REASSESSING THE LINK BETWEEN PREMARITAL COHABITATION AND MARITAL INSTABILITY*

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Premarital cohabitation has been found to be positively correlated with the likelihood of marital dissolution in the United States. To reassess this link, I estimate proportional hazard models of marital dissolution for first marriages by using pooled data from the 1988, 1995, and 2002 surveys of the National Survey of Family Growth (NSFG). These results suggest that the positive relationship between premarital cohabitation and marital instability has weakened for more recent birth and marriage cohorts. Using multiple marital outcomes for a person to account for one source of unobserved heterogeneity, panel models suggest that cohabitation is not selective of individuals with higher risk of marital dissolution and may be a stabilizing factor for higher-order marriages. Further research with more recent data is needed to assess whether these results are statistical artifacts caused by data weaknesses in the NSFG.

Industrial countries have witnessed rising cohabitation rates, while their first-marriage and remarriage rates have declined (Bumpass and Lu 2000; Bumpass, Sweet, and Cherlin 1991; Bumpass and Sweet 1989).[1] Social scientists are interested in cohabitation and marriage because the reasons individuals enter and leave committed relationships have large welfare implications on both the individual and societal levels. At the same time, welfare policies and tax policies may give individuals incentives to enter one form of relationship or the other (Moffitt, Reville, and Winkler 1998).

Cohabitation is a common experience in the United States. In 2002, more than one-half of all women aged 19–44 had ever cohabited. When cohabitation first emerged in the United States, it was mainly a phenomenon of the less-educated and economically disadvantaged, but it has extended to the American middle class. This study investigates the effect of these trends on the relationship between cohabitation and marital instability. Earlier empirical studies found that marriages preceded by premarital cohabitation are less stable in the United States (Booth and Johnson 1988; DeMaris and Rao 1992; Teachman and Polonko 1990) and Western Europe (Bennett, Klimas Blanc, and Bloom 1988).

The idea that couples learn about the match-specific quality during cohabitation goes back at least to Becker, Landes, and Michael (1977). Because cohabitors have a more precise estimate of their match quality, they should experience fewer bad surprises during marriage. Based on this theoretical argument, one would expect former cohabitors to have more stable marriages. However, earlier empirical evidence points in the opposite direction. Self-selection is now an accepted explanation for these counterintuitive results (Lillard, Brien, and Waite 1995; Schoen 1992). Brien, Lillard, and Stern (2006) formalized this

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1. *Cohabitation* here refers to living together under the same usual address and having an intimate sexual relationship.

idea: according to their search model of marriage and cohabitation, couples learn about the quality of their relationship during cohabitation, and some of them decide not to go through with their marriage. Brien et al. showed that couples with a lower initial estimate of their match quality are more likely to cohabit than to get married right away but have higher dissolution rates after they eventually marry.

According to the view of cohabitors as a select group, individuals who are at a higher risk of marital disruption also tend to cohabit before their marriage. In line with this view, cohabitors often have other elevated risk factors for marital disruption, such as lower education, unstable family background (Bumpass and Sweet 1989), and lower commitment to the institution of marriage (Bennett et al. 1988). However, to the extent that premarital cohabitation has become integrated into the regular courtship process, it may have become less signifying of individuals with elevated risk factors (Teachman 2003). If the nature of cohabitation changes, the use of a binary indicator for premarital cohabitation may not be completely adequate; indeed, the use of a binary measure has recently been criticized by sociologists because it may hide some important qualitative differences (Manning and Smock 2005). Nonetheless, the empirical relationship is robust to the exact definition of cohabitation in the particular data set, and the definition of cohabitation used in this study is similar to those used in the earlier empirical literature.

As cohabitation has become more common, there might be less self-selection on unobservable factors in the group of premarital cohabitors. However, then the apparent positive relationship between premarital cohabitation and marital instability may weaken or may become negative, as suggested by the recent experience in Denmark (Svarer 2004) or Australia (De Vaus, Qu, and Weston 2003), where premarital cohabitation does not predict higher rates of marital dissolution. Furthermore, Liefbroer and Dourleijn (2006) studied 16 European countries and found that premarital cohabitation is associated with marital dissolution only in countries with either very high or very low rates of premarital cohabitation. In the United States, Phillips and Sweeney (2005) documented variation in the association between premarital cohabitation and marital instability between racial/ethnic groups. Premarital cohabitation is associated with greater marital instability only for non-Hispanic white women but not for Mexican American women and non-Hispanic blacks, groups in which cohabitation is more common than among non-Hispanic white women. Teachman (2002) studied whether the effects of risk factors for divorce stayed constant in the United States for marriages formed between 1950 and 1984. Because of data limitations, he could use only marriages formed after 1969 when studying premarital cohabitation, and he concluded that the effect of cohabitation had not changed for the more recent marriage cohorts in his data set. Similarly, Kamp Dush, Cohan, and Amato (2003) studied whether the relationship between cohabitation and marital instability changed across U.S. cohorts. They compared the cohorts of couples married between 1964 and 1980 with those married between 1981 and 1997. Although they found that cohabitation was less strongly associated with divorce in the more recent cohort, the change was not significant, perhaps because their sample size was relatively small.

