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Participation in Further Education in England and Wales:
An Analysis of Post-War Trends

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Abstract

The paper examines the time-series evidence relating to participation rates in further education in England and Wales, and uses cointegration analysis to identify a long-run statistical relationship in the data consistent with an augmented human-capital model. The recent rapid growth of participation is attributable largely to the improvements in GCSE attainment of the last decade, coupled with the expansion of higher education. Fluctuations in labour demand play a significant role in determining movements in participation rates over time, and the substantial rise in youth unemployment of the early notes was a contributed to the rapid growth of participation at this time.

JEL Classification: I21, J24
Keywords: human capital, participation in further education, cointegration analysis.

1. Introduction.

There is a long-standing perception that shortcomings in the education system lie at the heart of the relatively poor performance of the British economy in recent decades. Historically, rates of participation in education and training following compulsory schooling in Great Britain have lagged far behind those in other OECD economies, resulting in a workforce with few skills and qualifications in comparison with that of other major industrial economies (Layard et al, 1995; O.E.C.D., 1995; Prais, 1995). Recently, there have been signs that Great Britain may be closing the skills gap; participation rates in full-time further education following compulsory schooling have increased substantially as Fig. 1 shows.[1] The purpose of this paper is to analyse this increase, and more specifically, to establish the extent to which it represents an increase in the underlying trend, rather than short-run adjustment to changing market conditions.

(Fig. 1 approx here)

Alternative explanations have been offered for the rapid growth of further education of recent years. Firstly, it has been asserted that attitudes to further education and training among young people in Great Britain have changed fundamentally as a consequence in part of recent education policies.[2] Measures such as the introduction of a national curriculum and a unified system of secondary school qualifications are held responsible for the marked improvement in average attainment levels among those completing compulsory schooling (Ashford et al, 1993; Gray et al, 1993). At the same time, reforms within the further education sector have broadened the curriculum to include more vocational qualifications and foundation courses, lessening the perception that further education is an option reserved for an academic elite.

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[1] Throughout the paper, the term further education is used to refer to education or training undertaken following compulsory schooling with a specialised educational establishment, either a school or college.

[2] To quote James Paice, then Minister of State for Education and Employment: "Over the last fifteen years, a dramatic revolution has taken place in the attitudes of young people....(i)t has become the norm for young people to continue their learning beyond the age of 16 years". (DFEE News, 25-1-1996.)

Secondly, the period has seen substantial changes in the labour market for young persons also. Since the mid-1970s, the trend in unemployment among young persons has risen steadily, punctuated by severe economic recessions exacerbating the lack of employment opportunities for young people. With a low probability of finding employment in the short-run, the opportunity costs of remaining in full-time education beyond age 16 years are relatively small. All the more so, following changes in benefit regulations in 1988 which effectively ended the payments to unemployed persons below the age of 18 years. The impact of high unemployment rates may have been offset to some degree by the expansion of government-sponsored training schemes designed to provide vocational training in the labour market for 16 and 17 year olds. However, many have argued that the paucity of job opportunities for young people has contributed to the rapid growth in participation in further education over recent years.

This paper uses post-war time series evidence to examine in detail the role played by economic factors in the growth of participation rates in further education. To date, there have been relatively few studies based on time-series data, although there is an extensive literature examining individual education choices using cross-section data (e.g. Rice, 1987,1999; Micklewright, 1989; Andrews and Bradley, 1997). Pissarides (1981), analysing time series data for England and Wales for the period 1955 to 1978, concludes that the growth of participation rates during the 1960s and the subsequent slowdown in the 1970s were the result of movements in the relative earnings of qualified workers, coupled with changes in real household income. In their later work, Whitfield and Wilson (1991) argue that the Pissarides' model lacks robustness when applied to more recent data. Their study of the period to 1985 identifies the rate of return to education, social class structure, the unemployment rate and the scale of youth training provision as the main determinants of the rate of participation in further education over the long-run.

The present paper advances this earlier work in a number of ways. First, it provides a more robust analysis of the long-run statistical relationships in the data by using full information methods rather than single equation techniques as in Whitfield and Wilson. If there is at most one cointegrating vector among the variables of interest then OLS estimation of the single static regression yields consistent parameter estimates. However, in this particular context, several of the variables of interest are potentially endogenous and there may be a multiplicity of cointegration vectors in which case single equation methods are no longer valid.[3] We use the full-information maximum likelihood method of Johansen (1991) that tests for the number of cointegrating vectors and, conditional on the correct number of cointegrating vectors, produces asymptotically optimal estimates. An additional advantage of this approach is that inclusion of short-run dynamics improves the small-sample properties of the estimates of the cointegrating vectors (Campbell and Perron, 1991, p.187).

In addition, this study extends the sample period to include the years of exceptionally rapid growth in participation rates in further education between 1988 and 1994. This allows us to address two questions concerning this recent experience. First, to what extent does it represent an increase in the long-run rate of participation, as distinct from short-run dynamic adjustment? Secondly, what are the relative contributions of social and of economic factors to the growth in participation, and hence to what extent can the higher rates of participation be expected to persist over the economic cycle?

Our results show that the rapid increase in rates of participation over this period were driven largely by improvements in the level of attainment at GCSE coupled with the expansion of the higher education sector. In addition, fluctuations in labour demand play a significant role in determining the movements of participation rates over time, and the sharp rise in unemployment in the early 1990s contributed to the growth of participation. For much of the 1990s, the actual rate of participation

[3] Endogeneity of the regressors in the static regression means that the asymptotic distribution of the OLS estimates depends on nuisance parameters. For further discussion, see Campbell and Perron, 1991, Hendry 1995.

appears to have exceeded its long-run level. Given this, and the relative stability of key determinants such as the level of GCSE attainment and the size of the higher education sector, participation rates in further education are unlikely to increase much beyond their present levels over the medium term.

In the next section of the paper, we discuss the key factors determining the rate of participation in further education over the long run and describe the variables used in the empirical analysis. Section 3 discusses the results of the maximum likelihood estimation of the VAR and the implied long-run relationships among the variables of interest. The short-run dynamics of the relationship between the variables are examined in section 4. In conclusion, we consider what light our findings can shed on the rapid expansion of further education over the last decade and the implications for future trends.

