Robert M.
La Follette School of Public Affairs
at the University of Wisconsin-Madison

Working Paper Series
La Follette School Working Paper No. 2006-020
http://www.lafollette.wisc.edu/publications/workingpapers

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Abstract

We hypothesize that the choices and behaviors of adolescent females that may result in a teen nonmarital birth event are influenced by their perceptions of the consequences of their choices. Two categories of such choice-conditioned longer-term effects are explored: (1) a teen's expected future marriage/cohabitation relationship pattern, and (2) the present value of future expected income. We also measure the effects of an extensive list of other factors, including the characteristics of the girl's family and their choices, the social and economic environment in which the adolescent lives (including policy-related factors, such as public expenditures by states on family planning programs), her neighborhood's characteristics and her own prior choices. The empirical work uses the Michigan Panel Study of Income Dynamics. The results provide evidence that choice-conditioned expected longer-term marriage/cohabitation relationships and expected economic outcomes influence the choices of unmarried teen girls, with a suggestion that expected marriage/cohabitation relationships may play a more important role than expected income. The estimates also suggest a significant effect of family planning efforts on the teen nonmarital childbearing outcome.

Key words: teen births; nonmarital fertility, marriage; cohabitation, family planning.

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I. INTRODUCTION

The prevalence of teen childbearing has been a topic of considerable public concern in the United States for the last two decades. Even though the rate of teen births in the U.S. has declined substantially over recent years, this rate remains highest among western industrialized countries. The U.S. teen birth rate stood at 61.8 births per thousand in 1991, experienced a drop of approximately a third over the subsequent dozen years, and stood at 41.2 births per thousand in 2004.[2] Even with this large drop, there are more than four hundred and twenty thousand births to U.S. teenagers each year; these births account for more than 10 percent of all births and seventeen percent of all African-American births. Even more staggering is the very high rate of births among unmarried teens; today, eighty-three percent of births to teenagers are out-of-wedlock; among Blacks nearly all (96 percent) teen births are out-of-wedlock. (Hamilton et al. 2005)

This teen nonmarital childbearing phenomenon is viewed as a social and economic problem because of the presumed adverse effects on the human capital and the future productivity of both the mothers and their children. While the question of longer-term impacts of early, nonmarital

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[1]The authors gratefully acknowledge the contributions of Scott Niemann, Kathryn Wilson, Yuichi Kitamura, Susan Lee and Dawn Duren. Special thanks go to Guilio Zanella for his assistance with calculations done for the final version of the paper. The views expressed in the paper are those of the authors, and should not be interpreted as those of the Congressional Budget Office.

[2]Then-President Bill Clinton, in his 1995 State of the Union Message referred to this problem as the nation's "most serious social problem." Today the comment is most appropriately applied to the Latino and black populations. In 20041998 the Latino teen birth rate was 93.82.6 and the black non-Latino teen birth rate was 85.462.7, compared to the overall rate of 41.251.1 per 1,000 females 15–19 (Hamilton, et al., 2005; Ventura et al., 2000).

childbearing on young mothers is unsettled among researchers,[3] the adverse effects on children born to a teen mother seem clear.[4]

In this paper, we hypothesize that the choices and behaviors of youths that may result in a teen nonmarital birth event are influenced by their perceptions of the consequences of their choices. We distinguish two categories of such choice-conditioned long-term effects for the potential mother—the effects on: (1) expectations regarding the attainment of a long-term and stable family-type relationship, and (2) her expected income stream. In particular, we hypothesize that the occurrence of a teen nonmarital birth has a negative effect on both the woman's expected future income and her chances of securing a long-term, stable family relationship. In turn, we hypothesize that the woman's expectations regarding these economic and family-based effects will influence her decisions related to sexual behavior (including the use of contraceptives), and hence the probability of a nonmarital birth event. In short, we seek to estimate the behavioral response of young unmarried women to perceptions of the economic and family-related incentives created by these consequences, as this behavior is revealed in the nonmarital birth outcome. In the process, we also measure the relationship to the nonmarital birth outcome 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-

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[3]While teen women who have a nonmarital birth tend to 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. Hotz, McElroy, and Sanders (1997), Hotz, Sanders McElroy (1999) 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. See also Geronimous and Korenman (1992,1993), Hoffman, Foster, and Furstenberg (1993), Brooks-Gunn, Duncan, Klebanov, and Sealand (1993), and Bronars and Grogger (1994) who also analyze teen fertility.

[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), Haveman, Wolfe, and Pence (2001), and Wolfe and Perozek (1996) and Wolfe and McHugh (2006). Rosenzweig and Wolpin (1995) also explore this issue.

related factors, such as public expenditures by states on family planning programs), her neighborhood's characteristics and her own prior choices.

II. A BRIEF LITERATURE REVIEW

Numerous early studies of the teen nonmarital birth outcome used cross-sectional and longitudinal data and reduced form estimation to relate the background, family and residence characteristics of young women to the non-marital birth outcome. None of them attempt to measure the response of teen unmarried women to the potential impacts of alternative nonmarital fertility-related choices that they make (e.g., choices regarding abstinence, contraception, or abortion), or study these choices in a dynamic framework, or as they interact with labor supply, schooling, and post-birth marital choices.[5]

Recent research has taken a more economic approach to understanding the determinants of unmarried teen women's fertility decisions, and has sought to relate these choices to the conditional opportunities and constraints with which these women are confronted. For example, because access to welfare benefits is available to single mothers, the generosity of public income support in the state that women live has been viewed as influencing the choice to give birth out of wedlock, which is one of the avenues into single motherhood.

Duncan and Hoffman (1990) were among the first to explore this approach. They estimate a two-stage logit model over African-American teenagers in which a teenage nonmarital birth event (joint with the receipt of welfare benefits) depends upon the woman's perceptions of the family income associated with alternative choices. Maximum state AFDC benefits are taken to be the income

[5]Haveman et al (2004) review these studies of the determinants of teenage pregnancy providing a qualitative meta-analysis of the empirical evidence. Mother's education, family moves, poverty, parent's marital status are factors that are consistently found to be important determinants of teen birth outcomes. In related work, Kalil and Kunz (1999) focus on cumulative risk factors that go beyond parental education, poverty and marital status to address whether the sheer number of risk factors facing an adolescent can lead to better understanding of the determinants of nonmarital childbearing. Upchurch, Lillard and Panis (2002) use a life cycle model that highlights joint decision making. Their focus, however, is not on adolescents and finds a significant effect of earlier childbearing on subsequent childbearing.

available in case of a birth, and predicted taxable family earnings at age 26 proxy for the income available without a birth. While both of these expected economic opportunity variables have the expected sign, only the variable indexing economic opportunities without a birth is statistically significant.