In light of this inconclusive evidence, I use data that are more recent than Teachman's (2002) to extend the time period covered and use a bigger sample than Kamp Dush et al.'s (2003). First, I investigate whether the relationship between cohabitation and marital instability in first marriages has weakened for more recent birth and marriage cohorts using the pooled data for the three most recent cycles of the National Survey of Family Growth (NSFG). In addition, I estimate a model interacting cohabitation status with education. Second, I investigate whether there is self-selection of divorce-prone individuals into premarital cohabitation by using information on higher-order marriages as well.

THEORETICAL CONSIDERATIONS

A Search Model of Marriage and Cohabitation

In the Brien et al. (2006) search model of marriage and cohabitation, couples learn about their mutual compatibility during cohabitation, yet their future marriages are less stable because there is self-selection on marital "quality" into premarital cohabitation.

In their model, single women meet a new potential partner in each period and get a first impression of their degree of compatibility. The woman then decides whether to continue searching for a partner or to enter a relationship, either cohabitation or marriage. While they are in a relationship, women continue learning more about whether they have found a good match. In addition, they also enjoy utility from being in a relationship and from underlying benefits of marriage and cohabitation. Because women learn about their relationship quality, some will realize that the mutual compatibility is not good, and they may decide to dissolve the relationship and be single again in the next period. On the other hand, some cohabiting women will decide to get married when they learn that they have found a good match.

There are separation costs in this model that differ between marriage and cohabitation: psychological costs and monetary costs (for instance, court costs). Brien et al. assumed that the benefits of marriage are higher than the benefits of cohabitation, and that the separation costs for a marriage are also higher. Both assumptions are necessary for the coexistence of cohabitation and marriage in equilibrium. The underlying benefits of marriage and cohabitation determine reservation values, at which individuals become indifferent about being single versus being in a relationship. These reservation values govern the decision to enter or end a relationship, and hence the degree of compatibility of married and cohabiting couples.

Brien et al. showed that the degree of compatibility of cohabitors is lower than that of couples who get married right away, leading to self-selection into premarital cohabitation. However, conditional on the lower degree of compatibility, the effect of premarital cohabitation on marital outcomes is the change in the separation probabilities if the couple cohabits and later marries versus if they immediately marry. This effect should be negative because only cohabiting couples who have learned something positive about their relationship quality during cohabitation get married. Brien et al. also showed that cohabitation would have a negative impact on marital instability if all couples were required to cohabit prior to getting married. In their model, cohabitation serves as a sort of screening device, weeding out matches with less compatibility between the partners. Overall, in the Brien et al. model, the self-selection effect dominates; hence, according to this model marriages preceded by cohabitation should be less stable than marriages that do not follow cohabitation.

Empirical studies that do not control for the unobserved degree of mutual compatibility will produce biased estimates of the causal effect of cohabitation on marital outcomes. The observed association between cohabitation and marital dissolution is the result of the causal effect of cohabitation and the self-selection of women with lower prospects of marital success into premarital cohabitation.

Brien et al. assumed that the draws for the degree of the partners' compatibility are uncorrelated and come from the same distribution for everyone. But an obvious extension of their model is one in which there are unobserved differences in this distribution across persons. In addition, there may be permanent unobserved differences in separation costs between people and other time-invariant factors affecting the stability of a relationship. This would introduce time-invariant, person-specific effects as another possible source of self-selection. This article addresses this problem by using Lillard et al.'s (1995) model but relaxing their strict distributional assumptions. Unfortunately, addressing the problem of match-specific heterogeneity is difficult because no credible instrument governs the decision to cohabit and can be safely excluded from the marital dissolution process.

Trends in Marriage and Divorce Rates and Self-Selection Into Premarital Cohabitation

The empirically observed decline of marriage rates and rise in divorce rates may be explained by declining benefits of marriage or rising benefits of cohabitation. If the benefits of marriage and cohabitation change, though, one would also expect a change in the process of self-selection. For example, Reinhold (2007) demonstrated that the average relationship quality of cohabitors getting married improves in the Brien et al. model if the benefits of marriage decline. When contemplating a marriage, cohabitors weigh the benefits of marriage against the potential costs of a later divorce. If the benefits of marriage decline, cohabitors require a higher relationship quality to get married because then the risk of a divorce—and, hence, the expected costs of divorce—gets smaller. Thus, if the benefits of marriage are low, only couples with a high degree of mutual compatibility get married. This explanation thus relies on a declining benefit to marriage as the key factor in increasing rates of cohabitation.

Abundant theoretical and empirical evidence exists on declining benefits of marriage. Most of these explanations are not mutually exclusive but rather reinforce one another. In Becker's (1973) model of marriage, the incentive to marry stems from the possibility to divide labor and to specialize on activities for which one is more productive than the spouse. One implication is that the gains to marriage are higher when the pay differential between males and females is wider. A decline in the gender pay differential would therefore erode the benefits of marriage, and Moffitt (2000) found evidence consistent with this view. In addition, the welfare system might encourage women not to marry but to cohabit instead (Moffitt et al. 1998).

