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THE BRITISH LABOUR MARKET IN DIFFERENT ECONOMIC ERAS 1857 - 1938

T J Hatton

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Centre for Economic Policy Research
6 Duke of York Street
London SW1Y 6LA

Tel: 01 930 2963

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ABSTRACT

In this paper a simple model of aggregate supply and demand for labour is developed which includes a "surprise" supply function and imposes labour market clearing. This model is estimated on British data for 1857 to 1938, an important period for the original Phillips curve estimates. For 1857-1913, the model specified in the nominal wage fits well and structural breaks are rejected but wage <u>change</u> outperforms wage <u>surprise</u>, supporting the traditional Phillips curve type view. The period 1921-38 is found to produce different results yielding a structural break from 1857-1913 in both supply and demand equations. These shifts cannot be accounted for by the introduction of national insurance though it had some effect on the labour demand curve.

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T. J. Hatton
Department of Economics
University of Essex
Wivenhoe Park
Colchester
Essex
C04 3SQ

0206 862286

SUMMARY

This study attempts to reexamine the aggregate labour market in Britain from 1857 to 1938. It uses a modern "classical" specification of labour supply model in which the proportion of workers supplied for work depends on the deviation between their local wage (and/or the prices they face) and their estimate of the economy wide average. Thus aggregate fluctuations depend on the difference between actual and expected values of these variables. The rest of the model consists of a standard labour demand curve and an assumption that the market clears each year.

Some variants of this model (particularly when prices are excluded from the labour supply equation) perform well for the period up to 1913. When the period is broken into three at 1874 and 1894 the functions change from one period to the next but these changes are not statistically significant. But when the 1857-1913 period as a whole is compared with 1921-38, significant differences do emerge. In particular, the interwar period does not support the classical interpretation as well as the prewar period.

Some efforts are made to see if the introduction of social insurance involving the right to unemployment benefit for employees and the contribution to the insurance fund by employers could account for the shift after the first World War. There is little evidence that this explains the break in supply and demand functions. Insofar as insurance did have direct effects, it appears to have been on the demand for labour rather than supply.

I

In this paper we examine some simple equilibrium models of the aggregate wage and employment rate of labour for various phases in British economic history from 1857 to 1938. Empirical work on the aggregate labour market for this period is dominated by the famous pioneering work of Phillips (1958) and Lipsey (1960). Though this has given rise to a large theoretical and empirical literature, the subject of their study - historical time series on the British labour market - has been somewhat neglected. In particular models of equilibrium unemployment which have developed from the original literature on the Phillips curve have not been generally applied to the period before 1945[1].

Studies which have concentrated on the pre-1945 and especially the pre-1914 period have followed directly in the tradition of Phillips. Proportionate changes in money wages are treated as a response to labour market disequilibrium and/or other variables[2]. The alternative interpretation of such a relation is as a labour supply function in which the supply of workers for employment (and hence the level of measured unemployment) depends on the deviation of actual from expected or perceived wages and/or prices. Such models may be based on search theory or inter-temporal labour/leisure substitution. For a proper market model one also needs a labour demand function - a feature largely ignored in the early literature on the Phillips curve. The clear separation

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[1] For recent studies along these lines on postwar data, see Batchelor and Sheriff (1979) and Beenstock and Warburton(1982).

[2] This includes the early literature which concentrated on testing the specification of and the inclusion of different additional variables in the Phillips curve. See Routh (1959), Kaldor (1959), Lipsey (1960), Lipsey and Steuer (1962). More recent studies which essentially follow the original Phillips line while testing various additional hypotheses, include Desai (1975), Gilbert (1976), and Smythe (1979).

and estimation of these functions should also shed light on the earlier debate between Keynes (1939) and Dunlop (1938) in which the two sides of the labour market were not clearly distinguished[1]. Thus, if there is market clearing, the labour demand curve will only be identified in a simultaneous equations context.

In section II we specify a basic model in the context of which a number of historical issues may be raised. The subsequent sections are devoted to the econometric tests of these hypotheses by using variants of the basic model.

In a previous paper one specific variant with a simple nominal wage surprise supply function was estimated and tested for the whole period 1857 - 1913. In that study the restrictions implied by rational expectations were tested and, on the whole, not rejected while those implied by the

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[1] Upon re-reading this debate it becomes clear that the real wage-employment relationship was not the central issue from the empirical view point as sometimes suggested. It originated with Keynes' statement in the General Theory that "in the short period falling money and rising real wages are each, for independent reasons, likely to accompany decreasing employment, labour being readier to accept cuts when employment is falling off, yet real wages inevitably rising under the same circumstances on account of the increasing marginal return to a given capital equipment when output is diminished"(1936, p.10). Thus there are two questions: (1) the cyclical relation between real and money wages and (2) the relation between real wages and employment. Since Keynes had explicitly argued that these were independent, there was no ground for inferring one from the other. Yet in their examination of the data, Dunlop (1938), Tarshis (1939) and Richardson (1939) stuck almost entirely to (1). Only in a post script to his paper did Tarshis address (2) directly, finding a strong negative relation between monthly U.S. data on man hours and hourly wages for 1931-38. Though the participants in the debate increasingly came to recognise the importance of this distinction, it was left to later writers to explore the statistical relationship between real wages and employment.

surprise supply function were rejected. Here, we take a more flexible approach and test several variants of the model. In particular, we examine the question of whether a model incorporating wage and price surprise out performs one based on wage and price change along the lines of the traditional Phillips curve literature. In the empirical work of section III we find that wage and price change tend to dominate wage and price surprise in the equations. A simple specification including only nominal wage change cannot be rejected against a variety of more general specifications[1].

