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Do Teens Make Rational Choices?
The Case of Teen Nonmarital Childbearing

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Abstract

With emphasis on the role of economic incentives, we explore the determinants of a woman's choice of whether or not to give birth as an unmarried teenager. Our data are taken from the Panel Study of Income Dynamics. Guided by a simple utility-maximization model, we represent the income possibilities available to teenaged women if they do and do not give birth out of wedlock. We estimate these choice-conditioned income possibilities through a two-stage probit procedure, relying on the observed incomes of a secondary sample of somewhat older women. The response of the young women in our primary sample to these income expectations is measured after controlling for the effects of a variety of other factors, including the characteristics of the girl's family, the social and economic environment in which she lives (including such policy-related factors as expenditures by states on family planning programs and education), and her own prior choices. We use the estimated structural parameters from our model to simulate the effects of a variety of policy interventions on the probability of becoming an unmarried teen mother. Our estimations provide evidence that income expectations have a persistent influence on the childbearing decision. They also provide evidence that the provision of public family planning expenditures and increases in parental education could reduce the prevalence of teen nonmarital births.

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I. INTRODUCTION

The prevalence of teen nonmarital childbearing has been described as the nation's "most serious social problem."[1] Annually, there are now more than one-half million births to U.S. teenagers, and this number grows by about 50,000 every four years. Teen births now account for about 13 percent of all births and nearly one-quarter of all African-American births. Even more dramatic is the very sharp increase in nonmarital births, shown in Figure 1. Today, nearly three-quarters of births to teenagers are out-of-wedlock; among non-Hispanic blacks nearly all (95 percent) teen births occur out-of-wedlock.

This pattern is viewed as a social and economic problem because of the presumed adverse effects on the human capital and the future productivity of both teen unmarried mothers and their children. While the longer-term impacts of early nonmarital childbearing on young mothers is unsettled among researchers,[2] the adverse effects of being born to a teen mother are not.[3]

The generosity and accessibility of welfare benefits have been cited as an important determinant of teen nonmarital childbearing, as have sexual education in the schools, the increased availability of child care assistance, high poverty incidence in the families and neighborhoods in which teen mothers

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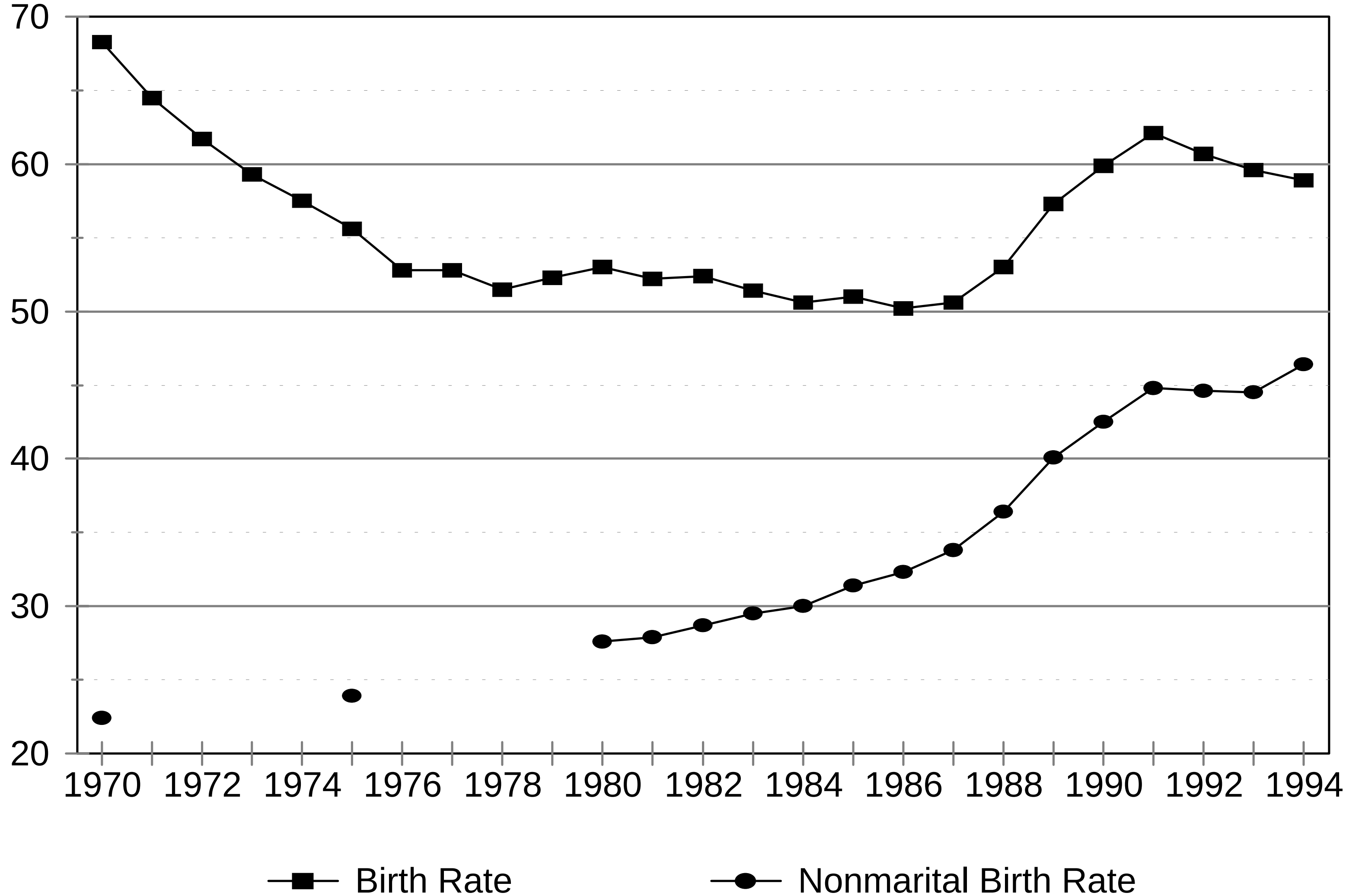
[1]President Bill Clinton, in his 1995 State of the Union Message.

[2]Although teen unwed mothers have less income, more marital instability, and lower educational attainment than those who these outcomes may be attributable to unmeasured adverse family background or personal characteristics. Maynard (1997) contains a set of studies exploring the consequences and costs of adolescent childbearing, including efforts to account for this selection problem. Geronimous and Korenman (1992, 1993), Hoffman, Foster, and Furstenberg (1993), Brooks-Gunn, Duncan, Klebanov, and Sealand (1993), and Bronars and Grogger (1994) have all attempted to account for this potential selection effect in estimates of the consequences for mothers of teenage and unwed childbearing, with varying results. Hotz, McElroy, and Sanders (1997) use a natural experiment—a comparison of teen mothers with women who became pregnant as teens but who experienced a miscarriage—to account for adverse unmeasured effects, and suggest that virtually all of the costs associated with early childbearing are a manifestation of this selection effect. Their conclusion, however, depends on the extent to which miscarriages are purely random events, and there are important reasons for believing that this is not the case.

[3]There is substantial evidence that the children born to teenage mothers (especially those who are not married) are more likely to grow up in a poor and mother-only family, live in a poor or underclass neighborhood, and experience high risks to both their health status and school achievements. See Haveman, Wolfe, and Peterson (1997) and Wolfe and Perozek (1997). Rosenzweig and Wolpin (1995) deal with this issue as well.

Figure 1


70
60
50
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30
20
1970
1972
1974
1976
1978
1980
1982
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1988
1990
1992
1994
Birth Rate
Nonmarital Birth Rate

grow up, and the poor labor market prospects available to those groups with the highest teen nonmarital birth rates. Unfortunately, knowledge of the relative strength of these potentially causal linkages is weak.

Here, we study the determinants of this individual choice, emphasizing the response of youths to economic incentives associated with alternative responses.[4] These incentives are represented by the income possibilities available to young women if they give birth as an unmarried teen or if they forgo childbearing while unmarried and an adolescent. We also measure the effects of a variety of other factors, including the characteristics of the girl's family and its choices, the social and economic environment in which she lives (including policy-related factors, such as expenditures by states on family planning programs and education), and her own prior choices.

Following a brief review of research on the determinants of the teen nonmarital birth decision, we present a simple utility-maximization model of an adolescent woman's choice regarding whether or not to have a birth out of wedlock. Expected utilities (incomes) conditional on either experiencing or not experiencing a teen nonmarital birth play a crucial role in this model, as do family characteristics and the neighborhood and policy environment. After describing our data, we explain our procedures for estimating the conditional income expectations that we attribute to older teenagers, and present our estimates of these expected incomes. We then estimate a model designed to capture the effects of conditional expectations on the childbearing decision, and use the estimated structural parameters to simulate the effects of a variety of policy interventions on the probability of experiencing nonmarital birth as a teen.