Changing attitudes and values are another possible explanation for the trends in marital behavior (Cherlin 1992). For instance, Amato and Booth (1995) showed that if wives adopt nontraditional gender roles, their perceived marital quality declines. Cherlin (2004) argued that the social norms governing expectations of behavior in marriage have weakened, adding a potential source of conflict between spouses. Lichter et al. (1992) proposed a "shortage of marriageable men" for some women, particularly for less-educated and African American women. Some factors affecting benefits of marriage, like the gender pay differential or the welfare system, also determine the benefits of cohabitation. However, there is reason to believe that the effect is asymmetric. One good example of symmetric effects is public assistance, as Moffitt et al. (1998) demonstrated. Song (2001) investigated labor supply and fertility patterns in marriage and cohabitation to find that labor supply for women is higher among cohabiting women than among married women. Thus, rising female wages for educated women might have an asymmetric effect on these living arrangements.

Testable Implications

In the theoretical search model discussed, declining benefits of marriage drive both an increase in the rates of premarital cohabitation and a rise in the average relationship quality of cohabitors, leading to reduced separation rates. Based on these theoretical considerations, two hypotheses can be tested. First, as more people cohabit, premarital cohabitation becomes less selective of individuals with high divorce risk; hence, for more recent birth or marriage cohorts, the association between premarital cohabitation and marital instability should weaken. This decline, however, could be attributed to either a change in the causal effect of premarital cohabitation or a change in the process of self-selection. For this reason, I also estimate a model accounting for person-specific heterogeneity to assess whether the process of self-selection differs from that in Lillard et al.'s study.

Second, in groups with high incidence of premarital cohabitation and possibly fewer benefits of marriage, such as women with low educational attainment, premarital cohabitation should be less selective of divorce-prone women. Therefore, premarital cohabitation should be less associated with increased risk of marital dissolution for women with low educational attainment than for women with high educational attainment.

DATA AND DESCRIPTIVE STATISTICS

National Survey of Family Growth

The National Survey of Family Growth (NSFG) was conducted by the National Center of Health Statistics (NCHS) for a representative sample of U.S. women aged 15–44 for the years 1973, 1976, 1988, 1995, and 2002. The NSFG provides information on marriages, divorces, fertility, and the health status of women. The survey includes information on important events, such as marriages and births, along with other socioeconomic and demographic information. The survey asks retrospective questions for the full history of marriages and divorces; in 1988, it also began including more detailed information on women's cohabitation histories.

Because I am interested in the effect of cohabitation on marital outcomes, women who never married are omitted. I analyze first marriages and the cohabitations that preceded them, leaving me with 5,030 first marriages using the NSFG 1988; 6,776 first marriages using the NSFG 1995; and 4,043 first marriages using the NSFG 2002. The pooled data thus consist of 15,849 observations on first marriages. Assuming that 15 years is the earliest age at which one can observe first marriages, the pooled data potentially cover marriages starting between 1959 and 2002. I define *marital dissolution* as the date of separation, as is common in other studies of marital instability. For most respondents, I use the self-reported date of separation (or of divorce, if these dates coincide). In the pooled sample, I interact premarital cohabitation with the year of birth and the year of marriage to study whether the effect of premarital cohabitation is different for more recent birth and marriage cohorts.

Unfortunately, there was a routing error in the survey instrument for the NSFG 2002 (Kennedy and Bumpass 2008; NCHS 2010): some respondents were not asked when their marriages ended, and this skip pattern was not random. For instance, women whose husbands had children from previous relationships were not asked when their marriage dissolved. This skipping pattern could be correlated with premarital cohabitation, thus rendering the estimates biased. In my final data set, 474 of 4,043 respondents in the NSFG 2002 were affected by this problem. For these individuals, dates for marital dissolution were imputed. In addition, the coverage of recent migrants was not constant between the NSFG 1995 and 2002. I address both issues in extensive robustness checks. Pooling all three surveys may mitigate the problem with the survey data of the NSFG 2002 because the problematic observations have less weight in the pooled sample. For my analysis, I construct new weights based on the original survey weights, reflecting the differences in the sample sizes across surveys, and use them for the descriptive statistics and pooled regression results.[2]

Summary Statistics and Survival Functions

Table 1 shows the means of selected variables for women who cohabited before their first marriage and for women who did not.[3] In the pooled sample, a bit more than one-third of first marriages were preceded by premarital cohabitation. Cohabitors have lower educational achievement than noncohabitors, showing the well-known association between socioeconomic background and cohabitation. Furthermore, cohabitors are (on average) younger; this reflects a cohort effect, with more recent cohorts more likely to

2. I also conducted regressions without using the survey weights, which does not qualitatively change the results. These results are available from the author upon request.

3. In the following, I refer to women who cohabited before their marriage as "cohabitors."

Table 1. Means of Variables Relative to First Marriage

Variable	Noncohabitors	Cohabitors
Percentage in Population	63.8	36.2
Education		
Less than high school (%)	14.9	18.9
High school diploma (%)	35.3	31.9
More than high school (%)	49.8	49.2
Year of Birth	1959.5	1962.9
Wife's Age at Marriage	21.3	23.3
Husband's Age at Marriage	24.1	26.5
Age Difference	2.8	3.2
Premarital Conception (%)	28.9	51.9
Premarital Birth (%)	7.7	23.2
Marital Birth (%)	85.3	75.4

Note: Sample weights are used.

