Intergenerational Mobility and Assortative Mating in the UK

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Abstract

In this paper I model family income mobility in Britain for the 1958 and 1970 cohorts. Previous research has shown an increase in the association between children's earnings and parental income over this period, particularly for sons. I consider the role of partnership formation in adding to intergenerational family income inequalities; and illustrate the role played by partners in contributing to trends in intergenerational mobility. I find that partners have a substantial role in generating family income persistence for daughters, with a strong association between husbands' earnings and parental income in both cohorts. For sons I find a striking change. In the first cohort there is no link between sons' wives' earnings and parental income; in the second cohort this is as strong as the relationship between parents and sons-in-law. Consequently, when I add partners' earnings to the story I find that the change in mobility for sons is even stronger than when it is measured on individual earnings.

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1. Introduction

In recent years economists have become increasingly interested in studying the links between children's success and their family background. Researchers have studied the extent to which parental income and wealth have influenced educational attainment, earnings and income. A strong relationship between parental income and children's outcomes is frequently interpreted as demonstrating an unacceptably closed society where individuals may not achieve their full potential.[1]

These studies have tended to consider adult children in isolation, however, and very few have taken into account the role of partnership formation in determining the economic and social status of the grown-up child (exceptions are Chadwick and Solon, 2002 and Peters, 1992 for the US and Ermisch et al, 2004 for the UK). As a result of economies of scale, household public goods and income pooling it is easy to argue that household income is a more important measure of welfare than individual earnings. Indeed, poverty is usually measured at the household, rather than individual, level. Consequently, if we are concerned with the persistence of welfare across generations we may think that the association between household incomes across generations is the most relevant measure.

It is immediately clear that an individuals' choice of partner will have an important bearing on intergenerational persistence at the household[2] level. If individuals choose partners with similar economic characteristics to their parents then the child's household income may be more strongly related to parental income than individual earnings are. In this way assortative mating (where individuals choose partners with similar characteristics to their own) can be an important mechanism adding to the persistence of income across generations. In this case, estimates of intergenerational mobility obtained at the individual level will underestimate the full extent of intergenerational income persistence.

This paper is the first to estimate changes in intergenerational income mobility at the household level in the UK, and to begin to understand how assortative mating contributes to intergenerational income links. I use data from two cohorts: one born in 1958 and one born in 1970.

[1] A summary of this general literature can be found in Solon (1999), while Corak (2004) explicitly considers the policy implications that can be gained through comparisons across countries.

[2] From this point on "household income" refers to the income of the household head and any partner, it does not, in fact, include the contributions of any other members of the household. This is due to the nature of the data available in the cohort studies.

Work by Blanden et al (2004) explores the change in individual mobility in the UK for the cohort datasets used in this paper. They find a substantial decrease in the extent of mobility in the UK when individual earnings (when cohort members are in their early 30s) are linked with parental income (at age 16), which is particularly pronounced for sons. Here I examine whether these changes are replicated once the partners of both sons and daughters are added to the picture.

The approach taken here is closest to those used by Chadwick and Solon (2002) and Ermisch et al (2004). Chadwick and Solon (2002) explore family income mobility and assortative mating in the US. They argue that the connection between intergenerational mobility and assortative mating is particularly important for women, where own earnings are frequently a minority contribution to family income. Chadwick and Solon examine assortative mating by estimating the relationship between daughters' husbands' earnings and parental income, they find this relationship to be as strong as the relationship between parental income and daughters' own earnings.

Ermisch et al (2004) use the British Household Panel Survey and the German Socio-Economic Panel to consider these issues. As the BHPS does not include measures of parental income, the analysis is based upon the relationship between own and spouses' occupation and recalled parents occupation at age 14 for both countries, while correlations in earnings are measured for Germany. In all cases partner-parent relationships are substantial, but smaller than the intergenerational relationships between parents and their own off-spring.

The very first study of intergenerational mobility in Britain, Atkinson et al (1983), also considered the role of assortative mating in intergenerational mobility. For the children of fathers in the 1950 York Rowntree Inquiry the association between fathers' and sons-in-law's earnings is slightly larger in magnitude than the one between fathers and sons, this demonstrates strong assortative mating.

The previous literature has focused either on the persistence between sons and their parents or sons and their parents-in-law, largely playing down the role of women's earnings. This is because the intergenerational mobility of women's earnings is difficult to measure, as women participate in the labour market less often than men, and when they do, they often work part-time. In some ways this problem can be ignored, if we want to know about intergenerational earnings persistence at a point in time it is reasonable to simply measure this without taking account of selection issues. In fact, measuring the way in which intergenerational mobility changes as patterns of female participation alters is an important focus of this paper. However, as we shall see, the theoretical background to these models is firmly couched in terms of permanent income. Therefore, I investigate how the results are being affected, and possibly distorted, by participation decisions.

In order to motivate the analysis, I present a simple model of intergenerational mobility and the marriage market. This demonstrates why persistence between an individual's parental income and his or her partner's earnings may vary. The conclusions of this simple approach are clear; the link between parental income and partner's earnings is increasing with the strength of assortative mating. If assortative mating is strong, the elasticity of partner's earnings with respect parental income will be similar to that for the child's own earnings.

In my results, I estimate the relationships described in the model to build up a picture of how assortative mating and intergenerational transmissions are related. I first develop a picture of how individuals match by education level. I find some evidence that assortative mating has increased. I also consider this the relationship between parental income and the education of partners; these relationships are strong in all cases, pointing to substantial assortative mating. However, I find less evidence of change by these measures.

The most important set of results show the regression coefficients and correlations between the earnings of sons, daughters-in-law, daughters and sons–in-law and their parents' (or parents-in-law's) income. I find, as expected, that the correlation between partners' earnings and parental incomes are strong. The most interesting results show that these relationships have changed quite substantially, especially for the earnings of daughters-in-law. In the earlier cohort, there was only a very weak relationship between daughters-in-law's earnings and parental income, whereas in the later cohort this is as almost as strong as the relationship between sons and their parents. There is a smaller, but still significant, increase in the relationship between daughters' parental incomes and their partners' earnings.

The challenge is to interpret these results. As emphasised above, we want to understand both the importance of changes in the underlying relationships and changes in the selection into employment. I therefore compare my uncorrected estimates for the employed sample with those which correct for the selection into employment. I find that in the second cohort, women's participation decisions are more strongly correlated to their potential wages. This introduces an upward bias on the change in intergenerational elasticities when women's earnings are the dependent variable. Indeed, this change in the selection into employment is in part responsible for the rise in the relationship between the earnings of daughters-in-law and the parental income of sons.

In the next section, I present the theoretical background and measurement approach used in this analysis. Section 3 describes the data, particularly focusing on the changes in household formation and employment between the two cohorts. In Section 4, I discuss how the cohorts match by education level while Section 5 considers the relationship between parental income and the education levels of the next generation. In Section 6, I present my main results on intergenerational mobility, and show how they are influenced by changes in participation. In Section 7, I discuss the interpretation of my findings while Section 8 concludes.

2. Theoretical Background and Measurement Issues

Theory

A model of how assortative mating contributes to intergenerational mobility must begin by characterising how couples match by their characteristics.

Sociologists have traditionally dominated research into marriage and a focus of their work has been the investigation of the extent to which characteristics influence who marries who. The main aspects explored in this literature are marriage within and between racial, religious and socioeconomic groups; in all cases individuals tend to marry individuals like themselves.

The early formal models of marital sorting were based on mathematical assignment models. The idea is akin to all singles being placed in the room together and then leaving at the end of the evening paired-off forever. In Gale and Shapely (1962) individuals have a single ranking of partners which is common to all in the marriage market. In this case a pure sorting equilibrium will result; the nth ranked woman and the nth ranked man will be matched, and so on throughout the distribution. Becker's model (1973, 1974) has a richer description of the benefits of marriage and is more strongly rooted in economic theory.

For Becker, all potential marriages have an output Z; Z includes the earnings of both partners, the gains from the division of labour within marriage, as well as the utility from rearing children and from receiving affection within the family. In a utility maximising framework all individuals will be seeking the marriage with the highest possible Z. In a sorting model with no frictions, pareto efficiency implies that men and women will sort into partnerships which maximise the total amount of Z. The mathematical properties of submodularity and supermodularity state that output is maximised if 'likes' are matched when male and female traits (A_h and A_w) are complements in producing Z.

$$\frac{\partial^2 Z(A_h, A_w)}{\partial A_h \partial A_w} > 0 \tag{1}$$

'Unlikes' are matched when male and female traits are substitutes in producing Z.

$$\frac{\partial^2 Z(A_h, A_w)}{\partial A_h \partial A_w} < 0 \tag{2}$$

From this follows the prediction that couples will be positively matched on characteristics like education and ability that are complements in the production of high quality children and negatively matched on wage rates as these are substitutes in the production of market goods. Of course, the strong correlation between education, ability and wages, means that it would be very difficult to separately identify a negative relationship between the wage rates of couples. Moreover, Lam (1988) argues that in the presence of household public goods, wage rates should be positively correlated, even conditional on other characteristics.