[4]We recognize that the childbearing outcome for an unmarried teenager reflects an extensive set of choices made by the woman, including, for example, whether or not to be sexually active, whether or not to use contraceptives, and if pregnant whether to have an abortion. For analytical convenience, we focus on this final decision of whether or not to have a child.

II. A TOUR OF RESEARCH STUDIES

The earliest microdata study of the determinants of adolescent nonmarital births is that of Hogan and Kitagawa (1985), who find that among a sample of 1,000 African-American teenage women in Chicago, nonmarital birth probabilities are positively related to a variety of adverse parental and background circumstances experienced during childhood. More recent studies of this outcome use longitudinal data, and thus include richer information on a child's background observed at several points during their formative years. These include Antel (1988), who uses 1979–1986 data from the National Longitudinal Survey of Youth (NLSY) in a bivariate probit specification designed to control for possible unobserved family-specific heterogeneity; Plotnick (1991) and Lundberg and Plotnick (1990), who also use the NLSY, but add information on state welfare policy, state family planning policy, and the socioeconomic environment proxied by characteristics of the girl's school.

More recently, Brooks-Gunn and Chase-Lansdale (1995) use a 20-year data set of about 300 low-income African-American families in Baltimore, and emphasize the critical roles of child care, the extended family, and the parenting ability of young mothers on the probability of this outcome.[5]

Parental economic resources and schooling are included as determinants of the teen nonmarital birth decision in nearly all of these studies. The estimated coefficient on the level of parental education is always negative and statistically significant; the estimated coefficient on parental income is negative and usually, but not always, significant. There is some evidence that the source of family income matters; parental welfare receipt generally has a positive effect on the probability that teens will choose to give birth out of wedlock. A number of other determinants of the nonmarital birth outcome are often statistically significant, including indicators of family structure, family stress factors (such as family

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[5]See also Akerlof, Yellen, and Katz (1996), who use a game theory framework to explore the effect of improved contraceptive and abortion technology to explain the increase in the prevalence of out-of-wedlock births.

disruptions and geographic moves during childhood), and parental attitude, expectations, monitoring and control of children, and contraceptive practice.

These studies vary widely in data, model specification, and estimation techniques. Some use ordinary least squares estimation methods, while most employ maximum likelihood techniques (e.g., probit, tobit); a few employ simultaneous estimation methods designed to characterize interrelated or joint outcomes (e.g., experiencing a teen nonmarital birth and subsequent welfare recipiency). Several of the studies view the teen out-of-wedlock birth outcome as an age-dependent probabilistic phenomenon, and employ hazard rate estimation methods. The extensiveness of variables describing social and parental investments in children ranges widely across the studies.

Studies that attempt to relate the girl's own decisions to the choice-conditioned opportunities and constraints with which she is confronted are rare. The earliest is Duncan and Hoffman (1990), who estimate a two-stage logit model in which the teenage nonmarital birth choice is viewed as dependent upon the girl's comparison of income opportunities associated with alternative choices; maximum state AFDC benefits and earned family income at age 26 are taken as crude proxies of these opportunities.

A more recent effort is Rosenzweig (1995), who models the initial fertility and marriage decisions of young women (incorporating concern for child quality and assortative mating) in an attempt to identify the independent effect of AFDC benefit levels on these choices.[6] Using eight cohorts of women in the National Longitudinal Survey of Youth and a fixed effects model to control for unobservable and permanent differences across cohorts and states, he relates three mutually exclusive marriage and fertility outcomes to variables reflecting expectations of future choice-conditioned opportunities (including welfare benefits) and a measure of the women's endowments. He finds that

[6]Murray (1984) has argued that the availability and generosity of welfare benefits play an important role in explaining teen childbearing choices. However, several studies have failed to support this claim. See Duncan and Hoffman (1990), An, Haveman, and Wolfe (1993), Lundberg and Plotnick (1990), Acs (1993), Moffitt (1994), Haveman and Wolfe (1995), and Clarke and Strauss (1995). Moffitt (1992) summarizes much of this evidence.

higher welfare benefits have a small but statistically significant overall effect, but a large effect on women with poor parents.[7]

A few studies have attempted to measure the effect of the social and policy environment on children's attainments, including Lundberg and Plotnick (1990, 1995) and Haveman, Wolfe, and Peterson (1997). They find that public family planning policies (measured by state-specific indicators of abortion accessibility/costs and contraceptive availability) have large and statistically significant effects on the probability of teen childbearing.

None of these studies attempts to measure the determinants of teen nonmarital fertility-related choices (e.g., contraception, abortion) in a dynamic framework, or as these choices interact with labor supply, schooling, and post-birth marital choices.[8] We also have opted for a more static approach, and emphasize the role of choice-conditioned expectations, extensive information on family characteristics and choices, neighborhood attributes, and a characterization of the policy environment in which the girl lives in explaining this outcome. Similarly, none of the prior studies includes information on the male partners of the women, as no longitudinal data set contains linked information on mothers' nonmarital male partners. This is an important limitation of these studies, and of the results presented here.

III. A SIMPLE UTILITY MODEL OF TEEN NONMARITAL CHILDBEARING

With few exceptions, prior studies have neglected the effect on teen childbearing decisions of youth expectations of the expected utility (or income) returns to having, or not having, a child out of

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[7]Rosenzweig's variable reflecting the "real" value of welfare may confound welfare generosity with time-related changes in state-specific earnings opportunities for low earnings, low ability, and minority youths, because this latter variable remains unmeasured. Hence, his large reported welfare effect could be interpreted as a response to market opportunities. A recent effort to replicate Rosenzweig's results using an alternative data source (PSID) failed to find large and significant welfare effects using samples based on more cohorts and superior information on welfare benefit levels, parental characteristics, and measures of nonmarital births. See Hoffman and Foster (1997).

[8]Hotz and Miller (1988) and Wolpin (1984) illustrate the application of dynamic programming techniques to fertility-related choices. See also Eckstein and Wolpin (1989), who also discuss the enormous computational burden of this approach and the extensive assumptions required for estimation of identifiable model parameters.

wedlock. As a result, available estimates may attribute to background and family characteristics effects properly attributed to differences in choice-conditioned expected incomes.

Here, we set out a simple model of the teen nonmarital birth decision that rests on the view that the childbearing choice of young unmarried women reflects the response of a rational utility maximizer to expectations of utility returns associated with alternative options which are available to her. We begin with a utility function which is separable in consumption (G) and childbearing (C):

$$U_i = f(C_i,x_i) + Z_g \ln G_i + \varepsilon_{ei}\ , \tag{1}$$

where $f(C_i,x_i)$ = the nonincome effects of teen unmarried childbearing

C_i = childbearing of individual I

x_i = variables which affect the nonincome effects of teen unmarried childbearing

G_i = lifetime discounted stream of consumption

Z_g = weight of consumption in utility

ε_{ei} = random utility term (conditional on childbearing)

The generalization of f(.) to be a function of both C_i and x_i allows young women to have different nonincome effects of teen nonmarital childbearing related to their family's characteristics and those of the larger community in which they live. These characteristics—for example, parental income—may affect the perceptions and aspirations of the young women, and hence their assessment of the nonincome consequences of nonmarital childbearing. The nonincome "costs" of avoiding childbearing—for example, the cost of acquiring and using contraceptives or obtaining an abortion—are also reflected in this component of the model.

The young woman maximizes utility subject to budget constraints relating childbearing and income:

budget constraints: $Y_i = \alpha(Q_i)c_i + \xi_i$ (2)

$$G_i \leq Y_i$$

where: Y_i = lifetime discounted income stream

$\alpha(Q_i)$ = returns to an individual conditional on unmarried teen childbearing

Q_i = variables which affect returns to unmarried teen childbearing

ξ_i = random component of income

Allowing α to vary with Q allows family characteristics, and those of the larger society in which the woman lives, to affect the income returns to the teen out-of-wedlock childbearing choice.

We presume that unmarried teens do not know their future income prospects with certainty, and form expectations regarding choice-conditioned economic position by observing the incomes of an older cohort of young women with similar characteristics who make similar choices.[9]

$$E[Y_i|Q_i,c_i = c_j] = \alpha\,(Q_j)c_j \quad \text{for } Q_i = Q_j \tag{3}$$

If i chooses to give birth while unmarried and a teen, her expected income will equal that of j, who carries the same set of characteristics that influence Y through the α(Q) transformation process.

If childbearing were continuous, the solution would have the following form:

$$\underset{(\text{mrt})}{\alpha(Q_i)} = \underset{(\text{mrs})}{-[U_f df(C_i,x_i)/dC_i]} \;/U_g \tag{4}$$

indicating that the teen woman chooses childbearing such that the marginal rate of transformation (mrt) of childbearing for income equals the marginal rate of substitution (mrs) of childbearing for consumption. The choice between childbearing options reflects these marginal benefits and costs.

