FAMILY STRUCTURE, EARLY SEXUAL BEHAVIOR, AND PREMARITAL BIRTHS*

Lawrence L. Wu
University of Wisconsin-Madison

Andrew J. Cherlin
Johns Hopkins University

Larry L. Bumpass
University of Wisconsin-Madison

January 1997

*Direct all correspondence to Lawrence L. Wu, Department of Sociology, 1180 Observatory Drive, University of Wisconsin, Madison, WI 53706, lwu@ssc.wisc.edu.
We gratefully acknowledge research funding to Cherlin from a NICHD Merit Award (HD 25936) and to Wu from NICHD (HD 29550), from the William T. Grant Faculty Scholars Program, and from funds provided by the Assistant Secretary for Program Evaluation to the Institute for Research on Poverty. Additional support was provided by a NICHD core grant to the Center for Demography and Ecology (HD 05876). We thank Kathleen Harris and James Walker for helpful comments and discussions, and Steven Martin for expert research assistance. An earlier version of this paper was presented at the Annual Meetings of the American Sociological Association, August 1996, New York, NY.

ABSTRACT

In this paper, we argue that entry into first sexual intercourse is a key process mediating the effects of family structure on premarital childbearing. We explicate three ways in which onset of sexual activity can mediate effects of family structure on premarital first births. First, the gross association between family structure and premarital birth risks may be due entirely to the effect of family structure on age at first intercourse. Second, the earlier the age at first intercourse, the longer the duration of exposure to the risk of a premarital first birth. Third, an early age at first intercourse may proxy unmeasured individual characteristics correlated with age at onset but uncorrelated with other variables in the model. We develop methods to assess such mediating effects and analyze data from two sources, the 1979–93 National Longitudinal Survey of Youth and the 1988 National Survey of Family Growth. We find that age at first intercourse partially mediates the effect on premarital birth risks of both snapshot measures of family structure at age 14 and a time-varying measure of the number of family transitions, but that significant effects of these variables remain net of age at first intercourse. Delaying age at intercourse by one year reduces the cumulative relative risk of a premarital first birth by a similar amount for both white and black women. For black women, the magnitude of this effect is roughly the same as that of residing in a mother-only family at age 14.

The association between young adult outcomes and variation in childhood and adolescent family structure is no longer in question. Many studies demonstrate an association between a measure of a person's family structure while growing up and outcomes such as high school graduation (McLanahan and Sandefur 1994), age at leaving home (Goldscheider and Goldscheider 1993), and childbearing outside of marriage (McLanahan and Bumpass 1988). More recently, researchers have begun to move beyond simply documenting that this association exists by investigating what it is about family structure that may affect affect the lives of children and adolescents (see, e.g., Astone and McLanahan 1991, 1994; Cherlin et al. 1991; Cherlin, Kiernan, and Chase-Lansdale 1995; Wojtkiewicz 1993; Wu and Martinson 1993; Wu 1996).

Yet we still know little about the presumed causal pathways through which family structure affects young adult outcomes. Knowing which characteristics of family structure are most closely associated with having a child prior to marriage—the outcome we examine—is an advance, but knowing about such characteristics says little about what might *mediate* the effects of family structure. In particular, we focus on mediating *events*—on process rather than structure—to clarify how family dynamics and processes experienced during childhood and adolescence might affect the subsequent life chances and trajectories of individuals.

In this paper, we argue that entry into first sexual intercourse is a key process mediating the effects of family structure on premarital childbearing. Prior research suggests several ways in which earlier sexual activity may be associated with living in nonintact family. Some authors stress the lower level of supervision and monitoring that single parents, on average, provide (Dornbusch et al. 1985); the differing norms and values their adolescent children develop (Thornton and Camburn 1987); and the cumulative stress on adolescents who experience multiple family transitions (McLanahan 1985; Wu and Martinson 1993; Wu 1996). Recent work by developmental psychologists suggests other mechanisms for such mediating effects, although data limitations preclude adequate tests of these effects in our analyses. A first argument concerns modeling: an adolescent whose mother is sexually active with partners other than the adolescent's biological father (or, equivalently but less commonly, an adolescent who lives with a father who is sexually active with partners other than the adolescent's mother) may view sexual activity outside of marriage as more acceptable. A second argument involves negativity: relationships between adolescents and their unmarried, divorced, or remarried parents may involve higher levels of noncompliant, acting-out behavior on the part of adolescents; or a lack of warmth and negative sanctions (yelling, denial of privileges) on the part of the parents.

Age at first intercourse could mediate effects of family structure on premarital childbearing in several ways. One possibility is that the gross association between family structure and premarital birth risks is mediated entirely by the effect of family structure on age at first intercourse. A second is that the longer the duration spent at risk of giving birth prior to marriage, the greater the likelihood of a premarital birth. Finally, an earlier age at first intercourse might affect premarital birth risks by proxying unmeasured characteristics of individuals that are correlated with age at onset. We contrast these alternative effects of age at sexual onset on premarital birth risks by using multistate hazard models and by assessing alternative effects of age at sexual onset on the cumulative relative risk of a premarital birth.

We analyze data from two surveys, the 1979–93 National Longitudinal Surveys of Youth (NLSY) and the 1988 National Survey of Family Growth (NSFG), to decompose the overall association between various measures of family structure on premarital birth risks into their component effects on entry into sexual activity and the subsequent risk of a premarital birth given entry into sexual activity. Our analyses thus extend previous research by directly modeling several ways in which age at onset of sexual intercourse might mediate effects of family structure and by comparing estimates of such mediating effects obtained from two surveys.

THEORY

In this section, we briefly review standard arguments linking family structure to early adult outcomes. We then discuss two additional developmental arguments stressing effects of modeling and negativity, which provide alternative mechanisms by which sexual activity might mediate effects of family structure on nonmarital childbearing.

Three standard arguments provide behavioral mechanisms that suggest how family structure might influence outcomes during early adulthood:

- *Socialization during childhood.* Within both developmental psychology and sociology, socialization theory suggests that the lessons children learn from their parents have lasting effects on their personalities. Early studies of the topic that followed Freud assumed that the absence of a father might inhibit healthy gender identity. Although few scholars make the Freudian argument today, it is sometimes claimed that growing up with one parent may influence children's norms and values in ways that make sexual activity and childbearing outside of marriage more acceptable (e.g., Thornton and Camburn 1987).
- *Parental supervision during adolescence.* Several studies present the view that, regardless of the efficacy of early socialization, lone parents have more difficulty supervising the behavior of adolescents than do two parents (e.g., Dornbusch et al. 1985). There is evidence that greater parental supervision is associated with lower sexual activity among adolescents (e.g., Hogan and Kitagawa 1985).
- *Stress and family instability.* A third possibility is that the greater the number of family transitions (parental divorce, cohabitation, or remarriage) an adolescent has experienced, the greater the cumulative stress, and the more likely the adolescent is to engage in sexual activity. Adolescents may disengage from a home continually in transition and seek emotional gratification elsewhere through impulsive and rebellious behavior such as sexual activity (Hetherington 1987).

In two recent papers, Wu and colleagues (Wu and Martinson 1993; Wu 1996) examined these arguments by measuring different aspects of family structure. Using data from the NLSY and National Survey of Families and Households, they found that socialization and parental supervision—as measured by prolonged exposure to a single mother family and by the types of parental figures present during adolescence—had no significant effects on premarital first birth risks, but that family instability—as measured by the number of family transitions experienced by an adolescent—had consistent and significant effects on premarital first birth risks.

Continuing in this vein, Wu and Thomson (1995) used similar family measures to analyze age at first intercourse for NLSY women. Previous research (see, e.g., Newcomer and Udry 1987) suggested that not living with two parents was associated with an earlier age at intercourse. As in their analyses of premarital childbearing, Wu and Thomson found no significant effects on age at first intercourse of prolonged exposure to a single-mother family, but consistently significant effects of the number of family transitions. However, they also found that women who lived as adolescents in family situations other with their two biological parents tended to have earlier ages at first intercourse. Although this latter finding might reflect less supervision in single-parent households, they argued that this explanation was insufficient because age at intercourse did not differ significantly for women in single-mother and mother-stepfather families, where two potentially supervisory adults are present.