2. A Model of the Participation Rate in Further Education.

Human capital theory provides a framework for the statistical analysis of participation rates in further education. The individual's decision to invest in further education is assumed to be based on a comparison of the expected returns to an additional period of full-time education with the expected returns to entering the labour market at a given age. Human capital theory identifies a wide range of factors that may be expected to affect the net return to investment in further education for an individual (e.g. see Freeman, 1986). Broadly speaking, these fall into two categories. The first consists of personal attributes related to the individual's skills and preferences. The second category consists of market variables which reflect the conditions, economic and social, faced by all members of a given cohort. Thus, in general terms, the expected net return to further education for individual i at date t may be considered as a function of the following form;

$$B_{it} = B(\mathbf{z}_{it}, \mathbf{w}_t) \tag{1}$$

where $\mathbf{z}_{it}$ is a vector of observable personal attributes of the ith individual at date t and $\mathbf{w}_t$ is a vector of factors at date t. It is assumed that the individual chooses to undertake further education if

B_{it} is non-negative. Aggregating, the demand for further education by a given age cohort of the population is determined in the long-run by a function of the form

$$P_t = F(\mathbf{Z}_t, \mathbf{W}_t) \tag{2}$$

where P_t denotes the proportion of the age cohort seeking to undertake full-time further education and $\mathbf{Z}_t$ $\mathbf{W}_t$ are the average values of the personal attributes and market factors for the age cohort.

The empirical analysis assumes that the rate of participation in further education in England and Wales is demand-determined throughout the sample period of 1954 to 1994. The 1944 Education Act established the right of all pupils to proceed to further education on completion of compulsory schooling if they so wished, and placed a duty on local education authorities to provide adequate facilities for further education (The Education Act 1944, clause 39). It is recognised that supply constraints may have been binding in certain areas, but the consensus is that for the further education sector as a whole, supply may be treated as perfectly elastic.

Our interest is in participation rates in full-time education among those who have completed their compulsory schooling. Legislation in 1972 raised the minimum age for completion of compulsory schooling from 15 to 16 years, and so care is required in defining the relevant age cohort for analysis. For present purposes, the participation rate in further education is defined as the proportion of the cohort aged 16 years at the start of an academic year (i.e. September 1st), in full-time education in the following January. Under the terms of the 1972 Education Act, all members of an age cohort so defined completed their compulsory schooling in the previous academic year. Prior to the 1972 Act, members of the age cohort completed their compulsory schooling one year earlier at age 15 years. In the statistical analysis, the possibility of a regime shift arising from this discrete change in the number of years of compulsory education is considered.

The choice of explanatory variables for empirical analysis is governed by two sets of considerations. The first is to specify a model that captures both the economic and social factors that have influenced participation in further education over time. The second set of considerations is essentially statistical. The sample consists of some 40 annual observations, and hence there is limited scope for obtaining reliable estimates of a large number of independent effects.

For an individual, the expected return to an investment in further education is an increasing function of their level of academic skills and abilities. For the cohort, a measure of average academic attainment may be derived from the distribution of grades achieved in the GCSE qualifications taken at the end of compulsory schooling. We use the proportion of the cohort achieving a minimum of five GCSE qualifications at higher grade (or the equivalent prior to 1988). The last decade has seen a marked improvement in GCSE attainment levels and this is widely believed to have contributed to the rapid growth in participation rates over this period (e.g. Ashford et al, 1993; Payne, 1998). This measure of academic attainment is potentially endogenous, depending as it does on past investments in human capital, as well as innate ability. Moreover, increases in the expected return to further education may serve as an incentive to expend greater effort on improving GCSE grades, as well as increasing the likelihood of participation.

Two further individual attributes are included in the analysis – family income and social class. The rate of participation in further education is expected to increase with average real household income for two reasons. First, further education is regarded as a normal good in consumption. Second, where capital markets are imperfect, family income provides an important source of finance for educational investments. The effects of social class are more nebulous, but are generally ascribed to the influence of socio-economic background on an individual's tastes and preferences. Cross-section studies have found that, other things being equal, individuals whose parents belong to the managerial, professional and related occupations are significantly more likely to undertake further education than individuals from other socio-economic backgrounds (Gray et al, 1993; Andrews and Bradley, 1997; Rice, 1999). Whitfield and Wilson found evidence in the time-series data of a strong positive relationship between the proportion of the workforce in white-collar occupation and participation rates over the long run. Here, we endeavour to focus more directly on the socio-economic composition of households by considering the proportion of household where the head of household belongs to a managerial, professional or related occupation.

Individual attributes aside, the expected return to investment in further education depends largely on the relative wages of different skill categories in the labour market. A first best measure would be based on a comparison of the present discounted value of lifetime earnings in occupations requiring further education qualifications with that offered in alternative occupations, but data limitations render this impractical. An alternative approach adopted by Pissarides (1981) is to include measures of the relative earnings of qualified and unskilled employees at different stages of the life cycle. However, given the small sample size available, this approach poses practical difficulties also. Instead, we employ a simple proxy in the form of the ratio of the average earnings of those employed in managerial, professional and related occupations relative to the average earnings of manual workers.

Demand conditions in the labour market affect the opportunity costs to the individual of a further period of full-time education. If job opportunities for 16 year-olds are limited then, all other things being equal, the opportunity costs of further education are relatively low and this is expected to be reflected in a higher rate of participation. The high proportion of new entrants means fluctuations in demand tend to have more pronounced effects on unemployment among young persons, than in the labour force as a whole. For this reason, the unemployment rate among those aged 18 to 20 years, rather than the overall unemployment rate, is used as an indicator of the state of demand. Until the early 1970s, this series remained relatively flat with unemployment rates averaging around 2% for both males and females. Since 1975, both the underlying trend in the unemployment series, and the magnitude of the fluctuations about this trend have increased sharply, with rates exceeding 20% between 1981 and 1987, and again between 1991 and 1994.

It is possible that the effects of the high unemployment rates since 1975 have been offset by the expansion of subsidised youth training schemes. First introduced in late 1970s, the measures were intended initially to ameliorate the problems of high unemployment among school-leavers by offering a short period of work experience. The 1980s saw a considerable expansion of these schemes and a shift in emphasis to providing work-based training. Raffe (1988) argues that school leavers regard these schemes as a last resort in the event of failure to find employment and not as a substitute for education or employment. If this assessment is correct then their impact on the decision to undertake further education is expected to be negligible. However, the results of Whitfield and Wilson appear to challenge this view, indicating that the expansion of these training schemes led to a decline in the rate of participation in further education over the long run.

