FAMILY STRUCTURE AND THE RISK OF A PREMARITAL BIRTH

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The positive association between growing up in a nonintact family and the risk of a first premarital birth has been interpreted by researchers as consistent with three hypotheses: (1) a childhood socialization hypothesis — that women who grow up in a mother-only family during early childhood are socialized in ways that result in a high risk of a premarital birth; (2) a social control hypothesis — that the supervision of adolescents is more difficult in single-parent families than in two-parent families; and (3) an instability and change hypothesis — that a premarital birth is a response to the stresses accompanying changes in a woman's family situation. Although these hypotheses imply distinct behavioral mechanisms, adjudicating between them has proven difficult, in part because researchers have relied on static measures of family structure. We use data from the National Survey of Families and Households and continuous-time hazard models to investigate the effects on premarital births of dynamic family measures that reflect a woman's family situation between birth and age 19. Our findings are consistent with the instability and change hypothesis, but provide little support for the socialization hypothesis and the social control hypothesis.

Recent changes in marriage and childbearing in the United States have led to a dramatic increase in the prevalence of families headed by single mothers. Increases in nonmarital childbearing and divorce suggest that over 50 percent of children born in recent birth cohorts will spend part of their childhood living with a single parent (Bumpass 1984; Bumpass and Rindfuss 1979). Chronic poverty is increasingly concentrated in single-parent households, particularly those headed by never-married women (Bane and Ellwood 1986; Duncan and Hoffman 1990). These trends have led to a resurgence of interest in

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1 The terms "intact" and "nonintact" are somewhat misleading because families always headed by a single mother are intact in the sense of never hav-
sponse by a woman to the stresses accompanying instability in her family situation.

Surprisingly, researchers have been unable to adjudicate among these hypotheses even though they imply distinctive, and in certain respects contradictory, behavioral mechanisms (Cherlin and Horiuchi 1980; Demo and Acocock 1988). We believe that the most immediate reason for this difficulty is the use of crude measures of family structure — typically a static snapshot of family situation at age 14. We construct dynamic measures of family structure using parent-history data from the National Survey of Families and Households.2 We use these data to examine the relative importance of family events during childhood, changes in family structure during childhood and adolescence, and durations spent in the modal family structures experienced by respondents while growing up. Our results suggest that these dynamic measures capture both theoretically and empirically distinct dimensions of family experience, and that distinguishing between these dimensions provides sufficient analytical leverage to test the socialization, social control, and instability and change hypotheses.

THEORY

Our review of theories on family structure and the risk of a first premarital birth departs somewhat from standard accounts by emphasizing how different perspectives stress different elements of a respondent's family situation between early childhood and adolescence. This discussion motivates the construction of family measures that tap theoretically distinct aspects of a respondent's family experience.

Socialization During Early Childhood

Specialists in early childhood development have generally held that parent-child interactions during early childhood have lasting consequences for later behaviors. For example, some authors have hypothesized that the absence of a father may inhibit the development of a healthy gender identity, appropriate sex-typed behaviors, and a positive sexual self image (Freud [1925] 1961; Goode 1956; Guidubaldi, Clemminshaw, Perry, Nastasi, and Lightel 1986; Hetherington 1972, 1981; McCord, Mc Cord, and Thurber 1962). Hetherington (1972) argued that these problems may increase the likelihood of sexual "acting out" during early adulthood and thus increase the risk of a premarital birth.

Other authors have argued that a young child may suffer from the presence of a single mother who makes single motherhood acceptable or desirable in the eyes of the child or adolescent (Kellam, Ensminger, and Turner 1977; McLanahan 1988; Mueller and Pope 1977; Thornton and Camburn 1987). This argument is more complicated than appears at first glance because a child can view the role model provided by a single mother in a positive or negative light depending on the "extent to which the parent is perceived as similar to what the child is and wants to be" (Maccoby and Martin 1983). These considerations place even more importance on very early childhood because the young child has little independent identity and few nonparental role models. Thus, living with a single mother during the first few years of life may have an especially strong influence on a child's acquisition of social norms, values, and sense of self; these factors may in turn form the basis of a child's subsequent selection of other role models.

The empirical evidence is inconsistent (Demo and Acocock 1988; Emery 1988; Herzog and Sudia 1973; Maccoby and Martin 1983; Shinn 1978). Some studies have found that severing parental attachments through divorce or desertion during the first few years of life is strongly associated with negative outcomes in later life (Rutter 1971; Rutter and Quinlan 1984). Others have found that the effects of parental absence vary in character and intensity with the age of the child, with an especially severe impact on very young children (Emery 1988; Hetherington, Cox, and Cox 1979; Krein and Beller 1988; Longfellow 1979; McLanahan 1985; McLanahan and Bumpass 1988; Shaw 1982; Wallerstein and Kelly 1980).

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2 The National Survey of Families and Households was designed and carried out at the University of Wisconsin, Madison under the direction of Larry Bumpass and James Sweet with funding from NICHD (HD 21009).
Social Control of Adolescents

A second hypothesis derives from the sociological literature on the social control of adolescents, a perspective rooted in studies of adolescent deviance (Gove 1985; Hirschi and Gottfredson 1983; Matsueda and Heimer 1987). Family researchers have argued that two-parent families exercise greater supervision and control than do single-parent families, and hence that single-parent families may be less able than two-parent families to prevent adolescent outcomes like a premarital birth (Dornbusch et al. 1985; Hogan and Kitagawa 1985; Matsueda and Heimer 1987; McLanahan and Bumpass 1988; Thomson, McLanahan, and Curtin 1992). An important distinction between the socialization and social control hypotheses is that the socialization hypothesis stresses the effect of prior experience on current behavior while the social control hypothesis stresses the effect of current family situation on current behavior. Thus, from a social control perspective, the number and types of adults present during adolescence are important predictors of the risk of premarital birth for adolescent women.

In a sense, the social control perspective views adolescence as a particularly trouble-prone period by presuming that adolescents are apt to engage in inappropriate behavior unless constrained by parents. A somewhat different viewpoint asks what family events might trigger such inappropriate behaviors. For example, Rubin (1976) suggested that "too-early" adult transitions (e.g., a premarital birth) may be a response to a particularly unfavorable family situation such as living with foster parents or with neither biological parent. Like the social control hypothesis, this argument supposes that adolescent behavior is affected primarily by current situation rather than past experience; however, it predicts an especially strong association between an unconventional family situation and the risk of a premarital birth.

