# Bank of England

## Discussion Papers

No 53

**The determination of average earnings in Great Britain**

by

M A S Joyce

*December 1990*

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The object of this series of papers is to give a wider circulation to econometric research being undertaken in the Bank and to invite comment upon it; and any comments should be sent to the author at the address given below. The views expressed are those of the author and not necessarily those of the Bank of England.

The author wishes to thank John Dorrington, Don Egginton, Neal Hatch, John Flemming, Paul Fisher, Brian Henry, Ian Michael, Bahram Pesaran and, particularly, to Stephen Hall for helpful comments and suggestions. Final responsibility for the contents rests solely with the author.

The author would also like to thank the Bulletin Group for editorial assistance in producing this paper.

Issued by the Economics Division, Bank of England, London, EC2R 8AH to which requests for individual copies and applications for mailing list facilities should be addressed; envelopes should be marked for the attention of the Bulletin Group. *(Telephone: 071-601-4030)*



ISBN 0 903315 55 6

**CONTENTS**

## ABSTRACT

This paper sets out to model real earnings growth in Great Britain, using a data set extending back over the last twenty five years. The empirical analysis (based principally on the two-stage co-integration framework developed by Engle and Granger) examines the long-run relevance of the factors most frequently put forward to explain the resilience of earnings growth during the 1980s. In particular, the paper examines the importance of 'wedges', unemployment duration, national and regional house prices, productivity and profits in explaining real consumption earnings (conditioned on the RPI). The concept of the wedge between consumption real wages and real labour costs is extended to include the wedge between retail and consumer prices. The preferred cointegrating vector incorporates a smoothed measure of labour productivity, average hours, logged unemployment, and terms representing various tax, real import price and retail price wedges. A dynamic equation based on this long-run solution is derived after a general to specific search. The equation allows real wages to be affected by price surprises in the short run, but in dynamic steady-state real wages are neutral with respect to price inflation. This restriction is found to be accepted by the data. The resulting equation passes a wide range of diagnostic statistics and provides a satisfactory explanation of wage inflation through the 1980s. The equation's performance is shown to compare favourably with the disaggregated earnings system currently on the Bank model; with the former dominating the latter, in a parsimonious encompassing sense.

## 1. INTRODUCTION

The resilience of earnings growth during the 1980s has been the subject of much debate among economists. During a period in which unemployment first rose sharply (from just over one million at the beginning 1980 to a peak of just over 3 million in 1986), then fell almost as rapidly (to 1.6 million in 1990), before turning up again in April of this year, annualised underlying earnings growth never fell below 7 1/2%, although RPI inflation fell to as low as 2 1/2%. In fact, on a year-on-year basis, real earnings grew in all but two quarters of the 1980s despite large swings in unemployment. Taking a longer view there has been a remarkably consistent upward trend in real earnings over the last twenty five years—with a trend growth rate averaging 2% per annum—which has been interrupted only rarely; the most noticeable period being during the 1975-77 period of incomes policy (see Figure 3.1).

A number of studies published during the middle to late 1980s advanced new empirical explanations for the strength of real wages growth in Britain.(1) Among these explanations, arguably the most influential related to the importance of the duration structure of unemployment [Layard and Nickell (1985,1986), Nickell (1987)], to house prices and regional house price differentials [Bover et al (1989)], and to profits [Carruth and Oswald (1989)]. In this paper we attempt to look at the empirical relevance of these factors, and others, for modelling the determination of real average earnings in Great Britain using a more recent data set extending back over the last twenty five years. The estimation approach adopted is based principally on the two stage co-integration framework developed by Engle and Granger (1987), which has been widely and successfully applied in recent empirical work by macro-modellers [see, for example Drobny and Hall (1989)]. This approach has the advantage that it enables us to examine separately the long and short-run properties of the data. However, it is not unproblematic, particularly where there is more than one long-run relationship between a set of variables, and we therefore also use the Johansen (1988) maximum-likelihood based approach to provide an additional check on the properties of our preferred co-integrating vector.

Apart from a desire to reappraise the evidence on real earnings, this study was also motivated by the need to find an aggregate wage equation to replace the disaggregated system of equations presently on the Bank of England's macro-model [see Mackie, Hatch (1990)]. This three equation system has a number of properties which have caused problems in model exercises and which a suitably specified aggregate equation could avoid. We therefore also report some encompassing tests of our preferred aggregate earnings equation against the current Bank disaggregated earnings system.

This paper is structured as follows. We begin in section 2 by briefly surveying some recent explanations put forward to explain British wage

(1) Following convention, we use the terms earnings and wages interchangeably in what follows.

inflation. In particular, we examine why unemployment duration, mismatch, profits and regional house prices might be expected to affect the determination of wages. We also consider in some detail the importance of tax and import price wedges. This analysis is then extended to the wedge between retail and consumer prices, a subject which has received relatively little explicit attention in previous work. Section 3 goes on to use co-integration techniques to develop a long-run model of real earnings. We find a role for productivity as expected. Unemployment in logarithmic form also enters the equation, which is consistent with the arguments about duration which have been advanced; although some additional work using short and long-term unemployment appears to contradict this interpretation. Tax, import price and retail price wedges are not found to be necessary to form a co-integrating vector, but they have plausible parameter values when included in a co-integrating regression and theoretical considerations argue for their inclusion. In our preferred co-integrating vector we assume that all the wedge terms are treated symmetrically by wage bargainers and therefore constrain all their coefficients to be equal. Section 4 embeds our preferred co-integrating vector in a dynamic equation which is shown to fit the data adequately and to forecast recent wage developments successfully. Section 5 then presents encompassing tests of this equation against the current Bank disaggregated earnings system. In a parsimonious encompassing sense we argue that the aggregate equation dominates the disaggregated system. Conclusions are presented in section 6.

## 2. EXPLAINING REAL WAGES

Most recent empirical work on wage determination in Britain is based, either explicitly or implicitly, on a union bargaining framework. It is generally assumed that either unions and firms bargain jointly over wages and employment (the efficient contract model), or that there is bargaining over the wage but the firm sets employment on the basis of its demand for labour curve (the right to manage model). We shall not set out the formal derivations of these two approaches here; they have been discussed extensively elsewhere [see eg Nickell (1984), Carruth and Oswald (1989)]. The important point to note is that empirically these theoretical approaches are identical in terms of the arguments which enter into the estimated wage equation, and indeed the sort of general wage equation we estimate here will also be consistent with a competitive framework.

Our primary purpose in this paper is not to set out another model of wage determination, but rather like Hall and Henry (1987) "to assess the degree of support that can be found for existing models". In this section we shall therefore set out some of the factors which have been used in recent empirical work to explain real wages, before moving on to our own empirical analysis in the remaining sections of the paper. We shall couch the discussion in terms of the long-run determinants of real wages to side-step the issues of dynamics and expectations formation which will be deferred until section 4.

### (i) A core model of wages

It is useful to begin by setting out a benchmark or 'core' model against which other explanations can be compared. We take as our core model what can be thought of as a variant of the well-known Sargan (1964) real wage resistance model:

$$\frac{W}{P} = f(U, T) \qquad \text{(A)}$$

where W denotes the nominal wage, P the price level, U is the unemployment rate and T denotes some measure of the target/warranted real wage. Since we are focusing on the long run, equation (A) imposes homogeneity with respect to prices, which we shall take to be axiomatic in the long run. We can, therefore, think of equation (A) as a model of the real wage.

The terms on the right-hand side of the equation can be justified in a variety of ways. The inclusion of the unemployment rate can be thought of as a proxy for excess supply in the labour market in the tradition of the Phillips curve, or justified by insider-outsider and efficiency wage theories as a measure of the probability of unemployment or outside opportunities. In the original Sargan model the target real wage, T, was measured by a time trend. However, more recent empirical work has often measured T by some indicator of productivity, e.g a smoothed measure of output per head [see Hall (1986), Rowlatt (1987)], even while sticking to an interpretation in terms of a union model of wage determination. This more general empirical model could equally be rationalised in terms of a competitive theory of wage formation.

When output per head is used to measure productivity, model (A) is sometimes extended by including a measure of hours worked as an additional regressor. The inclusion of hours allows for the possibility that either productivity is measured hourly or that there are overtime effects [see e.g. Hall (1986)]. This version of equation (A) is particularly interesting from our perspective because Hall finds that this set of variables forms a cointegrating vector. Our empirical analysis in section 3 below will therefore focus initially on variants of this model.

### (ii) Unemployment duration

The correct method of measuring unemployment in wage equations has been disputed since the time of the Phillips curve. One of the main findings of recent empirical work in modelling wages has been the importance of allowing for the duration structure of unemployment, by using either the rate of short-term unemployment, logged unemployment [Layard and Nickell (1985)(1986)] or both the aggregate unemployment rate (logged or unlogged) and the ratio of short-term to long-term unemployed [Nickell (1987)]. The argument for allowing for the duration structure of unemployment is that the short-term unemployed exert a far greater moderating influence on wage inflation than do the long-term unemployed. A number of plausible explanations can be advanced for expecting this result. For example, insider/outsider models suggest that unions will be more concerned with the probability of their members becoming unemployed than with the level of unemployment, *per se*, and the short-term unemployment rate is likely to provide a better measure of this. Another argument is that the extent to which the unemployed put downward pressure on wages depends on how effectively and intensively they search for jobs. The long-term unemployed, however, are likely to become discouraged from searching for employment, a tendency reinforced by the fact that employers are more reluctant to employ them (the signalling effect). This again suggests that short-term unemployment may provide a better measure of labour slack.

The duration argument is sometimes used to justify the inclusion of the logarithm of the unemployment rate in wage equations [Layard and Nickell (1985),(1986)]. The presence of the logged measure of unemployment implies that, at higher levels of unemployment, a larger absolute increase in unemployment is necessary to have the same restraining influence on real wages, which is consistent with the observation that the proportion of long-term unemployed tends to increase with the level of unemployment. However, concave unemployent effects can be justified in other ways besides the effects of duration. For example, Nickell (1987) following Lipsey (1960) shows that concave effects can emerge when there is aggregation across regionally distinct labour markets.

### (iii) Tax, import price and retail price wedges

Most recent empirical work on wages makes careful allowance for the factors which drive a wedge between the real consumption wage workers receive and the real labour cost paid by employers. It is the former that workers bargain over and they can be expected to resist any increases in the wedge being passed through into lower real consumption wages. This view is most closely associated with various economists at the former Centre for Labour Economics [for a recent exposition of this view see Jackman et al (1989)], but the view that wages are struck in net terms was also present in much earlier work [see e.g. Henry et al (1976)].

