**Change in the Stability of U.S. Marital and Cohabiting Unions Following the Birth of a Child**

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**Abstract.** The share of births to cohabiting couples has increased dramatically in recent decades. These families tend to be less stable than those formed in marriage, with potential implications for the well-being of parents and children. We use data from the 1995 and 2006-10 National Survey of Family Growth (NSFG) to: 1) describe change in the characteristics of couples having children together, paying attention to trajectories of cohabitation and marriage around the couple's first birth; 2) compare change in union stability following the birth of a child across four distinct union-birth trajectories; and 3) illustrate change in patterns of stability using simple simulations. Relying on multivariate event history models, we find evidence of a weakening association between cohabitation and instability, given marriage occurs at some point *before or after* the couple's first birth. The more recent data show statistically indistinguishable separation risks for couples who have a birth in marriage without ever cohabiting, who cohabit and then have a birth in marriage, and who have a birth in cohabitation and then marry. Cohabitating unions with children are significantly less stable when de-coupled from marriage, although the parents in this group also differ most from others on observed (and likely, unobserved) characteristics.

The share of births to unmarried women has almost doubled over the past 25 years: from 22% in 1985 to 41% in 2009 (Martin et al. 2011). The shift from marital to cohabiting births accounts for much of the increase over this period, particularly in the past decade (Kennedy and Bumpass 2011; Martinez, Chandra, and Daniels 2012; Raley 2001). As of the mid-2000s, 59% of nonmarital births—or 21% of all births—were to cohabiting parents (Lichter 2012). How we evaluate the implications of these changes for child well-being depends on our understanding of cohabitation as a normative setting for having and raising children, which in turn depends critically on the stability of cohabiting families for children.

From the perspective of children, living with two cohabiting parents in many ways resembles living with two married parents, with two potential earners and care-takers in the household. But family transitions are negatively associated with child well-being (Fomby and Cherlin 2007; Wu 1996; Wu and Martinson 1993), and cohabiting families are on average significantly less stable than married-parent families (Manning, Smock, and Majumdar 2004; Wu and Musick 2008). Children living with a cohabiting mother express greater ambiguity about their families, as reflected in differences in their reports of family structure as compared to their mothers' (Brown and Manning 2009). Children in cohabiting families also tend to exhibit elevated behavioral and emotional problems and less engagement with school than those in married-parent families (Brown 2004). Thus while the living arrangements of cohabiting and married-parent families may be similar, evidence points to differences in who selects into these arrangements and in what they mean to the families involved.

Change in the stability of cohabitation is important in assessing the family contexts of children; it also speaks to broader discussions about the evolving role of cohabitation in the family system. From the perspective of the second demographic transition theory, cohabiting families should become more normative and look increasingly like marriage, for example, in their chances of staying together (van de Kaa 1987). A contrasting view frames cohabitation as a second-best alternative when marriage's steep emotional and financial prerequisites cannot be met (Cherlin 2009; Furstenberg 19996), suggesting growing differences between cohabitation and the increasingly privileged state of marriage. Prior work has shown that the instability of cohabitation—after increasing in the 80s and early 90s—appears to have leveled off since 1995 (Kennedy and Bumpass 2008; Kennedy and Bumpass 2011). But multivariate analyses have not been done to assess the extent to which observed trends reflect change in the meaning of cohabitation versus change in the factors selecting men and women into cohabitation.

We draw on data from the 1995 and 2006-10 National Survey of Family Growth (NSFG) to examine change in the stability of marital and cohabiting unions following the birth of a child, paying attention to trajectories of cohabitation and marriage around the couple's first birth. Our analysis has three steps: First, we describe change in the characteristics of couples having children over time, looking in particular at their shifting sociodemographic composition and prior union and childbearing histories. Second, we use discrete-time event history analysis to compare change in stability following childbirth across four distinct union-birth trajectories. Although standard to think of the risk of union dissolution from the start of coresidence, we begin clocking risk at the time of the couple's first birth in a coresidential union, reflecting our primary interest in the stability of unions for children (see also Manning et al. 2004; Wu and Musick 2008). Finally, the paper illustrates change in patterns of stability using simple simulations, generating predicted probabilities of union dissolution altering assumptions about union formation and the composition of unions over time.

## BACKGROUND

### Why might we expect changes in the stability of cohabitation over time?

Conceptually, it is unclear how the meaning of cohabitation *should be* changing. In analyzing patterns of family change across Europe, Kiernan (2000) posited a series of stages in the second demographic transition, with cohabitation emerging as a marginalized behavior and gradually becoming an accepted family form. Along the way, distinctions between cohabitation and marriage fade, and cohabitation transitions from a short-term and largely childless state to a much more stable arrangement in which having and raising children is commonplace. Predictions based on the second demographic transition are consistent with Cherlin's (2004) deinstitutionalization hypothesis, although the mechanisms differ. Cherlin argues that marriage is undergoing a process of deinstitutionalization while the social norms defining partners' behavior in cohabitation are becoming stronger, implying blurring boundaries between the two. The process of deinstitutionalization implies convergence in expectations around marriage and cohabitation, leading to growing similarity in childbearing behavior and relationship stability.

An opposing view suggests persistent differences and potentially divergence in the experiences of marriage and cohabitation. This draws more heavily on ideas emphasizing the growing symbolic significance of marriage as a marker of prestige (Cherlin 2009). Qualitative and quantitative accounts report that men and women of all education levels place a high value on marriage but perceive substantial economic prerequisites (Carlson et al. 2004; Edin and Kefalas 2005; Gibson-Davis 2009; Gibson-Davis, Edin, and McLanahan 2005; Smock, Manning, and Porter, 2005). Short of these prerequisites, couples opt into cohabitation, and cohabitation thus becomes the "budget" route into family formation (Furstenberg 1996). This is consistent with McLanahan's (2004) discussion of the differential impact of the second demographic transition on women, with associated economic and ideational changes undermining stable relationships for women at the bottom of the education distribution and strengthening them for women at the top. In their cross-national investigation of cohabiting fertility, Perelli-Harris et al. (2010) emphasize the link between economic instability and the impermanence of cohabitation. Together, these strands of research suggest that despite increases in cohabiting fertility, the experiences of marital and cohabiting families may remain distinct – and potentially even diverge over time. In particular, cohabitation may remain a less stable union form, and potentially less stable over time relative to marriage.

