INSIGHTS FROM A SEQUENTIAL HAZARD MODEL OF ENTRY INTO SEXUAL ACTIVITY AND PREMARITAL FIRST BIRTHS

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ABSTRACT

In this paper, we depart from most previous research on premarital first births by holding that the typical woman is not at risk of a premarital first birth prior to becoming sexually active. Our model lets us examine how family background and other social and demographic factors influence premarital first births via two sequential processes: (1) a young woman's entry into sexual activity and (2) her subsequent risk of a premarital first birth in the period following onset. We use estimated model coefficients to decompose the probability of a premarital first birth into components reflecting group differences in exposure to risk generated by earlier or later entry into sexual activity and group differences in premarital first birth risks in the period following onset. Our analyses, using data from a nationally representative sample of women aged 14–21 in 1979, confirm previous findings that women from disadvantaged backgrounds initiate sexual activity earlier and have higher premarital first birth risks than more advantaged women. However, our decompositions indicate that differences in the timing of first intercourse have a far smaller influence on premarital first birth probabilities than do differences in risks following onset. We close by speculating on the possible substantive and policy implications of these results, particularly with respect to ongoing debates between proponents and critics of abstinence education.

Women in the United States who bear their first child outside of formal marriage often come from backgrounds marked by distinct social disadvantages (McLanahan and Bumpass 1988; Wu and Martinson 1993; McLanahan and Sandefur 1994; Edin and Lein 1996; Wu 1996; Hoffman and Foster 1997; McLanahan, Garfinkel, Reichman, and Teitler 2001; Sigle-Rushton and McLanahan 2004; Edin and Kefalas 2005; Currie 2006; England and Edin 2007). These prior disparities, coupled with hardships many unmarried mothers encounter following birth, compound the difficulties experienced by children born to unmarried mothers (Furstenberg 1976; Hofferth and Hayes 1987; Wu 2002; McLanahan 2004; Preston 2004; Furstenberg 2007). Together with continuing increases in nonmarital fertility, these issues have generated considerable interest in the ways social disadvantage is linked to nonmarital fertility.

A key antecedent of nonmarital fertility is sexual behavior. Differences in sexual behavior between disadvantaged and advantaged teens are commonly thought to contribute to the greater numbers of premarital births among disadvantaged youth. Surprisingly, past research says little about the precise nature of the linkages between early onset, sexual behaviors following onset, and premarital first births.

In this paper, we pay particular attention to the timing of first sexual intercourse. We model how social factors influence the timing of onset and a woman's risk of a premarital first birth following onset. This strategy presumes that the typical woman is not at risk of a birth prior to becoming sexually active. Our approach departs from past studies that have typically ignored data on the timing of first sexual intercourse and that thus assume implicitly that women are at risk of a birth both before and after becoming sexually active.

Our analyses use data from the 1979 National Longitudinal Survey of Youth (NLSY), a nationally representative sample of women aged 14–21 in 1979. These data contain highly detailed information that let us determine, to the nearest month, a woman's age at first sexual intercourse, a first birth, and a first marriage. Our analyses let us obtain explicit estimates of how family background and other social and demographic factors influence: (1) a woman's age-specific risk of entry into sexual activity as operationalized by her age at first sexual intercourse and (2) her subsequent age-specific risk of a premarital first birth in the period following onset. We use estimated coefficients from the resulting sequential hazard models to decompose the probability of a premarital first birth into components reflecting group differences in exposure to risk generated by earlier or later entry into sexual activity and group differences in premarital first birth risks in the period following onset. We pose three questions. First, what influences the timing of sexual onset? Second, what influences subsequent premarital first birth risks following onset? Third, how do differences in onset timing influence the probability that a woman proceeds to a premarital first birth?

The organization of this paper is as follows. We begin by reviewing the theoretical arguments and findings concerning the timing of sexual onset and premarital first births. Next, we review our sequential hazard modeling approach, our decomposition methods, and our data. We then turn to our empirical analyses and findings. We close by speculating on the possible substantive and policy implications of these results, particularly with respect to ongoing debates between proponents and critics of abstinence education.

THEORY

Although previous research often discusses how early sexual activity may be linked to heightened risks of a teen or nonmarital birth, only recently have researchers attempted to provide empirical evidence that speaks directly to this issue. Nevertheless, the theoretical linkage between early entry into sexual activity and the subsequent risk of a premarital pregnancy and birth is especially explicit in arguments made by abstinence proponents and critics. In this section, we review the arguments and empirical evidence linking a variety of social factors to premarital first birth risks. Many of these arguments and empirical findings apply as well to the timing of entry into sexual activity.

[Maybe a few short paragraph mapping theory section? Informal acknowledgement Timing of first sex and past research on nonmarital births

Past research has typically acknowledged, often informally, the linkages between early sexual activity and the risk of a teen or nonmarital birth. Several studies have documented that offspring of parents who have more permissive sexual attitudes are less likely to believe that nonmarital sexual intercourse is wrong, with these beliefs in turn associated with earlier entry into sexual activity and an increased risk of a teen or unplanned birth (see, e.g., Newcomer and Udry 1984; Thornton and Camburn 1987; Weinstein and Thornton 1989). Others have argued that teen sexual behaviors follow the model provided by parents, siblings, and other salient figures (Gagnon and Simon 1973; Haurin and Mott 1990; Hogan and Kitagawa 1985 Billy, Brewster, and Grady 1994; Brewster 1994a,b). Modeling in turn is often invoked to explain the association between earlier sexual onset, on the one hand, and increased premarital first birth risks, for children in nonintact families, relative to those in intact families. That is, since many single mothers or fathers engage in nonmarital sexual intercourse or are in cohabiting unions, children may conclude that sexual activity during adolescence is acceptable (Inazu and Fox 1980; McLanahan and Sandefur 1994; Thornton and Camburn 1987). Similarly, it is often argued that two biologial parents can better monitor teen behaviors, including dating and sexual activity (Dornbusch et al. 1985; Hogan and Kitagawa 1985; Matsueda and Heimer 1987; McLanahan and Bumpass 1988; Thomson, McLanahan, and Curtin 1992), than can single parents or parents in step-families. Several studies report findings consistent with such a hypothesis, for example, that greater parental supervision during adolescence is associated with lower levels of teen sexual activity (Hogan and Kitagawa 1985; Inazu and Fox 1980; Jessor and Jessor 1975; Miller et al. 1986; Small and Luster 1994).

The theoretical link between teen sexual activity, teen pregnancies, and teen premarital births is especially explicit in arguments made by abstinence proponents and critics. For example, Kim and Rector (2008) note delaying their entry into sexual activity will decrease exposure to the risk of nonmarital childbearing and hence that an emphasis on abstinence is "crucial to efforts aimed at reducing unwed childbearing and improving youth well-being." By contrast, abstinence skeptics often place emphasis on factors such as teen access to timely and comprehensive information on human sexuality and reproductive health, arguing that teens who are more knowledgeable in these ways will have lower the risk of teen pregnancies, HIV, and other sexually transmitted diseases by increasing effective contraception among teens who resist admonitions to be abstinent (see, e.g., Furstenberg, Moore, and Peterson 1985; Marsiglio and Mott 1986; Mauldon and Luker 1996; Darroch, Landry, and Singh 2000; Lieberman et al. 2000; O'Donnell, O'Donnell, and Stueve 2001; Lindberg, Santelli, and Singh 2006; Furstenberg 2007).

These expectations also are reflected in language contained in key provisions of the 1996 Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA), which in addition to ending AFDC also allocated funding to states "to provide abstinence education . . . with a focus on those groups most likely to bear chlidren out of wedlock" and to teach "that abstinence from sexual activity is the only certain way to avoid out-of-wedlock pregnancy" (PRWORA 1996, Section 510).[1] In response, many states and local school districts substantially revised their sex education curricula for middle and high school students (Sonfeld and Benson 2001), with declines in instruction about birth control methods and increases in instruction following federal "abstinence-only" guidelines (Lindberg, Santelli, and Sigh 2006).

