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The Wage Curve*

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Abstract

This paper, which follows in an LSE tradition begun by Phillips and Sargan, examines the role of unemployment in shaping pay. In contrast to most of the literature, it (i) uses micro-econometric data on individuals and workplaces, (ii) examines a variety of data sets as a check on the robustness of results, and (iii) studies the effects of unemployment on the real wage level (not on the rate of change of pay or prices). Evidence is found — on British and U.S. data — of a wage curve. The curve has a negative gradient at low levels of unemployment, but becomes horizontal at relatively high levels of unemployment.

I. Introduction

There exists much evidence that capitalist economies periodically suffer from high and persistent unemployment. Many economists have argued that, in a way not yet fully understood, normal wage adjustment mechanisms fail to operate when unemployment is high.

The object of this paper is to provide microeconomic evidence related to this failure and how it occurs. A *wage curve*, defined more precisely below, is estimated which describes the way in which unemployment acts to depress the level of pay. Unlike the large literature stemming from the work of Phillips (1958), the estimation is done on microeconomic data,

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and the focus is on the relationship between unemployment and the wage level rather than between unemployment and wage inflation. The paper uses one U.S. and three British data sets — each to provide a check against the others.

A central finding of the empirical analysis is that there is a wage curve which becomes flat at moderately high levels of unemployment. At such levels, therefore, the equilibrating forces of the labour market can break down. Similar wage curves emerge from the four data sets, which suggests that the statistical results are robust.

It is also argued that one well-known explanation for persistent British unemployment appears to be incorrect. The Layard and Nickell (1987) hypothesis, that it is a high proportion of long-term unemployment which nullifies downward wage adjustment in a slump, is tested. It is found that the effect of long-term unemployment disappears once non-linear unemployment terms are included. This suggests that the role found for long-term unemployment in previous wage equations — all estimated on short time-series data sets — may have been the result merely of a correlation between high unemployment and high long-term unemployment.

Section II of the paper concerns new work on the unemployment elasticity of pay (the proportional responsiveness of the real wage to the level of unemployment). Section III outlines a model of wage determination. It suggests that unemployment depresses pay by weakening workers' bargaining power, and that the exact shape of the resulting wage function cannot be determined on theoretical grounds. Section IV examines microeconomic data and, after adjustment for many individual and workplace control variables, reveals evidence of a wage curve linking workers' pay to the unemployment rate in their local area or industry. The estimated wage curve becomes flat at relatively high levels of unemployment. Section V concludes, and the Appendix provides background information.

II. Earlier Work on the Unemployment Elasticity of Pay

The work of Phillips (1958) has produced an unusually large literature.[1] Although Phillips' empirical evidence was greeted enthusiastically by many economists, observers became progressively more sceptical. The

[1] See, for example, Lipsey (1960), the survey by Laidler and Parkin (1975), Wadhwani (1985), the debate between Desai (1975, 1984) and Gilbert (1976), and the work on cross-country Phillips Curves in Grubb, Jackman and Layard (1983), Paldam (1980), Newell and Symons (1985) and Grubb (1986).

Phillips Curve is generally thought of as a fragile empirical relationship; see, for example, the recent negative results of Christofides *et al.* (1980) and Beckerman and Jenkinson (1986). The work of Phelps (1967) and Friedman (1968) exposed the theoretical weaknesses of early formulations. A long-run Phillips Curve, according to modern theory, is vertical.

Sargan (1964) was one of the first economists to point out that the Phillips Curve could be thought of as an adjustment mechanism around a long-run relationship in which the wage level depends upon the unemployment level. Sargan saw this as a function calibrating the way in which wage bargainers' demands are shaped by the extent of joblessness in the whole economy. In Sargan (1964), which has been widely overlooked by all but a small group of (time-series) econometrics specialists, the author estimated the average long-run elasticity of pay at −0.03.

This intellectual tradition is carried on today in work such as Layard and Nickell (1986). The authors estimate a real wage equation on British data between 1950 and 1983. Their estimate of the unemployment elasticity of pay is −0.06. Carruth and Oswald (1987, 1989) and Holly and Smith (1987) obtain slightly larger elasticities of −0.1 or over. In contrast to these British results, Sneesens and Drèze (1986) find a statistically insignificant elasticity for Belgium. Very small elasticities emerge from studies of Scandinavian pay, such as Hoel and Nymoen (1988) and Andersen and Risager (1988).

A difficulty with these kinds of studies is that small numbers of degrees of freedom are inevitable. Time series analysis based on highly aggregated data has other well-known limitations. In an attempt to apply a different empirical method Section IV of this paper uses cross-section data to explore the connections between unemployment and the wage level. This builds upon work reported in Blanchflower, Oswald and Garrett (1990).

