DOES A 'TEEN-BIRTH' HAVE LONGER-TERM IMPACTS ON THE MOTHER?

SUGGESTIVE EVIDENCE FROM THE BRITISH HOUSEHOLD PANEL STUDY

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ISER Working Papers
Number 2003-32

Institute for Social and Economic Research

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The support of both the Economic and Social Research Council (ESRC) and the University of Essex is gratefully acknowledged. The work reported in this paper is part of the scientific programme of the Institute for Social and Economic Research.

Acknowledgement:

Readers wishing to cite this document are asked to use the following form of words:

Ermisch, John, E. (2003) 'Does a 'Teen-birth' have Longer-term Impacts on the Mother? Suggestive Evidence from the British Household Panel Study', ***Working Papers of the Institute for Social and Economic Research*****, paper 2003-32. Colchester: University of Essex.**

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ABSTRACT

The paper studies associations between a woman's age at becoming a mother and subsequent 'outcomes', such as her living standard, when she is aged 30-51. The data come from the British Household Panel Survey over the years 1991-2001. The analysis suggests that having a teen-birth, particularly when aged under 18, constrains a woman's opportunities in the 'marriage market' in the sense that she finds it more difficult to find and retain a partner, and she partners with more unemployment-prone and lower earning men. Teenage mothers are much less likely to be a homeowner later in life, and her living standard, as measured by equivalent household income, is about 20% lower.

NON-TECHNICAL SUMMARY

The paper studies associations between a woman's age at becoming a mother and subsequent 'outcomes', such as household income, when she is aged 30-51. The data come from the British Household Panel Survey (BHPS) over the years 1991-2001. The advantages of these data are that they allow measurement of many outcomes, some of which have been difficult to measure in other British studies, such as 'living standards' (e.g. equivalent household income), and they provide multiple observations of these for each woman (up to 11 years). The disadvantage is that the family background variables that can be used as 'controls' are very limited, namely father's occupation at age 14 and whether or not the woman came from a one-parent family. Furthermore, we must assume that a woman's age of motherhood is not correlated with unobserved influences on the subsequent outcomes that are studied in order to interpret the estimated associations as causal effects

This analysis in this paper complements that in ISER Working Paper 2003-28, which studied the impact of teenage motherhood on 'outcomes' at age 30 using the British Cohort Study 1970 (BCS70). That study was able to use an 'instrumental variable' method that obtained estimates of causal effects under much weaker assumptions than used in this paper. The summary focuses on the results for outcomes for which the BCS70 analysis suggests that the estimate under the assumptions that must be used in this paper are not 'badly biased', in the sense that its 95% confidence interval lies fully within the confidence interval of the estimate using the instrumental variable method. The analysis focuses on women who became mothers by 1991 and who were born in 1950 or later, and we measure outcomes at ages 30-51.

The results indicate that women having a teen-birth are less likely to be living with a partner in their thirties and forties, particularly if they started childbearing before their 18th birthday. This association appears to reflect mainly the fact that women starting childbearing as a teenager are much less likely to be married at the time of their first birth. For women who have a partner in their thirties and forties, teen-mothers' partners are less likely to have a job, and if he has one, his pay is much lower, particularly if she became a mother before her 18th birthday. These results suggest that having a teen-birth, particularly when aged under 18, constrains a woman's opportunities in the 'marriage market' in the sense that she finds it more difficult to find and retain a partner, and she partners with more unemployment-prone and lower earning men.

Perhaps related to the above associations with a teen-birth, teenage mothers are much less likely to be a homeowner later in life, and her living standard, as measured by equivalent household income, is about 20% lower. Her probability of being 'poor' (in the bottom quartile of the equivalent income distribution) is also much higher. Finally, and probably related to the association of a teen-birth and her living standards, she is much more likely to suffer from common mental illness in her thirties and forties.

There is evidence that women starting childbearing in their early twenties are also less likely to be a homeowner and are more likely to have a lower living standard and to suffer from poorer mental health compared with women becoming mothers at ages 24 and older. But teen-mothers are nevertheless even more disadvantaged in these respects than early-twenties' mothers.

Does a 'Teen-birth' have Longer-term Impacts on the Mother?

There is considerable concern that having a child as a teenager, or more generally, earlier in life, may have longer-term consequences for the mother in terms of her earnings and standard of living, and of course these also entail consequences for the children living with her. These are usually thought to arise because having a child as a teenager disrupts her human capital investment, by causing her to curtail her formal education and by keeping her out of employment for a time, thereby depriving her of valuable work experience. It is difficult to measure these consequences because we do not know what the woman would have done if she did not have a child as a teenager. A comparison of teenage mothers with women starting childbearing later will usually not identify these consequences because the women who became teenage mothers may have had different 'outcomes' (e.g. household income at a later age) anyway, even if they had not given birth as a teenager. Unfortunately, the current analysis must use this comparison, and furthermore there is very limited information about other family background factors that may both influence a woman's age at first birth and outcomes later in life.[1] The advantage of the data used in the present paper is that they allow measurement of many outcomes, some of which have been difficult to measure in other British studies, such as 'living standards' (e.g. equivalent household income), and they provide multiple observations of these for each woman (up to 11 years). Furthermore, we can examine the associations between teenage childbearing and later outcomes for different birth cohorts. But it should be stressed that these associations can only be interpreted as causal 'effects' or 'impacts' under strong conditions that are not likely to be satisfied for many of the outcomes studied (see Ermisch and Pevalin (2003b).

[1] See Ermisch and Pevalin (2003b), which uses pregnancy history information to derive 'better' estimators of the impact of teenage childbearing.

1. Data

The data for this analysis come from the British Household Panel Survey (BHPS) over the years 1991-2001. Retrospective partnership and childbearing histories were collected from panel members in the 1992 wave of the BHPS, and these are used to determine the age of a woman's first birth and her marital status at that time. We can also derive from the BHPS two family background factors that are associated with a woman's age at first birth (see Ermisch and Pevalin 2003a) and also with at least some of the later life outcomes that are investigated. These are the father's occupation when the woman was aged 14 (from the 1991 wave) and whether or not she spent her entire childhood with both parents (from the 1996 wave). Father's occupation is represented by the Hope-Goldthorpe (prestige) score of the father's occupation. This score is strongly correlated with the earnings in that occupation.[2]

Thus, these data have very limited information on 'control variables' that may affect both outcomes and the age at first birth, but they possess the advantage of having multiple observations on outcomes for each woman (up to 11 years). Conditional on the exogeneity of age-at-first-birth, this produces more efficient estimates of the longer-term impacts of the timing of motherhood compared to observing outcomes at one point (e.g. such as in Ermisch and Pevalin 2003b and many other studies). Unfortunately, we are unable to relax the exogeneity assumption and use the instrumental variable estimator employed by Ermisch and Pevalin (2003b) because of the absence of pregnancy history information. Their analysis suggests that the estimator that we do use, which compares women with different ages at first birth, conditional on the two background variables, is likely to be biased toward overstating

[2] In the BHPS, the correlation coefficient between gross monthly earnings and the Hope-Goldthorpe score was 0.7 during the 1990s.

adverse impacts of early childbearing, and this should be taken into account in interpreting the estimates.