We suggest two other explanations for the effects of family structure on age at first intercourse. The first concerns modeling. In a recent study, Capaldi, Crosby, and Stoolmiller (1996) used data from an intensive longitudinal study of about 200 boys from higher crime neighborhoods in an Oregon city to analyze age at first intercourse. Consistent with findings by Wu and colleagues, Capaldi, Crosby, and Stoolmiller found that the greater the number of family transitions, the earlier the age at intercourse. However, their data also contained measures for several more proximate causes (antisocial behavior; parental supervision and monitoring) for the association between instability and earlier intercourse. Yet the number of family transitions continued to have significant effects on age at intercourse even after controlling for these variables. Reflecting on their results, these researchers suggested that:

> ... the direct effect [of transitions] may be due to modeling of parental dating behavior. When a new partner moves in to the household and the relationship is still in its early phase, a child—and especially an adolescent—is likely to be highly aware of the sexual nature of the relationship. Heightened

> awareness ... may make the adolescent feel that early sexual involvement is more acceptable and, therefore, accelerate sexual involvement. (p. 355)

Similarly, a 16-year old girl was quoted in an Australian study of sexuality during adolescence as saying:

> If your parents are divorced or separated, and your mum or dad brings home different people on weekends and each night of the week and stuff, you sort of think that [having sex] is no big deal... But if your parents are married and stuff like that, you sort of see it as a big deal and should only share it if you love the person. (Moore and Rosenthal 1993, p. 65)

Because both parental dating and parental cohabitation involve extramarital sexual activity, either may accelerate the onset of sexual activity by offspring.

Developmentalists argue that family structure may also influence age at first intercourse through negativity. For example, a 26-month study of 200 families conducted by Hetherington and Clingempeel (1992) examined two-biological parent, divorced, and remarried families, all of whom had 9- to 13-year-old children at the beginning of study. On average, Hetherington and Clingempeel found that relationships in the divorced and remarried families exhibited greater conflict and less warmth than in the non-divorced families. Adolescents in divorced and remarried families tended to be verbally aggressive and less compliant; correspondingly, their parents tended toward harsher responses that involved nagging, physical punishment, and denial of privileges, coupled with less overt affection, joking, and joint activities.

These behaviors are tied to the developmental tasks of adolescence. Adolescents typically test limits and seek autonomy as they confront their emerging sexuality. Although many divorced or remarried families cope well with these issues, a parental divorce or remarriage may make adolescence a more troubled period than is the case for intact families. For example, because single parents are more prone to anger and depression, in part because of financial strains due to lower income following divorce (Weiss 1975), such families may be more prone to cycles in which negative behaviors by adolescents provoke harsh or inconsistent responses by custodial parents, which may in turn reinforce adolescent behaviors (Hetherington and Clingempeel 1992).

Remarriage does not necessarily improve the parent-child relationship (Cherlin and Furstenberg 1994). Adolescents confronting their emerging sexuality may be uncomfortable with the addition of a stepparent engaged in sexual relations with the custodial parent. Hetherington and Clingempeel (1992) reported that conflict between adolescents and stepparents seemed especially acute among young adolescent girls whose custodial mothers remarried. They also report that stepparents found their initial efforts at parenting resisted by stepchildren; stepparents often disengaged from parenting as a result (see also Amato 1987; Cherlin 1978; Furstenberg 1987).

Issues of both modeling and negativity in adolescence are raised in a series of articles on the effects of parental divorce on children, based mainly on a large, longitudinal British survey of individuals born in 1958 and followed through age 33. In a first paper, Cherlin et al. (1991) found that effects of parental disruption and its aftermath at age 11 were substantially reduced after controlling for pre-disruption measures of behavior problems and academic achievement. Similar analyses of the children of respondents in the National Longitudinal Survey of Youth found that effects of parental disruption on pre-adolescent children were greatly reduced after an initial crisis period following the disruption (Morrison and Cherlin 1995). However, these researchers found significant effects at age 23 for the British respondents of parental divorce on demographic outcomes such as leaving home, cohabitation, and childbearing outside of marriage (Cherlin, Kiernan, and Chase-Lansdale 1995) and on a scale of mental health (Chase-Lansdale, Cherlin, and Kiernan 1995), even after controlling for pre-disruption behavior problems and academic achievement.

Overall, these findings suggest that the effects of a parental disruption and its aftermath were modest among pre-adolescents after the immediate crisis following a parental disruption, but reemerged during young adulthood in terms of earlier homeleaving, poorer mental health, and more premarital childbearing. In order to explain this pattern of findings, Cherlin, Kiernan, and Chase-Lansdale (1995) speculated that the effects became manifest only during adolescence and young adulthood because they involved sexual issues that only became salient after puberty:

> Early sexual activity may be a key to the effects of divorce on the transition to adulthood for two reasons. First, and most directly, the obvious sexual activity of divorced parents may stimulate nonmarital sexual activity in their children... Second, adolescents' early sexual activity could constitute rebellious or acting-out behavior begun in reaction to an unwelcome parental divorce or to the introduction of an unwelcome new stepparent. (p. 313)

Thus, early sexual activity could result not only from early childhood socialization, lack of supervision, or family instability, but also from modeling or negativity.

DATA

We analyze data from two household-based national probability samples: the 1979–87 National Longitudinal Survey of Youth (NLSY), a prospective survey of young adults aged 14–21 in 1979, and the 1988 National Survey of Family Growth (NSFG), a retrospective survey of young women aged 15–44 in 1988. Both surveys provide fertility and marriage histories, as well as data on age (to the nearest month) at first sexual intercourse and snapshot measures of family structure when the respondent was age 14. To facilitate comparisons between surveys, we restrict attention throughout to those birth cohorts in NSFG that match the birth cohorts sampled in the NLSY.

Outcomes. We examine three related outcomes for women in the NLSY and NSFG: (1) age at first intercourse; (2) age at a premarital first birth, ignoring when a woman becomes sexually active; and (3) age of a premarital first birth conditioned on age at first intercourse. Data on the calendar year and month of first sexual intercourse were asked of NSFG respondents in 1988 and of NLSY respondents in the 1985 and 1986 interviews. We converted these data into age (in months) at first intercourse and used a hot-deck procedure to impute calendar month at onset when calendar month of onset was missing. We censored data on first intercourse in two ways: (1)

at age at the relevant interview if the respondent reported never having experienced sexual intercourse, and (2) at age of first marriage if the respondent reported initiating sexual activity on or after the date of marriage. NLSY respondents were, on average, somewhat younger than respondents in the NSFG subsample when retrospecting about age at first intercourse (ages 19 to 27 in 1985–86 versus ages 23 to 30 in 1988 for NLSY and NSFG women, respectively). Despite these differences, Table 1 shows that the distribution of age at first intercourse are similar in both surveys.

[Table 1 about here.]

We constructed data on premarital first births using data on first births and first marriages in the NSFG and NLSY. First birth histories were constructed from retrospective fertility data in the NSFG and by combining retrospective and prospective fertility data from the 1979 and 1980–93 waves of the NLSY. First marriage histories were constructed similarly. We then censored data on premarital first births either at a woman's age at first marriage or her age at last valid interview. Finally, we dropped cases: (1) who reported a first sexual intercourse prior to age 10, (2) with missing data on first intercourse, or (3) who reported a birth prior to first intercourse.[1]

Measures of family structure. Both the NLSY and NSFG contain snapshot measures of the respondent's family structure at age 14. We used these data to construct a standard set of measures to contrast women in two-biological parent families, mother-only families, step families, and a residual category of other types of families. In addition, the NLSY contains a retrospective parental history administered in 1987. The parental history provides data from birth to age 18+ on whether the respondent lived with biological parent(s), stepparent(s), adoptive parent(s), or in

[1]We handled missing data differently for a limited number of cases in the NLSY. Because of the prospective nature of the NLSY fertility and marriage histories, data for some respondents can be nonmissing at one interview, but missing at the next valid interview. When this occurred, we censored the histories for the respondent at the last interview for which we have nonmissing data.

some other situation (e.g., with grandparents or other relatives). We merged these data with data on when the respondent first left home, which was constructed from an item in the parental history and with data from the annual household rosters to determine the family structures in which respondents lived before first leaving home.