The last few years have seen an expansion in the numbers of cross-section inquiries into the effect of unemployment upon pay. However, few attempts[2] have been made to draw together and compare the various estimates. Table 1 does this. First, it reveals that there is extensive evidence that unemployment depresses the real wage level. Second, it suggests that the unemployment elasticity of pay is small. Numbers close to −0.1 are the norm but estimates insignificantly different from zero exist. Adams (1985) and Beckerman and Jenkinson (1988) even obtain a positive elasticity on one unemployment rate and a negative one on another (when two rates are entered simultaneously). Third, fairly similar estimates of the unemployment elasticity of pay emerge from studies of the U.S., Britain, Sweden and Canada.

[2] Oswald (1986b) discusses some of the literature and concludes that the unemployment elasticity of pay is around −0.1.

Table 1. *Estimates of the unemployment elasticity of real wages from cross-section and panel data*

Study	Data	Notes	Unemployment elasticity
1. Bils (1985)	U.S. NLS Panel, 1970s, 5,000 young males.	Aggregate annual U.S. unemployment used as independent variable. Few annual observations.	−0.1 (approx.)
2. Rayack (1987)	U.S. PSID Panel, 1968–80, 27,000 white males.	Aggregate annual U.S. unemployment rates.	−0.1 (approx.)
3. Adams (1985)	U.S. PSID Panel, 1970–76, various samples.	State and industry unemployment rates.	−0.02 to −0.11 (industry rates) 0.13 to 0.20 (state rates)
4. Beckerman and Jenkinson (1986)	Panel of 12 OECD countries, 1963–83.	National unemployment rates.	Approx. zero
5. Beckerman and Jenkinson (1988)	Panel of 14 U.K. manufacturing industries 1972–86.	Unemployment by industry and nationally. Data on 1983–86 constructed by authors.	−0.13 (aggregate rates) +0.18 (industry rates)
6. Blanchflower (1989)	British BSA, 1983–86. 3,800 adult workers.	Regional unemployment.	−0.1
7. Blackaby and Manning (1987)	British General Household Survey, 1975, 7,300 white males.	Regional unemployment.	−0.16
8. McConnell (1988)	U.S. union contract data, 1970–81, 3,000 contracts.	State unemployment.	Approx. zero
9. Holmlund and Skedinger (1988)	Panel on Swedish timber industry. 70 regions. 1969–85.	Regional and national unemployment.	Zero to −0.04

10. Blanchflower, Oswald and Garrett (1990)	British 1984 WIRS, manual workers in 1,200 establishments.	County unemployment.	Zero to −0.14
11. Blanchflower and Oswald (1990)	British 1984 and 1980 WIRS. Non-manual workers in 800 establishments.	Regional unemployment. Regional wage included as a control.	Zero to −0.08
12. Nickell and Wadhwani (1987, 1988)	Panel of 219 U.K. firms. 1974–82.	Industry and national unemployment.	−0.05 (industry) −0.05 (national)
13. Christofides and Oswald (1988)	Canadian union contract data.	Provincial unemployment.	−0.03 to −0.12
14. Card (1988)	Canadian union contract data, 1,293 contracts, 1966–83.	Provincial unemployment. National unemployment for some provinces.	−0.05 to −0.1
15. Freeman (1988)	U.S. state data. British county data. Changes from 1979–85.	State and county unemployment.	Zero to −0.1 (approx.)
16. Symons and Walker (1988)	British FES data, 6,500 married males, 1979–84. Various samples.	Monthly regional unemployment.	Zero to −0.2

Notes:

We are grateful to Mark Bils for calculating for us the elasticity implicit in Bils (1985). Rayack (1987) does not report an elasticity explicitly. We have calculated the figure "−0.1 approx." by inserting our best estimate of the unemployment rate in his data set. Similarly for Card (1988).

We thank Ian Walker for helpful discussions about the elasticities in Symons and Walker (1988), and David Blackaby for the same on Blackaby and Manning (1987).

It is not possible to calculate the elasticity in Freeman (1988) so we have inserted our estimate of the British and U.S. means. The U.S. elasticity is insignificant at 5 per cent confidence, which is why the table gives zero as the lower bound.

III. Theoretical Issues

To place our statistical results in context, a theoretical framework is required. Although this paper does not attempt to test one specific model against another, it suggests a way to rationalize the correlations observed in the data.

Consider the following model. Assume that a profit maximizing firm bargains with a utility maximizing trade union.[3] Assume that the firm has a maximum profit function

$$\pi(w, p) = \max_{n} pf(n) - wn, \tag{1}$$

where w is the wage, p is the (exogenous) price of output, n is employment and $f(n)$ is a well-behaved production function. Assume that the union's utility function can be described locally by $u = w$, so that the union is risk neutral and assigns no weight to employment.[4] These assumptions are stronger than necessary for later results.