**What has recent data shown?**

Cohabitations tend to be short in duration, with most transitioning to marriage or dissolving within two years (Kennedy and Bumpass 2011, Table 4; Lichter, Qian, and Mellott 2006). Marriage used to be the more common exit out of cohabitation, but in recent years the link between cohabitation and marriage has weakened, and dissolution now accounts for the greater share of exits from cohabitation (Bumpass and Lu 2000; Kennedy and Bumpass 2011; Lichter, Qian, and Mellott 2006). Dissolution risks from cohabitation increased over the 1980s and early 1990s (Bumpass and Lu 2000), but appear to have leveled off between 1990-94 and 2002-07 (Kennedy and Bumpass 2001). This leveling off has also stalled the amount of union instability experienced by children: Kennedy and Bumpass (2011) report that despite increases in the share of children born to cohabiting parents, the percent of children experiencing parental separation by age 12 has remained stable. This recent leveling off of instability and the longer-term declines in transitions into marriage are consistent with ideas from the second demographic transition theory that cohabitation is becoming a more normative family context—one that may grow to more closely resemble marriage (e.g., in terms of stability) in the longer term.

In addressing potential explanations for the growing instability of cohabitation, Bumpass and Lu (2000) speculated that as cohabitation became more widespread, there would be fewer barriers to entry and couples would select into cohabitation at lower levels of commitment. As noted, others have reflected on how the institutionalization of cohabitation within the family system may be evolving (e.g., Cherlin 2009; Kiernan 2000). The papers cited above provide a rich descriptive portrait of changing family life, but they do not provide a multivariate framework for assessing period change, and thus questions remain as to what accounts for observed patterns. We focus specifically on couples having children together and address the extent to which observed trends in cohabitation reflect change in the factors selecting men and women into cohabitation versus change in the meaning of cohabitation.

**Changes in the composition of cohabiting families**

The composition of men and women opting into cohabitation affects the stability of these relationships, net of broader shifts in shared understandings of the nature and meaning of cohabitation. Education is a critical factor, associated both with selection into cohabitation and the stability of unions (Bumpass and Lu 2000; Kennedy and Bumpass 2008; Martin 2006). Cohabitation has always been somewhat more common among less advantaged men and women, although distinctions are much more sharply graded when childbearing is involved. In recent years, distinctions in cohabiting fertility by education have blurred along the lower end of the education distribution but not at the very top: between 1997-2001 and 2002-07, the proportion of births occurring within cohabitation increased 40% among women with some college (from 15% to 21%), 43% among women with a high school degree (from 23% to 33%), and 13% among those with no high school degree (from 32% to 36%) (Kennedy and Bumpass 2011, Table 6). This represents a significant shift in the education distribution of cohabiting fertility *up to* the ranks of the college-educated, for whom any childbearing out of marriage continues to be very rare: In both periods, just 3% of all births to college graduates were to cohabiting women.

The implications of changes in education patterns for the stability of families are not entirely clear. College graduates are increasingly distinct in their hold on childbearing in marriage, and the association between college and marital stability has strengthened over time (Raley and Bumpass 2003; Martin 2006). Nonetheless, relative to women who do not complete high school, those with a high school degree or some college should have more stable unions, and cohabiting family formation has moved especially rapidly into these educational ranks. Higher average levels of education among cohabiting parents may promote stability, although perhaps not relative to married parents, who are increasingly selected on college graduation.

Changes in the composition of cohabitors on the basis of prior union and childbearing histories may also factor into changes in the stability of cohabiting families—as well as the relative stability of marriage and cohabitation. The prevalence of cohabitation with more than one partner prior to marriage (or "serial cohabitation") has risen over time (Lichter, Turner, and Sassler 2010). Men and women with a history of more than one cohabitation tend to be disadvantaged socioeconomically and report lower marriage expectations and chances of marriage (Lichter et al. 2010; Lichter and Qian 2008; Cohen and Manning 2010). Cohabiting parents may be more likely to have a history of cohabitation with another partner, and research on marital dissolution suggests that this may affect the stability of the current union (Lichter and Qian 2008; Teachman 2003). The presence of children from another relationship (or "multipartnered fertility") has also risen and is more prevalent among unmarried parents (Carlson and Furstenberg 2006; Guzzo and Furstenberg 2007a, 2007b). Based on what we know about marital dissolution, the growing complexity of families formed out of marriage may lead to greater instability among cohabiting versus married-parent families. Yet as noted by Edin and Tach (2012:199), we know little about the stability of unions when unmarried mothers repartner.

The characteristics of cohabiting parents may be changing relative to married parents in other ways that could account for change in the relative stability of these families. In addition to education and family complexity, we account for race and ethnicity, family background characteristics, parents' ages at birth, birth intendedness, and subsequent fertility within the union, all factors associated with union stability (e.g., Phillips and Sweeney 2006; Teachman 2002).

**APPROACH**

Recent reports have highlighted the rise of births to cohabiting women (e.g., Martinez, Daniels, and Chandra 2012), and understanding the implications of this change for child well-being and more broadly for cohabitation's evolving place in the family system depends critically on the stability of cohabiting families. We examine change in the stability of cohabiting and marital unions following the couple's first birth. We recognize that union status at the time of the birth represents only a snapshot of family life, and thus we examine union-birth trajectories combining information on the couple's union status at the start of the union, time of the birth, and up to ten years following the birth. We account for a host of respondent and union characteristics; where possible, we also account for partner characteristics. There have been important shifts in the educational distribution of births to cohabitors and growing concern over increases in the complexity of families and implications for subsequent life histories (Kennedy and Bumpass 2011; Lichter et al. 2010; Lichter and Qian 2008). But these have not been examined in the context of change in the stability of unions for children. Using an event history framework and other descriptive tools, ours is the first analysis to our knowledge to examine change in the stability of marital and cohabiting families and the factors contributing to this change over time.

## DATA AND METHOD

### NSFG

We use data from the 1995 and 2006-10 NSFG, nationally representative fertility surveys of reproductive-age women ages 15–44 (Abma et al. 1997; NCHS 2011). Interviews were conducted in person and include retrospective histories of childbirth, marriage, cohabitation, education (to varying degrees), and pregnancy intentions. The NSFG was conducted six times between 1973 and 2002. In 2006, the National Center for Health Statistics moved to continuous interviewing, and our analyses rely on the 2006-10 release of these data. Marriage and fertility histories have long been a part of the NSFG; full cohabitation histories were collected starting in 1995. Men were added to the NSFG as of 2002. Unfortunately for our purposes, this round contained an error in skip patterns resulting in substantial missing data on dates of marital separation (Kennedy and Bumpass 2008), making it unsuitable for an analysis of union dissolution. We thus rely on data on women from the 1995 and 2006-10 interviews. The 1995 NSFG includes 10,847 women (79% response rate). Interviewing for the 2006-10 release was conducted from June 2006 through June 2010 and includes 12,279 women (78% response rate). The 1995 NSFG oversampled Hispanics and blacks, and in addition to these groups, the 2006-10 NSFG oversampled respondents ages 15–24. Sampling weights adjust for differences in sampling rates, response rates, and coverage rates and are applied in all analyses (using the *SVY* commands in STATA 12.0).