Lundberg and Plotnick (1995) have also studied the economic determinants of the teen nonmaritial birth outcome, recognizing that women's decisions reflect a sequence of premarital outcomes: premarital pregnancy, pregnancy resolution, and the occurrence of marriage prior to birth. They jointly estimate these stages of the decision process using a three-stage nested logit model. While they focus on state welfare benefits, recognizing that higher benefits reduce the costs of a nonmaritial birth, differences in the availability of family planning services and state abortion policies are also hypothesized to affect choices in this sequence. For whites, they find that the level of welfare benefits is positively and significantly related to the birth outcome, and that other state policy indicators also significantly influence relevant fertility-related choices. However, for blacks they find that none of the economic or policy variables are "significant in a manner consistent with an economic model" (p. 190), a result that agrees with that of Duncan and Hoffman.[6]

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, and statistically insignificant (but of larger magnitude) for African-

[6]The authors attribute these unexpected results to the small sample of black women, to the potential for under-reporting of several of events in some of the stages of the sequence, or to the existence of different racial responses to incentives.

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.[7]

Rosenzweig (1999) employs a model of the initial childbearing and marriage decisions of young women that incorporates concern for child quality and assortative mating. A primary objective 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, positive, but statistically significant overall effect and a large effect on women from low income families.[8]

Our approach extends these recent efforts to model the choices and behaviors of unmarried young women that may lead to a teen nonmarital birth event. Like these studies, we also view the choices of young unmarried women to be influenced by their expectations of the effects of alternative decisions among the options that are open to them. However, whereas prior research has focused on effects of alternative choices that derive from financial (welfare, personal, or family income) gains or

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[7]Moffitt (1998) criticizes the specification of the Clarke-Strauss model arguing that their instrument, state per capita income, probably belongs in their core or main equation.

[8]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 (2000) reexamine the effects of AFDC benefits on nonmarital childbearing through age 22. They use 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. While they are able to reproduce Rosenzweig's main finding, they fail to find a 'welfare effect' on teen nonmarital births, but a large effect on the choices of women in their early 20s. The finding of significant welfare effects when both cohort and state fixed effects are controlled for is at odds with other research relying on fixed effects estimation (Moffitt, 1994; Hoynes, 1997).

losses, we also analyze the response of young women to the implications of having a nonmaritial birth on the likelihood that they will ultimately establish a long-term and stable family-type relationship.

While we emphasize the role of the choice-conditioned expectations regarding marriage/cohabitation relationships and income, we also include more extensive information on family characteristics and choices, neighborhood attributes and the policy environment in which the young women lives than used in prior studies. 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. Unlike Lundberg and Plotnick (1995), we do not jointly estimate the response to incentives of several sequential choices that result in a nonmarital birth; all of these prior choices are reflected in the nonmarital childbearing outcome which we study.

III. OVERVIEW OF ESTIMATION APPROACH

The specification of our empirical model assumes that young unmarried women have knowledge of the longer-term marital/cohabitation relationship and economic (income) implications of nonmarital childbearing when making choices related to the probability of this outcome. In forming estimates of these relationship and income consequences, these young women assess the experiences of girls ten years their senior who are similar to themselves in terms of observed characteristics, some of whom had an out of wedlock birth as a teenager and some of whom did not.

We express this intuition through a two-stage econometric model. In the first stage, we relate the longer-term income and marriage/cohabitation relationship experiences of the older cohort of girls to their demographic characteristics and to the occurrence of a teen nonmarital birth. In the second stage, based on the coefficients from these first stage regressions, we predict the marriage/cohabitation relationship and income consequences for the younger girls conditional on having or not having a teen birth. We then estimate the probability of an out of wedlock birth for the younger cohort of teenage girls as a function of the predicted income and marriage/cohabitation consequences and other covariates. As indicated above, we hypothesize that perceptions of adverse income and marriage/cohabitation relationship consequences attributable to a teen nonmarital birth are related to a reduced probability of this event.

Our estimation strategy assumes that adolescent unmarried girls perceive the income and marriage/cohabitation outcomes of those individuals in the older cohort with observed characteristics similar to their own who did and did not experience a nonmarital birth. That is, we assume that these teens base their expectations on the experience of the older women taking account of an extensive set of observed characteristics, but assuming random assignment of members of the reference group to the with and without a teen nonmarital birth categories aside from these characteristics.[9]

A second identification assumption, common in two-stage models such as ours, is that the parameters in our model are identified by a broad array of exclusion and functional form restrictions that we describe below.

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).[10] The first data

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[9]An alternative view is that assignment of the reference sample to the two nonmarital birth outcomes is the result of a selection process that reflects all relevant determinants of this choice, and that individuals in the primary sample know this selection process and account for it in forming their economic and marriage/cohabitation expectations conditional on the nonmarital birth choice. In this view, young single women are presumed to reliably discern the effects of determinants of the nonmarital childbearing choices that are unobserved by the researcher, but which may have influenced the fertility choice of individuals in the secondary (older) sample. We judge that assigning this level of insight to the teenage girls whose outcomes we study seems unwarrantedly strong; hence our preferred results are based on economic and marriage/cohabitation expectations that are not adjusted for such potential selectivity. However, in related work we have compared the results from a model in which expectations are directly estimated and an alternative model in which the estimated income variables are selectivity-adjusted. We first estimate models that assume that youths do and do not perceive the effects of unobserved factors in forming expectations. The former model statistically controls for this selection process in estimates of expected personal income conditional on the choice that is made, using a two-stage Heckman-type selectivity correction model (See Heckman, 1979). See Haveman, Wolfe, and Wilson (2001). The estimated effects of expected income in selectivity-adjusted specification are very similar to those of the preferred specification that does not reflect the assumption that youths perceive the effects of unobserved factors in forming expectations.

[10]The PSID data provides longitudinal information on over 5000 families beginning in 1968. We use available data covering 25 years of information. The choice of ages (0–9 in 1970) reflects a balance among several objectives and constraints, including the need for a sample size sufficient to secure a required number of premarital births, the need to constrain the projection of neighborhood data beyond those from the 1970 and

set—our primary sample composed of young women whose choices we model—includes 1,172 teen age women who were ages 0–9 in 1970; they were followed until 1992, at which time they are young adults, ranging in age from 22 to 31 years.[11] The weighted proportion of observations in this sample with a nonmarital birth by age 18[12] is 8.6 percent; among the Black members of the sample, the weighted proportion is a far greater 29.6 percent.[13] A secondary sample is a somewhat older cohort of females who were aged 9–16 years in 1970, and who were 31 to 38 years old in 1992.[14]

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, 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 1998 dollars using the Consumer Price Index for all items.[15]

We merged onto both data sets an extensive array of year-specific state or county data designed to characterize the policy environment or community attitudes within which the individual

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1980 Census records that were available at the time we assembled our data set, and the decision to attach neighborhood data to each child's record beginning at age 6 (requiring a maximum age of 9 years in 1970).