Abstinence skeptics typically argue that teen access to timely and comprehensive information on human sexuality and reproductive health is likely to lower the risk of teen pregnancies, HIV, and other sexually transmitted diseases by increasing effective contraception among teens who

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[1]Section 510 of PRWORA states that "the term 'abstinence education' means an educational or motivational program that:

- Has as its exclusive purpose, teaching the social, psychological, and health gains to be realized by abstaining from sexual activity;
- Teaches abstinence from sexual activity outside marriage as the expected standard for all school-age children;
- Teaches that abstinence from sexual activity is the only certain way to avoid out-of-wedlock pregnancy, sexually transmitted diseases, and other associated health problems;
- Teaches that a mutually faithful monogamous relationship in the context of marriage is the expected standard of human sexual activity;
- Teaches that sexual activity outside of the context of marriage is likely to have harmful psychological and physical effects;
- Teaches that bearing children out-of-wedlock is likely to have harmful consequences for the child, the child's parents, and society;
- Teaches young people how to reject sexual advances and how alcohol and drug use increase vulnerability to sexual advances; and
- Teaches the importance of attaining self-sufficiency before engaging in sexual activity.

resist admonitions to be abstinent (see, e.g., Furstenberg, Moore, and Peterson 1985; Marsiglio and Mott 1986; Mauldon and Luker 1996; Darroch, Landry, and Singh 2000; Lieberman et al. 2000; O'Donnell, O'Donnell, and Stueve 2001; Lindberg, Santelli, and Singh 2006; Furstenberg 2007). They note that firm majorities of parents, teachers, and the public at large favor classroom instruction that both sends firm messages about abstinence but also provides students who are sexually active with information about how to reduce pregnancy, avoid teen and nonmarital births, and how to reduce STD and other health risks. Conversely, they argue that abstinence-only educational efforts may result in greater numbers of pregnancies and thus lead to greater numbers teen and premarital births; likewise, they view abstinence-only instruction as ignoring well-documented statistics (see, e.g., Abma et al. 2004) on the substantial numbers of teens and young adults who do not delay sexual activity until marriage

Although many social scientists have focused attention on the highly charged nature of debate between abstinence proponents and critics (see, e.g., Nathanson 1991, Luker 1996, Levine 2002; Furstenberg 2007), we wish instead to highlight the nature of the analytical arguments made by abstinence proponents and critics. Note, for example, that the logic of the argument typically made by abstinence proponents is identical to a classic social science insight—that all else being equal, decreased exposure to risk of an outcome will lower the probability of this outcome, and hence that factors delaying sexual onset will, all else being equal, lower the probability of a premarital first birth. Conversely, the logic of the argument typically made by abstinence critics is that premarital birth risks are more greatly influenced by by factors after women become sexually active. Stated in this way, both sets of arguments correspond to testable hypotheses for a given covariate x within the framework provided by in our sequential model of premarital first birth risks.

A recent evaluation of four sites featuring random assignment of students to courses following PWRORA abstinence-only guidelines found no difference between treatment and controls four to six years after treatment for a wide array of outcomes, including whether the student had initiated sexual activity, age at first intercourse, recent sexual activity, sexually transmitted diseases, pregnancies, and whether the student had given birth or fathered a child (Trenholm et al. 2007).

Although these findings have renewed skepticism among abstinence critics, proponents have countered that such conclusions are premature, in part because the experimental interventions were short in duration and administered in early grades, thus targetting students at ages before most become sexually active.

These recent findings thus shift the questions of interest to one of identifying potential factors that might exert more persistent influences on when teens initiate sexual activity and on their behaviors following onset. Thus, in response to skeptics, abstinence proponents hypothesize that a variety of factors are likely to be persistent influences on teen behaviors, with factors that delay onset in turn implying fewer out-of-wedlock births. Likewise, critics hypothesize that factors that are persistent influences on teen behaviors will be more important on risk behaviors following onset. These hypotheses thus raise several questions that can be examined empirically. What factors might be thought to be persistent influences on teen behaviors? By how much might such factors hasten or delay onset and how many more or fewer premarital first births are implied by the resulting variations in exposure to risk? Similarly, by how much do such factors increase or decrease premarital first birth risks following onset and what does this imply for the proportions who subsequently have a first birth outside of formal marriage?

[Rest not yet written. Paragraph noting that prior research has identified a number of factors that are, on theoretical grounds, plausible candidates for “persistent influences.” Review of empirical findings on age at first intercourse and on premarital first birth risks.]

[Review of empirical findings relevant to abstinence debate. Mixed evidence on effects of comprehensive sex education. Kirby review; Marsiglio and Mott; Furstenberg, Moore, and Peterson; Oettinger; Mauldon and Luker, etc. Reliance of many studies on cross-sectional data, variability of curriculum covered in sex education across teachers and schools, timing of when students are exposed to instruction vs. when they become sexually active. Mixed evidence regarding virginity pledges (Bearman and Bruckner; Bruckner and Bearman), instability of pledges (Hollander 2006).]

MODEL

Following Wu and Martin (2009), let T_1 and T_2 denote the random variables for a woman's age at first intercourse and a premarital first birth, respectively. Nearly all previous work on premarital first births proceeds by modeling T_2 but ignoring T_1 using, for example, a conventional proportional hazard specification

$$r_{2c}(t|\mathbf{x}_i) = q_{2c}(t)\exp(\mathbf{b}_{2c}\mathbf{x}_i)\,, \tag{1}$$

where i indexes women, $q_{2c}(t)$ denotes the age-specific baseline for T_2, $\mathbf{x}_i$ a set of observed covariates for woman i, $\mathbf{b}_{2c}$ the corresponding set of estimated coefficients, and the subscript "c" use of a "conventional" specification. Note that because (1) ignores the timing of first intercourse, it implicitly assumes that a woman is at risk of a birth both before and after she becomes sexually active. In a rare exception, Kiernan and Hobcraft (1997) incorporate onset timing into (1) via

$$r^*_{2c}(t) = q^*_{2c}(t)\exp(\alpha t_{1i} + \mathbf{b}^*_{2c}\mathbf{x}_i)\,, \tag{2}$$

where t_{1i} denotes woman i's age at onset of sexual activity. A difficulty with (2) is that it implicitly assumes knowledge of onset timing at *all* ages, including $t < t_1$.[2]

To avoid the behaviorally implausible assumptions in (1) and (2), we follow Wu and Martin (2009) by posing this problem in terms of two transitions—the transition $0 \to T_1$ for the onset of sexual activity, and the transition $T_1 \to T_2$ for a woman's ensuing risk of a premarital first birth. We model the age-specific risk of the $0 \to T_1$ transition using a conventional proportional hazard specification:

$$r_{1i}(t, \mathbf{x}_i) = q_1(t)\exp(\mathbf{b}_1\mathbf{x}_i)\,, \tag{3}$$

where t denotes age and $r_1(t)$ denotes the baseline age-specific risk for the $0 \to T_1$ transition. We then specify the $T_1 \to T_2$ transition *conditional* on t_1, a woman's observed age at onset

$$r_2(t, \mathbf{x}_i|t_{1i}) = q_{21}(t|t_{1i})\; q_{22}(u)\;\; \exp(\alpha t_{1i} + \mathbf{b}_2\mathbf{x}_i)\,, \tag{4}$$

[2] That is, consider a woman who initiates sexual activity at 17; then (2) models her birth risks at age 14 with knowledge that she will initiate sexual activity three years in the future.

where $u = t - t_1$ denotes duration since onset and $q_{21}(t|t_1)$ and $q_{22}(u)$ are baseline functions for the dependence of T_2 risks on age and duration. Note in particular that (4) differs from (1) and (2) by assuming that women do not become at risk of a premarital first birth until they become sexually active and by allowing T_2 risks to depend on both age and duration since onset.

