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Discussion Papers

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DP # 926-90

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September 1990

The authors share equal responsibility for this paper. Lundberg was a visitor at the Industrial Relations Section, Princeton University, and Plotnick was a visiting scholar at the Russell Sage Foundation when it was written. We thank Harriet Stranahan for research assistance. Kristin Moore and participants in workshops at the Russell Sage Foundation, Columbia University, the University of Michigan and the Alan Guttmacher Institute provided helpful comments on earlier versions of this paper. This study was supported in part by grant 88ASPE195A from the U.S. Department of Health and Human Services, Office of Assistant Secretary for Planning and Evaluation. The opinions expressed in this paper are the authors' and do not necessarily reflect those of DHHS. An earlier version of this paper was presented at the May 1990 Population Association of America conference in Toronto, Ontario.

Abstract

This study develops an empirical model of adolescent premarital childbearing which emphasizes the influence of opportunity costs. The model estimates determinants of premarital pregnancy, the choice to abort or carry to term, and whether a marriage occurs before the birth. The sample is from the National Longitudinal Survey of Youth.

The long-run opportunity costs are the predicted effects of premarital childbearing on own future wages and welfare benefits. State variables on abortion and family planning policy and availability, which are proxies for the costs of abortion and avoiding pregnancy, represent short-run costs.

For white adolescents, the long-run wage measure has statistically significant effects on abortion and pregnancy outcomes that are consistent with theoretical expectations. Their behavior also is associated with welfare, abortion, and family planning policy variables in directions consistent with an opportunity-cost model of behavior. Black adolescents' behavior shows no association with the opportunity-cost or policy variables. This may be a function of sample size. It may also be that there are important unmeasured racial differences in the factors that influence fertility and marital behavior.

ADOLESCENT PREMARITAL CHILDBEARING: DO OPPORTUNITY COSTS MATTER?

1. INTRODUCTION

Teenage childbearing emerged as an issue of national importance during the 1980s, and continues to be the focus of confused and heated debate among policymakers, researchers, and the general public. Providers of health and educational services have responded to the public's sense of a crisis in teenage fertility, treating the problem as one of imperfect information or irrational behavior. Sex education, family life classes, and "life option" and self-esteem programs in schools and community centers seek to improve either teens' knowledge of their alternatives or their decisionmaking skills.

An alternative, but not necessarily incompatible, approach treats premarital childbearing as a response to the incentives and constraints facing teenagers, particularly black and disadvantaged teenagers. Geronimus (1987) argues: "Policies that do not attempt to alter this social reality, but aim only to affect directly the fertility behavior of those teenagers subject to it, are likely to fail as they are counteracted by the incentives to early childbearing to which these teenagers respond" (p. 266).

What are the incentives that encourage some teens to become single parents and deter others? Teenage childbearing has been associated with a number of adverse social and economic consequences, including fewer years of schooling, a higher risk of marital disruption, lower earnings and family income, and a higher risk of poverty and welfare receipt. Care of a young child is a time-consuming responsibility that diverts time and energy away from other activities, including school attendance and, in the short run, market work. This reallocation of time is likely to have long run consequences, since school and work are investment activities that increase market productivity and wages. Future husband's income may also be affected by early fertility if market productivity and wages. Future husband's income may also be affected by early fertility if the presence of a child born out-of-wedlock, or an early marriage with a high probability of dissolution, alters the pool of potential spouses.

Is it possible that these consequences also play crucial causal roles in leading some young women to become, and others to avoid becoming, unwed mothers? Other things equal, perhaps girls who perceive that early motherhood will entail little loss of long run earnings and marriage opportunities see little reason to avoid becoming an adolescent parent, while those who expect early motherhood to greatly damage their labor market and marriage prospects will take steps to avoid this fate. Anderson's (1989) ethnographic research leads him to conclude "middle-class youths take a strong interest in their future and know what a pregnancy can do to derail that future. In contrast, the ghetto adolescent sees no future to derail, no hope for a tomorrow very different from today, hence, little to lose by having an out-of-wedlock child" (p. 76). Wilson (1987) argues that rates of black premarital childbearing are high, in part, because adolescent black females face such poor marriage opportunities that they sacrifice little in the way of long run income by not postponing motherhood. Set against this low cost are the benefits of childbearing, which may include acceptance as an adult member of the community. This perspective on the determinants of early and premarital childbearing is the "opportunity cost" or "nothing to lose" hypothesis.

The opportunity cost hypothesis is intuitively plausible and consistent with some anecdotal and journalistic evidence (e.g., Dash, 1989), but has just begun to receive careful study. Lundberg and Plotnick's (1990) estimates of the future wage losses associated with early and premarital childbearing show moderate reductions in future wages for white teenage mothers, but none for blacks. In fact, there are significant wage premia for black teenage mothers from disadvantaged backgrounds. These findings are consistent with the observed racial and socioeconomic patterns of premarital childbearing.

While the research literature on adolescent sexual behavior, pregnancy, and childbearing is extensive (see Hofferth and Hayes, 1987, for a comprehensive review), few studies have adopted an economic approach.[1] Most empirical studies of premarital childbearing have been grounded in sociological and psychological models of behavior. They focus on the effects of personal and family background variables rather than cost or policy variables.[2] Duncan and Hoffman (1990) begin a more sophisticated exploration of the economic approach. They find that a predicted measure of taxable family income in early adulthood affects the probability that a black teenager will have a premarital birth and receive AFDC. They do not, however, find a significant effect of AFDC benefits on this outcome. Their results suggest that further examination of opportunity costs is likely to be fruitful and informative for policy purposes.

We have developed an empirical model of premarital childbearing which emphasizes the influence of opportunity costs. The model explicitly recognizes that a teenager's route to single motherhood involves a multistage process of choices. This process is marked by three major decision points which fall logically into a hierarchical order: becoming pregnant; given a pregnancy, the choice to abort or carry to term; and given the choice to carry to term, the outcome of having the birth premaritally or marrying sometime before the birth.[3] We examine separately the factors affecting the probability of a premarital pregnancy, and those affecting its resolution.

Figure 1 illustrates the sequence of choices and outcomes we analyze in this paper. Since the abortion decision is conditional on the occurrence of a pregnancy, and the marriage decision is conditional on continuing the pregnancy, the three stages of the decision process must be jointly estimated. A three stage nested logit model is a natural candidate for estimating the determinants