To many observers, the period between the 1850s, and the first World War is far from homogeneous. The period is sometimes divided by historians into three phases which might be termed the "Mid-Victorian Boom", "the Great Depression" and "Edwardian Retardation". Although there is doubt about these as distinct phases in economic growth[2], they are most clearly marked out by different trends in wages and prices. On this criterion, the periods are traditionally marked out at 1873 and 1896 and, according to Hobsbawm, "the development of the labour market falls into much the same periods" (1964 p. 318). More recently Tarling and Wilkinson have associated this with changes in the operation of the labour market, stressing the importance of social and political forces in wage determination. Referring to the work of Phillips and others, they argued that "[e]conomists are used to regarding the social and

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[1] This is consistent with our earlier finding that labour supply was influenced by the systematic as well as surprise element in the nominal wage. See Hatton (1983b) p. 16.

[2] For surveys of the first two of these periods, see Church (1975) and Saul (1969).

political forces, as peripheral to the process of the determination of such 'economic' variables: an understanding of the events of the 50 years or so leading up the 1920s shows how empty and dangerous is this misconception" (1982 p. 23)[1]. In section IV we attempt to come to grips with one aspect of this argument by estimating our equations for the three sub periods and testing for structural breaks in the fundamental relationships[2]. Though there are substantial changes in some of the coefficients, we find that these do not amount to significant structural breaks.

Many observers see the major break as occurring across the divide of the first World War with the emergence of labour market disequilibrium leading to mass unemployment in the interwar period. One problem with such comparisons is that the unemployment data available for the interwar years which derives from the operation of the labour exchanges and the administration of the National Insurance system is totally different in scope and coverage from the pre-war trade union data. Even so, the comparison of pre and post war labour markets using these data may not

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[1] This conclusion was reached from a study of the growth of trade unionism and developments in the nature and institutions of collective bargaining up to the first World War. However, it is not a view which is universally shared. Pollard, for instance, in his survey of 1870-1914 found that "Where the market itself changed rapidly or violently, the 'market component' of the wage determinant was much greater than any possible organisational component and the unions themselves frequently appear as little more than playthings of the market situation"(1965 p. 111).

[2] This is, of course, not a test for the inclusion of specific social or political variables. Such forces are in any case, likely to be difficult to measure especially before 1914. Time series data on trades union membership and the number of strikes began only in the 1890's. When Hines estimates his function with trade union density explaining wage change, he found the period 1843-1913 did not support this causal relation (1964, p. 234). If such "non-economic" forces were becoming more important, one might nevertheless expect a simple market model to perform progressively less well and perhaps exhibit structural breaks as the market process breaks down.

be as misleading as sometimes supposed[1]. Since little else is available, it is essentially these series which have been used both in econometric work by Phillips and his followers and by both contemporaries and historians. Thus, conceptions of changes in the labour market have been largely conditioned by the use of these series and hence, quite apart from differences in representativeness, it is important to see if the two periods conform to a common structure. In section V we estimate the same models for 1921-38 and find that, on the whole, the models perform poorly relative to the period 1857-1913 taken as a whole. Furthermore, there are substantial structural breaks even when intercept shifts are admitted.

One might well suspect that these findings are not simply an artefact of data but a reflection that the interwar labour market did operate differently as compared with pre-war. One obvious feature is the extension of national collective bargaining (which was virtually absent before the War) via national representative bodies of unions and employers, and in unorganised industries, wage boards. This, coupled with structural readjustment imposed on the economy in the 1920s and 1930s are likely to have driven the economy off its supply and perhaps even demand curve for labour leading to labour market disequilibrium.

An alternative view inspired by Benjamin and Kochin (1979) is that the introduction of national insurance benefits and contributions served to mask the operation of the equilibrium labour market by shifting supply

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1 Of the trade union series Garside concludes that, despite its narrow base the series is "fairly representative of conditions generally over a period of years and not only in the trades directly represented". (1980, p. 23). For discussion of the comparability of the trade union and insurance series, see Hilton (1923), pp. 182-3 and Appendix 1 pp. 190-1, Feinstein (1972), p. 225 and Garside (1980), p. 22.

and demand curves for labour[1]. The econometric results of section V suggest that including these variables does little to repair the structural break. Finally, since it is sometimes suggested that the availability of unemployment benefits led to systematic dishoarding of labour manifested in temporary layoffs, the benefit to wage ratio was included in the demand function where it appears to have an effect despite its poor showing in the supply function.

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[1] The insurance system was introduced in 1911 but only became general after the first World War. Thus Benjamin and Kochin argued that "in 1927-9 and 1936-8 unemployment would have been at or near normal levels, if the dole had been no more generous than when it was first set up in 1913" (1979, p. 464). This view has attracted much criticism largely on empirical grounds (Collins (1982), Cross (1982) Metcalf et al. (1982), Ormerod and Worswick (1982), Broadberry (1983), Hatton (1983a)). It has also been emphasised that the model is inadequate as a model of supply and demand for labour (Hatton, (1980), pp. 5-12).

II

The model of the aggregate labour market is set up in terms of the employment rate ER which is related to the percentage unemployed (U) by ER = 100 - U. From the supply side this measures the rate at which labour is supplied for a given labour force. Its logarithm is decomposed into permanent and transitory components M and n respectively

$$\ln ER_t = \ln M_t + \ln n_t \qquad (1)$$

Since no time series measure of hours is available, this dimension of labour supply is not modelled directly.