Whitfield and Wilson use the proportion of those aged 16 years enrolled on Youth Training or related schemes as a measure of scale of youth training provision in a given year. This approach assumes that the supply of training places is inelastic in the short-run. This assumption is questionable, particularly for the period since 1986 when, in principle at least, all 16 years olds have been guaranteed a two-year training placement. An alternative approach is to allow for the effects of the training schemes through step dummy variables, on the assumption that the provision of subsidised work-based training produces discrete shifts in the opportunity costs of further education. Both approaches are investigated in the econometric analysis that follows.

Finally, a study of these issues should not ignore the considerable changes that have occurred in the higher education sector in Great Britain over the last decade, with the number of entrants as a proportion of the relevant age cohort more than doubling. There are a number of reasons for expecting the expansion of higher education to feedback into the decision of young persons to enter further education. First, it has led to lower entry requirements, thereby increasing the probability of entry to higher education for individuals who complete further education. Added to this, it may have induced significant 'role model' effects across successive cohorts, increasing the value placed on the benefits to education beyond compulsory schooling. We try to capture these effects by including a measure of higher education provision defined as the number of initial entrants to higher education as a proportion of the age cohort of 18/19 year olds in a given year.

While an augmented human capital model yields strong priors regarding the determinants of participation rates in further education over the long run, it provides little information regarding the short-run dynamics of such a relationship. A number of studies of individual behaviour have indicated the presence of significant role model effects across successive cohorts with higher rates of participation in further education among the older cohort increasing the likelihood of their younger colleagues choosing to remain in full-time education (e.g. Cheng, 1995; Feinstein and Symons, 1999). This suggests the presence of significant positive autocorrelation in the participation rate variable in the short-run. In addition, imperfect information may be expected to introduce lags into the relationship between participation rate and the measures of labour market conditions. The nature of the short-run dynamics of the relationship is explored in Section 4.

3. Estimation of the VAR: The Properties of the Long-run Relationship.

The system to be analysed consists of up to eight stochastic variables as follows: the rate of participation in full-time education for the cohort aged 16 years (P); average academic attainment (Q); social class structure (S); average real household income (I); the relative earnings of qualified workers (W); the unemployment rate for younger workers (U); the provision of youth training (YT); the provision of higher education (HE). Full definitions together with details of the compilation of the data series are provided in the Appendix. A number of the variables are defined as proportions of the relevant population, and in these cases log transformations are taken to avoid problems associated with bounded variables. The individual time-series for the levels of the variables exhibit strong trends over the sample period, and augmented Dickey-Fuller tests applied recursively fail to reject the null hypothesis of a unit root against the alternative of a trend stationary series in all cases. Taking first differences, the null hypothesis of a unit root is rejected at the 90% level for all variables, with the possible exception of the HE variable.

Proceeding on the basis that the stochastic variables are I (1), we consider the vector error-correction form of the VAR i.e.

$$\Delta \mathbf{y}_t = \mathbf{a}_0 + \mathbf{a}_1 t + \mathbf{A}\mathbf{d}_t - \Pi \mathbf{y}_{t-1} + \Gamma \Delta \mathbf{y}_{t-1} + \mathbf{u}_t \tag{3}$$

$\mathbf{y}_t$ denotes the vector of stochastic variables and $\Delta \mathbf{y}_t$, the vector of their first differences. In addition to the stochastic variables, the VAR includes a time trend, an intercept term and a vector $\mathbf{d}_t$ of dummy variables to allow for possible intercept shifts due to policy changes. $\mathbf{d}_t$ includes a step dummy CS_t that takes the value 1 after 1973/74 to allow for the effects of the increase in the number of years of compulsory schooling following the 1972 Education Act. Where the measure of youth training provision is not included in the set of stochastic variables $\mathbf{d}_t$ includes also step dummies to allow for the effects of youth training schemes. We find that the two approaches to incorporating the effects of youth training lead to similar conclusions regarding the relationships between the other stochastic variables in the system. In what follows, we focus on the results obtained with YT included among the set of stochastic variables, in part because these tend to be somewhat better determined, but also for comparability with earlier studies.

Johansen's maximum likelihood procedures are used to estimate (3) and test for cointegration among the set of stochastic variables (Johansen 1991).[4] For these purposes, the time trend is restricted to lie in the cointegrating space, while the intercept term and dummy variables are left unrestricted. Initially, all the stochastic variables in the VAR are treated as jointly endogenous. However, *a priori*, variables such as real household income, social class structure and relative wages

[4] The estimation was done using the cointegration analysis procedures in Microfit 4 (Pesaran and Pesaran, 1997).

are believed to play primarily an explanatory role, rather than to be determined by the other variables in the system. If an I(1) variable in the VAR is exogenous then the level terms, $\mathbf{y}_{t-1}$, do not enter the equation of the VAR for the variable in question, and the corresponding row of the parameter matrix Π is zero. Preliminary estimates indicate that this parameter restriction is not rejected by the data for the household income, social class and relative wage variables, and so in subsequent analysis these are treated as conditioning variables.[5]

Table 1 reports the cointegration test results with the maximal eigenvalue and trace statistics adjusted for degrees of freedom as proposed in Reimers (1992) (the unadjusted figures are given in parentheses). The adjusted trace statistics indicate at most one cointegrating vector in the data for both males and females. However, given the magnitude of the estimated eigenvalues, the adjustment may be excessive and we proceed on the basis of two cointegrating vectors in each case. The Johansen maximum likelihood estimates of the cointegrating vectors are given in cols (1.i) and (1.ii) of Table 1. For both males and females, the first cointegrating vector appears consistent with the augmented human capital model outlined in Section 2, aside from the negative coefficient on the real income variable. Moreover, the parameter estimates for males and females are broadly comparable in terms of signs and orders of magnitude. In the case of the second cointegrating vector, however, there are more substantive differences between males and females, and the interpretation in each case is less obvious. The next stage is to consider possible (over-) identifying restrictions on the parameters of the cointegrating vectors.

