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FEAR, UNEMPLOYMENT AND PAY FLEXIBILITY*

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The labour market plays a central role in macroeconomic analysis, and the theory of wage determination plays a central role in models of the labour market. Despite this, the behaviour of pay is not something economists can claim to understand fully.

Research in the field is divided across three different avenues. First, since the turn of the decade there has been much work on the theory of wage formation. Although multifarious in its approaches, this literature has been particularly concerned with the issue of why wages fail to clear the labour market. However, little of the theoretical work has been combined with a strongly empirical component.[1] Thus labour contract theory, trade union models, efficiency wage theories, search models and insider-outsider analysis *inter alia* have not been extensively tested. One possible defence is that the theories are too new to have been subjected to serious scrutiny by applied economists. Yet many of the ideas have been around rather longer than is always apparent ('the distinction between insiders and outsiders in wage discussions is as old as the hills (Dunlop 1944, chapter III)' – John Dunlop (1988), p. 69) and the nature of the argument about wage inflexibility is unchanged from Pigou's time.

The second form of research on wages is of a different character and has almost the opposite attributes from the first form. Cross-section research on pay, using micro-econometric data sets, and stemming largely from Mincer (1962), 'has been one of the great success stories of modern labour economics. It has been used in hundreds of studies using data from virtually every historical period and country...', Willis (1986), p. 526. The strength of this empirical research is not the coherence of its theoretical underpinnings – though Mincer (1962), Becker (1967) and others have suggested some – but the fact that its chief findings have been replicated countless times. The adherents to cross-section research stress its scientific credentials (see Freeman and Medoff (1984), for example). What is less commonly noted, however, is that most of the literature is not designed to answer the major questions which concern theorists and policy-makers. Relatively little progress has been made on the issue of how labour markets work and why wages do not seem to clear Western labour markets. Instead the focus has been on empirically valuable but conceptually narrow matters of economic measurement (How much do

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[1] This is least true for trade union analysis (see the empirical work in Pencavel (1985), for example), and Krueger and Summers (1988) recently consider empirical evidence consistent with efficiency wage models. This seems to be reasonable as a generalisation.

unions raise wages? What is the size of the return to education? How large are gender differentials?).

A third form of inquiry was begun by the work of Phillips (1958) and his contemporaries. It tackles the analysis of pay by estimating time series wage equations using aggregate data; its modern equivalent appears in papers such as Layard and Nickell (1986). Unlike the cross-section literature, the focus of this current of research has been on macroeconomic questions and the construction of empirically reliable models of the market for labour. A principal concern has been the role of excess supply in shaping rates of pay. The research findings remain almost as controversial as in Phillips' time, however, and there are some who see inferences based on small time series data sets as fragile.

The object of this paper is to blend aspects of these three – conventionally distinct – approaches. It takes microeconomic data on individuals for the United Kingdom between 1983 and 1989 and augments a Mincerian cross-section wage equation by adding a range of variables related to the extent of excess supply in the labour market. These variables are suggested by theoretical analysis developed in Section I (which is developed more formally in Appendix A) as well as in the earlier literature.

I. WAGES AND UNEMPLOYMENT

If a firm sells in a risky product market its employees will typically also face some risk: they may be made redundant if demand conditions are poor. This *fear* of unemployment will have an effect in the competitive model which is different from that in a model in which employees earn non-competitive rents. Under perfect competition in the labour market the risk of lay-off generates higher wages. Fear of unemployment has to be compensated, like any other disutility, by greater remuneration. By contrast, when wages are above their reservation level, and determined as if in some form of bilateral bargain, the risk of being fired, because, say, the workplace is going to close, will typically generate lower wages. Employees who earn rents will wish to retain their jobs: the threat of loosing their jobs may therefore induce workers to forego a portion of those rents.

The paper bears a close relationship to recent attempts (Blanchflower and Oswald, 1990*a*, Blanchflower *et al.*, 1990, Nickell and Wadhwani, 1990, Beckerman and Jenkinson, 1988) to test for the relative strength of 'insider' and 'outsider' forces in pay determination. These use, respectively, data on establishments, firms and industries in Great Britain. The present paper explores similar issues with data on individuals.

