Adolescent Fertility and the Educational Attainment of Young Women

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Analyses based on a sample of 2,795 women interviewed annually from 1979 through 1991 in the National Longitudinal Survey of Youth show that early childbearing lowers the educational attainment of young women. After controls for an extensive set of personal and community characteristics are taken into account, having a child before age 20 significantly reduces schooling attained by almost three years among whites, blacks and Hispanics. Having a child before age 18 has a significant effect only among blacks, reducing years of schooling by 1.2 years.

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Adolescent childbearing has been the focus of intense research activity in recent years, yet the causes and consequences of this phenomenon remain subjects of debate. An important question that has not been conclusively answered is: How do early family responsibilities affect the future opportunities and economic status of young women? After an exhaustive review of the research literature, a panel of the National Research Council concluded:

> Women who become parents as teenagers are at greater risk of social and economic disadvantage throughout their lives than those who delay childbearing until their twenties. They are less likely to complete their education, to be employed, to earn high wages, and to be happily married; and they are more likely to have larger families and to receive welfare.[1]

The assertion that teenage childbearing leads to adverse social and economic outcomes is both intuitively plausible and widely accepted by the public and many social scientists. Challengers of this conventional wisdom question whether early childbearing itself is the primary cause of the adverse outcomes, or whether it is the relatively disadvantaged backgrounds of teenage mothers and other observed and unobserved differences between them and women who avoid early motherhood that lead to those outcomes.[2] If it is the latter, then public policies that focus on reducing teenage pregnancy and childbearing may do little to improve the life chances of many disadvantaged young women. Proving, or disproving, the theory that teenage childbearing undermines a woman's chances of social and economic success in adulthood is essential if policies to reduce poverty and welfare dependence are to be grounded in an appropriate conceptual framework.

This article presents new estimates of the relationship between teenage childbearing and educational attainment, a central issue in the scholarly and public policy debates. Recognizing that adolescent fertility is endogenous with respect to educational attainment because it is likely to be related to the expected costs of and returns from investing in education, we take such endogeneity into account in our analyses. Because the relationship between fertility and education may vary by race and ethnicity, we conduct separate analyses for non-Hispanic whites, non-Hispanic blacks (hereafter referred to as "whites" and "blacks") and Hispanics. Our analyses are based on data from the National Longitudinal Survey of Youth (NLSY).

Recent Research

The ability of young mothers to support themselves and their children is affected by the employment opportunities available to them. These opportunities, in turn, are largely determined by the qualifications young mothers bring to the market. It seems reasonable to assume that caring for young children will conflict with and possibly reduce a woman's investment of time and effort in high school completion, college attendance, postsecondary training and early work experience. If reductions in these early investments occur, they are likely to have profound, long-term consequences for the earnings and employability of the mother and, hence, for the economic well-being of both the mother and her children.

Empirical evidence on the relationship between teenage childbearing and educational attainment that both confirms the conventional wisdom and meets rigorous methodological standards is, perhaps surprisingly, scanty. Path-breaking early research reported large negative effects of early childbearing on educational attainment after controlling for a variety of individual and family background factors.[3] That work, as well as some more recent studies,[4] treated fertility as exogenous to dropping out of high school or other educational decisions, an approach now widely recognized as likely to lead to biased estimates.

Other recent studies have followed one of two improved methodological paths. Some have used an instrumental variables approach to model the reciprocal relationship between fertility and schooling. Marini's sociological research has found that age at first birth significantly affects educational attainment, but that the impact is much smaller than that reported in the earlier literature.[5] Other studies using an instrumental variables approach have revealed no significant effects,[6] casting further doubt on models that treat fertility as exogenous. A hazard model of dropout behavior developed by Olsen and Farkas[7] found that when the endogeneity of pregnancy is accounted for, pregnancy has no effect on dropping out among poor black female high school students. Using a pooled sample of whites, blacks and Hispanics from the NLSY, Ribar found that teenage fertility does not affect the likelihood of dropping out of high school by age 20.[8] In another study using NLSY data, Moore and colleagues reported no effect of age at first birth on highest grade completed for whites and blacks.[9] They did, however, find a significant positive relationship between these variables for Hispanics.