Source: Pooled data from the 1988, 1995, and 2002 NSFG.

have cohabited before entering marriage. The recent rise in cohabitation rates have been described elsewhere (Bumpass et al. 1991; Bumpass and Lu 2000; Bumpass and Sweet 1989; Kennedy and Bumpass 2008), and similar results can be found by using the pooled data (Reinhold 2009). At the same time, cohabitors are older at first marriage, partly reflecting the time spent in cohabitation before marriage. Young age has been shown to be a predictor of marital dissolution (Teachman 2002), giving cohabitors a potential advantage. However, at the same time, the age difference between spouses is bigger for cohabitors, which is a potential risk factor for marital dissolution. Finally, there is an important difference between cohabitors and noncohabitors: cohabitors are much more likely to have children outside of marriage and are less likely to have children within the marriage.

Figure 1 displays survivor functions of first marriages for women who cohabited with their future spouse and women who did not cohabit for the pooled sample of the 1988, 1995, and 2002 NSFG.[4] The survivor function shows the proportion of surviving marriages at each duration. The survivor function for noncohabitors lies above the survivor function of cohabitors, showing that the latter marriages are less stable, thus replicating previous empirical results from this data source.

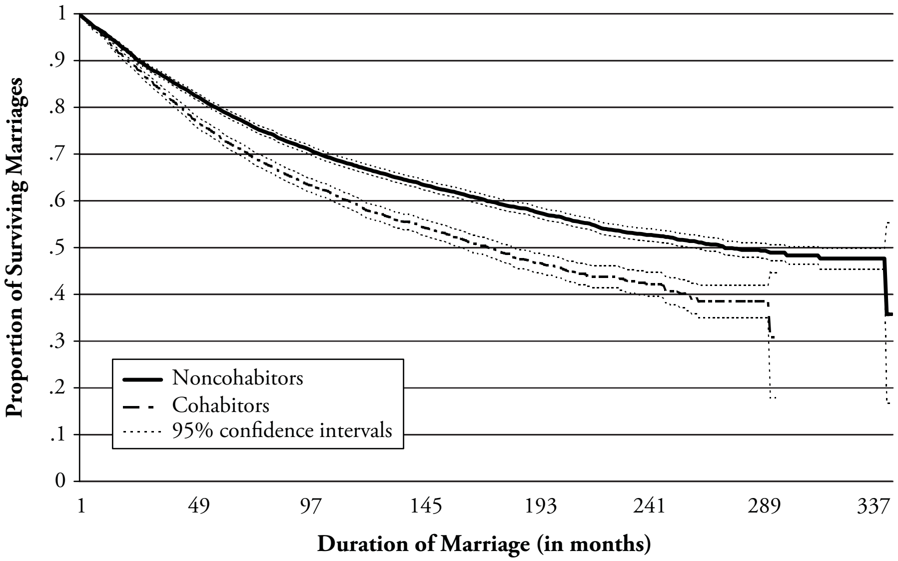
EMPIRICAL MODELS OF MARITAL INSTABILITY

Proportional Hazard Models

I first use proportional hazard regressions (Cox 1972). These models are estimated by pooling all three cycles of the NSFG. Premarital cohabitation is interacted with the year of birth and with the year of the marriage, allowing me to assess whether the association

4. These estimates of survival functions of cohabitors and noncohabitors can be compared with those in a recent NCHS report (NCHS 2010) that shows survival functions for the 2002 NSFG alone. The report suggests that premarital cohabitation is associated with increased marital instability, contrary to my regression results for recent cohorts. However, the NCHS study did not control for covariates; when I do not control for covariates, premarital cohabitation does lead to increased marital instability. In addition, the coefficient in 2002 without covariates is smaller than that in previous cohorts, which is the same pattern I find for the coefficients with covariates.

Figure 1. Survival Functions of First Marriages for Cohabitors and Noncohabitors, and 95% Confidence Intervals: 1988, 1995, and 2002 NSFG



of premarital cohabitation and marital instability has changed for more recent birth and marriage cohorts. Cohabitation is also interacted with education in a different model. In addition, I include other explanatory variables that were found as predictors of marital success in earlier studies: education, race, religion, fertility indicators, family background, migration status, and age at marriage.

The hazard in the simple proportional hazard model can be written as follows:

$$h(t \mid X(t)) = h_0(t)\exp(X(t)'\beta). \tag{1}$$

That is, the hazard at each point in time factors into two components: one that depends only on time ($h_0(t)$), and the other that depends only on the value of the covariates, $\exp(X(t)'\beta)$. The proportional hazard model is semiparametric, and the baseline hazard ($h_0(t)$) does not need to be specified but is estimated nonparametrically. Furthermore, notice that there is no unobserved heterogeneity in this specification. Therefore, the coefficient on cohabitation incorporates both the potential causal effect of cohabitation on marital duration and possibly the self-selection of high-risk individuals into cohabitation.

In Table 2, I present proportional hazard regressions for the pooled data. (See Reinhold [2009] for more detailed results with coefficient estimates for all covariates.) The dependent variable is the hazard of marital dissolution for the first marriage. All coefficients are reported as hazard ratios: a coefficient greater than 1 indicates that this regressor increases the risk of marital dissolution, and a coefficient less than 1 indicates a decrease in risk.

