

Cohabitation Experience and Cohabitation's Association with Marital Dissolution

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Draft paper

7/5/2018

Accepted Contingent in Journal of Marriage and Family

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ABSTRACT:

The literature on the association between premarital cohabitation and marital dissolution presents several puzzles. Some scholars find that premarital cohabitation has persistently been associated with higher rates of divorce, while other scholars find that in the most recent marriage cohorts, the association between premarital cohabitation and marital dissolution has disappeared. Data from a harmonized dataset of National Surveys of Family Growth (with 216,455 couple-years) show that in the first year of marriages, couples who have cohabited before have a lower breakup rate than couples who have never cohabited, which may be due to the initial experiential advantage that couples who have already lived together enter into marriage with. The experiential advantage of cohabitation lasts only for the first year marriage, which partly explains why some authors have found that premarital cohabitation's impact on divorce appears to dissipate in the most recent marriage cohorts. Different metrics for measuring the significance of changes over time are discussed.

Cohabitation Experience and Cohabitation's Association with Marital Dissolution

INTRODUCTION:

The association between premarital cohabitation and divorce in the U.S. is academically contested terrain. Scholars disagree about *why* couples who cohabit before marriage have had higher divorce rates. Scholars also disagree about *whether* the association between premarital cohabitation and divorce has diminished over time. Premarital cohabitation has been associated with higher divorce rates in the past (Bramlett and Mosher 2002, Smock 2000, Cherlin 1992). Premarital cohabitation was rare and stigmatized before 1970 in the US, but by the 2000s the novelty of and stigma against premarital cohabitation had both worn off (Smock 2000). As the stigma against premarital cohabitation has worn off, and as those who cohabit premaritally have become a broader and less selective subset of all marriages, one might expect the divorce rate for couples who cohabited before marriage to converge with divorce rate for married couples who did not cohabit before marriage.

We offer a new explanation for why premarital cohabitation has appeared in some analyses to be less predictive of divorce in the most recent marriage cohorts: in the first year of marriage, couples who cohabited before marriage have a lower marital dissolution rate than couples who did not cohabit before marriage. We hypothesize that premarital cohabitation confers a practical advantage in the experience of how to live together with the partner. After the

first year of marriage, the couples who had not cohabited before marriage have caught up in the practical experience of living with their partner, and after that point, the hazard of marital dissolution is substantially higher for couples who cohabited before marriage. Our analysis sheds new light on the short term and long term ways in which premarital cohabitation appears to effect marital stability.

THEORIES OF THE ASSOCIATION BETWEEN PREMARITAL COHABITATION AND DIVORCE:
SELECTION AND THE COHABITATION EXPERIENCE OF IMPERMANENCE

In the 1970s, when premarital cohabitation was new and unusual and less well accepted in the U.S. than it is now, scholars tended to assume that couples who had cohabited before marriage would have more stable marriages (Macklin 1978, Smock 2000), but there was a lack of available data in the 1970s to test the association between premarital cohabitation and later marital stability. Cohabitation was seen as a trial period before marriage, so scholars assumed that only the most compatible couples would transition from cohabitation to marriage, which, if true, would have meant that premarital cohabitation would have been associated with greater marital stability and *lower* hazard of divorce. Cohabitors themselves saw being sure they were compatible before marriage as the primary benefit of cohabitation (Bumpass et al. 1991). The potential of cohabitation to filter out incompatible relationships before marriage is what Smock (2000) referred to as the “common sense” understanding of how premarital cohabitation would affect later marital stability.

The “common sense” understanding that premarital cohabitation would lead to more stable marital unions was soon upended by research. When data first became available in the

1980s to study the association between premarital cohabitation and divorce, scholars were surprised to find that married couples who had cohabited before marriage had *higher* rates of divorce (Booth and Edwards 1992, DeMaris and Rao 1992, Bumpass and Sweet 1989) and worse marital adjustment (DeMaris and Leslie 1984) compared to married couples who had not cohabited before marriage. Studies in other countries also found that couples who cohabited before marriage had higher rates of marital dissolution (Bennett et al. 1988, Hall and Zhao 1995).

The finding that married couples who had cohabited before marriage had a higher divorce rate than married couples who had not cohabited with each other before marriage yielded two classes of explanations. The first kind of explanation was *selection*: couples who cohabited before marriage were, even before the cohabitation experience, more liberal, less religious, and more prone to divorce if the relationship turned sour (Cherlin 1992, Dush et al. 2003, Smock 2000). The selection explanation also implied that the kind of person who would never consider premarital cohabitation was perhaps also the kind of person who would not later consider divorce, even if the marriage was less than satisfactory. Lillard et al (1995) offered empirical support for the selection explanation for the higher rate of divorce among couples who cohabited before marriage. Booth and Johnson (1988), in contrast, found that the association between premarital cohabitation and later divorce remained intact even after measures of liberalism and religiosity were controlled for.

The second potential explanation for the association between premarital cohabitation and higher risk of divorce is *experience*, specifically the way that the experience of cohabitation teaches people that romantic relationships are impermanent and disposable, which we refer to as the Cohabitation Experience of Impermanence. If the Cohabitation Experience of Impermanence theory is correct, then as couples advance from cohabitation to marriage, they advance with their

relationship commitment already eroded by the casual and informal nature of cohabitation. Consistent with the Cohabitation Experience of Impermanence, Axinn and Thornton (1992) found that individuals were more accepting of the idea of divorce after they had cohabited, compared to before they had cohabited.

Cohabitation is a less institutionalized, less formal relationship than marriage. Marriage involves more gendered expectations and traditions that are different from the experience of cohabitation. As a result of the differences between cohabitation and marriage, the transition from cohabitation to marriage can introduce unanticipated challenges into relationships (Bass 2015, Sassler and Miller 2011) that might partly explain why cohabiters have higher divorce rates than couples who did not cohabit before marriage.

Teachman (2003) found that it was not the experience of premarital cohabitation with the marriage partner, but rather cohabitation and nonmarital sex (before marriage) with other men that was associated with a woman's higher risk of divorce. Throughout this paper, when we refer to nonmarital cohabitation, we mean prior cohabitation with men who the woman did not marry. Premarital cohabitation refers to cohabitation with the man who went on to become the woman's first husband. The implication of Teachman's (2003) finding is that it may be the breakups of prior nonmarital cohabiting relationships, rather than the experience of premarital cohabitation with the marriage partner, that imparts expectations about the impermanence of relationships to cohabiters. Women who cohabit with the future marriage partner are more likely to have also previously cohabited with other partners. Following Teachman (2003), if we fail to control for nonmarital cohabitation, the association between premarital cohabitation (with the future marital partner) and marital breakup could be at least partly spurious.

SELECTION INTO COHABITATION AND ITS CHANGE OVER TIME

As premarital cohabitation has gone from about 10% of first marriages in 1970 to more than 60% of first marriages after 2000 (see Figure 1 below), the selectiveness of cohabitation has necessarily diminished over time. Dush et al (2003) argued that if selection (of liberals and less religious people, for example) into cohabitation was the reason that premarital cohabitation was observed to be associated with higher rates of divorce, then we would expect the divorce rate of couples who cohabited before marriage to converge (over time) with the divorce rate of couples who did not cohabit before marriage. The selection explanation (for premarital cohabitation's effect on divorce risk), should have become less salient as cohabitation has become less selective. In other words, it was possible that in the past only people with certain attributes (attributes that might have been correlated with divorce) would have considered premarital cohabitation. Now that cohabitation is common and normative, it is more difficult to imagine that premarital cohabitation would select for individual traits that would be associated with higher risk of divorce. We use the National Surveys of Family Growth which are retrospective surveys, and therefore do not lend themselves to a direct analysis of the selectivity of cohabitation decisions at the time the cohabitation decisions were made.

A corollary to the decline of the selectiveness of cohabitation is the decline of stigma against premarital sex, premarital sex being a key component of cohabitation. Since 1972, the percentage of Americans who say that premarital sex is "always wrong" has declined sharply (Treas 2002). Declining stigma against cohabitation should be associated with increasing support of cohabiting couples (or less opposition and hostility towards cohabiting couples) from friends, family, and strangers (Rosenfeld 2007). Stigma has been shown to affect individuals' physical

and mental health (Riggle and Rostosky 2007, Hatzenbuehler et al. 2013) and might reasonably affect their marital satisfaction as well. If the stigma against premarital cohabitation (and its presumed corollary, premarital sex) was the reason that premarital cohabitation was observed to be associated with higher rates of divorce, then we would expect the divorce rate of couples who cohabited before marriage to converge (over time) with the divorce rate of couples who did not cohabit before marriage.

We refer to the hypothesized convergence of divorce rates between premarital cohabiters and other married couples as the Normalization Hypothesis. According to the Normalization Hypothesis, as cohabitation and premarital sex have become more common and normalized and less stigmatized over time, the cost (in terms of higher divorce rates) for carrying the formerly stigmatized characteristic should decline as well. The Normalization Hypothesis predicts a convergence in marital dissolution rates over time between couples who cohabited before marriage and couples who did not cohabit before marriage.

THE PRACTICAL EXPERIENCE OF COHABITATION:

Before it became clear in the 1980s that premarital cohabitation was associated with higher rates of divorce in the U.S., scholars assumed that the experience of cohabitation would teach couples important practical lessons about how to live together. We refer to this kind of experience as the Practical Experience of Cohabitation. DeMaris and Leslie (1984), for instance, expected that cohabiting couples would learn how to manage joint chores, how to accommodate themselves to each other's housekeeping habits, how to share time, how much sex to expect, and so on. In the initial phase of a marriage, cohabiting couples would in theory have the practical advantage of the prior experience of living together. Newlyweds have many practical household

decisions to make and issues to decide on, and couples who have the practical experience of cohabitation with each other have the advantage of already having worked out some or all of the initial issues that comprise living under one roof in a romantic union.