Assume that wage determination may be modelled as a Nash bargain. This may be justified axiomatically, as in Nash (1953), or strategically, as in Binmore, Rubinstein and Wolinsky (1986), and might even be used as a model of rent-sharing in the non-union sector.

Let the firm's and union's fall-back or delay utilities be, respectively, π^* and u^*. The former may be thought of as profit during a strike; the latter can be seen as a worker's income while on strike. The value of u^* will depend upon the availability of, and wage paid in, temporary work.[5] By the assumption of risk neutrality, the expected utility of a worker on strike may be assumed to be

$$u^* = w^* = s(U)\, y + (1 - s(U))\, z, \tag{2}$$

where $s(U)$ is the probability of finding temporary work, U is the unemployment rate in the economy, y is the income paid in the temporary

[3] Among the early studies, both Sargan (1964) and Desai (1973) rely implicitly on a bargaining framework. Recent empirical work on bargaining models includes Nickell and Andrews (1983), Blanchflower, Oswald and Garrett (1990), Hoel and Nymoen (1988), Holmlund and Skedinger (1988), Nickell and Wadhwani (1987) and Rowlatt (1987). Empirical work on the simpler monopoly union model includes Farber (1978), Pencavel (1984), Carruth and Oswald (1985), Pencavel (1985) and Hersoug, Kjær and Rødseth (1986).

[4] This is a special case of the seniority model in Oswald (1987); a state-contingent version is contained in Oswald (1986a). The flat indifference curve model is criticized in Holmlund and Skedinger (1988). The model of Section III can be generalized to a union utility function such as that in Drèze and Modigliani (1981), McDonald and Solow (1981) or Oswald (1982), by allowing unemployment to affect the worker's alternative income.

[5] During a strike the worker may derive income from other sources, such as spouse's earnings. This income is likely to depend upon the availability of jobs in the outside labour market.

job, and z is the income — equivalent value of leisure if no temporary work can be found. It is assumed that $y - z > 0$.

The function $s(U)$ is of some importance. It captures the probability that the striking employee will be successful in finding a temporary source of income. The function is assumed to be declining and convex in unemployment, U, and to have the characteristics

$$\lim_{U \to U'} s = 0 \tag{3a}$$

$$\lim_{U \to 0} s = 1. \tag{3b}$$

Thus when unemployment is U' the individual is certain to be unable to find temporary work, whereas when unemployment is zero the individual is always able to find such work. Once unemployment reaches U', therefore, a striking worker has no chance of obtaining additional income from the labour market. The worker's bargaining power reaches a minimum at this level, and remains there as unemployment rises above U'.

The Nash bargain solves the problem

$$\max_{w}(\pi(w, p) - \pi^*)(w - w^*), \tag{4}$$

so that an interior optimum requires that the wage be given by

$$w = w^* + \frac{\pi - \pi^*}{n}. \tag{5}$$

This presupposes that the problem is concave (it can be checked that the second-order condition relies on the restriction that the elasticity of labour demand be lower than 2). It should be stressed that equation (5) is easily generalized.

Equation (5), the wage formula, states that the outcome of the wage negotiations depends upon the sum of two components. The first, w^*, is the delay wage, namely, the level of income the individual earns during a breakdown in wage negotiations. By equation (2) this is taken to be a convex combination of the wage in temporary work and the value of leisure. The second component, $(\pi - \pi^*)/n$, is the level of (adjusted) profit per employee. Put loosely, the equilibrium level of pay is shaped by a mixture of external and internal forces.[6]

The next issue is that of how the unemployment rate affects wage determination. The first-order condition for the Nash maximization may

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[6] This is consistent with industrial relations surveys of managers' views on the forces determining pay; see, for example, Blanchflower and Oswald (1988).

be written

$$\pi(w, p) - \pi^* + (w - s(U)\, y - (1 - s(U))\, z)\, \pi_w = 0. \tag{6}$$

By differentiation, at this optimum,

$$\mathrm{d}w\{2\pi_w + (w - s(U)\, y - (1 - s(U))\, z)\, \pi_{ww}\} - \mathrm{d}U\{\pi_w s'(U)\,(y - z)\} = 0. \tag{7}$$

At a maximum the first term within curly brackets is negative. The second of the terms in curly brackets is positive.

Equation (6) defines an equation linking the price of labour to the level of unemployment. Intuitively, higher unemployment in the outside labour market weakens the union's bargaining strength, because it reduces workers' chances of finding temporary income during a delay (such as a strike) in reaching a wage agreement.