In the first column of Table 2, I show the basic specification without interactions between cohabitation and cohorts or education. According to this estimate, premarital cohabitation increases the risk of marital dissolution by about 30%, which is in line with

Table 2. Proportional Hazard Regressions in Pooled Sample: Dependent Variable = Hazard of Dissolution of First Marriage

Variable	Model 1	Model 2	Model 3	Model 4
Cohabitation	1.293** (0.053)	1.299** (0.052)	1.245** (0.051)	
Cohabitation × Year of Birth		0.982** (0.004)		
Cohabitation × Year of Marriage			0.979** (0.004)	
No High School × Cohabitation				1.158† (0.092)
High School Diploma × Cohabitation				1.414** (0.088)
More Than High School × Cohabitation				1.270** (0.070)
Education (ref. = more than high school)				
No high school	1.008 (0.051)	1.012 (0.052)	1.018 (0.052)	1.046 (0.065)
High school diploma	1.021 (0.042)	1.025 (0.043)	1.032 (0.043)	0.985 (0.050)
Year of Birth	1.008** (0.003)	1.014** (0.003)		1.008** (0.003)
Year of Marriage			1.015** (0.003)	

Notes: N = 15,849. Sample weights, adjusted for different sample sizes, are used. Estimates are reported as hazard ratios. A coefficient greater than 1 indicates an increase in the hazard of marital dissolution; a coefficient of less than 1 indicates a decrease in the hazard. Additional control variables are indicators for premarital conception, premarital childbearing, marital childbirth, religious affiliation, race, family background, migration status, both partners' age at marriage, and dummy variables for the different surveys of the NSFG. Standard errors are in parentheses.

$^{\dagger}p < .10$; $^{**}p < .01$

previous studies for the United States.[5] Thus, despite some potential weaknesses in the data, they can be used to reproduce previous empirical research.

The second column of Table 2 shows the specification in which premarital cohabitation is interacted with the year of birth. The coefficient on this interaction is less than 1 and highly significant: for each successive birth cohort, the risk of marital dissolution decreases by about 1.8% for cohabitors. Based on this specification, cohabitation is associated with an increase in the risk of marital dissolution for women born before approximately 1975; for younger women, it is associated with a decrease in the risk of marital dissolution.

A similar picture emerges for the interaction of premarital cohabitation with marriage cohorts. For cohabitors in later marriage cohorts, the risk of marital dissolution is reduced by about 2% per year. This indicates that for marriages contracted before approximately 1993, premarital cohabitation is associated with an increased risk of marital dissolution; for later marriage cohorts, cohabitation is associated with a decreased risk. Both specifications interacting cohabitation with cohorts show that for more recent cohorts, cohabitation is not associated with higher rates of marital dissolution, and these cohort effects are statistically significant.

5. Teachman (2002), for example, reported an increase in the risk of marital dissolution by around 35%.

The fourth column of Table 2 shows the results for the specification with interactions between premarital cohabitation and education. Cohabitation is a risk factor for marital break-up only for women with high school education or better. For women without a high school diploma, premarital cohabitation is not associated with a strong increase in the risk of marital dissolution. Testing the equality of the three coefficients on the interactions of cohabitation with education, I can reject the null hypothesis at the 10% level.

Robustness Checks for the Pooled Data

Because of weaknesses in the 2002 NSFG, Kennedy and Bumpass (2008) cautioned that the data may not be reliable for analyzing marital dissolution. For this reason, I conduct extensive robustness checks. The two main problems with the data are (1) a routing error in the survey instrument for the 2002 NSFG, resulting in a nonrandom skip pattern for marital dissolution dates, and (2) changes in the inclusion of recent migrants in the surveys. I conduct three robustness checks on the pooled data: I estimate models of marital dissolution similar to the ones discussed earlier by using (1) a subsample excluding observations with imputed dates, (2) a subsample excluding migrants in the 2002 NSFG, and (3) the pooled data without the 2002 NSFG. All three robustness checks lead to very similar qualitative conclusions as the models discussed in the previous section, even when all 2002 observations are discarded.[6] See Reinhold (2009) for detailed results on all robustness checks.

Accounting for Unobserved Heterogeneity

The proportional hazard models in the previous sections do not allow for conclusions about the causal effect of premarital cohabitation on marital stability because the coefficient on premarital cohabitation is likely to be tainted by self-selection. Changes in the association between premarital cohabitation and marital instability may reflect a change in the selection process into premarital cohabitation or a change in the causal effect of premarital cohabitation. To address at least one source of heterogeneity, I employ a panel model, using the first three marriages to account for unobserved person-specific effects (see Lillard et al. 1995). These panel models can be estimated only with the 2002 NSFG 2002 because detailed information on cohabitation histories are not available for all higher-order marriages for early NSFG cycles.