The common sense (in Smock's 2000 terms) possibility that the experience of premarital cohabitation might be useful experience for married couples has been mostly overlooked in the literature because the association between premarital cohabitation and higher divorce rates has led scholars to look for negative impacts rather than positive impacts of premarital cohabitation on later marital stability. We revisit the venerable (and more recently overlooked) common sense idea that the Practical Experience of Cohabitation may be, in some circumstances, beneficial for marital stability.

Cohabitation could confer both the Practical Experience of Cohabitation (positive for marital stability, and short acting) as well as the Cohabitation Experience of Impermanence (negative for marital stability, and longer acting). We find that the experience of cohabitation does appear to confer a marital stability benefit, but the experiential benefit of cohabitation is short acting. We find that married couples who have cohabited have a lower rate of marital dissolution in the calendar year of marriage. After a year of marriage, the experiential benefit of cohabitation dissipates. We make no attempt in this paper to isolate the causal mechanisms at the root of the association between premarital cohabitation and divorce, we merely identify marital dissolution trends that are consistent with different kinds of experience and selection.

Whereas the selection explanations for cohabitation's effect on marital dissolution imply a convergence of divorce rates between the former cohabiters and the non-cohabiters as premarital cohabitation became more common and less selective (the Normalization Hypothesis), the experiential explanations imply no such convergence over time. The experience

of cohabitation (and therefore the association between the cohabitation experience and marital dissolution) need not have changed over time as cohabitation became more popular. Therefore, stability over time in the odds ratio association between premarital cohabitation and marital dissolution would be more consistent with experiential explanations, while convergence in the association between premarital cohabitation and marital dissolution would be more consistent with selection explanations (Dush et al. 2003).

PREMARITAL COHABITATION AND DIVORCE: CHANGE VERSUS STABILITY OVER TIME:

Prior to the most recent waves of the National Surveys of Family Growth (NSFG), the NSFG single wave reports had consistently found that premarital cohabitation was associated with a greater hazard of marital dissolution (Bramlett and Mosher 2002, Cherlin 1992, Goodwin et al. 2010). Copen et al. (2012, table 7), using the 2006-10 wave of NSFG, found that premarital cohabitation was a significant predictor of marital dissolution for couples who had been married for 10, 15, and 20 years (consistent with prior literature using NSFG and other datasets), but that for couples who had been married only 5 years or less, there was no apparent relationship between premarital cohabitation and marital dissolution. One interpretation of Copen et al.'s results is that the most recent marriage cohorts no longer experience higher risk of marital dissolution after premarital cohabitation, a conclusion consistent with Reinhold (2012). An alternate interpretation of Copen et al.'s results is that premarital cohabitation is less of a risk for marital dissolution in the first few years of marriage, since the experience of marriage is less of a transition for couples who were living together already. Couples married in the most recent years before a survey have the shortest marital duration and are from the most recent marital cohort. If

there is an association between marital duration and the way in which premarital cohabitation affects marital dissolution risk, short marital duration could easily be mistaken for a marital cohort effect.

Reinhold (2012), using the 1988, 1995, and 2002 waves of NSFG, interacted premarital cohabitation with marital cohort and found a significant decline in the power of premarital cohabitation to predict the risk of marital dissolution, based on the significance of interactions at the coefficient level. Other scholars found no historical change in the strength of the association between premarital cohabitation and the later hazard of divorce in the US. Dush, Cohan, and Amato (2003) used a multistage US phone survey to determine the predictors of marital dissolution rates of couples who were married in 1980, compared to couples who were married in the 1990s. They found that the association between premarital cohabitation and divorce was the same across the marriage cohorts, despite the sharp rise in premarital cohabitation across marriage cohorts. Teachman (2002) found no significant change in the effect of premarital cohabitation on marital dissolution across marriage cohorts using NSFG waves 1988 and 1995. The literature on change over time in the effects of premarital cohabitation on marital dissolution in the U.S. yields divergent conclusions.

DATA AND METHODS

We generated a new harmonized event history dataset using the nine available cycles of the National Survey of Family Growth (National Center for Health Statistics 2016, hereafter NSFG, Copen et al. 2012). In this analysis of premarital cohabitation and marital dissolution, we use the 6 waves starting with 1988 (1988, 1995, 2002, 2006-10, 2011-13, and 2013-15) which included questions about premarital cohabitation, to analyze marital dissolution risk for first

marriages for women age 15-44 (NSFG did not add male respondents until the 2002 wave). NSFG was designed to study fertility, hence the age restriction to subjects still in the childbearing years. We examine women in first marriages exclusively because second and third marriages occur later in life, and marriage duration is heavily truncated for second and third marriages in the age-restricted NSFG.

All marriages recorded in NSFG were heterosexual marriages, i.e. marriages between a man and a woman. The NSFG surveys are retrospective surveys, which has advantages as well as limitations. One advantage of retrospective surveys is that there is no loss to follow-up, as subjects were interviewed only once. One disadvantage to the retrospective design of NSFG is that the age window of the subjects who experienced events becomes more constricted as one examines events further in the past before the survey. An additional limitation of the retrospective design of NSFG is that information about individual or household income is generally only available at the time of the NSFG survey, which means NSFG lacks times series data on household incomes, and therefore lacks the ability to predict marital dissolution from household income.

The following variables are available in every wave and are used as controls in every event history model below: wife's race (white, black, or other), wife's education (time varying), wife's mother's education, wife's age at marriage, whether wife grew up in an intact family of two parents or not, marital duration of wife's first marriage (time varying), and the presence of minor children in the home (time varying). Spouse's race is not included in the analyses because spouse's race was not available in the 1988 wave of NSFG. Information on non-marital cohabitation (i.e. cohabitation that did not lead to marriage) was included in NSFG beginning with the 1995 wave.

Our descriptive statistics from NSFG are weighted by a cross-wave harmonized analytic weight (weights rescaled to have mean equal to 1 within each wave). In our analyses we use discrete time event history logistic regression with NSFG data in a couple-year format (Yamaguchi 1991). Our dependent variable is marital dissolution, which transitions from zero to 1 in the year of divorce or separation, whichever comes first. Separate analyses with divorce only as the dependent variable (not shown below) yield similar substantive results. The two main differences between divorce as an outcome and marital dissolution as an outcome (including divorce and separation without divorce, whichever occurs first) is that black married couples were more likely to separate without getting divorced (Raley and Bumpass 2003, Sweeney and Phillips 2004), and divorce generally occurs after separation. In the analyses below that examine the early years of marriages, a slightly longer definition of the early years of marriage would have to be used if divorce were the only outcome, but the substantive results would be the same.

The 18,674 of the 24,888 married women in our NSFG dataset were white married women, and only 4,182 of the women in the dataset were black (2,032 had 'Other' race). The results of our analyses apply to white American women, but the results do not apply in the same way to black women, whose marriage patterns are quite distinct (see Cherlin 1992). Our results suggest that the association between premarital cohabitation and later marital dissolution is substantially less for black women than for white women (see Appendix 2).

Our event history logistic regressions are unweighted, to preserve likelihood maximization and the associated BIC tests. All event history regressions include predictors of the NSFG weights, race and dummy variables for wave (Winship and Radbill 1994). The full event history dataset (including control variables that were available in all waves) from the 1988 wave

forward has 24,888 women in first marriages, and 216,455 couple-years, for couples without missing data, and 8,488 marital dissolutions.

Because NSFG is a dataset with substantial sample size and therefore substantial statistical power to identify modest changes, the statistical significance of coefficients (at the traditional 5% level) and the statistical significance of likelihood ratio tests can misleadingly identify non-substantive effects as significant (Raftery 1995). The Bayesian Information Criteria is defined as $BIC = -LRT + (\ln(N))df$, where LRT is the Likelihood Ratio Test between two nested models, df is the number of degrees of freedom difference between the models, and N is the sample size. In the case of discrete time event history models, Raftery recommends the number of events (i.e. the number of marital dissolutions) for N . An alternate choice for N , the number of first marriages, would be roughly 3 times larger than the number of marital dissolutions in NSFG, and this larger N would make the BIC even more conservative and parsimony favoring, though the difference in BIC statistics would not change any of the substantive results below. Negative values of BIC are associated with better fit. The larger the N , the more dramatic a difference in LRT has to be in order to be significant by BIC. Raftery considers BIC values more negative than -10 to be indicative of statistically significant improvement in goodness of fit. In this paper when we refer to BIC we will be referring to comparisons between a substantively interesting model and the constant-only model. For models based on the same data subset, the model with the lowest BIC is best. In the comparison of substantive models to each other, we take the difference of the two models' BICs and we will refer to that difference as ΔBIC .

An alternate to the LRT and the BIC is the Akaike Information Criterion, or AIC (Akaike 1974). The AIC is defined as $AIC = -LRT + 2df$, with smaller values indicating better fit. Since

the NSFG data have N of events in the thousands, $\ln(N)$ is always substantially larger than 2, and therefore the BIC is more parsimony favoring than AIC.

