

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

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**Abstract:** This paper models regional earnings and unemployment in the ten regions of Great Britain between 1972 and 1995, paying particular attention to their interaction and to the important influence of the housing market. In contrast to Blanchard and Katz (1992, 1997) for the United States, we find less persistence in British regional earnings differentials but greater persistence in regional unemployment rates. We find no evidence of a negative effect of the overall unemployment rate on the earnings of men in non-manual, or women in full-time, employment and find a positive effect for women in part-time employment. However, for manual men, we find a significant elasticity of around  $-0.07$ , comparable with Blanchflower and Oswald (1994).

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

**JEL Classification:** C33;E24;R23

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## Non-Technical Summary:

In contrast to the United States, British regions display persistent differences in both earnings and unemployment rates. 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. 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, labour quality differences, differences in private or public capital per worker, or lack of regional amenity.

In this paper we analyse 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.

The use of relative regional earnings and unemployment 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.

The paper also explores the interactions of labour and housing markets in the determination of earnings and unemployment outcomes. The mechanisms at work are diverse and include the influence of housing tenure on labour mobility and migration; the effect of house prices on migration, commuting and firm location, and hence on regional mismatch; and the effect of house prices on the cost of living, on the demand side via wealth effects and perhaps on expectations.

In contrast to Blanchard and Katz (1992, 1997) for the United States, we find less persistence in British regional earnings differentials but greater persistence in regional unemployment rates. Regarding the relationship between earnings and unemployment (the 'wage curve'), 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 manual and non-manual men showed there to be no effect for the latter and a coefficient for manual men of around two-thirds of that estimated by Blanchflower and Oswald. 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. We argue that this reflects a mix of the unemployment trap, hours effects which have a large influence on weekly part-time earnings, and composition effects.

Thus, if the 'wage curve' is formulated in terms of the unemployment rate for all workers, it is not a universal phenomenon in the British labour market. Moreover, our evidence is that any failure to take into account the feedback from earnings to unemployment, results in an over-estimate of the negative effect of unemployment on earnings.

We also investigate the effect of the housing market on earnings. Our findings suggest that a one per cent rise in relative house prices will raise the relative earnings of full-time men by around around 0.075 per cent and those of full-time women by 0.10 per cent. 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. These results suggest the possibility that Blackaby and Manning 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 no significant effect for manual men. These appear to reflect the well-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.

Regarding the determination of regional differences in unemployment rates, we find a strong positive effect of lagged earnings on unemployment. 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 also helps to explain the positive long-run relative house price effect on relative unemployment rates. Net job migration rates are likely to respond negatively to high relative regional land costs, suggesting a positive relative house price effect on relative unemployment rates. We expect the effect of people migration to be in the opposite direction, since people migration is sensitive to relative house prices (Cameron and Muellbauer, 1998). However, there are good reasons to think that the effect of people migration on unemployment is small, since most of regional migration is by non-manual workers and these are more likely to be sensitive to house prices. They also have low unemployment rates and so have less impact on average unemployment rates for all workers. 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. This makes an important contribution to understanding the 'North-South divide' as reflected in low unemployment and higher earnings in the South East and its contiguous regions, particularly in the late 1980s. Our equation predicts that the current over-valuation of sterling is leading to a significant widening of the divide.

There is indirect evidence that from about 1989, job migration rates have tended to increase, which we believe is the results of new and cheaper information technology reducing the need for service sector firms to be located close to their customers. We find significant evidence that this has resulted in a narrowing of long-run regional unemployment differentials and has increased responses of relative unemployment rates to regional differences in labour and housing costs.

Our work has potential implications for the interpretation of the results of Blanchard and Katz (1992 and 1997). It seems likely that their finding of a high degree of wage persistence over time is the result of the omission of important explanatory variables, including the interaction of macroeconomic shocks with regional characteristics, as well as housing market variables. The UK evidence is that housing market fluctuations are far from being simple functions of earnings and employment shocks. For the US, even though housing supplies are more elastic than in the UK, it would be surprising if these fluctuations did not also contribute to regional evolutions.

## 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 (an alternative approach that produces similar results is to use time dummies). 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. 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. 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.

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,<sup>1</sup> the effect of house prices on migration, commuting and firm location, and hence on regional mismatch<sup>2</sup>, the effects of house prices on the cost of living, on the demand side via wealth effects and perhaps on expectations.<sup>3</sup>

Our third aim is to derive explicit one-year ahead forecasting models. Such models have multiple uses, including the derivation of efficient instruments for modelling regional variables such as consumption and house prices. Another major application is to the burgeoning econometric literature on convergence of

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

<sup>2</sup> See Jackman and Savouri (1992) and Cameron and Muellbauer (1998).

incomes between regions and between nations, see Sala-i-Martin (1996) 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 and discusses our econometric method. For the panel data used in this paper, the simplest approach is to restrict the coefficients to be the same across regions using either the dynamic fixed effects estimator (DFE) or the mean group (MG) estimator. Pesaran, Shin and Smith (1999) have recently proposed a pooled mean group (PMG) estimator which allows short run coefficients and error variances to differ across groups in a panel, but constrains the long run coefficients to be identical. This paper estimates models using both a dynamic fixed effects estimator (based on Generalised Least Squares, which allows error variances to differ across groups) and also the Pesaran et al. PMG estimator. Section 4 presents our models of earnings and unemployment for a panel of British regions and compares these results with models based on the previous literature, especially the contributions of Blanchard and Katz, Blackaby and Manning, and Blanchflower and Oswald. Section 5 draws conclusions. The data appendix reviews the sources of the data used in the paper.

## **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) anticipates the later literature on the expectations augmented Phillips curve by emphasising that, in the long-run, workers and firms should 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. Indeed, Sargan emphasizes the long-run solution for the real wage. Such a model can be trivially

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<sup>3</sup> See Bover, Muellbauer and Murphy (1989) and McCormick (1997).

<sup>4</sup> Indeed, Sargan emphasizes the long-run solution for the real wage.

re-parameterised<sup>5</sup> into a real wage model of the type that Layard and Nickell (1986) estimate for the UK, and that 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 work of Layard and Nickell gives good grounds for supposing that unemployment has a negative effect on wages at the aggregate level. At the regional level, a compensating differentials argument suggests 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, labour quality differences, differences in private or public capital per worker, or lack of regional amenity.

The evidence is mixed. For the United States, it appears 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 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.<sup>6</sup> However, as Card (1995) notes, the wage curve may be very different for different groups of workers. In general, younger, less well-educated, less unionised, men are more likely to have a significant unemployment elasticity of pay. 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).

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<sup>5</sup> By adding a term in the rate of change of current inflation.

<sup>6</sup> 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).

For the United States, it has been argued 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). For the United Kingdom, many commentators have suggested that movements of labour play only a 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.

This leads naturally to the role of housing market influences neglected in almost all the literature on national wages and unemployment as well as much of the regional literature (e.g. Blanchard and Katz, 1992 and 1997; Jackman, Layard and Savouri, 1991; Blanchflower and Oswald, 1994 and 1995; Pissarides and McMaster, 1990).<sup>7</sup>

Bover et al. (1989) argue that there are as many as five potential housing market channels of influence on national wages and unemployment. The first is the effect of housing tenure on mobility rates, which is likely to affect regional labour mis-match. In the UK, even after correction for individual, occupational and local characteristics, the highest mobility rates are for individuals in the private rented sector, especially after abolition of rent controls. The lowest rates are for those resident in public housing, particularly when rents are far below market levels, while those in owner-occupation have intermediate mobility rates, see Hughes and McCormick (1987, 2000). The second is the cost of living effect on workers already in the region and also on potential migrants from house and, indirectly, land prices.<sup>8</sup> The third is the cost of location effect from land prices on firms in the region and potential movers to the region. The fourth is a wealth effect on regional spending, in part via the collateral role of housing equity on credit constrained households and small businesses. The fifth is an expectations effect in that future movements of earnings may be capitalized in house and land prices.

