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FEAR, UNEMPLOYMENT AND PAY FLEXIBILITY

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ABSTRACT

The paper uses newly available cross-section data to study wage determination in the United Kingdom in the 1980s. The results are contrasted with those from a comparable sample from the US from 1977-1988.

1) Fear of unemployment substantially depresses pay in both countries.

2) There is some evidence of a wage ratchet in the UK whereby rates of pay are more flexible upwards than downwards.

3) The unemployment elasticity of pay averages -0.1 in the UK and apparently zero in the US.

4) Wages are almost twice as flexible in non-union and small workplaces in the UK.

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1. INTRODUCTION

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[2] (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 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 with 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 i) for Britain between 1983 and 1989 and ii) for the US from 1977 to 1988 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 2 and in the earlier literature.

2. WAGES AND UNEMPLOYMENT.

A prerequisite for a convincing account of labour market behaviour is an answer to the long-standing question of whether the classical competitive model provides a satisfactory framework.

Consider a firm which sells in a risky product market and must take on workers before product demand is known. Employees themselves will then typically face some risk: they may be made redundant if demand conditions are poor. The 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 lay-off will typically generate lower wages. Employees who earn rents will wish to retain their jobs: the threat of redundancy may therefore induce workers to forego a portion of those rents. This difference in predictions suggests a way to discriminate empirically between competitive and non-competitive theory.

The paper bears a close relationship to recent attempts (Blanchflower, Oswald and Garrett, 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 (1987a), Black and Bulkley (1984), Gottfries and Horn (1987), Begg, Lindbeck, Martin and Snower (1989), and Lindbeck and Snower (1987). 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(3).

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.

Another issue addressed in the paper is that of how flexibility varies across different parts of the economy. Some models of the economy (including Hicks (1974), Oswald (1982), Minford (1982) and McDonald and Solow (1984)) assume a unionised sector with relatively sticky wages and a non-unionised sector in which pay is more responsive to unemployment. There is also a large literature (see Doeringer and Piore (1971), the survey by Taubman and Wachter (1986), and papers such as Wachtel and Betsy (1972) and Dickens and Lang (1985)) on the view

that an economy can be seen as divided between a primary labour sector and a secondary labour sector. This paper attempts to discover whether there is empirical support for these ideas: it examines a set of disaggregated wage equations.

The later analysis allows the estimation of the unemployment elasticity of wages (Blanchflower, Oswald and Garrett (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[4] on individual data that calculates this elasticity is Blackaby and Manning (1987, 1988) and Symons and Walker (1988). Studies for the US include Bils (1985), Rayack (1987) and Adams (1985).

3. FEAR OF REDUNDANCY.

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 firm closes down and pays only its fixed costs. The slump occurs with what the firm believes to be probability ß.

Assume that workers' preferences can be represented by an expected utility function. In the boom state the representative worker's utility[5] is u(w). In the slump it is u(b), where b is the level of unemployment benefit or an equivalent income level in another job. The slump occurs with what the representative worker believes to be some different probability α.

The two sides are assumed to act as if solving an asymmetric Nash (1953) bargaining problem[6]. Workers' relative bargaining strength is denoted $s/(1-s)$. Each party has an outside option: it can withdraw entirely. Assume that the representative worker receives wage w^a 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:

$$\text{Maximize}_{w} \; N \equiv s \log [(1 - \alpha)u(w) + \alpha u(b) - u^*] + (1 - s) \log [(1 - \beta)\pi(w,k) - \beta k - \pi^*] \quad (1)$$

subject to

$$w - w^a \geq 0 \quad (2a)$$

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

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 $[w^a, w^b]$ provides a formalization 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 $\S$ and μ for the two constraints, the Lagrangean may be written

$$L = N(s, \alpha, w, b, u^*, \beta, k, \pi^*) + \S[w - w^a] + \mu(\pi(w,k) - \pi(w^b,k)) \quad (4)$$

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

$$N_w + \$ + \mu\pi_w(w,k) = 0 \qquad (5)$$

This defines a wage bargaining function

$$w = w\ (s,\ \alpha,\ b,\ u^*,\ \beta,\ k,\ \pi^*,\ w_a,\ w_b). \qquad (6)$$

The signs of the derivatives of the wage function with respect to perceived shut-down probabilities are of particular interest. In the interval $(w^a,\ w^b)$, the constraints are not binding, so that, for example,

$$\text{sign}\quad \partial w/\partial\alpha = \text{sign}\ N_{w\alpha}, \qquad (7)$$

which uses the second-order condition $N_{ww} < 0$. It is straightforward to show that

$$\text{sign}\ N_{w\alpha} = \text{sign} - s(u(b) - u^*)u'(w)[(1 - \alpha)u(w) + \alpha u(b) - u^*]^{-2} < 0 \qquad (8)$$

Thus the equilibrium wage is lower the greater is the likelihood (as perceived by the workers' side) of the bad state of nature in which the firm is forced to close down. In the special case of a competitive labour market, the feasible band $[w^a,\ w^b]$ is degenerate at a point given by the going market wage.

