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REASSESSING THE LINK BETWEEN PREMARITAL COHABITATION AND MARITAL INSTABILITY

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Abstract

It has been found that premarital cohabitation is positively correlated with the likelihood of marital dissolution in the U.S. To reassess this link, I estimate proportional hazard models of marital dissolution for first marriages using pooled data from three surveys of the NSFG 1988, 1995, and 2002 and demonstrate 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 parity marriages.

Keywords: Cohabitation, Marriage, Marital Stability, Self-selection, Duration Models

JEL Classification: C410, D830, J120

1 Introduction

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

At the present, cohabitation is a common experience in the United States. In 2002, more than half of all women aged 19-44 have ever cohabited in their lives. When cohabitation first emerged in the USA, it was mainly a phenomenon of the less educated and economically disadvantaged, but by now 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 have found that marriages preceded by premarital cohabitation are less stable both for the United States (Booth and Johnson 1988; Teachman and Polonko 1990; DeMaris and Rao 1992) and Western Europe (Bennett et al. 1988). I show that this empirical regularity has broken down over the last twenty years, and I hypothesize that a change in the process of self-selection into cohabitation drives this result.

The idea that couples learn about the match-specific quality during cohabitation goes back at least to Becker (1973) and Becker et al. (1977). Since cohabitors have a more precise estimate of their match quality, there should be fewer bad surprises during marriage. Based on this theoretical argument, one expects that former cohabitors lead more stable marriages. However, earlier empirical evidence points in the opposite direction. Self-selection is now an accepted explanation for these counterintuitive results (Schoen 1992; Lillard et al. 1995). Brien et al. (2006) formalize this idea. In their theoretical search model of marriage and

[1] Cohabitation is understood here as living together under the same usual address and having an intimate sexual relationship.

cohabitation, couples learn about the quality of their relationship during cohabitation, and some of them decide not to go through with their marriage. They show that couples with a lower initial estimate of their match quality are more likely to cohabit than to get married right away.

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. Supportive of this view is the fact that cohabitors often have other elevated risk factors for marital disruption as for example 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 in the regular courtship process, it may have become less signifying of individuals with elevated risk factors (Teachman 2003). As cohabitation has become more common there might be less self-selection on unobservables in the group of premarital cohabitors. But then the apparent positive relationship between premarital cohabitation and marital instability may weaken or even reverse its sign as the recent experience in Denmark (Svarer 2004) or Australia (De Vaus et al. 2003) suggests. Furthermore, Liefbroer and Dourleijn (2006) study 16 European countries and find 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) document that there is variation in the association between premarital cohabitation and marital instability between 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, a group in which cohabitation is more common than among White women. For the US, Teachman (2002) studies whether the effects of risk factors for divorce stayed constant between 1950 and 1984. Because of data limitations he can only use a more restricted time period when studying premarital cohabitation, and he concludes that the effect of cohabitation has not changed for more recent marriage cohorts. Similarly, Kamp Dush et al. (2003) study whether the relationship between cohabitation and marital instability changed across cohorts. They compare the cohorts of couples married between 1964 and 1980 with those married between 1981 and 1997. While they find that cohabitation is less strongly associated with divorce in the more recent cohort, the change is not significant, perhaps because the sample size was relatively small.

In light of this inconclusive evidence, I investigate more recent data than Teachman to extend the time period covered and a bigger sample than Kamp Dush et al. I pool the three most recent cycles of the National Survey of Family Growth (NSFG) in 1988, 1995, and 2002 to study the evolution of the relationship between cohabitation and subsequent marital instability. In this pooled sample, I interact the dummy for premarital cohabitation with marriage cohorts and birth cohorts. The interaction terms can be analyzed to assess whether for more recent cohorts the association between premarital cohabitation and marital instability has changed. In addition, I estimate proportional hazard models with interactions between education and premarital cohabitation to see whether the coefficient on cohabitation is different across educational groups. Premarital cohabitation is more common among the less educated and one could conjecture that within this group premarital cohabitation is less selective of divorce prone individuals. Thus, assessing the relationship of premarital cohabitation and marital stability within educational groups may shed light on the question whether differences in self-selection into premarital cohabitation translate into differences in the association between premarital cohabitation and marital instability.

The estimates from these proportional hazard models do not uncover the causal effect of cohabitation on marital instability because of self-selection into cohabitation and consequent endogeneity of the cohabitation variable. The conventional approach to estimating causal effects in this type of model would be to seek a valid and relevant instrument which is correlated with cohabitation but uncorrelated with the unobservables affecting marital stability. Unfortunately, instruments in this case are difficult to find because cohabitation and marital instability are points of an intrinsically joint process and variables that affect the latter will generally always affect the former. Nor do the data used here contain any information on credible instruments. Instead, I adopt the Lillard et al. (1995) model to address the narrower problem of self-selection on the basis of individual-specific, time-invariant unobservables in a random effects model using data on multiple marriages. This approach does not account for selection on time-varying unobservables such as the match-specific draws in the Brien et al. (2006) model, for example, but accounts for persistent differences across individuals in those draws or in other variables affecting cohabitation and marital instability. The Lillard et al. model also used a restrictive bivariate normal assumption for the error terms in the model, but this is relaxed by using a semi-parametric specification for their distribution.

2 Theoretical Considerations

In the Brien et al. (BLS) search model of marriage and cohabitation couples learn about their mutual compatibility during cohabitation, yet at the same time 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 receive a noisy signal of the unknown match quality. After receiving the signal the woman decides whether to continue searching for a partner or to enter a relationship, either cohabitation or marriage. In each period they are in a relationship, women derive flow utility from additional signals of their match quality which they also use to update their information. In addition, to these relationship specific signals they also enjoy utility from underlying benefits of marriage and cohabitation. Since women learn about their relationship quality in the model there may be 'bad' surprises and they may decide to dissolve the relationship and be single again in the next period. After 'positive' surprises about their relationship cohabiting women have the additional option to marry their current cohabiting partner.