There were no missing values for subject's race, the time-varying presence of children, or age at first marriage after NSFG imputation of missing values. Family of origin stability and subject's education were each missing in less than 1% of first married women. As the time axis for historical change, we use either calendar year or year of marriage. Calendar year minus years of marital duration equals year of marriage, so only two of the three predictors can be included linearly in a model. We prefer calendar year as the time axis for Figure 2. Year of marriage (i.e. marriage cohort) shows a slight tendency to interact more strongly with predictors of marital dissolution, for reasons we discuss below.

RESULTS

[Figure 1 here]

We begin with an examination of the extraordinary change in the percentage of women who cohabited before marriage with the man who became their first husband. Eleven percent of women who married for the first time in 1970 had cohabited with the marital partner before marriage, according to NSFG. The percentage of women who cohabited with the marriage partner before first marriage rose dramatically in the subsequent years, reaching 34% in 1980, 46% in 1990, 60% in 2000, and peaking at 70% in 2011. The prevalence, and therefore the selectivity of premarital cohabitation has changed dramatically over time.

[Figure 2 here]

Figure 2 shows the raw odds ratios of breakup (smoothed by 5 year moving averages) for married couples who cohabited before marriage, compared to married couples who did not cohabit before marriage. The Y axis scale is a log scale, because the natural log of the odds ratio is asymptotically Normal. Along with raw odds ratios, Figure 2 also plots the adjusted odds ratios, adjusted by event history logistic regressions controlling for marital duration, age at marriage, presence of minor children, education, race, family of origin stability, calendar decade, and NSFG wave. The adjusted odds ratios were between 1.2 and 1.4 for each decade in Figure 2. The adjusted odds ratio for marital dissolution appeared to decline a little in the 2010s in Figure 2, but the 95% confidence interval was much wider in the 2010s because the NSFG had relatively few marriages and marital dissolutions reported at the end of the time series.

Figure 2 shows that, for the years in which NSFG has substantial numbers of marriages and breakups, there was no apparent trend over time in the raw or adjusted odds ratios of breakup for premarital cohabitation. Given the enormous changes over time in the prevalence of premarital cohabitation (see Figure 1), Figure 2 shows a surprising stability in the association between premarital cohabitation and marital dissolution over time.

[Table 1 here]

The top panel of Table 1 tests the significance of the association between premarital cohabitation and marital dissolution, across the whole NSFG dataset. The tests are highly significant, because, as the literature has generally shown, premarital cohabitation has been

associated with higher odds of marital dissolution, 1.37 times higher than couples who never cohabited (Model 1, without controls), or 1.31 times higher with controls (Model 2). The Likelihood Ratio Test of premarital cohabitation's impact, including controls (Model 2), was 117.54 on 1 degree of freedom, which corresponded to a P value of approximately 2×10^{-27} . The Δ BIC value of -108.49 for this test corresponds to a probability that the model without premarital cohabitation predicts marital dissolution better of $e^{(-108.49/2)} = 2.7 \times 10^{-24}$ (Raftery 1995). The top panel of Table 1 shows that premarital cohabitation predicted marital dissolution (across all NSFG waves combined) to a high degree of statistical certainty, which is consistent with a broad literature.

The bottom panel of Table 1 tests the effect of premarital cohabitation on marital dissolution interacted with different operationalizations of time. Whether time was operationalized as categorical decades (Model 3), linear calendar year (Model 4), or linear marriage year (Model 5), the tests of premarital cohabitation interacted with time were mostly not significant. Figure 2 already showed how flat the interaction between premarital cohabitation and marital dissolution has been across calendar years. The Likelihood Ratio Test (LRT) of premarital cohabitation's interaction with linear calendar year was 0.3 on 1 degree of freedom, consistent with the null hypothesis of no change over time for premarital cohabitation's effect on marital dissolution. The Δ AIC and Δ BIC values (1.7 and 8.7, respectively) similarly rejected premarital cohabitation's effect on marital dissolution changing across calendar years. The LRT, AIC, and BIC tests also firmly rejected significant changes in Table 1, Model 3, which tested for non-linear changes (across decades) in the risk of marital dissolution associated with premarital cohabitation.

Further evidence for the stability (over marriage cohorts) of the association between premarital cohabitation and marriage dissolution can be found in Appendix Figure 2. Appendix Figure 2 shows Kaplan- Meier (1958) survival curves for couples who cohabited compared to couples who never cohabited before marriage, separately for three marriage cohorts (married pre 1980; married 1981-1995; and married 1996-2015). The marital survival curves show that married couples who did not cohabit before marriage had more stable unions after 5 years of marriage consistently across the three marriage cohort categories.

The lack of significant findings of change in premarital cohabitation's association with marital dissolution over time is not due to a lack of power in the NSFG. The NSFG data for Table 1 contains 216,455 couple-years of first marriages. Of these 216,455 couple years, 78,575 were for couples who had cohabited before marriage. If we partition the 78,575 married couple-years of previously cohabiting couples in half, and assign the first half the average annual breakup rate of couples who cohabited before marriage (4.70% breakup rate) and assign the second half a breakup rate equal to the non-cohabiters (3.47%), the power to distinguish between the two breakup rates would be 1 with a two tailed alpha of 0.05, and the power would still be 1 with a two tailed alpha of 0.001, (alpha of 0.001 more closely approximates a difference that would be significant by BIC). In other words, the harmonized NSFG event history dataset was large enough to allow for powerful tests of even small changes over time.

The interaction between premarital cohabitation and linear marriage year in Model 5 of Table 2 yielded ambiguous results about the change of premarital cohabitation's association with marital dissolution over time. The LRT (3.78 on 1 degree of freedom) was significant at the $P < 0.10$ level suggesting a significant change in the association between premarital cohabitation and marital dissolution across marital cohorts. The Model 5 AIC statistic of -1.8 was also

consistent with a change in the association between premarital cohabitation and marital dissolution across marital cohorts. The parsimony-favoring BIC in Model 5 preferred the model without the interaction with marriage year (Δ BIC value of 5.27), and rejected change over marital cohorts in the association between premarital cohabitation and marital dissolution. The BIC test yielded a substantively different result than the LRT and AIC tests in Model 5.

The literature that has reported evidence of a declining effect of premarital cohabitation on marital dissolution has used marriage year, i.e. marriage cohort (rather than calendar year) as the time axis (Copen et al. 2012, Reinhold 2012). If one were to use the traditional LRT as the criteria for accepting or rejecting the null hypothesis, the LRT (which yielded a P value in this case the same as the P value of the coefficient for interaction between premarital cohabitation and linear marriage year) yielded results consistent with prior literature that found a decline in the effect of premarital cohabitation over marriage cohorts.

Why would premarital cohabitation appear to have a (marginally significant by the LRT test) association with marriage year, but not with calendar year? The apparent association between premarital cohabitation, marriage cohort, and marital dissolution requires investigation.

[Figure 3 here]

RECONCILING THE DIVERGENT FINDINGS ON PREMARITAL COHABITATION AS A PREDICTOR OF BREAKUP

In this section we offer an explanation for the divergent findings in the empirical literature regarding the consistency or change in premarital cohabitation's association with marital dissolution. Figure 3 shows that premarital cohabitation does not impact the chances of

marital dissolution early in marriages the same way it does later in marriages. Taking all NSFG waves together since premarital cohabitation was first measured in the NSFG wave of 1988, Figure 3 shows that in first year of marriage, the breakup rate was higher for couples who had not cohabited than it was for couples who did cohabit. The raw difference in breakup rates (between formerly cohabiting couples and non-cohabiting couples) was not statistically significant at 12 months, but the difference was statistically significant up to 6 months of marital duration (from Kaplan-Meier survival analysis, not shown). The risk of quick separation was higher for newly married couples who never cohabited than for couples who did cohabit before marriage. In multivariable analyses below, we document the significant effect of cohabitation in reducing marital breakup in the very early stage of marriage.

In the first year of marriage, couples who had not cohabited had a breakup rate of 4.1%, whereas couples who had cohabited had breakup rate of 3.9%. The couples who did not cohabit before marriage showed the classic pattern of steadily falling marital dissolution rates for the duration of the marriage (until the number of marital dissolution events becomes sparse in the age-constrained NSFG at marital duration of 14 years). In contrast, the couples who cohabited before marriage did not reach their peak marital dissolution rate until marital duration between 2 and 5 years. The different pattern of breakup by marital duration for couples who cohabited before marriage versus couples who did not cohabit before marriage entirely explains why Copen et al (2012) found that there was no measurable difference in cumulative marital dissolution rates between the cohabiters and the non-cohabiters in the first five years of marriage using NSFG 2006-10.

Considering that the median duration of cohabitation before marriage in NSFG was 15 months (with a mean of 24 months), it is logical that it would have taken about one year for the

married couples who never cohabited to catch up in their practical experience of living together with the couples who had cohabited before marriage. It is probably also the case that the learning curve of how to live with a partner was steep at the beginning, and declining over time. In other words, most of the difficult and important practical partner-specific learning about how to live together took place at the beginning of cohabitation or the beginning of marriages that were not preceded by cohabitation. Bumpass, Sweet, and Cherlin (1991) showed that belief in the practical experiential benefit of cohabitation was the primary reason that couples cohabited. Given that the median premarital cohabitation duration was slightly more than a year, it may be inferred that cohabiting couples believed that a year's practical experience of living together is a sufficient amount of experience to be ready for an expected lifetime together.

Lillard, Brien and Waite's (1995) finding that the duration of premarital cohabitation did not predict later marital dissolution is also consistent with most of the practical experience of cohabitation occurring in the first year of cohabitation, so that the Practical Experience of Cohabitation does not increase substantially after the first year. Consistent with Lillard, Brien, and Waite, we also found that duration of cohabitation was not statistically significant and did not improve the goodness of fit significantly when added to any of the models predicting marital dissolution in Table 1 (results available from the authors).