Spelling out the exposition in Jackman et al (1989), the wedge terms can be derived in the following way. We begin with the real product wage of concern to employers. In logs this may be defined as:

$$ln\left(\frac{W(1+T_e)}{P}\right) \approx w + t_e - p \qquad \text{(i)}$$

where W = nominal wage
P = value-added deflator
$t_e$ = employment tax rate

and the lower-case denotes variables expressed in natural logarithms (except for the tax rate terms).

The real consumption wage of concern to workers is:

$$ln\left(\frac{W(1-t_d)}{PC}\right) \approx w - t_d - pc \qquad \text{(ii)}$$

where PC = consumer price deflator
$t_d$ = direct tax rate

Additionally, value-added prices can be related to consumer prices by the following two identities:

$$pc = ln\,[PP\,(1+t_i)] \approx pp + t_i \qquad \text{(iii)}$$

$$pp = ln\left(P^{1/(1+v)} . P_m^{v/(1+v)}\right) \qquad \text{(iv)}$$

$$= \frac{p}{(1+v)} + \frac{v.p_m}{(1+v)}$$

or $p = (1+v)\,pp - vp_m$

where PP = producer prices
$t_i$ = indirect tax rate
$P_m$ = import prices
v = share of imports in GDP

Subtracting real take-home pay (ii) from real labour costs (i) and substituting in from (iii) and (iv) we are then left with the following wedge terms:

$$t_d + t_e + t_i + v\,(p_m - pp) \qquad \text{(v)}$$

If any element of this wedge increases we would expect there to be upward pressure on real labour costs as workers attempt to resist any fall in the real purchasing power of their wages. What is unclear is if real wage resistance of this sort can be sustained in the long run and therefore whether any of these wedge terms belong in the co-integrating vector for real earnings. Jackman et al (1989) argue that the implications of most theoretical models is that rises in proportional taxes and relative import prices will be passed on into lower wages. Only taxes that are a fixed amount per worker can be resisted in the long-run. However, notice that if we are modelling the real consumption wage rather than the real producer wage (as e.g. Layard and Nickell do), then the coefficients on any indirect tax or (weighted) import price terms need to be non-positive in order to offset their impact on consumer prices. So excluding the wedge terms from an equation for real wages conditioned on consumer prices would imply the strong conclusion that workers were able to fully compensate themselves for indirect tax increases and devaluations in the real exchange rate. This would be the implication of estimating equation (A) as it stands with P measured by consumer prices.

The concept of the wedge can be extended further. If we are concerned with the determination of the real consumption wage then we need to decide which measure of prices is most relevant to the bundle of goods consumed by workers. Here the two obvious choices are either the consumers' expenditure deflator (CED) or the retail price index (RPI). *A priori,* it would seem more reasonable to use the latter since it is this measure which is commonly presented in public discussion as measuring the 'cost of living', although it is the former which is more common in empirical work on earnings.[1] However, since we shall also wish to impose homogeneity with respect to prices, use of the RPI rather than the CED is problematic. This is principally because the RPI is affected by mortgage interest payments which do not affect the CED and in the long-run it is unclear that firms will be willing to let wages rise in line with RPI increases stemming from this source. We can therefore think of there being a further wedge between retail and consumer prices which will enter into the wage equation. If the measure of prices most relevant to workers is the RPI then equation (ii) should be rewritten as

$$ln\left(\frac{W(1-t_d)}{RPI}\right) \approx w - t_d - rpi \qquad \text{(ii)'}$$

and, if we ignore all the other differences between the two price measures besides mortgage interest payments, we can write down the following additional identity

$$rpi = ln\,[\,PC\,(\,1 + MIP\,)] \approx pc + MIP \qquad \text{(iv)'}$$

where $MIP$ = mortgage interest rate payment component expressed as a proportion of the CED

It then follows that there is an additional wedge term so that the difference between (i) and (ii)' is now

$$t_d + t_e + t_i + v\,(\,p_m - pp\,) + MIP \qquad \text{(v)'}$$

In principle, the effect of this additional wedge term is identical to the others. The outcome, as in the case of the other wedges discussed above, will depend on the relative strengths of the two sides of the wage bargain. If workers and firms treat all wedges symmetrically then we would expect the sign on these terms to have equal magnitudes, but we examine this empirically below.

### (iv) House prices

In an influential paper Bover et al (1989) argued that national house prices as well as regional house price differentials have had an important role to play in explaining the persistence of high wage inflation in the face of high unemployment. They advance three main arguments for why this may be the case. The first is based on a model with regionally distinct labour markets, where there can be an asymmetry between the upward and downward wage pressure exerted by a given amount of excess demand/supply. They go on to argue that regional differences in the house price/wage ratio may provide a proxy for regional mismatch and by analogy that the national house price/wage ratio may act as a proxy for national demand shocks. This implies a positive association between these measures and wage inflation. The second argument is based on the role of either interregional or international migration. High regional house price/wage ratios or national house price/wage ratios may both act as a disincentive to migration and therefore put upward pressure on wage inflation. The third argument is based on the wedge concept described in section (iii) above. If there are non-linearities in the response of wages to excess demand, and regional house price differentials provide a good proxy for regional differences in the wedge between consumer and producer real wages, then this is another reason for expecting a positive association with wages. A related argument can also be applied to national house prices because, to the extent that house prices mirror land prices (which do not enter into value-added prices), they drive an additional wedge between consumer and producer wages. (This argument could be used to justify the inclusion of the retail price wedge above.)

(1) Rowlatt (1987) is an exception. Layard and Nickell (1985,1986) focus on the real product wage, which leads to them using a value-added price index to measure prices.

**(v) Profits**

The argument that profits affect wages has a long history dating back to the controversy surrounding the Phillips curve [eg see Kaldor (1959)]. More recently, Carruth and Oswald (1989) have argued that profits provide the missing link for the strength of earnings growth during the 1980s. The argument for including profits follows most naturally from a bargaining theory of wage determination and this is how it is usually justified. However, a link between pay and profits can also be justified by appealing to the psychological literature on 'fairness' (see Carruth and Oswald op cit).

The main theoretical objection to the profits argument is that profits are endogenous and really a proxy for other factors like labour productivity and excess demand. This question is difficult to decide empirically because profits and productivity are linked by identity with real wages.

**(vi) Other factors**

A number of other factors are commonly included in wage equations and we briefly mention them here:

There are strong theoretical arguments for including some measure of the **replacement ratio** (ie the ratio of benefits to post-tax real wages) in models explaining the determination of wages. In most union-based models it is obvious that an increase in the replacement ratio will improve the bargaining strength of the union since the costs of unemployment will be lower to its members, with the result that the wage increases in equilibrium. In the competitive paradigm there is of course an effect on labour supply.

It is widely held that union power has an influence on wage bargaining. During the 1960s and 70s empirical estimates of the Phillips curve often used measures of union density to proxy this, although there have always been doubts over whether this represented the best measure. More recently, Layard and Nickell (1986) have used an index of union power they term the **union mark-up** based on industry cross-section regressions of earnings on a number of explanatory variables including union coverage, with the coefficient on union coverage representing the mark-up.

The work of Layard and Nickell has also been influential in suggesting that a measure of **mismatch** is important in explaining wage growth. The argument is that as jobs and potential employees become more mismatched, a given level of excess demand in the labour market becomes less effective in constraining wage pressure. The difficulty has been in how to measure this concept. Layard and Nickell (1986) use the absolute change in the ratio of employees in the production industries as their measure of mismatch (elsewhere this has been termed a 'turbulence index') with some success. Below we investigate each of these terms.

## 3. TIME SERIES PROPERTIES AND CO-INTEGRATION RESULTS

The empirical analysis reported in this paper is based principally on the two stage co-integration framework developed by Engle and Granger (1987). This involves first estimating a co-integrating levels equation by ordinary least squares and then at the second stage including these residuals (lagged one period) in a dynamic equation. [For a recent empirical application, see eg Drobny and Hall (1989).] This approach has the advantage that it is easily understood and enables us to separately examine the long and short-run properties of the data. However, it does have several problems. Besides the well known problem of small sample bias [see Banerjee et al (1986)], there is a problem when there is no unique co-integrating vector. Where this occurs, estimation by OLS may produce a complicated linear combination of the distinct co-integrating vectors which may be hard to interpret [Hall et al (1989)]. The advantage of using the Johansen (1988) maximum-likelihood based approach to co-integration is that it is able to allow for the case where there is more than one long-run relationship between a given set of variables. Moreover, the procedure enables us to identify these distinct co-integrating vectors and where appropriate it provides weights which can be used to combine these vectors. Since the Johansen approach is still being developed and its properties not fully understood [for a simplified exposition see Eitrheim (1990)], the analysis in this paper will concentrate on the more familiar Engle and Granger approach. However, we do present Johansen estimates for our