The labour supply function used is based on that derived by Lucas (1973) suitably adapted to the labour market. The logarithm of the transitory component of labour supply for the individual micro labour market, z, is

$$\ln n_t(z) = \gamma_1\left[\ln W_t(z) - E(\ln W_t/I_t(z))\right] + \gamma_2\left[\ln C_t(z) - E(\ln C_t/I_t(z))\right] + \sum_{i=1}^{k} \gamma_{2+i} \ln n_{t-i} \qquad (2)$$

where $W_t(z)$ and $C_t(z)$ are respectively the nominal wage rate and the index of consumption prices in market z at time t and W_t and C_t are the economy wide (geometric) means of these variables. The latter are not known at t but expectations are formed based on the information set $I_t(z)$ which includes: $W_t(z)$ and $C_t(z)$; their estimated variances across the individual markets ψ_w^2 and ψ_c^2; W_{t-i} and C_{t-i} for $i > 0$; and the estimated variances of time series forecasts of W_t and C_t based on their past values σ_w^2 and σ_c^2. If z is taken to index markets by percentage deviations from the aggregate then the local wage and price are

$$\ln W_t(z) = \ln W_t + z_w \qquad (3)$$

$$\ln C_t(z) = \ln C_t + z_c \qquad (4)$$

The expectations of aggregate wage and consumption price levels at t are formed by

$$E(\ln W_t | \ln W_t(z), \ln W_t^*) = (1-\theta_w)\ln W_t(z) - \theta_w \ln W_t^* \qquad (5)$$

$$E(\ln C_t | \ln C_t(z), \ln C_t^*) = (1-\theta_c)\ln C_t(z) - \theta_c \ln C_t^* \qquad (6)$$

where W_t^* and C_t^* are the time series forecasts of aggregate variables referred to above. This estimate and the local wage or price are weighted according to their relative variances to form the expectation of the current economy wide variables. Thus $\theta_w = \psi_w^2/(\sigma_w^2 + \psi_w^2)$ and $\theta_c = \psi_c^2/(\sigma_c^2 + \psi_c^2)$.

Substituting (5) and (6) into (2) gives

$$\ln n_t(z) = \gamma_1\left[\ln W_t(z) - (1-\theta_w)\ln W_t(z) + \theta_w \ln W_t^*\right] + \gamma_2\left[\ln C_t(z)-(1-\theta_c)\ln C_t(z) + \theta_c \ln C^*\right] + \sum_{i=1}^{k} \gamma_{2+i} \ln n_{t-i} \qquad (7)$$

Aggregating over markets and substituting into (1) gives the aggregate labour supply equation

$$\ln ER^s = \alpha_o + \alpha_1(\ln W_t - \ln W_t^*) + \alpha_2(\ln C_t - \ln C_t^*) + \sum_{i=1}^{k} \alpha_{2+i} \ln ER_{t-i} \qquad (8)$$

where

$$\alpha_o = (1 - \sum_{i=1}^{k} \gamma_{2+i})M \qquad \alpha_1 = \gamma_1\theta_w, \quad \alpha_2 = \gamma_2\theta_c, \quad \alpha_i = \gamma_i (i > 2).$$

If labour supply is determined by the real wage, then $\gamma_1 = -\gamma_2$ but, even so, α_1 and α_2 will not, in general, be equal. If cross sectional variance in prices were zero, then $\alpha_2 = \theta_c = 0$ and only the nominal wage will enter the equation. Similarly, differences in α_1 and α_2 over time may be due to changes in either θ or γ. One special case of this model is where $\ln W_t^* = \ln W_{t-1}$ and $\ln C_t^* = \ln C_{t-1}$. This formulation is the simple Phillips curve in which unemployment is related to the rate of change of wages and the cost of living.

On the labour demand side, it is assumed that the short run marginal productivity condition for profit maximisation is continuously met. Competitive firms in market z are concerned only with maximimising profits, given their capital stock and the prices and wages facing them. The underlying production function is taken to be C.E.S. of the form

$$Q_t = Ae^{\tau t} \left[\bar{\Lambda}E_t^{-\eta} + (1-\bar{\Lambda})K_t^{-\eta}\right]^{\frac{\mu}{\eta}} \qquad (9)$$

where Q is output, E employment and K the capital stock. τ is the parameter representing the rate of disembodied technical progress, $\frac{1}{1+\eta}$ gives the elasticity of factor substitution and μ the degree of returns to scale. Taking the capital stock as predetermined, the first order condition for profit maximisation gives the labour demand function

$$E_t^D = A^{\frac{\eta}{(1+\eta)\mu}} (\mu\bar{\Lambda})^{\frac{1}{1+\eta}} e^{\frac{\eta\tau}{(1+\eta)\mu} t} \left(\frac{W}{P}\right)^{\frac{-1}{1+\eta}} Q_t^{\frac{1+\eta/\mu}{1+\eta}} \qquad (10)$$

where P is an index of product prices and the maximised value of output has been substituted back into the expression by solving out for K_t to give an expression linear in logs. To convert this to an employment rate, the trend in the growth of the labour force is represented as

$$L_t = L_o e^{\phi t} \tag{11}$$

Dividing (11) into (10) and taking logs gives the employment rate demand equation for $\ln(E/L) = \ln ER$.

$$\ln ER^D = \beta_o + \beta_1 \ln(\frac{W}{P})_t + \beta_2 \ln Q_t + \beta_3 t \tag{12}$$

where $\beta_1 = \frac{1}{1+\eta}$, $\beta_2 = \frac{1+\eta/\mu}{1+\eta}$, $\beta_3 = \frac{\eta\tau}{(1+\eta)\mu} - \phi$

We take the two equation system formed by (8) and (12) as the basis for estimation in the following sections.

