The effect of teenage childbearing on educational attainment: evidence from two samples of twins and their sibling sisters

Dinand Webbink*, Nick Martin**, Peter M. Visscher**

Abstract:

This paper analyses the effect of teenage childbearing on educational attainment. We extend previous studies, which compare the outcomes of sisters, by comparing the outcomes of twin sisters. We use two samples of Australian twins and their relatives. For the oldest cohort we find that the adverse effect of teenage childbearing reduces when moving from cross sectional estimates to fixed effect estimates on samples of siblings, and further reduces in the samples of twins and identical twins only. The estimates on the sample of identical twins are statistically not significant. For the youngest cohort we find no difference in educational attainment between teen mothers and their twin sisters. The findings are robust for different age cutoffs and for the exclusion of twin pairs who report early separation. For a smaller sample of teen mothers we have information about the educational attainment of both their twin sister and their sibling sister. We find that the within twin estimate is approximately one third smaller than the within sibling estimate.

Keywords: teenage childbearing, education attainment, twins

* CPB Netherlands Bureau for Economic Policy Analysis and SCHOLAR, University of Amsterdam
** Queensland Institute of Medical Research, Brisbane

1. Introduction

Teenage childbearing is considered a major problem in many countries.[1] A large literature documents that early fertility appears to be associated with a number of adverse economic and social outcomes such as lower educational attainment and family income, higher numbers of households headed by single women, higher rates of poverty and higher rates of low birth weight and infant mortality. However, these findings have been challenged in several studies that address the endogeneity of teenage childbearing by using within family estimation (Geronimus et al., 1992, Holmlund, 2005) or instrumental variable approaches (Ribar, 1994, Hotz, et al., 2005, Ashcraft et al., 2006, Bradbury, 2006). These studies typically find modest negative effects or even positive effects of teenage childbearing (Hotz, et al., 2005).

This paper analyses the effects of teenage childbearing on two cohorts of Australian twins and their relatives. Our paper makes several contributions to the literature. First, one of the main approaches to address the endogeneity of teenage childbearing is within family estimation using samples of siblings (Geronimus, et al., 1992, Holmlund, 2005). We extend this approach by using pairs of twins and pairs of identical twins only. In addition, we have two measures of educational attainment for two samples of twins and their relatives. Second, Our data enable us to compare outcomes within pairs of siblings with outcomes within pairs of twins for the same sample of teen mothers. Hence, we can make a comparison of the within sibling estimate with the within twin estimate. A further contribution is that we can control for several differences between twins.

Our main finding is that the effect of teenage childbearing on educational attainment appears to be small or not significantly different from zero, especially for the most recent cohort of twins. As in previous studies, we find that within family estimates of the difference in educational attainment are much smaller than cross sectional estimates. For the oldest cohort we find that teenage mothers have 0.5 years less education than their twin sisters, for the identical twins this difference is 0.2 years and statistically not significant. For the youngest cohort we find no difference in educational attainment between teen mothers and their twin sisters. These findings are robust for different age cutoffs and exclusion of pairs of twins who report early separation or a large difference in education. The comparison of the difference between teen mothers and their twin sisters with the difference between the same teen mothers and their sibling sister shows that the former difference is one third smaller.

[1] The highest rates for the developed countries are found in the US (52 per 1,000), UK (31), New Zealand (30), Canada (20) and Australia (18) (Unicef, 2001, data for 1998).

The remainder of this paper is organized as follows. The next section reviews previous studies on the effects of teenage childbearing and explains the methodology used in this paper. Section three describes the data. The main estimation results are shown in section 4. Section 5 concludes.