We generate a union-level file, including all marital and cohabiting unions bearing a child within 10 years of the 1995 and 2006-10 interviews. Although uncommon, women may contribute more than one union to the analysis file. Our union sample includes 2,656 unions from the 1995 survey (2,562 women) and 3,046 unions (2,907 women) from 2006-10. We are explicitly interested in union stability following the birth of a child, and thus we run our baseline duration "clock" from the time of the couple's first birth together. (This strategy excludes the many marriages and cohabitations without children, some of which are entered into as family states and some of which are meant to be short-term relationships of convenience.) To examine the monthly risk of separation in a multivariate discrete-time event history framework, we transform our union-level file into a union-month file, with one record for every month at risk of union dissolution from the time of birth. Our final union-month sample includes 136,955 months from the 1995 survey and 145,456 months from the 2006-10 survey. The two surveys cover unions bearing a first child in 1985-1995 and 1997-2010, respectively.

**Measures**

*Union-birth trajectories.* We focus on union status and transitions around the time of a couples' first birth together. We create 3 time-invariant indicators: married at union start, cohabiting at union start and married at birth, and cohabiting at birth. We also have a time-varying dummy indicating marriage at union duration *t* among the group cohabiting at birth. Together, these allow us to compare 4 distinct union-birth trajectories: 1) married at union start and birth (M→B); 2) cohabiting at union start and married at birth (C→M→B); 3) cohabiting at birth and married at some time *t* following the birth (C→B→M); and 4) cohabiting at birth without ever marrying (C→B). Trajectories 1 and 2 are estimated directly by the model; we combine model estimates for cohabiting at birth and marriage at duration *t* to construct trajectories 3 and 4 (for a similar approach, see Wu and Musick 2008). This can be seen in the chart below:

| Union-birth trajectory | Married at union start (1=yes) | Cohabiting at union start and married at birth (1=yes) | Cohabiting at birth (1=yes) | Married at duration *t* among those cohabiting at birth (time-varying) |
|---|---|---|---|---|
| M→B | 1 | 0 | 0 | 0 |
| C→M→B | 0 | 1 | 0 | 0 |
| C→B→M | 0 | 0 | 1 | 1 from time of marriage |
| C→B | 0 | 0 | 1 | 0 |

There may be quite a lot of fluidity between the time women become pregnant and birth (e.g., Lichter 2012; Manning 2001; Raley 2001). This adds a layer of complexity not factored into this analysis.

*Education.* The 1995 NSFG contains complete education histories, making it possible to map transitions into and out of schooling onto first birth and union transitions. Considerably less information on education is available in 2006-10 (only the date of high school graduation and, for the later years of interviewing, college graduation), precluding the possibility of precisely dating births relative to schooling transitions. We thus rely on education at interview (which potentially overstates to some degree education at birth by including educational upgrading between birth and interview).

*Family complexity.* Relying on the full cohabitation and marriage histories, we are able to construct indicators for whether the respondent was previously married and whether she ever lived with another partner outside of marriage. Comparing union and fertility histories, we generate an indicator for whether the respondent had any children prior to moving in with or marrying her partner. For women with children born prior to the current cohabitation or marriage, we indicate whether she has a child less than a year old at union start, 1-2 years old, or older than 2. Because the NSFG has no information on non-coresidential relationship histories for our sample, we cannot be sure that pre-union children are actually children from a prior relationship, i.e., they could be joint children born prior to coresidence. That scenario would be more likely among children born in the last year versus children born much earlier.

The NSFG includes limited information on partners' characteristics, especially in 1995. In 2006-10, for each prior coresidential partner (marital or cohabiting), women are asked about partners' previous marriages and children from prior relationships. No information is available on partners' past cohabitations. In 1995, partners' prior marriages are ascertained only for a subset of unions (current husbands and cohabiting partners and all past husbands, but not past cohabiting partners), and there is no information on partners' children from prior relationships. We are thus limited by data availability in the 1995 survey in the extent to which we can compare partners' union and birth histories and their association with union dissolution over time. We do, however, examine some supplementary data from the 2002 NSFG, as well as supplementary analyses including information on partners from 2006-10. (We also would have liked to have included other information on partners, such as race and education. But unfortunately these data are not available on all prior partners, nor are they available in the pregnancy histories on all birth fathers, as is father's age.)

*Other controls.* We control for several background characteristics of the respondent, including racial and ethnic background, mother's and father's educational attainment, whether she grew up with both biological parents, and whether she grew up attending church on a weekly basis. We include mother's and father's age at birth (for each pregnancy, women are asked the age of the birth father). We also control for whether the pregnancy leading to birth was intended, i.e., according to the respondent, whether she stopped contracepting around the time of the birth because she wanted to get pregnant, or reported that she wanted a baby at some time and her pregnancy came too late or at the right time. Finally, a time-varying indicator accounts for whether the couple had a second birth within their union.

### Method

After describing basic patterns of change, we examine them more closely in a multivariate framework. We run discrete-time event history models of the log-odds of separation, with the period of risk extending from the month of birth until separation or censoring at interview, for up to 120 months or 10 years. A union-month file allows for the precise timing of transitions into marriage and separation and is especially important in analyses of cohabitation, as transitions may occur at short durations. We cluster on women to account for the possibility of multiple events (unions) per woman. Models are run separately for 1995 and 2006-10 surveys, and differences in parameter estimates are tested across models.

Relying on model estimates, we next run a set of simulations to flesh out the implications of our findings in easy-to-interpret quantities. We transform our discrete-time logits into monthly predicted probabilities of separation, varying key characteristics and holding others at their weighted mean values. We multiply the monthly predicted probabilities (i.e., *conditional* predicted probabilities) to generate the probability of separation within 5 years of birth—a more intuitive measure of risk than either an estimated odds ratio or predicted monthly probability of risk. Model results are applied to various compositional assumptions to illustrate the substantive implications of our results.