This sequential approach provides a richer empirical structure for modeling premarital first births than the more conventional model in (2). It assumes that all never-married women have an identifiable period during which their premarital first birth risks are negligible—the ages prior to when they become sexually active, that never-married women will vary considerably in when they become sexually active, that premarital first birth risks will also vary systematically for women in the period following the onset of sexual activity, and that observed factors will influence both age at onset and the risk of a premarital birth subsequent to onset.[3]

The models in (3) and (4) also carry implications for the probability of initiating sexual activity and of having a first birth prior to a first marriage. Under (3), the probability that woman i will have initiated sexual activity by age t is given by

$$\begin{aligned} \Pr(T_{1i} \leq t|\mathbf{x}_i) &= 1 - \Pr(T_{1i} > t|\mathbf{x}_i) \\ &= 1 - S_1(t|\mathbf{x}_i) \\ &= 1 - \exp\Big[- \int_0^t r_1(s|\mathbf{x}_i)\, ds\Big] \end{aligned} \tag{5}$$

where S_1 denotes the so-called survivor function. Similarly the probability that woman i has a premarital first birth by age t, conditional on onset at age t_1, is given by:

$$\begin{aligned} \Pr(T_{2i} \leq t|\mathbf{x}_i, T_1 = t_1) &= 1 - \Pr(T_{2i} > t|\mathbf{x}_i, T_1 = t_1) \\ &= 1 - S_2(t, \mathbf{x}|t_1) \\ &= 1 - \exp\Big[- \exp(\alpha t_1 + \mathbf{b}_2\mathbf{x}) \int_{t_1}^t q_{21}(v)\, dv \int_0^u q_{22}(w)\, dw\Big] \end{aligned} \tag{6}$$

with the timing of entry into T_2 risk reflected in the lower limits of integration in (6).

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[3] An additional key assumption is that T_1 and T_2 are not jointly determined, that is, that the typical woman does not seek to become pregnant and carry the pregnancy to term when first initiating sexual activity. This assumption allows us to model T_1 and T_2 sequentially.

Because a covariate x will in general influence both the $0 \rightarrow T_1$ and $T_1 \rightarrow T_2$ transitions, one can show that x will have both direct and indirect effects on the probability that a woman has a premarital first birth, with the direct effect of x corresponding how x affects premarital birth risks following onset and the indirect effect of x corresponding to whether x delays or hastens onset and thus indirectly influencing the probability of a birth via the duration of exposure to risk. To provide some intuition into why this might be so, we begin by considering the transition to sexual activity. A first issue is that in a hazard model setting, there will be an implied distribution for the timing of any event even for a completely homogeneous population. This differs from a linear regression setting for an outcome y, where in a homogeneous population there will be a single expected value, E[Y]. Because of this, because some women may not have initiated sexual activity by interview, and because some woman may never become sexually active, it is more natural to focus on percentiles of the T_1 distribution than the expectation of T_1.

Figure 1 illustrates issues when the selected percentile is the median. Consider a following a hypothetical sample of women from intact and nonintact families as they initiate sexual activity. As noted above, the percentage of women who have initiated sexual activity by age t is given by the expression in (5). Because a standard finding is that women from intact families initiate sexual activity at somewhat later ages than those from nonintact families, Figure 1 shows differences at every age between the solid and dashed curves, with higher proportions initiating sexual activity for those in nonintact families comparedto those in intact families. Figure 1 also shows how one can obtain the predicted median age at onset for each group, which then also provides the estimated difference in median age at onset for the two groups.

[Figure 1 about here]

As noted above, abstinence proponents argue that delaying the onset of sexual activity will, all else being equal, lead to fewer teen and premarital first births by virtue of decreased exposure to risk. The two graphs in Panel A of Figure 2 illustrates what is implied when women from nonintact families have a median age at onset of t_1 while those from intact families have a later median age at onset of $t_1 + \Delta$. The shaded areas depict the integrals in the age baseline in (6) and show how risks cumulate with age for women in intact and nonintact families following onset. Because women in intact families delay onset relative to women from nonintact families, they have less exposure to the risk of a premarital first birth and hence lower cumulative risks. Then all else being equal, delayed onset implies lower cumulative risks, which in turn implies a lower probability of a premarital birth by age t, with the magnitude of this difference given by (6).

[Figure 2 about here]

In Panel A, we implicitly assumed that the only difference between women in intact and nonintact families was in age at onset; hence, the curves in Panel A, which represent a woman's age-specific risks of a premarital first birth following onset, are identical for women who grew up intact or nonintact families. However, a covariate that delays or hastens onset will often also influence the risk of a premarital first birth following onset. Panel B of Figure 2 illustrates this possibility by supposing that women from nonintact families have higher risks following onset than do women from intact families, with these higher cumulative risks implying an even higher probability of a premarital first birth for those who grew up in a nonintact family. Thus, Figure 2 illustrates two ways in which covariates can influence the probability of a premarital first birth, a first following from earlier or delayed onset of sexual activity and a second following from lower or higher birth risks following onset.

Finally, note that (4) includes t_1 as a ordinary right-hand-side covariate. This parallels the logic of status attainment models (see, e.g., Featherman and Hauser 1978) in which family of origin affects both an individual's years of schooling completed and occupational attainment, and in which years of schooling completed also affects occupational attainment.[4] Substantively, one might include t_1 as an ordinary right-hand-side variable in the $T_1 \rightarrow T_2$ equation if T_1 were thought to have a causal effect on T_2 or if T_1 were to be correlated with unobserved covariates that in turn affect T_2.

To summarize, under our sequential model, a covariate x will have both direct and indirect

[4] More formally, when modeling T_2, one can condition on any relevant aspect of an individual's past history, including the timing of the event T_1 (Aalen 1978; Tuma and Hannan 1984).

effects on the probability of a premarital first birth. We trace the influence of x on T_1 timing by examining selected percentiles of the predicted T_1 distribution, as illustrated in Figure 1. Under (4), our model for the $T_1 \rightarrow T_2$ transition, x will influence the probability of a premarital first birth in expression (6) in three ways: (i) an indirect effect in which x alters a woman's duration of exposure to risk, as depicted in both the upper and lower panels of Figure 2; (ii) a second indirect in which x affects T_2 risks through t_1 when t_1 is included as an ordinary right-hand-side covariate in the $T_1 \rightarrow T_2$ equation in (4); and (iii) a direct effect of x when x is specified as a ordinary right-hand-side covariate in (4), as depicted in the lower panels of Figure 2.

DATA

We use data from the 1979 National Longitudinal Survey of Youth (NLSY), a household-based national probability sample of persons aged 14-21 in 1979. The original 12,686 cases consist of a main sample of 6,111 respondents, an oversample of 5,295 minorities and poor whites, and a sample of 1,280 Armed Forces personnel. The military sample was suspended in 1985, with 1,079 (out of the original 1,280) cases affected. Retention has been high in the NLSY, with for example, 10,485 (90.3 percent) Of the 11,607 non-military respondents reinterviewed in the 1987 wave, for a retention rate of 98.8 percent.

Of the 6,283 women present at the initial 1979 interview, we excluded women: (1) with missing data on race and ethnicity (n = 45); (2) who reported not knowing their biological mother (n = 9); (3) with missing data on the timing of first menstruation (n = 254); (4) with missing data on number of siblings (n = 9); or (5) with missing first intercourse, first birth, first marriage histories (n = 371). These selection criteria yielded a sample of n = $5,595$ women.

Data on age at first sexual intercourse were obtained in the 1984–1986 interviews, when all respondents were at least 18 years old. In the 1984 wave, age at first intercourse was obtained to the nearest year. In the 1985 wave, questions on the calendar month and year of menarche and first sexual intercourse were administered to all female respondents; these questions were repeated in 1986 for 1985 female nonrespondents. We computed the young woman's age in months at first premarital sexual intercourse using data from the 1985 and 1986 waves, using a hot-deck procedure to impute missing calendar month at first sexual intercourse. Wu, Martin, and Long (2001) find that these self-reports are of reasonable quality, with comparisons of these data in close agreement with data on sexual onset for a comparable birth cohort of women from the 1995 National Survey of Family.

For women who report never having engaged in sexual activity, we censored their first sexual intercourse history at their age at interview in 1985 or 1986, depending on the year in which they were asked the question. We likewise censored women's first sexual intercourse history at their age at first marriage if they reported that they had initiated sexual intercourse on or after the date of first marriage. We similarly censored a woman's premarital birth history at either her age at last interview or at her age at first marriage if she did not report a first birth prior to last survey observation or first marriage.