Here I use the model from Ermisch at al (2004), where assortative mating occurs on the basis of human capital. In Chadwick and Solon (2002) mating occurs on the basis of full income. Simplifying the matching process to human capital serves to reduce the number of parameters relevant to the model and seems a reasonable simplification of the above discussion of marital matching.

Assortative mating is modelled as a positive correlation between the human capital of wives (H_{wi}) and husbands (H_{hi}); s .

$$Corr(H_{wi}, H_{hi}) = s \tag{3}$$

For both husbands and wives income is positively related to human capital, although the return to human capital may vary across genders as in equations (4) and (5) below.

$$\ln Y_{wi} = t_w + g_w H_{wi} + v_{wi} \tag{4}$$

$$\ln Y_{hi} = t_h + g_h H_{hi} + v_{hi} \tag{5}$$

In this formulation the intergenerational relationship is driven by parents' optimising behaviour, but this is not essential for the general conclusions of the model, in a similar model in Lam and Schoeni (1994), the mechanism behind the correlation between education across generations is left ambiguous and similar conclusions are reached.

The parental utility function in Ermisch et al includes parental consumption and the child's household income, so that their child's partner's income is also included, p indicates the extent to which parents are altruistic and care about their child's income. From now on I shall couch the model in terms of the wife's parents' income, so parameters and variables relating to the parents are subscripted w, however the model is fully symmetric for husbands and wives.

$$U_{wi}^{parents} = (1-p)\ln C_{wi}^{parents} + p \ln E(Y_{wit} + Y_{hit}) \tag{6}$$

Parents solve this model subject to their budget constraint. In this example debt and bequests are not permitted, so parents must spend all their available income on their own consumption and the education of their children. Each unit of human capital has a price p_H .

Solving the model gives the following solution for the intergenerational parameter, b , the coefficient from a log-log regression of child's income on parents' income. Intergenerational persistence is positively related to parental altruism and the returns to education for women, but negatively related to the cost of investment.

$$\ln Y_{wi} = a + b_w \ln Y_{wi}^{parents} + e_{wi} \text{ where } b_w = pg_w / p_H \quad (7)$$

Similar factors are important for the relationship between husband's income and his wife's parental income. In this case, the male return to education is important and the relationship is moderated by assortative mating and the differences in the distribution of education between husbands and wives.

$$\ln Y_{hi} = a + d_w \ln Y_{iw}^{parents} + e_{hi} \text{ where } d_w = spg_h / p_H \frac{SD^{H_h}}{SD^{H_w}} \quad (8)$$

Putting b and d together enables us to understand more about the expected relationship between these two parameters. If the model is worked through in terms of son's parental income, the relationship is symmetric so that:

$$\frac{d_w}{b_w} = s \frac{SD^{H_h}}{SD^{H_w}} \frac{g_h}{g_w} \text{ and } \frac{d_h}{b_h} = s \frac{SD^{H_w}}{SD^{H_h}} \frac{g_w}{g_h} \quad (9)$$

As shown, if the returns to education and the distributions of human capital are equal for men and women, the ratio of d_w and b_w, will enable the identification of s ; the elasticity between the income of the daughter's partner and her parents' income over the elasticity of her own income with respect to her parents' income will be equal to the extent of assortative mating. In this paper, the focus is on how these relationships change across cohorts. The implication of equation (9) is that increases in b for sons are likely to lead to increases in d for daughters-in-law, ceteris paribus. In addition, increases in s will lead to a rise in d relative to b .

It is also clear that changes in the returns to human capital for men and women have a part to play in this model. If we think of the incomes in this model as permanent, participation will influence the return to human capital over the lifetime. The implication is that if daughters participate more, g_w will increase and b_w will rise relative to d_w, and d_h will rise relative to b_h. This provides an illustration of how changing patterns of participation can influence intergenerational mobility. Of course, it is difficult to observe the implications of these lifetime factors when incomes for the children's generation are observed at only one point in time.

A further implication of the model is that in order to understand changes in the relationship between parental income and partners' earnings, it will be necessary to try to unpick changes in assortative mating (s) and changes in returns, which, as noted, may come through changes in participation. To express this in another way, if the association between daughters-in-law's earnings and parental incomes increases we want to know if this is because individuals have changed the way they match or if the match has remained the same but wives' working patterns have changed.

In my estimations I also measure the link between the family incomes of parents and children. In this case the dependent variable is the combined earnings of the daughter and her partner.

$$\ln(Y_{hi} + Y_{wi}) = a + m \ln Y*_{wi}^{parent} + e_{wi} \tag{10}$$

As shown by Chadwick and Solon (2002) there is an intimate relationship between m, b and d.

$$\text{As } m = \frac{\partial(Y_{hi} + Y_{wi})}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{wi}^{parents}}{(Y_{hi} + Y_{wi})} \tag{11}$$

$$\text{And } b_w = \frac{\partial Y_{wi}}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{wi}^{parents}}{Y_{wi}} \text{ and } d_w = \frac{\partial Y_{hi}}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{hi}^{parents}}{Y_{wi}} \tag{12}$$

It is simple to show that

$$\frac{\partial(Y_{wi} + Y_{hi})}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{wi}^{parents}}{(Y_{hi} + Y_{wi})} = \tag{13}$$

$$\frac{\partial Y_{wi}}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{wi}^{parents}}{Y_{wi}} \cdot \frac{Y_{wit}}{(Y_{hi} + Y_{wi})} + \frac{\partial Y_{hi}}{\partial Y_{wi}^{parents}} \cdot \frac{Y_{wi}^{parents}}{Y_{hi}} \cdot \frac{Y_{hi}}{(Y_{hi} + Y_{wi})}$$

Which is equivalent to $m_w = (1-s)b_w + sd_w$, where s is the share of husband's income in $(Y_{iw} + Y_{ih})$. In an estimation setting all of these variables will be expected values and therefore the decomposition will not be precise, nevertheless this is a useful concept to keep in mind. Household earnings mobility will be a weighted average of the elasticity of the child's earnings with respect to parental income and the elasticity of partner's earnings with respect to parental income, where the weight depends on the share of earnings contributed by each partner. It is clear that changes in the share of earnings contributed by men and women will alter the relative importance of b and d in m, as will increases in female participation. I shall return to this in the results section.

Measurement Issues

In my empirical work I concentrate on estimating b , d and m for both sons and daughters. These are obtained by running linear regression models of log earnings for children when they are in their early 30s on the log income of parents at age 16, controlling for the age and age-squared of both generations.

The primary difficulty in estimating any models of intergenerational persistence is that short-run proxies for income and earnings are being used when the model implies permanent measures. Solon (1989, 1992) and Mazumder (2001) have both highlighted the importance of obtaining multiple measures of parental income and using the average of these over a long period as the explanatory variable. It is not possible here as the only information on parental income which is comparable across both datasets refers only to a single week. In Blanden (2005) I discuss the impact of measurement error for estimates of intergenerational mobility in some detail. The conclusions reached here on changes over time rest on the assumption that the extent of measurement error does not change substantially between the two cohorts; an assumption which is explored in Blanden et al (2004).

I report two estimates for each parameter of interest, the regression coefficient and the partial correlation. In a regression which includes controls for age the partial correlation will be equal to the coefficient on parental earnings times the ratio of the residual standard deviations. This will account for the different variance of earnings due to life-cycle effects, (as parents' incomes are measured when they are relatively older, see Grawe 2003), as well as due to gender and secular changes in earnings inequality.

$$(\text{Corr}_{\ln Y^{\text{parent}}|\text{age},\, \ln Y^{\text{child}}|\text{age}}) = ß*(\frac{SD^{\ln Y^{\text{parent}}|\text{age}}}{SD^{\ln Y^{\text{child}}|\text{age}}}) \tag{14}$$

The interpretation of the coefficient b (or d) is that it describes the proportion of fathers' earnings that are transmitted between generations. If b =.4 then comparing two sets of parents, one with double the income of the other, the child of the richer parents will earn 40% more than the child of the poorer parents. The absolute size of this earnings advantage will obviously depend on how wide the earnings distribution is. The partial correlation measure is based on standardised distributions. A partial correlation of .4 means that if the first father earns one standard deviation more than the second father; the first son will earn .4 of a standard deviation more than the first.

These general issues are common to all papers which measure intergenerational earnings mobility. What is more unique to this paper are the difficulties caused by lower participation and part-time work among women. This generates additional difficulties in estimating intergenerational parameters with women's earnings as the dependent variable (b_w and d_h, in this notation).

