Earnings, Unemployment, and Housing: Evidence from a panel of British regions

Gavin Cameron and John Muellbauer*
Nuffield College, Oxford
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Abstract:

British regions display persistent differences in both earnings and unemployment rates. A number of studies have found that in general, regions that have high unemployment tend to have low wages. This runs contrary to a compensating differentials argument that high wages should compensate for high unemployment. However, levels of labour mobility in Britain, and especially levels of labour migration, are surprisingly low. The housing market therefore has an important impact on regional convergence. This paper discusses the determination of regional earnings and unemployment in the ten regions of Great Britain between 1972 and 1995, paying particular attention to their joint determination and to the influence of the housing market. We conclude that there is no wage-curve for non-manual men nor for full-time women, and that the wage-curve appears to be positively sloped for part-time women. However, for manual men, we find a significant elasticity of around –0.07, contrasting with Blanchflower and Oswald's –0.1. For full-time men and women, we find highly significant but somewhat smaller long-run housing market effects than Blackaby and Manning (1992), but with particularly strong effects for non-manual men compared with manual men. For unemployment, we confirm the important positive effect of lagged earnings on unemployment.

Key Words: Earnings, Unemployment, Housing Markets, Wage-Curves, Regions.

JEL Classification: C33;E24;R23

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1. Introduction

This paper has three main aims. The first is to use regional data to illuminate classic debates about the relationship between unemployment and earnings, which concerned Dennis Sargan himself (Sargan 1964, 1971, and 1980). See also Layard, Nickell and Jackman (1991) and the new twists given to the issue by Blanchflower and Oswald (1994). The latter argue that a 'wage-curve' exists, implying a negative relationship between unemployment and real wages which is fairly stable over time and between countries.

The use of regional earnings and unemployment differentials has important advantages in removing difficult to model national features of the data, such as national expectations of prices and changes in national legislation. For example, it is likely that regional differentials are less contaminated by the effect of equal pay legislation, incomes policies, and the many industrial and labour market reforms of the Thatcher era. A less attractive aspect is that models of regional earnings differentials can have little to say about the relevant macrovariables entering national wages and salary negotiations. This limits the conclusions which can be drawn on the macroeconomics of pay determination.

When the effects of structural changes have a homogenous additive effect on the variables of interest, we have a simple example of 'co-breaking', a concept introduced by Hendry (1997). Co-breaking occurs when a linear combination of variables subject to a structural break removes the effect of that break. Taking the difference between the regional and national average is a simple example of such a linear combination.

Second, the paper explores the interactions of labour and housing markets in the determination of earnings and unemployment outcomes. The mechanisms at work include, in principle, the influence of housing tenure on labour mobility and migration,[1] the effect of house prices on migration and commuting and hence on regional mismatch[2], the effects of house prices on the cost of living, on the demand side and perhaps on expectations.[3] The paper confirms the findings of Blackaby and Manning (1992) that, at the regional level, lagged house prices have a major influence on earnings. It also establishes an influence of mortgage costs on earnings. We also examine the hypothesis advanced by Oswald (1998) of an important link between a high rate of owner-occupation and a high unemployment rate.

The third aspect concerns the study of regional inequalities, a classic issue in regional economics, see Armstrong and Taylor (1994) and more recently, studies in the burgeoning econometric literature on convergence of incomes between regions and between nations, see Barro and Sala-i-Martin (1991) and Quah (1996). Typical studies of regional convergence examine the convergence of GDP per worker between regions, as measured in the national and regional accounts. We examine the most important component of GDP per worker, namely earnings per worker.

The plan of the paper is as follows. Section 2 reviews the theoretical background. Section 3 considers regional labour market data issues, presents our models of unemployment and wages in a panel of British regions and compares these results with models based on the previous literature, especially the contributions of Blackaby and Manning, and Blanchflower and Oswald. Section 4 draws conclusions. The data appendix reviews the sources of the data used in the paper.

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1 See, for example, Hughes and McCormick (1987), Minford, Peel and Ashton (1987) and McCormick (1997).

2 See Jackman and Savouri (1992) and Cameron and Muellbauer (1998).

3 See Bover, Muellbauer and Murphy (1989) and McCormick (1997).

2. Background

The work of Phillips (1958) stimulated a great deal of research on the connection between nominal wage inflation and unemployment. Apart from its many important contributions to econometric methodology, the paper by Sargan (1964) anticipated the later literature on the expectations augmented Phillips curve by emphasising that, in the long-run, workers and firms would be concerned with real wages (see Phelps, 1968, and Friedman, 1968). Sargan therefore posits an equilibrium correction model where the rate of change of nominal wages depends on the lagged rates of changes of prices, on the lagged real wage, on the lagged unemployment rate and on shifting trends reflecting productivity, government policy and union power.[4] Such a model can be trivially re-parameterised[5] into a real wage model of the type Layard and Nickell (1986) estimated for the UK and Layard, Nickell and Jackman (1991, chapter 9), estimate for the OECD countries. These authors include further terms such as the long-term unemployment rate, proxies for union power and labour market mismatch, components of the 'wedge' between product prices relevant for firms and consumer prices relevant for workers and the replacement rate (the ratio of unemployment benefits to the average wage).

Bover et al (1989) argued that such equations omitted three potentially important housing market influences. The first arises from the different mobility rates associated with different housing tenures. As the tenure structure changes, the average rate of labour mobility and hence labour market mismatch will alter. Evidence based on aggregate data in Bover et al was consistent with this interpretation with a quantitatively large effect on the unemployment/vacancies trade-off.

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4 Indeed, Sargan emphasizes the long-run solution for the real wage.
5 By adding a term in the rate of change of current inflation.

The second effect is part of the wedge between the cost of living relevant for workers and the prices received by firms. Since land is not a produced good, house prices largely reflect such a wedge. Empirically the lag between wages and house prices appears to be from one to three years.

The last effect arises in an aggregate equation from temporarily high labour market mismatch arising because of high regional differentials between ratios of house prices to earnings. In the late 1980s, the record differentials between the South East and other regions encouraged net migration from the South East, as confirmed by the analysis of regional migration data in Jackman and Savouri (1992) and Cameron and Muellbauer (1998). This worsened regional mismatch, increasing wage pressure and aggregate unemployment relative to aggregate vacancies.

Though the work of Layard and Nickell needs augmenting in these respects, it gives us good grounds for supposing that unemployment has a negative effect on wages at the aggregate level. However, at the regional level, a compensating differentials argument could suggest the opposite: high wages may be necessary to compensate for high unemployment. In the influential model of Harris and Todaro (1970), which comprises a developed urban and undeveloped rural sector, a migration equilibrium is achieved through unemployment in the developed urban sector. Of course, while high wages may serve to compensate for higher unemployment, in a developed economy they might just be compensating for some other factor, such as high housing costs or lack of regional amenity, rather than unemployment.[6]

As is to be expected, the evidence is mixed. For the United States, it does appear that high wage states tend to have higher unemployment rates (see Blanchard and Katz, 1992, and Hall, 1970). In contrast, Blanchflower and Oswald (1994, 1995) claim to find an empirical regularity (the 'wage curve') of a robust negative correlation between wages and log unemployment for a wide range of different countries and datasets, with a typical elasticity of around –0.10. Their approach, which has received a great deal of publicity, is to use repeated cross-sections of data on individuals in a range of countries. They argue therefore that employees in areas of high unemployment earn lower wages, other things being equal, than those in low unemployment areas.[7] Blanchflower and Oswald also challenge the orthodoxy of the Phillips curve by suggesting that dynamics are unimportant in the wage equation and conclude that once region and time effects are included, there is 'little sign of autoregression in wage equations' (1994, pp. 284).