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<sup>7</sup> Blanchard and Katz (1992) acknowledge the potential relevance of land prices, p. 24, but confine empirical analysis to a briefly summarised bivariate model of employment and house prices, p.45.

<sup>8</sup> This can give rise to temporarily high regional labour market mismatch associated with temporarily high house price differentials relative to earnings differentials, which constrain net migration to booming regions such as the South East of England in the 1980's. The evidence in Cameron and Muellbauer (1998) is for large relative house price effects in Britain on net interregional migration rates, given earnings and unemployment differentials.



For models of British regional earnings and unemployment differentials, much of the first channel is likely to wash out since changes in tenure structure tend to be national, e.g. the removal of rent controls, the sale of public housing, the rise of owner-occupation. Moreover, while at the national level, higher proportions of market rented housing should be associated with greater mobility and lower average unemployment, this is far from obvious at the regional level. Suppose that in the South East unemployment is unusually low and house prices unusually high, the latter encouraging out-migration and tending to block in-migration of the credit constrained into owner-occupation. A larger market rented sector in the South East will remove some of the in-migration blocks and so bring down unemployment rates in the other regions. It is therefore far from clear that *within a region* there will be a negative association between the tenure proportion in the market rented sector (or a positive association with the owner-occupation rate<sup>9</sup>) and the unemployment rate.

The remaining four channels could all, however, be manifested in relative house price effects on regional earnings and unemployment differentials. The cost of living, wealth and expectations channels imply positive effects on relative earnings from lagged relative house prices, while the cost of location effect partly offsets these. For relative unemployment, the forces are more evenly balanced. The cost of living effect of relative house prices on higher future relative earnings is likely to repel firms but credit constrained potential migrants are also likely to be repelled, leaving the effect on relative unemployment rates ambiguous.<sup>10</sup> The cost of location effect pushes up relative unemployment while the wealth effect is in the reverse. The expectations channel is ambiguous since labour costs and profit expectations have opposite effects on firms' location decisions.

At any rate, it is clear that the relative earnings, employment and unemployment system considered by Blanchard and Katz (1992, p.21), needs considerable extension. Relative house prices

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<sup>9</sup> 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.

<sup>10</sup> In practice, the migration effect is likely to be small since we know that migration rates are lowest among manual workers (see Hughes and McCormick, 1987), who in the 1980s and 1990s, according to the Labour Force Survey, have had unemployment rates twice as high as non-manuals, with even greater discrepancies for men.

should enter all four equations: the labour demand equation, the 'wage curve', the labour force growth equation (largely migration) and the equation governing the evolution of regional demand function intercepts. To complete the new system, two more equations are needed, one for relative house prices (see Muellbauer and Murphy, 1994, for one specification) and one for relative housing stocks, since these are significant determinants of relative house prices. In what follows, we model only relative earnings and relative unemployment rates in a broadly conceived subset of a VAR system. However, another difference from Blanchard and Katz and other previous work is that we make explicit the channels by which key macroeconomic variables have distinct regional impacts.

### **3. Regional Earnings Data and Previous Work**

#### *3.1 Regional Earnings Data*

At the regional level, there are two main sources of data: the Regional Accounts and the New Earnings Survey.<sup>11</sup> 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. It has four advantages as a data source. The first is that the place of employment is arguably more appropriate to 'locate' 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 full-time men and women employees and part-time women employees.

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. Cameron and Muellbauer (2000) provide evidence that between 1979 and 1994, the Regional Accounts data are significantly biased. Finally, the longitudinal element in the NES sample is substantial, with an overlap of between seventy and eighty per cent for all workers in adjacent annual samples.

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<sup>11</sup> 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).

Figure 1 displays the regional deviation of the South East from Great Britain of our three measures of log earnings (NES gross weekly earnings of workers not affected by absence for full-time men, full-time women and part-time women) and the unemployment rate.

### 3.2 *Econometric Method*

In the following section we estimate models of relative regional log earnings for three types of workers and relative regional unemployment rates using data for 1972 to 1995 for the ten regions of Great Britain. The models take the following general form<sup>12</sup>:

$$(1) \quad \Delta y_{it} = \mathbf{a}_i \left( \mathbf{q}_i + \sum_{j=1}^k \mathbf{b}_j x_{jit-1} - y_{it-1} \right) + \sum_{s=1}^m \mathbf{g}_s \Delta y_{it-s} + \sum_{s=1}^m \sum_{j=1}^k \mathbf{g}_{js} \Delta x_{jit-s}$$

Thus, the long-run solution has the form

$$(2) \quad y_i = \mathbf{q}_i + \sum_{j=1}^k \mathbf{b}_j x_{ji}$$

where, the four  $y$  variables are, respectively, relative log-earnings for full-time men and women and part-time women<sup>13</sup>, and the unemployment rate in each region minus the rate in Great Britain. For each  $y$  variable, the  $x$  variables include the relevant complementary elements of the other three  $y$  variables; relative house prices weighted by owner-occupation in the UK for the earnings equations and unweighted for the unemployment equation; the relative proportion of employment in the production sector and its interaction with the log real exchange rate; the relative proportion of employment in banking and financial services and its interactions with bank base rate, a proxy for financial liberalisation, real UK house prices, and with a trend beginning in 1989, reflecting a shift in technology further discussed in section 4.2; the relative proportion of women in part-time employment; and the relative proportion of manual workers. The  $x$  variables also include interactions between the relative mortgage debt to earnings ratio with the average mortgage interest rate. All the  $x$  variables test as being  $I(1)$  in levels and  $I(0)$  in differences using the panel unit root test suggested by Im, Pesaran and Shin (1997), as do three of the  $y$  variables (with relative log earnings of part-time women being a borderline case), see data appendix.

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<sup>12</sup> Also, in the earnings equations, we test the effect of the change in the contemporaneous relative unemployment rates, and in the unemployment equation we test the effect of the change in contemporaneous relative earnings (instrumenting appropriately).

Two features of the  $\Delta x$  variables should be noted. Where  $x_{jit-1}$  appears in the form  $w_{it-1} x_{jit-1}^*$ , where  $w_{it-1}$  is a regional weight and  $\sum_i w_{it-1} = 0$ ,  $\Delta x_{jit-s}$  actually denotes  $w_{it-1} \Delta x_{jit-s}^*$ . Otherwise  $\Delta x_{jit-s}$  has the conventional meaning. Second, given the sophistication of the banking and financial services sector and its sensitivity to interest rates, the set of  $\Delta x$  variables includes last year's expectation of the contemporaneous change in bank base rate, weighted by the regional deviation of the sector's lagged employment share.