The paper also tests for the existence of a wage 'ratchet'. A number of authors have recently suggested models – particularly of trade union behaviour – in which there is such a phenomenon. The literature includes Blanchard and Summers (1986), Carruth and Oswald (1987*a*), and Lindbeck and Snower (1986, 1988). Although these accounts of the wage ratchet differ, they share a common principle. An expansion in demand leads to a larger pay increase than a decline creates a decrease. There is therefore an asymmetry in wage determination[2].

The macroeconomic implications of such behaviour are potentially of importance. Booms and slumps, one after another, may lead to a net contractionary movement in aggregate employment. This effect is likely to be most marked in, for example, the union sector. The theoretical model of Appendix A extends this idea and generates the empirical prediction that wages should be a declining function of the probability of firm closure. This is true whatever is the relative bargaining strength of the workers. The analysis can be seen as a formalisation of the long standing idea of 'concession bargaining'.[3] Fear of redundancy can reduce workers' wage demands. Although the model relies on the extreme assumption of closure in the poor state of nature, the ideas apply more generally to layoffs.

The later analysis allows the estimation of the unemployment elasticity of wages (Blanchflower *et al.* (1990) surveys the literature). It does so by incorporating, at the regional level, the unemployment rate within a cross-section equation on individual workers' rates of pay. The only other British research on individual data of which I am aware that calculates this elasticity is Blackaby and Manning (1987, 1990) and Symons and Walker (1988). Studies for the United States include Bils (1985), and Adams (1985).

II. EMPIRICAL IMPLEMENTATION

The series of surveys used in this paper – the British Social Attitudes Surveys (BSA) are unusual in that they contain information on workers' past experience of unemployment and their perceptions of the chance of losing their job. It is possible to distinguish four 'unemployment' variables that can be used in our estimation. Table 1 provides the details. First, workers were asked about their experiences of unemployment. Between 18% and 23% of workers said they had experienced an unemployment spell over the preceding five years. Although various interpretations are possible it is conceivable that this measures an otherwise unobservable level of worker quality.

Second, workers were asked whether they would leave their employer in the following year. If the respondent answered either 'very likely' or 'quite unlikely', they were asked 'why do you think you will leave?'. Answers to these questions are provided in Table 1. Depending upon the year, between 19% and 26% of workers were in the former two categories. In 1986, for example, approximately one in ten of them said that this was because their plant would close and a further two in ten because of redundancy.

Third, workers were also asked whether, over the following year, they expected that their workplace would increase or decrease in size. The proportion of individuals reporting that they expected that their workplace

[2] I am grateful to a referee for pointing out that 'if there is a cyclically repeating stationary economy with no productivity growth, aggregate employment should tend to zero according to this argument. In a macroeconomic framework, however, this is unlikely to persist indefinitely since the structure of the economy and the reduced form nature of the earnings equations estimated here will all change over time'.

[3] Shultz and Myers (1950) were among the first to document the phenomenon.

Table 1

Workers' attitudes and unemployment experience (%)

	1983	1984	1985	1986	1987	1989
(i) Unemployed last 5 years?	18	23	22	20	22	21
(ii) Leave employer?						
(*a*) very likely	9	13	11	10	n.a.	11
(*b*) quite likely	10	13	•12	11	n.a.	13
(*c*) not very likely	26	26	31	30	n.a.	28
(*d*) not at all likely	55	48	46	48	n.a.	48
(iii) Why leave employer?						
(*a*) Firm close down	2	2	1	2	n.a.	1
(*b*) Declared redundant	4	5	4	4	n.a.	3
(*c*) Reach retirement age	2	2	1	1	n.a.	1
(*d*) Contract end	1	1	1	1	n.a.	1
(*e*) Early retirement	1	2	1	1	n.a.	1
(*f*) New employer	7	12	11	10	n.a.	14
(*g*) Other	3	5	5	6	n.a.	4
(iv) Workforce size						
(*a*) Increase	16	18	22	21	24	27
(*b*) fall	30	30	25	24	23	20
(*c*) constant	55	52	53	55	54	53
Number of employees	817	778	968	1,532	1,381	1,353

Notes: all data reported in this Table are weighted.

would increase in size grew over the period from 16% to 27%, as UK's economic climate improved. Such workers presumably assign a low value to their chance of redundancy.[4]

Finally, it is possible to identify the region or state in which each individual lives. A natural measure of excess labour supply is the rate of unemployment in that local area. Within a bargaining framework the outside unemployment rate is likely to work through its effects upon alternative wage rates. It may also provide information about employees' long-term probability of joblessness.