For the United States, it appears that most of the regional adjustment to shocks is through movements of labour rather than through job creation or job migration (see Blanchard and Katz, 1992). This reflects the different labour and housing market institutions of the US. For the United Kingdom, a wide range of commentators have suggested that movements of labour play only a very small role in regional adjustment. For example, McCormick (1997) suggests that differences in UK regional unemployment rates show little sign of disappearing and that this is exacerbated by low levels of migration among manual workers. Jackman and Savouri (1992) argue that it is also important to examine regional commuting, since this enables workers to change their jobs without changing their home.

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6 See Ghatak, Levine, and Price (1996) for a survey of the literature following Harris and Todaro.

7 Note that this conclusion is only strictly true in models with no regional fixed effects. If regional fixed effects are included, the models cannot be interpreted as saying anything about the absolute level of wages. Instead, it is more relevant to say that other things being equal, when unemployment rises in a region, wages fall (Bell, 1997).

Cameron and Muellbauer (1998) examine the determination of commuting and migration in the UK regions and find that migration responds strongly to relative earnings and relative employment prospects. However, high relative house prices discourage migration into a region, although this may be offset a little by the expectation of fast future house price growth in that region. Recent experience of negative returns in housing also acts as a strong disincentive against migration into a region. Overall, even if high house prices and the fear of negative returns deter migrants into a region, commuting can take the place of migration to some extent. Of course, commuting comes at a cost, and a sharply increasing one with travel time and distance.

While high relative earnings and good employment opportunities in the South East should encourage in-migration, higher house prices and demand for housing as a portfolio investment crowd out people who wish to migrate to the South East. This helps to explain why net out-migration from the South East peaked in 1987-89 when the South East labour market was relatively buoyant. This net out-migration is difficult to explain in a compensating differentials framework unless one assumes there was a dramatic fall in the amenity value of living in the South East. A more plausible explanation seems to lie in the operation of the housing market.

In addition to the migration/commuting trade-off, the housing market is likely to have an effect on regional wages in a number of other ways, such as through cost of living effects, expectations, and through its effect on wealth and hence demand in a region (see Muellbauer and Murphy, 1995). Clearly, a region with higher house prices will tend have a higher overall cost of living which may affect the level of wage bargains. For regions with a higher level of mortgage debt, changes in interest rates are also likely to have differential impacts upon wage bargaining, to the extent that there is local rather than national bargaining.

Oswald (1998) has recently argued that high levels of home-ownership and the collapse of the private rented sector in the UK has led to higher structural unemployment. Using data from the British Social Attitudes Survey between 1983 and 1994, he finds that home-owners are generally less willing to move to a different area to find a job, other things being equal. Furthermore, there does appear to be a correlation between home-ownership and unemployment at a cross-country level. Oswald's formal analysis of nineteen OECD countries concludes that the rise in home-ownership in the UK since the 1960s added around 4 percentage points to the UK unemployment rate.[8]

3. Modelling Regional Earnings Differentials & Unemployment

3.1 Regional Earnings Data

At the regional level, there are two main sources of data: the Regional Accounts and the New Earnings Survey.[9] For the Regional Accounts, national estimates of wages and salaries are distributed across the regions using data from a one percent sample of National Insurance records. Since these records are on a place of residence basis, the regional accounts measure of earnings is therefore also on a residence basis. The New Earnings Survey is on a place of employment basis. This has three advantages as a data source. The first is that the place of employment is arguably more appropriate to 'locate'

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8 Though the decline in Bover et al's mobility index since the early 1970s appears to have only a small effect on unemployment relative to vacancies.

9 Though regional data on male manual earnings in manufacturing and a small group of other sectors are available back to 1960, see British Labour Statistics Historical Abstract 1886-1968 (1971).

a labour market than the place of residence since it is where employers and employees meet. The second is that we can disaggregate the NES data into male and female full-time and female part-time categories.

The third is that there are discrepancies between the NES and Regional Accounts measures of earnings, even allowing for the difference between place of employment and place of residence (see Cameron and Muellbauer, 1998, for a discussion of UK regional commuting). Until these have been explained, we regard the NES as the more reliable source.

3.2 Previous Work

There is a large literature on modelling regional earnings or wage differentials. Jackman, Layard and Savouri (1991) estimate the following equation for 1967-1987 for manual men's wage differentials:

$$(2) \qquad \ln w_{it} - \ln w_t = -\underset{(5.6)}{0.045}(\ln u_{it} - \ln u_t) + \underset{(12.9)}{0.68}(\ln w_{it-1} - \ln w_{t-1}) + \underset{(4.2)}{0.16}\left(\frac{ltu_{it}}{u_{it}} - \frac{ltu_t}{u_t}\right)$$
$$+\text{region fixed effects} + \text{regional time trend}$$

Here w_i is the real hourly wage for full time manual male workers in region i and w that in Great Britain, u_i and u are the respective regional and national unemployment rates and ltu_i and ltu are the long-term unemployment rates. Estimation is by pooled unweighted OLS and the equation standard error is 0.0070. One obvious problem with this equation is that the estimates are likely to be inconsistent due to contemporaneous correlation between unemployment shocks and the error term.

Blackaby and Manning (1990, 1992) follow a more sophisticated approach. They model wp_i, the log deviation between men's full-time gross weekly earnings in region i and `predicted' earnings based on applying national earnings patterns by industry to the employment patterns of the ith region. Effectively, this uses shift-share analysis to remove the effect of differential employment patterns from regional earnings variations. Using NES data for 1972-1988, they begin with the following general specification, in their notation:

$$(3) \qquad \Delta wp = a_0 + a_1 \Delta ru + a_2 \Delta wp_{-1} + a_3 \Delta ru_{-1} + a_4 \Delta ru52_{-1} + a_5 \Delta rh_{-1} + a_6 \Delta lh_{-1} + a_7 wp_{-1} + a_8 ru_{-1} + a_9 ru52_{-1} + a_{10} lh_{-1} + a_{11} rh_{-1}$$

where region subscripts are suppressed. All variables are in regional differential form where ru refers to the unemployment rate, ru52 the unemployment rate for those over 52 weeks unemployed, rh refers to the regional cost of living index[10], and lh refers to the regional house price index. This model is estimated by unweighted least squares.

The long-run solution has the form

$$(4) \qquad wp = fixedeffect - 0.0045(ru - ru52) + 0.15rh + 0.15lh$$

In the short-run dynamics α_1=-0.0053 (t=2.92) reflecting a negative effect from the change in the current unemployment rate (estimation is by IV to reflect the potential endogeneity of this variable but the OLS results are very similar) and α_6=-0.033 (t=2.24) reflecting a negative effect from last year's house price change or implying that the level of house prices from 2 years ago, as well as last year has a positive effect on earnings. The lagged level of house prices has a t-ratio of 6.13 in the dynamic specification. The equation standard error is 0.00704. Blackaby and Manning do not impose symmetry between regions but allow a few region specific effects, particularly for Scotland. These are selected on a goodness of fit basis.

Blackaby and Manning (1992) write their unemployment equation in the general form

$$(5) \qquad \Delta ru = \beta_0 + \beta_1 \Delta ru_{-1} + \beta_2 \Delta ru52_{-1} + \beta_3 \Delta wp + \beta_4 \Delta wp_{-1} + \beta_5 (ru_{-1} - ru52_{-1}) + \beta_6 wp_{-1} + \beta_7 rg_{-1}$$

where ru, ru52 and wp are as above and rg is the log deviation of regional GDP per head from the national average.

The empirically estimated form of (5) as of (3) contains some important data-selected terms specific to particular regions and has an equation standard error of 0.281. Neglecting these asymmetries, the long run solution takes the form

$$(6) \qquad ru = \text{constant} + ru52 + 19.4\, wp - 7.4\, rg$$

Thus, the regional rate of unemployment of those unemployed less than one year increases with regional male earnings and decreases with regional GDP/head. Blackaby and Manning do not indicate whether they tested the long run coefficient of one on long-term unemployment r52.