The maximum lag considered in all four equations was  $m=2$  and 'general to simple' testing procedures were used to select parsimonious specifications. The estimation technique is two step GLS using the 'seemingly unrelated regression' option in the Hall et al. (1995) Time Series Processor (TSP) for the first step (see Zellner, 1962), computing the covariance matrix of residuals across the ten regions and fixing this for GLS at the second step. Where relevant, the covariance matrix is also corrected for instrumented contemporaneous effects. Where F-tests are reported these hold constant the covariance matrix of the null hypothesis against the more general alternative. The equation standard errors reported are the square root of the average of the squared residuals, corrected for degrees of freedom.

We also consider several kinds of generalisations of these models. The first concerns the de-meaning procedure, i.e. setting up our models in terms of regional deviations from the means for Great Britain. There are plausible arguments to suggest that regional interactions are different for contiguous than for more widely spatially separated regions. These arise, for example, in the context of migration and commuting flows, see Jackman and Savouri (1992) and Cameron and Muellbauer (1998). In the latter paper, contiguity effects are modelled in detail, using employment weighted average values of the variables of regions contiguous to region  $i$ .

The long-run solution for relative log earnings and the unemployment rate, equation (2), can be rewritten as

$$(3) \quad y_i = Y_i - Y_{GB} = \mathbf{q}_i + \sum_{j=1}^k \mathbf{b}_j (X_{ji} - X_{jGB})$$

where  $x_{ji} = X_{ji} - X_{jGB}$ . Formulating an analogous equation for deviations from contiguous regions gives

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<sup>13</sup> These are defined as  $\ln e_i - \ln e_{GB}$ , where  $e_i$  and  $e_{GB}$  are the average earnings of the group in region  $i$  and Great Britain respectively.

$$(4) \quad y_{ci} = Y_i - Y_{ci} = \mathbf{q}_i^* + \sum_{j=1}^k \mathbf{b}_j^* (X_{ji} - X_{jci})$$

where  $ci$  refers to the average for the regions contiguous to region  $i$ . Weighting the two equations by  $(1-\alpha)$  and  $\alpha$  respectively gives a formulation nesting both. We generalise equation (1) by replacing the error correction term based on a one year lag of (3) by one based on a one year lag of this nested formulation.<sup>14</sup> However, for all four dependent variables, the estimated values of  $\alpha$  and the  $\mathbf{Ib}_j^*$  terms are jointly insignificant, suggesting that this formulation of contiguity effects offers no significant improvement for either the earnings or the unemployment equations. However, note that our formulation imposes the same short-run dynamics, irrespective of contiguity.

A second generalisation is to relax the restriction that the speeds of adjustment and the short-run dynamics are identical in each region, the pooled mean group (PMG) estimator as recommended by Pesaran, Shin and Smith (1999). This generalizes equation (1) by allowing  $\mathbf{a}_1, \mathbf{g}$ , and  $\mathbf{g}_{js}$  but not  $\mathbf{b}_{js}$ , albeit in the parsimonious forms of the models, to differ across regions. The PMG  $\mathbf{a}_1, \mathbf{g}$ , and  $\mathbf{g}_{js}$  coefficients are calculated from the simple average of the region coefficients and the estimated standard errors reported are the standard deviations of the coefficients across regions. We also consider a third generalisation where all parameters are region-specific.

We model the deviation of regional log earnings from the log of average earnings for Great Britain rather than the deviation between regional log earnings from their predicted value when the region's industrial structure is applied to national average earnings by industry as in Blackaby and Manning (1992). There are two reasons for this. One is that we want to derive a one-year ahead forecasting model for regional differentials. The other is that we prefer a more 'structural' model in which the interactions of macroeconomic variables with regional economic structures are made explicit.

We exclude long-term unemployment, being unable to find a significant effect in any of the earnings equations. 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

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<sup>14</sup> Where the  $\mathbf{q}_i^*$  are not identifiable.

Nickell (1986) and in most of the country studies in Layard, Nickell and Jackman (1991). However, short-term unemployment effects fit better in terms of changes in levels rather than changes in logs.

## **4. Empirical Results**

### *4.1 Three Models of Relative Regional Earnings*

#### *Full-time Men's Relative Earnings*

The men's earnings version of equation (1) was reduced to the more parsimonious specification reported in Table 1, column 1. There is a significant equilibrium-correction term with  $a_1$  around 0.45 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 has a significant and negative effect with a long-run coefficient around -0.034, one third of the -0.1 figure claimed by Blanchflower and Oswald as a robust order of magnitude of the slope of the wage curve. Contemporaneous changes in unemployment were instrumented using the forecast changes from the model described in the next section, but were found to be insignificant. 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 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 than if relative log house prices are unweighted. At a UK owner-occupation proportion of 0.68, the long-run effect of relative log house prices on men's relative full-time earnings is 0.075. 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 earnings 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. Last year's expectation of a rise in this year's bank base rate, interacted with the relative proportion of employment in banking and financial services, has a negative effect on relative earnings. There is no evidence of a spill-over effect from women's relative regional earnings.

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, or that a high proportion of part-time women signals weak labour demand. 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. There is no evidence for men's relative earnings of an effect from a post-1989 trend, interacted with either the proportion of employment in banking and financial services or a range of other variables.

Column 2 of Table 1 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,  $F_{80,138}=1.10$  [ $P=0.31$ ].

Having found no significant effect for the forecast from the unemployment equation, we can compare these results with those using the actual unemployment data, see column 3 of Table 1. The coefficient on the contemporaneous change in the unemployment rate is now significantly negative, ( $t=6.1$ ), evidence consistent with a notable endogeneity bias. Also the estimated long-run slope of the wage curve is  $-0.050$  as opposed to  $-0.034$  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).

We also test the Table 1, column 1, baseline model against the pooled mean groups estimator of Pesaran et al. (1999), where speeds of adjustment and short-run dynamics are allowed to differ across regions, as shown in column 4. A test of the baseline model against this alternative, given the error covariance matrix of the baseline model, cannot reject the baseline model,  $F_{35,183}=1.35$  [P=0.11]. A test of the baseline model against the even more general alternative that all parameters are permitted to vary by region also cannot reject the baseline model,  $F_{108,110}=1.03$  [P=0.44].

One of the most interesting results from Table 1 is that the long-run coefficient on relative house prices, 0.075 at a UK owner-occupation proportion of 0.68, 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 their omission of the interaction between financial liberalization and the proportion of employment in the banking sector. 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 influence both house prices and earnings.<sup>15</sup>

For non-manual men's full-time earnings, we expect to find stronger house price effects than for men's full-time earnings, as discussed in section 2. These expectations are confirmed by the empirical findings. The long-run coefficient on relative house prices at a UK owner-occupation proportion of 0.68 is 0.12 for non-manual men compared with 0.075 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 Blanchflower and Oswald figure of -0.1 and

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<sup>15</sup> Blackaby and Manning (1990, 1992) model  $wp$ , the log deviation between men's full-time gross weekly NES earnings in each region from 'predicted' earnings based on applying national earnings by industry to the region's industrial structure. Over the period 1972-1988, they find a long-run solution of  $wp = -0.0045(ru - ru52) + 0.15rh + 0.15lh + \text{fixed effects}$ , where  $ru$  is the regional differential in the unemployment rate,  $ru52$  is the regional differential in the unemployment rate of those out of work for more than 52 weeks,  $rh$  is a relative regional cost of living index and  $lh$  is the relative regional house price index. They instrument the contemporaneous effect of changes in the relative unemployment rate and take the covariance matrix of errors across regions to be the identity matrix. Their equation standard error is 0.00704.



close to the estimate of Jackman, Layard, and Savouri (1992) of -0.067.<sup>16</sup> There is no significant long-run house price effect.