The analysis can be seen as a formalisation of the long standing idea of 'concession bargaining'(7). Fear of redundancy reduces workers' wage demands. Whilst the model above relies on the extreme assumption of closure in the poor state of nature, the ideas apply more generally to layoffs.

In the competitive model of the labour market, risk of unemployment has quite a different effect upon wages. It has been noted by authors from Smith (1776) to Rosen (1986) that, ceteris paribus, the probability of lay-off and the level of pay should be positively associated in atomistic equilibrium. Workers who face a higher than average chance of becoming unemployed must, in a wage-taking (or, more generally, utility-taking) framework, receive a wage premium as a compensating differential.

In a competitive equilibrium all workers receive constant expected utility Eu. Where the notation is as before, therefore,

$$Eu = (1 - \alpha)\, u(w) + \alpha u(b).$$

This is common to all forms of (comparable) labour. The equation defines a locus linking wages and lay-off probability, for a given unemployment benefit level, with gradient

$$\partial w/\partial \alpha = \frac{u(w) - u(b)}{(1 - \alpha)\, u'(w)} \geq 0 \qquad (9)$$

This is weakly positive[8], not negative as in equations (7) and (8).

In a bargaining framework, external labour market pressure must operate through the delay or strike wage, w*. Aggregate unemployment then depresses pay if it reduces alternative sources of workers' incomes during a strike or lockout. When workers have a large amount of power, so that s is near unity, the effects of external pressures (such as unemployment) are small.

4. EMPIRICAL IMPLEMENTATION.

The theory of Section 3 indicates that wages will depend on, among other things, the probability of plant closure. More generally, α might

be taken to be the workers' perception of the probability of large scale job losses, in which the representative worker loses his or her job even when the plant continues in existence.

The two series of surveys used in this paper - the British Social Attitudes Surveys (BSA) and the US General Social Surveys (GSS) - are unusual in that they both 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; Part A of the Table relates to the UK and Part B to the US. First, workers in both countries were asked about their experience of unemployment. Between a quarter and one fifth of workers in the two countries 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, in the case of the UK, workers were asked whether they would leave their employer in the following year. If the respondent answered either 'very likely' or 'quite likely', 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. In the US, workers were asked a similar question - 'how likely is it next year that you will lose your job or you will be laid-off?'. Between 7% and 13% of workers per year reported either that this was 'very likely' or 'quite likely'.

Table 1. Workers' attitudes and unemployment experience (%)

A) **United Kingdom**

	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) Workplace 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	1532	1381	1353

b) **United States**

	1977	1978	1982	1983	1985	1986	1988
i) Unemployed last 5 years?	n/a	24	n/a	27	n/a	26	26
ii) Why leave employer?							
a) very likely	4	4	7	5	7	4	4
b) quite likely	6	3	8	8	5	7	4
c) not very likely	24	19	28	25	23	22	25
d) not at all likely	66	74	57	61	65	67	66
Number of employees	753	896	1013	918	934	846	608

Notes: all data reported in this Table are weighted

Third, in the UK alone, 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 would increase in size grew over the period from 16% to 27%, as the UK's economic climate improved. Such workers presumably assign a low value to their chance of redundancy[9].

Finally, in both countries 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

$$\ln w = \beta_0 + \beta_1 u + \beta_2 x + \varepsilon$$

which w is earnings, β_0 is a constant, β_1 is a vector of parameters, u is the set of unemployment variables discussed above, β_2 is a vector of parameters, x is a set of conventional control variables and ε is an error term. The x variables, which are familiar from other cross-section work, include

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, size, union status)

iv) a range of dummy variables (for industries and years).

A description of these variables and of the data sets is contained in the Appendices.

In the case of the UK 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 a 1988. Hence the next section presents results only on 1983 to 1986 and 1989. For the US, the paper pools surveys from the GSS for the years 1977, 1978, 1982, 1983, 1985, 1986 and 1989. These are the only years that respondents were asked if they expected to lose their jobs.