Although the three-way interaction between early marital duration, premarital cohabitation, and marital dissolution was not exactly the same in every wave of NSFG, the fundamental pattern was not significantly different across waves (according to models not shown that tested the 4-way interaction of premarital cohabitation, marital duration, marital dissolution, and NSFG wave). In the first year of marriage, the experience advantage of the premarital cohabiters appears to offset the association between premarital cohabitation and higher rates of

marital dissolution. As marriages progressed beyond 2 years of duration, and as couples who did not cohabit before marriage gained specific experience in how to coreside with their spouses, the couples who had not cohabited before (across all marriage cohorts in NSFG) came to have lower rates of marital dissolution. In Kaplan-Meier (1958) survival analysis (see Appendix Figures 1 and 2) for the combined NSFG data, the cumulative survival rate as married couples became significantly higher for non-cohabiters compared to cohabiters at marital duration of 5 years, with a widening gap in cumulative marital dissolution thereafter.

An alternate version of Figure 3 (available from the authors), with divorce rather than marital dissolution as the outcome, shows a substantively similar pattern. Because divorce takes time to accomplish (whereas breakup or separation can be implemented quickly), the divorce rates for all couples (regardless of premarital cohabitation) were low in the first year of marriage, and in the subsequent year. By marital duration of 5 years (and for every year thereafter), the couples who had cohabited before marriage had a substantially higher rate of divorce than the couples who never cohabited.

[Table 2 here]

In Table 2, we explore the associations between premarital cohabitation, marital dissolution, and marital duration in more detail. Table 2 reports log odds ratio coefficients and summary statistics for event history models predicting marital dissolution. In Table 2, we report log odds ratios rather than odds ratios because the small differences in the coefficients are easier to identify in log odds ratio form. Model one includes no interactions between premarital cohabitation and time. Whereas Figure 1 used calendar year as the time axis, in Table 2 we use

marriage cohort as the time axis, because marriage cohort has been reported to be significantly associated with declining effect of premarital cohabitation on marital dissolution in some previous literature (Copen et al. 2012, Reinhold 2012).

Model 2 of Table 2 introduces an interaction between premarital cohabitation and the marriage cohort (in continuous years). In Model 2, the interaction between premarital cohabitation and marriage cohort was negative, and very nearly significant at the 0.05 level (-0.043/0.022 yields a Z score of -1.95, and a P value of 0.051). Model 2, in other words, comes very close to endorsing a significant decline in the effect of premarital cohabitation on marital dissolution over time, by the criteria of significant coefficients (using the standard two-tailed $\alpha=0.05$). The likelihood ratio test comparison of Models 1 and 2 has the same P value of 0.051, coming very close to significantly preferring Model 2 to Model 1. The parsimony-favoring BIC, however, firmly prefers Model 1 to Model 2 (as Model 1's BIC is 5.27 points more negative than Model 2's BIC). The comparison of Model 2 and Model 1 helps to illustrate what is at stake in choosing different criteria for goodness of fit. In a large N study such as the harmonized NSFG, small and potentially fragile interactions can more easily achieve significant coefficients than significant goodness of fit improvements by BIC.

Model 3 adds a simple dummy variable interaction between premarital cohabitation and the calendar year of a couple's marriage. As seen in Figure 3, couples who cohabited before marriage had a much lower rate of breakup than expected in the calendar year of marriage. The coefficient for the interaction between premarital cohabitation and the year of marriage in Model 3 was a highly significant -0.39, more than offsetting the log odds ratio of 0.32 for premarital cohabitation in general. Model 3 also fits better than Model 2 by the BIC criteria, improving (i.e. making more negative) Model 2's BIC by a significant 10.14 points. Importantly, the inclusion

of the interaction between the calendar year of marriage and premarital cohabitation reduced the interaction between cohabitation and marriage cohort to a firmly insignificant -0.0035 (standard error 0.0022 ; Z-score -1.58 ; P value 0.115). A full set of model coefficients for Models 3 and 5 can be found in Appendix 1.

Once the two-way interaction between early marriage stage and premarital cohabitation was accounted for, the interaction between premarital cohabitation and marriage cohort was diminished to insignificance. Model 4 adds a second (insignificant) interaction between premarital cohabitation and marital durations of 5 years or less. Although this new coefficient worsened the goodness of fit, it also further eroded the coefficient for the key interaction between premarital cohabitation and marriage cohort (from -0.0035 in Model 3, to -0.0027 in Model 4). Model 5, which included no interaction between premarital cohabitation and marriage cohort, was the best fitting of the models in Table 2 by BIC ($-2,406.31$ is the lowest of the 5 BICs), showing that the BIC rejected change over marital cohorts in premarital cohabitation's association with marital dissolution. If instead of marital dissolution (divorce or separation), we used divorce alone as the outcome variable, Table 2 would yield similar substantive findings (alternate results available from the authors)

Table 2 illustrates the benefit of using a conservative criterion such as BIC with large datasets, in the interest of avoiding the endorsement of statistically fragile findings. In this case, the apparent reduction of premarital cohabitation's association with marital dissolution across marriage cohorts was a fragile finding. Table 2 also illustrates how a failure to account for interaction between early marriage duration and premarital cohabitation can yield an over-estimate of the change across marriage cohorts in the effect of premarital cohabitation on marital dissolution risk. Previous analyses that found a declining effect of premarital cohabitation on

divorce over time failed to take account of the Practical Experience of Cohabitation, acting to preserve marriages in their first year.

[Table 3 here]

PREMARITAL AND NONMARITAL COHABITATIONS:

For the 1995 wave, NSFG began asking questions about nonmarital cohabitations, i.e. cohabitations that did not lead to marriage. Table 3 incorporates the nonmarital (cohabitations with men other than the future first husband) as well as premarital cohabitations, as predictors of marital dissolution, dropping the 1988 wave of NSFG and using the 1995 and later waves. Model 1 of Table 3 shows a highly significant association between premarital cohabitation and marital dissolution (log odds of 0.35) which is entirely reversed in the first year of marriages (log odds of -0.36). Model 2 adds the nonmarital cohabitations and their interactions with marriage cohort as predictors. Model 2 improves the goodness of fit over Model 1 dramatically, as nonmarital cohabitation before marriage is an especially powerful predictor of marital dissolution. The coefficient for premarital cohabitation in Model 2, 0.18 was only half as large as Model 1 (0.35), because a large part of the association between premarital cohabitation and marital dissolution was accounted for by the association between premarital cohabitation (with the eventual first spouse) and nonmarital cohabitation (with other partners). According to Model 2, the odds of marital dissolution in any given year of marriage were $e^{0.18}=1.19$ times higher for women who cohabited with their husband before marriage. The odds of marital dissolution were $e^{1.18}=3.25$ times higher for women who had (before marriage) one nonmarital cohabiting partner, and $e^{1.54}=4.66$ times higher for women who had two or more nonmarital cohabitations before

marrying their first husband. Consistent with Teachman (2003), we found that nonmarital cohabitation was a much stronger predictor of marital dissolution than premarital cohabitation with the marriage partner.

Table 3, Model 3 shows that nonmarital cohabitation (with a partner other than the first husband) yielded no significant benefit (in reducing marital dissolution) in the first year of marriage, whereas premarital cohabitation (with the future husband) yielded a consistently significant benefit (in reducing marital dissolution) in the first year of marriage. The Practical Experience of Cohabitation is partner-specific.

Model 4 of Table 3 dispenses with the non-significant and marginally significant predictors of marital dissolution, to yield the best-fitting model of Table 3 by the parsimony-favoring BIC (-3693.0). The BIC rejected the interaction between premarital cohabitation and marriage cohort. An alternate version (not shown) of Table 3 with the time axis of calendar years instead of marriage cohorts showed substantively similar findings, except that the standard coefficient significance test for the interaction of premarital cohabitation with calendar year was not even marginally significant.

[Figure 4 here]

Figure 4 illustrates several of the key findings of Table 3. First, the breakup rate of married couples whose wife cohabited with men other than their future husband (i.e. nonmarital cohabitation) was dramatically higher than the breakup rate for women who cohabited only with the future husband (premarital cohabitation). Because the marital breakup rate after nonmarital cohabitation was so much higher, the Y axis scale of Figure 4 covers a wider range of values

than the Y axis scale of Figure 3. Because the Y axis scale of Figure 4 covers such a wide range of values, it is difficult to visually distinguish the bottom two series, the breakup rate for couples who cohabited premaritally, and married couples whose wives never cohabited (with the husband or with other men). Nonetheless, across all marital durations, the breakup rate in this subset of the data was significantly higher (odds ratio of 1.31, corresponding to a log odds ratio coefficient of 0.27) for the couples who cohabited before marriage (but had no other cohabitations) compared to the couples whose wives never cohabited before marriage.

Figure 4 also demonstrates the distinct first year pattern of breakup that we associate with the Practical Experience of Cohabitation. Married couples who cohabited before marriage had a sharply lower first year breakup rate, whether the wife had other prior cohabitation partners or not.

CONCLUSION AND DISCUSSION

We find little support for the Normalization Hypothesis. The Normalization Hypothesis argues that as the formerly rare and stigmatized status of premarital cohabitation became dramatically more common, the penalty in higher marital dissolution rates for former cohabiters should have diminished. The Normalization Hypothesis predicts a convergence over time in marital dissolution rates between groups. Instead of convergence over time, we find that the marital stability disadvantage of premarital cohabitation emerges most strongly after 5 years of marital duration, and has remained roughly constant over time and over marriage cohorts.