We focus on women who became mothers by 1991 and who were born in 1950 or later, and we measure outcomes at ages 30 and above. In these data, 5.7% of the 1,329 women had a birth before their 18th birthday (i.e. 76 women), 13.1% had a first birth when aged 18-19 (174 women), thereby implying 18.8% had a teen-birth.[3] In some of the analyses, we split the sample into three sets of birth cohorts: 1950-55, 1956-61 and 1962-71. As the panel data covers 1991-2001, outcomes for the first cohort span the ages 36-51, outcomes for the second cohort span ages 30-45, and outcomes for the third span ages 30-39. Some of the analyses split the sample by quartiles of the father's Hope-Goldthorpe score.

The analyses divide the mother's age at first birth into four intervals: 15-19, 20-21, 22-23 and 24 or older. It was ascertained that, with three exceptions, the association of a teen-birth with outcomes did not differ significantly between women giving birth when aged under 18 and women with a first birth when aged 18-19. These exceptions are noted below when relevant. Table 1 shows the means of the continuous outcome measures computed using all woman-year observations for the entire sample (born in 1950 or later and aged 30 or older) by these age-at-motherhood groups, and Table 2 shows the means for dichotomous outcomes. The first of these is the logarithm of 'equivalent household income', which is defined here as a woman's monthly household income divided by the square root of household size. This can be viewed as a general measure of a woman's economic welfare, which has been difficult to measure in other British studies of the impact of teenage childbearing. It is

[3] In terms of the 11,130 woman-year observations, 4.7% are from women having a birth aged less than 18, and 11% when aged 18-19.

lowest for women who became a teenage mother and highest for those who started childbearing at age 24 or older.

There may be more interest in the chances of a woman being 'poor' than in average differences in equivalent household income. Defining poverty as being in the bottom quartile of the equivalent household income distribution, Table 2 shows that percentage of women living in poverty falls as their age at first birth rises. The percent poor is twice as large for teenage mothers as for those becoming mothers beyond their 24^{th} birthday. Another indicator of poverty is the percentage of women receiving Income Support (IS), which is the main means-tested welfare benefit in Britain. Compared to women having their first child when aged 24 or older, the percentage receiving IS is nearly three times higher for teenage mothers. More generally this percentage falls with age at first birth.

Table 2 shows that the percentage of women in employment is much lower for women who became teenage mothers. Conditional on having a job, there is little difference in a woman's monthly pay between teenage mothers and those having their first birth when aged 20-21, and then it rises with age at first birth (Table 1). Table 3 shows that women who have their first child later tend to have higher educational qualifications (here the number of observations refers to the number of women).

The percentage of women living with a partner is lowest for those who became teenage mothers and highest for women who became mothers at age 24 or older (Table 2). Among women with a partner, the percentage with a partner in employment rises with age at first birth. Table 1 shows that, among women with an employed partner, there is a clear gradient in their partner's monthly pay, rising with age at first birth.

Whether or not a person is an owner-occupier is indicative of more choice in housing consumption, thereby tending to increase living standards, and also of wealth accumulation, or at least the potential for it. Table 2 shows a large and steady increase in the percentage of women who are living in owner-occupied housing as the age at first birth increases from a woman's teens to her mid-twenties.

The 12-item General Health Questionnaire (GHQ) can be used as an indicator of minor psychiatric morbidity, or 'common mental illness' (e.g. see Pevalin and Goldberg 2003). The 12 subjective indicators are: (i) loss of concentration; (ii) loss of sleep; (iii) playing a useful role; (iv) capable of making decisions; (v) constantly under strain; (vi) problem overcoming difficulties; (vii) enjoy day-to-day activities; (viii) ability to face problems; (ix) unhappy or depressed; (x) losing confidence; (xi) believe in self-worth; (xii) general happiness. Responses to these in the BHPS are obtained from a self-completion questionnaire. In the Likert version of the GHQ score, each indicator is measured over a scale that runs from 0 to 3, implying a range of 0-36, with higher scores indicating poorer mental health. Despite the fact that these measures are taken long after starting childbearing in most cases, the mean GHQ score falls (mental health improves) with later ages of first birth (Table 1).

An alternative scale in the health literature is known as *caseness* (see Cox *et al*. 1994). It re-codes values of 0 and 1 on the individual indicators above to zero, and values of 2 and 3 to one, and sums over all indicators to give a new scale ranging from 0-12. The scores on this scale are grouped into four categories in the last four rows of Table 2. The percentage of women reporting a high GHQ score declines as the age at first birth increases, while those reporting low scores increases, again suggesting better mental health for later ages at first birth.

2. Econometric analysis

Overall it appears that women who became mothers earlier in their lives, particularly as teenagers, are 'worse off' at ages 30 and older than women who became mothers beyond their 23rd birthday. We now estimate some multivariate models for these outcomes that control for father's occupation and explicitly take into account the multiple measures of the outcome variable for the same woman. After controlling for father's occupation and the presence of informaion on the father, whether or not a woman lived with both parents throughout her childhood was rarely a significant influence on these outcomes, and so it is not considered further.[4]

The statistical model takes the following form:

$$y_{it} = c + \mu_i + F_i\alpha + B_i\delta + x_{it}\beta + u_{it} \tag{1}$$

where y_{it} is an outcome measure for woman i at year t; F_i is a vector of the three youngest age-at-first-birth intervals (24 or older is the reference category); B_i is a vector of background variables, including three dichotomous variables for each of the top three quartiles of the Hope-Goldthorpe score of the father's occupation when the woman was aged 14 (the bottom quartile is the reference category), an indicator variable indicating that the father's occupational information was missing, two dichotomous indicators of a woman's birth cohort (the 1950-55 cohort is the reference category) and, in some specifications, two dichotomous indicators of whether or not the first birth was outside a live-in partnership and whether or not it was in a cohabiting union (a birth in marriage being the reference category); x_{it} is a vector containing age and its square; μ_i is a woman-specific, zero-mean random variable distributed independently of F_i, B_i, x_{it} and u_{it}; α, δ and β are vectors of parameters to be estimated, with α being the parameters of primary interest; c is an intercept parameter and u_{it} is an identically and independently distributed random variable over women and years.[5] No particular distributional assumption is made when the outcome variable is continuous, but when the outcome variable is dichotomous, y_{it} is a latent variable and μ_i and u_{it} are assumed to be normally distributed. Define ρ to be the proportion of variance attributable to variation in μ_i (i.e. $\rho = var(\mu_i)/[var(\mu_i) + var(u_{it})]$)

The assumption of independence of μ_i and F_i means that we treat F_i as exogenous. As noted above, analysis by Ermisch and Pevalin (2003b) strongly suggests that our estimate of α under this assumption is likely to be biased toward overstating adverse impacts of early childbearing for many outcomes, and this should be taken into account in interpreting the estimates.

