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ABSTRACT

Does a nonmarital birth ignite an “engine” of subsequent nonmarital fertility? Or does a marital first birth serve as a “brake” on subsequent nonmarital fertility? To answer these questions, we compare the pace of subsequent marital and nonmarital childbearing, given a marital or nonmarital first birth, for currently married versus currently unmarried women. We employ parity-specific, competing-risk hazard models that specify the duration, age, and period structure of higher-order marital and nonmarital births occurring between 1970 and 1995, using retrospective marital and fertility histories from the June 1990 and 1995 Current Population Surveys. Our findings provide little support for an “engine” hypothesis. We find that the interval between a nonmarital first and nonmarital second birth is significantly *longer* than the interval between a marital first and marital second birth for both white and black women, a result that holds across observed demographic characteristics for these women. For higher-order births, our results also run counter to expectations, with some exceptions. For black women, the interval between a nonmarital second and third birth or between a nonmarital third and fourth birth is generally longer than the corresponding intervals between successive marital births, although these differences are rarely statistically significant. Results for white third births follow a similar pattern. The sole exception to these results concern white fourth births, in which we observed a quicker pace of nonmarital childbearing for white women in some demographic subgroups. Thus if there exists an engine of nonmarital fertility, we find evidence for it only for selected demographic subgroups of white women in their progression from a third to a fourth birth. By contrast, our results are uniformly consistent with predictions from a “brake” hypothesis. We find that women who initiate childbearing within marriage but who then experience a marital separation are significantly less likely to have a subsequent birth outside of marriage. This finding holds for black and white women, across all birth orders, and across all demographic subgroups. Taken together, these empirical findings shed light on the evolving linkage between marriage and fertility, which has drawn increasing attention from social scientists and policy makers concerned with the childbearing of poor and socially disadvantaged women.

“All else equal, growing up in a single-parent family is a handicap. But having a lot of brothers and sisters is also a handicap, for many of the same economic and psychological reasons.”
Christopher Jencks, *Rethinking Social Policy* (1992, p. 194)

Any understanding of the social welfare of children and families in the United States can no longer ignore the fact that a substantial proportion of children are born outside of marriage. Although approximately half of all children spend some time in a single-parent family while growing up, it is less well recognized that nearly two-fifths of such families are begun by an *unmarried* birth (Bumpass and Raley 1995, Bumpass and Sweet 1989). Moreover, nonmarital births continue to increase as a proportion of all births. As of the late 1990s, fully one-third of all births were nonmarital; for whites, 1 out of every 4 births occurred to an unmarried woman, whereas for blacks, 7 out of 10 births were outside of marriage (Wu, Bumpass, and Musick 2001). These high rates reflect persistently high nonmarital birth rates among increasing numbers of unmarried women (Martin, Hamilton, and Ventura 2001), signaling that a rising proportion of children in single-parent families were born into such families.

These transformations in family life are mirrored in the evolving nature of nonmarital childbearing itself. In recent decades, nonmarital fertility has shifted in character, from a phenomenon primarily involving first births to teen women to one involving higher-order births to older women. For example, about one half of all nonmarital births were to teenage women in 1970, but by the early 1990s, nonmarital births to non-teen women accounted for more than two-thirds of all nonmarital births (Ventura et al. 1995). This shift is mirrored in higher-order births. As of the early 1990s, second and higher-order nonmarital births accounted for 48 percent of births to unmarried white women and fully 60 percent of births to unmarried black women (Ventura 1995).

Households characterized by repeated nonmarital childbearing may also play an important role in the growing complexity of family life (Furstenberg and Cherlin 1991; Martinson and Wu 1992). Since children in families marked by repeated nonmarital births may not share the same biological father, the economic, social, and psychological obligations and relationships between children and nonresident (and multiple) biological fathers are likely to be even more complicated (see, e.g., Mott 1990) than the complex patterns of nonresident biological parent/child relationships found in separated, divorced, and stepfamilies (Hetherington and Clingempeel 1992; Maccoby and Mnookin 1992; Seltzer 1991; Weiss and Willis 1985).

Despite the marked nature of these trends, we know surprisingly little about the microlevel dynamics underlying higher-order births to unmarried women. This gap in our knowledge is particularly noteworthy given the considerable speculation about higher-order nonmarital fertility in the policy debates leading up to welfare reform (see, e.g., Murray 1984). Knowledge about the pace, levels, and prevalence of higher-order nonmarital births is important for yet other reasons. For example, although social disadvantage is likely to be greater for children whose parents never marry than for children raised in families in which both biological parents are present at birth (see, e.g., Wu and Wolfe 2001), these disadvantages are likely compounded when such children are raised in households in which they share resources with many siblings (Jencks 1992). Thus, identifying the salient characteristics of families that might be at particularly high risk of repeated childbearing outside of marriage may provide important clues about the socioeconomic disadvantages faced by a growing number of families and children in the United States.

In this paper, we pose two questions: (1) whether a nonmarital birth triggers an “engine” of subsequent nonmarital fertility and (2) whether a marital first birth serves as a “brake” on subsequent nonmarital fertility. The first question asks if a woman who initiated childbearing outside of marriage and who remains unmarried is likely to bear her subsequent children at a quicker pace than if she had initiated childbearing inside marriage. The second question asks if a woman who initiated childbearing within marriage but who subsequently exits marriage is less likely to give birth while unmarried than a woman who bore her first child outside of marriage. These questions need not be mutually exclusive; indeed, as we show below, one can identify conditions under which both, one, or neither hypothesis holds.

Our approach to studying these issues draws upon a now classic demographic literature that asked whether a teen birth ignited an engine of subsequent fertility (Bumpass, Rindfuss, and Janosik 1978; Trussell and Menken 1978; Mott 1986; Rindfuss and Parnell 1989; Morgan 1996; Morgan and Rindfuss 1999). Following this literature, we adopt an analytic strategy focusing on the *pace* of childbearing, but instead of comparing the pace of childbearing for teen and nonteen mothers, we instead engage in three sets of simultaneous comparisons, one comparing the pace of childbearing for women who initiate childbearing inside versus outside of marriage, a second comparing the pace of childbearing for currently married and currently unmarried women, and the last comparing the pace of childbearing for women who initiated childbearing during the teen and nonteen years. We carry out these comparisons for both second and later birth intervals, thus examining not only the transition between a first and second birth, but also the transition from a second to a third birth, and from a third to a fourth birth. A final complication is that what we study has itself been subject to rapid social transformation. This motivates a final set of analytic issues—whether period trends in the progression to a subsequent nonmarital birth differ from those for the progression to a subsequent marital birth.

Our empirical analyses examine transitions between successive births occurring between 1970 and 1995. To track women's trajectories through joint marital and parity statuses, we employ competing-risk, continuous-time hazard models that specify the duration, age, and period structure of childbearing at a particular birth order. Our analyses utilize retrospective marital and fertility histories in the June 1990 and June 1995 Current Population Surveys (CPS). These data contain large representative samples of women residing in the United States at survey but contain relatively little covariate information. As a result, while these data permit us to provide solid evidence concerning the demographic behavior of women with respect to marital and nonmarital fertility, they do not let us establish the causal mechanisms producing the associations we observe. Hence, our analyses can be seen as providing, in some sense, higher-order descriptions, but not firm evidence, concerning causal linkages.

THEORY

Much of what is known about various factors underlying nonmarital fertility derives from research on: (1) time-series or state-level variation in *aggregate* ratios or rates of nonmarital fertility, and (2) individual-level determinants of teen births, nonmarital teen births, or premarital first births. Studies of the effects of state-level variation in AFDC benefits on nonmarital births (Acs 1996; Duncan and Hoffman 1990; Hoynes 1997; Jackson and Klerman 1994; Lichter, McLaughlin, and Ribar 1997; Lundberg and Plotnick 1990, 1995; Plotnick 1990; Robins and Fronstin 1996; see Moffitt 1998 for a review of findings) say little about the individual-level factors associated with *higher-order* nonmarital births, let alone factors associated with *trajectories* of nonmarital childbearing. Indeed, only a handful of studies within this literature have examined the effects of incremental AFDC benefits on subsequent nonmarital births (Acs 1996; Fairlie and London 1997; Grogger and Bronars 2001; Robins and Fronstin 1996). Similarly, studies examining variation at the metropolitan level in nonmarital fertility ratios (Fossett and Kiecolt 1993) and in nonmarital fertility rates (South and Lloyd 1992) say little about parity-specific factors associated with higher-order nonmarital births.

By contrast, numerous studies have examined a variety of social factors thought to affect entry into a nonmarital *first* birth (An, Haveman, and Wolfe 1993; Astone and Washington 1994; Bumpass and McLanahan 1989; Bumpass and Raley 1995; Cherlin, Kiernan, and Chase-Lansdale 1995; Garfinkel and McLanahan 1986; Hogan and Kitagawa 1985; McLanahan 1985; Kiernan and Hobcraft 1997; McLanahan and Sandefur 1994; Powers and Hsueh 1997; Wu 1996; Wu and Martinson 1993). Although valuable, these studies by design do not examine childbearing beyond a first birth and hence are silent on how marital status at first birth might influence subsequent childbearing in and outside of marriage.

Does a Nonmarital First Birth Ignite an Engine of Subsequent Nonmarital Fertility?

The void in the literature concerning our empirical knowledge of higher-order nonmarital fertility may well reflect a presumption that a first nonmarital birth is but a prelude to many subsequent nonmarital births. Indeed, there exist various empirical and theoretical grounds that generate expectations that fertility will be higher among some women, and lower among others. For example, higher fertility and shorter birth intervals have long been characteristic of women from less advantaged backgrounds; this might suggest larger completed families and a quicker pace of childbearing among unmarried women, relative to married women.

The empirical association between social disadvantage and a nonmarital first birth also plays an important role in the so-called quality/quantity tradeoff posited by Becker (1991). Consider two childless women, both of whom desire children, but who differ in their expected income over their life span by virtue of their own earnings, those of a potential spouse, or those provided via government social services. Because of the highly labor intensive nature of child care, children—particularly infants and young children—represent a substantial cost to families, costs which are assumed to be borne disproportionately by women. Under the assumption that time represents a scarce resource and that the allocation of scarce resources is the result of deliberate actions by rational decision makers, it follows that the cost of bearing and caring for young children will vary between families, depending largely on the woman's price of time, with differences in a woman's price of time varying with her potential life cycle earnings.