The classic analysis of the problems caused by selection bias is presented in Heckman (1979). There are two equations governing the processes, an earnings equation for all women (where, in this case, the explanatory variable would be parental income) and a latent variable relationship governing the decision to participate.

$$Y_i = a + b X_i + u_i \tag{15}$$

$$Z_i = x_o + x_1 Q_i + e_i \tag{16}$$

The woman participates only if Z_i>0. Therefore the regression of the observed Y_i on X_i will be biased by an additional error term, similar to an omitted variable bias. If those with higher earnings are more likely to work, and X_i is positively correlated with earnings, b will be upward biased.

$$Y_i = a + b X_i + E(u_i \mid e_i > -x_0 - x_1 Q_i) \text{ for the employed sample.} \tag{17}$$

As always, it is the change in selection bias which will be important when making comparisons across cohorts. In recent papers, Mulligan and Rubinstein (2004, 2005) discuss the implications of the changing selection of women into work for the gender wage gap. They argue that as the returns to skill have increased, potential wages have become increasingly important in determining the selection into work, leading to an increasing positive selection bias and an observed reduction in the gender wage gap. It is clear that selection bias may have changed for the daughters and daughters-in-law I observe in my data, particularly given the changes in characteristics across cohorts which I highlight in the data section. It is therefore necessary to attempt to model the influence of endogenous selection.

Heckman's framework provides an obvious route to exploring the implications of the changing selection into work. The bias in equation (6.10) can be shown to equal

$$Y_i = a + b X_i + \frac{s_{eu}}{s_e} l_i \text{ where } l_i = \frac{f(v_i)}{F(v_i)} \text{ and } v_i = -x_0 - x_1 Q_i \qquad (18)$$

The bias will be larger the stronger is the correlation between the unobserved determinants of wages and participation. Inclusion of the inverse Mills ratio (l_i), demonstrates that the selection bias will be stronger when participation is low. While the extent of the bias can be estimated by making distributional assumptions it is more convincing to estimate the parameters of the correction using a Probit model of employment. In order to do this, it is necessary to have an exclusion restriction (i.e. a variable which determines employment but not earnings), so that the employment equation can be identified separately from the earnings model.

In their model of intergenerational occupational mobility, Ermisch et al (2004) account for the selection of women into work by incorporating a Heckman selection correction in their estimates. Ermisch et al use a number of variables to predict employment and then use the cubic of the predicted probability of employment index generated to identify the selection. I follow their approach, while recognising that without a truly persuasive identification strategy the results must be treated with some caution.

Heckman's correction is very powerful as it provides a point estimate of the parameter of interest in the context of missing information. However Manski (forthcoming *inter alia*) believes that the distributional and exclusion restrictions invoked by this approach are too strong. In Manski's formation each woman in the sample is characterised by (y, x, z), where y is the dependent variable (the woman's permanent income), x is the explanatory variable (either her own or her husband's parents' incomes), z is a binary variable which takes the form of 1 if the woman is employed and 0 if she is not working. The regressions that are being estimated aim to reveal the relationship between y and x, ($E[y|x]$) for the full sample of women. The law of total probability implies that

$$E[y|x] = E[y|x, z=1]P(z=1|x) + E[y|x, z=0]P(z=0|x) \qquad (19)$$

In other words, the true parameter will be a weighted average of $E[y|x,z=1]$ and $E[y|x,z=0]$, where the weights depend on the proportion of women who are not working. There is no information which can identify $E[y|x,z=0]$ and therefore one cannot identify a point estimate of $E[y|x]$. Again, it is clear that the problem is exacerbated if a large proportion of the sample is not working.

Manski has developed a method to partially identify $E[y|x]$ by using the information that is available to derive bounds for the expected value. Minicozzi (2002) considers the intergenerational mobility of women using this approach. Initally she calculates the 'worst case' bounds, assuming no information is available about earnings for those not working full time. Minicozzi then narrows the bounds by making assumptions about the upper and lower bounds of earnings for individuals according to their characteristics and current work status. Unfortunately the bounds which result from these assumptions are still wide at .12 to .53. Having such wide bounds would make it very difficult to draw conclusions on the relative magnitudes of b and d, which is why I prefer to use Heckman's approach.

I present results which show the uncorrected regression coefficients for the sample of employed individuals and also results which correct for selection. Both are informative. The uncorrected estimates will show how parental income is related to the earnings of daughters and daughters-in-law with the current patterns of employment. However, the selectivity corrected results enable me to try to separate the influence of changes in the selection into employment from changes in assortative mating.

An additional way of exploring the influence of assortative mating versus changes in participation is to estimate some of the other parameters from the model for the full sample. To begin with, I measure the relationship between the educational attainments of partners, as a direct measure of s for the two cohorts. Of course, couples will match on broader measures of human capital than educational attainment, so this will not provide a perfect estimate of s. An advantage of using education is that it is observed for the full population.

An alternative approach recognises the fact that matching and intergenerational investments are both modified through human capital. If we return to the model, it is clear that it also yields strong predictions about the relationship between human capital and parental education.

$$H_{wi} = a + y_w \ln Y_{wi}^{parents} + e_{wi} \text{ where } y = p / p_H \tag{20}$$

and

$$H_{hi} = a + v_w \ln Y_{wi}^{parents} + e_{hi} \text{ where } v = s \frac{SD^{H_h}}{SD^{H_w}} \frac{p}{p_H} \tag{21}$$

Consequently the relative magnitudes of y and v will be informative about the extent of assortative mating. Again, this relies on the premise that the measures of educational attainment used here are good proxies for human capital.

3. Data

The data used in this paper is taken from the two British cohort studies, the National Child Development Study (NCDS) and British Cohort Study (BCS). The NCDS includes all individuals born in a week in March 1958 and the BCS includes all individuals born in a week in April 1970. The surveys are ongoing and so far detailed data has been collected about many aspects of the cohort members' lives at birth and ages 7, 11, 16, 23, 33 and 42 for the NCDS and at birth, ages 5, 10, 16 and 30 for the BCS.

The parental income data used here is taken from both surveys at age 16, and it is based on a weekly measure of the income category that parents fall in to. It is necessary to manipulate the data slightly to ensure full comparability, and more detail and further robustness checks can be found in Blanden et al (2005). As mentioned previously, the reliance on a single week's income as the explanatory variable is a concern, and comparisons over time reply on the assumption that the bias introduced is similar for both cohorts.

Information about the earnings of cohort members is available at age 33 in the NCDS (1958 cohort) and at age 30 in the BCS (1970 cohort). At the same time information was obtained about the cohort member's partner, for both married and cohabitating couples. I use the questions on partner's sex (I drop the few same sex couples in the sample), partner's age, employment status, education and net earnings. The education variable available for partners is not very satisfactory, we only know the age at which partners completed their full-time education, this is much less useful than variables detailing qualification attainment, as educational outcomes for individuals who left school at the same age are quite dispersed in the UK system, especially for those leaving school at 16.

The data I use on the partner were obtained as part of the cohort members' main interview, so partners were not necessarily involved in answering the questions about themselves[3]. To help us to understand more about how this may affect the accuracy of the data collection, we know both whether the cohort member's partner was present in the room while the questions were being asked and whether the partner helped to answer these questions. In about 80 to 90 percent of the cases where partners were present, they helped to answer the questions. However, there was a lot of variation by cohort and sex in the proportion of partners who were present when the questions were asked.

Female partners are more likely to be involved than male partners, although this difference narrowed across the cohorts. Around 50 percent of the wives/partners of NCDS men were present, while only 30 percent of BCS wives/partners were in the room. The female cohort members are less likely to have their partners present, with about 20 percent of them doing so in each cohort. I check the implications of these differences for my results.

The need for information on partners means that I drop some cohort members for whom this data is invalid. I discard observations with invalid information on partner's earnings and employment (e.g. partners are working but no earnings are reported for them), and also where the partner is self-employed.

In Table 1, I describe the main variables used in my samples, by cohort, sex and partnership status. The first thing to note is that rather more of the individuals from the earlier cohort have partners at the time of the survey. When the NCDS cohort is observed 78 percent of males have partners and 79 percent of females. For the BCS this has declined to 61 percent of males and 68 percent of females. There has clearly been a strong shift towards later partnership and this is compounded by the fact that the 1970 cohort is observed when three years younger than the 1958 cohort.

[3] A separate questionnaire was administrated to the cohort members in the NCDS at 33, but this information is not used here, as it was not also collected for the BCS.

Figures 1 and 2 provide a stark illustration of this point by graphing the age at which individuals moved in with their current partner (for those that have one) in the two surveys. It is obvious that BCS individuals are forming partnerships later, and that for many of those who do not currently have partners it is probably just a matter of time as the sample appears to be truncated. Those who are observed in partnerships in the BCS are likely to be those who have formed partnerships relatively early, implying that the selection into the sample of couples has probably changed between the cohorts. As with the selection into employment, it is hard to deal with this difficulty convincingly, but it is important to keep it in mind.