Living standards later in life

The estimation strategy pursued is best illustrated by discussion of the impact of the timing of motherhood on the logarithm of equivalent household income at age 30 and beyond. The estimates of α for this outcome are shown in Table 4, and when multiplied by 100 these coefficients can be interpreted as the percentage reduction in equivalent household income associated with the particular age-at-first-birth category relative to having a first child at age 24 or older. For example, from the first row of panel A, relative to having her first child at age 24 or older, having a teen-birth reduces a woman's equivalent household income by 22%; starting childbearing at ages 20-21 reduces it by 14%; and entering motherhood when aged 22-23 reduces it by 10%. A chi-square test of the hypothesis that the parameters associated with a

[4] Inclusion of the variable indicating the presence of parents through a woman's childhood also produces a considerable reduction in sample size, because only panel members present at wave 6 (1996) were asked this question.

teen-birth and a birth when aged 20-21 are equal cannot be rejected at the 0.10 level or less, nor can we reject the hypothesis that the aged 20-21 and aged 22-23 parameters are equal. But the parameters associated with a teen-birth and a birth when aged 22-23 are significantly different (*p-value*=0.019).

Panel B of Table 4 addresses the issue of whether apparent adverse impacts of early childbearing are mainly reflecting the fact that early births are disproportionately outside marriage, either in a cohabiting union or outside a live-in partnership altogether. For instance, among women born during 1962-71, only 33% of teen-first-births were inside marriage, 30% were born into a cohabiting union and 37% were born outside a live-in partnership. As women giving birth outside marriage are much more likely to spend time as a single parent, this may depress future living standards. Comparing the first rows of panels A and B indicates that the effects of early childbearing are indeed smaller when we control for partnership status at the time of the first birth, and they are not significantly different from each other. They do, however, show that childbearing before the age of 24 is associated with a significant reduction in equivalent household income later in life, by the order of 10%. The partnership status coefficients (not shown) indicate that having a first child outside a live-in partnership is associated with a statistically significant 22% reduction in equivalent household income relative to having a marital first birth. Initiating childbearing in a cohabiting union reduces it by 10% relative to a marital birth, but this is not statistically significant (at the 0.05 level or less). Comparison of corresponding rows in the two panels indicates that the age-at-first-birth parameters are usually lower in panel B, but the patterns of statistical significance of these are similar in the two panels of Table 4.

[5] In the cohort-specific equations, the cohort variables are dropped, and in the father's occupational quartile-specific equations, the quartile indicators are dropped.

Thus, controlling for partnership status at birth only moderates age-at-first-birth associations with equivalent household income. Furthermore, the exogeneity of partnership status at birth is at least as questionable as that of age-at-first-birth, and so the discussion focuses on the results in panel A. Comparing across the next three rows for different sets of birth cohorts, teen-births there are only significant negative associations for latter two sets of cohorts, whose equivalent household income is measured over the ages 30-45 and ages 30-39 respectively. But confining the estimation for the first set of cohorts (born 1950-55) to ages 30-45 or aged 30-39 produces similar results to those in panel A, suggesting different impacts in this set of cohorts. For the middle set of cohorts, starting childbearing at ages 20-23 has equivalent effects as teen-births. Comparing across the four quartiles of the father's occupational score, significant associations are confined to the bottom two quartiles, and in the bottom quartile, having a first birth when aged 20-21 has nearly the same adverse 'impact' as having a teen-birth.

Regarding the background influences (B_i) on equivalent household income later in life, there is a tendency for it to be larger for women whose fathers were in better earning occupations, and for more recent birth cohorts. It also tends to increase with age. Also, 50-60% of the residual variance is attributable to woman-specific influences (μ_i); put differently, the coefficient of correlation in equivalent household income between any two years for a particular woman, conditional on F_i, B_i and x_{it}, is in the range 0.5 to 0.6 (see the estimate of ρ in Table 4).

The association of the age at first birth with the risk of being poor (i.e. in the bottom quartile of the equivalent household income distribution) is examined in Table 5. The estimates of α shown there correspond to the association of being in the particular age-at-first-birth category on the 'latent poverty index' (y_{it}), and the ‘marginal effects’ in the first rows of panels A and B are calculated (numerically) for a change from 0 to 1 in the particular age-at-first-birth category, and evaluated at the mean values of the other explanatory variables B_i and x_{it}, F_i=0 and μ_i=0 (i.e. at the mean woman-specific unobserved influence).[6] Thus, for the full sample in panel A, having a teen-birth increases the probability of being poor by 0.216 compared to a woman having her first child when aged 24 or older. Starting childbearing when aged 20-21 increases this probability by 0.114 relative to this reference group. These are large ‘effects’ in light of the predicted probability of being poor of 0.055 for the reference woman, who has a birth when aged 24 or older, mean values of the other explanatory variables B_i and x_{it} and μ_i=0.[7]

Looking down the rows of panel A of Table 5, there is evidence of strong association of a teen-first-birth with the risk of poverty within each of the sets of birth cohorts and each of the father’s occupational groups. About 70% of the residual variance is attributable to woman-specific influences on the probability of being in poverty.

Controlling for partnership status at first birth reduces the associations of a teen-birth with poverty, as shown in Panel B of Table 5. These estimates also indicate that, if the first child is born outside a live-in partnership, the probability of being poor increases by 0.08 relative to a first child born in marriage, and if the child is born in a cohabiting union, this probability increases by 0.10. Both of these associations are statistically significant.

Receipt of Income Support (IS) benefits is another indicator of poverty. Table 6 shows the associations of age at first birth on the probability with receiving IS

[6] In panel B, the marginal effects are evaluated for the reference partnership status, a marital first birth.

