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The relationship between employment and unemployment

by

M J Dicks
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The object of this series is to give a wider circulation to research being undertaken in the Bank and to invite comment upon it; and any comments should be sent to the authors at the address below.

The authors are grateful for comments from Bank economists John Flemming, Roger Clews, John Dorrington, Gary Dunn, Don Egginton, Brian Henry and Kerry Patterson and from Paul Gregg of the National Institute of Economic and Social Research, Bryan Cress of HM Treasury and Alfred Webb of the Department of Employment. Any errors remain the authors' responsibility. The views expressed are those of the authors and not necessarily those of the Bank of England.

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CONTENTS

Section 1: Introduction

The purpose of this paper is to examine changes in the relationship between employment and unemployment in the United Kingdom over the past two decades. In particular we concentrate on the question of why unemployment remained high throughout the first half of the 1980s despite strong growth in employment since 1983 (unemployment only beginning to fall from mid-1986 onwards) and why, subsequently, the claiment total has fallen so quickly. We model changes in unemployment taking levels of employment as given. At first blush this might seem to be a fairly straightforward task. In fact, however, we find that it is important to distinguish between employment in different sectors, between full-and part-time workers and between male and female workers. Our results also suggest a role for Special Employment Measures in reducing the claimant count. Especially important in recent years have been the Restart programme and the extension of availability-for-work tests.

The paper is organised as follows. In Section 2 we discuss recent trends in the labour force, highlighting the persistence of high unemployment rates during the 1980s despite strong employment growth. Next, in Section 3, we show the inadequacies of the simple "rules-of-thumb" unemployment models used in many macro-models (including the Bank's) during the 1960s and 1970s to link changes in unemployment with changes in employment (it was the failure of the Bank's model to track either the rise in unemployment which occurred during the late 1970s and early 1980s or its subsequent decline which motivated this study). We then detail our first attempts to improve upon this type of model by estimating an unemployment equation containing a number of employment variables disaggregated by sector. Although the resulting model is a great improvement on its predecessor it does not distinguish between full and part-time workers or between male and female employees.[1] Hence, we also report our more recent attempts to improve the model further in these respects (Section 4). Using our preferred specification we next analyse the persistence of unemployment during the 1980s (Section 5), finding that changes in the sectoral composition of the workforce and the rise in the number of part-time workers (mainly women and concentrated in the service industries) are crucial to our understanding of this phenomenon. Finally, in Section 6, we report our conclusions.

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1 The new model also had some interesting problems regarding its dynamic properties which we also felt were worth looking at more closely. Details are given in Section 3 below.

Section 2: Recent trends in unemployment, employment and the labour force

Two decades ago unemployment[2] stood at close to 500,000. Since then, however, there have been just six years during which unemployment has fallen but fourteen during which it has increased. Moreover, the magnitude of the rises which have occurred have far outweighed those of the falls. As a result, unemployment now stands at close to 1 3/4 million, having reached nearly 3 1/4 million in June 1986. Chart 1 illustrates these developments, indicating that much of the increase in unemployment occurred in two short 'bursts', in 1975/6 and between 1980 and 1983. On each of these occasions most of the rise was associated with an increase in unemployment duration, so that the share of long-term unemployed (defined for our purposes as those out of work for more than six months) increased from below 15% in 1974 to around 30% at the end of 1976 and from close to 25% at the beginning of 1980 to 55% at the end of 1983. This does not mean, of course, that changes in unemployment have not been matched by changes in employment of broadly equal magnitude (although, obviously, of the opposite sign). Chart 2 shows annual movements in the total employed labour force (which comprises employees in employment, the self-employed and the Armed Forces) and compares them with annual changes in unemployment.

Changes in unemployment would be matched one-for-one by changes in employment were it not for two important facts: first, their sum is not constant because the relevant population is not constant. Moreover, certain definitions of unemployment - such as those actively seeking work - mean that the relevant populations may vary systematically and the factors causing these variations may also affect employment. Second, and underlying this observation, is the fact that there are not two but three states, the third being that of not participating in the labour force. The allocation of individuals to these three states depends on the definition of unemployed that is used.

In recent years the UK definition has changed from a registration to a claimant basis. Many other countries (and the Labour Force Survey) use a definition which includes as unemployed only those both available for work and seeking it (they, together with those in employment, constitute the labour

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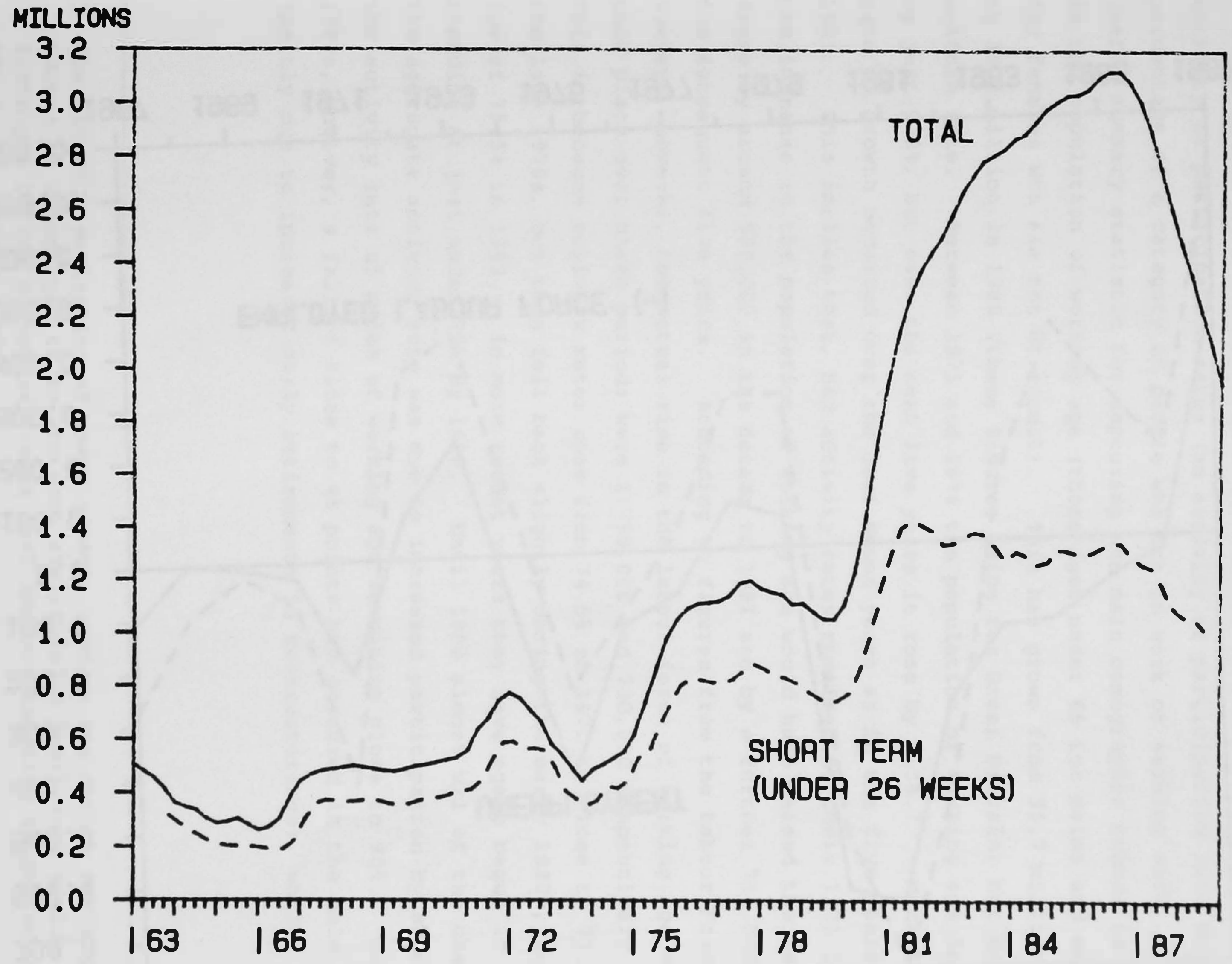
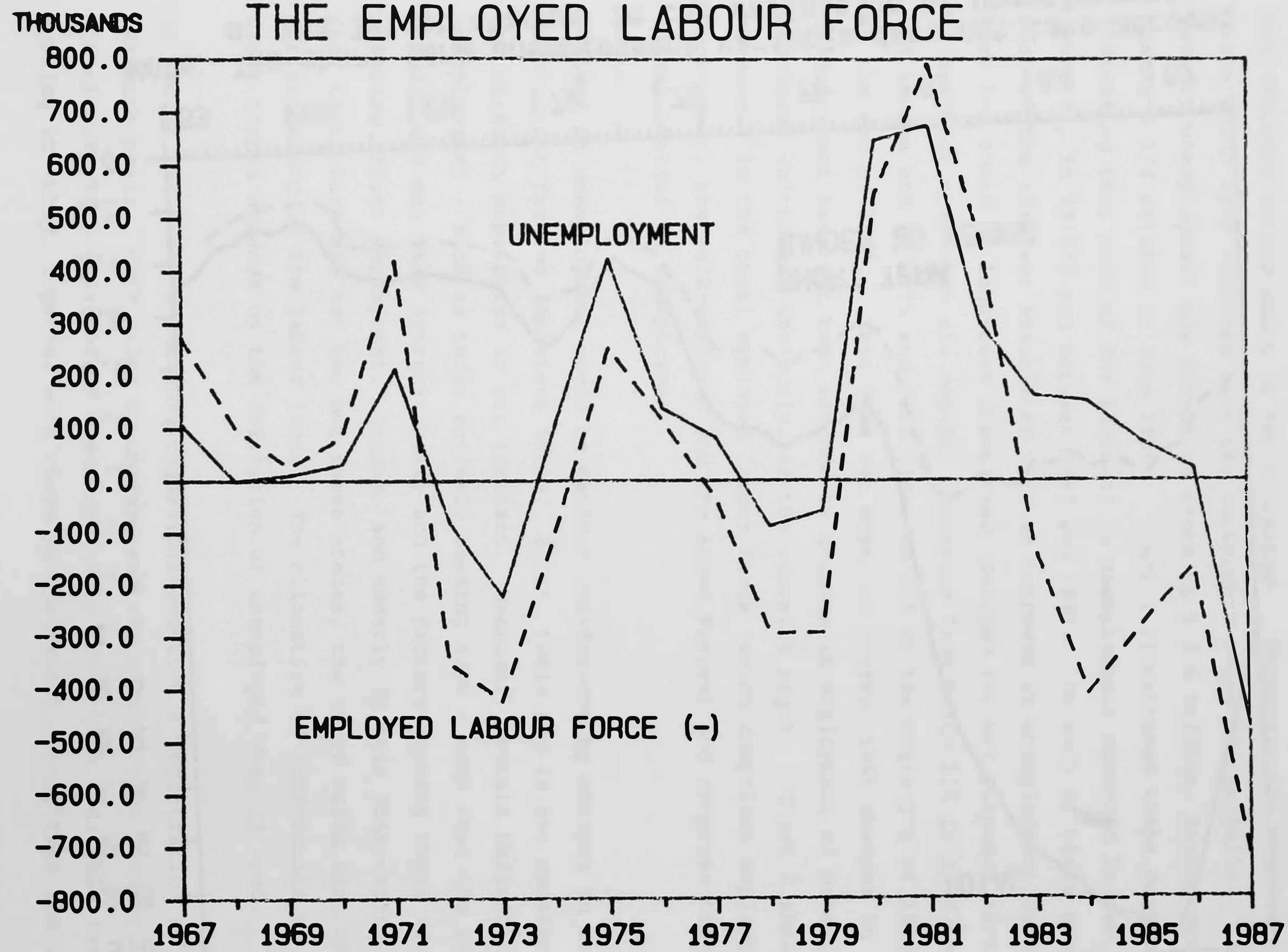
2 In this note we use unemployment to refer to the total number of unemployed claiming benefit, seasonally adjusted.

