0:433 [¡ 0:60]	
$wedge_{Rt}$		1:524 [2:73]		¡ 0:649 [¡ 2:45]
REG ¤ t			0:0010 [2:71]	
constant				6:142 [8:89]
$\bar{R}^2$	0.912	0.894	0.911	0.898
F test	41:90 (119;349)	31:58 (109;286)	41:16 (120;348)	33:54 (127;341)
Log likelihood	802.5	659.7	801.0	773.9
Autocor. coef.	-0.009	-0.079	-0.023	-0.030
J (regional)	2:841 [0:005]			
J (aggregate)	5:017 [0:000]			
Time dummies	no	no	no	yes
Observations	469	396	469	469
Bargaining units	112	98	112	112

Table 2: Comparison of regional and aggregate e®ects

Notes: Wald (regional)/(aggregate): test of joint hypothesis of zero coe±cients on, respectively, regional or aggregate variables (aggregate variables excluding $UR52_t$) ($\hat{A}^2_1$). J (regional)/(aggregate): Davidson-MacKinnon (1981) non-nested test for superiority of, respectively, regional or aggregate speci¯cations. See also notes to Table 1.

can probably rule out the possibility that there are any speci¯cally regional variations in the product market(s), so the regional dummy is picking up di®erences in local labour market conditions (for example, imbalances in the distribution of vacancies and unemployment across the country), and also possibly di®erences in costs of living across regions (including regional price di®erentials, which are captured in the regional speci¯cation in the de°ator used for the sample average regional wage). Because of the nature of the data used, we can rule out the possibilities that regional di®erences re°ect uneven distribution of skill attributes of workers (the wage variable relates to jobs that require similar skills), and that they re°ect di®erent industry or technology mixes across the country (bargaining units come from a single industry). The importance of local conditions persists because of low migration between regions: there is convincing evidence from other sources that mobility is very low in the UK.[31] The ability to capture such e®ects is a major advantage of working with data at this disaggregated level. The counterpart to the inclusion of a regional dummy in the aggregate equation is the addition of time dummies to the regional equation (column [10]).[32] The addition of time e®ects leaves estimates largely unchanged (compare column [1] in Table 1), adding weight to the argument that regional factors are in°uential in wage determination.

5.3. Company performance

Company performance is measured by the natural log of pro¯t per employee.[33] It is sometimes maintained that \for companies with a pro¯t maximisation objective this is the primary ratio"[34] and this measure of performance can be shown to appear in the ¯rst order condition of the Nash bargaining problem.[35] Results consistently

[31]For example, in 1986, according to the OECD, only 1.1 per cent of the UK population changed its region of residence. And Hughes and McCormick (1987) estimated that US manual workers are ¯ve and a half times as likely to migrate between regions as UK manual workers.

[32]$UR52_t$ is dropped as it would be collinear with the time dummy.

[33]A logarithmic form was found to ¯t the data better than non-logs, although it necessitates dropping observations where pro¯t is negative.

[34]Norkett (1986), p. 94.

[35]See the model in the Appendix. Statistical tests indicate that a pro¯t measure of performance is preferred. Non-nested J-tests (Davidson and MacKinnon 1981) reject sales per employee in favour of pro¯t per employee: the ¯tted values from a regression including sales were insigni¯cantly di®erent from zero at the 5% con¯dence level when included in the regression with pro¯ts, whereas pro¯ts were signi¯cant at 5% in the sales regression. Results are available from the author on request.

show significant positive effects from profitability to wages, in accordance with the predictions of bargaining models (see Tables 1 and 2). Among other things, this suggests that the method adopted to control for the simultaneous determination of profit and wages, involving instrumenting the first lag of profit measures with its second and third lags, has been successful. If it had not been, we would have expected a negative relationship between profit and wages.

The elasticity of wages with respect to profit per employee of roughly 0.04 implies that a 10% increase in log profits per worker (a rise in profit from $8,979 to $22,312 per annum) would lead to a wage increase of $4.09 a week (4%) at the average-wage bargaining group, from $115.44 to $119.53.

Our estimate of the elasticity of wages with respect to profit per employee is similar to that found by Hildreth and Oswald (forthcoming) in their establishment panel, although they found the profitability effect poorly defined (coefficient 0.04, standard error 0.07). The similarity of estimates occurs despite a substantial difference in approach to endogeneity and in detail of specification of the model. In addition to a lagged dependent variable Hildreth and Oswald include the current value and up to three lags of profit or profit per employee (thus not allowing for the endogeneity of profit). Their specification excludes measures of bargaining power and local labour market influences which we find important in wage determination.