We used the NLSY parent history data to construct several additional measures of family structure. A first set of variables consists of three measures capturing the exposure of the respondent to a mother-only family: whether the respondent was born out-of-wedlock, the proportion of life spent in a mother-only family during early childhood (ages 0–5), and a time-varying measure reflecting the proportion of life spent in a mother-only family between birth and age t. These variables correspond to hypotheses that predict behavioral differences for children and adolescents who experience prolonged exposure to a mother-only family (Guidubaldi et al. 1986; Hetherington 1972, 1981; Kellam, Ensminger, and Turner 1977; McLanahan 1988; Mueller and Pope 1977; Thornton 1991; Thornton and Camburn 1987).

A second set of variables consists of several measures reflecting family structure during adolescence, following hypotheses concerning the control that parents can exert over the behaviors of their adolescent children (Dornbusch et al. 1985; Hogan and Kitagawa 1985; Matsueda and Heimer 1987; McLanahan and Bumpass 1988). These variables are identical to the standard snapshot measures of family structure except that they are time-varying and are constrained to have effects only during the period of adolescence (ages 10–18 or 120 to 227 months).

A third set of variables consists of two time-varying dummy variables equal to one at age t if the respondent has experienced a parental divorce or (re)marriage by age t. These variables provide rough measures of modeling effects by adolescents of the nonmarital sexual activity of parents (Axinn and Thornton 1996; Thornton 1991; Thornton and Camburn 1987). Because many parents are sexually active after divorce and many parents who gave birth out-of-wedlock subsequently marry and were sexually active prior to marriage, one would expect a parental divorce or (re)marriage to be positively associated with offspring sexual activity and premarital childbearing.

A final measure of family structure is a time-varying variable for the cumulative number of family transitions experienced by the respondent by age t. This variable is motivated by hypotheses related to the stresses accompanying changes in family structure (Furstenberg and Cherlin 1991; Hetherington 1987; Hetherington, Camara, and Featherman 1983; McLanahan 1985; Rutter 1983; Wu and Martinson 1993) and has been shown to be consistently associated with both age at first intercourse and premarital first births (Wu and Martinson 1993; Wu and Thomson 1995; Wu 1996).

Controls for background characteristics. In a first set of analyses, we contrast results from the NLSY and NSFG controlling for a limited set of background characteristics common to both; in these analyses, we also restrict the ages of NSFG women to match the birth cohorts sampled in the NLSY. A second set of analyses restrict attention to the NLSY to exploit the more extensive set of covariates, particularly with respect to data on family structure obtained from the parent histories available in the NLSY (Wu 1996). Background variables common to both data are race and ethnicity, mother's education, mother's age at first birth, a dummy variable for Catholic religion, a time-varying covariate equal to 1 at all ages greater than or equal to the respondent's age at first menstruation and a time-varying covariate equal to 1 if the respondent had become sexually active within the previous 12 months.[2] We include dummy variables for missing mother's education and whether we employed a hot-deck procedure to impute the calendar month at first intercourse. In further analyses of data from the NLSY, we include controls for the following additional set of background characteristics: SEI of respondent's father or of the adult male in the

[2] We use the time-varying covariate indicating if sexual activity had begun during the past 12 months as a rough control for ineffective contraception during early periods of sexual activity. The NLSY variable for religion refers to the religion in which respondents were raised, while the NSFG item refers to respondents' religious affiliation at the time of interview.

respondent's household at age 14; number of siblings; educational expectations; the respondent's score on the Armed Forces Qualifying Test (AFQT); a composite index summing dummy variables for the presence of magazines, newspapers, and library cards; and a time-varying dummy variable equal to one if the respondent had left the parental household. In this second set of analyses, we also include dummy variables for missing AFQT and if the respondent's father did not work, was not present in the household, or if father's SEI was otherwise missing.

METHODS

As noted above, age at first intercourse could mediate effects of family structure on premarital childbearing in several ways. One possibility is that the gross association between family structure and premarital birth risks may be mediated entirely by the effect of family structure on age at first intercourse. Two other possibilities involve effects of age at first intercourse on premarital first births. For example, *ceteris paribus*, the earlier age at intercourse, the greater the duration spent sexually active and hence the greater the likelihood of a premarital birth (Bumpass and McLanahan 1989). We term this the *exposure effect* of age at first intercourse. In modeling this effect, we assume that the decisions regarding sexual activity are separate from decisions regarding first marriage. This assumption would have been questionable at mid-century, when the sexual activity most young women and many young men was restricted to the person they would marry (May 1988). In those circumstances, earlier onset of intercourse might have brought about an earlier marriage, leaving unchanged the length of exposure to the risk of a premarital birth. However, over the past few decades, the proportion of adolescents who are sexually active prior to marriage has increased greatly (Hayes 1987; Forrest and Singh 1990), as has age at first marriage (Cherlin 1992). In addition, unmarried pregnant women have become much less likely to marry before giving birth (U.S. Bureau of the Census 1991). Together, these trends have loosened the connection between sexual activity and marriage and increased the average durations of exposure to the risk of premarital childbearing.

It is possible that age at first intercourse could have an effect on premarital births that is due not only to exposure but also to other correlates. Young women and men who initiate sex earlier may live in different social contexts. For instance, they may internalize different norms concerning the acceptability of childbearing outside of marriage; or they may have less opportunity to marry. Earlier intercourse could also be a marker for rebellious or acting-out behavior that could also be associated with premarital childbearing. As distinct from an exposure effect, we term this the *correlate effect* of age at first intercourse. This latter "effect" is difficult to interpret causally because it is likely to reflect unmeasured characteristics of the young woman, her family, or her social environment that are correlated with earlier intercourse.

Our statistical models differentiate between these three possibilities. Let 0 denote the state for unmarried childless women prior to first intercourse, 1 the state for unmarried childless women who have initiated first intercourse, and 2 the state for women who have experienced a premarital first birth. We analyze three related outcomes: (1) $r_{02}(t)$, the age-specific rate of a premarital first birth ignoring whether or not a woman has experienced sexual intercourse; (2) $r_{01}(t)$, the age-specific rate of first sexual intercourse; and (3) $r_{12}(t|t_1)$, the age-specific rate of a premarital first birth conditioned on t_1, the age at which a woman has experienced first sexual intercourse. For the moment, we set aside conditioning on t_1. Let $r_{jk}(t)$ denote the hazard rate from state j to state k; then the specification used in all analyses is given by:

$$r_{jk}(t|\mathbf{x}_{ijk}(t)) = q_{jk}(t)\exp(b_1 x_{1i}(t) + b_2 x_{2i}(t) + \cdots)\,, \tag{1}$$

where $\mathbf{x}_i(t)$ denotes a vector of covariates (possibly time-varying) for case i, and

$$\log q_{jk}(t) = \begin{cases} \alpha_{jk1} + \delta_{jk1}t, & \text{ages less than 16.5 years } (< 198 \text{ months}), \\ \alpha_{jk2} + \delta_{jk2}t, & \text{ages 16.5–18.5 years (198–222 months)}, \\ \alpha_{jk3} + \delta_{jk3}t, & \text{ages 18.5–21.0 years (223–252 months)}, \\ \alpha_{jk4} + \delta_{jk4}t, & \text{ages greater than 21.5 years } (> 252 \text{ months}), \end{cases} \tag{2}$$

with $\log q_{jk}(t)$ subject to three spline constraints:

$$\begin{aligned} \alpha_{jk1} + 198 \times \delta_{jk1} &= \alpha_{jk2} + 198 \times \delta_{jk2}\,, \\ \alpha_{jk2} + 222 \times \delta_{jk2} &= \alpha_{jk3} + 222 \times \delta_{jk3}\,, \\ \alpha_{jk3} + 252 \times \delta_{jk3} &= \alpha_{jk4} + 252 \times \delta_{jk4}\,. \end{aligned} \tag{3}$$

Throughout our analyses, we assume women become at risk of sexual activity and a premarital first birth at age 10 (120 months); hence, for $r_{02}(t)$, the age-specific rate of a premarital first birth ignoring whether or not a woman has experienced sexual intercourse, and $r_{01}(t)$, the age-specific rate of first sexual intercourse, woman i contributes to the log likelihood function as follows:

$$\ln L_{ijk} = (1 - c_{ijk}) \ln r_{ijk}(t) - \int_{120}^{t} r_{ijk}(s)ds\,, \tag{4}$$

where c_{ijk} is a censoring indicator equal to 1 if data for woman i are right-censored for the transition from state j to state k. For $r_{12}(t|t_{1i})$, the age-specific rate of a premarital first birth conditioned on t_{1i}, the age at which a woman has experienced first sexual intercourse, woman i contributes to the log likelihood function as follows:

$$\ln L_{i12} = (1 - c_{i12}) \ln r_{i12}(t) - \int_{t_{1i}}^{t} r_{i12}(s)ds\,. \tag{5}$$

Equation (5) differs from Equation (4) by specifying a so-called left-truncation time t_{1i} in the lower limit of integration, which lets individuals vary in their duration of exposure to age-specific premarital birth risks following the onset of sexual activity.