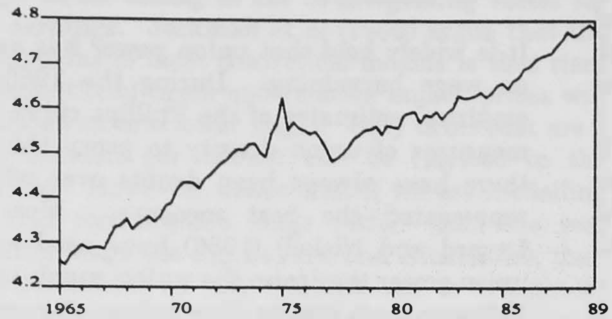



**Fig 3.1: REAL EARNINGS (ln WR)**

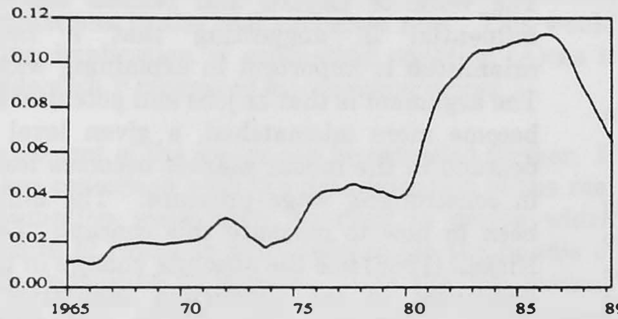



**Fig 3.2: UNEMPLOYMENT RATE (UR)**

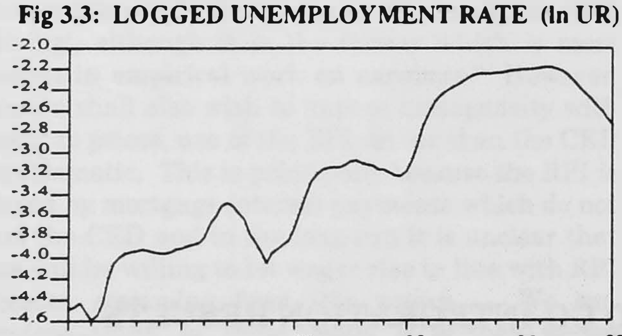



**Fig 3.3: LOGGED UNEMPLOYMENT RATE (ln UR)**

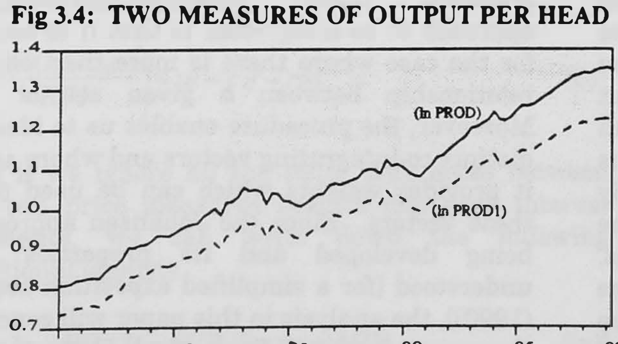



**Fig 3.4: TWO MEASURES OF OUTPUT PER HEAD**

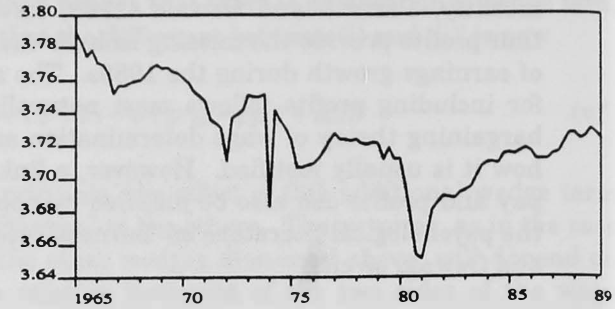



**Fig 3.5: AVERAGE HOURS (ln HMF)**

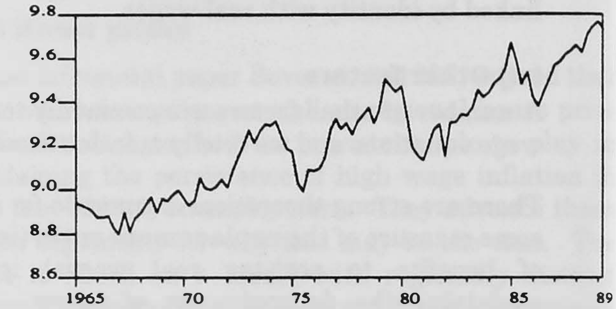



**Fig 3.6: REAL PROFITS (ln RPROF)**

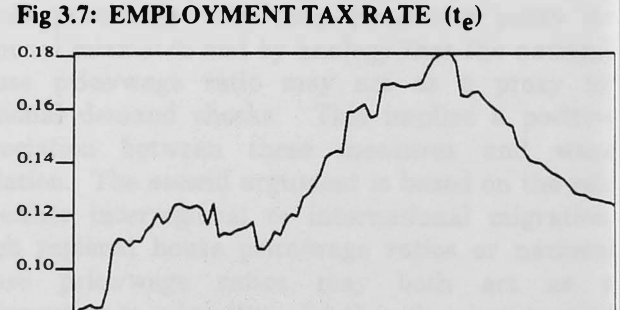



**Fig 3.7: EMPLOYMENT TAX RATE ($t_e$)**

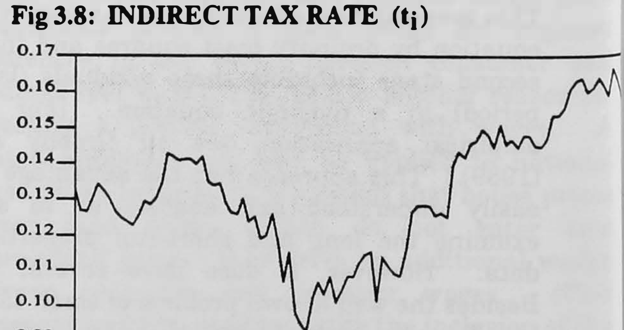



**Fig 3.8: INDIRECT TAX RATE ($t_i$)**

Fig 3.9: RETENTION RATIO (ln RR)

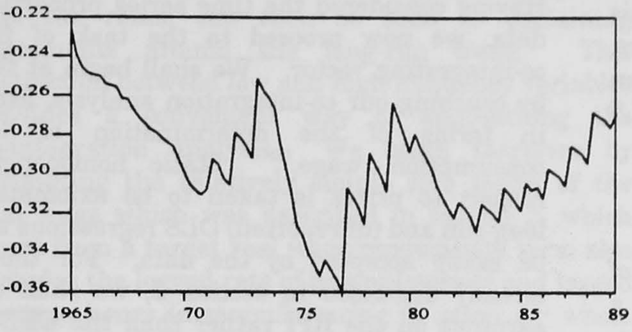



Fig 3.10: WEIGHTED REAL IMPORT PRICES (V ln PMP)

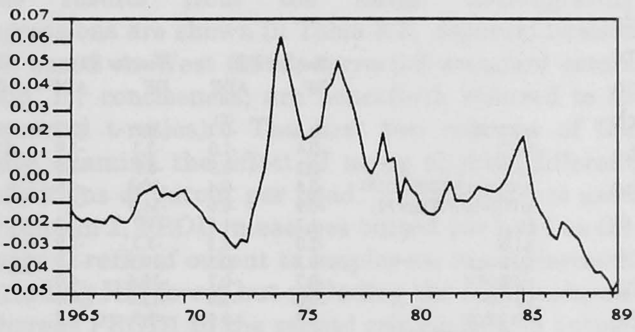



Fig 3.11: SHORT-TERM UNEMPLOYMENT RATE (STUR)

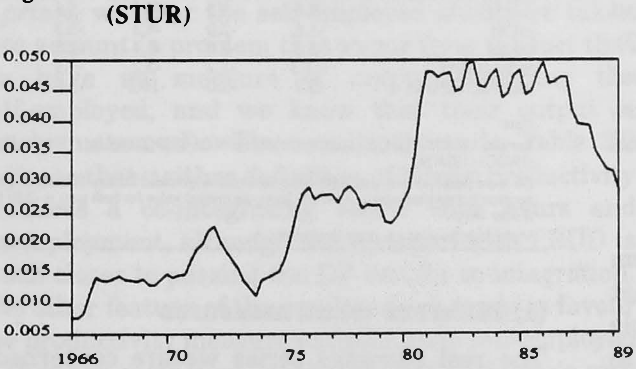



Fig 3.12: PROPORTION OF MALE UNEMPLOYMENT OVER 52 WEEKS (PUO52)

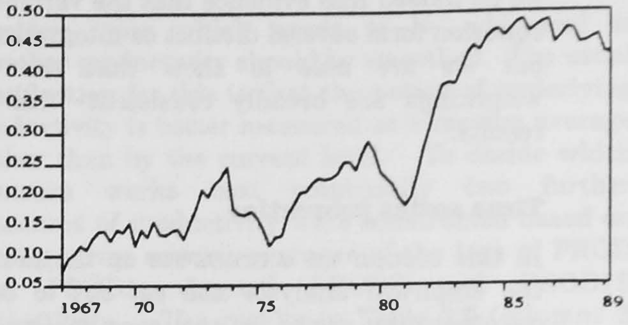



Fig 3.13: MISMATCH (ln MM)

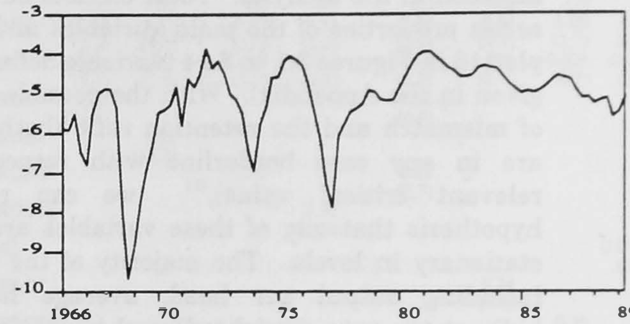
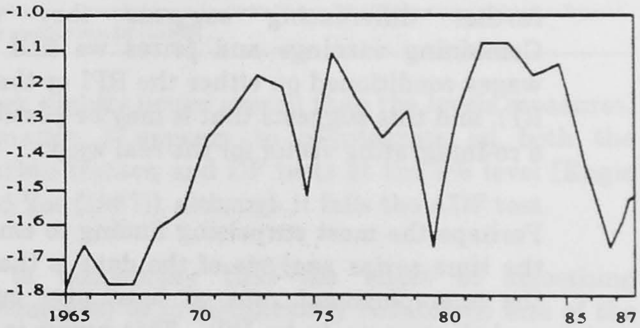



Fig 3.14: UNION MARK-UP (ln UM)



preferred co-integrating vector as an additional check on its properties. To anticipate our findings slightly, we do indeed find evidence that the variables in this equation form several distinct co-integrating vectors, but we are able to show that the Johansen weightings are broadly consistent with the OLS results.

## Time series properties

In this section we concentrate on the first stage of the empirical analysis and set out to estimate a co-integrating vector for real consumption earnings. The existence of a co-integrating vector between a set of variables depends on them all being integrated of the same order, and we begin therefore by examining the orders of integration of the variables to be included in the analysis. Table 3.1 sets out the time series properties of the main variables and they are plotted in Figures 3.1 to 3.14 (variable definitions are given in the Appendix). With the possible exception of mismatch and the retention ratio (both of which are in any case borderline with respect to the relevant critical value),[1] we can reject the hypothesis that any of these variables are I(0), i.e. stationary in levels. The majority of the variables, including output per head, average hours, the indirect tax rate, (weighted) real import prices, the union mark-up and logged unemployment, are all clearly I(1). However, we cannot reject the hypothesis that both prices and earnings are random walks in first differences [a common finding see, eg Hall (1986), Moghadam and Wren-Lewis (1989)], and further differencing suggests they are I(2). Combining earnings and prices we find that real wages conditioned on either the RPI or the CED are I(1), and this suggests that it may be possible to find a co-integrating vector for the real wage.

Perhaps the most surprising finding to emerge from the time series analysis of the data is the fact that the unemployment rate, at least over our sample period, appears to be I(2). This result is robust to whether we include or exclude a constant or time trend in the Dickey-Fuller/Augmented Dickey-Fuller regressions. The implication of this would appear to be that on statistical grounds it is the first difference rather than the level of the unemployment rate which belongs in the co-integrating vector for real earnings, a result which seems to tie in with recent insider-outsider theories of wage determination. We use the logarithm of the unemployment rate [which is I(1)] rather than the rate itself throughout the co-integrating analysis presented below.

## A co-integrating vector for real consumption earnings

Having considered the time series properties of the data, we now proceed to the task of finding a co-integrating vector. We shall begin at the outset by couching our co-integration analysis, exclusively, in terms of the determination of the real consumption wage.[2] Static homogeneity with respect to prices is taken to be axiomatic in the long-run and (unreported) OLS regressions show it to be easily accepted by the data. For the reasons already discussed in section 2, we shall condition earnings on the RPI rather than the CED, but we examine this assumption more closely in section (c) below.