2 Previous studies and methodology

The traditional approach to estimate the causal effect of teenage childbearing on adult outcomes is to control for observables factors, especially measures of the socioeconomic status, using regression models (see Card and Wise 1978, Hofferth and Moore 1979, Upchurch and McCarthy 1990 and McElroy 1996). These studies typically find substantial adverse outcomes of teenage childbearing. However, the estimates with this approach will be biased if unobserved factors are correlated with teenage childbearing and adult outcomes. Several studies have attempted to improve on these estimates by explicitly taking the endogeneity of teenage childbearing into account. For instance, bivariate probit models have been used to estimate the joint process of the woman's decision to bear a child as a teenager as well as the maternal outcome of interest, such as, education (Ribar 1994), labor supply, or poverty status (Lundberg and Plotnick, 1989). These studies find at most modest negative outcomes of teenage childbearing. Several recent papers use an instrumental variable approach. Hotz et al. (2005) exploit the fact that some women who become pregnant experience a miscarriage and do not have a live birth. If miscarriages are random and all miscarriages occur before abortions have taken place this provides a consistent estimate of the effect of teenage childbearing. They find that teenage childbearing increases the probability to complete high school and also has a positive effect on later earnings. Ermisch and Pevalin (2003, 2005) employ the same technique for the UK. They do not find adverse effects of teenage childbearing on education or labour market outcomes, but do find that women who were teenage mothers are more likely to have unemployed or low-income husbands at age 30. Bradbury (2006) uses the same approach for Australian data and finds no adverse impact of early childbearing on education, employment and income, and a lower probability of having a partner. However, miscarriages occur both early and late in pregnancy and that some abortions prevent miscarriages, which rejects the assumptions of the IV-estimator (Ashcraft et al., 2006). Using a competing risk model they find modest adverse outcomes of teenage childbearing. The third approach to reduce bias by unobserved effects is based on within-family estimation. We follow this approach in this paper.

Methodology

The typical econometric model used for within-family estimation is:

$$y_{ij} = \alpha + \beta T_{ij} + \gamma X_{ij} + f_j + \varepsilon_{ij} \qquad (1)$$

where y_{ij} is the outcome of individual i in family j, T_{ij} a dummy variable that takes the value one if the individual is a teenage mother and zero otherwise, X_{ij} a vector of covariates, f_j is an unobserved family effect common to all siblings within the same family and ε_{ij} is a random error term. In this model the family fixed effect is removed by differencing (demeaning) between siblings.

An application of this approach is Geronimus and Korenman (1992) (GK) who compare the socioeconomic outcomes of sisters who timed their first birth at different ages. In two of the three data sets that they examine they find that the adverse outcomes of teenage childbearing are much smaller when family fixed effects are taken into account. GK note that the comparison of sisters provides an improved way of accounting for family background characteristics but these estimates might still be biased by heterogeneity within families. There may be differences between siblings in genetic endowments or in the way parents treat them (Rosenzweig and Wolpin, 1988). In addition, the socioeconomic conditions facing sisters and the parental inputs received by sisters may differ if family circumstances change over time and with the childrearing experiences of their parents (Hotz, et al, 2005).
This concern about the validity of the within-family estimates has also been expressed in the context of estimating returns to schooling. Bound and Solon (1999) show that the bias in the within-family estimator is not always smaller than the bias in the cross-sectional estimator. This depends on the importance of the fixed family component in the unobservables that both affect teenage fertility and the outcome variable. If the family component accounts for a larger fraction of the variance in those unobservables then the bias of the within-estimator is smaller than the bias in the cross-sectional estimator. A recent paper on the effect of teenage childbearing on educational attainment attempts to reduce the within family heterogeneity by adding several additional controls in the within models, especially premotherhood school performance (Holmlund, 2005). Using the within sibling approach she finds a negative effect

of teenage childbearing which is similar to the effect in the traditional cross sectional approach.

In this paper we estimate the consequences of teenage childbearing on educational attainment using 'within-family' estimation on two samples of Australian twins and their relatives. The sample used in the analysis consists of women who all have at least one child and one sister in the sample. In line with GK and Holmlund (2005) we start with cross-sectional estimates followed by within-sibling estimates. As mentioned above, the within-sibling estimates might be biased by within-family heterogeneity. We extend previous research by addressing this possible bias in three ways.

First, we replicate the within-estimation for samples which consists of twins and identical (MZ) twins. It is likely that family circumstances for twins will be more equal than for siblings, which differ in age. For instance, the socioeconomic conditions facing the sisters in the twin sample are expected to be more comparable than the conditions in the sample that also includes the relatives. In addition, the within-estimates of identical twins also control for differences in genetic endowments. Hence, we may expect that the within-twin estimates will be less biased by heterogeneity within families than the within-sibling estimates.