To clarify the behavioral assumptions in Equation (5), consider two identical women, one of whom initiates intercourse at age t_1 and the other of whom initiates intercourse at some later age $t_1 + \Delta$, where $\Delta > 0$. Equation (5) states that both women are subject to identical age-graded risks of a premarital first birth for all ages greater than $t_1 + \Delta$, but that the woman who initiates intercourse earlier is subject to an additional age-graded component generated by her exposure to the age-specific risk of a premarital first birth between ages t_1 and $t_1 + \Delta$.

One of our analytical goals is to compare the magnitude of estimated effects for the exposure and correlate effects of t_1, the age at first intercourse, on first premarital birth risks. We specify these effects as follows:

$$r_i(t|t_1i) = q(t)\exp(b_1 t_{1i} + b_2 x_{2i}(t) + \cdots), \tag{6}$$

where for clarity we have dropped the subscripts $jk = 12$ from various quantities in Equation (6). The exposure effect is captured in how the left-truncation time, t_1, affects the log likelihood in (5), while b_1 captures the correlate effect of age at first intercourse on the age-specific premarital birth risks, where $100 \times [\exp(b_j\Delta) - 1]$ has the usual interpretation of the percentage change in the relative risk corresponding to a shift of Δ in the effect of a covariate x_j.

Because a shift of Δ in a woman's age at first intercourse influences her likelihood of a premarital first birth through both correlate and exposure effects, assessing the magnitude of these effects requires a common metric by which to base comparisons. Let $H_{i12}(t_2|t_{1i})$ denote the cumulative risk of a first premarital birth evaluated at t_2:

$$H_{i12}(t_2|t_{1i}) = \int_{t_{1i}}^{t_2} r(s)ds\,, \quad t_{1i} < t_2\,. \tag{7}$$

As above, consider two women, i and j, who differ only in their age at first sexual intercourse, and let t_1 and $t_1 + \Delta$ denote the ages at first intercourse for i and j, respectively. Consider the *cumulative relative risk* of a premarital birth, $H_i(t_2|t_1 + \Delta)/H_j(t_2|t_1)$, and note that, by assumption, the quantities $\exp(b_2 x_2(t) + \cdots)$ and $q(t)$ are identical for women i and j (more precisely, the latter quantity is identical during the ages during which both women are at risk). Then:

$$\begin{aligned}
\frac{H_i(t_2|t_1+\Delta)}{H_j(t_2|t_1)} &= \frac{\int_{t_1+\Delta}^{t_2} q(s)\exp[b_1(t_1+\Delta)]\exp(b_2x_2(t)+\cdots)ds}{\int_{t_1}^{t_2}\exp(b_1t_1)\exp(b_2x_2(t)+\cdots)ds} \\
&= \frac{\exp[b_1(t_1+\Delta)]\int_{t_1+\Delta}^{t_2} q(s)ds}{\exp(b_1t_1)\int_{t_1}^{t_2} q(s)ds} \\
&= \exp(b_1\Delta)\left[\frac{\int_{t_1+\Delta}^{t_2} q(s)ds}{\int_{t_1}^{t_2} q(s)ds}\right].
\end{aligned} \tag{8}$$

Equation (8) illustrates that, under a proportional hazard specification, the correlate effect of a Δ change in t_1 on the cumulative relative risk is given by the usual quantity $\exp(b_1\Delta)$, while the "exposure" effect of a Δ change in t_1 on the cumulative relative risk, evaluated at t_2, is given by the bracketed ratio of integrals involving the baseline hazard function, $q(t)$. Inspection of (8) shows that the correlate effect is time-invariant but that the exposure effect varies with time and, in particular, depends on the age t_2 at which one evaluates the cumulative relative risk. Consequently, we report exposure effects for a range of t_2 in our analyses.

RESULTS

We begin by contrasting estimated effects of the snapshot measures of family structure on the transition to first intercourse and to a premarital first birth for white women in the NLSY and NSFG. These results control for background factors common to both surveys; for brevity, we do not report estimated effects of control variables.

[Table 2 about here.]

Analyses of women in the NLSY and NSFG. Models 1 and 2 of Table 2 report estimated effects of the snapshot measures of family structure on the age-specific rate of first sexual intercourse for NLSY and NSFG women, respectively. Consistent with findings in prior research, we observe higher risks for women in mother-only, step, and other types of families at age 14 relative to women in two-biological parent families at age 14. Overall, estimated effects of family structure are larger in magnitude in the NSFG than NLSY, but these differences are statistically significant only for women in the residual category for other types of families at age 14.

Models 3 and 4 give estimated effects of the snapshot measures of family structure on the age-specific rate of a premarital first birth when one ignores all information on whether a woman has begun sexual activity. Models 5 and 6 differ from Models 3 and 4 by specifying a "correlate" effect of age at intercourse (e.g., by adding age at first intercourse as a right-hand-side covariate). The effects of the snapshot measures of family structure decline in magnitude when controlling for for the correlate effect of age at intercourse (compare estimates in Models 3 and 5, and in Models 4 and 6). As before, estimated coefficients are similar in magnitude across surveys (compare estimates in Models 3 and 4, and in Models 5 and 6), although estimated standard errors are somewhat smaller in the NLSY than in the NSFG, due in part to the greater number of premarital first births in the NLSY.

Models 7 and 8 differ from Models 3–6 by controlling for differences in the duration of exposure produced by variation in age at first intercourse; Models 9 and 10 specify both exposure and correlate effects of age at first intercourse. The correlate effect of age at first intercourse in Models 9 and 10 is larger in the NSFG than in the NLSY; however, this difference is not significant. The other results in Models 7–10 are otherwise qualitatively similar to those in Models 3–6, with the effects of the snapshot measures of family structure declining in magnitude when controlling for for the correlate effect of age at intercourse (compare estimates in Models 7 and 9, and in Models 8 and 10), but similar in magnitude across surveys (compare estimates in Models 7 and 8, and in Models 9 and 10).

Table 3 gives results for black women in the NLSY and NSFG. As expected, women in mother-only, step, and the residual category for other types of families at age 14 have higher rates of entry into first intercourse activity and a premarital first birth than do black women in two-biological parent families at age 14; however, these effects are generally smaller in magnitude for blacks than for whites. As in Table 2, more estimated effects are significant for black women in the NLSY than for those in the NLSY, but as for whites, none of the differences in the magnitude of estimated effects between surveys is significant.

[Table 3 about here.]

Both Tables 2 and 3 suggest that the process of entry into first sexual intercourse does indeed mediate the effects of snapshot measures of family structure on premarital childbearing. For both white and black women in both the NLSY and NSFG, the effects of family structure at age 14 in Models 9 and 10, which control for both the exposure and correlate effects of first intercourse, are smaller than those in Models 3 and 4, which control for neither effect. In some instances, the decline in the magnitude of effects is substantial—for example, the effect of residing in a mother-only family at age 14. This pattern of results is consistent with theoretical expectations: because living in a nonintact family at age 14 hastens entry into sexual activity for both white and black women in the NLSY and NSFG, women in such families have longer durations of exposure to the risk of nonmarital childbearing. Thus, models that adjust for these longer durations of exposure suggest smaller effects of residing in a nonintact family at age 14 than do models that ignore such exposure effects.

The correlate effects of first intercourse in Tables 2 and 3 suggest that women with earlier ages at first intercourse have higher premarital birth risks than those with later ages at onset in ways that are not captured by family structure, duration of exposure, or the background variables. These results provide indirect evidence that women with earlier ages at onset differ systematically from those with later ages at onset in terms of characteristics that are not observed in these data.

Table 4 compares the magnitude of exposure and correlate effects for white women in the NLSY and NSFG through a series of simulations that examine these effects on the cumulative relative risk of a premarital first birth as generated by differences in age at onset of first intercourse. These comparisons require choosing a period of observation, $[t_1, t_2]$, during which to assess cumulative relative risks (see Equation 8 above). The baseline group in Table 4 consists of white women who resided in a mother-only family at age 14; thus, these simulations can be interpreted as suggesting how the cumulative relative risk of a premarital first birth would vary for women who resided in a mother-only family at age 14 if the onset of first intercourse occurred at an earlier or later age than the median age at onset for this group of women.[3]

[Table 4 about here.]