RESULTS

Figure 3 presents smoothed nonparametric estimates using a procedure described in Wu (1989) for the age-graded risk of entry into sexual activity, the age-graded risk of a premarital first birth, and the duration-graded risk of a premarital first birth conditional on entry into sexual activity. The top panel of Figure 3 plots smoothed nonparametric estimates of the logarithm of the hazard rate of first sexual intercourse by age, the middle panel plots two different estimates of the logarithm of the hazard rate for a premarital first birth, and the bottom panel plots estimates of the logarithm of the hazard rate for a premarital first birth by duration since sexual onset. In the upper two panels, the curves for the logarithm of the rate rise in a roughly linear fashion to about age 18.5, after which the curves decline, again in a roughly linear fashion.

[Figure 3 about here]

In the middle panel of Figure 3, the two curves differ in the assumptions they make about when women become at risk of a premarital first birth. The solid curve presents estimates that do not place a woman at risk of a premarital first birth until she reports becoming sexually active; hence, for this curve, we use a woman's report of age at first intercourse to left-truncate her premarital birth history. The dotted curve presents estimates that ignore this left truncation; hence, while this curve can be viewed as the average of the logarithm of premarital first birth risks in the population, it ignores variation in onset of sexual activity and implicitly assumes that women are at risk of a premarital first birth even if they have not initiated sexual activity, an implausible assumption.

A comparison of the two curves in the lower panel of Figure 3 shows that left truncation affects estimates substantially, with the curve ignoring left truncation systematically underestimating premarital first birth risks relative to the curve that incorporates left truncation. Differences between these two curves are especially apparent at younger ages, reflecting the tendency for premarital births risks to be especially high for teen women in the period following the initiation of sexual activity.

The nonparametric estimates in the bottom panel of Figure 3 exhibit a non-monotonic pattern of duration dependence in which premarital first birth risks first rise and then decline. Based on these nonparametric results, we model age dependence in both the T_1 and T_2 equations using a splined piecewise Gompertz specification with nodes at ages 16, 18.5 and 20 (e.g. Wu and Tuma 1990, Lillard 1993). For the T_2 equation, we modeled duration dependence using a piecewise constant specification for durations 0 to 6, 7 to 14, 14 to 35, and 36+ months. Estimates from these models are presented in Table 1.

[Table 1 about here]

The first two columns in Table 1 adopt a conventional approach to modeling premarital first birth risks by examining women's age-specific risks of a premarital first birth but ignoring the timing of first sexual intercourse. We present estimates from two proportional hazard specifications, the Cox proportional hazard model and a piecewise splined Gompertz model with proportional effects of covariates. Estimates from these models reveal substantially higher relative risks for blacks compared to whites, but no significant difference in relative risks for white and Hispanic women. The next four columns present corresponding estimates for the transition to first sexual intercourse and the transition to a premarital first birth conditional on entry into sexual activity. Compared to white women, black women have significantly higher risks of first sexual intercourse (corresponding to earlier ages at onset) as well as significantly higher premarital first birth risks following onset. However, the Hispanic/white contrasts are opposite in sign for the two transitions, with significantly *lower* risks of first sexual intercourse but significantly *higher* premarital first birth risks following onset for Hispanic women relative to white women,

Results for family structure, religion, and ability are reported in the next three rows of Table 1. These associations show qualitative agreement between approaches, with the signs and significance levels similar for estimated coefficients of the risks for the unconditional transition to a premarital first birth, to first intercourse, and to a premarital first birth conditional on sexual initiation.

The next several rows present estimated coefficients for mother's education, number of siblings, and income-to-needs. All three variables have associations in the expected directions with unconditional premarital first birth risks, small and statistically insignificant associations with age at first intercourse, and associations in the expected directions with premarital first birth risks conditional on sexual onset. Thus, our results suggest that conclusions obtained from our sequential approach can yield qualitatively different insights than those obtained from a more conventional approach.[5]

The results in Table 1 also show close agreement between estimates the Cox and piecewise splined Gompertz specifications. As noted above, our decomposition derivations require explicit estimates of the various baseline hazards, which are not easily obtained from a Cox specification; hence, we henceforth restrict our discussion to estimated coefficients from the piecewise splined Gompertz models.

We now turn to results for selected decompositions. Table 2 presents decomposition results comparing black and white women. Predicted median ages at onset of sexual activity are reported in Panel A of Table 2 and are calculated using the estimated coefficients in column 2 of Table 2,

[5]Table 1 also reports estimated coefficients for a time-varying dummy variable equal to one at all ages after first menses. In our sequential model, we specify this variable only when modeling onset because the baseline counterfactual for premarital birth risks following onset are women who have initiated sexual activity but who have not yet reached sexual maturity and who thus face negligible birth risks (and who are very rare in these data).

with other covariates set to their sample means. The values of the predicted medians are 17.55 and 17.73 (210.6 and 212.7 months) for black and white women, respectively. The resulting difference, while in the expected direction, is thus relatively small, corresponding to the black coefficient (.10) in Table 1. Although observed black/white differences in age at first sexual intercourse are larger, these differences do not control for other variables; thus, our results suggest that much of the unconditional difference in age at onset of sexual activity can be attributed to the association of variables other than race on women's age at onset of sexual activity.

[Table 2 about here]

As noted above, decomposition results will vary with duration of exposure; hence, Panel B presents results for 60 and 90 months. Panel B reports the arithmetic difference in the predicted percentage of premarital first births, obtained using the expression in (18) and using the estimated coefficients in columns 4 and 6 of Table 1. As expected, there are substantial differences in prevalence, even holding constant other variables, with a predicted black/white difference of 11.7 and 13.2% in the percentage of women having a premarital birth at 60 and 90 months, respectively, following sexual onset. As noted above, the derivations of the previous show that the 11.7 and 13.2 coefficients can be decomposed into three components, a direct component (labeled "D"), corresponding to the estimated coefficients in columns 4 and 6 of Table 1, and two indirect components, one corresponding to the estimated right-hand-side coefficient for age at first intercourse in Table 1 ("E") and a second due to black/white differences in exposure to risk ("F"). Thus at 60 months of exposure, the predicted black/white difference of 11.7% can be decomposed into a direct effect of 10.3% and two indirect effects of .3 and 1.0%, respectively. This shows that the direct effect is substantially larger than either of the two indirect effects, which is in qualitatively agreement with the estimated coefficients in Table 1 (black/white coefficient of .11 for onset of sexual activity and .75 for premarital first birth risks conditional on onset).[6]

[6] Because a premarital first birth and first marriage are competing risks, our predicted probabilities at 60 and 90 months of duration should be interpreted under the counterfactual in which women cannot marry during these durations of exposure. This counterfactual is substantively most appropriate when the variables examined in our decompositions do not have a strong association with the competing risk of first marriage.

Table 3 presents parallel decompositions for Hispanic and white women, with all other covariates set to their sample means. Recall that the parallel Hispanic/white coefficients in Table 1 for the piecewise splined Gompertz model were negative and significant for age at first sexual (−.37), but positive and significant for premarital first births conditional on age at onset (.53). The differences in predicted median ages at onset of sexual activity correspond to the −.37 coefficient in Table 1, with predicted values of 18.65 and 17.73 (223.8 and 212.7 months) for Hispanic and white women, respectively

[Table 3 about here]

Because the Hispanic/white contrasts in Table 1 take opposite signs, the sign of the arithmetic difference in Table 3 could in principle be positive or negative, depending on the magnitude of estimated coefficients in Table 1. Our empirical results show lower prevalence for Hispanic women relative to their white counterparts (−6.9 and −4.9% at 60 and 90 months of exposure, respectively), conditional on onset of sexual activity and after setting all other covariates to their sample means. The magnitude of these differences is roughly half that of the corresponding black/white differences in the decompositions in Table 2.

It is informative to contrast the above results with the black/white (.88) and Hispanic/white (.14) coefficients in Table 1 for the unconditional transition to a premarital first birth. That is, adopting a conventional approach that examines a woman's age at a first premarital birth but ignores the timing of first sexual intercourse suggests large black/white differences but small Hispanic/white differences in the percentage with a premarital first birth, holding other covariates. If, however, premarital first birth risks are assumed to be negligible prior to onset of sexual activity, our results suggest more premarital first births to blacks relative to whites (about 12 or 13% for 60 and 90 months of exposure), but *fewer* premarital first births to Hispanics relative to whites (between 5 and 7% for 60 and 90 months of exposure). These comparisons show that the results from our sequential model generate insights that are qualitatively different from more conventional

These cautions affect possible interpretations of our results, an issue especially important for our black/white decompositions.

approaches.