[7] It should be noted that the calculation of these marginal effects is sensitive to the particular constellation of explanatory variables assumed, because it affects the base value of the probability.

sometime during the year. In panel A, estimates from the full sample indicate that the probability that the reference woman (defined as for the poverty outcome) receiving IS increases by 0.026 if the woman starts childbearing as a teenager compared with starting when aged 24 or older. Becoming a mother when aged 20-21 (22-23) increases the probability of IS-receipt by a statistically significant 0.005 (0.001), which is significantly lower than the 'marginal effect' a teen-birth. The marginal effects of childbearing under the age of 22 on the probability of IS-receipt are halved when we control for partnership status at the time of the birth and evaluate them for marital first births, as shown in panel B of Table 6. But the marginal effects of early childbearing remain statistically significant. Having her first birth outside a partnership increases the probability of receiving IS by 0.013 relative to a marital first birth, but the impact of starting childbearing in a cohabiting union is small and insignificant relative to marriage.

Comparing the rows in panel A, childbearing as a teenager is associated with higher chances of IS-receipt for each set of cohorts, and a first birth when aged 20-21 also increases these chances for the two most recent sets of cohorts. There is evidence of strong positive associations of motherhood before the age of 24 with IS-receipt in the top and the two bottom quartiles of father's occupations.

Women's employment and earnings

The rules for the receipt of Income Support make the probability of its receipt intimately related to whether or not a woman has an employed partner and whether or not she has a job herself. More generally, the risk of being poor and the amount of equivalent household income depend on the chances that a woman has a job, her earnings if she has one, her chances of having a partner and his earnings. This section examines the first two of these factors, and the next the last two.

Although not always statistically significant, Table 7 shows that becoming a mother as a teenager is associated with a lower probability of employment compared with the reference group of women, who became mothers when aged 24 or older. This is despite the fact that a teen-mother's oldest child is older than those starting childbearing later and the ages and number of children are known to influence the probability that a mother has a job. In line with this reasoning, women becoming a mother when aged 20-23 have a higher probability of having a job than mothers who were older when they started childbearing. Thus, teenage mothers stand out in this regard amongst the 'early motherhood' group. The 'marginal effects' in the first row of panel A indicate that, relative to the reference group, teen-mothers' probability of employment is 0.051 lower (although only significant at the 0.10 level), while the probability is 0.049 and 0.064 higher for women starting childbearing at ages 20-21 and 22-23 respectively. When teen-births aged under 18 are distinguished from those when the mother was aged 18-19, the marginal effects are significantly different between these two groups of teen-mothers. Having a birth aged under 18, reduces the probability of employment by a statistically insignificant 0.035, while starting childbearing when aged 18-19 reduces the probability of employment by 0.16.

Panel B indicates that the associations of age at first birth with mother's employment are even larger when we control for partnership status at birth. Also, having a birth outside a partnership is associated with a much lower probability of employment, reducing it by 0.388, while starting childbearing in a cohabiting union increases this probability by 0.066. The negative association between employment and having a teen-birth is also evident in all of the sub-samples in panel A of Table 7.

Does this represent a causal impact of a teen-birth, or non-random selection into the population of teen mothers? In order to see how badly biased the estimate of the impact of a teen-birth on mother's later employment might be using the method of this paper, it is compared with the instrumental variables' estimator used by Ermisch and Pevalin (2003b), which is consistent under weaker conditions than the estimator used in Table 7. The top panel of Table A in the appendix shows the confidence interval for the equivalent estimator to that in Table 7 using the British Cohort Survey 1970 (BCS70) data. denoted as α_0, and also the confidence interval for the IV estimator (from the BCS70 data). The former lies inside the latter, suggesting that the estimator used in this paper is not badly biased (relative to the imprecise IV estimator). The second panel compares the confidence interval of the estimate from the first row of panel A in Table 7 with that of the IV estimate from the BCS70. This also falls in the IV estimator's confidence interval. We should nevertheless be cautious in interpreting the marginal effects in Table 7 causally.

Conditional on having a job in a particular year, Table 8 indicates that there is no evidence that the timing of age at first birth affects a woman's pay. Neither does the partnership status at the time of the first birth; while being outside a live-in partnership at the time of the birth is estimated to increase pay by 5% and being in a cohabiting union at that time is estimated to increase pay by 12%, neither estimate is significantly different from zero at a level of 0.10 or less. Ermisch and Pevalin's (2003b) instrumental variable estimate also indicated no significant impact of a teen-birth on a woman's pay at age 30, although their results from an estimator like that used in Table 8 suggested large and significant negative impacts.

Partnering, partner's employment and earnings

It generally appears to be the case from Table 9 that starting childbearing before the age of 24 is associated with a much smaller probability of living with a partner when the woman is in her 30s and 40s. Teenage childbearing *per se* does not stand out as being particularly different, although some of the sub-samples in panel A suggest variation with age at first birth, some of these being quite odd patterns.[8] Distinguishing teen-births under 18 from other teen-births, produces quite a different pattern of marginal effects: -0.147 (under 18), -0.054 (18-19), -0.097 (20-21) and – 0.009 (22-23). Thus, a woman becoming a mother before her 18th birthday, is much less likely to be living with a partner in her thirties and forties, even compared with women starting childbearing when aged 18-19 (the difference is statistically significant).

Panel B suggests that these large associations between early childbearing and having a partner arise in large part because of the larger prevalence of non-marital childbearing at these ages. For those having a first birth outside marriage, in either a cohabiting union or outside a live-in partnership, the probability of living with a partner is lower, by a statistically significant 0.078 and 0.072 respectively (relative to a marital birth), and the association between early childbearing and having a partner in her 30s and 40s is much smaller in panel B.

Teenage childbearing is association with a substantially lower probability that a woman's partner has a job later in life, if she has a partner. As shown in Table 10, in the full sample a teen-birth reduces this probability by 0.05 relative to women having their first child when aged 24 or older. While first childbearing when aged 20-23 also

[8] In the first row of each panel of Table 9, the age-at-first-birth effects are not significantly different from one another. As Appendix Table A shows, Ermisch and Pevalin (2003b) find a statistically insignificant effect of a teen-birth on the probability of having a partner at age 30 with their instrumental variable estimator, and a reduction with the estimator used in Table 9.

reduces the probability of having an employed partner, the associations are significantly smaller. Adding the partnership status at birth categories does not significantly improve the model, and not only do the associations with age-at-first birth change very little but the associations with partnership status at birth are virtually zero. These strong negative associations of teenage childbearing with a partner being employed are evident in most of the sub-samples, and they are broadly consistent with the instrumental variable estimate in Ermisch and Pevalin (2003b), as Appendix Table A shows. It indicates that a teen-birth reduces the probability that the partner is employed at age 30 (if she has a partner) by 0.18, but the 95% confidence interval around this estimate does not contain the estimate in the Table 10; it indeed indicates a larger effect than Table 10 suggests. The associations of having a first birth at ages 20-23 with the partner's probability of employment are more mixed across the sub-samples.