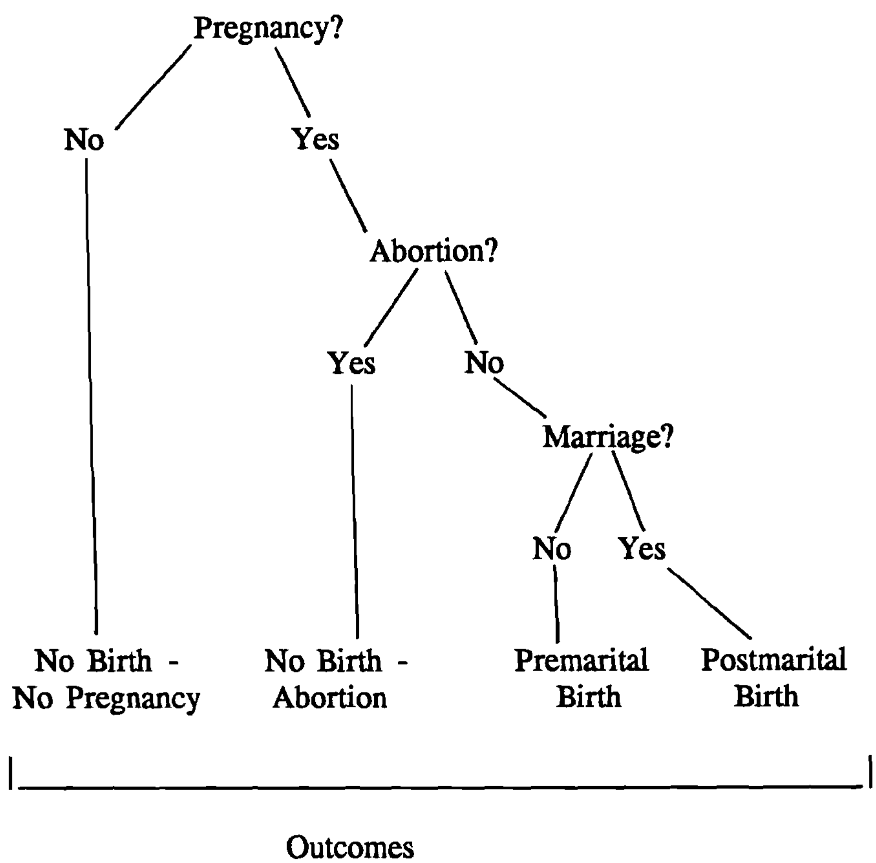

Figure 1. Sequential Decisionmaking in Premarital Childbearing

of such a series of outcomes. The model uses data for the youngest cohorts of the National Longitudinal Survey of Youth (NLSY).

An economic approach suggests that the choice at each stage of the process will be influenced by both expected long run opportunity costs and short run costs. One long run cost we focus upon is the predicted effect of premarital childbearing on own future wages. A second is the level of welfare benefits. Higher benefits reduce the financial costs of an out-of-wedlock birth and so may affect both marital and fertility behavior. Other public policies can affect the short run costs of using contraceptives or obtaining an abortion -- actions that affect the likelihood of pregnancy or giving birth. The model, therefore, includes measures of state abortion and family planning policy and availability.

Racial differences in teenage fertility and marital behavior are substantial. In 1987 the black teenage birth rate was twice the white rate, while the birth rate to unmarried black teens was more than four times the white rate. Birth rates for unmarried black teens have fallen in recent years, but because they have fallen less rapidly than the birth rates for married black teens, the proportion of births to black teens which are out-of-wedlock rose to .91 in 1987 from .64 in 1970. The proportion of births to white teens which are out-of-wedlock was .51 in 1987, a sharp rise from the 1970 value of .18. Because these racial patterns are of great interest, we analyze separately non-Hispanic whites and non-Hispanic blacks (hereafter referred to simply as whites and blacks).[4]

Section 2 discusses the data and the statistical model. Section 3 presents the estimates of the nested logit model and some illustrative cases. Section 4 is a brief conclusion.

2. AN EMPIRICAL MODEL OF PREMARITAL CHILDBEARING

A. The Sample

The sample consists of 1718 black and white girls in the NLSY who were age 14 to 16 in 1979, and who have sufficient information in their fertility and marital records that we could identify whether and when they first became premaritally pregnant, whether they aborted or carried to term, and whether they married between conception and delivery. We followed their fertility and marital histories from 1979 through 1986, but examine only first premarital pregnancies in this study.[5]

Table 1 presents basic statistics on the fertility and marital behavior of the sample, adjusting for sampling weights. The pregnancy rate among whites was 24 percent. Of these, 11 percent ended by miscarriage or still birth. Of the pregnancies that continued, 37 percent were aborted. Of the live births, 48 percent were to unmarried women; the others married between conception and birth. The net result was a premarital birth rate of 6.6 percent. The corresponding figures for blacks were 48, 12, 12, 94 and a net premarital birth rate of 34.8 percent.

The reported abortion rate for whites is about 90 percent of the actual teenage abortion rate derived from Vital Statistics data and information compiled by the Alan Guttmacher Institute (Moore, 1989). Thus, the model of white abortion choice is unlikely to suffer from any problem associated with underreporting. Since abortions are not seriously underreported, neither are pregnancies, so the pregnancy model also should not suffer from such problems. The rate of marriage among whites in the NLSY who become premaritally pregnant and do not abort is slightly higher than that obtained from the Current Population Survey for 1980-1981 (reported in Hofferth and Hayes, 1987, p. 450).

Table 1

Pregnancy, Abortion, and Marital Outcomes of Samples
(weighted observations, in thousands)

	Whites	Blacks
Number of persons	4543.4	805.7
Premarital pregnancies	1103.0	388.2
(percentage)	(24.3)	(48.2)
Less stillbirths and miscarriages	122.1	48.2
= Pregnancies to be resolved by choice	980.9	340.0
Abortions	363.7	40.5
(percentage)	(37.1)	(11.9)
Live births:	617.2	299.5
Mother married before birth	321.0	19.2
Mother did not marry before birth	296.2	280.3
(percentage)	(48.0)	(93.6)
Percentage of all cases with out-of-wedlock birth	6.6	34.8

Source: Tabulations of NLSY.

The reported abortion rate for blacks, however, is only about 30 percent of their actual rate. Since the reported stillbirth and miscarriage rate nearly matches that based on medical records, it appears that few black abortions are being reported as one of these socially more acceptable terminations. They are simply not being reported in the NLSY. The consequence is a reduction in the number of cases in the black abortion model, and an especially large reduction in the number with the abortion outcome. This may hinder estimation of the abortion model and, because unreported abortions are also unreported pregnancies, the pregnancy model. Underreporting may lead to biased estimates of the behavioral parameters if the choice to underreport is systematically correlated with the variables in the model, but we are unaware of evidence on such correlations (after controlling for race and marital status, as here). The rate of marriage among blacks in the NLSY who become premaritally pregnant and do not abort is nearly the same as that obtained from the Current Population Survey.

B. The Nested Logit Model

The hierarchical sequence of outcomes illustrated in Figure 1 fits naturally into a three stage nested logit framework. One can estimate a nested logit model by considering the decision stages sequentially. Denoting these as i (pregnancy), j (abortion), and k (marriage), the probability of a final outcome is

$$P_{ijk} = P_{k|ij} * P_{j|i} * P_{i}.$$

Let the value of a final outcome be

$$U_{ijk} = V_{ijk} + e_{ijk}$$

where V_{ijk} is a function of measured characteristics and e_{ijk} is a residual. In nested logit, the e's are assumed to have a generalized extreme value distribution rather than the independent extreme value distribution of multinomial and conditional logit. This specification permits correlation among the residuals of alternative choices and, thus, avoids the independence of irrelevant alternatives property of standard logit models.

In our model there are four possible final outcomes: no pregnancy, abortion, marital birth, and premarital birth. In a standard conditional logit framework, their values can be specified as follows:

No teenage pregnancy	$V_N = aX_N$
Pregnancy - Abortion	$V_A = aX_A + cZ$
Pregnancy - Carry to Term - Marriage	$V_M = aX_M + bY + cZ$
Pregnancy - Carry to Term - Single	$V_S = aX_S + bY + cZ$

where Y represents variables common to the outcomes in which a pregnancy is carried to term, and Z is a vector of variables common to all outcomes beginning with a teenage pregnancy.