We begin by discussing results for white women in the NLSY. Lines 1–3 in Table 4 show the effect of lowering the age at first intercourse by 24 months. The results in line 1 show that, when evaluated at 60 months after the median age at onset, a 24 month decrease in age at onset generates a 38.5 percent increase in the cumulative relative risk due to longer durations of exposure and a 20.6 percent increase due to unmeasured differences in such women. These exposure and correlate effects combine for a total effect corresponding to a 67.0 percent increase ($1.385 \times 1.206 = 1.670$) in the cumulative risk. The exposure and total effects for women in the NLSY diminish when evaluated at later ages (lines 2 and 3). Lowering age at onset by 12 months relative to the median age at onset (lines 4–6) increases cumulative risks, but by a smaller amount, while increasing age at onset by 12 months (lines 7–9) or 24 months (lines 10–12) results in lower cumulative relative risks.

Simulations for white women in the NSFG suggest similar conclusions. The exposure effects are slightly smaller, and the correlate effects somewhat larger, in the NSFG than the NLSY; these patterns yield larger total effects in the NSFG than the NLSY. Nevertheless, there is broad agreement in estimated effects on cumulative risks for white women in both the NLSY and NSFG.

Table 5 gives a parallel set of results for black women in the NLSY and NSFG. As for whites, the magnitude of correlate effects is slightly larger in the NSFG relative to the NLSY, and these patterns produce total effects that are marginally larger in the NSFG than the NLSY. Nevertheless, estimated effects for black women show

[3]The median age at first intercourse for white women in a mother-only family at age 14 is 205 and 204 months (17.1 and 17.0 years) in the NLSY and NSFG, respectively. For white women in two-biological parent families, the corresponding ages are 218 and 221 months (18.2 and 18.4 years) in the NLSY and NSFG, respectively.

substantial agreement in both surveys.

[Table 5 about here.]

Overall, Tables 4 and 5 suggest that the exposure effect of delaying age at intercourse by one year reduces the cumulative relative risk of a premarital first birth by a similar amount for both white and black women. For black women, the magnitude of the exposure effect corresponding to a one-year delay ranges between a 15.0 and 24.9 percent reduction in the cumulative relative risk of a premarital first birth, an effect that would roughly offset the $12.7 = 1 - \exp(.12)$ to $20.9 = 1 - \exp(.19)$ percent increase in the cumulative relative risk of residing in a mother-only family at age 14 (see estimates in Table 3, Models 9 and 10). For white women, the corresponding exposure effect ranges between a 12.5 and 23.7 reduction in the cumulative relative risk, values that are substantially smaller than the $93.5 = 1 - \exp(.66)$ percent increase in risk of residing in a mother-only family at age 14.

Which effect is larger, the exposure effect generated by increased durations of exposure accompanying an earlier age at onset of first intercourse, or the correlate effect of earlier age at onset reflecting unmeasured characteristics of such women? Our analyses do not answer this question unambiguously. For both white and black women in the NLSY, the exposure effects are uniformly larger in magnitude than the corresponding correlate effects. For black women in the NSFG, the exposure effects are also uniformly larger than the corresponding correlate effects, but for white women in the NSFG, exposure effects tend to be smaller than the corresponding correlate effects.

Additional analyses of women in the NLSY. The analyses in Tables 2–5 restrict attention on snapshot measures of family structure when a woman was age 14 because of data limitations in the NSFG. But increases in marriage, divorce, and nonmarital childbearing over the past several decades have entailed increasingly complex family trajectories for a substantial fraction of children and adolescents (Martinson and Wu

1992). In particular, snapshot measures of family structure confound several family experiences commonly hypothesized to have distinct consequences for adolescent and young adult outcomes. For example, a woman residing in a mother-only family at age 14 (a) may have experienced the divorce of her two biological parents only recently, (b) may have been born out-of-wedlock and remained in such a family throughout childhood and early adolescence, and (c) may have experienced multiple family transitions—for example, a parental divorce, remarriage, redivorce, and so forth. As a consequence, snapshot measures of family structure shed little about the presumed causal mechanisms through which family structure influences adolescent and early adult outcomes.

Following work by Wu and colleagues (Wu and Martinson 1993; Wu and Thomson 1995; Wu 1996), we investigate alternative family measures to investigate possible linkages between family structure and premarital birth risks. These analyses restrict attention to the NLSY to exploit the retrospective parental histories available in these data. Table 6 presents estimated effects of a variety of measures of family structure on premarital birth risks for white women; models in Table 6 parallel those in Models 3, 7, and 9 in Table 2. For example, Model a.3 estimates the effect of the number of family transitions using a specification of premarital birth risks similar to that in Model 3 in Table 2 that ignores whether a woman has begun sexual activity. Model a.7 estimates the effect of the number of family transitions using the specification of Model 7 in Table 2 that specifies an exposure effect of age at first intercourse. Similarly, Model a.9 parallels Model 9 in Table 2 by specifying both exposure and correlate effects of age at first intercourse. The models in subsequent rows of Table 6 add effects of other family measures to Models a.3, a.7, and a.9.

[Table 6 about here.]

The first column of Table 6 gives the estimated effect of the number of family transitions on premarital birth risks. Comparing estimated effects in Models a.3, a.7, and a.9 shows that controlling for the exposure and correlate effects of age at first intercourse reduces the size of the effect of the number of family transitions on premarital birth risks; similar patterns emerge in the remaining rows of estimates in Table 6. Still, the effect of the number of family transitions is significant in all models except for Models f.7 and f.9.

The second column of Table 6 gives estimated effects for the correlate effect of age at first intercourse. As expected, the correlate effect corresponding to a later age at first intercourse reduces premarital birth risks, as expected, but the correlate effect is not significant in any of the models estimated. A comparison of Models a.7 and a.9 shows that controlling for the correlate effect of age at first intercourse reduces the magnitude of the effect for number of family transitions only slightly. The remaining models in Table 6 estimate the effects of the other family structure measures. Although effects are generally in the expected directions, none of these variables have significant effects on the risk of a premarital first birth net of the effect of the number of family transitions and of the exposure and correlate effects of age at first intercourse.

Table 7 presents results for black women in the NLSY. Results are roughly similar to those for whites. The estimated effect of the number of family transitions is significant in all models, and the magnitude of this effect decreases when controlling for the exposure and correlate effects of age at first intercourse. The correlate effect corresponding to a later age at first intercourse reduces premarital birth risks, as expected, but these effects are significant in all models, unlike the case for white women. None of the remaining measures of family structure have significant effects on premarital birth risks, with the exception of the time-varying measure for women residing in the residual category of other types of families during adolescence.

[Table 7 about here.]

DISCUSSION

Previous research on the effects of family structure on premarital childbearing typically assumes that women are at risk of a premarital birth even prior to the onset of sexual activity. In this paper, we develop models for the transition to a premarital first birth that explicitly account for the ways in which the transition into sexual activity may operate to mediate the effects of family structure on premarital first births. Our models are, in some formal respects, similar to recursive structural equation models for a static metric outcome (see, e.g., Heckman and Walker 1990), except that our mediating variable is an event—the onset of sexual activity—which in turn implies a substantive focus on mediating processes, rather than mediating attributes.

Our analyses investigate how family dynamics and processes experienced during childhood and adolescence might affect the subsequent life chances and trajectories of individuals. In particular, we examine three ways in which the onset of sexual activity might mediate the effects of family structure on a woman's age-specific risk of a premarital first birth. One possibility is that the gross association between family structure and premarital birth risks may be due entirely to the effect of family structure on age at first intercourse. Our results, which utilize data from two sources, the 1979–93 National Longitudinal Survey of Youth and the 1988 National Survey of Family Growth, show that the process of entry into sexual activity does indeed partially mediate the effect on premarital birth risks of both snapshot measures of family structure at age 14 and a time-varying measure of the number of family transitions. Nevertheless, we observe significant effects of both sets of family structure variables on premarital birth risks even after conditioning on various mediating effects of age at first intercourse.