Given that a woman has a partner in a job, the current partner of women who started childbearing as a teenager earns about 30% less than that of woman who started at 24 or older, according to the full sample results in Table 11.[9] A sizeable negative impact of a teen-birth on partner's pay is also evident in all of the sub-samples in Table 11. There is also evidence of a significant difference between the impacts of having a birth when aged under 18 and when aged 18-19. Estimates of a model making this distinction (results not shown) imply marginal effects of –0.52 and –0.24 respectively.

As Appendix Table A illustrates, these results are consistent with the findings in Ermisch and Pevalin (2003b). They find that the estimator used in Table 11 and an instrumental variable estimator indicate that having a teen-birth is associated with

[9] Partnership status at first birth does not have a significant effect.

respectively 15% and 28% lower current partner's pay when the woman is aged 30, although the latter (IV) estimator is not statistically significant. All of these results suggest that having a teen-birth, particularly when aged under 18, constrains a woman's opportunities in the 'marriage market' in the sense that she finds it more difficult to find and retain a partner, and she partners with more unemployment-prone and lower earning men.

Housing tenure

Whether people become homeowners is important for a number of reasons. It is usually associated with more choice in housing consumption, thereby contributing to a higher standard of living for the woman. Second, British homeowners have received relatively large, tax-exempt returns on their investment, thereby contributing directly to their wealth accumulation. Third, home ownership is indicative of asset ownership more generally. Those who are owner-occupiers are much more likely to have other financial assets, particularly riskier investments, and they also have higher average levels of wealth (Banks and Tanner, 1999, Tables 5.2 and 5.5). For example, among those working age individuals who do not contribute to a private pension, 22% of those who are not homeowners own no financial assets compared with 4% of homeowners in this group. The non-homeowners in this group have mean financial wealth of £1,200 compared with £6,900 for homeowners. It appears that those who do not accumulate housing wealth do not compensate by accumulating more of other types of wealth.

Table 12 indicates that women who became mothers as teenagers are much less likely to become homeowners. For example, from the full sample in panel A, having a teen-birth is associated with a 0.225 lower the probability of being an owner-occupier relative to a woman becoming a mother at age 24 or later. As Appendix

Table A shows, this result is consistent with the findings in Ermisch and Pevalin (2003b), in which their IV and other estimators indicate at reduction about 0.3 in the probability of being an owner-occupier at age 30.[10] A woman starting childbearing in her early 20s also has a lower probability of being a homeowner, but comparison with panel B indicates that the pattern of effects of age-at-first birth is sensitive to whether or not there are controls for partnership status at the time of the birth. Having a first birth outside a partnership reduces the probability of being a homeowner by 0.054 and having it in a cohabiting union reduces it by 0.013 (relative to a marital first birth in each case). Taking account of the partnership context of the first birth reduces the marginal effect of a teen-birth to –0.189 and that of a first birth when aged 22-23 to virtually zero.

Mental health in later life

It is possible that the adverse effects of teen-births on living standards later in life may also affect a woman's mental health by causing stress and other problems (e.g. Goldberg *et al* 1990). The main measure of 'common mental illness' employed in the multivariate analysis is a 'high caseness' indicator variable, namely a caseness score of 10 or higher. Table 13 shows persistent positive associations between a teen-birth and the probability of poor mental wealth across all sub-samples. For instance, in the full sample having a teen-birth increases this probability by 0.036 (first row of panel A). Starting childbearing when aged 20-21 also increases the risk of poor mental health in the full sample, although this effect is significantly smaller than the effect of a teen birth. Statistically significant associations when aged 20-21 are, however, only evident in two of the sub-samples, and they are generally much smaller than the

[10] If we treat the panel data as a pooled cross-section and estimate a conventional probit model, the estimated marginal effects of the age-at-first birth variables at the mean values of the regressors are

associations with a teen-birth. Partnership status at the time of the first birth does not have a significant association with the probability poor mental health. Finally, note that there is somewhat less time persistence in poor mental health compared to many of the other variables—i..e. the proportion of variance accounted for by the woman-specific component is generally of the order of 0.4. As the GHQ measure offers a comparison of wellbeing in comparison with 'usual conditions', this to be expected.

An alternative dichotomous measure of 'common mental illness' that is often used (Goldberg *et al* 1998, Pevalin and Goldberg 2003) takes the value of one if her caseness scale has a value of 4 or more, and zero otherwise. Thus, it represents a broader definition of poor mental health. The results are, however, similar, and because a higher proportion of women fall into this poor health category, the 'marginal effects' of a teen-birth on the probability of poor health are larger. For instance, in the estimates from the full sample without controls for partnership status at the time of the birth (panel A of Table 14), a teen-birth increases this probability by 0.13 relative to a first birth when aged 24 or older, and having a first birth when aged 20-21 increases it by 0.062. Controlling for partnership status at birth, these effects (panel B) are marginally smaller, and a first birth in a cohabiting union raises the probability by 0.062 relative to a marital first birth.

Interestingly, a teen-birth had no significant effect on this mental health indicator in an analysis of the BCS70, even when using the estimator employed in the present analysis.[11] This difference could conceivably arise because of the single observation of this outcome at age 30 in the BCS70, in contrast to the multiple observations on women aged 30 and over in the present analysis. Re-doing the analysis in Table 14 using only information on the GHQ score of the woman at the

–0.331, -0.211 and –0.122, respectively.

age of 30, the results are very similar to those in Tables 13 and 14, although the standard errors are, of course, larger.

Another approach, estimates the impacts on the latent variable y_{it} in equation (1) using all four categories of 'caseness' given in the last four rows of Table 2 (i.e, 0-3, 4-6, 7-9 and 10-12). Assuming a standard normal distribution for $\mu_i + u_{it}$, we obtain ordered probit estimates. These produce comparable marginal effects to those in Tables 13 and 14 for the top and bottom categories of caseness scores respectively. In light of the skewness in the distribution of caseness scores evident in Table 2 (e.g. 75% have scores of 0-3), a 'negative binomial count data' model for the 12-point caseness score may be more appropriate.[12] Estimation of such a model (not shown) produces a strong positive association of starting childbearing with the caseness score.