5. 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 the weekly wage. Due to the possibility that the dependent variable could be corrupted by movements in hours, these equations were also estimated using the logarithm of hourly earnings as the dependent variable. The results were little different and are omitted. They are available from the author on request.

Four out of five of the unemployment variables are statistically significant. Regional unemployment (across ten regions and entered as a log of the percentage) has an elasticity of approximately -0.11. This is similar to the aggregate time series results of Layard and Nickell (1986) and Carruth and Oswald (1987b), the panel data findings of Bils (1987), Nickell and Wadhwani (1990) and Christofides and Oswald (1989), and cross-section estimates by Blackaby

Table 2. United Kingdom, 1983-1989 and US, 1977-1988 - Weekly Wages

	United Kingdom				United States
	(1)	(2)	(3)	(4)	(5)
	All	All	Union	Non-union	All
Experience	.0257	.0257	.0248	.0256	.0515
	(12.86)	(12.81)	(8.75)	(9.00)	(16.93)
Experience^2	-.0004	-.0004	-.0004	-.0004	-.0007
	(11.90)	(11.85)	(8.14)	(8.29)	(13.06)
Schooling	.0728	.0734	.0798	.0656	.0886
	(15.28)	(15.38)	(11.84)	(9.73)	(19.96)
Male	.4456	.4452	.3819	.5014	.5129
	(29.84)	(29.80)	(19.23)	(22.89)	(21.34)
Married	.1059	.1065	.0948	.1267	.0602
	(5.72)	(5.75)	(3.97)	(4.55)	(2.16)
Separated	.1672	.1662	.1687	.1618	.0188
	(5.59)	(5.55)	(4.31)	(3.66)	(0.33)
Divorced					.0855
					(2.23)
Widowed	.2309	.2270	.1333	.2937	-
	(5.35)	(5.27)	(2.42)	(4.52)	
Self-employed last 5years	-.0063	-.0060	-.1471	.0366	-
	(0.19)	(0.18)	(2.57)	(0.86)	
Supervisor	.1736	.1743	.1691	.1755	.2192
	(13.05)	(13.11)	(9.85)	(8.74)	(7.27)
Part-time	-.8727	-.8731	-.8014	-.8965	-1.0229
	(47.18)	(47.18)	(29.82)	(34.88)	(33.60)
Nonmanual	.2630	.2641	.2640	.2657	.1305
	(17.98)	(18.06)	(12.91)	(12.56)	(4.40)
Union member	.0963	.0949	-	-	.1665
	(6.54)	(6.44)			(4.58)
Union recognition	.0468	.0473	.0815	.0569	-
	(2.61)	(2.64)	(0.39)	(2.36)	
London	.1372	.1365	.1560	.1240	-
	(7.03)	(6.99)	(5.86)	(4.43)	
Unemployment Variables					
Unemployed last 5 years	-.0997	-.1013	-.1435	-.0829	
	(6.54)	(6.64)	(6.31)	(4.10)	
Regional unemployment rate	-.1172	-.1155	-.0068	-.1960	-.0075
	(4.78)	(4.78)	(0.20)	(5.51)	(0.14)
Employment rise	.0868	.0878	.0462	.1134	-
	(5.69)	(5.76)	(2.10)	(5.38)	
Employment fall	.0153	.0072	-.0190	.0397	-
	(1.00)	(0.49)	(1.12)	(1.53)	
Redundancy expected	-.0865	-	-	-	-.2179
	(3.06)				(6.55)
Plant closure expected	-	-.1231	.0111	-.2922	-
		(2.17)	(0.16)	(3.17)	
Industry dummies	60	60	60	60	30
Year dummies	5	5	5	5	7
Constant	7.2079	7.1976	7.1368	7.5356	7.3842
	(67.12)	(67.02)	(38.71)	(38.42)	(49.42)
Adjusted R^2	.7039	.7037	.6594	.7189	.4903
Degrees of Freedom	5276	5276	2434	2768	5416
F	152.70	152.51	64.13	92.05	102.17

and Manning (1987, 1990). The estimate is also close to those derived from establishment data in our earlier work (1987, 1990c).