Age at first intercourse may mediate effects of family structure on premarital birth risks in at least two other ways. One is that the earlier the age at first intercourse, the longer the duration of exposure to the risk of a premarital first birth, an effect that we term the "exposure" effect of first intercourse. Another is that an early age at first intercourse may proxy unmeasured individual characteristics correlated with age at onset but uncorrelated with other variables in the model, an effect that we term the "correlate" effect of first intercourse.

Our analyses reveal close agreement in estimates of exposure and correlate effects across our two data sources for both white and black women. For black women, the increase in the cumulative relative risk due to the exposure effect of delaying age at first intercourse by one year past the median age at first intercourse is roughly comparable to the increased cumulative risk of residing in a mother-only family at age 14 relative to residing with both biological parents at age 14. For white women, estimated exposure effects are similar in magnitude to those for black women, but unlike for blacks, estimated exposure effects are smaller in magnitude than the effect for whites of residing in a mother-only family at age 14. For both white and black women in the NLSY, the exposure effects are uniformly larger in magnitude than the corresponding correlate effects. For black women in the NSFG, the exposure effects are also uniformly larger than the corresponding correlate effects, but for white women in the NSFG, exposure effects tend to be smaller than the corresponding correlate effects.

More generally, our analyses show that estimates of exposure and correlate effects can tell one much about the dynamic interplay between onset of sexual activity and premarital first births. For example, our results make clear that exposure effects vary with age—that is, delaying age at first intercourse from 16 to 17 will, in general, yield a different exposure effect than delaying age at first intercourse from 14 to 15—in ways that are intimately related to the age-specific variation in the risk of a premarital first birth. Similarly, because women who initiate intercourse earlier may differ systematically from those who initiate intercourse later in ways not observed by the researcher, the gross effect of delaying age at first intercourse cannot be simply equated with the effect that might be obtained if one were able to alter age at onset exogenously, as in a classical experimental treatment. These observations provide important cautionary notes regarding the subtleties that arise when attempting to evaluate programs that seek to delay sexual activity or to promote more effective contraceptive practices in teen or high-risk populations.

REFERENCES

Amato, Paul R. 1987. "Family Processes in One-Parent, Stepparent, and Intact Families: The Child's Point of View." *Journal of Marriage and the Family* 49: 327–37.

Astone, Nan Marie, and Sara S. McLanahan. 1991. "Family Structure and High School Completion: The Role of Parental Practices." *American Sociological Review* 56(3): 309–20.

Astone, Nan Marie, and Sara S. McLanahan. 1994. "Family Structure, Residential Mobility, and School Report: A Research Note." *Demography* 31(4): 575–84.

Axinn, William G., and Arland Thornton. 1996. "The Influence of Parents' Marital Dissolutions on Children's Attitudes Towards Family Formation." *Demography* 33(1): 66–81.

Bumpass, Larry L., and Sara S. McLanahan. 1989. "Unmarried Motherhood: Recent Trends, Composition, and Black-White Differences." *Demography* 26(2): 279–286.

Capaldi, Deborah M., Lynn Crosby, and Mike Stoolmiller. 1996. "Predicting the Timing of First Sexual Intercourse for At-Risk Adolescent Males." *Child Development* 67: 344–59.

Chase-Lansdale, P. Lindsay, Andrew J. Cherlin, and Kathleen E. Kiernan. 1995. "The Long-Term Effects of Parental Divorce on the Mental Health of Young Adults: A Developmental Perspective." *Child Development* 66: 1614–34.

Cherlin, Andrew J. 1978. "Remarriage as an Incomplete Institution." *American Journal of Sociology* 83: 634–50.

Cherlin, Andrew J. 1992. *Marriage, Divorce, Remarriage*. Revised and enlarged edition. Cambridge, MA: Harvard University Press.

Cherlin, Andrew J., and Frank F. Furstenberg, Jr. 1994. "Stepfamilies in the United States: A Reconsideration." *Annual Review of Sociology* 0: 359–81.

Cherlin, Andrew J., Frank F. Furstenberg, Jr., P. Lindsay Chase-Lansdale, Kathleen E. Kiernan, Philip K. Robins, Donna Ruane Morrison, and Julien O. Teitler. 1991. "Longitudinal Studies of Effects of Divorce on Children in Great Britain and the United States." *Science* 252(5011): 1386–89.

Cherlin, Andrew J., Kathleen E. Kiernan, and P. Lindsay Chase-Lansdale. 1995. "Parental Divorce in Childhood and Demographic Outcomes in Young Adulthood." *Demography* 32(3): 299–318.

Dornbusch, Sanford M., J. Merrill Carlsmith, Steven J. Bushwall, Philip L. Ritter, Herbert Leiderman, Albert H. Hastorf, and Ruth T. Gross. 1985. "Single Parents, Extended Households and the Control of Adolescents." *Child Development* 56(2): 326–41.

Forrest, Jacqueline Darroch, and Susheela Singh. 1990. "The Sexual and Reproductive Behavior of American Women, 1982–1988." *Family Planning Perspectives* 22: 206–14.

Furstenberg, Frank F. Jr. 1987. "The New Extended Family: The Experience of Parents and Children after Remarriage." Pp. 42–61 in *Remarriage and Stepparenting: Current Research and Theory*, edited by K. Pasley and M. Ihinger-Tallman. New York: Guilford.

Furstenberg, Frank F. Jr., and Andrew J. Cherlin. 1991. *Divided Families: What Happens to Children when Parents Part.* Cambridge, MA: Harvard University Press.

Goldscheider, Frances K., and Calvin Goldscheider. 1993. *Leaving Home Before Marriage: Ethnicity, Familism, and Generational Relationships.* Madison, WI: University of Wisconsin.

Guidubaldi, J., H. K. Cleminshaw, J. D. Perry, B. K. Nastasi, and J. Lightel. 1986. "The Role of Selected Family Environment Factors in Children's Post-Divorce Adjustment." *Family Relations* 35(1): 141–51.

Hayes, Cheryl D. (Ed.) 1987. *Risking the Future: Adolescent Sexuality, Pregnancy, and Childbearing.* Washington, D.C.: National Academy Press.

Heckman, James J., and James R. Walker. 1990. "The Third Birth in Sweden." *Journal of Population Economics* 3: 235–75.

Hetherington, E. Mavis. 1972. "Effects of Father Absence on Personality Development in Adolescent Daughters." *Developmental Psychology* 7(3): 313–26.

Hetherington, E. Mavis. 1981. "Children and Divorce." Pp. 33–58 in R. W. Henderson (Ed.), *Parent-Child Interaction: Theory, Research, and Prospects*. New York: Academic Press.

Hetherington, E. Mavis. 1987. "Family Relations Six Years After Divorce." Pp. 185–205 in in Kay Pasley and Marilyn Ihinger-Tallman (Eds.), *Remarriage and Stepparenting: Current Research and Theory.* New York: Guilford.

Hetherington, E. Mavis, Kathleen A. Camara, and David L. Featherman. 1983. "Achievement and Intellectual Functioning of Children in One Parent Households." Pp. 205–84 in Janet T. Spence (Ed.), *Achievement and Achievement Motives: Psychological and Sociological Approaches*. San Francisco: W. H. Freeman.

Hetherington, E. Mavis, and W. Glenn Clingempeel. 1992. *Coping with Marital Tran-*

sitions. Chicago: Society for Research in Child Development.

Hogan, Dennis P., and Evelyn M. Kitagawa. 1985. "The Impact of Social Status, Family Structure, and Neighborhood on the Fertility of Black Adolescents." *American Journal of Sociology* 90(4): 825–55.

Kellam, Sheppard G., Margaret E. Ensminger, and R. Jay Turner. 1977. "Family Structure and the Mental Health of Children: Concurrent and Longitudinal Community-Wide Studies." *Archives of General Psychiatry* 34(September): 1012–22.

Martinson, Brian C., and Lawrence L. Wu. 1992. "Parent Histories: Patterns of Change in Early Life." *Journal of Family Issues* 13(3): 351–77.

Matsueda, Ross L., and Karen Heimer. 1987. "Race, Family Structure, and Delinquency: A Test of Differential Association and Social Control Theories." *American Sociological Review* 52(6): 826–40.

May, Elaine Tyler. 1988. *Homeward Bound: American Families in the Cold War Era*. New York: Basic.

McLanahan, Sara S. 1985. "Family Structure and the Reproduction of Poverty." *American Journal of Sociology* 90(4): 873–901.