In the calendar year of marriage, couples who had previously cohabited actually had lower rates of breakup compared to couples who did not cohabit before marriage. We hypothesize that the practical experience of having lived together gives cohabiting couples who transition to marriage an early advantage (reflected in the lower breakup rate), which lasts a year, over newlyweds who never lived together; this is the hypothesized advantage of the Practical Experience of Cohabitation.

The Practical Experience of Cohabitation is partner-specific. Compared to premarital cohabitation (cohabitation with the future spouse), nonmarital cohabitations (cohabitations with partners other than the future spouse) appear to offer none of the short term experiential advantages (in marital stability) that premarital cohabitation offers. Nonmarital cohabitation has a much stronger association (compared with premarital cohabitation) with later marital dissolution. The Cohabitation Experience of Impermanence is more strongly associated with the breakups of prior nonmarital unions than with the experience of the premarital cohabitation with the future spouse.

The literature on cohabitation has usually described the experience of cohabitation as a negative for marital stability (i.e. the Cohabitation Experience of Impermanence). Both the (positive for marital stability, and short-acting) Practical Experience of Cohabitation and the (negative for marital stability, and longer acting) Cohabitation Experience of Impermanence are consistent with the associations we observed between marital duration, premarital cohabitation, and marital dissolution rates. Failure to account for the interaction between marital duration and premarital cohabitation explains why some researchers have claimed that the most recent marriage cohorts show no association between premarital cohabitation and marital dissolution.

The rate of cohabitation continues to rise in the U.S. due to the delay of first marriage, lower stigma around sex outside of marriage, and expensive housing in urban markets that makes living alone very expensive. Despite decades of research on cohabitation and marriage, it turns out there is much we still need to learn about how cohabitation, marriage, and divorce interrelate. In this paper we show that the effects of premarital cohabitation are positive for marital stability in the first year of marriages. The underlying questions about *why* premarital cohabitation has the effects that it has are not answerable with retrospective surveys such as NSFG. Other kinds of data (prospective, qualitative) about couples and their decisions are needed.

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Table 1: Tests of Premarital Cohabitation's effect on Marital Dissolution, and Interactions with Time

Predictor of Marital Dissolution	Model	Controls applied	OR (and 95% CI) of breakup	LRT	Δ AIC	Δ BIC
Premarital Cohabitation	1	No	1.37 (1.31, 1.43)			
	2	Yes	1.31 (1.25, 1.38)	117.54*** (1 df)	-115.5	-108.49***
Premarital Cohabitation \times Time	3	Yes	Cohab \times decade	3.1 (5 df)	7.0	42.2
	4	Yes	Cohab \times linear calendar year	0.3 (1df)	1.7	8.7
	5	Yes	Cohab \times linear marriage year	3.78+ (1 df)	-1.8	5.27

Source: NSFG data on first marriages, female subjects age 15-44, waves 1988 and later.

Note: Results from unweighted discrete time event history models in logistic form. All predictors significant by LRT. For BIC, the more negative the value, the greater the significance. Controls: marital duration, age of first marriage or cohabitation, minor children, education, race, stable parents, decade, wave, mother's education. Results are from unweighted discrete time event history models in logistic form. For BIC, values smaller than -10 are statistically significant. LRT is Likelihood Ratio Test, AIC is Akaike Information Criterion, BIC is Bayesian Information Criterion. AIC values less than zero are preferred. Number of breakup events, N= 8,488 of which 3,697 occurred to women who had cohabited with their first husband before marriage. N of couple years is 216,455. Model 2 tests are tests of premarital cohabitation's significance as a predictor of marital dissolution. Models 3,4, and 5 tests are tests of the interactions of premarital cohabitation with different constructions of time.

Δ AIC=-LRT+2df. Δ BIC=-LRT+(ln(N))df.

+ P<0.10; * P<0.05; ** P<0.01; ***P<0.001

Table 2: Testing Interactions with premarital cohabitation: Log Odds Ratio Coefficients and Summary Statistics from Discrete Time Event History Models Predicting Marital Dissolution (SE in Parentheses)

	M1	M2	M3	M4	M5
Premarital Cohabitation	0.27*** (0.025)	0.30*** (0.021)	0.32*** (0.031)	0.36*** (0.038)	0.29*** (0.026)
Premarital Cohab × Marriage Cohort (year)		-0.0043+ (0.0022)	-0.0035 (0.0022)	-0.0027 (0.0023)	
Premarital Cohab × Calendar Year of Marriage			-0.39*** (0.090)	-0.37*** (0.092)	-0.40*** (0.090)
Premarital Cohab × First 5 Years of Marriage				-0.067 (0.043)	
df	27	28	29	31	28
LRT compared to constant only	2,639.12	2,642.90	2,662.09	2,664.58	2,659.61
Δ LRT compared to previous model		3.78	19.19	2.49	
Chisquare P(Δ LRT, 1df)		0.051	1.18 × 10 ⁻⁵	0.288	
BIC Compared to Constant Only	-2,394.87	-2,389.60	-2,399.74	-2384.14	-2,406.31
Δ BIC from previous model		5.27	-10.14	15.60	-22.17

Source: NSFG data on first marriages for women age 44 and under, NSFG wave 1988 and later.

Note: Additional controls not reported above: marital duration (1df), year of marriage (1df) marital duration first calendar year dummy variable (1df), age at marriage (categorical, 3df), presence of children under 18 (1df), calendar decade (5df), educational attainment (3df), race (2df), stable family of origin (1df), mother's education (3df), NSFG wave (5 df). N of events (marital breakups) is 8,488 (used as the N for BIC), of which 3,697 occurred to women who had cohabited with before marrying their first husband. N of couple-years is 216,455. N of couples is 24,888.

*** P<0.001; ** P<0.01; * P<0.05; + P<0.10, two tailed tests.

Table 3: Comparing the effect of Premarital and other nonmarital cohabitations on marital dissolution: Log Odds Ratio Coefficients and Summary Statistics from Discrete Time Event History Models (SE in Parentheses)

	M1	M2	M3	M4
Premarital Cohabitation	0.35*** (0.04)	0.18*** (0.04)	0.18*** (0.04)	0.13*** (0.03)
Premarital Cohab × Marriage Cohort (year)	-0.0052* (0.0026)	-0.005+ (0.003)	-0.005+ (0.003)	
Premarital Cohab × Calendar Year of Marriage	-0.36*** (.10)	-0.37*** (0.10)	-0.36*** (0.10)	-0.39*** (0.10)
Nonmarital cohab (1)		1.18*** (0.04)	1.18*** (0.04)	1.11*** (0.03)
Nonmarital cohab (2 or more)		1.54*** (0.06)	1.54*** (0.06)	1.46*** (0.04)
Nonmarital cohab (1) × Marriage Cohort		-0.007* (0.003)	-0.006* (0.003)	
Nonmarital cohab (2 or more) × Marriage Cohort		-0.008+ (0.004)	-0.008* (0.004)	
Nonmarital cohab (1) × Calendar year of Marriage			-0.08 (0.12)	
Nonmarital cohab (2 or more) × Calendar year of Marriage			0.07 (0.15)	
df	27	31	33	28
LRT compared to constant only	2222.1	3953.2	3954.2	3940.2
BIC Compared to Constant Only	-1983.8	-3679.5	-3662.9	-3693.0
Δ BIC from previous model		-1695.8	16.7	-30.1

Source: NSFG data on first marriages for women age 44 and under, NSFG wave 1995 and later.

Note: Additional controls not reported above: marital duration (1df), year of marriage (1df) marital duration first year dummy variable (1df), age at marriage (categorical, 3df), presence of children under 18 (1df), calendar decade (4df), educational attainment (3df), race (2df), stable family of origin (1df), mother's education (3df), NSFG wave (4 df). N of events (marital breakups) is 6,820 (used as the N for BIC), of which 3,304 breakups were recorded by women who cohabited with their first husbands before marriage. Premarital cohabitation is cohabitation with the woman's first husband before marriage. Nonmarital cohabitation is prior cohabitation with men other than the man who would later become the woman's first husband. N of couple-years is 167,723. N of couples is 19,777.