The wage function described by (6) is, in general, nonlinear in the level of unemployment. The properties of this wage curve matter at the macroeconomic level; see e.g. Layard and Nickell (1986).

Define the elasticity of labour demand as

$$\alpha = -w\pi_{ww}/\pi_w. \tag{8}$$

Then

$$\frac{\mathrm{d}w}{\mathrm{d}U} = s'(U)\,(y - z)\Big/\left\{2 - \left(1 - s(U)\,\frac{y}{w} - (1 - s(U))\,\frac{z}{w}\right)\alpha\right\} < 0. \tag{9}$$

Under these assumptions, the wage curve has a negative gradient.

Beyond this it is difficult to make clear predictions about the structure of the wage curve. However, by assumptions made earlier about the $s(U)$ function, and the monotonicity of the curve, it is necessary that

$$\lim_{U \to U'} w = w^{\min} = z + \frac{\pi - \pi^*}{n} \tag{10}$$

$$\lim_{U \to 0} w = w^{\max} = y + \frac{\pi - \pi^*}{n}. \tag{11}$$

Beyond unemployment levels of U', therefore, the wage curve is flat. A somewhat similar argument has been made recently by Manning (1988).

Differentiation of equation (9) produces a complicated mixture of positive and negative terms. Even under the restriction that α, the labour demands elasticity, is a constant, there appears no way to sign unambiguously the second derivative of the wage curve.[7]

[7] We would like to thank David Soskice and Meghnad Desai for detailed suggestions on the theory of the wage curve, which will be explored in our forthcoming book.

The wage curve is bounded below. No matter how high the level of unemployment, the individual employee on strike enjoys some positive value from leisure. This imparts a minimum degree of bargaining power.

The thrust of this approach is different from that associated with papers stemming from Hall (1972), such as Marston (1985) and Topel (1986), in which wages and unemployment are positively correlated in equilibrium. Under the assumption of imperfectly competitive labour markets the traditional compensating wage differential argument need not apply.

IV. Estimating the Wage Curve on Micro Data

The possibility of complicated nonlinearities has rarely been considered in empirical work. However, both Carruth and Oswald (1987, 1988) and Nickell (1987) find that, on British time series data, the best equation is more complex than a loglinear function. Nickell (1987), for example, shows that the wage equation reaches a minimum, in unemployment space, at 19 per cent unemployment (male rate, pre-1982 definition). Similar evidence is presented in Layard and Nickell (1987), where the authors argue that it is high long-term unemployment which pushes up wage pressure at large levels of total unemployment. Carruth and Oswald (1987) report wage equations in which the coefficients of both the natural logarithm of unemployment and its cube are statistically significant. Whilst these studies find evidence of a highly nonlinear wage curve, they suffer inevitably from a shortage of observations. This is liable to be a particular difficulty when the object is to estimate nonlinearities.

An alternative approach is to use microeconomic data sets which allow for effects from outside unemployment. In contrast to the macroeconomic Phillips Curve tradition, an aggregate link from unemployment to wages can be investigated by studying the existence of such a relationship at the microeconomic level.

The method adopted in this paper is to use data on unemployment rates by geographical area or industry. These are inserted into microeconomic wage equations. The underlying assumption is that — if imperfectly — unemployment in the firm's local area or particular industry can be used to proxy external labour market forces.

The empirical work uses four microeconomic data sets:

(1) The Workplace Industrial Relations Survey of 1980 (WIRS), which provides information on approximately 2,000 British establishments.

(2) The National Child Development Study (NCDS) of 1981, which provides information on approximately 6,000 British 23 year old employees.

(3) The British Social Attitude (BSA) Surveys of 1983–87, which when pooled provide information on approximately 5,000 British adult employees.
(4) The International Social Survey Programme (ISSP) data for 1985–87, which when pooled provide information on approximately 2,000 U.S. employees.

Details of surveys 1–3 are given in Blanchflower and Oswald (1989b), and of survey 4 in Blanchflower and Oswald (1989a).

For WIRS and NCDS, county unemployment rates (across approximately 65 countries) were grafted onto the data sets. For BSA, regional unemployment rates across 11 regions by 5 years were added to the survey data. In the case of the ISSP survey, 30 industry unemployment rates by 3 years were added. In all cases the unemployment variables were total unemployment rates. To allow an examination of the Layard and Nickell (1987) hypothesis that long-term unemployment influences wage pressure, long-term unemployment rates were also entered, by geographical area, into the British data sets.

For each data set the wage equation

$$w = f(x, U)$$

was estimated, where x is a vector of individual or establishment variables, and U is the unemployment percentage in the relevant county or region. The x variables are of a kind conventional in the literature on cross-section wage equations.

