Background: Before data existed on premarital cohabitation and divorce, scholars assumed that the experience of premarital cohabitation would select compatible couples into marriage and lead to lower rates of divorce. The advent of data on premarital cohabitation and divorce overturned the early preconceptions, as premarital cohabitation was found to be associated with higher rates of divorce. Premarital cohabitation has risen dramatically in the United States. Scholars disagree about whether the divorce rates of premarital cohabiters and noncohabiters have converged.

Method: A harmonized data set of 6 waves of the retrospective National Surveys of Family Growth (with 216,455 couple-years) is analyzed with discrete time-event history methods to predict marital dissolution. The data are nationally representative of women aged 44 years and younger in first marriages in the United States for 1970 to 2015. Different criteria for model selection are discussed.

Results: The results show that in the first year of marriages, couples who cohabited before marriage have a lower marital dissolution rate than couples who did not cohabit before marriage, a difference that may be due to the practical experience of cohabitation, as couples who have cohabited learned to adapt to each other. We find that the association between marital dissolution and premarital cohabitation has not changed over time or across marriage cohorts. The benefits of cohabitation experience in the first year of marriage has misled scholars into thinking that the most recent marriage cohorts will not experience heightened marital dissolution due to premarital cohabitation.

Conclusion: Premarital cohabitation has short-term benefits and longer term costs for marital stability.

Introduction
The association between premarital cohabitation and divorce in the United States 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 & Mosher, 2002; Cherlin, 1992; Smock, 2000). Premarital cohabitation was rare and stigmatized before 1970 in the United States, but by the 2000s the novelty of and stigma against premarital cohabitation had 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 affect 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 United States 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 a lower hazard of divorce. Cohabiters themselves regarded being certain they were compatible before marriage as the primary benefit of cohabitation (Bumpass, Sweet, & Cherlin, 1991). The potential of cohabitation to filter out incompatible relationships before marriage is what Smock (2000, p. 6) 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 & Edwards, 1992; Bumpass & Sweet, 1989; DeMaris & Rao, 1992) and worse marital adjustment (DeMaris & Leslie, 1984) when compared with 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, Blanc, & Bloom, 1988; Hall & 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, Cohan, & Amato, 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, Brien, and Waite (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 term 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 & 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 article, 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 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), the selectiveness of cohabitation has necessarily diminished over time. Dush et al. (2003) argued that if selection (e.g., of liberals and less religious people) into cohabitation was the reason that premarital cohabitation was observed to be associated with higher rates of divorce, then the rise in popularity of premarital cohabitation would diminish the selectiveness of premarital cohabitation for characteristics that might predict divorce. A decline in the selectivity of premarital cohabitation should yield convergence (over time) with the divorce rate of couples who did not cohabit before marriage if the selection hypothesis is correct. 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 toward cohabiting couples) from friends, family, and strangers (Rosenfeld, 2007). Stigma...
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has been shown to affect individuals’ physical and mental health (Hatzenbuehler, Phelan, & Link, 2013; Riggle & Rostosky, 2007) 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 United States, scholars assumed that the experience of cohabitation would teach couples important practical lessons about how to live together. To refer to this kind of experience, we use the term *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 in some circumstances may be 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 article 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 noncohabiters 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, whereas 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 & Mosher, 2002; Cherlin, 1992; Goodwin, Mosher, & Chandra, 2010). Copen, Daniels, Vespa, and Mosher (2012, table 7), using the 2006–2010 wave of the 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 the NSFG and other data sets), 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 the results of Copen et al. (2010) is that the most recent marriage cohorts no longer experience a 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 because 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 the 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 United States. Dush et al. (2003) used a multistage U.S. phone survey to determine the predictors of marital dissolution rates of couples who were married in 1980 when compared with 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 the NSFG 1988 and 1995 waves. The literature on change over time in the effects of premarital cohabitation on marital dissolution in the United States yields divergent conclusions.