The role of family and broader community characteristics and prior choices is reflected in $\alpha(Q_i)$ and $-[U_f df(C_i,x_i)/dC_i]$. Variables in Q_i that increase α will increase childbearing by increasing the mrt of that choice; that is, factors that increase the returns to childbearing will increase its level. Similarly, variables that influence the nonincome effects of childbearing (mrs) will alter the level of this outcome that is observed.

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[9]This assumes that the expectation formation process is constant across all individuals. If this process is, in fact, unknown, the estimation of expected income streams may be impossible. See Manski (1993).

A discrete analogue of this continuous framework recognizes that the childbearing decision is dichotomous, such that the woman will give birth if the expected utility of childbearing exceeds the expected utility of not having a child while unmarried. Substituting the budget constraint into the utility function altered to reflect this discrete choice, and rearranging terms, reveals that the woman will bear a child if:

$$(F_1 - F_0)x + \{E[\ln Y_1] - E[\ln Y_0]\}\ Z_g > \varepsilon_0 - \varepsilon_1 \qquad (5)$$

where $F_{1,0}x = f(C,x)$. Therefore, a young, unmarried woman's decision to give birth is not entirely random, and would be repeated given the information available.

The probability that the individual will choose to give birth is:

$$Pr_1 = Pr[\ \varepsilon < Fx + \{E[\ln Y_1] - E[\ln Y_0]\}Z_g] \qquad (6)$$

where $\varepsilon = \varepsilon_0 - \varepsilon_1$ and $F = F_1 - F_0$.

Hence, the posture that we adopt focuses on the effect of the expected income possibilities available to the woman under both birth outcome possibilities. These expectations reflect the economic incentives that will influence her decision, and the characteristics and prior choices of herself, her family, and the larger society.[10] These characteristics and prior choices can also directly affect the young woman's decision, apart from any influence they have on the income expectations. As noted above, while the teen woman's choice of whether or not to give birth is often made simultaneously with schooling (e.g., whether to graduate from high school) or marital decisions, our model does not reflect the complex joint and dynamic interaction among these decisions.

[10]And, again, ideally, the characteristics and prior decisions of her sexual partners.

IV. ESTIMATION OF THE TEEN CHILDBEARING CHOICE MODEL

Empirical specification of our model relies heavily on estimation of the choice-conditioned income expectations that we attribute to teen female decision makers. However, prior to describing these expectation variables and our estimation of them and the full model, we describe the data on which our analysis relies.

A. Data on Teen Women, Their Families, and Their Neighborhoods

Our estimates are based on two large longitudinal data sets constructed from a national stratified sample of families, the Michigan Panel Study of Income Dynamics (PSID).[11] The first data set—our primary sample composed of younger individuals whose choices we model—includes 873 girls who were ages 0–6 in the beginning year of the survey; they were followed until 1988, at which time they are young adults, ranging in age from 21 to 27 years.[12] A secondary sample consisting of a somewhat older cohort includes 720 females who were aged 8–12 years in 1968, and who were 30 to 34 years old in 1988; this sample is used to estimate the choice-specific income expectation variables employed in our choice model.

For individuals in both data sets, we have extensive information on family status, income and source of income, parental education, neighborhood characteristics, and background characteristics such as race, religion, and location. In order to make comparisons of individuals with different birth years, we

[11]The PSID data provides longitudinal information on 5,000 families beginning in 1968. We use available data covering 21 years of information—from 1968 to 1988.

[12]Only those females who remained in the survey for each year until 1988 are included. In Haveman and Wolfe (1994), we studied the effect of attrition on our sample, and concluded that, with the exception of race, those who attrited do not appear to differ from the remaining sample. (Previous studies of attrition in the PSID also find little reason for concern that attrition has reduced the representativeness of the sample. See Becketti, Gould, Lillard, and Welch, 1988, and Lillard and Panis, 1994.) In a few cases, observations could not be used and are excluded from the analysis. These include persons with two or more contiguous years of missing data. Those observations with but one (contiguous) year of missing data were retained and the missing data were filled in by averaging the data for the two years contiguous to the year of missing data. For the first and last years of the sample, this averaging of the contiguous years is not possible. In this case, the contiguous year's value is assigned, adjusted if appropriate using other information that is reported.

indexed the time-varying data elements in each data set by age. All monetary values are expressed in 1976 dollars using the Consumer Price Index for all items.

We merged onto both the primary and secondary data sets state-specific policy information, including an annual, state-specific series of welfare generosity,[13] average state unemployment rates, the average per capita expenditures on family planning by the public sector in the state,[14] whether or not the state required parental consent for abortions, whether or not the state Medicaid program funds abortions, average county expenditures per capita for education, state prevalence of belonging to religious organizations, and average state median income. These jurisdiction-based policy variables are measured during the teenage years of the girls in our sample, and are matched on an annual basis to the state (county) in which the child resided each year.

Finally, we added neighborhood information to our primary data set, constructed by matching small area data from the 1970 and 1980 Censuses to the location of the children.[15] The merged neighborhood data include information on the unemployment rate, the proportion of persons in high status occupations, the proportion of youths that drop out of high school, and the proportion of families

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[13]For each state, we have annual data from 1968 to 1988 on the state maximum benefits for the Aid to Families with Dependent Children (AFDC) program, the maximum Food Stamp benefit, and the average Medicaid expenditures for AFDC families. In incorporating this information into our basic data set, we match maximum benefits (the maximum amount paid by the state as of July of that year to a family of four with no other income), in 1976 dollars (deflated by the personal consumption expenditure deflator) for the years when the child is ages 6 to 21. For food stamps, the benefit is the amount of the allotment (or the allotment minus the purchase requirement) for a family of four with no other income, again measured as of July of that year. Finally, average Medicaid expenditures for each state equal three times the state-specific fiscal year per child Medicaid expenditures for dependent children under 21 who are in categorically needy families plus the state-specific average per person annual Medicaid payments for adults in categorically needy families. These are deflated into 1976 dollars using the Current Price Index for medical care.

[14]1984 values are an average of 1983 and 1985 values for each observation; 1986 values are an average of 1985 and 1987 values.

[15]The matching was done by combining geographic codes added to the annual PSID data over the years 1968 to 1985 by the Michigan Survey Research Center to 1970 and 1980 Census data. Using 1970 and 1980 Census data, we assign neighborhood values to the neighborhood in which each family in the PSID lived to Census data. In most cases, this link is based on a match of the location of our observations to the relevant Census tract or block numbering area (67.8 percent for 1970 and 71.5 percent for 1980). For years prior to 1970 we use 1970 data; for years after 1980 we use 1980 data while for years 1971–1979 we used a weighted combination of 1970 and 1980 data (weights are .9 (1970) and .1 (1980) for 1971; .8 (1970) and .2 (1980) for 1972 and so on).

that are female headed in the neighborhood in which the family of each child in our primary sample lived for each of the years from 1968 to 1985.

B. Income Expectations with Alternative Childbearing Choices

We rely on our secondary, older cohort sample in estimating the choice-specific expected income variables for each girl in our primary sample. These expectation variables are obtained from estimated parameters of a series of personal income[16] equations fit over observations in the secondary sample, together with the relevant characteristics of the girls in our primary sample.

As a first step, we divide our older cohort into those who have and those who have not given birth through age 18 and are unmarried. We then estimate a reduced form probit equation with this dichotomous childbearing outcome as the dependent variable.[17] (In our sample of 720 older females, 128 gave birth as a teen and 592 did not.) From this equation, we calculate an inverse Mills ratio selectivity correction (lambda) variable for each person in the older sample.

Finally, we estimate 11 tobit equations (one for each year from ages 19 to 29) for each of the two childbearing groups—the group of teens who gave birth while unmarried and the group that did

[16]Personal income is defined as the sum of the person's own earnings, transfer benefits, and unearned income from all other sources. We use it rather than individual earnings, because transfer income (including welfare benefits) is not contained in earnings and therefore omits an important component of the relevant expected economic well-being concept specified in our model. We use personal income rather than family income, because the latter incorporates issues of family composition and allocation which are outside of our model and, for the most part, our observation.