In all four data sets there is evidence of a nonlinear association between pay and unemployment. This was investigated by fitting different polynomial structures in unemployment. After some experimentation it was found that two specifications worked well:

(i) $w = g(x, U, U^2)$
(ii) $w = h(x, \log U, (\log U)^3)$

In general these give similar results. Checks using a series of dummy variables confirmed that the curvature was not being forced on these data by the functional forms.

The detailed wage equations are set out in Tables 2–5.[8] As these and Table 6 show, there is a well-defined wage curve which becomes horizontal between 9 per cent and 15 per cent unemployment.[9] This is the paper's

[8] This paper reports wage equations which include very large numbers of control variables. Given the focus of the paper, and constraints of space, it is not possible to provide a proper discussion of them here. The tables are constructed to be as self-explanatory as possible -- Blanchflower and Oswald (1989b) report the complete results. A full description of the variables' effects will be given in a forthcoming book.

[9] The wage is at a minimum at somewhat different levels of unemployment across the four data sets. Current work is concerned with explanations of the different minima.

Table 2. *British WIRS1 wage equations, 1980*

	(1)	(2)	(3)	(4)	(5)	(6)
Log unemployment				−0.0827	−0.2353	−0.2442
				(4.35)	(2.65)	(2.75)
(Log unemployment)^3					0.0129	0.0162
					(1.76)	(2.10)
Unemployment	−0.0096	−0.0387	−0.0371			
	(3.88)	(3.00)	(2.87)			
Unemployment^2		0.0017	0.0019			
		(2.29)	(2.52)			
Long-term unemployment			−0.0044			−0.0037
			(1.50)			(1.30)
Constant	4.0560	4.1646	4.2202	4.1432	4.3353	4.3932
	(118.50)	(71.30)	(60.97)	(90.96)	(36.68)	(34.79)
Adjusted R^2	0.4915	0.4930	0.4935	0.4929	0.4936	0.4939
F	26.542	26.252	25.867	26.680	26.312	25.907
Degrees of freedom	1,449	1,448	1,447	1,449	1,448	1,447

Notes: The following control variables were also included:

(1) Per cent part-time employees; (2) per cent female manuals; (3) per cent manuals; (4) per cent skilled; (5) shiftworking dummy; (6) financial performance dummies; (7) quadratic in establishment size; (8) age of plant dummies; (9) union and closed shop dummies; (10) private sector dummy; (11) 60 industry dummies.

The equation uses 65 county unemployment rates. For full details of these equations see Blanchflower and Oswald (1989b).

Table 3. *British NCDS wage equations, 1981*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log unempt.	−0.246130	−0.111986	−0.086468	−0.223615	−0.039634	−0.081034		−0.309471	−0.755016
	(8.81)	(2.82)	(2.25)	(8.27)	(1.27)	(2.34)		(2.58)	(6.75)
(Log unempt.)^3								0.014042	0.032134
								(2.00)	(4.88)
Unemployment							−0.029279		
							(2.89)		
(Unemployment)^2							0.000983		
							(2.50)		
Long-term unempt.	0.004220	0.002097	0.001784	0.004130	0.000789	0.001819	0.000209	0.000604	0.002432
	(2.46)	(1.00)	(0.89)	(2.49)	(0.42)	(0.95)	(0.10)	(0.30)	(1.45)
South East		*	*		0.066016	0.041895	0.035944	0.038938	
					(5.00)	(2.64)	(2.22)	(2.46)	
Greater London		0.108234	0.107189		0.180265	0.149586	0.146926	0.148973	
		(6.72)	(6.88)		(11.99)	(8.02)	(7.85)	(8.07)	
Regional dummies (9)	No	Yes	Yes	No	No	No	No	No	No
"Worst" regions†						−0.029829	−0.024511	−0.021808	
						(2.77)	(2.28)	(2.01)	
Industry dummies	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Constant	4.618071	4.366995	4.282714	4.516647	4.131207	4.229183	4.343349	4.679752	5.449024
	(62.92)	(49.33)	(46.65)	(57.15)	(48.63)	(45.99)	(45.52)	(22.36)	(28.54)
Adjusted R^2	0.47720	0.48916	0.53101	0.52037	0.53088	0.53137	0.53419	0.53418	0.52473
F	114.06192	100.47902	60.29091	62.84190	64.38182	63.96073	64.66702	63.71940	62.97289
D. of freedom	6,389	6,379	6,318	6,328	6,326	6,325	6,323	6,282	6,285

Notes:
*Excluded category.
†Category includes E. Anglia, S. West, E. Midlands, Yorshire and Humberside, and Scotland.