*** P<0.001; ** P<0.01; * P<0.05; + P<0.10, two tailed tests.

**Premarital Cohabitation at time of marriage from NSFG
for women under age 45.
Moving Average of Weighted data**

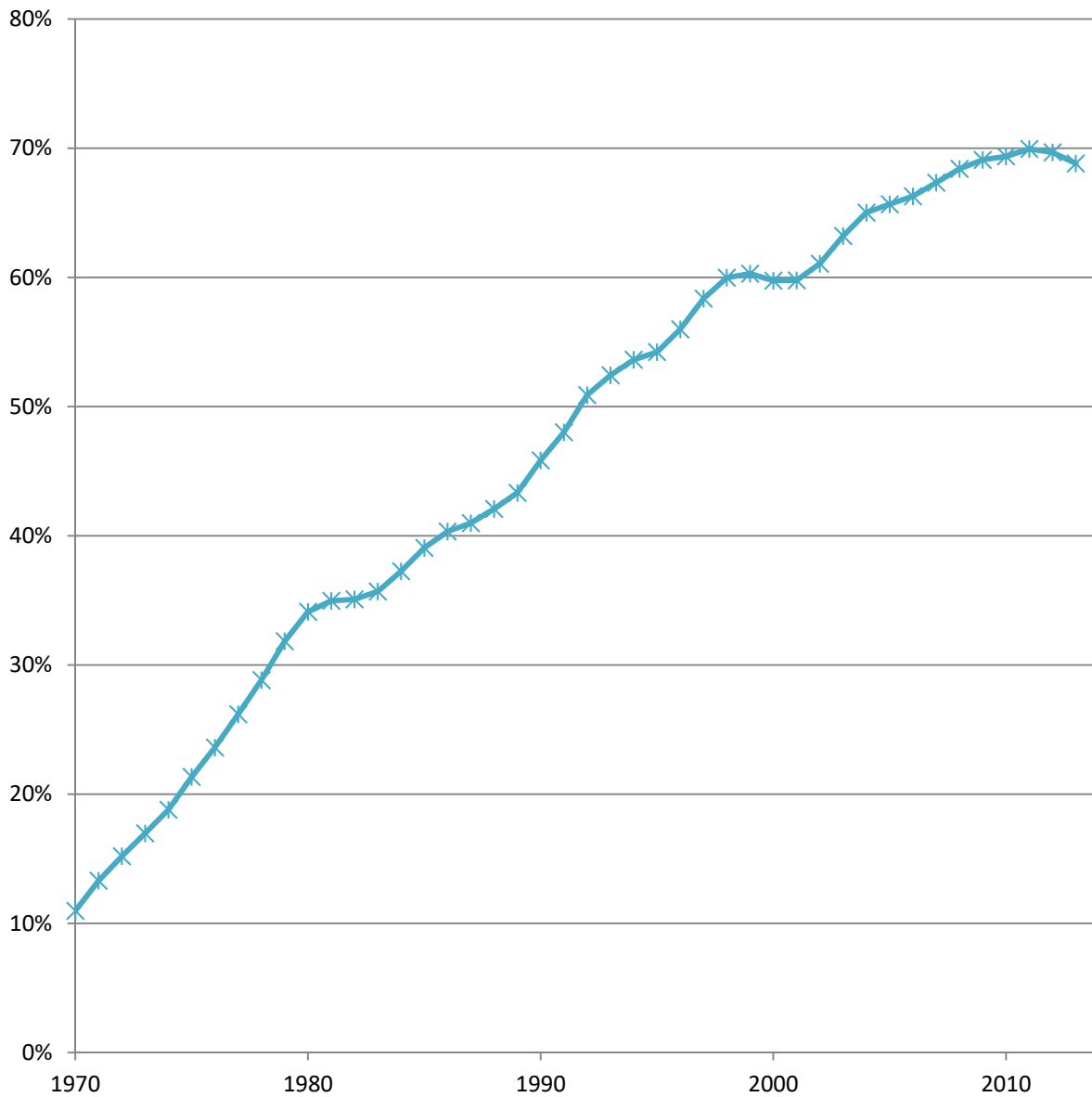


Figure 1: Weighted NSFG data on first marriages, female subjects age 15-44, smoothed with 5 year moving average. Last two years, 2014 and 2015 not shown because the number of new marriages was too small.

Premarital Cohabitation's stable association with marital dissolution across calendar years

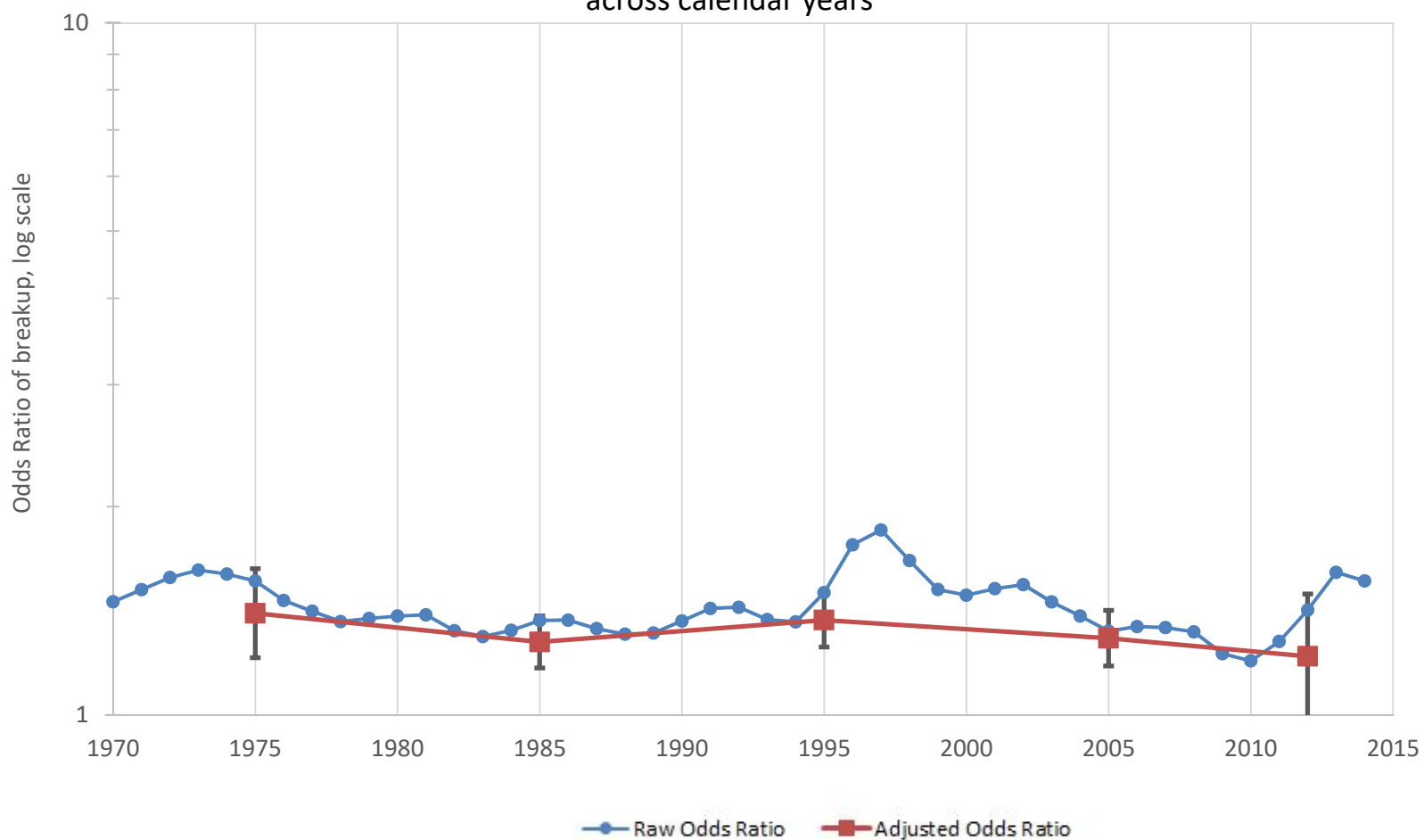


Figure 2: Raw Odds Ratios of Breakup, and Adjusted Odds Ratios of Breakup (with 95% CI) for First Marriages by Calendar Year

Source: NSFG data on first marriages, female subjects age 15-44. Unadjusted Odds Ratios are weighted, and smoothed with 5 year moving average. Adjusted Odds Ratios interacted with decade are derived from unweighted discrete time event history logistic regressions, controlling for: marital duration, age at marriage, presence of minor children, subject's education, subject's race, family of origin stability, mother's educational attainment, NSFG wave.

Annual Breakup Rate for First marriages, with and without premarital cohabitation