**TABLE 3.1: Time series properties of t variables**

| | Test for I(0) | | Test for I(1) | | Test for I(2) | |
|---|---|---|---|---|---|---|
| | DF | ADF | DF | ADF | DF | ADF |
| lnETDE | -0.4 | -0.7 | -6.1 | -2.7 | -18.8 | -6.1 |
| lnPC | -0.4 | -1.0 | -3.2 | -1.9 | -11.9 | -5.4 |
| lnRPI | -0.2 | -0.9 | -5.3 | -2.3 | -15.5 | -5.[illegible] |
| lnWP=ln(ETDE/PC) | -0.5 | -0.4 | -10.3 | -4.8 | | |
| lnWR=ln(ETDE/RPI) | -0.9 | -0.6 | -11.3 | -4.7 | | |
| UR | -1.2 | -1.6 | -1.7 | -2.0 | -7.3 | -5.[illegible] |
| lnUR | -2.0 | -2.3 | -3.3 | -2.9 | | |
| lnPROD | -0.3 | -0.5 | -10.5 | -3.8 | | |
| lnPROD1 | -0.7 | -0.7 | -10.6 | -4.0 | | |
| lnHMF | -2.6 | -2.0 | -12.9 | -4.9 | | |
| lnRPROF =ln(PROF/PC) | -0.7 | -0.9 | -9.9 | -4.8 | | |
| lnRPROF1 =ln(RPROF/LE) | -0.7 | -0.8 | -10.1 | -4.9 | | |
| $t_e$ | -2.3 | -2.1 | -8.9 | -3.2 | | |
| $t_i$ | -0.8 | -0.7 | -11.2 | -3.9 | | |
| lnRR | -2.9 | -3.3 | -10.9 | -3.0 | | |
| lnPMP=ln(PM/PPOX) | -1.1 | -2.0 | -6.6 | -4.4 | | |
| vLPMP | -1.1 | -1.8 | -6.5 | -4.8 | | |
| lnMM | -2.8 | -3.2 | -7.0 | -5.8 | | |
| lnUM | -1.8 | -2.0 | -4.3 | -3.3 | | |
| lnRER | -1.2 | -2.6 | -3.4 | -2.9 | | |
| ln(PAHM/PC) | 1.3 | -1.2 | -5.2 | -3.8 | | |
| ln(HPSE/HPUK) | -0.7 | -2.0 | -10.5 | -3.0 | | |

Note:

Sample: 1965Q1-1989Q2; except for lnMM (1966Q1-1989Q2), lnUM (1965Q1-1987Q4) and lnRER (1965Q1–1985Q4).

DF is the Dickey-Fuller test; ADF is the 4th order Augmented Dickey-Fuller test (calculated with a constant but no time trend); at the 5% level, the critical value for both tests is -2.89 for 100 observatio (Fuller 1976).

Variable definitions are given in appendix A.

### (a) Different trend measures

The real earnings series we are concerned with is shown in Figure 3.1. We have already made the point in section 1 that it exhibits a strong upward trend over time, with exception of the 1975–77 period of incomes policy when real earnings fell back. (The spike in real earnings at the beginning of 1975 was due to a large back-dated pay award received by teachers.) Hall and Henry (1987) make the point that a strongly trended variable like this is more likely to be explained by other variables that also exhibit strong trend-like behaviour. The obvious contenders in this context are various proxies for the

(1) Since the correlogram for both of these variables dies out very slowly over eight quarters, it seems more appropriate to regard them as being non-stationary in levels.

(2) This is not strictly true in the sense that we use gross average earnings deflated by the RPI as our dependent variable, so direct taxes are not netted off. Nevertheless, implicitly it is the real consumption wage which is being determined.

target real wage/productivity: the capital-labour ratio, output per head, real profits or a time trend. The role of other variables, such as unemployment and tax rates, can then be seen as explaining movements around the long-run trend. This distinction between low and high frequency variables provides a convenient way of approaching the co-integration analysis. We begin therefore by estimating the relatively simple core model of the real wage which was described in section 2, which apart from a target real wage/productivity term also includes the logged rate of unemployment and logged average hours in manufacturing (a proxy for whole economy hours, which are not available on a quarterly basis).(1) The idea here was to examine the properties of this simple model as different trend variables were included.

The results from the initial co-integrating regressions are shown in Table 3.2; reported t-ratios are based on West (1988) corrected standard errors (and, for conciseness, are henceforth referred to as corrected t-ratios).(2) The first two columns of the table examine the effect of using slightly different definitions of output per head.(3) The measure used in column 1, PROD, measures output per head as the (logged) ratio of output to employees in employment including HM forces but excluding the self-employed, whereas PROD1 in the second column relates output to a wider definition of employment which includes the self-employed. Since the measure of average earnings we are attempting to explain refers exclusively to employees in employment it is unclear, *a priori*, whether the self-employed should be taken into account (a problem that stems from the fact that we have no measure of output excluding the self-employed, and we know that their output is poorly measured). The results shown in Table 3.2 indicate that neither definition of labour productivity produces a co-integrating vector with hours and unemployment, although the equation with PROD is much closer to passing the DF test for co-integration. The other feature of the results which tends to favour the productivity measure excluding the self-employed is that the coefficient on the PROD1 variable is well above unity. Such a result would imply a constantly rising labour share which seems both inconsistent with the past, and undesirable theoretically.(4)

Another issue which needs to be addressed is whether productivity should be smoothed. The usual justification for this is that the notion of underlying productivity is better measured as a moving average rather than by the current level. To decide which measure works best empirically two further measures of productivity were constructed based on eight-quarter moving averages of the logs of PROD and PROD1, denoted PRODS and PROD1S respectively. The results in Table 3.2 (columns 3 and 4) suggest that the moving average measures

**TABLE 3.2: Testing for different trend measures**

Dependent Variable = ln WR

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| lnHMF | -1.33 | -1.11 | -0.77 | -0.52 | -0.84 | -0.88 | -1.00 | -0.38 |
| | (5.4) | (4.2) | (3.5) | (2.2) | (2.9) | (1.4) | (1.7) | (0.8) |
| lnUR | -0.076 | -0.067 | -0.039 | -0.034 | -0.044 | 0.046 | 0.026 | -0.080 |
| | (4.1) | (3.4) | (2.7) | (2.2) | (2.2) | (1.4) | (0.7) | (2.0) |
| lnPROD | 1.07 | | | | 0.13 | | | |
| | (14.6) | | | | (0.4) | | | |
| lnPROD1 | | 1.18 | | | | | | |
| | | (13.2) | | | | | | |
| lnPRODS | | | 0.99 | | 0.88 | | | |
| | | | (16.3) | | (2.7) | | | |
| lnPROD1S | | | | 1.10 | | | | |
| | | | | (14.9) | | | | |
| TIME | | | | | | | | 0.0065 |
| | | | | | | | | (6.6) |
| lnRPROF | | | | | | 0.34 | | |
| | | | | | | (4.6) | | |
| lnRPROFH | | | | | | | 0.36 | |
| | | | | | | | (4.8) | |
| $R^2$ | 0.97 | 0.96 | 0.98 | 0.97 | 0.98 | 0.90 | 0.90 | 0.95 |
| CRDW | 0.61 | 0.50 | 0.68 | 0.56 | 0.67 | 0.36 | 0.39 | 0.30 |
| DF | -4.13 | -3.71 | -4.37 | -3.93 | -4.35 | -3.10 | -3.21 | -2.80 |
| ADF | -3.31 | -3.41 | -3.39 | -3.34 | -3.38 | -2.43 | -2.52 | -2.51 |

Sample: 1965Q1 to 1989Q2

West corrected t-ratios are in parentheses.
CRDW is the co-integrating regression Durbin Watson; DF/ADF statistics are explained in Table 3.1.
Each regression included a constant.

work slightly better overall than the levels measures. Equation 3 appears to co-integrate on both the Durbin-Watson and DF tests at the 5% level [Engle and Yoo (1987)], although it fails the ADF test.

It is noteworthy that the effect of smoothing productivity is to significantly reduce the size of the coefficients on both the unemployment rate and

(1) By including hours we allow for an overtime effect and the possibility that productivity may be measured on a hourly basis when output per head is included as a regressor. It was not possible to cyclically adjust earnings as, for example, Layard and Nickell (1985,1986) do because our earnings series has a broader coverage than theirs which referred to male manual workers only. They were therefore able to adjust their figures using data on actual and normal hours, which are not available for the whole economy.

(2) We decided at the outset not to include the capital-labour ratio in our analysis because of the problems there are in constructing a quarterly capital stock series. The published data are only available on an annual basis and have been criticised by many researchers. We note in passing, however, that in their analysis of the long-run determinants of real earnings Hall and Henry (1987) found that the capital-labour ratio was clearly dominated by output per head.

(3) We followed convention in using the output measure of GDP as the numerator. Oil production was not excluded because this restriction would imply that the long-run effect of the discovery of North Sea oil was to reduce the share of labour in value-added.

(4) The effect on the model's NAIRU depends on the size of the coefficient on productivity in the price equation. If the coefficient in the wage equation is larger, then the NAIRU will depend on trend productivity. However, provided both coefficients are contrained to have equal magnitudes then the effect on the NAIRU nets out. This constraint is imposed in the Layard-Nickell model (1985,1986) and in macro-models, like the NIESR NIDEM model, which are based on broadly the same approach [see Joyce and Wren-Lewis (forthcoming, 1991)].

hours. The equations reported in columns 1 and 2 seem to imply that a unit coefficient could be imposed on hours, which would imply that hourly productivity is the relevant measure [Moghadam and Wren-Lewis (1989) use this measure], but once productivity is smooothed this restriction looks less plausible. Constraining hours to have a unit coeficient implies that there is no overtime effect, although in a long-run equation this restriction could be justified. However, the fact that the hours term refers only to manufacturing, and therefore only proxies whole economy hours, perhaps suggests that its coefficient should be decided by the data rather than imposed.

We looked at two main measures of profits: real profits, RPROF, defined as the gross trading surplus deflated by the CED, and real profits per head, RPROFH, where employment is measured by employees in employment. Columns 6 and 7 in Table 3.2 show the results obtained when these profit measures, in logarithmic fonn,(1) are included separately in our core wage equation, excluding labour productivity. Both of these regressions fail to co-integrate and the positive sign on the unemployment term is clearly implausible.(2) For these reasons we are led to the conclusion that output per head provides a better measure of productivity. If we ignore the possible misspecification and include both profits and labour productivity in the regression, the coefficient on profits is reduced substantially in size and it is clear that the productivity term dominates.