Second, we investigate the robustness of the estimates after excluding pairs who report a separation of their co-twin of at least one year during childhood or pairs of twins with large differences in education. This separation or large education difference might indicate that these twins are different.

Third, for a smaller sample of teen mothers we have information about the educational attainment of both their twin sister and of their sibling sister. This enables us to compare the within sibling estimates with the within twin estimates at the individual level.

Another concern in within-family models is measurement error. It is well-known that the within-family estimator exacerbates measurement error, which is likely to bias the estimates towards zero. Several studies on the returns to schooling using samples of twins addres measurement error in schooling by instrumenting with a second independent measure of schooling (Ashenfelter, et al., 1994, Miller, et al. 1995). These studies typically show that instrumenting leads to higher estimates of the returns to schooling. Hence, it might be expected that the within-family coefficient provides a lower bound of the estimated coefficient of interest (Bound, et al., 1999, Ribar, 1999).

3. Data

In this study we analyze data from two cohorts of twins (and their relatives) of the Australian Twin Register. The data of the first cohort, which is called the older cohort (or the Canberra sample), were gathered in two mail surveys, in 1980-1982 and 1988-1989. The sample consists of all 5967 twin pairs aged over 18 years enrolled in the Australian National Health and Medical Research Council Twin Registry at the time of the first survey. In the first survey 3810 complete pairs participated, in the second survey 2934 twin pairs responded. (Miller, et al. , 1995). In addition to these surveys, data were gathered for the relatives of these twins, including parents, siblings and children, in a survey in 1989-1991. The total number of siblings in this data set is 4832 of which 2434 is female.

The data for the second cohort, which is called the young cohort, were gathered in two surveys, in 1989-1990 and 1996-2000. They constitute a volunteer twin panel born between 1961 and 1974. Nearly all were first registered with the panel between 1980 and 1982 by their parents. A total of 4264 twins pairs were recruited at that time. The data were collected by means of a telephone interview conducted by lay interviewers.

The surveys gathered information on the respondent's family background (parents, siblings, marital status, and children), socioeconomic status (education, employment status and income), health behavior (body size, smoking and drinking habits), personality, and feelings and attitudes. In this paper we restrict the sample to women who are mothers who have at least one (twin)sister in the sample. To avoid sample selection, for instance due to retirement or natural death, we exclude women above the age of 60 at the time of the second survey.

The main independent variable in the analysis is a dummy variable which has value 1 if the women had a child before the age of twenty and has value 0 if the women had her first child at the age of twenty or at a later age. For the older cohort we use information from both surveys to construct this dummy. The first survey only asked the year of first birth, whereas in the second survey the exact date has been asked. Therefore, we use information from the second survey. If this information is missing, we add information from the first survey (this applies to 12 twin pairs). As covariates we use mothers and fathers education, age, birth order and birth weight. Educational attainment is measured with a seven point scale and translated into years of education (Miller, et al. , 1995). This variable is measured in both surveys.

For the second cohort we also use data from both surveys to construct a variable for teenage mothers. If both surveys indicate that the women had her first birth before the age of

twenty the teen mother dummy has value 1. If one survey indicates that the women had a baby before the age of twenty and this information is missing in the other survey we also give this dummy the value 1.

Table 1 shows sample means and proportions for background characteristics and outcome variables. The top panel shows the statistics for the older cohort, the bottom panel for the younger cohort. The cross section sample consists of mothers with at least one child and a sister in the sample. The number of teenage mothers in the older cohort is 299 and the number of non teen mothers is 4462. The within family samples consist of sisters of which at least one is a teen mother. With this restriction the sample size reduces to 241 (143) teenage mothers and (twin) sisters. The number of teen mothers is smaller than in the cross section sample because of losing pairs which completely consists of teen mothers. The number of non teen mothers in the sibling sample exceeds the number of teen mothers because of pairs that include three sisters or more. Previous studies included between 50 and 125 pairs of siblings (GK, 1992) and 322 pairs of siblings (Holmlund, 2005). The sample of (identical) twins consists of 143 (80) teenage mothers and their twin sisters. The sample of the younger cohort consists of 104 (48) teenage mothers and their twin sisters.