The model estimated here is

$$\ln w = \beta_0 + \boldsymbol{\beta}_1 \mathbf{u} + \boldsymbol{\beta}_2 \mathbf{x} + \epsilon \tag{1}$$

which w is annual earnings, β_0 is a constant, $\boldsymbol{\beta}_1$ is a vector of parameters, $\mathbf{u}$ is the set of unemployment variables discussed above, $\boldsymbol{\beta}_2$ is a vector of parameters, $\mathbf{x}$ is a set of conventional control variables and ϵ is an error term. The $\mathbf{x}$ variables, which are familiar from other cross-section work, include the following.

(i) Human capital variables (experience and its square, years of schooling).

(ii) Personal characteristics (gender, marital status, union status, among others).

(iii) Workplace characteristics (industrial activity and union status).

(iv) A range of dummy variables (for industries, regions and years).

[4] Evidence consistent with this is provided in Oswald (1989). It is also worth noting that, of all those in the sample who thought that they would become redundant, 80% were in workplaces where employment was predicted to decline.

A description of these variables and of the data set is contained in Appendix B.

The paper pools five of the six BSA cross-sections from 1983–1989. Unfortunately, the 1987 survey dropped the question about redundancy; there was not a survey in 1988. Hence the next section presents results only on 1983 to 1986 and 1989. For the years 1983–1986 the data covers Great Britain only: in 1989 the survey also included Northern Ireland (305 cases).

III. THE RESULTS

Aggregate results for the United Kingdom are given in the first column of Table 2. These use data on approximately 5,300 employees and produce an adjusted R^2 of approximately 0·70. The dependent variable is the natural logarithm of annual earnings. The earnings are grouped and open-ended. In order to put these data into tractable form the standard practice of allocating midpoints to all of the bands was followed. A series of sensitivity tests were undertaken which showed that the results were relatively stable to changes in values allotted to the open-end categories. This is to be expected given the small number of observations in these end groupings.

To assess the overall specification of the equations a number of statistical tests were conducted which were designed to test for heteroscedasticity in the error variance, appropriateness of functional form and normality in the residuals. First, the standard errors in Table 2 are adjusted for heteroscedasticity using the White (1980) method. Second, a series of F-tests were conducted to test the pooling restrictions across the five years. For 1983, 1986 and 1989 one cannot reject the null hypothesis of equality of coefficients.[5] However, the F-tests for 1984 and 1985 were significantly different from zero. Experimentation with a series of year interaction variables produced remarkably similar results on the main variables of interest to those reported here. Third, the Lilliefors test statistic[6] provides some evidence of non-normality in the residuals. However, when sample sizes are relatively large, detecting mispecification of this type is almost inevitable (see Baker *et al.*, 1989). Examination of the residuals, however, suggested that, in every case, departure from normality was small.

All four of the unemployment variables are statistically significant in column 1 of Table 2. Regional unemployment (across ten regions and entered as a log of the percentage) is significant even controlling for a full set of regional fixed effects and has an elasticity of approximately −0·10. This is similar to the aggregate time series results of Layard and Nickell (1986) and Carruth and Oswald (1987*b*), the panel data findings of Bils (1987), Nickell and Wadhwani (1990) and Christofides and Oswald (1989), and cross-section estimates by Blackaby and Manning (1987, 1990). The estimate is also close to those derived

[5] The results of the individual F-tests were as follows; 1983, 1·18; 1984, 1·46; 1985, 1·44; 1986, 1·24; 1989, 1·16. The critical value of the F-statistic with 70 degrees of freedom in the numerator and 5200 in the denominator, at the 5% significance level is 1·30.

[6] For details see Conover (1980) pp. 357–61. The critical value for rejection of the test at the 5% significance level is 0·015. It should be noted that this procedure should be interpreted cautiously as it is based on the unproved conjecture that the statistic approaches its limiting asymptotic distribution as a function of the square root of the sample size.