## RESULTS

### Describing patterns of change

As a first step, we describe basic patterns of change: first in the distribution of union-birth trajectories over time; next, in the characteristics of these trajectories, focusing in particular on shifts in the distributions of education and family complexity. **Table 1** shows characteristics of our sample of unions with children for the 1995 and 2006-10 survey periods, for the full samples and separately by union-birth trajectory. Trajectories represented here include: married at union start and birth (M→B); 2) cohabiting at union start and married at birth (C→M→B); and 3) cohabiting at birth. When we move to an event history framework and incorporate time-varying covariates, we will distinguish further between the cohabitors at birth who do and do not marry subsequent to having a child (C→B→M vs. C→B). Important to note here is that many cohabitors in this group will go on to marry. According to competing risk life table estimates (not shown here but available upon request), 59% of those cohabiting at birth married prior to separating in the earlier period and 48% in the later period.

The shift from marriage to cohabitation from the 1995 to 2006-10 surveys is striking. Among unions bearing children in the 10 years prior to interview, the share married at union start dropped from half to 30% and the share cohabiting at birth more than doubled from 17% to 36% (there was a slight increase in the share cohabiting at union start and married at birth, from 33 to 35%). For those cohabiting at birth, there was an increase in the average duration to marriage, from 18.9 to 23.1.

How has the composition of these union-birth trajectories changed? There have been substantial shifts in the education distributions of women across union-birth trajectories. Births to married couples—whether those who married directly (M→B ) or cohabited premaritally (C→M→B)—are increasingly concentrated among the college educated. Half of these married mothers are now college graduates (compared to 31% of those who married directly and 23% who cohabited premaritally in the 1995 period). Cohabiting mothers have moved up the educational ranks as well, but the progression stops short of college in both time periods: Of those cohabiting at birth, there has been a shift from mothers with a high school degree to some college (with the some college group increasing from 17% to 29%). College graduates accounted for 5% or less of all cohabiting births in both periods.

Changes in prior union and childbearing experiences have been much less significant, to the extent that we are able to capture them. Consistent with prior literature, we see evidence of increasing serial cohabitation; less discussed, however, we also see evidence of declining serial marriage. Overall, the proportion with prior cohabitations has increased to about the same degree that the proportion with prior marriages has declined, such that in both periods about 20% of our sample had a prior relationship. The proportion with prior childbearing experience increased from 10 to 11%. Differences across union-birth trajectories are substantial (e.g., 11% of mothers who married directly vs. 26% of mothers who cohabited premaritally vs. 38% of those cohabiting at birth had either prior union or childbearing experience at the start of their current union). Looking within trajectories, however, the story remains one of no substantial change. The only group that experienced statistically significant increases in serial cohabitation were those cohabiting at birth, and these increases were (as reflected in the overall totals) completely offset by declines in prior marriage. Further, the small *increase* in the share of women bringing children with them from prior relationships masks *declines* within union-birth trajectories. Both those marrying directly and those cohabiting at birth experienced declines in the share with prior children (from 6% to 3% among the mothers marrying directly and from 27% to 21% among the cohabiting mothers). The overall increase stems from the substantial shift into cohabitation, in which prior childbearing is more common (despite declines over time).

As noted in our discussion of measures, the 1995 survey limits the extent to which we are able to account for partners' prior relationships and children. (And more generally, the NSFG does not allow us to assess increasing complexity arising from non-coresidential relationships.) We include here a supplementary table (Table S1) comparing the 2006-10 NSFG Women to the 2002 NSFG Women. This comparison allows us to incorporate information on partner's prior marriages and children (but not cohabitations). Here, if anything, we find evidence of *declining* family complexity (although, as noted, we are unable to compare prior cohabitations over time).

All other controls are also shown in Table 1. Those marrying directly were the only group to experience increases in weekly church attendance (from 48% to 56%, compared to about 20-25% among other mothers). Mothers were somewhat older at birth in the later versus earlier period, and here change appears to be due entirely to increases in age at birth among married mothers (who are significantly older than cohabiting mothers, e.g., in the later period: 27.6 among those married directly, 28.8 among the premarital cohabitors, and 23.1 among the cohabiting mothers). Birth intendedness also changed differentially over time by union-birth trajectory: the share of intended first births among those married at first birth remained nearly unchanged (at about 83% among those marrying directly and premaritally cohabiting); the share to cohabiting couples declined significantly, from 56% to 48%.

**Event history analysis**

Table 2 presents results from discrete-time event history models of union dissolution over 10 years as a function of months duration since the couple's first birth together. Model 1 includes only our union status indicators, and Model 2 includes all controls. We explore the substantive implications of these models in greater detail in Table 3 and thus provide only a cursory review here.

Coefficients comparing the odds of separation to those married at union start are all statistically significant and strong in magnitude, with the exception of the coefficient on cohabitation at union start followed by a marital birth (C→M→B) as estimated from the 2006-10 survey. In 1995, premarital cohabitation followed by a marital birth was associated with about a 50% increase in the odds of separation relative to marrying directly based on Model 1 (without controls) and a 33% increase based on Model 2 (with controls). In the later period, in Models 1 and 2 (with and without controls), the odds of dissolution among premarital cohabitors are statistically indistinguishable from those of couples marrying directly. This is consistent with recent work by Manning and Cohen (2012) finding a decline over time in the risk to marital disruption associated with living together prior to marriage. In Model 2, the difference in this coefficient is statistically significant across survey years.

Cohabiting at births is associated with greatly increased odds of separation: 5.48 and 6.79 times the odds of couples marrying directly in the 1995 and 2006-10 periods, respectively, based on Model 1. Coefficients drop substantially in magnitude when the full set of controls is added (in Model 1 vs. 2). But associations nonetheless remain strong, at 2.82 and 2.40 times the odds of couples marrying directly in the earlier and later periods, respectively. The time-varying indicator for marriage among cohabiting parents is also highly significant, suggesting a strong reduction in the odds of separation upon marriage. Below, we will combine these coefficients to examine union stability associated with two distinct trajectories: cohabiting at birth and then marrying versus cohabiting at birth and never marrying.

Education and family complexity appear to be associated with union disruption in much the same way in the 1995 versus 2006-10 periods; we find no statistically significant differences in Model 2 coefficients tested across models run separately by survey year. Unions involving women with college degrees are between 40-45% less likely to dissolve in any month, in both time periods. We find some evidence that unions involving a mother with prior relationships are less stable, although estimated associations are weaker than expected. Whether the respondent was previously married is not statistically significant in either time period; whether she had a previous cohabiting partner is associated with a 43% higher odds of monthly union dissolution, statistically significant in the earlier period only. Whether the respondent had a child older than 2 at the start of the current union is associated with an 84% higher odds of monthly union dissolution, statistically significant in the later period only. In supplementary models run on the 2006-10 data only (available upon request), we found that including additional measures of partners' prior marriages and children did not change our results.