The model can be estimated in three stages, beginning with the lowest level and working up. This sequential procedure is simple to compute and yields consistent, though inefficient, estimates of the parameters. In the first stage the probability of marriage (M), conditional on a pregnancy (P) that is carried to term (C), is a function of only the parameters in vector a.

$$Prob\ (M \mid P, C) = \frac{exp[\ aX_M/(1-f)\ ]}{exp[\ aX_M/(1-f)\ ] + exp[\ aX_S/(1-f)]}$$

where f is an index of similarity of the choices at this level of the tree (Maddala, 1983) and the parameters in vector a can be estimated only up to a scale factor $1 - f$.

The results from the marriage level of the decision tree are used to calculate an inclusive value for the carry-to-term option:

$$I_C = \ln\left(exp[\ aX_M/(1 - f)\ ] \ + \ exp[\ aX_S/(1 - f)]\ \right).$$

Then, from the abortion-carry decision, we estimate the parameters in vector b. The conditional probability of carrying to term, conditional on pregnancy, is

$$Prob\ (C \mid P) \ = \ \frac{exp[\ bY + (1 - f)I_C\ ]}{exp[\ bY + (1 - f)I_C\ ] \ + \ exp[\ aX_A\ ]} \ .$$

Since the parameter f is approximately equal to the correlation between e_M and e_S (the errors in the marriage and single values), the coefficient on the inclusive value should lie between zero and one. If it equals one, the model reduces to a standard multinomial logit, so our particular nesting structure can be tested against this alternative.

From the second stage we calculate a new inclusive value for the joint pregnancy outcomes:

$$I_P = \ln\left(exp[\ bY + (1 - f)I_C\ ] + exp\ [\ aX_A\ ]\ \right).$$

At the highest level of the decision tree, the unconditional probability of pregnancy is

$$Prob\ (\ P\ ) \ = \ \frac{exp[\ cZ + (1 - d)I_P\ ]}{exp[\ cZ + (1 - d)I_P\ ] \ + \ exp[\ aX_N\ ]}$$

where d is an index of similarity of the abort-carry outcomes.[6]

The probability of pregnancy will be a function of the pregnancy avoidance measures chosen by each young woman, which in turn will depend upon the costs of such measures and the anticipated costs of becoming pregnant. The Z vector thus includes measures of the availability of family planning services. The anticipated costs of becoming pregnant depend upon the choices in the succeeding two stages. The specification, therefore, also includes variables that influence the abortion and marriage choices. With the nested logit method, these are indirectly included through the inclusive value.

Similarly, the probability of carrying to term will depend on financial and psychic costs of obtaining an abortion and the anticipated costs of continuing the pregnancy and having the child. The Y vector includes indicators of state policy on abortions. Because the anticipated costs of abortion depend upon the choices in the final stage, the specification also includes variables that influence that stage. Again, these enter through the inclusive value.

For each outcome the X vector contains variables affecting financial and psychic well-being in that outcome. These include personal characteristics such as religion and frequency of religious attendance, family background characteristics, and expected wages specific to that fertility-marital outcome.[7] Measures of public assistance benefits are also included for the out-of-wedlock birth outcome.

C. Opportunity Costs and Other Variables

The set of independent variables at each stage of the nested logit model consists of the long run opportunity cost and short run cost variables which may plausibly influence young woman's choice at each stage and exogenous family background variables measured when each girl was age 14. The specifications for the second and third stage also contain the inclusive value derived from the prior stage. Every stage has a constant term as well.

The opportunity cost variables are central to the study. In the marriage stage they are the potential wages over a ten-year period if the young woman either marries or has the child out-of-wedlock and the welfare guarantee available if she bears the child out-of-wedlock. Both of these measure long run consequences of adolescent fertility and marriage outcomes. In the main sets of results, we measure the guarantee as the AFDC cash benefit provided to a family with no other income plus the amount of food stamps it would receive.[8]

The impact of adolescent fertility and marriage outcomes on own future wages is not observable and must be estimated. Details of the estimation procedure are in Appendix A. In brief, we developed potential long run wage functions using data on members of the older cohorts of the NLSY, for whom we can observe both adolescent fertility-marriage outcomes and earnings as young adults. Separate wage functions were estimated for three subsamples -- women who had a premarital birth before age 19, women who married and bore a child before 19, and women who were childless before 19 -- correcting for endogenous selection. The explanatory variables in these functions are family background characteristics measured at age 14 and local labor market characteristics. Because such characteristics are also reported for members of the younger cohorts used in our sample, we can use the wage functions to impute the potential wage for each fertility-marriage outcome for each sample member. Using exogenous variables measured at age 14 permits prediction of the expected wages that an adolescent could anticipate from alternative marital and fertility choices before she makes those choices.

The independent variables in the model of the abort-carry choice include potential wages if abortion is chosen and the short run cost of obtaining an abortion. As proxies for this cost, we include three measures of the availability of abortion and of state funding policies for abortion. These measures vary by state and, where data are available, over time. Appendix B provides details on the source and nature of these three variables and the family planning variables discussed below.[9]

The first is an index of policy on state funding of abortions for needy women. The index equals one if a state funded all or all medically necessary abortions and rises to four if it refused to pay for any reason. One anticipates that women living in states with higher values to be less likely to abort, so a negative coefficient is expected. The second is an index of the restrictiveness of state abortion law. One would predict more restrictive policies to be associated with lower chances of abortion. Here, too, a negative coefficient would accord with expectation. Third is an indicator of the availability of abortion, proxied by the percentage of counties within a state with large providers of abortions. A positive relationship with the likelihood of abortion is expected.

The model of premarital pregnancy includes potential wages if pregnancy is avoided and the short run cost of obtaining contraceptive services and supplies. The proxies for this cost are four measures of availability of and state policy on contraceptive services. These measures also vary by state and, where data are available, over time.

Three measures indicate access to family planning services within a young woman's state of residence. One is the percentage of teenage women who are at risk of an unintended pregnancy and obtain family planning services.[10] The magnitude of such a variable depends both on availability of such services and women's choices to use them and, hence, is not truly exogenous to the behavior this study is modeling. Its association with out-of-wedlock childbearing is worth exploring nonetheless. Two similar variables are the percentage of Medicaid eligible women at risk of unintended pregnancy served by organized family planning clinics and the percentage of counties with family planning clinics serving less than 50 percent of low income women at risk of unintended pregnancy. One would anticipate a negative coefficient on the first two variables and a positive one on the third. These variables are not available by race.

The fourth variable is a dummy for the presence of state laws, regulations, and policies which restrict the advertisement, sale, or licensing of contraceptives. A positive value indicates the presence of at least one restriction. A positive association with the likelihood of becoming pregnant would accord with expectation.[11]

We use the same set of background variables across all three levels. These variables partly capture differences in psychic benefits and costs and in family resources associated with outcomes at each stage. There are three family structure dummy variables: lived with mother and stepfather, with mother alone, or in a residual "other" family structure. The omitted category is "lived with both natural parents." There are dummy variables for whether an adult female, usually the mother, worked in the paid labor force, for whether a foreign language was spoken at home, and whether the girl was born or lived in the south. Variables for mother's education and number of siblings are included. There are dummy variables for religious identification: Protestant (excluding Baptist), Baptist, Catholic, and Jewish or other. The omitted category is "no religious identification." Religiosity is gauged by dummies for frequency of attendance at religious services: none, rarely, or occasionally. The omitted category is "frequently." The pregnancy model also includes age of menarche. Females who reach physical maturity earlier may be more likely to begin sexual activity earlier, thereby running a higher risk of becoming pregnant.