Table 2

Annual Earnings Equations, 1983–86 and 1989

	(1) All workers	(2) All workers	(3) Union Plant	(4) Non-union plant	(5) Union plant	(6) Non-union plant
Unemployment Variables						
Unemployed last 5 years	−0·0989	−0·1004	−0·1218	−0·0780	−0·1216	−0·0780
	(6·38)	(6·47)	(5·72)	(3·54)	(5·72)	(3·61)
Employment rise	0·816	0·0844	0·0907	0·0741	0·0912	0·0741
	(5·54)	(5·75)	(4·73)	(3·34)	(4·82)	(3·34)
Redundancy expected	−0·0832	—	—	—	—	—
	(3·12)					
Plant closure expected	—	−0·0842	−0·0076	−0·1890	−0·0095	−0·1874
		(1·86)	(0·18)	(2·37)	(0·19)	(2·35)
Regional unemployment rate	−0·1021	−0·1025	−0·0208	−0·1613	0·0082	−0·2004
	(4·22)	(4·23)	(0·68)	(5·49)	(0·18)	(2·29)
Regional wage rate	—	—	—	—	0·1259	−0·1822
			(0·89)	(0·47)		
Other variables						
Experience	0·0258	0·0257	0·0221	0·0288	0·0221	0·0288
	(11·50)	(11·43)	(7·71)	(8·35)	(7·67)	(8·32)
Experience^2	−0·0004	−0·0004	−0·0004	−0·0005	−0·0004	−0·0005
	(10·22)	(10·16)	(7·18)	(7·33)	(7·18)	(7·33)
Schooling	0·0734	0·0738	0·790	0·0663	0·0790	0·0665
	(11·17)	(11·24)	(9·82)	(6·37)	(9·78)	(6·40)
Male	0·4456	0·4455	0·4433	0·4479	0·4432	0·4482
	(30·21)	(30·18)	(25·39)	(18·07)	(25·48)	(18·07)
Married	0·1068	0·1075	0·1052	0·1101	0·1059	0·1103
	(6·00)	(6·04)	(4·90)	(3·83)	(4·93)	(3·83)
Separated	0·1708	0·1706	0·1758	0·1697	0·1765	0·1700
	(5·91)	(5·89)	(4·99)	(3·64)	(5·01)	(3·64)
Widowed	0·2346	0·2326	0·1733	0·3122	0·1750	0·3121
	(4·64)	(4·62)	(2·73)	(3·80)	(2·75)	(3·79)
Supervisor	0·1754	0·1760	0·1709	0·1886	0·1708	0·1883
	(13·79)	(13·83)	(11·19)	(8·73)	(11·18)	(8·71)
Part-time	−0·8730	−0·8729	−0·8263	−0·9060	−0·8265	−0·9059
	(36·89)	(36·85)	(27·78)	(24·31)	(27·76)	(24·31)
Nonmanual	0·2617	0·2632	0·2879	0·2349	0·2875	0·2350
	(16·75)	(16·80)	(14·48)	(9·32)	(14·43)	(9·32)
Union member	0·0949	0·0928	0·0811	0·1127	0·0817	0·1134
	(6·12)	(5·97)	(4·61)	(3·23)	(4·64)	(3·24)
Union recognition	0·0413	0·0406	—	—	—	—
	(2·44)	(2·40)				
Industry dummies	60	60	60	60	60	60
Year dummies	4	4	4	4	4	4
Regional dummies	11	11	2	8	2	8
Constant	7·1941	7·1872	7·0462	7·4026	6·3318	8·4335
	(61·60)	(61·43)	(36·59)	(44·99)	(7·53)	(3·86)
Adjusted R^2	0·7046	0·7043	0·6969	0·7101	0·6969	0·7102
Degrees of freedom	5270	5270	2993	2032	2992	2202
F	143·08	142·84	91·51	66·940	90·361	67·755
Lilliefors test statistic	0·0329	0·0313	0·0324	0·0280	0·0320	0·0294

Note: Standard errors, in parenthesis, are adjusted for heteroscedasticity using the White (1980) method.

in earlier work (Blanchflower and Oswald (1990*b*) and Blanchflower *et al.* (1990)).