CHART 1 : UNEMPLOYMENT (U.K., SEASONALLY ADJUSTED)



Note: The short term unemployment variable does not take account of the latest changes in the definition of unemployment.

CHART 2 : ANNUAL CHANGES IN UNEMPLOYMENT AND THE EMPLOYED LABOUR FORCE



force).[3] The 1988 Labour Force Survey suggests that these two approaches give similar totals as about 800,000 claimants were either inactive or already had jobs while around 750,000 job seekers were not claiming any benefit.

Changes in the labour force will reflect both changes in demographics and in activity or participation rates (an activity or participation rate being the percentage of a category of people who are in work or seeking work). A useful summary statistic for capturing the main demographic trends is changes in the population of working age (those aged under 65 for males and under 60 for females who are not at school). This has grown from 31.7 million in 1971 to 34.2 million in 1988 (these figures being for Great Britain) but not at a uniform rate. Between 1971 and 1976 the population of working age increased by just 0.8%, but over the next five years it rose by 3.1%.[4] Much the same rate of growth occurred over the last seven years as in the five years to 1981. This implies that, had activity rates remained at their 1971 levels, the increase in the population of working age would have raised the labour force by around 900,000 in the decade to 1981 and by a further 750,000 during the subsequent five years. According to figures from the Labour Force Surveys, however, the actual rise in the labour force of working age which took place over these periods were 1,750,000 and 700,000 respectively.[5] This is because activity rates rose from 74.5% in 1971 to close to 77.5% in the late 1970s, but then fell back slightly during the early 1980s, reaching a low of 75.5% in 1983. In more recent years they have again begun to rise, standing at just under 79% by 1988. Until 1980 almost all of the change in the aggregate activity rate was due to increased participation by women, with the activity rate of males of working age remaining close to 90%. During the 1980s, however, a fall of close to 4% points has occurred in the male rate (partly due to increased early retirements in restructurings), whilst that of

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3 The ILO/OECD definition of unemployment, used in the Labour Force Survey, counts as unemployed those who were available to start work within the next fortnight of the survey reference week and who had also sought work within the previous four weeks.

4 Note that the rise between 1971 and 1976 has been calculated on the basis of assuming the school leaving age was 16 throughout the period. In fact, however, it was only 15 in 1971. (Because of this the actual population available for work is likely to have fallen considerably over this period).

5 These figures relate to the civilian labour force (ie they exclude those in the Armed Forces) and again assume the school leaving age was 16 throughout the period.

females remained fairly stable in the early 1980s (falling from 64.3% in 1980 to 63.5% in 1983) and more recently has begun to grow quickly again (reaching 69.4% by 1988). As a result, although the employed labour force increased from 23.5 million at the beginning of 1983 to 25.4 million in the middle of 1988, unemployment still continued to rise (although at a decreasing rate). Only in the last two and a half years has growth in the labour force been sufficiently slow that further increases in employment have resulted in falling unemployment.

Section 3: A simple model of the relationship between employment and unemployment

The fact that participation is endogenous means that neither of the two identities:

$$\Delta LU_{SW} = \Delta LF - \Delta LE \qquad (1)$$

or $$\Delta LU_C = \Delta WP - \Delta LE \qquad (2)$$

provides a satisfactory basis for modelling (where LU_{SW} is the number of people without jobs seeking work, LF is the labour force, LE is total employment, LU_C is claimant unemployment and WP the working population). While models can generate forecasts of employment they have problems with both the labour force and the working population. Since we are interested in explaining and predicting the claimant measure it is (2) and the problem of modelling the working population that concern us.

A very simple model suggests that a constant normal proportion α, where $0<\alpha<1$, of an increase in the relevant population will join the working population which will also be increased to the extent $(1-\beta)$, where $0<(1-\beta)<1$, that additional employment draws those previously not participating either into jobs or into job seeking (and benefit claiming - perhaps by trying to meet availability criteria). Thus:

$$\Delta WP = \alpha\ \Delta POWA + (1-\beta)\ \Delta LE \qquad (3)$$

where POWA is the population of working age.

Together (2) and (3) imply that:

$$\Delta LU_C = \alpha\ \Delta POWA - \beta\ \Delta LE \qquad (4)$$

The proportion α of an increase in the population of working age joining the working population would not necessarily be the same for different sexes, regions etc; and participation is also a function of age within the range covered of POWA. Similarly, and indeed partly for this reason, additional jobs of different types are likely to have different effects on participation and the working population.

Thus (4) could be improved by disaggregating the explanatory variables:[6]

$$\Delta LU_c = \Sigma\, \alpha_{ijk}\, \Delta POWA_{ijk} - \Sigma\, \beta_{ijklm}\, \Delta LE_{ijklm} \qquad (5)$$

where i relates to sex, j to age, k to region, l to industrial sector and m to full/part-time.

Availability of data and limited variation of POWA by sex, region and age mean what we actually estimate is:

$$\Delta LU_c = \alpha\, \Delta POWA - \Sigma\, \beta_{ilm}\, \Delta LE_{ilm} \qquad (6)$$

Throughout much of the 1960s and 1970s most of the macro-models used simple "rule-of-thumb" equations of this type to link changes in employment and unemployment. Thus, for example, the Bank model used an equation disaggregated only by employment sectors of the form;

$$\Delta LU = \alpha\, \Delta POWA - \beta\, \Delta(LEMF + LHMF) - \gamma\, \Delta(LEG + LOTH + LSE) \qquad (7)$$

where LU refers to registered unemployment, LEMF is employment in manufacturing, LHMF is employment in the armed forces, LEG is employment in general government (including the armed forces),[7] LOTH is employment in non-manufacturing (excluding general government) and LSE is self-employment.

None of the coefficients in this model were estimated. Rather they had been imposed at $\alpha = 0.5$, $\beta = 0.8$ and $\gamma = 0.65$. Nevertheless they gave a "reasonably good" description of unemployment during the 1960s and early 1970s in terms of tracking performance - as is shown by Chart 3. During the 1980s, however, the model failed to predict either the persistence of high unemployment in the period 1983 to 1985 or the sharpness of recent falls. The simplest explanation for this breakdown is that the imposed coefficients

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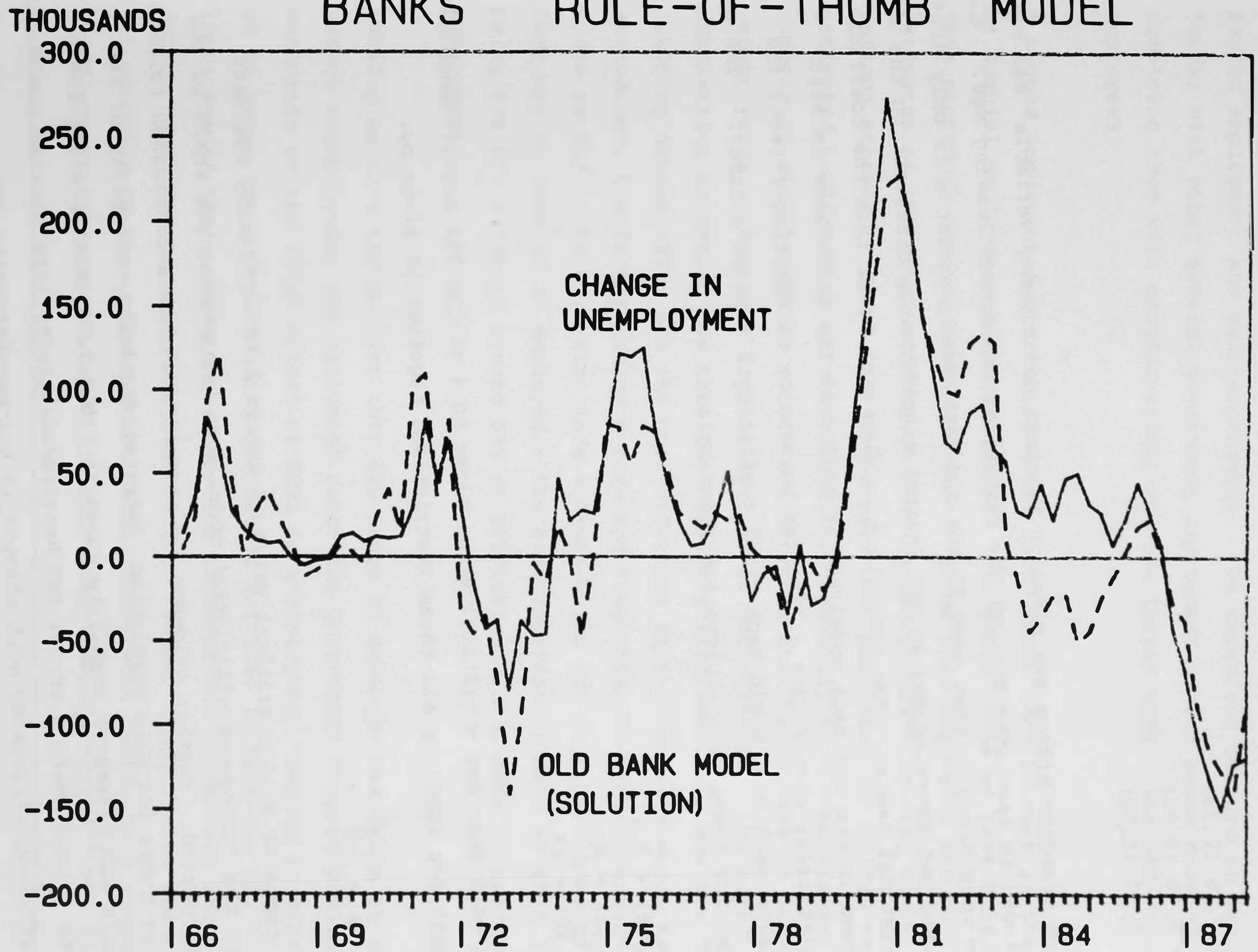
6 In practice it will also be necessary to make some adjustments for the effects of Special Employment Measures. These are discussed below.