Empirical support for these hypotheses is mixed. Hogan and Kitagawa (1985) found strong effects of parental supervision on the risk of premarital pregnancy and early sexual activity for black adolescent women in Chicago. Dornbusch et al. (1985) compared adolescents in mother-headed families in which no other adult was present, mother-headed families with another adult present, and stepfamilies. A social control argument predicts that deviance should be high for adolescents in mother-headed families with no other adults present, and low in the other types of families. However, they found that deviance was similar in mother-only families and stepfamilies, and lower in mother-headed families with another adult present. Steinberg (1987) also found no difference in adolescent susceptibility to peer pressure between mother-only families and stepparent families.

Instability and Change

A third hypothesis focuses on the responses of children and adolescents to the stresses accompanying a major disruption in the family environment. While many researchers are skeptical of research on stress because of measurement difficulties and ambiguities in the concept, most would agree that a parental divorce or remarriage constitutes a major stressor in the lives of children and adolescents (Garmezy 1983). Such a major change often disrupts a child’s or adolescent’s sense of emotional security in ways that distinguish such an event from other stressors like a major childhood illness, a geographical move, or a change in economic well-being. Similarly, the departure of a parent after divorce or the addition of a new parental figure after remarriage may create uncertainty in children and adolescents about whether they can rely on a parent for stability and emotional support.

These arguments suggest that various stresses may accompany a major family disruption, and that changes other than divorce may trigger responses by children and adolescents. Researchers have drawn from these ideas in several ways. Some have argued that frequent changes in family situation increase the likelihood that a woman prematurely assumes adult statuses, including motherhood prior to marriage (McLanahan 1985). Others have argued that the disequilibrium following a major family disruption may push offspring out of the parental home (White and Booth 1985; McLanahan 1988). Finally, some have found that the disequilibrium following a disruption causes some adolescents to disengage from the parental household, for example, by spending less time in the home or by engaging in impulsive, rebellious, or aggressive actions (Hetherington 1987; Wallerstein and Kelly 1980;
Wallerstein and Blakeslee 1989). Disengagement could have short-term consequences if an adolescent woman seeks emotional support or intimacy outside the home following a disruption, but could also have long-term consequences if it alters the way a woman construes (or behaves in) intimate relationships (Wallerstein and Blakeslee 1989).

These arguments suggest that the instability accompanying frequent family changes is an important factor increasing the risk of a premarital birth. An important distinction between the social control hypothesis and the instability and change hypothesis is that the social control hypothesis emphasizes the effects of particular family statuses, while the instability and change hypothesis emphasizes the effects of particular family events (McLanahan 1988). The above arguments also suggest that a disruption could have a short-term effect during the period immediately following a change in family situation (McLanahan 1988) or a long-term effect with consequences for behavior much later in life (Longfellow 1979; Maccoby and Martin 1983). These possibilities are not mutually exclusive since the risk of a premarital birth could increase rapidly in the short term following a change in family situation, but decline slowly (or not at all) in the long term.

The empirical evidence is largely indirect, with particularly little known about the long-term consequences of family instability (Hetherington 1987; McLanahan 1988; Rutter 1983). Most research has focused on the consequences of parent/child separation early in life or marital disruption during childhood or adolescence. Rutter (1971) found strong short-term effects, but only weak long-term effects, of parent/child separation. He attributed this finding to the fact that children separated from both parents often came from disturbed homes and had experienced many spells of foster care following various family crises. Rutter (1983) later found that the adverse effect of a stressful event is greater for children who experience chronic stress than for children who experience only a single stressful event. Hetherington, Camara, and Featherman (1983), Wallerstein and Blakeslee (1989), and Hetherington (1987) found that young children behave less well, study less effectively, and are more apt to be truant during a two- to three-year period before and after a marital disruption. Others have found that after a disruption, parenting abilities may deteriorate (Weiss 1979) as both parents and children slowly adjust to new circumstances (Chase-Lansdale and Hetherington 1990; Hetherington, Cox, and Cox 1982). Cherlin et al. (1991) reported that the effect of a disruption on school behaviors and achievement is reduced after controlling for prior behavioral problems, achievement, and family discord.

Although most researchers have focused on the consequences of a disruption, some have suggested that other changes in parental situations may have negative consequences. For example, the incomplete institutionalization of "reconstituted" families means that individuals in stepfamilies encounter numerous problems that do not arise in intact families — the roles of stepparents and noncustodial parents in disciplining children, the forging of emotional attachments between nonkin, the fiscal and legal responsibilities of stepparents and nonkin children, and even the proper terms to describe nonkin (Cherlin 1978; Price-Bonham and Balswick 1980; see also Amato 1987; Furstenberg and Nord 1985; Walker and Messinger 1979). Others have found that children typically maintain close ties with only one or two parental figures in a pattern of "serial parenthood" following the divorce and remarriage of parents (Seltzer and Bianchi 1988; see also Furstenberg, Nord, Peterson, and Zill 1983), and that children tend not to identify stepfamily members when asked about family composition (Furstenberg 1987).

MEASURING FAMILY STRUCTURE


Snapshot measures of family structure have several difficulties. First, they conflate clearly distinct groups, including children in (1) never-married mother-only families, (2) recently dis-
ruptured mother-only families, and (3) mother-only families with multiple disruptions. Based on the preceding discussion, similar outcomes would be expected for all three groups from a social control perspective, different outcomes for groups 1 and 2 from a childhood socialization perspective, and different outcomes for groups 1 and 3 based on a family instability and change perspective. Past research typically has confounded these effects into a single estimated effect of nonintactness. Second, family situation at age 14 is only a rough proxy for family situation during adolescence because it ignores changes after age 14 and equates the experiences of children in a recently disrupted family with those with extended periods in a disrupted family. Finally, parental absence may have different consequences as a child's cognitive and social skills change with age (Allison and Furstenberg 1989; Emery 1988; Longfellow 1979; Maccoby and Martin 1983; Wallerstein and Kelly 1980). Past research typically ignores these issues, for example, by confounding disruption during late adolescence with father absence during early childhood.