Table 1 demonstrates a number of other differences across the cohorts and between those individuals with partners compared to those without. In both datasets men with partners are more likely to be employed and have higher wages (there is a growing literature on understanding this married-man wage premium, for example, Korenman and Sanders, 1991) The pattern with respect to education has switched however; in the NCDS partnered men are more likely to have higher education, whereas in the BCS those who do not have partners are more likely to be highly educated. This is likely to be due to the fact that the BCS sample of those with partners will include more of those who formed partnerships relatively early.

For women, there is little difference in the education levels of those with and without partners in the NCDS, while in the BCS, those with partners are less likely to have either very low or very high education. Women with partners have lower earnings in both cohorts. It seems that this is related to different employment patterns. The overall employment rates are higher in the BCS than the NCDS for both groups. But the relatively small differences in employment rates mask larger differences for full and part-time work. Women with partners are much less likely to work full-time and much more likely to work part-time than single women, although this gap has closed somewhat between the two cohorts. Figures from the Labour Force Survey (tabulated in the Annual Abstract of Statistics, 2002 and 2004) show that the members of the cohort studies reflect a general trend: between 1992 and 2003 the employment rate of women aged between 25 and 34 increased from 64 percent to 71 percent. I also report full-time equivalent earnings for women; this closes part of the gender gap but by no means all of it.

The choice of full-time or part-time work is closely associated with the presence of children. Once again, there are marked differences between the cohorts with almost 70 percent of men with partners having children in the household in the NCDS compared with 60 percent in the BCS. For women with partners the proportion with children is 77 percent in the NCDS and 65 percent in the BCS. Many women without partners also have children in the household in both cohorts; this is 48 percent in the NCDS and 34 percent in the BCS.

A further difference between the cohorts is the proportion of couples who are legally married. In the NCDS this is 87 percent for men and 89 percent for women, in the BCS it is much smaller at 60 percent for men and 69 percent for women. This change is a potential worry as the degree of commitment in a cohabiting relationship has a considerable amount of variability. Ermisch and Francesconi (2000) explore patterns of cohabitation using data from the British Household Panel. The evidence that cohabitation is a very temporary state is mixed. Cohabiting unions do tend to be short with 70 percent lasting less than 3 years, but 62 percent of those who end their cohabitation are moving into marriage. There is a however a strong negative relationship between the age at which the cohabitation began and the chances of dissolution, this means that the relatively young sample in the BCS are more likely to have temporary cohabitations. This issue is explored further in the robustness checks.

4. Changes in Assortative Mating

I begin my empirical analysis by using the data to consider the extent of assortative mating directly. I measure the similarity of education levels within couples. As stated above, information available on partners' education is limited, so I am only able to present the association between the education-leaving age within couples. Results are presented in Tables 2A (for sons) and 2B (for daughters), which show cross-tabulations of the education levels of the cohort members and their partners.

It is immediately clear that school-leaving ages are heavily clustered around age 16 in the UK, which limits the power of this approach. Also, a general increase in educational attainment is plain. In almost half of all couples in the NCDS, both partners left school at or below age 16, while in the BCS, this is just below 40 percent.

There are several ways of using these cross tabulations to infer the extent of assortative mating on education levels. A simple (but potentially misleading) approach is to add up the value of the cells for which couples have the same education group, or where they have the same or adjacent education groups. Using this approach, it appears that there has been a rise in assortative mating for sons and a fall for daughters. The proportion in the samples marrying someone in the same education group (educational homogamy) rose from 60 percent in the NCDS to 64 percent in the BCS for sons while it fell from 57 percent to 54 percent for daughters.

As noted, the education levels of the cohorts have risen; individuals are now more likely to stay in school beyond age 16 and consequently the education distribution has become more dispersed. This means that if couples match randomly, we would expect to find fewer couples with the same education level in the BCS compared with the NCDS. The implications of this are shown by the figures in parentheses; these show the likelihood of each combination of education levels if partners' education levels are independent. Assortative matching on education is demonstrated by the fact that the actual probabilities are higher than these along the diagonal.

An alternative measure of assortative mating is generated by dividing the actual proportion of couples with the same education group by the expected random proportion. This approach reveals a rise in assortative mating. There are 1.409 times more NCDS sons with the same education group as their partner than would be predicted by random matching. In the BCS this number has risen to 1.805. For daughters there has also been a rise in assortative mating, although it is slightly smaller. The relative odds of the daughter being in the same education group as her partner has increased from 1.228 to 1.485 between the cohorts[4].

These results for the UK therefore show a rise in assortative mating by education group. This is similar to the results of similar exercises found in Pencavel (1998) and Mare (1991) for the US. Both Pencavel and Mare use data on young husbands and wives from the 1940 census onwards to consider the association of educational levels within couples. Mare takes care to use models which take account

[4] We would not necessarily expect men and women to follow exactly the same patterns, because women tend to marry men slightly older than themselves. We can think of men and women born at the same time as being part of slightly different (although overlapping) marriage markets.

of the changing distributions of education and finds evidence that part of the rise in homogamy can be explained by the falling gap between the age when young people leave education and the age of marriage.

Chan and Halpin (2003) use data from the General Household Surveys in 1973, 1986 and 1995 to consider educational matching within marriage in the UK, and compare this with data from a number of sources for Ireland. Like Mare, Chan and Halpin use log-linear models to account for the changes in overall educational distributions. Chan and Halpin find a decrease in educational assortative mating for the UK, although their data focuses on earlier cohorts than those considered here. The authors argue that that this may be explained by the rise in the gap between school leaving and first marriage from the 1970s onwards in the UK (meaning that individuals are less likely to marry their class-mates), but do not offer further evidence on this. It seems unlikely that the reversal of this trend over the 1990s is a result of a closing of the gap between education and marriage as although education has lengthened on average, Figures 1 and 2 show that partnership formation is also increasingly delayed.

The evidence on educational matching therefore suggests that assortative mating has increased somewhat across the cohorts. This suggests that we might expect to see an increasing relationship between parental income and the education and earnings of sons-in-law and daughters-in-laws.

5. Education and Parental Income

The next stage of my empirical analysis considers the relationship between educational attainment and parental income for the cohort members and their partners.

As I have shown in Section 2, if educational attainment is a proxy for human capital, comparing these relationships can provide additional information about the extent of assortative mating. To reiterate,

$$H_{wi} = a + y_{w} \ln Y_{wi}^{parents} + e_{wi} \quad \text{where } y = p / p_{H} \tag{22}$$

and

$$H_{hi} = a + v_{w} \ln Y_{hi}^{parents} + e_{hi} \text{ where } v = s \; \frac{p}{p_H} \frac{SD^{H_h}}{SD^{H_w}} \qquad (23)$$

So $\frac{v_{w}}{y_{w}} = s \; \frac{SD^{H_h}}{SD^{H_w}}$. If the relationship between parental income and education is similar for the cohort member and their partner, this implies that assortative mating is strong.

In Table 3 I use Probit models to estimate the relationship between age left education and parental income for the cohort members and their partners. I use two dependent variables; leaving school after age 16 (similar, but not identical, to staying on after the compulsory leaving age) and leaving education at age 20 or older (close to university participation). Here I measure the linear relationship between log parental income and these two outcomes, and I report the marginal effect of log income on the probability of the two outcomes.

To provide a comparison, I report the models for single cohort members first and then for couples. The first two panels of Table 3 report these relationships for single sons and daughters in the cohorts. There is no strong evidence of a rise in the relationship between family income and educational attainment for single sons or daughters. For sons and daughters in couples, the strengthening relationship between parental income and educational attainment which was observed by Blanden, Gregg and Machin (2005) is more apparent, with a strong rise in the impact of family income on higher education participation.

Results for children's partners indicate strong assortative mating, with strong relationships between parental income and partners' education levels for both sexes and in both cohorts. Notably, these relationships have not changed between the cohorts, suggesting no increase in assortative mating, in contrast to the evidence in the previous section on educational matching.

6. Results on Changes in Intergenerational Mobility

Intergenerational Mobility of Sons and Daughters in the UK

I begin my empirical analysis of earnings mobility by investigating the evidence on changes in individual intergenerational mobility for sons and daughters by partnership status. There are two motivations behind this exercise: the first is to understand more about intergenerational persistence for women and the second is to compare results for single individuals with those in couples.

Table 4 provides results for both the elasticity and partial correlation measures of intergenerational persistence. It is clear that there is a very strong rise in intergenerational mobility for sons; both singles and those in couples. The partial correlation between sons' earnings and parental income rises by .079 for single sons and .111 for sons with partners when comparing those born in 1970 with those born in 1958[5].