### *Full-time Women's Relative Earnings*

For full-time women's earnings, we formulated a model according to equation (1) and again followed a general to simple model selection procedure, see Table 2. In contrast to full-time men's earnings, we could find no influence from mortgage costs (though the point estimate is positive,  $t=1.3$ ), the proportion of women's employment that is part-time, and no wage curve. However, the contemporaneous change in the unemployment rate (appropriately instrumented) has a very significant negative effect, presumably capturing demand shocks. The speed of adjustment is 0.46 and there is also a negative response to the previous year's change in relative log earnings for full-time women.

There is no spill-over effect from men's relative regional earnings. As in the earnings equation for men, the proportion of employment in the production sector (interacted with the real exchange rate, but at a lag of two years) has a negative effect on relative women's earnings. The interaction of financial liberalization with the proportion of employment in banking has a similarly large positive effect on women's as on men's earnings. A trend beginning in 1989, representing new technology, interacted with the regional deviation in the proportion of employment in the banking and financial services sector, has a positive effect of relative earnings of full-time women, unlike its insignificant effect on the relative earnings of men.<sup>17</sup>

Finally, although the relative house price effect is 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 0.10 at an owner-occupation proportion of 0.68, even larger than for full-time men.

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<sup>16</sup> Jackman, Layard and Savouri (1991) estimate a model of the log hourly wages of full-time manual men in the ten regions of Great Britain relative to the national average for 1967-1987, using data from the Department of Employment October survey. Their explanatory variables are the relative log unemployment rates and relative proportions of the long-term unemployed, as well as regional fixed effects and regional time trends. Estimation is by OLS and the equation standard error is 0.0070. They do not control for the contemporaneous correlation between unemployment shocks and the error term.

Table 2, column 2, 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 cannot be rejected,  $F_{80,120}=1.11$  [P=0.31]

Table 2, column 3, shows results for the pooled mean groups version of the model allowing speeds of adjustment and dynamics to differ across regions. The baseline model (column 1) cannot be rejected against this alternative,  $F_{54,146}=0.82$  [P=0.54], or against the alternative that all coefficients differ across regions,  $F_{90,110}=1.25$  [P=0.13].

### *Part-time Women's Relative Earnings*

The general equation for relative earnings of part-time women again follows equation (1), included lagged effects both from full-time men's and women's earnings. Of these only the positive effect from last year's change in the relative earnings of full-time women survives in the parsimonious specification shown in Table 3, column 1.

There are some striking differences from 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. 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. There is no effect from the current change in the unemployment rate (whether instrumented or not) but a negative response to the lagged change.

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<sup>17</sup> To the extent that the new technology increases job migration, we might expect it to decrease the extent to which higher house prices translate into higher earnings and to reduce the impact of the fixed effects, As with the full-time men's earnings equation, we find no significant effect of this kind for women.

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. In part, this is likely to reflect an hours response to financial pressure.

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 at a two year lag, 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 change in relative proportion of employment in the banking sector has a positive effect on relative part-time earnings.

Table 3, column 2, suggests that the parameter estimates are fairly stable over the 1974-1987 subsample and an F-test confirms this result,  $F_{79,142}=0.95$  [P=0.60]. Table 3, column 3, shows results for the pooled mean groups version of the model allowing speeds of adjustment and dynamics to differ across regions. The baseline model (column 1) cannot be rejected against this alternative,  $F_{27,194}=1.23$  [P=0.21], or against the alternative that all coefficients differ across regions,  $F_{81,140}=1.12$  [P=0.28].

## 4.2 *A Model of Relative Regional Unemployment*

In the long run, we expect the relative unemployment rate to increase with the relative earnings of full-time men and women, via the effect on regional labour demand, other things being equal. Section 2 argues that the effect of relative house prices is, in principle, ambiguous, but on balance, more likely to raise relative unemployment in the long-run.

Relative unemployment rates can also be expected to be affected in the long-run by industrial structure and its interaction with macroeconomic variables such as international competitiveness (measured by the real exchange rate), financial liberalisation, UK real house prices, and changes in technology. In the long-run, the share of employment in production industries has declined, while that in services, particularly financial services, has risen across OECD countries. However, since the late 1980s, the fall in the price of computing power, the availability of new software and, probably, lower telephone charges, has led to a rise in on-line financial service provision and in 'call centres', where large numbers of employees are engaged in

selling and providing information or responding to customer requests. It has therefore become less important for these activities to be located close to the customers. One way of representing this effect is to interact the proportion of employment in banking and financial services in region  $i$  with the national proportion of employment in call centres. Such data as exist on the latter<sup>18</sup> suggest an approximately linear trend from 1989 to 1995 and we use this trend interaction in one formulation. Such a linear trend is unlikely to persist indefinitely, and with more data it would be preferable to estimate a logistic trend.

An alternative formulation of the interaction effect appears at least as plausible. The changes in technology discussed above suggest greater job mobility: a bigger rôle for relative labour and housing cost variables in the location of businesses and a smaller rôle for the locational advantages represented by the fixed effects, the  $\gamma_i$ , in our model. We capture these effects in a single parameter, which represents the proportionate fall beginning in 1989 of the  $\gamma_i$ , and the proportionate rise in the coefficients on relative men's and women's earnings and relative house prices.

After a series of simplifications, we arrive at two parsimonious models with these two respective interactions. To check parameter stability we estimate the models for 1972-1987 using the same covariance matrix as for the full sample. An F-test of no parameter shift between the short and long samples gives  $F_{80,137}=1.42$  [ $P=0.036$ ] for the first model, and  $F_{80,137}=1.42$  [ $P=0.036$ ] for the second model, both of which are significant at the 5 per cent level, casting doubt on the model's validity.

As noted in the introduction to section 4, we then considered several kinds of generalisations of these models. The first concerned the de-meaning procedure, i.e. testing for contiguity effects, but these were found to be insignificant. The second generalisation is to relax the restriction that the speeds of adjustment and the short-run dynamics are identical in each region, see Table 4, column 4. We test the baseline models against this generalization holding constant the covariance matrix of errors in the two baseline models,  $F_{34,181}=1.46$  [ $P=0.058$ ], and  $F_{34,181}=1.46$  [ $P=0.057$ ]. These p-values suggest that there is something to be learnt from the more general specification. The most striking feature of the results is the significantly higher speed of adjustment for Wales and Scotland than for the other regions.

This suggests a slight generalization of the baseline model where  $\mathbf{a}_1$  is allowed to differ for Wales and Scotland but is the same for the other eight regions, but otherwise all the other short-run

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<sup>18</sup> See the article by Seamus Milne in The Guardian, 26 November 1999.

dynamics are restricted to be the same across regions. The results for the two variants of this extended base-line model are shown in columns 1 and 2 of Table 4.

The new base-line model in column 1 shows a speed of adjustment,  $\hat{\alpha}_1 = 0.16$ , but 0.35 in Wales and 0.28 in Scotland. There is significant persistence in relative unemployment changes,  $\hat{\xi} = 0.35$ . Of the  $x_j$  variables listed above, only the relative proportion of manual workers and the interaction of the relative size of the banking and financial services sector with the financial liberalisation proxy proved insignificant. To interpret some of the results, a rise of 1 per cent in relative earnings of full-time men and women eventually raises the relative unemployment rate by 0.16 percentage points, other things being equal. A rise of 10 per cent in relative house prices eventually raises the relative unemployment rate by 0.42 percentage points, though the lags are long. A 10 per cent rise in the real exchange rate eventually raises the relative unemployment rate in a region such as the West Midlands, which in 1980 had a percentage of employment in the production sector 9.6 per cent higher than in Great Britain, by 0.80 percentage points.