If a woman can derive satisfaction both from the numbers of children she bears *and* from the resources she devotes to each child, then equivalent satisfaction may be obtained from different combinations of numbers of children and the average investment per child. If so, then a standard result is that rational actors will engage in tradeoffs between these two factors—the quantity of chldren and the amount of resources devoted to each child—yielding the so-called quantity/quality tradeoff. Differences in a woman's price of time—for example, the opportunity cost represented by her foregone wages in the labor market, resulting from time spent out of the labor force to bear and care for an infant—will thus affect the numbers of children she will choose to bear (Becker and Lewis 1973; Becker 1991). An implication of this argument is that two women with different potential earnings can derive equivalent satisfaction by bearing different numbers of children, with the woman with higher earnings potential bearing fewer children but devoting more resources to each child and the woman with lower earnings potential bearing more children but devoting fewer resources to each child.[1]

This logic also underlies the emphasis by economists on the "incentive" effects of AFDC on nonmarital fertility. Under standard economic models, one would expect little fertility among young, unmarried women with poor current and future earnings capacities, who face poor marital prospects, or who cannot "borrow" against future earnings; thus, these arguments predict that women will not bear children in the absence of income from sources such as AFDC (see, e.g., Murray 1994). Proponents of this line of reasoning further note that in many states, AFDC benefits were limited to unmarried women with dependent children, providing a disincentive to marriage, in critics' eyes, for unmarried women. Recent theoretical results (see, e.g., Willis 1999) also provide comparative static conditions under which men would be expected to bear many children by different women outside of marriage.[2]

Despite the strong nature of these theoretical predictions, most studies to date have found that AFDC had only small effects on nonmarital fertility (see Moffitt 1998 for a review and Rosenzweig 1999 for an exception). Moffitt (2001) concludes from an analysis of time-series variation in AFDC benefits that labor market factors affecting low-wage men and women, rather than government policies such as AFDC, were primarily responsible for the rise, in recent decades, in female headship. Thus, whereas empirical findings are in the expected direction, the findings

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[1]Becker's argument provides implications for the comparative statics for family size, but does not speak directly to issues concerning the pace of childbearing. Empirically, however, most women space their births closely, with most subsequent births following within five years of a previous birth. Thus, theoretical predictions concerning variation in completed family size also carry empirical implications about the pace of childbearing.

[2]In addition the quality/quantity tradeoff, a standard economic argument posits a "pure" income effect on fertility, in which higher-income families are expected to have higher completed family sizes than lower-income families; however, economic theory provides no guidance on the relative magnitude of these factors. Thus, if the effect of quality/quantity factors were larger than the pure income effect, one would expect higher fertility among disadvantaged groups; conversely, if the effect of pure income were larger, one would expect higher fertility among the socially advantaged.

in the AFDC literature provide at best ambiguous guidance concerning the size of differentials in completed family size and in the pace of childbearing for married and unmarried women.

The quality/quantity tradeoff faced by women also plays a prominent role in the historical and demographic literatures. Researchers in these fields typically place far greater emphasis than economists on the role of values and norms in determining historical variation in outcomes such as desired or completed family size. Thus, classic historical accounts by Ariès (1962), Shorter (1975), and Stone (1977) argue that Western European conceptions of childhood, motherhood, and the individual evolved in ways that profoundly affected demographic behaviors in the realms of childbearing, marriage, and sexual intimacy. Preston (1986) articulates a standard set of demographic expectations by noting that long-term patterns of economic change have raised the "socially constructed 'costs' of children and the opportunity costs of pursuing a lifetime of responsible parenting." These costs now fall most heavily on women with higher educational attainments, expected earnings, and consumption aspirations. Conversely, Thornton (1979) finds that intended family size for women with low aspirations for child quality (as measured by responses to items such as education and inputs of maternal time) is between .40 and .47 higher than for those with high aspirations for child quality.

Recent demographic trends have also been taken by some commentators as evidence of an engine of nonmarital fertility (Gilder 1986; Murray 1994). For example, the demographic literature has long posited that teen mothers are likely to differ in various ways from those who delay childbearing; early childbearers are likely to have poorer contraceptive practices, higher fecundability, and higher fertility desires. These expectations led researchers in the late 1970s to posit an engine of teen fertility (Bumpass, Rindfuss, and Janosik 1978; Trussell and Menken 1978; Mott 1986; Rindfuss and Parnell 1989; Morgan 1996; Morgan and Rindfuss 1999), in which the occurrence of a teen first birth produced a quicker pace of subsequent childbearing. Although the association between a teen first birth and a woman's pace of subsequent childbearing appears to have weakened substantially in recent years (see, e.g., Rindfuss and Parnell 1989), the *initiation* of nonmarital fertility remains highly concentrated in the teen and early adult years (Wu, Bumpass, and Musick 2001). Moreover, while such nonmarital first births remain prevalent among teenage women, overall nonmarital childbearing has shifted increasingly toward higher-order births to nonteen women; second and higher-order nonmarital births now account for nearly half of births to unmarried white women and fully three out of five births to unmarried black women (Ventura 1995). Although increases in the proportion of births to unmarried women and increases in the numbers of higher-order nonmarital births need not necessarily imply an engine of nonmarital fertility, the converse would be true—a more powerful engine of nonmarital fertility would, *ceteris paribus*, lead to increases in the proportion of births to unmarried women and to increases in higher-order nonmarital births. Thus, such trends provide clues, albeit tentative ones, that may signal the presence of an engine of nonmarital fertility and may suggest increases in the propensity of successive cohorts of women with a nonmarital first birth to proceed to higher-order births outside of marriage.

Thus far, we have reviewed arguments positing possible mechanisms that might underlie an engine of nonmarital fertility. However, there also exist grounds on which one might expect a *slower* pace of childbearing among unmarried women relative to that for married women. For example, classic discussions in the demographic literature (see, e.g., Davis and Blake 1956; Bongaarts 1978) have long identified sexual activity as an important (albeit typically unobserved) proximate determinant of fertility. These accounts, focusing on *marital* fertility, typically cast the proximate influence of sexual activity in terms of the coital frequency of married couples. By contrast, our focus on contrasting nonmarital and marital fertility places the proximate influence of sexual activity in a different light; in particular, the arguments we review below suggest that sexual activity might well be expected to be lower, on average, among single, non-cohabiting women, relative to married women.

The basis for such an argument rests on the extent to which men are present in the lives of unmarried women. Although many states conditioned AFDC receipt on the absence of a man in the household, numerous ethnographic accounts suggest that men in fact play a critical role in the lives of unmarried mothers. Nevertheless, the available ethnographic evidence provides widely varying depictions of the relationship between unmarried women and the men in their lives. Some depict young black males in urban ghettos as sexually predatory (Anderson 1990), whereas others describe young black women as actively resisting attempts by welfare officers to sever their ties with "homeboys," whom welfare officers—often themselves working class black women—view as contributing to dependency and delinquency among their AFDC clients (Haney 1996). Similarly, accounts in which AFDC mothers express their view of the men in their lives as unreliable breadwinners (Edin and Lein 1997; Waller 200x) may appear inconsistent with other accounts describing the majority of such women and men as romantically involved and optimistic about their futures (McLanahan et al. 2001).

Whatever the nature of the relationships between unmarried women and men, it does seem likely that such relationships are marked by greater instability than those of couples in more advantaged circumstances. For example, while McLanahan and colleagues report that substantial numbers of unmarried fathers and mothers are romantically involved, that many unmarried fathers are present at birth, and that many fathers express their willingness to be actively involved in raising their children, they caution that father involvement is likely to decline over time, with some unmarried couples breaking off their relationship as time passes. Similarly, Furstenberg (1995), in summarizing findings from his Baltimore study of young teen mothers, concluded that "the intentions of [unmarried] fathers far outstrip their ability to make good on their goal of becoming involved caretakers. Whether by design, desire, or default many fathers retreat—some almost immediately but most after their initial efforts end in frustration or their motivation lags." As a result, the greater instability of such relationships would be likely to be reflected in the availability of sexual partners for unmarried women; this in turn, all else being equal, would be expected to lead to lower fertility of unmarried women relative to their married counterparts.

A complicating factor in this argument is that a substantial fraction of unmarried births occur to parents residing in a cohabiting union (Bumpass and Lu 2000). However, accumulating evidence suggests that such cohabiting unions are substantially less stable relative to marital unions. For example, several studies document higher dissolution rates among married couples who cohabited prior to marriage (Lillard, Brien, and Waite 1995) relative to those who did not cohabit prior to marriage. There is also evidence for higher rates of union dissolution among couples with a cohabiting first birth than among couples with a marital first birth (Ermisch 2001; Wu, Bumpass, and Musick 2001). Thus, the greater instability of cohabiting unions relative to marital unions provides additional grounds for expecting that sexual activity will be lower, on average, for unmarried women than married women.[3]

Does a Marital First Birth Act As a "Brake" on Subsequent Nonmarital Fertility?

We now turn to our second question—whether a marital first birth might act as a "brake" on a woman's propensity to bear a subsequent birth outside of marriage. In posing this question, we ask if a woman who bears her first child within marriage, but who subsequently experiences a marital separation before a second birth, is less likely to have a child during the periods when she is unmarried than a woman who has had a nonmarital first birth (Rindfuss and Parnell 1989). Clearly, the "engine" and "brake" hypotheses pose fundamentally different questions. For example, the "brake" hypothesis involves comparisons of the fertility of women with marital and nonmarital first births when both groups of women are *unmarried*. By contrast, the "engine" hypothesis might be thought of as comparing an engine of nonmarital fertility with an engine of marital fertility, thus involving comparisons of the fertility of women with marital first births while they are *married* with the fertility of women with nonmarital fertility while they are *not married*.

[3] It is noteworthy that this argument—that sexual activity is likely to be lower among unmarried women than among married women—departs considerably from more traditional views, for example, that of historians who have often regarded the incidence of nonmarital childbearing as a rough empirical proxy for illicit sexual activity in the population at large (Laslett, Oosterveen, and Smith 1980). Although such assumptions are clearly implausible in modern societies in which sexual activity outside of formal marriage is commonplace, Gordon (1994) notes that such seemingly outdated assumptions about the sexual promiscuity of women with illegitimate children may have contributed to continuing public perceptions of such women as morally deficient and hence as "undeserving" of public assistance.

Why might the occurrence of a marital first birth act as a “brake” on the propensity to bear a child outside of marriage among women that experience a marital dissolution? Arguments by social scientists as varied as Wilson (1987) and Becker (1991) suggest that such a “brake” is likely to reflect the nonrandom selection of women into marriage—i.e., differences among women that are likely to predate the entry into marriage. Several potential mechanisms underlying such selection have been proposed, including differences in the pools of “marriageable” males (Wilson 1987), the relative wages of men and women (Becker 1991; Schultz 1994; Moffitt 2001), or a woman’s forecast of access to subsequent marriage. This last proposes that women with a marital first birth anticipate higher remarriage prospects than women with a nonmarital first birth (Lichter et al. 1992; South and Lloyd 1992; Lichter and Graefe 2001; Upchurch, Lillard, and Panis 2001).