There are separation costs in this model which differ between marriage and cohabitation. BLS assume that the benefits of marriage are higher than the benefits of cohabitation, and that the separation costs for a marriage are also higher. Both these assumptions are necessary for the coexistence of cohabitation and marriage in equilibrium. The underlying benefits of marriage and cohabitation determine reservation values governing the decision to enter or end a relationship. BLS show that the match quality of cohabitors is lower than of couples who get married right away leading to self-selection into premarital cohabitation. However, conditional on the unobserved match quality, the effect of premarital cohabitation on marital outcomes is the change in the separation probabilities if the person cohabits and marries after a while versus if she immediately marries. This effect should be negative because only cohabiting couples who have experienced positive surprises during cohabitation get married. BLS also show 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 bad matches. Overall, in the BLS model the self-selection effect dominates, and hence one would observe that marriages preceded by cohabitation are less stable. Empirical studies that do not control for the unobserved match quality effects will deliver 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.

BLS assume that the draws for match-quality are uncorrelated and come from the same distribution for everyone. But obviously, one could think of an extension of their model where there are unobserved differences in the distribution of match quality 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 paper addresses this problem by using Lillard et al.'s (1995) model but relaxing their strict distributional assumptions. Unfortunately, it is difficult to address the problem of match-specific heterogeneity because there is no credible instrument which governs the decision to cohabit but can be safely excluded from the marital dissolution process.

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. But if the benefits of marriage and cohabitation change one would also expect a change in the process of self-selection. For example, Reinhold (2007) demonstrates that the average match quality of cohabitors getting married improves in the BLS model if the benefits of marriage decline. Cohabitors trade off the benefits of marriage against the potential costs of a divorce. If the benefits of marriage decline then cohabitors require a higher match quality to get married because then the risk of a divorce, and hence the expected costs of divorce , gets smaller. This explanation thus relies on a declining benefit to marriage as the key factor in increasing rates of cohabitation.

There has been abundant theoretical and empirical evidence on declining benefits of marriage. Most of these explanations are not mutually exclusive but rather reinforce each other. In Becker's (1973, 1981) model of marriage, the incentive to marry stems from the possibility to divide labor and to specialize on activities where one is more productive than the spouse. One implication is that the gains to marriage are higher in a situation where the pay differential between males and females is higher. A decline in the gender pay differential would therefore erode the benefits of marriage, and Moffitt (2000) finds evidence consistent with this view. In addition, the welfare system might encourage women not to marry and to cohabit instead (Moffitt et al. 1998). Changing attitudes and values are another possible explanation for the trends in marital behavior (Cherlin 1991). For instance, Amato and Booth (1995) have shown that if wives adopt non-traditional gender roles their perceived marital quality declines. Cherlin (2004) argues that the social norms governing expectations of behavior in marriage have weakened adding a potential source of conflict between spouses. Lichter et al. (1992) proposes a 'shortage of marriageable men' for some women, particularly for less educated and African-American women. Some of the factors affecting benefits of marriage will also determine the benefits of cohabitation. However, there is reason to believe that the effect is asymmetric. One good example is public assistance as Moffitt et al. (1998) demonstrate. Song (2001) investigates labor supply and fertility patterns in marriage and cohabitation. She found 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.

In the theoretical search model discussed declining benefits of marriage are both driving an increase in the rates of premarital cohabitation and a rise in the average match quality of cohabitors. Based on these theoretical considerations there are two hypotheses that can be tested. First, as more people start cohabiting premarital cohabitation becomes less selective of individuals with high divorce risk. Hence, one should find that for more recent cohorts the association between premarital cohabitation and marital instability has weakened. This decline could in principle be attributed to either a change in the causal effect of premarital cohabitation or a change in the process of self-selection. Given the recent dramatic changes in cohabitation and marriage behavior it is reasonable to think that changes in the process of self-selection into premarital cohabitation dominate any changes in the true causal effect and would drive my empirical results. Furthermore, one should expect no selection on unobservable characteristics into premarital cohabitation for more recent cohorts. This latter hypothesis can be tested using a model accounting for unobserved heterogeneity (Lillard et al. 1995). Second, in groups with high incidence of premarital cohabitation and possible lower benefits of marriage such as women with low educational attainment, one expects that premarital cohabitation is less selective of divorce prone women. Hence, if one interacts premarital cohabitation with educational attainment one expects that premarital cohabitation is not associated with increased risk of marital dissolution for women with low educational attainment.

3 Data and Descriptive Statistics

The National Survey of Family Growth (NSFG) was conducted by the National Center of Health Statistics (NCHS) for a representative sample of women aged 15-44 for the years 1973, 1976, 1988, 1995, and 2002. Its main purpose is to provide information on marriages, divorces, fertility, and the health status of women and their children. The survey includes information on important events such as marriages and child-births along with other socioeconomic and demographic information. The survey asks retrospective questions for the full history of marriages and divorces; but only starting in 1988 it also included more detailed information on women's cohabitation history.