Individuals with a history of unemployment earn less, *ceteris paribus*. A spell of unemployment in the previous five years lowers pay, according to Table 2, by approximately 10%. A similar result has been observed by Chowdhury and Nickell (1985) using US panel data and Nickell (1982) using British panel data. This variable is presumably a proxy for poorer quality workers.

The probability of job loss appears to have a powerful effect upon earnings. Workers who stated that they expected to be made redundant did not receive a compensating differential but were paid, on average, approximately 8% less, *ceteris paribus*. This ties in with much industrial relations evidence on concession bargaining (Cappelli, 1985 describes the literature and presents modern results for the United States), but appears to be the first estimate based on individual microeconomic data for the United Kingdom.

One possibility is that bad workers have a relatively high fear of redundancy because of their poor performance. However, this paper argues that fear of unemployment itself, and not poor worker quality, is the explanation for the significant coefficient on the redundancy dummy. One possible way around this problem is to exploit the fact that when plants close both good and bad workers lose their jobs. Thus, as a check the 'Redundancy expected' variable for the United Kingdom was replaced with one relating to the expectation of plant closure. As can be seen from column 2 of Table 2, fear of plant closure *lowers* pay by 8%, *ceteris paribus*. This seems to support the idea that fear of unemployment is not primarily a proxy for worker quality.

Workers who reported that they expected employment to grow at their workplace received a wage premium of approximately 8%. (A dummy variable was also included where workers expected employment to decline at their workplace, but this was always insignificant and was excluded.) This may be an example of what Solow (1985) has described as 'the willingness and ability of insiders to convert higher demand into higher wages for themselves', p. 285. A closely related interpretation follows the theory set out in Appendix A. If, as seems plausible, workers feel secure when their workplace is growing, they may feel able – and be able – to extract higher remuneration from their employer.

When taken together these findings suggest that unemployment works – through a variety of channels – to depress wages. There is no evidence for the competitive model's prediction that fear of unemployment produces a compensating wage premium. The reverse appears to be true. The results are consistent with the idea that pay is fixed in a bilateral bargain where unemployment acts to weaken workers' bargaining position.

Post-war US economists such as Lester (1952) and Slichter (1950) believed in a band of wages within which employers had to pay. They argued that those with the highest ability to pay set the top of this range, whilst those close to bankruptcy tend to fix pay at the bottom of the range.[7] Blanchflower *et al.*

[7] Further evidence and discussion is provided by Dickens and Katz (1987), Krueger and Summers (1987) and MacKay *et al.* (1971).

(1990), using 1984 data on British establishments, estimate the range at between 8% and 22% of the wage. Similar findings from individual data emerge from Table 2. *Ceteris paribus*, the spread of wages from the top ('employment rise expected') to the bottom ('plant closure expected') is approximately 21% of average income.

Table 2 reveals evidence of an asymmetry in UK wage behaviour. Workplaces where employment was expected to rise paid a significant wage premium; those facing a decline in employment did not set lower pay. This is consistent with the prediction of the literature cited earlier of the existence of a wage ratchet. Employers facing alternate booms and slumps might, according to these results, progressively raise their wage rates. The finding is compatible with the fairly common but unproven idea that wages are flexible upwards but sticky downwards.

Columns 3 and 4 of Table 2 test the empirical prediction of the earlier theoretical model in the case where workers have different degrees of bargaining power.[8] The coefficient of the closure expected variable is significantly negative in the non-union sector but insignificant in the union sector (t-statistic on the difference = 2·01). Moreover, the unemployment elasticity of pay is strikingly different across the sectors. In the union part of the economy it is −0·02, but not significantly different from zero at normal confidence levels. By contrast, the coefficient in the non-union sector is −0·161. (t-statistic on the difference between the two coefficients is 3·31). Furthermore, the unemployment rate in the non-union equation is robust to the inclusion of both regional dummies and the regional wage (column 6) whereas it has no role in the union equation (column 5). Despite the fact that the coefficients on the 'Employment rise' and 'Unemployed last five years' variables are not significantly different from each other in the two sectors (t-statistics on the differences are 0·57 and 1·43 respectively) pay in the United Kingdom does seem to be more responsive to outside unemployment pressure in workplaces without trade unions.[9]

Table 2 also produces estimates of the influence of human capital and workplace variables. The conventional hump-shaped earnings/experience structure is confirmed. The earnings profile maximises after 29 years of experience (31 years in the union sector and 27 in the non-union). The Mincerian schooling variables are highly significant and imply that the rate of return to schooling is approximately 7%. Being a white-collar worker or a supervisor raises pay by approximately 25% and 18% respectively.