The following control variables were also included:
(1) Fourteen qualification dummies; (2) 2 numeracy/literacy problems dummy variables; (3) 4 health dummies; (4) 8 dummies for workers' attitudes to jobs; (5) gender, marital status, children dummies; (6) dummies for second job, previous unemployment, unsocial hours; (7) no. of jobs since leaving school; (8) tenure current job; (9) 2 unionization dummies; (10) dummy for a move of location between 1974/81; (11) 5 plant size dummies; (12) branch office and limited company dummies; (13) 63 industry dummies.
The equations use 65 county unemployment rates. For full details of these equations see Blanchflower and Oswald (1989b).

principal result. It suggests that, when unemployment is sufficiently large, downward pressure on pay reaches a maximum. A further increase in unemployment then has no depressive effect on wage rates. At low levels of unemployment, however, the unemployment elasticity of pay is negative rather than zero.

The empirical evidence presented here suggests that the recent literature on real wage equations has been wrong to assume constancy of the unemployment elasticity of pay. This assumption has rarely been tested. The paper's results seem to imply that it is only at low levels of unemployment that the real wage is flexible.

Table 2 reports wage equations for semi-skilled manual workers using data from the 1980 Workplace Industrial Relations Survey (WIRS), which are based upon our earlier work in Blanchflower (1984) and Blanchflower and Oswald (1990). The dependent variable is the "gross (weekly) pay of the typical semi-skilled employee". The wage data are grouped and open-ended. We follow the standard practice of allocating midpoints to the wage bands. A series of sensitivity tests were undertaken which showed that the results were stable to changes in the values allotted to the end categories.

Table 3 presents wage equations using data on approximately six thousand young people — all of whom were born between March 3 and March 9, 1958 — who were interviewed in 1981. They form part of a large scale cohort study — the National Child Development Study (NCDS). In this case the respondent reported to the nearest £ his or her "gross (weekly) pay before deductions for tax and National Insurance including any overtime, bonus, commission and tips" on the last occasion they were paid.

Table 4 reports wage equations using a pooled set of cross-sections from the British Social Attitude Surveys of 1983–86. The dependent variable is "gross annual earnings before deductions of income tax and National Insurance". Once again the wage data are grouped and open-ended — in this case into 13 categories. The same method described in the paragraph above was used to allocate values to each of the wage bands.

A difficulty with the results on WIRS, BSA and NCDS is that it is not possible to control adequately for region specific fixed effects. It is only in BSA that a full set of area dummies can be entered, and — perhaps unsurprisingly given the small number of data points — when this is done (see equations (4) and (8) of Table 4) the unemployment rates become insignificant. The NCDS results do allow for region-level fixed effects: for example equation (7) of Table 3 reveals that the wage curve is robust to the inclusion of regional dummies.[10]

[10] The 10 regional dummies in column 3 of Table 5 are grouped in 3 categories, on the basis of F-tests.

Table 4. *British BSA wage equations, 1983–87*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log unemployment	−0.1128	−1.0184	−1.0212	−0.3186				
	(4.03)	(2.80)	(2.81)	(0.52)				
(Log unemployment)[3]		0.0520	0.0506	0.0533				
		(2.50)	(2.42)	(1.48)				
Unemployment					−0.0094	−0.0791	−0.0840	0.0205
					(3.79)	(2.99)	(3.09)	(0.33)
(Unemployment)[2]						0.0030	0.0031	0.0014
						(2.64)	(2.71)	(0.64)
Long-term unemployment			0.0017				0.0021	
			(0.68)				(0.83)	
Regional dummies	No	No	No	Yes	No	No	No	Yes
Constant	6.5900	8.0270	7.9875	6.5853	6.4248	6.8120	6.7763	6.1511
	(57.74)	(13.69)	(13.56)	(6.50)	(67.24)	(38.94)	(37.60)	(15.83)
Adjusted R^2	0.6764	0.6767	0.6767	0.6781	0.6762	0.6766	0.6766	0.6781
F	342.86	332.69	322.59	261.56	342.67	332.58	322.50	261.60
Degrees of freedom	5,040	5,039	5,038	5,030	5,040	5,039	5,038	5,030

Notes:
The following control variables were also included:

(1) Dummies for employment expected to rise/fall; (2) previous unemployment dummy; (3) dummy for redundancy expected; (4) dummies for gender, part-time and marital status; (5) a quadratic in age; (6) years of schooling; (7) dummies for private sector, unionization, supervisor, manual and workers' attitudes; (8) dummy for self-employment history; (9) 3 year dummies; (10) 9 industry dummies.
The equations use 11 regional unemployment rates by 5 years. For further details of these equations see Blanchflower and Oswald (1989b) and Blanchflower (1989).