III

In this section some versions of the model set out previously are estimated on annual time series for 1857 to 1913. The data is taken from Feinstein (1972). The wage index W is the Bowley-Wood index of average weekly wage rates, and C the cost of living index for working class families (T140). For U the series for trade union unemployment is used (T125) and for Q the GDP output index (T18). P is the GDP deflator and since this is unavailable before 1870 the index for 1855-1869 was estimated from the coefficients of a regression of the GDP deflator on four other price series for 1870-1913.[1]

In order to obtain an empirical specification of equation (8) it is necessary to provide suitable forecasts of wages and consumption prices W* and C*. Since these depend only on lagged values of the aggregate variables they are obatined from a first stage regression of lagged values of all the variables in the model W,C,P,Q, a constant and time trend. This ignores the restrictions on the expectation generating process if expectations are fully rational, i.e. that they are generated according to the same structure as the model itself. However unless a full econometric

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[1] The four series were the cost of living index C already mentioned, indices for import prices (P_M) and export prices (P_X) given by Imlah (1959) Table 8, pp94-98 and the Rousseaux index of the prices of principal industrial products P_i given in Mitchell and Deane (1962), p. 471. The equation estimated was logarithmic and included a time trend. This gave the following coefficients which were used as weights in generating the composite price series for 1855-1869

$$\ln P = 1.8 + 0.0016\text{TIME} - 0.261 \ln P_M + 0.461 \ln P_X + 0.441 \ln C - 0.051 \ln P_i$$

model is specified, generating a value for W^* involves making unrestricted forecasts of P_t and Q_t.

Following some experimentation the lag length in (8) is restricted to $k = 2$ and the forecasting equations are also restricted to two lags. Having obtained values for W^* and C^* the equations were estimated using two-stage least squares. Clearly, all the current values of variables in the model are endogenous or would be in a full macro model and should therefore be instrumented. The set of instruments must include all the lagged variables used to generate W^* and C^*, but it must also include some current dated variables since otherwise $\ln \hat{W}_t = \ln W^*_t$ where $\ln \hat{W}_t$ is the instrumental variable for $\ln W_t$. It is difficult enough to obtain exogenous current variables and the choice is narrowed by the availability of data. In the event, the volume indices of exports and fixed investment were used on the grounds that these have been frequently seen as determining economic fluctuations in the period (see Aldcroft and Fearon (eds).1972) together with import and export price indices on the grounds that these are, at least in part, determined externally.

The procedure followed was first to obtain values of $\ln \hat{W}$, $\ln \hat{C}$, $\ln W^*$ and $\ln C^*$ as the fitted values from a first stage regression. Since the first two include current dated information while the second two do not, the surprise or innovation in the wage and cost of living (purged of the effects of movements along the demand curve) are $\ln(\hat{W}_t/W^*_t)$ and $\ln(\hat{C}_t/C^*_t)$. The systemmatic parts of proportionate wage and price change are $\ln(W^*_t/W_{t-1})$ and $\ln(C^*_t/C_{t-1})$. An alternative and much simpler assumption consistent with an adaptive expectations scheme is to simply use the actual changes:

$$\ln(\hat{W}_t/W_{t-1}) = \ln(\hat{W}_t/W^*_t) + \ln(W^*_t/W_{t-1}),\ \ln(\hat{C}_t/C_{t-1}) = \ln(\hat{C}_t/C^*_t)+\ln(C^*_t/C_{t-1}).$$

In testing various restricted forms of the model against unrestricted forms the likelihood ratio test is used. The F test was preferred to the χ^2 since it appeared to give slightly narrower confidence intervals. Before moving to structural estimates of the supply equation this test was performed for the regressions generating the expected variables against those in generating the instrumental variables. For both ln W and ln C the inclusion of the current variables cannot be rejected even at the 1% level in either case so that the innovation in each of these variables reflects the significant influence of current macroeconomic shocks[1].

The results for the supply function are given in Table 1 where the first two equations compare assumptions about expectations formation. In both cases the variables in the wage and cost of living terms are just on the border of 5% significance but the latter consistently gives the wrong sign. In other respects the equations appear to be well specified and give a good fit without serial correlation of first or 4th order. The two equations were tested against a more general version with all four terms and the inclusion of the equation 1 variables was easily rejected while that for the inclusion of the equation 2 variables was on the borderline of 5% significance[2].

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1 The computed F values were 24.344 and 8.975 compared with the critical value at 1% of 3.82.

2 The computed F values were 0.065 and 3.043 compared with the critical value at 5% of 3.19.

Table 1

	C	$\ln(\hat{W}_t/W^*_t)$	$\ln(\hat{C}_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS.	D.W./B.P.
(1)	2.6957	0.4685	0.2866	0.8751	−0.4663	0.6017	2.0321
	(0.4337)	(0.2276)	(0.1418)	(0.1055)	(0.1056)	0.0129	0.7915
	C	$\ln(\hat{W}_t/W_{t-1})$	$\ln(\hat{C}_t/C_{t-1})$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS.	D.W./B.P.
(2)	2.8234	0.3847	0.2366	0.6210	−0.2045	0.6318	0.6318
	(0.4211)	(0.1674)	(0.1214)	(0.1174)	(0.1127)	0.0119	2.0596
	C	$\ln(\hat{W}_t/W_{t-1})$	$\ln(\hat{W}_t/W^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS.	D.W./B.P.
(3)	2.8496	0.4513	0.2790	0.6693	−0.2948	0.6132	1.9702
	(0.4322)	(0.1883)	(0.2636)	(0.1348)	(0.1263)	0.0125	1.1052
	C	$\ln\hat{W}_t$	$\ln W_{t-1}$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	D.W./B.P.
(4)	2.8954	0.5943	−0.5933	0.6042	−0.2408	0.6049	1.8436
	(0.4382)	(0.1338)	(0.1334)	(0.1215)	(0.1167)	0.0128	1.2659
	C	$\ln(\hat{W}_t/W_{t-1})$		$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	D.W./B.P.
(5)	2.8981	0.5937		0.6044	−0.2408	0.6123	1.8443
	(0.4303)	(0.1319)		(0.1202)	(0.1156)	0.0128	1.2673
	C	$\ln\hat{W}_t$	$\ln P_t$	$\ln\hat{Q}_t$	t	$\bar{R}^2$/RSS	D.W./B.P.
(6)	1.0266	−0.1612	0.1371	1.0215	0.0176	0.6306	1.3007
	(0.3585)	(0.0632)	(0.0432)	(0.1237)	(0.0021)	0.0119	9.4249
	C	$\ln(\hat{W}_t/\hat{P}_t)$		$\ln\hat{Q}_t$	t	$\bar{R}^2$/RSS	D.W./B.P.
(7)	1.0242	−0.1419		0.9925	−0.0171	0.6364	1.2832
	(0.3556)	(0.0412)		(0.1001)	(0.0017)	0.0120	9.5680