The quality/quantity tradeoff posited by economists also yields predictions about the association between a marital or nonmarital first birth and the propensity of unmarried women to bear a child outside of marriage. Under this hypothesis, a marital first birth is likely to signal a preexisting preference by women to invest greater resources in such a birth; hence, a woman with a marital first birth may choose to forgo childbearing while unmarried or separated, until such time as she can make comparable investments in a second child—that is, when she remarries. Alternatively, differences in such preferences may emerge during marriage. For example, new parents often express surprise at both the amount of effort needed to raise children and the pleasures that accompany parenthood (Gerson 1985), suggesting that adults who have not yet become parents do not fully anticipate either the rigors or rewards of a first birth. Thus, the preference of a woman with a marital first birth not to bear a subsequent child outside of marriage could reflect either a preference that predates marriage or a preference—for example, for a stable family environment—that is shaped by the experience of raising a first child within a marital union.

One can also motivate arguments that predict no difference in the childbearing of unmarried women with marital and nonmarital first births. For example, conservative commentators often express concern over the apparent decline in social stigma attached to nonmarital childbearing.

If social stigma were a powerful factor in shaping women's fertility decisions and if the stigma attached to a nonmarital birth were to have declined equally for all women, then those women with a marital first birth who then subsequently exit a marriage might be as likely to bear a child outside of marriage as women with a nonmarital first birth. Similarly, because divorce is typically unanticipated, economists might view a divorce as posing a random shock to a woman's economic situation; this might in turn lower a woman's expected investment in a second child. Alternatively, the arguments made by proponents of marriage (Waite 1995; Popenoe 1996) could be taken to imply that the pool of those who divorce selects on those parents who are willing to accept lower investments in their children after divorce, given that the divorced parents could have chosen to remain married, at a cost to the parents but with presumed benefits to the children. If so, then such an argument might imply small differentials in the pace of childbearing among unmarried women with a marital and nonmarital first birth.

DATA AND METHODS

The questions we pose require large samples to obtain stable estimates at higher birth orders and for births within and outside of marriage. For this reason, we analyze data from the June 1990 and 1995 Current Population Surveys (CPS). These data provide a representative cross-section of the U.S. population aged 15 and over; retrospective reports from female respondents let us reconstruct their marital and fertility histories. We report weighted results throughout.

The June CPS maintained reasonably consistent items on marriage and fertility from 1990 to 1995. The sample universe for fertility questions consists of married women aged 15 or older and never-married women aged 18 or older. As a result, the CPS lacks fertility data for never-married women who were younger than 18 at survey. The resulting truncation bias should not seriously affect our analyses because births at very young ages will be available for women who began childbearing prior to age 18 but who were older than 18 at survey. A screener question obtains the number of children born (excluding stillbirths), followed by items on the date of birth (calendar month and year) for the first four and the last child. We used these data to construct a retrospective birth history through the fifth parity for women with five or fewer births at interview; however, because the transition from a fourth to fifth child is relatively rare, we restrict our empirical analyses through the transition from a third to fourth child only.

The CPS marital history is asked of ever-married women aged 15 or older. Dates (to nearest month and year) of marriage, separation, and widowhood/divorce are obtained for the first two marriages and for the most recent marriage. A screener question obtains the respondent's number of marriages (0, 1, 2, 3+ in 1990; 0, 1, 2, 3, and 4+ in 1995). This means that is not possible to distinguish, in 1990, between women with exactly three marriages or with more than three marriages or, in 1995, between women with exactly or more than four marriages. This problem affects only a handful of women and we censor a woman's marital and fertility histories when her marital history become ambiguous in this way.

An advantage of pooling data from these two surveys is that they provide very large samples of both white and black women, even for parity transitions at higher birth orders. Nevertheless, these data have two distinct drawbacks. First, they lack the rich array of covariate information for respondents available in other longitudinal surveys such as the 1979 National Longitudinal Survey or the Panel Study of Income Dynamics. Second, they lack information that would permit us to construct a retrospective cohabitation history, detailing entries and exits from cohabiting and marital unions. As a result, the CPS lets us distinguish only between married and unmarried births and we cannot further distinguish between unmarried births that occurred within or outside of a cohabiting union.

In the analyses reported below, we define a nonmarital birth as a birth occurring prior to entry into marriage or after a marital separation. If the reported month and year of a birth and entry into (or separation from) a marriage coincide, we classify the birth as marital.[4] A small

[4] One possible ambiguity is that births occurring immediately following a marital separation may have been conceived within marriage. In analyses not reported, we added a control for a time-varying dummy variable

number of women at parity p experience a subsequent twin birth. In our analyses, we pool women with singleton second births with women who experience twin births after a first birth. We then increment a woman's parity status accordingly, putting women with singleton second births at risk of a third birth, and putting those with a twin second birth at risk of a fourth birth. Higher-order multiple births are treated similarly.

To increase the number of observed parity transitions, we examine births occurring in the period 1970–1995. We condition on cohorts of birth at a given achieved parity p; thus, in analyses of the transition from a first to second birth, we restrict our sample to women with first births that occurred between 1970 and 1995, while for the transition from a second to third birth, we restrict our sample to second births occurring between 1970 and 1995. As a consequence, our analyses of second births and third births contain women who initiated childbearing prior to 1970.

Methods

Figure 1 illustrates transitions defined by joint marital and parity statuses. Thus, a woman with a nonmarital first birth can transit to a second nonmarital birth, to marriage prior to a second birth, to neither a second birth nor marriage prior to survey, and so forth. Movements through these statuses define the relevant risk sets for marital and nonmarital second and higher-order births.

Predictions from the "engine" and "brake" hypothesis can be stated in terms of comparisons among the transitions in Figure 1. Both hypotheses share a concern with the progression from a pth birth to a $p + 1$ birth outside of marriage, but they differ by directing attention to distinct sets of comparisons. For women who have achieved parity 1, the "engine" hypothesis focuses on a comparison between marital and nonmarital fertility, positing that, all else being equal, the pace of fertility is higher for the $01 \rightarrow 02$ transition than for the $11 \rightarrow 12$ transition. By contrast, the "brake" hypothesis focuses on comparing nonmarital fertility among women who initiated

equal to 1 in the first ten months following a marital separation, with the ten month period chosen to represent a typical 38 week gestational period. Results from these analyses differ only in small ways from the results reported below.

$$\begin{tikzcd}
& \text{parity 1} & \text{parity 2} & \text{parity 3} & \text{parity 4} & \\
\text{not married} & \boxed{01} \arrow[r] \arrow[d, shift right] & \boxed{02} \arrow[r] \arrow[d, shift right] & \boxed{03} \arrow[r] \arrow[d, shift right] & \boxed{04} \arrow[d, shift right] & \cdots \\
\text{married} & \boxed{11} \arrow[r] \arrow[u, shift right] & \boxed{12} \arrow[r] \arrow[u, shift right] & \boxed{13} \arrow[r] \arrow[u, shift right] & \boxed{14} & \cdots
\end{tikzcd}$$

Figure 1: Transitions between joint marital and fertility statuses of women.

childbearing within and outside of marriage, thus positing that, all else being equal, the pace of fertility for the 01 $\rightarrow$ 02 transition is lower for women with a marital first birth than for those with a nonmarital first birth. Similar statements hold for higher-order parity transitions.

Because both hypotheses are concerned with nonmarital childbearing, another way in which to juxtapose these hypotheses is to build on the simple observation that women are not at risk of a nonmarital birth unless they are unmarried at some time u, where u denotes the duration since a previous birth. Thus, a woman can be unmarried at duration u if she has had her previous birth while unmarried and if she has not married in the interim, or if she had her previous birth while married but has had a marital separation in the interim, and so forth. Unmarried women who have achieved parity p are thus at risk of a nonmarital p + 1 birth, whereas married women who have achieved parity p are not at risk of such a birth.

Figure 2 depicts a 2×2 cross-classification of marital status at first birth with marital status at u, where u denotes duration since birth p and where the quantities $r_a(u)$, $r_b(u)$, $r_c(u)$, and $r_d(u)$ denote the four transition intensities for the transition between birth p and birth $p+1$. The transition intensities for marital fertility are thus given by $r_a(u)$ and $r_c(u)$, while those for nonmarital fertility are given by $r_b(u)$ and $r_d(u)$. Consider, by way of illustration, the transition from a first to second birth; then in terms of Figure 2, the engine hypothesis posits that $r_d(u) > r_a(u)$, whereas the brake

	married at u	not married at u
marital first birth	$r_a(u)$	$r_b(u)$
nonmarital first birth	$r_c(u)$	$r_d(u)$

Figure 2: A 2 cross-classification of parity-specific transition rates for marital status at first birth and marital status at duration u since the birth of child p.

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hypothesis posits that $r_b(u) < r_d(u)$.

To formalize ideas, let $r_p^n(t, u)$ and $r_p^m(t, u)$ denote the age- and duration-specific transition intensities for nonmarital and marital fertility, respectively, for women who have achieved parity p, $p = 1, \ldots, P$, where the superscripts m and n denote marital and nonmarital parity progressions, respectively; t denotes a woman's age; and u duration since previous birth. We proceed conditional on the occurrence of a marital or nonmarital first birth. Given these assumptions and under a proportional hazard specification for the effects of covariates, the transition rates for nonmarital and marital fertility, respectively, can be written as:

$$\begin{aligned} r_p^n(t, u) &= \lambda_p^n(t) q_p^n(u) \exp[\mathbf{b}_p^n \mathbf{x}_p(t)] \\ r_p^m(t, u) &= \lambda_p^m(t) q_p^m(u) \exp[\mathbf{b}_p^m \mathbf{x}_p(t)] \end{aligned} \tag{1}$$

where $\lambda_p^n(t)$ and $\lambda_p^m(t)$ model age dependence at each parity; $q_p^n(u)$ and $q_p^m(u)$ parameterize duration dependence in the interval between births p and $p+1$; $\mathbf{b}_p^n$, and $\mathbf{b}_p^m$ are vectors of parameters to be estimated; and $\mathbf{x}_p(t)$ is a vector of covariates (some time-varying) assumed to affect both marital and nonmarital fertility in the interval between the p and $p+1$ birth. Throughout, we treat a woman with a marital pth birth who separates or divorces at some duration u' following birth p as at risk of a *marital* $p + 1$ birth at durations 0 to u', and at risk of a *nonmarital* birth at durations u' and later—that is, women can have staggered entries into the risk of a marital or nonmarital birth, as defined by the left-truncation time given by a change in marital status at duration u'.