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 9% less, *ceteris paribus*. This ties in with much industrial relations evidence on concession bargaining (Cappelli (1985, 1988) describes the literature and presents modern results for the US), but appears to be the first estimate based on individual microeconomic data for Great Britain. Column 5 reports a similar result for the US, although here the coefficient is more than twice as large as for the UK. Workers in the US who report that they expect to lose their job in the following year or be laid off have approximately 20% lower pay, *ceteris paribus*. Interestingly enough, the coefficient on the US state unemployment rate is insignificant. In other work we have found that non-linear terms in unemployment are significant (see Blanchflower and Oswald, 1990b), but this possibility is not explored further here.

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 UK was replaced with one relating to the expectation of plant closure. (This information is not available for the US). As can be seen from column 2 of Table 2, fear of plant closure lowers pay by 12%, *ceteris paribus*. This seems to support the idea that fear of unemployment is not primarily a proxy for worker quality.

British workers who reported that they expected employment to grow at their workplace received a wage premium of approximately 9%. 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 Section 2. If, as seems plausible, workers feel secure when their workplace is growing, they may feel able - and be able - to extract higher remuneration[10] 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.

There is one caveat that should be noted at this point. The results discussed here examine, among other things, the effects upon wage determination of the level of unemployment in an individual's regional labour market. It is inevitably difficult to disentangle

regional unemployment effects from other kinds of regional influences. As might be expected, given the relatively small number of regional observations (11 * 5 years), the insertion into the British equation of a full set of (11) regional and (5) year dummy variables drives the regional unemployment coefficient insignificant. The other unemployment variables remain unchanged.

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[11]. Blanchflower, Oswald and Garrett (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 1. 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.

An obvious distinction to draw is between the union and non-union sectors. Columns 3 and 4 of Table 2 present estimates splitting the UK sample into those workers who do, and those who do not, report being

a member of a trade union. Four out of five of the unemployment variables have significantly larger coefficients in absolute terms in the non-union sector than in the union. The unemployment elasticity of pay is strikingly different across the sectors. In the union part of the economy it is -0.01, but not significantly different from zero at normal confidence levels. By contrast, the coefficient in the non-union part of Britain's private sector is -0.196. (t-statistic on the difference between the two coefficients is 3.84). Similarly, fear of plant closure is significant in the non-union sector but not the union (t-statistic on the difference = 2.63). The coefficients on the 'Employment rise' and 'Unemployed last five years' variables are also significantly different from each other in the two sectors (t-statistics on the difference are 2.21 and 1.99 respectively). It appears, therefore, that in Britain pay is more responsive to outside unemployment pressure in workplaces without trade unions. We could find no such differences between the union and non-union sectors for the US (results not reported)[12].

Table 2 also produces estimates of the influence of human capital and workplace variables. The conventional hump-shaped earnings/experience structure in both the UK and the US is confirmed. The profile is steepest in the US where it maximises after 37 years of experience compared with 32 years in the UK As known (Greenhalgh, 1980), marital status enters significantly in an earnings equation. The Mincerian schooling variables are highly significant and of the same order of magnitude in both countries. The coefficients on both the 'Married' and the 'Non-manual' variables are approximately twice as large in the UK than in the US. Brunello and Wadhwani (1989) have

recently suggested that employees' remuneration is least flexible in large firms. The results of Table 3, which uses data for 1984-6 and 1989 (because establishment size data is not available for 1983), tends to support this view. It is only in the smallest workplaces that fear of plant closure significantly reduces pay. Workers with a history of unemployment are paid significantly less in workplaces with at least 100 workers. Outside unemployment does not have a significant influence in the largest establishments of 500+ workers.

Union membership, according to the results in Table 2, leads to a wage premium of approximately 10% in the UK. This is only slightly above the existing estimates in Stewart (1983), Shah (1984) and Blanchflower (1984), inter alia. Disaggregated estimates of the union/non union wage differential or wage gap for the 1980s are reported in Table 4. Rather surprisingly, there is no evidence of a 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(13).

6. CONCLUSIONS

This paper studies pay determination in Britain and the US in the 1970s and 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.