Since I am interested in the effect of cohabitation on marital outcomes, women who never married are omitted. I analyze first marriages and cohabitation that preceded them. This left me with 5030 first marriages using the NSFG 1988, 6776 first marriages using the NSFG 1995, and 4043 first marriages using the NSFG 2002. After pooling the data set consists of 15849 observations on first marriages. Assuming 15 years is the earliest age one can observe first marriages then the pooled data potentially covers 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 divorce if these dates coincide). In the pooled sample, I interact premarital cohabitation with the age of the respondent[2] and the year of marriage and 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 and some respondents were not asked when their marriages have 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 rendering the estimates biased. In my final data set 474 out of 4043 respondents in the NSFG 2002 are affected by this problem. For these individuals dates for marital dissolution were imputed. In addition, the coverage of recent migrants has not stayed 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.[3]

Figures 1-3 display survivor functions separately for the NSFG 1988, 1995, and 2002 of

[2] In this pooled data set this is the age of the respondent in 2002.

[3] I also conduct regressions without using the survey weights which does not qualitatively change the results. These results are available from the author upon request.

first marriages for women who have cohabited with their future spouse and women who have not. The survivor function shows the proportion of surviving marriages at each duration. In all three graphs, the survivor function for non-cohabitors lies above the survivor function of cohabitors showing that the latter marriages are less stable. However, these differences have become smaller for the initial years of the marriage indicating that the differences in dissolution behavior across these two groups have become more similar. The 95% confidence intervals for the survivor functions of the two groups overlap during the first 45 months of the first marriages in the NSFG 2002. Statistically, there are hardly any differences in the dissolution behavior of first marriages during the first four years of a marriage between the two groups in the NSFG 2002. Over the three cycles of the NSFG, marriages have become less stable for both groups of women reflecting a general upward trend in marital instability for all groups.

The main variable of interest is a dummy for premarital cohabitation where women were asked whether they have cohabited with their future spouse before they got married. The use of a binary indicator for premarital cohabitation may not be completely adequate and has recently been criticized by sociologists because it may hide some important qualitative differences (Manning and Smock 2005). For example, a formal engagement before cohabitation with a clear understanding that a marriage is planned might change the expectations and behavior of the couple during this phase. In the NSFG 2002 women were asked whether they were engaged while cohabiting. I find that engagement and cohabitation combined increases marital stability in the NSFG 2002. I do not use this measure further because this particular question was not asked in earlier cycles of the NSFG making it impossible to study this effect over time. Similarly, one might worry that the average length of premarital cohabitation is important in determining its effect on marital stability. For this reason, I examined whether a cohabitation shorter than three months has a different effect than longer cohabitation, but I did not find differences in the coefficients on these two measures of cohabitation. Also, I do not find that cohort effects were important in determining whether marriage was preceded by a long or short cohabitation. My justification to use the binary measure is that the positive effect of cohabitation on marital instability has been found in many different datasets in which cohabitation was measured differently. Thus, the empirical relationship is robust to the exact definition of cohabitation in the particular data set, and my definition of cohabitation is similar to those used in the literature.

Because the rise in cohabitation over time and its composition is critical to the explanations provided in this paper, it is useful to begin with a brief descriptive analysis of these trends. There is now ample evidence that cohabitation rates have risen in the past (Bumpass and Sweet 1989; Bumpass et al. 1991; Bumpass and Lu 2000; Kennedy and Bumpass 2008). Cohabitation has by now become a common experience among women in the United States. Besides concentrating on premarital cohabitation, there are at least two ways to measure this rise in cohabitation rates. Bumpass and Lu (2000) use both the National Survey of Families and Households (NSFH 1987/1988) and the NSFG 1995 to calculate percentages of women ever cohabiting and currently cohabiting. I combined these results with newer results from Kennedy and Bumpass (2008) in tables 1 and 2.

In table 1, cohabitation rates by age group are shown. The first three columns show the percentage of ever cohabiting women by age group in 1987, 1995, and 2002. Overall the percentage of ever cohabiting women has risen steadily from a third in 1987 to well over half of all women in 2002. This rise in cohabitation rates was most marked for women aged 35-44. While in this group prevalence of cohabitation was below average in 1987, it is now well above average. Furthermore, while cohabitation was more common among younger women in earlier years, older women are now very likely to experience cohabitation. The last three columns of table 1 show rates of currently cohabiting women (of not currently married) for the same years by age group. While in 1987 younger women aged 25-29 were the most likely to cohabit, the rate of currently cohabiting women rose quicker for women aged 35-39 while the growth was more modest for other age groups. In empirical studies, a common finding is that a young age increases the risk of union disruption. The rise of cohabitation among older women would therefore be one additional factor stabilizing the relationships of cohabitors.

Table 2 shows percentages of ever cohabiting women by education and race for 1987, 1995, and 2002. While cohabitation is still more common among the less educated, cohabitation ceases to be a fringe phenomenon among the well-educated. Among highly educated women with a college degree, almost half have ever cohabited. Bumpass and Lu (2000) conclude that economic constraints could not explain the differentials in cohabitation rates among the different educational groups since cohabitation is so common for all groups. Table 2 also shows rates of ever cohabiting women by race. Cohabitation is most common among blacks but the racial divide in prevalence of cohabitation has been slightly reduced. The strongest increase in cohabitation rates was among whites, continuing the trend identified in Bumpass and Lu (2000). To the extent that economic disadvantages are still associated with race in the United States this trend supports the argument that cohabitation has now extended to the middle class.

Table 3 shows the means of selected variables for women who cohabited before their first marriage and for women who did not.[4] In the pooled sample, a bit more than a third of first marriages are preceded by premarital cohabitation. Cohabitors have lower educational achievement than non-cohabitors showing the well-known association between socioeconomic background and cohabitation. Furthermore, cohabitors are on average younger reflecting a cohort effect with more recent cohorts more likely to have cohabited before entering marriage. The recent rise in cohabitation rates have been described in other papers (Bumpass and Sweet 1989; Bumpass et al. 1991; Bumpass and Lu 2000; Kennedy and Bumpass 2008), and similar results can be found using the pooled data (Reinhold 2009). At the same time, cohabitors are older at the day of their marriage, partly reflecting the time spent in cohabitation before marriage. It has been shown that young age is a predictor of marital dissolution, 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. There is an important difference in fertility behavior between cohabitors and non-cohabitors. Cohabitors are much more likely to have children outside of marriage and are less likely to have children within the marriage.