Standard errors in parentheses; RSS = sum of squared residuals (reported below $\bar{R}^2$); BP = Box-Pierce statistic for 4th order serial correlation (reported below the Durbin Watson statistic).

In view of the wrong sign on the price terms and their borderline significance these were dropped and the two wage variables compared in equation 3. Judging by the t value, this clearly suggests that wage change dominated wage surprise as an explanatory variable. Furthermore when the surprise term is excluded the t value on the wage change term rises to 4.5 giving strong support for a Phillips cruve type formulation. This is reinforced when another restriction is tested: that the signs on current and lagged wage should be equal and opposite. As is shown in equation 4 not only is the restriction not rejected but the variables take coefficients which are virtually the same numerically. A similar result was obtained when this restriction was tested for the wage surprise.[1]

As a final test the most restricted model which just included wage change was tested against the most unrestricted model which included the levels of current, lagged and expected wages and consumer prices. The joint significance of the additional variables is easily rejected at the 5% level.[2] Thus although the significance tests are not very decisive between different variants there is strong support for a simple model relating the employment rate to the proportionate change in wage rates and its own lagged values. The long run values for the unemployment rate when wage change is set to zero is equivalent to almost exactly 5% unemployment compared with the mean of the series of 4.45.

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[1] The computed F value was 0.409 compared with the critical 5% value of 4.04.

[2] The computed value was 1.48 compared with the critical 5% value of 2.49.

The final equations in Table 1 are the labour demand function the instrumental variables $\hat{P}_t$ and $\hat{Q}_t$ were generated from first stage regressions on the same set of variables as for $\hat{W}_t$ and $\hat{C}_t$. One obvious restriction to test is that the wage and price terms give equal and opposite signs. From equations 6 and 7 it can be seen that this restriction does not come close to being rejected. Each of the terms is individually significant and t values of over 10 are obtained on output and time. One problem is that the Durbin-Watson statistic is near to the lower bound (though not below it). It was thought that some simple dynamics should be added by including the lagged dependent variable but this took a very small t value and hence was omitted from the equation.

If the coeffficients are interpreted in terms of the model of Section II then the real wage term implies an elasticity of factor substitution of 0.14 which is a little on the low side. The coefficient on output is not significantly different from 1 and the point estimate of the returns to scale parameter μ is 1.01. The time trend gives the expected sign and implies that with the real wage and output held constant, the employment rate would fall by 1.7% per annum. With the labour force growing at about 1% ($\phi = 0.01$) this implies a value of τ, the technical progress parameter of 0.03 which is slightly on the high side.

IV

In this section we turn to examining changes in the structure of the labour market equations in the period up to the first world war. Following the earlier discussion of wage and price trends the period is split into three sub periods: the mid Victorian Boom, 1857-1874; the Great Depression, 1875-1894; the Edwardian Retardation 1895-1913. This periodisation slightly attenuates the great depression but gives three periods more or less equal in length. The annual averages for the unemployment rate and wage and price change over these periods were as follows

	U(%)	$\dot{W}$(%)	$\dot{C}$(%)
1857 - 1874	3.60	1.30	0.05
1875 - 1894	5.49	-0.17	-1.54
1895 - 1913	4.15	0.67	0.98
1857 - 1913	4.45	0.58	-0.20

This clearly shows that the sharp differences in wage and price trends are captured by the periodisation chosen.

Given these differences, one might suspect that first stage regression generating expected wage and consumer price would be subject to structural changes. Each of the first stage regressions was run for each sub period and the test for structural breaks performed against the period as a whole. These tests massively rejected the null hypothesis of no structural break[1].

The same test was also applied to the first stage regressions used to generate the instrumental variables for $\ln W_t$, $\ln C_t$, $\ln P_t$, $\ln Q_t$. The hypothesis of no structural break was overwhelmingly rejected in each case. Thus each of the sub period structural equations was estimated using

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[1] The computed F values were for $\ln W^*$ 27.881 and for $\ln C^*$ 24.598. The 1% critical value is 3.82.

first stage regressions estimated within the period only. A selection of the results is given in Table 2. In order to perform the tests for structural breaks across these periods, a whole period regression was run for each equation using the first stage regressions for each sub-period pieced together. Thus since the first stage estimates are the same in each case, the structural break refers to the second stage estimate only.

Altogether five variants of the supply equations were estimated and the same five variants run again without consumer prices, making ten in all. In none of these cases could the null hypothesis of no structural break be rejected at the 5% level[1]. This stands in sharp contrast to the forecasting equations though the result is not changed if the first stage regressions are taken for the period as a whole. As the first panel of Table 2 illustrates, this result is due not so much to the stability of the coefficients as their lack of significance for the individual subperiods. In these equations the perverse sign on the surprise in the cost of living arise from a dominant positive coefficient for the Great Depression outweighing insignificant negative coefficients for the other two periods.

The same within period tests between different specifications as were performed for the whole period were repeated for each sub-period. On the whole the results were rather similar as for the whole period. The most naive version representing the simple Phillips curve is given for each sub-period in the middle panel of Table 2. When this is tested for each sub-period against the most general model (as before including ln W_t,

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1 For the two variants of the supply equation given in the top two panels of Table 2, the computed F values were 1.122 and 0.524 compared with 5% critical values of 2.07 and 2.16.