We turn now to the most controversial finding of this version of the model. Up to 1988, when the percentage of employment in the banking and financial services sector in the South East was 5 percent greater than in Great Britain, this accounted in our model (column 1 of Table 4) for a 4.0 percentage points lower unemployment rate. But from 1989, this benefit decays, so that, by 1995, only a 0.5 percentage point unemployment rate advantage accrues to the South East. These effects are surprisingly large.

Column 2 presents the alternative formulation where the trend beginning in 1989 results in a proportional shift in the fixed effects and the relative earnings and housing cost effects. The parameter estimate suggests a 9.4 percent per annum decline in the fixed effects beginning in 1989 and a corresponding rise in the relative earnings and housing cost effects. The speeds of adjustment are a little higher than in the model shown in column 1 but, with some exceptions, the other parameter estimates are very similar. One exception is the coefficient on the proportion of employment in banking and insurance which is around 45 per cent of the column 1 estimate and is now not subject to a regime shift. The coefficients for relative house price and relative earnings of women fall compared with column 1.

The estimated short-run dynamics of the two models are very similar. There are two significant interest rate effects. The first weights the two-year rise in the mortgage interest rate, lagged two years, by last year's difference from Great Britain in the mortgage debt to earnings ratio in each region. This

suggests there are substantial lags before a rise in mortgage interest rates feeds, via cash flows and housing collateral values, into lower spending and higher unemployment. The other interest rate effect weights the expected value of the current change in bank base rate by the relative proportion of employment in banking and financial services. This suggests that hiring decisions in this sector respond quickly to current circumstances or have a forward looking element. The weighted interest change lagged one year is not significant using either set of weights, though the point estimate is also positive.

Parameter stability tests for 1972-87 compared with 1972-1995 for the two variants give  $F_{80,135}=1.30$  [P=0.09] for the column 1 variant and  $F_{80,135}=1.28$  [P=0.10] for the column 2 variant, suggesting parameter stability can now be accepted at the 5 per cent level. The estimates for the column 2 variant for 1972-87 are shown in column 3.

Taking the column 2 specification as the baseline, we then relax the restrictions across regions on  $\mathbf{a}_1, \mathbf{g}, \mathbf{g}_{js}$  but not  $\mathbf{b}_{js}$ , see column 4. These results suggest that the column 2 specification gives estimates which well represent the averages of the more general pooled mean groups estimator. An F-test of the column 2 specification, fixing its covariance matrix of errors, against the more general alternative, gives  $F_{34,181}=1.08$  [P=0.36], implying that the column 2 specification can be accepted. Finally, we carried out the analogous test of the column 2 model against the model where all parameters differ across regions. This gives  $F_{115,100}=1.03$  [P=0.43] implying the parsimonious model can be accepted against the most general alternative.

Why the speed of adjustment should be higher for Wales and Scotland remains a puzzle. One hypothesis is that regional policy has been more active in these regions and has reacted against rises in unemployment rates in these regions, bringing a speedier convergence to long-run levels than for the remaining regions.

## 5. 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 manual and non-manual men 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. However, note that we have used the unemployment rate for all workers in these models. This will tend to under-estimate the size of the wage curve when the elasticity of group unemployment with respect to aggregate unemployment is less than one. Formally, if  $a$  is the wage curve for group unemployment, and  $e$  is the elasticity of group unemployment with respect to aggregate unemployment, then the wage curve using aggregate unemployment will be  $ae$ , conditional on other variables.

For part-time women, there is a strong positive association between relative regional earnings and relative regional unemployment rates. We argued in section 4.1 that this reflected a mix of the unemployment trap, hours effects which have a large influence on weekly part-time earnings, and composition effects.

Thus, if the wage curve is formulated in terms of the unemployment rate for all workers, it is not a universal phenomenon in the British labour market. Moreover, our evidence is that any failure to instrument the current unemployment rate results in an over-estimate 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 0.075 for full-time men and around 0.10 for full-time 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. As noted in section 4.1, 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 manual men reveals a striking difference in the house price effects: 0.12 for non-manuals and no significant effect 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). 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.

Regarding the determination of regional differences in unemployment rates, we find a strong positive effect of lagged earnings on unemployment. 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 also helps to explain the positive long-run relative house price effect on relative unemployment rates. 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, suggesting a positive relative house price effect on relative unemployment rates. We expect the effect of people migration to be in the opposite direction, since people migration is sensitive to relative house prices (see Cameron and Muellbauer, 1998). However, there are good reasons to think that the effect on unemployment is small. Most of regional migration is by non-manual workers and these are more likely to be sensitive to house prices. They also have low unemployment rates and so have less impact on average unemployment rates for all workers. Indeed, if manual workers and non-manual workers are joint inputs in production, the net migration of non-manuals from a region in response to high relative house prices in that region may destroy jobs or impede job creation for manual workers in that region. 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. This makes an important contribution to understanding the 'North-South divide' as reflected in low unemployment and higher earnings in the South East and its contiguous regions, particularly in the late 1980s and resurfacing recently. There also appear to be effects from interactions between the proportion of employment in the banking and financial services sector with current changes in interest rates and past levels of real UK house prices, which are associated with cyclical fluctuations in this sector.

There is indirect evidence that from about 1989, job migration rates have tended to increase, which we believe is the results of new and cheaper information technology reducing the need for service sector firms to be located close to their customers. We find significant evidence that this has resulted in a narrowing of long-run regional unemployment differentials and has increased responses of relative unemployment rates to regional differences in labour and housing costs.

We also investigated the connection between high levels of owner-occupation (or low levels of private-tenancy) and high levels of unemployment emphasized on theoretical and empirical grounds by Oswald (1998). However, as forewarned in section 2, we were unable to find significant effects either in our most general or our most parsimonious specification of our model of regional unemployment rates.

Our work has potential implications for the interpretation of the results of Blanchard and Katz (1992 and 1997). One is that omitted variable bias is probably a substantial contributor to their finding of a high degree of wage persistence. In our model, we also find high levels of persistence when relative house price effects and interactions of macroeconomic variables with regional characteristics are omitted. More generally, the UK evidence is that housing market fluctuations are far from being simple functions of earnings and employment shocks (see Muellbauer and Murphy, 1994, 1997). For the US, even though housing supplies are more elastic than in the UK, it would be surprising if these fluctuations did not also contribute to regional evolutions.