Table 1 shows two columns of means and proportions for each sample. The first column shows the statistics for women who have a child as a teenager. The second column shows the statistics for women who did not have a child before the age of 20.

Table 1. Sample means (standard deviations) and proportions

	Cross section		Within family					
			All siblings		All Twins		MZ Twins	
Older cohort	(1) Teen	(2) Non teen	(3) Teen	(4) Non teen	(5) Teen	(6) Non teen	(7) Teen	(8) Non teen
Years of education 1980	9.9	11.4			9.9	10.3	10.0	10.3
	(1.7)	(2.0)			(1.6)	(1.7)	(1.4)	(1.4)
Years of education 1988	10.0	11.7	10.1	11.1	10.0	10.4	10.1	10.4
	(1.8)	(2.4)	(1.8)	(2.4)	(1.7)	(2.1)	(1.6)	(1.7)
Years of education mother	8.9	9.6	9.0	9.0	8.7	8.7	8.7	8.3
	(2.4)	(2.5)	(2.3)	(2.5)	(2.4)	(2.4)	(2.6)	(2.4)
Years of education father	8.9	9.9	9.1	9.1	8.8	8.6	9.2	8.6
	(2.9)	(3.1)	(2.8)	(2.8)	(2.8)	(2.6)	(2.9)	(2.7)
Age in 1988	39.3	37.9	38.7	37.8	40.3	40.3	40.7	40.7
	(8.6)	(9.3)	(8.4)	(9.0)	(8.7)	(8.7)	(9.1)	(9.1)
Age at first birth	18.5	25.7	18.5	24.5	18.5	23.9	18.5	23.7
	(1.2)	(3.6)	(1.2)	(3.4)	(1.2)	(3.3)	(1.2)	(3.1)
Own birth weight	509.4	503.0	504.5	508.0	507.7	525.5	499.2	520.5
	(175.3)	(146.7)	(151.2)	(159.1)	(156.5)	(167.4)	(148.0)	(159.9)
Birth order	1.6	1.5	1.6	1.4	1.6	1.4	1.6	1.4
N	299	4462	241	355	143	143	80	80
Younger cohort								
Years of education 1989	10.5	12.1			10.7	10.5	11.1	10.3
	(1.7)	(1.9)			(1.7)	(1.4)	(1.4)	(1.3)
Years of education 1996	10.2	12.2			10.5	10.4	10.5	10.4
	(1.7)	(2.4)			(1.8)	(1.8)	(1.8)	(1.9)
Years of education mother	8.9	10.4			8.9	9.3	8.9	8.8
	(2.1)	(2.5)			(2.0)	(2.3)	(2.0)	(1.6)
Years of education father	9.0	10.8			9.0	9.2	8.7	8.6
	(2.7)	(2.9)			(2.7)	(2.8)	(2.3)	(2.4)
Age in 1996	30.3	29.9			30.1	29.9	30.0	29.8
	(2.6)	(2.5)			(2.6)	(2.5)	(2.1)	(2.2)
Age at first birth	18.6	27.0			18.7	25.2	18.7	25.0
	(1.0)	(3.0)			(0.9)	(3.2)	(0.9)	(3.1)
Own birth weight								
Birth order	1.6	1.5			1.6	1.5	1.5	1.5
N	134	2525			104	104	48	48

Starting with the cross sectional comparisons in column (1) and (2) we observe that teen mothers have lower educated parents and are lower educated that non-teen mothers. The differences are larger for the younger cohort. Column (3) to (8) show the within family comparison of teen mothers and their sisters. This comparison eliminates the differences in social background in the first rows of table 1.[2] For the older cohort we observe that the difference in educational attainment between teen mothers and their sisters becomes smaller

[2] The difference in parental education in the samples twins are due to missing values.

when we move to the right. For the younger cohort the difference in educational attainment reduces to 0.1 year in 1996.