10 From Rewards Regional Surveys Ltd.

3.3 A Model of Relative Regional Earnings

Our model of relative regional earnings differs in a variety of ways from those of Jackman et al (1991) and Blackaby and Manning (1992). First, we model the full-time earnings of both men and women and test for lagged links between them. We also model women's part-time earnings. Secondly, we include a measure of the regional difference in mortgage interest costs using data on the ratio of the average regional mortgage stock to average full-time earnings in the region. However, we exclude the regional cost of living (reported by Rewards Regional Surveys Ltd) used by Blackaby and Manning given concerns about its methodology of construction and its accuracy. Thirdly, unlike Blackaby and Manning (1992) we model the deviation of log earnings from the average for Great Britain rather than the deviation between log earnings in the region from what they would have been had that region's industrial structure being applied to national average earnings by sector. Fourthly, unlike Blackaby and Manning (1992), we impose symmetry on all slope parameters but include interaction effects between two indicators of regional employment structure and a small set of macro-variables. These indicators are the proportion of employment in the production sector (i.e., manufacturing and mining) and the proportion of employment in the banking and financial services sector. The former is interacted with the real exchange rate or 'competitiveness' since the production sector is heavily exposed to international competition and hence exchange rate movements. The latter is interacted with an indicator of financial liberalization which is zero up to 1980 and rises to a peak of unity in 1989-90.[11] We also include the

[11] The time profile is broadly consistent with regional and national evidence from loan-to-value ratios on mortgages for first-time buyers, see Muellbauer (1997).

proportion of employment accounted for by part-time women workers, competition from whom might be expected to exert downward pressure on men's earnings.

A fifth difference is that we exclude long-term unemployment, being unable to find a significant effect. However, we investigate whether the unemployment effects enter in log or levels form. The evidence is consistent with a long-run log unemployment effect as in Sargan (1964), Layard and Nickell (1986) and in most of the country studies in Layard, Nickell and Jackman (1991). However, the instrumented contemporaneous effect, where relevant, is stronger in the change in the unemployment rate rather than its log.

Finally, we estimate using GLS to allow for heteroscedastic and correlated stochastic disturbances across regions, but like Blackaby and Manning (1992) instrument the contemporaneous unemployment rate.

Let $lrme_i$ be relative log-earnings for full-time men in the ith region. Our model takes the following fairly general form:

$$
\begin{aligned}
(7) \qquad \Delta lrme = {} & a_0 + a_1\, lrme_{-1} + a_{11}\Delta lrme_{-1} + a_{12} lrfe_{-1} + a_{13}\Delta lrfe_{-1} \\
& + a_2\Delta ru + \alpha_{21}\Delta ru_{-1} + \alpha_{22} lru_{-1} + \alpha_3 poouk_{-1}.lrhp_{-1} + \alpha_{31}\, poouk_{-1}.\Delta lrhp_{-1} \\
& + \alpha_4 wwr_{-1} + \beta_{41} w\Delta rr_{-1} + \beta_{42} w\Delta rr_{-2} \\
& + \beta_5 pfpt_{-1} + \beta_{51}\Delta pfpt_{-1} + \beta_6 pp_{-1} + \beta_{61}\Delta pp_{-1} \\
& + \beta_{62}(rer_{-1})pp_{-1} + \beta_{63}(\Delta rer_{-1})pp_{-1} \\
& + \beta_7 pb_{-1} + \beta_{71}(aflib_{-1})pb_{-1} + \beta_{72}(\Delta aflib_{-1})pb_{-1}
\end{aligned}
$$

where lrfe is relative log-earnings for full-time women, ru is the regional deviation in unemployment rates, lru is the regional deviation in log unemployment rates, lrhp is the deviation in log house prices and poouk is the national percentage of owner-occupiers, wrr is the deviation in the average mortgage debt to income ratio in region i multiplied by the mortgage interest rate. $w\Delta rr_{-1}$ applies the same weight to last year's change in the mortgage interest rate while $w\Delta rr_{-2}$ applies it to the change two years previously.[12] The remaining variables are various employment composition proportions and their interactions. Thus pfpt is the proportion of part-timers women, pp is the proportion of employment in the production sector, pb is the proportion of employment in banking and financial services. There are two interactions. The first is between production sector employment and the real exchange rate rer. Thus, when Britain loses international competitiveness, as in 1979-81 and 1988-92, it seems likely that regions with large employment shares in this sector would suffer disproportionately. The second is between the employment share in banking and an index of financial liberalization which is approximately zero before 1980 and rises to 1 in 1988-9.[13]

To test whether the response of relative log-earnings is to relativities in levels or logs of the regional unemployment rates, eq. (7) was nested in an even more general model containing Δlru, Δru, Δlru_{-1}, Δru_{-1}, ru_{-1} and lru_{-1} effects. The data favour a long-run response to log-unemployment but short run responses to changes in the levels of unemployment. We also investigated an alternative specification analogous to Blackaby and Manning's using ru52, the unemployment rate for those unemployed one year or more. However, this proved inferior.

12 This is the only variable where an effect from as long as three years earlier enters the general specification. Macroevidence suggests quite long lags from interest rates on output. Allowing for a somewhat longer lag here gives greater scope to the possibly negative effect of interest rates on earnings.

13 We also experimented with an alternative measure of financial liberalization: the proportion of mortgages to first-time buyers with a loan-to-value ratio greater than 90 per cent. This variable gave very similar results to the index of financial liberalization in both our general and parsimonious specifications.

Following a general to specific testing down procedure, eq. (7) was reduced to the more parsimonious specification reported in Table 1. Contemporaneous changes in unemployment were instrumented using the forecast changes from the model described in the next section. Estimations using the SUR procedure in Hall et al (1996) Time Series Processor (TSP), but iterating once on the contemporaneous covariance matrix of errors across the 10 regions, which is also corrected for the use of fitted instead of actual changes in the unemployment rate, where this appears.

Table 1 summarizes our results for men's earnings. There is a significant equilibrium-correction term which implies that earnings return quickly to their equilibrium levels. There are two points to note here. First, that the presence of this term means that the equation can be interpreted as a conditional convergence regression, with the implication that relative steady-state income is determined by the other variables in the model. Second, that this is evidence against Blanchflower and Oswald's contention that autoregression is unimportant in wage equations (1994, pp. 284).

The lagged log level of the deviation of regional unemployment from the national level also has a significant and negative effect with a long-run coefficient around -0.02, one fifth of the -0.1 figure claimed by Blanchflower and Oswald as a robust order of magnitude of the slope of the wage curve. The contemporaneous fitted value of the change in relative unemployment (taken from the unemployment equation) is not significant. One possible reason for this is that the NES earnings data are observed in April while the unemployment rates are annual averages. However, there is a strong negative effect from the change in the previous year's relative unemployment rate.

Turning now to the housing market effects, lagged relative house prices have a positive and highly significant effect on relative earnings, and this effect has become stronger as the proportion of owner-occupation in the UK has risen. Note that the proportion of owner-occupation in the UK produces a more significant coefficient here than if the proportion of owner-occupation in the region is used or then if relative log house prices are unweighted. At a 68% rate of UK owner-occupation, the long-run effect of relative log house prices on men's relative full-time earnings is 0.07. Relative mortgage costs in the previous year have a positive effect on relative earnings. But a rise in mortgage interest rates 2 years earlier has a temporary negative effect on relative earning in regions with high ratios of mortgage debt to earnings. This appears to reflect the long-term recessionary implications of higher interest rates on regions with high debt to income ratios.