Figure 3 presents smoothed estimates of the logarithm of transition rates for a marital second birth (solid curve) and for a nonmarital second birth (dotted curve) using a nonparametric procedure described in Wu (1989). The solid curve thus corresponds to the transition 11 $\rightarrow$ 12 and the dotted curve to the transition 01 $\rightarrow$ 02 as depicted in Figure 1. Progressions involving higher-order birth intervals are given in the next two panels and correspond to the transitions 12 $\rightarrow$ 13, 13 $\rightarrow$ 14, 01 $\rightarrow$ 02, 02 $\rightarrow$ 03, and 03 $\rightarrow$ 04. We examine the nonparametric estimates for these transitions because they provide crude but useful information on the outcomes we examine in our more parametric models and because they provide important clues on how to specify these parametric models.

[Figure 3 about here]

One way in which these estimates are somewhat crude is to note that they do not condition on whether a woman was married or unmarried *at the time of first birth*, which is central to testing the engine and brake hypotheses. To clarify the estimates in Figure 3, first recall that throughout, we condition on the occurrence of a first birth; hence, there are no childless women in our analyses. Then consider the transition 11 $\rightarrow$ 12 in Figure 1, corresponding to the solid curve in the upper panel of Figure 3. In estimating the solid curve, we place women at risk of a marital second birth if they are at parity 1 and are currently married. This can occur in a number of ways. The most common example will consist of women with a marital first birth who remain married. A woman with a nonmarital first birth who experiences a marriage will also be at risk of a marital second birth at the time of marriage. Similarly, a woman with a marital first birth who experiences a marital dissolution, followed by a remarriage will be at risk of a second marital birth at the time of remarriage. Other examples involving multiple marriage transitions are also possible. In estimating the dotted curve, we place women at risk of a nonmarital second birth if they are at parity 1 and are not currently married. This can also occur in a number of ways. Curves in the middle and lower panels of Figure 3 proceed in a similar way.

The above examples make clear that while Figure 3 provides simple comparisons of marital and nonmarital fertility, such comparisons do not condition on marital status at the time of first birth and thus do not constitute tests of the engine and brake hypotheses. (Our parametric models, discussed below, incorporate such conditioning explicitly.) Nevertheless, Figure 3 provides important clues about the dependent variables in our models—that is, the transition rates $r_{01\to02}, r_{11\to12}, r_{02\to03}, r_{12\to13}$, and so forth. In particular, it provides important hints on duration dependence—how the curves in Figure 3 vary with duration—and nonproportionality—whether the difference between marital and nonmarital fertility is constant or varies with duration.

An important finding from Figure 3 is that nonmarital fertility is similar to (or may exceed) marital fertility immediately following a pth birth, but is generally lower at longer durations. For example, the curves in the upper panel of Figure 3 resemble one another closely immediately following a first birth, but then diverge sharply, with rates of a nonmarital second birth substantially lower than rates of marital second birth. Patterns at higher birth orders are similar; transition rates for nonmarital births higher are than those for marital births immediately following a previous birth, and slightly lower than or equal to those for marital births at later durations.

That the difference between marital and nonmarital fertility appears to vary with duration creates difficulties both technical and substantive. A technical point is that a nonconstant effect of marital status on fertility violates the standard proportionality assumption—that the difference between two groups can be summarized by a single proportionality constant or, equivalently, by a single additive constant on a logarithmic scale. Stated substantively, the engine hypothesis implicitly assumes that the true difference can be summarized by a single number—for example, that the difference in access by married and unmarried women to a sexual partner does not vary over time. Figure 3 calls this assumption into question. We return to these issues below.

Finally, Figure 3 helps guide our choice of models for duration dependence in nonmarital and marital fertility. We observe a roughly unimodal pattern in Figure 3, with $\log r(t)$ rising in a roughly linear way at short durations and declining in a roughly linear way at longer durations.

	married at u	not married at u
marital first birth	omitted	$b_1 x_1(u)$
nonmarital first birth	$b_2 x_2$	$b_1 x_1(u) + b_2 x_2 + b_3 x_1(u) x_2$

Figure 4: Parametric specification of the transition rate from birth p to $p + 1$ by marital status at first birth and marital status at duration u since birth p.

We model this pattern using a splined piecewise Gompertz model (see, e.g., Lillard 1993; Wu and Martinson 1993; Wu 1996), which models $\log r(t)$ using a piecewise linear spline specification for dependence. This model requires prespecifying nodes for the spline. Inspection of Figure 3 suggests placing these nodes in the duration splines at 9, 17, 36, and 60 months.

We now turn to our parametric models. Consider births to women who have achieved parity p and suppose (contrary to Figure 3) that fertility differentials between women who bear their first child in or out of marriage is a constant that does not vary with the duration since a previous birth. Figure 4 formalizes these ideas by specifying main effects of marital status at first birth, marital status at duration u, and the interaction of these two variables.

Let H_1 and H_2 denote the engine and brake hypotheses, respectively. It is possible to show that H_1 and H_2 are not mutually exclusive; indeed, one can identify conditions under which both H_1 and H_2 are true or under which both H_1 and H_2 can be false. Recall from Figure 4 that

$$\begin{aligned} H_1 :& \qquad b_1 + b_2 + b_3 > 0 \\ H_2 :& \qquad b_2 + b_3 > 0 \end{aligned} \tag{2}$$

Then assuming proportionality, both H_1 and H_2 will be false if

$$b_2 + b_3 \leq 0 \quad \text{and} \quad b_1 \leq -(b_2 + b_3)\,; \tag{3}$$

similarly, both H_1 and H_2 will be true if

$$b_2 + b_3 > 0 \quad \text{and} \quad b_1 > -(b_2 + b_3)\,. \tag{4}$$

Because the CPS lacks rich covariate information, our analyses contain only a limited set of controls; see Table 1. To control for age variation in parity-specific birth rates, we specify several dummy variables for age at first birth (dummy variables equal to 1 if the respondent is aged 18 or less, age 19, age 20 to 24, age 25 to 29, and age 30 or older). To control for period variation in the transition from birth p to birth $p + 1$, we specify several dummy variables capturing the calendar years (1970–74, 1975–79, 1980–84, 1985–89, and 1990 or later) in which birth p occurred. In some models, we also introduce controls for the woman's highest grade completed (less than high school, high school, some college, and a college degree or higher). Note, however, that this variable reflects completed education at time of interview; hence, some women may have completed additional years of schooling after the birth of a child. We present separate analyses for white and black women.

[Table 1 about here]

To compensate in part for the relatively small set of controls, we have included in all of our models interactions of the time-varying dummy variable for a woman's marital status at duration u with the covariates for age at first birth, the period dummy variables, and (in relevant models) education. As a result, these interactions let the magnitude of the parameters relevant to tests of the engine and brake hypotheses vary across all observed demographic characteristics. Thus, despite the lack of a rich array of control variables, this specification provides a quite flexible parameterization for our tests of the engine and brake hypotheses.

RESULTS

Table 2 presents weighted counts of marital or nonmarital births by birth order for black and white women. As noted earlier, for the transition from a first to second birth, we condition on first births that occurred during the period 1970–95. In this period, 41,339.4 white women had a first birth, with 17,093.0 of these women proceeding to a marital second birth and 2,161.1 to a nonmarital second birth. Similarly, in our analyses of the transition from a second to third birth for white women, we condition on the 32,525.8 second births that occurred during the period 1970–95. Of these white women, 7,002.7 proceeded to a marital third birth, and 1,002.5 to a nonmarital third birth. The subsequent columns give the remaining weighted frequencies for white and black women. Note that from 1970 to 1995 and at each achieved parity, marital births outnumber nonmarital births for white women, whereas for black women, nonmarital births outnumber marital births.

[Table 2 about here]

Table 3 presents estimates from the proportional hazard specification depicted in Figure 4. Note that Figure 3 suggests that the assumption of proportionality in marital/nonmarital fertility differentials by duration is violated in these data. We nevertheless begin our discussion of results using this model specification, for two reasons. First, the estimates in Table 3 have the desirable property of summarizing fertility differentials of unmarried and married women in terms of a single number. Although formal tests (not reported) suggest strongly rejecting this assumption—that fertility differentials between married and unmarried women do not vary with duration since a previous birth—a proportional specification has the virtue of simplicity in that it uses a single parameter to summarize the differential between marital and nonmarital fertility.[5] Second, we use the simpler estimates presented in Table 3 to motivate subsequent analyses, where we relax the proportionality assumption. Thus, these estimates provide a baseline for comparison and help gauge the sensitivity of results to violations of proportionality.

[Table 3 about here]

[5]Note also that these estimates can be viewed as providing the average value of this differential, with the average taken over the observed durations in these data, as adjusted for right censoring. Let $S(t)$ denote the baseline survival probability; then recall that the exponentiated value of a coefficient $\gamma_1 = \exp(b_1)$ in a proportional hazard model affects baseline survivor as $[S(t)]^{\gamma_1}$. Empirically, estimates from proportional and nonproportional hazard models tend to agree most closely at the median of the event distribution. It is in this sense that a proportional hazard estimate provides a rough estimate of the expected value of differentials, as evaluated around the median event time, even under violations of the proportionality assumption.

The baseline model specification contains only three variables: a time-varying dummy variable equal to 1 if the respondent is unmarried at duration u, a non-time-varying dummy variable equal to 1 if the respondent was unmarried at first birth, and a final dummy variable interacting these two variables. The omitted category thus consists of currently married women with marital first births—that is, women at risk of a marital second birth. Under this specification, the baseline hazard (specified as a splined piecewise Gompertz model) refers to women in the omitted category, with the risks for other women defined by the main and interaction effects of the other three variables. The first column of results presents estimates from a model for the transition from a first to second birth for white women. The remaining columns give estimates for third and fourth births to white women and estimates for second, third, and fourth births to black women.

The main effect of marital status at a duration u following a previous birth, given in the rows labelled 1, contrasts the fertility of currently married and currently unmarried women at specific parities. The first row of estimates indicates that in our baseline model, currently unmarried white women have a slower pace of childbearing than currently married white women at all parities. This effect is large and significant (-1.24) for the transition from a first to second birth for white women. The next two columns show that birth risks for currently unmarried white women remain significantly lower than for currently married white women ($-.29$ and $-.22$ for third and fourth births, respectively), but estimated coefficients are substantially smaller than that for a second birth. The next three columns present results for black women. Overall, estimates suggest that, as for white women, currently unmarried black women have a slower pace of childbearing than currently married black women at all parities. Estimated coefficients are smaller in magnitude than for white women and are significant only for the transition from a first to second birth.