Table 4 present decomposition results for AFQT. Recall that AFQT had significant effects in Table 1 for both onset of sexual activity ($-.14$) and premarital first birth risks given onset ($-.38$). In these decompositions, we compare women with low and high AFQT scores, defined as a score half a standard deviation below or above the mean. Panel A of Table 4 shows that varying AFQT in this way corresponds to just under a 3 month difference in the predicted median age at first sexual intercourse (215.1 vs. 213.2 months). Differences in prevalence, even holding constant other variables, are 8.5 and 9.5% for the percentage of women having a premarital birth at 60 and 90 months, respectively, following sexual onset. These correspond to direct and indirect effects of 6.7 (direct), 0.3, and 1.4% (indirect) at 60 months following sexual onset and 7.9, 0.4, and 1.3% at 90 months following sexual onset. Thus, these decompositions show substantially smaller indirect effects of AFQT on premarital birth risks, and a far larger direct effect, holding constant all other variables.

[Table 4 about here]

Table 5 present decomposition results for women from intact and nonintact families at age 14. In Table 1, results from the piecewise splined Gompertz specification were that residing in a nonintact family at age 14 was associated with a .34 higher risk of onset of sexual activity and a .26 higher risk of a premarital first birth conditional on onset of sexual activity. Panel A of Table 1 shows that these results yield a predicted difference in the median age at first sexual intercourse of 7 months (209.6 vs. 216.6). The corresponding differences in the probability of a premarital first birth are 15.8 and 17.7% for 60 and 90 months of exposure following sexual onset, respectively, holding all other covariates at their sample means. The decompositions for the 15.8% difference at 60 months of exposure show that the largest portion comes from the direct effect (11.1%), with the next largest portion stemming from the indirect effect of differential exposure (3.8%). The results for 90 months of exposure are similar, with the largest portion of the overall 17.7% difference stemming from the direct effect (13.3%) and far smaller portions from the two indirect effects (3.4% from the indirect exposure effect and 1.0% from the indirect right-hand-side covariate effect of age at onset). Thus, these decomposition results show that direct effects dominate indirect effects even though the relative risks for nonintact family structure in Table 1 are larger for first sexual intercourse than for premarital first births.

[Table 5 about here]

[Paragraphs not written. Tables 6–7 present parallel decompositions for mother's education, and income-to-needs in the woman's family of origin; Table 8 presents a summary of the decomposition results.]

[Tables 6–8 about here]

DISCUSSION

In this paper, we have proposed a sequential hazard model of premarital first births in which we model women's entry into sexual activity and her subsequent risk of a premarital first birth conditional on age at onset of sexual activity. This sequential model differs from more conventional approaches by supposing that the typical woman is not at risk of a premarital first birth until she becomes sexually active. Although highly stylized in that sexual activity will vary in intensity and frequency following first intercourse, this sequential approach nevertheless highlights key periods during which premarital first birth risks can be expected to vary substantially. It also follows a long demographic tradition that holds that better approximating durations of exposure to risk is a central task in understanding demographic phenomena such as fertility.

Does our sequential hazard model yield insights different from those obtained from a more conventional approach? The answer is yes. Our empirical results suggest numerous examples in which the effects of covariates on the transition into sexual activity and to a premarital first birth conditional on onset are close to zero for one transition but substantial and statistically significant for the other transition, or in which coefficients are substantial in magnitude and statistically significant but opposite in sign.

Following Wu and Martin (2009), we also decompose the difference in the probability of a premarital birth into direct and indirect components of covariates. That is, black/white differences in the probability of a premarital first birth can reflect, for example, a direct effect reflecting higher premarital first birth risks in the period following sexual initiation for blacks relative to whites. But black/white differences in the probability of a premarital first birth can also result from an indirect effect reflecting, for example, earlier entry into sexual activity among black women, which in turn will imply longer durations of exposure to risk for blacks relative to whites.

Our decompositions provide additional insights not easily obtained from a simple inspection of estimated coefficients in our sequential hazard model. In particular, the results from our decompositions followed a pattern in which direct effects typically outweigh indirect effects, a finding that holds for 6 of the 7 decompositions we examined. For most social scientists, this result will not be surprising given that proximate determinants are typically more influential than more distal determinants. Nevertheless, recent policies targeting teen and nonmarital fertility have often assumed that delaying sexual activity or encouraging abstinence will produce substantial reductions in these outcomes; conversely, policies targeting how premarital first birth risks might be reduced in the period following initiation of sexual activity have been pursued far less aggressively, at least in recent years.

Although our empirical results provide no firm causal estimates of the effects of covariates on either sexual initiation or premarital first birth risks, they nevertheless generally run counter to arguments that delaying sexual onset will lead to substantial reductions in nonmarital fertility. The one exception is our comparison of white and Hispanic women. Our hazard regressions indicate a substantially later onset, net of the other covariates in our models, for Hispanic women than for white women, but also substantially higher premarital first birth risks following onset for Hispanics relative to whites. We estimate that this cohort of Hispanic women had a median age at onset that was roughly 12 months later than that for white women. Our decompositions suggest that a delay of this magnitude would lower the probability of a premarital first birth substantially relative to whites, but that roughly half of this reduction offset by the higher premarital first birth risks following onset for Hispanic women.

A distinct advantage of our sequential hazard framework and our decomposition methods is that it both mirrors closely the behavioral hypotheses of abstinence proponents and critics and provides estimates of hypothesized effects, thus letting us assess competing claims. More generally, our analyses represent a first step toward reconceptualizing premarital first births in terms of a series of underlying behaviors, and then asking what factors might influence these behaviors. We have proceeded in a highly stylized way by examining what might influence when women become sexually active and what might influence premarital first birth risks following the initiation of sexual activity. This approach, we argue, improves substantially upon past research by recognizing that the typical woman has negligible birth risks prior to becoming sexually active. Nevertheless, our approach, like past research, proceeds quite crudely with respect to a number of fertility-related behaviors following the initiation of sexual activity, including the frequency of sexual activity following onset; contraceptive knowledge and use and, conditional on use, effort, method choice, and contraceptive efficacy; changes in fecundability with age and with exposure to possible health-related impairments to fecundability; the formation (and dissolution) of romantic relationships and cohabiting unions; the occurrence of a nonmarital pregnancy and, conditional on such a pregnancy, the factors influencing how such a pregnancy might be resolved.

APPENDIX

Let t_{1p} denote the pth percentile of T_1 and let $\pi = 1 - (p/100)$; then

$$\begin{aligned} p/100 &= 1 - S_1(t_{1p}|\mathbf{x}) \\ 1 - (p/100) = \pi &= S_1(t_{1p}|\mathbf{x}) \\ &= \exp\left[-\exp(\mathbf{ax})\int_0^{t_{1p}} q_1(s)\, ds\right] \\ &= \exp[-\exp(\mathbf{ax})Q_1(t_{1p})] \\ Q_1(t_{1p}) &= -\log(\pi)/\exp(\mathbf{ax}) \\ t_{1p} &= Q_1^{-1}[-\log(\pi)/\exp(\mathbf{ax})] \end{aligned} \tag{A1}$$