7 Note that this means that LHMF effectively enters the equation twice, once with a coefficient of 0.8 (β) and once with a coefficient of 0.65 (γ). Possibly this reflects the fact that when armed forces employment rises some of the increase is based abroad and these employees may take their (previously unemployed) spouses with them! However, we know of no (even anecdotal) evidence to support this proposition. Perhaps a more likely explanation is that this was simply a mistake.

CHART 3 : THE TRACKING PERFORMANCE OF THE BANKS' "RULE-OF-THUMB" MODEL



were wrong. Of course, these reflected the priors of those constructing the model, as regards registration rates and participation effects, and so are hard to defend. Nevertheless, as a first step we have estimated (7) using seasonally adjusted data covering the period 1964 to 1987. The freely estimated coefficients are;

$\hat{\alpha} = 0.25\ (0.14)$

$\hat{\beta} = 0.90\ (0.06)$

$\hat{\gamma} = 0.37\ (0.05)$

where standard errors are given in brackets and a dummy variable[8] has been included to take into account the raising of the school leaving age. Clearly $\hat{\beta}$ is very close to that used in the old Bank model, whilst $\hat{\alpha}$ is only half that imposed but nevertheless is not (quite) significantly different on the usual statistical tests. The estimate $\hat{\gamma}$ is also much lower than the coefficient imposed in the old Bank model but in this case the difference is also statistically significant. This can be seen by calculating F-tests for imposing each of the old Bank model coefficients (one at a time). The test values for these restrictions are shown below;

$F\ (1,90) = 3.3$

$F\ (1,91) = 0.5$

$F\ (1,92) = 24.4$

where each test has a critical value close to 4 at the 95% significance level. The test for imposing all three restrictions together is given by;

$F\ (3,86) = 7.5$

which compares with a critical value of around 2.75. Clearly the data rejects the 'imposed-coefficients' model when tested over the whole of the 1964-87 period. Further evidence (if needed) for such a conclusion follows from our tests for autocorrelation, heteroscedasticity, normality of the residuals and parameter stability shown in Table 1, Column A, since these indicate autocorrelated errors and heteroscedasticity (the squared residuals being strongly correlated with changes in non-manufacturing employment). Moreover, although the old Bank model predicted well throughout the period 1985:1 to 1986:2, it cannot forecast the more recent falls.

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8 The dummy variable takes the value 1 in 1973 Q3 and 0 elsewhere.

One obvious way of improving on (7) might be to lift the restrictions implied by including the sum of manufacturing and armed forces employment as one term and the sum of self-employment, general government employment and other non-manufacturing employment as another. Unfortunately, however, when this was tried it was found difficult to identify sensible coefficients on armed forces employment and self-employment. We therefore decided to include the former with other general government employment (which seems more natural than combining them with manufacturing) and the latter with other non-manufacturing employees.

A second important factor which needs to be taken into account is the effects of Special Employment Measures (SEMs). To gauge these we have followed the same procedure used by Keating (1986) - ie we have regressed the published estimates of the total effects of SEMs on unemployment against numbers on the various schemes. Since the former were only published until 1986 the estimated coefficients were combined with the still-published figures for the numbers on each scheme to obtain estimates of the total effects in more recent years. Appendix 2 gives details of our workings - the main conclusions to emerge are that the Community Programme (CP) and the Youth Training Scheme (YTS) are the most effective of the measures which have been introduced, the former having a combined "registration-effectiveness"[9] of close to 0.9. We have also made allowance for the fact that some of those on SEMs may be counted as employed - the most important scheme in this respect being the CP, although around 15% of YTS trainees also have contracts of employment.

Finally we have the problem that the range of measures available to help reduce unemployment has broadened since the Government stopped publishing estimates of the total effect of SEMs on unemployment (making it difficult for us to gauge their impact using the method outlined in Appendix 2). The most important of these changes has been the introduction of Restart interviews (which began in the second half of 1986). These, coupled with the introduction of more rigorous availability-for-work tests, help reduce unemployment because they not only help long-term unemployed[10] to find jobs,

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9 ie what happens to unemployment if the number of people on the scheme is increased by one. Note that the unemployment figures published during the period for which the SEMS effects were available was the registered unemployment series.

10 Originally the interviews were of those unemployed for more than a year. Now the scheme has been extended, however, to include all those unemployed for more than six months.

but they increase the numbers going on to other schemes and remove from the register those who are ineligible for benefits. In the Pilot Schemes around 22% of those interviewed were offered places on the CP, whilst around 10% left the register. Keating argues that it is reasonable to assume that the nationwide schemes might have similar effects, but we choose instead to include as a regressor the number of Restart interviews carried out each quarter (denoted RES). Although in our main model this gives us very few observations with which we can gauge the impact of the scheme on unemployment the results we obtain accord with those found using much the same method but monthly data and hence around two dozen observations (see Appendix 3 for details).

Using our estimates of the effects of SEMs on employment[11] and on unemployment (each prefixed by SEMS) we define the following terms in our model;

ULU = LU + SEMSLU

ULEMF = LEMF - SEMSLEMF

ULOTHA = LOTH + LSE - SEMSOTH

ULOTHB = LOTH + LSE + LEG - SEMSOTH

Where LU refers to claimant unemployment (this notation is maintained throughout the remainder of the paper). ULU can be thought of as "underlying" unemployment, in the sense that it is the level of unemployment which would result were there no SEMs. (Similarly, ULEMF and ULOTHA are "underlying" levels of employment.) The unemployment equation we seek to estimate takes the form;

$$\Delta ULU = \alpha\ \Delta POWA + \beta\ \Delta ULEMF + \gamma\ \Delta ULOTHA + \psi\ \Delta LEG + \phi\ RES \qquad (8)$$

In practice we found that when estimating equations like (8), first order autocorrelation was a common problem. For this reason we have included lagged values of both the dependent variable and the regressors in our model. A (potential) problem arises if we want to include lags, however, in that we might expect an asymmetric response from changes in employment according to the sign of these changes. When employment is rising we might expect quick falls in unemployment (with little in the way of lagged effects) but when jobs are being lost there may be considerable delays before those being made

11 Again Appendix 2 gives details.

redundant actually register as unemployed. In practice we found that adding a lagged dependent variable improved the specification considerably but did not eliminate the problem of autocorrelation. To do this a lag on manufacturing employment was also needed.

Table 2 shows our results. Column A shows the simplest version of (8) (with no "dynamic" effects) whilst Column B shows what happens if we add a lagged dependent variable. Finally, in Column C, we show the model we prefer. This has been obtained by testing down using the "general-to-specific" methodology made popular by Hendry (1978). Originally a sample period ending in 1984 Q4 was used to search for the "best" model (and this is what we have used to carry out the relevant parameter stability (forecast) tests). Then we have tried extending the sample period to include more recent data, adding in the Restart variable.[12]

The specification shown in Column A suggests very low registration rates for both the self-employed and for employees working outside manufacturing. For those in manufacturing, however, it suggests a rate close to one. This marked contrast could perhaps be due to the fact that the proportion of male workers (who would, in the main, be full-time employees) varies across sectors. Alternatively, it may be associated with different industrial composition in terms of skill levels - manufacturing industries employing a higher proportion of manual workers, for example. Despite its improved standard error model A is clearly inadequate, since it suffers from autocorrelation. It also fails to forecast the period 1985 to 1987 accurately, although this is due to our having to drop the Restart variable from the model when carrying out such a test. (Hence a six-period forecast from 1985 to mid-1986 is easily passed.) One further point concerns the effect of SEMs. We have constrained these to equal our estimates (outlined in Appendix 2). In fact, however, we find the relevant F-test pertaining to the coefficient on SEMSOTH is failed at the 95% significance level (the $F(1,86)$ test of 7.5 compares with a critical value close to 4).[13] The freely estimated coefficients imply a larger effect than we had gauged. Nevertheless, given that there are other problems with the model (which may be giving us biased coefficients) we decided to keep our effects imposed on the model.

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12 Obviously this had to be excluded from our previous search since Restart only began in 1986.

13 This is sufficiently high to ensure that the test on all three restrictions taken together is also failed ($F(3,85)=4.0$).

Model B has only slightly better statistical properties than A - a slightly smaller standard error and not quite such strong evidence of autocorrelation (although tests regarding the latter are still significant at the 95% level). Moreover, new problems are evident in that several terms are found to have insignificant coefficients. Given that the effect of a change in general government employment appears to be very close to that of a change in non-manufacturing employment it was decided to amalgamate these two categories. The failure to identify a strongly significant effect on the dummy variable (designed to pick-up the change in the school leaving age) is rather more surprising. Here we found that the data preferred long lags on the population of working age variable. Making these two changes improved the model considerably. Nevertheless, the problem of autocorrelation remained - the relevant LM test just being failed at the 95% level. For this reason we decided to add lags to all of the explanatory variables and test-down using the Hendry "general-to-specific" method. Of course this approach implies that not only will activity and "claimant" rates vary across industry but also that the dynamics of the response of unemployment to changes in employment will depend upon industry.

Column C shows our preferred specification - only a lag on manufacturing employment proved to be significant, apart from the lagged dependent variable. Nevertheless, the new specification performs well. Its standard error is around one-half of that relating to the old Bank model, there are no problems of parameter stability,[14] autocorrelation, heteroscedasticity or non-normality of the residuals and the long-run activity and "claimant" rates accord well with those obtained earlier. For non-manufacturing we still find a very small long-run "claimant" rate (of below 20%), whilst for manufacturing the corresponding figure is very close to one (0.96). Moreover, we still find that under 40% of any changes in the population of working age feeds through to unemployment and that close to one in five of those interviewed under the Restart programme leave unemployment. To test whether these effects are sensitive to the sample period chosen we have re-estimated the model using recursive least squares (dropping the Restart variable obviously). The results are fairly encouraging regarding the stability of the model - sequential Chow tests indicating very few significant values. Serial correlation is slightly more of a problem, however, with 9 of the 67 Von

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14 The 3-year forecast simply serving to emphasise that a model without a Restart effect is inadequate.