Lastly, we have the composition effects. Since this is a model of men's earnings, it is interesting that regions with high proportions of part-time women have lower relative men's earnings, suggesting an element of substitution between these groups of workers. We also find that men's earnings in regions with more production workers suffer more when competitiveness falls (that is, the log real exchange rate rises) and that men's earnings in regions with more banking and financial sector workers do better when there is financial liberalisation.

Since these estimates use Generalised Least Squares, each region's equation has a different standard error, which as expected, tends to be inversely related to employment in each region. The quoted equation standard error is the standard deviation of all the residuals for comparison with the OLS or unweighted IV estimates of Jackman et al.

(1991) and Blackaby and Manning (1992). Tests for heteroscedasticity and residual autocorrelation within regions are satisfactory. But the normality test suggests that the residuals are close to being non-normal.

Table 1b presents the short-sample estimate of the model, from 1972 to 1987. This omits the peak of the 1980s house price boom and the 1990s housing market crisis, a severe test of parameter stability. There are no significant differences in estimated coefficients, though the point estimate of the lagged weighted change in mortgage interest rates falls, and the restriction of no structural break cannot be rejected at the five percent level.[14]

Having estimated the relationship using forecast from our unemployment equations, and not being able to reject the hypothesis of a zero coefficient on the current unemployment rate, we can compare these results with using the actual unemployment data. These results are given by table 2. The coefficient on the contemporaneous change in the unemployment rate is now significantly negative, (t=4.8), evidence consistent with a notable endogeneity bias. Also the estimated long-run slope of the wage curve is -0.03 as opposed to -0.02 when current unemployment is instrumented or omitted. Part of the estimated negative effect of current unemployment on current earnings is probably due to common demand shocks which raise unemployment and depress wages. This suggests that Blanchflower and Oswald’s claims of a slope of -0.1 for the wage curve may be exaggerated because of endogeneity bias, a possibility noted by Card (1995).

One of the most interesting results from Table 1 is that the long-run coefficient on relative house prices, 0.07 at a 68% rate of owner occupation, though highly significant, is only around one half Blackaby and Manning's estimate of 0.15. Of the various differences in specification between their model and Table 1, an important one is the interaction between financial liberalization and the proportion of employment in the banking sector included in our model.[15] Excluding this variable pushes the long-run coefficient on relative house prices up to 0.11. If we further exclude the interaction of the real exchange rate with the proportion of employment in the production sector and exclude the proportion of employment accounted for by part-time women, the long-run house price coefficient rises to 0.19. This suggests that, in Blackaby and Manning's model, house prices may, in part, be proxying more fundamental structural changes in the economy which influenced both house prices and earnings.

For non-manual men's full-time earnings, we expect to find stronger house price effects than for men's full-time earnings. Owner-occupancy rates are higher among non-manual workers and from the work of Hughes and McCormick (1987) and McCormick (1997), we know that non-manuals have higher migration rates than manual workers and that these are likely to respond more strongly to regional house price differentials. These expectations are confirmed by the empirical findings. The long-run coefficient on relative house prices at a UK rate of owner-occupation of 68% is 0.12 for non-manual men compared with 0.07 for all full-time men. The long-run log-unemployment effect is zero: there is no wage curve for non-manual men. Again, this is not a big surprise since manual workers dominate the aggregate unemployment rate. For manual men, however, we estimate the long-run slope of the wage curve to be -0.07, now much closer to the

[14] We impose the full-sample residual covariance matrix in the SUR procedure for the short-sample, and also hold the residual covariance matrix constant in the two samples when conducting the F-test for parameter stability (see Greene, 1993).

Blanchflower and Oswald figure of -0.1, and the long-run house price effect to be roughly 0.02, but not very accurately estimated, at a UK rate of owner-occupation of 68%.

Full-time Women's Earnings

For full-time women's earnings, we formulated a model symmetric to eq (7) and again followed a general to specific model selection procedure. The results are shown in Table 3. In contrast to full-time men's earnings, we could find no influence from mortgage costs, the proportion of employment defined by part-time women, the interaction of competitiveness with the proportion of employment in the production sector and no wage curve. However, the contemporaneous change in the unemployment rate has a very significant negative effect, presumably capturing demand shocks.

There is a positive spill-over effect from men's earnings: if relative men's earnings increased last year, there is a small positive effect on relative women's earnings this year. There is also a small negative reaction to a positive shock to last year's relative women's earnings. The coefficient of -0.64 on the lagged level of relative earnings suggest a relatively high speed of adjustment.

Unlike the earnings equation for men, the proportion of employment in the production sector (measured as the lagged 2 year moving average) has a positive effect on relative women's earnings. The interaction of financial liberalization with the proportion of employment in banking has a similarly large effect on women's as on men's earnings.

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[15] As indicated in footnote 13, a simple alternative indicator of financial liberalization gives very

Finally, although the relative house price effect is roughly represented by the lagged two year moving average, i.e., on average, half a year further back than for men, the long-run coefficient is around 0.07, as for full-time men.

Table 3b also shows results for the 1972-87 subsample with a somewhat weaker effect from the change in the unemployment rate but otherwise very similar parameters. An F-test for parameter stability is easily accepted.

Women's Part-time Earnings

The general equation for relative earnings of part-time women, otherwise symmetric to equation (7), included lagged effects both from full-time men's and women's earnings. Of these only the temporary effect from last year's change in the relative earnings of full-time women survive in the parsimonious specification shown in Table 4.

Otherwise there are some striking differences with the full-time earnings models. First, the wage curve is perversely sloped: the effect of relative log-unemployment on relative part-time women's earnings is positive and very significant. We comment further below. There is no effect from the current change in the unemployment rate but a negative response to the lagged change.

Secondly, there is no relative house price effect but a strong positive effect both from last year's level and last year's change in relative mortgage costs. Thus higher mortgage interest rates raise the relative earnings of part-time women in regions with high mortgage debt-to-income ratios.

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similar results, including a similar long-run house price effect.

Thirdly, the lagged two-year moving average of the relative proportion last year of employment accounted for by part-time women has a large positive effect on relative part-time women's earnings. Fourthly, the interaction of the real exchange rate entering as a lagged two-year moving average, with the proportion of employment in the production sector has a large negative effect on relative part-time women's earnings. This proportion itself also has a negative effect while the relative proportion of employment in the banking sector has a positive effect on relative part-time earnings.

To interpret these results, note that weekly earnings of part-time women are sensitive to the number of hours worked. The data may also be sensitive to a selection problem in that the NES data omit the earnings of many individuals below the national insurance floor. The sample composition of part-time women included in the survey may therefore fluctuate.

In the conclusion of the paper, we suggest an interpretation of the perversely sloped wage curve for part-time females in terms of the unemployment trap and other mechanisms. The mortgage cost effects also suggest a relative hours response to higher mortgage rates in regions with high mortgage debt to income ratios.

Table 4 suggest that the parameter estimates are fairly stable over the 1974-1987 subsample and an F-test confirms this result.