Table 2

Eq.No/ Period	C	$\ln(W_t/W^*_t)$	$\ln(C_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
1	3.7496	0.5649	-0.1081	0.7129	-0.5338	0.2868	2.0122
(1857-74)	(1.1609)	(0.6898)	(0.3908)	(0.2338)	(0.2457)	0.0045	4.6760
2	2.7497	0.9517	0.5997	0.8622	-0.4666	0.6459	2.3608
(1875-94)	(0.6588)	(0.6091)	(0.2475)	(0.1671)	(0.1626)	0.0047	4.1692
3	2.6599	1.7746	-0.1116	0.8695	-0.4525	0.4679	1.6980
(1895-1913)	(0.8263)	(1.0730)	(0.4302)	(0.2156)	(0.2018)	0.0024	3.0225

	C	$\ln(W_t/W_{t-1})$		$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
4	3.8756	0.4375		0.4733	-0.3230	0.5442	2.3596
(1857-74)	(0.9291)	(0.1603)		(0.2065)	(0.2110)	0.0031	3.4457
5	2.5137	0.6282		0.6886	-0.2411	0.5081	1.8684
(1875-94)	(0.7872)	(0.3437)		(0.2186)	(0.2279)	0.0070	3.2371
6	2.8592	0.1837		0.7677	-0.3947	0.4092	1.8294
(1895-1913)	(0.9083)	(0.2383)		(0.2628)	(0.2254)	(0.0028)	2.7694

	C	$\ln(W_t/P_t)$		$\ln Q_{t-1}$	t	$\bar{R}^2$/RSS	DW/BP
7	0.9226	-0.1354		1.0229	-0.0179	0.5876	1.9897
(1857-74)	(0.8364)	(0.1305)		(0.2312)	(0.0042)	0.0028	4.9346
8	0.9777	0.1147		1.0207	-0.0189	0.7979	1.6822
(1875-94)	(0.4121)	(0.2009)		(0.1182)	(0.0029)	0.0029	1.3318
9	1.5541	-0.2621		0.8506	-0.0151	0.5998	1.1247
(1895-1913)	(0.6820)	(0.4262)		(0.1994)	(0.0041)	0.0019	4.6983

Standard errors in parentheses; RSS = sum of squared residuals (reported below $\bar{R}^2$); B.P. = Box-Pierce statistic for 4th order serial correlation (reported below the Durbin-Watson statistic).

$\ln W_{t-1}$, $\ln W^*_t$, $\ln C_t$, $\ln C_{t-1}$, $\ln C^*_t$) the null hypothesis is easily rejected for the Great Depression but does not come close to rejection for the other two period[1]. This clearly reflects the stronger role of consumer prices since, in this period, in all but one of the equations, the inclusion of price terms cannot be rejected at the 5% level.

The results for the demand equation in the bottom panel follow a similar pattern. This time the real wage term gives the expected sign in the first and third period, and the opposite sign in the second but all three are insignificant. Output and the time trend remain strongly significant throughout and although the coefficients tend to drop in the final period. The null hypothesis of no structural break cannot be rejected[2].

It appears that for both supply and demand equations, the well determined coefficients are relatively stable as between periods while the poorly determined ones are not. Even though this is just what might have been expected, the wage and price terms give disappointingly low levels of significance. It must be remembered, however, that wage and price trends were used as the criterion for choosing the period, thus most of the variation comes between periods rather than within periods. The results indicate that examining annual time series for periods as short as twenty years may be unreliable as a guide to identifying the underlying structure. In part this may be due to the poor quality of the data for this period. In any event, it seems that more information is required either through more observations or richer data before sharp distinctions can be made either within or between periods. With this qualification in mind, we move on to examining results for the interwar period.

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[1] The computed values of F for this test for the three periods were respectively 0.459, 11.868 and 0.685. 5% critical values are 3.48, 3.20 and 3.33 respectively.

[2] The computed value of F was 0.691 compared with the critical 5% value of 2.16.

V

In order to compare the interwar period with the prewar periods, exactly the same set of equations was run on inter-war data for 1921-38. All the variables appearing in the structural equations are from the same source and are continued as index numbers from the pre-war period. As before, the first-stage estimates were estimated separately for this period alone.

The first equation gives results which are similar to the pre-war period with both wage and price surprise entering positively and with a similar lag structure. But the coefficients on wage and price terms are insignificant. Perhaps this suggests a different regime in the inter-war years but in view of the poor results for pre 1913 sub periods, one should not press this. The next equation - the standard Phillips curve - now gives the wrong sign on the wage change term and the third equation for labour demand gives a positive sign on the real wage. To some extent these results resemble those of the other 'perverse' period 1875-94[1].

These results look quite different from the pre 1913 period taken as a whole but, as has been emphasised, it is whether there are significant differences that matters. In order to perform tests of structural change, we take the pre 1914 period as a single period for both the first and second stage regressions. The tests were then performed using these and the set from which the first three equations in table 3 were taken by estimating a third equation for the whole period 1857-1913, 1921-38 taking

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[1] One particularly perverse result was obtained when the supply function was estimated with wage and price change terms. Both entered with the wrong sign and t values in excess of 4!

the first stage regressions as two segments. As was pointed out earlier, the comparison of overall levels of unemployment is open to question because of data differences, so the whole period regressions include an intercept shift dummy between the two periods (D = 1 for 1921-38)[1]. Hence the F tests performed are only for the joint significance of shifts in the slope coefficients.