**Data and Methods**

We generated a new harmonized event history data set using the nine available cycles of the NSFG (National Center for Health Statistics, 2016; Copen et al., 2012). In this analysis of premarital cohabitation and marital dissolution, we use the six waves starting with 1988 (1988, 1995, 2002, 2006–2010, 2011–2013, and 2013–2015), which included questions about premarital cohabitation, to analyze marital dissolution risk for first marriages for women aged 15 to 44 years (NSFG did not add male respondents until the 2002 wave). The NSFG was designed to study fertility, hence the age restriction to respondents 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, that is, 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 the respondents were interviewed only once. One disadvantage to the retrospective design of the NSFG is that the age window of the respondents 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 the NSFG is that information about individual or household income is generally only available at the time of the NSFG survey, which means that the NSFG lacks time-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: 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, 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 the NSFG. Information on nonmarital cohabitation (i.e., cohabitation that did not lead to marriage) was included in the NSFG beginning with the 1995 wave.

Our descriptive statistics from the NSFG are weighted by a cross-wave harmonized analytic weight (weights rescaled to have mean equal to one 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 0 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 & Bumpass, 2003; Sweeney & Phillips, 2004), and divorce generally occurs after separation. In the analyses that follow 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.

Of the 24,888 married women in our NSFG data set, 18,674 were White married women, and 4,182 of the women in the data set 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 Table S2).

Our event history logistic regressions are unweighted, to preserve likelihood maximization and the associated Bayesian Information Criterion (BIC) tests. All event history regressions include predictors of the NSFG weights, race and dummy variables for wave (Winship & Radbill, 1994). The full event history data set (including control variables that were available in all waves) from the 1988 wave forward has 24,888 women in first marriages; 216,455 couple-years, for couples without missing data; and 8,488 marital dissolutions.

Because the NSFG is a data set 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 nonsubstantive effects as significant (Raftery, 1995). The BIC is defined as \( \text{BIC} = -\text{LRT} + (\ln(N)df) \), where \( \text{LRT} \) 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 (1995) 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 three times larger than the number of marital dissolutions in the NSFG, and this larger \( N \) would make the BIC even more conservative and parsimony favoring, although the difference in BIC statistics would not change any of the substantive results. Negative values of BIC are associated with better fit. The larger the \( N \), the more dramatic a difference in LRT has to be to be significant by BIC. Raftery (1995) considers BIC values more negative than −10 to be indicative of statistically significant improvement in goodness of fit. In this article, when we refer to BIC we are 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 refer to that difference as ABIC.

An alternate to the LRT and the BIC is the Akaike information criterion, or AIC (Akaike, 1974). The AIC is defined as \( AIC = -\text{LRT} + 2df \), with smaller values indicating better fit. Because the NSFG data have \( N \) of events in the thousands, \( \ln(N) \) is always substantially larger than two, and therefore the BIC is more parsimony favoring than AIC.

There were no missing values for respondent’s race, the time-varying presence of children, or age at first marriage after NSFG imputation of missing values. Family-of-origin stability and respondent’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
Figure 2: Premarital Cohabitation’s Stable Association With Marital Dissolution Across Calendar Years.

Source. Raw odds ratios of breakup and adjusted odds ratios of breakup (with 95% CI) for first marriages by calendar year. National Surveys of Family Growth data on first marriages, female respondents aged 15 to 44 years.

Note. Unadjusted odds ratios are weighted and smoothed with 5-year moving average. Comparison group is married couples who did not cohabit before marriage. 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, respondent’s education, respondent’s race, family-of-origin stability, mother’s educational attainment, National Surveys of Family Growth wave.

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 next.

Results

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. Of the women who married for the first time in 1970, 11% had cohabited with the marital partner before marriage according to the 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 shows the raw odds ratios of breakup (smoothed by 5-year moving averages) for married couples who cohabited before marriage when compared with 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 the 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.