Other controls are largely associated with union dissolution in expected ways. African Americans have a higher odds of union dissolution than Whites (although differentials are significantly smaller in the more recent period); Hispanics have a lower odds. An additional year of mother's age at birth is associated with a 7-9% reduction in the odds of dissolution; father's age at birth appears to provide no additional protection above and beyond mother's age. An intended birth is associated with greater subsequent union stability (36% lower odds of monthly disruption for intended vs. unintended births; significant in the later period only and significantly different from the 1995 estimate). Also associated with reductions in the odds of disruption: growing up with both parents (statistically significant in the later period only), church attendance, and having another child.

**Simple simulations**

Table 3 facilitates the interpretation of union status indicators, showing predicted probabilities of separation within 5 years estimated from the discrete-time event history models just described (Model 1-2, Table 2). These incorporate information on time-varying marriage among cohabiting parents, and thus we are able to make comparisons across four union-birth trajectories: 1) married at union start and birth (M→B); 2) cohabiting at union start and married at birth (C→M→B); 3) cohabiting at birth and married at some time *t* following the birth (C→B→M); and 4) cohabiting at birth without ever marrying (C→B). Supplementary Table S2 produces odds ratios comparisons across our four union-birth trajectories. Recall trajectories 3 and 4 are constructed by combining and exponentiating coefficients from Models 1 and 2 on cohabiting at birth and marriage following the birth (as described in the methods section); comparisons across these trajectories are tested using the standard error of sums of estimated coefficients.

The predicted probabilities of separation within 5 years derived from models with no controls are shown in Table 3, which uses information from estimates in Table 2. In both periods, by far the highest levels of instability are among cohabiting parents who never marry. In the 2006-10 period, the estimated proportion separating within 5 years is 11% among those marrying directly (M→B), 14% among those premaritally cohabiting and then having a birth in marriage (C→M→B), 28% among those having a birth in cohabitation and then marrying (C→B→M), and 56% among those having a birth in cohabitation and never marrying (C→B; fully 4.89 times the proportion separating in the M→B group). Overall patterns are similar over time although the unions involving marriage are somewhat more stable in the later period. In 1995, all union-birth trajectories are significantly different from each other. In the later period, as apparent in Table 2, the separation risks for those who marry directly are not significantly different from those who cohabit premaritally.

Accounting for the observed ways in which cohabitors differ from married couples substantially reduces gaps in separation risks across union-birth trajectories (Table 3, results derived from Model 2 in Table 2 and descriptives in Table 1). In the 2006-10 period, including all controls, the estimated proportion separating within 5 years is 15% among those marrying directly (M→B), 14% among those premaritally cohabiting and then having a birth in marriage (C→M→B), 20% among those having a birth in cohabitation and then marrying (C→B→M), and 32% among those having a birth in cohabitation and never marrying (C→B; 2.16 times the proportion separating in the M→B group). In 1995, all union-birth contrasts are statistically significant *but* the difference between cohabitors who marry before versus after their first birth (with estimated proportions separating of 18% vs. 25%, respectively). In the later period, none of the trajectories involving marriage (direct marriage, premarital cohabitation, or marriage following a cohabiting birth) differ significantly from each other; the *only* union-birth trajectory that is distinct is cohabiting without marriage.

Finally, Figure 1 shows results of a set of simulations examining separation risks derived from the full model (Model 2, Table 2), varying the assignment of sample characteristics from the 1995 and 2006-10 survey periods. This exercise addresses the following hypothetical: Had the characteristics of unions not changed over time, how would union stability in the most recent period compare to what was actually observed? As above, predicted probabilities are estimated for each month under various conditions and then multiplied to generate the estimated probability of separation within 5 years. The first column, using model results and weighted mean characteristics for the 1995 survey, shows a predicted probability of separation within 5 years of 16.7%. The second column, using model results and weighted mean characteristics for the 2006-10 survey, shows a similar (statistically indistinguishable) level of separation risk at 18.1%. Keeping the 2006-10 model parameters and assigning union status means from 1995 but all other covariates their means from 2006-10, the predicted probability drops to 16.1%. This suggests that had nothing changed but union status across these periods, the probability of separating would have been 11% lower than what was actually observed in the later period (accounting for more than the total difference in estimates of instability across periods). Keeping the 2006-10 model parameters and assigning all covariates the means from 1995, the predicted probability of separation within 5 years drops to 15.8% or 13% of what was actually observed in the later period.

## (PRELIMINARY AND INCOMPLETE) DISCUSSION

The composition of couples having children has changed in striking ways over the past decade and a half. Most notably, births within cohabitation more than doubled from 17% to 36%. All else equal, this would tend to increase the overall instability of unions over time, whereas our evidence suggests little change overall. This appears due in part to offsetting compositional changes and in part to greater stability among some union-birth trajectories, net of changing characteristics. The education of respondents improved, and evidence is somewhat mixed with respect to the growing complexity of families with children: We reported increases in serial cohabitation (although only among the cohabiting parents in our sample), but also declines serial marriage. Women in unions with children were overall modestly more likely to have children from a prior relationship, stemming entirely from an increase in the *share* of cohabitors (who are *more likely* to have children from prior relationships than married parents, even though *less likely* to do so over time). Associations were reasonably modest between our indicators of family complexity and union instability.

Both unions marked by a first birth in marriage and those marked by a first birth in cohabitation became more stable over time—assuming that marriage followed birth among the cohabiting parents. For never-married cohabitors, estimates of proportions separating over 5 years were remarkably similar over time. A part of this story is a weakening association between cohabitation and instability, so long as marriage occurs at some point (before *or* after the birth). A few take-home points to develop further, together suggesting a continued evolution not in the meaning but *meanings* of cohabitation (which may also vary across groups, e.g., race/ethnicity and education): 1) cohabiting as precursor to marriage and childbearing involves little selection on socioeconomic status and no discernible risk to stability; 2) accounting for selection, the timing of cohabitation and marriage relative to birth has little implication for stability; 3) cohabitation is less stable when de-coupled from marriage, although selection factors are most important here.