Educational qualifications

Examination of the impact of the teen-births on educational attainments has been common in previous studies, which indicate that those having a teen-birth are more likely to have no or lower educational qualifications. The analysis in Ermisch and Pevalin (2003b) also finds this when they treat a teen-birth as exogenous (as done in the present analysis), but these associations disappear in their instrumental variable estimates. Indeed, the confidence interval from the estimate assuming exogenous teen-birth is often not contained in the confidence interval of the IV estimate. This strong suggests that the results presented in Table 15 should not be treated as causal effects. Rather they include the influence of non-random selection into age-at-first-birth categories with respect to eventual educational qualifications. That is, women

[11] Their IV estimator indeed indicated a *reduction* in the probability of a caseness score of 4 or more of 0.16, although this only significant at the 0.10 level.

[12] The negative binomial model deals with circumstances in which there is more variation than would be expected were the count process Poisson.

who become pregnant as teenagers or who do not abort their pregnancy would have lower qualifications in any case.

Table 15 shows the ‘marginal effects’ implied by an ordered probit model applied to the educational qualifications accumulated at the oldest age that we observe the woman (beyond the age of 30). These indicate a strong gradient with age at first birth, with the probability of having no qualifications and O-level qualifications decreasing with age at first birth, while the probability of having qualifications of A-level or above increases with age at first birth. Partnership status at the time of the birth does not have a statistically significant influence on educational attainments.

3. Conclusions

These analyses suggest that, relative to women starting a family when age 24 or older, women having a teen-birth are less likely to be employed later in life (in their thirties and forties), but their pay is not affected if they have a job. They are also less likely to be living with a partner at these ages, particularly if they started childbearing before their 18th birthday. This association appears to mainly reflect the fact that women starting childbearing as a teenager are much less likely to be married at the time of their first birth. For women who have a partner in their thirties and forties, it is less likely that her partner has a job, and if he has one, his pay is much lower, particularly if she became a mother before her 18th birthday. These results suggest that having a teen-birth, particularly when aged under 18, constrains a woman’s opportunities in the ‘marriage market’ in the sense that she finds it more difficult to find and retain a partner, and she partners with more unemployment-prone and lower earning men.

No doubt related to the above influences of a teen-birth, teenage mothers are much less likely to be a homeowner later in life (the probability is reduced by about 0.2), and her living standard, as measured by equivalent household income, is about

20% lower. Her probability of being ‘poor’ (in the bottom quartile of the equivalent income distribution) is about 0.2 higher, and the probability that she receives Income Support is also higher (by 0.02). Finally, and probably related to the impacts of a teen-birth on her living standards, she is much more likely to suffer from common mental illness in her thirties and forties.

As noted at the outset, it is not clear the extent to which these associations represent ‘causal effects’; that is, would the women becoming mothers as teenagers have had these poorer outcomes anyway? But the fact that the adverse impacts of a teen-birth on the probabilities of being a homeowner and of having an employed partner estimated here are in line with the statistically significant instrumental variable estimates of these impacts in Ermisch and Pevalin (2003b) increases our confidence that these associations measure, in large part, causal effects. The estimated impacts of a teen-birth on the probabilities that the woman has a partner, and the impact on her partner’s pay (if she has one) are also contained in the 95% confidence interval associated with the corresponding instrumental variable estimates. This increases our confidence in the estimates of the effect of a teen-birth on living standards later in life, which are measured much better in the BHPS data.

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Table 1: Means of outcome variables by woman's age at first birth

Outcome variable	15-19	20-21	22-23	24 or older	F-test (3 df)
ln(Equiv. HH income)	6.562	6.687	6.698	6.838	84.85
ln(own pay)	6.085	6.082	6.178	6.236	14.87
ln(partner's pay)	7.071	7.153	7.177	7.372	13.45
GHQ score	13.623	12.751	12.079	11.670	58.10

Table 2: Percentages in particular states by woman's age at first birth

Outcome variable	15-19	20-21	22-23	24 or older	Chi-sq.test (3 df)
% poor	39.1	28.1	25.7	19.5	280.80
% receiving IS	19.0	12.1	10.0	6.5	243.16
% employed	58.8	70.1	73.0	71.0	112.79
% with a partner	75.9	80.5	81.8	86.8	126.95
% partner employed	77.2	85.4	90.2	93.6	313.52
% owner-occupier	51.6	65.9	74.8	86.4	978.27
GHQ case-ness 10-12	10.6	7.3	6.0	4.3	
GHQ case-ness 7-9	9.4	8.7	7.5	6.5	
GHQ case-ness 4-6	13.3	11.9	11.1	11.3	
GHQ case-ness 0-3	66.7	72.1	75.4	77.8	135.82 (9 df)

Table 3: Educational qualifications (%) by woman's age at first birth

Qualifi-cation level	15-19	20-21	22-23	24 or older
None	37.5	26.5	16.6	11.2
O-level	37.5	33.7	36.5	31.0
A-level & nursing	6.0	7.1	11.6	11.4
Teaching and other higher qual	16.8	29.1	29.1	31.2
Degree	2.2	3.6	6.2	15.3
N	232	196	241	564

Chi-square (12df)=130.84

Table 4: Woman's age at first birth (α) and the logarithm of equivalent household income, aged 30 and over

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample	-0.215*	-0.135*	-0.102*	0.565	10,710 1,318
Born 1950-55	-0.128	-0.048	-0.016	0.597	4,124 470
Born 1956-61	-0.243*	-0.235*	-0.219*	0.536	4,170 474
Born 1962-71	-0.263*	-0.122	-0.102	0.554	2,416 409
Father, top quartile HGS	-0.077	-0.061	0.020	0.504	2,711 311
Father, 2nd quartile HGS	-0.164	-0.079	-0.180*	0.544	2,426 296
Father, 3nd quartile HGS	-0.271*	-0.072	-0.100	0.628	2,227 262
Father, 4th quartile HGS	-0.257*	-0.217*	-0.044	0.598	2,047 259

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample	-0.158*	-0.105*	-0.096*	0.561	10,710 1,318
Born 1950-55	-0.063	0.008	-0.009	0.585	4,124 470
Born 1956-61	-0.186*	-0.212*	-0.221*	0.534	4,170 474
Born 1962-71	-0.230*	-0.103	-0.094	0.554	2,416 409
Father, top quartile HGS	-0.135	-0.092	0.021	0.504	2,711 311
Father, 2nd quartile HGS	-0.073	-0.038	-0.167*	0.533	2,426 296
Father, 3nd quartile HGS	-0.243*	-0.058	-0.099	0.629	2,227 262
Father, 4th quartile HGS	-0.187*	-0.160*	-0.056	0.587	2,047 259

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.