[17]The variables included in this equation include mother's education (dummy variables indicating high school graduation, some college or college completion); race, family position (first born, average number of siblings), whether the mother gave birth out of wedlock as a teen, and background variables measured over ages 12–15 (average of dummies for years in each of three regions, years lived in an urban area, years head of household was disabled, years lived with one parent, years mother worked, years family received AFDC, and the years the families post-tax income was below the poverty line). Also included (and used as identifiers) are the person's own health (two dummy variables indicating if fair or poor; if excellent), mother's own teen childbearing, prevalence of religious membership in the state, and state variables related to the availability of abortions and family planning. The means and standard deviations of these variables are shown in Appendix 1. The values for ages 12–15 are used because they are the longest period available for the secondary sample.

not—with personal income as the dependent variable.[18] In each equation, we include the inverse Mills ratio selectivity term (lambda) estimated from the first stage probit estimation in order to control for self-selection into one of the childbearing outcomes. The results of the first and second stage estimates are generally as expected and are available from the authors.[19]

We use the relevant individual characteristics of each girl in our primary sample, together with the coefficients from the two sets of 11 regressions to predict income values (for each of the ages from 19 to 29) for each primary sample observation.[20] Two 11-year series of predicted expectations are obtained; one series representing income expectations should each girl give birth as an unmarried teen, and another 11-year series of expected incomes assuming that she chooses not to have a teen nonmarital birth.[21]

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[18]We included in these equations variables likely to be related to the personal income dependent variable [race (African-American = 1), family position (if first born and average number of siblings while 12–15), parental education (high school graduate, some college or college graduate dummy variables for the highest level of education each parent)], and a variety of background variables, including proportion of years lived with one parent, proportion of years mother worked, proportion of years lived in an urban area, proportion of years head of household was disabled, the average family income relative to needs, proportion of years received AFDC, the average median family income in the state in which the girl lived while age 12–15, and the average state unemployment rate in the state in which the girl lived while age 12–15. The last two variables are included to capture the variation in state labor market opportunities. For the group who gave birth as teens, the maximum AFDC payment in the state in which they lived at each age is also included. The means and standard deviations of these variables are shown in Appendix 1. The 12–15 age range for most background variables is determined by the period of observation for the older sample.

[19]In the income regressions, growing up in a family where there are many siblings is associated with a reduction in income for those who have a teen birth, but an increase in income for those who do not have a teen birth. Similarly, the coefficient estimate on the state median income variable is negative and often significant in the teen-birth income estimates, but positive and statistically significant in the without-a-teen-birth estimates. Growing up in a family that received AFDC is associated with higher incomes among individuals who have a teen birth, but has no statistically significant (and usually negative) association with income among individuals who do not have a teen birth. Although state AFDC generosity (included only in the income equations for those with a teen birth) is positive and significant in about half the equations, at older ages (above age 25) it has a negative but insignificant effect on income. The family income/needs ratio is positive and significant in about half of the no-birth equations, but not significant in any of the teen birth equations and generally small in magnitude. Being African American is positive in most of the income equations for those with a teen birth (but generally not significant), while in the no birth equations it is negative in the early years but positive in the later years (and occasionally significant). The lambda term is significant in almost all equations for those without a teen birth, and three of the equations for those with a teen birth. The other variables generally have the expected sign, but are not consistently statistically significant.

[20]We have 21 years of information on each individual in each sample. Hence, we are constrained from using incomes beyond age 29 because of the need to include childhood experience variables as predictors of incomes as an adult.

[21]In using the coefficients to predict income at each age, conditional on the individuals in our primary sample choosing to give birth out of wedlock as a teen or not, we do not use the lambda term—that is, we make an unconditional prediction. In order to avoid reducing or increasing the expected income by omitting this term, we add

The predicted mean values of these personal income expectations (and the standard deviation for each mean value) are shown in Table 1 for each of the 11 years for each of the assumed childbearing outcomes. These predicted values are shown for the entire primary sample, and for each of the two childbearing groups in that sample. The childbearing-conditioned expected income patterns are revealing. For ages 19 and 20, predicted income if the teen gives birth while unmarried is higher than if she does not give birth. However, the income trajectory in the teen-mother option shows virtually no real growth after age 20. The drop in the mid-twenties is consistent with the observed pattern of early receipt of AFDC followed by an attempt to shift to the labor market.[22] Mean expected income assuming no teen unmarried birth generally increases over the 11 years, and because of its steeper slope exceeds predicted income of a teenage mother in her twenties. (Note that the predicted incomes with and without a birth are for the same individuals.) Beginning at age 21, the relative predicted income trajectories suggest substantial gains to not giving birth as an unmarried teen.

In the second two panels of Table 1, expected incomes for the teens who did not have a nonmarital birth (with their characteristics) can be compared to those of teens who did give birth out of wedlock. After age 23, assuming a teen nonmarital birth, the predicted incomes of the girls who did not have a nonmarital birth diverges rapidly from those of the teen unmarried mothers (assuming no teen nonmarital birth, there is no difference). Perhaps most interestingly, girls who are in fact unmarried mothers have substantially higher expected incomes with that childbearing outcome than those girls who in fact did not have an unmarried birth would have had if they had chosen to give birth.[23]

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the mean expected value of the lambda times its coefficient to the constant term in predicting income for each age and for both childbearing outcomes.

[22]For example, Hotz, McElroy, and Sanders (1997) find that receipt of AFDC declines among teen mothers in their mid-twenties. They also find that controlling for background, teen mothers work few hours as teens but by their mid-twenties work more hours than if they had delayed childbearing.

[23]These predictions can be interpreted as the impact of simulating that all of the young women in the sample did not have a teen nonmarital birth. If those who did bear a child out of wedlock are simulated as having chosen not to have a birth while unmarried, they would increase their income substantially. Their incomes in early years would be greater than those of the women who actually did not give birth, but they would face a lower rate of income growth.

TABLE 1
Predicted Incomes of Young Women Who Become and Do Not Become Teenaged Mothers
(1976 dollars)

	Teen Birth		No Teen Birth	
	Mean	St. Dev.	Mean	St. Dev.
Whole Sample				
Age 19	$2,907.8	$912.8	$2,184.9	639.6
Age 20	3,069.9	1,306.9	2,703.3	905.2
Age 21	2,226.3	805.7	3,017.2	950.4
Age 22	2,100.9	1,530.5	3,845.1	1,000.1
Age 23	2,471.1	1,582.3	4,469.0	1,065.1
Age 24	1,705.1	1,624.4	4,592.1	744.0
Age 25	1,815.4	1,398.6	4,998.4	1,207.8
Age 26	2,235.8	857.7	5,370.3	1,701.3
Age 27	992.1	940.8	6,002.3	1,533.9
Age 28	2,570.4	1,100.6	5,966.7	1,301.2
Age 29	2,086.0	1,408.1	5,869.5	1,271.3
Net Present Value	**19,419.0**	**7,351.2**	**37,828.0**	**7,437.2**
Those without Teen Birth				
Age 19	2,973.6	829.0	2,203.9	630.8
Age 20	3,043.2	1,305.4	2,718.0	921.7
Age 21	2,162.5	782.9	2,991.4	957.2
Age 22	1,992.4	1,485.2	3,839.5	1,015.2
Age 23	2,363.0	1,550.7	4,490.2	1,086.3
Age 24	1,552.5	1,510.6	4,598.0	755.7
Age 25	1,700.1	1,322.5	5,011.2	1,238.2
Age 26	2,216.0	847.6	5,409.2	1,747.7
Age 27	914.7	868.1	6,036.2	1,570.6
Age 28	2,501.7	1,019.6	5,986.0	1,325.2
Age 29	1,970.1	1,317.9	5,863.5	1,287.3
Net Present Value	**18,815.0**	**6,921.4**	**39,726.0**	**7,580.7**
Those with Teen Birth				
Age 19	2,190.6	828.9	1,977.6	791.3
Age 20	3,360.8	1,292.9	2,543.0	683.6
Age 21	2,922.4	721.2	3,298.1	825.7
Age 22	3,284.2	1,524.0	3,906.4	818.6
Age 23	3,649.7	1,444.3	4,237.5	764.7
Age 24	3,369.4	1,880.0	4,527.7	600.4
Age 25	3,072.9	1,583.7	4,859.7	796.3
Age 26	2,452.4	937.6	4,946.4	982.5
Age 27	1,836.2	1,245.6	5,632.5	986.2
Age 28	3,320.5	1,577.4	5,755.0	983.4
Age 29	3,350.5	1,715.0	5,934.8	1,084.9
Net Present Value	**26,006.0**	**8,610.4**	**36,758.0**	**5,558.4**

We discount each of the choice-conditioned, age-19-to-29 expected income streams for each girl in the primary sample to age 16 (a likely age for making decisions that influence whether or not to give birth as a teen) using a discount rate of 3 percent. This procedure implicitly assumes that at age 16, each young woman in our primary sample forms her expectations of future childbearing-conditioned incomes by observing the realized incomes of persons like themselves who are in their late teens and twenties—hence, our use of conditional income terms constructed from the experiences of individuals in our secondary sample while they were ages 19 to 29. In this sense, our expected income terms may be superior to estimates of full lifetime incomes. Note that by including incomes during the late teens and early twenties, we capture the opportunity costs in the form of income forgone due to postponed working, or delayed marriage, that may be associated with early childbearing.