The samples used in each stage of the nested logit estimation drop observations with missing values for any one of the explanatory variables in that stage or a prior stage. In these samples the rates of premarital pregnancy and abortion and the proportion of women who did not marry before the birth are similar to those in Table 1. Pregnancies ending in miscarriages/stillbirths are dropped from the abortion and marriage samples since the women did not make explicit pregnancy resolution choices.

3. FINDINGS ON THE DETERMINANTS OF PREGNANCY AND PREGNANCY RESOLUTION

Tables 2 and 3 present estimated coefficients for a model that is representative of the range of models we examined. We then summarize findings from models which vary the specification of

Table 2

Three-Stage Model of Premarital Childbearing:
Nested Logit Estimates for Whites
(standard errors in parentheses)

	Marital Birth (versus Premarital Birth)	Abort (versus Carry to Term)	Become Pregnant (versus Not Become Pregnant)
Expected earnings/1000	.017 (.028)	.416*** (.065)	.211*** (.047)
Welfare guarantee/100	.595*** (.123)	NI	NI
Abortion funding policy	NI	-.268** (.106)	NI
Restrictiveness of abortion law	NI	-.080 (.136)	NI
Abortion availability	NI	.019*** (.005)	NI
Contraceptive laws	NI	NI	.460 *** (.109)
Foreign language at home	1.283** (.621)	-2.773*** (.694)	1.578 *** (.282)
Mother and stepfather	.280 (.328)	1.003** (.408)	-.213 (.185)
Mother only	-1.442*** (.333)	1.588*** (.419)	-.624*** (.178)
Other family structure	.730 (.671)	5.314*** (.938)	-2.194*** (.441)
Mother worked	1.046*** (.242)	.282 (.265)	.197** (.085)
Mother's education	-.011 (.062)	.113 (.080)	-.018 (.030)
Number of siblings	-.065 (.062)	.093* (.056)	-.054** (.025)

Table 2, continued

	Marital Birth (versus Premarital Birth)	Abort (versus Carry to Term)	Become Pregnant (versus Not Become Pregnant)
Protestant	-.168 (.446)	-.807** (.337)	-.120 (.163)
Baptist	-.550 (.399)	-1.270*** (.361)	.403** (.173)
Catholic	.126 (.426)	-2.006*** (.370)	.231 (.183)
Jewish and other	-.894** (.458)	-.825* (.432)	.157 (.176)
Never attends services	.655* (.348)	-.723** (.326)	.327** (.133)
Rarely attends services	-.665** (.282)	1.035*** (.257)	-.234** (.112)
Occasionally attends services	-.585** (.283)	.083 (.294)	.221** (.110)
Age of menarche	NI	NI	-.017 (.027)
Inclusive value	NI	-.356 (.362)	.577** (.248)
Constant	4.105*** (.998)	-38.889*** (5.271)	17.564** (2.475)
Log-likelihood	-303	-419	-2041
Chi-squared	122	339	1340
N of persons =	161	239	997

*** = significant at 1%. ** = significant at 5%. * = significant at 10%.
NI = Not included.

Model is estimated using the sample weights. The weights are rescaled to sum to the raw sample size to avoid inflating the t-statistics. Standard errors in columns 2 and 3 are corrected according to McFadden (1981) because the inclusive values, derived from regressions, are stochastic variables.

Table 3

Three-Stage Model of Premarital Childbearing:
Nested Logit Estimates for Blacks
(standard errors in parentheses)

	Marital Birth (versus Premarital Birth)	Abort (versus Carry to Term)	Become Pregnant (versus Not Become Pregnant)
Expected earnings/1000	-.064 (.103)	-.014 (.576)	-.443 (.536)
Welfare guarantee/100	.311 (.433)	NI	NI
Abortion funding policy	NI	.510* (.307)	NI
Restrictiveness of abortion law	NI	-.399 (.350)	NI
Abortion availability	NI	-.028** (.014)	NI
Contraceptive laws	NI	NI	.279 (.496)
Southern residence	.680 (1.112)	.554 (2.245)	-.036 (.903)
Southern birth	-1.848* (1.061)	-.923 (3.391)	-1.185 (1.271)
Mother and stepfather	-.378 (1.185)	-4.395 (5.769)	-1.791 (3.251)
Mother only	-1.857*** (.678)	-2.070 (3.202)	.135 (1.851)
Other family structure	-2.192* (1.227)	.634 (1.444)	.473 (.683)
Mother worked	-.072 (.553)	1.206 (1.917)	2.160** (1.014)
Mother's education	-.020 (.117)	.750** (.354)	.429** (.186)

Table 3, continued

	Marital Birth (versus Premarital Birth)	Abort (versus Carry to Term)	Become Pregnant (versus Not Become Pregnant)
Number of siblings	-.097	-.110	-.178
	(.106)	(.212)	(.113)
Baptist	-.381	-.008	-.296
	(.542)	(.576)	(.339)
Never or rarely attends services	-1.356	-.163	-.497
	(.985)	(.821)	(.414)
Occasionally attends services	1.059*	.542	-.199
	(.587)	(.703)	(.471)
Age of menarche	NI	NI	-.012
			(.109)
Inclusive value	NI	-3.406	-1.038
		(4.882)	(2.060)
Constant	1.327	-1.078	-14.341
	(3.092)	(23.500)	(13.930)
Log-likelihood	-58	-76	-134
Chi-squared	237	247	656
N of persons =	172	191	444

*** = significant at 1%.
** = significant at 5%.
* = significant at 10%.
NI = Not included.

Model is estimated using the sample weights. The weights are rescaled to sum to the raw sample size to avoid inflating the t-statistics. Standard errors in columns 2 and 3 are corrected according to McFadden (1981) because the inclusive values, derived from regressions, are stochastic variables.

the policy variables. We use the results to prepare illustrative cases that show how the probability of having a premarital child varies for different combinations of personal characteristics and the policy and opportunity cost variables.

Table 2 displays the basic model for whites for all three stages. Consider first the key opportunity cost variable, own potential earnings. Its coefficient in column one is positive, as anticipated, but insignificant. This result implies that the choice between marrying and not marrying of pregnant girls who carry to term does not depend on the long run consequences of that choice for their own earnings.

Own potential earnings do appear related to outcomes in the abortion and pregnancy stages, however. Both coefficients are significant at the one percent level. The positive coefficients imply that the higher the expected potential earnings associated with an outcome, the more likely it is to occur. Such findings are precisely what an opportunity cost model predicts.

The coefficient on the welfare guarantee is strongly significant. The positive coefficient in a conditional nested logit model implies that a pregnant teen is more likely to choose the alternative with the higher guarantee. Given that welfare eligibility criteria generally restrict AFDC benefits to single parents with children, this clearly means that the higher the benefit, the more likely the mother is to remain unmarried and have a premarital child.

We find statistically significant effects of abortion policy and availability. The negative coefficient on the "abortion funding policy" variable means that young women living in states with more restrictive policies on public funding of abortions are less likely to abort. The positive coefficient on the availability measure implies that greater availability is associated with a greater likelihood of aborting. The measure of restrictiveness of abortion law is not significant.[12] Because these three variables represent sources of variation in the short run cost of aborting versus carrying to term, the findings provide further evidence that costs affect white teenagers' fertility behavior.