The next two rows of estimates provide additional contrasts by giving fertility differentials for women with nonmarital first births and for currently unmarried women with nonmarital first births. We observe positive values for the second row of coefficients; this indicates that women with a nonmarital first birth who are observed to subsequently marry have somewhat higher rates of marital fertility compared to women who were married at first birth. The third row of estimates is most easily interpreted in combination with the other estimates; these interactions, together with the other estimates, provide tests of the "engine" and "brake" hypotheses, as we explain shortly.

The remaining sets of estimates in Table 3 introduce additional controls for period trends, age at first birth, and education. In these models, we observe negative values for the main effect giving the contrast between the duration-specific fertility of married and unmarried women. For white women, this effect is significant in all models and at all parities. For black women, this effect is also significant in all models and at all parities, with one exception—the transition from a third to fourth birth in the model controlling only for birth cohort. Similarly, we observe generally positive values for the effect of a nonmarital first birth and for the interaction giving the effect for currently unmarried women with nonmarital first births. However, no clear pattern of results for these two variables emerges across parities.

Does a nonmarital first birth ignite an engine of subsequent nonmarital fertility? As noted above, contrasting the pace of subsequent childbearing among those with marital and nonmarital first births involves the sum of the three parameters reported in Table 3. Estimates of $b_1 + b_2 + b_3$, taken from the models in Table 3, are reported in Table 4. We obtained tests of significance for the parameter $b_1 + b_2 + b_3$ by comparing the likelihoods from two models, one in which all three parameters varied freely, and a second in which the sum was constrained to equal zero. The resulting likelihood ratio test involves one degree of freedom, corresponding to the single equality constraint. Table 4 reports the corresponding χ^2 values in parentheses below the corresponding point estimates for $b_1 + b_2 + b_3$.[6]

[Table 4 about here]

If a nonmarital first birth ignited an engine of subsequent nonmarital fertility relative to an engine of marital fertility, then we should observe positive values for the quantity $b_1 + b_2 + b_3$. The

[6] If a random variable x is χ^2 distributed on 1 df, then $\sqrt{x}$ is normally distributed (see, e.g., Yule and Kendall 1950); hence, the square root of the reported χ^2 values can be evaluated using the usual normal test statistic.

results in Table 4 suggest little support for this hypothesis. For whites, we observe a large, negative, and significant coefficient for the transition from a first to a second birth; hence, for this transition and this model, results are the opposite of expectations. Only in the baseline model for third and fourth births to white women (next two columns) do we observe positive and significant values for this parameter. Note, in addition, that in this model, the absolute magnitude of $b_1 + b_2 + b_3$ for third and fourth births (.13 and .30, respectively) is substantially smaller than that for second births ($-.76$). As controls are added, the coefficients for third births flip in sign, and the coefficients for fourth births decrease in magnitude and are no longer significant. Overall, then, our results suggest that the pace of a nonmarital second birth is significantly *lower* than the pace of a marital second birth. For higher-order births, we find that the pace of a nonmarital birth at higher parities is no higher than (and sometimes significantly lower than) the pace of a marital birth in models that control for the period in which the birth occurred, for the age of the mother at first birth, or for education.

The results for blacks are similar to those for whites. For the transition from a first to second birth, the pace of nonmarital fertility is significantly lower than the pace of marital fertility; moreover, the magnitude of this difference increases when controls are added. For higher order birth transitions, we observe a quicker pace of nonmarital fertility than marital fertility in our baseline model, but this effect reverses in sign and loses statistical significance as controls are added. Thus, for both white and black women, the results in Table 4 suggest little support for an engine of nonmarital fertility.

The results in Table 4 refer to birth risks for the omitted categories in our hazard regressions—for example, the 1970–74 childbearing cohort for women aged 20–24. Table 5 repeats the analyses using the controls in the third row of Table 4, but asks if an engine of nonmarital fertility can be detected for women of different childbearing cohorts or at different ages. The results for second and third births for white and black women in Table 5 indicate little support for the engine hypothesis. Coefficients are typically negative and those coefficients that are positive are not significantly different from zero, except for third births to white women who had a second child between 1990 and 1995. Fourth births to black women follow a similar pattern. The major exception to this pattern of results concerns fourth births to white women; for these births, we find positive and statistically significant coefficients in several demographic subgroups—for white women who had a third child in 1980–84, 1985–89, and 1990–95, and for white women who were 30 or older at first birth. These results suggest that if there is an engine of nonmarital fertility, it would appear to be confined to fourth births to white women of particular childbearing cohorts and ages. Moreover, the absence of such an engine for lower-order births to black and white women across a wide range of childbearing cohorts and ages would appear to pose difficulties for proponents of such a hypothesis—it is difficult to construct plausible arguments that would generate no engine of nonmarital fertility at lower birth orders but would let such an engine emerge at higher birth orders for white—but not black—women. Overall, then, we interpret the evidence in Tables 4 and 5 as providing little support for the hypothesis that a nonmarital first birth ignites an engine of subsequent nonmarital fertility.[7]

[Table 5 about here]

As we noted earlier, the results in Tables 4–5 assume proportionality in the differentials between the fertility of married and unmarried women—in particular, that this differential does not vary with duration—whereas Figure 3 raises questions about this assumption. To examine the sensitivity of our results to proportionality, we report results from two additional models that relax this assumption by letting the differential between marital and nonmarital fertility at each parity vary with duration. One model specifies the differential between marital and nonmarital fertility using a series of piecewise constant contrasts for the periods 0–9, 10-17, 18–36, 37–60, and 61+ months; a second model specifies these differentials using a piecewise linear spline on

[7]We include controls for education in Table 5 to parallel the progression of models in Table 4. Excluding educational controls in Table 5 and in subsequent tables yields qualitatively similar findings (results available upon request).

a log scale for the hazard rate. Estimated coefficients from these models (not reported) show the patterns observed in Figure 3, with rates of nonmarital childbearing significantly lower than rates of marital childbearing at longer durations. At shorter durations, results from the piecewise linear models suggest rates of nonmarital childbearing at short durations that begin at a higher level, but proceed at a slower pace, relative to marital childbearing. Note that a drawback of such alternative specifications is that the differential in marital and nonmarital fertility is no longer easily summarized by a single non-time-varying quantity such as $b_1 + b_2 + b_3$; hence, a consequence of relaxing the proportionality assumption is that one can no longer specify a test of the engine hypothesis in terms of a single quantity.

Does a marital first birth act as a "brake" on nonmarital fertility by being associated with lower risks of a nonmarital birth at higher-order parities? We argued above that answering this question requires estimates of $b_2 + b_3$, since this quantity compares the pace of childbearing among two groups of women—those with marital and nonmarital first births—during the period when both groups are at risk of a nonmarital birth—that is, at durations when both groups are unmarried. Table 6 contrasts three alternative estimates of the quantity $b_2 + b_3$ to assess the sensitivity of our results to the proportionality assumption for differences between marital and nonmarital fertility. In rows labelled "1," we report estimates of $b_2 + b_3$ drawn from the proportional specification in Table 3. In rows labelled "2" and "3," we relax this assumption by letting the difference between marital and nonmarital fertility vary with duration by using a piecewise constant specification and the piecewise linear spline, respectively, as discussed above. As in Tables 4 and 5, likelihood ratio χ^2 values (on 1 df) are reported in parentheses.

[Table 6 about here]

Our results suggest that currently unmarried women who bore their first child outside of marriage have generally higher birth risks than currently unmarried women who bore their first child within marriage. In our baseline model for whites, this association for the transition from a first to second birth is .48 under a proportional hazard specification; hence, among women who are currently unmarried, those who bore their first child outside of marriage have a 61 percent higher risk (exp(.48) = 1.61) of a nonmarital second birth than those who bore their first child within marriage. This model also yields estimates for white women at higher parities that are strikingly similar in magnitude to those for the transition from a first to second birth. Relaxing the proportionality assumption (results in rows 2 and 3) yields slightly smaller estimates for whites, with changes in magnitude largest for second births and smallest for fourth births. For blacks, estimates from the baseline model are also positive and usually statistically significant.

The results in Table 6 suggest that estimates of $b_2 + b_3$ are slightly larger in the proportional hazard specification than in the two nonproportional specifications (rows labelled “2” and “3”), with the magnitude of this parameter declining slightly when controls are added in successive models. Nevertheless, we observe positive coefficients in all model specifications, at all parities, and for both white and black women. Estimates are statistically significant in most models and for most parities, except for the highest-order parities for black women.

The results in Table 7 parallel those in Table 5 by examining the brake hypothesis across different demographic subgroups of women. As in Table 6, we observe positive coefficients across most model specifications and parities, with estimates generally statistically significant.[8] Overall, we interpret the results in Tables 6 and 7 as generally consistent with predictions from the brake hypothesis.

[Table 7 about here]

DISCUSSION

It is widely presumed that a nonmarital first birth is but a prelude to many other births outside of marriage. For example:

[8] The most notable exceptions occur for third births to white women with 16 or more years of education and for second births to black women with 16 or more years of education.

> "On the right and among large numbers of blue-collar Democrats, there was increasing resentment at the permanency of welfare. It was acceptable to provide for the aged and disabled, they agreed. It was acceptable that a worker get unemployment checks while looking for a new job. But it was quite another thing for society to be supporting a healthy adult year after year.
>
> [As] the focal point for [this] resentment . . . AFDC evolved into the *bête noire* of the social welfare system. By the fifties it had become embarassingly, outrageously clear that most of these women were not widows. Many of them had not even been married. Worst of all, they didn't stop having babies after the first lapse. They kept having more."
>
> Charles Murray (1984), *Losing Ground*, p. 18.

Similarly:

> "[To] sociologists the astonishing rise in illegitimacy among the poor in recent decades is a great mystery, one of the points on which more research is required. . . . But if you are a fifteen-year-old girl in the ghetto, doing poorly at school, fighting with your mother, . . . you will want to escape . . . [Here is how] it will be possible.
>
> On your sixteenth birthday, the government will offer you a chance for independence, in an apartment of your own: free housing, medicine, legal assistance, and a combination of welfare payments and food stamps worth several hundred dollars a month. It may not seem much to a sociologist, but it is a package hugely beyond the pittance allowed you by your mother and far beyond the earnings capacity of any of your male acquaintances. It is all offered on one crucial condition. You must bear an illegitimate child. After three children, the payments in New York State will rise to $8,333, an amount 45 percent higher than the after-tax earnings of a full-time job at the minimum wage.
>
> If you are that fifteen-year-old girl, what do you do? Do you go for it? Or do you go through a rigamarole of inserting diaphragms or taking pills?"
>
> George Gilder (1986), *Men and Marriage*, pp. 91–92.