These present-value estimates are shown at the bottom of each panel in Table 1. The expected present value of income for the average young woman in the sample, if she were not to give birth as an unmarried teen, is $93,284 (in 1992 dollars; $36,758 on table); the average expected present value if they chose to have a nonmarital birth is $47,887, for a difference of $45,397 (in 1992 dollars). Interestingly, the gain from not giving birth as a teen is far greater for whites ($47,732) than for African Americans ($15,575; not on table).

C. The Effect of Income Expectations on Teen Childbearing Choices

For each individual, the difference in the natural logarithms of the present value of their income predictions—the one for “if no teen birth” minus that for “if a teen birth”; $E[\ln Y_1] - E[\ln Y_0]$—is taken to reflect the expected net opportunity gain associated with deciding to not bear a child out of wedlock, and included in our structural model of the decision of whether or not to give birth as a teen. This approach is similar to models involving switching with endogenous selection in Manski (1987) and Lee (1979).[24]

Several assumptions are implicit in our use of this indicator of the expected gain from choosing to forgo an unmarried birth as a proxy for the difference in her utility in the two states. First, since the consumption term is separable, we are assuming that the utility from income is the same if a woman does or does not give birth as a teen. By using the personal income of the young women, rather than their future family income relative to needs, we implicitly assume that the effects of the teen childbearing decision on the future living arrangements of the mother (including cohabitation, marriage, or living with parents) have expected benefits that are just equal to the costs (see also note 16).[25]

The estimates are presented in Table 2. The dependent variable in the models is equal to 1 if the young woman bore a child while unmarried and 18 years of age or less, and 0 otherwise.[26] Unweighted

[24]There is substantial overlap in the characteristics of the teen women who do and do not give birth out of wedlock. Our reduced form model predicting this choice fails to explain a high proportion of the choices made, suggesting but limited self-selection in terms of the economic opportunities facing young women in their choices of child birth options; adolescents with both low forgone income associated with giving birth and high forgone income are observed to both give birth and to refrain from giving birth. This avoids a potential identification problem in the use of these income expectation variables to explain the choices observed.

[25]For example, while a change in the probability of marriage associated with the teen nonmarital childbearing decision might increase the income of the household in which the woman lives, it would also increase needs and involve personal nonpecuniary benefits and costs; we are assuming that these benefits and costs net out to zero. Use of the expected difference in family income-to-needs would require a quite different set of implicit assumptions. Because children increase the level of the family needs measure, we would be assuming that children entail a reduction of the mother's utility apart from any effect on her expected future income. Further, using family income relative to needs to proxy for utility would entail assuming that, if the young woman lives with her parents, their income would tend to raise her utility, and that there are no costs associated with her living with her parents apart from those reflected in the family income relative to needs measure. Similarly, if the woman would marry or cohabit, this procedure would implicitly assume that all of the benefits of this living arrangement are reflected in the partner's income and that any costs are reflected in the increase in the measure of family needs due to the addition of another adult.

[26]There are a variety of criteria that could be used to define "teen births." We have chosen age 18 as the cutoff because most of the policy concern is directed at childbearing during ages when high school attendance is expected. We could focus on only younger ages, but the relatively rare occurrence of births at ages 15 and 16 limits our ability to explore the determinants of this outcome.

TABLE 2
Reduced Form and Structural Teen Out-of-Wedlock Childbearing Models
(Dependent Variable: Gave Birth Out of Wedlock as an Unmarried Teenager = 1)
N = 873

Variable	Reduced-Form Probit Coefficient (Std. Error)	Structural Model Coefficient[a] (Std. Error)	Structural Model Coefficient[b] (Std. Error)
LN[Pred. income if no teen birth] - LN[Pred. income if teen birth]		-1.10 (0.16)***	-0.66 (0.30)***
Average education expenditures per capita in county, ages 6–15	0.00 (0.00)	0.11 (0.38)	-0.21 (0.50)
Average of maximum state welfare benefits per month, ages 15–18	-0.03** (0.01)	-0.03 (0.01)**	-0.02 (0.02)
Average public family planning expenditures per capita, ages 13–19	-0.34 (0.18)*	-0.63 (0.18)***	-0.52 (0.22)**
Whether state Medicaid funds abortion, age 17	0.02 (0.14)	-0.11 (0.15)	-0.00 (0.19)
Whether state required parental consent for abortion, age 16	-0.37 (0.30)	-0.45 (0.31)	-0.29 (0.36)
Percentage of individuals in state who belong to a religious organization, ages 12–15	-0.02 (0.01)**	-0.03 (0.01)***	-0.03 (0.01)***
Average state median family income, ages 12–15		-0.08 (0.05)	
Average state unemployment rate, ages 12–15		-0.06 (0.06)	
African-American = 1			0.30 (0.20)
Mother high school graduate			-0.63 (0.16)***
Mother attended college			-0.99 (0.35)***
Mother gave birth as a teen			-0.04 (0.13)
Percentage of neighborhood youth who are high school dropouts, ages 6–15			0.01 (0.01)

(table continues)

TABLE 2, continued

Variable	Reduced-Form Probit Coefficient (Std. Error)	Structural Model Coefficient[a] (Std. Error)	Structural Model Coefficient[b] (Std. Error)
Proportion of years mother worked, ages 6–15			0.19 (0.20)
Proportion of years in poverty, ages 6–15			-0.22 (0.31)
Proportion of years lived with one parent, ages 6–15			0.61 (0.24)***
Average number of siblings, ages 6–15			0.15 (0.04)***
Average family after-tax income, ages 6–15			-0.01 (0.02)
Proportion of time received AFDC 10–30%, ages 6–15			-0.01 (0.17)
Proportion of time received AFDC 40–70%, ages 6–15			-0.48 (0.27)*
Proportion of time received AFDC > 70%, ages 6–15			-0.25 (0.31)
Firstborn			-0.11 (0.18)
Percentage of neighborhood adults in labor force unemployed, ages 6–15			-0.01 (0.02)
Percentage of neighborhood families headed by a female, ages 6–15			0.01 (0.01)
Percentage of neighborhood households headed by a person with a high status occupation, ages 6–15			0.00 (0.01)
Any religion = 1			-0.37 (0.23)
Proportion of years with a location move, ages 6–15			0.57 (0.38)
Proportion of years lived in SMSA, ages 6–15			0.07 (0.18)

(table continues)

TABLE 2, continued

Variable	Reduced-Form Probit Coefficient (Std. Error)	Structural Model Coefficient[a] (Std. Error)	Structural Model Coefficient[b] (Std. Error)
Average years in the Northeast, ages 12–15			-0.34 (0.25)
Average years in the South, ages 12–15			-0.28 (0.27)
Average years in the West, ages 12–15			-0.03 (0.30)
Proportion of years family head is disabled, ages 6–15			0.13 (0.22)
Two parents not present in 1968 (missing parental education)			-0.35 (0.18)*
Constant	2.15 (1.27)**	1.99 (0.68)**	1.21 (1.17)
Log-Likelihood	-350.46	-326.22	-277.88

*Significant at 10 percent level.
** Significant at 5 percent level.
*** Significant at 1 percent level.

[a]Includes difference in logs of expected income, plus policy variables.

[b]Adds family and neighborhood variables.

data are used for estimation; 125, or 14.3 percent, of the young women in our sample gave birth while an unmarried teen.[27] The means and standard deviations of these variables are shown in Appendix 2.

The first column presents a simple reduced form probit regression run using the 873 young women in our sample. This estimate does not view the childbearing decision as being influenced by the expected economic returns to either the decision to bear a child out of wedlock as a teen or the decision not to bear a child.