Neumann ratios being significant at the 95% level.[15] Interestingly, nearly all of the significant values occurred when we used samples which ended before 1980; thereafter we rarely found a test which failed. Given the small "claimant" rates implied by our model outside manufacturing we have also checked whether or not this is a common phenomenon throughout our sample period. The coefficient pertaining to LOTHB changes from close to -0.21 in samples ending in the early 1970s to around -0.08 in one which ends in 1977. (Here we have started with a window of 25 observations beginning in 1964 and gradually added more observations to the sample.) As the sample is extended still further the coefficient begins to fall, reaching a nadir of -0.16 around 1981. As more recent data is added, however, it again begins to approach zero. By 1985 the coefficient is again close to -0.09. Trends in the coefficient on the lagged dependent variable offset movements in the coefficient on LOTHB to some extent. Nevertheless we find that the model's implied long-run "claimant" rate for this group changes from close to 0.25 in the early 1970s to around 0.16 in a sample ending in 1977. It then increases sharply again as more observations are added (to a high of 0.30 around 1981) before falling to close to 0.17 by the mid-1980s. Much of the shift in these rates during the 1980s is likely to be related to the faster growth of female part-time employment in service industries over this period since few of the new workers are likely to have been able to claim benefits when out of work. Changes over the earlier period, however, are less easy to explain.

One of the reasons suggested to us for our having found instability in the coefficients might be that we have used the population of working age instead of the civilian labour force. The civilian labour force might be thought to be preferable to the population of working age because it more closely approximates the labour supply. However, by choosing to use it instead of POWA one might simply introduce more problems since it is based not on claimant unemployed but on those seeking work (see the discussion on page 5 above). It is important to note that, although using the civilian labour force might enable us to explain more accurately past changes in unemployment, given changes in employment, than if we used the population of working age, it is by no means obvious that it is a better approach to take in trying to develop a model for use in forecasting, since one is still left with the

15 Although we discovered that our results were sensitive to the size of the "window" (here we used 25 observations). A larger window (of 40 observations) gave stronger evidence of serial correlation, there being 24 significant values out of a total of 52 tests.

problem of forecasting future movements in the labour force. Although this could be remedied by using instead the Department of Employment's projections (see, for example, Department of Employment (1988)) this too has its problems, since their projections assume a "roughly constant" level of demand in the labour market. Since this amounts to assuming that the unemployment rate will remain broadly constant, then this could clearly be an inappropriate assumption to make if the forecast from the macro-model and the labour force projections are to be internally consistent. At one time the Treasury did use this approach in forecasting unemployment, since the Department of Employment were able to revise their published projections according to what the Treasury were forecasting for labour demand. In practice, however, iterating between forecasts from the macro-model and different sets of projections from the Department of Employment proved to be a time-consuming process and has since been abandoned.

It was also suggested to us that our having used the population of working age rather than the civilian labour force might also explain why the coefficient on ULEMF is so high whilst that on ULOTHB is close to zero. In order to test these propositions we have tried substituting estimates of the civilian labour force (all ages) made by the Department of Employment (see Department of Employment (1987) and (1988)) for POWA. Interestingly our results are very similar to those reported in Table 2, although the long-run elasticity regarding LOTHB does increase slightly, from 0.15 to 0.30. The offset to this change, however, is a rise in the long-run elasticity on ULEMF from 1.00 to 1.05. In the remainder of our work we retain POWA as our measure of labour supply. Our results so far rather suggest that we need to consider further disaggregation of our employment variables if we are to explain the changes in "claimant" rates which have occurred during the past two decades.

Section 4: Improving the basic model

Although the methodology described in the last section enabled us to improve upon the old "rule-of-thumb" unemployment equation, there are still some further problems that need to be resolved. The first of these is to determine the importance of part-time workers in explaining unemployment. The second is to consider whether an explicit male/female split is relevant. Finally, a closer examination of the equation's error dynamics is required. We have already seen that the split between manufacturing employment and other employment is important, but we must also determine whether the composition of employment within a sector is relevant. In particular, we need to ask if the incidence of part-time working helps explain why the coefficient on manufacturing employment is close to unity while the coefficient on non-manufacturing employment is low.

Chart 4 shows cumulative changes in employment since the beginning of the sample period.[16] Clearly there have been large sustained falls in manufacturing employment throughout the entire period, with periods of fast growth in the economy or recession being marked by a slight fluctuations around this long-run trend. Outside manufacturing, however, employment has generally been rising, or at worst flat. For the purpose of illustration in Chart 4 non-manufacturing has been defined as 'other' employees plus the self-employed and the armed forces, so that the sum of the two lines gives cumulative changes in the employed labour force. Looked at in this light many of the gains and losses in sectoral employment appear to cancel, with little net change in the employed labour force over the sample period, the main exception to this being the falls recorded during the period from 1980 to 1982. Nevertheless, the fall in non-manufacturing employment at this time was not as large as that experienced in manufacturing and was reversed by the end of 1984. In fact, growth has continued strongly in these 'other' sectors since early 1983 mainly because of the strength of the private service sector.

Chart 5 shows cumulative changes in claimant unemployment over the last two decades. The close correlation between falls in manufacturing employment and rises in unemployment is clear. However, for much of the period employment outside manufacturing rises while unemployment rises. Only in the period 1980 to 1983 does 'other' employment fall significantly while unemployment increases.

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16 Although earlier results were estimated over the period 1964:3 to 1987:4, data limitations require us to estimate subsequent regressions over a shorter period from 1971:3 onwards.

Clearly there is very little evidence in these charts to suggest why unemployment has fallen so strongly over the last 18 months of the sample period. Manufacturing employment has risen very slightly, but not sufficiently fast to explain the recent changes in unemployment, whilst in the 'other' sector employment growth has accelerated through 1986 and 1987, compared with the early 1980s, but again not sufficiently fast to explain the change in unemployment. In fact, between 1983 and 1986 'other' employment was already growing strongly at a time when unemployment was rising. Hence, if we are to explain the recent fall in unemployment then what would be required is either a change in the composition of that growth or a change in magnitude of the coefficients (ie parameter instability of some form).

Ideally, we would like to be able to disaggregate by sex and between full and part-time workers. However, data on male part-time employees are not available on the industry split we require, at least for the period up to 1984, and so we have concentrated our efforts on discriminating between female full and part-time employees (for which data are available). Since fewer than 20% of part-time employees are males (17% in December 1987) this simplification is unlikely to make much difference to our results.

Employment outside manufacturing has been broken down into male employees, female full-time employees, female part-time employees, and 'other'.[17] This sector as a whole comprises some 20 million people, of which approximately 17 million were employees. These are divided almost equally between men and women and of the women just under half are part-time employees. This implies that at the end of 1987 approximately 90% of all female part-timers outside manufacturing. In manufacturing, by way of contrast, there are some 5 million employees, of whom around 70% are men and 30% women, but of the women only 20% are part-time. This implies that approximately 6% of manufacturing employees are part-time. It would therefore seem reasonable to assume that the incidence of part-time working in manufacturing industry has remained much lower than in other sectors, particularly services, where part-time working has grown in significance. In fact we found, by disaggregating the manufacturing employment figures, that movement in female part-time and full-time workers has mirrored movement in male employment much more closely than in the non-manufacturing sector, suggesting that even the sex split may not be very important. For this reason we retain ULEMF as the sole regressor pertaining to manufacturing employment.

17 Primarily self-employed plus HM Forces.

Chart 6 illustrates the cumulative changes that have taken place in male and female employment over the last decade and a half. Growth in female part-time employment has been almost continuous with the exception of 1980-83 when it fell slightly, with this component accounting for around 50% of total growth in non-manufacturing employment over the period. Male employees accounted for around 10% of the rise while the self-employed and female full-time employees each accounted for close to 20%.

Within these overall trends there are some interesting features. Male employees are the only group whose numbers have fallen strongly during the early 1980s, although a small decline also occurred in the number of female full-time employees. Since 1983 the full-time males and females categories have grown broadly in line (in terms of numbers) although much more slowly than has female part-timers. The number of self-employed tended to decline between 1971 and 1978, but since then has risen. Around 1983 growth in self-employment rose sharply and, since then, it has mirrored growth in female part-time employees closely.

In order to use these components we have chosen to estimate an unemployment equation of the form;

$$\Delta ULU = a_0\Delta POWA + a_1\Delta ULEMF + a_2RES + a_3\Delta UFMFT + a_4\Delta UFMPT + a_5\Delta UMALE + a_6\Delta UOTHER + a_7\Delta SEMSOTH \qquad (9)$$

where the new terms are defined as follows;[18]

UFMFT = Female full-time employees in non-manufacturing
UFMPT = Female part-time employees in non-manufacturing
UMALE = Male employees in non-manufacturing
UOTHER = Self-employment plus the Armed Forces

Note that the new employment variables are defined in such a way as to sum to the total of self-employment plus non-manufacturing employees in employment;

$$UFMFT + UFMPT + UMALE + UOTHER = LOTH + LEG + LSE$$

The new preferred equation was again arrived at by testing down using the "general-to-specific" methodology. Again this raises a number of problems

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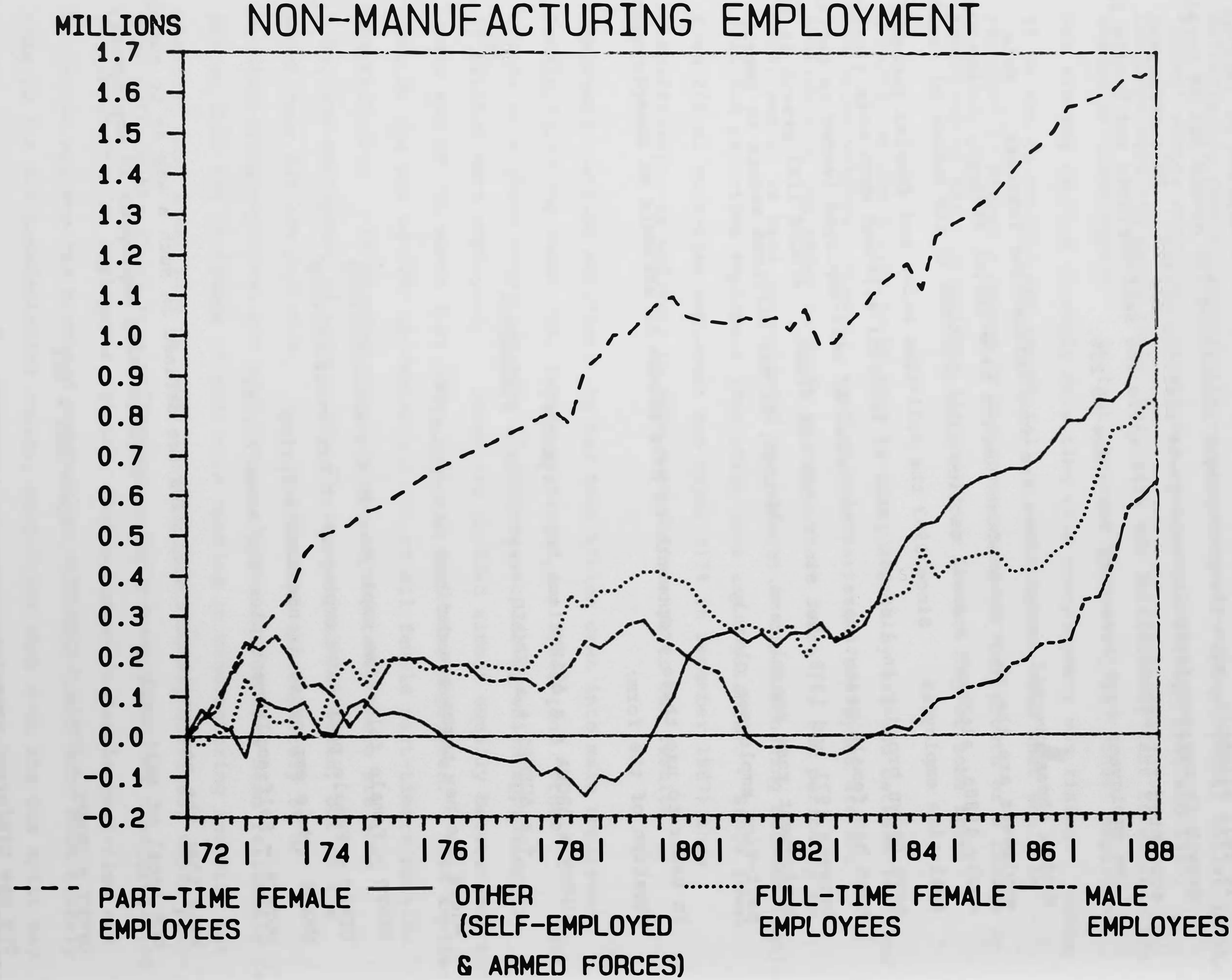
18 We discuss the effects of SEMs below.