A second set of studies has focused on accounting for unobserved family heterogeneity in estimating the consequences of teenage childbearing. An innovative study

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Table 1. Means or percentages for sample and community variables, by race and ethnicity

Variable	White	Black	Hispanic
ENDOGENOUS†			
Yrs. of schooling at age 25	13.2	12.6	12.1
Birth before age 18 (%)	7	22	15
Birth before age 20 (%)	16	40	30
EXOGENOUS (BOTH MODELS)†			
Mother's education (yrs.)	12.0	10.5	7.5
Father's education (yrs.)	12.3	9.9	7.6
Father's education missing (%)	5	22	15
Living arrangements at age 14			
Mother only (%)	8	33	17
Mother and stepfather (%)	7	7	5
Other (%)	5	11	6
Both parents (%)	80	49	72
Yrs. with mother only	.66	3.40	1.37
Yrs. with mother and stepfather	.51	.72	.73
Yrs. in other living arrangements	.28	.72	.32
Experienced parental divorce (%)	12	18	14
No. of siblings	3.2	5.0	4.8
No. of older siblings	1.9	2.9	2.6
Mother (or surrogate) worked (%)	52	57	46
Foreign born (%)	2	‡	19
Mother foreign born (%)	4	‡	38
Father foreign born (%)	3	‡	36
Foreign language at home (%)	7	3	96
Born in South (%)	24	61	30
Lived in South at age 14 (%)	25	59	29
Urban residence at age 14 (%)	75	80	91
Magazines in home at age 14 (%)	75	39	37
Newspapers in home at age 14 (%)	90	64	52
Library card at age 14 (%)	80	63	68
Employment in state of residence at age 14			
% in services	18	17	19
% in wholesale/retail trade	22	22	23
% in other	60	61	58
Religion			
Baptist (%)	16	63	‡
Other protestant (%)	30	‡	‡
Catholic (%)	32	‡	85
Jewish/other (%)	14	‡	‡
None (%)	8	‡	‡
Attendance at religious services			
Never (%)	17	8	9
Rare (%)	28	21	23
Occasional (%)	19	30	23
Often (%)	36	41	45
EXOGENOUS (FERTILITY MODEL ONLY)			
Age at menarche†	12.9	12.8	12.6
State level§			
Maximum AFDC payment to two-person family††	$211	$162	$205
Restrictive abortion provisions (%)	7	15	5
Restrictive laws on contraceptive sales/advertisements (%)	41	30	56
Restrictions on Medicaid funding of abortion (%)	20	15	51
Maximum % of state median income for eligibility for Title XX family planning services	55	51	55
No maximum for Title XX eligibility (%)	2	12	0
Age of consent for abortion	16.4	16.7	15.0
No age of consent for abortion (%)	64	48	48
Age of consent for contraception	16.5	16.6	14.8
No age of consent for contraception (%)	69	62	50
County level‡‡			
No. of abortions per 1,000 women	26.9	44.5	37.9
Abortion provider performing more than 400 abortions (%)	52	62	79
Presence of abortion clinic (%)	54	61	77
Presence of hospital abortion provider (%)	66	68	90
Presence of physician abortion provider (%)	47	61	70
Presence of Planned Parenthood clinic (%)	55	47	77
% of women 15–19 using family planning services	14	17	14
% of women 15–44 using family planning services	8	10	10
% of family planning patients aged 15–19	37	33	28
No. of marital births per 1,000 women aged 15–19	371	605	428
No. of nonmarital births per 1,000 women aged 15–19	16	91	30
No. of women aged 15–19 per family planning clinic	4,019	3,099	3,021
No. of patients aged 15–19 per family planning clinic	517	442	382
No. of patients aged 15–44 per family planning clinic	1,452	1,382	1,367

†Data from National Longitudinal Survey of Youth (NLSY), based on 1,565 whites, 952 blacks and 493 Hispanics. ‡Insufficient number of cases. §Data prepared for the United States Department of Health, Education and Welfare (DHEW) by The Alan Guttmacher Institute (AGI). ††Data provided by DHEW. ‡‡Data provided by AGI.

by Geronimus and Korenman used three major data sets—the NLSY, the Panel Study of Income Dynamics (PSID) and the National Longitudinal Survey of Young Women—to compare the experiences of sisters whose first births occurred at different ages.[10] Comparing sisters controls for unobserved heterogeneity across families and, the authors argued, yields estimates (from "fixed-effect" regressions) likely to be more reliable than the cross-sectional findings typically reported in the literature. Hoffman, Foster and Furstenberg replicated this study using a different PSID sample.[11] Neither study disaggregated the sample by race because of the small number of sibling pairs typically obtained for such analyses of qualitative outcomes.