As discussed earlier, there are two sources of unobserved heterogeneity. The first source comprises time-varying, person-specific factors, including the unobserved quality of the match in the Brien et al. model. Instrumental variables are one way to deal with this sort of endogeneity. To be valid, instrumental variables must be correlated with the endogenous regressor cohabitation and must not be correlated with the error term in the main regression. It is very difficult to find a variable for cohabitation that would satisfy these conditions, since cohabitation and marriage are similar interdependent decision problems. (Recall that in the Brien et al. model, all variables that affect the probability of marital dissolution are also likely to affect the probability of premarital cohabitation.) To my knowledge, no

6. The results for the 1988 and 1995 NSFG can also be compared with those of Teachman (2002). However, the focus of his study is not premarital cohabitation but rather other risk factors; he covers a much different time period, during which not much information on premarital cohabitation was available in the NSFG. In a basic model for the 1988 and 1995 NSFG, he found that cohabitation increases marital instability by about 35%, which is the same as my result. In addition, he tested a specification including an interaction between marriage cohorts and premarital cohabitation. Although he found evidence supporting a trend, he could not reject the null hypothesis of no change across cohorts. For this test, however, he used the Bayesian information criterion (BIC) for model selection, which requires a much stronger t value. The value of the difference in BIC between a baseline specification and a specification including the interaction term between marriage cohorts and premarital cohabitation indicates that this coefficient would be statistically significant on conventional levels. Unfortunately, he did not report the coefficient on the interaction between marriage cohorts and premarital cohabitation or even the sign of this coefficient, nor did he test for interactions between age cohorts and premarital stability.

previous study has attempted to implement an instrumental-variable estimator in the context of cohabitation and marriage.

The second form of heterogeneity includes time-invariant, person-specific effects: for instance, unobserved and permanent differences in separation costs. Panel estimators are one way to deal with this correlation of multiple outcomes for one person. Such estimators use data on multiple marriages and essentially difference across marriages, correlating differences in marital dissolution with differences in premarital cohabitation. Lillard et al. (1995) employed such techniques to model the decision to cohabit jointly with the marriage dissolution process. For identification, they relied on the presence of multiple marriage outcomes for some women. With a random-effects assumption, they could identify the correlation between unobserved person-specific characteristics in the cohabitation and the marital dissolution process.

The Lillard et al. (1995) model. Lillard et al. modeled the decision to cohabit before marriage and the marital dissolution process simultaneously. There is an unobserved heterogeneity term in both processes that may be correlated. Conditional on all other covariates and the person-specific components, cohabitation is independent of idiosyncratic match-specific quality.

Because the heterogeneity term is assumed to be permanent for a person, the correlation can be identified by using multiple marriage outcomes for a person. A positive correlation between the heterogeneity terms indicates self-selection of individuals with a high risk of marital disruption into premarital cohabitation. Lillard et al. interpreted the coefficient on premarital cohabitation as the causal effect of cohabitation on marital stability because it is purged from any self-selection if the model is correct.

The marginal likelihood contribution of all marriages of a given woman is given by

$$\int_{\delta}\int_{\varepsilon}\frac{1}{\sigma_\delta\sigma_\varepsilon}\varphi\left(\frac{\delta}{\sigma_\delta},\frac{\varepsilon}{\sigma_\delta}\middle|\rho_{\delta\varepsilon}\right)\prod_{m=1}^{M}\left[S_m\left(X_m^d,Coh_m,t,\delta\right)h_m\left(X_m^d,Coh_m,t,\delta\right)^{D_m}\right.$$
$$\left.\times\,\Phi\left((2Coh_m-1)\left(\beta_0+\beta_1' X_m^c+\varepsilon\right)\right)\right]d\delta d\varepsilon. \tag{2}$$

The expression $S_m\left(X_m^d,Coh_m,t,\delta\right)h_m\left(X_m^d,Coh_m,t,\delta\right)^{D_m}$ is just a parametric survival model for the *m*th marriage of a woman, where $D^m = 1$ indicates a completed marriage spell and $D^m = 0$ indicates a censored spell. $\Phi\left((2Coh_m-1)\left(\beta_0+\beta_1' X_m^c+\varepsilon\right)\right)$ is a probit model for premarital cohabitation prior to the *m*th marriage. One takes the product over the first three marriages of a woman and integrates out the unobserved heterogeneity components. Lillard et al. assumed that the heterogeneity components δ and ε are drawn from a bivariate normal distribution. That is,

$$\begin{pmatrix}\delta\\ \varepsilon\end{pmatrix}\sim N\left(\begin{pmatrix}0\\0\end{pmatrix},\begin{pmatrix}\sigma_\delta^2 & \sigma_{\delta\varepsilon}\\ \sigma_{\delta\varepsilon} & \sigma_\varepsilon^2\end{pmatrix}\right). \tag{3}$$

Conditional on the unobserved heterogeneity and the covariates, there is no correlation in outcomes across marriages for a given woman. The unobserved heterogeneity and the correlation between the heterogeneity component is identified even without exclusion restrictions because some women have more than one marriage. See Lillard et al. (1995) for more details on the derivation of the marginal likelihood function.