CHART 6 : CUMULATIVE CHANGES IN COMPONENTS OF NON-MANUFACTURING EMPLOYMENT



with regard to the dynamics of the equation, however, since by including a lagged dependent variable we ensure that all of the variables in the equation are subject to dynamic effects. This is particularly a concern with respect to Restart since we have few data observations with which to test whether Restart interviews have continued effects in subsequent periods.[19] Instead, it seems likely that a Restart interview should have a one-off effect in the period in which it occurs. There are, however, a number of possible routes through which dynamic Restart effects could occur - it is possible, for example, that some long-term unemployed might choose to leave the register before they are interviewed having observed the effects of previous interviews. This might occur if, for example, they were not entitled to benefit or not searching for work.

We have already shown how inclusion of a lagged dependent variable and one lag on manufacturing employment produced equation C (Table 2). This model has the following dynamic responses after shocks to its exogenous regressors;

Percentage of long-run effect after x quarters

No of quarters x	ΔULEMF	ΔULOTHB, RES, POWAA
1	77%	48%
2	86%	72%
3	91%	86%
4	92%	93%
5	94%	96%
6	95%	98%

Clearly the impact effects of changes in the regressors other than that relating to manufacturing employment are small relative to their long-run effects, which seems rather implausible. One method of eliminating this type of problem is the use of autoregressive (AR) error structures, whereby the regressors and dependent variable enter the equation with no lagged effects and the error term follows an AR process. General-to-specific testing showed that no second lags entered the regression with significant coefficients, implying that we could simplify the structure to AR(1). Another potential solution is to impose a common factor restriction on the Restart variable (which can, of course, be tested) but with other equation dynamics being allowed to operate freely.

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19 Monthly data might be more useful for testing dynamic effects. A problem arises here, however, in that most of the employment series are not available at this frequency.

To illustrate these ideas consider first the following general dynamic model;

$$\Delta ULU = a_0\Delta ULU_{-1} + a_1\Delta POWA + a_2\Delta POWA_{-1} + a_3\ RES + a_4\ RES_{-1} + a_5\Delta ULEMF + a_6\Delta ULEMF + a_7\Delta UFMFT + a_8\Delta UFMFT_{-1} + a_9\Delta UFMPT + a_{10}\Delta UFMPT_{-1} + a_{11}\Delta UMALE + a_{12}\Delta UMALE_{-1} + a_{13}\Delta UOTHER + a_{14}\Delta UOTHER_{-1} + a_{15}\Delta SEMSOTH + a_{16}\Delta SEMSOTH_{-1} + \epsilon_t \quad (10)$$

where ϵ_t is white noise error.

Now assume a common factor within the Restart dynamics such that;

$$a_4 = -a_3 a_0$$

If this restriction were to hold then the long-run coefficient on Restart would equal a_3 and the long-run effect would be equal to the impact effect. A simple Wald test can be used to test for the common factor. Furthermore, we could impose common factors on all exogenous regressors. Equation (10) could then be written as;

$$\Delta ULU = a_1\Delta POWA + a_3 RES + a_5\Delta ULEMF + a_7\Delta UFMFT + a_9\Delta UFMPT + a_{11}\Delta UMALE + a_{13}\Delta UOTHER + a_{15}\Delta SEMSOTH + U_t \quad (11)$$

where $U_t = a_0 U_{t-1} + \epsilon_t$

A complex dynamic equation has thus been transformed into a simple model with dynamic errors by imposing common factor restrictions.

The first stage of our estimation process was to estimate equation (10) in unrestricted dynamic form. A Wald test was then performed on each of the common factor restrictions with the following results;

Test	t-value
a_2	-0.44
a_4	0.73
a_6	1.19
a_8	0.34
a_{10}	0.49
a_{12}	0.46
a_{14}	-0.35
a_{16}	-0.14

The Wald test for joint imposition of all the common factors took the value of 4.16. This has a chi-squared distribution with a critical value of 15.5.

All of the tests are passed at conventional levels of significance, the largest t-value being found on the manufacturing employment variable. The fact that (jointly) the common factor restrictions pass is not surprising given the low t-values on many of the employment variables. It does, however, suggest that the dynamic equation can be simplified to a simple AR(1) process.

At a fairly early stage it was decided to simplify the model by attributing the SEMs effects on employment among the variables so that we could retain the notion of "underlying" employment suggested earlier. There was assumed to be little effect upon self-employment or part-time employment, since the bulk of YTS and Community Programme places fall primarily within the service sector. Hence, SEMSOTH was allocated to full-time females and to full-time males in other sectors, by defining;

$UFMFTS = UFMFT - \delta\ SEMSOTH$

$UMALES = UMALE - (1-\delta)\ SEMSOTH$

Choice of δ is discussed below.

Table 3 sets out some of the estimates we have obtained, with each of the models being static equations with AR(1) error structures. As regards equation D, the coefficients on female employees and 'other' (mainly self-employment) all turn out to be small in absolute terms and insignificant. Male employees, however, attract a large (ie more negative) coefficient which is strongly significant. The standard errors on all four of these variables are of broadly similar magnitude implying that it is the coefficients' small (absolute) size that is producing the low t-value in the case of female employees and the self-employed. Hence, one might argue that a more appropriate test is whether the coefficients are significantly greater than <u>minus</u> one. The results of carrying out these tests are shown below;

Variable	t test
ΔUFMFTS	11.73
ΔUFMPT	9.30
ΔUMALES	8.82
ΔUOTHER	11.59

All four variables are found to have coefficients which are significantly greater than minus one.

We have also tried imposing the restriction that the coefficients on female full-time and part-time employees and self-employed are equal. This improves marginally the standard error of the equation but, not surprisingly, the coefficient pertaining to the female employees plus self-employed category is still insignificantly different from zero.[20] Hence Equation E (Table 3) drops these employment categories completely.[21] All the remaining coefficients are significantly greater than zero with the exception of that pertaining to POWA. (It should be noted that the latter is still very small in absolute terms.) The F-test for the restrictions from D to E is passed with a value of $F(3,57) = 0.90$.

The equations shown in Table 3 both assume that SEMSOTH falls entirely on male employees. Different values of δ were tried but the results proved to be fairly insensitive to our choice. Given the decision to eliminate the female employees variable from the model it was therefore decided to retain the SEMs employment variable within the male employees category, rather than drop it altogether.

So far we have not tested whether the coefficients on the other employment variables are significantly different from each other. If they are not then we could accept an equation which retains the same employment split used earlier (column C of Table 2) although keeping the auto-regressive error structure. In fact, however, the relevant test statistic is significant at the 1% level $F(3,57) = 4.19$ implying that we should retain the disaggregated equation.

Equation E is also preferred over the equation F shown in Table 4 (in which we have re-estimated equation C over the shorter sample period) since the standard error of E is lower than that of F. E also has the advantages that its dynamic properties are more acceptable than that of F and its 'other' sector employment variable is more significant.

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20 Note that retaining just female full-timers also proved not to be possible.

21 One possibility suggested to us, which we have not yet tried exploring but which might enable us to retain these variables, is to try modelling male and female unemployment separately.

A number of issues relevant to our modelling approach have been raised by Turner, Wallis and Whitley (1988). The first concerns the simulation properties of our model[22] vis-à-vis those of the National Institute (NI), the London Business School (LBS) and HM Treasury (HMT). Both the Bank and NI equations model unemployment by taking as given employment and a measure of labour supply (such as the population of working age), which is treated as exogenous, whilst the LBS and HMT models include equations for the working population (thus modelling the participation rate) and calculate unemployment by residual. The different approaches result in models with very different properties - with both the Bank and NI models implying that, following a rise in government expenditure, unemployment falls by much less than is suggested by the LBS or HMT models (because of the larger "encouraged worker" effects in the former - ie as employment increases more people are encouraged to look for work).

In the Bank and NI models this effect arises primarily because of the small coefficients relating to non-manufacturing employment. Turner, Wallis and Whitley feel that these properties are "unduly pessimistic". Clearly this is largely an empirical question, however, which should be judged according to which models best explain past developments (and perhaps also which turn out to be best at forecasting). Unfortunately their attempts to evaluate the models fail to provide adequate information on which to make such a judgement. Instead they show how the HMT and LBS models of working population perform (in terms of explaining changes in working population, not unemployment) - which is fairly well, admittedly - thus giving some credence to these models. They also show that the NI unemployment equation appears to be poorly specified, since it has a number of insignificant coefficients and "fails tests for instrument validity, simultaneous equation bias and parameter stability", and dismiss the Bank model because they found that the equation is difficult to replicate if one ignores Special Employment Measures (by assuming that they have no effect).[23] These results are hardly surprising; dropping a number

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22 At the time that Turner, Wallis and Whitley did their research the Bank model equation was very similar in structure to the models presented in this paper (being closest to that shown in Table 2, column c).