In all four sibling samples, cross-sectional regressions showed that early childbearing reduced both the probability of completing high school and the probability of obtaining any postsecondary schooling. In contrast, the fixed-effect approach yielded mixed results: The effect of early childbearing on these educational outcomes was in some cases insignificant and in other cases substantially less than in the equivalent cross-sectional analyses. The concern that unobserved family heterogeneity biases upward the estimated effects of early childbearing appears warranted, yet significant negative effects persist in most of the samples.

Grogger and Bronars[12] used census data to compare socioeconomic outcomes of women experiencing twin first births to those of women experiencing single first births. Twin births are exogenous events with respect to unobserved variation in opportunity costs and in preferences for childbearing. With this rather different approach to dealing with unobserved heterogeneity, Grogger and Bronars found that teenage childbearing has no significant effect on the likelihood of high school graduation among whites but has significant negative effects on the likelihood of graduation among blacks.

None of these methodologically clever studies takes the endogeneity of fertility into account. Nor do they control for differences between sisters in many personal characteristics that may affect educational attainment and childbearing (e.g., attitudes toward school and parenthood). Hence, these three studies may not have estimated unbiased effects of teenage childbearing.

Data and Methods

The data were obtained from the NLSY, other public sources and The Alan Guttmacher Institute (AGI). In 1979, the

NLSY interviewed 12,686 males and females who had been between the ages of 14 and 21 on January 1 of that year. Blacks, Hispanics and economically disadvantaged whites were oversampled. The NLSY reinterviewed the respondents annually through 1991.

The sample for this analysis includes all women aged 14 to 20 in 1979 except for those in the special military subsample or the oversample of economically disadvantaged whites. We conduct separate analyses for whites, blacks and Hispanics because preliminary empirical results and prior research have shown that results vary substantially by race and ethnicity. After exclusion for missing values, our sample includes 1,445 whites, 906 blacks and 444 Hispanics.*

Our measure of educational attainment is completed years of schooling at time of interview in the year the respondent reached age 25. Reductions in investments in schooling during the teenage years because of the demands of parenting may be partially replaced by later investments. By examining education levels at age 25, when most people will have completed their formal schooling, or at least will have begun college if they intend to do so, we capture most delayed (as opposed to permanently foregone) investment in schooling. Given the age range of the respondents, schooling at age 25 is measured during the 1984–1990 period. In instances in which the education measures were missing for the interview year in which respondents reached age 25, we substituted outcome measures recorded at the time of interview in the year they turned age 26. We follow this strategy to reduce potential bias due to nonresponse and because education is relatively stable at these ages.

We measure the occurrence of an early birth in two ways: whether the respondent had had a birth before her 18th birthday, and whether she had had a birth before her 20th birthday. Comparing findings for the two measures will indicate the importance of the timing of a teenage birth. The number of Hispanics who had had a child before they reached age 18 was too small to estimate fertility regressions, so for Hispanics, we present only the effects of having a birth before age 20.

The education equations include a large set of personal and family background characteristics as well as measures of employment opportunities. Table 1 lists all of the variables used and their means. Personal and family background variables include highest school grade completed by mother and father, a set of variables specifying the living arrangements the respondent experienced as a child, number of siblings and of older siblings, whether the household included an adult female working for pay when the respondent was aged 14, whether the respondent or her parents were born outside the United States, whether the respondent was born in the South, whether the respondent lived in the South or an urban area at age 14, whether a language other than English was spoken in the respondent's home when she was aged 14, whether her household subscribed to magazines or newspapers, whether anyone in her household had a library card, the respondent's religious affiliation and frequency of attendance at religious services.

We measure employment opportunities open to adolescents by the percentage of workers employed in services and in wholesale and retail trade for the state in which the respondent lived at age 14. In early regressions, the family background variables also included the ratio of family income to the poverty line. Because it did not have a significant effect in any of the three groups and because many cases lacked income data, we exclude it in the results reported in this article.