The top panel of Table 1 tested the significance of the association between premarital cohabitation and marital dissolution across the entire NSFG data set. The tests were 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 LRT of premarital cohabitation’s impact, including controls (Model 2), was 117.54 on one degree of freedom, which corresponded to a p value of approximately $2 \times 10^{-27}$. The ΔBIC value of $-108.49$ for this test corresponded to a probability that the model without premarital cohabitation predicted 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 was consistent with a broad literature.

The bottom panel of Table 1 tested 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 LRT of premarital cohabitation’s interaction with linear calendar year was 0.3 on one degree of freedom, consistent with the null hypothesis of no change over time for premarital cohabitation’s effect on marital dissolution. The ΔAIC and ΔBIC 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 nonlinear 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 Figure S2. Figure S2 shows Kaplan and Meier (1958) survival curves for

<table>
<thead>
<tr>
<th>Predictor of marital dissolution</th>
<th>Model</th>
<th>Controls applied</th>
<th>OR [95% CI] of breakup</th>
<th>LRT</th>
<th>ΔAIC</th>
<th>ΔBIC</th>
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<tr>
<td>Premarital cohabitation</td>
<td>1</td>
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<td>1.37 [1.31–1.43]</td>
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<tr>
<td>Premarital Cohabitation × Time</td>
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<td>Yes</td>
<td>1.31 [1.25–1.38]</td>
<td>117.54*** (1 df)</td>
<td>−115.5</td>
<td>−108.49***</td>
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<tr>
<td></td>
<td>3</td>
<td>Yes</td>
<td>Cohab × Decade</td>
<td>3.1 (5 df)</td>
<td>7.0</td>
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<td>4</td>
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<td>Cohab × Linear Calendar Year</td>
<td>0.3 (1 df)</td>
<td>1.7</td>
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<td></td>
<td>5</td>
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<td>Cohab × Linear Marriage Year</td>
<td>3.78+ (1 df)</td>
<td>−1.8</td>
<td>5.27</td>
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</table>

**Source.** National Surveys of Family Growth data on first marriages, female respondents aged 15 to 44 years, waves from 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. The results are from unweighted discrete time-event history models in logistic form. For BIC, values smaller than $-10$ are statistically significant. 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 are tests of the interactions of premarital cohabitation with different constructions of time. AIC = Akaike information criterion, BIC = Bayesian information criterion; Cohab = cohabitation; LRT = likelihood ratio test. ΔAIC = −LRT + 2df. ΔBIC = −LRT+(ln(N))df. ***p < .001; **p < .01; *p < .05; +p < .10.
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 showed 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 was not due to a lack of power in the NSFG. The NSFG data for Table 1 contained 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 noncohabiters (3.47%), the power to distinguish between the two breakup rates would be 1 with a two-tailed \( \alpha \) of .05, and the power would still be 1 with a two-tailed \( \alpha \) of .001, (\( \alpha \) of .001 more closely approximates a difference that would be significant by BIC). In other words, the harmonized NSFG event-history data set 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 df) was significant at the \( p < .10 \) 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 (ΔBIC 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.

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 noncohabiting 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 the multivariable analyses, 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
Table 2. Testing Interactions With Premarital Cohabitation: Log Odds Ratio Coefficients and Summary Statistics From Discrete Time-Event History Models Predicting Marital Dissolution (Standard Errors in Parentheses)

<table>
<thead>
<tr>
<th>Predictors of marital dissolution</th>
<th>Model 1</th>
<th>Model 2</th>
<th>Model 3</th>
<th>Model 4</th>
<th>Model 5</th>
</tr>
</thead>
<tbody>
<tr>
<td>Premarital cohabitation</td>
<td>0.27***</td>
<td>0.30***</td>
<td>0.32***</td>
<td>0.36***</td>
<td>0.29***</td>
</tr>
<tr>
<td></td>
<td>(0.025)</td>
<td>(0.021)</td>
<td>(0.031)</td>
<td>(0.038)</td>
<td>(0.026)</td>
</tr>
<tr>
<td>Premarital Cohab × Marriage Cohort</td>
<td>−0.0043+</td>
<td>−0.0035</td>
<td>−0.0027</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.0022)</td>
<td>(0.0022)</td>
<td>(0.0023)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Premarital Cohab × Calendar Year of Marriage</td>
<td>−0.39***</td>
<td>−0.37***</td>
<td>−0.40***</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.090)</td>
<td>(0.092)</td>
<td>(0.090)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Premarital Cohab × First 5 Years of Marriage</td>
<td>−0.067</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
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</table>