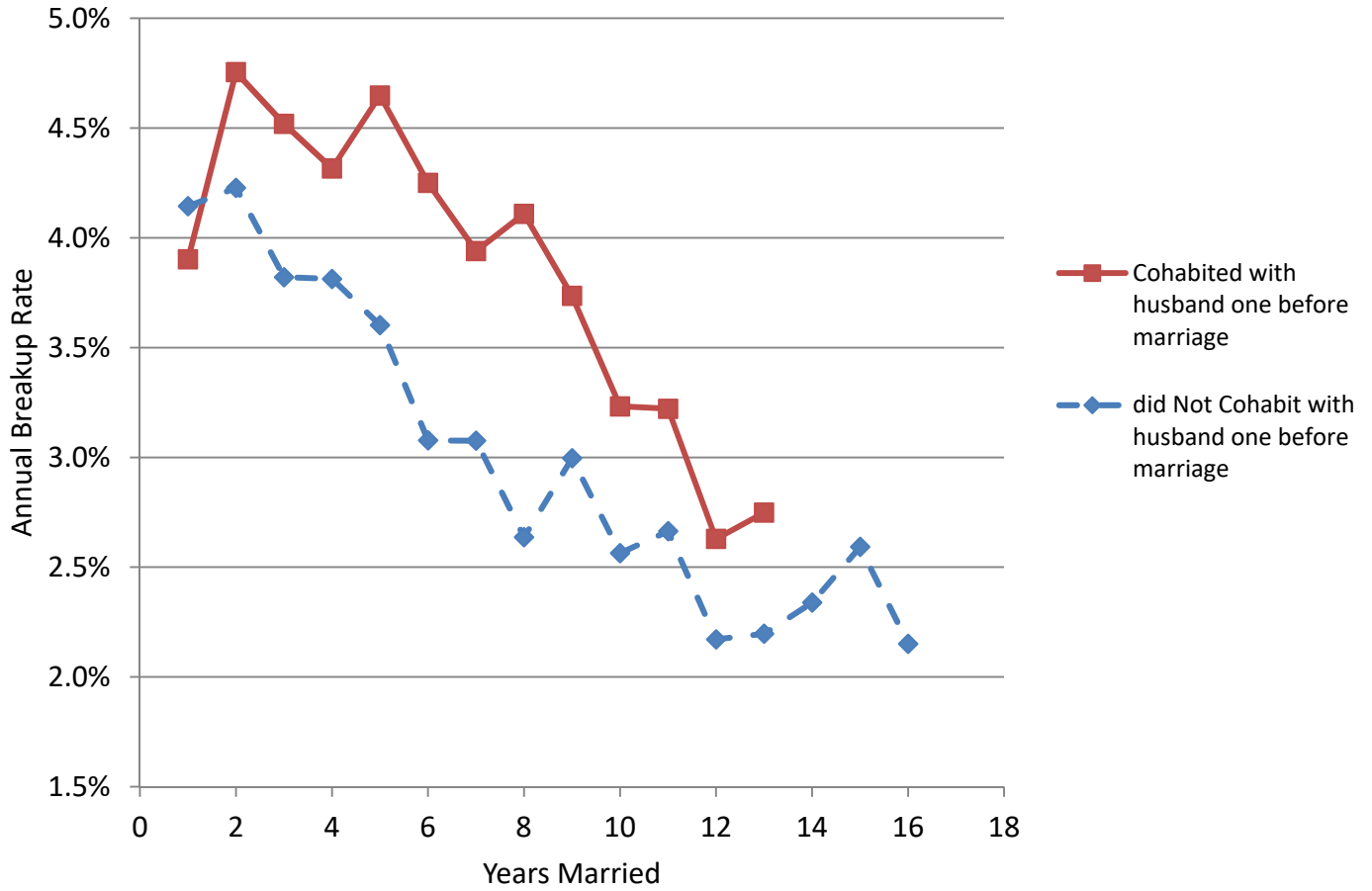


Figure 3. Weighted NSFG data on first marriages, female subjects age 15-44, waves 1988 and later. In this graph, the 12 months between marriage and the first anniversary is year 1, and the next year is year 2, and so on.

Annual Breakup Rate for First marriages, by Nonmarital and Premarital Cohabitation

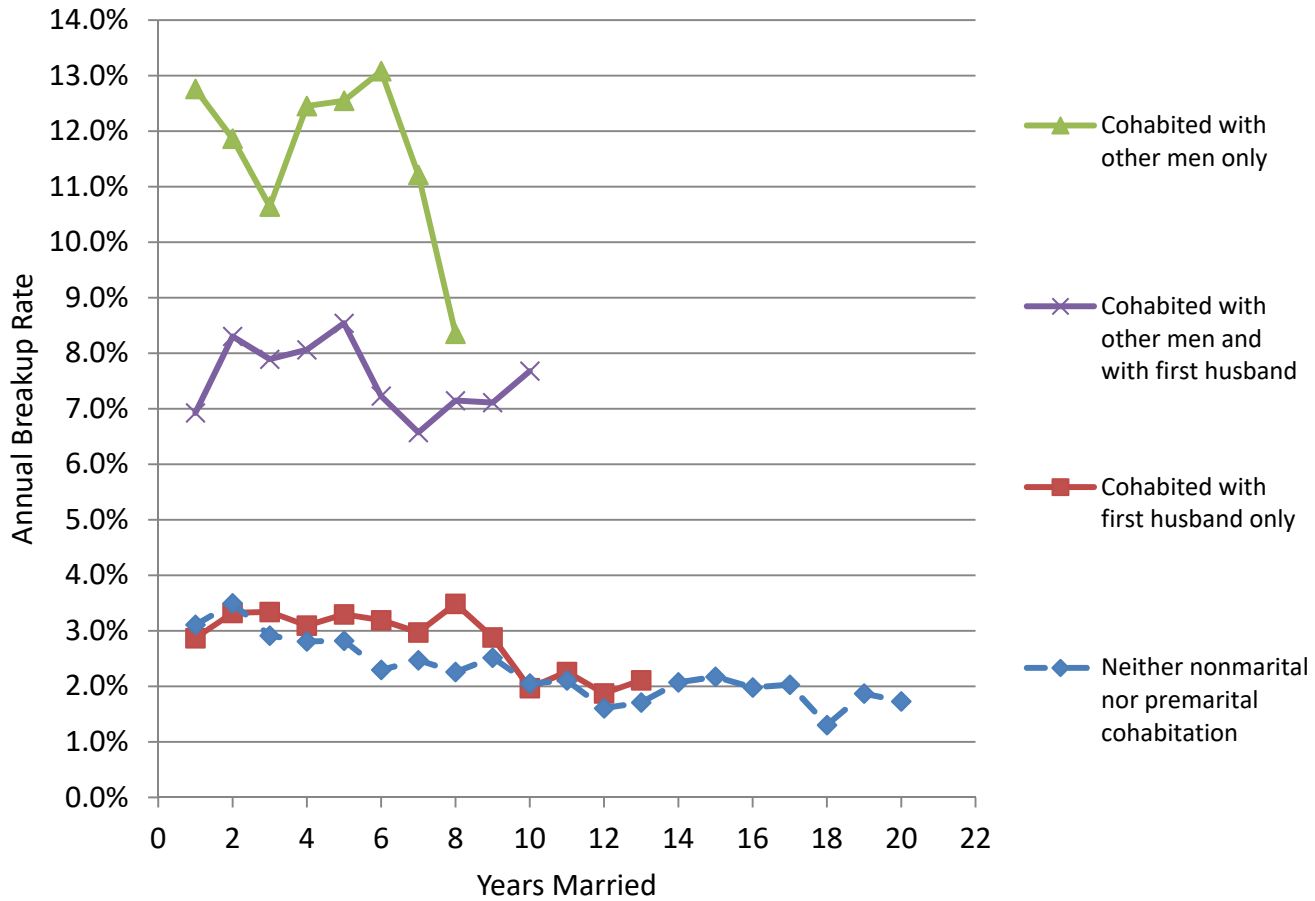


Figure 4. Weighted NSFG data on first marriages, female subjects age 15-44, waves 1995 and later. In this graph, the 12 months between marriage and the first anniversary is year 1, and the next year is year 2, and so on. Sample size of wives who cohabitated only with other men was 1443; there were 2950 wives who cohabitated with first husband and with other men; there were 6801 wives who cohabitated only with first husband; there were 8920 wives who never cohabitated before marriage with anyone. Time series are trimmed where the number of couples becomes too sparse.

Appendix 1: Full set of controls shown for two models from Table 2. Discrete time logistic models predicting marital dissolution.

	M3	M5
Premarital Cohabitation	0.32*** (0.031)	0.29*** (0.026)
Premarital Cohabitation × Marriage Cohort (year)	-0.0035 (0.0022)	
Premarital Cohab × Calendar Year of Marriage	-0.39*** (0.090)	-0.40*** (0.090)
Marriage Cohort (year)	0.021*** (0.005)	0.019*** (0.004)
Calendar year of marriage	-0.83*** (0.06)	-0.83*** (0.06)
Marital duration, in years	-0.029*** (0.005)	-0.039*** (0.005)
Age at marriage (ref <18): 18-19	-0.21***	-0.21***
Age at marriage: 20-24	-0.61***	-0.61***
Age at marriage: ≥25	-1.05***	-1.05***
Minor children in HH	-0.18***	-0.18***
Calendar Decade (ref 1980s): 1960s	-0.06	-0.08
1970s	0.03	0.02
1990s	-0.30***	-0.30***
2000s	-0.42***	-0.43***
2010s	-0.38***	-0.39***
Respondent's education (ref<HS): HS	0.015	0.013
Education: Some college	0.08*	0.08*
Education: BA+	-0.34***	-0.34***
Respondent's race (ref: black): white	-0.40***	-0.40***
Respondent's race: Other	-0.66***	-0.66***
Respondent from Stable Family	-0.37***	-0.37***
NSFG wave (ref 1988): 1995	0.07+	0.07+
NSFG wave 2002	0.028	0.03
NSFG wave 2006-10	0.27***	0.27***
NSFG wave 2011-13	0.38***	0.38***
NSFG wave 2013-15	0.32***	0.32***
Mother's education (ref: <HS): HS	0.13***	0.13***
Mother's education: Some College	0.15***	0.15***
Mother's education: BA+	0.18***	0.18***
df	29	28
LRT compared to constant only	2,662.09	2,659.61
BIC	-2,399.74	-2,406.31

Source: NSFG data on first marriages for women age 44 and under, NSFG wave 1988 and later.