To test whether teenage childbearing affects educational attainment, we include dummy variables for early fertility in a regression model of years of schooling completed. The primary estimation issue raised by this procedure is the potential endogeneity of fertility. Through abstinence and the use of contraception, adolescents can control the likelihood that they will experience a pregnancy and through abortion determine whether a pregnancy is carried to term. Consequently, if adolescents perceive that childbearing will affect their schooling and work opportunities, fertility will be determined jointly with those outcomes. To control for this potential source of bias, we estimate the impact of teenage childbearing on schooling using an instrumental variables approach.

We first estimate a probit model of the probability that a young woman had a teenage birth. The model includes all exogenous variables listed in Table 1. Using the estimated parameter values and the regressors, we calculate the predicted probability of a teenage birth for each woman in the sample. We then estimate an instrumental variables regression for years of education, with the predicted probability that a woman had a teenage birth included among the variables for endogenous fertility. For comparison purposes we also report results from ordinary least squares education models.

The instrumental variables models permit fertility to be determined jointly with education. The effect of fertility on education is determined by exclusion restrictions: A set of variables included in the fertility equation is excluded from the education equation. These variables are shown in the lower part of Table 1 along with their means and sources. One such variable—age at menarche—is an individual characteristic likely to affect fertility but not educational attainment. This category also includes state policy variables likely to affect fertility, such as the maximum Aid to Families with Dependent Children (AFDC) payment for a family of two, the presence of restrictive abortion provisions, and the ages at which parental consent is no longer needed for a young woman to have an abortion or use contraceptives. These variables are measured for the state in which the respondent resided at age 14, an age at which residential location can be regarded as exogenous. As a woman matures, she may be more likely to choose her residence partly on the basis of the state policy environment.

The set of variables omitted from the education equation also includes indicators of the availability of abortion and family planning services and of the social context within which fertility decisions are made. A substantial body of research shows that such variables exert important influences on fertility.[13] We measure these variables for the county in which the respondent was living at the time of interview in 1979 (or in 1980 if data were not available for 1979). For these variables, we would prefer to use measures taken at a uniform early age (as we did for the state-level ones), but county of residence was not available for years prior to 1979. The county-level variables include the abortion rate, whether there is an abortion clinic performing more than 400 abortions, whether there are any Planned Parenthood clinics, marital and nonmarital fertility rates for women aged 15–19, the proportion of women aged 15–19 using family planning services, and similar variables. The appendix summarizes results from the fertility probit regressions.

We apply two types of specification tests to the instrumental variables models—Hausman tests of the first-stage probit specification and tests of overidentifying restrictions. In the Hausman tests, we compare the estimates produced by a consistent instrumental variables estima-

*Tables 1 and 2 are based on a slightly larger sample that includes all observations not missing values for education and fertility.

Table 2. Measures of educational attainment at age 25, by race and ethnicity, according to age at first birth, NLSY

Race and measure of attainment	<18	18–20	>20
White	(N=128)	(N=178)	(N=1,597)
Mean yrs. of schooling completed	10.7	11.5	13.5
% graduated from high school	28.9	60.1	91.5
% with some college education	6.3	9.6	51.3
Black	(N=264)	(N=215)	(N=753)
Mean yrs. of schooling completed	11.4	12.0	13.3
% graduated from high school	50.0	72.1	89.6
% with some college education	14.4	19.1	52.9
Hispanic	(N=102)	(N=122)	(N=460)
Mean yrs. of schooling completed	9.7	11.3	12.4
% graduated from high school	21.6	54.9	76.3
% with some college education	8.8	11.5	42.4

tor (linear two-stage least squares) with those produced by an estimator efficient under the null hypothesis of no misspecification but inconsistent under the alternative that the model is misspecified (instrumental variables with a first-stage probit).[14] All the models passed this test at a 95% confidence level with the exception of the model for white births before age 20. Evidence that the fertility probit is misspecified in this case calls into question the consistency of the estimated parameters in the second stage, so we have reported linear two-stage least squares estimates for this model.

The tests of overidentifying restrictions are tests of the joint hypothesis that the model is correctly specified and that the instrumental variables are valid. One possible reason for these restrictions to be rejected is that one of the excluded variables in fact belongs in the education equation, or is correlated with an unmeasured determinant of educational attainment. We have applied a test that involves regressing the estimated residuals* from the education equation on the full set of instruments.[15] If the instrumental variables are valid, they should have no explanatory power in this regression, and a simple chi-square test is appropriate. In four of our five models, passing this test required that we exclude a small number of variables (these exclusions are noted in the footnotes to Table 3), but these modifications did not substantively change our estimations of the effects of early childbearing.