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## Data Appendix

### Descriptions of Variables

- LFTME<sub>i</sub>** Log real average weekly earnings of full-time men. Source: New Earnings Survey.
- LFTFE<sub>i</sub>** Log real average weekly earnings of full-time women. Source: New Earnings Survey.
- LPTFE<sub>i</sub>** Log real average weekly earnings of part-time women. Source: New Earnings Survey.
- U<sub>i</sub>** Regional unemployment rate/100(all workers, claimant count). Source: Employment Gazette, various issues.
- UH<sub>i</sub>** Regional unemployment rate/100(all workers, claimant count), fitted value from unemployment equation.
- PP<sub>i</sub>** Share of production workers in total employment. Source: Office for National Statistics, Earnings and Employment Division.
- PB<sub>i</sub>** Share of banking, finance and real estate workers in total employment. Source: Office for National Statistics, Earnings and Employment Division.
- PFPT<sub>i</sub>** Share of part-time women in total women workers. Source: Office for National Statistics, Earnings and Employment Division.
- LHP<sub>i</sub>** Log mix-adjusted second-hand house prices. Source: Department of Environment, Transport and the Regions.
- LRHPUK** Log mix-adjusted real UK house prices. Source: Department of Environment, Transport and the Regions.
- POO** Proportion of owner-occupiers in UK. Source: Housing and Construction Statistics, Department of Environment, Transport and the Regions.
- WRR<sub>i</sub>** Average UK Building Society mortgage interest rate weighted by mortgage debt to income ratio in region (relative to GB). Source: Cameron and Muellbauer (1998).
- RER** The real exchange rate. Source: Economic Trends
- DMLRH** Fitted value for change in UK bank base rates from a simple autoregressive forecasting model with US 3 month Treasury Bill interest rate, inflation, the trade balance to GDP ratio and the change in personal income.
- AFLIB** Financial liberalisation proxy defined as the moving average of the proportion of first-time buyers (excluding buyers at subsidized prices of social housing by former tenants) obtaining mortgages of ninety per cent or more. This rose from an average of 0.319 between 1972 and 1977 to 0.675 between 1985 and 1990. The variable is defined as the deviation from its mean over 1972 to 1995. Source: special tabulations from the 5 per cent Survey of Building Society Mortgages, DETR.
- Trend89** Split time trend which is zero until 1988, one in 1989, two in 1990, etc.

Table A1  
Dynamic Panel Data Unit Root Tests

	Levels	Differences
LFTME	2.51	11.51
LFTFE	1.58	14.41
LPTFE	5.91	13.77
U	0.68	5.99
PP	2.41	10.10
PB	2.46	9.92
PFPT	2.96	12.54
LHP	2.20	8.86
WRR	1.11	7.87

*Notes:*

5% Critical value (approx.): 3.96 from Im, Pesaran and Shin (1997, table 3). Sample Period is 1972-1995.

Table 1  
Full-Time Men's Relative Earnings Model

Parameter		Model 1 Baseline	Model 2 Short Sample	Model 3 Actual DU (it)	Model 4 PMG
Rel. log men's full-time earnings (it-1)	$\alpha$	0.446 (.033)	0.487 (.055)	0.449 (.032)	0.515 (.136)
Rel. log unemployment (it-1)	$\beta$	-0.034 (.009)	-0.042 (.011)	-0.050 (.009)	-0.032 (.009)
DRel. unemployment/100 (it) (not instrumented)	?			-0.742 (0.121)	
DRel. unemployment/100 (it-1)	?	-0.397 (.123)	-0.345 (.180)	-0.703 (.116)	-0.113 (.977)
Rel. log house prices(it-1)*UK prop. owner-occupied(t-1)	$\beta$	0.114 (.015)	0.120 (.033)	0.136 (.157)	0.119 (.017)
Weighted av. Building Society mortgage rate /100 (it-1)	$\beta$	0.283 (.106)	0.252 (.206)	0.461 (.108)	0.252 (.133)
Weighted D av. Building Society mortgage rate /100 (it-2)	?	-0.771 (.104)	-0.337 (.224)	-0.801 (.091)	-0.967 (1.282)
Rel. prop. Part-time women in total women(it-1)	$\beta$	-0.700 (.120)	-0.656 (.156)	-0.691 (.123)	-0.699 (.130)
Rel. prop. Production workers(it-1)	$\beta$	0.036 (.083)	0.013 (.135)	0.129 (.083)	0.111 (.102)
Rel. prop. Bank workers(it-1)	$\beta$	0.542 (.184)	0.522 (.314)	0.454 (.187)	0.489 (.231)
Rel. prop. Production workers(it-1)*log real ex rate(t-1)	$\beta$	-0.518 (.210)	-0.526 (.230)	-0.053 (.219)	-0.484 (.223)
Rel. prop. Bank workers(it-1)*financial liberalization(t-1)	$\beta$	1.776 (.377)	1.769 (.475)	2.175 (.393)	1.695 (.437)
Rel. prop. Bank workers(MA,t-1)* fittedDbase rate/100 (UK,t)	?	-2.098 (.816)	-2.598 (1.194)	-2.492 (.796)	-3.030 (2.730)
North	?	0.007 (.009)	0.011 (.017)	0.015 (.008)	0.002 (.011)
North-West	?	-0.004 (.005)	-0.001 (.001)	-0.005 (.005)	-0.007 (.006)
Yorkshire & Humberside	?	0.004 (.008)	0.007 (.019)	0.006 (.008)	0.001 (.011)
West Midlands	?	-0.038 (.010)	-0.036 (.017)	-0.045 (.009)	-0.049 (.012)
East Midlands	?	-0.042 (.010)	-0.044 (.019)	-0.052 (.010)	-0.049 (.011)
East Anglia	?	-0.039 (.006)	-0.043 (.008)	-0.043 (.006)	-0.037 (.006)
South East	?	0.043 (.010)	0.037 (.021)	0.039 (.010)	0.053 (.014)
South West	?	-0.042 (.007)	-0.046 (.009)	-0.044 (.007)	-0.038 (.008)
Wales	?	-0.031 (.007)	-0.025 (.011)	-0.026 (.007)	-0.031 (.009)
Scotland	?	-0.015 (.007)	-0.010 (.008)	-0.006 (.006)	-0.017 (.008)
Number of Observations		240	160	240	240
Sample Period		1972-1995	1972-1987	1972-1995	1972-1995
Equation Standard Error		0.00833	0.00843	0.00813	0.00820
AR Test (F)		1.26 [0.29]	3.52 [0.03]	0.62 [0.54]	1.67 [0.19]
HS Test (F)		0.15 [0.70]	1.07 [0.30]	0.01 [0.91]	0.70 [0.40]
$\overline{LM}$		12.13	8.61	11.60	12.83

Notes: Heteroscedasticity-Consistent Standard Errors in parentheses. Dependent Variable is the change in log relative full-time men's earnings. Estimation is by SUR in TSP (Hall et al., 1995). Equation standard error is the square root of the average of the squared residuals, corrected for degrees of freedom. Following the notation in equation (1),  $\alpha$  is the speed of adjustment;  $\beta$  are long-run coefficients; ? are dynamic terms; ? are long-run fixed effects. Pooled Mean Group (PMG)  $\alpha$  and ? coefficients are calculated from the simple average of the region coefficients and their estimated standard error is the standard deviation of the coefficients across regions.

AR F-test for 1st and 2nd order serial correlation, Breusch-Pagan (1980).

HS F-test for heteroscedasticity, White (1980).