A series of policy variables, including those most often cited in the public debate over the causes of the rapid increase in teen nonmarital births, are included in this specification. The trends from 1976 to 1988 in these central policy variables—welfare generosity, public expenditures on family planning, and whether the Medicaid program funds abortions in the state—are shown in Figure 2.[28] The decreasing trend in the family planning/abortion variables is consistent with the recorded increase in teen nonmarital birth rates; the decline in the real value of welfare benefits over time suggests that this policy variable is unlikely to have played a sizable role in accounting for the increase in teen birth rates.[29]

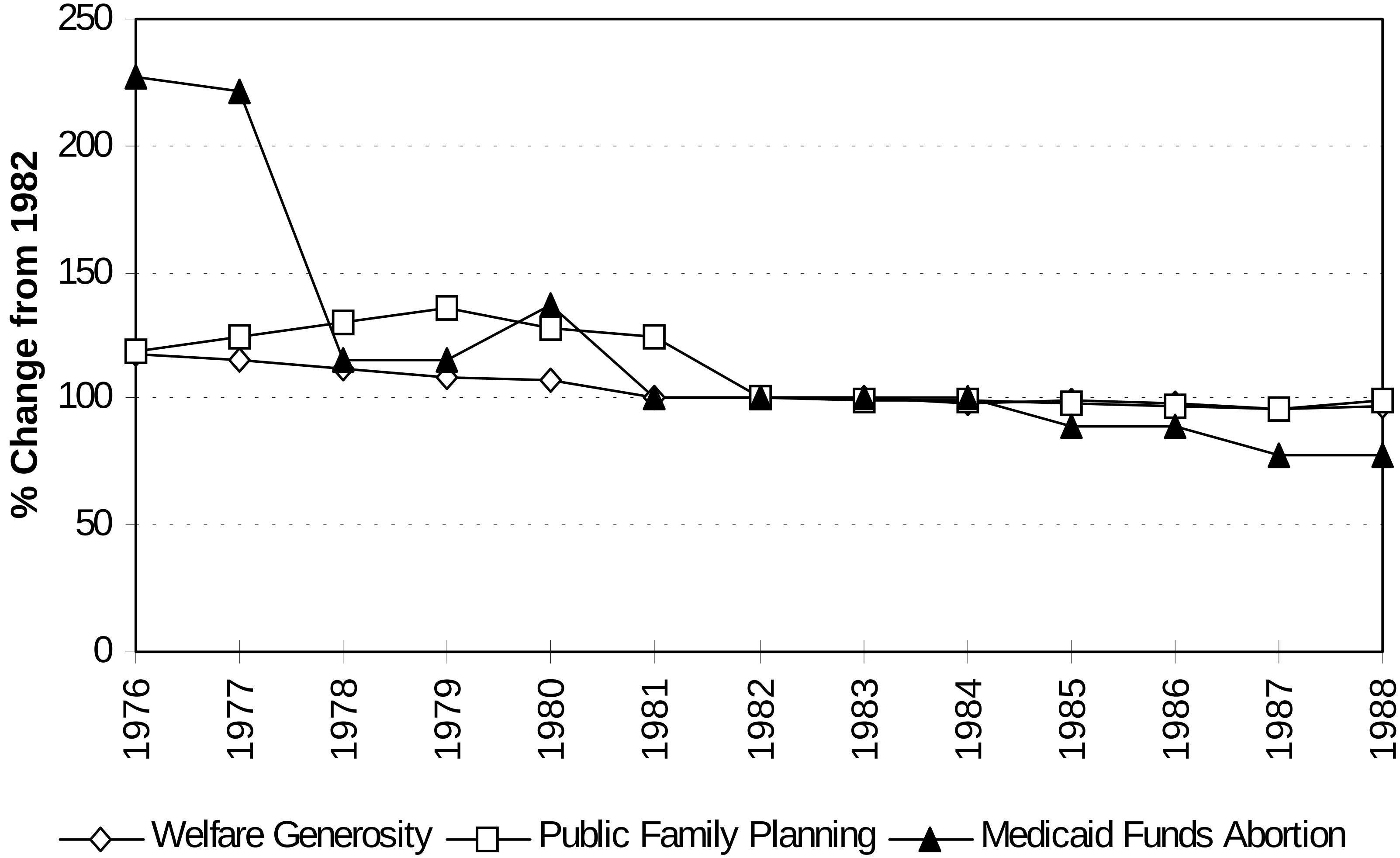
The effect of these policy variables on the individual choices of the teen women in our sample reflects cross-sectional and cohort variation rather than intertemporal variation. Both the level of family planning expenditures and the prevalence of religious membership in the state in which the woman lived while aged 13–19 are negatively related to the probability of the nonmarital birth outcome, and are statistically significant. Consistent with the observed trends in Figure 2, there is a negative and

[27]The period over which our sample would be giving birth is 1978–1986. During this time the nonmarital birth rate among girls aged less than 20 was about 11 percent (Mosher and Bachrach, 1996). The rate among African-American teens was far higher than among whites. Since our data oversamples African Americans, our higher rates are consistent with the observed rates.

[28]The period described by the state variables reflects the maximum number of years during the relevant period for which information was available. For the welfare variable, only generosity prior to the girl's childbearing choice is included. For the unemployment rate variable, we use the rate for the year in which personal income is measured for the older sample.

[29]These state policy variables are exogenous to the girl's childbearing decision. While all of them influence and are determined by a broader set of the populations within each state, the age group we study is not likely to include the median voter (for most years they are too young to vote). Moreover, most of the policy variables included in our estimates predate the decision period regarding motherhood among these young women.

Figure 2
Trends in Public Policy Variables

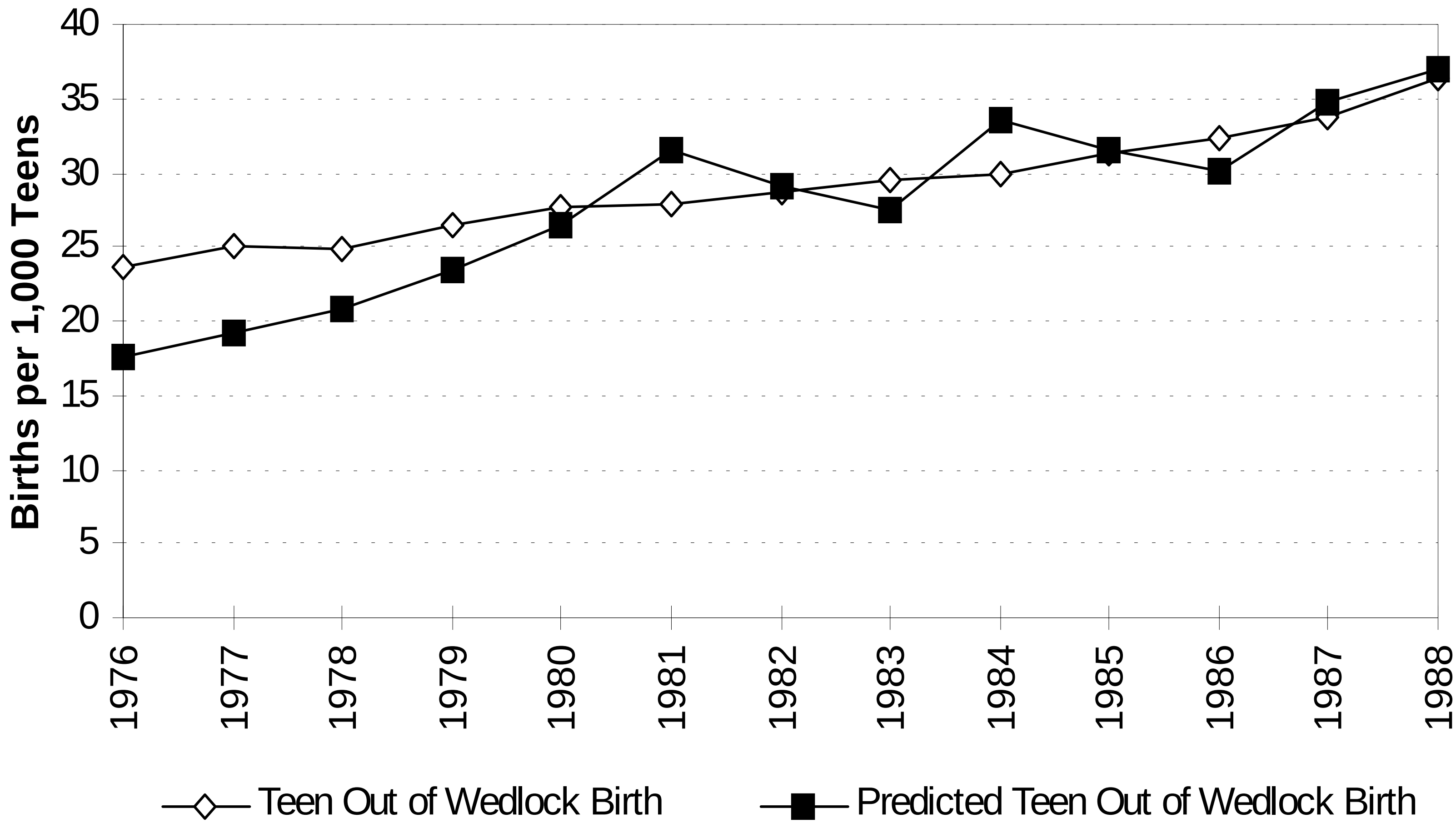


significant coefficient on the generosity of the state's welfare benefits (youths ages 15–18). The negative but insignificant coefficient on the state unemployment rate (ages 12–15) is unexpected. The other policy-related variables (education spending, Medicaid funding of abortion, or a parental consent requirement) are not statistically significant in this specification. These temporal changes in policy variables are able to track the observed trend in teen nonmarital fertility rates. In Figure 3 we used estimated coefficients from the reduced form estimate reported in column 1 of Table 2, which included only these policy variables and their time trends to predict the trend in the teen out-of-wedlock birth rate. As shown in Figure 3, the actual and predicted teen fertility variables track quite closely, even though expected income differences, background, family, and neighborhood characteristics are neglected in this exercise.