<table>
<thead>
<tr>
<th>df</th>
<th>27</th>
<th>28</th>
<th>29</th>
<th>31</th>
<th>28</th>
</tr>
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<tbody>
<tr>
<td>LRT compared to constant only</td>
<td>2,639.12</td>
<td>2,642.90</td>
<td>2,662.09</td>
<td>2,664.58</td>
<td>2,659.61</td>
</tr>
<tr>
<td>Δ LRT compared to previous model</td>
<td>3.78</td>
<td>19.19</td>
<td>2.49</td>
<td></td>
<td></td>
</tr>
<tr>
<td>χ² (P(Δ LRT, 1 df))</td>
<td>0.051</td>
<td>1.18 x 10⁻⁵</td>
<td>0.288</td>
<td></td>
<td></td>
</tr>
<tr>
<td>BIC compared to constant only</td>
<td>−2,394.87</td>
<td>−2,389.60</td>
<td>−2,399.74</td>
<td>−2,384.14</td>
<td>−2,406.31</td>
</tr>
<tr>
<td>Δ BIC from previous model</td>
<td>5.27</td>
<td>−10.14</td>
<td>15.60</td>
<td>−22.17</td>
<td></td>
</tr>
</tbody>
</table>

Source. National Surveys of Family Growth data on first marriages for women aged 44 years and younger, National Surveys of Family Growth 1988 wave and later. Note. Additional controls not reported: marital duration (1 df), year of marriage (1 df), marital duration first calendar year dummy variable (1 df), age at marriage (categorical, 3 df), presence of children younger than age 18 (1 df), calendar decade (5 df), educational attainment (3 df), race (2 df), stable family of origin (1 df), mother’s education (3 df), 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. BIC = Bayesian information criterion; Cohab = cohabitation; LRT = likelihood ratio test. ***p < .001; **p < .01; *p < .05; +p < .10, 2-tailed tests.

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 noncohabiters in the first 5 years of marriage using NSFG 2006–2010.

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 1 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 et al. (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 et al.’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 et al. (1995), 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 four-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 and Meier (1958) survival analysis (see Figures S1 and S2 for the combined NSFG data, the cumulative survival rate as married couples became significantly higher for noncohabiters when compared with 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.

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
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 .05 level (−0.043/0.022 yields a Z score of −1.95, and a p value of .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 α = .05). The LRT comparison of Models 1 and 2 has the same p value of .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 (SE = 0.0022, Z score = −1.58, p = .115). A full set of model coefficients for Models 3 and 5 can be found in Online Supplement, Table S1.

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 five 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 data sets 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 overestimate 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.

**Premarital and Nonmarital Cohabitations**

For the 1995 wave, the NSFG began asking questions about nonmarital cohabitations, that is, 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
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 (Standard Errors in Parentheses)