[4] In the following, I refer to women who cohabited before their marriage as 'cohabitors.'

4 Empirical Models of Marital Instability

I estimate proportional hazard models of dissolution of first marriages controling for other observable differences. These models are estimated pooling all three cycles of the NSFG. Premarital cohabitation is interacted with the age of the individual and with the year of the marriage allowing to assess whether the association of premarital cohabitation and marital instability has changed for more recent birth and marriage cohorts. Robustness checks to problems with the NSFG 2002 are reported below.

Changes in the association between premarital cohabitation and marital instability may either 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 three first marriages to account for unobserved person-specific effects (see Lillard et al. 1995). These panel models can only be estimated using the NSFG 2002 because detailed information on cohabitation histories are not available for all higher parity marriages prior to the NSFG 2002.

4.1 Proportional Hazard Models

I first use proportional hazard regressions (Cox 1972) to facilitate comparison with earlier empirical work and with the panel models using only NSFG 2002 data.

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

$$h\left(t|X\left(t\right)\right) = h_0\left(t\right) * \exp(X\left(t\right)'\beta) \tag{1}$$

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

In table 4, I present proportional hazard regressions for the pooled data. The dependent variable is the hazard of marital dissolution for the first marriage. All coefficients are reported as hazard ratios: a coefficient of greater than one indicates that this regressor increases the risk of marital dissolution while a coefficient smaller than one indicates a decrease in risk. I choose other explanatory variables that were found as predictors of marital success in earlier studies. These include education, race, religion, fertility indicators, and age at marriage. In addition, I interact premarital cohabitation with the age of the respondent and the date of the marriage.

In the first column, 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 previous studies for the United States.[5] Thus, despite some potential flaws one is able to reproduce previous empirical research. In the second column, I show the specification where premarital cohabitation is interacted with the age of the woman in 2002. The coefficient on this interaction is positive and highly significant. This means that for an additional year in age the risk of marital dissolution increases by about 1.4% for cohabitors. Based on this specification, cohabitation is associated with an increase in the risk of marital dissolution for women born before about 1981 while for younger women it is associated with a decrease in the risk of marital dissolution. Figure 4 plots graphically the relationship between the hazard ratio for premarital cohabitation and age cohorts. A similar picture emerges when analyzing the interaction of premarital cohabitation with marriage cohorts (see figure 5). For cohabitors in later marriage cohorts the risk of of marital dissolution is reduced by around 2% per year. This indicates that for marriages contracted before 1993 premarital cohabitation is associated with an increase in the risk of marital dissolution while for later marriage cohorts cohabitation is associated with a decrease in risk. Both specifications interacting cohabitation with cohorts show that for more recent cohorts cohabitation is not associated with higher rates

[5] Teachman 2002, for example, reports an increase in the risk of marital dissolution by around 35%

of marital dissolution and these cohort effects are statistically significant.

In the fourth column, I show 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. One explanation for this finding is that premarital cohabitation is not selective of divorce prone individuals in this educational group.

One can attribute the change in the coefficient on premarital cohabitation to either a change in the causal effect of cohabitation or to a change in the process of self-selection, and the proportional hazard models do not allow to distinguish between those alternatives. However, given the dramatic changes in cohabitation behavior, one would expect strong changes in the process of self-selection. It is therefore plausible to attribute the change in the coefficient mainly to a change in the process of self-selection. My findings on the role of education is consistent with this view: For less educated women cohabitation has always been more common than for other socioeconomic groups. For this reason, self-selection has not been as severe within this educational group even for earlier cohorts where 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 the small minority of well-educated cohabiting women was perhaps more selective of divorce-prone individuals. In the next section, I further investigate whether the process of self-selection into premarital cohabitation has changed by accounting for the unobserved person-specific effects as in Lillard et al. (1995).

Other results are in line with previous studies: Premarital conception increases the risk of marital dissolution while a marital birth decreases this risk. Religious affiliations decrease the risk of marital dissolution. Race plays a role: White and other non-Black respondents have more stable marriages than Black respondents. Respondents with an intact family background are less likely to get separated. A higher age for wives at wedding reduces the risk of separation greatly as it is found in many other empirical studies. Because of the potential data flaws in the NSFG 2002 it is important to note that none of the dummies for the different surveys are statistically significant.

4.2 Robustness Checks

Because of data flaws in the NSFG 2002 Kennedy and Bumpass (2008) caution that the data may not be reliable for analyzing marital dissolution. For this reason, I conduct extensive robustness checks. There are two main problems with the data: (i) a routing error in the survey instrument for the NSFG 2002 resulting in a non-random skip pattern for marital dissolution dates (ii) changes in the inclusion of recent migrants in the surveys.

Because of these potential flaws in the data I conduct several robustness checks on the pooled data: I estimate models of marital dissolution using (i) a subsample excluding observations with imputed dates (ii) a subsample excluding migrants in the NSFG 2002 (iii) the pooled data without the NSFG 2002. In a separate analysis for the NSFG 2002 I investigate which covariates are correlated with having an imputed value for the date of marital dissolution. In addition, I investigated whether the results using the NSFG 2002 separately are different from the NSFG 1995 for the same birth cohorts.[6] For the NSFG 1995 I estimate a model for the age groups 15 to 37 years while for the NSFG 2002 I estimate a model for the age groups 22 to 44 years covering the same birth cohorts for both surveys. I artificially censor the NSFG 2002 in 1995.