Union membership, according to the results in Table 2, leads to a wage premium of approximately 10%. This is only slightly above the existing estimates in Blanchflower (1984), Stewart (1983) and Shah (1984), *inter alia*. Interestingly, union membership conveys a significant differential in *both* the union and the non-union sectors (Table 2 columns 3–6). In the former case this

[8] On the basis of a series of t-tests the number of regional dummies in the union sector in Table 2 was reduced to two (South West and London). In the case of the non-union sector to it was reduced to eight (including all areas except Scotland, Wales and Northern Ireland).

[9] Layard and Nickell (1987) argue theoretically that a 'key variable will be "fear" – the fear of job loss.' Our results support this hypothesis.

is presumably identifying the influence of the closed shop (see Blanchflower *et al.* 1990 for a discussion). Union membership appears to convey a significant wage premium even where there are no recognised trade unions. Disaggregated estimates of the union/non-union wage differential or wage gap for the 1980s are reported in Table 3. Rather surprisingly, there is no evidence of a

Table 3
Disaggregated Union Wage Gap Estimates (%)

Group	Estimate	Group	Estimate
All	10	Age < 25 yrs	6
1983	11	Age 25–49 yrs	10
1984	11	Age ⩾ 50 yrs	10
1985	1*	Experience 0–10 yrs	3*
1986	13	Experience 10–29 yrs	10
1987	9	Experience ⩾ 30 yrs	9
1989	11	Manufacturing	8
Male	1*	Services	10
Female	19	Private sector	8
< 25 workers	16	Public sector	11
25–99 workers	10	< 10% unemployment	8
100–499 workers	7	⩾ 10% unemployment	13
⩾ 500 workers	1*	North	11
Manual	12	South†	7
Non-manual	6	No qualifications	10
Part-time	24	CSE	9
Full-time	7	O/A-levels	8
		Degree/higher degree	5*

Notes: Union wage gap estimates obtained from running separate regressions for the indicated group and calculating the natural anti-logarithms of the coefficient on the union membership variable and deducting 1.

Workplace size data missing in the 1983 survey, whilst highest qualification data are missing in the 1983 and 1984 surveys.

* Insignificantly different from zero at the 1% level.

† 'South' = South East including London, the South West and E. Anglia.

significant differential for males; this contrasts with the very substantial estimate (19%) obtained for females. The differential is highest for manual workers, part-timers, those living in the North and/or in high unemployment areas and those working in small plants. In all years except 1985 the differential is approximately 10%. The 1985 result is clearly a puzzle for which I have no explanation.[10]

IV. CONCLUSIONS

This paper studies pay determination in the United Kingdom in the 1980s. It is based upon a series of surveys which provide psychological data on variables such as perceived chance of redundancy. The object of the inquiry is to use cross-section methods to address issues traditionally tackled with small time-series data sets. The paper suggests that risk of plant closure can be expected

[10] Having done various checks, I can only conclude that there was some problem with the 1985 data collection.

to enter negatively in a microeconomic wage bargaining equation. The main empirical results can be summarised as follows.

1. Fear of unemployment appears to depress pay substantially. Workers who expect to be made redundant earn 9% less, *ceteris paribus*.[11] Workers in non-union workplaces who say they expect their plant to close earn 19% less than those who do not. No evidence could be found for such an effect in the union sector.

2. There is some evidence of an asymmetry or 'wage ratchet' in the United Kingdom. Workers in expanding plants receive a pay premium; those in contracting plants suffer no pay disadvantage. This is consistent with the claim that wages are more flexible upwards than downwards.

3. Unemployment in the individual's region depresses pay with an average elasticity of $-0{\cdot}1$.

4. Pay is more responsive to changes in the unemployment rate in the non-union sector than it is in the union sector.

5. A history of personal unemployment depresses pay 10% on average. Being a supervisor raises pay by approximately 18% on average.