The whole period regressions are given in the bottom part of Table 3. The surprise supply function is similar to that for the interwar period alone while the Phillips curve version now gives a coefficient on wage change close to zero. The effect of the dummy in each case is to shift the employment rate down significantly and by about 3-4%. In the demand function the real wage term is now small but negative and the dummy shifts the employment rate up (as might have been expected) by about 10%. Not surprisingly, out of the ten variants of the labour supply function, eight indicate structural breaks at the 5% level, seven of these at the 1% level. Both the supply equations in Table 3 and the demand equations exhibit structural breaks in the slope coefficients at the 1% level.

One possible reason for shifts in intercept and, perhaps, slope coefficients between the two periods is the impact of the unemployment insurance system. This may have affected the labour supply through the effects on search and leisure choice and on the demand for labour via the effect of the employers' contribution to both the unemployment insurance and health insurance systems. The data for benefits was taken as a weighted average of different rates as calculated in Hatton (1980)

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[1] One might also attribute such a shift to the changing structure of labour supply due to the war or, in the demand function, the retardation of technical progress.

Table 3

1921-1938

	C						
	C	$\ln(\hat{W}_t/W^*_t)$	$\ln(\hat{C}_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
1.	2.5580 (0.7710)	0.8768 (2.3727)	0.5715 (0.8651)	0.6321 (0.2163)	−0.2009 (0.2000)	0.3664 0.0079	1.4290 1.4928
	C	$\ln(\hat{W}_t/W_{t-1})$		$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
2.	2.7034 (0.7784)	−0.2067 (0.1273)		0.9278 (0.2833)	−0.5296 (0.2850)	0.3627 0.0854	1.3512 4.0279
	C	$\ln(\hat{W}_t/\hat{P}_t)$		$\ln Q_t$	t	$\bar{R}$ /RSS	DW/BP
3.	2.4137 (0.1628)	0.1382 (0.0915)		0.7457 (0.0559)	−0.0184 (0.0015)	0.9233 0.0010	1.9826 1.6181

1857-1938

	C	D	$\ln\hat{W}/W^*_t$	$\ln(\hat{C}_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}$ /RSS	DW/BP
4.	2.6217 (0.3701)	−0.0356 (0.0065)	0.4682 (0.2549)	0.3459 (0.1510)	0.7866 (0.0950)	−0.3615 (0.0919)	0.7670 0.2191	2.0061 2.8357
	C	D	$\ln(\hat{W}_t/W_t)$		$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
5.	2.6318 (0.4187)	−0.0366 (0.0077)	−0.0401 (0.0789)		0.8190 (0.1249)	−0.3962 (0.1243)	0.7025 0.0284	2.0110 3.8984
	C	D	$\ln(\hat{W}_t/\hat{P}_t)$		$\ln(Q_t)$	t	$\bar{R}$ /RSS	DW/BP
6.	2.1000 (0.2705)	0.1013 (0.0219)	−0.0738 (0.0465)		0.5942 (0.0766)	−0.0123 (0.0014)	0.7583 0.0231	0.9022 33.7838

Standard errors in parentheses; RSS = the sum of squared residuals (reported below $\bar{R}^2$); B.P. = Box-Pierce statistic for 4th order serial correlation (reported below the Durbin-Watson statistic).

and expressed as a ratio with average weekly wages. This ratio, R, was adjusted with the following way to give $\hat{R} = \exp(\ln R - \ln(W/\hat{W}))$ which was entered into the supply equation as the absolute ratio (since before 1914 it is taken as zero). Since the dependent variable is the log of the employment rate, the predicted sign on this variable should be negative.

On the demand side, following Harrison and Hart (1982) the average ratio of contribution to wages was obtained from data in Chapman (1953). This ratio plus one (S) is the factor by which labour costs are raised by the insurance contributions. This ratio was also adjusted to give $\ln \hat{S} = (\ln S - \ln W/\hat{W})$ which was then entered logarithmically into the demand equation. This variable should therefore enter the equation with the same coefficient as the real wage variables when first entered into the equations for 1921-38 and then for the whole period.

The results for the same set of equations are given in Table 4. The inclusion of $\hat{R}$ does not markedly change either version of the supply equation and the variable itself gives a perverse sign. In the demand equation $\ln \hat{S}$ gives a plausible coefficient, unlike the real wage but is not significant and the hypothesis that the coefficients are equal cannot be rejected. When the whole period is taken, the coefficients on the supply functions are somewhat similar and $\hat{R}$ still gives positive signs. However, with the change in sign of the real wage in the demand equation, the coefficient on $\ln \hat{S}$ becomes significantly negative but implausibly large. The test for the equality of coefficients is now rejected at the 1% level[1].

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[1] Values of F were for 1921-38, 0.237 and for 1857-1913, 1921-38, 20.301 compared with critical 5% values of 4.60 and 4.00.

Table 4

1921-1938

	C	$\hat{R}$	$\ln(\hat{W}_t/W^*_t)$	$\ln(\hat{C}_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
1.	2.0620	0.1482	0.7070	0.4693	0.6465	-0.1187	0.3438	1.2369
	(0.9551)	(0.1647)	(2.2576)	(0.8736)	(0.2207)	(0.2231)	0.0075	2.2852

	C	$\hat{R}$	$\ln(\hat{W}_t/W_{t-1})$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
2.	1.5330	0.3906	-0.3786	1.2119	-0.5862	0.5183	1.8803
	(0.8419)	(0.1670)	(0.1326)	(0.2743)	(0.2490)	0.0060	1.286

	C	$\hat{S}$	$\ln(\hat{W}_t/\hat{P}_t)$	$\ln Q_t$	t	$\bar{R}^2$/RSS	DW/BP
3.	2.4178	-0.2260	0.1307	0.7404	-0.0180	0.9054	2.2180
	(0.1815)	(0.7043)	(0.1126)	(0.0638)	(0.0020)	0.0012	2.7902