Perhaps more unexpected are equally unflattering views of the "undeserving poor" as expressed by welfare recipients themselves:

> "Howard Smith [an instructor in a welfare demonstration project and himself a past welfare recipient] asks, 'Does welfare encourage alcoholism?'

> 'When you're bored and don't know how to use your idle time, people want to escape from reality,' Timothy Wilson says.
>
> [Pearl Dawson, a welfare mother] is asked by Smith whether it is true that some of those on welfare prefer welfare, and have more children to qualify for more benefits.
>
> 'Yes. Have baby behind baby behind baby. . . Some people don't want to do no better,' Pearl says."
>
> Ken Auletta (1999), *The Underclass*, p. 90–91

In this paper, we ask if such widespread expectations are in fact reflected in the actual fertility behavior of recent cohorts of U.S. women. Before summarizing our empirical findings, however, we think it important to acknowledge that to some, our analyses may be beside the point. Consider, for example, the question of how one might deem nonmarital childbearing to be "high" or "low." One possible answer is to invoke an absolute standard—for example, one could hold that any level of nonmarital childbearing is too high. Another is to adopt a relative standard—for example, assessing levels of nonmarital fertility relative to the fertility of some comparison group. In this paper, we have adopted such a relative standard, but we acknowledge that the empirical findings we report could logically be held irrelevant by those who deem current levels of nonmarital childbearing to be "too high" relative to some absolute standard.

Most social scientists, however, do not adopt such a position. For this audience, we pose two essentially descriptive questions about nonmarital childbearing. A first asks if there an "engine" of nonmarital fertility. Answering this question, we argued, requires comparing the pace of childbearing for women who bear successive children within marriage with the pace of childbearing for women who bear successive children outside of marriage. A second question asks if a marital first birth functions, in effect, as a "brake" on the likelihood that a woman bears a later child outside of marriage. We view these questions as essentially descriptive because they do not identify a specific causal mechanism that might generate the hypothesized associations; similarly, our answers to these questions are likewise descriptive in nature because they also do not shed direct light on what mechanisms might underlie the empirical associations we observe.

We argued that answering these questions requires juxtaposing the risk of nonmarital childbearing for women who initiated motherhood inside versus outside marriage, comparing rates of nonmarital childbearing for these two groups of women during the times when they are unmarried. Thus, we compare the fertility of women with marital first births when they are observed to be unmarried—for example, when they experience a marital separation—with the fertility of women with nonmarital first births during the periods when they are observed to be unmarried. A central methodological point, then, is that the comparisons needed to evaluate the "engine" hypothesis are different from those required to evaluate the "brake" hypothesis.

Does a nonmarital first birth ignite an "engine" of subsequent nonmarital fertility? Our results suggest the answer is no. We find that the interval between a nonmarital first and nonmarital second birth is significantly longer than the interval between a marital first and marital second birth for both white and black women. These findings suggest a *slower* pace of nonmarital second births relative to the pace of marital second births—that is, a finding *opposite* to that predicted by the "engine" hypothesis. At higher birth orders, the interval between successive nonmarital births tends to be longer than that between successive marital births, particularly after adjusting for compositional differences. Our results show that a quicker pace of nonmarital births can be found, but only for fourth births to white women who were of particular childbearing cohorts and of particular ages. These results would appear difficult to reconcile with an engine hypothesis—proponents of such a hypothesis would need to explain why no engine of nonmarital fertility occurs at lower birth orders but emerges only after a third birth and only among certain demographic groups of white women. In the absence of such arguments, our findings would appear inconsistent with the assertion that a nonmarital first birth sparks an engine of subsequent nonmarital fertility.

Although this conclusion—that there appears to be no general engine of nonmarital fertility—runs counter to many common expectations—it is consistent with arguments made in other venues. For example, the finding that fertility is higher among married women than among unmarried women may reflect the lack of access to a regular sexual partner by unmarried women relative to married women. That is, because men in unmarried women's lives are more likely to depart than are men in married women's lives, all else being equal we might expect that the fertility of unmarried women will on average be lower than that for married women. Similarly, although the consensus emerging from the economic literature examining the linkage between welfare and demographic behaviors is that there exists a real but small effect of AFDC on demographic outcomes, most observers would place equal emphasis on "small" and "real" (Moffitt 1998). These issues take on additional weight when viewed in light of findings regarding the serious economic constraints faced by many unmarried women in making ends meet from month to month (Edin and Lein 1997). It is worth emphasizing that while our results are consistent with the hypothesis that the sexual activity of unmarried mothers is, on average, lower than that of married mothers, such an argument represents a considerable departure from more traditional depictions of unmarried mothers as sexually promiscuous, views which may have contributed historically to perceptions of such women as "undeserving" of assistance (Gordon 1994). Although our analyses do not shed any direct light on these issues, our findings nevertheless suggest that a variety of factors, including the greater instability of intimate relationships the economic and social disincentives to bear a child outside of marriage, may outweigh incentive effects on nonmarital fertility of factors such as AFDC.

Does a marital first birth function as a "brake" on subsequent nonmarital fertility? Our answer is yes. We find that women who initiated childbearing within marriage but who then separated have rates of a second birth while unmarried that are moderately lower relative to their unmarried counterparts who initiated childbearing outside of marriage. These differentials persist beyond a second birth, holding for third and fourth births. Moreover, our models suggest similar effects for both white and black women; indeed, our estimates are strikingly similar across parities *and* race. Taken as a whole, we interpret these results as consistent with the view that a marital first birth acts as a “brake” on subsequent nonmarital fertility.

What might generate this association between a marital first birth and lower rates of subsequent nonmarital childbearing? We have no answer to this question—our data do not let us say anything about the possible mechanisms underlying such an association. It may be, for example, that the experience of a birth within marriage alters the preferences of women, in ways that lead such women to postpone their fertility following a marital dissolution, at least until they have entered a subsequent marriage. This might occur if having a child within marriage alters a woman’s perception of the social and economic resources needed or that would be desirable to bear and raise additional children. Or it may be that a preference to bear a child within marriage is reflective of norms or views formed prior to a first marriage, so that marriage effectively selects on those women who, for various reasons, choose to bear their children within marriage. Thus, our finding of an apparent “brake” on nonmarital fertility, while provocative, does not shed light on what mechanisms might generate this association. Still, our results make clear that a marital first birth is unlikely to be followed by a second or higher-order nonmarital birth.

Do our findings carry any implications for marriage? One way to pose this question is to ask what our results say about the assertion that childbearing has become decoupled from marriage (Murray 1984; Whitehead 1993). To us, our results raise doubts about such assertions, and we find little evidence that this decoupling occurs even for women who experience a nonmarital first birth. For women with a marital first birth, our results are consistent with the view that such women prefer to bear later children within marriage, with our results indicating that such women lower their fertility upon marital separation or dissolution. For women with a nonmarital first birth, our results show that such women do *not* have high levels of fertility while unmarried, when their parity- and duration-specific fertility rates are compared to those for women who are married. Thus, one plausible, albeit speculative, interpretation of our results is that both groups of women appear to view marriage as the preferred venue for childbearing, and that such a view is reflected in observed behavior, with married women having higher fertility rates relative to their unmarried counterparts.

While such results might appear at first glance to be difficult to square with well-known increases in the proportion of births to unmarried women (Ventura et al. 1995; Wu, Bumpass, and Musick 2001), it is important to emphasize that our findings proceed conditional on marital status at first birth and speak to women's underlying fertility rates. Because the so-called nonmarital fertility ratio is given by the ratio of births to unmarried women to births to all women, increases in this proportion could be produced by increases in in the proportion of women with a nonmarital first birth even if nonmarital fertility was lower than marital fertility at higher birth orders. Similarly, increases in the nonmarital fertility ratio could be produced by declines in the probability of marriage following a nonmarital first birth. Finally, it is important to emphasize that the *initiation* of nonmarital childbearing remains heavily concentrated among young women—those in their teen years or in their early 20s. One consequence of this differential in the age at which childbearing begins is that women with a nonmarital first birth will, on average and all else being equal, be exposed to the risk of a second birth over a longer span than women with a marital first birth, simply by virtue of their earlier age at entry into motherhood. This implies that even if the *pace* of childbearing is lower for unmarried women than for married women, completed family size for women with a nonmarital first birth may equal, or even exceed, that for women with a marital first birth. Thus, larger numbers of births can result from a longer exposure to risk, even if underlying fertility rates are lower. Finally, our results for higher-order births differ somewhat across observed demographic characteristics of women. While our findings are necessarily somewhat tentative given smaller sample sizes at higher-order births, our results nevertheless suggest that high third and fourth birth rates may be influenced as much by age at first birth as by a nonmarital first birth. If so, these factors could also lead unmarried women to have more children than married women. Thus, our findings—that nonmarital fertility rates are generally lower than marital fertility rates—need not be inconsistent with the possibility that substantial numbers of women who are unmarried at first birth will proceed to higher birth orders.

Our results concerning the variation in birth rates with duration since a previous birth suggest yet other complications. At most durations after a previous birth, nonmarital second birth rates are lower than marital second birth rates, and nonmarital third and fourth birth rates are about the same as marital third and fourth birth rates. However, during the first 20 months or so after a previous birth, nonmarital second birth rates are comparable to marital birth rates, and nonmarital third and fourth birth rates are higher than corresponding marital birth rates. Thus, although overall nonmarital birth rates may be low, a significant fraction of unmarried mothers experience rapid and repeated childbearing postpartum. Some portion of this differential may be due to the poorer quality of the data for these women—for example, reports of births at unlikely intervals—but some portion of this differential may represent behavioral differences among the pool of unmarried women.

Although our observations about variation in birth rates with duration and parity, and about variation in exposure to risk, all carry implications for the average completed fertility of unmarried mothers, we think it at least as important to emphasize that these factors will also affect the *range* of fertility-related outcomes. For example, our estimates of marital birth rates suggest that the completed fertility of married women who have had one birth will be two to three children, and that the interval between successive births typically ranges from two to four years. By contrast, our estimates for nonmarital births indicate much greater heterogeneity in the timing and number of children for unmarried women. This finding is consistent with the descriptive results in Wu, Bumpass, and Musick (2001), who report striking diversity in fertility and marital trajectories of unmarried women in the period following a first out-of-wedlock birth. As a result, women's marital and fertility trajectories following an out-of-wedlock first birth span a very wide range of circumstances: some of these women may experience a closely-spaced nonmarital second birth followed by additional closely-spaced births, others a marriage after a first birth followed by a second birth within this marriage, still others a long interval between a first and second nonmarital birth, and yet another group experiencing no second birth at all. This observation leads us to speculate that such diversity may imply a correspondingly wide range of consequences for women and children, with some situations likely conferring significant disadvantage, but others conceivably carrying much more positive consequences. Finally, we note that while even a casual inspection of available data makes apparent the very wide variety of women's situations following a nonmarital first birth, some of these situations appear to contribute more to public perceptions of nonmarital fertility than others.