Finally, equation 8 demonstrates the effect of including a time trend in the regression rather than either output per head or profits. The results clearly show that this specification fails to co-integrate. Adding a time trend to any of the other regressions also fails to produce co-integration, with the time trend, typically, insignificant according to its corrected t-ratio.

We may conclude then that, of the target/warranted real wage variables surveyed here, output per head appears to dominate either real profits, profits per head or a simple time trend. Since the moving average measure of productivity defined in this way dominates the level of productivity we shall concentrate on the latter measure in what follows. The measure calculated using employees in employment in the denominator is preferred to the self-employment based measure because it has a coefficient on productivity almost exactly equal to one. Since the unit restriction is consistent with the data we impose it to get the following equation

$$\ln WR - \ln PRODS = \underset{(8.3)}{6.22} \; - \underset{(3.7)}{0.77} \ln HMF \; - \underset{(5.2)}{0.041} \ln UR \qquad (1)$$

CRDW=0.68 DF=-4.37 ADF=-3.44 $R^2$=0.44

RCO: 0.66, 0.43, 0.19, 0.04, -0.00, -0.06, 0.09, -0.03

This equation has plausible coefficients, passes the DW and DF statistics and has a correlogram which dies away relatively quickly. Although it narrowly fails the ADF test [critical value 3.62 at the 5% level (Engle and Yoo, 1987)], this test is known to have low power and it seems reasonable to conclude that this rather simple model co-integrates. However, the model as it stands has some rather undesirable theoretical properties, which argue against its acceptance. Since the dependent variable is real earnings and conditions on the RPI, the model implies real exchange rate depreciations, and rises in either indirect taxation, employment taxes or mortgage interest payments feed through fully into higher wages, with no effects on the real purchasing power of workers. Yet the same equation also implies that workers have no success in resisting rises in direct taxes since no terms in direct taxation appear. We shall consider the evidence for wedge effects below, but we first look at how different measures of unemployment affect our results.

**(b) Unemployment duration, alternative unemployment definitions**

In Table 3.3 column 2 we show the results from including the rate of short-term unemployment in our model for real earnings (the level of the rate is

**TABLE 3.3: Testing for different measures of unemployment**

Dependent Variable = ln WR - ln PRODS

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| lnHMF | -0.74 | -0.82 | -0.85 | -0.73 | -0.44 | -0.31 | -0.27 |
| | (3.6) | (3.8) | (3.1) | (3.1) | (1.5) | (2.2) | (1.6) |
| lnUR | -0.04 | | | -0.05 | -0.01 | | |
| | (5.3) | | | (1.8) | (0.6) | | |
| STUR | | -2.25 | | 0.23 | | | |
| | | (5.2) | | (0.2) | | | |
| lnSTUR | | | -0.06 | | | | |
| | | | (4.1) | | | | |
| PUO52 | | | | | -0.13 | | |
| | | | | | (1.5) | | |
| lnLEP | | | | | | 0.47 | |
| | | | | | | (5.4) | |
| lnLEP1 | | | | | | | 0.65 |
| | | | | | | | (4.3) |
| $R^2$ | 0.45 | 0.40 | 0.35 | 0.45 | 0.51 | 0.45 | 0.37 |
| CRDW | 0.67 | 0.71 | 0.63 | 0.66 | 0.69 | 0.58 | 0.50 |
| DF | -4.26 | -4.27 | -4.06 | -4.23 | -4.43 | -3.95 | -3.43 |
| ADF | -3.60 | -3.50 | -3.56 | -3.58 | -3.00 | -3.39 | -3.03 |

Sample: (1)-(4) 1966Q1-89Q2
(5) 1967Q1-89Q2
(6)-(7) 1965Q1-89Q2

West corrected t-ratios are in parentheses.
CRDW is the co-integrating regression Durbin Watson; DF/ADF statistics are explained in Table 3.1
Each regression included a constant.

(1) Hall and Henry (1987) make the point that, since profits can be negative in principle, taking the log is technically a misspecification. On a practical level no problem arises here because aggregate profits have never been negative over the period.
(2) If North Sea oil profits are excluded from our definition of profits, there is a modest improvement in the properties of these equations, but they remain inferior to the equation with labour productivity and the coefficient on unemployment remains implausible.

I(1) so there is no statistical problem with including the rate). The properties of this equation are very similar to the equation with the log of the unemployment rate (column 1). The equation appears to co-integrate according to the DW and DF statistics and the coefficient on hours is similar (we have imposed a unit restriction on the smoothed labour productivity term). Including the logarithm of the short-term unemployment rate [as in Moghadam and Wren-Lewis (1989)] instead, shown in column 3, produces slightly inferior results and the equation now fails the DF test.

We have already discussed the fact that inclusion of the logged measure of the unemployment rate can often justified by appealing to duration arguments. The results in Table 3.3 suggest that the log formulation seems to perform quite adequately as a proxy for the short-term rate. Indeed, when we include both measures in the equation (column 4) it is clear that the logged measure dominates the short-term measure, at least according to the relative corrected t-ratios. To try to shed some more light on this question, we ran an equation including the logged unemployment rate and the proportion of unemployed males with over 52 weeks on the register (the series is shown in Figure 3.12). However, the results shown in column 5 are rather counter-intuitive, showing a negative coefficient on the ratio variable. This would imply that it is the long-term unemployed, rather than the short-term, who have the greatest impact on real wages (this result is unaffected if we include the unemployment rate rather than the logged unemployment rate in the equation). Given the wealth of evidence on the importance of duration effects, this result clearly requires further investigation. Taken at face value, it appears to suggest that we need to look at other reasons for the non-linear response of real wages to unemployent.

The use of the claimant count measure of unemployment in a wage equation might also be questioned. The Restart programme in particular may have caused a discontinuity in the unemployment figures over recent years [see Dicks and Hatch (1989)]. The latest work at NIESR uses the proportion of the working population not working in preference to recorded unemployment as a determinant of real earnings. We constructed two labour demand measures based on employment and population of working age statistics: LEP, defined as the (logged) ratio of employees in employment to the population of working age and, LEP1, the (logged) ratio of the workforce in employment to the population of working age. However, the resulting equations (columns 6 and 7) are inferior to those which include the logged unemployment rate and it was decided therefore to keep with the official measure. It might be argued, in any case, that wage negotiators will be more aware of the officially published figures on unemployment.

### (c) Tax, import and retail price wedges

We now turn to an investigation of the effect of allowing for the various wedges between real consumption wages and real labour costs. We shall first confine our analysis to the wedges specified in the Jackman et al (1989) analysis described in section 2(iii) above, before considering the additional retail price wedge. The wedge terms used were defined so as to tie in as closely as possible with the variables in equation (v), with the exception that we used the log of the retention ratio—the proportion of pre-tax earnings which are retained after tax—rather than the direct tax rate itself (this can be thought of as an approximation to minus $t_d$).[1] The definition of the retention ratio is the same as the one currently used in the manufacturing equation on the Bank of England model [see Mackie, Hatch (1990)]. It incorporates the effects of basic rate personal taxes, employees' national insurance contributions and personal tax allowances. Full definitions of all the wedge terms are contained in the Appendix;[2] Figures 3.7 to 3.10 plot the series over the sample period.

The results from adding the tax and import price wedges to the basic model (1) above are shown in columns 1 and 2 of Table 3.4. The coefficients on all the wedge terms are correctly signed and of plausible

**TABLE 3.4: Testing for wedge effects**

Dependent Variable = ln WR - ln PRODS

| | (1) | (2) | (3) |
|---|---|---|---|
| lnHMF | -0.61 | -0.63 | -0.68 |
| | (5.5) | (5.8) | (5.6) |
| lnUR | -0.03 | -0.03 | -0.02 |
| | (4.7) | (2.2) | (4.7) |
| $t_e$ | -0.46 | -0.49 | -0.51 |
| | (4.3) | (4.7) | (4.5) |
| $t_i$ | -0.82 | -0.88 | -0.91 |
| | (4.6) | (5.0) | (4.8) |
| lnRR | -0.59 | -0.60 | -0.53 |
| | (7.1) | (7.4) | (5.5) |
| lnPMP | -0.13 | | |
| | (4.4) | | |
| vlnPMP | | -0.52 | -0.49 |
| | | (5.0) | (4.4) |
| ln(RPI/PC) | | | -0.37 |
| | | | (1.9) |
| $R^2$ | 0.71 | 0.72 | 0.73 |
| CRDW | 1.06 | 1.11 | 1.08 |
| DF | -5.92 | -6.09 | -6.01 |
| ADF | -3.88 | -3.85 | -3.73 |

Sample 1965Q1 to 1989Q2

West corrected t-ratios are in parentheses.
CRDW is the co-integrating regression Durbin Watson; DF/ADF statistics are explained in Table 3.1.
Each regression included a constant.

(1) One advantage of using this definition is that both its sign and, in principle, the size of its coefficient will be the same as the other wedges.
(2) Our definition of $t_e$ includes employers' contributions to pension funds which are not strictly speaking a tax. Nevertheless, these contributions do add to labour costs, and by including these payments we are merely following the practice of vitually all empirical work in this area [this includes Layard and Nickell (1985,1986) and Moghadam and Wren-Lewis (1989)].

magnitudes. In contrast to recent work by Moghadam and Wren-Lewis (1989), we find plausible and and quite powerful effects from both employment taxes and the real exchange rate. However, resistance to indirect taxes appears rather weak with the coefficient on $t_i$ quite close to minus one. In general, the coefficients on employers' taxes, (implicit) direct taxes and weighted real import prices are all of a broadly similar size. We experimented with both weighted and unweighted measures of real import prices, the latter being the measure preferred in the early work by Layard and Nickell (1985). There is little to choose between these measures but the higher corrected t-ratio on the weighted measure argues in its favour.

In column 3 we expand the model reported in column 2 by including a retail price wedge term following the arguments advanced above in section 2 (iii). Since the RPI and CED differ in a number of other ways besides mortgage interest payments, the logged ratio of the two was used to measure the wedge. This term enters the equation with a plausible negative coefficient, although its corrected t-ratio would suggest that it is not strongly significant. On theoretical grounds, however, there seem strong reasons for including it in the model with the other wedge terms.