<table>
<thead>
<tr>
<th>Predictors of marital dissolution</th>
<th>Model 1</th>
<th>Model 2</th>
<th>Model 3</th>
<th>Model 4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Premarital cohabitation</td>
<td>0.35***</td>
<td>0.18***</td>
<td>0.18***</td>
<td>0.13***</td>
</tr>
<tr>
<td></td>
<td>(0.04)</td>
<td>(0.04)</td>
<td>(0.04)</td>
<td>(0.03)</td>
</tr>
<tr>
<td>Prema marital Cohab × Marriage Cohort</td>
<td>0.0052*</td>
<td>0.005+</td>
<td>0.005+</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.0026)</td>
<td>(0.003)</td>
<td>(0.003)</td>
<td></td>
</tr>
<tr>
<td>Prema marital Cohab × Calendar Year of Marriage</td>
<td>−0.36***</td>
<td>−0.37***</td>
<td>−0.36***</td>
<td>−0.39***</td>
</tr>
<tr>
<td></td>
<td>(.10)</td>
<td>(0.10)</td>
<td>(0.10)</td>
<td>(0.10)</td>
</tr>
<tr>
<td>Nonmarital cohab, 1</td>
<td>1.18***</td>
<td>1.18***</td>
<td>1.11***</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.04)</td>
<td>(0.04)</td>
<td>(0.03)</td>
<td></td>
</tr>
<tr>
<td>Nonmarital cohab, 2 or more</td>
<td>1.54***</td>
<td>1.54***</td>
<td>1.46***</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.06)</td>
<td>(0.06)</td>
<td>(0.04)</td>
<td></td>
</tr>
<tr>
<td>Nonmarital Cohab (1) × Marriage Cohort</td>
<td>−0.007*</td>
<td>−0.006*</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.003)</td>
<td>(0.003)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Nonmarital Cohab (2 or More) × Marriage Cohort</td>
<td>−0.008+</td>
<td>−0.008*</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.004)</td>
<td>(0.004)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Nonmarital Cohab (1) × Calendar Year of Marriage</td>
<td>−0.08</td>
<td></td>
<td>0.12</td>
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<tr>
<td></td>
<td></td>
<td></td>
<td>(0.12)</td>
<td></td>
</tr>
<tr>
<td>Nonmarital Cohab (2 or More) × Calendar Year of Marriage</td>
<td>0.07</td>
<td></td>
<td>(0.15)</td>
<td></td>
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<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>df</strong></td>
<td>27</td>
<td>31</td>
<td>33</td>
<td>28</td>
</tr>
<tr>
<td>LRT compared to constant only</td>
<td>2,222.1</td>
<td>3,953.2</td>
<td>3,954.2</td>
<td>3,940.2</td>
</tr>
<tr>
<td>BIC compared to constant only</td>
<td>−1,983.8</td>
<td>−3,679.5</td>
<td>−3,662.9</td>
<td>−3,693.0</td>
</tr>
<tr>
<td>Δ BIC from previous model</td>
<td>−1,695.8</td>
<td>167.6</td>
<td>−30.1</td>
<td>94.2</td>
</tr>
</tbody>
</table>

Source. National Surveys of Family Growth data on first marriages for women aged 44 years and younger. National Surveys of Family Growth 1995 wave and later. Note. Additional controls not reported: marital duration (1 df), year of marriage (1 df), marital duration first year dummy variable (1 df), age at marriage (categorical, 3 df), calendar decade (4 df), educational attainment (3 df), race (2 df), stable family of origin (1 df), mother’s education (3 df), 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. BIC = Bayesian information criterion; Cohab = cohabitation; LRT = likelihood ratio test ***p < .001; **p < .01; *p < .05; +p < .10, 2-tailed tests.

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 (−3,693.0). 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 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, regardless of whether the wife had prior cohabitation partners.

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 when compared with 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. When compared with 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 United States 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 article, 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 the NSFG. Other kinds of data (prospective, qualitative) about couples and their decisions are needed.

NOTE

Roesler, with feedback from Rosenfeld, prepared (with substantial commitment of time) the harmonized National Surveys of Family Growth event history data set used in the article. Rosenfeld, with feedback from Roesler, performed the analyses and wrote the article. Thanks to Soomin Kim, Amy Johnson, and Stanford’s Graduate Family Workshop for comments on previous drafts.

SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of the article.

Table S1. Full set of controls shown for two models from Table 2. Discrete time logistic models predicting marital dissolution.

Table S2. Predicting marital dissolution: Log odds ratio coefficients and summary statistics from discrete time-event history models.

Figure S1. Kaplan and Meier (1958) survival (as intact married couple) estimates based on unweighted data on first
marriages, female respondents aged 15 to 44 years, Waves 1995 and later.

**Figure S2.** First marriage survival by cohabitation history and marriage cohort.

**References**