There are exceptions to the widespread use of snapshot measures. Using data from the Panel Study of Income Dynamics (PSID), McLanahan (1985) examined the effect of duration since marital disruption on schooling outcomes and found that children in recently disrupted households are especially likely to drop out of high school. Unfortunately, intact families and stepfamilies could not be distinguished at first interview; hence, the duration in a disrupted family is biased downward for some respondents. McLanahan and Bumpass (1988), using data from the 1982 National Survey of Family Growth, found no effect of age at disruption on the risk of a premarital birth for white women and only a weak effect at ages 5 through 9 for black women. Allison and Furstenberg (1989), using data from two panels of the National Survey of Children, found small but significant effects of a marital dissolution on children's well-being, especially for young children. However, the panel design of these data made it impossible to disentangle the effects of age and duration. Cherlin et al. (1991), analyzing schooling outcomes in two panels of the National Survey of Children and the British National Child Development Study, found that the effect of divorce was significant for girls, but not for boys, after controlling for problem behaviors at the first time point. Astone and McLanahan (1991) analyzed two panels of the High School and Beyond Survey; they found that a marital disruption had a negative effect on school attendance, parental monitoring, and parental supervision, but no significant effect on grades or college aspirations. Finally, Sandefur, McLanahan, and Wojtkiewicz (1992), using data from the 1979–1985 waves of the National Longitudinal Surveys of Youth, found that a change in family situation between ages 14 and 17 had a significant negative effect on high school graduation, even after controlling for family income.

DATA AND METHODS

The National Survey of Families and Households (NSFH) provides a national probability sample of persons ages 19 and over who resided in the United States in 1987 and 1988 (Sweet, Bumpass, and Call 1988). There are 13,017 respondents consisting of a main sample of 9,643 cases and an oversample of 3,374 cases of minorities, single-parent families, families with stepchildren, cohabiting couples, and recently married persons.

The survey provides detailed information on a woman's parental, fertility, marital, and home-leaving histories. Unfortunately, information on family income during a respondent's childhood and adolescent years is not available, owing in part to the retrospective design of the NSFH. This limits our ability to contrast hypotheses on economic deprivation and family structure, although other background variables provide rough controls for economic status during a respondent's formative years.

Of the 7,790 women in the NSFH, we excluded (1) those born before 1937 (N = 2,609), (2) those in the oversample who were cohabiting at the time of interview (N = 144), and (3) those with invalid marital, fertility, or parental histories, and minorities other than blacks and Hispanics (N = 261). Imposing these restrictions resulted in a sample size of 3,372 white women, 978 black women, and 426 Hispanic women. Roughly 10 percent (313) of white

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3 Women who reported a first birth before age 13 were considered to have invalid fertility data and were excluded. Women in the oversample who were cohabiting at the time of interview were excluded to avoid oversampling on a potentially en-
women, 50 percent (496) of black women, and 25 percent (106) of Hispanic women experienced a premarital birth.

Data on our outcome — age (in months) at a first premarital birth — are obtained from fertility and marriage histories. Women who had not married or given birth were censored at the date of interview. Similarly, women who legitimated their first birth or who married without (or before) giving birth were censored at the date of marriage. Finally, we took age 13 (156 months) to be the age at which a woman first becomes at risk of a birth.

Measures

We constructed measures of family situation from a retrospective "parent calendar" which provides data at each age from birth to age 19+ on whether a respondent lived with biological parent(s), stepparent(s), adopted parent(s), or in some other situation (e.g., with grandparents or other relatives). We merged these data with a home-leaving history to obtain the family situations of respondents before they left home. Table 1 reports the 20 family situations defined for these data.

Our stress on dynamic measures of family structure might seem inappropriate if there were little flux in respondents' family situations. Table 2 reports the percentage distribution for selected family sequences between birth and age 19 (see also Martinson and Wu 1992). These sequences describe the family trajectories of respondents ignoring the timing of changes. For example, 73.0 percent of white women born between 1938 and 1969 lived in an intact family from birth until home-leaving or age 19, 7.7 percent were born into an intact family but then lived with a biological mother only, and 3.3 percent were born into an intact family, then lived with a biological mother only, and then with a biological mother and stepfather.

Most respondents lived in an intact family from birth until age 19 or leaving home. Still, about one-fourth of white women, one-half of black women, and one-third of Hispanic women fall outside this traditional pattern. Black women are much more likely than white women to have lived in a mother-only family, but for black women, entering a mother-only family through a disruption is as likely as entering a mother-only family at birth — 19 percent of the 978 black women in our sample were born into a mother-only family, while an equal percentage lived first in an intact family and later in a mother-only family (data not shown in Table 2).

Table 2 reveals a striking degree of diversity in the family situations of respondents (see also Hofferth 1985; Martinson and Wu 1992; Wojtkiewicz 1992). About 90 percent of white women fall into one of the 12 sequences listed, but the remaining 10 percent require 175 sequences to characterize their family experiences. Black and Hispanic women show even greater variability — for example, about 17 percent of black women cannot be characterized by the 12 sequences listed. Thus, while a few sequences summarize the family situations of most respondents, many patterns are necessary to summarize the situations of the remaining respondents. Sequence length also varies considerably. Short sequences are typical of the
dogenous variable. We retained all other women in the oversample of single families, stepfamilies, and recently married families because we adjust directly for these factors by treating them as censored alternatives to a premarital birth.
vast majority of black and white women, but some respondents have extremely long sequences indicating frequent changes.

*Static and Dynamic Measures of Family Situation*

Table 2 shows that the parent history data are, in some sense, too rich. This problem is compounded by the fact that the sequences in Table 2 ignore the timing of changes in family situation, but including information on timing (e.g., by distinguishing between changes in early childhood and adolescence) would only complicate matters further.

We address this problem by constructing measures motivated by our theoretical concerns. The adolescent social control hypothesis contends that an adolescent woman's risk of a premarital birth is (1) influenced by her current, not prior, family situation,4 and (2) varies with the number and types of adults currently present in her household. We operationalize these ideas by creating four age-varying dummy variables for respondents in intact families, mother-only families, families with one biological parent and one stepparent, and all other types of families. These variables are thus identical to those used by other researchers, except that we use age-varying measures instead of a static snapshot measure.