Table 1a Men's Full-Time Earnings Model - Long Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RLFTMEi(-1)	-.443864	.036049	-12.3128	[.000]
LRUi(-1)	-.865848E-02	.405652E-02	-2.13446	[.033]
DRU i(-1)	-.358203E-02	.124559E-02	-2.87577	[.004]
RLHPi(-1)*POOUK(-1)	.471869E-03	.683664E-04	6.90206	[.000]
Wrri(-1)	.113308E-02	.433598E-03	2.61321	[.009]
WDrri(-2)	-.697308E-02	.925717E-03	7.53263	[.000]
PFPTi(-1)	-.296017	.052060	-5.68609	[.000]
PPi(-1)*LRER(-1)	-.280771	.075215	-3.73292	[.000]
PBi(-1)*AFLIB(-1)	.334727	.065603	5.10228	[.000]
Northern	-.359974E-02	.256687E-02	-1.40238	[.000]
North-West	-.572528E-02	.156157E-02	-3.66635	[.001]
Yorkshire & Humberside	-.322776E-02	.284133E-02	-1.13600	[.256]
West Midlands	-.021129	.285425E-02	-7.40283	[.000]
East Midlands	-.022708	.308165E-02	-7.36893	[.000]
East Anglia	-.017158	.308071E-02	-5.56959	[.000]
South East	.026619	.313361E-02	8.49478	[.000]
South West	-.018538	.355429E-02	-5.21573	[.000]
Wales	-.018998	.247134E-02	-7.68730	[.000]
Scotland	-.959501E-02	.278746E-02	-3.44220	[.001]

Number of Observations	24	**Sample Period**	1972-1995
Equation Standard Error	0.00796648	**Autocorrelation Test (F)**	1.45 [0.24]
Heteroskedasticity Test (F)	0.48 [0.49]	**Jarque-Bera Normality**	5.96 [0.06]
Parameter stability (F)	1.06 [0.38]		

Notes:
Dependent Variable is DRLMEi, that is, the change in log relative men's earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

Table 1b Men's Full-Time Earnings Model - Short Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RLFTMEi(-1)	-.491314	.055860	-8.79545	[.000]
LRUi(-1)	-.013623	.604156E-02	-2.25488	[.024]
DRUi(-1)	-.323711E-02	.171762E-02	-1.88465	[.059]
RLHPi(-1)*POOUK(-1)	.491719E-03	.162233E-03	3.03094	[.002]
Wrri(-1)	.122378E-02	.791051E-03	1.54703	[.122]
WDrri(-2)	-.275786E-02	.187862E-02	1.46803	[.142]
PFPTi(-1)	-.308981	.068718	-4.49634	[.000]
PPi(-1)*LRER(-1)	-.291650	.093551	-3.11756	[.002]
PBi(-1)*AFLIB(-1)	.359324	.096878	3.70906	[.000]
Northern	-.294530E-02	.379377E-02	-.776352	[.438]
North-West	-.578897E-02	.250902E-02	-2.30727	[.021]
Yorkshire & Humberside	-.316971E-02	.457342E-02	-.693074	[.488]
West Midlands	-.022975	.373894E-02	-6.14482	[.000]
East Midlands	-.027470	.468899E-02	-5.85847	[.000]
East Anglia	-.021040	.473459E-02	-4.44390	[.000]
South East	.027342	.436568E-02	6.26296	[.000]
South West	-.022553	.485793E-02	-4.64254	[.000]
Wales	-.017912	.311378E-02	-5.75243	[.000]
Scotland	-.810821E-02	.381014E-02	-2.12806	[.033]

Number of Observations	16	**Sample Period**	1972-1987
Equation Standard Error	0.0079616	**Autocorrelation Test (F)**	3.13 [0.05]
Heteroskedasticity Test (F)	5.02 [0.03]	**Jarque-Bera Normality**	21.51 [0.00]

Notes:
Dependent Variable is DRLME, that is, the change in log relative men's earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

Table 2 Men's Full-Time Earnings Model with actual unemployment

Parameter	Estimate	Standard Error	t-statistic	P-value
RLFTMEi(-1)	-.444065	.034878	-12.7321	[.000]
LRUi(-1)	-.014051	.378648E-02	-3.71085	[.020]
DRUi(-1)	-.117913E-02	.116801E-02	-1.00952	[.001]
DRUi	-.594497E-02	.123813E-02	-4.80158	[.000]
RLHPi(-1)*POOUK(-1)	.513740E-03	.665460E-04	7.72007	[.000]
Wrri(-1)	.197726E-02	.449256E-03	4.40118	[.000]
WDrri(-2)	-.733211E-02	.844713E-03	8.68000	[.000]
PFPTi(-1)	-.329779	.052820	-6.24349	[.000]
PPi(-1)*LRER(-1)	-.147645	.078639	-1.87750	[.000]
PBi(-1)*AFLIB(-1)	.344865	.066002	5.22507	[.000]
Northern	.333436E-02	.274752E-02	1.21359	[.225]
North-West	-.178136E-02	.165676E-02	-1.07521	[.282]
Yorkshire & Humberside	.239414E-02	.295980E-02	.808888	[.419]
West Midlands	-.018070	.279071E-02	-6.47517	[.000]
East Midlands	-.020695	.310202E-02	-6.67159	[.000]
East Anglia	-.017825	.298009E-02	-5.98131	[.000]
South East	.019532	.333915E-02	5.84938	[.000]
South West	-.019812	.351510E-02	-5.63625	[.000]
Wales	-.015173	.263687E-02	-5.75421	[.000]
Scotland	-.688017E-02	.273169E-02	-2.51865	[.012]

Number of Observations	24	**Sample Period**	1972-1995
Equation Standard Error	0.00776164	**Autocorrelation Test (F)**	1.11 [0.33]
Heteroskedasticity Test (F)	0.94 [0.33]	**Jarque-Bera Normality**	5.32 [0.07]
Parameter stability (F)	1.01 [0.42]		

Notes:
Dependent Variable is DRLMEi, that is, the change in log relative men's earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

Table 3a Women's Full-Time Earnings Model - Long Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
DRLFTMEi(-1)	.099543	.044089	2.25776	[.024]
RLFTFEi(-1)	-.640462	.068340	-9.37166	[.000]
RLFTFEi(-2)	-.143259	.060294	-2.37599	[.018]
DRUHi	-.015369	.358317E-02	-4.28931	[.000]
DRUi(-1)	.520756E-02	.185257E-02	2.81099	[.005]
DRLHPi(-1)*POOUK(-1)	-.387382E-03	.167678E-03	-2.31028	[.021]
RLHPi(-2)*POOUK(-1)	.690720E-03	.111071E-03	6.21870	[.000]
Wrri(-2)	-.792556E-03	.585717E-03	1.35314	[.176]
PPi(-1)* DLRER(-1)	-.488633	.097796	4.99646	[.000]
PBi(-1)*AFLIB(-1)	.484421	.068160	7.10710	[.000]
Northern	-.021787	.433915E-02	-5.02109	[.000]
North-West	-.018948	.312208E-02	-6.06912	[.000]
Yorkshire & Humberside	-.029022	.440880E-02	-6.58285	[.000]
West Midlands	-.034504	.410163E-02	-8.41216	[.000]
East Midlands	-.037993	.505216E-02	-7.52015	[.000]
East Anglia	-.026155	.382487E-02	-6.83820	[.000]
South East	.050308	.555491E-02	9.05654	[.000]
South West	-.022733	.374303E-02	-6.07355	[.000]
Wales	-.015720	.357644E-02	-4.39540	[.000]
Scotland	-.022623	.335754E-02	-6.73792	[.000]

Number of Observations	24	**Sample Period**	1972-1995
Equation Standard Error	0.00893314	**Autocorrelation Test (F)**	0.50 [0.61]
Heteroskedasticity Test (F)	15.39 [0.00]	**Jarque-Bera Normality**	0.12 [0.94]
Parameter stability (F)	1.07 [0.36]		

Notes:
Dependent Variable is DRLFEi(-1), that is, the change in log relative women's full-time earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals. DRUHi is the contemporaneous value of fitted unemployment from the unemployment model discussed in section 3.4.