The level of intergenerational persistence for men is very similar whether they have partners or not. This is not the case for women. The correlation between women's earnings and their parental income at age 16 is considerably stronger for daughters who are single in their early 30s, in the BCS the partial correlation is .327 for single daughters and .181 for those in couples; this difference is statistically significant. This indicates that the presence of a husband in the household weakens the intergenerational link for daughters, perhaps because joint labour supply decisions mean that her own earnings are now more weakly tied to her own capabilities. What is interesting to see is whether the intergenerational link is strong between the daughter's husband and her parents; leading to a continued persistence in household income for daughters in couples.

The partial correlation measure of intergenerational persistence for single daughters shows a similar rise to that observed for sons. However due to the small sample sizes, the change is not statistically significant. There is essentially no change in intergenerational mobility at the individual level for daughters in couples.

There is a large difference between the *ß* coefficients and partial correlations for daughters. For both groups of daughters, the coefficients are considerably larger than the partial correlation; a feature not observed for sons. The contrast between these results shows the importance of adjusting for the changing variance of income for women. Equation (24) provides a reminder about the relationship between the elasticity and the partial correlation.

[5] The data used here differs slightly from that used in Blanden et al (2004), the sample sizes are not the same as individuals with invalid information on partner's earnings are dropped. Additionally net earnings are used as the dependent variable while gross earnings were used in the earlier work.

$$(\text{Corr}_{\ln Y^{\text{parent}}|\text{age},\, \ln Y^{\text{child}}|\text{age}}) = ß\left(\frac{SD^{\ln Y^{\text{parent}}|\text{age}}}{SD^{\ln Y^{\text{child}}|\text{age}}}\right) \tag{24}$$

Therefore, the reason that the partial correlation is lower than the elasticity for both cohorts is because there is a wide dispersion of earnings among daughters. It falls further for single women in the NCDS because the dispersion of earnings is very wide in this early cohort. As Table 1 showed, more single women in the NCDS have child care responsibilities and a higher proportion work part-time.

Taking the Table as a whole demonstrates that there is a large fall in intergenerational mobility observed for employed sons, but less evidence of this for employed daughters. In a later section, I shall assess how robust this conclusion is when I take account of endogenous participation.

Intergenerational Mobility for Couples

I now provide the substantive results of the paper, showing estimates of the intergenerational persistence of earnings for sons and daughters, their partners and for the couple as a whole (which I describe as family mobility). In Section 2, I described how the elasticity of couples' earnings with respect to parental income (defined as m) can be decomposed, to demonstrate the contribution from the earnings of the cohort member and those of their partner. For couples where both partners are working, $m = (1-s)b + sd$, where b is the elasticity between the child's earnings and parental income and d is the elasticity between the partner's earnings and parental income, and s is the share of earnings contributed by the partner. This decomposition makes it clear that a rise in the share of earnings contributed by the female partner will have implications for m. If we assume that parental income is more strongly associated with the child's earnings than that of the partner, an increase in the woman's share will result in a fall in m for sons and a rise in m for daughters.

The contribution of b and d to family mobility for all couples will also depend on the patterns of employment among couples. $m = (1-s)b + sd$ will only be the case for couples where both partners work, while $m = b$ if only the cohort member works, and $m = d$ if the partner is the only member of the couple working. Table 5 shows the employment patterns for the couples in my sample, and the share of income provided by partners when both work. This Table makes it clear that the proportion of households where the female partners work has increased, as has the share of household earnings contributed by women when they do work. As a result, the relationship between partners' earnings and parental income has become more important in determining the extent of intergenerational inequality for men, and less important in determining intergenerational persistence for women.

Table 6 provides results for b , d and m by cohort and sex. The results found in Table 4 for individual persistence are reiterated here. There is a strong rise in b and in the partial correlation for sons, but no rise in intergenerational persistence for daughters. The crucial results in this Table are for the relationships between parental income and partners' earnings. The results for d_w show that for daughters the relationship between partners' earnings and parental income is very strong, and also that it has increased significantly over time. The partial correlations show that the relationship between partners' earnings and daughters' parental incomes is stronger than that between parents and their daughters. This result is in line with others in the literature and suggests strong assortative mating. The partial correlations confirm the picture of a rise in d_w: it increases by .062 from .168 to .230 (a change which is significant at the 6 percent level).

The parent to daughter-in-law[6] relationship has been less frequently studied in the literature. For the NCDS, this lack of attention seems justified as there is no significant relationship between parents' incomes and the earnings of their daughters-in-law. However, this changes dramatically for the second cohort when the relationship between the parental income of the son and his partner's earnings are of the same magnitude as they are between the parental income of the daughter and her partner's earnings[7]. This is a very strong result, and implies that marriage is now an important way of generating persistence in economic status for men in a way which

[6] From now on son-in-law and daughter-in-law also refer to daughter's and son's cohabitees.

[7] In the data section, I discussed the accuracy of partners' earnings reports, as in many cases these are given by the cohort member rather than by the partner themselves. I have checked if results differ depending upon who reports the partner's earnings. There is one significant result. In the NCDS the partner-parents elasticity is stronger for sons' partners if the partner was not present when the earnings question was asked. This implies that men tend to over-estimate the similarity between their wives' earnings and their parental income. This implies that the NCDS sons' partner elasticities, which are very low, may even be over-estimates.

has long been considered to be the case for women. For example, the sociology study by Glenn, Ross and Tully (1974) describes female mobility entirely in terms of marriage and male mobility in terms of occupational change. This division clearly no longer holds.

The results for family mobility demonstrate that income persistence from parents to partners does contribute to intergenerational persistence. The rise in individual persistence observed for sons is magnified when his partner's earnings are added. The rise in the partial correlation is .111 for his own earnings, and an even larger .179 for his earnings and his partner's earnings combined. The influence of the rise in d is magnified by the fact that partners are contributing a larger share of income in the second cohort. The increase in d for the daughter's partner has also led to an increase in family income persistence for daughters, although to a much smaller extent than for sons. The partial correlation associated with m increases by a statistically significant .066. It is clear that partners' earnings make an important contribution to the intergenerational persistence of incomes across generations.

In Section 2, I showed that the persistence between partner's earnings and parental income will increase with assortative mating. Results from Table 6 make it clear that this relationship has indeed increased, for both men and women. However, in order to distinguish the influence of assortative mating from changes in participation it is important to investigate how selection bias is influencing the results, and it is to this issue which I now turn.

Changes in Female Participation and Family Characteristics

It is very clear that the changing participation behaviour of women may influence my results for the intergenerational persistence of employed daughters and daughters-in-law; a self-selected sub-sample of the full population. Table 1 showed a rise in the proportion of women employed in the BCS compared with the NCDS and an increase in the extent to which women work full-time when they do work. In addition, the first cohort is slightly older and more likely to have children. Consequently it may be that 'non-wage' factors are more important in their decision to work. As noted earlier, Mulligan and Rubinstein (2004, 2005) find that the correlation between employment and skills has strengthened considerably over time in the US. This evidence suggests that we might expect to find an increase in the extent of positive selection bias between the cohorts.

In Table 7, I present the relationship between participation and parental income for women (daughters and daughters-in-law). There is evidence for all women that the participation decision is increasingly determined by the income of parents (or parents-in-law). For the BCS at least, selection into the intergenerational mobility sample is endogenous. This is likely to affect the results for women's earnings mobility in Tables 4 and 6; with the implication that the BCS results may be over-estimated when women's earnings are the dependent variable.

In Table 8, I compare the estimated b coefficients reported in Tables 4 and 6 with selectivity-corrected results for all the regressions for which women's earnings is the dependent variable. The upper panel reiterates the uncorrected results. To recap, the uncorrected regressions show that there has been no significant change in the intergenerational relationship between daughters and their parents; while there has been a strong rise in the relationship between parental income and the earnings of daughters-in-law (this is even more pronounced when measured by the partial correlation, rather than the regression coefficient as shown here). In the lower panel parallel estimations are shown from using a Heckman selection model, where the model is identified using the cubic index of the predicted probability of employment from a Probit[8].

The magnitude of the selection corrections are consistent with the models of selection presented in Table 7. In every case for which Table 7 reported a significant positive association between parental income and women's participation, the intergenerational coefficient is substantially reduced when the selection into employment is taken into account. This means that the estimates for single daughters are substantially reduced in both cohorts, as are the estimates from the second cohort for daughters with partners and daughters-in-law. The increased selection in the BCS has implications for changes over time; the corrected point estimates indicate that b for women with partners has declined, while the magnitude of the rise in the relationship for daughters-in-law has more than halved.

[8] This probit includes parental income, own education, marital status and the number of children in the household. The results are robust to the specification used.