Table 3. United Kingdom, 1985-1989 - Weekly Wages by Workplace Size

	(1) < 25	(2) 25-99	(3) 100-499	(4) 500+
Experience	.0285	.0241	.0192	.0300
	(7.01)	(5.37)	(4.62)	(35.11)
Experience^2	-.0005	-.0004	-.0004	-.0005
	(6.29)	(4.38)	(4.55)	(6.13)
Schooling	.0734	.0713	.0638	.0706
	(7.10)	(7.49)	(6.70)	(6.95)
Male	.4420	.4638	.4746	.4212
	(13.56)	(14.60)	(16.16)	(12.60)
Married	.0642	.1251	.0972	.1469
	(1.62)	(3.05)	(2.64)	(3.61)
Separated	.0846	.2014	.2567	.1424
	(1.36)	(3.00)	(4.04)	(2.17)
Widowed	.2035	.2367	.1055	.1991
	(2.35)	(2.49)	(1.13)	(1.87)
Self-employed last 5years	.0069	.0111	-.0346	-.1027
	(0.11)	(0.16)	(0.41)	(1.18)
Supervisor	.2230	.1258	.1789	.1745
	(7.78)	(4.33)	(6.76)	(6.20)
Part-time	-.9102	-.8472	-.7604	-.8028
	(26.44)	(20.64)	(18.52)	(15.69)
Nonmanual	.2608	.2833	.2714	.2621
	(8.14)	(8.99)	(9.01)	(8.10)
Union member	.1499	.0963	.0665	-.0104
	(4.28)	(3.10)	(2.33)	(0.32)
Union recognition	.0633	-.0085	-.0402	-.0204
	(1.60)	(0.21)	(1.00)	(0.35)
London	.1458	.1056	.1060	.1120
	(3.26)	(2.58)	(2.61)	(2.65)
Unemployment Variables				
Unemployed last 5 years	-.0417	-.0692	-.1741	-.1775
	(1.39)	(2.16)	(5.34)	(4.35)
Regional unemployment rate	-.1068	-.0496	-.1109	-.0599
	(2.12)	(0.95)	(2.21)	(1.10)
Employment rise	.0688	.1046	.0897	-.0012
	(2.087)	(3.23)	(3.01)	(0.04)
Employment fall	-.0042	.0276	.0260	-.0974
	(0.11)	(0.84)	(0.92)	(3.33)
Plant closure expected	-.2380	-.1204	.1335	-.0705
	(2.03)	(0.84)	(1.02)	(0.11)
Industry dummies	60	60	60	60
Year dummies	4	4	4	4
Constant	7.5458	7.5726	7.8003	7.8005
	(35.65)	(34.31)	(22.14)	(35.11)
Adjusted R^2	.6897	.6629	.6857	.7061
Degrees of Freedom	1462	1091	1065	724
F	48.35	31.19	35.01	27.92

Notes: 1. US equation in Table 2 covers years 1977-8, 1982-3, 1985-6 and 1989. The equation also includes 2 race dummies and a children dummy.

2. UK weekly wage equations in Table 2 are for years 1983, 1984, 1985, 1986 and 1989 In Table 3 they cover years 1984, 1985, 1986 and 1989.

3. In 1989 the BSA included an additional area - Northern Ireland.

Table 4. 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.

* insignificantly different from zero at the 1% level.

The paper suggests that risk of plant closure can be expected 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 in the UK, and 22% less in the US, *ceteris paribus*(14). UK workers who say they expect their plant to close earn 12% less than those who do not.

2. There is some evidence of an asymmetry or 'wage ratchet' in the UK. Workers in expanding plants receive a pay premium; those in contracting plants suffer no pay disadvantage. (The one exception to this is found in the biggest plants). 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 -0.1 in the UK and apparently zero in the US.

4. In the UK a history of personal unemployment depresses pay 10% on average. Being a supervisor raises pay by approximately 17% on average. Londoners earn 12% more on average.

5. The union wage gap (or mark-up) in the UK in the 1980s 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 1980s.

There is no indication, from these equations, that fear of unemployment is compensated by higher pay: my judgement is that the competitive model of the labour market is an unreliable guide. It seems more appropriate to see unemployment as a force in a non-competitive world which acts to weaken workers' negotiating power. This emasculation is clearest in Britain's non-union sector.

Appendix A - Data Sources

1. British Social Attitudes Survey Series, 1983-1989

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.

2. US General Social Surveys, 1972-1988

The General Social Surveys have been conducted every year since 1972, with the exception of 1979 and 1981. The surveys have been funded by the Russell Sage Foundation and the National Science Foundation. Each survey is a sample of English speaking persons 18 years of age or over, living in non-institutional arrangements within the United States. The sample is a multi-stage probability sample, with the primary sampling units being Standard Metropolitan Areas or non-metropolitan counties. For details of the sample design etc., see General Social Surveys, 1972-1988: Cumulative Codebook, National Opinion Research Center, Chicago.

Appendix B.