Table 3b Women's Full-Time Earnings Model - Short Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
DRLFTMEi(-1)	.068124	.058546	1.16360	[.245]
RLFTFEi(-1)	-.717997	.097565	-7.35917	[.000]
RLFTFEi(-2)	-.087278	.075101	-1.16215	[.245]
DRUHi	-.806851E-02	.504297E-02	-1.59995	[.110]
DRU i(-1)	458998E-03	.298192E-02	.153927	[.878]
DRLHPi(-1)*POOUK(-1)	-.226801E-03	.296261E-03	-.765545	[.444]
RLHPi(-2)*POOUK(-1)	.646324E-03	.279909E-03	2.30905	[.021]
Wrri(-2)	-.219842E-02	.145180E-02	1.51427	[.130]
PPi(-1)* DLRER(-1)	-.420284	.118832	3.53678	[.000]
PBi(-1)*AFLIB(-1)	.579863	.114366	5.07024	[.000]
Northern	-.024957	.683333E-02	-3.65230	[.000]
North-West	-.024281	.505370E-02	-4.80464	[.000]
Yorkshire & Humberside	-.038323	.779566E-02	-4.91590	[.000]
West Midlands	-.035322	.546368E-02	-6.46489	[.000]
East Midlands	-.045131	.749739E-02	-6.01956	[.000]
East Anglia	-.029732	.595485E-02	-4.99287	[.000]
South East	.061344	.994473E-02	6.16851	[.000]
South West	-.027015	.547868E-02	-4.93088	[.000]
Wales	-.020423	.522649E-02	-3.90768	[.000]
Scotland	-.029853	.550760E-02	-5.42030	[.000]

Number of Observations	16	**Sample Period**	1972-1987
Equation Standard Error	0.00857148	**Autocorrelation Test (F)**	1.85 [0.16]
Heteroskedasticity Test (F)	2.27 [0.14]	**Jarque-Bera Normality**	1.08 [0.59]

Notes:
Dependent Variable is DRLFE, that is, the change in log relative women's full-time earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals. DRUHi is the contemporaneous value of fitted unemployment from the unemployment model discussed in section 3.4.

Table 4a Women’s Part-Time Earnings Model - Long Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RLFTFEi(-2)	.258655	.050845	5.08710	[.000]
RLFTPEi(-1)	-.573714	.045613	-12.5777	[.000]
LRUi(-1)	.041507	.664106E-02	6.25006	[.000]
DRUi(-1)	-.635639E-02	.214232E-02	-2.96706	[.003]
Wrri(-1)	.167338E-02	.626706E-03	2.67012	[.008]
WDrri(-2)	.013528	.159129E-02	8.50159	[.000]
MAPFPTi(-1)	.530761	.139111	3.81539	[.000]
PPi(-1)	-.443329	.124594	-3.55819	[.000]
Pbi(-2)	.852236	.289583	2.94298	[.003]
PPi(-1)*MALRER(-1)	-.995981	.123954	8.03508	[.000]
Northern	-.043550	.891401E-02	-4.88555	[.000]
North-West	-.011916	.628128E-02	-1.89700	[.058]
Yorkshire & Humberside	-.015168	.763159E-02	-1.98751	[.047]
West Midlands	.311725E-02	.011555	.269776	[.787]
East Midlands	.874217E-02	.010792	.810094	[.418]
East Anglia	-.016426	.661400E-02	-2.48350	[.013]
South East	.033843	.010724	3.15582	[.002]
South West	-.044444	.711163E-02	-6.24955	[.000]
Wales	-.031405	.703164E-02	-4.46630	[.000]
Scotland	.179849E-02	.514373E-02	.349647	[.727]

Number of Observations	22	**Sample Period**	1974-1995
Equation Standard Error	0.017850	**Autocorrelation Test (F)**	0.93 [0.40]
Heteroskedasticity Test (F)	0.87 [0.35]	**Jarque-Bera Normality**	17.1 [0.00]
Parameter stability (F)	1.27 [0.11]		

Notes:
Dependent Variable is DRLPEi, that is, the change in log relative women’s part-time earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

Table 4b Women's Part-Time Earnings Model - Short Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RLFTFEi(-2)	.303445	.062595	4.84772	[.000]
RLFTPEi(-1)	-.663395	.062242	-10.6583	[.000]
LRUi(-1)	.027817	.010059	2.76537	[.006]
DRUi(-1)	-.640214E-02	.383958E-02	-1.66741	[.095]
Wrri(-1)	.376871E-02	.157827E-02	2.38787	[.017]
WDrri(-2)	.011458	.288008E-02	3.97826	[.000]
MAPFPTi(-1)	.682926	.211552	3.22817	[.001]
PPi(-1)	-.481810	.203988	-2.36196	[.018]
PBi(-2)	1.29871	.459813	2.82442	[.005]
PPi(-1)*MALRER(-1)	-1.06796	.177305	6.02330	[.000]
Northern	-.035169	.013950	-2.52098	[.012]
North-West	-.014253	.953064E-02	-1.49548	[.135]
Yorkshire & Humberside	-.016789	.013195	-1.27242	[.203]
West Midlands	-.002845	.018691	-0.15226	[.879]
East Midlands	.011552	.017958	.643270	[.520]
East Anglia	-.023034	.896655E-02	-2.56893	[.010]
South East	.028256	.017866	1.58158	[.114]
South West	-.049699	.010683	-4.65221	[.000]
Wales	-.028825	.923222E-02	-3.12222	[.002]
Scotland	.010336	.718068E-02	1.43937	[.150]

Number of Observations	14	**Sample Period**	1974-1987
Equation Standard Error	0.017277	**Autocorrelation Test (F)**	0.47 [0.63]
Heteroskedasticity Test (F)	0.00 [0.97]	**Jarque-Bera Normality**	2.51 [0.29]

Notes:
Dependent Variable is DRLPE, that is, the change in log relative women's part-time earnings.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

3.4 A Model of Relative Regional Unemployment

Let us consider a model for the change in relative regional unemployment. Let ru_i be relative unemployment in the ith region. Our model takes the following form:

$$
\begin{aligned}
(8)\quad \Delta ru = {} & \gamma_0 + \beta_1 ru_{-1} + \beta_{11}\Delta ru_{-1} + \beta_{12}\Delta ru_{-2} + \beta_{21}(pooUK_{-1})\, lrhp_{-1} \\
& + \beta_{22}(pooUK_{-1})\Delta lrhp_{-1} + \beta_{23}(pooUK_{-1})\Delta lrhp_{-2} \\
& + \beta_{31} lrfe_{-1} + \beta_{32}\Delta lrfe + \beta_{33}\Delta lrfe_{-1} + \beta_{34} lrme_{-1} + \beta_{35}\Delta lrme + \beta_{36}\Delta lrme_{-1} \\
& + \beta_{41} wrr_{-1} + \beta_{42} w\Delta rr_{-1} + \beta_{43} w\Delta rr_{-2} + \beta_5 pp_{-1} + \beta_{51}(lrer_{-1})pp_{-1} \\
& + \beta_{52}(\Delta lrer_{-1})pp_{-1} + \beta_{53}(\Delta lrer_{-2})pp_{-1} + \beta_{61}\, pb_{-1} + \beta_{62}(\Delta mlr_{-1})\, pb_{-1} \\
& + \beta_{63}(\Delta mlr_{-2})\, pb_{-1} + \beta_{64}(\Delta mlr_{-3})\, pb_{-1} + \beta_{71}(\Delta lhpruk)_{-1}\, pb_{-1} \\
& + \beta_{72}(\Delta lhpruk)_{-2}\, pb_{-1} + \beta_{81}\,(aflib_{-1})\, pb_{-1} + \beta_{82}\,(\Delta aflib_{-1})\, pb_{-1}
\end{aligned}
$$

where ru is the regional deviation in unemployment rates, lrhp is the deviation in log house prices, wrr is deviation in the average mortgage debt to income ratio in region i weighting the mortgage interest rate, $w\Delta rr_{-1}$ applies the same weight to last year's change in the mortgage interest rate while $w\Delta rr_{-2}$ applies to the change two years previously. Lrfe is relative women's log earnings in the region, and lrme is relative men's log earnings.