The descriptive statistics in Table 3 contrast the snapshot measures and dynamic measures of family situation during adolescence. The top panel presents percentages for the snapshot measure at age 14; the second panel reports percentages who ever fell into these categories between ages 13 and 19. The two sets of distributions are similar except that the percentages ever living in each family situation are 2 to 5 percent higher than the comparable percentages for family situation at age 14.

We operationalized the socialization hypothesis using three alternative measures of exposure to a mother-only family. The first measure examines socialization during early childhood using a dummy variable coded 1 if the respondent was born into a mother-only family and 0 otherwise. The second measure also examines socialization during early childhood, but uses a dummy variable coded 1 if the respondent spent at least 75 percent of the first six years of life in a mother-only family and 0 otherwise. The third measure examines socialization over all ages using an age-varying dummy variable coded 1 at age t if the respondent spent 75 percent or more of life through age t in a mother-only family, and 0 otherwise. Because the parent calendar gathers data to the nearest year, the second measure contrasts those who spent five or six years of early life in a mother-only family with those who spent four years or less in a mother-only family. The third measure can change value at any (monthly) age as the percentage of life spent by a respondent in a mother-only family changes with age.

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4 The influence of a woman's current family situation on the risk of a premarital birth refers to its effect at age t (or, more generally, at age t - Δ for some specified positive lag Δ) on the risk of a premarital birth at age t.
Large race/ethnic differences emerge for these measures (third panel of Table 3). Percentages for the first measure show that 3.7 percent of white women, 8.7 percent of Hispanic women, and 18.8 percent of black women were born into a mother-only family. The percentages for the second and third measures of exposure to a mother-only family are similar and vary strongly with race and ethnicity, but are lower than the percentages for the first measure.

We operationalized the instability and change hypothesis using an age-varying variable for the number of changes in family situation experienced by a respondent using all 20 family situations in Table 1. Note that this variable captures some changes (e.g., the change from living with a biological mother and stepfather to living with a biological father and stepmother) that are not reflected in the other age-varying family variables. Because this measure varies with age, we report the percentage distribution for 0, 1, 2, and 3 or more changes for those who ever fell into these categories between ages 13 and 19; hence, percentages sum to more than 100 percent. The bottom panel shows that black women are substantially more likely to have experienced one or more changes in family situation than are white women or Hispanic women.

Background Variables

Background variables are: whether the respondent was raised in the Catholic faith; socioeconomic index of the respondent’s father; dummy variables for years of schooling completed (0–11, 12, and 13 and over) by the respondent and by the respondent’s mother; number of full siblings; a time-varying variable for historical period coded 1 on or after July 1973; dummy variables for respondent’s year of birth (1938–1947, 1948–1957, and 1958–1969); a time-varying dummy variable indicating whether the respondent was living in the parental home; mother’s age at respondent’s birth; and an interaction between mother’s age at respondent’s birth and a dummy variable coded 1 if the respondent is the oldest child. The last two variables are crude controls for the association between early childbearing by the respondent and her mother.

Models

One problem in analyzing the timing of a first birth is that some women have not given birth by the date of interview, in part because women of different ages have different exposures to the risk of a birth. We use hazard regression methods (Cox and Oakes 1984; Tuma and Hannan 1984) to deal with such right-censored observations, to allow the risk of a premarital birth to depend on individual characteristics of the woman.
teristics, and to adjust for the dependence of the hazard rate on age, duration, historical period, and cohort. Because age at first premarital birth is measured to the nearest month, we use continuous-time hazard models and maximum-likelihood estimation methods to make full use of the temporal precision in these data. The models incorporate proportional effects of the covariates and lag all time-varying variables by 12 months to adjust for the 9-month period between conception and birth.6 (For details, see Appendix.)

RESULTS

Figure 1 presents smoothed nonparametric estimates of the logarithm of the premarital birthrate using a procedure described by Wu (1989). The curve for black women is substantially higher than the curve for white women, with the curve for Hispanic women falling midway between these two groups.

Figure 1 suggests a relatively simple pattern of age dependence, with a roughly linear rise in the logarithm of the premarital birthrate from age 13 to 16, a flatter linear rise from age 16 until about age 19, followed by a roughly linear decline. The curves for the three race and ethnic groups are approximately parallel when plotted on a logarithmic scale, as would be expected under the assumption of proportionality. We observed similar patterns when plotting the logarithm of the hazard rate by other observed factors and were unable to reject the proportionality assumption in our parametric models. Based on Figure 1, we modeled age dependence using a four-period piecewise exponential specification using the age intervals 13–15, 16–18, 19–24, and 25+ years (i.e., 156–191, 192–227, 228–299, and 300+ months).

Snapshot Analyses

We begin by examining the effects of the standard snapshot measures. The first panel of Table 4 reports results for the simplest and most widely used measure — a dummy variable contrasting women in intact families and
the effect of living in some other family situation is positive, but significant only for black women. For both groups, the effect of living in a stepfamily is not significant.

These results pose something of a puzzle. It might be tempting to conclude that most of the effect of nonintactness is attributable to living in a mother-only family. However, the effects of living in a mother-only family do not differ significantly from the effects of the two other family categories for any of the race/ethnic groups. Thus, another possible conclusion is that living in any nonintact family situation increases the risk of a premarital birth. The problem is that the associations in Table 4 say little about the behavioral mechanisms governing a premarital birth or exactly what it is about nonintactness that matters.

Dynamic Analyses I

Table 5 reports estimates for each race and ethnic group using dynamic measures of family structure. The zero-order estimates (Model 0) report the effects of a single family variable in order to examine each family hypothesis separately. Models 1 and 2 provide simultaneous tests of the socialization, social control, and instability and change hypotheses. All models include controls for age, historical period, cohort, home-leaving, and the background variables. To simplify the discussion, we report only the estimated effects of family situation.

White women. We begin with the zero-order effects of the exposure measures estimated from three equations, one for each measure of exposure to a mother-only family.8 Surprisingly, there are no significant effects on the risk of a premarital birth at age \( t \) for white women who were born into a mother-only family, who spent at least 75 percent of early childhood in a mother-only family, or who spent at least 75 percent of life in a mother-only family between birth and age \( t - 12 \) months.

We obtained the zero-order effects of current family situation by estimating a single equa-

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7 Estimated standard errors are not reported but are available for all tables on request.