[11]Only those females who remained in the survey until age 19 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 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.

[12]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.

[13]14.4 percent of the unweighted observations, 169 out of 1172, had a nonmarital birth at some point during their teen years.

[14]For the secondary sample, we use women who remained in the survey until age 29. An older sample of 727 women is used for estimating the expected 'relationship stability' variables; the sample for estimated expected income is 733. Missing data account for the difference in sample size.

[15]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.

makes choices or forms income or marriage/cohabitation relationship expectations. These variables include:

- state maximum welfare benefits per month,[16]
- state unemployment rates,
- state median family income,
- per capita state public expenditures on family planning,[17]
- state requires parental consent for abortions,
- state Medicaid program funds abortions,
- state restricted abortions pre Roe v. Wade,
- state teen birth rate,
- percent of state residents belonging to a religious organization,
- state divorce rate,
- state has a no fault divorce law,
- percent of state births out-of-wedlock.

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:

- unemployment rate,

[16]For each state, we have annual data from 1968 to 1992 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.

[17]1984 values are an average of 1983 and 1985 values for each observation; 1986 values are an average of 1985 and 1987 values.

- percent of workers in high status (professional/managerial) occupations,
- median family income,
- percent of families with low income (less than $10,000 in 1970 dollars)
- percent of families that are female headed,
- male/female ratio,
- the employed male/female ratio,
- percent of persons belonging to a religious organization.

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.[18]

A number of these policy/community environment variables were added to the data set as they are likely to influence choices related to the nonmarital birth outcome. For example, state welfare generosity and the prevalence of teen or nonmarital births in the state or of female-headed families in the neighborhood are likely to be positively associated with this outcome. On the other hand, women living in states with restrictive state abortion laws or a strong religious orientation may be less likely to make choices leading to a nonmarital birth.

Other of these policy/community variables (e.g., the state/neighborhood unemployment rates, welfare generosity, state/neighborhood median income, the prevalence of poverty, the share of workers employed in high status occupations, and relative female employment) are measures of the overall economic status of the community or of labor market connections that are likely to influence the personal income expectations of these young women. Similarly, the prevalence of divorce or ease of obtaining a divorce, the preponderance of female-headed families and the religious orientation of

[18]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).

state residents (as indicators of attitudes toward marital instability), and the (employed) male/female ratios (as measures of potential partners) will influence expectations regarding future marriage/cohabitation relationships. Consistent with these expected patterns of effect, we selectively include these as control variables in the nonmarital choice, income and marriage/cohabitation relationship models (see below). Appendix Tables 1 and 2 present a complete set of statistics for both samples.

V. LONGER-TERM INCOME AND MARITAL/COHABITATION EXPECTATIONS WITH ALTERNATIVE CHILDBEARING CHOICES

We assume that the unmarried teen women in the primary sample form their expectations regarding the effects of alternative childbearing choices by observing the choice-conditioned outcomes and experiences of older cohorts; hence, we derive the choice-specific income and relationship expectation variables required for our model from information on the experiences of an older cohort of young women. Whereas all prior studies of the teen nonmarital birth outcome based on a choice-conditioned expectations approach assume that the choices of young women respond to choice-specific expectations of a single variable (based on some income definition; see Haveman, Wolfe, and Wilson, 2001), we employ two utility-based components—expected income and an expected longer-term marriage/cohabitation relationship variable.

A. Choice-Conditioned Marriage/Cohabitational Expectations

We posit that choices regarding nonmarital childbearing affect a woman's well-being by influencing her future lifestyle and living arrangements, in addition to her future income trajectory. Perception of these choice-dependent effects will, in turn, influence the childbearing decisions that she makes. While numerous aspects of a teenage woman's future life-course are likely to be affected by a nonmarital birth event, we focus on future marriage-type relationships (defined to include both cohabitation and marriage), and account for the probability of their occurrence, stability and durability.

The variables that we construct to reflect expectations regarding longer term marriage/cohabitation relationships are unique to our approach. There is little research on the value placed by young people on having a longer-term marital or cohabitational partner. More importantly, there is no evidence on the effect of conditional expectations regarding marital/cohabitational relationships on fertility choices. The standard microeconomic approach to this issue views marriage (or cohabitational living arrangements) as an institution which allows two individuals to merge their individual attributes in the joint production of household goods so as to maximize expected joint lifetime utility or well-being (see for example Becker, 1974; Becker, Landes and Michael, 1977; McElroy, 1985). Complementary to this analytical approach is a more empirical and demographic literature relating to cohabitation, and the probability of marriage of women who give birth out of wedlock.[19]

We construct alternative measures of such longer-term relationships that reflect both the utility gains from a marriage/cohabitation relationship, and the utility costs associated with the dissolution of such a relationship. We estimate separate conditional expected values of marriage/cohabition relationship variables for each women in our primary sample relying on both economic and demographic perspectives. These estimates are based on several years of longitudinal data from the older cohort of women, some of whom did and others of whom did not have a teen nonmarital birth.

The indicator of longer-term marriage/cohabitation relationships that we employ is defined as follows:

$$R = (10 + Y_m + 0.6\ Y_c) / [(1 + N_d + .6\ N_t)^2],$$

where:

Y_m = number of years married

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[19]For example, Bumpass and Sweet (1989) report that while most cohabiting relationships are relatively short lived, approximately 60 percent end in marriage. Bumpass, Sweet and Cherlin (1991) and Bumpass and Lu (1998) also address this issue, and indicate that while the percentage of first cohabitations that result in marriage to the same person is declining, the proportion of first unions begun by cohabitation has been steadily increasing.

Y_c = number of years cohabiting

N_d = number of divorces

N_t = number of cohabitations terminated

We measure this indicator for each woman in our secondary sample over the 11 year period covering their ages 19–29, using annual information on marriage, cohabitation, and year over year changes in these statuses. Each woman gets a base value of 10; each year of marriage increases this value by one unit and each year of cohabitation increases the value by 0.6. This "weight" is based on the long term probability a cohabitating relationship evolves into a marriage. A divorce reduces the value of the indicator more than does the termination of a cohabiting relationship, and multiple disruptions reduce the value at an increasing rate. The indicator is increasing with the years in a relationship: a woman who is single over the 11 years has an indicator value of 10; a woman who is married for all of the 11 years and has no divorce has the maximum stability value of 21; a woman who cohabitates over the 11 years without terminating the relationship has a value of 16.6. Conversely, the indicator is decreasing in the number of divorces or terminated cohabitations: a woman who is married for one year, and the marriage ends in a divorce has a value of 9 which is less than the value if she had remained single throughout the period; a woman who is married for 5 years and has one divorce has a stability value of 11; with two divorces the value falls to 6. While the parameters of this indicator are arbitrary, it captures aspects of the creation, existence, and termination of relationships that are consistent with the sketchy literature on this issue. We test the robustness of our results to a number of alternative parameterizations of this indicator.