If wage bargainers are rational we would expect them to treat all the wedge terms symmetrically, at least in the long run. If this was not the case then it would be possible for tax switches to affect the level of real wages.(1) The results from our co-intergration analysis do not appear to provide any strong justification for allowing for different long-run responses; Although there may be a possible inconsistancy problem given that the tax bases differ slightly, we therefore impose the restriction that all the wedge terms are equal to get the following preferred equation:

$$\ln WR - \ln PRODS =$$
$$\underset{(6.8)}{6.75} \; \underset{(6.9)}{-0.92} \ln HMF \; \underset{(4.5)}{-0.025} \ln UR \; \underset{(6.2)}{-0.53} WEDGE \quad (2)$$

where WEDGE =

$$\{t_e + t_i + \ln RR + v \ln LPMP + \ln (RPI/PC)\}$$

CRDW=1.05; DF=-5.88; ADF=-3.47; $R^2$=0.70

RCO: 0.47, 0.27, 0.05, -0.05, 0.03, -0.03, -0.02, 0.04

This equation has plausible parameter estimates, suggesting that about one half of any increase in the wedge is passed though into higher wages, and appears to represent a co-integrating vector. It passes the DW and DF tests and has a correlogram which dies away quickly. It fails the ADF test as reported here [critical value 4.02 at the 5% level (Engle and Yoo, 1987)], but this test has been calculated by including four lagged difference terms, none of which turn out to be statistically significant using a joint test. When they are dropped the DF regression still passes the test for autocorrelation, so we may conclude that it is the DF statistic which is more relevant. Recursive estimates of the equation suggest that the parameter estimates are stable.

To investigate the properties of the equation more closely, the Johansen (1988) procedure was then applied to the same set of variables. Using an unrestricted VAR model with a maximum lag of four quarters, the likelihood ratio test that there are at most r co-integrating vectors gave the following results.

| r | LR test | 5% critical value |
|---|---|---|
| 0 | 76.19 | 47.21 |
| 1 | 33.77 | 29.68 |
| 2 | 15.61 | 15.41 |
| 3 | 2.083 | 3.762 |

These results suggest that there may be as many as three co-integrating vectors among this set of variables, which might at first appear to cast doubt over the OLS estimates. Restricting our attention to the eigenvectors with significant eigenvalues and then normalising the values of the associated eigenvectors to the variable we are interested in explaining (the real wage adjusted for smoothed productivity) we get the following:

| Eigenvalue | Eigenvector | | | |
|---|---|---|---|---|
| | lnWR-lnPRODS | lnUR | lnHMF | WEDGE |
| 0.36 | -1 | -0.058 | -1.62 | -0.42 |
| 0.18 | -1 | 0.004 | -1.08 | -1.05 |
| 0.13 | -1 | 0.032 | 3.70 | 1.63 |

Of these eigenvectors only the one with the largest eigenvalue could be plausibly thought of as a co-integrating vector for earnings. However, although all the parameter estimates are of the correct sign, the absolute magnitude of the coefficient on the hours term is implausibly large.

| Implied earnings eigenvector, weighted by alpha matrix | | | |
|---|---|---|---|
| lnWR-lnPRODS | lnUR | lnHMF | WEDGE |
| -1 | -0.043 | -1.02 | -0.26 |

However, the Johansen procedure enables us to weight these vectors together using the alpha matrix which is also produced by the optimisation. If we normalise on this term and examine the solution for earnings we get the following parameter estimates:

(1) This is not to imply that changes in the structure of taxation could not have important effects on wage determination. For example, 'pure' reductions in marginal and average rates of tax will have opposite effects on unions' wage demands [Creedy and MacDonald (1989)].

These estimates are now theoretically plausible (the hours term is virtually unity) and, furthermore, much closer to the OLS results reported in equation 2 above. So although it appears that there are a number of co-integrating vectors which could be formed from the set of variables in our preferred model, the OLS estimates appear to correspond reasonably closely to the vector for real earnings generated using the Johansen procedure. This gives some added justification for using equation 2 as our preferred cointegrating vector.

### (d) Mismatch, the union mark-up, house prices and replacement ratios

A number of other variables were also investigated in the co-integration analysis, but subsequently rejected. We investigated the role of house prices using both a measure of regional differences and a measure of national house prices. The regional measure was defined as the logged ratio of South East to UK house prices, based on the measure used by Bover et al (1989). We found, however, that when it was included in our preferred co-integrating regression it was both wrongly signed (negative) and insignificant according to its corrected t-ratio. The aggregate house price measure we investigated was a measure of real house prices, where nominal house prices were deflated by the CED. This variable was correctly signed (positive) when included in our preferred co-integrating vector, but only weakly significant according to its corrected t-ratio. It had little effect on the co-integration properties of the equation and it was therefore dropped from the analysis.

We also looked at the effect of including the mismatch and union mark-up variables used by Layard and Nickell (1985,1986). Both these variables were correctly signed when included as additional regressors in our preferred co-integrating regression, but neither appeared significant according to its corrected t-ratio. Neither variable had any significant impact on the co-integrating properties of the equation nor on the coefficients of the other explanatory variables in the model.

To examine the impact of the replacement ratio on earnings we used the measure used in Layard and Nickell (1986). However, we found this variable to be wrongly signed in our preferred co-integrating regression; a finding consistent with the other applied work in this area [e.g. Henry and Hall (1987), Wren-Lewis and Moghadam (1989)].

## 4. A DYNAMIC MODEL

In this section we turn to the estimation of a dynamic earnings equation based on our preferred co-integrating vector for real earnings, equation (2) in section 3(c) above. However, before getting on to the empirical results there are a number of further issues which need to be discussed.

The first of these is derivative or **dynamic homogeneity**. We have already imposed static homogeneity of earnings with repect to prices. This assumption seems unexceptional. It merely implies that there is no long-run money illusion and it is easily accepted by the data. More contentious perhaps, is the stronger assumption that wages are dynamically homogeneous with respect to prices. However, on theoretical grounds there seem to be strong reasons for imposing this restriction. Unless we do so the implication is that in long-run steady-state the real consumption wage will depend on the rate of inflation, which, in the context of a macro-model, will normally imply a long-run trade-off between unemployment and inflation. To avoid this result we impose dynamic homogeneity on our equation;(1) although we shall also show below that in our preferred equation this restriction is, in fact, accepted by the data. (2)

Another issue which needs to be mentioned is the approach taken to modelling **price expectations**. Although we are imposing the restriction that wages are dynamically homogeneous with respect to prices we want to allow for the possibility that real wages can differ from their intended level in the short run due to price surprises. Why this might be the case is clear from the fact that wage negotiations typically decide the nominal rather than the real wage. If prices turn out to be higher than anticipated we would expect the real wage to be lower than

(1) Clearly, this is not a sufficient condition for the model NAIRU to be independent of the rate of inflation. To ensure this result dynamic homogeneity needs to hold in the price equation.

(2) The imposition of dynamic homogeneity means using (the first difference of) the real wage as our dependent variable. However, since we also include lagged dependent variables in the equation, this means that in dynamic steady-state the real consumption wage depends on the growth of the real wage [this is also true for the work by Layard and Nickell (1985,1986)]. Dynamic homogeneity with respect to the real wage is rejected by the data and there seems to be no strong theoretical reason for imposing it.

intended, and vice versa, and to allow for this means including some measure of price expectations in our equation. There are a number of approaches which can be adopted here. We could, for example, use survey data on expectations or model them explicitly using the assumption of rational expectations [e.g as in Moghadam and Wren-Lewis (1989)]. Here we follow the approach taken by Layard and Nickell (1985,1986) among others and make the assumption that it is only when price inflation rises that prices turn out to be higher than expected, and vice versa. We therefore include a term in the current double difference of prices as an additional regressor, which is then instrumented in estimation.

Finally, it is worth mentioning the fact that our preferred dynamic equation does not follow the common practice of including dummy variables to control for the effect of **incomes policies**. There are well-known problems with following this procedure because it does not allow for the degree of severity of different policies, nor does it not allow for their likely endogeneity, and it reduces significantly the degrees of freedom available. An adequate treatment of all of these problems is not attempted here, but it is perhaps worth reporting that we were unable to find any statistically significant incomes policy effects in our initial estimation work. An intuitive explanation for this finding may be the fact that we are modelling the real rather than the nominal wage. If the aim of past incomes policies has been to reduce nominal rather than real wages then our results do not seem surprising and may still be consistent with incomes policies having been successful. Many periods of income policy have, of course, involved controls on both wages and prices and, even in cases where prices have not been explicitly controlled, it seems plausible to suggest that price increases will have been reduced in line with lower inflationary expectations.

To derive a dynamic equation we followed the conventional general to specific modelling strategy, letting the data decide the form of the equation dynamics. We started from a general equation which included on the right-hand side all the terms contained in the co-integrating vector, entered in differences up to the fourth lag, as well as lagged dependent variables, the current period price surprise term, the lagged residuals from the co-integrating vector and four seasonal dummy variables.[1] The equation also included a dummy variable set equal to 1 in 1975Q1 and -1 in 1975 Q2 to take out the effect of the spike in real wages in 1975Q1 which occurred due to a large back-dated pay award to teachers (see Figure 3.1).

Testing down from the general model we arrived at the following specification, estimated by instrumental variables

$$\begin{aligned}\Delta lnWR = {} & \underset{(4.0)}{0.324}\Delta lnWR_{t-1} + \underset{(4.2)}{0.222}\Delta lnWR_{t-2} \\ & + \underset{(3.3)}{0.316}\Delta lnHMF_{t-1} - \underset{(2.8)}{0.559}\Delta lnt_{it-1} \\ & - \underset{(3.4)}{0.266}\Delta lnRR - \underset{(3.2)}{0.057}\Delta\Delta_2 lnUR \\ & - \underset{(2.5)}{0.038}\Delta lnUR_{t-3} - \underset{(4.2)}{0.563}\Delta\Delta lnRPI \\ & - \underset{(5.5)}{0.339}Z_{t-1} - \underset{(8.1)}{0.049}D75 + \underset{(1.8)}{0.005}\,Q3 \\ & + \underset{(3.8)}{0.007}\,Q4\end{aligned}$$

where $Z_{t-1}$ are the lagged residuals from the cointegrating regression, equation 2, described in section 3(c). Instrumented variable: $\Delta\Delta$ln RPI; additional instruments were $\Delta\Delta$ln $RPI_{t-1}$ - $\Delta\Delta$ln $RPI_{t-4}$, $\Delta$vlnPMP$_{t-1}$ – $\Delta$vlnPMP$_{t-4}$. (Asymptotic t-ratios in parentheses.)

$R^2$=0.798 SE=0.00809 DW=2.04
LM(1)=0.2 LM(2)=1.6 LM(4)=1.7
RESET(1)=0.9 NORMALITY(2)=0.3
HETEROSCED(1)=0.1
ARCH(1)=1.6 CHISQ-S(8)=4.2 CHISQ-D(8)=4.0
MISSPEC(9)=3.7

Note: LM = Sargan's test of serial correlation of IV residuals; RESET = Ramsey's RESET test using the square of the fitted values; NORMALITY = Based on a test of skewness and kurtosis of residuals; HETEROSCED = Based on the regression of squared residuals on squared fitted values; MISSPEC = Sargan's test of misspecification in the case of IV estimation; CHISQ-S/D = Static and dynamic forecast tests.