It is likely that the restrictiveness of laws about sale, licensing, or advertisement of contraceptives most strongly affects teenagers. Such laws would tend to raise the full costs of obtaining contraceptives, thereby reducing their use. The results in column three show that more restrictive laws are associated with higher chances of pregnancy.[13]

The estimates for the black sample are in Table 3. Neither the potential earnings measure nor any policy variable is significant in a manner consistent with an opportunity cost model. This is a sharp contrast to the results for whites. Most are not significant. Unexpectedly, more restrictive policy on funding of abortions and greater availability of abortion have significant coefficients that are opposite in sign to what was expected. We also find far fewer significant coefficients on the personal and family background variables than in the white model.

For the abortion and pregnancy stages, the smaller samples may be responsible for the weak results. The marriage stage sample is actually slightly larger for blacks. But as Table 1 showed, few blacks marry before giving birth, so difficulty in finding variables associated with this choice is not surprising. Also, the severe underreporting of abortions by blacks means that a small proportion of the sample for the abortion model reports an abortion. The differences between the results for whites and blacks may reflect either data problems in the black sample or a genuine difference in responsiveness to the variables in the NLSY or to the policy variables we added.

A. Findings for Alternative Specifications

To test the robustness of the findings on welfare benefits, we re-estimated the model with three different measures in place of the AFDC plus food stamp guarantee. They are the cash

AFDC guarantee alone, the AFDC and food stamp guarantee plus the expected value of the Medicaid benefits used by an AFDC family, measured at market prices, and the same three program guarantee, but with Medicaid benefits measured at their cash equivalent value.[14] Another variation used the logarithm of the cash-food stamp measure. We also estimated the model after deflating the cash-food stamp measure by Fournier and Rasmussen's (1986) state cost-of-living index.

For whites the welfare measure's coefficient was positive and statistically significant at the 5 percent level or higher in all variations. This strengthens our confidence in the finding that larger welfare benefits are related to a greater likelihood of out-of-wedlock childbearing among whites. For blacks none of the measures was significant.

Tables 2 and 3 report the impact of state laws and regulations on contraceptives. We substituted the three other family planning variables in the pregnancy stage to see whether the findings were sensitive to the choice of policy indicator. Table 4 shows the results for whites. When we enter each variable separately, the coefficients on all three have the anticipated sign and two are statistically significant. When all four appear in the same equation, all coefficients retain their sign, and two of the four are significant. Thus, the evidence linking family planning services and white behavior leading to premarital childbearing outcomes appears robust. For blacks no family planning variable was significant.

B. Illustrative Cases

Table 5 uses the findings for whites to illustrate in a simple way the effects of changes in policy and individual variables. It shows how the probability of pregnancy, the conditional probability of carrying to term, given a pregnancy, and the conditional probability of bearing the child out-of-wedlock change as the teenager's policy environment and personal characteristics

Table 4

Coefficients on Family Planning Variables under Alternative Specifications, for Whites[a]

	Entered Singly	Entered Jointly
Contraceptive laws	.460***	.329***
	(.109)	(.096)
Family planning services for teenage women	-.001	-.004
	(.004)	(.005)
Family planning services for Medicaid eligible women	-.013***	-.007**
	(.004)	(.003)
Low availability of family planning services	.011***	.004
	(.003)	(.003)

[a]Complete model includes all the opportunity cost, policy and personal variables shown in Table 3.

*** = significant at 1%.
** = significant at 5%.

Table 5

Illustrative Effects of Explanatory Variables on the Probability of Outcomes Linked to Premarital Childbearing among White Adolescents

	Probability of Pregnancy	Probability of Carrying to Term	Probability of Being Unmarried at Birth	Unconditional Probability of Premarital Birth
Effects of Policy Variables				
1. Base case[a]	.163	.485	.256	.020
Base case except:				
2. No child wage up 5%	.141	.139	.256	.005
3. No child wage down 5%	.256	.846	.256	.055
4. Welfare = $465	.165	.497	.149	.012
5. Welfare = $693	.160	.465	.404	.030
6. Abortion funding index =4 (most restrictive)	.145	.616	.256	.023
7. Low abortion availability	.149	.579	.256	.022
8. Conservative abortion policy climate[b]	.134	.718	.256	.025
9. Liberal abortion policy climate[c]	.200	.312	.256	.016
10. Some restrictions on contraceptives	.235	.485	.256	.029
11. Conservative welfare and abortion policies[d]	.200	.728	.149	.022
12. Liberal welfare and abortion policies[e]	.198	.296	.404	.024

Table 5, continued

	Probability of Pregnancy	Probability of Carrying to Term	Probability of Being Unmarried at Birth	Unconditional Probability of Premarital Birth
<u>Effects of Family Background Variables</u>				
Base case[a]	.163	.485	.256	.020
Base case except:				
Mother only	.209	.591	.581	.072
Mother/stepfather	.233	.626	.219	.032
Mother's educ. = 8	.242	.869	.235	.049
Mother's educ. = 16	.161	.110	.279	.005
Mother did not work	.146	.663	.514	.050
Baptist	.231	.622	.335	.048
Jewish and other	.212	.532	.416	.047
Catholic	.171	.742	.204	.026
No religion	.227	.405	.131	.012
Siblings = 1	.161	.451	.240	.017
Siblings = 5	.165	.518	.273	.023

[a]Lived with both natural parents, mother's schooling = 12, mother worked, three siblings, non-South residence and birth, no foreign language spoken at home, frequently attends services, Protestant, age of menarche = 13, welfare = $579, abortion funding index = 2, index of abortion law = 1, abortion availability = 55 (percent of counties with large providers of abortion), no restrictions on contraceptives.
[b]Funding index = 4, law index = 3, availability = 35.
[c]Funding index = 1, law index = 0, availability = 75.
[d]Same as b, plus welfare guarantee = $465 and some restrictions on contraceptives.
[e]Same as c, plus welfare guarantee = $693 and no restrictions on contraceptives.

change. Multiplying these three probabilities yields the unconditional probability of having an out-of-wedlock child.

The base case assumes that the young woman lived with her mother and father when she was 14, her mother completed high school, she has three siblings, an adult female in the household (usually her mother) worked when she was 14, no foreign language was spoken at home, she attends religious services frequently and age of menarche was 13. Religious identification is non-Baptist Protestant. We set the welfare guarantee at the mean of $579, the abortion funding policy index at two (slightly less restrictive than the mean), the abortion law index at one (also sightly less restrictive than the mean), the abortion availability index at the mean of 55 percent, and the dummy for restrictions on contraceptive sales and advertising at its modal value of zero. Most other rows show the effects on the conditional probabilities and the unconditional probability of a premarital birth as one policy or personal variable changes and all others remain at their base case values. In a few cases, we vary a combination of the policy variables.

For the base case, the probabilities of getting pregnant, carrying to term, and not marrying before the child is born are .163, .485 and .256. The unconditional probability of becoming an unwed teen mother is .020.[15]

The next two rows of the table illustrate the strength of the association between the wage variable and premarital childbearing. A 5 percent increase in the expected wage if the girl avoids childbearing reduces the unconditional probability to .005, primarily through a reduction in the likelihood of carrying to term to only .139. That is, raising the return to avoiding motherhood substantially raises the likelihood of choosing abortion, holding constant the wages associated with the other outcomes. The likelihood of becoming pregnant falls slightly as well. Conversely, lowering the wage associated with avoiding childbearing by 5 percent raises the chances of premarital childbearing to .055.[16]

Rows 4 to 12 of the table illustrate the strength of the association between the policy variables and premarital childbearing. Lowering welfare benefits by one standard deviation to $465 lowers the likelihood of a premarital birth from .020 to .012. Raising them by the same amount increases the likelihood to .030.