The remaining variables are various employment composition proportions and their interactions. As before, pp is the proportion of employment in the production sector and pb is the proportion of employment in banking and financial services. There are a variety of interactions. The first is between production sector employment and the real exchange rate, rer. The second is between the employment share in banking and changes in bank base rate. The third is between the employment share in banking and lagged changes in real UK house prices. And the fourth is between the employment share in banking and an index of financial liberalization.

Table 5 presents our parsimonious model for the full sample period of 1972 to 1995. We find a relatively small negative and significant coefficient of –0.2 on the lagged level of relative unemployment, suggesting that an unemployment shock will persist for some time, although it does not suggest hysteresis in the sense of there being a unit root in relative unemployment (see Blanchard and Katz, 1997). The lagged change in relative unemployment rates has a positive effect, however, suggesting equilibrium correction to the relative unemployment rate of two years ago. In addition, there are significant effects from lagged relative earnings, with higher relative wages for both full-time men and women leading to higher relative unemployment. The men's effect is at a two year lag. The women's effect is much smaller in the long-run, though there is a positive effect from the lagged change in women's earnings.[16]

Turning to the housing market effects, there appears to be no long-run effect of higher house prices on unemployment. We interpret this finding in the conclusion of the paper.

Finally, we have a variety of effects of the industrial structure of the region on unemployment. The first is an interaction between the share of workers in the production section and changes in UK competitiveness. Therefore, when competitiveness falls, regions with more workers in the production sector suffer larger rises in unemployment. The second is an interaction between the share of workers in the banking and finance sector and changes in the base rate three years ago. Regions with a larger exposure to the banking sector are found to be more vulnerable to interest rate changes. The third is an interaction between the banking sector and changes in UK log real house prices two years ago. Regions with a larger exposure to the banking and financial services sector benefit more from rises in UK house prices. Mortgage lending and associated activities are an important part of the business of this sector and will tend to be weak when real national house prices are falling.

We found no significant evidence for a direct effect of the regional rate of owner-occupation on relative unemployment, in contrast to the results of Oswald (1998). If lagged log relative owner-occupation was included in the regression it had a coefficient of -.09 (t-statistic -0.2) and if we included the second lag, it had a coefficient of -0.16 (t-statistic -0.4). Inclusion of both together yields similar results. An F-test accepts the null hypothesis that relative owner-occupation had no significant effect.

Table 5 presents the short-sample estimate of the model, from 1972 to 1987. There are no significant differences in estimated coefficients, and the restriction of no structural break can be accepted at the five percent level.

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[16] The inclusion of fitted contemporaneous men's and women's full-time earnings from their respective equations does not yield a significant coefficient and all the other results are robust to this change.

Table 5a Unemployment Model - Long Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RUi(-1)	-.205141	.027862	-7.36268	[.000]
DRUi(-1)	.342632	.055424	6.18204	[.000]
RLFTFEi(-1)	2.87826	.919683	3.12962	[.002]
RLFTFEi(-2)	-2.09920	.892905	-2.35098	[.019]
RLFTMEi(-2)	4.20498	.923518	4.55322	[.000]
DRLHPi(-1)	.532419	.281223	1.89323	[.058]
PPi*LRER(-1)	12.8841	2.33602	5.51542	[.000]
PBi* DMLR(-3)	.700798	.208554	3.36027	[.001]
PBi* DLRHPUK(-2)	-9.17242	4.83342	-1.89771	[.058]
Northern	.976457	.121782	8.01808	[.000]
North-West	.606786	.082519	7.35334	[.000]
Yorkshire&Humberside	.568708	.076548	7.42946	[.000]
West Midlands	.563472	.094277	5.97675	[.000]
East Midlands	.380833	.079120	4.81334	[.000]
East Anglia	-.061582	.075330	-.817499	[.414]
South East	-.918848	.114798	-8.00407	[.000]
South West	.092164	.060356	1.52701	[.127]
Wales	.683330	.089694	7.61850	[.000]
Scotland	.469760	.099708	4.71136	[.000]

Number of Observations	24	**Sample Period**	1972-1995
Equation Standard Error	0.254091	**Autocorrelation Test (F)**	0.71 [0.49]
Heteroskedasticity Test (F)	0.53 [0.47]	**Jarque-Bera Normality**	25.93 [0.00]
Parameter stability (F)	1.06 [0.38]		

Notes:
Dependent Variable is DRUi, that is, the change in the relative unemployment rate.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

Table 5b Base Unemployment Model - Short Sample

Parameter	Estimate	Standard Error	t-statistic	P-value
RUi(-1)	-.194102	.042875	-4.52720	[.000]
DRUi(-1)	.252450	.074126	3.40569	[.001]
RLFTFEi(-1)	2.63255	1.21744	2.16236	[.031]
RLFTFEi(-2)	-1.72914	1.05159	-1.64431	[.100]
RLFTMEi(-2)	4.15880	1.31676	3.15836	[.002]
DRLHPi(-1)	.483251	1.33561	1.33561	[.182]
PPi*LRER(-1)	11.7636	2.59309	4.53651	[.000]
PBi* DMLR(-3)	.401458	.311627	1.28826	[.198]
PBi* DLRHPUK(-2)	-9.55019	5.91511	-1.61454	[.106]
Northern	.950074	.151030	6.29062	[.000]
North-West	.640036	.103213	6.20110	[.000]
Yorkshire & Humberside	.547029	.095909	5.70362	[.000]
West Midlands	.584355	.116535	5.01440	[.000]
East Midlands	.346760	.105840	3.27627	[.001]
East Anglia	-.054005	.117738	-.458687	[.646]
South East	-.964834	.140852	-6.84996	[.000]
South West	.075347	.093692	.804203	[.421]
Wales	.666571	.107626	6.19340	[.000]
Scotland	.537489	.128157	4.19398	[.000]

Number of Observations	16	**Sample Period**	1972-1987
Equation Standard Error	0.244243	**Autocorrelation Test (F)**	0.70 [0.49]
Heteroskedasticity Test (F)	0.93 [0.34]	**Jarque-Bera Normality**	31.14 [0.00]

Notes:
Dependent Variable is DRUi, that is, the change in the relative unemployment rate.
Estimation is by SUR in TSP (Hall, 1996). Equation standard error is the unweighted average of all the residuals.

4. Conclusions

In this paper we have analysed panel data for 1972-1995 on relative regional earnings for full-time men and women and part-time women and for the relative unemployment rates for the ten standard regions of Great Britain. The earnings data come from the New Earnings Survey and, for full-time workers, refer to gross weekly earnings of workers not affected by absence.

Regarding the earnings - unemployment relationship, a negative long-run effect for log unemployment on log earnings as in Sargan (1964) and Blanchflower and Oswald (1994) was confirmed for full-time men. A further split between manuals and non-manuals showed there to be no effect for the latter and a coefficient of around -0.07 for manual men. For full-time women, there is also no evidence of a negative long-run effect though rises in relative regional unemployment rates are associated with declines in relative regional full-time earnings of women. For part-time women, there is a strong positive association between relative regional earnings and relative regional unemployment rates. This may, in part, reflect the unemployment trap: if male partners are unemployed, the effective marginal tax rate on the earnings of women at low levels of earnings can be close to one as means-tested benefits are withdrawn. Thus, women's participation rates, particularly at lower levels of earnings, are likely to drop with a rise in the men's unemployment rate. With fewer women prepared to work, wages rates need to be higher. There is then also likely to be a composition effect: those women who continue to participate are those with higher earnings, driving up average observed part-time women's earnings. Finally, since weekly part-time earnings are quite sensitive to the number of hours worked, the positive association with unemployment also suggests that those part-time women who participate work longer hours to replace lost men's earnings.