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8 We examined other exposure measures using durations in family categories defined by the number and types of parents present during childhood and adolescence. These alternative measures had weaker effects than the measures reported here. However, the effects of other family measures were stable across specifications employing different exposure measures.

<table>
<thead>
<tr>
<th>Family Situation</th>
<th>White</th>
<th>Black</th>
<th>Hispanic</th>
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<tr>
<td></td>
<td>Model 0</td>
<td>Model 1</td>
<td>Model 2</td>
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<tr>
<td>Exposure to a Mother-Only Family</td>
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<td></td>
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<td>.44</td>
<td>—</td>
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<tr>
<td>At least 75 percent of life, all ages</td>
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<td>—</td>
<td>.22</td>
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<tr>
<td>Current Family Situation</td>
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<tr>
<td>Mother-only family</td>
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<tr>
<td>Stepfamily</td>
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<td>-.64</td>
<td>-.62</td>
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<td>Instability and Change in Family Situation</td>
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<tr>
<td>Number of changes in family situation</td>
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<td>.22***</td>
<td>.22***</td>
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<tr>
<td>Degrees of freedom</td>
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<td>24</td>
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</tr>
</tbody>
</table>

*p < .05          **p < .01           ***p < .001 (one-tailed test)

Note: All models control for age, historical period, cohort, home-leaving, and the background variables.

The zero-order estimates treat each hypothesis generously by not estimating the effects predicted by alternative hypotheses. Models 1 and 2, which operationalize the socialization hypothesis differently, provide direct tests of competing perspectives.

Model 1 contrasts the effects of the socialization, social control, and instability and change hypotheses using exposure to a mother-only family at birth and exposure to a mother-only family between ages 0 and 5 to operationalize the socialization hypothesis. Results are broadly similar to those in Model 0. Being born into a mother-only family and prolonged exposure to a mother-only family during early childhood continue to have no significant effect for white women. None of the estimates of the effects of current family situation is significant. The estimated effect of frequent changes in family situation remains positive and highly significant.

Model 2 tests the socialization hypothesis using exposure to a mother-only family at birth and at all ages. Estimates resemble those in Models 0 and 1. For white women, the effect...
of changes in family situation remains positive and highly significant, while the effects of all other family variables are not significant.

For white women, the major surprise concerns the effect of exposure to a mother-only family and the effect of currently living in a mother-only family. None of these effects is significant; hence, we cannot reject the null hypothesis that risks are the same for white women who were born into, who have had lengthy exposures to, or who currently live in a mother-only family relative to those in intact families. Taken together, these results are somewhat provocative given the strong (although sometimes controversial) emphasis in the literature on the disadvantages of living in a mother-only family.

**Black women and Hispanic women.** The results for black women and Hispanic women are broadly similar to those for white women, despite the different sample sizes. The effect of number of changes in family situation is positive and significant, while the effects of being born into a mother-only family, prolonged exposure to a mother-only family during early childhood, or prolonged exposure to a mother-only family during all ages are not significant. There is a positive and significant effect of currently living in a mother-only family in Model 0, but this effect is smaller and no longer significant in Models 1 and 2, which control for the other family variables. The other effects of current family situation are weak and not significant in all three models.

**Dynamic Analyses II**

We noted earlier that the instability and change hypothesis can also be operationalized by modeling the duration since the most recent change in family situation. Table 6 presents results from three alternative model specifications of the instability and change hypothesis that model the effect of duration since the most recent change in family situation for women who experienced at least one family change. As before, the zero-order estimates (Model 0) report the effects of a single family variable in order to examine each family hypothesis separately, while Models 1 and 2 estimate effects predicted by the socialization, social control, and instability and change hypotheses. All models control for age, historical period, cohort, home-leaving, and the background variables.

**Two-period exponential model.** The two-period exponential model specifies short- and long-term effects of family change by letting the level of the premarital birthrate differ for short durations and long durations since the most recent change in family situation for women who experienced at least one change in family situation.⁹

Model 0 for white women indicates that the risk of a premarital birth at age t is higher for both short durations and long durations since a change in family situation at age t = 12 months relative to white women who experienced no change in family situation (.24 or an exp(.24) ≈ 27 percent higher risk during the first 23.4 months and .62 or a 83 percent higher relative risk after 23.5 months). The estimate at short durations is not significant, while the estimate at long durations is highly significant. The effects of duration for white women are virtually identical in Models 1 and 2. The risk of a premarital birth is high (.57 or a 77 percent higher relative risk) during the first 23.4 months following a change in family situation, but this effect is not significant. After 23.5 months, the risk of a premarital birth is even higher (.77 or a 116 percent higher relative risk) and this effect is highly significant.

For black women, all duration effects in the two-period exponential model are significant, but the pattern of duration dependence differs from that for white women. Black women who experienced one or more changes in family situation have a significantly higher risk of a premarital birth than do black women who experienced no change in family situation. However, the risk is highest at short durations and less at long durations, a pattern that differs from that for white women.

The pattern of duration dependence for Hispanic women resembles that for white women, although effects are generally smaller for Hispanics than for whites. As expected, risks for Hispanic women who experienced one or more changes are higher at both short durations and long durations than those for Hispanic women.

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⁹We model duration dependence in the logarithm of the hazard rate in terms of two level parameters, β₁ and β₂ for durations of 0 to 23.4 months, and β₃ for durations over 23.5 months. A duration effect is expected only for women who have experienced one or more changes in family situation; hence, we constrain β₁ and β₂ to be 0 until a change in family situation is observed. (See Appendix.)

<table>
<thead>
<tr>
<th>Model and Duration Parameter</th>
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<th>Black</th>
<th>Hispanic</th>
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<tr>
<td></td>
<td>Model 0</td>
<td>Model 1</td>
<td>Model 2</td>
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<tr>
<td>Two-Period Exponential Model</td>
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<tr>
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<td>.57</td>
<td>.57</td>
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<tr>
<td>$\beta_2$, duration 23.5+ months</td>
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<td>.77***</td>
<td>.77***</td>
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<td>Two-Period Splined Gompertz Model I</td>
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<td>-.61</td>
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<td>.076*</td>
<td>.078*</td>
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<td>-.005**</td>
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<td>-.005**</td>
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<tr>
<td>Number of changes in family situation</td>
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<td>-.06</td>
<td>-.07</td>
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</table>

* $p < .05$  ** $p < .01$  *** $p < .001$ (one-tailed test)

No standard errors are computed for bracketed estimates because of constraints on these parameters.