Row 6 shows that those in states with the most restrictions on public funding of abortions are about 15 percent more likely to become unwed mothers (.023 versus .020). Greater restrictiveness significantly raises the probability of carrying to term, as the coefficient from table 2 implied. Observe that greater restrictiveness also lowers to probability of becoming pregnant. This is an indirect effect that operates through the inclusive value. As restrictiveness rises, the value of one of the outcomes associated with getting pregnant falls. Hence, girls are less likely to become pregnant as well. Row seven shows similar effects of lowering the availability of abortion (percentage of counties with large providers of abortions) by 20 percentage points.

States with more restrictive values for one abortion policy variable tend to have more restrictive values on the others. To assess the effect of the overall climate of abortion policy in a state, row eight shows predicted outcomes for a state with relatively restrictive values on all three abortion variables. The joint effect raises the unconditional probability to .025, roughly double the change predicted in row six or seven. Row nine contains similar findings for a state with more liberal values on all three and shows a fall in the probability to .016.

Restrictions on the sale and advertising of contraceptives raise the likelihood of becoming pregnant to .235, a jump of 44 percent. This implies a likelihood of premarital childbearing of .029.

The final two policy simulations compare predicted outcomes for young women living in archetypical conservative and liberal states. The conservative state has restrictive values on all abortion variables, restrictions on the sale and advertising of contraceptives, and low AFDC benefits. The liberal state has the reverse. The combinations of incentives favorable and unfavorable to becoming an unwed mother roughly cancel out. On net, our estimates show that the incentive not to marry created by higher AFDC slightly dominates the incentives to avoid pregnancy and to abort, so the likelihood of premarital childbearing is marginally higher in the liberal state (.024 versus .022.)

While the effects of personal and family background characteristics on premarital childbearing are not the focus of this study, they merit brief consideration.[17] Note the importance of family structure. Living with only one's mother more than triples the probability of premarital childbearing to .072. Living with mother and stepfather raises the overall probability to .032, a 60 percent jump. The chances of having a premarital child rise to .049 for daughters of mothers who only finished eighth grade and fall to .005 for daughters of college graduates. Daughters whose mothers did not work for pay have a likelihood of premarital childbearing of .050 -- 150 percent higher than the base case. This increase largely stems from the sharply increased probability that daughters from such a background will not marry.

Baptists and those in the "Jewish and other" category are more than twice as likely to become unwed mothers as Protestants. For Baptists the principal source of the increase is the greater likelihood of getting pregnant. For Jews and others, it is the greater likelihood of not marrying. Catholics are slightly more likely to become unwed mothers, mainly because they are less likely to abort. Girls claiming no religious identification are the least likely to have premarital births. Last, the substantive importance of number of siblings is small even though this variable showed a statistically significant association with pregnancy and abortion outcomes. The predicted outcomes hardly differ between having one sibling or five.

C. <u>Comparison of Opportunity Cost and Policy Variable Results to Related Studies</u>

Of studies which use individual level data to analyze determinants of premarital pregnancy and pregnancy resolution, only Moore and Caldwell (1977) and Serrato (1989) analyze policy variables. No such study develops estimates of the labor market consequences of having a premarital birth and includes them as explanatory variables. Among reduced form studies which analyze the probability of a premarital birth, only Duncan and Hoffman (1990) carefully assess the effect of opportunity costs.

Moore and Caldwell's analysis of pregnancy shows that better family planning services reduce pregnancies among blacks, but not whites, while we find the reverse. They combine blacks and whites to analyze pregnancy resolution choices. They report no effect of welfare benefits, which is inconsistent with our findings for whites. They find that greater availability of abortion is positively associated with the likelihood of abortion, as we did for whites and the combined sample.

There are several possible sources of the conflicting findings. Moore and Caldwell's data, gathered in 1971, measure behavior during 1971, 1970, and the late 1960s. We examine the 1979 to 1986 period -- a different social era in many respects and one that postdates Roe v. Wade. Responsiveness to policy variables may have changed. Also, the studies differ in statistical methods and in the measurement of policy variables.

Serrato (1989) estimates a nested logit model of premarital pregnancy resolution on a sample of the NLSY not restricted to young teenagers, but does not examine determinants of pregnancy. He finds, as we do, that measures of abortion availability significantly affect the probability of abortion. Living in a state which offers Medicaid-funded abortions increases the probability of an abortion for a pooled white/Hispanic sample, but has no significant impact for blacks. The number of high-volume abortion providers per 100,000 women in the state has a significant positive effect on abortions. (No race-specific effects were estimated.) Serrato also finds significant effects of welfare variables on the probability of marriage: the presence of an AFDC-UP program for two-parent families increases marriages for blacks only, and the expected AFDC payment has a large and significantly negative effect on marriage for all groups. These results suggest that the age composition of the sample may be important for estimates of black marriage behavior.

Duncan and Hoffman (1990) examine the probability of AFDC-related premarital births among a sample of black teenagers from the Panel Study of Income Dynamics. The results of their reduced form model agree with ours in finding statistically insignificant effects of AFDC benefit levels. They differ in that their measure of predicted future income has a significant negative impact on the joint probability of AFDC receipt and a premarital birth.

4. CONCLUSION

Do the opportunity costs of becoming a teenage unwed mother matter? Our empirical results do not provide the kind of evidence that permits a clear "yes" or "no," but instead offer a mixed picture.

White females' behavior appears to systematically respond to differences in long run opportunity costs associated with different teenage fertility and marital outcomes. All three coefficients on the long run wage measure are positive and those in the abortion and pregnancy models are statistically significant. White behavior also appears sensitive to other costs. It is associated with welfare, abortion, and family planning policy variables in directions consistent with a rational choice model of behavior. These findings suggest that an opportunity cost perspective does contribute to our understanding of premarital childbearing. The simulations based on the estimates should be viewed as illustrative. It would be premature to draw strong conclusions or offer policy recommendations that hinge on the precise results of this exercise.

It is possible that state differences in policies, especially those on abortion and family planning, are serving to some extent as indicators of more fundamental differences in state attitudes and social mores on abortion, family planning, and adolescent sexual behavior. Perhaps these hard-to-measure elements of the social reality exert the principal influences on adolescent behavior. If this indicator effect is dominant, changing the policy parameters by, say, federal legislation that overrides state preferences, would in itself have little effect on premarital pregnancies. If, however, the policy variables really operate through changes in individual costs, then changing the policy parameters would tend to change individual behavior.

Black behavior shows no association with the opportunity cost or policy variables. This may be a function of sample size or the underreporting of abortions and pregnancies. It may be that the behavior of minorities is not as strongly related to the variables available in the NLSY or to the policy variables we added. Or it may be another demonstration that there are important unmeasured racial differences in the factors that influence fertility, marital, and other social behavior.

Early childbearing and marriage behavior may influence a young woman's chances of being married in the future and the earnings potential of the men she can expect to marry. This study did not include measures of long run marriage prospects and husband's earnings. Developing measures of these opportunity costs and adding them to the model are logical next steps.

To argue that opportunity costs influence adolescent premarital childbearing outcomes is not to deny that social, cultural, and attitudinal factors also influence this behavior. Rather, our findings, like those of Duncan and Hoffman (1990), begin to make the case that responses to the costs of individual actions also play a role in the complex set of choices that result in adolescent premarital childbearing. For those interested in developing public policies to reduce premarital childbearing, evidence that opportunity costs matter is good news. Our knowledge of how to change earnings and other conventional opportunity costs and incentives is far from complete. But it is greater than our knowledge of how to change attitudes and culture.