The estimates reported here are also similar to those found by Carruth and Oswald (1989) in their examination of wage determination in the UK at the aggregate level. They estimated the elasticity of average earnings with respect to aggregate profit per employee to be 0.05 in the long run on the basis of annual data. Beckerman and Jenkinson (1990) report a long-run elasticity of UK industry-level wages with respect to profit per employee of 0.044. Hildreth and Oswald (1994) estimated the long-run profit elasticity of pay to be 0.02 on the basis of a panel of company-level data. But profit elasticities reported elsewhere differ: some are smaller by a factor of up to ten. In a company-level study for the UK, Denny and Machin (1991) report elasticity estimates of 0.01 after adjustment for inclusion of an outside wage variable in addition to the lagged dependent variable, 0.004 in the short run, and 0.005 in the long run.

We conclude that company performance has a positive influence on the determination of basic wages. Workers are able to grab, or are given, a larger reward when rents are higher. This is consistent with `pure' bargaining models and some types of efficiency wage model — for example, where effort is related to the perceived fairness of rewards (see Akerlof 1982; Summers 1988).

6. Conclusions

In this paper the results of applying a bargaining model of wage setting to a unique panel of data covering bargaining units in a major UK industry over the last decade have been presented. A bargaining model appears to fit the data reasonably well. It is important that we have been able to confirm this on the basis of disaggregated data at a level that accords with the theory and that avoids data-related problems which might have affected previous studies.

Empirical findings concerning the signs of coefficients on explanatory variables are presented in Table 3 together with predictions derived from bargaining models. To summarise: certain factors which are interpreted as determinants of the alternative options and bargaining powers of the bargaining parties are found to be important in determining wages. As predicted by strategic bargaining models, the level of stocks held by the firm is an important component of their bargaining power: a higher level of stocks enables the firm to resist wage demands. Wages are higher where a greater proportion of the workers in the settlement group are union members. A greater number of unions bargaining jointly is able to push for higher wages, which might partly explain an apparent lack of inter-union competition for membership. The identity of the major trade union in the bargaining group also matters.

Regional factors are shown to be important in affecting wage levels, although apart from the local wage rate, regional labour market variables are found to have relatively little influence on wage determination. Results indicate that company performance has a significant positive impact on basic pay. This is in accordance with theoretically-derived wage equations which are log-linear approximations to first-order conditions of bargaining models, and with `rent-sharing' considerations.

Mnemonic	Definition	Predicted sign	Estimated sign
$DENS_i$	average union density	+	+
TU_i	identity of main trade union (dummy)	n/a	n/a
$TUNO_i$	number of trade unions (dummy)	− or +	+
Π/N_{jt}	pre-tax profit per employee	+	+
K/N_{jt}	capital/employment ratio	+	+
LIQ/N_{jt}	liquid assets per employee	− or +	+
STK/N_{jt}	total stocks per employee	−	−
FIN/N_{jt}	stocks of finished goods per employee	−	+
WIP/N_{jt}	stocks of work in progress per employee	−	+
RAW/N_{jt}	stocks of raw materials per employee	−	−
REG	region (dummy)	n/a	n/a
UR_{Rt}*	unemployment rate	−	+
V/U_{Rt}*	vacancies/unemployment	+	+
W_{Rt}*	sample average regional wage	+	+
$WEDGE_{Rt}$*	ratio of product to consumption wage	−	−
$UR52_t$	long-term/total unemployment	+	−

Table 3: Predicted and estimated signs of coefficients

Notes: i: bargaining unit; j: company; R: region; t: time. * Variables also examined at the aggregate level. Real variables expressed in constant January 1987 prices. Deflator for regional variables is the regional consumer price index excluding housing costs (source: Reward Group). Deflator for other variables is the retail price index excluding mortgage interest payments (source: CSO).