Note: All models control for age, historical period, cohort, home-leaving, and the background variables. Models 1 and 2 control for exposure to a mother-only family and adolescent family situation.

who experienced no change in family situation; however, only the effect at long durations is significant.

Two-period splined Gompertz model I. The two-period exponential model tests one version of the instability and change hypothesis by letting the risk of a premarital birth differ at short and long durations since the most recent change in family situation. The two-period splined Gompertz model tests yet another version of this hypothesis — that risks rise at short durations and decline at long durations — by specifying the logarithm of the premarital birthrate as one linear function of duration at short durations and another linear function of duration at long durations. Although we expect rising risks at short durations and declining risks at long durations, this model permits risks to be monotonically increasing (both slope parameters positive), monotonically decreasing (both slope parameters negative), unimodal as hypothesized (positive slope at short durations, negative slope at long durations), or V-shaped (negative slope at short durations, positive slope at long durations).

The results for white women and Hispanic women are consistent with expectations — rates rise steeply at short durations (positive slope parameter $\delta_1$) and decline slowly at long durations (negative slope parameter $\delta_2$). For white women, the estimated slope at short durations and long durations is statistically sig-

---

10 We model duration dependence in the logarithm of the hazard rate as $\beta_1 + \delta_1 \times u$ for durations between 0 and 23.4 months, and $\beta_2 + \delta_2 \times u$ for durations after 23.5 months, where $u$ denotes duration; $\beta_1$ and $\beta_2$ denote the intercept parameters at short durations and long durations, respectively; and $\delta_1$ and $\delta_2$ denote the slope parameters at short durations and long durations, respectively. We specify a duration effect only for women who experienced one or more changes in family situation by constraining $\beta_1$, $\beta_2$, $\delta_1$, and $\delta_2$ to be 0 until a family change occurs. Finally, we impose a spline constraint that $\beta_1 + \delta_1 \times u = \beta_2 + \delta_2 \times u$ for $u = 23.5$ months to make the hazard rate a continuous function of duration. Because of this constraint, standard errors were not calculated for $\beta_2$ (See Appendix.)
nificant in all three models. For Hispanic women, the estimated slope at short durations is not significant, while the estimated slope at long durations approaches significance in Models 0 and 1 ($p = .061$ and .078, respectively) and is significant in Model 2 ($p = .046$). There are large differences in the magnitudes of the slope parameters at short durations and long durations. For white women, the estimated slope at short durations is about 15 times larger than the estimated slope at long durations; for Hispanics, the estimated slope at short durations is about 6 times larger than the estimated slope at long durations.

Among black women, risks decline monotonically with duration, and this decline is slightly faster at short durations than at long durations. Only the estimated slope at long durations is significant.

Two-period splined Gompertz model II. The second two-period splined Gompertz model adds the variable for number of changes in family situation to the first two-period splined Gompertz model. For all three race/ethnic groups, the effect of number of changes in family situation is no longer significant.\(^\text{11}\)

Figure 2 shows how the logarithm of the premarital birthrate varies with age and duration for the Model 2 estimates. Recall that this model lets the premarital birthrate vary with age for all women but constrains duration parameters to be 0 until a change in family situation occurs. The upper graph shows how the logarithm of the predicted premarital birthrate varies with age for the reference group — women who experienced no change in family situation. The bottom graph shows the predicted pattern of duration dependence, i.e., how risks vary with duration for women who experienced at least one change in family situation relative to those who experienced no change in family situation.

\(^{11}\) The two-period exponential model and the first two-period splined Gompertz model are semi-Markovian with respect to duration in that they constrain the premarital birthrate to depend only on a woman's age and duration in her current family status, after controlling for background variables and other family variables. The second two-period splined Gompertz model relaxes this assumption by letting the logarithm of the premarital birthrate vary linearly with number of changes in family situation. The results provide no grounds for rejecting the semi-Markovian assumption because the effect of number of changes is small and not significant, particularly for white women and black women.

The bottom panel clarifies the short- and long-term effects of a change in family situation. For white women and Hispanic women, the logarithm of the rate rises at short durations, with both curves lying mostly above 0. Thus, for white women and Hispanic women, the short-term effect of a change in family situation is to increase risks relative to women who experienced no change in family situation. At long durations, the logarithm of the rate for white women and Hispanic women declines slowly, which suggests a surprisingly persistent effect of a change in family situation.

Figure 2 helps to reconcile an apparent anomaly in findings from the two-period exponential and two-period splined Gompertz models. Results for the two-period exponential model in Table 6 suggested that risks for white women and Hispanic women are low at short durations and high at long durations, while results for the two-period splined Gompertz models suggested that risks for white women and Hispanic women rise at short durations and decline at long durations. Figure 2 shows that there is no inconsistency: At short durations risks are low but rise quickly, whereas at long durations risks are high and decline slowly. Indeed, the point estimates from the two-period exponential model (.57 and .77 for short and long durations, respectively, for white women; .25 and .56 for short and long durations, respectively, for Hispanic women) are similar to the overall levels shown in Figure 2.

Table 7 reports estimates for the other family variables for the second two-period splined Gompertz model. The zero-order estimates are identical in Tables 5 and 7, but the estimates in Models 1 and 2 are also similar despite the fact that the model in Table 5 assumes no duration structure for the premarital birthrate, while the model in Table 7 lets the premarital birthrate vary with duration according to a two-period splined Gompertz model.

We continue to observe only a small effect or a negative effect of being born into a mother-only family. As expected, prolonged exposure to a mother-only family during early childhood and prolonged exposure to a mother-only family during all ages have positive effects; however, these effects are not significant.\(^\text{12}\)

\(^{12}\) We obtain similar results when the variable for exposure to a mother-only family at birth is omitted from Models 1 and 2. (Results available from the authors on request.)
For current family situation, several effects that were not significant and that were in the opposite direction of expectations remain in the opposite direction of expectations, but are now significant. This is true for currently being in a mother-only family or in some other family for white women, and (in Model 2) for being in a stepfamily for black women.