Table 5: Woman's age at first birth and the probability of being in the bottom quartile of the equivalent household income distribution, aged 30 and over

A. No control for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	0.988* **0.216***	0.639* **0.114***	0.291 **0.041**	0.698	10,710 1,318
Born 1950-55: α	0.663*	0.396*	-0.520*	0.719	4,124 470
Born 1956-61: α	0.893*	1.038*	0.623	0.697	4,170 474
Born 1962-71: α	1.322*	0.529*	0.494*	0.731	2,416 409
Father, top quartile HGS: α	0.510*	0.032	-0.435*	0.726	2,711 311
Father, 2nd quartile HGS: α	1.034*	0.519*	0.504*	0.731	2,426 296
Father, 3nd quartile HGS: α	1.252*	0.686*	0.709*	0.750	2,227 262
Father, 4th quartile HGS: α	1.381?	1.108*	0.553?	0.694	2,047 259

*Statistically significant at 0.05 level

**First row is number of woman-year observations, second row is number of women.

? Denotes inability to compute standard error of the parameter estimate.

B. Controls for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	0.862* **0.168***	0.601* **0.099***	0.324* **0.044***	0.708	10,710 1,318

Table 6: Woman's age at first birth and the probability of receiving Income Support, aged 30 and over

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	1.391* **0.026***	0.790* **0.005***	0.427* **0.001***	0.794	10,957 1,322
Born 1950-55: α	0.984*	0.398	0.168	0.802	4,227 472
Born 1956-61: α	1.317*	0.812*	0.973 ?	0.813	4,246 472
Born 1962-71: α	1.250*	0.816*	0.495	0.807	2,484 413
Father, top quartile HGS: α	1.132*	1.343*	0.654*	0.799	2,787 311
Father, 2nd quartile HGS: α	0.344	0.311	0.401	0.748	2,426 296
Father, 3nd quartile HGS: α	1.214*	1.063 ?	1.171*	0.787	2,279 265
Father, 4th quartile HGS: α	1.723 ?	0.735 ?	1.264 ?	0.786	2,088 260

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.
? Denotes inability to compute standard error of the parameter estimate.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	1.316* **0.012***	0.744* **0.002***	0.578* **0.001***	0.800	10,957 1,322

Table 7: Woman's age at first birth and the probability of being in employment, aged 30 and over

A. No control for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	-0.193 **-0.051**	0.227 **0.049**	0.310* **0.064***	0.752	11,130 1,329
Born 1950-55: α	-0.396*	0.302	-0.223	0.814	4,291 473
Born 1956-61: α	-0.572*	0.131	0.343*	0.716	4,319 473
Born 1962-71: α	-0.453	0.339	0.295	0.761	2,520 420
Father, top quartile HGS: α	-0.205	0.624*	0.494*	0.715	2,813 313
Father, 2nd quartile HGS: α	-1.351*	0.132	0.305 ?	0.788	2,521 296
Father, 3nd quartile HGS: α	-0.554*	-0.419 ?	0.156	0.764	2,322 267
Father, 4th quartile HGS: α	-0.648*	-0.205	-0.910*	0.807	2,130 262

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.
? Denotes inability to compute standard error of the parameter estimate.

B. Controls for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	-0.180 **-0.051**	0.313* **0.070***	0.380* **0.082***	0.749	11,130 1,329

Table 8: Woman's age at first birth (α) and the logarithm of a woman's pay, aged 30 and over and with a job

A. No control for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	**ρ**	***N*********
Full sample	-0.027	0.0111	0.057	0.771	7,063 1,074
Born 1950-55	0.021	0.060	-0.022	0.819	2,866 398
Born 1956-61	0.041	-0.018	0.106	0.723	2,731 387
Born 1962-71	-0.109	-0.035	0.085	0.757	1,466 314
Father, top quartile HGS	0.142	0.247	0.304*	0.748	1,836 261
Father, 2nd quartile HGS	0.003	0.033	0.023	0.724	1,630 241
Father, 3nd quartile HGS	-0.169	0.002	-0.084	0.815	1,488 218
Father, 4th quartile HGS	-0.199	-0.099	-0.009	0.785	1,330 215

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.

B. Controls for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	**ρ**	***N*********
Full sample	-0.049	0.002	0.054	0.771	7,063 1,074

Table 9: Woman's age at first birth and the probability of living with a partner, aged 30 and over

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-1.530* **-0.029***	-1.459* **-0.024**	-1.662* **-0.039***	0.839	10,964 1,321
Born 1950-55: α	-0.465*	-0.494*	0.800*	0.876	4,190 467
Born 1956-61: α	-1.041*	0.240	-1.673 ?	0.815	4,280 471
Born 1962-71: α	-1.547*	-1.695*	-0.232	0.847	2,494 419
Father, top quartile HGS: α	0.715*	-1.688*	-1.549*	0.801	2,770 312
Father, 2nd quartile HGS: α	-0.113	-0.876*	-1.828*	0.866	2,493 294
Father, 3nd quartile HGS: α	-3.090*	-0.254	-2.268*	0.881	2,281 266
Father, 4th quartile HGS: α	-0.157	-0.648*	-1.091*	0.894	2,083 258

*Statistically significant at 0.05 level

**First row is number of woman-year observations, second row is number of women.

? Denotes inability to compute standard error of the parameter estimate.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-0.544 **-0.003***	-0.513* **-0.003***	-0.488* **-0.002***	0.839	10,964 1,321

Table 10: Woman's age at first birth and the probability of her partner being in employment, aged 30 and over with a partner

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-1.609* **-0.050***	-0.729* **-0.005**	-0.646* **-0.004***	0.779	9,085 1,194
Born 1950-55: α	-1.335 ?	-0.506*	-0.061	0.814	3,528 427
Born 1956-61: α	-2.091*	-1.306*	-0.895*	0.774	3,615 434
Born 1962-71: α	0.137	0.050	0.790	0.837	1,942 363
Father, top quartile HGS: α	-1.127*	0.632	0.127	0.666	2,385 301
Father, 2nd quartile HGS: α	-2.160*	-1.128*	0.276	0.812	2,138 270
Father, 3nd quartile HGS: α	-2.723*	-3.112*	-1.733*	0.847	1,916 242
Father, 4th quartile HGS: α	-1.627*	-0.287	0.359	0.813	1,666 231

*Statistically significant at 0.05 level

**First row is number of woman-year observations, second row is number of women.

? Denotes inability to compute standard error of the parameter estimate.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-1.785* **-0.055**	-0.834* **0.005***	-0.664* **0.003***	0.784	9,085 1,194

Table 11: Woman's age at first birth (α) and the logarithm of her partner's pay, aged 30 and over and with a partner in a job