Table 5 shows the robustness checks for the pooled data. In the first three columns, results for the data without the potentially flawed data from the NSFG 2002 are shown. The first column shows the basic specification without an interaction term between premarital cohabitation and marriage or age cohorts. The coefficient on premarital cohabitation only changes slightly, and there are no large differences in the coefficients on the other coefficients except for the dummy for the 2002 survey. The second column shows the model where premarital cohabitation is interacted with age cohorts. The coefficient on this interaction increases in sign compared to the whole sample showing an even stronger trend in the

[6] I thank a reviewer for this suggestion

relationship between premarital cohabitation and marital instability. For each additional year in age, cohabitors experience an increase in risk by about 1.8% in the restricted sample compared to 1.4% in the whole sample. The qualitative conclusion, however, remains the same. For more recent birth cohorts, cohabitation is not associated with an increase in risk for marital dissolution. In the third column, I present the results for the model including an interaction term between the marriage cohort and premarital cohabitation. The only difference in this restricted sample is that the dummy on the survey 2002 becomes statistically significant but for the coefficients of interest there is almost no change in comparison to the whole sample. In the next three columns the results for a restricted sample without migrants from the NSFG 2002 are shown. For this restricted sample, there are no remarkable differences to the complete sample, and thus, it is unlikely that differences in coverage of recent migrants or potential data flaws in the NSFG 2002 drive the results for the trend in the association between premarital cohabitation and marital instability.

Finally, in the last three columns I show the results for the pooled data for the NSFG 1988 and 1995. Column VII shows the result for the basic specification. According to this estimate, premarital cohabitation increases marital instability by around 35% which is somewhat higher than in the pooled sample without the NSFG 2002. For both the age cohorts and the marriage cohorts, I find even stronger trends in the association between premarital cohabitation and marital instability than in the whole sample including the NSFG 2002. However, the same qualitative conclusion as in the whole sample emerges. For more recent age and marriage cohorts premarital cohabitation is no longer associated with an increase in the risk of marital dissolution. From all three robustness checks, I conclude that my results are not driven by data flaws in the NSFG 2002 or differences in the inclusion of recent migrants.[7]

[7] The results for the NSFG 1988 and 1995 can also be compared to Teachman (2002). One should note that the focus of his study is not premarital cohabitation but other risk factors because he covers a much different time period where not much information on premarital cohabitation was available in the NSFG. In a basic model for the NSFG 1988 and 1995, he finds that cohabitation increases marital instability by around 35% which is exactly the same as my result. In addition, he tests a specification including an interaction between marriage cohorts and premarital cohabitation. While he finds evidence supporting a trend he cannot reject the null of no change across cohorts. For this test, however, he uses the Bayesian Information Criterion (BIC) for model selection which requires a much stronger t-value. The value of the

For the panel models in the next section, I only use the NSFG 2002 because information on higher order marriages is not available in previous cycles of the NSFG. As a first robustness check, I investigate whether the probability of having an imputed value for the date of marital dissolution is correlated with premarital cohabitation (see Appendix). I do find that there is positive correlation between premarital cohabitation and the probability of having an imputed value. In a separate model, I include a dummy for having an imputed value of the date of marital dissolution into the proportional hazard models. I find that women with imputed values for the date of marital dissolution are at greatly increased risk of marital breakup. At the same time the coefficient on premarital cohabitation decreases in sign. One would expect that in a proportional hazard model without the dummy for an imputation the coefficient on premarital cohabitation should be biased upwards, and this is what I find in this robustness check. Thus, the estimates using all observations from the NSFG 2002 are conservative in the sense that one is more likely to find a positive relationship between premarital cohabitation and marital breakup when this data set is used confirming the results from robustness checks using the pooled data. Excluding the observations with imputed values or all observations from the NSFG 2002 from the pooled sample one finds an even stronger decline for more recent birth and marriage cohorts in the association between premarital cohabitation and marital instability.

I conduct an additional robustness check where I compare the results using the the same birth cohorts in the NSFG 1995 and 2002 but artificially censor the NSFG 2002 data in 1995 (see Appendix). Since the two surveys are independent I can conduct a simple t-test of whether the coefficients on premarital cohabitation are the same (p-value=0.257). Thus, I cannot reject the null that the coefficients on premarital cohabitation are the same across the two surveys if one analyzes the same birth cohorts. However, one should note that one finds some differences in the coefficients on other covariates as for instance religion.

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 does not report the coefficient on the interaction between marriage cohorts and premarital cohabitation or even the sign of this coefficient, nor does he test for interactions between age cohorts and premarital stability.

From these robustness checks, I conclude that it is possible to cautiously use the NSFG 2002 to analyze the relationship between premarital cohabitation and marital instability. One should, however, be careful to use the NSFG 2002 to analyze other risk factors of marital dissolution as they might be more biased.