**Predicted probabilities by age 18.** Table 8 presents the predicted probability of a premarital birth by age 18 (216 months) using estimates from the second two-period splined Gompertz model. In computing these probabilities, we assigned the mean or modal values for the background variables within each racial and ethnic group. We also assumed that women were born after 1958 and were at risk of a premarital birth in the period after July 1973.

For women who always lived in an intact family, the models predict that approximately 2 percent of white women, 25 percent of black women, and 12 to 13 percent of Hispanic women will have had a premarital birth by age 18. For women who lived in an intact family from birth through age 13 and in a mother-only family from age 14 to age 18, the predicted probabilities rise to 3 percent for white women, 37 percent for black women, and 26 to 29 percent for Hispanic women.

The next three sets of estimates illustrate the influence of the timing and sequence of family changes. For women who lived in an intact family from birth through age 13, in a mother-only family from age 14 through age 15, and in a stepfamily age 16 and after, the predicted probabilities are somewhat higher than for women who always lived in an intact family. For women with the same sequence of family situations but who experienced changes earlier rather than later in life, the predicted probabilities are higher for white women and Hispanic women, but lower for black women.

For women who always lived in a mother-only family, the predicted probabilities are 1 percent for white women, 28 to 29 percent for black women, and 20 to 21 percent for His-
panic women. For women who lived in a mother-only family from birth to age 13 and in a stepfamily age 14 and after, the probabilities rise to 6 percent for white women, 30 to 34 percent for black women, and 24 to 30 percent for Hispanic women.

Our analytical models found that prolonged exposure to a mother-only family had a relatively weak effect on the risk of a premarital birth, a finding mirrored in the predicted probabilities. Predictions for black women and Hispanic women who always lived in a mother-only family are, as expected, higher when compared to women who always lived in an intact family; however, they are lower than those for women who lived in an intact family at ages 0 to 13 and in a mother-only family at ages 14 to 18. For white women, the results are even more unexpected: Women who always lived in an intact family do not have a lower predicted probability of a premarital birth by age 18 than those who always lived in a mother-only family. Predictions for black women are also consistent with the instability and change hypothesis, but they differ in important respects from those for whites and Hispanics. Figure 2 shows that, all else being equal, black women who experienced one or more changes in family situation have a higher predicted risk of a premarital birth than black women who experienced no change in family situation. Table 8 shows that black women whose last change in family situation was 2 to 4 years ago have a higher predicted probability of a premarital birth than black women who experienced no change. However, our analytical models also found that black women who experienced one or more changes in family situation have a high initial rate that declines with duration. Table 8 shows that black women born into an intact family and whose last change in family situation was six to eight years ago have predicted probabilities that converge to those for black women who always lived in an intact family.

* Predicted probabilities by family intactness at age 14. Clearly, our results challenge conventional wisdom by suggesting a stronger effect of family instability and a weaker effect of prolonged exposure to a mother-only family than previous research would suggest. One

<table>
<thead>
<tr>
<th>Family Situation</th>
<th>White</th>
<th></th>
<th>Black</th>
<th></th>
<th>Hispanic</th>
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<tbody>
<tr>
<td></td>
<td>Model 1</td>
<td>Model 2</td>
<td>Model 1</td>
<td>Model 2</td>
<td>Model 1</td>
<td>Model 2</td>
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<tr>
<td>Intact family at ages 0–18</td>
<td>2.0</td>
<td>2.0</td>
<td>24.7</td>
<td>24.6</td>
<td>13.0</td>
<td>12.3</td>
</tr>
<tr>
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<td>3.0</td>
<td>36.9</td>
<td>36.5</td>
<td>28.8</td>
<td>26.1</td>
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<tr>
<td>Intact family at ages 0–13, mother-only at ages 14–15, stepfamily at ages 16–18</td>
<td>2.2</td>
<td>2.3</td>
<td>30.0</td>
<td>30.4</td>
<td>21.0</td>
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<tr>
<td>Intact family at ages 0–5, mother-only at ages 6–9, stepfamily at ages 10–18</td>
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<td>23.5</td>
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<tr>
<td>Intact family at ages 0–5, mother-only at ages 6–11, stepfamily at ages 12–18</td>
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<td>24.1</td>
<td>25.0</td>
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<tr>
<td>Mother-only at ages 0–18</td>
<td>1.3</td>
<td>1.3</td>
<td>27.9</td>
<td>28.8</td>
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<td>20.9</td>
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<tr>
<td>Mother-only at ages 0–13, stepfamily at ages 14–18</td>
<td>5.5</td>
<td>6.0</td>
<td>29.9</td>
<td>33.9</td>
<td>23.6</td>
<td>29.9</td>
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</table>

Possibility is that these differences stem from the use of snapshot measures, which forced previous researchers to confound distinct effects of family structure into a single estimate for nonintactness.

Table 9 examines this assertion empirically by computing predicted probabilities of a premarital birth for subsamples of the data. We computed the probability of a premarital birth by age 18 (216 months) for each white woman, black woman, and Hispanic woman using her observed characteristics. We then averaged predictions using estimates from the models in Tables 4 through 7 for the entire sample, for women living in intact families at age 14, and for women in nonintact families at age 14. Finally, we calculated the nonparametric estimate of the probability of a premarital birth by age 18 for these three groups using the Kaplan-Meier estimator (Kaplan and Meier 1958), which provides unbiased estimates of the probability of a premarital birth.

These estimates address three distinct issues. First, they provide probabilities predicted by the parametric models after accounting for compositional differences between subgroups. Second, they indicate whether the intact/nonintact family differential reported in previous research can be reproduced by our models. Finally, they provide a rough assessment of absolute fit because model predictions can be compared to the unbiased nonparametric estimates (Wu and Tuma 1991).

For black women and Hispanic women, all models underpredict relative to the nonparametric estimates. For white women, all models overpredict slightly for those who lived in intact families at age 14 and underpredict slightly for those who lived in nonintact families at age 14. Not surprisingly, predictions for the snapshot model of family situation are close to the nonparametric estimates for intact and nonintact families at age 14 because this model controls for nonintactness at age 14. By contrast, there are no a priori grounds for expecting agreement between the other models and the nonparametric estimates because these models do not control for nonintactness at age 14. Thus, the close agreement between the nonparametric estimates and the predictions from all models suggests that all models provide an adequate fit to the data and that our alternative models reproduce the differential in the risk of a premarital birth of living in an intact family at age 14 and living in a nonintact family at age 14 observed in prior research.