Column 2 presents a simple structural model including the difference in logs of the expected incomes together with a number of policy variables. The strongest influence of the expected utility (income) gain variable is in this equation, and suggests that young women do respond to economic opportunities. Column 3 includes the expected gain term together with a full set of policy, family, and neighborhood variables. Consistent with many social science models of children's attainments, these variables are likely to be related to the probability that a girl will give birth out of wedlock as a teenager. African-American girls and those whose parents have low education levels are more likely to have a teen nonmarital birth than are girls without these characteristics.[30] A number of these family and neighborhood variables are related in a statistically significant way to the birth outcome, including the number of years the youth lived with one parent while growing up and the average number of siblings (indicating family size). The number of times the girl changed geographic location when growing up, not

[30]Our race variable is significant in our final probit equation on teen nonmarital births when the difference in expected income is not included; however, once we control for the difference in expected incomes, race is not significant at standard levels of significance. This suggests that a major factor behind the far higher rate of teen birth among African Americans is the difference in expected incomes compared to non-African-American teenagers.

Figure 3
Trends in Teen Out of Wedlock Births



having both parents present in 1968, and whether or not she had religious affiliation are related to the probability of the nonmarital birth outcome, but are only marginally significant. The expected income variable remains negative; while the value of the t-statistic for the income gain variable is reduced in this specification, it remains significant.[31] We conclude that increasing the expected gain from not giving birth out of wedlock as a teen will reduce the rate of teen nonmarital childbearing.

The results for the policy variables are robust to the inclusion of the economic gain indicator. As in the column 1 results, family planning expenditures are statistically significant, suggesting an important potential role of this intervention in decreasing the incidence of nonmarital childbearing. None of the other state policy variables are statistically significant determinants of teen fertility patterns. Religious membership is again negatively and significantly associated with the probability of a nonmarital teen birth.[32]

D. Some Alternative Specifications

In order to test the robustness of these results, we altered the specification of the model in a variety of dimensions. First, to control for unobservable but exogenous differences across states, we estimated a fixed effects version of both the base and the extended model, including a dummy variable for each state in which a respondent resided at age 16.[33] The estimated coefficients on the remaining variables differed little from those reported in columns 2 and 3. The predicted income variable was marginally significant (t-statistic = 1.65) in the extended model, but remains statistically significant at the 1 percent level in the base model of column 2. The sign on the welfare generosity variable remains

[31]The standard errors in the probit estimations of columns 2 and 3 have not been corrected for the use of a predicted value, and hence should be interpreted with caution.

[32]We accurately predict the outcome for 86 percent of the observations. The proportion is very high (98–100 percent) for those who did not give birth, however, we correctly predict the nonbirth outcome in only 13.6 percent of the cases.

[33]Because the state dummy variables are collinear with the state religion variable, this version of the model excluded the religion variable.

negative in the base specification, has a very small positive coefficient in the full model, but is not at all significant in either the base or extended model, implying that in the base (column 2) specification the welfare variable may be capturing other, unobserved state characteristics. The coefficient on the state family planning expenditures variable remains negative and significant in the base specification of column 2, and marginally significant in the column 3 specification. A log-likelihood test to determine if the addition of the state variables adds to the fit of the model fails to reject the null hypothesis that the fit is not improved relative to the model reported in the final column of Table 2.

As an alternative to defining the income expectation term as the difference in the logarithm of the present value of expected incomes, we specified these expectation variables both as the ratio of the with-birth to the without-birth present values, and as the absolute difference between these two values. We also created a present-value term for a longer period of time by extending the ages over which we could create expected income values. Unfortunately, that required us to use the values over ages 19–29 described above and to either extrapolate the two streams from the 11 observation trend or to take the difference in the two income values observed at age 29 as continuing until retirement. The results of these specifications are similar to those reported in Table 2.

Finally, we estimated the model using the woman's income relative to needs in her family unit, instead of the personal income variables. Again, the results of these estimations are similar to those reported. The estimated coefficients on the various expectations terms constructed from the family income-to-needs variables (e.g., differences and ratios of the logs and the absolute values of these family income-to-needs variables) are consistently negative and highly statistically significant in the models shown in Table 2. This is true for the total sample, and for the African-American and white subsamples, with one exception; the estimated coefficient on the variable indicating whether the state allowed

Medicaid funding for abortions becomes statistically significant at the 10 percent level in all of the estimated models for the white subsample.[34]

We also explored the impact of these policy variables on those who grew up in a poor family—the population most likely to be affected by them. More specifically, we interacted AFDC generosity and state family planning expenditures with the variable indicating whether the girl grew up in a poor family. The interaction with welfare generosity has effectively zero statistical significance. The interaction with family planning expenditures is also not significant; the negative sign on this coefficient does suggest that family planning efforts may have a greater negative effect on the probability of a teen birth for girls who grew up in poor families relative to those who grew up in nonpoor families.

E. A Note on Model Identification

Several aspects of our structural model provide for its identification. First, we secure identification through exclusion restrictions in the first part of the model. The initial probit equation (estimated to obtain the sample selection variables) requires variables which affect the probability that individuals in the secondary sample will have a teen nonmarital birth, but which do not have an effect on the expected incomes of these young women *except* through the teen nonmarital birth outcome. We use variables describing the abortion and teen fertility prevalence in the state in which the youth lives, dummy variables for health status and whether the girl's mother gave birth as a teen, and the percentage of people in the girl's state who are members of a religious organization to gain identification in this part of the model.

Second, for the income terms to be identified, at least one variable expected to affect income expectations but not the probability of choosing to have an unmarried teen birth (other than through the income terms) must be included in the income estimations. In our estimation, five variables provide this

[34]Results from all of these estimations are available from the authors upon request.

identification for each age over which income is estimated; they are the median income of the state, the state unemployment rate, and three dummy variables indicating father's level of education. In an alternative specification in which we rely only on functional form and timing for identification, and have no exclusion restrictions, the income expectations term remains negative, but the t-statistic falls below unity. In this specification, the coefficient on median state income is negative and insignificant, those on father's education are also insignificant, while the coefficient for the state unemployment variable is negative and significant at the 10 percent level. In a set of specifications that sequentially omit identifying variables (the father's education dummy variables, the state unemployment rate, and the state median income) the results are robust, including in particular that of the income difference term, which is statistically significant at the 5 percent level in all of these tests for sensitivity to identifying restrictions.

Identification is also achieved through the nonlinear functional forms utilized in the estimation. Since consumption is entered nonlinearly into the utility function, assuming decreasing marginal utility from consumption, the predicted income terms in the final estimation stage are the natural logs of income and thus are not a linear combination of the other independent variables.[35] Finally, the timing of independent variables (over ages 12–15 in the income predictions; over ages 6–15 in the birth choice equation, a choice based on data availability) also provide identification of the model.[36]

[35] If utility is assumed to be linear in consumption, and the resulting model is estimated without functional form providing any identification, the results are very consistent with those reported. The coefficient estimate on the predicted income term has the same sign and is significant at the same levels. In addition, the other coefficient estimates are largely unchanged.

[36] Because data, rather than strong a priori theoretical justification, constrain the use of background data to ages 12–15 for the secondary sample, the model was also estimated without using the timing to provide identification. In the final stage estimation, the variables measured over ages 12–15 were used for the primary sample rather than the complete information over ages 6–15. The results from this specification are similar to those reported.

V. SIMULATIONS OF THE IMPACT OF POLICY VARIABLES

Although the coefficient estimates of our structural model indicate the sign and statistical significance of a number of policy-related variables, they reveal little regarding the quantitative impact of changes in these variables on the probability of a teen nonmarital birth. However, using these estimates, together with our estimates of the determinants of the predicted income variables, it is possible to simulate the quantitative effect of changes in a variety of variables of interest on the teen birth outcome.

The results reported in Table 3 combine the direct effect of simulated changes in the policy variables (based on the full structural model coefficient estimates of Table 2), and their indirect effect (computed through measuring the impact of the variables on expected incomes, and then translating these income changes into changes in the probability of a teen nonmarital birth using the coefficient estimate on the income gain variable in the full structural model).

The largest simulated impact is for parental education; if it is assumed that all parents with less than a high school education were to be high school graduates, our model predicts that teen nonmarital childbearing prevalence would be cut in half.[37] If we truncate at one the number of location moves per family while the child is ages 6–15, the predicted probability of a teen nonmarital birth is reduced by nearly 10 percent. Similarly, if all children were to grow up in a two-parent household, our model predicts a 25 percent reduction in the probability of giving birth as an unmarried teen.

