# Do Teens Make Rational Childbearing Choices?: Family, Neighborhood, and Expected Income Determinants of Teen Nonmarital Childbearing

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## I. INTRODUCTION

The prevalence of teen nonmarital childbearing has been described as the nation's "most serious social problem."[2] Even though prevalence has declined somewhat in recent years, the United States still maintains the highest rate of teen births among western industrialized countries. Annually, there are still nearly one-half million births to U.S. teenagers; they account for twelve and a half percent of all births and more than twenty one percent of all African-American births. Even more dramatic is the very high rate of nonmarital births among teens, shown in Figure 1. Today, more than three-quarters of births to teenagers are out-of-wedlock; among Non-Hispanic Blacks nearly all (95 percent) teen births are 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

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[1]While the authors are listed in reverse alphabetical order, all contributed equally to this paper.

[2]President Bill Clinton, in his 1995 State of the Union Message. Today the comment is most appropriately applied to the Latino and black populations. In 1998 the Latino teen birth rate was 93.7 and the black teen birth rate was 85.3, compared to the overall rate of 51.1 per 1,000 females 15–19 (*National Vital Statistics Report* 47:25, 1999).

the question of longer-term impacts of early, nonmarital childbearing on young mothers is unsettled among researchers,[3] the adverse effects of being born to a teen mother seem clear.[4]

The generosity and accessibility of welfare benefits have been cited as an important determinant of teen nonmarital childbearing, as has sexual education in the schools, the increased availability of child care assistance, high poverty incidence in the families and neighborhoods in which teen mothers grow up, and the poor labor market prospects available to those groups with the highest teen nonmarital birth rates.[5] Unfortunately, knowledge of the relative strength of these potentially causal linkages is weak; below, we discuss the findings of some of these studies.

In this paper, we hypothesize that those behaviors of youths that are likely to result in a teen nonmarital birth event are influenced by their expectations of the net economic benefits or costs associated with the occurrence of such a birth.[6] In short, we posit that an increase in the perceived costs associated with the occurrence of a nonmarital birth would lead to a reduction in these risky behaviors, and to a lower

[3]While teen women who have a nonmarital birth have less income, more marital instability, and lower educational attainment than those who do not, some portion of these outcomes may be attributable to unmeasured adverse family background or personal characteristics. A number of the studies in Maynard (1997) attempt to account for this selection problem in studying the consequences of adolescent childbearing. Geronimous and Korenman (1992,1993), Hoffman, Foster, and Furstenberg (1993), Brooks-Gunn, Duncan, Klebanov, and Sealand (1993), and Bronars and Grogger (1994) have also attempted to account for this potential selection effect in estimating the consequences of giving birth while an unmarried teenager for the mothers; the estimates in these studies vary widely. Hotz, McElroy, and Sanders (1997) and Hotz, Mullin, 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.

[4]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 (1996) and Wolfe and Perozek (1996). Rosenzweig and Wolpin (1995) also explore this issue.

[5]Current discussions of the sources of the recent decline in both teen pregnancy and birth rates focus on the effectiveness of abstinence programs, improved contraceptive use, and the improved labor market.

[6]We recognize that the childbearing outcome for an unmarried teenager reflects an extensive set of choices made by the woman, including whether or not to be sexually active, whether or not to use contraceptives, if pregnant whether to have an abortion, and whether or not to marry before giving birth. All of these prior choices are reflected in the nonmarital childbearing outcome which we study.

probability of a nonmarital birth event. We represent these economic incentives by the income expectations of young women conditional on either giving birth as an unmarried teen or foregoing childbearing while unmarried and an adolescent. We also measure the effects of an extensive list 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 AFDC generosity), 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 unmarried adolescent woman's childbearing choice. Her expected utility (income) conditional on either having or not having a teen nonmarital birth play a crucial role in this framework, as do family characteristics, and the neighborhood and policy environment in which she lives. After describing our data, we explain the procedures used in estimating the conditional income expectations that we attribute to the sample of teenagers whose behavior we study, and present our estimates of these conditional expectations. We then estimate our preferred model relating conditional expectations to the childbearing decision, and test the robustness of the estimates to several alternative specifications. Estimated structural parameters are then employed to simulate the effects of a variety of policy interventions on the probability of having a nonmarital birth as a teen. We find that conditional income expectations are important determinants of the childbearing choice. Since AFDC benefits are a component of the expected income terms, we also use our results to simulate the expected impact of an increase or decrease in AFDC generosity on the rate of teen nonmarital childbearing.

## II. A TOUR OF RESEARCH STUDIES

Early studies of the teen nonmarital birth outcome used cross-sectional data and reduced form estimation to relate background and family characteristics to a non-marital birth variable.[7] These studies found positive relationships between nonmarital birth probabilities and a variety of adverse parental and background circumstances experienced during childhood. A large second generation of studies used longitudinal data, with more extensive information on children's background throughout their formative years, to more fully explore the effects of background and family characteristics and choices on this outcome. Some of these studies also included information on welfare and family planning policies in states where the teen women live in order to explore the potential role of these public programs on teen nonmarital childbearing.[8] Nearly all of these studies employ cross-sectional data and reduced form estimation techniques.

The influence of family characteristics on the teen nonmarital birth outcome is substantial in these studies. The estimated coefficient on the level of mother's 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 daughters 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 disruptions and geographic moves during childhood), and parental attitude, expectations, monitoring and control of children, and contraceptive practice. The results on the effects of AFDC generosity vary widely among these studies. Those studies that have attempted to measure the effect of public family planning policies (measured by state-specific indicators of

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[7]See, for example, Hogan and Kitigawa (1985).

[8]These include Lundberg and Plotnick (1990, 1995), Plotnick (1992), An, Haveman, and Wolfe (1993), Moffitt (1994), Haveman and Wolfe (1994), and Acs (1996).

abortion accessibility/costs and contraceptive availability) tend to find large and statistically significant effects on the probability of teen childbearing.[9]

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, logit); a few employ simultaneous estimation methods designed to characterize interrelated or joint outcomes (e.g., having 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, and concerns with issues of endogeneity, unobserved variables, and model identification plague all of them. No studies attempt 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.[10] Haveman and Wolfe (1995), Robins and Fronstin (1996), and Moffitt (1992, 1998) present comprehensive descriptions of these studies, and summarize and critique their findings.

The third generation of research involves structural model estimation relating unmarried teen women's fertility decisions to the choice-conditioned opportunities and constraints with which they are confronted—including the generosity of welfare benefits. The earliest is Duncan and Hoffman (1990), who estimate a two-stage logit model in which the probability of a teenage nonmarital birth (associated with receipt of welfare) is viewed as dependent upon the girl's comparison of income expectations associated with alternative choices; maximum state AFDC benefits in case of a birth and family earnings at age 26

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[9]See also Akerlof, Yellin, and Katz, 1996, who use a game theoretic framework to explore the effect of improved contraceptive and abortion technology on the prevalence of out of wedlock births.

[10]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.

without a birth are taken as crude proxies of these choice-specific expectations. Both of these expected economic opportunity variables have the expected sign, though only the variable indexing economic opportunities without a birth is statistically significant.

A more recent effort is Rosenzweig (1998), whose model of the initial childbearing and marriage decisions of young women incorporates concern for child quality and assortative mating. A primary objective of this study is to identify the independent effect of AFDC benefit levels on these choices. 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 through age 22 to variables reflecting expectations of future choice-conditioned opportunities (including welfare benefits available to the young woman during her teenage years), and a measure of the woman's endowments. He finds that higher welfare benefits have a small but statistically significant overall effect, but a large effect on women from low income families.[11]

Clarke and Strauss (1998) use aggregate state-level data to study the effect of the level of welfare benefits, female wages, and male wages on the illegitimacy rate. After controlling for the potential endogeneity of the state's AFDC generosity using a fixed effects approach that adds controls for state and year, they find real welfare benefits have a strong and robust positive relation to teen illegitimacy, while female wages are negatively related to this outcome. The wage effect is statistically significant for whites,

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[11]Rosenzweig's variable reflecting the "real" value of welfare benefits is plagued by missing values due to the NLSY data that he uses, and hence may mismeasure the benefits available to women who move during their teenage years. That variable may also 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 reported welfare effect could also be interpreted as a response to market opportunities. Hoffman and Foster (1998) have attempted to replicate Rosenzweig's results using an alternative data source—the Michigan Panel Study of Income Dynamics (PSID)—allowing analysis that includes more cohorts and superior information on welfare benefit levels, parental characteristics, and measures of nonmarital births. They also find significant welfare effects, but only when both cohort and state fixed effects are controlled for, a finding that is at odds with other research relying on fixed effects estimation (Moffitt, 1994; Hoynes, 1997). This result is sensitive to specification, and when behavioral responses for unmarried teenagers (as opposed to those in their young-20s) are studied, no welfare effect is found. Moffitt (1998) observes that analyses by Hoynes (1997), Moffitt (1994), and Robins and Fronstin (1996) do not find stronger effects for the low-income subpopulation. (p.73).

and statistically insignificant (but of larger magnitude) for African-Americans. The wages of males, a proxy for spouse wages, are not significantly related to illegitimacy rates. The results are not robust but belong only to the fixed effects model; Moffitt (1998) criticizes the specification of their model of aggregate state-level fertility arguing that their instrument, state per capita income, probably belongs in their core or main equation.[12]

The present study extends this structural, choice-conditioned expectations approach to understanding teen nonmarital birth choices. We use separate conditional expected present value of discounted income variables for the women in our primary sample, rather than proxies for this value such as AFDC benefits, wages, or family income. These expected income values are estimated from several years of longitudinal data on a slightly older cohort of women, some of whom did and others of whom did not have a teen nonmarital birth. While we emphasize the role of the choice-conditioned income expectations, we also include more extensive information on family characteristics and choices than used in prior studies; as noted above, these factors have been found to be quantitatively important in prior research. We also extend prior research by emphasizing the potential role of neighborhood attributes and the policy environment in which the girl lives in explaining this outcome. Finally, we employ our estimates of the effects of choice-conditioned income expectations on this outcome to indicate the potential role of welfare program generosity in explaining teen girls' fertility choices. None of the earlier studies include information on the male partners of the women, as no longitudinal data set contains linked information on mothers' non-spousal male partners; our data are subject to the same limitation.

[12]Moffitt (1998) reviews the studies of the effects of state policy changes on individual behavior. He finds that these studies do not have consistent results, and are sensitive to the years studied and state-level "instruments" chosen. These estimates have also been criticized for overidentification, as well as the very restrictive assumption—that all state heterogeneity contributes to differences in teen nonmarital birth rates—on which they rest.

## III. A SIMPLE UTILITY MODEL OF TEEN NONMARITAL CHILDBEARING

Here we set out a simple, static model designed to illustrate the potential role of choice-conditioned income expectations in understanding the childbearing choices of unmarried teenage women. We presume that these young women are rational utility maximizers who form and respond to expectations of utility returns associated with the various options open to them.

Consider a utility function for an unmarried teenage woman which is separable in the nonincome and income utility of childbearing:

$$U_{ci} = B_c X_i + Z \ln G_{ci} + \varepsilon_{ci} , \tag{1}$$

where $B_c X_i$ = the nonincome effects of the childbearing choice

$X_i$ = a vector of background, family, and community variables that directly affect utility conditional on childbearing

$G_{ci}$ = lifetime discounted stream of consumption conditional on childbearing

$Z$ = weight of consumption in utility

$\varepsilon_{ci}$ = random utility term conditional on childbearing

This utility function allows the utility effect of nonmarital childbearing to differ among the young women depending on their family and community characteristics. These characteristics—for example, parental education—may affect the perceptions and aspirations of the young women, and hence their assessment of the utility effects of nonmarital childbearing, apart from the income consequences of this choice.[13] Since the consumption term, $G_i$, is separable, we are assuming that the utility from income is the same irrespective of the birth outcome.