Table 1 (a) and (b) approx here

The restricted estimates are reported in cols (2.i) and (2.ii) together with their standard errors. The first vector in each case corresponds to the long-run relationship determining rates of participation in further education. The results show a strong positive relationship between the average level of GCSE attainment and the rate of participation in further education for both males

[5] We are grateful to an anonymous referee of this journal for pointing out this approach to us.

and females, as has been found in many studies based on individual data (Cheng, 1995; Andrews and Bradley, 1997; Rice, 1999). Neither the real income variable nor the social class variable is found to exert a significant direct effect on participation rates in further education for males or females. At first sight, this appears to conflict with earlier findings, in particular Whitfield and Wilson (1991) report a significant positive relationship between the social class variable and the rate of participation in further education for both males and females. However, social class and real household income enter the second cointegrating vector with statistically significant coefficients, and so this earlier result may be a consequence of the failure to allow for the possibility of more than one cointegrating vector.

Of the labour market variables, increases in the relative earnings of those in professional and related occupations raise the proportion of the male cohort undertaking further education as expected, but the corresponding effect for young females is not significant. For both cohorts, unemployment has a small, but significant, positive effect on rates of participation in the long run, consistent with the argument that increases in unemployment reduce the opportunity cost of investment in further education. Participation rates for males appear to be somewhat more responsive to fluctuations in unemployment than those for their female counterparts, and this too is supported by the results from micro-level studies (Rice, 1999). We find some evidence that the provision of youth training offsets the effects of higher unemployment rates, but this effect is significant in the long run only for females, not for males.

The interpretation of the second cointegrating vector remains unclear. For the female cohort, it may be plausibly argued that this vector identifies the relationship determining the average level of academic attainment over the long run. The observed positive relationship between GCSE attainment levels and the real income, social class and relative wage variables is consistent with the hypothesis that academic attainment at age 16 years is positively related to past levels of investment in human capital. The positive relationship between the participation rate and the average level of academic attainment is consistent with a role model effect whereby higher participation rates encourage improved attainment levels in subsequent cohorts. It is possible also that higher participation rates tend to bid up entry requirements in some areas of further education and this induces higher average levels of attainment.

Unfortunately, this interpretation is less convincing when we consider the estimates obtained in the case of the male cohort, and in particular the implied negative relationship between the average level of attainment level and real income. An alternative explanation is that in the case of young males, the second vector is picking up the long-run relationship determining the unemployment rate. To investigate this further it would be necessary to expand the information set to include variables that are more directly related to the determination of the youth unemployment rate.

4. The Dynamic Behaviour of the Rate of Participation in Further Education.

In this section. we examine the dynamic behaviour of the rate of participation in further education over the sample period. Our concern here is to assess how well the dynamic specification is able to account for movements in the participation rate. It is not our intention in this paper to offer a complete account of all variables in the system, and the VAR in (3) is unlikely to provide a satisfactory description of the behaviour of variables such as the unemployment rate. To mitigate the effects of possible misspecification elsewhere in the system, we consider a single dynamic equation for the participation rate with contemporaneous terms in the other stochastic variables included as regressors. Given that the contemporaneous regressors may not satisfy the conditions for weak exogeneity, a generalised instrumental variable (GIV) estimator is considered, as well as the OLS estimator.

The starting point is a general dynamic specification of the form

$$\Delta \ln\left(\frac{P}{1-P}\right)_t = \beta_0 + \beta_1 d_t + \beta_2 \Delta \ln\left(\frac{P}{1-P}\right)_{t-1} \\ \mathbf{b}_0{}'\Delta\mathbf{x}_t + \mathbf{b}_1{}'\Delta\mathbf{x}_{t-1} + \mathbf{a}'\mathbf{z}_{t-1} + \varepsilon_t \qquad (4)$$

$\Delta\mathbf{x}$ denotes the vector of stochastic variables excluding the participation rate P, and d is the step dummy for the raising of the minimum school-leaving age. $\mathbf{z}_{t-1}$ denotes the vector of I(0) 'equilibrium errors' computed from the estimates of the cointegrating vectors reported in Table 1. From this general form, a parsimonious specification is obtained through testing a sequence of restrictions on individual parameters. The parameter estimates for the preferred restricted form, together with their standard errors and a range of diagnostic statistics are reported in Table 2. The OLS and GIV estimators yield similar results in terms of the individual parameter estimates; the GIV estimates tend to be somewhat larger in absolute terms, but less well determined.

(Table 2(a) and (b) approx. here)

Overall, the ECMs for both males and females appear to be well specified. The diagnostic statistics based on the residuals are insignificant and the model's forecasts show little evidence of parameter instability. The equilibrium error corresponding to the first cointegrating vector enters the dynamic equations with a large positive coefficient consistent with deviations from the long-run participation equation inducing substantial equilibrating adjustments in the rate of participation, conditional on the contemporaneous values of the other stochastic variables. The second equilibrium error terms enter with a positive coefficient also but the implied adjustment coefficient is smaller in magnitude and in the case of males, is not statistically significant.

A better understanding of the nature of the dynamics may be gained by re-writing the ECMs in Table 2 in terms of the levels of the variables as follows:

Females:

$$
\begin{aligned}
Ln\left(\frac{P}{1-P}\right)_t = {} & 0.078 + 0.07CS_t + 0.35Ln\left(\frac{P}{1-P}\right)_{t-1} + 0.41Ln\left(\frac{Q}{1-Q}\right)_{t-1} - 0.30Ln\left(\frac{Q}{1-Q}\right)_{t-2} \\
& + 0.13Ln\left(\frac{U}{1-U}\right)_t - 0.01Ln\left(\frac{U}{1-U}\right)_{t-1} - 0.69LnYT_{t-1} + 0.37Ln\left(\frac{HE}{1-HE}\right)_t \\
& + 0.25Ln\left(\frac{HE}{1-HE}\right)_{t-1} - 0.5Ln\left(\frac{HE}{1-HE}\right)_{t-2} - 0.48LnI_t + 0.91LnI_{t-1} \\
& + 0.38LnW_{t-1} + 0.21Ln\left(\frac{S}{1-S}\right)_{t-1} + 0.01trend
\end{aligned}
$$

Males

$$Ln\left(\frac{P}{1-P}\right)_t = -1.16 + 0.1CS_t + 0.16Ln\left(\frac{P}{1-P}\right)_{t-1} - 0.30Ln\left(\frac{P}{1-P}\right)_{t-2} + 0.9Ln\left(\frac{Q}{1-Q}\right)_t$$

$$- 0.14Ln\left(\frac{Q}{1-Q}\right)_{t-1} + 0.1Ln\left(\frac{U}{1-U}\right)_t + 0.04Ln\left(\frac{U}{1-U}\right)_{t-1} + 0.45Ln\left(\frac{HE}{1-HE}\right)_t$$

$$- 0.1Ln\left(\frac{HE}{1-HE}\right)_{t-1} - 0.87\Delta LnYT_{t-1} - 0.54LnW_t + 0.28LnW_{t-1} + 0.71LnW_{t-2}$$

Consistent with the hypothesis of role model effects across successive cohorts, the rate of participation shows positive first-order autocorrelation for both males and females. The effect is less strong in the case of males and offset by a negative second order term, which is not evident in the data for females. This may explain why the 1990s have seen more rapid growth of participation rates among females than among their male counterparts.