However, the other notable contrast between the two panels of Table 8 is a large increase in the size of the standard errors. This means that unfortunately it is impossible to distinguish between the corrected and uncorrected results. Indeed the results in the lower panel show no significant changes for women between the cohorts. The significant rise in the daughter-in-law's elasticity is wiped out in the corrected model, although the magnitude of the increase is .071. It is clear that the change in the daughter-in-law's elasticity is in large part a consequence of a change in the participation behaviour of women.

Another difference across the datasets is the increase in the proportion of women working full-time, which is stronger than the growth in overall employment. The proportion of employed daughters who work full-time increases between the cohorts from 56 to 72 percent while for sons' partners, full-time participation increased from 52 to 70 percent of those working. I have checked for changes in the relationship between full-time employment and parental income in the cohorts and find that there is a strong relationship between daughters-in-law's full-time work and parental income in the second cohort, while there is none for the NCDS. The other relationships for daughters are stable across cohorts conditional on the participation decision. As female partners who work part-time earn substantially less (approximately £500 a month in the BCS compared to £1200 for full-timers) this may provide an additional explanation for the rise in the relationship between parental income and daughter-in-law's earnings. In order to investigate this properly, a more complex model of participation would be required, and I leave this to further research.

As stressed in my data section, there are other differences between the two cohorts which may have an impact upon my results. Table 9 tries to address the consequences of the increase in cohabitation. It is difficult to believe that the switch to more informal partnerships could be responsible for the growing importance of partners in intergenerational mechanisms - if anything, we might expect the effect to work the other way. Nonetheless, this Table repeats the analysis of Table 6 just for those couples who are legally married. There are some slight differences between married and cohabiting couples, but in general the patterns are very similar: there has been a strong rise in family income persistence for sons and a smaller rise for daughters.

One result which does stand out is that the correlation between the daughter-in-law's earnings and her husband's parental income is weaker for sons who are married in the second cohort. The difference between d_h for married and cohabiting couples is significant at the 11 percent level. Ceteris paribus, the growth in cohabitation has contributed towards the increased importance of sons' partners in leading to intergenerational persistence; however this is likely to be related to changes in participation as cohabiting partners are more likely to work and to work full-time.

7. Discussion

The main empirical findings in this paper are as follows:

- Evidence on assortative mating by education suggests that this has risen between the cohorts.
- However, there has been no rise in the relationship between parental income and partners' education; this is strong for both men and women and in both cohorts.
- Intergenerational persistence has not increased as much for the sample of employed daughters as it has for the sample of employed sons. Indeed when an adjustment is made for the selection into employment, there is a decline in persistence for daughters with partners, although this is not significant.
- For daughters, the relationship between partners' earnings and parental income is strong in both periods and it also has increased slightly over time.
- For sons, there has been a very sharp rise in the relationship between their partners' earnings and their parental income. However, this is partially explained by the stronger association between parental income and daughters-in-law's participation in the second cohort.
- As it is coupled with a rise in the number of female partners working and an increase in the share of income they contribute, the rise in the daughter-in-law's elasticity leads to a large increase in the persistence of family income across generations for sons.

In order to interpret my results I return to the model presented in Section 2. In this model I discussed how mobility can be interpreted in terms of the structural parameters of the intergenerational mobility model. For example, $b_w = g_w p / p_H$, where b_w is earnings persistence for daughters, g_w is the return to human capital for women, p is the weight placed on the daughter's income in the parental utility function and p_H is the cost of human capital. Partner mobility for sons (i.e. the elasticity for daughters-in-law) is $d_h = g_w \frac{p}{p_H} s \frac{SD^{H_w}}{SD^{H_h}}$, where s is the correlation between the human capital of couples and SD^H measures the dispersion of human capital.

One of the implications of this framework is that an increase in assortative mating will lead to an increase in d relative to b, ceteris paribus. For daughters there is certainly an increase in d_w relative to b_w; while b_w is flat or declining, d_w has increased somewhat. This suggests a rise in assortative mating.

For sons and daughter-in-laws the increase in d_h is very strong compared to b_h in the uncorrected sample. However, this change is much lower when the change in endogenous participation is taken into account. Indeed it appears that d_h has fallen relative to b_h when results are corrected for endogenous participation, although this difference is not significant. This implies that changes in the selection into employment are behind the increasing relationship between the earnings of daughters-in-law and their husbands' parental income, rather than changes in the pattern of marital matching.

The results so far indicate that any increase in assortative mating which has occurred through the 1990s in the UK is fairly weak. This is confirmed by the evidence using the education data. While matching on education has increased somewhat when measured directly, there is no increase in the association between parental income and partner's education level.

There is one final piece of evidence to add to the jigsaw. An initially puzzling result is the smaller rise (or even fall) in the intergenerational persistence of daughters when compared with the strong increase in the persistence of income for sons. Within the setup of the simple model, this must be accounted for by a fall in the relative return to income for women. A decline in the return to human capital for women would also explain the smaller selectivity corrected change in persistence for daughters-in-law. The small increase in assortative mating is counteracted by the relative decline in the returns to education for women.

The return to education in the model refers to a permanent return rather than the one-off return observed in an earnings regression. Nonetheless an investigation of the return to education does show a fall in the earnings return for women. This is found in both simple regressions and in selectivity corrected models, while there is no evidence that there has been a fall in returns for men with partners[9].

The result that earnings differentials by education level have declined for women is also found in these cohorts by Dearden, Goodman and Saunders (2003). Evidence from the Labour Force Survey shows no such change, with returns to education for 25-40 year old women extremely constant across the 1990s[10]. This implies that the fall in the intergenerational mobility for women may be cohort-specific, perhaps due to particular life-cycle effects and the three year age gap between the data collection. It will be interesting to observe if this persists for the next wave of data.

Taken together, my results point to a fairly modest increase in the extent of assortative mating in the UK through the 1990s. This is confirmed by the evidence using the education data. While matching on education has increased somewhat there is no increase in the association between parental income and partner's education level. More important has been the growing association between potential wages, and by extension family background, in women's participation decisions. This has been primarily responsible for the new-found importance of wives' earnings in contributing to the intergenerational persistence of sons incomes.

8. Conclusion

[9] Although it should be pointed out that women's returns to education are much higher than men's returns for both cohorts.
[10] Thanks to Steve McIntosh for supplying these results.

This paper makes a number of contributions to the literature on intergenerational mobility in the UK. The first is to study changes in intergenerational mobility for sons and daughters by comparing the 1958 and 1970 cohorts. I find that trends for women do not show the large falls in mobility found for sons, and that when selection is taken into account mobility for women may have even increased between the cohorts. The second is to provide an up-to-date analysis of the contribution of assortative mating to intergenerational persistence. Previous studies have focused only on the contribution of women's partners. This is clearly misplaced. For the cohort born in 1970, the wives and partners of sons are making a substantial contribution to the intergenerational persistence of incomes across families. Marriage is now as strong a mechanism for securing economic and social advantage for men as it is for women, and this change has led to an additional fall in the family income mobility of sons.

The evidence presented here suggests that partnership formation magnifies the changes in individual earnings persistence found in previous work (Blanden et al. 2004), leading to even greater intergenerational inequalities in family incomes. There is evidence that this is partly due to a small rise in assortative mating, while also a consequence of the growing influence of potential wages and family background on participation decisions. These changes may also have implications for cross-sectional household income inequality. Previous research has indicated that women's earnings have an equalising effect on family incomes, (Cancian and Reed, 1996, for the US and Harkness, Machin and Waldfogel, 1997, for the UK). However, my evidence suggests that this may be reversing. Not only are wives earnings more strongly linked with their husbands' family backgrounds, but partners of men from well-off backgrounds are more likely to work, and given participation, are likely to work longer hours. An investigation of these trends for household income inequality is firmly on the agenda for further research.