McLanahan, Sara S. 1988. "Family Structure and Dependency: Early Transitions to Female Household Headship." *Demography* 25(1): 1–16.

McLanahan, Sara S., and Larry L. Bumpass. 1988. "Intergenerational Consequences of Family Disruption." *American Journal of Sociology* 94(1): 130–52.

McLanahan, Sara S., and Gary Sandefur. 1994. *Growing Up with a Single Parent*. Cambridge, MA: Harvard University Press.

Moore, Susan, and Doreen Rosenthal. 1993. *Sexuality in Adolescence*. London: Routledge.

Morrison, Donna Ruane, and Andrew J. Cherlin. 1995. "The Divorce Process and Young Children's Well-Being: A Prospective Analysis." *Journal of Marriage and the Family* 57: 800–12.

Mueller, Charles W., and Hallowell Pope. 1977. "Marital Instability: A Study of its Transmission Between Generations." *Journal of Marriage and the Family* 39(1): 83–94.

Newcomer, Susan F., and J. Richard Udry. 1987. "Parental Marital Status Effects on Adolescent Sexual Behavior ." *Journal of Marriage and the Family* 49: 235–40.

Rutter, Michael. 1983. "Stress, Coping, and Development: Some Issues and Some Questions." Pp. 1–42 in Norman Garmezy and Michael Rutter (Eds.), *Stress, Coping, and Development in Children*. New York: McGraw Hill.

Thornton, Arland. 1991. "Influence of the Marital History of Parents on the Marital and Cohabitational Experiences of Children." *American Journal of Sociology* 96(4): 868–94.

Thornton, Arland, and Donald Camburn. 1987. "The Influence of the Family on Premarital Sexual Attitudes and Behavior." *Demography* 24(3): 323–40.

U.S. Bureau of the Census. 1991. "Fertility of American Women: June 1990." *Current Population Reports* P20–454. Washington, D.C.: U.S. Government Printing Office.

Weiss, Robert S. 1975. *Marital Separation.* New York: Basic Books.

Wojtkiewicz, Roger. 1993. "Simplicity and Complexity in the Effects of Parental Structure on High School Graduation." *Demography* 30(4): 701–18.

Wu, Lawrence L. 1996. "Effects of Family Instability, Income, and Income Instability on the Risk of a Premarital Birth." *American Sociological Review* 61(3): 386–404.

Wu, Lawrence L., and Brian C. Martinson. 1993. "Family Structure and the Risk of a Premarital Birth." *American Sociological Review* 58(2): 210–32.

Wu, Lawrence L., and Elizabeth Thomson. 1995. "Family Change and Early Sexual Initiation." CDE Working Paper No. 95-26. Center for Demography and Ecology, University of Wisconsin-Madison.

Table 1. Kaplan-Meier estimates of the cumulative percent experiencing first sexual intercourse by age for white and black women born 1958–65. National Longitudinal Survey of Youth, 1979–93, and National Survey of Family Growth 1988.

Age	13	14	15	16	17	18	19	20	21	22	n
White women, NLSY	1	3	7	15	32	50	68	76	81	86	2367
White women, NSFG	2	4	7	17	31	47	62	72	78	83	1353
Black women, NLSY	2	6	12	25	49	67	82	88	91	93	1189
Black women, NSFG	2	7	17	29	48	66	80	87	92	94	784

Table 2. Effects of family structure variables on the age-specific rate of: first sexual intercourse, a premarital first birth ignoring the duration of exposure since onset of first intercourse, and a premarital first birth controlling for the duration of exposure since onset of first intercourse. White women, National Longitudinal Survey of Youth, 1979–1993 (n = 2367), and National Survey of Family Growth, 1988 (n = 1353).

	First intercourse		Premarital first birth							
			Not controlling for duration of exposure since onset of first intercourse				Controlling for duration of exposure since onset of first intercourse			
	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Respondent's family structure at age 14										
mother-only	.36***	.57***	.96***	1.07***	.68***	.71**	.71***	.76**	.66***	.66*
	(.07)	(.09)	(.15)	(.23)	(.15)	(.23)	(.15)	(.23)	(.15)	(.23)
step	.39***	.55***	.77***	.81*	.46*	.43	.53**	.54	.48*	.47
	(.08)	(.12)	(.17)	(.32)	(.17)	(.33)	(.17)	(.32)	(.17)	(.32)
other	.20	.64***	.48*	.90*	.19	.57	.31	.62	.26	.52
	(.11)	(.13)	(.24)	(.35)	(.24)	(.35)	(.24)	(.35)	(.24)	(.35)
Correlate effect of age at first intercourse										
age at onset					−.024***	−.028***			−.008**	−.015***
					(.002)	(.003)			(.003)	(.004)

Note: All models control for age and the background variables common to both data sources. See text for model specifications for duration of exposure to the age-graded risk of a premarital first birth since age of onset at first sexual intercourse and for the correlate effect of age at first sexual intercourse on the age-specific rate of a premarital first birth.

* $p < .05$ ** $p < .005$ *** $p < .0005$ (two-tailed tests)

Table 3. Effects of family structure variables on the age-specific rate of: first sexual intercourse, a premarital first birth ignoring the duration of exposure since onset of first intercourse, and a premarital first birth controlling for the duration of exposure since onset of first intercourse. Black women, National Longitudinal Survey of Youth, 1979–1993 (n = 1189), and National Survey of Family Growth, 1988 (n = 784).

	First intercourse		Premarital first birth							
			Not controlling for duration of exposure since onset of first intercourse				Controlling for duration of exposure since onset of first intercourse			
	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG	NLSY	NSFG
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Respondent's family structure at age 14										
mother-only	.32***	.36***	.39***	.38***	.22*	.11	.22*	.16	.19*	.12
	(.07)	(.08)	(.09)	(.11)	(.09)	(.11)	(.09)	(.11)	(.09)	(.11)
step	.23*	.32*	.48***	.23	.43**	−.01	.38*	.02	.38*	−.01
	(.11)	(.16)	(.14)	(.21)	(.14)	(.21)	(.14)	(.21)	(.14)	(.21)
other	.28*	.00	.52***	.22	.46**	.25	.43**	.21	.43**	.21
	(.11)	(.12)	(.14)	(.15)	(.14)	(.15)	(.14)	(.15)	(.14)	(.15)
Correlate effect of age at first intercourse										
age at onset					−.023***	−.027***			−.007**	−.008**
					(.002)	(.002)			(.002)	(.002)

Note: All models control for age and the background variables common to both data sources. See text for model specifications for duration of exposure to the age-graded risk of a premarital first birth since age of onset at first sexual intercourse and for the correlate effect of age at first sexual intercourse on the age-specific rate of a premarital first birth.

* $p < .05$ ** $p < .005$ *** $p < .0005$ (two-tailed tests)

Table 4. Percentage change in the cumulative relative risk of a premarital first birth under alternative simulations for age at first sexual intercourse for white women in a mother-only family at age 14. National Longitudinal Survey of Youth, 1979–1993, and National Survey of Family Growth, 1988.

Simulation	Baseline group		Comparison group		Percentage change in cumulative relative risk					
					NLSY			NSFG		
	t_1	t_2	t_1	t_2	exposure	correlate	total	exposure	correlate	total
1.	$\hat{t}_1$	$\hat{t}_1+60$	$\hat{t}_1-24$	$\hat{t}_1+60$	38.5	20.6	67.0	27.0	42.0	80.4
2.	$\hat{t}_1$	$\hat{t}_1+90$	$\hat{t}_1-24$	$\hat{t}_1+90$	27.7	20.6	54.0	19.3	42.0	69.4
3.	$\hat{t}_1$	$\hat{t}_1+120$	$\hat{t}_1-24$	$\hat{t}_1+120$	22.2	20.6	47.3	15.7	42.0	64.3
4.	$\hat{t}_1$	$\hat{t}_1+60$	$\hat{t}_1-12$	$\hat{t}_1+60$	24.1	9.8	36.2	19.8	19.2	42.8
5.	$\hat{t}_1$	$\hat{t}_1+90$	$\hat{t}_1-12$	$\hat{t}_1+90$	17.3	9.8	28.8	14.2	19.2	36.1
6.	$\hat{t}_1$	$\hat{t}_1+120$	$\hat{t}_1-12$	$\hat{t}_1+120$	13.9	9.8	25.0	11.5	19.2	32.9
7.	$\hat{t}_1$	$\hat{t}_1+60$	$\hat{t}_1+12$	$\hat{t}_1+60$	−23.7	−8.9	−30.5	−21.5	−16.1	−34.1
8.	$\hat{t}_1$	$\hat{t}_1+90$	$\hat{t}_1+12$	$\hat{t}_1+90$	−17.1	−8.9	−24.5	−15.4	−16.1	−29.0
9.	$\hat{t}_1$	$\hat{t}_1+120$	$\hat{t}_1+12$	$\hat{t}_1+120$	−13.6	−8.9	−21.3	−12.5	−16.1	−26.6
10.	$\hat{t}_1$	$\hat{t}_1+60$	$\hat{t}_1+24$	$\hat{t}_1+60$	−45.5	−17.1	−54.8	−41.9	−29.6	−59.1
11.	$\hat{t}_1$	$\hat{t}_1+90$	$\hat{t}_1+24$	$\hat{t}_1+90$	−32.8	−17.1	−44.3	−29.9	−29.6	−50.7
12.	$\hat{t}_1$	$\hat{t}_1+120$	$\hat{t}_1+24$	$\hat{t}_1+120$	−26.2	−17.1	−38.8	−24.4	−29.6	−46.7