Among young males, improvements in the average level of academic attainment produce a large contemporaneous increase in the rate of participation causing the actual rate to overshoot its steady state value and producing a negative effect one period later. The same pattern is observed in the female data, but lagged one period. This may come about because young girls are more likely to make their decisions about whether to continue to further education in advance of their GCSE results, and on the basis of expected levels of attainment. A number of the explanatory variables – the higher education and relative wages - display a similar patterns of alternating signs in the successive lagged values with the result that the actual path for the rate of participation tends to oscillate around its steady state path.

5. Growth of Further Education During the 1990s: Conclusions.

We conclude with a closer examination of the behaviour of the rate of participation in further education over the last decade. Figure 2 illustrates the estimated long-run path for the rate of participation in further education, together with that of the actual rate for the years since 1980. In the early 1980s, the rapidly rising unemployment among young persons contributed to faster growth in the long-run participation rate in further education, particularly among young males. Among young women, the effects of higher unemployment were offset in part by the rapid expansion of youth training measures. Other key determinants - the average level of academic attainment, the size of the higher education sector - increased only slowly over this period. After 1984, unemployment rates among young workers fell reducing participation rates over the long run. Continued modest growth in the other variables was sufficient to offset these effects and the long-run participation rate remained stable between 1983 and 1988.

(Fig. 2 approx. here)

After 1988, conditions combine to produce a period of very rapid growth in participation rates. Following the introduction of the GCSE qualification, the average level of academic attainment within a cohort improves significantly year-on-year. This is accompanied by the rapid expansion of the higher education sector. Initially, the effects on participation rates are tempered by falling unemployment, but with unemployment rates among younger workers increasing again from 1990, the growth in long-run participation rates tends to accelerate.

The movements of the actual rate of participation around its long-run path reflect the relative speeds of adjustment to shocks. The evidence on the dynamics suggests that the short-run impact of changes in the average level of GCSE attainment in the cohort, and the size of the higher education sector exceed the long-run effects, leading to over-shooting, and causing the actual path to oscillate around the long-run path. As a result, the actual rate of participation grew even more rapidly than the long-run rate after 1988. Towards the end of the period, however, we detect signs of the actual rate of participation converging towards the long-run path.

What distinguishes the years since 1988 from the earlier period is the rapid improvement in the average level of qualification achieved on completion of compulsory schooling, coupled with the substantial expansion of higher education. The average level of GCSE attainment continues to improve but at a more modest rate than in recent years and the expansion of the higher education sector has ground to a halt. Moreover, since 1994, unemployment rates among young people have been falling once again. As a result, the long-run path for the rate of participation in further education is declining in the late 1990s. With the actual rate of participation above the steady-state rate, and the latter declining, the short-run dynamics imply falling participation rates in further education during the latter half of the 1990s. Recently published estimates confirm this with participation among 16 year-olds in England and Wales falling from their 1994 peak.

This paper set out to establish to factors responsible for the rapid growth in further education of the last decade. The results of our analysis of the time series evidence are broadly consistent with an augmented human capital model of investment in further education over the long run. In the short-run, however, other important factors are at work with strong role model effects across successive cohorts, particularly it would appear among young women, and these would seem to explain why female rates of participation have outstripped those of their male counterparts in recent years. The improvement in attainment levels that followed the introduction of the GCSE qualification has played a major role in encouraging greater participation for both males and females. Our results suggest that for females at least this improvement is not simply a consequence of exogenous changes in assessment methods or school curriculum, but is itself a response to the greater returns in the market to workers with further educational qualifications.

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Table 1. Participation Rates in Full-time Further Education – Cointegration Analysis.
(a) Females.
Endogenous variables: P^f, Q^f, U^f, YT, HE. Exogenous variables: I, S, W^f.

		Eigenvalue	*Trace statistic*	**Maximal statistic**
H_0:r=0;	H_1:r>=1;	0.8931	142.1 (194.7)	60.4 (82.7)
H_0:r<=1;	H_1:r>=2;	0.7096	81.7 (112.0)	33.4 (45.7)
H_0:r<=2;	H_1:r>=3;	0.6010	48.3 (66.2)	24.8 (34.0)
H_0:r<=3;	H_1:r>=4;	0.4530	23.5 (32.2)	16.3 (22.3)
H_0:r<=4;	H_1:r>=5;	0.2353	7.3 (9.9)	7.3 (9.9)

Maximum Likelihood Estimates of Cointegrating Vectors

	(1.i)	(1.ii)	(2.i)	(2.ii)
$Ln(P^f/(1-P^f))$	-1.0000	0.3974	-1.0000	0.8973 (0.0752)
$Ln(Q^f/(1-Q^f))$	0.2579	-1.0000	0.6268 (0.1271)	-1.0000
$Ln(U^f/(1-U^f))$	0.0141	0.1057	0.0923 (0.0376)	-
YT	-0.8130	-0.6393	-0.5540 (0.2312)	-
Ln(HE/(1-HE))	0.1992	0.1113	0.1001 (0.0504)	-
LnI	-1.0223	1.1037	-	0.6395 (0.1982)
Ln(S/(1-S))	0.3924	0.5288	-	0.3067 (0.0882)
LnW^f	0.0228	0.9763	-	0.5663 (0.1395)
Trend	0.0530	-0.0306	0.0277 (0.0065)	-0.0363 (0.0039)

$\chi^2(4) = 4.8407$

* Standard errors of parameter estimates are given in parentheses.

Table 1. Participation Rates in Full-time Further Education – Cointegration Analysis.
(b) Males.
Endogenous variables: P^m, Q^m, U^m, YT, HE. Exogenous variables: I, S, W^m.