$\overline{LM}$  Test for non-stationary residuals based on Im, Pesaran and Shin (1997), 5% Critical value (approx.): 3.96

Table 2  
Full-Time Women's Relative Earnings Model

Parameter		Model 1 Baseline	Model 2 Short Sample	Model 3 PMG
Rel. log women's full-time earnings (it-1)	$\alpha$	0.456 (.050)	0.547 (.070)	0.567 (.115)
DRel. log women's full-time earnings (it-1)	?	-0.231 (.050)	-0.187 (.061)	-0.192 (.132)
DRel. unemployment/100 (it) (instrumented)	?	-0.559 (.152)	-0.505 (.239)	-0.566 (.923)
Rel. log house prices(it-1) *UK prop. owner-occupied(t-1)	$\beta$	0.179 (.023)	0.170 (.036)	0.159 (.023)
Rel. prop. production workers(it-1)	$\beta$	0.025 (.089)	-0.021 (.111)	0.098 (.105)
Rel. prop. banking workers(it-1)	$\beta$	-0.458 (.243)	-0.095 (.291)	-0.644 (.283)
Rel. prop. production workers(it-1)*log real ex rate(t-2)	$\beta$	-0.790 (.166)	-0.700 (.161)	-0.678 (.161)
Rel. prop. bank workers(it-1)*financial liberalization(t-1)	$\beta$	1.980 (.304)	1.571 (.306)	2.308 (.391)
Rel. prop. banking workers(MA,t-1)*trend 1989	?	0.094 (.018)		0.083 (.021)
<b>North</b>	?	-0.046 (.010)	-0.029 (.014)	-0.039 (.010)
<b>North-West</b>	?	-0.033 (.006)	-0.028 (.008)	-0.021 (.007)
<b>Yorkshire &amp; Humberside</b>	?	-0.057 (.010)	-0.047 (.014)	-0.038 (.009)
<b>West Midlands</b>	?	-0.066 (.012)	-0.049 (.015)	-0.039 (.011)
<b>East Midlands</b>	?	-0.080 (.014)	-0.065 (.017)	-0.054 (.022)
<b>East Anglia</b>	?	-0.061 (.005)	-0.052 (.006)	-0.037 (.010)
<b>South East</b>	?	0.108 (.013)	0.088 (.017)	0.057 (.010)
<b>South West</b>	?	-0.056 (.005)	-0.054 (.005)	-0.024 (.006)
<b>Wales</b>	?	-0.047 (.010)	-0.034 (.011)	-0.044 (.010)
<b>Scotland</b>	?	-0.040 (.005)	-0.039 (.005)	-0.028 (.010)
<b>Number of Observations</b>		240	160	240
<b>Sample Period</b>		1972-1995	1972-1987	1972-1995
<b>Equation Standard Error</b>		0.00951	0.00943	0.00948
<b>AR Test (F)</b>		1.56 [0.21]	2.88 [0.06]	2.80 [0.06]
<b>HS Test (F)</b>		2.30 [0.13]	1.08 [0.30]	4.90 [0.03]
<b><math>\overline{LM}</math></b>		11.47	7.37	11.46

Notes: Heteroscedasticity-Consistent Standard Errors in parentheses. Dependent Variable is the change in log relative full-time women's earnings. Estimation is by SUR in TSP (Hall et al., 1995). Equation standard error is the square root of the average of the squared residuals, corrected for degrees of freedom. Following the notation in equation (1),  $\alpha$  is the speed of adjustment;  $\beta$  are long-run coefficients; ? are dynamic terms; ? are long-run fixed effects. Pooled Mean Group (PMG)  $\alpha$  and ? coefficients are calculated from the simple average of the region coefficients and their estimated standard error is the standard deviation of the coefficients across regions.

AR F-test for 1st and 2nd order serial correlation, Breusch-Pagan (1980).

HS F-test for heteroscedasticity, White (1980).

$\overline{LM}$  Test for non-stationary residuals based on Im, Pesaran and Shin (1997), 5% Critical value (approx.): 3.96

Table 3  
Part-Time Women's Relative Earnings Model

Parameter		Model 1 Baseline	Model 2 Short Sample	Model 3 PMG
Rel. log women's part-time earnings (it-1)	a	0.565 (.046)	0.648 (.064)	0.574 (.173)
DRel. log women's full-time earnings (it-1)	?	0.264 (0.054)	0.294 (0.068)	0.220 (0.315)
Rel. log unemployment (it-1)	B	0.076 (.015)	0.044 (.019)	0.070 (.018)
DRel. unemployment/100 (it)	?			
DRel. unemployment/100 (it-1)	?	-0.393 (.223)	-0.288 (.358)	-0.531 (.836)
Weighted av. Building Society mortgage rate/100 (it-1)	B	0.221 (.112)	0.451 (.246)	0.255 (.154)
Weighted D av. Building Society mortgage rate/100 (it-1)	?	1.409 (.166)	1.195 (.313)	0.578 (.242)
DRel. prop. part-time women in total women(it-1)	?	0.505 (.143)	0.604 (.219)	0.501 (.851)
Rel. prop. production workers(it-1)	B	-0.407 (.201)	-0.362 (.290)	-0.663 (.228)
DRel. prop. banking workers(it-1)	?	0.978 (.300)	1.460 (.462)	0.694 (1.556)
Rel. prop. production workers(it-1)*log real ex rate(t-2)	B	-1.674 (.217)	-1.527 (.273)	-2.075 (.345)
<b>North</b>	?	-0.087 (.016)	-0.065 (.024)	-0.068 (.018)
<b>North-West</b>	?	-0.028 (.012)	-0.029 (.016)	-0.018 (.015)
<b>Yorkshire &amp; Humberside</b>	?	-0.037 (.014)	-0.037 (.020)	-0.021 (.015)
<b>West Midlands</b>	?	-0.010 (.020)	-0.021 (.029)	-0.013 (.023)
<b>East Midlands</b>	?	0.001 (.019)	0.002 (.027)	0.005 (.008)
<b>East Anglia</b>	?	-0.025 (.011)	-0.030 (.014)	-0.032 (.012)
<b>South East</b>	?	0.074 (.020)	0.060 (.030)	0.052 (.022)
<b>South West</b>	?	-0.071 (.011)	-0.068 (.015)	-0.082 (.012)
<b>Wales</b>	?	-0.060 (.013)	-0.049 (.015)	-0.056 (.013)
<b>Scotland</b>	?	0.001 (.009)	0.013 (.011)	0.001 (.010)
<b>Number of Observations</b>		220	140	220
<b>Sample Period</b>		1974-1995	1974-1987	1974-1995
<b>Equation Standard Error</b>		0.01864	0.01841	0.02075
<b>AR Test (F)</b>		2.20 [0.11]	0.12 [0.89]	1.97 [0.14]
<b>HS Test (F)</b>		2.26 [0.13]	0.39 [0.53]	8.01 [0.00]
<b><math>\overline{LM}</math></b>		11.38	7.84	12.20

Notes: Heteroscedasticity-Consistent Standard Errors in parentheses. Dependent Variable is the change in log relative part-time women's earnings. Estimation is by SUR in TSP (Hall et al., 1995). Equation standard error is the square root of the average of the squared residuals, corrected for degrees of freedom. Following the notation in equation (1), a is the speed of adjustment;  $\beta$  are long-run coefficients; ? are dynamic terms; ? are long-run fixed effects. Pooled Mean Group (PMG) a and ? coefficients are calculated from the simple average of the region coefficients and their estimated standard error is the standard deviation of the coefficients across regions.

AR F-test for 1st and 2nd order serial correlation, Breusch-Pagan (1980).

HS F-test for heteroscedasticity, White (1980).