Results

Table 2 presents bivariate relationships between adolescent childbearing and educational attainment. Consistent with our expectations and with other descriptive findings, the data reveal a strong positive relationship between age at first birth and schooling. Whites who had a child before reaching age 18 completed an average of 10.7 years of schooling. Those who had no children until they were aged 18 or 19 or until they were aged 20 averaged 11.5 and 13.5 years of schooling, respectively. Mean educational levels for blacks and Hispanics exhibited the same pattern, although the average level of schooling completed by black teenage mothers was somewhat higher than that of whites, and the schooling level of Hispanics in each fertility category was lower.

High school completion rates also varied with fertility. Only 29% of whites who had a child before reaching age 18 completed high school, compared with 60% of those who did not have their first birth until they were aged 18 or 19 and 92% of those who had no children until after age 20. The table suggests that early childbearing has less of an impact on blacks: Fifty percent of those who had a child before they were 18, and 72% of those who did not have a birth until they were aged 18 or 19, completed high school. However, these completion rates fall well below the 90% rate for black women who did not become mothers until they were aged 20 or older. Hispanics exhibited the lowest rate of high school completion in each fertility category. The observed association between teenage childbearing and high school completion among Hispanics was similar to that observed for whites.

Table 2 shows an even stronger relationship between age at first birth and the likelihood of attending college. Among whites, women who began childbearing after age 20 were five to eight times as likely as teenage mothers to attend college. Among blacks and Hispanics, women who began childbearing after age 20 were three to five times as likely to attend college as were their counterparts who became mothers during their teenage years.

Table 3 shows the effects of adolescent fertility on educational attainment estimated using ordinary least squares regressions, which do not control for the endogeneity of fertility, and instrumental variables regressions, which do. With one exception, the models explain 15–40% of the variance in completed years of schooling, a range typical for models of educational attainment. All ordinary least squares estimates (columns 1 and 3) show a significant and large negative impact of early childbearing on years of schooling completed by age 25. Among whites and Hispanics, the difference is roughly 1.5 years; among blacks, it is about 1.1 years.

Four of the five instrumental variables models show that early childbearing has a statistically significant adverse impact on educational attainment, even after taking account of the endogeneity of early fertility. All of the significant coefficients are larger than in the corresponding ordinary least squares regression. For whites, having a child before age 20 reduced schooling by 2.8 years. However, having a child before age 18 (an event experienced by less than 7% of the sample) had no significant effect. Among blacks, a birth before age 18 was associated with a decrease in schooling of one year. Among both blacks and Hispanics, a birth before age 20 had a significant negative effect—reducing educational attainment by nearly three years for each group. Our estimates do not suggest that childbearing before the usual age of high school completion creates a major obstacle to educational attainment for

*The difference between actual educational attainment and educational attainment predicted using the instrumental variables coefficients.

Table 3. At age 25, years of schooling lost (and t-statistics) as a result of a teenage birth, by race and ethnicity, according to age at first birth and whether controlled for endogeneity

Race	<18		<20	
	Uncontrolled†	Controlled‡	Uncontrolled†	Controlled‡
White	–1.562* (–8.502)	–0.436 (–0.466)	–1.470* (–11.790)	–2.766*§ (–2.959)
Black	–1.015* (–7.430)	–1.234* (–2.121)	–1.193* (–10.569)	–2.971* (–4.041)
Hispanic	††	††	–1.467* (–6.807)	–2.831* (–3.676)

*Significant at p=.10. †Ordinary least squares equations include all variables listed in Table 1 for both models with the following exclusions because of small cell sizes: **Black sample**—Foreign birth for self, mother or father, all religious categories except Baptist (reference category is all others); **Hispanic sample**—All religious categories except Catholic (reference category is all others). ‡Instrumental variables equations include all exogenous variables in the education model, plus all variables listed in Table 1 for the fertility model, with the following exclusions to avoid rejection of overidentifying restrictions: **White (<18)**—County marital and nonmarital fertility rates for women aged 15–19; **White (<20)**—County nonmarital fertility rate for women aged 15–19; **Black (<18)**—Title XX eligibility, proportion of women aged 15–19 and proportion of women aged 15–44 using family planning clinics, proportion of family planning patients aged 15–19, presence of abortion facility providing ≥400 abortions; **Black (<20)**—Title XX eligibility, proportion of women aged 15–19 and proportion of women aged 15–44 using family planning clinics. §Estimated as linear two-stage least squares because of evidence of misspecification in fertility probit. ††Insufficient number of cases for analysis.