For both women in the secondary sample who gave birth while an unmarried teen, and women who did not have a teen nonmarital birth,[20] we regress this marriage/cohabitation relationship

[20]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 728 women in the older sample, 132 gave birth as an unmarried teenager, and 596 did not.

indicator on a set of variables that are expected to covary with relationship.[21] Then, using the estimated parameters from this pair of regressions, together with the relevant characteristics of the girls in our primary sample, we predict two indicators of expected marriage/cohabitation relationships for each primary observation—one representing her expected relationship index conditional on giving birth, and the other representing her expected relationship indicator if she does not give birth.

The (weighted) mean values of these predicted indicators (and the standard deviation for each mean value) are shown in Table 2 for each of the childbearing outcomes. These values are shown for the entire primary sample, separately for those who did and did not give birth in that sample and by race. The childbearing-conditioned expected relationship patterns indicate that the value of the expected relationship variable decreases substantially because of the teen nonmarital birth experience for all women and especially for nonblack women. For those women who did not give birth (85.6 percent of the total) the index is 13.94 (standard deviation = 1.91); the indicator is 7.20 (4.65) for those that did give birth. For black women who gave birth, the value of the index is 10.18 (3.66), with an expected increase of 2.04 had they not given birth. The value of the index is 11.42 (2.21) for those who did not give birth, with a difference of 2.23 had they given birth. The average differences are far greater for nonblack women: for those who gave birth the mean difference is 6.74 and for those that did not, the mean difference is 7.48.

[21]We included in these equations variables likely to covary with relationship stability, including race, if first born, parental education, family structure, mother's employment, urban residence, region, family location changes, disability status of family head, family income, family welfare recipiency, being Catholic, whether mother was divorced, number of times mother was married, whether the state has no fault divorce laws, divorce rates in the state and ratios of males to females. Most of these variables are measured over the girl's ages 12–15. This range is determined by the 25 years of observations that are available. The vectors of independent variables used in the relationship regressions are virtually the same between the two teen childbearing groups (father a college graduate was omitted in the with teen birth regression due to insufficient observations). Most of the variables have similar and expected effects on relationship stability for the two groups. The definitions, means, and standard deviations of these variables are shown in Appendix Table 1. The estimated relationships are available from the authors.

B. <u>Choice-Conditioned Personal Income Expectations</u>

The second variable hypothesized to affect the choice of young unmarried women regarding childbearing reflects the income expectations associated with and without such a nonmarital birth. In particular, we estimate separate conditional personal income variables for each woman in our analysis sample.[22] The conditional (expected present value of discounted) personal income values are also estimated from several years of longitudinal data on the slightly older cohort of women (whose experiences also underlie the relationship expectations).

We use personal income rather than family income (or family income relative to needs) since these latter variables incorporate issues of family composition and allocation, which are outside of our model and, for the most part, our observation. These include a young woman living at home, with friends or at school. Both measures require a quite different set of implicit assumptions than use of the personal income variable.[23] For example, a family income relative to needs variable implicitly assumes that all of the benefits of a marriage/cohabitation relationship 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.

One could argue that our marriage/cohabitation variable is simply a proxy for the income that is derived from having a spouse or mate living in the home, and that this additional income is the primary or only source of utility associated with a marital or cohabitational relationship. In this case,

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[22]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 personal income rather than individual earnings since neither transfer income (including welfare benefits) nor child support payments are contained in earnings, hence omitting an important component of the relevant expected economic well-being concept specified in our model.

[23]Because additional children increase the level of family needs, the family income relative to needs variable assumes 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.

use of a family (as opposed to personal) income variable would cause our marriage/cohabitation variables to be irrelevant in the estimation of the determinants of fertility behavior. Although we view marriage/cohabitational relationships to have utility apart from the additional income derived from the spouse/mate, we test for whether or not the use of family income rather than personal income changes our results; these results are reported in sensitivity tests described below.

The two choice-specific expected personal income variables for each of the 1,172 young women in the primary sample 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 observations with no income, especially at younger ages.[24] Estimates are conducted for each of the 11 years from ages 19 to 29[25] 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 birth),[26] for a total of 22 estimated equations.[27] The results of these estimations are available from the authors.

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[24]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 18 percent at age 19. Among the women who did not give birth, an average of 15 percent report no personal income, with a maximum of nearly 29 percent at age 19.

[25]Given our desire to use neighborhood and childhood data, we are constrained from using information for ages greater than 29 years.

[26]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 733 women in the older sample over whom the income equations are estimated, 132 gave birth as an unmarried teenager, and 601 did not.

[27]We included in these equations variables likely to be related to the personal income dependent variable, including race, family position (if first born), 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 percent of neighborhood residents in high status occupations, percent neighborhood residents with low income, neighborhood unemployment rate 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 25 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, with two exceptions: (1) parental education is limited to a high school or more in the with-birth estimation rather than the high school graduate and more than high school used in the no-birth estimations, and (2) state AFDC generosity is included only in the with-birth equations. 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. Father's education is positively associated with income only in the estimates for those without a teen birth. Beyond these two examples, the variables have similar and expected effects on personal income for the two groups. The

We use the relevant individual characteristics of each female adolescent 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. 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 not give birth.[28]

The (weighted) 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.[29] The childbearing-conditioned expected income patterns are 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 bottom two panels of Table 1, expected incomes for unmarried teens who did not have a birth are 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

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definitions, means, and standard deviations of these variables are shown in Appendix Table 1. The estimated relationships are available from the authors.

[28]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.

[29]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.

income penalty for the women who did have a birth. These patterns suggest that the young women in our sample are choosing rationally.

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. By including incomes during the late-teens and 20s, we capture the income foregone due to postponed working (or delayed marriage) associated with early childbearing; in this sense, our expected income terms may be superior to estimates of full lifetime incomes.[30]

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 $123,126; the average expected present value assuming a birth is $75,798, for a difference of $47,328. As noted earlier, the gain from not giving birth as a teen is far greater for those girls who did not, in fact, give birth ($48,977) than for those girls that did ($29,771).