One shortfall of this approach is that different dynamics might be at play in higher-order marriages compared with first marriages and that these may be correlated with the decision to cohabit. Teachman (2008), for example, found that premarital cohabitation in the second marriage does not raise the risk of marital dissolution. For this reason, I also use specifications that include interactions between premarital cohabitation and higher-order marriages, the main coefficient of interest. Teachman also discussed other risk factors that may have a different influence in second marriages compared with first marriages. For sake of simplicity and to allow for better comparability with previous results, I do not model those additional interactions. Lillard et al. also assumed a bivariate normal distribution for the unobserved effects. In this study, I estimate the model but relax their distributional assumption and replace it with a finite discrete distribution (see also Svarer [2005], who estimated a similar model with Danish data). The main difference between our distributional assumptions is that in mine, no symmetry of the distribution around its mean is imposed around its mean, and the finite discrete distribution can assign more mass to the tails of the distribution.

The random effects in the hazard and probit equation are given by the following equations:

$$e_h = v_0 \tag{4}$$

$$e_p = \rho_p v_0 + v_1. \tag{5}$$

I model the correlation between the unobserved random effects across the two processes through the values of ρ_p. The components v_0 and v_1 are independent, and each follows a two-point distribution. That is

$$v_i = \begin{cases} m_{i,1} \text{ with probability } w_{i,1} \\ m_{i,2} \text{ with probability } w_{i,2} \end{cases}. \tag{6}$$

I impose no restrictions on the support points and the weights of these distributions. Because, in general, the expectation of these random effects is nonzero, there is no constant term in either process.

Results of random-effects model. The 2002 NSFG contains complete information on 4,021 individuals. A little less than one-half of first marriages are preceded by premarital cohabitation. The prevalence of cohabitation with the future spouse is higher for second and third marriages. The data contain 689 observations on second marriages, around two-thirds of which are preceded by cohabitation. Furthermore, the data contain 108 observations on third marriages, around three-fourths of which are preceded by cohabitation.

Table 3 presents the results using a finite discrete mixture. (See Reinhold [2009] for more detailed results for other covariates and for a model using bivariate normality.) The first two columns of Table 3 display results for a model without unobserved heterogeneity; the first column displays results from a model in which the effect of cohabitation is restricted to be the same across all marriages, while the second column shows results of a model in which the effect is allowed to be different for higher-order marriages. When interactions between premarital cohabitation and parity of marriage are not included (column 1), these results imply that cohabitation is associated with higher marital stability. However, this result is entirely driven by the coefficient on higher-order marriages. In first marriages, cohabitation is associated with a slightly increased risk of marital dissolution (column 2).

By introducing random effects, I find that in each process, one could approximate the unobserved heterogeneity with a two-point discrete distribution. Introducing more support points does not significantly improve the fit of the model. In columns 3 and 4 of Table 3, I restrict the correlation between the heterogeneity components to zero. The coefficients on premarital cohabitation and its interactions with parity of marriage are hardly affected by this form of heterogeneity.

In the last two columns of Table 3, I present the results allowing correlation between the unobserved heterogeneity components. In the model without interactions between cohabitation and higher-order marriages (column 5), the correlation between the unobserved heterogeneity components is statistically significant, indicating that cohabitors have unobserved characteristics that make them more likely to also have less-stable marriages. This result is in line with Lillard et al.'s finding of this form of self-selection. The coefficient on premarital cohabitation becomes smaller than 1 and is statistically significant when

Table 3. Maximum Likelihood Estimation of Duration Model With and Without Heterogeneity (finite discrete mixture): Dependent Variable = Hazard of Dissolution of First Three Marriages

Variable	No Heterogeneity		Finite Discrete Mixture, ρ = 0		Finite Discrete Mixture, No Restriction on ρ	
Premarital Cohabitation	0.871* (0.048)	1.014 (0.059)	0.824** (0.050)	1.009 (0.072)	0.639** (0.059)	0.902 (0.099)
Premarital Cohabitation × Second Marriage		0.480** (0.062)		0.406** (0.060)		0.426** (0.064)
Premarital Cohabitation × Third Marriage		0.273** (0.077)		0.175** (0.049)		0.190** (0.054)
More Than One Marriage	1.874** (0.150)	2.941** (0.284)	1.534** (0.157)	2.537** (0.340)	1.605** (0.167)	2.563** (0.348)
ρ					0.375** (0.127)	0.127 (0.102)
Point 1			–3.687** (0.399)	–3.782** (0.392)	–3.56** (0.399)	–3.743** (0.396)
Point 2			–1.776** (0.349)	–1.760** (0.351)	–1.540** (0.349)	–1.691** (0.354)
Weight 1[a]			–0.380** (0.148)	–0.338** (0.125)	–0.516** (0.131)	–0.409** (0.128)
Log-Likelihood	–13,593.09	–13,574.70	–13,537.35	–13,511.39	–13,532.95	–13,510.92

Notes: N = 4,021. Sample weights are used. Estimates are reported as hazard ratios. A coefficient greater than 1 indicates an increase in the hazard of marital dissolution; a coefficient less than 1 indicates a decrease in the hazard. Standard errors based on numerical standard errors are in parentheses. Additional control variables are marital duration, duration since first marital birth, marital birth, premarital conception, premarital birth, education, race, religion, family background, migration status, both partners' age at marriage, year of birth, indicators for children from previous relationships, a dummy variable for higher-parity marriages, and divorced husband.