4.3 Accounting for Unobserved Heterogeneity

The proportional hazard models in the previous sections do not allow conclusions about the causal effect of premarital cohabitation on marital stability since the coefficient on premarital cohabitation is likely to be tainted by self-selection. As discussed earlier there are two sources of unobserved heterogeneity: The first source are time-varying person-specific factors including the unobserved match-quality in the BLS model. Instrumental variables are, in general, one way to deal with this sort of endogeneity. Instrumental variables must be correlated with the endogenous regressor and must not be correlated with the error term in the main regression to be valid. For this reason, it is very difficult to think of a good instrument in the context of cohabitation. Since cohabitation and marriage are similar interdependent decision problems, one would not expect to find a variable satisfying the necessary condition, for in the BLS model, all variables that affect the probability of marital dissolution are also likely to affect the probability of premarital cohabitation. To the knowledge of the author, no previous study has attempted to implement an instrumental variable estimator in the context of cohabitation and marriage. The second form of heterogeneity are time-invariant, person-specific effects, as for instance unobserved 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. One shortfall of this approach is that there may be different dynamics at play in higher order marriages compared to first marriages and that these are correlated with the decision to cohabit. Teachman (2008), for example, finds that premarital cohabitation in the second marriage does not raise the risk of marital dissolution. For this reason, I also use specifications including interactions between premarital cohabitation and higher order marriages since this is the main coefficient of interest. Teachman also discusses other risk factors which may have a different influence in second marriages compared to first marriages. For sake of simplicity and to allow for better comparability with previous results, however, I do not model those additional interactions.

Lillard et al. (1995) account for this second source of endogeneity by modeling the decision to cohabit jointly with the marriage dissolution process. For identification they rely on the presence of multiple marriages for one woman. With a random effects assumption, they can identify the correlation between unobserved person-specific characteristics in the cohabitation and the marital dissolution process. In this study I estimate the model with their assumption of bivariate normality and extend by relaxing the distributional assumption (see also Svarer 2005 who estimated a similar model on Danish data). This allows a direct comparison of the results.

4.3.1 The Lillard, Brien, and Waite (1995) model[8]

LBW model the decision to cohabit before marriage and the marital dissolution process simultaneously. There is an unobserved heterogeneity term in both of these processes that may be correlated. Conditional on all other covariates and the person-specific components cohabitation is independent of idiosyncratic match-specific quality.

The heterogeneity term is assumed to be permanent for a person so that 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. LBW interpret the coefficient on premarital cohabitation as the causal effect of cohabitation on marital stability since it is purged from any self-selection if the model is correct. I discuss the marital dissolution process and the decision to cohabit separately.

LBW use a continuous time duration model of marriage. The logarithm of the instan-

[8] LBW henceforth.

taneous probability of dissolution at time t for the m^{th} marriage conditional on not having dissolved before that time (log-hazard) is given by the following equation:

$$\begin{aligned}\ln h_m\left(X_m^d, Coh_m, t, \delta\right) &= \alpha_0 + \alpha_1^{'} DurMar\left(t\right) + \alpha_2^{'} DurBirth\left(t\right) + \\ &\quad \alpha_3^{'} X_m^d + \alpha_4^{'} Coh_m + \delta \end{aligned} \tag{2}$$

In this equation $DurMar$ and $DurBirth$ represent duration splines starting at the beginning of the marriage and the birth of the first child respectively. The time-dependent part can also be written as:

$$h_{0m}\left(t\right) = \exp\left(\alpha_0 + \alpha_1^{'} DurMar\left(t\right) + \alpha_2^{'} DurBirth\left(t\right)\right) \tag{3}$$

This is the baseline hazard. All the other regressors will shift this baseline hazard proportionally. The baseline survivor function is given as:

$$S_{0m}\left(t\right) = \exp\left(-\int_{t_0}^{t} h_{0m}\left(t\right) dt\right) \tag{4}$$

The survival function is then given as:

$$S_m\left(X_m^d, Coh_m, t, \delta\right) = \prod_{i=1}^{I} \left(\frac{S_{0m}\left(t_{i+1}\right)}{S_{0m}\left(t_i\right)}\right)^{-\exp\left(\alpha_3^{'} X_m^d + \alpha_4^{'} Coh_m + \delta\right)} \tag{5}$$

where I is the number of periods in which the covariates are constant. The regressor set X_m^d includes regressors that are fixed for a given marriage but may vary across one individual's marriages including dummies for higher order marriages, education, age at wedding, and other socioeconomic variables. The coefficient α_4 measures the effect of premarital cohabitation (with the future spouse). Finally, there is an unobserved component δ which is assumed to be fixed for all marriages of a given woman.

The decision to cohabit before the m^{th} marriage is modeled as a probit model:

$$I_m = \beta_0 + \beta_1' X_m^c + \epsilon + \eta_m \tag{6}$$

$$Coh_m = \begin{cases} 1 \text{ if } I_m > 0 \\ 0 \text{ otherwise} \end{cases} \tag{7}$$

The set of regressors X_m^c again includes socioeconomic variables, ϵ is again the unobserved heterogeneity. It is constant across all marriages of an individual. Finally, η_m is distributed i.i.d. according to a standard normal distribution.

LBW assume that the heterogeneity components δ and ϵ are drawn from a bivariate normal distribution. That is

$$\begin{pmatrix} \delta \\ \epsilon \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_\delta^2 & \sigma_{\delta\epsilon} \\ \sigma_{\delta\epsilon} & \sigma_\epsilon^2 \end{pmatrix} \right) \tag{8}$$

A positive correlation, that is $\sigma_{\delta\epsilon} > 0$, would indicate self-selection of individuals with a high risk of marital disruption into premarital cohabitation The unobserved heterogeneity is integrated out so that the joint marginal likelihood contribution of all marriages of a given woman is then given by:

$$\int_\delta \int_\epsilon \frac{1}{\sigma_\delta \sigma_\epsilon} \phi\left(\frac{\delta}{\sigma_\delta}, \frac{\epsilon}{\sigma_\delta} \mid \rho_{\delta\epsilon} \right) \prod_{m=1}^{M} \Big[S_m\left(X_m^d, Coh_m, t, \delta\right) h_m\left(X_m^d, Coh_m, t, \delta\right)^{D_m} \times \Phi\left((2Coh_m - 1)\left(\beta_0 + \beta_1' X_m^c + \epsilon\right)\right) \Big] \, d\delta d\epsilon \tag{9}$$

where the product is taken over the first three marriages of a woman. $D^m = 1$ indicates a completed marriage spell while $D^m = 0$ indicates a censored spell. 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 since one observes more than one marriage for some women.