6. The union wage gap (or mark-up) in the United Kingdom in the 1980s was approximately 10%. It was highest for women, part-timers, those who lived in the North, in high unemployment areas and worked in small plants. The differential appears to have stayed broadly constant in the 1980.

There is no indication, from these equations or the theoretical model, that fear of unemployment is compensated by higher pay. It seems more appropriate to see unemployment as a force which acts to weaken workers' negotiating power. This emasculation is clearest in the United Kingdom's non-union sector.

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Appendix A. Fear of Redundancy: Theory

Consider the wage bargain between a profit maximising firm and its employees. Assume for simplicity that there are only two possible states of nature. In the boom state, the selling price of the firm's product price is unity, output is $f(n)$, employment is n, the wage is w and fixed costs are k. In the other state of nature, the slump, the price of the product is zero, so the firm closes down and pays only its fixed costs. The slump occurs with probability θ.

Assume that workers' preferences can be represented by an expected utility function. In the boom state the representative worker, who may be the one with median seniority, has utility $u(w)$, (Oswald (1987) explores a model related to the one developed here). In the slump utility is $u(b)$, where b is the level of unemployment benefit or an equivalent income level in another job.

The two sides are assumed to act as if solving an asymmetric Nash (1953)

[11] This contrasts with some recent findings by Holzer and Montgomery (1990). They estimated a wage *growth* equation using firm level panel data for the US between 1980 and 1982 and found that non-union wages are sticky downwards but flexible upwards. In contrast, union wages were found to be sticky in both directions.

bargaining problem. This assumption can be justified axiomatically as in Nash (1953) or strategically as in Binmore *et al.* (1987). Workers' relative bargaining strength is denoted $s/(1-s)$. Each party has an outside option: it can withdraw entirely. The representative worker receives wage b in that case. If the firm withdraws (perhaps by moving its operations elsewhere) it can produce at wage w^b.

The Nash bargain can be represented by the following problem:

$$\underset{w}{\text{Maximise}}\; N \equiv s \log[(1-\theta)\, u(w) + \theta u(b) - u^*] + (1-s) \log[(1-\theta)\, \pi(w,k) - \theta k - \pi^*] \quad (1)$$

subject to

$$w - b \geqslant 0 \quad (2a)$$

$$\pi(w,k) - \pi(w^b,k) \geqslant 0. \quad (2b)$$

The delay or strike utility of the worker is u^*, whilst the equivalent profit level is π^*. It is assumed that $u(b) > u^*$ (strikers are unable to draw benefits b). The above formulation uses the maximum profit function

$$\pi(w,k) = \max_n f(n) - wn - k, \quad (3)$$

which is decreasing and convex in the wage rate.

The closed interval $[b, w^b]$ provides a formalisation of Lester's (1952) feasible 'range' of wages. Where within this interval the firm will set pay depends upon demand, production and utility parameters.

Defining multipliers Φ and μ for the two constraints, the Lagrangean may be written

$$L = N(s, w, b, u^*, \theta, k, \pi^*) + \Phi(w-b) + \mu[\pi(w,k) - \pi(w^b,k)] \quad (4)$$

which is assumed appropriately differentiable and concave. At the maximum,

$$N_w + \Phi + \mu\pi_w(w,k) = 0. \quad (5)$$

This defines a wage bargaining function

$$w = w(s, b, u^*, \theta, k, \pi^*, w^b). \quad (6)$$

The sign of the derivative of the wage function with respect to θ is of particular importance. This shut-down probability affects the worker's expected utility and the firm's expected profit. A rise in θ makes both parties worse off, and in general changes their relative utility.

In the interval (b, w^b) the two inequality constraints are not binding. Then it follows from normal methods that

$$\operatorname{sign}\frac{\partial w}{\partial \theta} = \operatorname{sign} N_{w\theta}, \quad (7)$$

which relies on the second-order condition $N_{ww} < 0$.