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Figure 1: Age Formed Current Partnership, Males

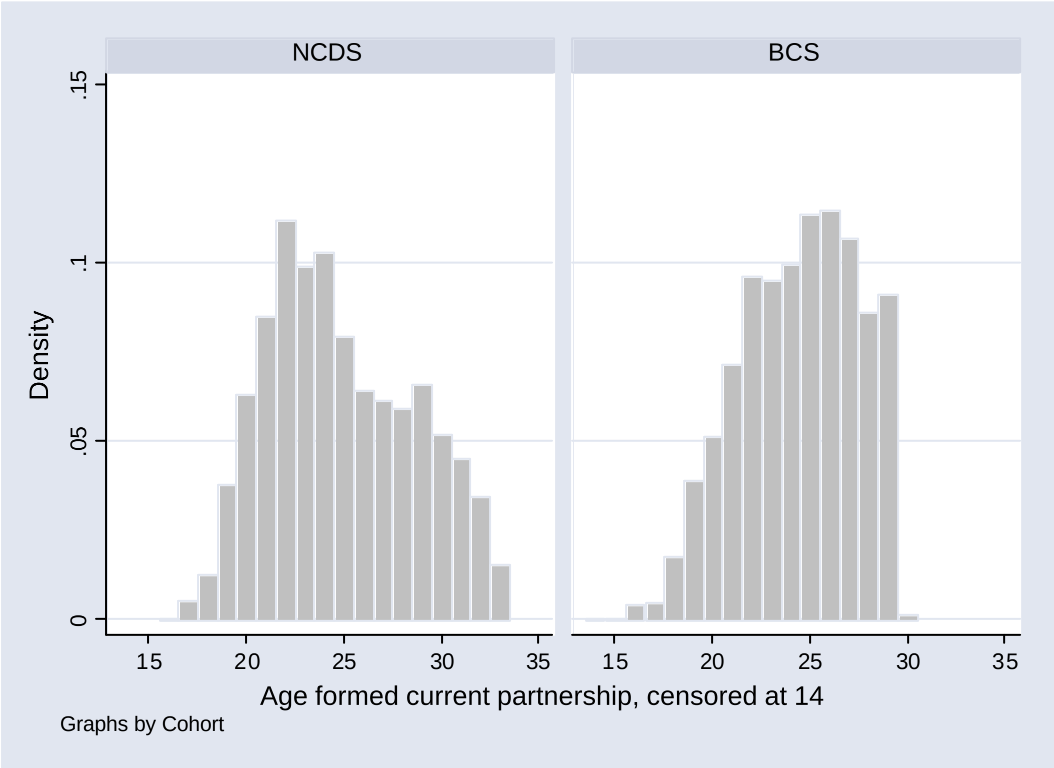

Figure 2: Age Formed Current Partnership, Females

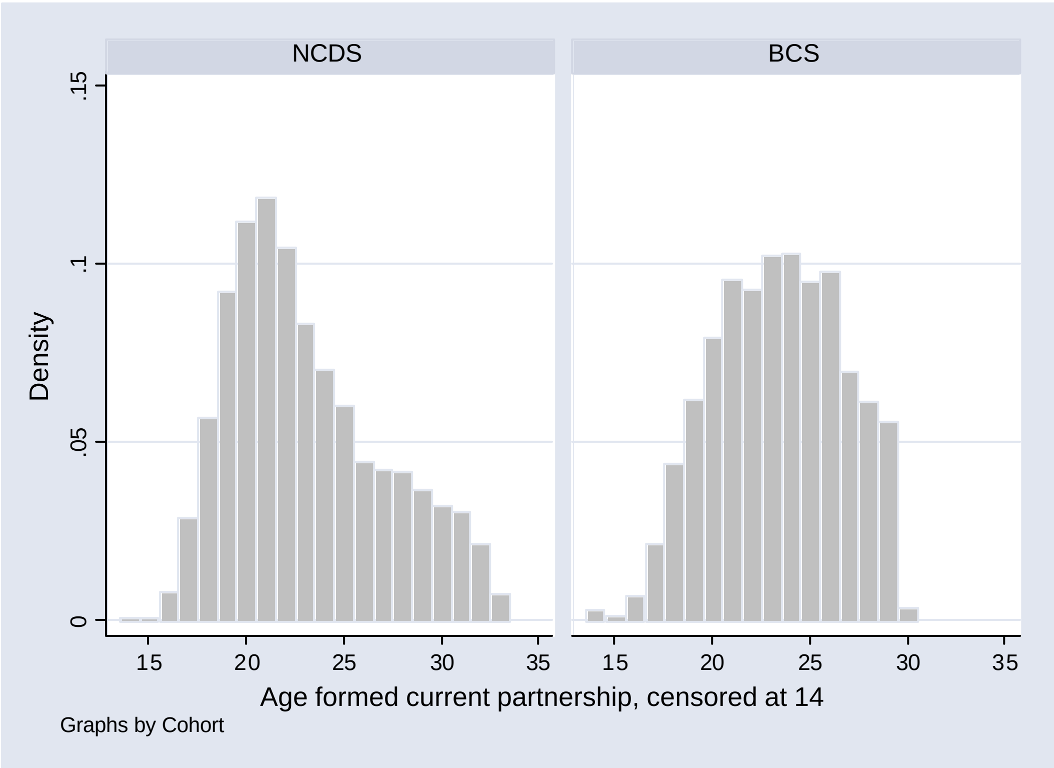

Table 1: Characteristics of Samples by Partnership Status

	Sons			
	NCDS		**BCS**	
	No Partner	Has Partner	No Partner	Has Partner
< GCSE A-C	.324	.245	.249	.259
GCSE A-C	.249	.231	.190	.201
A level	.291	.360	.309	.298
Degree	.136	.164	.254	.242
Proportion employed	.738	.918	.835	.933
Monthly Earnings	1140 (612)	1342 (667)	1319 (987)	1474 (990)
Monthly Parental Income	1334 (496)	1384 (492)	1460 (762)	1457 (725)
Married	-	.873	-	.606
Has kids in the household	.075	.682	.200	.609
Sample size	507	1783	853	1322
	Daughters			
	NCDS		**BCS**	
	No Partner	Has Partner	No Partner	Has Partner
< GCSE A-C	.318	.264	.297	.267
GCSE A-C	.302	.333	.203	.239
A level	.279	.293	.259	.279
Degree	.101	.111	.241	.215
Proportion employed	.665	.641	.722	.747
Proportion full-time	.352	.206	.436	.367
Proportion part-time	.313	.435	.286	.380
Monthly Earnings	616 (429)	806 (912)	1057 (577)	909 (650)
Full time equivalent earnings	911 (987)	862 (432)	1137 (565)	1334 (495)
Monthly Parental Income	1364 (552)	1385 (523)	1419 (756)	1431 (684)
Married	-	.898	-	.689
Has kids in the household	.481	.776	.344	.647
Sample size	516	1889	766	1651

Notes:

1. Standard deviations are in parentheses.
2. Full time equivalent earnings are defined as (weekly wage/hours)*40.
3. Earnings and income are expressed in 2000 pounds

Table 2A: Assortative Matching on Age Left Full Time Education, Sons

	NCDS sons				
	Age partner left full-time education				
Age cohort Member left education	16 or younger	17	18 or19	20 or above	All
16 or younger	.473 (.372)	.060 (.037)	.069 (.088)	.021 (.090)	.624
17	.052 (.058)	.017 (.012)	.018 (.014)	.010 (.014)	.098
18 or 19	.047 (.081)	.025 (.016)	.029 (.019)	.035 (.020)	.136
20 or above	.025 (.086)	.016 (.017)	.025 (.020)	.077 (.020)	.144
All	.596	.119	.141	.144	
	BCS sons				
	Age partner left full-time education				
Age cohort Member left education	16 or younger	17	18 or19	20 or above	All
16 or younger	.477 (.276)	.029 (.061)	.036 (.112)	.030 (.123)	.572
17	.059 (.045)	.012 (.009)	.013 (.018)	.009 (.020)	.094
18 or 19	.072 (.073)	.008 (.016)	.033 (.029)	.037 (.032)	.151
20 or above	.035 (.089)	.009 (.019)	.024 (.036)	.116 (.040)	.184
All	.484	.106	.195	.215	

Notes:

1. In parentheses are the probabilities of each outcome if the partners' education levels are independent.
2. Assortative mating index based on couples with the same education level is 1.409 for the NCDS and 1.805 for the BCS.
3. Assortative mating index based on couples with the same or adjacent education level is 1.379 for the NCDS and 1.439 for the BCS.
4. Sample sizes are 1546 for the NCDS and 1305 for the BCS.

Table 2B: Assortative Matching on Age Left Full Time Education, Daughters

	NCDS daughters				
	Age partner left full-time education				
Age cohort Member left education	16 or younger	17	18 or19	20 or above	All
16 or younger	.390 (.405)	.056 (.036)	.102 (.063)	.051 (.094)	.598
17	.044 (.083)	.033 (.007)	.027 (.013)	.018 (.019)	.122
18 or 19	.060 (.110)	.017 (.010)	.050 (.017)	.035 (.025)	.162
20 or above	.014 (.081)	.008 (.007)	.020 (.013)	.077 (.019)	.119
All	.677	.061	.106	.157	
	BCS daughters				
	Age partner left full-time education				
Age cohort Member left education	16 or younger	17	18 or19	20 or above	All
16 or younger	.375 (.295)	.024 (.032)	.051 (.063)	.027 (.088)	.478
17	.086 (.074)	.016 (.008)	.010 (.016)	.007 (.022)	.120
18 or 19	.112 (.130)	.015 (.014)	.040 (.027)	.039 (.039)	.210
20 or above	.043 (.118)	.011 (.013)	.027 (.025)	.111 (.035)	.192
All	.617	.067	.132	.185	

Notes:

1. In parentheses are the probabilities of each outcome if the partners' education levels are independent.
2. Assortative mating index based on couples with the same education level is 1.288 for the NCDS and 1.485 for the BCS.
3. Assortative mating index based on couples with the same or adjacent education level is 1.193 for the NCDS and 1.315 for the BCS.
4. Sample sizes are 1682 for the NCDS and 1622 for the BCS.