Our analytic models of nonmarital and marital fertility have focused primary attention on the dynamics underlying women's progression from a first to second birth, and from second births to yet higher order births; by contrast, we devote much less effort to modeling women's entries and exits from marriage. It is thus somewhat paradoxical that our results appear to provide little support for the contention that a nonmarital birth either alters or reveals a woman's "desire" to bear additional children outside of marriage, but instead point to the large role played by marriage in influencing women's trajectories of childbearing, both within and outside of marriage. As such, these results join a growing body of empirical evidence suggesting that transformations in marriage, not childbearing, are key to understanding continuing trends in nonmarital fertility.

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Table 1: Variables in parametric models.

Current marital status and marital status at first birth

1. 1 if unmarried at duration u; else 0 (time-varying)
2. 1 if unmarried at first birth; else 0
3. 1 if unmarried at duration u and unmarried at first birth; else 0 (time-varying)

Period

4. 1 if first/second/third birth in 1970–74; else 0 (omitted category)
5. 1 if first/second/third birth in 1975–79; else 0
6. 1 if first/second/third birth in 1980–84; else 0
7. 1 if first/second/third birth in 1985–89; else 0
8. 1 if first/second/third birth in 1990 or later; else 0

Age

9. 1 if 18 or younger at first birth; else 0
10. 1 if 19 or younger at first birth; else 0
11. 1 if 20–24 at first birth; else 0 (omitted category)
12. 1 if 25–29 at first birth; else 0
13. 1 if 30 or older at first birth; else 0

Highest grade completed

14. 1 if less than 12 years of schooling completed; else 0
15. 1 if 12 years of schooling; else 0 (omitted category)
16. 1 if 13–15 years of schooling; else 0
17. 1 if 16 or more years of schooling; else 0

Interactions

Interactions of variable 1 by the period, age, and education variables.

Table 2: Marital and nonmarital fertility transitions through the fourth birth. Births to white and black women, 1970–1995. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \to 2$	$2 \to 3$	$3 \to 4$	$1 \to 2$	$2 \to 3$	$3 \to 4$
marital next birth	17093.0	7002.7	2371.7	1648.2	969.8	467.7
nonmarital next birth	2161.1	1002.5	355.1	2113.7	1063.0	470.1
weighted n at risk	41339.4	32525.8	14439.4	8045.5	6215.4	3385.0

Table 3: Proportional hazard estimates of marital status at first birth and current marital status. Second and higher-order births during 1970–74 to women aged 20–24 with 12 years of completed schooling. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \to 2$	$2 \to 3$	$3 \to 4$	$1 \to 2$	$2 \to 3$	$3 \to 4$
Baseline model						
1. unmarried at u	−1.24***	−.29***	−.22**	−.66***	−.18†	−.15
	(.04)	(.05)	(.08)	(.10)	(.11)	(.14)
2. nonmarital first birth	.13***	.44***	.15*	.18**	.37***	.09
	(.03)	(.03)	(.06)	(.05)	(.06)	(.09)
3. unmarried at u × nonmarital first birth	.35***	−.02	.36**	.26*	.13	.16
	(.05)	(.07)	(.12)	(.11)	(.12)	(.16)
Baseline model and birth cohorts						
1. unmarried at u	−1.25***	−.39***	−.41**	−.79***	−.27*	−.17
	(.06)	(.08)	(.13)	(.11)	(.13)	(.18)
2. nonmarital first birth	.12***	.44***	.16**	.17**	.37***	.12
	(.03)	(.03)	(.06)	(.05)	(.06)	(.09)
3. unmarried at u × nonmarital first birth	.35***	−.07	.29*	.24*	.12	.16
	(.06)	(.07)	(.12)	(.11)	(.12)	(.16)
Baseline, birth cohorts, and age						
1. unmarried at u	−1.31***	−.51***	−.42*	−.93***	−.40*	−.59*
	(.06)	(.09)	(.15)	(.12)	(.15)	(.22)
2. nonmarital first birth	.02	.27***	.10	.01	.20*	.06
	(.03)	(.04)	(.06)	(.06)	(.07)	(.10)
3. unmarried at u × nonmarital first birth	.39***	.02	.36**	.26*	.10	.18
	(.06)	(.08)	(.13)	(.11)	(.13)	(.17)
Baseline, birth cohorts, age, and education						
1. unmarried at u	−1.21***	−.34***	−.43*	−.78***	−.34*	−.53*
	(.07)	(.10)	(.16)	(.13)	(.15)	(.23)
2. nonmarital first birth	.02	.26***	.08	.00	.16*	.02
	(.03)	(.04)	(.06)	(.06)	(.07)	(.10)
3. unmarried at u × nonmarital first birth	.36***	.00	.34*	.22†	.09	.15
	(.06)	(.08)	(.13)	(.11)	(.13)	(.17)

† $p < .10$ * $p < .05$ ** $p < .005$ *** $p < .0005$ (two-tailed test; se's in parentheses)

Table 4: Proportional hazard estimates of the parameter $b_1+b_2+b_3$. Second and higher-order births during 1970–74 to women aged 20–24 with 12 years of completed schooling. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$
Baseline model						
$b_1 + b_2 + b_3$	$-.76^{***}$	$.13^{*}$	$.30^{***}$	$-.22^{***}$	$.32^{***}$	.10
	(1012.35)	(7.83)	(12.72)	(34.19)	(34.01)	(1.47)
Baseline model and birth cohorts						
$b_1 + b_2 + b_3$	$-.78^{***}$	$-.02$	.05	$-.38^{***}$	$.22^{*}$	.10
	(257.67)	(.04)	(.11)	(28.58)	(4.49)	(.45)
Baseline, birth cohorts, and age						
$b_1 + b_2 + b_3$	$-.90^{***}$	$-.23^{*}$	.03	$-.66^{***}$	$-.10$	$-.35^{\dagger}$
	(242.80)	(5.48)	(.04)	(59.43)	(.62)	(3.04)
Baseline, birth cohorts, age, and education						
$b_1 + b_2 + b_3$	$-.83^{***}$	$-.09$	.00	$-.56^{***}$	$-.09$	$-.36^{\dagger}$
	(168.55)	(.69)	(.00)	(35.51)	(.41)	(2.93)

Note: Likelihood ratio χ^2 (on 1 df) reported in parentheses.

$^{\dagger}\ p < .10$ $\quad ^{*}\ p < .05$ $\quad ^{**}\ p < .005$ $\quad ^{***}\ p < .0005$

Table 5: Proportional hazard estimates of the parameter $b_1 + b_2 + b_3$ for other socio-demographic groups. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$
Year of birth						
1975–79	−.80***	−.01	−.21	−.38***	−.09	−.28
	(176.23)	(.02)	(1.21)	(17.46)	(.41)	(1.78)
1980–84	−.91***	−.12	.39*	−.42***	.03	−.40†
	(257.64)	(1.64)	(5.69)	(20.88)	(.06)	(3.75)
1985–89	−.94***	−.05	.42*	−.45***	−.24†	−.20
	(226.23)	(.22)	(5.95)	(18.80)	(3.15)	(.90)
1990–95	−.68***	.43**	.62**	−.38*	.17	−.14
	(60.35)	(10.98)	(8.38)	(5.36)	(.67)	(.23)
Age at first birth						
12–19	−.83***	−.09	−.14	−.48***	.06	.05
	(182.67)	(.87)	(.73)	(28.54)	(.23)	(.08)
25–29	−.95***	−.02	.12	−.64***	−.13	.32
	(117.80)	(.01)	(.23)	(22.23)	(.34)	(1.10)
30 or older	−.80***	.59*	.94**	−.32†	.08	.57
	(41.41)	(6.98)	(8.37)	(3.18)	(.07)	(1.65)
Highest grade completed						
0–11	−.36***	.14	.07	−.50***	−.04	−.25
	(20.60)	(1.22)	(.14)	(18.33)	(.07)	(1.13)
13–15	−1.03***	−.38**	−.04	−.76***	−.21	−.81**
	(232.58)	(11.86)	(.05)	(56.92)	(1.92)	(10.75)
16 or more	−1.23***	−.84***	−.05	−1.15***	−.39†	−.51
	(213.55)	(32.33)	(.04)	(72.27)	(2.78)	(1.92)

Note: Likelihood ratio χ^2 (on 1 df) reported in parentheses. See also text for additional details.

† $p < .10$ * $p < .05$ ** $p < .005$ *** $p < .0005$

Table 6: Estimates of $b_2 + b_3$ from a proportional and two nonproportional hazard specifications. Second and higher-order births during 1970–74 to women aged 20–24 with 12 years of completed schooling. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \to 2$	$2 \to 3$	$3 \to 4$	$1 \to 2$	$2 \to 3$	$3 \to 4$
Baseline model						
1. $b_2 + b_3$	.48***	.42***	.52***	.44***	.50***	.25†
	(104.83)	(43.87)	(23.25)	(23.28)	(26.28)	(3.77)
2. $b_2 + b_3$	.35***	.35***	.47***	.40***	.48***	.25†
	(52.50)	(29.87)	(19.37)	(18.66)	(23.85)	(3.69)
3. $b_2 + b_3$	.37***	.35***	.48***	.40***	.50***	.26*
	(56.75)	(28.58)	(19.42)	(19.50)	(25.63)	(3.99)
Baseline model and birth cohorts						
1. $b_2 + b_3$	.47***	.37***	.45***	.41***	.49***	.28*
	(97.74)	(32.97)	(17.68)	(20.19)	(24.44)	(4.37)
2. $b_2 + b_3$	.35***	.32***	.42***	.37***	.47***	.28*
	(51.16)	(24.06)	(15.15)	(16.48)	(22.60)	(4.36)
3. $b_2 + b_3$	.36***	.32***	.42***	.38***	.49***	.28*
	(54.37)	(23.15)	(15.21)	(17.34)	(24.16)	(4.64)
Baseline, birth cohorts, and age						
1. $b_2 + b_3$	.38***	.25***	.43***	.22*	.25*	.17
	(60.80)	(14.51)	(15.09)	(5.32)	(5.64)	(1.49)
2. $b_2 + b_3$	.29***	.24***	.42***	.24*	.29*	.24†
	(35.05)	(12.59)	(14.38)	(6.16)	(7.48)	(3.16)
3. $b_2 + b_3$	.36***	.32***	.42***	.38***	.49***	.28*
	(54.37)	(23.15)	(15.21)	(17.34)	(24.16)	(4.64)
Baseline, birth cohorts, age, and education						
1. $b_2 + b_3$	.38***	.25***	.43***	.22*	.25*	.18
	(60.98)	(14.24)	(14.99)	(5.52)	(5.80)	(1.79)
2. $b_2 + b_3$	.27***	.21**	.40***	.19*	.24*	.17
	(29.38)	(9.95)	(12.98)	(3.89)	(5.12)	(1.57)
3. $b_2 + b_3$	.31***	.23**	.42***	.25*	.30**	.25†
	(37.45)	(11.99)	(14.48)	(6.69)	(8.28)	(3.38)

Note: Likelihood ratio χ^2 (on 1 df) reported in parentheses. See also text for additional details.