$\overline{LM}$  Test for non-stationary residuals based on Im, Pesaran and Shin (1997), 5% Critical value (approx.): 3.96

Table 4  
Relative Unemployment Rate Model

Parameter		Model 1 Baseline 1	Model 2 Baseline 2	Model 3 Short Sample	Model 4 PMG
Rel. unemployment rate (it-1)	a	0.164 (.025)	0.201 (.023)	0.189 (.037)	0.203 (.039)
Rel. unemployment rate (Wales, t-1)	a	0.345 (.059)	0.536 (.077)	0.379 (.091)	0.412 (0.093)
Rel. unemployment rate (Scotland, t-1)	a	0.280 (.059)	0.377 (.062)	0.382 (.123)	0.376 (0.069)
Rel. log women's full-time earnings (it-1)	$\beta$	18.90 (5.00)	10.90 (2.90)	17.52 (6.34)	9.73 (3.90)
Rel. log men's full-time earnings (it-2)	$\beta$	14.09 (4.07)	9.75 (2.67)	9.63 (4.88)	14.12 (4.33)
DRel. unemployment(it-1)	?	0.353 (.042)	0.351 (.043)	0.318 (.063)	0.285 (.132)
Rel. log house prices(it-3)	$\beta$	4.24 (1.08)	1.62 (0.62)	4.25 (1.56)	1.98 (.86)
Weighted D2 av. Building Society mortgage rate (it-2)	?	0.189 (.020)	0.209 (.023)	0.119 (.047)	0.153 (.167)
Rel. prop. part-time women in total women(it-1)	$\beta$	16.46 (4.42)	23.66 (3.23)	14.63 (5.30)	20.24 (4.64)
Rel. prop. production workers(MA,t-1)	$\beta$	4.77 (9.25)	-7.49 (5.96)	-8.60 (10.03)	5.81 (9.56)
Rel. prop. banking workers(MA,t-1)	$\beta$	-80.43 (18.51)	-35.97 (9.49)	-69.69 (21.30)	-52.79 (15.54)
Rel. prop. production workers(MA,t-1)*log real ex rate(t-1)	$\beta$	84.17 (15.84)	71.31 (10.97)	69.88 (15.96)	88.55 (17.47)
Rel. prop. bank workers(it-1)*log real house prices (UK,t-2)	$\beta$	-57.66 (15.01)	-33.35 (11.21)	-38.32 (18.34)	-28.48 (15.89)
Rel. prop. banking workers(MA,t-1)* fittedDbase rate (UK,t)	?	0.525 (.147)	0.692 (.169)	0.788 (.229)	0.684 (1.159)
Rel. prop. banking workers(MA,t-1)*trend 1989	$\beta$	11.74 (1.39)			
Trend 1989 * f(fixed effects, earnings, house prices)			0.094 (0.014)		0.091 (0.022)
North	?	0.038 (.007)	0.042 (.005)	0.044 (.009)	0.033 (.007)
North-West	?	0.028 (.005)	0.031 (.004)	0.037 (.006)	0.025 (.005)
Yorkshire & Humberside	?	0.016 (.006)	0.014 (.005)	0.023 (.009)	0.006 (.007)
West Midlands	?	0.012 (.010)	0.024 (.007)	0.024 (.012)	0.010 (.010)
East Midlands	?	0.002 (.009)	0.009 (.007)	0.012 (.011)	-0.006 (.010)
East Anglia	?	-0.009 (.005)	-0.019 (.004)	-0.015 (.006)	-0.017 (.006)
South East	?	-0.025 (.009)	-0.034 (.006)	-0.036 (.010)	-0.020 (.009)
South West	?	0.002 (.006)	-0.011 (.004)	-0.008 (.006)	-0.003 (.006)
Wales	?	0.024 (.005)	0.027 (.003)	0.026 (.006)	0.021 (.005)
Scotland	?	0.028 (.004)	0.033 (.003)	0.030 (.004)	0.030 (.004)
Number of Observations		240	240	160	240
Sample Period		1972-1995	1972-1995	1972-1987	1972-1995
Equation Standard Error		0.0026023	0.0025040	0.0023825	0.0024607
AR Test (F)		2.10 [0.13]	2.70 [0.07]	3.75 [0.03]	1.67 [0.19]
HS Test (F)		0.43 [0.51]	0.11 [0.75]	0.71 [0.40]	0.01 [1.00]
LM		9.85	10.13	6.01	10.59

Notes: Heteroscedasticity-Consistent Standard Errors in parentheses. Dependent Variable is the change in relative unemployment rate/100. Estimation is by SUR in TSP (Hall et al., 1995). Equation standard error is the square root of the average of the squared residuals, corrected for degrees of freedom. Following the notation in equation (1),  $\alpha$  is the speed of adjustment;  $\beta$  are long-run coefficients; ? are dynamic terms;  $\gamma$  are long-run fixed effects. Pooled Mean Group (PMG)  $\alpha$  and ? coefficients are calculated from the simple average of the region coefficients and their estimated standard error is the standard deviation of the coefficients across regions.

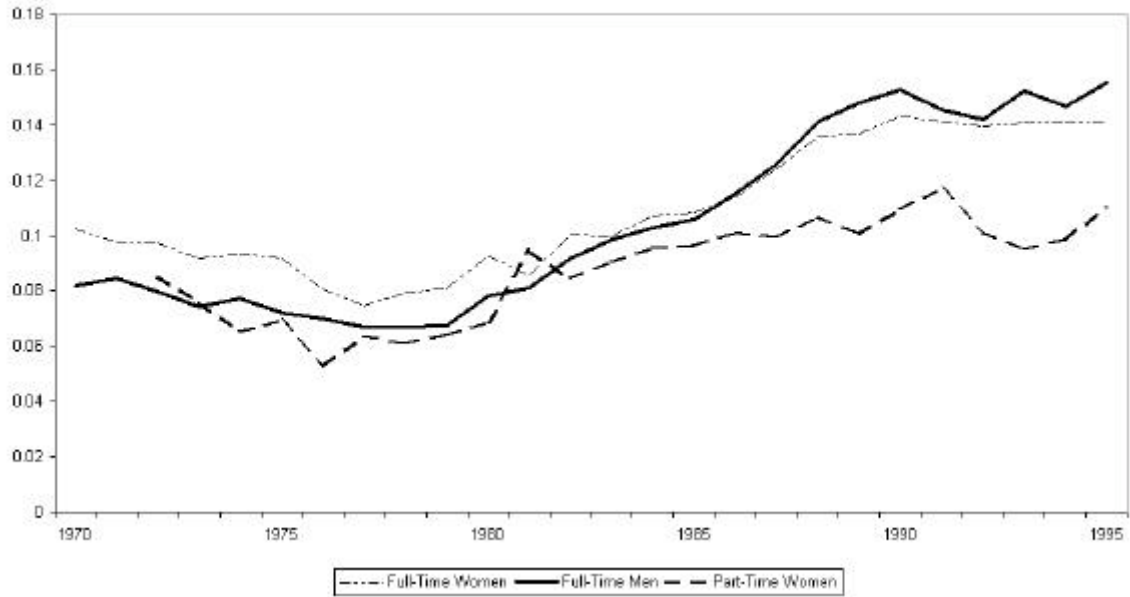
AR F-test for 1st and 2nd order serial correlation, Breusch-Pagan (1980).

HS F-test for heteroscedasticity, White (1980).

LM Test for non-stationary residuals based on Im, Pesaran and Shin (1997), 5% Critical value (approx.): 3.96



**Figure 1a**  
**Log Earnings in South East Relative to Great Britain**



**Figure 1b**  
**Unemployment Rate/100 in South East Relative to Great Britain**