Table 3: Relationships between Education and Parental Income

	NCDS	BCS	Change	Sample
	Single Sons			
Probit marginal effect of income on leaving after 16	.259 (.062)	.322 (.038)	.063 (.073)	NCDS: 441 BCS: 852
Probit marginal effect of income on leaving at age 20 or later	.131 (.045)	.160 (.027)	.029 (.052)	
	Single Daughters			
Probit marginal effect of income on leaving after 16	.257 (.058)	.272 (.039)	.015 (.069)	NCDS: 460 BCS: 766
Probit marginal effect of income on leaving at age 20 or later	.152 (.036)	.168 (.027)	.016 (.045)	
	Sons with Partners			
Probit marginal effect of income on son leaving after 16	.247 (.034)	.316 (.032)	.069 (.046)	NCDS: 1559 BCS: 1322
Probit marginal effect of income on son leaving at age 20 or later	.149 (.024)	.222 (.023)	.073 (.033)	
Probit marginal effect of income on partner leaving after 16	.218 (.032)	.202 (.031)	-.016 (.044)	NCDS: 1765 BCS: 1305
Probit marginal effect of income on partner leaving at age 20 or later	.174 (.023)	.199 (.024)	.025 (.033)	
	Daughters with Partners			
Probit marginal effect of income on daughter leaving after 16	.240 (.032)	.267 (.029)	.027 (.043)	NCDS: 1701 BCS: 1651
Probit marginal effect of income on daughter leaving at age 20 or later	.140 (.020)	.254 (.021)	.114 (.029)	
Probit marginal effect of income on partner leaving after 16	.211 (.029)	.197 (.027)	-.014 (.040)	NCDS: 1864 BCS: 1623
Probit marginal effect of income on partner leaving at age 20 or later	.144 (.022)	.162 (.021)	.018 (.030)	

Note:

The coefficients shown are for separate Probit models of log parental education at age 16 on leaving school after age 16 (close to staying on) and of log parental education at age 16 on leaving education after age 20 (close to university participation).

Table 4: Estimates of Earnings Mobility by Gender and Partnership Status

		Single Sons					
Dependent variable		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Son's earnings	ß	.191 (.056)	.249 (.043)	.178 (.052)	.257 (.044)	.079 (.068)	NCDS: 374 BCS: 712
		Sons with Partners					
		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Son's earnings	ß	.186 (.025)	.270 (.029)	.176 (.024)	.267 (.030)	.111 (.039)	NCDS: 1637 BCS: 1234
		Single Daughters					
		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Daughter's earnings	ß	.429 (.102)	.449 (.061)	.243 (.057)	.327 (.044)	.083 (.072)	NCDS: 343 BCS: 553
		Daughters with Partners					
		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Daughter's earnings	ß	.287 (.053)	.262 (.040)	.154 (.028)	.186 (.028)	.032 (.040)	NCDS: 1211 BCS: 1233

Note:
Estimates are from regressions of log earnings at age 33/30 on log parental income at age 16 with controls for parental age and age-squared.

Table 5: Household Composition and Earnings Shares

Proportion of Households	**NCDS Sons**	**BCS Sons**	**NCDS Daughters**	**BCS Daughters**
Partner, only self works	.364	.245	.045	.034
Partner, both works	.608	.732	.646	.751
Partner, only partner works	.027	.024	.308	.215
Sample size	1683	1264	1751	1571
Share of partners earnings when both work	.317	.386	.680	.609

Table 6: Household Earnings Mobility for those with Partners

		Sons with Partners					
Dependent variable		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Sons' earnings	$ß_h$	.186 (.025)	.270 (.029)	.176 (.024)	.267 (.030)	.111 (.039)	NCDS: 1637 BCS: 1234
Partners' earnings	d_h	.097 (.056)	.306 (.037)	.054 (.031)	.233 (.028)	.179 (.042)	NCDS: 1070 BCS: 955
Couples' earnings	μ_h	.174 (.033)	.342 (.031)	.132 (.025)	.311 (.028)	.179 (.038)	NCDS: 1683 BCS: 1264
		Daughters with Partners					
		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Daughters' earnings	$ß_w$	.287 (.053)	.266 (.040)	.154 (.028)	.186 (.028)	.032 (.040)	NCDS: 1211 BCS: 1233
Partners' earnings	d_w	.206 (.027)	.239 (.026)	.168 (.022)	.230 (.025)	.062 (.033)	NCDS: 1672 BCS: 1518
Couples' earnings	μ_w	.252 (.030)	.302 (.028)	.182 (.022)	.252 (.023)	.070 (.032)	NCDS: 1751 BCS: 1571

Notes:

1. Estimates are from regressions of log earnings at age 33/30 on log parental income at age 16.
2. Partners age and age-squared are also added to the regression of partners earnings.

Table 7: Parental Income and Participation

	NCDS	BCS	Change	Sample Size
	Single Daughters			
Probit marginal effect of income on employment	.158 (.052)	.223 (.034)	.065 (.062)	NCDS: 516 BCS: 766
	Daughters with Partners			
Probit marginal effect of income on daughter's employment	.015 (.028)	.094 (.024)	.079 (.036)	NCDS: 1889 BCS: 1650
	Sons' Partners			
Probit marginal effect of income on partner's employment	-.009 (.030)	.084 (.027)	.093 (.040)	NCDS: 1783 BCS: 1322

Table 8: The Earnings Mobility of Women – Correcting for Endogenous Participation

Results for Employed Samples (as Tables 6.4 and 6.6)					
Dependent variable		**NCDS**	**BCS**	**Change**	**Sample Size**
		Single Daughters			
Daughters' earnings	*ß*	.429 (.102)	.449 (.061)	.020 (.119)	NCDS: 343 BCS: 553
		Daughters with Partners			
Daughters' earnings	$ß_w$	.287 (.053)	.266 (.040)	-.021 (.066)	NCDS: 1211 BCS: 1233
		Sons' Partners			
Sons' partners' earnings	d_h	.174 (.033)	.343 (.031)	.169 (.045)	NCDS: 1070 BCS: 955

Selectivity Corrected Results					
Dependent variable		**NCDS**	**BCS**	**Change**	**Sample Size**
		Single Daughters			
Daughters' earnings	*ß*	.130 (.167)	.092 (.110)	-.038 (.167)	NCDS: 512 BCS: 766
		Daughters with Partners			
Daughters' earnings	$ß_w$	.250 (.083)	.132 (.070)	-.118 (.110)	NCDS: 1873 BCS: 1646
		Sons' Partners			
Sons' partners' earnings	d_h	.090 (.091)	.161 (.072)	.071 (.116)	NCDS: 1757 BCS: 1319

Notes:

1. Parameters are estimated using a two-step Heckman correction for selectivity.
2. The participation equation is identified using a cubic index of employment probability where the employment probability is modelled as a function of the number of children in five age groups, marital status, parents' (or parents-in-law's) income.

Table 9: Estimates of Earnings Mobility for Cohort Members and Their Households, Married Sample

		Sons					
Dependent variable		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample Size
Sons' earnings	$ß_h$	.200 (.027)	.278 (.038)	.184 (.025)	.298 (.040)	.114 (.047)	NCDS: 1420 BCS: 753
Partners' earnings	d_h	.098 (.061)	.271 (.051)	.053 (.033)	.195 (.037)	.142 (.049)	NCDS: 896 BCS: 554
Couples' earnings	μ_h	.196 (.034)	.325 (.039)	.150 (.026)	.310 (.038)	.160 (.045)	NCDS: 1452 BCS0: 768
		Daughters					
		Coefficient		Partial Correlation			
		1958 Cohort	1970 Cohort	1958 Cohort	1970 Cohort	Change	Sample
Daughters' earnings	$ß_w$	.252 (.055)	.222 (.053)	.136 (.030)	.149 (.035)	.013 (.046)	NCDS: 1066 BCS: 832
Partners' earnings	d_w	.209 (.028)	.243 (.033)	.171 (.023)	.234 (.032)	.063 (.040)	NCDS: 1508 BCS: 1062
Couples' earnings	μ_w	.243 (.032)	.264 (.033)	.177 (.225)	.225 (.028)	.048 (.037)	NCDS: 1571 BCS: 1089

Note:
Estimates are from regressions of log earnings at age 33/30 on log parental income at age 16.