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Table 1. Characteristics of unions with children, 1995 and 2006-10 NSFG Women

| | All | | M→B | | C→M→B | | Cohabiting at birth | |
|---|---|---|---|---|---|---|---|---|
| | 1995 | 2006-10 | 1995 | 2006-10 | 1995 | 2006-10 | 1995 | 2006-10 |
| Union duration (months from first birth in union) | 51.21 | 51.88 | 56.72 | <u>62.30</u> | 50.23 | <u>55.70</u> | 36.40 | <u>39.48</u> |
| *Union-birth trajectory* | | | | | | | | |
| M→B (married at union start) | 0.50 | <u>0.30</u> | 1.00 | 1.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| C→M→B (cohabiting at union start, married at birth) | 0.33 | 0.35 | 0.00 | 0.00 | 1.00 | 1.00 | 0.00 | 0.00 |
| Cohabiting at birth | 0.17 | <u>0.36</u> | 0.00 | 0.00 | 0.00 | 0.00 | 1.00 | 1.00 |
| Time to marriage among those cohabiting at birth | 18.86 | <u>23.07</u> | 0.00 | 0.00 | 0.00 | 0.00 | 18.86 | <u>23.07</u> |
| *R's education (highest grade at interview)* | | | | | | | | |
| Less than HS | 0.13 | <u>0.16</u> | 0.08 | 0.10 | 0.11 | 0.08 | 0.30 | 0.30 |
| HS degree | 0.39 | <u>0.24</u> | 0.35 | <u>0.16</u> | 0.39 | <u>0.19</u> | 0.49 | <u>0.36</u> |
| Some college | 0.25 | 0.26 | 0.26 | 0.23 | 0.27 | 0.26 | 0.17 | <u>0.29</u> |
| College + | 0.24 | <u>0.33</u> | 0.31 | <u>0.51</u> | 0.23 | <u>0.47</u> | 0.04 | 0.05 |
| *Family complexity (union and birth histories)* | | | | | | | | |
| R married previously | 0.10 | <u>0.06</u> | 0.05 | 0.04 | 0.16 | <u>0.08</u> | 0.13 | <u>0.07</u> |
| R cohabited previously | 0.10 | <u>0.15</u> | 0.03 | 0.05 | 0.17 | 0.17 | 0.16 | <u>0.23</u> |
| R had other child(ren) at start of this union | 0.10 | <u>0.11</u> | 0.06 | <u>0.03</u> | 0.08 | 0.07 | 0.27 | <u>0.21</u> |
| Age of oldest child (R's with other children) | | | | | | | | |
| < 1 | 0.19 | 0.24 | 0.27 | 0.34 | 0.16 | 0.17 | 0.22 | 0.24 |
| 1-2 years | 0.20 | 0.20 | 0.23 | 0.19 | 0.20 | 0.21 | 0.21 | 0.20 |
| 2+ years | 0.61 | 0.56 | 0.50 | 0.46 | 0.64 | 0.62 | 0.57 | 0.56 |
| *Summary measures of family complexity* | | | | | | | | |
| R had any previous union | 0.19 | 0.22 | 0.08 | 0.08 | 0.31 | <u>0.23</u> | 0.28 | 0.28 |
| R had a prior union or children | 0.23 | <u>0.28</u> | 0.11 | 0.11 | 0.34 | <u>0.26</u> | 0.41 | 0.38 |
| *Sociodemographic background* | | | | | | | | |
| Racial-ethnic background | | | | | | | | |
| Non-Hispanic White | 0.74 | 0.65 | 0.72 | <u>0.64</u> | 0.84 | <u>0.78</u> | 0.58 | 0.54 |
| Non-Hispanic Black | 0.07 | <u>0.10</u> | 0.06 | 0.07 | 0.04 | <u>0.07</u> | 0.17 | 0.16 |
| Hispanic | 0.14 | <u>0.18</u> | 0.15 | 0.18 | 0.08 | 0.11 | 0.22 | 0.25 |
| Other | 0.05 | 0.07 | 0.08 | <u>0.11</u> | 0.03 | 0.04 | 0.03 | 0.05 |
| Father's education (highest grade) | | | | | | | | |
| Less than HS | 0.18 | <u>0.22</u> | 0.19 | 0.20 | 0.14 | 0.17 | 0.22 | <u>0.27</u> |
| HS degree | 0.43 | <u>0.31</u> | 0.41 | <u>0.26</u> | 0.44 | <u>0.30</u> | 0.45 | <u>0.35</u> |
| Some college + | 0.30 | <u>0.38</u> | 0.32 | <u>0.48</u> | 0.32 | <u>0.44</u> | 0.16 | <u>0.25</u> |
| Missing | 0.10 | 0.09 | 0.07 | 0.06 | 0.10 | 0.08 | 0.17 | <u>0.13</u> |
| Mother's education (highest grade) | | | | | | | | |
| Less than HS | 0.17 | <u>0.22</u> | 0.17 | <u>0.25</u> | 0.12 | 0.15 | 0.24 | 0.27 |
| HS degree | 0.54 | <u>0.35</u> | 0.53 | <u>0.29</u> | 0.58 | <u>0.39</u> | 0.52 | <u>0.37</u> |
| Some college + | 0.25 | <u>0.41</u> | 0.27 | <u>0.46</u> | 0.27 | <u>0.46</u> | 0.16 | <u>0.34</u> |
| Missing | 0.04 | 0.01 | 0.03 | <u>0.01</u> | 0.03 | <u>0.01</u> | 0.08 | <u>0.01</u> |
| Grew up with both parents | 0.63 | 0.62 | 0.73 | <u>0.78</u> | 0.60 | 0.63 | 0.39 | <u>0.47</u> |
| Attended church weekly | 0.36 | <u>0.32</u> | 0.48 | <u>0.56</u> | 0.25 | 0.24 | 0.23 | 0.20 |
| *Characteristics of first birth in this union* | | | | | | | | |
| R's age | 26.04 | <u>26.38</u> | 26.30 | <u>27.56</u> | 27.19 | <u>28.75</u> | 22.95 | 23.11 |
| Partner's age | 27.26 | <u>28.34</u> | 27.26 | <u>29.58</u> | 28.50 | <u>30.34</u> | 24.74 | <u>25.38</u> |
| Pregnancy was intended | 0.78 | <u>0.71</u> | 0.83 | 0.84 | 0.82 | 0.83 | 0.56 | <u>0.48</u> |
| Couple had another child in this union (time-varying) | 0.38 | 0.38 | 0.40 | <u>0.43</u> | 0.36 | <u>0.39</u> | 0.32 | 0.32 |
| Number of unions | 2656 | 3046 | 1312 | 812 | 832 | 886 | 512 | 1348 |
| Number of women | 2562 | 2907 | 1288 | 804 | 812 | 874 | 464 | 1247 |
| Number of union-months | 136955 | 145456 | 75728 | 48326 | 42041 | 45735 | 19186 | 51395 |

Source: 1995 and 2006-10 NSFG. Sample limited to couples having a first child together within 10 years of interview.