Table 4. At age 25, years of schooling lost or gained (and t-statistics) as a result of selected background variables, according to race and ethnicity

Variable	White (N=1,445)	Black (N=906)	Hispanic (N=444)
Birth before age 20	–2.766* (–2.959)	–2.971* (–4.041)	–2.831* (–3.676)
Mother's education	0.186* (6.107)	0.092* (2.953)	0.107* (3.075)
Father's education	0.150* (7.508)	0.081* (3.806)	0.039 (1.279)
Father's education missing	1.612* (5.390)	0.483* (1.879)	0.438 (1.188)
Living arrangements at age 14			
Mother only	–0.104 (–0.440)	0.151 (0.766)	0.247 (0.689)
Mother and stepfather	–0.295 (–1.048)	0.152 (0.555)	–0.051 (–0.088)
Other	–0.273 (–0.906)	0.220 (0.772)	–0.972* (–1.917)
Experienced parental divorce	–0.123 (–0.643)	–0.004 (–0.023)	–0.405 (–1.197)
Yrs. with mother only	–0.015 (–0.528)	–0.013 (–0.864)	–0.011 (–0.295)
Yrs. with mother and stepfather	–0.002 (–0.073)	–0.013 (–0.506)	–0.018 (–0.368)
Yrs. in other living arrangements	–0.051 (–1.559)	–0.012 (–0.431)	0.104* (1.711)
No. of siblings	–0.105* (–2.800)	0.028 (0.761)	–0.060 (–1.100)
No. of older siblings	0.075* (1.946)	–0.050 (–1.276)	0.061 (1.015)
Mother (or surrogate) worked	–0.065 (–0.701)	–0.007 (–0.057)	0.089 (0.430)
Foreign born	–0.523 (–1.439)	†	–0.649* (–1.963)
Mother foreign born	0.573* (2.160)	†	1.055* (3.715)
Father foreign born	0.595* (2.168)	†	–0.225 (–0.769)
Foreign language at home	0.578* (2.851)	0.340 (0.905)	–0.006 (–0.012)
Born in South	0.110 (0.611)	–0.019 (–0.095)	0.003 (0.009)
Lived in South at age 14	0.034 (0.183)	0.328 (1.536)	0.541 (1.501)
Urban residence at age 14	–0.129 (–1.179)	–0.094 (–0.551)	–0.235 (–0.637)
Magazines in home at age 14	0.166 (1.230)	0.106 (0.645)	–0.003 (–0.011)
Newspapers in home at age 14	0.068 (0.412)	0.077 (0.552)	0.378* (1.648)
Library card at age 14	0.148 (1.229)	0.140 (0.961)	0.343 (1.506)
Employment in state of residence at age 14			
In services (%)	0.008 (0.304)	–0.026 (–1.009)	0.008 (0.118)
In wholesale/retail trade (%)	–0.022 (–0.795)	0.027 (1.007)	–0.091 (–1.030)
Religion			
Baptist	–0.156 (–0.727)	–0.111 (–0.869)	†
Other protestant	0.315* (1.720)	†	†
Catholic	–0.143 (–0.738)	†	0.067 (0.232)
Jewish and other	–0.045 (–0.218)	†	†
Attendance at religious services			
Rare	0.218 (1.301)	–0.246 (–0.948)	–0.148 (–0.363)
Occasional	0.369* (2.227)	–0.128 (–0.503)	–0.121 (–0.299)
Often	0.640* (3.984)	0.023 (0.081)	0.111 (0.290)
Constant	*9.547* (11.393)*	*11.831* (13.129)*	*13.373* (5.529)*
Adjusted R^2	*0.373*	*0.118*	*0.179*

*Coefficient is significant at p=.10. †Insufficient number of cases for analysis.

young white women but, since most teenage childbearing in this group occurs at age 18 or 19, our sample includes very few early childbearers. Our estimates of the effects of teenage childbearing are among the largest reported in the literature and are remarkably consistent across racial and ethnic groups.