[13]The nonincome "costs" of avoiding childbearing—for example, the disutility associated with 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: $Y_{cit} = \alpha_c Q_{it} + \xi_{cit}$ (2)

$$G_i \leq Y_i$$

where: $Y_{cit}$ = income of individual i in year t conditional on childbearing c

$Y_i$ = lifetime discounted income stream

$\alpha_c Q_{it}$ = income returns to the individual conditional on childbearing

$Q_{it}$ = variables which affect choice-conditioned income returns

$\xi_{cit}$ = random component of income for each year conditional on childbearing

In this framework, the future income stream expected by the young woman is conditioned on her decision regarding childbearing out of wedlock. For example, relative to not giving birth out of wedlock, having a nonmarital birth is likely to affect future labor supply, the extent of human capital investment, marriage outcomes, and the probability of receiving public government welfare benefits. All of these effects suggest that the future income expected by a young unmarried woman will be different if she bears a child than if she forgoes childbearing.[14] The error terms in (1) and (2) [$\varepsilon_{ci}$ and $\xi_{cit}$] reflect heterogeneity in tastes and other unobserved determinants of utility and income, respectively.[15]

We presume that unmarried teens do not know their future income prospects with certainty, and form expectations regarding their choice-conditioned income stream by observing the incomes of an older

[14]While we do not model the choice-specific human capital accumulation and marriage outcomes separately, and thus are not able to differentiate between them, their total effects (along with those of labor supply and public transfer recipiency) are captured in the choice-specific income terms.

[15]If these vectors of unobserved factors include elements that affect both the utility associated with childbearing (net of income) and income, empirical estimation will have to account for the resulting endogeneity. In one of our estimates presented below, we consider this possibility and adjust for selectivity into childbearing in estimating expected choice-conditioned incomes.

cohort of young women. Some of the women in this older cohort have chosen to have a nonmarital birth while a teenager, while others have chosen to forego childbearing while an unmarried teen.[16]

$$E[Y_i|Q_i,C_i = C_j] = \alpha_c Q_j \quad \text{for } Q_i = Q_j \tag{3}$$

If *i*, in our primary cohort, chooses to (chooses not to) give birth while unmarried, her expected income will equal that of *j*, in the older cohort, who carries the same set of characteristics that influence Y through the $\alpha_c Q$ transformation process, and who has been observed as having given (having not given) birth out of wedlock as a teen.

Within this framework, a teen will chose to have a child if the utility with a child is greater than utility without a child.[17] Substituting the budget constraint into the utility function, and rearranging terms, reveals that the woman will bear a child if:

$$(B_1 - B_0)X + Z\ \{E[\ln Y_1] - E[\ln Y_0]\} > \varepsilon_0 - \varepsilon_1 \tag{4}$$

Therefore, a young, unmarried woman's decision to have a birth is not random, and would be repeated given the information available. The direct role of family and broader community characteristics are prior choices which are reflected in B; the nonincome effects of childbearing differ depending on these other characteristics. The indirect role of family and community is reflected in a combination of $\alpha_c$ and Z; these factors affect the income stream conditional on childbearing ($\alpha_c$), which streams affect her childbearing decision (Z).

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

$$Pr_1 = Pr[\ \varepsilon < Bx + Z\{E[\ln Y_1] - E[\ln Y_0]\}] \tag{5}$$

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[16]We assume here that this expectation formation process is constant across all individuals—for example, that all young women look to the income experiences of an explicit older cohort in forming expectations of their own choice-conditioned incomes. If the "true" process of expectation formation differs from that which we assume, our estimated responses to choice-specific expectations may be biased. Without knowledge of how expectations are formed, the estimation of responses to incentives may be impossible. See Manski (1993).

[17]Because childbearing is contingent on factors such as fertility that are beyond the girls' control, it is perhaps more accurate to think of the teen's choice as whether or not to engage in risky sexual behavior, than choosing to have a teen birth.

where $\varepsilon = \varepsilon_0 - \varepsilon_1$ and $B = B_1 - B_0$.

As this model illustrates, we focus on the effect of choice-conditioned expected incomes on the nonmarital childbearing decision of teenage women. These expected income variables reflect her own characteristics and prior choices, and the characteristics of her family and the larger society in which she lives (including public policy measures such as welfare programs that provide income grants to unmarried mothers, among other groups).[18] They serve as measures of economic incentives that may influence her behavior. In this framework, personal, family, and community characteristics can also directly affect the young woman's decision, apart from any influence they have on income expectations. While this framework reflects the schooling and marital choices that may be made jointly with the nonmarital childbearing decision, it does not explain these simultaneous or sequential choices.

## IV. DATA ON THE TEEN UNMARRIED AND OLDER WOMEN COHORTS

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).[19] The first data set—our primary sample composed of young women whose choices we model—includes 873 teen age women who were ages 0–6 in the beginning year of the survey (1968); they were followed until 1988, at which time they are young adults, ranging in age from 21 to 27 years.[20] A secondary sample is a somewhat older

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

[19]The PSID data provides longitudinal information on over 5000 families beginning in 1968. We use available data covering 22 years of information—from 1968 to 1989.

[20]Only those females who remained in the survey for each year until 1988 are included. 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 year of missing data were retained and the missing information was 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. Studies of attrition in the PSID indicate that erosion of the sample has reduced its representativeness. See Becketti, Gould, Lillard, and Welch (1988), Lillard and Panis (1994), and Haveman and Wolfe (1994) . A recent study by

cohort of 720 females who were aged 8–12 years in 1968, and who were 30 to 34 years old in 1989. We assume that the teen unmarried women in the primary sample form their expectations regarding the effects of alternative childbearing choices by observing the choice-conditioned outcomes and experiences of this older cohort; hence, we derive the choice-specific income expectation variables required for our model from information on this older cohort.

For individuals in both cohorts, we have extensive longitudinal information on the status, characteristics, and choices of family members, family income (by source), living arrangements, neighborhood characteristics, and background characteristics such as race, religion, and location. In order to make comparisons of individuals with different birth years, we index the time-varying data elements in each data set by age. All monetary values are expressed in 1997 dollars using the Consumer Price Index for all items.[21]

We merged onto both data sets an extensive array of year-specific state/county policy information, characterizing the community environment within which the individual makes choices, including:

- state-specific welfare generosity,[22]

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Fitzgerald, Gottschalk, and Moffitt (1998), however, finds that, while "dropouts" from the PSID panel do differ systematically from those observations retained, behavioral responses estimated from the data do not appear to be significantly affected.

[21]While alternative indexes could be used, the Census Bureau describes the CPI as the best measure for adjusting payments to consumers when the intent is to allow them to purchase, at today's prices, the same market basket of consumer goods and services that they could purchase in an earlier reference period. "It is also the best measure to use to translate hourly and weekly earnings into inflation free dollars." http://stats.bls.gov/cpifaq.htm Question 1.

[22]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), for the years when the child is ages 6 to 21 (deflated by the personal consumption expenditure deflator). 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 using the Current Price Index for medical care. We thank Robert Moffitt for providing these data.

- state unemployment rates,
- state median income,
- per capita state expenditures on family planning,[23]
- whether or not the state required parental consent for abortions,
- whether or not the state Medicaid program funds abortions,
- whether or not the state allowed abortions pre Roe v. Wade,
- average per capita county expenditures for education, and
- the percent of state residents who belong to a religious organization.

These jurisdiction-based policy variables are matched to individuals during each of the teenage years depending on the jurisdiction of the girl's residence in each year.

Finally, we merged the following neighborhood-specific data to each person in our samples depending on their residence in each year from 1968 to 1985:

- the unemployment rate,
- the proportion of persons in high status occupations,
- median family income,
- the proportion of families with income below the poverty line,
- the proportion of youths that drop out of high school, and
- the proportion of families that are female headed.

This was accomplished by matching small area data from the 1970 and 1980 Censuses to each individual on the basis of their year-specific location.[24]

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[23]1984 values are an average of 1983 and 1985 values for each observation; 1986 values are an average of 1985 and 1987 values.

[24]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).

## V. INCOME EXPECTATIONS WITH ALTERNATIVE CHILDBEARING CHOICES

We rely on our secondary, older cohort to estimate the two choice-specific expected personal income[25] variables for each girl in the primary sample. These expectation variables are obtained from estimated parameters of a series of personal income tobit equations fit over observations in the secondary sample, together with the relevant characteristics of the girls in our primary sample. Tobit maximum likelihood is used due to a sizable number of observations with no income, especially at younger ages.[26] We estimate equation (2) for each of the 11 years from ages 19 to 29[27] for each of two groups in the older cohort (women who gave birth while an unmarried teen, and women who did not have a teen nonmarital

[25]Personal income is defined as the sum of the person's own earnings, asset income, transfer benefits (AFDC, SSI, other welfare, Social Security, veterans benefits, other retirement/pensions, unemployment insurance, and worker compensation) and unearned income from all other sources (child support, help from relatives, and "other" income). We use it rather than individual earnings since neither transfer income (including welfare benefits) nor child support is contained in earnings, hence omitting an important component of the relevant expected economic well-being concept specified in our model. We use personal income rather than family income since the latter incorporates issues of family composition and allocation which are outside of our model and, for the most part, our observation. For example, while a change in the probability of marriage associated with the teen nonmarital childbearing decision might increase the income to which the woman has access, it would also increase needs and involve personal non-pecuniary benefits and costs; we are assuming that these benefits and costs net out to zero. An alternative measure might be the expected difference in family income-to-needs. This measure would require a quite different set of implicit assumptions than use of the personal income variable. Because additional children increase the level of family needs, we would be assuming that these children reduce the mother's utility if there were no associated change in her expected income. Further, using family income relative to needs to proxy for utility in those cases in which the young woman lives with her parents implies that parental income increases the young woman's utility, and that there are no other utility costs associated with living in her parents home. 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 family needs due to the addition of another adult. In one of a series of tests of the sensitivity of our results to alternative specifications, we use an income measure that includes the income of a spouse or cohabiting partner (See Section D., below).

[26]Across the years, an average of about 11 percent of the young women in the birth subgroup had no reported personal income; the maximum with no report is 16 percent at age 20. Among the women who did not give birth, an average of 14 percent report no personal income, with a maximum of nearly 24 percent at age 19.

[27]Given the 21 years of available information, we are constrained from using incomes beyond age 29 because of the need to include childhood experience variables as regressors in the income estimates.

birth),[28] for a total of 22 estimated equations.[29] The results of these estimations are available from the authors.

We use the relevant individual characteristics of each girl in our primary sample, together with the coefficients from the two sets of 11 Tobit estimations, to predict income values (for each of the ages from 19 to 29) for each primary sample observation [equation (3)]. Two 11-year series of predicted income expectations are obtained for each teenage unmarried girl; one series representing her expected income trajectory conditional on giving birth, and another 11-year series representing her expected income trajectory if she does give birth.[30]

The mean values of these predicted 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 mean predicted values are shown for the entire primary sample, and separately for those who did and did not give birth in that sample.[31] The childbearing-conditioned expected income patterns are

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[28]Women in the older cohort who had a nonmarital birth before age 19 are included in the childbearing group; women who did not have a nonmarital birth before age 19 form the no childbearing group. Of the 720 women in the older sample, 128 gave birth as an unmarried teenager, and 592 did not.