The wage-curve, as Blanchflower and Oswald term it, is thus far from a universal phenomenon even in the British labour market. Moreover, our evidence is that the failure to instrument the current unemployment rate results in over-estimates of the negative effect of unemployment on earnings.

The second objective of the paper was to investigate housing market effects on earnings. Our findings suggest a long-run coefficient of around 0.07 for both full-time men and women of relative regional house prices on relative regional earnings. Though the effects are highly significant, they contrast with a coefficient of 0.15 estimated by Blackaby and Manning (1992) on data for full-time men for 1972-1988. Differences in specification account for the differences. One important variable in our model is the interaction between an indicator of financial liberalization and the proportion of employment in banking and financial services. This rose in the 1980s. Since the South East had the biggest (and rising) share of employment in banking and financial services, the rise in this variable broadly matches the relative rise in house prices in the South East. Omission of this variable raises our estimate of the long-run house price effect for full-time men's earnings from 0.07 to 0.11.

Another important interaction effect in our model is between the real exchange rate and the proportion of employment in the production sector. Omitting this effect further raises the estimated house price effect on men's earnings. These results suggest the possibility that Blackaby and Manning (1992) may have over-estimated the house price effects by omitting more fundamental variables which help determine both house prices and earnings.

Further disaggregation of men into non-manuals and manuals reveals a striking difference in the house price effects: 0.12 for non-manuals and only around 0.02 for manual men. These appear to reflect the known differences in migration rates and in rates of owner-occupation between the two groups of workers. These house price effects on wages exclude effects operating via national bargaining which tend to be eliminated in taking regional deviations. At the macroeconomic level, the effects are therefore likely to be larger.

Many studies have examined convergence in incomes between regions, usually by looking at convergence of GDP per worker (see Quah, 1996). Although regional convergence is not the main focus of this paper, our results make three contributions to the literature. First, we find significant equilibrium corrections in our models of earnings, which imply that earnings return quickly to their equilibrium levels. For the UK, we suggest that regions are usually fairly close their steady-states, but that these steady-states are fairly different. Since we have twenty four annual observations in our panel and also find significant and large coefficients on the equilibrium correction terms, it seems unlikely that our conclusions are being driven to any appreciable extent by finite sample bias.

Second, we conclude that the determinants of the steady-state levels of relative earnings are different for men than they are for women in the UK. Notably, the labour market effects and the effect of international competitiveness work quite differently for men’s full-time earnings than for women’s full-time earnings. This suggests that regional convergence studies that look at total earnings may be misleading.

Third, this paper has used data on earnings from the New Earnings Survey. It is very likely that the behaviour of other components of personal income will have a different regional pattern, for example, social security transfers. But this pattern is not likely to reflect regional differences in productivity as clearly as the earnings data. Furthermore, the data from the New Earnings Survey are on a place of employment basis, which seems to be the obvious location for regional convergence. We would therefore argue that studies of regional convergence in the UK that look at GDP per capita are likely to be misleading because they aggregate men’s and women’s earnings together, because they include components of income that do not reflect differences in productivity, and because the data in the regional accounts are on a place of residence basis.

Regarding the determination of regional differences in unemployment rates, we find a strong positive effect of lagged earnings on unemployment, particularly for men’s earnings. Thus, regions with higher labour costs, tend, other things being equal, to suffer higher unemployment. Job migration may well be a significant aspect of this tendency.

Job migration may also help to explain the <u>absence</u> of a negative long-run relative house price effect on relative unemployment rates. Note that since net migration rates to a region are strongly sensitive to regional house price differentials, one might have expected a negative effect of regional house price differentials on unemployment rate differentials. Net job migration rates, on which we lack data, are likely to respond negatively to high relative regional land costs, which will be correlated with house prices, and thus offset the effect on unemployment of inter-regional migration.

Furthermore, we know that non-manual workers have higher migration rates and these are more likely to be sensitive to house prices. They also have low unemployment rates and so have little impact on average unemployment rates for all workers. Indeed, if manual workers and non-manual workers are joint inputs with production, the net migration of non-manuals from a region in response to high relative house prices may destroy jobs or impede job creation for manual workers to whom the observed unemployment rates are particularly sensitive.

These interpretations are consistent with the striking differences in the effect of relative house prices on relative earnings discussed above.

Our unemployment equation, like our earnings equations for manual men and for women shows a strong effect from the lagged real exchange rate interacted with the proportion of employment in the production sector. There also appear to be effects at somewhat longer lags from interactions with the proportion of employment in the banking and financial services sector between past changes in interest rates and past changes of real UK house prices.

We also investigated the connection between high levels of owner-occupation and high levels of unemployment emphasized on theoretical and empirical grounds by Oswald (1998). However, we were unable to find significant effects either in our most general or our most parsimonious specification of our model of regional unemployment rates.

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Data Appendix

Descriptions of Variables

$LFTME_i$ Log real average weekly earnings of full-time male employees. Source: New Earnings Survey.
$LFTFE_i$ Log real average weekly earnings of full-time female employees. Source: New Earnings Survey.
$LFTFE_i$ Log real average weekly earnings of part-time female employees. Source: New Earnings Survey.
U_i Regional unemployment rate. Source: Employment Gazette, various issues.
UH_i Regional unemployment rate, fitted value from unemployment equation.
PP_i Share of production workers in total employment. Source: Office for National Statistics, Earnings and Employment Division.
PB_i Share of banking, finance and real estate workers in total employment. Source: Office for National Statistics, Earnings and Employment Division.
$PFPT_i$ Share of female part-timers in female total workers. Source: Office for National Statistics, Earnings and Employment Division.
LHP_i Log mix-adjusted second-hand house prices. Source: Department of Environment, Transport and the Regions.
DLRHPUK The change in log mix-adjusted real UK house prices. Source: Department of Environment, Transport and the Regions.
$LPOO_i$ Log percentage of owner-occupiers. Source: Department of Environment, Transport and the Regions.
WRR_i Average mortage debt to income ratio in region (relative to GB) weighted by the mortgage interest rate. Source: Cameron and Muellbauer (1998).
RER The real exchange rate. Source: Economic Trends
DMLR Change in UK bank base rates. Source: Financial Statistics.
AFLIB Financial liberalisation dummy, normalised to between 0 and 1, where 1 represents full liberalisation. Source: Muellbauer and Murphy, 1995.

Notes: The following convention for variable names is used in the paper. For example, u_{it} is the regional unemployment rate, ru_{it} is the regional unemployment rate minus the GB unemployment rate, lru_{it} is the regional log unemployment rate minus the GB log unemployment rate

Data Appendix Table A1 Summary Statistics for South East Variables

Variable Name	Mean	Std Deviation	Minimum	Maximum
RLFTMESE	0.106	0.033	0.067	0.153
RLFTFESE	0.109	0.024	0.075	0.143
RLPTFESE	0.087	0.018	0.053	0.117
DRUSE	0.043	0.328	-0.725	0.650
RUSE	-1.628	0.945	-2.850	-0.050
RPPSE	-0.073	0.009	-0.090	-0.064
RPBSE	0.044	0.009	0.035	0.062
RPFPTSE	-0.020	0.016	-0.039	0.004
RLHPSE	0.174	0.072	0.081	0.340
DRLHPSE	0.003	0.039	-0.083	0.060
DLRHPUK	0.004	0.096	-0.174	0.230
RLPOOSE	0.034	0.012	0.004	0.048
WRR	3.734	2.028	1.348	8.253
WDRR	0.013	0.666	-1.241	1.667
LRER	-0.105	0.108	-0.316	0.016
AFLIB	0.518	0.429	0.000	1.000
DMLR	-0.02046	2.47736	-3.55	4.62

Notes: See text for description of variables. Housing market variables in this table are not weighted by the proportion of owner-occupiers in the UK. Sample period is 1972 to 1995.