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Appendix: The model

We present here a simple theoretical framework that generates our estimable equation and clarifies the role of company performance and bargaining power in wage determination. Assume that the wage is determined by bargaining between a union which cares about the wage alone[36] and a firm which has the 'right to manage'. In other words, there is bargaining over wages in the knowledge that the firm will then choose employment to maximise profits. The standard symmetric Nash bargaining problem for this model (which it is possible to generalise to the asymmetric case) can be written:

$$\max_{W} (\Omega(W) - \bar{\Omega})(\Pi(W;N;P;A) - \bar{\Pi}); \tag{6.1}$$

where W is the wage, Π is profit, P is the firm's output price and A is a productivity shift parameter.[37] $\Pi(W;N;P;A)$ is a maximum profit function that accords with the usual assumptions[38] and Ω represents the union's utility function. The firm chooses the level of employment so as to maximise profit subject to the bargained wage:[39]

$$\Pi(W;N;P;A) = \max_{N} APf(N) - WN.$$

The first order condition for the Nash maximisation in (6.1) is:

$$\Omega_W (\Pi(W;N;P;A) - \bar{\Pi}) + (\Omega(W) - \bar{\Omega}) \Pi_W = 0. \tag{6.2}$$

The first order condition (6.2) can be rewritten as a function connecting wages and profits. By duality theory, employment is given by the derivative of the profit function with respect to the wage: $N = -\Pi_W$. Substituting this into (6.2):

$$\Omega_W (\Pi(W;N;P;A) - \bar{\Pi}) - (\Omega(W) - \bar{\Omega}) N = 0.$$

Dividing by N and rearranging gives:

[36] This is known as an 'insider-dominated' union, and can arise under last-in-first-out firing rules, for example. See, amongst others, Carruth and Oswald (1989).

[37] A semi-colon indicates a conditional. Variables to its right are taken as fixed (exogenous).

[38] In other words the function is homogeneous of degree one in $(W;P)$, convex, decreasing in W and increasing in P, and twice differentiable.

[39] An obvious consequence of the firm choosing the level of employment subject to the bargained wage is that the wage-employment combination will always lie on the firm's labour demand curve.

Mnemonic	Mean	Std. Dev.	Min.	Max.	Cases
W_{it}	5,216.12	1,710.28	2,290.08	11,986.00	1,325
$DENS_i$	88.2%	13.8%	32.9%	100.0%	175
TU_i	10.3	3.2	1	17	2,383
$TUNO_i$	2.5	2.0	1	9	1,454
FIN_{jt}*	109,190	217,210	104	1,434,000	2,262
K_{jt}*	931,380	2,777,200	6	33,860,000	3,021
LIQ_{jt}*	477,520	1,174,100	14	14,740,000	3,169
N_{jt}	16,846	28,547	3	219,000	2,818
π_{jt}*	291.2	912.6	-683	10,620	2,869
RAW_{jt}*	54,517	111,100	3	669,000	2,332
STK_{jt}*	317,330	732,410	13	7,330,000	3,080
WIP_{jt}*	40,704	86,848	7,000	826,900	2,034
REG	5.88	3.31	1	11	5,136
UR_{Rt}	8.41%	3.64%	2.64%	15.40%	4,648
$(V/U)_{Rt}$	7.13%	4.29%	2.33%	23.6%	3,652
W_{Rt}	6,123.70	721.29	4,593.00	8,973.00	4,617
$wedge_{Rt}$	0.4333	0.0260	0.3514	0.4833	1,325
UR_t	7.56%	2.58%	4.05%	11.10%	5,312
$UR52_t$	37.52%	5.94%	27.33%	45.57%	4,316
V/U_t	7.37%	2.78%	4.16%	12.09%	3,652
W_t	6,107.00	469.06	5,487.00	6,943.00	5312
$wedge_t$	0.4425	0.0292	0.3715	0.4836	1,325

Table 4: Summary statistics from the full sample

Notes: For mnemonic definitions see Table 3. *: values expressed in $'000.

$$\Omega(W) = \bar{\omega} + \omega_W \frac{(\pi(W; N; P; A) - \bar{\pi})}{N}: \qquad (6.3)$$

There is a wage equation implicit in (6.3), in which the wage is affected by (factors which determine) the union's fall back option, (factors which determine) the firm's fall back option (divided by employment), and profit per employee. In the asymmetric-bargaining-power case, the wage will also be a function of (factors which affect) the firm's and the union's bargaining powers. Thus we have equation (4.1):

$$W = W\left(\bar{\omega}; \frac{\bar{\pi}}{N}; \frac{\pi}{N}; \theta\right);$$

where θ is the ratio of the union's bargaining power to that of the firm.