[29]We included in these equations variables likely to be related to the personal income dependent variable, including race, family position (if first born and average number of siblings), parental education, family structure, mother's employment, urban residence, region, family location changes, disability status of family head, family income, family poverty status, family welfare recipiency, neighborhood median income and neighborhood percent in high status occupations characteristics, and state welfare generosity, median income and unemployment rates. Most of these variables are measured over the girl's ages 12–15, which range is determined by the 21 years of observations that are available. The vectors of independent variables used in the income regressions do not vary between the two teen childbearing groups. The signs on some of the coefficients in the income regressions differ between the two groups. For example, growing up in a family which consistently received welfare benefits is associated with higher income for those women who had a teen nonmarital birth, but is negatively related to income for those who did not have a teen birth. Most of the variables have similar and expected effects on personal income for the two groups. The definitions, means, and standard deviations of these variables are shown in Appendix I.

[30]These income terms are pre-tax income. It would be ideal if we could obtain estimates of disposable income by adjusting for taxes, particularly since welfare income (which is likely a larger component of personal income if the woman chooses the childbearing option than if she does not) and earned income are subject to different tax regimes. However, while we recognize this shortcoming, we are unable to reliably adjust for tax liability with the available data.

[31]The predicted incomes with and without giving birth are for the same individuals within each of the three panels—whole sample, those without teen birth and those with teen birth.

revealing. For early ages, predicted income if the unmarried teen gives birth is lower than but similar to income if she does not give birth. However, the income trajectory in the birth option shows virtually no real growth. Mean expected income assuming no birth generally increases over the 11 years, and in all years exceeds predicted income if the unmarried girl gives birth. The predicted income trajectories suggest substantial gains to not giving birth as an unmarried teen beginning at age 21.

In the second two panels of Table 1, expected incomes for the unmarried teens who did not have a birth (with their characteristics) can be compared to those who did give birth. Interestingly, the income loss associated with choosing to give birth is substantially greater for those girls who did not, in fact, give birth out of wedlock as a teen. Giving birth as an unmarried teenager does not appear to carry as substantial an income penalty for the women who did in fact have a birth.

We discount each of the choice-conditioned age 19 to 29 expected personal income streams for each girl in the primary sample to age 16 (a likely age for making decisions that influence whether or not a teen birth occurs) using a discount rate of three percent. This procedure implicitly assumes that at age 16, each young unmarried woman in our primary sample forms her expectations of future childbearing-conditioned incomes by observing the realized incomes of women with her same characteristics who are 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 20s, we capture the opportunity costs in the form of income foregone due to postponed working, or delayed marriage, that may be associated with early childbearing.[32]

[32]We performed a variety of tests of the accuracy of our predicted incomes. First, we compared the means of the predicted and actual income distributions for the reference (older) sample. In general, the predicted mean income value approximates closely actual mean income across the 11 age groups. Over the groups, the predicted mean is within 3 percent of the actual value. We also ran a series of regressions specifying actual income at each age as the dependent variable with predicted income and a constant as independent variables. We expect the coefficient on the predicted income term to be unity and the constant to be zero. The value of the constant is never statistically significantly different from zero, but is negative for ten of the eleven ages. This suggests some overestimation of the income values, but we cannot reject the null hypothesis that the actual value equals the predicted value. The coefficient estimate on the predicted income term is always very close to one, and we cannot

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, unmarried woman in the sample assuming no teen birth is $134,800; the average expected present value assuming a birth is $84,109, for a difference of $50,708. The gain from not giving birth as a teen is far greater for whites ($55,582) than for African-Americans ($23,874).

## VI. ESTIMATION OF THE TEEN CHILDBEARING CHOICE MODEL

Empirical estimation of the determinants of the teen nonmarital childbearing decision focuses on the role played by the income expectations variable. 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 is included in our structural model of the decision of whether or not to give birth as a teen. We estimate this model using a probit specification, similar to switching models in Manski (1987) and Lee (1982).[33] The dependent variable is equal to 1 if the young unmarried woman gives birth before she is 19 years old, and 0 otherwise.[34] Unweighted data are used for

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reject the null hypothesis that the coefficient is one at any age. Finally, in the regressions of predicted income on actual income, we included individual variables one at a time to see if we overestimate income systematically for some groups of individuals. There is some indication of slight overestimation for African-Americans; the coefficient on this variable is consistently negative, the value is about $100, but the t-statistic never exceeds 0.5. The coefficient estimate on regressions including family income is always positive, but again it is never close to being significant. We conclude that the specification of the income equations yields choice-conditioned predicted values that are reliable estimates of their actual counterparts.

[33]There is substantial overlap in the characteristics of the teen women who do and who do not give birth out of wedlock. A 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 foregone income associated with giving birth and high foregone 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 observed childbearing choice.

[34]Numerous criteria 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

estimation; 125, or 14.3 percent, of the young women in our primary sample gave birth while an unmarried teen.[35]

When the expected income difference variable is taken to be the only factor influencing the teen birth decision, the coefficient is negative (-.81) and statistically significant (t-statistic = 4.1). While this strong effect suggests that young women respond to economic opportunities in making childbearing choices, it does not control for family and neighborhood characteristics, nor for the direct influence of the policy environment on this choice.

### A. Our Preferred Model

The first column of Table 2 presents our preferred estimate of the determinants of the teen childbearing choice, equation (5). It is consistent with the utility maximization model of Section III, in which the error terms in (1) and (2) [$\varepsilon_{ci}$ and $\xi_{cit}$] reflect heterogeneity in tastes and other unobserved factors, and are specific to each equation. This specification includes an extensive set of family, neighborhood and policy environment variables, in addition to the expected income variable.[36] The coefficient on the net income gain from foregoing childbearing variable is negative and statistically significant, which supports the hypothesis that expectations regarding the economic consequences of nonmarital childbearing do influence the choices of young women; increasing the expected gain to foregoing childbearing while an unmarried teen seems likely to reduce the prevalence of this behavior.[37]

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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.

[35]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.

[36]The means and standard deviations of these variables are shown in Appendix II.

[37]The standard errors in these probit estimations have not been corrected for the use of a predicted value. According to Hsiao (1986) if the income equations are estimated over a sample that is independent of the sample used for the teen birth probit, conditional on the regressors, then the standard errors do not need to be corrected.

The results of this specification are also consistent with many social science models of children's attainments that indicate that family characteristics are strongly related to the probability that a girl will give birth out of wedlock as a teenager. African-American girls, those whose mother has little education, those growing up in a lower income or single parent family or a family receiving public welfare benefits or experiencing numerous location moves, and those with a large number of siblings are more likely to have a teen nonmarital birth than are girls without these characteristics; all of these family variables are related in a statistically significant way to the birth outcome. None of the location or neighborhood variables are statistically significant.

The results for the direct effects of the policy variables are also consistent with previous research. The state welfare generosity variable is not at all significant.[38] The family planning expenditure variable is negatively and significantly related to the nonmarital childbearing outcome, suggesting an important potential role of this intervention in decreasing the prevalence of nonmarital childbearing. Similarly, the prevalence of religious organization membership in the state is negatively and significantly associated with the probability of a nonmarital teen birth.

[38]Because the regression includes the expected income term, the insignificant coefficient estimate on the welfare generosity variable indicates that there is no additional effect of welfare generosity controlling for income. However, welfare generosity can affect income and thereby affect teen births; this is explored in our simulation estimates presented below. In addition, a number of studies that have attempted to estimate the effect of welfare benefits on the teen nonmarital childbearing outcome have used state (and cohort) fixed effects specifications in order to control for unobservable but exogenous differences across states (or cohorts), and have found that the welfare generosity variable has the expected positive sign only when a fixed state specification is employed (see footnote 11). Hence, we also estimated a fixed effects version of our preferred model, including a dummy variable for each state in which an observation resided at age 16. In our fixed effects estimate, the sign on the welfare generosity variable is positive though very small (.07) and not at all statistically significant (t-statistic=0.60). A log-likelihood test to determine if the addition of state variables adds to the fit of the model fails to reject the null hypothesis that fit is not improved relative to the preferred model.

*A Note on Model Identification*

Several aspects of this model provide for its identification. Identification is secured through exclusion restrictions in the first stage income model, through the nonlinear specification of the income difference term, and through timing. A standard approach to identify the first stage income model is to include at least one variable expected to affect income expectations but not the birth choice (other than through the income terms) in the income estimations. In our estimation, multiple variables which are traditionally related to earnings provide this identification for each age over which income is estimated. State level variables include the unemployment rate and median family income, and neighborhood variables include median family income and proportion in high status occupations. Identification is also achieved through the nonlinear functional forms utilized in the estimation. Since consumption is entered non-linearly into the utility function, assuming decreasing marginal utility from consumption, the predicted income term in the final stage estimation is the difference in the natural logs of predicted income and hence is not a linear combination of the other independent variables. Finally, the period over which family variables are measured differs between the analyses of the secondary sample of older women and the primary sample estimate, and this difference also contributes to model identification.[39]

In order for the exclusion restrictions to validly provide identification, the variables must be correlated with income but must not be correlated with the error term of the teen birth equation. The economic intuition for the correlation between our instruments and income is very straightforward: state income, the state unemployment rate, neighborhood median income, and the proportion of persons in the neighborhood in high status occupations are all labor market measures that are closely related to earnings

[39]The family variables used in the income estimation tobit equations are measured over ages 12–15 of the women in the secondary sample, while these family variables are measured over the longer ages 6–15 period in the final stage estimation. We are assuming that childhood environment during the girl's entire childhood affects teen childbearing, but only late childhood environment (ages 12 to 15) is related to her future income.

prospects and income.[40] There is little theoretical or observational reason to expect them to be related to the teen nonmarital birth outcome.

The presence of more than one identifying variable allows for the testing of overidentification restrictions for correlation between the instruments and the error term in the teen birth equation. The test involves regressing the residuals from the teen birth probit on the instruments with the hypothesis that the instruments are uncorrelated with the residuals. The test statistic is the uncentered R-squared multiplied by the number of observations, and this test statistic is distributed chi-squared with degrees of freedom equal to the number of overidentifying restrictions.[41] Using the residuals from the preferred model, the uncentered R-square is .003958, resulting in a test-statistic value of 3.46, well below the critical value for statistical significance, and indicating that the instruments are uncorrelated with the residuals from this final stage estimate.

B. Model with Selection

The estimates from the preferred specification rest on the assumption that the unmarried teenage women in our primary sample take the incomes of like individuals in the older cohort at face value, in effect presuming that members of the reference group had been randomly assigned to the birth or no birth groups. However, membership of the older women in the two childbearing groups may not be random, but rather

[40]Bound, Jaeger, and Baker (1995) indicate the importance of having instruments that are not just weakly correlated with the endogenous variable. Focusing on Angrist and Krueger's (1992) use of quarter of birth as an instrument for years of education, they show how biased estimates can result if the correlation between the instruments and the endogenous variable is weak, even thought the estimated relationship is statistically significant (because of, say, large sample size). Thus, economic significance is important as well as statistical significance. The economic significance of our instruments in determining income is well supported by the literature [see Datcher(1982); Corcoran and Adams (1997)]. In OLS estimation of our income equations at each age, the R-square indicator of correlation increases by an average of .05 (18.1 percent) for income with a teen birth and an average of .01 (4.2 percent) for income without a teen birth when the labor market variables used for identification are included in the regression. Alternatively, comparing the OLS estimations including only the instruments to the full specifications shows that on average the instruments account for about one-third of the full R-square indicator for income with a teen birth and 21 percent for income if there was no teen birth.

[41]See Johnston and DiNardo (1997), pages 336–338. This is asymptotically equivalent to a Basmann test.

the result of a selection process that reflects their own utility maximizing decisions. An alternative specification would attempt to statistically control for this selection process implying that the unmarried teenagers in the primary sample understand the process which leads some of the women in the secondary sample to have given birth out of wedlock, while others did not.