23 They also mention the difficulty of knowing how to allow correctly for the school leaving age having changed, preferring to assume that it had no effect on labour supply, rather than allow the data to decide if there is evidence of any effects. In fact, however, only a tiny fraction of the discrepancy between their attempt at replicating our model and our results are due to the different ways that the change is treated.

of highly significant variables is almost bound to result in large changes to the remaining coefficients. We feel it is important, however, to emphasise that, unless one attempts to take on board these effects (and, in particular, the effects of Restart and more rigorous availability-for-work tests) one cannot hope to explain recent developments at all accurately. (Haskel and Jackman (1987) provide corroborative evidence that the Community Programme, in particular, has helped reduce long-term unemployment.)

To facilitate a better comparison of the various models used to explain changes in unemployment we have done two things. First, we have re-estimated the NI model, finding (as did Warwick) that the equation is poorly specified.[24] Thus it is hardly surprising that it explains the past rather poorly - with a standard error some 70% bigger than that relating to our model (36 compared to 22). Second, in order to make a fair comparison between our model and those used by the LBS and HMT we have calculated the standard error

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24 As a comparison between our equation with a levels equation the May 1988 vintage of the National Institute unemployment equation was re-estimated over the sample period 1971:4 to 1988:1 (see "National Institute Model 10.9" May 1988). The National Institute equation can be interpreted as an error-correction model, the dependent variable being the first difference of unemployment. The regressors enter both as difference and lagged level terms, plus the lagged level of unemployment. The regressors are manufacturing and non-manufacturing employment, a constant and a demographic labour supply variable (civilian labour force and population of working age were both tried). The equation was estimated both by OLS and IV (using the National Institute's suggested instruments). In addition, the SEMS effects were both included and excluded. The results tended to give coefficients of the wrong sign that were also insignificant. Indeed, the published version of the equation has a t-value of only 1.3 on the non-manufacturing employment category. The most important point, however, is that the standard error of the re-estimated National Institute equation, at around 36, is at least 70% larger than our preferred difference equation, even when allowance has been made for the effect of Special Employment Measures. The most recent version of the National Institute model drops the distinction between manufacturing and non-manufacturing employment growth in the long run. We believe this is wrong, as is easily illustrated by looking at the incidence of part-time working and distribution of male and female employees. Moreover, the tests performed above support our hypothesis. It is also clear that, whatever level of disaggregation of employment is used, it is not possible to reconcile all of the changes in unemployment without reference to SEMs and Restart effects.

of the current LBS equation in terms of unemployment (ie by substituting fitted values for the male and female working populations and actual values of employment into the identity linking working population with employment and unemployment in order to calculate fitted values for unemployment). This, at close to 90 over the period 1979-86, is much higher than that of our preferred model, so casting some doubt on Turner, Wallis and Whitleys' conclusions that the LBS and HMT approach is "preferable on economic and statistical grounds". It is especially clear that, as regards statistical grounds, the Bank model performs very well.

As for its economic basis, it certainly is the case that the Bank model does not identify a levels solution (as Turner, Wallis and Whitley make clear). Whilst this does not prevent the model from being used successfully in forecasting or simulation exercises, it will mean that the equation cannot be used to discuss concepts which require a solution in the level of unemployment (for example, it does not tell us what the "natural rate" of unemployment is, assuming that one believes in the usefulness of this type of concept).[25]

Despite our misgivings as to the usefulness of such an approach we decided to test the sensitivity of our previous results by including an error correction term in our model (defined in terms of the levels of unemployment, employment and the population of working age). Rather than impose coefficients in this term, we have followed the two-step Granger and Engle estimation procedure - whereby we first attempt to estimate a co-integrating vector (in terms of the levels of unemployment etc) and then use the residuals from this equation as the error correction term (denoted ERROR) in the full dynamic equation (which can then be estimated using the usual "general-to-specific" methodology).[26]

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25 The main problem, in practice, with these models is that the equilibrium unemployment rate appears to depend on the history of the actual unemployment rate. Layard and Nickell (1986), for example, present estimates which suggest that a large proportion of the rise in the actual male unemployment rate which took place between 1956-66 and 1980-83 (from 1.96% to 13.79%) occurred as a result of the natural rate rising (from 1.96% to 10.47%). These strong hysteresis effects suggest that a model with a unique equilibrium level of unemployment is unlikely to provide a realistic description of the UK economy. For a discussion of these issues see, for example, Blanchard and Summers (1986).

26 Hendry (1986) provides a useful overview of these techniques.

Thus we aim to estimate two models. Firstly a levels equation of the form;

$$ULU = b_0 + b_1 ULEMF + b_2 UMALES + b_3 UFMFTS + b_4 UFMPT + b_5 UOTHER + b_6 CUMRES + b_7 POWA \qquad (12)$$

and secondly the full dynamic model;

$$\Delta ULU = + a_1 \Delta POWA + a_2 \Delta ULEMF + a_3 \Delta ULEMF_{-1} + a_4 RES + a_5 \Delta UMALES + a_6 \Delta UFMFTS + a_7 \Delta UFMPT + a_8 \Delta UOTHER + a_9 \Delta ULU_{-1} + a_{10} ERROR \qquad (13)$$

$$\text{where } ERROR = \left[ULU - \hat{b}_0 - \hat{b}_1 ULEMF - \hat{b}_2 UMALE - \hat{b}_3 UFMFTS - \hat{b}_4 UFMPT - \hat{b}_5 UOTHER - \hat{b}_6 CUMRES - \hat{b}_7 POWA \right]_{-1}$$

where CUMRES is the stock of (cumulative) Restart interviews (so that ΔCUMRES=RES) and ^ refers to the estimated coefficients.

In searching for a co-integrating vector (12) we had difficulty in identifying separate roles for the various employment terms outside manufacturing. For this reason UMALE, UFMFTS, UFMPT and UOTHER were amalgamated into one (defined as ULOTHB). The resulting regression provides at best weak evidence of a "long-run" solution between unemployment, employment and the civilian labour force - with the Augmented Dickey-Fuller test not being particularly close to its critical value (see Table 5). A particularly interesting feature of our results, however, is the close correspondence between the parameters relating to the levels terms and those found earlier on the dynamic (growth) terms. Thus, the long-run coefficient on manufacturing employment is fairly close to one, that on non-manufacturing is "small" (at less than one quarter), whilst that on the population of working age is around one-half.

When we tested down to find a new dynamic equation (including ERROR as the error correction term) we again found it necessary to amalgamate some of the non-manufacturing employment terms together (see Table 5). The new equation has broadly the same characteristics as has our still preferred model (from Table 4), and a very similar sized standard error (of close to 21). Thus we have demonstrated how it is possible to provide a model which explains the past well, forecasts accurately and gives a long-run solution in the <u>level</u> of unemployment.

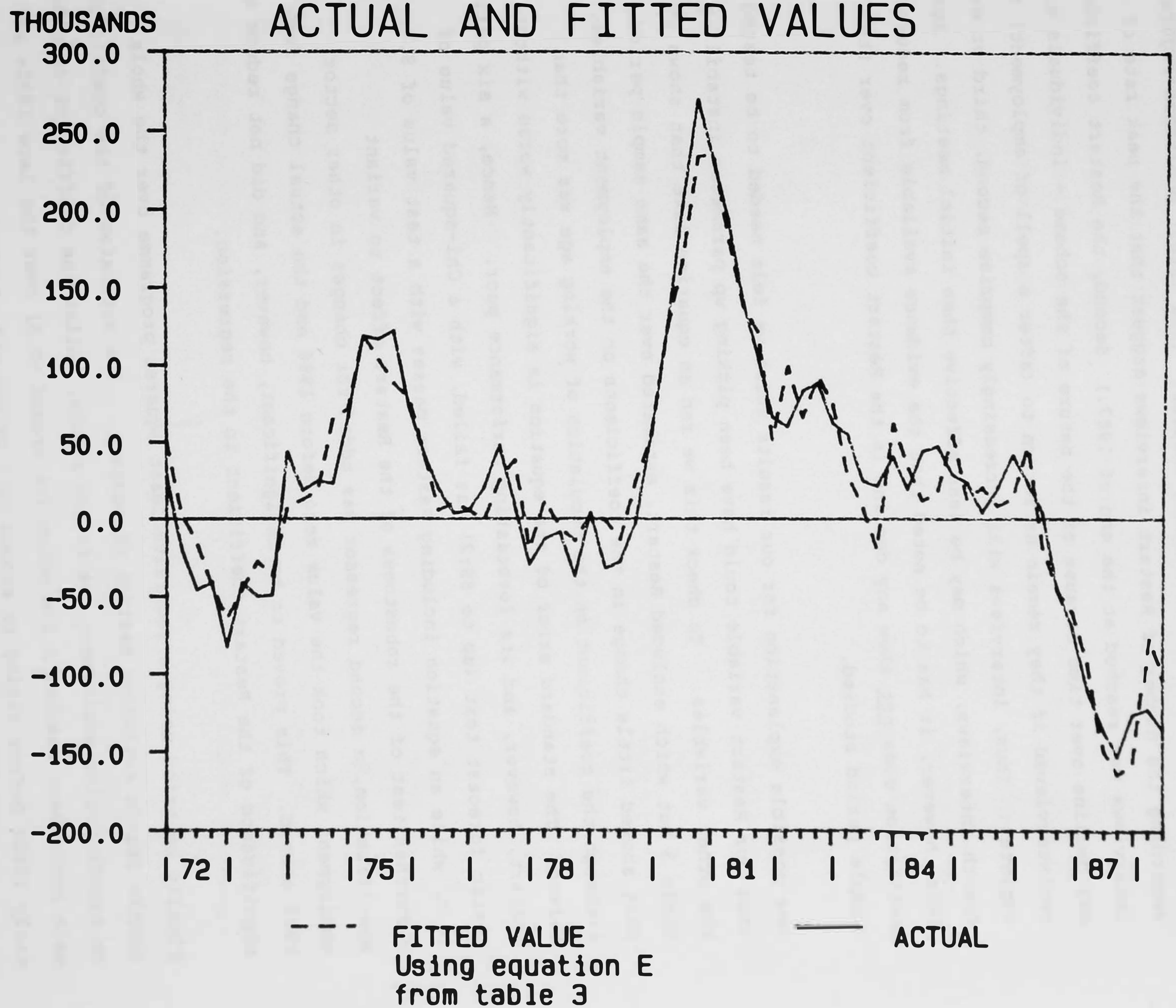
Clearly it might be possible to impose long-run coefficients other than those obtained from the first stage regression on the employment variables. For example, one might want to assume that there are no long-run effects from Restart. However, our results show that there are clear differences between the impact of increased employment on unemployment depending on the industrial mix of the increase in employment and also on the proportion of the gain that is due to full or part-time employment. Large Restart and SEMs effects are also still required if we are to track recent movements in unemployment at all accurately.