	Eigenvalue	*Trace statistic*	**Maximal statistic**
H_0:r=0; H_1:r>=1;	0.8320	149.9 (204.7)	48.2 (66.0)
H_0:r<=1; H_1:r>=2;	0.7520	101.2 (138.7)	37.7 (51.6)
H_0:r<=2; H_1:r>=3;	0.6534	63.6 (87.1)	28.6 (39.2)
H_0:r<=3; H_1:r>=4;	0.5450	35.0 (47.9)	21.3 (29.1)
H_0:r<=4; H_1:r>=5;	0.3984	13.7 (18.8)	13.7 (18.8)

Maximum Likelihood Estimates of Cointegrating Vectors

	(1.i)	(1.ii)	(2.i)*	(2.ii)*
$Ln(P^m/(1-P^m))$	-1.0000	0.8804	-1.0000	0.7016 (0.0643)
$Ln(Q^m/(1-Q^m))$	0.4169	-1.0000	0.6690 (0.1421)	-1.0000
$Ln(U^m/(1-U^m))$	0.0861	-0.5320	0.1229 (0.0129)	-0.4392 (0.2245)
YT	0.1201	0.4491	-	-
Ln(HE/(1-HE))	0.5438	-0.0910	0.3903 (0.0841)	-
LnI	-1.4022	-2.2424	-	-2.4399 (1.6345)
Ln(S/(1-S))	0.2298	0.4658	-	0.3779 (0.2459)
LnW^m	0.5658	0.4930	0.4198 (0.1321)	-
Trend	0.0385	0.0631	-	0.0708 (0.0365)
			$\chi^2(5) = 1.7414$	

* Standard errors of parameter estimates are given in parentheses.

Table 2. Participation Rates in Full time Further Education – Estimates of the Error Correction Model. (a) Females

	OLS estimates	*IV estimates*
$\Delta Ln(Q^f/(1-Q^f))_{t-1}$	0.3343 (0.1314)	0.3012 (0.1385)
$\Delta Ln(U^f/(1-U^f))_t$	0.1393 (0.0192)	0.1283 (0.0263)
$\Delta Ln(HE/(1-HE))_t$	0.2329 (0.1304)	0.3738 (0.1925)
$\Delta Ln(HE/(1-HE))_{t-1}$	0.6217 (0.1475)	0.5002 (0.1863)
ΔLnI_t	-0.5045 (0.2263)	-0.4795 (0.2363)
$Z^f_{1,t-1}$	1.2848 (0.1712)	1.2475 (0.1925)
$Z^f_{2,t-1}$	0.6868 (0.1697)	0.6710 (0.1962)
CS_t	0.0736 (0.0147)	0.0703 (0.0154)
Intercept	-0.0372 (0.0146)	-0.0347 (0.0164)
S.E of regression.	0.0320	0.0329
Serial correlation: 1st order	$F(1,27)=1.8769$	$\chi^2(1)=1.3994$
2nd order	$F(2,26)=3.1082$	$\chi^2(2)=3.9281$
Functional form.	$F(1,27)=0.4001$	$\chi^2(1)=0.1988$
Normality.	$\chi^2(2)=2.7649$	$\chi^2(2)=1.8828$
DF test for residuals.	-7.1613	
Parameter restrictions.	$F(10,17)=0.5980$	
Instrument validity*		$\chi^2(6)=3.2261$
Prediction Errors:		
1985-1994		
Sum sqs pred errors	0.0015	0.0017
	$F(10,18)=0.7866$	
1989-1994		
Sum sqs pred errors	0.0009	0.0009
	$F(6,22)=0.4612$	
1992-1994		
Sum sqs pred errors	0.0020	0.0034
	$F(2,26)=0.6975$	

*Additional instruments are $\Delta Ln(U^f/(1-U^f))_{t-1}$; $\Delta Ln(U^f/(1-U^f))_{t-2}$, $\Delta Ln(P^f/(1-P^f))_{t-1}$;$\Delta Ln(Q^f/(1-Q^f))_{t-2}$; ΔLnI_{t-1}, $\Delta Ln(S/(1-S))_{t-1}$; ΔLnW^f_{t-1}; ΔLnW^f_{t-2}..

Table 2. Participation Rates in Full time Further Education – Estimates of the Error Correction Model. (b) Males

	OLS estimates	***IV estimates***
$\Delta Ln(P^m/(1-P^m))_{t-1}$	0.3282 (0.0712)	0.3041 (0.0781)
$\Delta Ln(Q^m/(1-Q^m))_t$	0.8788 (0.1336)	0.8992 (0.1952)
$\Delta Ln(U^m/(1-U^m))_t$	0.0928 (0.0135)	0.1033 (0.0155)
ΔYT_{t-1}	-0.9618 (0.2680)	-0.8733 (0.2835)
$\Delta Ln(HE/(1-HE))_t$	0.3803 (0.0869)	0.4539 (0.1263)
ΔLnW^m_t	-0.4678 (0.2060)	-0.5357 (0.2185)
ΔLnW^m_{t-1}	-0.6583 (0.2154)	-0.7093 (0.2239)
$Z^m_{1,t-1}$	1.0966 (0.1133)	1.1400 (0.1217)
CS_t	0.0946 (0.0116)	0.0980 (0.0121)
Intercept	-0.0680 (0.0108)	-0.0740 (0.0119)
S.E of regression.	0.0244	0.0253
Serial correlation: 1st order	$F(1,26)=2.2116$	$\chi^2(1)=2.4509$
2nd order	$F(2,25)=2.5273$	$\chi^2(2)=3.1334$
Functional form.	$F(1,26)=0.9549$	$\chi^2(1)=0.9187$
Normality.	$\chi^2(2)=1.2108$	$\chi^2(1)=1.1146$
DF test for residuals.	-7.5241	
Parameter restrictions.	$F(9,18)=0.9895$	
Instrument validity*		$\chi^2(7)=1.5582$
Prediction Errors:		
1985-1994		
Sum sqs pred errors	0.0013	0.0018
	$F(10,17)=1.3948$	
1989-1994		
Sum sqs pred errors	0.0014	0.0029
	$F(6,21)=1.2261$	
1992-1994		
Sum sqs pred errors	0.0026	0.0044
	$F(2,25)=1.6494$	