Notes

[1]These include Bernstam and Swan (1986); Duncan and Hoffman; (1990); Ellwood and Bane (1985); Eisen et al. (1983); Field (1981); Leibowitz et al. (1986); Moore (1980); Plotnick (1990); and Serrato (1989).

[2]The major studies containing multivariate analyses of microdata in which the outcome of having a premarital child is a dependent variable include Abrahamse, Morrison, and Waite (1988); Cooksey (1990); Devaney and Hubley (1981); Hanson, Myers, and Ginsburg (1987); McLanahan and Bumpass (1988); Moore and Caldwell (1977); Yamaguchi and Kandel (1987); and Zelnik, Kantner, and Ford (1981).

[3]We do not consider the choice between raising the child as an unmarried mother and placing it for adoption since the placement option is now rarely exercised.

[4]We also estimated a separate model for Hispanics, but the sample size was so small that we have little confidence in the findings.

[5]We do not use older females in the NLSY since many of them will have had a premarital pregnancy before 1979. Since state of residence is not available in the NLSY before 1979, we would not be able to identify where they lived when they became pregnant and, thus, would not be able to assign state policy variables to the observation; yet dropping such pregnancies from analysis would create a sample biased toward women with stronger tendencies to avoid pregnancy. Those age 14 to 16 in 1979, in contrast, will have had relatively few pregnancies before 1979.

[6]In a simultaneous estimation of this model, the parameter vector a could be constrained to be equal across all outcomes. In practice, we allow these parameters to vary freely at each stage.

[7]The standard nested logit model is derived from the conditional logit model. In the standard conditional logit model all explanatory variables are characteristics of the alternative <u>choices</u> facing an individual. As is well known, the characteristics of an <u>individual</u> also affect her fertility and

marriage behavior. Because personal characteristics do not vary across the possible choices, a standard nested (conditional) logit model is not adequate. We resort instead to a "mixed"nested logit model, which allows inclusion of characteristics of both the individuals and the available outcomes.

[8]We used the guarantee for a family of four since it was readily available. The adult plus one child benefit would be more appropriate, but because its correlation with the four person value is about .98, little problem arises.

[9]A measure of the average medical bill for an abortion would also be appropriate, but no such measure was available.

[10]The Alan Guttmacher Institute, which provided the data for this measure, defines a woman "at risk" if she has been sexually active, believes she is fecund, is not currently pregnant, and does not want to get pregnant.

[11]One issue arises in assigning values of family planning variables. If a female becomes pregnant, we could assign values of her state's family planning variables for the year she became pregnant. But what values should we assign to those who do not become premaritally pregnant? To be consistent, for all young women in the sample used to analyze determinants of premarital pregnancy, we assign values for state of residence at age 16. This is not a problem at the abortion stage since, once pregnant, each young woman has three or four months to abort or not. Thus, for the abortion variables we can assign values which applied for the year she became pregnant. Similar logic means we can assign welfare guarantees for that year as well.

[12]It was significant, though, when entered as the only abortion variable. The other two variables also were significant when entered singly.

[13]Such laws and regulations may neither be well publicized nor enforced. As indicators of state differences in conservatism in matters of sexual behavior, they may be proxies for the costs imposed by social disapproval.

[14]Medicaid benefits are deflated by Grannemann and Pauly's (1983) price index for medical care. Based on Smeeding's (1982) estimate, we multiply Medicaid benefits by .37 to obtain their cash equivalent value. We thank Robert Moffitt for sharing his data on AFDC, food stamp and Medicaid benefits.

[15]This is smaller than the sample statistic in Table 1. We use modal values for the base case demographic variables, which generally represent characteristics less likely to be associated with a premarital birth.

[16]Neither change affects the probability of being unmarried at birth conditional on carrying to term because this outcome depends only on the wages associated with childbearing within and outside of marriage.

[17]We use the long run potential wage regressions to impute the level of expected wages associated with each set of personal and family background characteristics. Thus, changes in probabilities associated with changes in personal and family background characteristics reflect both the direct effect of such characteristics on fertility and marital behavior and an indirect effect operating through the wage term. The wage regressions also include local labor market variables. We use the same mean values for these variables in all imputations.

APPENDIX A

Measuring the Effects of Early Childbearing and Marriage on Own Potential Wages

The nested logit model presented in Tables 2 and 3 includes alternative wages each young woman could expect to receive in her twenties, given her adolescent fertility-marital outcome. These predicted wages are derived using the labor market experiences and fertility-marital histories of the oldest cohorts of the NLSY (those aged 19 to 21 in 1979). Separate wage functions were estimated on three subsamples: women who had a premarital birth before age 19, women who married and bore a child before 19, and women who were childless before 19. The predicted wages of these older cohorts were then used as proxies for the expectations of similar individuals in the younger sample. "Similarity" is based on a set of background characteristics measured as of age 14 and collected retrospectively for all cohorts of the NLSY. Details of the estimation procedure can be found in Lundberg and Plotnick (1990); a brief summary and the final results are presented below.

The dependent variable in these regressions is a summary measure of real wages from ages 19 to 28, based on the age 14-variables and specific to a fertility-marital outcome. Since wages are observed only for labor market participants and the observation period for each individual in the sample does not coincide exactly with the chosen age range, an imputed wage series is constructed to fill in the gaps in the data record. A conventional wage equation is estimated on the pooled cross-section time-series of the panel, with a correction for selection bias due to nonparticipation. The wage equation includes current human capital variables and cumulative measures of the fertility and marital history, as well as family background variables and age interaction terms. The participation equation includes all variables in the wage equation, plus measures of spouse's income and unemployment, an indicator of young children in the household, and current local unemployment rates. The imputed wages are converted to potential full-time earnings by multiplying them by 2000, discounted at 3 percent, and summed over the age range 19 to 28.

To measure the expected loss in potential earnings due to early childbearing and marriage, we regress the summary wage measure on the age 14-variables separately for each of the three fertility-marital subsamples. The results are presented in Table A-1 for whites and in Table A-2 for blacks. Since women do not sort themselves randomly into these three subsamples, we incorporate a correction for selection bias based on a multinomial logit model of the fertility-marital outcome. The selection term from this nonnormal model is calculated by the method described in Lee (1983). For identification purposes, variables affecting fertility and marriage decisions but not directly affecting wages are required: we include age of menarche and dummy variables for religion and frequency of attendance at religious services in the logit selection model for this purpose.

Table A-3 summarizes the results, presenting predicted wages by fertility-marital status at age 19 for young women with fixed family background characteristics. The percentage difference between the predicted wage given the birth of a premarital child and the predicted wage in the "no child" case provides a measure of the impact on adult wages of early out-of-wedlock childbirth. A similar calculation yields the predicted impact on wages of becoming a married teenage mother, relative to remaining childless during the teenage years. The base case presents wage predictions for a girl with representative family background characteristics. Local labor market variables are set at their sample means. We also present one "advantaged" case and two "disadvantaged" cases, each a "worst case" for either blacks or whites.