Our instrumental variables results agree with the findings of Moore and colleagues that teenage childbearing has significant effects among Hispanics, but their results did not show significant effects among whites or blacks. Our finding of significant effects for all three groups also conflicts with the results of studies by Ribar and by Olsen and Farkas, both of which took the endogeneity of fertility into account. These differences probably result from our use of a far more extensive set of identifying variables with substantial predictive power. Moore and colleagues used six identifying variables, Ribar used three, and Olsen and Farkas used only one. In exploratory work with a more limited set of policy and contextual instrumental variables, we found little evidence of a significant fertility effect in any group. This suggests that an inability to predict fertility well may be responsible for a failure to find significant effects on schooling.

Table 4 shows the full instrumental variables regression estimates with a birth before age 20 as the fertility variable. The estimated effects of demographic control variables are not the focus of this article, but bear brief mention. In all groups, a mother's education exerted a major influence on her daughter's years of schooling. The father's education was also important among whites and blacks, although the coefficient is smaller. Living arrangements and having a working mother had surprisingly little direct effect on schooling in any group. Such variables are important determinants of early childbearing, which may have mediated their effects on schooling. Among whites only, the amount of schooling declined as the number of siblings rose, but it increased with the number of older siblings. Hispanic women who were foreign born tended to complete fewer years of schooling than their counterparts born in the United States, but those whose mother was foreign born tended to complete more years of schooling. Whites with foreign-born parents also tended to complete more schooling. The frequency of attendance at religious services had a strong positive relationship with schooling among whites.

Conclusion

This study, which used an instrumental variables approach to estimate the effect of adolescent fertility on education, shows that early childbearing has large negative effects on young women's years of schooling after accounting for the endogeneity of fertility. Other recent research has reported that the social and economic effects of teenage childbearing are not as great as early studies of the relationship between teenage childbearing and adult outcomes had suggested. The results in this study, however, suggest that the "revisionist" findings, while methodologically superior to the early research, are open to challenge.

After controlling for both observed and unobserved differences in background and personal characteristics, we found that early childbearing reduced the educational attainment of young women by one to three years. These strong negative effects held for white, black and Hispanic women. Our results suggest that public policies that succeed in reducing teenage pregnancy and childbearing would also increase the educational attainment of disadvantaged young women and improve their chances for economic self-sufficiency.

Appendix

The fertility probits have substantial explanatory power and predict mean probabilities of early childbearing well. Table A-1 reports chi-square statistics for the explanatory power of the model as a whole, and for the incremental explanatory power of the excluded instruments. The fertility model as a whole is significant at a 99% confidence level in all cases. The null hypothesis that the entire set of excluded instruments has no explanatory power can be rejected at a significance level of 30% or less for all models, but in only two of the models does the joint test approach conventional significance levels.

All of the equations contain at least two significant instrumental variables—age at menarche was significant in four of the five models, the county abortion rate was significant in three, and the ages of consent for abortion and contraception were important determinants of early fertility in most models. Abortion provider and family planning service variables were occasionally, but not consistently, significant. Few of the instrumental variables measured at the state level, such as AFDC benefits, were significant predictors of early childbearing. The exceptions were the age of consent for abortion and contraception and,

Table A-1. Chi-square values for rejection of null hypothesis, by race and ethnicity

Race/ethnicity	All variables		Excluded variables	
	Birth <18	Birth <20	Birth <18	Birth <20
White				
Chi-square	134.9	214.6	26.1	28.1
Significance	0	0	0.30	0.30
DF	55	57	23	25
Black				
Chi-square	154.0	136.4	37.7	23.8
Significance	0	0	0.01	0.30
DF	45	47	19	21
Hispanic				
Chi-square	u	78.7	u	32.6
Significance	u	0.01	u	0.11
DF	u	53	u	24

Note: u=unavailable

for the Hispanic sample only, restrictions on Medicaid funding of abortions.

In addition to the instrumental variables, mother's education, some living arrangements variables, magazine subscriptions and attendance at religious services were significant in most of the fertility equations. For the black sample only, number of siblings, birth order and urban residence also had sizable effects on the probability of early childbearing.

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