*Additional instruments are $\Delta Ln(P^m/(1-P^m))_{t-2}$; $\Delta Ln(Q^m/(1-Q^m))_{t-1}$; $\Delta Ln(Q^m/(1-Q^m))_{t-2}$; $\Delta Ln(U^m/(1-U^m))_{t-1}$;$\Delta Ln(U^m/(1-U^m))_{t-2}$, ΔLnI_{t-1} ; ΔLnI_{t-2}; $Ln(S/(1-S))_{t-1}$; $\Delta Ln(HE/(1-HE))_{t-1}$

Fig. 1. Proportion of the cohort aged 16 years in further education in England and Wales

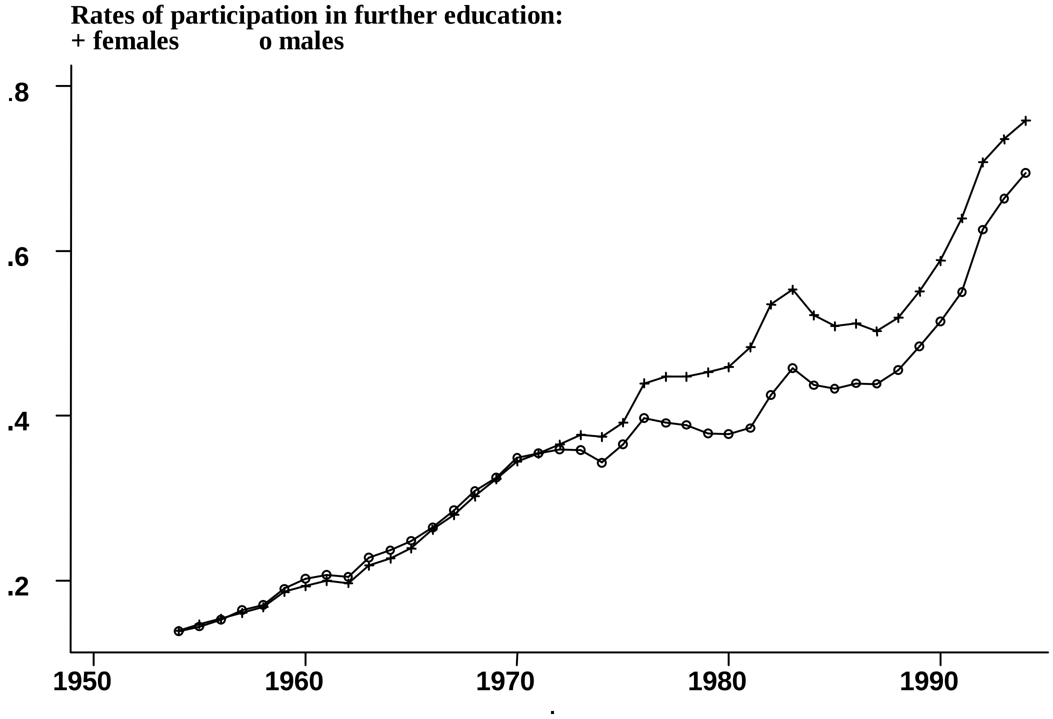
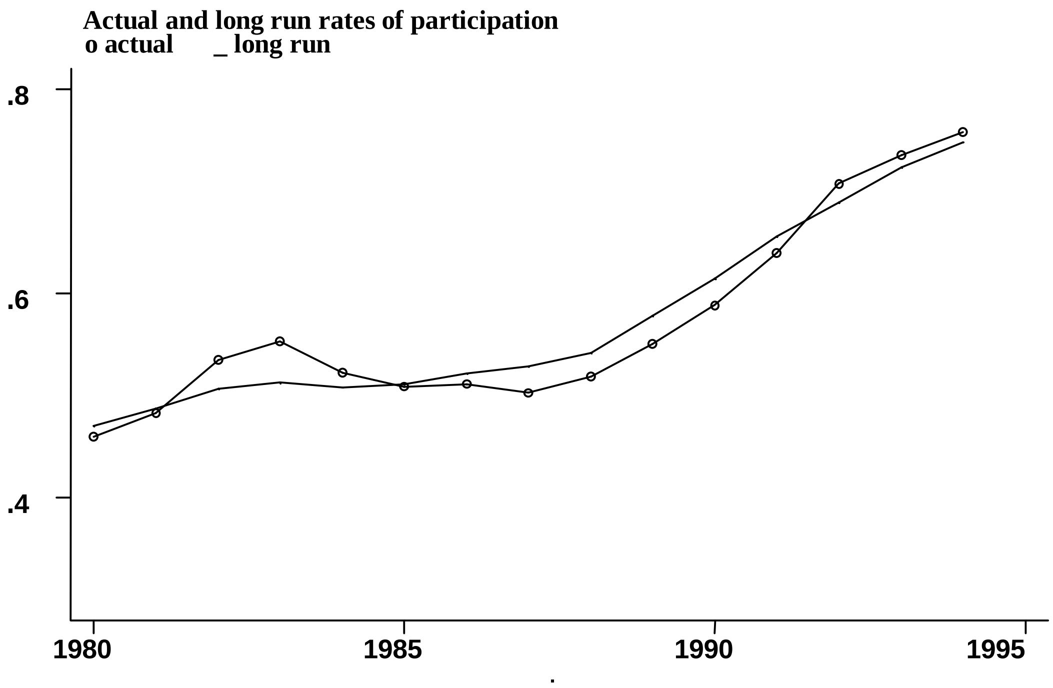
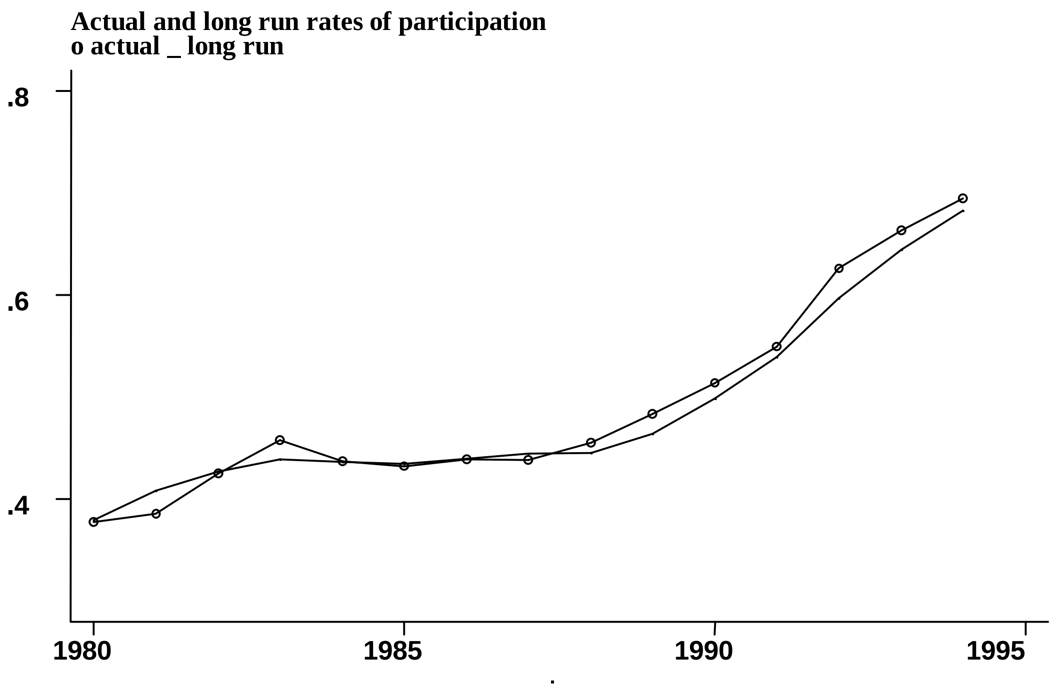