Notes: *N*'s unweighted. All means weighted using *SVY* procedures in STATA 12.0. Underlined terms significantly different from 1995 at p<.05. R=respondent.

Table 2. Odds ratios from discrete-time event history models of union dissolution within 10 years of birth, 1995 & 2006-10 NSFG

| | Model 1 | | Model 2 | |
|---|---|---|---|---|
| | 1995 | 2006-10 | 1995 | 2006-10 |
| Union duration (months from first birth in union) | 1.00 | 1.01 † | 1.01 | 1.02 *** |
| Union duration squared | 1.00 | 1.00 ** | 1.00 | 1.00 ** |
| *Union-birth trajectory* | | | | |
| M→B (married at union start) | 1.00 | 1.00 | 1.00 | 1.00 |
| C→M→B (cohabiting at union start, married at birth) | 1.47 ** | 1.21 | 1.33 * | 0.95 |
| Cohabiting at birth | 5.48 *** | 6.79 *** | 2.82 *** | 2.40 *** |
| Marriage among those cohabiting at birth (time-varying) | 0.54 ** | 0.41 *** | 0.59 ** | 0.43 *** |
| *R's education (highest grade at interview)* | | | | |
| Less than HS | | | 0.94 | 1.18 |
| HS degree (reference) | | | 1.00 | 1.00 |
| Some college | | | 0.87 | 0.93 |
| College + | | | 0.60 ** | 0.55 *** |
| *Family complexity(union and birth histories)* | | | | |
| R married previously | | | 1.21 | 1.11 |
| R cohabited previously | | | 1.43 * | 1.14 |
| R had no child(ren) at start of this union (reference) | | | 1.00 | 1.00 |
| R had child age <1 year | | | 0.81 | 1.03 |
| R had child age 1-2 years | | | 1.40 | 1.46 |
| R had child age >2 years | | | 1.37 | 1.84 ** |
| *Sociodemographic background* | | | | |
| Racial-ethnic background | | | | |
| Non-Hispanic White (reference) | | | | |
| Non-Hispanic Black | | | 1.77 *** | 1.28 † |
| Hispanic | | | 0.79 | 0.74 * |
| Other | | | 0.95 | 0.78 |
| Father's education (highest grade) | | | | |
| Less than HS | | | 0.84 | 0.77 |
| HS degree (reference) | | | 1.00 | 1.00 |
| Some college + | | | 0.94 | 1.27 † |
| Missing | | | 0.77 † | 1.03 |
| Mother's education (highest grade) | | | | |
| Less than HS | | | 0.89 | 1.05 |
| HS degree (reference) | | | 1.00 | 1.00 |
| Some college + | | | 1.07 | 1.10 |
| Missing | | | 0.71 | 0.96 |
| Grew up with both parents | | | 0.87 | 0.65 ** |
| Attended church weekly | | | 0.65 *** | 0.65 ** |
| *Characteristics of first births in this union* | | | | |
| R's age | | | 0.91 *** | 0.93 ** |
| Partner's age | | | 0.99 | 1.01 |
| Pregnancy was intended | | | 0.89 | 0.64 *** |
| Couple had another child in this union (time-varying) | | | 0.78 † | 0.57 *** |
| | | | | |
| Constant | 0.00 | 0.00 | 0.05 | 0.02 |
| Observations | 136955 | 145456 | 136241 | 145436 |

Source: 1995 and 2006-10 NSFG. Sample limited to couples having a first child together within 10 years of interview.
Notes: *N*'s unweighted. All models weighted using *SVY* procedures in STATA 12.0. Underlined terms significantly different from 1995 at p<.05. Asterisks indicate differences from 1.00 at †p<.10, *p<.05, **p<.01, ***p<.001. R=respondent.

Table 3. Predicted probabilities of separation within 5 years derived from discrete-time event history models, varying union status

| | Model 1 | | Model 2 | |
|---|---|---|---|---|
| | 1995 | 2006-10 | 1995 | 2006-10 |
| Overall | 0.19 | 0.21 | 0.17 | 0.18 |
| | | | | |
| M→B (married at union start) | 0.14 | 0.11 | 0.13 | 0.15 |
| C→M→B (cohabiting at union start, married at birth) | 0.19 | 0.14 | 0.18 | 0.14 |
| C→B→M (cohabiting at birth then married) | 0.35 | 0.28 | 0.25 | 0.20 |
| C→B (cohabiting at birth and never married) | 0.55 | 0.56 | 0.32 | 0.32 |
| | | | | |
| Ratio C→B / M→B | 4.03 | 4.89 | 2.38 | 2.16 |

Source: 1995 and 2006-10 NSFG. Sample limited to couples having a first child together within 10 years of interview.

Notes: Predicted probabilities of monthly separation risk derived from event history models shown in Table 2, varying union status and holding all other covariates at weighted mean values shown in Table 1. Monthly conditional probabilities of separation risk multiplied to generate estimated proportions separating over 5 years. *N*'s unweighted. All descriptives and models weighted using *SVY* procedures in STATA 12.0.

Supplementary Table S1. Additional measures of family complexity, 2002 and 2006-10 NSFG Women

| | All | | M→B | | C→M→B | | Cohabiting at birth | |
|---|---|---|---|---|---|---|---|---|
| | 2002 | 2006-10 | 2002 | 2006-10 | 2002 | 2006-10 | 2002 | 2006-10 |
| Partner previously married | 0.16 | 0.14 | 0.10 | 0.10 | 0.21 | 0.14 | 0.18 | 0.17 |
| Partner had other child(ren) at start of this union | 0.17 | 0.17 | 0.09 | 0.07 | 0.17 | 0.15 | 0.28 | 0.29 |
| | | | | | | | | |
| Both partners married before | 0.04 | 0.03 | 0.02 | 0.01 | 0.06 | 0.03 | 0.03 | 0.03 |
| Either partner married before | 0.22 | 0.18 | 0.14 | 0.12 | 0.28 | 0.19 | 0.24 | 0.21 |
| Neither partner married before | 0.78 | 0.82 | 0.86 | 0.88 | 0.72 | 0.81 | 0.76 | 0.79 |
| | | | | | | | | |
| Both partners had kids before | 0.03 | 0.04 | 0.01 | 0.01 | 0.03 | 0.02 | 0.08 | 0.07 |
| Either partner had kids before | 0.26 | 0.25 | 0.13 | 0.09 | 0.25 | 0.19 | 0.46 | 0.43 |
| Neither partner had kids before | 0.74 | 0.75 | 0.87 | 0.91 | 0.75 | 0.81 | 0.54 | 0.57 |
| | | | | | | | | |
| Number of unions | 2052 | 3046 | 739 | 812 | 666 | 886 | 648 | 1348 |
| Number of women | 1866 | 2907 | 729 | 804 | 648 | 874 | 560 | 1247 |
| Number of union-months | 101709 | 145456 | 75728 | 48326 | 42041 | 45735 | 25570 | 51395 |

Source: 2002 and 2006-10 NSFG. Samples limited to couples having a first child together within 10 years of interview.
Notes: *N*'s unweighted. All means weighted using *SVY* procedures in STATA 12.0. Underlined terms significantly different from 2002 at p<.05.