Note: Calculations based on estimates in column (10) of Table 2 and on $\hat{t}_1$ = median age at first sexual intercourse for white women in a mother-only family at age 14 in the relevant survey. See text for definitions of the percentage change in the cumulative relative risk due to the duration of exposure to the age-graded risk of a premarital first birth since onset of first sexual intercourse and due to the correlate effect of onset of first sexual intercourse on premarital first birth risks.

Table 5. Percentage change in the cumulative relative risk of a premarital first birth under alternative simulations for age at first sexual intercourse for black women in a mother-only family at age 14. National Longitudinal Survey of Youth, 1979–1993, and National Survey of Family Growth, 1988.

Simulation	Baseline group		Comparison group		Percentage change in cumulative relative risk					
					NLSY			NSFG		
	t_1	t_2	t_1	t_2	exposure	correlate	total	exposure	correlate	total
1.	$\hat{t}_1$	$\hat{t}_1 + 60$	$\hat{t}_1 - 24$	$\hat{t}_1 + 60$	36.4	16.9	59.4	38.0	21.4	67.4
2.	$\hat{t}_1$	$\hat{t}_1 + 90$	$\hat{t}_1 - 24$	$\hat{t}_1 + 90$	26.9	16.9	48.4	28.4	21.4	55.8
3.	$\hat{t}_1$	$\hat{t}_1 + 120$	$\hat{t}_1 - 24$	$\hat{t}_1 + 120$	22.7	16.9	43.4	24.0	21.4	50.5
4.	$\hat{t}_1$	$\hat{t}_1 + 60$	$\hat{t}_1 - 12$	$\hat{t}_1 + 60$	22.0	8.1	31.9	21.7	10.2	34.1
5.	$\hat{t}_1$	$\hat{t}_1 + 90$	$\hat{t}_1 - 12$	$\hat{t}_1 + 90$	16.3	8.1	25.7	16.2	10.2	28.0
6.	$\hat{t}_1$	$\hat{t}_1 + 120$	$\hat{t}_1 - 12$	$\hat{t}_1 + 120$	13.7	8.1	22.9	13.7	10.2	25.3
7.	$\hat{t}_1$	$\hat{t}_1 + 60$	$\hat{t}_1 + 12$	$\hat{t}_1 + 60$	–24.9	–7.5	–30.5	–23.7	–9.2	–30.8
8.	$\hat{t}_1$	$\hat{t}_1 + 90$	$\hat{t}_1 + 12$	$\hat{t}_1 + 90$	–18.4	–7.5	–24.5	–17.7	–9.2	–25.3
9.	$\hat{t}_1$	$\hat{t}_1 + 120$	$\hat{t}_1 + 12$	$\hat{t}_1 + 120$	–15.5	–7.5	–21.8	–15.0	–9.2	–22.9
10.	$\hat{t}_1$	$\hat{t}_1 + 60$	$\hat{t}_1 + 24$	$\hat{t}_1 + 60$	–45.7	–14.5	–53.6	–45.0	–17.6	–54.7
11.	$\hat{t}_1$	$\hat{t}_1 + 90$	$\hat{t}_1 + 24$	$\hat{t}_1 + 90$	–33.8	–14.5	–43.4	–33.6	–17.6	–45.3
12.	$\hat{t}_1$	$\hat{t}_1 + 120$	$\hat{t}_1 + 24$	$\hat{t}_1 + 120$	–28.5	–14.5	–38.8	–28.5	–17.6	–41.1

Note: Calculations based on estimates in column (10) of Table 3 and on $\hat{t}_1$ = median age at first sexual intercourse for black women in a mother-only family at age 14 in the relevant survey. See text for definitions of the percentage change in the cumulative relative risk due to the duration of exposure to the age-graded risk of a premarital first birth since onset of first sexual intercourse and due to the correlate effect of onset of first sexual intercourse on premarital first birth risks.

Table 6. Effects of alternative family structure variables on the age-specific rate of a premarital first birth. White women, National Longitudinal Survey of Youth, 1979–1993 ($n = 2367$).

	number of family transitions	age at first sexual intercourse	born out-of-wedlock	pct of early life, mom-only family	pct of all life, mom-only family	mom-only family, adol. years	step family, adol. years	other family, adol. years	ever in divorced family	ever in remarried family
a.3	.26***									
a.7	.17**									
a.9	.16**	−.005								
b.3	.26***		.06							
b.7	.17**		−.10							
b.9	.16**	−.006	−.17							
c.3	.25***			.19						
c.7	.17**			.06						
c.9	.16**	−.005		.04						
d.3	.23***				.57					
d.7	.16**				.24					
d.9	.15*	−.005			.23					
e.3	.23***					.18	.16	.18		
e.7	.17**					−.08	.02	−.02		
e.9	.16*	−.005				−.06	.02	−.02		
f.3	.18*								.05	.25
f.7	.08								.02	.36
f.9	.08	−.005							.01	.35

Note: All models control for age and the extended set of background variables in the NLSY. Estimates in rows labelled a.3–f.3 parallel estimates in Model 3 of Table 2, which omit controls for duration since onset of first intercourse; estimates in rows labelled a.7–f.7 parallel estimates in Model 7 of Table 2, which include controls for duration since onset; estimates in rows labelled a.9–f.9 in Model 9 of Table 2, which include controls for duration since onset of first intercourse and for a correlate effect of age at first intercourse. See text for additional details about model specification. Estimated standard errors are available upon request.

* $p < .05$ ** $p < .005$ *** $p < .0005$ (two-tailed tests)

Table 7. Effects of alternative family structure variables on the age-specific rate of a premarital first birth. Black women, National Longitudinal Survey of Youth, 1979–1993 ($n = 1189$)

	number of family transitions	age at first sexual intercourse	born out-of-wedlock	pct of early life, mom-only family	pct of all life, mom-only family	mom-only family, adol. years	step family, adol. years	other family, adol. years	ever in divorced family	ever in remarried family
a.3	.23***									
a.7	.17***									
a.9	.17***	−.005*								
b.3	.23***		.17							
b.7	.17***		.10							
b.9	.17***	−.005*	.09							
c.3	.23***			.11						
c.7	.17***			.00						
c.9	.17***	−.005*		−.01						
d.3	.23***				−.03					
d.7	.18***				−.16					
d.9	.17***	−.005*			−.18					
e.3	.16**					.24	.09	.66***		
e.7	.13*					.01	.05	.49**		
e.9	.12*	−.005*				.01	.07	.50**		
f.3	.30***								−.12	−.11
f.7	.23**								−.17	.01
f.9	.22**	−.005*							−.17	.02

Note: All models control for age and the background variables. Estimates in rows labelled a.3–f.3 parallel estimates in Model 3 of Table 3, which omit controls for duration since onset of first intercourse; estimates in rows labelled a.7–f.7 parallel estimates in Model 7 of Table 3, which include controls for duration since onset; estimates in rows labelled a.9–f.9 in Model 9 of Table 3, which include controls for duration since onset of first intercourse and for a correlate effect of age at first intercourse. See text for additional details about model specification. Estimated standard errors are available upon request.

* $p < .05$ ** $p < .005$ *** $p < .0005$ (two-tailed tests)