This study relaxes the assumption of bivariate normality and models the heterogeneity components as being drawn from a finite, discrete distribution (see also Svarer 2005). The random effects in the hazard and probit equation are given by:

$$e_h = v_0 \tag{10}$$

$$e_p = \rho_p * v_0 + v_1 \tag{11}$$

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

$$v_i = \begin{cases} m_{i,1} \text{ with probability } w_{i,1} \\ m_{i,2} \text{ with probability } w_{i,2} \end{cases} \tag{12}$$

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

4.3.2 Results of Random Effects Model

After controlling for other exogenous covariates and unobserved person-specific heterogeneity, LBW interpret the coefficient on cohabitation as the true causal effect of cohabitation on marital instability. The correlation between the heterogeneity components sheds light on the process of self-selection into cohabitation and marriage where a positive correlation indicates that cohabitors lead also less stable marriages. LBW found a strong and positive correlation between these heterogeneity components and no causal effect of premarital cohabitation on marital instability using the National Longitudinal Study of the High School Class of 1972 with its follow-up in 1986. A comparison of their estimates with mine allows to study whether the process of self-selection has changed during the last decades. Unfortunately, detailed cohabitation data relative to all higher order marriages was not collected for the NSFG 1988 and 1995. Therefore, I cannot study the time-evolution of the correlation between the heterogeneity components with my datasets. To facilitate comparison, I estimate one specification similar to LBW where the unobserved heterogeneity is bivariate normal and where the coefficient on premarital cohabitation is constrained to be the same for all marriages with the only difference in the specifications arising from a different set of covariates.

In addition, in this study the decision to cohabit is interacted with dummies for higher order marriages. In proportional hazard models[9], I found that the effect of cohabitation is different in second and third marriages, and the interaction effects dramatically improves the fit of the models. Furthermore, premarital cohabitation is much more common in higher order marriages, and the incentives to cohabit may be different in first marriages than in later marriages. This study also relaxes the assumption of bivariate normal heterogeneity and approximates the underlying distribution with a finite discrete mixture. Table 6 presents the estimation results for the models with bivariate normal heterogeneity and table 7 presents the results using a finite discrete mixture. For a comparison, both tables include specifications without unobserved heterogeneity.

Bivariate Normal Heterogeneity The results for the models with bivariate normal heterogeneity are shown in table 6 for the marital dissolution process[10]. Coefficients are again reported as hazard ratios where a coefficient greater than one indicates that the variable increases marital instability while a coefficient smaller than one indicates a reduction in the hazard of marital dissolution.

In the first two columns of table 6, the results for a parametric hazard model without heterogeneity are shown. The first column shows a coefficient on premarital cohabitation of 0.871 indicating a reduced hazard of marital dissolution for cohabitors. The coefficient on the dummy for higher order marriages indicates that first marriages are more stable than later marriages. The coefficients on the other controls lead to similar qualitative conclusions than the coefficients in the proportional hazard models. In the second column, the model includes

[9]These results are available from the author upon request

[10]The results for the probit model are available upon request

the interactions of premarital cohabitation with dummies for second and third marriages. According to these estimates, premarital cohabitation is not associated with a reduced hazard of marital dissolution in first marriages, but it dramatically reduces the hazard in higher order marriages. The coefficient on premarital cohabitation is 1.014 which is of comparable size than the coefficient of the proportional hazard model using only first marriages for the NSFG 2002. Like in the proportional hazard models one finds that with the new NSFG 2002 dataset there is only a very weak association between premarital cohabitation and the marital dissolution process. The interactions of premarital cohabitation and dummies for second and third marriages, however, are statistically significant and their impact are of considerable size. At the same time, the coefficient on higher order marriages alone has also increased in size. One interpretation is that higher order marriages are less stable only for couples who have not cohabited before their marriage. Like in the proportional hazard models, cohabitation with partners other than the future spouse is still associated with a higher hazard of marital dissolution. The other coefficients are broadly in line with the results of the proportional hazard regressions.

In columns 3 and 4 of table 6 the specifications with unobserved heterogeneity but no correlation between the unobserved components are presented. Introducing this form of heterogeneity improves greatly the overall fit of the model as judged by a comparison of the Log-Likelihood statistic for both models with the models without unobserved heterogeneity. In column 3 the model without interactions between premarital cohabitation and higher order marriages is presented. The coefficient on premarital cohabitation itself is not greatly affected by introducing unobserved heterogeneity. Introducing these random effects has a big impact on the duration dependency of the marital dissolution process. In addition, the coefficient on higher order-marriages is reduced somewhat. Therefore, the instability of higher order marriages is at least partly explained by selection of more divorce-prone individuals into higher order marriages. In column 4 the model with interactions is shown. Again, introducing unobserved heterogeneity without correlation between the unobserved heterogeneity components does not greatly affect the estimated coefficients on premarital cohabitation and its interactions.