† $p < .10$ * $p < .05$ ** $p < .005$ *** $p < .0005$

Table 7: Estimates from proportional and two nonproportional hazard specifications for the parameter $b_2 + b_3$ for other socio-demographic groups. June 1990 and 1995 Current Population Surveys.

	Whites			Blacks		
	$1 \to 2$	$2 \to 3$	$3 \to 4$	$1 \to 2$	$2 \to 3$	$3 \to 4$
Year of birth						
1975–79						
1.	.41***	.33*	.22	.40**	.25	.25
	(24.19)	(7.14)	(.95)	(9.06)	(2.09)	(1.01)
2.	.29**	.28*	.19	.37*	.24	.25
	(11.91)	(5.38)	(.70)	(7.63)	(1.96)	(1.03)
3.	.32***	.28*	.20	.38**	.26	.26
	(14.05)	(5.27)	(.77)	(8.16)	(2.26)	(1.10)
1980–84						
1.	.29***	.22†	.82***	.36*	.37*	.12
	(13.94)	(3.41)	(17.96)	(7.59)	(5.01)	(.29)
2.	.18*	.17	.79***	.31*	.35*	.14
	(5.17)	(2.13)	(16.32)	(5.64)	(4.62)	(.36)
3.	.22*	.17	.81***	.33*	.39*	.15
	(7.59)	(2.09)	(16.98)	(6.36)	(5.42)	(.41)
1985–89						
1.	.26**	.29*	.84***	.33*	.10	.33
	(9.86)	(5.90)	(18.45)	(5.79)	(.32)	(1.90)
2.	.09	.21†	.81***	.24†	.06	.35
	(1.11)	(2.94)	(16.43)	(3.03)	(.12)	(2.11)
3.	.13	.21†	.83***	.28†	.11	.37
	(2.44)	(2.78)	(17.19)	(3.79)	(.41)	(2.42)
1990–95						
1.	.53***	.77***	1.05***	.40*	.50*	.39
	(24.58)	(28.68)	(19.42)	(4.48)	(5.10)	(1.65)
2.	.27*	.62***	.93***	.25	.44†	.42
	(6.10)	(18.03)	(14.74)	(1.72)	(3.80)	(1.88)
3.	.31**	.61***	.95***	.30	.51*	.45
	(7.88)	(17.32)	(15.23)	(2.36)	(4.97)	(2.20)

Table 7: (continued)

	Whites			Blacks		
	$1 \to 2$	$2 \to 3$	$3 \to 4$	$1 \to 2$	$2 \to 3$	$3 \to 4$
Age at first birth						
12–19						
1.	.38***	.25*	.28†	.30*	.39*	.58*
	(31.42)	(6.94)	(3.29)	(6.37)	(7.85)	(7.63)
2.	.26***	.20*	.26†	.25*	.38*	.58*
	(14.74)	(4.66)	(2.82)	(4.47)	(7.26)	(7.64)
3.	.28***	.20*	.26†	.27*	.40**	.59**
	(16.38)	(4.42)	(2.83)	(4.92)	(7.93)	(7.93)
25–29						
1.	.26*	.32*	.54*	.14	.21	.84*
	(7.49)	(4.82)	(5.20)	(.75)	(.76)	(6.04)
2.	.16†	.29†	.53*	.12	.20	.86*
	(2.94)	(3.80)	(4.92)	(.54)	(.72)	(6.30)
3.	.17†	.28†	.53*	.12	.22	.86*
	(3.18)	(3.67)	(4.93)	(.61)	(.82)	(6.25)
30 or older						
1.	.41**	.93***	1.36***	.46*	.41	1.10*
	(8.58)	(16.86)	(17.04)	(5.35)	(1.93)	(5.42)
2.	.34*	.89***	1.31***	.45*	.40	1.09*
	(5.91)	(15.52)	(15.97)	(5.17)	(1.83)	(5.36)
3.	.34*	.88***	1.31***	.46*	.42	1.11*
	(5.87)	(15.22)	(15.79)	(5.19)	(1.96)	(5.55)

Table 7: (continued)

	Whites			Blacks		
	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$
Highest grade completed						
0–11						
1.	.84***	.47***	.50*	.28*	.29†	.28
	(109.94)	(18.84)	(7.77)	(4.34)	(3.66)	(1.89)
2.	.71***	.42***	.45*	.23†	.28†	.28
	(75.66)	(14.29)	(6.31)	(2.83)	(3.19)	(1.86)
3.	.72***	.41***	.46*	.24†	.30†	.30
	(78.34)	(14.08)	(6.41)	(3.25)	(3.75)	(2.12)
13–15						
1.	.17*	−.04	.38*	.02	.13	−.29
	(5.08)	(.16)	(4.32)	(.02)	(.65)	(1.45)
2.	.07	−.08	.35†	−.01	.12	−.28
	(.75)	(.50)	(3.66)	(.00)	(.59)	(1.39)
3.	.08	−.08	.36†	.00	.14	−.28
	(1.06)	(.56)	(3.72)	(.00)	(.72)	(1.34)
16 or more						
1.	−.02	−.50**	.38	−.37*	−.05	.02
	(.06)	(10.93)	(2.40)	(5.00)	(.04)	(.00)
2.	−.12	−.53***	.34	−.39*	−.05	.01
	(1.61)	(12.46)	(1.97)	(5.66)	(.05)	(.00)
3.	−.11	−.54***	.34	−.39*	−.04	.02
	(1.37)	(12.61)	(1.97)	(5.44)	(.03)	(.00)

Note: Likelihood ratio χ^2 (on 1 df) reported in parentheses. See also text for additional details.

† $p < .10$ * $p < .05$ ** $p < .005$ *** $p < .0005$

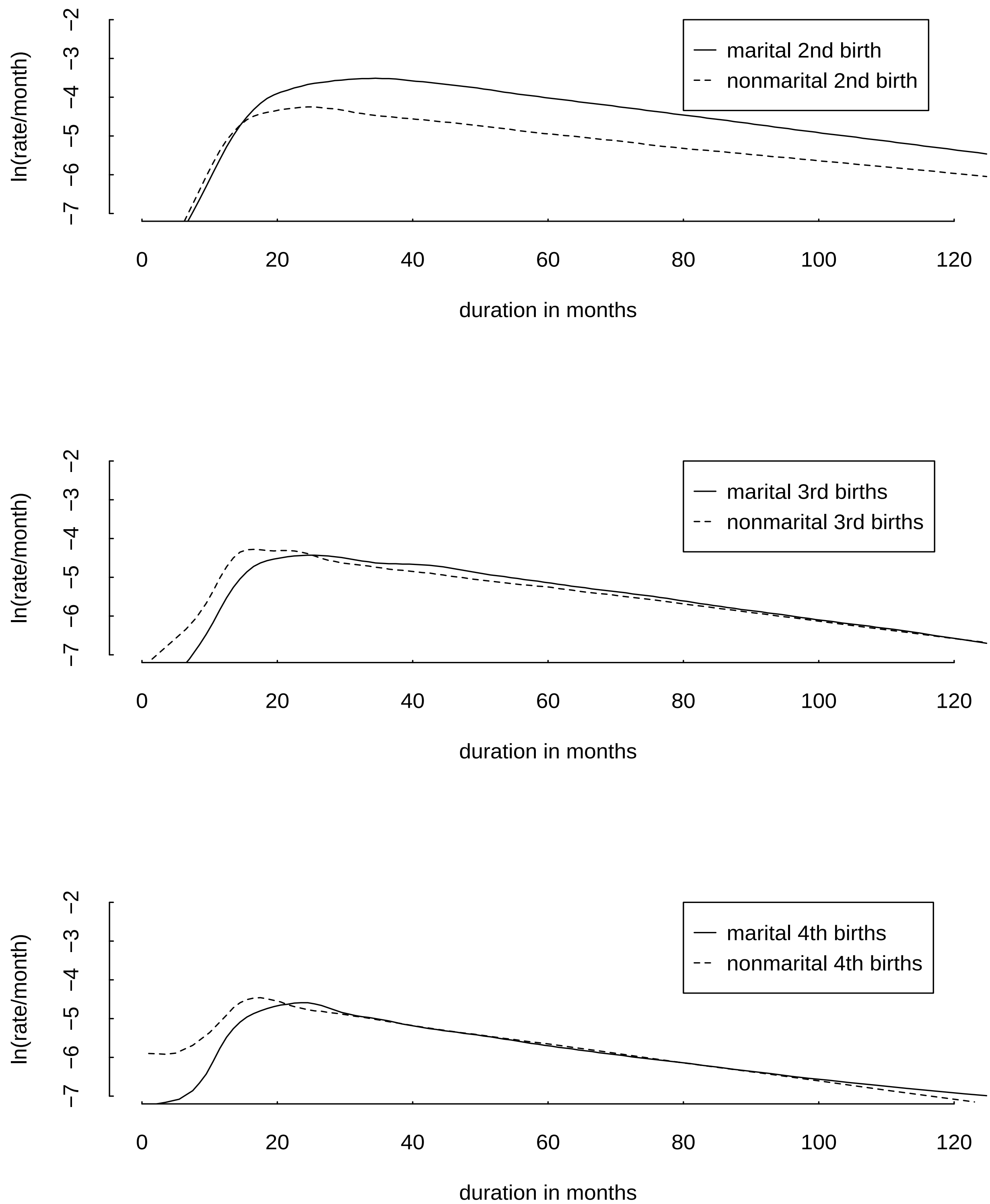

Figure 3: Transitions to a second, third, and fourth marital and nonmarital birth. Births to U.S. women between 1970–1995.

Source: June 1990 and 1995 Current Population Surveys.

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