VI. ESTIMATION OF THE TEEN CHILDBEARING CHOICE MODEL

Empirical estimation of the determinants of the teen nonmarital childbearing decision focuses on the effects of the two expectations variables, one for expected marriage/cohabitation patterns and the other for expected personal income. For each individual, the choice-conditioned differences ("if no teen birth" minus "if a teen birth") in marriage/cohabitation and personal expectations are taken to reflect expected net opportunity gains associated with deciding to **not** bear a child out of wedlock. We

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[30]We also experimented with assigning age 29 income for an additional ten years and using the earnings growth over the 10 years with observed income to create income for subsequent years. The basic results are invariant to these alternative measures of conditional income.

estimate this model using a switching model, probit specification.[31] The dependent variable is equal to 1 if the young unmarried woman gives birth before age 19 years, and 0 otherwise. Unweighted data are used for estimation; 169 (14.4 percent) of the young women in our primary sample gave birth while an unmarried teen.[32]

When the expected income difference and expected relationship variables are taken to be the only factors influencing the teen birth decision, both coefficients are negative and statistically significant.

	coefficient	t-statistic
Income difference	-0.345	-3.07
Marriage/cohabitation relationship difference	-0.045	-4.47

The marginal probability on the income difference variable is -0.075, while that on the relationship difference is -0.01. To indicate the quantitative magnitude of this response, we simulate the change in the probability of a teen nonmarital birth from a one standard deviation increase in the marriage/cohabitation and income difference variables. On average, a one standard deviation increase in the income difference variable is associated with a 0.331 standard deviation decline in the probability of the teen birth outcome; a one standard deviation increase in the marriage/cohabitation variable is associated with a one-fifth (0.044) standard deviation decline in the probability of the birth outcome.[33] While these effects suggest that young women respond in expected ways to both increased

[31]See Manski (1987) and Lee (1982). 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 and high foregone income associated with giving birth 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.

[32]The women in our sample are teens with risk of a nonmarital teen birth during years 1982–1990. During this period, 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. Our statistic is not rate per year but a cumulative rate over the teen years. African-Americans are oversampled in our data; we conclude that our higher rates are consistent with the observed rates.

[33]These estimates standardize the dependent variable but not the regressor (Long and Freese, 2006). Alternatively, a twenty five percent increase in the marriage/cohabitation difference is associated with a

economic opportunities and gains in marriage/cohabitation outcomes in making childbearing choices, they control for neither family and neighborhood characteristics nor for the direct influence of the policy environment on this choice.

Table 3 presents our preferred estimate of the determinants of the teen childbearing choice. This specification includes a number of family, neighborhood and policy environment variables, in addition to the expected income and expected relationship variables.[34] The log-likelihood test statistic indicates that the entire equation is significant at the .01 level.

The coefficient on the expected difference in the marriage/cohabitation variable is negative and marginally significant (t-statistic = 1.80), indicating that concerns over the future structure of partner and family relationships are likely to play some role in teen unmarried girls' choices regarding sexual activities and behavior. The marginal effect on the overall probability of a teen nonmarital birth associated with the coefficient is -0.01. The coefficient on the variable reflecting the expected net income gain from foregoing childbearing is negative and statistically significant at about the 5 percent level, 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.[35] Converting the income coefficient into a marginal probability, the estimate is -0.053.[36]

The coefficient estimates of our preferred model reveal little regarding the quantitative impact of changes in these variables on the fertility choice. We convert the coefficient estimates to more

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decrease of nearly 12 percent in the probability of a teen nonmarital birth; a twenty five percent increase in the log income difference is associated with a decrease of about 7 percent in the teen nonmarital birth outcome.

[34]The means and standard deviations of these variables are shown in Appendix Table 2.

[35]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.

[36]As noted above, since we define our income variable to be personal income, which excludes income from a partner, this variable may also be reflecting the added income of a spouse if married or a mate if cohabiting. As noted above, we have also estimated the effects of using a family income variable in sensitivity tests reported below.

intuitive measures; since the relationship index is an ordinal measure, we present selected estimates from the preferred model in terms of standard deviation changes.[37]

The effect of increasing the marriage/cohabitation variable by one standard deviation results in a 0.025 standard deviation decline in the probability of a teen nonmarital birth. Alternatively, when the expected income difference is assumed to increase by one standard deviation, teen nonmarital births are estimated to decrease by one-quarter of a standard deviation (0.250). We also convert the estimates of expected income and relationship differences for the sample of black women.[38] An increase of one standard deviation in the expected income difference variable lowers the probability of a nonmarital birth by 0.111 of a standard deviation, less than half the impact estimated for the full sample. An increase in the relationship variable yields a reduction of around 0.029 of a standard deviation, only slightly larger than the impact on the full sample.

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 (see Haveman and Wolfe, 1995; Haveman et al, 2004). Those whose mother has little education, those who are African-American, and those growing up in a single parent family (t-statistic = 1.33) are more likely to have a teen nonmarital birth than are girls without these characteristics; girls whose religion is Catholic are less likely to have a teen birth out of wedlock.

The results for the direct effects of the policy variables are also consistent with previous research. The prevalence of religious organization membership in the community is negatively and significantly associated with the probability of a nonmarital teen birth. Similarly, 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

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[37]For this conversion methodology, see Long and Freese (2006).

[38]As described in Section VII, below, we have also estimated our preferred model over the sample of black women only.

childbearing. Our model suggests that if state family planning expenditures were increased by a full standard deviation, the rate of nonmarital childbearing would decline by almost one-half a standard deviation (0.497); for black women only, the probability of a nonmarital teen birth would fall by more than two-thirds standard deviation (0.693).[39]

A. Note on Model Identification

We identify this model through exclusion restrictions in the first stage income and relationship models, through the nonlinear specification of the income difference terms, and through timing. A standard approach to identifying the first stage (income and relationship) models is to include in these specifications at least one variable expected to affect expectations but not the birth choice (other than through the income and relationship terms). In our personal income estimation, multiple variables that 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; neighborhood variables include median family income, percent with low income, the unemployment rate, and proportion in high status occupations. In the first stage relationship estimation, whether the girl's mother was ever divorced, the number of times the mother was married, whether the state has a no fault divorce law, the divorce rate in the state and the two male/female ratios are included in the estimation of the value of the relationship indicator, but not in the final stage teen birth equation.

Identification is also achieved through the nonlinear functional forms utilized in the estimation. 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. In addition, the period over which family variables are measured differs between the analyses of the

[39] Alternatively, the model suggests that if state family planning expenditures were increased by 25 percent, the rate of nonmarital childbearing would decrease from 0.078 to about 0.0627, a reduction of over 19 percent; for black women, the childbearing rate would fall from 0.270 to 0.214, or over 20 percent.

secondary sample of older women and the primary sample, and this difference also contributes to model identification.[40]

In order for the exclusion restrictions to provide valid identification, the variables must be correlated with income (relationship indicator) 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, proportion with low income, and the proportion of persons in the neighborhood in high status occupations are all labor market measures that are closely related to earnings prospects and income. A similar rationale stands behind our choice of variables in the relationship equations. The mother's own divorce experience and number of marriages capture the effect that the mother's marriage/cohabitation relationship has on the daughter's relationship including attitudes toward divorce. State divorce laws and divorce rates influence the probability of divorce while male to female ratios represent opportunities for current and future relationships. There is little theoretical or observational reason to expect these variables to be related to the teen nonmarital birth outcome.[41]

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[40]The family variables used in the income and stability 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, and future relationship stability.