Variable Definitions - UK Sample

	Mean	SD	
Male	.545	.498	a (1,0) dummy variable for gender.
Experience	22.07	12.992	(age - schooling) + 5
Part-time	.180	.384	a (1,0) dummy variable if the respondent reported that they normally worked less than 30 hours per week.
Separated	.047	.213	a (1,0) dummy variable if the respondent was separated or divorced.
Widow	.020	.139	a (1,0) dummy variable if the respondent was widowed.
Married	.727	.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	.469	.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	.215	.411	a (1,0) dummy variable if the respondent expected their workplace to increase its number of employees over the coming year.
Employment fall	.238	.426	a (1,0) dummy variable if the respondent expected their workplace to decrease its number of employees over the coming year.
Supervisor	.362	.481	a (1,0) dummy variable if the respondent was a supervisor.
Union member	.464	.499	a (1,0) dummy variable if the respondent was a member of a trade union or a staff association.
Non-manual	.548	.498	a (1,0) dummy variable if the respondent's occupation was non-manual.
Self-employed last 5 yrs.	.035	.184	a (1,0) dummy variable if the respondent reported that they had ever been self-employed over the previous five years as their main job.
Unemployed last 5 years	.208	.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	.050	.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	.011	.104	a (1,0) dummy variable if the respondent expected that during the next year they would lose their job because of firm closure
Regional unemployment	2.272	.395	unemployment rate in the Standard Region, entered in in natural logarithms.
London	.109	.311	a (1,0) dummy representing Greater London.
Year84/5/6/7/9			5 (1,0) year dummies
Industry			60 (1,0) dummy variables at the two digit SIC level dummies

Dependent Variable

Annual earnings	8.710	.762	Gross annual earnings before deductions of income tax and national insurance (natural logarithm). Grouped data in 13 categories with open-ends. Mid-points allocated.

Variable Definitions - US Sample

Independent Variables

	Mean	SD	
Male	.531	.499	a (1,0) dummy variable for gender.
Experience	20.86	13.69	(age - schooling) + 5
Part-time	.161	.368	a (1,0) dummy variable if the respondent reported that they worked part-time in the preceding week.
Separated	.039	.193	a (1,0) dummy variable if the respondent was separated.
Divorced	.121	.326	a (1,0) dummy variable if the respondent was divorced.
Married	.606	.489	a (1,0) dummy variable if the respondent was married or living as married.
School	13.025	2.879	number of years of schooling.
Supervisor	.191	.393	a (1,0) dummy variable if the respondent was a supervisor.
Union member	.095	.293	a (1,0) dummy variable if the respondent was a member of a trade union or a staff association.
Non-manual	.692	.462	a (1,0) dummy variable if the respondent's occupation was non-manual.
Redundancy expected	.112	.316	a (1,0) dummy variable if the respondent expected that during the next year they would lose their job or be laid off.
Regional unemployment	1.989	.274	unemployment rate in the State, entered in natural logarithms.
Year77/78/82/83/85/86/88			7 (1,0) year dummies
Industry			30 (1,0) dummy variables at the two digit SIC level dummies

Dependent Variable

Annual earnings	9.324	1.044	Gross annual earnings before tax or other deductions (natural logarithm). Grouped data. Mid-points allocated.

ENDNOTES

(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. But it appears to be reasonable as a generalisation.

(2) It should be noted that cross-section analysis generally rests upon untested assumptions about model identification and causality. At the least, however, such results provide a set of stylised econometric facts about key labour market variables, and are therefore complementary to theoretical model building.

(3) I am grateful to a referee for pointing out that if there is a cyclically repeating stationary economy with no productivity gowth, 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.

(4) As far as I am aware.

(5) The representative worker may be the one with median seniority. Oswald (1987) explores a model related to the one developed here.

(6) This can be justified axiomatically as in Nash (1953) or axiomatically as in Binmore, Rubinstein and Wolinsky (1987).

(7) Shultz and Myers (1950) were among the first to document the phenomenon.

(8) It can certainly be zero, as in simple frictionless models.

(9) 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.

(10) Part of the effect may be overtime working, but that can be viewed as an expression of insider power.

(11) Further evidence and discussion is provided by Dickens and Katz (1987), Krueger and Summers (1986) and MacKay et. al. (1971).

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

(13) Having done various checks, I can only conclude that there was some problem with the 1985 data collection.

(14) 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. They found that non-union wages are sticky downwards but flexible upwards. Union wages were found to be sticky in both directions.

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