In this alternative specification, a two-stage Heckman-type selectivity correction model is fit over the secondary, older cohort to estimate the two, choice-specific expected personal income variables for each girl in the primary sample.[42] Having divided our older cohort into the two teen nonmarital childbearing groups (women who gave birth while an unmarried teen, and women who did not have a teen birth), we estimate a reduced form probit equation with the dichotomous childbearing outcome as the dependent variable.[43] From this equation, an inverse Mills ratio selectivity correction variable (lambda) is calculated for each of the 720 women in the older sample. The appropriate lambda selectivity correction variable is then included in each of the 11 age specific personal income equations for each of the childbearing groups [equation (3)] in order to control for selection into one of the childbearing groups.[44] Other than this lambda

[42]See Heckman (1979) and Lee (1979). This specification assumes that unobserved variables included in $\varepsilon_{ci}$ and $\xi_{cit}$ affect both childbearing choices and personal incomes. Tastes for a professional career (ambition) or aversion to statuses carrying social disapproval (stigma) such as being a teen unmarried mother or not being economically independent are examples. If this is the case, an endogeneity problem exists in obtaining reliable income predictions as incomes observed for women who did and did not have a nonmarital birth depend on factors beyond those observed and measured. In effect, parameter estimates (and hence predicted values) from income equations fit over the two groups will be biased because of the selection process that assigns observations to the groups, which process reflects the unobserved variables included in the error terms. The Heckman technique accounts for this process and yields unbiased parameter estimates in the income equations. This specification presumes that the girls in the primary sample perceive this selection process in framing their prediction of expected personal income conditional on the teen nonmarital childbearing choice.

[43]The independent variables included in this selection equation are those expected to affect the teen birth decision [equation (6)], as well as the variables expected to affect teen birth through their effect on income [equation (2)]. They include all of the variables in the full model of Table 2 (except, of course, the predicted income variable), plus the additional variables included in the income prediction equations, namely median state income, state unemployment, neighborhood median income, and proportion of persons in the neighborhood in high status occupations. Again, these variables are typically measured over the girl’s ages 12–15. The definitions, means, and standard deviations of these variables are shown in Appendix I.

[44]In the model with selection, we estimate the income equations as OLS rather than Tobit. We attempted to estimate a Tobit model with selection using maximum likelihood (see Greene, 1997), which is our preferred specification; however, the model would not converge for every age and the estimated variance matrix was singular

variable, the vector of independent variables included in these personal income equations is the same as that used in preferred model (see note 29). The results of the first and second stage estimates are generally as expected and are available from the authors. As in the preferred model, we then use the coefficient estimates from the income equations together with the relevant characteristics of each girl in the primary sample to form two 11-year streams of predicted personal income—one series if each girl should give birth as a teen and another series assuming that she chooses not to have a teen nonmarital birth.[45] Again, the difference in the present value of the logs of these two conditional income streams is the predicted income term of interest.

Column 2 of Table 2 presents the results from this three-stage model with sample selection. The coefficient estimate on the predicted income term is again negative, and is statistically significant at the 10 percent level. The coefficient estimates for the other variables are similar to those in column 1. We conclude that irrespective of the assumption made regarding the expectation formation process of unmarried teenage women, the expected income term remains an important determinant of the ultimate childbearing choice.

In this model, additional identification is necessary beyond that in the preferred model. The initial probit equation (estimated to obtain the sample selection term) requires variables which affect the probability that individuals in the secondary sample will have a teen nonmarital birth, but which do not

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for some ages. Therefore, we opted to sacrifice the Tobit specification for the consistency of the OLS estimation with the Heckman selection term. We also estimated the model using Tobit for income and including a Heckman selection term, recognizing that this could lead to inconsistent results, and obtained estimates for the final teen birth equation that are virtually identical to the results using OLS with selection. These estimates are available upon request from the authors.

[45]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 use a fitted value of the lambda term. This is appropriate provided the teen knows things about herself that are unobservable to the analyst, and these unobservable characteristics are captured in the lambda term. We are grateful to an anonymous referee for this suggestion. Controlling for sample selection results in a lower predicted income stream with a birth and a slightly higher expected income stream without a teen birth. The average expected income difference associated with not having a teen birth is higher than in the model without selection (Table 1).

have an effect on the expected incomes of these young women, except through the teen nonmarital birth outcome. We use variables describing abortion availability and family planning expenditures in the state in which the girl lives, the attitude of the state's residents toward abortions (as reflected in abortion laws prior to Roe v. Wade), the percent of the families in the teen's neighborhood that are headed by a female, a dummy variable indicating if her family is religious, the percent of people in the girl's state who are members of a religious organization, and whether the girl's mother gave birth as a teen. These are directly related to the fertility outcomes of the teen women in our primary sample, but not to their future labor market opportunities. The test of overidentifying restrictions is also passed at conventional significance levels in this model with selection.

## VII. ROBUSTNESS TESTS

The specification of our structural model rests on numerous judgements and assumptions on which responsible researchers can differ. While the model that we have designated as 'preferred' reflects our best judgments on a variety of theoretical and empirical estimation issues, it is important to determine if our estimates are robust to alternative reasonable assumptions. Hence, we have tested the robustness of our preferred model (shown in column 1, Table 2) by systematically altering the specification of the teen nonmarital childbearing model in a variety of dimensions. These include:

- Alternative timing and functional form assumptions for identification
- Alternative exclusion restrictions for model identification
- Allowing the effect of income to differ conditional on giving birth as a teen or not
- Expansion of the expected income concept to include partner's income
- Alternative discount rates
- Estimation with OLS instead of Tobit
- Estimation over the African-American subsample only

The results of these robustness tests are summarized in Table 3. In this table we present the t-statistic on the income variable(s); the log-likelihood test statistic for the significance of the entire equation and a simulation of the expected response to an increase in the expected income variable(s). The response simulations allow a comparison of the effects of specified income changes using coefficients estimated across the alternative specifications.

The first row of Table 3 shows these statistics for the preferred model, presented in Table 2 (column 1). The t-statistic on the difference in the log of the difference in the expected personal income variables—$[E(\ln Y_1) - E(\ln Y_0)]$—is significant at the .05 level, as is the log-likelihood test statistic for the significance of the entire equation. The effect of a simulated increase of 10 percent in the expected income difference variable[46] (reflecting an increase in the opportunity cost of choosing the birth option) reduces the base probability of a teen nonmarital birth by .008, from .0794 to .0714, or about 10 percent (implying an elasticity of about unity).

The second row summarizes the results of the alternative sample selection specification, shown in Table 2 (column 2). The coefficient on the expected income variable in this specification assumes that the young women in the primary sample understand the process by which women in the older cohort select between the two options, and it is also statistically significant. We estimate that a 10 percent increase in the costs of choosing the birth option would reduce the probability of a teen nonmarital birth by .005 points, or about 6 percent.

[46]In this and the other simulations reported in Table 3, we increase the expected income variable for the without birth choice by 5 percent, and reduce the expected income associated with choosing a teen birth by 5 percent, while holding all other variables at their actual level. The simulated probability is the average probability of a nonmarital teen birth for the weighted sample using the simulated values of the independent variables—that is, Simulated Probability = $1/n^* \sum F(B'X^*)$.

A. Alternative Timing and Functional Form Assumptions for Identification

Our model is identified by exclusion restrictions, timing of the measurement of childhood experiences, and the functional form of the utility function. In panel A of Table 3, we report results for alternative function form and timing specifications.

Although relating utility to the log of consumption is a common specification, it lacks a strong theoretical justification. To ensure that our nonlinear functional form assumption is not driving our results, we estimated a model that assumes utility linear in consumption. The coefficient estimate on the predicted income term is similar in magnitude to that of the preferred model, and is statistically significant at the 10 percent level (t-statistic=1.74); the other coefficient estimates are largely unchanged.

As an alternative to defining the income expectation term as the difference in the logarithm of the present value of expected incomes, we also created a ratio variable of the two conditional expected values, placing the without birth expected income value as the numerator. This ratio is also an indicator of the opportunity cost of choosing the nonmarital birth option. As Table 3 indicates, the model statistics are nearly identical to those of the preferred model (t-statistic=2.15), and the simulated effect of increasing the relative payoff to foregoing a birth is only slightly lower than in the base model. We conclude that while functional form contributes to the identification of the preferred model, the results are robust to both a linear income specification which eliminates the functional form basis for identification and alternative functional forms.

In the preferred model, the variables used in the income regressions are for the teen years, from age 12 to 15, while the variables in the teen birth equation are measured over a longer childhood span, from age 6 to 15. We judge that early childhood events are more closely related to teenage behaviors and choices than to income during the 20s; in addition, data constraints do not allow us to observe both the early childhood and later earnings for the reference group. To exclude the contribution of timing to identification, we use variables measured over ages 12 to 15 in the teen birth equation. The model statistics from this specification are similar to those of the preferred model, as is the simulated change in probability.

B. Alternative Exclusion Restrictions for Model Identification

In our discussion of the preferred model, we presented and justified the exclusion restrictions adopted to provide model identification, including both tests of instrument validity and tests of overidentification. Because alternative judgements regarding these restrictions are possible, we have examined the robustness of this specification to alternative plausible exclusion restrictions; the results are in panel B of Table 3.

In the first, and stiffest, test, we rely only on functional form and timing for identification, and include all of the identifying variables in the final stage equation of the preferred model. Given the results in panel A of Table 3, that functional form and timing are not driving the model, it is not surprising that the coefficient on the income expectation term, while negative, is not statistically significant. A log-likelihood test to determine if the addition of the identifying variables adds to the fit of the model fails to reject the null hypothesis that the fit is not improved. These results indicate that when timing and functional form alone are used for model identification, the standard errors expand substantially, and that exclusion restrictions are required to adequately identify the income term.[47]

In a series of other specifications, we sequentially add the individual identifying variables (the state unemployment and the state median income, neighborhood median income, and neighborhood high status occupation) to the teen birth specification. This procedure allows these variables to affect not only income

[47]The fact that the variables are not statistically significant in the final stage estimate does not imply that they are valid identifiers. To be valid, the variables would have to be uncorrelated with the error term of the teen birth equation. Rather, the regression of the instruments on the residuals from the teen birth equation, discussed in the identification section, provides the evidence that the instruments pass overidentification tests. Hence, the purpose of this specification is to show that timing and functional form alone are not sufficient for model identification.

but also the non-income utility associated with a teen birth. The results are quite robust to these alternative sets of exclusion restrictions. The income difference term is statistically significant at the 10 percent level in all of these specifications, save that for the model that removes state income as a restriction.[48]

We conclude that, while timing and functional form do contribute to model identification, the exclusion restrictions are the primary source of identification. In addition, while the precision of the estimation of the coefficients varies somewhat with the identification assumptions, the simulated change in the probability of a teen nonmarital birth stays within the range of -.006 to -.009—or an elasticity of .75 to 1.00—under all of these exclusion restriction alternatives.

## C. Allow the Effect of Income on Utility to Vary by Childbearing

In the theoretical model, the effect of income on utility is independent of whether the individual gave birth as a teen. As an alternative, we allow the effect of income to vary depending on the childbearing outcome by entering the two expected income variables separately into the final equation. Panel C of Table 3 indicates that the coefficient on expected income with a teen birth is positively and significantly related to the probability of a teen birth; the level of expected income if the woman foregoes childbearing is negatively related to the probability of a teen birth, but it is not significant.[49] This pattern implies that expected income opportunities conditional on having a birth are more important in the decision of a teen unmarried woman than expected income opportunities without a teen birth. This differs from the pattern estimated by Duncan and Hoffman (1990) who find a statistically significant and negative effect of income at age 26 without a teen birth and an insignificant positive effect of income at age 26 with a teen birth.[50]

[48]In this case, neither the expected income difference variable nor state income are statistically significant in the final stage equation.