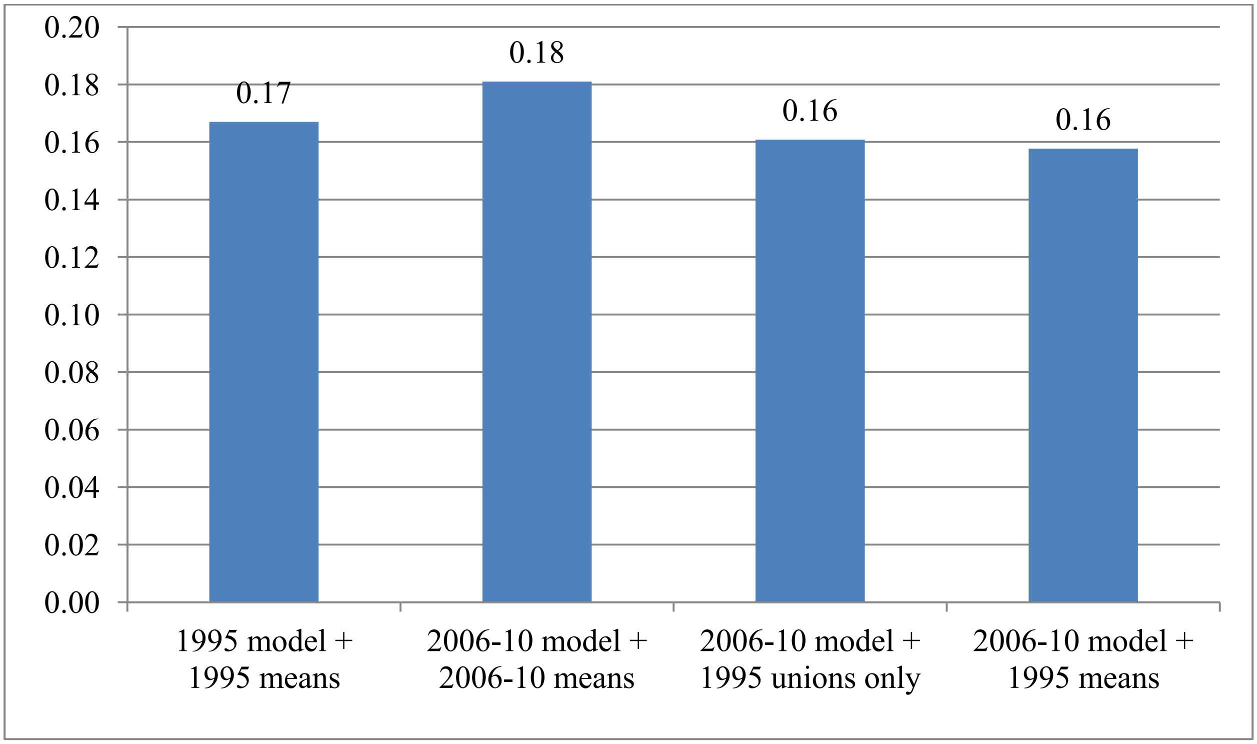
Supplementary Table S2. Estimated odds ratios of monthly separation derived from Models 1-2, Table 2

| 1995 | M→B | C→M→B | C→B→M | C→B |
|---|---|---|---|---|
| *Model 1* | | | | |
| M→B (married at union start) | 1.00 | 1.47 *** | 2.97 *** | 5.48 *** |
| C→M→B (cohabiting at union start, married at birth) | | 1.00 | 2.02 *** | 3.74 *** |
| C→B→M (cohabiting at birth then married) | | | 1.00 | 1.85 *** |
| C→B (cohabiting at birth and never married) | | | | 1.00 |
| *Model 2* | | | | |
| M→B (married at union start) | 1.00 | 1.33 * | 1.66 *** | 2.82 *** |
| C→M→B (cohabiting at union start, married at birth) | | 1.00 | 1.25 | 2.12 *** |
| C→B→M (cohabiting at birth then married) | | | 1.00 | 1.70 ** |
| C→B (cohabiting at birth and never married) | | | | 1.00 |
| | | | | |
| 2006-10 | | | | |
| *Model 1* | | | | |
| M→B (married at union start) | 1.00 | 1.21 | 2.75 *** | 6.79 *** |
| C→M→B (cohabiting at union start, married at birth) | | 1.00 | 2.27 *** | 5.60 *** |
| C→B→M (cohabiting at birth then married) | | | 1.00 | 2.47 *** |
| C→B (cohabiting at birth and never married) | | | | 1.00 |
| *Model 2* | | | | |
| M→B (married at union start) | 1.00 | <u>0.95</u> | <u>1.03</u> | 2.40 *** |
| C→M→B (cohabiting at union start, married at birth) | | 1.00 | <u>1.08</u> | 2.52 *** |
| C→B→M (cohabiting at birth then married) | | | 1.00 | 2.34 *** |
| C→B (cohabiting at birth and never married) | | | | 1.00 |

Source: 1995 and 2006-10 NSFG. Sample limited to couples having a first child together within 10 years of interview.

Notes: Coefficients on cohabiting at birth and time-varying marriage combined to generate distinct odds ratios for the trajectories: C→B→M and C→B. Statistical significance tested across 4 union-birth trajectories. All models weighted using *SVY* procedures in STATA 12.0. Underlined terms significantly different from 1995 at p<.05. Asterisks indicate differences from 1.00 at *p<.05, **p<.01, ***p<.001.

Figure 1. Predicted probability of separation within 5 years derived from discrete-time event history models **(M2)**, altering model parameters and covariate means




*Source:* 1995 and 2006-10 NSFG. Sample limited to couples having a first child together within 10 years of interview.

*Notes:* Predicted probabilities of monthly separation risk derived from event history models shown in Table 2, varying union status and holding all other covariates at weighted mean values shown in Table 1. Monthly conditional probabilities of separation risk multiplied to generate estimated proportions separating over 5 years. N's unweighted. All descriptives and models weighted using SVY procedures in STATA 12.0.