Consider two groups, A and B, with covariates $\mathbf{x_A}$ and $\mathbf{x_B}$; then from (7), we have

$$\begin{aligned} \Delta t_{1p} &= t_{1p}^B - t_{1p}^A \\ &= Q_1^{-1}[-\log(\pi)/\exp(\mathbf{ax}_B)] - Q_1^{-1}[-\log(\pi)/\exp(\mathbf{ax}_A)]\,, \end{aligned} \tag{A2}$$

$$\begin{aligned} t_{1B} &= t_1 + \Delta t_{1p} \\ u_{1B} &= t - t_{1B} = t - (t_1 + \Delta t_{1p}) = (t - t_1) - \Delta t_{1p} = u - \Delta t_{1p} \end{aligned}$$

$$\begin{aligned} P_{2A} - P_{2B} &= 1 - S_{2A} - (1 - S_{2B}) \\ &= S_{2B} - S_{2A} \\ &= S_2(t, u_B|t_{1B}, \mathbf{x}_B)] - S_2(t, u|t_1, \mathbf{x}_A) \\ &= \exp[-H_2(t, u_B|t_{1B}, \mathbf{x}_B)] - \exp[-H_2(t, u|t_1, \mathbf{x}_A)] \\ &= \exp\left[-\exp(\alpha[t_1 + \Delta t_{1p}] + \mathbf{bx}_B)\int_{t_1+\Delta t_{1p}}^{t} q_{21}(s)\, ds \int_0^{u-\Delta t_{1p}} q_{22}(v)\, dv\right] \\ &\quad - \exp\left[-\exp(\alpha t_1 + \mathbf{bx}_A)\int_{t_1}^{t} q_{21}(s)\, ds \int_0^{u} q_{22}(v)\, dv\right]. \end{aligned} \tag{A3}$$

To gain some intuition into the above, consider two groups, A and B, with covariates $\mathbf{x_A}$ and $\mathbf{x_B}$; then let

$$\begin{aligned} t_{1B} &= t_1 + \Delta t_{1p} \\ u_{1B} &= t - t_{1B} = t - (t_1 + \Delta t_{1p}) = (t - t_1) - \Delta t_{1p} = u - \Delta t_{1p} \end{aligned}$$

Note that the ratio of the two cumulated hazards is given by:

$$\frac{H_{2B}(t)}{H_{2A}(t)} = \frac{\int_{t_{1B}}^{t} \int_{0}^{u_B} q_{21}(s)\; q_{22}(v) \exp[\alpha t_{1B} + b_1(x_1 + \Delta) + \cdots]\, ds\, dv}{\int_{t_1}^{t} \int_{0}^{u} q_{21}(s)\; q_{22}(v) \exp[\alpha t_1 + b_1 x_1 + \cdots]\, ds}$$

$$= \frac{\exp[\alpha(t_1 + \Delta t_{1p}) + b_1(x_1 + \Delta) + \cdots] \int_{t_1+\Delta t_{1p}}^{t} q_{21}(s) \; \int_{0}^{u-\Delta t_{1p}} q_{22}(v)}{\exp[\alpha t_1 + b_1 x_1 + \cdots] \int_{t_1}^{t} q_{21}(s) \; \int_{0}^{u} q_{22}(v)}$$

$$\begin{aligned} \frac{H_{2B}(t)}{H_{2A}(t)} &= \frac{\exp[\alpha(t_1 + \Delta t_{1p}) + b_1(x_1 + \Delta) + \cdots] \int_{t_1+\Delta t_{1p}}^{t} q_{21}(s) \; \int_{0}^{u-\Delta t_{1p}} q_{22}(v)}{\exp[\alpha t_1 + b_1 x_1 + \cdots] \int_{t_1}^{t} q_{21}(s) \; \int_{0}^{u} q_{22}(v)} \\ &= \frac{\exp[\alpha(t_1 + \Delta t_{1p})]}{\exp[\alpha t_1]} \times \frac{\exp[b_1(x_1 + \Delta) + \cdots]}{\exp[b_1 x_1 + \cdots]} \times \text{[ratio of integrals]} \\ &= \exp(\alpha \Delta t_{1p}) \exp(b_1 \Delta)\text{[ratio of integrals]} \end{aligned} \qquad (A4)$$

where

$$\exp(b_1\Delta) = \text{direct effect of } \mathbf{x}$$

$$\exp(\alpha\Delta t_{1p}) = \text{RHS indirect effect of } \mathbf{x}$$

$$\text{ratio of integrals} = \text{indirect exposure effect of } \mathbf{x}$$

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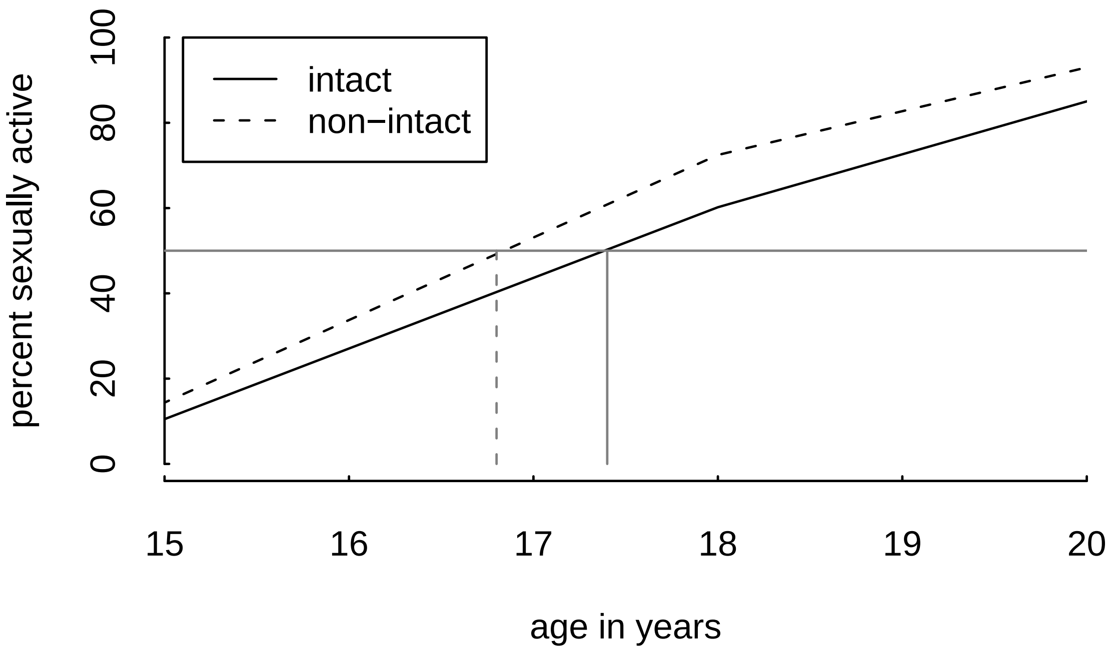

Figure 1: Hypothetical effect of non-intact family structure on the predicted median timing of first sexual intercourse.