Section 5: The persistence of high unemployment as employment rose during the 1980s

Finally, we have used the preferred equation (from Table 3 rather than Table 5) to examine changes in unemployment over the last decade. Chart 7 shows the actual changes in unemployment between 1972:1 and 1988:3 together with the fitted values from our model. The model clearly tracks the data well throughout the period, even over the final nine quarters of the sample period when the number of registered unemployed fell strongly. As we would expect, much of the rise in unemployment in the early 1980s was due to the rapid decline in manufacturing employment. More recently employment in this sector has risen, though only very slowly, but this has been overshadowed to a great extent by growth in employment in other sectors of the economy. Although in numerical terms additional non-manufacturing employment has outweighed the recent rise in manufacturing employment it has not helped to reduce unemployment much because a large proportion of the growth has been in female and part-time employment. The effect from growth in employment in other sectors has only begun to counterbalance the effect of a long-term decline in manufacturing employment.

The sharper fall in unemployment than rise in employment in recent years has been due to the introduction of Restart interviews and stricter availability-for-work tests. Thus the Restart variable has since early 1986 contributed approximately 3/4 million to the fall in unemployment. In addition to the time series evidence on the effect of Restart (Appendix 3), figures published in the Employment Gazette show that, for the period May to October 1986, of the 523 thousand individuals then invited to attend an interview some 366 thousand attended and, as a result of the interviews, 86 thousand ceased to claim. Some of these people will have found jobs (which, of course, will be counted in our employment terms) while others will have entered Government schemes such as the Community Programme (and so again will be reflected in our SEMs figures). Nevertheless, a proportion of the individuals interviewed will have left the register, either because they were not entitled to claim benefits or because they have stopped looking for work. The number ceasing to claim expressed as a percentage of the number of interviews was 23%, a figure which is somewhat higher than our estimated long-run coefficient of 11-19% (see Tables 3 and 5). However, the estimated coefficient should be smaller than 23%, due to the fact that some individuals will be placed on schemes or in jobs following their interviews and thus captured elsewhere in the regression. Thus our results are not at all surprising.

CHART 7 : QUARTERLY CHANGES IN UNEMPLOYMENT, ACTUAL AND FITTED VALUES



Although the Restart programme has been very successful in helping to reduce registered unemployment there are several reasons to believe that its effects may soon begin to wane. First, the number of Restart interviews is likely to decline, mainly because of the fall in unemployment itself - the pool of potential interviewees is becoming smaller. (The latest published figures recording the number of Restart interviews suggest that the peak rate of interviews was reached at the end of 1987.) Second, the Restart coefficient may decline over time because of the nature of the scheme - individuals are re-interviewed if they remain or return to (after a spell of employment) the register. Thus, interviews will increasingly comprise second, third or even fourth interviews, which may be less effective than initial meetings. Against this, however, it has to be noted that the evidence available from recursive estimation does not show any decline in the Restart coefficient over the sample period studied.

One possible explanation for our results which we felt needed to be tested is that our Restart variable could have been picking up parameter instability in the other variables. To check this we ran an equation like that shown in Table 3 but which excluded Restart, estimated over the same sample period. This showed little change in the coefficients on the employment variables, although the coefficient on the population of working age was more than halved. The standard error of the equation is significantly worse without Restart, however, and its forecasting performance poor. Hence, a six period static forecast test (up to 88:2) was failed, with a Chi-squared value of 18.0, while an equation including Restart passes with a test value of 8.7. As a further test of the robustness of the Restart effect to variant specification, a second regressor was added for changes in other sector employment which took the value zero before 1986 and the actual change from 1986 onward. This proved to be insignificant, however, and did not reduce the significance of the Restart coefficient in the regression.

Finally we tried using a recursive least squares programme over the whole sample (again excluding Restart of course). The estimates of the coefficient on manufacturing employment are fairly stable, whilst the coefficient on other male employment reaches a low point (of around -0.5) over the late 1970s and early 1980s before rising to around -0.1 by the end of the sample. These results are sensitive to the size of the window chosen, however, the results presented being based on a window comprising 25 observations. The slight deterioration in stability of the coefficients witnessed towards the end of the sample provides some (limited) evidence of parameter instability (indeed the labour supply coefficient becomes negative towards the end of the sample). However, the failure of the model to explain changes in unemployment in recent years provides additional support for the Restart variable, leading us to conclude that, without such a variable, one cannot fully explain the fall in unemployment.

Section 6: Conclusions

In this paper we have tried to improve our understanding of the changes in the relationship between unemployment and employment over the past two decades. In doing so we have shown that it is important to distinguish between employment by industrial sector and by sex and between full and part-time female workers. Our results suggest that, although unemployment is falling because there are more jobs, it is also true that much of the decline in the claimant count which has occurred since mid-1986 has been due to a shift in the unemployment/employment relationship resulting from changes in the Government's range of Special Employment Measures - especially the introduction of more rigorous availability-for-work tests and the rapid growth in the Restart programme.

Appendix 1: Test Statistics

We use the following tests;

$\eta_1(.)$ Modified Lagrange-multiplier statistic for testing against residual autocorrelation (see Harvey (1981)).

$\eta_2(.)$ Wald statistic for testing against the relevant unrestricted maintained model (see Harvey (1981)).

$\eta_3(.)$ Chow's statistic for testing parameter constancy (see Chow (1960)).

$\eta_4(.)$ White's statistic for testing against residual heteroscedasticity (see White (1980)).

$\eta_4^*(.)$ Engle's ARCH statistic for testing against residual heteroscedasticity (see Engle (1982)).

$\epsilon_1(.)$ Lagrange multiplier statistic for testing against residual autocorrelation (see Harvey (1981)).

$\epsilon_1^*(.)$ Ljung-Box statistic for testing against residual autocorrelation (see Ljung and Box (1978)).

$\epsilon_3(.)$ Hendry's static 'forecast' statistic for testing parameter constancy (see Davidson, Hendry, Srba, and Yeo (1978)).

$\epsilon_3^*(.)$ Hendry's dynamic 'forecast' statistic for testing parameter constancy. This is equivalent to $\epsilon_3(.)$ except that when calculating the forecast predicted values (rather than actuals) are used for all lagged dependent variables.

$\epsilon_4(.)$ Breusch and Pagan's statistic for testing against residual heteroscedasticity (see Breusch and Pagan (1979)).

$\epsilon_4^*(.)$ Engle's ARCH statistic for testing against residual heteroscedasticity (see Engle (1982)).

$\epsilon_5(.)$ Jarque and Bera's statistic for testing against non-normality in the residuals (based on skewness and excess kurtosis) (see Jarque and Bera (1980)).

All the $\eta(.)$ tests are F-tests, whilst the $\epsilon(.)$ tests have chi-squared distributions.

Ψ_7, Ψ_8 and Ψ_9 are tests for normality. Ψ_7 is the Shapiro-Wilk statistic (see Maddala (1979)), whilst Ψ_8 and Ψ_9 are tests for skewness and kurtosis (see Kiefer and Salmon (1982)).

Appendix 2: Estimates of SEMs Effects

Until 1986 the Department of Employment published monthly estimates of the effects of special employment measures on registered unemployment. Keating (1986) regressed changes in the published registered effect upon changes in the numbers supported by each of the various schemes. In this way he hoped to retrieve the formulae by which the Department of Employment made its estimates.[27] Some of the schemes were estimated to have little effect. However, five schemes - the Youth Opportunity Programme, Community Programme, Youth Training Scheme, Young Workers Scheme and Temporary Short Time Working Compensation Scheme - were found to have contributed significantly to the total effect attributed to SEMs. Although the total estimated effectiveness of the schemes ceased to be published after early 1986, because the numbers supported by the measures are still available,[28] then the coefficients representing the individual register effects of these schemes can be applied to these numbers to obtain estimates of the total register effects. There are, however, a number of additional problems;

1 The published total register effect includes the effect on youth unemployment. The unemployment regressions described in this paper use seasonally adjusted unemployment excluding school leavers.[29]

2 Since 1986 there have been a number of new schemes, for example, the New Workers Scheme, Jobstart etc. We have no direct way of estimating the register effectiveness of these schemes.

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27 In fact, this is a rather grand way to describe what is a very informal process. The Department of Employment do use coefficients for each scheme with many of these having been published in the Public Expenditure White Paper for 1985/86. The coefficients are based on evidence from administrative data, postal monitoring surveys and interview surveys with employers, participants, etc. This process has been described at various times for different schemes in articles in the Employment Gazette (eg an article on the Young Workers Scheme appeared in May 1986). The Department of Employment stopped publishing count effects in 1986, following a decision to cease estimating count effects for YTS which was regarded as a training programme rather than an employment measure.

28 Published monthly in the Employment Gazette Tables 9.1 and 9.2.

29 The unemployment count was redefined in October 1988 to exclude those aged under 18.

3 The register effectiveness of any one scheme is likely to vary over time. A simple OLS regression can not take account of this. However, a Kalman filter based approach can be used to update the register effectiveness coefficients over the period to early 1986 for which the Department of Employment estimate of effectiveness has been published.

4 With respect to individual schemes, there are possible problems concerning the YTS and Community Programme coefficients estimated by Keating. The Youth Training Scheme has several effects. In many cases a YTS participant will fill a newly created job. In some cases, however, the participant will be filling a job that would have existed anyway. In either case the effect on the adult unemployment count is likely to be much smaller than the 60% estimated by Keating and with the change in regulations concerning under 18 year olds last Autumn, the effect is likely to become even smaller. On the other hand, individuals participating in the Community Programme are likely to have almost a one-to-one effect on the unemployment register, higher than suggested by Keating's estimates (although one also needs to take into account the fact that the count effect includes supervisors and participants from other groups who are given waivers to exempt them from benefit regulations - these may reduce the count effect to a little below one).

5 Due to the size of the YTS and the Community Programme it is clear that these two schemes form the bulk of the SEMs effect on unemployment. In the period since 1986, however, we believe that Restart and the activities of claimant advisors have also had large effects. Neither Restart nor claimant advisors can be assessed using the above method and, therefore, the number of Restart interviews enters directly into the estimated employment/unemployment relationship.

6 It is possible to obtain a reasonably accurate time series for the total effect of SEMs on unemployment by using estimates of the individual register effectiveness of schemes together with the numbers actually participating. It is, however, also necessary to obtain a measure of the effect of SEMs on employment. This is more difficult since no published figures exist to guide us and so we must, therefore, apply sensible estimates. However, occasional articles in the Employment Gazette outline the economic effects of various schemes (the YTS scheme was assessed in October 1987 and the August 1988 issue contains a useful discussion of recent changes to employment statistics) and it is possible to use these results to estimate the effect on employment.

These studies show that the majority of SEMs places are in the non-manufacturing sector. Thus, we make an assumption as to the proportion of total SEMs employment effects which fall on the non-manufacturing sector.