[41]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 age 29 for girls who had a birth, the R-squared indicator of correlation is .225 when only the instruments are included in the specification. The R-squared statistic is .293 for the regression with the full of set of explanatory variables. Likewise, R-square is .107 when only the instruments are included in the income without a birth equation, compared with .109 with the full set of variables. In the relationship stability with a teen birth equation, the R-squared statistic is 0.04 when we regress the stability index on the instruments alone. In contrast, the R-square statistic is 0.23 when all the explanatory variables are included; the instruments account for about 17 percent of the full R-square indicator. Our instruments have even more explanatory power in the relationship stability without a teen birth equation. In that equation, the R-square statistic is 0.07 when we use only the instruments as explanatory variables, compared with 0.11 with the full set of variables. In this case, the instruments account for about 60 percent of the full R-square indicator.

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.[42] Using the residuals from the marriage/cohabitation relationship model, the test-statistic has a value of 10.20, well below the critical value of 18.31 (chi-squared with ten degrees of freedom) for statistical significance, and indicates that the instruments in the relationship model are uncorrelated with the residuals from this final stage estimate. Similarly the residuals from our predicted income term give a test-statistic of 12.89 (chi-squared with eighteen degrees of freedom), again well below the critical value of 28.87 for statistical significance.

VII. ROBUSTNESS TESTS

While our preferred model reflects our best judgment 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 Table 3 estimates by systematically altering the specification of the model in a variety of dimensions. The results of these robustness tests are summarized in Table 4, where we present the coefficient and t-statistic on both the income and relationship variables. For each test, the specification of the family, policy and neighborhood variables are the same as in Table 3.The first row shows these statistics for the preferred model, presented in Table 3. The t-statistic on the log of the difference in the expected personal income variables indicates that this variable is statistically significant at about the .05 level; that for the relationship stability variable indicates significance at the .07 percent level.

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

In our first set of robustness tests, we modify our index of marriage/cohabitation relationships. In alternative (1) we modify this relationship measure to assign a weight of 5 to being single. In this case, the income difference is statistically significant at the .05 level, the stability difference is negative and significant at the .10 level. The coefficients are similar to those in the preferred specification. In alternative (2) we assign single a weight of 2. In this case, the log income difference term is statistically significant at the .05 level, but the stability difference term is not quite significant at the .10 level. In alternative (3), we assign a weight of zero to being single; the results are similar although the statistical significance of the relationship variable is again reduced. The values of the coefficients on these variables are similar across these 3 alternatives.

In alternative (4) we replace our difference in expected marriage/cohabitation relationships measure with the ratio of the expected measure without a birth to that with a birth. In this estimate, the income difference term is robust and statistically significant at nearly the .05 level, but the relationship ratio variable is not at all statistically significant, and instead has an unexpected positive coefficient.

In alternative (5), we allow the effect of the relationship variable to vary depending on the childbearing outcome by entering the two predicted relationship variables separately into the childbearing equation. The coefficient on expected relationship if no birth is negative but not at all statistically significant, while the coefficient on the expected relationship variable if the girl has a teen birth is positive and marginally statistically significant (.07 level of significance). The pattern implies that the expected relationship if a teen birth occurs is more important in the decision making of young unmarried women than the expected relationship with no teen birth; a plausible result.

Finally we alter the relationship measure by replacing the index defined above with the first principal factor of a factor analysis. This factor combines a variety of dimensions of marriage/cohabitational relationships without consideration of relationship demise. The results of the estimated childbearing equation using this factor is shown as alternative (6). In this case, the relationship variable has the expected negative sign but it is not at all significant. This suggests that it is not only the probability of having a longer-term relationship that is important for teenagers but also avoiding divorce or separation. The coefficients on the other variables remain largely unchanged.

In the next set of tests, we focus on the income term. Although the log specification is a common one, it lacks a strong theoretical justification. To ensure that our nonlinear functional form assumption is not driving our results, robustness test (7) takes the absolute difference rather than the difference in the logs. The estimate of the coefficient on the income variable is statistically significant at the 5 percent level (t-statistic=2.41); the significance of the stability variable is little changed. 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. As alternative (8) of Table 4 indicates, the results are quite consistent with our preferred model and the income variable is statistically significant at the 10 percent level as in our preferred model (t-statistic=1.61). 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 base results, the relationship of income to the teen childbearing outcome is independent of whether the individual gave birth as a teen. In alternative (9), we allow the effect of the income term to vary depending on childbearing outcome by entering the two expected income variables separately into the final equation. The results indicate that the coefficient on expected income without a teen birth is negatively and significantly related to the probability of a teen birth; the level of expected income if the woman has a child is positively related to the probability of a teen birth, but it is not at all significant. This pattern implies that expected income opportunities conditional on not having a birth are more important in the decision of a teen unmarried woman than expected income opportunities with a teen birth. This is consistent with the pattern estimated by Duncan and Hoffman (1990).[43]

To test the sensitivity of the results to the race of the teen woman, we have estimated the model over black women only; the results are shown in alternative (10). We note that the rate of teen nonmarital childbearing is far higher among blacks than nonblacks (except for Hispanics).[44] The results on the expected relationship variable are nearly identical to those of the model estimated over the entire population: the expected relationship difference measure is marginally statistically significant at the .10 level of significance and the coefficient is negative and similar (though somewhat smaller) in magnitude. Alternatively the difference in expected incomes is not at all statistically significant for this group of women. Lundberg and Plotnick also fail to find an income incentive effect for Blacks [also see Moffitt (1998)].

Finally, in alternative (11) we estimate our preferred model using family income rather than personal income. Earlier, we argued that personal income is preferred to family income since the latter encompasses issues of family composition and allocation, which we do not observe in our data. However, it is possible that young women only care about the economic resources available to their living unit, in which case the use of total family income in place of personal income might generate different results. Indeed, as we indicated above, if the marriage/cohabitation relationship measure is primarily reflecting the income provided by a spouse/mate to the living unit, use of the family income variable would result in a statistically insignificant coefficient on the relationship measure. The estimation results shown in alternative (11) indicate that this is not the case; the relationship variable is statistically significant at the 5 percent level and is slightly larger than in the preferred model (-0.043). The income difference variable is also larger than in the preferred specification (-0.514) and is marginally statistically significant at the 10 percent level.

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[43]An alternative estimate in which both expected income and relationship stability are entered in separately shows a similar pattern.

[44]Other studies of fertility behavior have found differences by race; see Moffitt (1998) for examples of these models and estimates based on them.