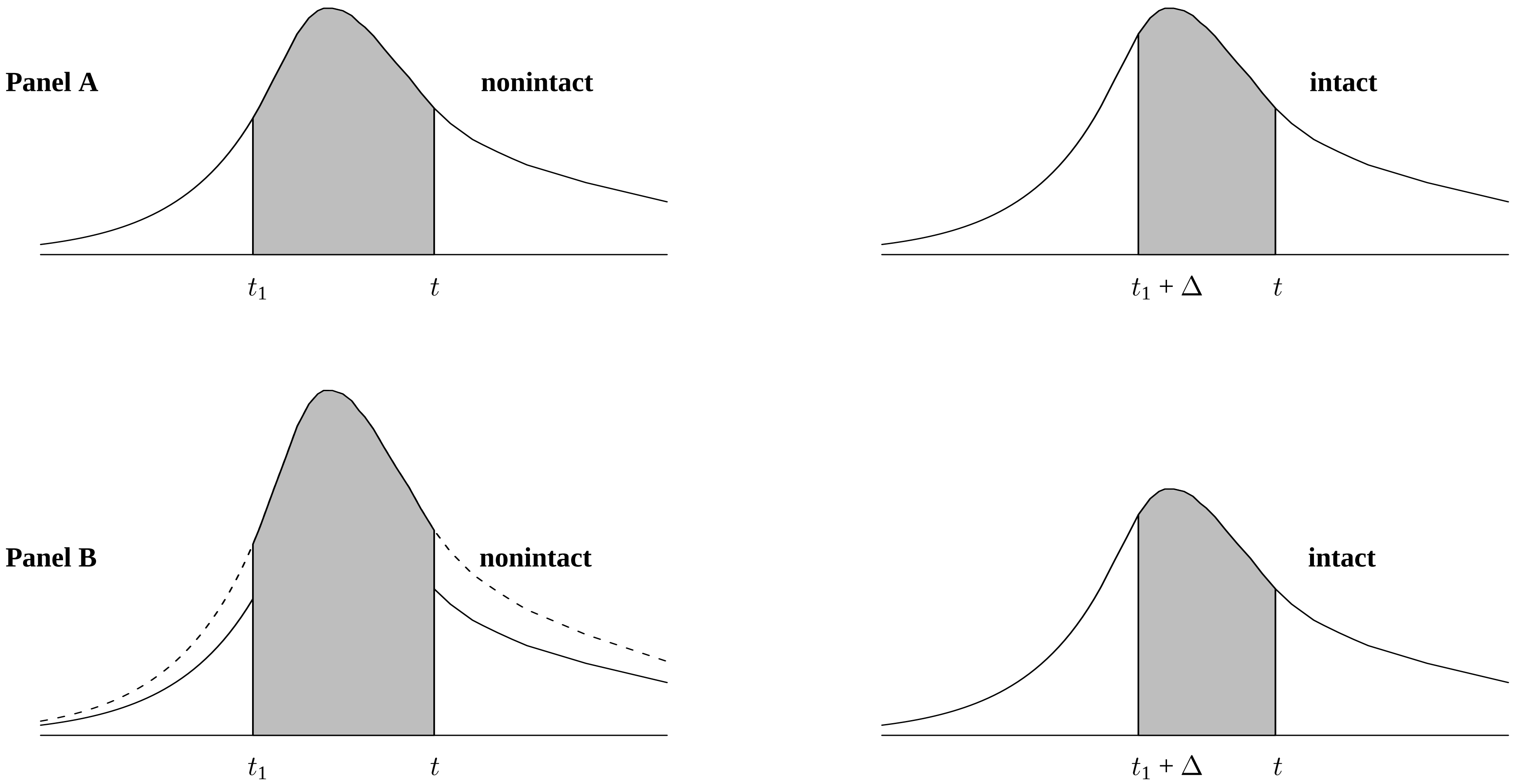

Figure 2. Panel A: Hypothetical example of how earlier onset increases cumulative risk for women from nonintact relative to women from intact families. **Panel B:** Hypothetical example of how earlier onset and higher risks following onset increases cumulative risk for women from nonintact relative to women from intact families. In both Panel A and B, the cumulative risk of a premarital first birth is given by the area in the shaded region of each curve, with onset of sexual activity at age t_1 for women from nonintact families and delayed onset at age $t_1 + \Delta$ for women from intact families.

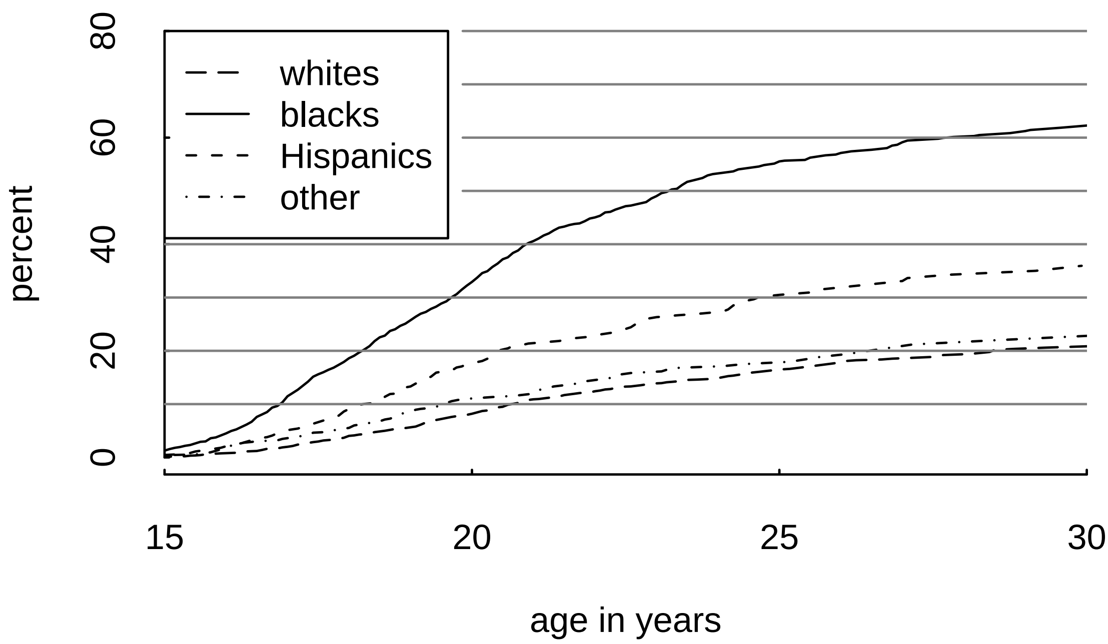

Figure 3: Probability of a premarital first birth, by race and ethnicity.

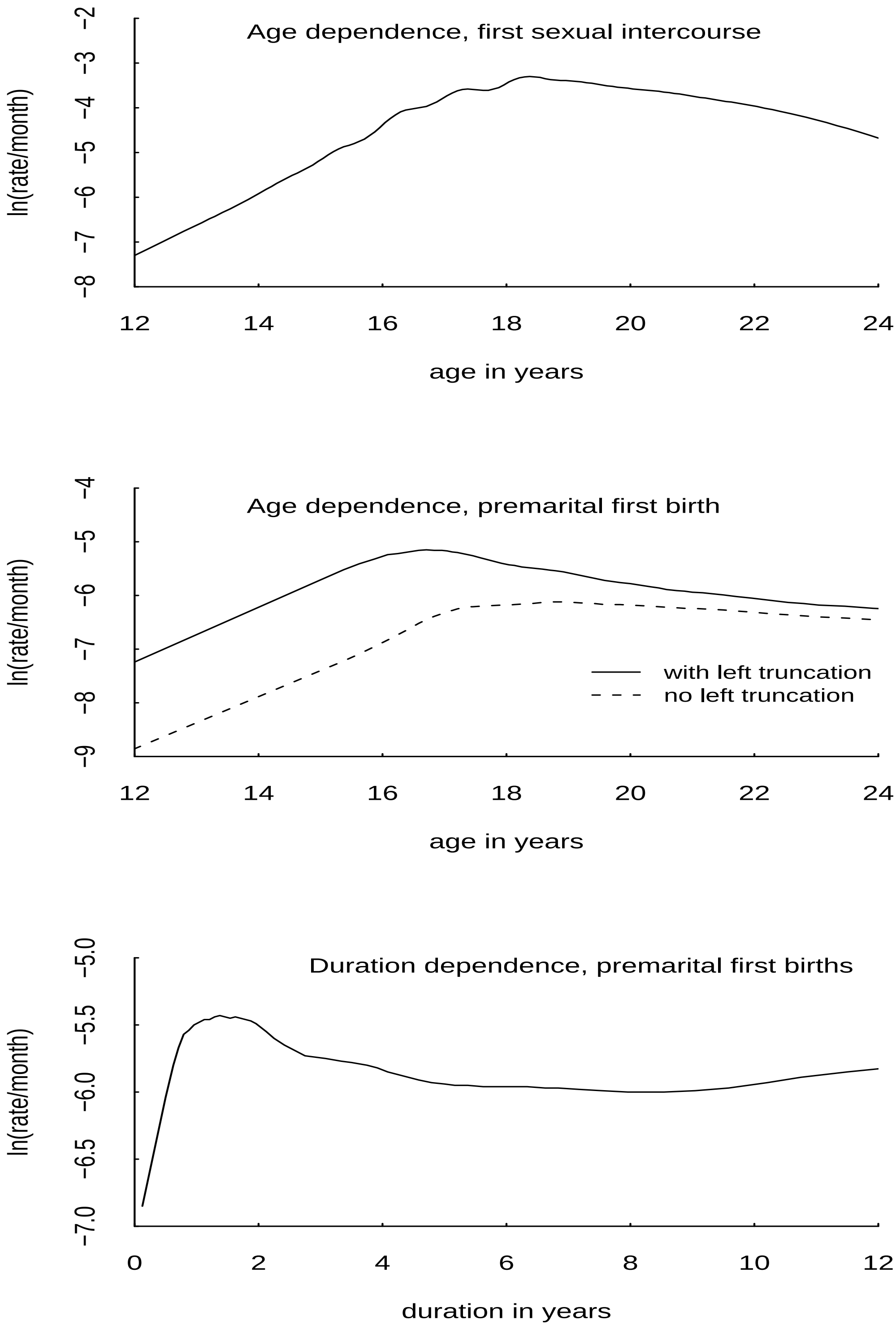

Figure 4: Smoothed nonparametric estimates of the logarithm of the hazard for: (a) age dependence in the transition to first sexual intercourse, (b) age dependence in the transition to a premarital first birth, and (c) duration dependence in the transition from sexual onset to a premarital first birth.