1857-1913, 1921-1938

	C	D	$\hat{R}$	$\ln(\hat{W}_t/W^*_t)$	$\ln(\hat{C}_t/C^*_t)$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
4.	2.5119	-0.0649	0.0835	0.4702	0.3349	0.7900	-0.3409	0.7667	1.8711
	(0.3947)	(0.0370)	(0.1039)	(0.2548)	(0.1492)	(0.0951)	(0.0955)	0.0216	2.4126

	C	D	$\hat{R}$	$\ln(\hat{W}_t/W_{t-1})$	$\ln ER_{t-1}$	$\ln ER_{t-2}$	$\bar{R}^2$/RSS	DW/BP
5.	2.4663	-0.0849	0.1345	-0.0850	0.8610	-0.4017	0.7025	1.9271
	(0.4502)	(0.0488)	(0.1343)	(0.0905)	(0.1316)	(0.1243)	0.0280	3.3333

	C	D	$\hat{S}$	$\ln(\hat{W}_t/\hat{P}_t)$	$\ln Q_t$	t	$\bar{R}^2$/RSS	DW/BP
6.	1.8276	0.2006	-3.3977	-0.1055	0.7680	-0.0134	0.8144	1.5051
	(0.2442)	(0.0283)	(0.7146)	(0.0413)	(0.0690)	(0.0013)	0.0175	8.2734

The inclusion of these additional variables does little to improve the performance of the supply functions. Out of the ten versions, seven still exhibit structural breaks between the two periods, 6 at the 5% level[1]. All four variants of the demand function gave structural breaks at the 5% level, three at the 1% level.

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[1] Values of F for the two versions reported in Table 4 are 0.747 and 7.947 compared with critical 5% values of 2.37 and 2.53.

VI

Let us summarise our conclusions so far. Over the prewar period up to 1914, the supply and demand equations seem to fit quite well and give plausible coefficients with the exception of the cost of living index which gives the wrong sign in the supply equation. In particular, a simple Phillips curve with labour supply simply depending on actual wage change outperforms most other models. When the period is broken down into three sub-periods, the coefficients are rather unstable though, in general, the structural breaks are not significant. When the period is extended to 1938, the structural breaks became much more significant and it appears that in most cases the observations were not generated by the same model. Including the effects of the national insurance acts from 1921 onwards does not alter this conclusion.

This sheds little light on the extent to which unemployment insurance raised the interwar level of unemployment compared with pre-war, since the underlying model does not seem to be stable between the two periods. If we want to address this issue, we can only do so if a stable model can be found. On the supply side one simple model which did not give a structural break and gives plausible coefficients is the version which included just the nominal wage surprise. The result from this, for 1857-1913, 1921-1938 was as follows

$$\underset{}{\ln ER_t} = \underset{(0.4041)}{2.5119} - \underset{(0.0381)}{0.0649D} + \underset{(0.1069)}{0.0835R_t} + \underset{(0.2091)}{0.8532\ \ln(W_t/W_t^*)}$$

$$+ \underset{(0.0978)}{0.7900\ \ln ER_{t-1}} - \underset{(0.9830)}{0.3409\ \ln ER_{t-2}}$$

$$\bar{R}^2 = 0.7530 \quad RSS = 0.0232 \quad DW = 1.9911 \quad BP4 = 1.8891$$

This equation fits well, is free of serial correlation and gives acceptable coefficients. The dummy variable indicates a downward shft of 6.5% in the interwar period. The benefit to wage ratio does not give the expected sign or a significant coefficient. Thus there is little evidence even on a supply model that is stable between the two periods that the labour supply schedule was shifted up by the unemployment insurance system.

On the demand side, we investigated one further possibility suggested by a previous finding that the benefit to wage ratio shifted the demand curve rather than the supply curve of labour (Hatton, 1983). This effect arises because the insurance system allows firms to dishoard labour and encourages temporary layoffs. When the benefit to wage ratio was included in the demand function the structural break disappears and the following was obtained

$$\begin{aligned} \ln ER_t = \; & \underset{(0.2555)}{1.3594} + \underset{(0.0476)}{0.3476D} - \underset{(0.1508)}{0.5709\hat{R}_t} - \underset{(1.0751)}{0.1683\ln \hat{S}} \\ & - \underset{(0.0385)}{0.1333\ln(\hat{W}_t/\hat{P}_t)} + \underset{(0.0719)}{0.8930\ln Q_t} - \underset{(0.0013)}{0.0155t} \end{aligned}$$

$\bar{R}^2 = 0.8445 \quad RSS = 0.0144 \quad DW = 1.3006 \quad BP = 11.987$

This equation fits well though there is evidence of serial correlation on both DW and BP statistics. Changes in the coefficients are interesting and suggestive however. $\hat{R}$ gives a large negative coefficient implying an elasticity of the employment rate of about 1.6 but the coefficient on the dummy rises to 0.35 which is very close to the mean of the $\hat{R}$ series for 1921-1938. Hence, there is no net downward shift in the demand curve but employment becomes sensitive to deviations of $\hat{R}$ around the mean.

Furthermore, the coefficient on ln $\hat{S}$ is now the same order of magnitude as that on the real wage although it has lost significance. Clearly, the restriction that these equal cannot be rejected.

Since labour market equations tend to be rather unstable over short periods with few observations, one might argue that the effect of introducing unemployment insurance could only by gauged by finding a model which appear to be robust over a long period of time and which spans the periodswhen insurance was absent and present. If we pick the best of our equations (or any others for that matter) we fail to find the benefit to wage ratio shifting the supply curve as has sometimes been suggested. On the other hand, this variable does seem to have had an effect in the demand curve by making employment responsive to the benefit to wage ratio but not shifting the function down appreciably.

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