[49]A test of joint significance indicates that the two variables are jointly significant at the 10 percent level when entered separately into the model.

[50]One potential reason for the discrepancy is that Duncan and Hoffman use an AFDC benefit indicator as their measure of income with a teen birth while our predicted income includes income from all sources. Indeed, in our sample, more than half of the income of teens who gave birth is from sources other than AFDC, including

D. Expanded Measure of Income to Include Partner's Income

The conditional expected income variables in our preferred specification reflect the level of personal income that teen unmarried women would expect if giving birth, or if foregoing birth. The choice of this income concept is consistent with our analytic framework, which excludes marriage/cohabitational effects of the childbearing choice.

However, it could be argued that the unmarried teen does not ignore the effects of her childbearing choice on the probability of future marital/cohabitational relationships and their implications for both the income that a possible partner might bring to the relationship and the effect of his presence on family income 'needs'. We test the robustness of our preferred model by using an expected income concept that captures both aspects of a partnering relationship.

This alternative measure of income includes both the personal income of the woman and that of a spouse or cohabiting male (if one were present).[51] In addition, it adjusts this sum by applying an economies-of-scale adjustment factor to account for the income requirements imposed by the spouse/cohabitor.[52] Use of this expanded income concept increases the levels of both predicted conditional

child support and earnings. This more inclusive measure of income with a teen birth may better capture the teen's expected income and thus be a reliable indicator of incentives related to the teen birth decision. A second potential reason for the discrepancy is in the extent to which family background characteristics are controlled. With a sparse set of family controls (only dummy variables indicating family income between $10,000–$20,000, income $20,000 or more, a dummy variable if income at age 14 was from AFDC, three region variables, and two variables for city size are included in the regression), the coefficient on the Duncan and Hoffman variable for earned family income if there is not a teen birth may be capturing unobserved family background. For example, if parental education is associated with earnings and also associated with teen birth, then not controlling for parental education in the teen birth equation will lead to a biased estimate on the expected income term. In fact, Duncan and Hoffman use mother's education as an identifying variable, assuming it is associated with income but not teen birth; in our model we allow mother's education to affect income and teen birth, and find that it is statistically significant for both.

[51]If the woman is defined by the PSID as "wife," which can either mean she is cohabitating or married, then the income term we employ is head plus wife taxable income plus the transfer income of the head and of the wife. If no partner is present, then income is defined as the female's personal income as in the preferred model.

[52]In order to adjust for economies of scale, the combined income of the woman and partner (if present) is divided by .6, consistent with the pattern of scale economies reported in Citro and Michael (1995). Use of the woman plus partner expected income-to-needs measure assumes that all of the benefits of this living arrangement

expected income variables from those shown in Table 1. The increase of expected income is substantially greater for those who do not give birth as an unmarried teen than for those who do, suggesting lower partner-income prospects for the latter group regardless of their childbearing choice.

As panel D in Table 3 shows, the expected family income variable is negative but not statistically significant in the extended model. Given the rather arbitrary definition of this variable,[53] and the implication that the unmarried teenager is accounting for both future personal income and marriage/cohabitation arrangements in assessing choice-specific economic expectations, it is not surprising that the standard error is larger in this model. The simulated change in the probability of a nonmarried teen birth (-.006) is well within the range of estimates from the other robustness tests.

E. Use of OLS rather than Tobit

While the number of observations with no personal income led us to estimate the 22 income equations using a Tobit specification, it is interesting to ask if an ordinary least squares (OLS) estimation would have altered our conclusions based on the preferred model. While the Tobit specification fits the data better than does the OLS model, the model statistics and simulated effects are virtually unchanged from the preferred model.

F. Alternative Discount Rates

The present discounted value of the expected income streams in the preferred model is obtained by discounting the 11-year trajectory of predicted values to age 16 for each girl, using a 3 percent rate. To test

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.

[53]This family income measure has a number of weaknesses. First, it includes the income of a partner only if the young woman is married or has reported living with a male for at least a year, hence excluding the contribution of a cohabitor who was not present for a full year. Moreover, use of this value implicitly assumes that those cohabitors whose incomes are included share that income with the woman. See Bumpass and Raley (1995) for a discussion of the role of cohabitation in living arrangements.

the robustness of conclusions from this model, we used a number of alternative rates, ranging from zero to 25 percent. The results from these specifications (panel F of Table 3) are quite consistent with those of the preferred model and the other robustness tests; only in the specification using the 25 percent rate is the expected income coefficient not significant.

### G. African-American Sample Only

Since the rate of teen nonmarital childbearing has been far greater among blacks than nonblacks, and since other researchers have found that the effect of welfare on fertility and marriage differs by race, (Moffitt, 1998), we also estimated the models over African-American women only. While the overall results are similar to those in the preferred model, the coefficient on the income difference term is only statistically significant at the 13 percent level. This pattern is consistent with other studies that tend to find welfare is less likely to have a statistically significant association with fertility and marriage decisions of African-American or nonwhite women compared to white women (see for example Clarke and Strauss, 1998). With the smaller size of the black sample (425 observations), the larger estimated coefficient on the choice-conditioned income term (relative to the preferred model) is dominated by a very large increase in the standard error; indeed, the simulated change in the probability of a nonmarital birth due to increasing the opportunity cost of nonmarital childbearing (-.016) is about double that of the preferred model.

## VIII. SIMULATED EFFECTS OF FAMILY VARIABLES

The coefficient estimates of our preferred model indicate the sign and statistical significance of the effect of a number family characteristics and choices on the probability of a teen nonmarital birth. Although these coefficients reveal little regarding the quantitative impact of changes in these variables on this choice, using these coefficient estimates together with assumed changes in these variables it is possible to simulate the quantitative effect of these changes on the teen birth outcome.

In Table 4, we present simulation results for those family characteristic and choice variables with statistically significant coefficients in the final stage probit equation. We present both the direct effect of simulated changes in these variables (based on the coefficient estimates of the preferred model of Table 2, column 1), and the indirect effect (computed through measuring the impact of changes in the variables on expected incomes, and in turn the effect of these income changes on the probability of a teen nonmarital birth using the coefficient estimate on the expected income difference variable). While these simulated results convey the magnitude of effect on the teen nonmarital birth outcome of changes in these characteristics, there is no implication that these characteristics are amenable to change in any reasonable period of time.

As the first row of Table 4 indicates, the base probability in our weighted sample is .0794, or nearly 8 percent. In order to compare the simulated results in this table with those in Table 3, the second row indicates that reducing the expected income associated with having a nonmarital birth—$E[\ln Y_1]$—by 25 percent decreases this probability by about 18 percent.[54]

A very large simulated impact is associated with truncating the bottom tail of the distribution of the schooling of the parents of these girls at the high school graduation level. This is a massive change in that nearly 50 percent of the girls in the sample have at least one parent with less than a high school degree.[55] Our preferred model suggests that, if such a change in parental schooling were attained, the probability of a teen nonmarital birth would fall from .08 to .046, or by 43 percent.[56] Similarly, if all of the girl's in our

[54]A similar change in expected income associated with foregoing a nonmarital birth—$E[\ln Y_0]$—has a symmetric, though opposite signed, effect on the predicted probability of a nonmarital birth. As with the simulation estimates in Table 3, these simulated effects are calculated by changing the relevant independent variable while holding all other variables at their actual level. See footnote 46.

[55]For our weighted sample, 35 percent have a mother who did not graduate high school, 38 percent have a father who did not graduate high school, and 48 percent had at least one parent that did not graduate high school.

[56]Because parental education is likely to associated with a variety of unmeasured parental characteristics (e.g., attitudes toward education, monitoring of behavior), attaining a minimum of a high school degree for every mother of a teen nonmarried woman would be unlikely to generate a change as large as that simulated. A similar caveat applies to our other simulated effects.

primary sample are assumed to have lived in an intact, two-parent family each year from ages 6–15—again, a massive change—the probability of the nonmarital birth outcome would fall to .057, or by 28 percent. Increasing average family income and decreasing the proportion of years that a girl experienced a location move by 10 percent over each girl's ages 6–15 is calculated to decrease the probability of a teen nonmarital birth by 5.9 percent and 1.8 percent, respectively. Reducing the number of siblings in the family of each girl who was not an only child is simulated to decrease the nonmarital birth outcome by 9 percent.

For all of the family variables, the direct effect is much larger than the indirect effect. While these factors affect the expected income of the teen conditional on childbearing, and thus their teen birth choice, most of the effect is a direct influence on the teen birth choice. For example, the effect of truncating the distribution of parental education at the high school graduation level increases the income difference variable, resulting in a predicted reduction in the likelihood of a teen birth by .012. In contrast the direct effect of this change in parental education is a reduction in the likelihood of a teen birth of .022. This suggests stronger effects of parental schooling through role model or information provision or monitoring channels, than through its effects on children's future expected income.

In addition to the impact of family variables, we use our preferred model to estimate the effects of simulated changes in family planning expenditures on the probability of a nonmarital birth. These results are shown in the bottom bank of Table 4. Our preferred model suggests that if state family planning expenditures were increased by 25 percent, the rate of nonmarital childbearing would decrease from about .08 to about .067, a reduction of over 15 percent. The large effect of family planning expenditures is particularly important as it is more amenable to change through policy intervention than are family-based choices and characteristics.

## IX. SIMULATED EFFECTS OF AFDC GENEROSITY

An important question concerns the potential impact of changes the generosity of welfare benefits on the teen nonmarital childbearing outcome. Simulating the effect of changing the generosity of AFDC benefits is not as straightforward as the prior simulations, and we adopt two procedures for estimating this effect.

A. Model-Based Simulation

In our first approach, we use the coefficient of the AFDC generosity variables in the income equations to simulate the effect of an increase of 25 percent in the level of welfare generosity on each girl's predicted expected income, conditional on having and not having a teen marital birth. These two revised estimates for each girl yield a revised expected income difference variable for each. We then simulate the resulting probability of giving birth as a nonmarried teen using this revised expected income difference and the coefficients from the final stage equation of our preferred model. The results of this calculation are shown in Table 4, and indicate a small increase in the probability of a teen nonmarital birth of about 1 percent attributable to the increase in AFDC benefits, or an elasticity of about .05.[57]

B. Computed Expected Income Simulation

The model-based simulation examines the marginal effect on income of an increase in AFDC generosity while holding constant the girl's family, neighborhood, and other policy characteristics in the estimated income equations. In a second estimate, we impose an increase in income proportional to the amount of income that comes from AFDC in both the with birth and without birth income terms, rather

[57]One could argue that a "direct effect" of increased welfare generosity could be obtained by increasing the value of the negatively signed and insignificant AFDC benefits variable by 25 percent in the final stage equation. Given that the teen birth equation is controlling for expected future income, and AFDC generosity is assumed to have its effect through income, we interpret this variable in the final stage equation as capturing other unobserved factors of the state that are correlated with AFDC generosity, such as social attitudes towards teen pregnancy. Because this variable is neither statistically significant nor appropriately interpreted as reflecting the effect of an increase in the monetary value of AFDC benefits, we do not shows this effect in the table. Were it to be used along with the simulated effect, the overall effect would be a reduction in the probability of a teen nonmarital birth of 15.9 percent.

than using the coefficient estimates from the income equation. We begin by calculating the percent of personal income accounted for by public welfare transfers for both those women in the reference sample who gave birth while unmarried, and those who did not have a nonmarital birth. (These percentages were 43.2 for the group who gave birth [ranging from 34.1 percent to 49.4 percent over the 11 years observed for each of these older women], and 12.1 percent for the group who did not give birth [ranging from 9.6 percent to 13.2 percent]. The transfer income component in each conditional expected income variable for each girl in the primary sample was then increased by 25 percent, after adjusting for the reduction in earned income that the increased welfare benefits would likely entail. For this offsetting labor supply effect, we relied on estimates in the previous literature, where the reduction in earned income was found to range from 15 cents to 35 cents for each dollar of additional welfare benefits.[58]

For the 15 percent replacement option, an increase in welfare generosity implies an increase in total expected income of 9.2 percent for the with birth option and an increase in expected income for the without birth option of 2.6 percent. The two expected income changes are 7 percent and 2 percent for the 35 percent replacement option. Then, using these predicted changes in expected incomes, we obtain the expected changes in the probability of a teen nonmarital birth shown in Table 4 for both the 15 and 35 earnings replacement rate assumptions.