Appendix 3: Time Series Estimates of the Effect of Restart

As well as using traditional regression analysis we have tried to assess the underlying path of unemployment by using a Kalman filter model on monthly data in order to take into account Restart interviews. The work suggests that falls in underlying unemployment during 1987 did not exceed 15,000 per month, which compares with actual falls of up to 50,000. The implied Restart effect was estimated at between 18 and 20%.

The technique involves regressing the change in unemployment on a linear trend and the number of Restart interviews, where both the trend and the effect of Restart are permitted to vary over time.[30] This allows us to ascertain whether the Restart coefficient has tended to fall over time, although no such evidence was found for the period under consideration (January 1981 to February 1988).

To test whether the Restart variable is proxying the activity of claimant advisors (as well as of Restart interviews themselves), we also tried estimating a trend for short-term unemployment with the number of Restart interviews included as a regressor. Again we found evidence of a significant role for Restart. Since it is only the long-term unemployed who are interviewed under the scheme the significance of the Restart variable in explaining short-term unemployment provides evidence that some other factor has been helping to reduce the count. The most obvious candidate is more rigorous availability-for-work tests. Hence, throughout our work we stress that our Restart variable (RES) may be picking up the effects of both the Restart programme and availability-for-work tests.

30 Details of the technique are given in West, Harrison and Pole (1988).

Appendix 4: Data Sources

All data are seasonally adjusted, measured in thousands and taken from the Department of Employment Gazette, unless otherwise stated.

LU	Unemployment (UK).
LEMF	Manufacturing employment (UK). An adjustment is made to GB figures in order to take into account employment in Northern Ireland.
LEG	Employment in the non-trading general government sector (including those in the Armed Forces). Source: Central Statistical Office.
LHMF	Number employed in the Armed Forces (UK, two quarter moving average).
LE	Employees in employment (UK, two quarter moving average).
LOTH	Employment in the non-manufacturing sector. Defined as LOTH = LE-LEG-LEMF+LHMF.
POWA	Population of working age (GB). Population of males aged between 16 and 64 and females aged between 16 and 59. Annual observations interpolated to obtain quarterly estimates. Source: Office of Population Censuses and Surveys.
CLF	Civilian Labour Force (GB). Civilians aged 16 and over who are either in paid work or actively seeking it. Annual observations interpolated to obtain quarterly estimates. (Note that the figures have been adjusted to allow for the change from the former labour force measure to the ILO/OECD measure. See Employment Gazette - April 1989.)
RES	Number of Restart interviews conducted within the period (GB).

UMALE Number of male employees in the non-manufacturing sector (UK). Defined using published figures relating to total number of male employees (UK) minus those in manufacturing, where the latter is calculated by subtracting the number of female employees in manufacturing (GB) from LEMF.

UFMPT Number of part-time female employees in the non-manufacturing sector (GB).

Defined by subtracting the number of female part-timers in manufacturing from the total number of female part-timers (GB).

UFMFT Number of female full-time employees in the non-manufacturing sector.

Defined by subtracting UFMPT from the number of female employees in non-manufacturing (the latter having been calculated as the total number of female employees (UK) minus those in manufacturing (GB)).

UOTHER Other non-manufacturing employment (primarily self-employed and Armed Forces).

Defined as UOTHER = (LOTH+LEG+LSE)-UMALE-UFEMPT-UFEMFT.

LSE Self-employed (UK).

SEMSLU Effect of special employment measures on the adult unemployment count (see Appendix 2).

SEMSLEMF Effect of special employment measures on manufacturing employment (see Appendix 2).

SEMSOTH Effect of special employment measures on non-manufacturing employment (see Appendix 2).

NOTE 1 Data on female part-time employment in Great Britain is only available quarterly since June 1978. Before 1978 we rely on annual observations from the census of employment interpolated to obtain quarterly data.

2 It was decided not to make arbitrary assumptions about composition of employment in Northern Ireland. Therefore some of the disaggregated employment categories contain elements of employment relating to Northern Ireland. These are not thought to significantly affect our results.

Table 1: Tests on Models based on Equation (7)

Column A refers to the old Bank model (coefficients imposed), whilst Column B is a freely estimated version of the same model.

		A	B
<u>Autocorrelation</u>			
η_1 (4,86)	=	12.6 ϕ	23.5 ϕ
ϵ_1 (4)	=	34.7 ϕ	49.1 ϕ
ϵ_1^* (4)	=	51.2 ϕ	88.3 ϕ
<u>Heteroscedasticity</u>			
ϵ_4 (1)	=	10.3 ϕ	14.3 ϕ
<u>Normality</u>			
Ψ_7	=	0.99	0.96
Ψ_8	=	-0.21	-0.88 ϕ
Ψ_9	=	3.24	5.50 ϕ
<u>Parameter Stability</u>[a]			
ϵ_3(6)	=	7.0	2.8
ϵ_3(12)	=	23.8 ϕ	86.4 ϕ

Details of the tests used are given in Appendix 1.

ϕ implies significant at the 95% level.

(a) The six-period forecast tests are based on the period 1985:1 - 1986:2.

Table 2: Some Models based on Equation (8)

Dependent variable; ΔULU

Explanatory Variables	A Coef	(t-value)	B Coef	(t-value)	C Coef	(t-value)
ΔULU_{-1}	-		0.36	(5.7)	0.53	(6.4)
ΔPOWA	0.50	(4.8)	0.24	(2.4)	-	
ΔPOWAA	-		-		0.18	(2.9)
ΔULEMF	-0.87	(18.4)	-0.62	(10.1)	-0.75	(10.7)
$\Delta ULEMF_{-1}$	-		-		0.30	(2.9)
ΔULOTHA	-0.24	(5.0)	-0.13	(3.0)	-	
ΔULOTHB	-		-		-0.08	(2.5)
ΔLEG	-0.20	(2.2)	-0.12	(1.6)	-	
RES	-0.18	(7.1)	-0.13	(5.9)	-0.11	(4.8)
DUMMY	336	(4.0)	134	(1.7)	-	
Sample Period	1964:3 - 1987:4		1964:3 - 1987:4		1965:1 - 1987:4	
R^2	0.88		0.92		0.93	
SEE	26.4		22.6		20.7	
DW	0.84		1.28		1.87	

where $POWAA = \frac{1}{8} \sum_{i=0}^{7} POWA_{-i}$

All other variables are as defined in the text. Data sources are given in Appendix 4.

Tests;

Autocorrelation			
η_1 (4,84)	= 11.2 ϕ	4.6 ϕ	0.6
ϵ_1 (4)	= 32.7 ϕ	17.2 ϕ	2.5
ϵ_1^* (4)	= 47.5 ϕ	17.3 ϕ	1.6
Heteroscedasticity			
ϵ_4 (1)	= 0.1	0.0	0.3
Normality			
Ψ_7	= 0.98	0.97	0.97
Ψ_8	= -0.43	0.46	0.54 ϕ
Ψ_9	= 4.66	3.04	3.21
Parameter Stability[(a)]			
ϵ_3 (6)	= 2.1	2.2	2.5
ϵ_3^* (6)	= -	2.3	1.2
ϵ_3 (12)	= 59.0 ϕ	47.4 ϕ	42.6 ϕ

Details of the tests used are given in Appendix 1

ϕ implies significant at the 95% level

(a) The six-period forecast tests are based on the period 1985:1 - 1986:2.

Table 3: Some Models based on equation (11)

Dependent variable; ΔULU

Explanatory Variables	D		E	
ΔPOWA	0.30	(1.7)	0.22	(1.3)
ΔULEMF	-0.90	(-11.9)	-0.91	(-12.0)
RES	-0.17	(-4.8)	-0.18	(-5.4)
ΔUMALES	-0.26	(-3.1)	-0.23	(-3.3)
ΔUFMFTS	-0.12	(-1.6)	-	-
ΔUFMPT	-0.07	(-0.7)	-	-
ΔUOTHER	-0.05	(-0.6)	-	-
$\hat{a}_o$	0.58	(5.5)	0.61	(6.1)
Sample Period	1972:1 - 1988:3		1972:1 - 1988:3	
SEE	22.2		22.09	
$\bar{R}^2$	0.77		0.76	
DW	1.83		1.92	

<u>Tests</u>:

<u>Autocorrelation</u>

$\epsilon_1(4)$	= 2.67	2.42

<u>Heteroscedasticity</u>

$\epsilon_4^*(1)$	= 0.02	0.01

<u>Normality</u>

$\epsilon_5(1)$	= 1.84	2.00
Ψ_8	= -0.11	-0.08
Ψ_9	= 2.22	2.17

Table 4: A model based on Model C of Table 2 using a shorter sample period

Dependent variable; ΔULU

Explanatory variables	F	
ΔULU_{-1}	0.57	(5.4)
$\Delta POWA$	0.13	(1.2)
$\Delta ULEMF$	-0.82	(-9.4)
$\Delta ULEMF_{-1}$	0.39	(3.0)
RES	-0.08	(-3.6)
$\Delta ULOTHB$	-0.08	(-1.7)
Sample Period	1972:1 - 1988:3	
SEE	22.2	
$\bar{R}^2$	0.93	
DW	1.96	

Tests

Autocorrelation

ϵ_1 (4) = 2.52

Heteroscedasticity

ϵ_4^* (1) = 2.93

Normality

ϵ_5 (1) = 2.68

Ψ_8 = 0.40

Ψ_9 = 2.42

Table 5: A Model based on (12) and (13)

Dependent variable; ULU

Explanatory variables		
ULEMF	- 0.86	(16.1)
ULOTHB	- 0.21	(-3.8)
POWA	0.47	(4.8)
CUMRES	- 0.11	(5.1)
C	-3985.0	(-1.4)

Sample period	1972:1 - 1988:3
SEE	91.2
$\bar{R}^2$	0.99
ADF	2.3

Dependent variable; ΔULU

Explanatory variables	G	
ΔULU_{-1}	0.60	(5.6)
ΔPOWA	0.05	(0.4)
ΔULEMF	-0.77	(8.6)
$\Delta ULEMF_{-1}$	0.37	(2.9)
RES	-0.08	(3.3)
ΔUFMFTS } ΔUFMPTS } ΔUOTHER }	-0.001	(0.02)
ΔUMALE	-0.14	(1.7)
ERROR	-0.06	(1.7)

Sample period	1972:2 - 1988:3
SEE	21.4
$\bar{R}^2$	0.94
DW	1.92

Tests

Autocorrelation

$\epsilon_1(4) = 2.89$

Heteroscedasticity

$\epsilon_4^*(1) = 2.93$

Normality

$\epsilon_5(1) = 2.45$

$\psi_8 = 0.33$

$\psi_9 = 2.32$

NOTE: Data used to estimate this model include the latest employment estimates (based on the 1988 Labour Force Survey).

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