4. Empirical findings

4.1 Main estimation results

Table 2 shows the estimated effects of teenage childbearing on several measurements of educational attainment. Column (1) is based on a linear regression of a dummy for teenage childbearing on educational attainment (standard errors are adjusted for clustering within families). Column (2) shows the results after including the education of the parents and age as covariates. Column (3), (4) and (5) show the within-family estimates of a linear regression model for respectively the sample of siblings, twins and identical twins. The model for the sample of siblings also controls for age. Each cell shows the results of a separate regression. The result for the older cohort are shown in the top panel, the results for the younger cohort in the bottom panel.

Table 2. Effect of teenage childbearing on educational attainment

	Cross section		Within family		
Older cohort			Siblings	Twins	MZ
	(1)	(2)	(3)	(4)	(5)
Years of education 1980	-1.551	-1.296		-0.388	-0.338
	(0.130)***	(0.133)***		(0.148)***	(0.152)**
N (groups)	4140	2885		286 (143)	160 (80)
Years of education 1988	-1.777	-1.353	-0.664	-0.457	-0.194
	(0.117)***	(0.123)***	(0.155)***	(0.169)***	(0.174)
N (groups)	4971	4038	559 (204)	256 (128)	144 (72)
Younger cohort					
Years of education 1989	-1.653	-1.473		-0.036	0.500
	(0.160)***	(0.275)***		(0.292)	(0.294)
N (groups)	2812	1687		112 (56)	50 (25)
Years of education 1996	-2.091	-1.777		0.175	0.132
	(0.154)***	(0.297)***		(0.201)	(0.280)
N (groups)	3461	1692		154 (77)	76 (38)

Notes: Controls: age, education of parents, standard errors in brackets.

The cross sectional estimates in column (1) and (2) show that teenage childbearing is associated with a lower educational attainment of 1.3 to 2.1 years. Controlling for the education of the parents and age reduces the estimates. The estimates strongly reduce when family fixed effects are taken into account in column (3), (4) and (5). The estimate for the sample of siblings from the older cohort is about halve the estimate in column (2). Moving to

the sample of twins and identical twins only, furthers reduces the estimated difference between teen mothers and their sisters. The estimate of the difference in educational attainment in 1988 for the sample of identical twins is 0.2 years, which is statistically not significant. For the younger cohort the within twin estimates indicate no difference in educational attainment between teen mothers and their twin sisters. We even find some positive, but statistically not significant, point estimates. These findings suggest that the effect of teenage childbearing on educational attainment is very small or zero, especially for the most recent cohorts. In addition, the within sibling comparisons of educational attainment appears to overestimate the lower educational attainment of teenage mothers.

4.2 Sensitivity analysis

We checked the robustness of these findings by excluding pairs of twins who reported a separation of at least one year in early childhood or pairs with large differences in education . Table 3 shows that the findings are robust for these reductions of the samples.

Table 3 Within twin effect of teenage childbearing on educational attainment excluding twin pairs with large differences in age at first birth or education

	Without twins reporting early separation		Difference in education<= 4 years	
	Twins	MZ	Twins	MZ
Older cohort	(1)	(2)	(3)	(4)
Years of education 1980	-0.397	-0.346	-0.336	-0.342
	(0.152)**	(0.155)**	(0.147)**	(0.154)**
N (groups)	272 (136)	156 (78)	274 (137)	158 (79)
Years of education 1988	-0.426	-0.197	-0.387	-0.197
	(0.174)**	(0.177)	(0.162)**	(0.177)
N (groups)	244 (122)	142 (71)	248 (124)	142 (71)
Younger cohort				
Years of education 1989	-0.071	0.543	0.194	0.500
	(0.311)	(0.319)	(0.253)	(0.294)
N (groups)	98 (49)	46 (23)	108 (54)	50 (25)
Years of education 1996	0.193	0.214	0.027	-0.014
	(0.215)	(0.294)	(0.176)	(0.246)
N (groups)	140 (70)	70 (35)	150 (75)	74 (37)

* standard errors in brackets.