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample	-0.318*	-0.182*	-0.197*	0.537	6,789 1,019
Born 1950-55	-0.327*	-0.262*	-0.175*	0.400	2,574 366
Born 1956-61	-0.295*	-0.222	-0.307*	0.559	2,767 384
Born 1962-71	-0.344*	-0.018	-0.146	0.747	1,448 288
Father, top quartile HGS	-0.586*	-0.258	-0.158	0.413	1,920 277
Father, 2nd quartile HGS	-0.393*	-0.123	-0.342*	0.739	1,678 232
Father, 3nd quartile HGS	-0.180	-0.174	-0.124	0.605	1,399 205
Father, 4th quartile HGS	-0.244*	-0.242*	-0.108	0.411	1,134 182

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample	-0.290*	-0.170*	-0.197*	0.537	6,789 1,019

Table 12: Woman's age at first birth and the probability of being a homeowner, aged 30 and over

A. No control for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-1.962* **-0.225***	-0.844* **-0.028***	-1.337* **-0.082***	0.909	10,969 1,322
Born 1950-55: α	-1.683*	-1.809*	-1.392*	0.913	4,232 472
Born 1956-61: α	-1.275*	-0.246	0.083	0.886	4,255 474
Born 1962-71: α	-2.386*	-0.639*	-0.036	0.904	2,482 413
Father, top quartile HGS: α	-2.439*	-2.383*	-0.603*	0.852	2,783 311
Father, 2nd quartile HGS: α	-0.639*	-0.862*	-2.166*	0.921	2,485 296
Father, 3nd quartile HGS: α	-3.143*	-0.819?	-2.905*	0.911	2,279 265
Father, 4th quartile HGS: α	-0.951?	-0.617?	-0.496?	0.859	2,092 260

*Statistically significant at 0.05 level

**First row is number of woman-year observations, second row is number of women.

? Denotes inability to compute standard error the parameter estimate.

B. Controls for partnership status at birth

Specification	15-19	20-21	22-23	ρ	*N***
Full sample: α **Marg. Effect**	-3.348* **-0.189***	-2.505* **-0.042***	-0.502* **-0.000**	0.902	10,969 1,322

Table 13: Woman's age at first birth and the probability of 'poor mental health' (caseness score of 10 or higher), aged 30 and over

A. No control for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	0.602* **0.036***	0.350* **0.016***	0.183 **0.007**	0.443	10,808 1,321
Born 1950-55: α	0.453*	0.024	-0.011	0.513	4,160 472
Born 1956-61: α	0.639*	0.577*	0.323	0.374	4,195 472
Born 1962-71: α	0.705*	0.425*	0.199	0.427	2,453 411
Father, top quartile HGS: α	0.528*	-0.232	-0.011	0.357	2,762 311
Father, 2nd quartile HGS: α	0.331	0.344	-0.015	0.337	2,443 296
Father, 3nd quartile HGS: α	0.833*	0.481	0.199	0.551	2,248 265
Father, 4th quartile HGS: α	0.925*	0.424	0.491	0.447	2,060 260

*Statistically significant at 0.05 level
**First row is number of woman-year observations, second row is number of women.

B. Controls for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α **Marg. Effect**	0.566* **0.032***	0.333* **0.014***	0.181 **0.007**	0.442	10,808 1,321

Table 14: Woman's age at first birth and the probability of 'poorer mental health' (caseness score of 4 or higher), aged 30 and over

A. No control for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α	0.431*	0.224*	0.098	0.391	10,808
Marg. Effect	**0.129***	**0.062***	**0.026**		1,321

B. Controls for partnership status at birth

Specification	**15-19**	**20-21**	**22-23**	ρ	***N*****
Full sample: α	0.406*	0.214*	0.097	0.390	10,808
Marg. Effect	**0.119***	**0.058***	**0.025**		1,321

*Statistically significant at 0.05 level

**First row is number of woman-year observations, second row is number of women.

Table 15: Woman's age at first birth and the probability of obtaining particular educational qualifications by oldest age observed, aged 30 or more

A. No control for partnership status at birth*

Qualification level	**15-19**	**20-21**	**22-23**	**Predicted Probability**
None	0.196*	0.104*	0.052*	0.115
O-level	0.075*	0.064*	0.041*	0.326
A-level & nursing	-0.023*	-0.009*	-0.003	0.109
Teaching and other higher qual	-0.165*	-0.099*	-0.053*	0.341
Degree	-0.082*	-0.060*	-0.037*	0.108

*Evaluated at 'first birth aged 24 or older' category and mean values for the other control variables. N=1,233 women aged 30 and older.

B. Controls for partnership status at birth*

Qualification level	**15-19**	**20-21**	**22-23**	**Predicted Probability**
None	0.193*	0.103*	0.052*	0.115
O-level	0.075*	0.063*	0.041*	0.326
A-level & nursing	-0.023*	-0.009*	-0.003	0.110
Teaching and other higher qual	-0.163*	-0.098*	-0.053*	0.341
Degree	-0.081*	-0.059*	-0.037*	0.108

*Evaluated at 'first birth aged 24 or older' category and mean values for the other control variables. N=1,233 women aged 30 and older.

Table A: 95% Confidence Intervals of Average Effect of a Teen-birth among Mothers having a Teen-birth

Panel I.

Outcome at age 30	**α_0* from BCS70**		**α_{IV} from BCS70**	
	Lower	**Upper**	**Lower**	**Upper**
Pr(In employment)	**-0.13**	**-0.02**	-0.25	0.17
Pr(Has partner)	**-0.14**	**0.04**	-0.18	0.19
Pr(Partner in employment)	**-0.20**	**-0.10**	-0.28	-0.08
Log(Partner's pay)	**-0.26**	**-0.05**	-0.68	0.12
Pr(Owner-occupier)	**-0.36**	**-0.25**	-0.52	-0.13

Panel II.

Outcome at ages 30-51	**α_0* 'Marginal effect' from BHPS**		**α_{IV} from BCS70**	
	Lower	**Upper**	**Lower**	**Upper**
Pr(In employment)	**-0.19**	**-0.05**	-0.25	0.17
Pr(Has partner)	**-0.03**	**-0.00**	-0.18	0.19
Pr(Partner in employment)	-0.08	-0.02	-0.28	-0.08
Log(Partner's pay)	**-0.37**	**-0.11**	-0.68	0.12
Pr(Owner-occupier)	**-0.33**	**-0.22**	-0.52	-0.13

Bold indicates that the confidence interval of α_0 lies fully within the α_{IV} confidence interval.