This equation seems to fit the data well and has sensible theoretical properties. The residuals from the co-integrating vector are correctly (negatively) signed and highly significant, as is the price surprise term. In addition, a number of the other elements in the co-integrating vector also enter separately in the dynamics. Terms in hours, indirect tax, the retention ratio and unemployment were all found to be statistically significant and correctly signed.

The dynamic structure of the equation is rather complicated, and the inclusion of two lagged

(1) The seasonal dummies were introduced primarily because data for the RPI is only available on unadjusted basis and consequently our real wage measure exhibits seasonality, but it also corrects for seasonality in any of the regressors. This issue could be avoided in the co-integration analysis, on the assumption that seasonality in the real wage is stationary.

dependent variables has the effect of lengthening the lag response to exogenous shocks. The positive sign on the hours term acts to offset the negative impact implied by the co-integrating vector, so that the impact effect from an increase in hours is to increase real earnings growth. The effect of unemployment on real earnings is quite complex, although it can be given an intuitive explanation. The included terms in unemployment suggest that in the short run the growth of real wages is depressed by both increases in the rate of unemployment and by increases in its rate of acceleration.[1]

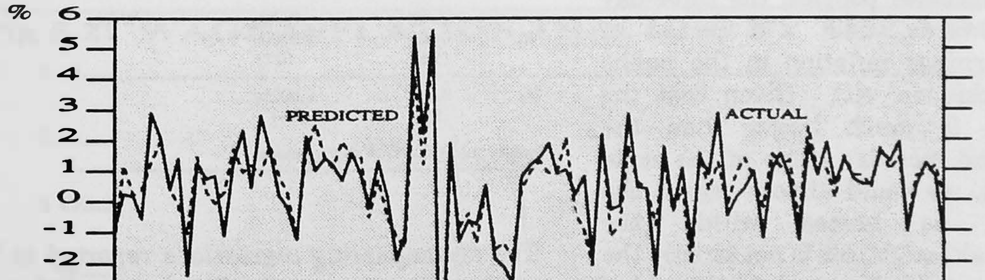
The price surprise term is correctly signed and highly significant, but the fact that the absolute size of its coefficient is significantly less than unity merits comment. *A priori*, we would expect a coefficient near to (but not greater than) unity for this variable since it is intended to proxy unexpected changes in inflation and therefore any one point acceleration in the rate of inflation would be expected to reduce the current period growth of real wages by the same (unintended) amount. One possible explanation for finding a coefficient significantly less than unity might be that agents forecast the future rate of inflation and so they are not totally fooled when prices accelerate. This would suggest the need to model the formation of expectations explicitly and this is an area which it is hoped can be investigated in future work.

On statistical grounds the equation appears to perform adequately. The equation fits the data reasonably well (Figure 4.1), has a low standard error of less than 1 per cent, and passes a wide range of diagnostic statistics for autocorrelation, normality, heteroscedasticity and misspecification. The forecasting performance of the equation is also reasonable according to the static and dynamic forecasting tests for 8 quarters which are both passed. In addition, the equation appears to be reasonably stable. According to Gujarati's dummy variable test on the joint stability of the model parameters [see Gujarati (1970)], the hypothesis that there has been a structural break in the 1980s can be rejected.

To examine the robustness of equation (2), and to see whether we could improve the fit of the equation, we also looked at the effect of including in the short-run dynamics some of the variables not included in the long-run solution of the equation.[2] In particular, we tried including additional terms in real house prices and profits, entered as first differences up to the fourth lag. We found no additional role for house prices, but the inclusion of terms in real profits was somewhat more successful in that we found a statistically significant and correctly signed effect on the difference term in real profits lagged three quarters. However, the fact that it is only the difference term at the third lag which appears to be statistically significant suggests that the profits term may be picking up some of the missing dynamics lost by smoothing the productivity term.[3] This might be taken to indicate that the effect from profits is spurious, although it could be equally argued that this result is consistent with profits being reported with a lag. Given that the equation with profits performs only marginally better than equation (3) we decided to stick with the simpler specification.

We now turn to consider whether our assumption of dynamic homogeneity is accepted in our preferred dynamic equation. A simple test for this would be to

**Fig 4.1: ACTUAL AND PREDICTED VALUES FROM THE DYNAMIC EQUATION**




(1) Current period unemployment is not instrumented in equation 4, although it could be argued that it should be to allow for simultaneity with real wages. We did not do so because instrumenting had little effect on the size or significance of this term.

(2) We also tried including a dummy variable for the three-day week (set equal to 1 in 1974Q1 and 0 elsewhere), but found that this variable could be rejected.

(3) Testing down from a general equation which includes real profits results in a very similar specification, with the profits effect restricted to the difference term at the third lag.

include additional terms in current and lagged inflation as regressors in the equation. So we have

$$\Delta lnWR = \ldots + a_0\Delta lnRPI_t + a_1\Delta lnRPI_{t-1} + a_2\Delta lnRPI_{t-2} + \ldots\ldots$$

The null hypothesis of dynamic homogeneity is then that these coefficients sum to zero, i.e. $H_o$: $a_0 + a_1 + a_2 + .. = 0$. An alternative reparameterisation of this test is to include additional lagged price acceleration terms in the preferred equation plus a single term in the level of price inflation. So, for example, we have

$$\Delta lnWR = \ldots + a_0\Delta lnRPI_{t-1} + a_1\Delta\Delta lnRPI_t + a_2\Delta\Delta lnRPI_{t-1} + \ldots\ldots$$

The null hypothesis of dynamic homogeneity is then more simply that the coefficient on the price inflation term is zero, i.e. $H_o$: $a_o=0$. This version of the test is easier to handle because it allows us to include a lagged term in inflation and therefore removes the problem of endogeneity. When this model was estimated we found the lagged inflation term to be statistically insignificant, with a t-ratio equal to 1.1, and this seems to suggest that the null hypothesis of dynamic homogeneity is accepted by the data. A possible problem with this conclusion is the fact that the terms on the price acceleration terms are also statistically insignificant and there may be a problem of multicollinearity. However, when these lagged terms were dropped the lagged inflation term remained statistically insignificant at the conventional 5% level (t-ratio equal to 1.4). We therefore conclude, perhaps surprisingly, that dynamic homogeneity is accepted by the data.

## 5. ENCOMPASSING TESTS

As a final test of the aggregate equation we attempted to see whether it encompassed or could be encompassed by the current system of disaggregated earnings equations on the Bank model [see Mackie, Hatch (1990)].

The disaggregated model is composed of three equations which determine manufacturing, public sector and 'other' earnings, and it allows for interactions between these sectors. The system contains a total of 38 explanatory variables, including a number of dummy variables designed to pick up effects from incomes policies, the three-day week, the Clegg awards, and "a period of particularly rapid earnings inflation in the public sector in 1975" (Mackie op. cit). Given that the disaggregated model is much larger than the aggregate equation and includes many of the same explanatory variables, we could almost regard the aggregate equation as nested within the disaggregated model (although this is not strictly the case). So the tests reported below are best thought of as parsimonious encompassing tests, i.e. as indicating whether the aggregate equation omits any significant factors contained in the disaggregated model.

To look at this question the Bank earnings system was first re-estimated over its original sample period using the same vintage of data as used in the empirical work described above. The results for the re-estimated model are not shown here but on the whole, the revised parameter estimates are broadly similar to those reported by Hatch (1990). Using these estimates implied fitted values for aggregate earnings were then derived by inverting the identity used to construct earnings in the 'other' sector. Figure 5.1 shows how the model tracks movements in quarterly wage inflation over its estimation period; Figure 5.2 provides the comparable picture for the aggregate equation.

### TABLE 5.1 ENCOMPASSING TESTS

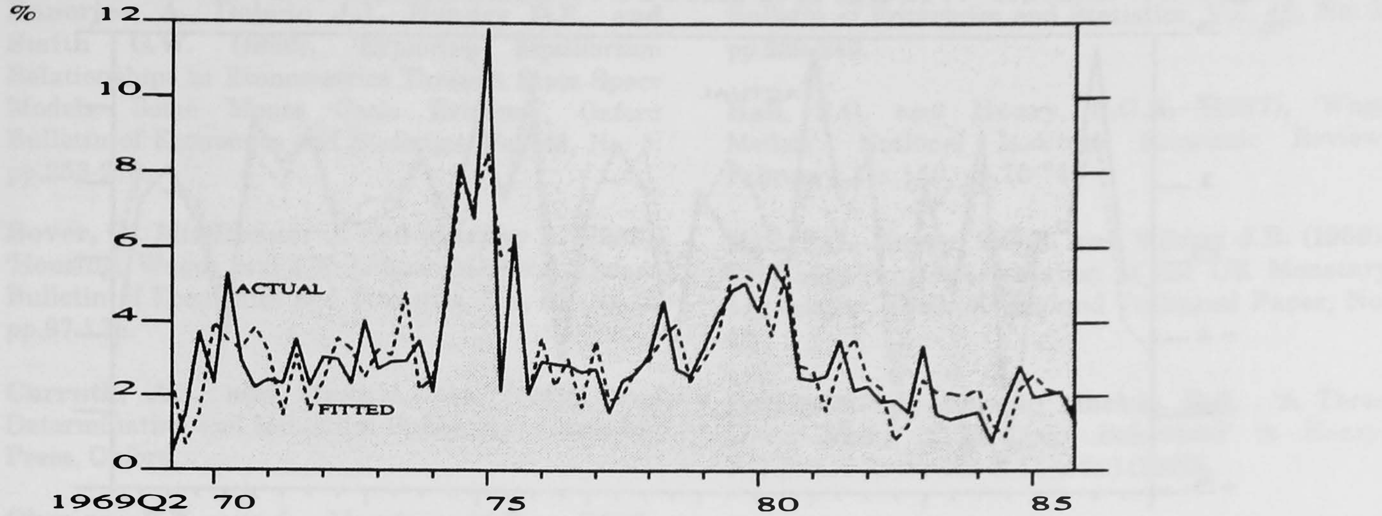
| | Dependent Variable | | |
|---|---|---|---|
| | W | W-W$_a$ | W-W$_d$ |
| W$_a$ | 0.506 | 0.00006 | |
| | (5.6) | (0.1) | |
| W$_d$ | 0.494 | - 0.00034 | |
| | (5.4) | (0.4) | |

Sample: 1969Q1 to 1985Q4; OLS regressions

The encompassing regressions reported in Table 5.1 attempt to compare the fitted values of each model with outturn. The regressions were run over the sample period used to estimate the disaggregated earnings system, 1969 Q2 to 1985 Q2. This is a rather shorter period than the one used to estimate the aggregate equation and it inevitably biases the results somewhat against the aggregate equation. Column 1 of Table 5.1 reports a simple regression of average earnings (W) on the predicted values from the aggregate and disaggregated models (denoted $W_a$ and $W_d$, respectively) i.e.

$$w = a.W_a + b.W_d + u$$

Fig 5.1: WAGE INFLATION TRACKED BY THE DISAGGREGATED MODEL




An insignificant t-ratio on either $W_a$ or $W_d$ could be interpreted as implying that model's predicted value added nothing to the prediction of the other. In fact, the fitted values from both models are statistically significant, with virtually identical coefficients and t-ratios, and on this test we cannot discriminate between them.