Fig. 2 Rates of participation in further education 1980-94: Actual and estimated long run:

(a) Females



(b) Males



Appendix 1.

P - proportion of the male/female cohort aged 16 years participating in full-time education in January of the academic year.

The cohort is defined as those aged 16 years at the start of the academic year (i.e. at August 31st), and therefore members of the cohort completed their compulsory schooling in the previous academic year. From 1979/80 onwards, the published data for numbers in full-time education is based on age at the start of the academic year, but prior to this date, the published figures are by age at the end of the calendar year. As a result, the published figures for the number of 16 year-olds in full-time education for 1973/4 to 1978/79 include some individuals who have yet to complete their compulsory schooling. To overcome this problem, the raw data is adjusted to obtain a consistent series based on age at August 31st for the whole of the sample period. Department of Education sources provide a limited amount information on the numbers in full-time education based on both age at August 31st and age at December 31st for the period prior to 1979/80. This is used to compute "age-adjustment" multipliers equal to the ratio of the number of 16 year-olds in full-time education (August 31st definition) to the number of 16 year-olds in full-time education (December 31st definition). Separate multipliers are computed for males and females, and for the maintained schools, independent schools and FE college sector. Further information on the adjustments made, together with the set of "age-adjustment" multipliers used, are available from the authors on request.

Department for Education, 'Participation in education by 16-18 year olds in England 1979/80 to 1993/4'. Statistical Bulletin, 16/93. Welsh Office, Statistics of Education in Wales, 1977/8-1993/4.

Department of Education & Science, Statistics of Education: Schools, 1953/4-1978/9.

Department of Education and Science, Statistics of Education: Further Education, 1961/2-1978/9.

Q - proportion of the male/female cohort achieving 5 or more GCSE qualifications at grade C or higher, or their equivalent.

Data relating to levels of attainment in the GCSE examinations among the cohort taking the examinations in year 11 of schooling is available from 1988. Prior to this, the annual School Leavers Survey provides information on the number of school-leavers of a given age with 5 or more GCE/CSE qualifications at grade C or higher, and this is used to derive estimates for a given age-cohort.

Department for Education, Statistics for Education: School Leavers and GCSE Examinations, 1993/4.

Department for Education; Statistics for Education: School Examinations Survey/School Leavers Survey, 1958/9-1990/1

I - average level of real net income of households in Great Britain.

The series is compiled from data on the average weekly income of all households in the Family Expenditure Survey sample, deflated by the all-items Retail Price Index.

Department of Employment, Family Expenditure Survey, 1954,1957-9,1960-94. National Food Survey Committee, Domestic Food Consumption and Expenditure: Annual Report of the National Food Survey Committee, 1950-7.

S - proportion of households in Great Britain with a household head in a managerial, professional or related occupation.

The series is compiled from information published in the FES relating to the occupations of heads of households in the survey. The measure adopted includes employers and managers, those in professional, technical and administrative occupations and teachers. After 1986, a broader classification covering employers and managers, those in professional occupation and those in occupations classified as intermediate non-manual is used.

Department of Employment, Family Expenditure Survey, 1954,1957-9,1960-94.

W - average gross weekly earnings of managerial, professional and related occupations relative to the average gross weekly earnings of manual workers for males/females in full-time employment.

For 1970 onwards, annual data on the average gross weekly earnings of those in full-time employment by occupational category is taken from the New Earnings Survey. British Labour Statistics Historical Abstract provides data on the average weekly earnings of administrative, technical and professional employees in the banking, insurance and public sector for the period 1955 to 1968.

Department of Employment, New Earnings Survey, 1970-94. Department of Employment, British Labour Statistics Historical Abstract.

U - the rate of unemployment among males/females aged between 18 and 20 years.

The series is compiled from Department of Employment data on the numbers registered as wholly unemployment and unemployment rates by age and sex.

Department of Employment, Department of Employment Gazette, 1955-94.

HE: Home initial entrants aged less than 21 years to full-time and sandwich courses in the higher education sector in Great Britain as a proportion of the age cohort (i.e. the average of the age cohorts aged 18 years and 19 years).

The above definition corresponds to the DfEE's Age Participation Index. For 1966/7 to 1993/4, the API is published in the DfEE Statistical Bulletins. Prior to 1966/7, the API is estimated from data on the numbers of initial entrants to higher education and the size of the age cohorts.

Department for Education, 'Students in higher education in Great Britain', Statistical Bulletins, 12/80,17/83,9/85,8/92,13/94. Central Statistical Office, Annual Abstract of Statistics, 1960-72.

YT - the proportion of the cohort aged 16 years (males and females combined) participating in YT or a related scheme.

For 1976/7 to 1989/90, data on the numbers of individuals participating in government youth training measures by age are obtained from published sources. The Statistical Services Division of the Department of Employment provided information on the age distribution of participants on YTS and related schemes for the later years.

Department of Education and Science, 'Educational and economic activity of young people aged 16 to 18 years in England from 1974/75 to 1989/90'. Statistical Bulletin, 13/91

CS – step dummy for extending the period of compulsory schooling under the 1972 Education Act. Variable takes the value 0 to 1973/74 and the value 1 thereafter.