*** P<0.001; ** P<0.01; * P<0.05; + P<0.10, two tailed tests.

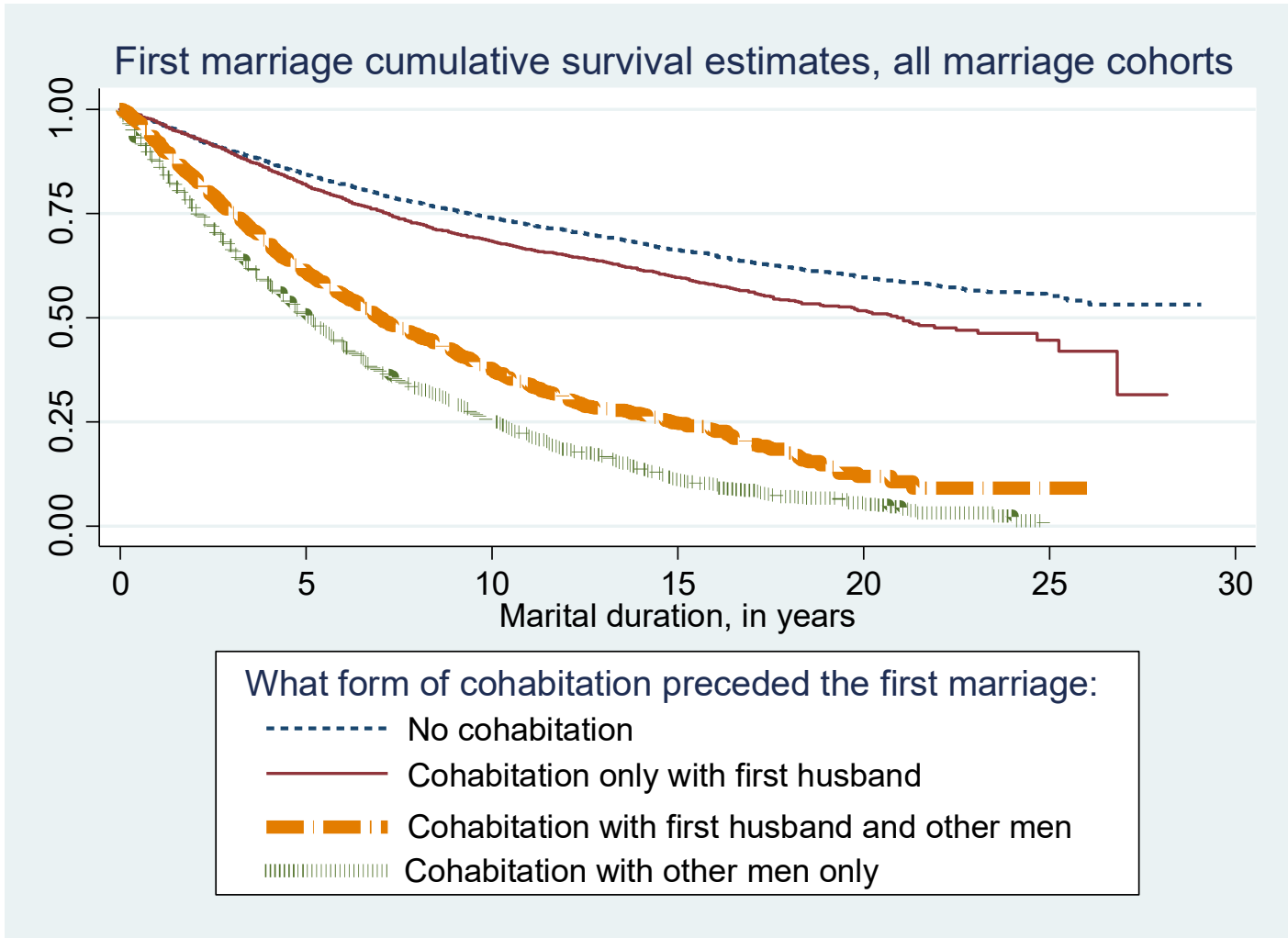
Appendix 2: Predicting marital dissolution: Log Odds Ratio Coefficients and Summary Statistics from Discrete Time Event History Models (SE in Parentheses)

		Table 3, Model 2			
		M2	M2a	M2b	M2c
Additional Filters		None	Married at age<=30	White women only	Black women only
	Premarital Cohabitation	0.18*** (0.04)	0.18*** (0.04)	0.22*** (0.05)	-0.08 (0.09)
	Premarital Cohab × Marriage Cohort (year)	-0.005+ (0.003)	-0.004 (0.003)	-0.006* (0.003)	0.002 (0.006)
	Premarital Cohab × Calendar Year of Marriage	-0.37*** (0.10)	-0.34*** (0.10)	-0.38*** (0.12)	-0.25 (0.22)
	Nonmarital cohab (1)	1.18*** (0.04)	1.18*** (0.04)	1.28*** (0.05)	0.92*** (0.10)
	Nonmarital cohab (2 or more)	1.54*** (0.06)	1.55*** (0.06)	1.55*** (0.07)	1.38*** (0.15)
	Nonmarital cohab (1) × Marriage Cohort	-0.007* (0.003)	-0.004 (0.003)	-0.009** (0.004)	0.002 (0.006)
	Nonmarital cohab (2 or more) × Marriage Cohort	-0.008+ (0.004)	-0.003 (0.004)	-0.006 (0.005)	-0.007 (0.009)
	df	31	31	29	29
	N of couple-years	167,723	159,708	130,059	22,126

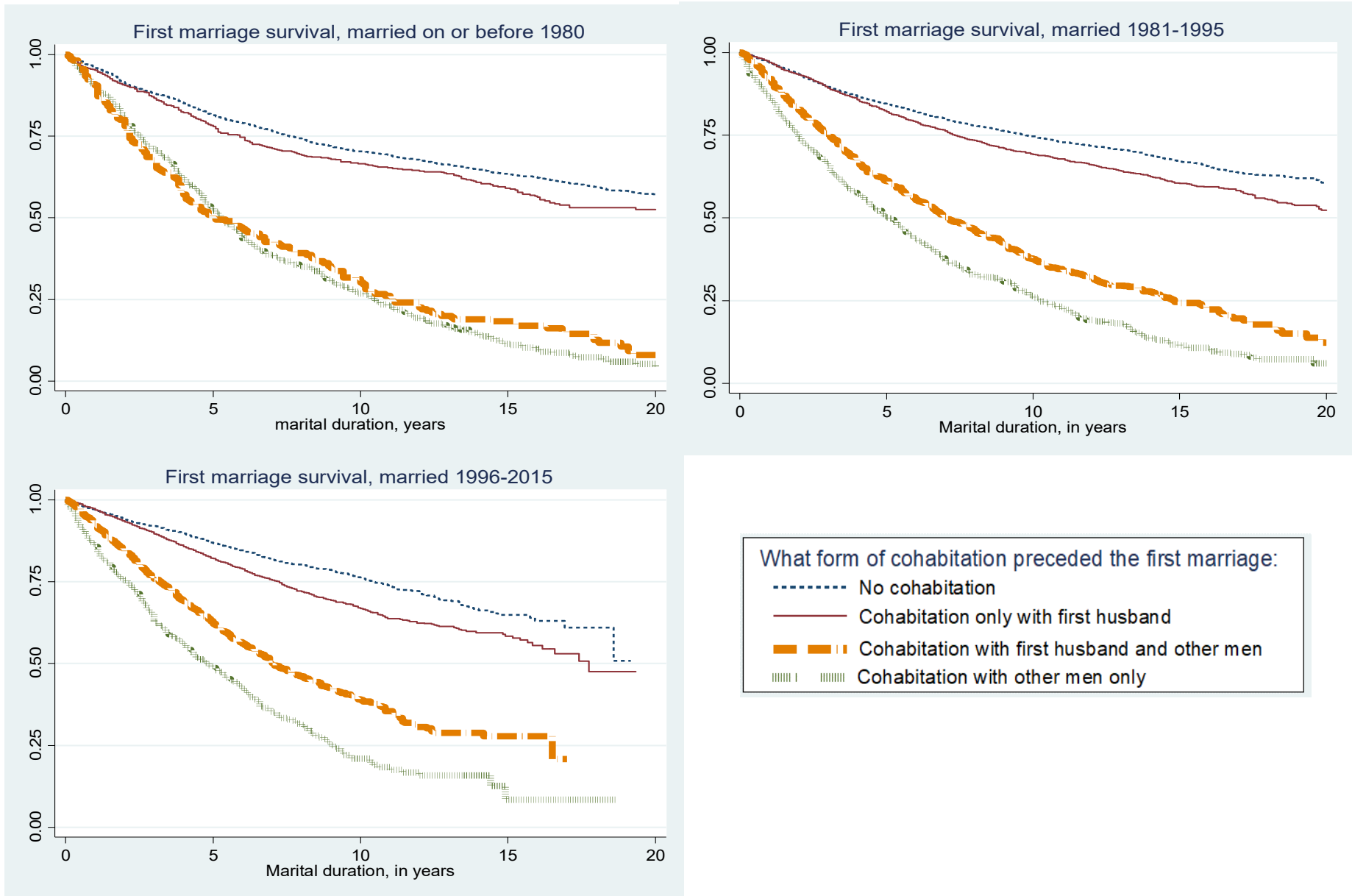
Source: NSFG data on first marriages for women age 44 and under, NSFG wave 1995 and later.

Note: Additional controls not reported above: marital duration (1df), year of marriage (1df) marital duration calendar year dummy variable (1df), age at marriage (categorical, 3df), presence of children under 18 (1df), calendar decade (4df), educational attainment (3df), race (2df, but not included as control in the single-race models M2b or M2c), stable family of origin (1df), mother's education (3df), NSFG wave (4 df). Each of the models above has a different sample of couples from NSFG, so the goodness of fit of the models cannot be compared.

*** P<0.001; ** P<0.01; * P<0.05; + P<0.10, two tailed tests.



Appendix Figure 1. Kaplan-Meier (1958) survival (as intact married couple) estimates based on unweighted data on first marriages, female subjects age 15-44, waves 1995 and later. Sample size of wives who cohabitated only with other men was 1,443; there were 2,950 wives who cohabitated with first husband and with other men; there were 6,801 wives who cohabitated only with first husband; there were 8,920 wives who never cohabited before marriage with anyone.



Appendix Figure 2: First marriage survival by cohabitation history and marriage cohort. Kaplan-Meier (1958) survival (as intact married couple) estimates based on unweighted data on first marriages, female subjects age 15-44, NSFG waves 1995 and later.