VIII. CONCLUSION

These estimation and simulation results suggest that choice-specific measures of expected marriage/cohabitation patterns and expected income 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 worthwhile interventions for securing reductions in teen nonmarital childbearing. Recent experience following the 1996 welfare reform is consistent with this finding. Somewhat ironically, but not unexpectedly, the results on the expected relationship variable suggest that decreasing the value of the expected relationship variable among young adults who gave birth as an nonmarried teenager would have a greater impact on reducing the probability of teen births than would increasing the expected relationship variable among those young women who did not give birth as a teen. The results also indicate that increased family planning efforts may be an effective instrument for reducing the prevalence of this problem.

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TABLE 1
Predicted Personal Incomes (in 1998 $)

	With Teen Birth		No Teen Birth	
	Mean	Std. Dev.	Mean	Std. Dev.
Whole Sample (N = 1172)				
Age 19	$6,818	$3,556	$7,719	$2,339
Age 20	3,945	2,591	8,527	2,515
Age 21	6,092	3,484	10,297	2,650
Age 22	7,409	2,737	12,097	3,466
Age 23	9,961	5,516	13,925	3,672
Age 24	8,729	3,941	15,102	4,332
Age 25	12,098	8,227	16,250	5,052
Age 26	8,841	3,789	18,227	5,761
Age 27	11,761	3,279	17,939	4,153
Age 28	11,357	3,812	19,693	4,429
Age 29	10,707	5,721	19,736	4,389
Net Present Value	$75,798	$22,395	$123,126	$28,862
Those With Teen Birth (N = 169)				
Age 19	$6,247	$3,430	$6,245	$2,168
Age 20	5,221	2,702	7,095	2,026
Age 21	6,259	3,331	8,614	2,187
Age 22	7,165	2,533	9,785	3,123
Age 23	9,020	4,580	11,402	3,259
Age 24	8,373	3,545	12,217	3,629
Age 25	10,033	5,958	13,035	3,870
Age 26	7,982	3,724	14,381	4,057
Age 27	11,399	3,862	15,013	3,830
Age 28	10,839	3,881	16,911	4,409
Age 29	9,591	5,583	16,906	4,379
Net Present Value	$71,698	$21,895	$101,469	$25,494
Those Without Teen Birth (N = 1003)				
Age 19	$6,872	$3,565	$7,857	$2,307
Age 20	3,825	2,549	8,662	2,516
Age 21	6,077	3,499	10,455	2,636
Age 22	7,431	2,756	12,314	3,418
Age 23	10,049	5,589	14,162	3,620
Age 24	8,762	3,976	15,372	4,295
Age 25	12,292	8,385	16,552	5,046
Age 26	8,922	3,786	18,589	5,767
Age 27	11,795	3,218	18,214	4,077
Age 28	11,406	3,803	19,954	4,342
Age 29	10,812	5,724	20,002	4,297
Net Present Value	$76,183	$22,411	$125,160	$28,329

TABLE 2
Expected Relationship Stability Indicators
(N = 1172)

	If No Birth		If Birth		
	Mean	Standard Deviation	Mean	Standard Deviation	No Birth - Birth Mean Difference
All	13.94	1.91	7.20	4.65	6.74
Those Gave Birth					
Black	12.22	1.97	10.18	3.66	2.04
Nonblack	14.09	1.66	7.35	5.12	6.74
Those with No Teen Birth					
Black	11.42	2.21	9.19	3.98	2.23
Nonblack	14.30	1.60	6.82	4.63	7.48

TABLE 3
Teen Nonmarital Childbearing Model
(N = 1172)
probit maximum likelihood

	Coefficient	Std. Err.
Expected Income LN[Predicted Income if No Birth] - LN[Predicted Income of Birth]	-0.297*	0.154
Expected Relationship Stability [Predicted Relationship Stability if No Birth] - [Predicted Relationship Stability if Birth]	-0.029*	0.016
Individual and Family Background Characteristics		
Race (African American = 1)	0.524**	0.167
Proportion of Years Lived with Single Parent, Ages 6–15	0.199	0.150
Mother High School Graduate = 1	-0.386**	0.133
Mother Attend College = 1	-0.846**	0.245
Missing Mother Education = 1	-0.558**	0.216
Mother Gave Birth as a Teen (yes = 1)	0.100	0.138
Mother Teen Birth Information Missing	-0.138	0.191
Proportion of Years in Poverty, Ages 6–15	0.132	0.193
Belong to Catholic Church = 1	-0.460**	0.202
State Choices		
Whether State Restricted Abortion pre Roe vs. Wade	0.229	0.313
Average Public Family Planning Expenditures per capita, Ages 13–18	-0.590**	0.209
Whether State Medicaid Funds Abortions, ages 13–18	-0.168	0.170
Whether State Required Parental Consent for Abortion, ages 13–18	-0.183	0.257
Average State Teen Birth Rate, Ages 13–18	0.006	0.007
Average State Percentage of Births Out-of-Wedlock, Ages 13–18	-0.003	0.009

(table continues)

TABLE 3, continued

	Coefficient	Std. Err.
Neighborhood Attributes		
Percent of Families Headed by a Female, 6–15	0.002	0.005
Percent of Individuals belong to Religious Organization	-1.769**	0.739
Constant	-0.119	0.478
LR chi2(19) =	158.970	
Prob > chi2 =	0.000	
Log likelihood =	-403.979	
Pseudo R2 =	0.164	

*t-statistic is between 1.80 and 1.96, and is interpreted as being marginally significant.
**t-statistic is 1.96 or more, and is interpreted as being significant.

TABLE 4
TESTS OF ROBUSTNESS

	Coeff	Std. Err.
Preferred Model from Table 3 (where weight on single status = 10)		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.297	0.154
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.029	0.016
TESTS ON ROBUSTNESS OF RELATIONSHIP STABILITY MEASURE		
(**1) Single Weight = 5**		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.305	0.154
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.037	0.022
(2) Single Weight = 2		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.311	0.154
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.041	0.027
(3) Single Weight = 0		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.314	0.155
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.042	0.032
(4) Using a Ratio of Stability-Without-Birth/Stability-With-Birth		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.296	0.154
[Relationship Stability if No Teen Birth]÷[Relationship Stability if Teen Birth]	0.003	0.006

(table continues)