The difficulty of controlling for region-specific fixed effects is circumvented in Table 5, which presents new and separate evidence for the U.S. In this case the regressions include not area unemployment but (2-digit) industry unemployment rates. Remarkably, despite this difference, and the fact that these data are from a different country, a similar wage curve is found.[11]

Table 5. *United States ISSP wage equations, 1985–87*

	(1)	(2)	(3)	(4)
Log unemployment	−0.2426	−0.7865		
	(5.26)	(3.90)		
(Log unemployment)^3		0.0583		
		(2.77)		
Unemployment			−0.0346	−0.1333
			(4.58)	(3.79)
(Unemployment)^2				0.0070
				(2.87)
Constant	6.4131	7.1411	6.2178	6.6332
	(9.48)	(9.99)	(9.28)	(9.69)
Adjusted R^2	0.3789	0.3809	0.3769	0.3791
F	38.19	37.44	37.87	37.16
Degrees of freedom	2,039	2,038	2,039	2,038

Notes: The following control variables were also included:

(1) Three marital status dummies; (2) dummies for part-time, union, male, supervisor, manual and manufacturing; (3) years of schooling; (4) quadratic in age; (5) 9 regional dummies; (6) 5 highest qualification dummies; (7) 6 city size dummies; (8) 2 year dummies. The equations use 30 industry unemployment rates by 3 years.

Table 6. *The unemployment percentage at which estimated wage curves minimize*

	Log cubic specification	Quadratic specification
(1) Great Britain		
WIRS	12%	12%
NCDS	15%	15%
BSA	13%	14%
(2) United States		
ISSP	9%	10%

Note: Based upon columns (2) and (5) of Table 2, columns (7) and (8) of Table 3, columns (2) and (6) of Table 4, and columns (2) and (4) of Table 5.

[11] This is reassuring for another reason. Data on area prices were unavailable for the British data sets, so the estimation implicitly imposes a national price deflator. The evidence of a U.S. wage curve suggests that the result is not being generated by some geographical misspecification of this type.

Figures 1 and 2 sketch the wage curves that emerge from the individual adult equations from BSA for Great Britain and ISSP for the U.S.[12] The British wage curve minimizes at 13 per cent compared with 10 per cent for the U.S. The other two British wage curves are not presented: they have the same shape.

The idea that the wage curve turns up significantly, and so takes a positive gradient, is not predicted by the earlier theoretical model and appears to go against commonsense. It may be that, because few unemployment observations occur over that range, the results there are unreliable.

The addition of long-term unemployment (as a proportion of total unemployment) to the British wage equations contributes nothing once nonlinear unemployment effects are incorporated. On its own, however, long-term unemployment is occasionally positive and significant, as predicted by Layard and Nickell (1987). See column 4 of the NCDS results in Table 3, for example. In column 9, the inclusion of a cubic term in unemployment, itself highly significant, halves the coefficient on long-term unemployment and drives it insignificant. Once regional dummies are

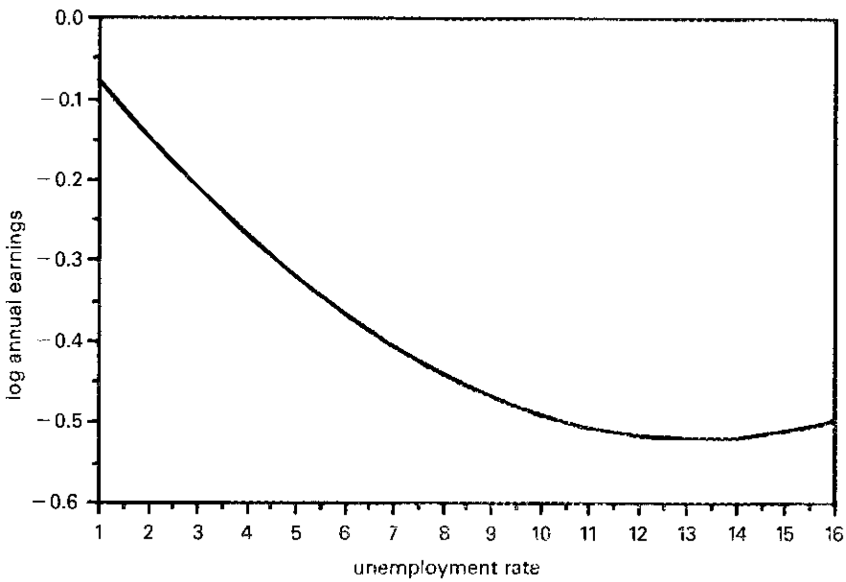

Fig. 1. U.K. wage curve.

[12] Whilst the WIRS, BSA, and NCDS data sources do not generate identical wage curves, they produce similar ones. In the figures we plot the level of unemployment and its square rather than the log formulation.