The relevant first-order condition and cross-partial derivative are therefore

$$N_w = \frac{su'(w)}{(1-\theta)\, u(w) + \theta u(b) - u^*} + \frac{(1-s)\, \pi_w}{(1-\theta)\, \pi(w,k) - \theta k - \pi^*} = 0, \quad (8)$$

and

$$\operatorname{sign} N_{w\theta} = \operatorname{sign}\frac{-su'(w)\,[u(b)-u(w)]}{[(1-\theta)\, u(w) + \theta u(b) - u^*]^2} + \frac{(1-s)\, \pi_w(\pi + k)}{[(1-\theta)\, \pi(w,k) - \theta k - \pi^*]^2}, \quad (9)$$

By deducting equation (8), which is equal to zero from the first order conditions, from $(1-\theta)$ multiplied by equation (9), ($(1-\theta)$ is unambiguously non-negative), the following expression is obtained by manipulation. (I am grateful to Donald Deere for this insight.)

$$\frac{-su'(w)}{[(1-\theta)\,u(w)+\theta u(b)-u^*)]^2}[u(b)-u^*] - \frac{(1-s)\,\pi_w}{[(1-\theta)\,\pi(w,k)-\theta k-\pi^*]^2}(-k-\pi^*) \Rightarrow \frac{dw}{d\theta} < 0. \quad (10)$$

Both of the two terms are unambiguously negative. The first term is negative as $su'(w)$ and $[u(b)-u^*]$ are both positive. The second term is also negative as π_w and $(-k-\pi^*)$ must also be negative. Therefore $dw/d\theta$ is less than zero.

Appendix B. British Social Attitudes Surveys, 1983–9

This series of surveys, core-funded by the Sainsbury Family Trusts, was designed to chart movements in a wide range of social attitudes in Britain. The data were collected by Social and Community Planning (SCPR) and derive from annual cross-sectional surveys from a representative sample of adults aged 18 or over living in private households in Great Britain whose addresses were on the electoral register. The first three surveys involved around 1800 adults; the numbers were increased to 3000 in 1986. The sampling in each year involved a stratified multi-stage design with four separate stages of selection. For further details of the survey designs, non-responses etc. see *British Social Attitudes*, 1983, 1984, 1985, 1986, 1987, 1989, edited by R. Jowell, S. Witherspoon and L. Brook, SCPR, Gower Press.

Variable	Mean	S.D.	Definitions
Male	0·545	0·498	A (1,0) dummy variable for gender
Experience	22·07	12·992	(Age − schooling) + 5
Part-time	0·180	0·384	A (1,0) dummy variable if the respondent reported that they normally worked less than 30 hours per week
Separated	0·047	0·213	A (1,0) dummy variable if the dependent was separated or divorced
Widow	0·020	0·139	A (1,0) dummy variable if the respondent was widowed
Married	0·727	0·446	A (1,0) dummy variable if the respondent was married or living as married
School	11·226	1·487	Number of years of schooling
Union recognition	0·469	0·499	A (1,0) dummy variable if trade unions or staff associations at the place of work are recognised by management for negotiating pay and conditions
Employment rise	0·215	0·411	A (1,0) dummy variable if the respondent expected their workplace to increase its number of employees over the coming year
Supervisor	0·362	0·481	A (1,0) dummy variable if the respondent was a supervisor

Variable	Mean	S.D.	Definitions
Union member	0·464	0·499	A (1, 0) dummy variable if the respondent was a member of a trade union or a staff association
Non-manual	0·548	0·498	A (1, 0) dummy variable if the respondent's occupation was non-manual
Unemployed last 5 years	0·208	0·406	A (1, 0) dummy variable if the respondent reported that they had ever been unemployed and seeking work in the preceding five years
Redundancy expected	0·050	0·219	A (1, 0) dummy variable if the respondent expected that during the next year they would lose their job because of firm closure or being fired
Plant closure expected	0·011	0·104	A (1, 0) dummy variable if the respondent expected that during the next year they would lose their job because of firm closure or being fired
Regional unemployment	2·272	0·395	Unemployment rate in the Standard Region, entered in natural logarithms. Source: *Employment Gazette*, various issues
Year 84/5/6/7/9			5 (1, 0) year dummies
Industry			60 (1, 0) dummy variables at the two digit SIC level dummies
Regional wage	5·360	0·180	Weekly earnings of male workers in the Standard Region, entered in natural logarithms. Source: *Regional Trends*, various issues
Dependent Variable			
Annual earnings	8·710	0·762	Gross annual earnings before deductions of income tax and national insurance (natural logarithm). Grouped data in 13 categories in each year with open-ends. Mid-points allocated

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