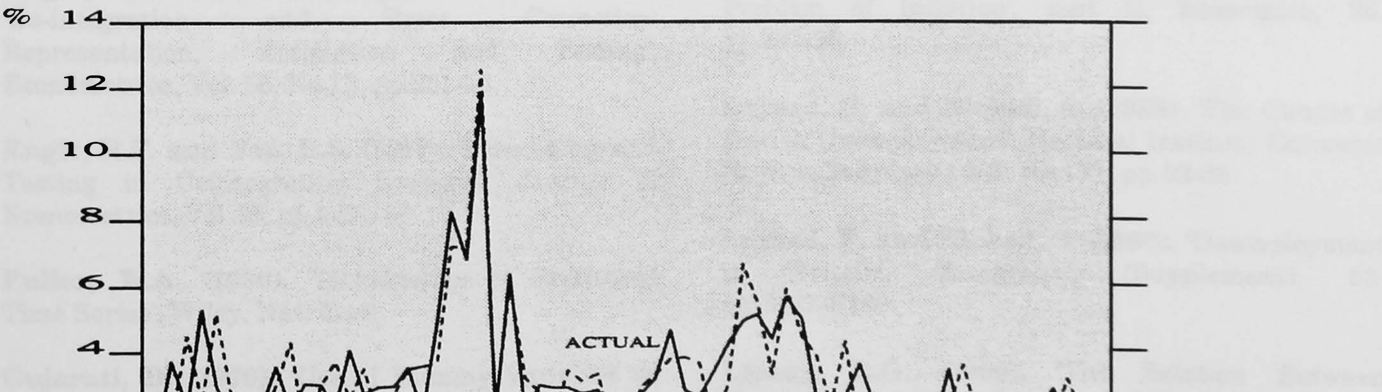
Columns (2) and (3) of Table 5.1 report the results from a regression of each model's prediction error on the predicted values of the other, i.e.

$$W - W_a = b.W_d + u$$

The idea here is to see whether the predictions of one model explain the errors of the other. A test of the hypothesis is provided by the t-ratio on the predicted value in each case, with a significant coefficient indicating that one model encompasses the other [Chong and Hendry (1986)]. Again, however, the tests reported in Table 5.1 are inconclusive with neither model encompassing the other.(1)

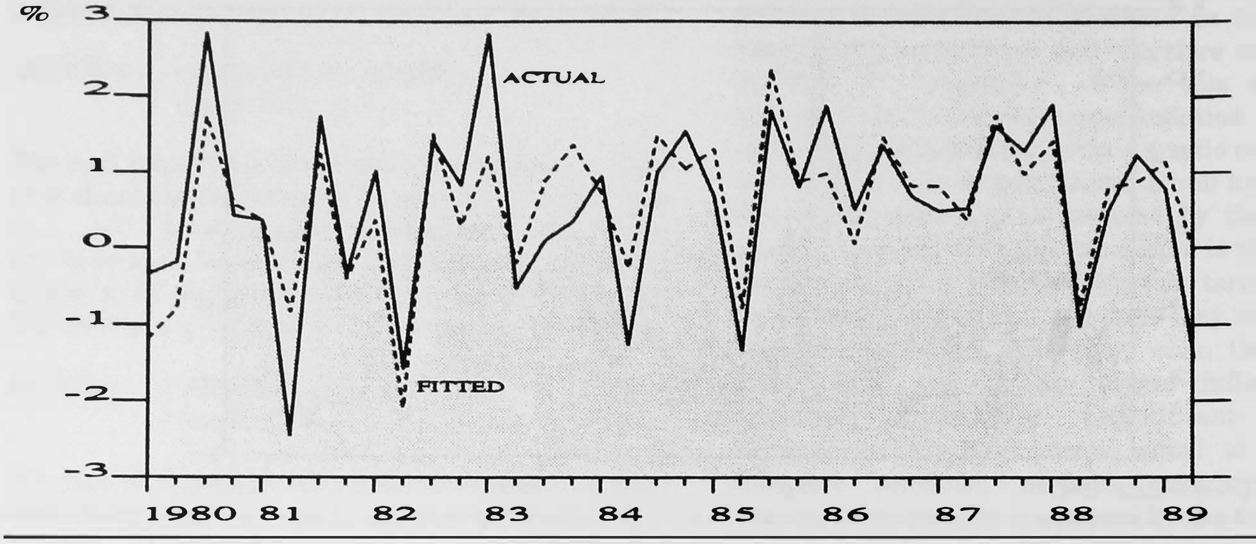
In summary, on none of these tests does the disaggregated model encompass the aggregate equation. Given that the aggregate equation can be thought of as almost a nested version of the disaggregated system and that the choice of sample period inevitably favours the larger system, these results appear to provide strong support for the aggregate equation, and, in a parsimonious encompassing sense, the aggregate equation could therefore be said to dominate the disaggregated system.

Fig 5.2: WAGE INFLATION TRACKED BY THE AGGREGATE MODEL




(1) We also tested the hypothesis that the error variances from the two models were equal using a standard F-test and found that equality could not be rejected.

**Fig 6.1: REAL WAGE INFLATION TRACKED DURING THE 1980S**




## 6. CONCLUSIONS

In this paper we set out to estimate a model of real earnings growth in Great Britain, using principally the two-stage co-integration framework developed by Engle and Granger. One of the purposes of this analysis was to examine the long-run relevance of the factors most frequently put forward to explain the resilience of earnings growth during the 1980s. In particular, we have examined in some detail the importance of 'wedges' between real consumption wages and real labour costs, unemployment duration, national and regional house prices, productivity and profits in explaining real consumption earnings conditioned on the RPI.

Our results suggest that a reasonably parsimonious dynamic model, based on a co-integrating vector containing terms in smoothed labour productivity, average hours, logged unemployment and various wedges, can adequately explain the path of earnings growth during the last twenty five years. Over the more recent period of the 1980s this model appears to track the data reasonably well (Figure 6.1) with no tendency for sustained underprediction. What is perhaps most interesting about this model is the variables it omits. In particular, there is no role for house prices, profits, mismatch, the replacement ratio or the union mark-up.

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## APPENDIX : DATA DEFINITIONS

| | |
|---|---|
| DEYW | Wages and salaries. £mn. Defined: 19.739837*ETDE*(LOTH+LEG+LEMF)/1000. |
| EP | Employees in employment as a proportion of population of working age.<br>Defined: LE/POWA |
| EP1 | Employed labour force as a proportion of population of working age<br>Defined: (LE+LSE+LWRT)/POWA<br>where:<br>LE = employees in employment and HM forces<br>LSE = self-employed<br>LWRT = participants on government work-related schemes |
| ETDE | Average whole economy earnings 1985=100. Seasonally adjusted.<br>Figures back to 1977 refer to DE measure of average weekly earnings (Great Britain). Earlier figures derived from CSO measure of wages and salaries divided by employees in employment (LE). |
| HMF | Average hours worked per operative in manufacturing industry. Hrs/wk. |
| HPSE | South-East second-hand house price series. Source: Anthony Murphy (Oxford) |
| HPUK | UK second-hand house price series. Source: Anthony Murphy (Oxford) |
| MM | Mismatch. Defined as in Layard & Nickell (1986).<br>Absolute change in proportion of employees in industry relative to total employees. Source: NIESR |
| PAHM | Price deflator for all houses mix adjusted. 1985 = 1. |
| PC | Price deflator for total consumption. 1985 = 1. |
| PM | Price deflator for imports of goods and services. 1985 = 1. |
| PROD | Output per head (based on GDP output estimate). Defined GDPO/LE. |
| PROD1 | Output per head (based on GDP output estimate). Defined GDPO/(LE+LSE). |
| PROF | ICC gross trading profits. £mn. |
| PUO52 | Proportion of male unemployment over 52 weeks duration. Source: Centre for Labour Economics. |
| RER | Replacement ratio. Defined as in Layard and Nickell (1986). A measure of average annual income on benefit to mid-year earnings. Source: Paul Kong (Oxford). |
| STUR | Short-term unemployment rate.<br>Defined:<br>SUN/(LE+LSE+LWRT+SUN)<br>SUN = Number unemployed less than 26 weeks: Source NIESR<br>LE = Number of employees in employment<br>LSE = Number of self-employed<br>LWRT = Number of participants on government work-related schemes. |
| $t_e$ | Employment tax rate.<br>Defined:<br>ln{(DEYW + YECO + YECN + YECS +TSET)/ DEYW}<br>Defined:<br>DEYW = Defined above<br>YECN = Employers' national insurance contributions. £mn<br>YECS = Accruals of national insurance surcharge. £mn<br>YECO = Employers' other contributions. £mn<br>TSET = Selective employment tax receipts. £mn. |

| | |
|---|---|
| $t_i$ | Indirect tax rate.<br>Defined:<br>ln{(GDP£ + FCA£ - YECS - TSET) / GDP £}<br>where:<br>GDP£ = GDP expenditure estimate. £mn.<br>FCA£ = Factor cost adjustment. £mn.<br>YECS = Accruals of national insurance surcharge. £mn<br>TSET = Selective employment tax receipts. £mn. |
| RR | Retention ratio.<br>Defined:<br>TXA + {(TXB*TRYE/100)/19.739837*ETDE)}<br>where:<br>TXA = 1- {(TRYE/100) + (YJCN/DEYW)}<br>TXB = {(100*TARR/TRYE) + TPAL}/NCPA<br>TRYE = TRY * {1-2/9 * (1-D73B)}<br>YJCN = National insurance contributions paid by employees and self-employed. £mn.<br>TRY = Standard rate of income tax. %<br>D73B = Dummy variable. Equal to 1 in 73Q2; otherwise 0.<br>TARR = Reduction in income tax due to the existence of reduced rates. £mn.<br>TPAL = Aggregate married, single and child allowances, £ mn/ qtr<br>NCPA = Total number claiming personal allowances. |
| UM | Union mark - up variable. Defined as in Layard & Nickell (1986). Source: Paul Kong (Oxford). |
| UR | Unemployment rate. %.<br>Defined LU/(LE+LSE+LWRT+LU)<br>where: LU = Number unemployed. |
| V | Share of imports in GDP.<br>Defined M£/GDPN<br>where<br>M£ = Imports of goods and services. £mn<br>GDPN = Nominal GDP at market prices (average measure). £mn. |

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| | Title | Author |
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