[a]Reported as $\Phi^{-1}(w1)$ of weight.

*$p < .05$; **$p < .01$

restricting the effect of cohabitation to be the same across all marriages, indicating that cohabitation stabilizes marriages. In the preferred estimate, including the interactions between cohabitation and higher-order marriages (column 6 of Table 3), the correlation between the unobserved heterogeneity components becomes smaller and insignificant. At the same time, premarital cohabitation has no effect on marital stability in first marriages.

These results are different in some respects to those of Lillard et al. They found a strong and positive correlation between these heterogeneity components and no causal effect of premarital cohabitation on marital instability when using the National Longitudinal Study of the High School Class of 1972 with its follow-up in 1986. Thus, the association between premarital cohabitation and marital instability is caused by self-selection and not by some true effect of cohabitation on marital stability. Similarly, the current study also finds a positive correlation between the heterogeneity components, indicating self-selection of more divorce-prone women into premarital cohabitation. However, when premarital cohabitation are interacted with indicators for higher-order marriages, the correlation becomes insignificant. In contrast to Lillard et al., I find some indications for a stabilizing effect of premarital cohabitation, particularly for higher-order marriages.

Robustness checks for the 2002 NSFG. The results of the previous section rely solely on the 2002 NSFG with its potential weaknesses. To assess whether the data weaknesses can lead to spurious conclusions, I investigate which covariates are correlated with having an imputed value for the date of marital dissolution. I do find a positive correlation between premarital cohabitation and the probability of having an imputed value. Thus, weaknesses in the data may potentially bias these results.

In addition, I investigated whether the results from using the 2002 NSFG separately are different from the 1995 NSFG for the same birth cohorts.[7] I estimate a model for the age group 15–37 years for the 1995 NSFG and a model for the age group 22–44 years for the 2002 NSFG, covering the same birth cohorts for both surveys. I artificially censor the 2002 NSFG in 1995. Because the two surveys are independent, I can conduct a simple t test of whether the coefficients on premarital cohabitation are the same. The point estimates are clearly different (1.250 vs. 1.046). However, because of the size of the standard errors, I cannot reject the null hypothesis that they are equal (p value = .257). There are some differences in the coefficients on other covariates (e.g., religion). (See Reinhold [2009] for more details on these results.) Overall, the results when using only the 2002 NSFG should be considered even more cautiously than the results using the pooled data, in which the tainted observations have much less influence.

CONCLUSION

In this article, I reassessed the question of the influence of premarital cohabitation on marital instability. A theoretical search model of marriage and cohabitation suggests that cohabitation should help couples learn about their match quality, decreasing their dissolution rates. On the other hand, there may be self-selection in the sense that the average match quality of couples who transform their cohabitation into a marriage is lower than for couples who marry without prior cohabitation. Self-selection of high-risk individuals could explain the established empirical evidence for the United States and other industrialized countries that shows that marriages preceded by cohabitation are less stable.

This article lends some support to the thesis that the once-strong association between premarital cohabitation and marital instability has weakened over time, and there no longer seems to be an association for the more recent birth and marriage cohorts. Given the rise in premarital cohabitation, changes in the process of self-selection could explain these findings. As cohabitation has become more common, it has ceased to be selective of individuals with high risk of marital break-up. The results for different educational groups support this view, showing an increase in risk for well-educated cohabitors only. One explanation for this finding is that premarital cohabitation is not selective of divorce-prone individuals in this educational group. For less-educated women, cohabitation has always been more common than for other socioeconomic groups. For this reason, self-selection may not have been as severe within this educational group even for earlier cohorts in which there was a strong overall association between premarital cohabitation and marital instability. On the other hand, premarital cohabitation was relatively uncommon for well-educated women in earlier years, suggesting that this small group of well-educated cohabiting women was perhaps more selective of divorce-prone individuals. One should be cautious in interpreting the results because of the weaknesses in the most recent cycle of the NSFG. However, because these results rely on the pooled data over the three cycles of the 1988, 1995, and 2002 NSFG, the impact of the problematic data should be somewhat mitigated because they have relatively little influence. In fact, the same qualitative picture emerges even when one only considers the 1988 and 1995 cycles. Nonetheless, it would be important to confirm these findings by using untainted data.

Using the Lillard et al. model, I find mixed evidence for the hypothesis that the process of self-selection has changed. There seems to be little self-selection based on unobservable person-specific characteristics and no effect of cohabitation in first marriages. In higher-order marriages, on the other hand, cohabitation is important in stabilizing marriages.

7. I thank a reviewer for this suggestion.

Therefore, the earlier finding of a positive self-selection into premarital cohabitation may be spurious and may be driven by the particular assumption and by different dynamics of premarital cohabitation in first and later marriages. Because these results are based solely on the problematic 2002 data, even more caution is warranted in interpreting the results. Only fresh data will show whether this empirical finding is robust.

Although my results are perhaps new and surprising for the United States, they are in line with more recent evidence from Denmark (Svarer 2004) and other European countries (Liefbroer and Dourleijn 2006). When about one-half of the population cohabits, cohabitation ceases to be selective of divorce-prone individuals in these European countries. Incidentally, the rates of premarital cohabitation in the United States have just reached this level.

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