Table 1. Estimated coefficients from Cox and piecewise splined Gompertz proportional hazard models for: (a) the unconditional transition to a premarital first birth; (b) the transition to first sexual intercourse; and (c) the transition from onset of sexual activity to a premarital first birth.

	Unconditional transition to a premarital first birth		Transition to first sexual intercourse		Transition to a premarital first birth given sexual onset	
	Cox	Gmp	Cox	Gmp	Cox	Gmp
Race and ethnicity						
black	.88***	.88***	.11*	.11*	.76***	.75***
	(.09)	(.09)	(.05)	(.05)	(.09)	(.09)
Hispanic	.14	.14	−.37***	−.37***	.53**	.53**
	(.16)	(.16)	(.08)	(.08)	(.16)	(.16)
other	.07	.06	.05	.05	.00	.01
	(.12)	(.12)	(.04)	(.04)	(.12)	(.12)
Family background						
mother's education	−.05***	−.05***	−.01	−.01	−.05***	−.05***
	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
number of siblings	.06***	.06***	.01	.01	.06***	.06***
	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
mother's age at first birth	−.05***	−.05***	−.03***	−.03***	−.04***	−.04***
	(.01)	(.01)	(.00)	(.00)	(.01)	(.01)
nonintact family at age 14	.62***	.62***	.31***	.31***	.34***	.34***
	(.07)	(.07)	(.03)	(.03)	(.07)	(.07)
catholic	−.20*	−.20*	.02	.02	−.19*	−.20*
	(.09)	(.09)	(.03)	(.03)	(.09)	(.09)
reading materials	−.12**	−.12**	−.03	−.03	−.12**	−.12**
	(.04)	(.04)	(.02)	(.02)	(.04)	(.04)
Ability						
AFQT	−.48***	−.49***	−.15***	−.15***	−.44***	−.45***
	(.04)	(.04)	(.02)	(.02)	(.04)	(.04)

Table 1. (continued)

	Unconditional transition to a premarital first birth		Transition to first sexual intercourse		Transition to a premarital first birth given sexual onset	
	Cox	Gmp	Cox	Gmp	Cox	Gmp
Sexual maturation and age at first intercourse						
menstruation (time-varying dummy variable)	1.15* (.56)	1.12* (.51)	1.06*** (.12)	.98*** (.11)		
age (in months) at first first intercourse					−.006* (.002)	−.008*** (.002)
Duration dependence						
0 to 6 months					−.92*** (.15)	−.92*** (.15)
14 to 35 months					−.14 (.11)	−.19 (.11)
36 months or more					−.09 (.15)	−.27 (.15)

All models also include dummy variables for missing values of: mother's education, mother's age at first birth, family structure at age 14, reading materials, AFQT, and calendar month of first sexual intercourse. See text for additional details.

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

Table 2. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of black and white women, holding other covariates at their respective means.

Panel A	age in months
Predicted median age at first intercourse	
non-Hispanic blacks	213.0
non-Hispanic whites	215.4

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: blacks	23.6	30.7
B: whites	11.0	15.0
Predicted difference, premarital first birth		
C: A − B	12.4	15.7
Decomposition of C		
D: direct component	11.3	14.5
E: indirect component, differential age at onset	.3	.4
F: indirect component, differential exposure	.8	.7

Table 3. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of Hispanic and white women, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
Hispanics	224.1
whites	215.4

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: Hispanics	15.6	21.5
B: whites	11.0	15.0
Predicted difference, premarital first birth		
C: A − B	4.6	6.5
Decomposition of C		
D: direct component	6.6	8.7
E: indirect component, differential age at onset	−0.9	−1.2
F: indirect component, differential exposure	−1.1	−1.0

Table 4. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of white women and women from other race and ethnicities, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
other race/ethnicity	214.2
white	215.4

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: other race/ethnicities	11.4	15.5
B: whites	11.0	15.0
Predicted difference, premarital first birth		
C: A − B	.4	.5
Decomposition of E		
D: direct component	.1	.1
E: indirect component, differential age at onset	.1	.1
F: indirect component, differential exposure	.3	.2

Table 5. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of women whose mothers had 10 and 14 years of completed schooling, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
mother's education = 10 years	215.0
mother's education = 14 years	215.8

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: mother's education = 10 years	13.5	18.4
B: mother's education = 14 years	11.0	15.1
Predicted difference, premarital first birth		
C: A – B	2.5	3.3
Decomposition of E		
D: direct component	2.4	3.1
E: indirect component, differential age at onset	.1	.1
F: indirect component, differential exposure	.1	.1

Table 6. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of [number of siblings], holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
number of siblings − .5 sd	???.?
number of siblings − .5 sd	???.?

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: number of siblings − .5 sd	??.?	??.?
B: number of siblings + .5 sd	??.?	??.?
Predicted difference, premarital first birth		
C: A − B	?.?	?.?
Decomposition of E		
D: direct component	?.?	?.?
E: indirect component, differential age at onset	.?	.?
F: indirect component, differential exposure	.?	.?

Table 7. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of women with mothers who had a first birth at early and later ages, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
mother's age at first birth = 20	214.2
mother's age at first birth = 25	217.1

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: mother's age at first birth = 20	13.6	18.5
B: mother's age at first birth = 25	10.6	14.5
Predicted difference, premarital first birth		
C: A − B	3.1	3.9
Decomposition of E		
D: direct component	2.3	3.1
E: indirect component, differential age at onset	.3	.4
F: indirect component, differential exposure	.5	.4

Table 8. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of women from nonintact and intact families, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
nonintact	210.6
intact	217.5

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: mother's education = 10 years	18.2	23.6
B: mother's education = 14 years	11.2	15.2
Predicted difference, premarital first birth		
C: A − B	7.0	8.4
Decomposition of E		
D: direct component	4.5	5.9
E: indirect component, differential age at onset	.8	1.0
F: indirect component, differential exposure	1.6	1.6

Table 9. Decomposition of the difference in the percentage of women predicted to have a premarital first birth: comparison of women with AFQT scores half a standard deviation below and above the mean, holding other covariates at their respective means.

Panel A	age in months
Predicted median, age at first intercourse	
mean AFQT −.5 s.d.	212.2
mean AFQT +.5 s.d.	215.1

Panel B	months of exposure	
	60	90
Predicted percentage, premarital first birth		
A: mean AFQT −.5 s.d.	24.0	28.7
B: mean AFQT +.5 s.d.	15.5	19.2
Predicted difference, premarital first birth		
C: A − B	8.5	9.5
Decomposition of E		
D: direct component	6.7	7.9
E: indirect component, differential age at onset	.3	.4
F: indirect component, differential exposure	1.4	1.3

Table 10. Summary of total, direct, and indirect effects. Decompositions of the difference in the percentage of women predicted to have a premarital first birth, holding other covariates at their respective means and with effects evaluated 60 months after onset of sexual activity for the baseline group,

	total	direct	indirect	
			onset age	exposure
black vs. white	12.4	11.3	.3	.8
AFQT $\mp$ .5 sd	7.0	5.8	.4	.9
nonintact vs. intact	7.0	4.5	.8	1.6
Hispanic vs. white	4.6	6.6	−.9	−1.1
early vs. late mother's age at first birth	3.1	2.3	.3	.5
low vs. hi reading materials	2.9	1.7	.1	.2
low vs. hi mother's education	2.5	2.4	.1	.1
non-Catholic vs. Catholic	2.2	2.2	−.0	−.0
other vs. white	.4	.1	.1	.3