TABLE 4, continued

	Coeff	Std. Err.
(5) Entering Relationship Stability With- and Without- a Birth Separately		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.292	0.154
E[Relationship Stability] if No Teen Birth	-0.017	0.033
E[Relationship Stability] if Teen Birth	0.032	0.018
(6) Using First Principal Component as Substitute Relationship Measure[a]		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.285	0.155
[Relationship Factor if No Teen Birth]-[Relationship Factor if Teen Birth]	-0.050	0.125
TESTS ON ROBUSTNESS OF EXPECTED INCOME		
(7) Expected Income		
[Predicted Income if No Teen Birth]-[Predicted Income if Teen Birth]	-4.590	0.190
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.028	0.016
(8) Ratio of LN Income if No Birth to LN Income if Teen Birth		
LN[Predicted Income if No Teen Birth]÷LN[Predicted Income if Teen Birth]	-0.151	0.019
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.030	0.016
(9) Enter Income With- and Without- a Birth Separately		
LN[Predicted Income if No Teen Birth]	-0.607	0.253
LN[Predicted Income if Teen Birth]	-0.094	0.202
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.023	0.017

(table continues)

TABLE 4, continued

	Coeff	Std. Err.
FURTHER TESTS OF ROBUSTNESS		
(10) Estimated Over Blacks Only		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.120	0.141
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.031	0.019
(11) Family Income instead of Personal Income[b]		
LN[Predicted Income if No Teen Birth]-LN[Predicted Income if Teen Birth]	-0.514	0.305
[Relationship Stability if No Teen Birth]-[Relationship Stability if Teen Birth]	-0.043	0.018

[a]The factor combines five measures of relationships including whether in a single intact relationship, relationship experiences from ages 24–34, and number of relationships.
[b]The family income variable is estimated for the 12–15 year old sample. Certain policy covariates are not included in these regressions (see footnotes 27 and 37).

APPENDIX TABLE 1
Variables Used in Estimation of Relationship Stability and Income Prediction Equations (weighted; N = 962)

Variable	Mean	Std. Dev.	Min	Max
Individual and Family Background Characteristics				
Gave birth as an unmarried teen	0.096	0.295	0	1
Father High School Graduate	0.282	0.450	0	1
Missing Father Education = 1	0.135	0.342	0	1
Mother Ever Divorced = 1	0.206	0.405	0	1
Number of Marriages of Mother	1.090	0.669	0	1
Mother Marital History Missing = 1	0.120	0.325	0	6
Belong to Catholic Church = 1	0.322	0.468	0	1
Belong to Religion other than Catholic = 1	0.619	0.486	0	1
Mother High School Graduate = 1	0.374	0.484	0	1
Mother Attend College = 1	0.207	0.406	0	1
Missing Mother Education = 1	0.035	0.183	0	1
Race (African American = 1)	0.181	0.385	0	1
Proportion of Years Mother Worked, Ages 12–15	0.517	0.417	0	1
Proportion of Years Lived with One Parent, Ages 12–15	0.158	0.339	0	1
Proportion of Years on AFDC, Ages 15–15	0.063	0.200	0	1
Firstborn = 1	0.220	0.415	0	1
Proportion of Years with a Locational Move, Ages 12–15	0.103	0.191	0	1
Proportion of Years Lived in SMSA, Ages 12–15	0.719	0.431	0	1
Proportion of Years Family Head is Disabled, Ages 12–15	0.171	0.335	0	1
Proportion of Years in Northeast, Ages 12–15	0.282	0.448	0	1
Proportion of Years in West, Ages 12–15	0.166	0.369	0	1
Proportion of Years in South, Ages 12–15	0.261	0.437	0	1

(table continues)

APPENDIX TABLE 1, continued

Variable	Mean	Std. Dev.	Min	Max
State Choices				
Average State Maximum Welfare Benefits per Month	302.47	97.52	60.07	455.79
Average Unemployment Rate	5.33	1.49	2.15	9.53
Average State Median Family Income	15,222	1,959	9,475	18,301
State has No Fault Divorce Law, Ages 19–30	0.588	0.438	0	1
Divorce Rate in State, Ages 19–30	4.753	1.134	3.046	7.650
Percent of Individuals belong to Religious Organization in State	0.207	0.106	0.0310	0.420
Neighborhood Attributes				
Average Neighborhood Unemployment Rate	4.992	2.878	0.225	24.625
Average Neighborhood Median Family Income	16,896.260	6,306.366	5,456.308	66,360.49
Proportion of Neighborhood Employed in High Status Occupations[a]	23.498	11.106	1.875	60.0
Proportion of Neighborhood with Low Income	0.192	0.138	0.013	0.772
Ratio Males to Females, Ages 19–30	98.41	12.94	57.61	216.92
Ratio Employed Males to Females, Ages 19–30	65.21	11.65	20.77	119.50
Percent of Families Headed by a Female, Ages 6–15	11.48	6.86	3.63	56.47

[a]Professional or Managerial.

APPENDIX TABLE 2
Variables Used in Teen Birth Model Estimates
(weighted; N = 1172)

Variable	Mean	Std. Dev.	Min	Max
Gave birth as an unmarried teen	0.086	0.280	0	1
Expected Income				
LN[Predicted Income if No Birth]-LN[Predicted Income of Birth]	0.501	0.385		
Expected Relationship Stability				
[Predicted Relationship Stability if No Birth]	13.944	1.909	6.095	20.438
[Predicted Relationship Stability if Birth]	7.197	4.651	-5.643	24.121
Individual and Family Background Characteristics				
Race (African American = 1)	0.132	0.338	0	1
Proportion of Years Lived with Single Parent, Ages 6–15	0.152	0.299	0	1
Mother High School Graduate = 1	0.456	0.498	0	1
Mother Attend College = 1	0.212	0.409	0	1
Missing Mother Education = 1	0.024	0.154	0	1
Whether State Restricted Abortion pre Roe vs. Wade	0.080	0.266	0	1
Mother Gave Birth as a Teen (yes = 1)	0.042	0.200	0	1
Mother Teen Birth Information Missing	0.054	0.226	0	1
Proportion of Years in Poverty, Ages 6–15	0.103	0.226	0	1
Father High School Graduate = 1	0.330	0.470	0	1
Father Attend College = 1	0.315	0.465	0	1
Missing Father Education = 1	0.085	0.280	0	1
Belong to Catholic Church = 1	0.267	0.443	0	1

(table continues)

APPENDIX TABLE 2, continued

Variable	Mean	Std. Dev.	Min	Max
State Choices				
Average Public Family Planning Expenditures per capita, ages, 13–18	1.175	0.365	0.356	3.102
Whether State Medicaid Funds Abortions, ages 13–18	0.534	0.421	0	1
Whether State Required Parental Consent for Abortion, ages 13–18	0.087	0.248	0	1
Average State Teen Birth Rate, Ages 13–18	50.670	12.923	29	83
Average State Percentage of Births Out-of-Wedlock, Ages 13–18	18.396	4.437	6.2	56.5
Neighborhood Attributes				
Percent of Families Headed by a Female, Ages 6–15	12.700	7.450	1.160	72.350
Percent of Individuals belong to Religious Organization in State	0.220	0.107	0.031	0.420