When earnings are assumed to be replaced by increased welfare benefits at the 35 percent rate, the 25 percent increase in welfare generosity is simulated to increase the probability of a teen nonmarital birth by about 5.5 percent; the smaller earnings replacement rate implies a larger effect of welfare benefit increases on expected earnings, and implies an increase in teen nonmarital births by about 7 percent.[59]

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[58]See Moffitt (1992), Robins (1985), and Keane and Moffitt (1998). These estimates span the range of reported elasticities in these articles. For example, Keane and Moffitt report income elasticities of labor supply (hours worked) of -.15 to -.27 for female family heads.

[59]Again, we interpret the statistically significant AFDC variable in the final equation as representing unobserved characteristics of the state that are correlated with AFDC generosity.

Irrespective of the approach, our simulations suggest that the reduced expected opportunity cost of teen nonmarital childbearing caused by such a large increase in welfare benefits leads to a small increase in the nonmarital childbearing outcome. The implied elasticities range from .05 to .29.

## C. Some Indirect Evidence

As a check on these alternative estimates of the effect of AFDC generosity on the teen nonmarital birth outcome, we also adopted a quite different approach. An interesting question concerns the overall relationship of trends in those policy variables most closely associated with teen nonmarital birth choices to the aggregate national trend in this outcome (holding constant other factors, such as the demographic composition of the population, macroeconomic performance, and the choices and circumstances of the families and communities in which teen women live). If the combination of policy variables most closely tied to the teen nonmarital birth problem accurately track the actual national rate of teen nonmarital childbearing, it is difficult to attribute a primary causal effect to any single policy variable (or to changes in the demographic and economic factors held constant).

In Figure 2, we present time series trends over the 1976–1988 period for indices of those policy variables that often enter the public debate on the causes of the teen nonmarital birth problem—welfare generosity, public expenditures on family planning, and whether the Medicaid program funds abortions.[60] The decreasing trend in the two family planning/abortion variables is consistent with the increase in teen nonmarital birth rates recorded over this period (Figure 1). However, the real value of welfare benefits has declined over time, a pattern that runs contrary to assertions that welfare generosity promotes behaviors that increase the probability of nonmarital childbearing.

[60]For each year, the mean of the variable is the average over the 873 girls in our sample of the value of the variable in the state in which the girls lived in that year. The values shown in the figure are the annual levels of the mean of each variable relative to the mean of each variable in 1982.

In order to answer the question posed above, we fit a reduced form probit model to the 873 young women in our primary sample choosing as regressors those policy and community variables most often cited in the public debate over the causes of the increase in teen nonmarital births.[61] We then use these regression results in a simulation exercise designed to track the combined effect of trends in these variables on the trend in the national rate of teen nonmarital childbearing.[62] Using the individual annual, state-specific level of these policy variables for each of the 873 girls in our sample from 1976 to 1988 (reflecting the actual intertemporal time trend in the level of each variable for each girl)[63] together with the reduced form coefficients (reflecting cross-section and cohort variation rather than intertemporal variation), we predict the annual average teen nonmarital birth rate for this sample over this period.[64] The results of this exercise are shown in Figure 3; the predicted rate rises from about 17 births per year per 1000 teens in 1976 to about 36 births per year in 1988. Figure 3 also shows the actual level of the annual teen birth rate over this period. The predicted effect of change in the full constellation of policy measures tracks closely the overall change in the national rate of teen nonmarital childbearing.

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[61]In this estimation, the policy variables are taken as exogenous to the girl's childbearing decision. While all of them influence and are determined the population of each jurisdiction, the median voter in the relevant jurisdiction is not likely to be in the age group of the young women that we study; indeed, for most of the years that we observe them, they are too young to vote. Moreover, most of the policy variables included in our estimates predate the childbearing decision period of these young women.

[62]The level of state-specific public family planning expenditures, the prevalence of religious organizations averages 13–19, and whether the abortion was permitted prior to the Supreme Court ruling on Roe v. Wade (1 = only permitted abortion in order to save the life of the mother) are all negatively related to the probability of the nonmarital birth outcome, and are statistically significant. Consistent with other estimates in the literature in which few other statistical controls are included (e.g., background and family characteristics, and in some cases cohort and state fixed effects), the generosity of the state's welfare benefits is negatively and significantly related to the nonmarital birth outcome. The other policy-related variables (state of residence, education spending, Medicaid funding of abortion, or a parental consent requirement) are not statistically significant in this specification. It should be noted that the coefficients on these variables in this simple model may reflect the effect of unmeasured policy variables and state characteristics that are correlated with them.

[63]These policy variables are measured over the teenage years of the young women in our sample. The period described by the variables reflects the maximum number of years during the relevant period for which information was available.

[64]The estimated coefficients are for the probability over ages 15–18 of a teen nonmarital birth. For this simulation, the coefficients were adjusted to yield estimates of the annual probability of a teen nonmarital birth.

Given this predicted effect of the full constellation of policy measures, we conclude that the decrease in welfare generosity over this period (which, according to some, would have been expected to reduce teen nonmarital childbearing) was more than offset by changes in other policy variables (or other economic or demographic factors) that may be associated with the observed increase in this outcome. These results would seem to make the assertion that generous welfare benefits have played a major role in explaining the increase in teen, nonmarital birth rates over this period difficult to sustain.

## X. CONCLUSION

These estimation and simulation results suggest that choice-specific income expectations have a persistent influence on the childbearing decisions of teen unmarried women. Policy measures designed to increase the net return to not having a birth out of wedlock—by either increasing expected income if a birth is foregone, or reducing income expectations conditional on having a birth—may be worth while interventions for securing 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 may be efficient instruments for reducing the prevalence of this problem.

Many observers have suggested that reducing AFDC generosity, including eliminating cash assistance for teenage mothers, 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; that reducing welfare generosity by itself may have only a very small impact on teen nonmarital birth rates. Our simulations suggested that the elasticity of the probability of teen nonmarital childbearing to increases in welfare benefit generosity ranges from .05 to .29.

On the other hand, our results on the response of the teen out-of-wedlock birth variable to public family planning expenditures increases suggest that the reduction in the real value of these expenditures and the concurrent increase in the prevalence of teen nonmarital childbearing may not be mere coincidence.

**TABLE 1**
**Predicted Incomes from Base Model (in 1997 $)**

| | With Teen Birth | | No Teen Birth | |
|---|---|---|---|---|
| | Mean | (St. Dev.) | Mean | (St. Dev.) |
| Whole Sample | | | | |
| Age 19 | $6868 | ($2954) | $8913 | ($2910) |
| Age 20 | 7916 | (3167) | 10472 | (2961) |
| Age 21 | 6085 | (3590) | 14127 | (3492) |
| Age 22 | 9512 | (5455) | 14836 | (3581) |
| Age 23 | 9821 | (6086) | 15489 | (4277) |
| Age 24 | 13587 | (6657) | 16563 | (4789) |
| Age 25 | 10613 | (4772) | 17065 | (4600) |
| Age 26 | 10361 | (4487) | 16936 | (4466) |
| Age 27 | 10524 | (5004) | 18718 | (4825) |
| Age 28 | 11726 | (6438) | 20296 | (4823) |
| Age 29 | 10563 | (4597) | 20003 | (5240) |
| Net Present Value | $84,109 | ($27,620) | $134,800 | ($29,664) |
| Those Without Teen Birth | | | | |
| Age 19 | $6813 | ($2928) | $9015 | ($2928) |
| Age 20 | 7851 | (3173) | 10589 | (2964) |
| Age 21 | 5995 | (3622) | 14303 | (3470) |
| Age 22 | 9481 | (5547) | 15046 | (3519) |
| Age 23 | 9864 | (6213) | 15758 | (4250) |
| Age 24 | 13862 | (6721) | 16833 | (4763) |
| Age 25 | 10667 | (4814) | 17352 | (4588) |
| Age 26 | 10346 | (4394) | 17225 | (4421) |
| Age 27 | 10411 | (5055) | 18990 | (4804) |
| Age 28 | 11692 | (6521) | 20585 | (4756) |
| Age 29 | 10623 | (4603) | 20317 | (5231) |
| Net Present Value | $84,113 | ($28,033) | $136,810 | ($29,145) |
| Those With Teen Birth | | | | |
| Age 19 | $7471 | ($3178) | $7797 | ($2463) |
| Age 20 | 8632 | (3025) | 9201 | (2619) |
| Age 21 | 7067 | (3066) | 12205 | (3156) |
| Age 22 | 9850 | (4339) | 12541 | (3464) |
| Age 23 | 9361 | (4475) | 12548 | (3385) |
| Age 24 | 10588 | (5036) | 13621 | (4041) |
| Age 25 | 10024 | (4266) | 13932 | (3426) |
| Age 26 | 10529 | (5414) | 13784 | (3688) |
| Age 27 | 11756 | (4244) | 15757 | (4011) |
| Age 28 | 12097 | (5466) | 17140 | (4422) |
| Age 29 | 9911 | (4500) | 16569 | (3964) |
| Net Present Value | $84,065 | ($22,742) | $112,920 | ($26,394) |

**TABLE 2**
**Reduced Form and Structural Teen Out-of-Wedlock Childbearing Models; Without and With Selection (Dependent Variable: Child Had an Out of Wedlock Birth While a Teenager = 1)**
**N = 873**

| Variable | Preferred Model Coefficient (Std. Error) | Model with Selection Coefficient (Std. Error) |
|---|---|---|
| LN[Predicted Income if no teen birth] - LN[Predicted Income if teen birth] | -0.81** (0.37) | -0.42* (0.26) |
| Average education expenditures per capita in county, ages 6–15 (in thousands) | 0.01 (0.03) | 0.01 (0.03) |
| Average of maximum state welfare benefits per month, ages 15–18 (in hundreds) | -0.05 (0.07) | -0.06 (0.07) |
| Average public family planning expenditures per capita, ages 13–19 | -0.20** (0.08) | -0.16** (0.08) |
| Whether state Medicaid funds abortion, age 17 | -0.11 (0.18) | -0.17 (0.19) |
| Whether state required parental consent for abortion, age 16 | -0.14 (0.32) | -0.12 (0.35) |
| Percent of individuals in state who belong to a religious organization, ages 12–15 | -0.03** (0.01) | -0.03** (0.01) |
| Whether state restricted abortions, pre-Roe v Wade | -0.25 (0.19) | -0.28 (0.19) |
| Race (African-American) = 1 | 0.36* (0.20) | 0.38* (0.20) |
| Mother high school graduate | -0.74*** (0.19) | -0.68*** (0.18) |
| Mother attended college | -0.88** (0.36) | -0.73** (0.36) |
| Father high school graduate | 0.36 (0.25) | 0.20 (0.23) |

(table continues)