In addition, we analyzed whether the findings are robust for the strict dividing line of the age of 20 for being a teenage mother. This generally used dividing line might be arbitrarily. We investigated cutoff's at the age of respectively 19.5, 19 and 18.5 years. The last cutoff approximately halves the sample used in table 2. Despite the reductions of the sample sizes the estimates appear to be robust for these new dividing lines. For the older cohort the estimates for the cutoff of 18.5 years are nearly the same as those in table 2. For the younger

cohort we observe more negative point estimates when moving to younger groups. However, these estimates are statistically not significant.

Table 4 Within twin effect of teenage childbearing on educational attainment using other definitions of teenage mothers

	Teen mum < 19.5 years		Teen mum < 19 years		Teen mum < 18.5 years	
	Twins	MZ	Twins	MZ	Twins	MZ
Older cohort	(1)	(2)	(3)	(4)	(5)	(6)
Years of education 1980	-0.415	-0.231	-0.324	-0.010	-0.411	-0.148
	(0.168)**	(0.175)	(0.199)	(0.219)	(0.223)*	(0.230)
N	236	134	182	102	146 (73)	88
Years of education 1988	-0.481	-0.195	-0.482	-0.098	-0.418	-0.075
	(0.198)**	(0.200)	(0.224)**	(0.273)	(0.254)	(0.299)
N	208	118	166	92	134	80
Younger cohort						
Years of education 1989	-0.192	0.156	-0.294	-0.267	-0.357	0.000
	(0.320)	(0.277)	(0.381)	(0.361)	(0.261)	(0.337)
N	78	32	68	30	42	22
Years of education 1996	0.095	-0.115	-0.010	-0.125	-0.266	-0.412
	(0.222)	(0.275)	(0.249)	(0.298)	(0.345)	(0.480)
N	116	104	98	48	64	34

4.3 Siblings versus twins

One of the main approaches used in the literature on teenage childbearing is within sibling estimation. Our data enable us to extend this approach to comparing teen mothers with their twin sisters. In addition, for a smaller sample of twins we have measures of educational attainment of both their sibling sister and their twin sister. This information allows us to compare the within sibling approach with the within twin approach for the same sample of teen mothers (table 5).

First, we compare the estimates for the fraternal twins with the estimates for the identical twins. This is shown in column (1) and (2). Next, we compare the within sibling approach with the within twin approach for the same sample of teen mothers.

Table 5. Siblings versus twins

Older cohort	DZ	MZ	Siblings	Twins
	(1)	(2)	(3)	(4)
Years of education 1980	-0.452	-0.338		
	(0.278)	(0.152)**		
N	126	160		
Years of education 1988	-0.795	-0.194	-1.194	-0.815
	(0.310)**	(0.174)	(0.387)***	(0.294)***
N	112	144	128 (54)	108 (54)
Younger cohort				
Years of education 1989	-0.468	0.500		
	(0.461)	(0.294)		
N	62	50		
Years of education 1996	0.218	0.132		
	(0.290)	(0.280)		
N	78	76		

For the older cohort the estimates for the identical twins are less negative than the estimates for the fraternal twins. This holds for both measures of educational attainment. For the younger cohort all estimates are statistically not significant.

Column (3) shows that the smaller sample of twins has a 1.2 years lower educational attainment than their sibling sisters. This difference reduces to 0.8 years when the control group consists of their twin sisters. Hence, the estimates of the within twin approach are approximately one third smaller than the estimates of the within sibling approach. The findings in table 5 show that the adverse outcomes for teen mothers reduce when the control group is expected to be more equal.

5. Conclusions

The effect of teenage childbearing on educational attainment appears to be small or not significantly different from zero, especially for the most recent cohort of twins. We find that within family estimates of the difference in educational attainment are much smaller than cross sectional estimates. For the oldest cohort we find that teenage mothers have 0.5 years less education than their twin sisters, for the identical twins this difference is 0.2 years and statistically not significant. For the youngest cohort we find no difference in educational attainment between teen mothers and their twin sisters. These findings are robust for different age cutoffs and exclusion of pairs of twins who report early separation or a large difference in education. The comparison of the difference between teen mothers and their twin sisters with the difference between the same teen mothers and their sibling sister shows that the former difference is one third smaller.

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