If we simulate a 25 percent reduction in the expected income associated with having a nonmarital birth—$E[\ln Y_1]$—the probability of teen nonmarital childbearing is reduced by about 24 percent. A similar change in expected income associated with forgoing a nonmarital birth—$E[\ln Y_0]$—has a

[37]Because parental education is likely to be associated with a variety of unmeasured parental characteristics (e.g., attitudes toward education, monitoring of behavior), the mandating of high school graduation would not be likely to generate a change as large as that simulated. A similar caveat applies to our other simulated effects.

TABLE 3

Impact of Policy Variables on the Probability That a Teen Gives Birth while Unmarried, Based on Model 3 of Table 2

	Probability	Percentage Change	Direct Impact[a]	Indirect Impact[b]
Base probability	.0794			
25% increase in family planning expenditures	.0659	-17.0%		
Parents are high school graduate	.0341	-57.1%	.0434	.0651
Proportion of time spent in single-parent household reduced to zero	.0597	-24.8%	.05884	.0807
Reduce expected income if a teen mother by 25%	.0603	-24.1%		
Constrain locational moves to a maximum of one	.0721	- 9.2%		

[a]Estimated using the coefficient in the final probit equation and modifying the value of the underlying independent variable for each of the 873 observations as specified.

[b]Estimated using the coefficients in each of the underlying tobit equations used to predict income and modifying the underlying independent variable for each of the 873 observations as specified.

symmetric, though opposite signed, effect on the predicted probability of a nonmarital birth.[38] Finally, if we assume that each state increases its funding for family planning services by 25 percent, the probability of nonmarital childbearing is decreased by about 17 percent.

VI. CONCLUSION

These estimation and simulation results suggest that choice-specific income expectations appear to have a persistent influence on the childbearing decisions of teen unmarried women. Policy measures designed to increase the net return to not giving birth out of wedlock—by either increasing expected income if a birth is forgone, or reducing it if a nonmarital birth occurs—could secure reductions in teen nonmarital childbearing. The results also suggest that increasing both family planning expenditures and support for interventions designed to increase parental education and maintain intact families could also reduce the prevalence of this problem.

A current view in the United States is that reducing AFDC generosity will lead to a reduction in teen nonmarital births. Our results suggest that the relationship between welfare generosity and teen nonmarital childbearing is substantially more complex than this simple statement. Our structural model indicates that a reduction in expected income conditional on a teen having a nonmarital birth—which reduction could be caused by a reduction in welfare generosity—would tend to reduce the probability of this outcome. Moreover, state welfare benefit levels themselves are not positively related to the teen

[38]We also simulated the direct impact of a 25 percent increase in state welfare generosity (including the change through the level of the girls' choice-specific income expectations), and predicted a **decrease** in the nonmarital teen birth outcome of more than 20 percent. This pattern is consistent with the trends in AFDC generosity and teen nonmarital childbearing over the past two decades (Figure 2). While this variable is statistically significant in the column 2 model of Table 2, the value of the t-statistic is 1.2 in the column 3 model used for simulation. This may indicate that young women are little influenced by the short-term income prospects that state welfare generosity reflects, rather than the longer-term effects reflected in the expected income variable. We would caution that this negative effect could be due to the association of the state welfare variable with unmeasured state characteristics; for example, states with more generous welfare benefits may also have a higher cost of living. However, see our discussion of the estimation with state fixed effects, above.

nonmarital birth rate. Whether this indicates that welfare generosity is packaged with other state characteristics and policies, that it is unadjusted for differences in cost of living, or that teens tend not to respond to short-run differences in one source of income support such as welfare is not clear. However, the results do suggest that reducing welfare generosity by itself would not reduce the prevalence of this early childbearing problem.[39] On the other hand, our finding that increasing expenditures on family planning by the public sector would tend to reduce the probability of teen nonmarital childbearing suggests that the reduction in the real value of these expenditures and the concurrent increase in the prevalence of teen nonmarital childbearing may not be merely coincidence.

[39]Moreover, the predicted time trend in the national rate of teen nonmarital childbearing (using our reduced form model based upon only the policy variables) closely mirrors the actual pattern (Figure 3), and increases in spite of the declining trend in welfare generosity.

APPENDIX 1
Variables Used in Estimation of Selection and Income Prediction Equations
(N=718)

State of Residence		Mean	St. Dev.
Average of maximum state welfare benefits per month, ages 19–29[a]		$227	99
Average state unemployment rate ages, 19–29[a]		7.53%	2.03
Average state median family income, ages 19–29[a]		15,060	2,040
Percentage of individuals belonging to religious organizations in state, ages 12–15		24.04%	11.46
Percentage of teens ages 15–19 in state who gave birth		5.59%	1.39
Percentage of pregnant teens ages 15–19 in state who had abortion		37.04%	10.66
Percentage of births in the state which were to teens		29.77%	5.82
Background			
African-American = 1		.51	.50
Health is excellent		.59	.49
Health is bad	Health is good is missing category	.08	.27
Parental Choice/Opportunities			
Mother high school graduate = 1		.32	.47
Mother some college = 1	Mother less than high school graduate is missing category	.06	.23
Mother college graduate = 1		.04	.19
Father high school graduate = 1		.17	.38
Father some college = 1	Father less than high school graduate is missing category	.07	.25
Father college graduate = 1		.08	.27
Two parents not present in 1968 (missing education) = 1		.26	.44
Average family after-tax income, ages 12–15		15,405	10,264
Average family income/family needs standard, ages 12–15		2.09	1.54
Proportion of years lived with one parent, ages 12–15		.29	.43
Proportion of years mother worked, ages 12–15		.53	.42
Average number of siblings, ages 12–15		3.03	2.00
Proportion of years lived in SMSA, ages 12–15		.75	.41
Mother gave birth as a teen (yes=1)		.39	.49
Proportion of years on AFDC, ages 12–15		.13	.28
Family Circumstances			
Proportion of years family head is disabled, ages 12–15		.23	.37
Firstborn = 1		.17	.37
Proportion of years in poverty, ages 12–15		.27	.37
Region			
Proportion of years in the South, ages 12–15		.46	.50
Proportion of years in the West, ages 12–15	Avg. years in Midwest is missing category	.13	.33
Proportion of years in the Northeast, ages 12–15		.17	.38

[a]In the income estimation, the value of the variable at that age is used. For example, for the equation predicting income at age 19, the value of the variable when the individual was age 19 is used. However, for ease in displaying descriptive statistics, this table includes the average of the variable over the ages 19 to 29.

APPENDIX 2
Variables Used in Reduced Form and Structural Model Estimates
(N=873)

	Mean	St. Dev.
State of Residence		
Average of maximum state welfare benefits per month, ages 15–18	\$354	77
Average public family planning expenditures per capita, ages 13–19	\$1.07	.36
Whether state Medicaid funds abortions, age 17 (yes = 1)	.49	.50
Whether state required parental consent for abortion, age 16 (yes = 1)	.04	.20
Average education expenditures per capita in county, ages 6–15	\$334	175
Average state unemployment rate, ages 12–15	6.96%	1.42
Average state median family income, ages 12–15[a]	\$15,150	2,020
Percentage of individuals belonging to religious organizations in state, ages 12–15	23.69%	11.21
Background		
African-American = 1	.49	.50
Parental Choice/Opportunities		
Mother high school graduate = 1	.38	.49
Mother some college = 1 (Mother less than high school graduate is missing category)	.08	.27
Mother college graduate = 1	.04	.19
Religion = 1	.93	.26
Average family after-tax income, ages 6–15	16,390	11,510
Proportion of years lived with one parent, ages 6–15	.28	.39
Proportion of years mother worked, ages 6–15	.57	.37
Average number of siblings, ages 6–15	2.15	1.54
Proportion of years lived in SMSA, ages 6–15	.72	.43
Proportion of years with a location move, ages 6–15	.15	.17
Mother gave birth as a teen (yes=1)	.45	.50
Proportion of time family received AFDC 10–30%, ages 6–15 (Family never received AFDC is missing category)	.16	.36
Proportion of time family received AFDC 40–70%, ages 6–15	.07	.26
Proportion of time family received AFDC >70%, ages 6–15	.05	.23
Two parents not present in 1968 (missing education)	.21	.41
Family Circumstances		
Proportion of years family head is disabled, ages 6–15	.18	.29
Firstborn = 1	.21	.41
Proportion of years in poverty, ages 6–15	.23	.32
Neighborhood Attributes		
Percentage of families headed by a female, ages 6–15	19.69	13.21
Percentage of adults in labor force unemployed, ages 6–15	7.14	4.10
Percentage of youth who are high school dropouts, ages 6–15	16.67	9.45
Percentage of households headed by a person with a high status (managerial or executive) occupation, ages 6–15	19.88	9.78

[a]The state median income data for these years are based on weighted averages of the state median income in 1969, 1979, and 1989.

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