**TABLE 2, continued**

| Variable | Preferred Model Coefficient (Std. Error) | Model with Selection Coefficient (Std. Error) |
|---|---|---|
| Father attended college | 0.14 | -0.04 |
| | (0.41) | (0.39) |
| Two parents not present in 1968 | -0.22 | -0.30 |
| (missing parental education) | (0.20) | (0.19) |
| Proportion of years mother worked, | 0.13 | 0.11 |
| ages 6–15 | (0.20) | (0.20) |
| Proportion of years lived with one parent, | 0.56** | 0.65*** |
| ages 6–15 | (0.24) | (0.23) |
| Average number of siblings, ages 6–15 | 0.13*** | 0.12** |
| | (0.05) | (0.05) |
| Firstborn | -0.16 | -0.11 |
| | (0.18) | (0.18) |
| Average family income-to-needs ratio, ages 6–15 | -0.22* | -0.20 |
| | (0.13) | (0.13) |
| Proportion of years in poverty, ages 6–15 | -0.44 | -0.37 |
| | (0.33) | (0.33) |
| Proportion of time received AFDC 10–30%, | -0.05 | -0.01 |
| ages 6–15 | (0.18) | (0.18) |
| Proportion of time received AFDC 31–70%, | -0.61** | -0.50* |
| ages 6–15 | (0.29) | (0.28) |
| Proportion of time received AFDC > 70%, | -0.35 | -0.23 |
| ages 6–15 | (0.33) | (0.32) |
| Any Religion = 1 | -0.34 | -0.41* |
| | (0.23) | (0.24) |
| Proportion of years family head is | 0.09 | 0.08 |
| disabled, ages 6–15 | (0.22) | (0.22) |

(table continues)

**TABLE 2, continued**

| Variable | Preferred Model Coefficient (Std. Error) | Model with Selection Coefficient (Std. Error) |
|---|---|---|
| Mother gave birth as a teen | 0.02 | 0.00 |
| | (0.13) | (0.13) |
| Proportion of years with a location move, ages 6–15 | 0.74* | 0.68* |
| | (0.40) | (0.40) |
| Proportion of years lived in SMSA, ages 6–15 | 0.28 | 0.24 |
| | (0.20) | (0.20) |
| Percent of neighborhood youth who are high school dropouts, ages 6–15 | 0.68 | 0.69 |
| | (0.83) | (0.83) |
| Percent of neighborhood families headed by a female, ages 6–15 | -0.42 | -0.79 |
| | (1.07) | (1.06) |
| Percent of neighborhood households in poverty, ages 6–15 | 0.82 | 1.10 |
| | (1.05) | (1.04) |
| Average years in the Northeast, ages 12–15 | -0.39 | -0.35 |
| | (0.25) | (0.25) |
| Average years in the South, ages 12–15 | -0.27 | -0.38 |
| | (0.28) | (0.27) |
| Average years in the West, ages 12–15 | -0.24 | -0.30 |
| | (0.34) | (0.35) |
| Constant | 1.13 | 1.17 |
| | (1.14) | (1.16) |
| Log-Likelihood | -275.21 | -276.27 |
| Mean of Dependent Variable = 0.143 | | |

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

**TABLE 3**

**Robustness Tests of Preferred Teen Childbearing Choice Model**

| | t-statistic | Log-Likelihood | Simulated Change in Probability |
|---|---|---|---|
| **Preferred Model** (column 1, Table 2) | -2.20 | -275.21 | -.008 |
| **Model with selection (column 2, Table 2)** | -1.65 | -276.27 | -.005 |
| **Robustness Tests** | | | |
| *A. Timing and Functional Form* | | | |
| Difference of absolute expected incomes | -1.74 | -276.67 | -.008 |
| Ratio of logs of expected incomes | -2.15 | -275.16 | -.007 |
| Timing of childhood events | -1.67 | -273.51 | -.008 |
| *B. Alternative Exclusion Restrictions for Model Identification* | | | |
| No exclusion restrictions (all identifiers) | -0.92 | -272.99 | -.005 |
| With neighborhood income not excluded | -2.17 | -275.16 | -.009 |
| With neighborhood high status not excluded | -2.12 | -275.02 | -.008 |
| With state income not excluded | -1.30 | -274.73 | -.006 |
| With state unemployment not excluded | -2.29 | -274.46 | -.009 |
| *C. Utility Effect of Income Differs by Childbearing Choice* | | | |
| Expected Income without birth | 2.29 | -274.93 | -.007 |
| Expected Income with birth | -0.64 | | |
| *D. Expected Income, including Partner's Income* | -1.15 | -276.98 | -.006 |
| *E. Estimation with OLS Rather than Tobit* | -1.99 | -275.62 | -.008 |
| *F. Alternative Discount Rates* | | | |
| Discount rate = 0% | -2.26 | -275.09 | -.009 |
| Discount rate = 5% | -2.16 | -275.30 | -.008 |
| Discount rate = 10% | -2.05 | -275.53 | -.008 |
| Discount rate = 25% | -1.70 | -276.21 | -.006 |
| *G. African-American Sample Only* | -1.53 | -191.16 | -.016 |

**TABLE 4**

**Impact of Significant Family, Neighborhood, and Policy Variables on Probability of a Teen Nonmarital Birth (Based on Preferred Model)**

| | Total Simulated Change in Probability | Simulated Direct Effect[a] | Simulated Indirect Effect[b] | Total Percentage Change |
|---|---|---|---|---|
| **Base Probability = .0794** | | | | |
| **Reduce expected income if a teen mother by 25 percent** | -.0146 | NA | NA | -18.5% |
| **Family Characteristics/Choices** | | | | |
| *Parents are high school graduates* | -.0338 | -.0216 | -.0115 | -42.6 |
| *Proportion of time spent in single parent household, ages 6–15 =0* | -.0225 | -.0192 | -.0045 | -28.3 |
| *Reduce number of siblings by 1* | -.0073 | -.0098 | .0026 | -9.2 |
| *Increase family income/needs, ages 6–15, by 10 percent* | -.0047 | -.0043 | -.0004 | -5.9 |
| *Reduce the proportion of years with a location move, ages 6–15, by 10 percent* | -.0014 | -.0016 | .0002 | -1.8% |
| **Policy Variables** | | | | |
| *Increase state family planning expenditures by 25 percent* | -.0123 | NA | NA | -15.5 |
| *Increase AFDC generosity by 25 percent* | | | | |
| Model based simulation | +.0009 | — | — | +1.1 |
| Computed expected income simulation | | | | |
| —15 percent earnings replacement | +.0057 | — | — | +7.2 |
| —35 percent earnings replacement | +.0044 | — | — | +5.5 |

[a]Estimated using the coefficient in the 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.

**Note**: Because of the non-linear nature of the probit, the total effect is not a linear combination of the direct and indirect effect.