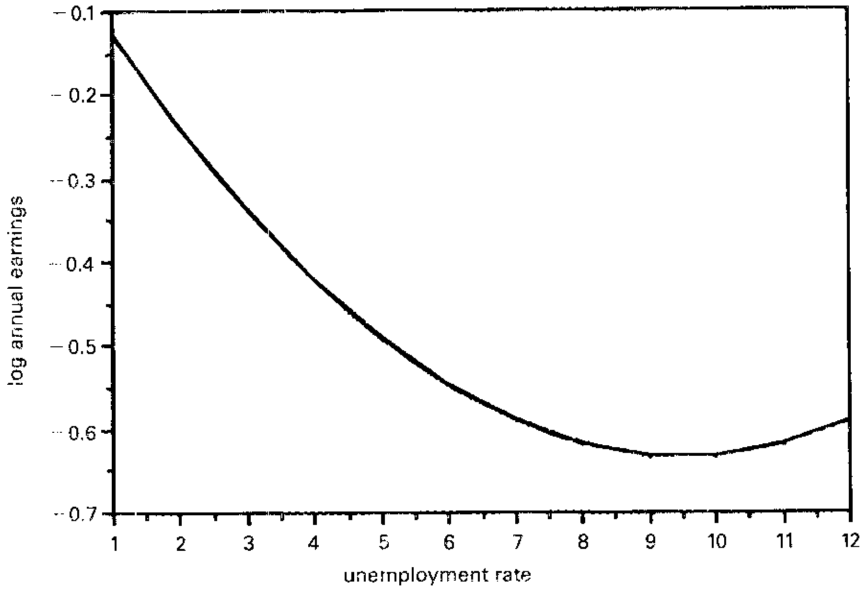

Fig. 2. U.S. wage curve.

incorporated, the long-term unemployment statistic goes down to 0.03, and the unemployment variables remain highly significant.

Across the first three data sets the statistical performance of long-term unemployment is consistently weak. Although the possibility of Type II error exists, the results suggest that long-term unemployment does not play an independent role in wage formation.

V. Conclusions

The purpose of this paper has been to estimate the shape of the wage curve linking pay to the level of unemployment. It uses British microeconomic data sets on establishments, adults and young workers, and a U.S. data set on individuals. They produce comparatively similar wage curves. These curves were estimated by fitting unrestricted polynomials and were not forced on to the data.

The paper should be seen as an attempt to identify an empirical regularity in microeconomic data on wages and unemployment in the 1980s. Various theoretical interpretations are possible, but we favour one based upon a bargaining framework.

Two major conclusions emerge from the empirical analysis. First, in both Britain and the U.S. there is a wage curve which has a negative gradient over low levels of unemployment. However, these curves become flat once sufficiently large unemployment rates are reached.[13] Second, contrary to the argument in Layard and Nickell (1987), the British evidence does not support the view that long-term unemployment[14] is an important element in the wage determination process.[15]

These findings have implications for macroeconomics. If the wage curve flattens out at moderate to high levels of unemployment, shocks to the economy over this range can produce little or no wage adjustment but substantial changes in unemployment. Wage flexibility is greatest when unemployment is low.

Appendix. Variable Definitions in the Four Data Sets

Great Britain

(1) WIRS		*Mean*
Wages	Log of weekly earnings — semi-skilled workers.	4.161
Unemployment rate	Total unemployment rate in the county, 1980. Data supplied by the Department of Employment.	7.106
Long-term unemployment	Proportion of the unemployed in the county who had been continuously unemployed for at least one year. Data supplied by the Department of Employment.	17.793
(2) BSA		
Wages	Log of annual earnings.	8.59
Unemployment rate	Unemployment rate in the Standard Region.	2.44
Long-term unemployment	Proportion of the unemployed who had been registered as unemployed and claiming benefit for at least one year.	43.22

[13] This accords well with Nickell's (1987) time-series results, despite the differences in data and methodology. The results are also consistent with Carruth and Oswald's (1987) conclusion that the cube of the log of unemployment enters a wage equation.

[14] A referee has pointed out that the lack of a long-term unemployment effect may be because the variable should be entered in a different way. This is plausible, so we do not rule out the possibility that future research will uncover positive evidence.

[15] This possibility was anticipated in Nickell's (1987) closing caveat.

(3) NCDS		
Wages	Log of weekly earnings.	4.508
Unemployment rate	Unemployment rate by county. Data supplied by the Department of Employment.	11.012
Long-term unemployment	Proportion of the unemployed in the county continuously unemployed for at least one year. Data supplied by the Department of Employment.	20.208

United States

(4) ISSP		
Wages	Log of annual earnings.	9.410
Unemployment rate	Unemployment rate by industry. Source *Employment and Earnings*, 1988, 1989 U.S. Bureau of Labour Statistics.	6.130

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