DISCUSSION

Although the increasing availability of longitudinal data and dynamic models has attracted considerable interest among social scientists, a review of the literature to date could lead one to conclude that it makes little difference whether temporal data and methods are used. One reaction is to be skeptical about the utility

<table>
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<th>Model and Subgroup</th>
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<th></th>
<th></th>
<th>Black</th>
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<td>Model 2</td>
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<td>Snapshot Model (Table 4)</td>
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of temporal analyses; another is to try to exploit the potential of temporal data and methods more fully. Although no single study can settle such an issue, this paper indicates that attending to the dynamics of change matters a great deal.

We employed dynamic measures of family structure that trace a respondent’s parental history between birth and late adolescence, and used these measures in models that allow the risk of a premarital birth to vary with age, historical period, cohort, and duration since a change in family situation. Using time-varying measures of family situation provides a dramatically different picture of the premarital birth process than is obtained from standard static measures of family situation. We began our analyses with a replication using snapshot measures of family situation to model the risk of a premarital birth. We then constructed a comparable set of dynamic analyses by letting the conventional snapshot measures vary with time in order to track changes in a respondent’s family situation during adolescence. The differences were substantial: Analyses using standard snapshot measures yielded large, positive, and highly significant effects on the risk of a premarital birth; analyses using dynamic family measures produced small and sometimes negative effects that usually were not significant.

Our main analyses attempted to adjudicate between three hypotheses linking family structure and the risk of a premarital birth. We investigated these hypotheses directly using alternative model specifications and indirectly using predictions from these models. Our findings were consistent with an instability and change hypothesis, but provided little support for a socialization hypothesis and a social control hypothesis.

Change in a woman’s family situation had a large and highly significant effect on the risk of a premarital birth for white women and His-
panic women, and a smaller but still significant effect for black women. Moreover, these effects were surprisingly persistent, particularly for white women and Hispanic women, even at long durations after a change in family situation. These results are noteworthy because we observed a strong effect of instability in all models and because the results provide a more direct test of the instability and change hypothesis than has been possible to date.

An unstable family environment may exert a strong and persistent effect on children or adolescents by altering their perceptions of relationships with adults and peers following a divorce or remarriage. If so, the cumulative effect of family instability may affect a woman's behavior even much later in life. A premarital birth may also represent a rational response by a young woman to an unstable family environment because bearing a child is one way to escape from undesirable family circumstances (Rubin 1976). However, this argument appears best suited to explaining a short-term effect of family instability. Yet another explanation is that an unstable family environment alters parenting practices and reduces parental control as adults and children adjust to new circumstances and roles following a change in family situation. Changes in parenting practices and parental control could have long-term consequences if they result in persistent problem behaviors by children or adolescents. A final possibility is that families marked by chronic instability represent a particularly problem-prone segment of the population. Under such a selection hypothesis, children in families marked by chronic instability could be at especially high risk of problem behaviors for reasons unrelated to family instability.

Our results provide little support for hypotheses emphasizing the effects of socialization. We tested three variants of this hypothesis by examining the consequences of being born into a mother-only family, spending large fractions of life in a mother-only family during very early childhood, and spending large fractions of life in a mother-only family during the entire period between birth and late adolescence. Contrary to expectations, the effects on the risk of a premarital birth of being born into a mother-only family were small, not significant, and often in the opposite direction expected for women in all three racial and ethnic groups.

Our analyses also revealed positive but not significant effects of prolonged exposure to a mother-only family during early childhood and prolonged exposure to a mother-only family during all ages.

These findings raise questions about the socialization hypothesis because they provide little support for the general expectation of disadvantage accruing to women in mother-only families. Still, the positive but not significant effects of lengthy exposures to a mother-only family observed in these data might attain significance if larger samples of women were available. Similarly, the significance of these effects in our models could be deflated if important differences among mother-only families were not captured well by our exposure measures. Yet another possibility is that detailed data on the interactions between parents and offspring may provide greater explanatory power than the exposure measures we examined.

Finally, our results provide little support for a hypothesis drawn from the literature on the social control of adolescents. Effects of current family situation were often small, not significant, and in the wrong direction for women in all three race and ethnic groups. We emphasize, however, that our operationalization of this hypothesis proceeded in a structural way by focusing on the number and types of parents present during adolescence. Although such an approach is typical of research in this area, parental control and authority undoubtedly vary across one-parent and two-parent families — some mother-only families may exercise strong control over the behavior of their adolescent children, while some two-parent families may have only weak control over their adolescent children. These considerations may explain our findings for black women, which would otherwise appear inconsistent with the results for black women reported by Hogan and Kitagawa (1985).

In considering racial and ethnic differences more generally, we have found it useful to speculate on cultural patterns guiding family life and the regulation of adolescent sexual activity. Demographically, a first premarital birth is the modal fertility pattern for recent cohorts of black women; it is somewhat less common for Hispanic women, and much less common for white women. Some researchers have as-
serted that family life and, in particular, the social meanings attached to a premarital birth, differ for blacks, whites, and Hispanics (Burton 1990; Furstenberg, Brooks-Gunn, and Morgan 1987; Goode 1960; Ooms 1981). Thus, some of the structural aspects of family situation that we have emphasized may be less salient for black women than for white women or Hispanic women. Ethnographic studies of the black community have documented that "absent" biological fathers often retain some contact with offspring and that grandparents and nonkin often play important roles in raising children, particularly in mother-only families (Burton 1990; Furstenberg 1976; Stack 1974). The smaller effect of changes in family situation for black women relative to white women or Hispanic women may reflect adaptations by black children to a more complex family environment produced by the greater numbers and types of adults present between birth and adolescence. This weaker effect may also be a result of growing up in an urban ghetto environment (Wilson 1987) or the effect of poverty and distinctive cultural practices on the process of family reproduction (Burton 1990).

Perhaps the clearest implication of our findings for future research is that the alternative family measures and models used in this paper appear to capture both theoretically and empirically distinct dimensions of a woman's family experience. Although our results should be replicated on other data and for other outcomes, they illustrate the feasibility and utility of adjudicating between alternative explanations for the association between family structure and a premarital birth.

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