# Figure 1
# Trends in Teen Birth Rates

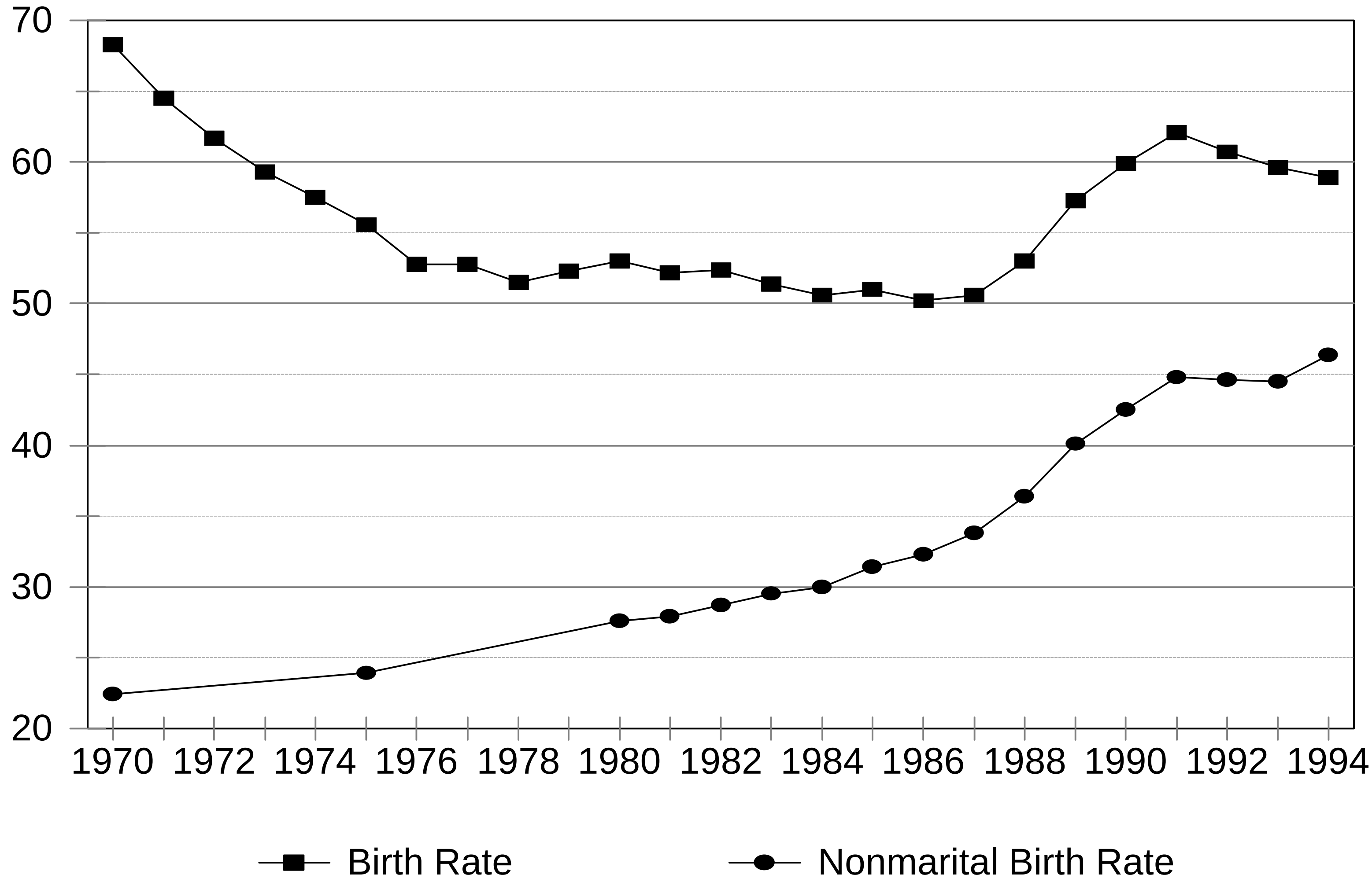
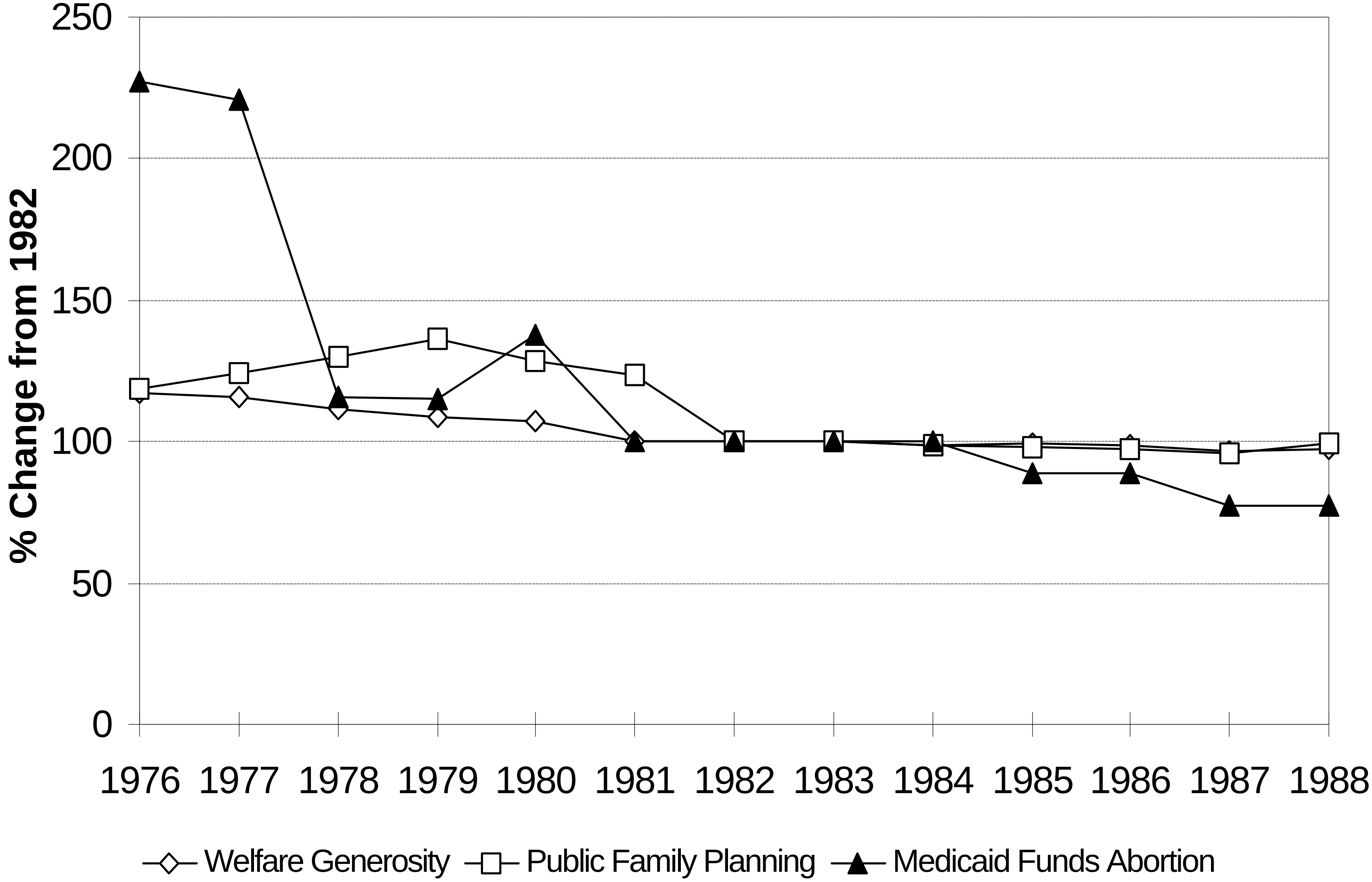
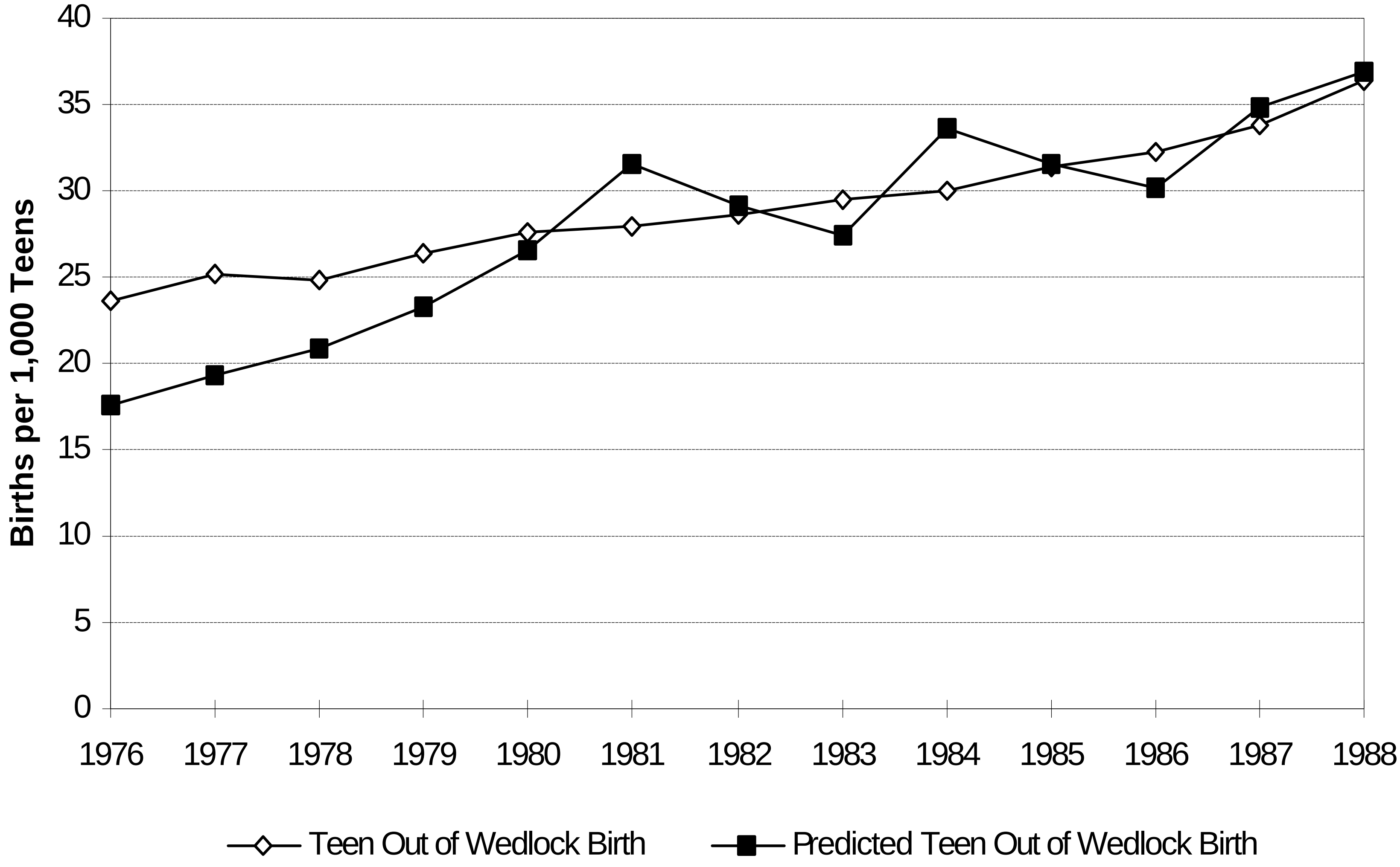


**Figure 2**
**Trends in Public Policy Variables**



## Figure 3
## Trends in Teen Out of Wedlock Births



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

| Variable | | Mean | St. Dev. |
|---|---|---|---|
| State Choices | | | |
| -Average of maximum state welfare benefits per month, ages 19–29[a] | | 640 | 279 |
| -Average state unemployment rate ages, 19–29[a] | | 7.53 | 2.03 |
| -Average state median family income, ages 19–29[a] | | 40,842 | 6,375 |
| -Average education expenditures per capita in county, ages 12–15 | | 872 | 288 |
| Background | | | |
| -Race (African-American = 1) | | .51 | .50 |
| Parental Choice/Opportunities | | | |
| -Mother high school graduate = 1[b] | | .33 | .47 |
| -Mother attended college = 1 | Mother less than high school graduate is omitted category | .10 | .29 |
| -Father high school graduate = 1[b] | | .18 | .38 |
| -Father attended college = 1 | Father less than high school graduate is omitted category | .14 | .35 |
| -Two parents not present in 1968 (missing education) = 1 | | .26 | .44 |
| -Ln Average family after-tax income, ages 12–15 | | | 10.470.66 |
| -Proportion of years lived with one parent, ages 12–15 | | .29 | .43 |
| -Proportion of years mother worked, ages 12–15 | | .53 | .42 |
| -Proportion of years with a location move, ages 12–15 | | 0.13 | 0.20 |
| -Average number of siblings, ages 12–15 | | 3.02 | 2.00 |
| -Proportion of years lived in SMSA, ages 12–15 | | .75 | .41 |
| -Proportion of years on AFDC 10–30%, ages 12–15 | Family never | .06 | .23 |
| -Proportion of years on AFDC 40–70%, ages 12–15 | received AFDC | .06 | .23 |
| -Proportion of years on AFDC > 70%, ages 12–15 | is omitted category | .11 | .31 |
| Family Circumstances | | | |
| -Proportion of years family head is disabled, ages 12–15 | | .22 | .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 omitted category | .13 | .33 |
| -Proportion of years in the Northeast, ages 12–15 | | .18 | .38 |
| Neighborhood Attributes | | | |
| -Percent of youths that dropped out of high school, ages 12–15 | | 18.09 | 10.50 |
| -Percent of families living in poverty, ages 12–15 | | 20.51 | 14.70 |
| -Median family income, ages 12–15 | | 40,050 | 15,768 |
| -Percent of households headed by a person with a high status (managerial or executive) occupation, ages 12–15 | | 18.85 | 10.49 |

[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 19–29.
[b]For the teen birth equations, parental education is just a single dummy variable per parent reflecting high school graduation or more.

**APPENDIX 2**
**Variables Used in Structural Teen Birth Model Estimates**
(N=873)

| | | Mean | St. Dev. |
|---|---|---|---|
| Expected Income | | | |
| -Ln [predicted income if no birth] - Ln [predicted income if birth] | | 0.40 | 0.39 |
| State Choices | | | |
| -Average education expenditures per capita in county, ages 6–15 | | 8,820 | 2,962 |
| -Average of maximum state welfare benefits per month, ages 15–18 | | 991 | 218 |
| -Average public family planning expenditures per capita, ages 13–19 | | 3.00 | 1.01 |
| -Whether state Medicaid funds abortions, age 17 (yes = 1) | | .49 | .50 |
| -Whether state required parental consent for abortion, age 16 (yes = 1) | | .04 | .20 |
| -Percent of individuals belonging to religious organizations in state, ages 12–15 | | 23.69 | 11.21 |
| -Whether state restricted abortions, pre-Roe v Wade | | 0.51 | 0.50 |
| Background | | | |
| -Race (African-American = 1) | | .49 | .50 |
| Parental Choice/Opportunities | | | |
| -Mother high school graduate = 1 | | .38 | .49 |
| -Mother attended college = 1 | Mother less than high school graduate is omitted category | .12 | .32 |
| -Father high school graduate = 1 | | .23 | .42 |
| -Father attended college = 1 | Father less than high school graduate is omitted category | .19 | .39 |
| -Mother gave birth as a teen ( yes=1) | | .45 | .50 |
| -Religion = 1 | | .93 | .26 |
| -Average family income/needs ratio, ages 6–15 | | 2.35 | 1.32 |
| -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 |
| -Proportion of years family received AFDC 10–30%, ages 6–15 | Family never | .16 | .36 |
| -Proportion of years family received AFDC 40–70%, ages 6–15 | received AFDC | .07 | .26 |
| -Proportion of years family received AFDC >70%, ages 6–15 | is omitted category | .05 | .23 |
| -Two parents not present in 1968 (missing education) = 1 | | .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 |
| Region | | | |
| -Proportion of years in the South, ages 6–15 | | .47 | .50 |
| -Proportion of years in the West, ages 6–15 | Avg. years in Midwest is omitted category | .11 | .31 |
| -Proportion of years in the Northeast, ages 6–15 | | .19 | .39 |
| Neighborhood Attributes | | | |
| -Percent of families headed by a female, ages 6–15 | | 19.69 | 13.21 |
| -Percent of youth who are high school dropouts, ages 6–15 | | 16.67 | 9.45 |
| -Percent of families living in poverty, ages 6-15 | | 19.48 | 13.19 |

[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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