The results for whites and blacks are dramatically different. For whites, premarital childbearing is associated with moderate wage losses of 11 to 14 percent; postmarital childbearing causes somewhat higher losses. For blacks, no reduction in potential wages is associated with teenage childbearing. Instead, there are wage premia which are higher for those from disadvantaged family backgrounds. The black postmarital birth sample is small, so imputations based on it may be unreliable. The coefficients in Tables A-1 and A-2 are used to impute state-specific wages for the younger cohorts whose adolescent behavior is examined in this study. Since the wages are based on exogenous variables measured at age 14, they can be regarded as proxies for the wages an adolescent could anticipate from alternative fertility and marital choices before she makes those choices.

Table A-1

Determinants of Future Potential Earnings by Fertility-Marital Status at Age 19: Whites

	No Child	Premarital Child	Postmarital Child
Constant	3.621*** (.038)	3.521*** (.140)	3.456*** (.080)
Trade employment	-.013*** (.004)	-.054*** (.018)	.010 (.010)
Government employment	.014*** (.004)	.025 (.017)	.003 (.013)
Per capita income	.019*** (.005)	.022 (.016)	.023 (.014)
Mother only	-.049*** (.015)	-.089* (.052)	-.055 (.045)
Mother and stepfather	-.048*** (.020)	-.004 (.049)	-.072* (.040)
Other family structure	-.159*** (.022)	-.094* (.056)	-.114** (.047)
Number of siblings	-.004 (.002)	-.017* (.009)	.000 (.006)
Foreign language spoken at home	.090*** (.015)	.133*** (.049)	.035 (.037)
Mother's schooling	.011*** (.003)	.011 (.011)	-.004 (.008)
Adult female in household worked	.006 (.009)	-.010 (.036)	.061** (.028)
Southern birth	.019 (.018)	-.076 (.060)	-.027 (.048)
Southern residence	-.009 (.018)	.020 (.063)	.038 (.043)
Sample selection	-.138*** (.043)	-.089 (.071)	.019 (.057)
R^2	.19	.26	.13
Mean of dependent variable	3.72	3.38	3.47
N	1436	117	183

*Significant at 10%.
**Significant at 5%.
***Significant at 1%.

Table A-2

Determinants of Future Potential Earnings by Fertility-Marital Status at Age 19: Blacks

	No Child	Premarital Child	Postmarital Child
Constant	3.281*** (.033)	3.428*** (.050)	3.336*** (.144)
Trade employment	-.010* (.005)	-.008 (.006)	-.051*** (.014)
Government employment	-.007* (.004)	-.011 (.007)	.003 (.019)
Per capita income	.006 (.005)	.010 (.008)	.014 (.015)
Mother only	-.130*** (.012)	-.121*** (.018)	-.126*** (.035)
Mother and stepfather	-.241*** (.022)	-.203*** (.029)	-.083* (.046)
Other family structure	.017 (.017)	.029 (.025)	-.002 (.035)
Number of siblings	.007*** (.002)	.005* (.003)	.008 (.006)
Foreign language spoken at home	.015 (.025)	.010 (.050)	-.039 (.057)
Mother's schooling	.011*** (.002)	.008** (.004)	.016** (.007)
Adult female in household worked	-.066*** (.010)	-.044*** (.013)	-.049 (.033)
Southern birth	-.071*** (.016)	.000 (.020)	-.053 (.052)
Southern residence	.003 (.016)	-.052*** (.019)	-.032 (.062)
Sample selection	.108*** (.032)	-.053 (.047)	-.084 (.053)
R^2	.53	.54	.85
Mean of dependent variable	3.37	3.37	3.22
N	408	194	29

*Significant at 10%.
**Significant at 5%.
***Significant at 1%.

Table A-3

Relative Future Potential Earnings by Fertility-Marital Status by Age 19

(Dependent variable: Present discounted value of the predicted potential full-time earnings, summed from age 19 to 28 in 1985 dollars)

	Earnings, No Child	Percentage Loss in Earnings, Relative to No Child	
		Premarital Child	Postmarital Child
<u>Whites</u>			
Base Case*	84,800	14	24
Disadvantaged: White	68,700	11	15
Advantaged	88,800	11	33
<u>Blacks</u>			
Base Case	58,000	-14	-14
Disadvantaged: Black	41,400	-21	-29
Advantaged	63,900	-10	-15

*Base Case: Lived with mother and father at 14, mother's education = 12 years, 3 siblings, foreign language not spoken at home at 14, non-south birth, non-south residence, adult female in household worked at 14.

Disadvantaged (black): Same as base case except lived with mother and stepfather at 14, mother's education = 8 years, 5 siblings, south birth, south residence.

Disadvantaged (white): Same as base case except other family structure at 14, mother's education = 8 years, 5 siblings.

Advantaged: Same as base case except mother's education = 16 years, 1 sibling, adult female in household did not work.

APPENDIX B

Abortion and Family Planning Variables: Sources and Details on Their Construction

1. State funding policy on abortions:

   Indexed as follows: 1 = funded all or all medically necessary abortions; 2 = under court order to pay for abortions for medical, and/or emotional reasons; 3 = state will pay only if full-term pregnancy may endanger life of woman, or if pregnancy is due to rape or incest; 4 = not paid by state for any reason.

   Source: R. Gold. "Public funded abortions in FY 1980 and 1981," Family Planning Perspectives 14 (1982): 205; and R. Gold and J. Macias. "Public funded abortions in FY 1985," Family Planning Perspectives 18 (1986): 263. Data for all three years were coded in identical fashion: 1980 data for observations from 1979 and 1980; 1981 data for observations from 1981-1983; 1985 data for observations from 1984-1986.

2. Restrictiveness of state abortion law:

   Index starts at zero and increases by one if abortion was legalized later than 1969, if whether, as of 1980, minors had to obtain parental consent, and if second trimester abortions had to be performed in a hospital. Maximum value is 3. High values mean more restrictive policies. Same value used for all years.

   Source: Unpublished data provided by Susheela Singh, Alan Guttmacher Institute.

3. Abortion availability:

   Percentage of population living in counties with large providers of abortions, as of 1980. Same value used for all years.

   Source: Unpublished data provided by Susheela Singh, Alan Guttmacher Institute.

4. Use of family planning services:

   Percentage of women under age 20 at risk of unintended pregnancy who obtained family planning services from clinics or physicians. Data are available for 1979 and 1981. Used average of these two variables for 1980; used 1981 values for later years.

   Source of 1979 data: A. Torres, J. Forrest, and Eisman. "US Organized Family Planning Services," Family Planning Perspectives 13: (1981) 138. Source of 1981 data: A. Torres and J. Forrest. "US Family Planning Clinic Services," Family Planning Perspectives 15 (1983): 276

APPENDIX B (continued)

5. Family planning services for Medicaid eligibles:

   Percentage of Medicaid eligible women at risk of unintended pregnancy served by organized family planning clinics. Same value used for all years.

   Source: M. Orr and L. Brenner. "Medicaid funding of family planning clinic services," Family Planning Perspectives 13 (1981): 282.

6. Family planning service availability for low income women:

   Percentage of counties who serve less than 50 percent of low income women at risk of unintended pregnancy in 1983. Same value used for all years.

   Source: A. Torres and J. Forrest. "Family planning clinic services in US counties in 1983," Family Planning Perspectives 19 (1987): 57.

7. Laws on contraceptives:

   Dummy for presence of state laws, regulations, and policies which restrict the advertisement, sale, or licensing of contraceptives, 1 = at least 1 restriction. Same value used for all years.

   Source: D. Bush. "Fertility related state laws enacted in 1982," Family Planning Perspectives 15 (1983): 115.

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