In columns 5 and 6 of table 6, the coefficients under the assumption of bivariate normality with no restrictions on the correlation between the unobserved components are shown. In the specification without interactions between premarital cohabitation and dummies for higher order marriage the correlation is positive and statistically significant and even higher than the correlation coefficient in LBW. More divorce-prone individuals are more likely to cohabit. At the same time, the coefficient on premarital cohabitation becomes smaller than one and statistically significant. While there is self-selection of divorce-prone individuals into premarital cohabitation the causal effect of cohabitation is a dramatic reduction in the hazard of marital dissolution by about a half. In comparison, LBW did not find a significant causal effect of premarital cohabitation on marital stability. Introducing, higher order marriages shows that the reduction in separation risk for cohabitors is not as strong for first marriages. At the same time, the positive correlation between the unobserved heterogeneity components is reduced after introducing these interactions. Neglecting the heterogeneous effects of premarital cohabitation in first and later marriages leads to an overestimate of the degree of self-selection into premarital cohabitation. The intuition for this is that premarital cohabitation is much more common in higher order marriages than in first marriages. At the same time, higher order marriages are selective of divorce-prone individuals resulting in a spurious overestimate of the self-selection of divorce-prone individuals into premarital cohabitation. Introducing the interactions of higher order marriages with premarital cohabitation partly alleviates this problem but overall one still finds that there is positive self-selection of divorce-prone individuals into premarital cohabitation.

Finite Discrete Mixture In a further step, I relax the distribution assumption on the unobserved person-specific effects and estimate them non-parametrically with a discrete finite mixture. The results are presented in table 7. For comparison, I report the results for the model without unobserved heterogeneity in the first two columns of this table.

I find that in each process one could approximate the unobserved heterogeneity with a two-point discrete distribution. Introducing more support points did not significantly improve the fit of the model. In columns 3 and 4, I restrict the correlation between the heterogeneity components to zero. The coefficient estimates on premarital cohabitation and its interactions are broadly similar to the ones under the assumption of normality and no correlation. In the model without interactions between premarital cohabitation and marriage order premarital cohabitation reduces the hazard of marital dissolution. In the model with interactions (column 4), premarital cohabitation has no effect on the hazard of marital dissolution in the first marriage, but it significantly reduces the hazard in the second and third marriage.

In the last two columns, I present the results allowing correlation between the unobserved heterogeneity components. In the model without interactions between cohabitation and higher order marriages (column 5), one finds statistically significant correlation between the unobserved heterogeneity components. The coefficient on premarital cohabitation becomes smaller than 1 and is statistically significant. In the preferred estimate including the interactions between cohabitation and higher order marriages (column 6) the correlation becomes smaller and insignificant. The interaction terms of premarital cohabitation and higher order marriages are statistically significant. At the same time, premarital cohabitation has no effect on marital stability in first marriages.

These new results are different in some points to LBW. When a bivariate normal distribution or no interactions between premarital cohabitation and higher order marriage are used, one finds self-selection of divorce-prone individuals into premarital cohabitation which is consistent with their earlier findings. However, in contrast to their earlier findings, in the new estimates the coefficient on premarital cohabitation becomes smaller than one. Premarital cohabitation reduces the risk of marital dissolution, and this effect is strong.

A further result of comparing the estimates based on the two distributional assumptions is that the results of the LBW model are sensitive to the distributional assumption. With the nonparametric assumption one finds no self-selection of divorce-prone individuals into premarital cohabitation while one finds very significant self-selection when one assumes Normality. Furthermore, premarital cohabitation only reduces the hazard of marital dissolution in the second and third marriage but not in the first. Probably, premarital cohabitation in higher order marriages follows a different dynamic than in the first marriage. Neglecting these interactions leads to an overestimation of the correlation between the unobserved heterogeneity components.

5 Conclusion

In this essay, I have 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 and should decrease 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 empirical evidence that has been established for the US and other industrialized countries showing that marriages preceded by cohabitation are less stable.

This essay demonstrates that the once-strong association between premarital cohabitation and marital instability has weakened over time, and there is no longer 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 ceased to be selective of individuals with high risk of marital break up. The results for different educational groups supports this view. For women with low educational attainment, who always have had high rates of premarital cohabitation, there is no increased risk of marital dissolution for cohabitors. Several robustness checks show that these results are not sensitive to potential data flaws in the NSFG 2002.

Using the LBW model, I find mixed evidence for the hypothesis that the process of self-selection has changed. When using their distributional assumption of bivariate normality for the unobserved heterogeneity components, I find self-selection of more divorce-prone individuals into premarital cohabitation. However, the nonparametric estimates indicate that the normality assumption may be a poor approximation. Estimating the distribution nonparametrically and with interactions between premarital cohabitation and higher order marriages, I do not find significant self-selection and no causal effect of cohabitation in the first marriage. In higher order marriages, on the other hand, cohabitation is important in stabilizing marriages. The earlier finding of a positive self-selection into premarital cohabitation may therefore be spurious and be driven by the particular assumption and by different dynamics of premarital cohabitation in first and later marriages.

While my results are 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 half of the population cohabits, cohabitation ceases to be selective of divorce-prone individuals. Another reason that possibly puts doubt on the thesis of a stable relationship between premarital cohabitation and marital instability is that the character of cohabitation might have changed over time. This change in character of cohabitation might be badly measured by a binary indicator of cohabitation. Cherlin (2004) cites work by the British demographer Kiernan suggesting that acceptance of cohabitation in society follows four steps. In the first step, cohabitation is a fringe phenomenon, after which it becomes acceptable as testing ground for marriage in a second step. In step three, cohabitation becomes an accepted alternative to marriage, and finally it even becomes virtually indistinguishable from marriage. If this thesis is true, then it would be even more likely that the self-selection process into cohabitation has changed fundamentally. Given the rapid changes in marriage and cohabitation behavior in the United States, the relationship between cohabitation and marital instability may not be stable yet.

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Table 1: Trends in Cohabitation Rates among non-Married Women

	Percentage ever cohabiting			Percentage currently cohabiting		
Age	1987	1995	2002	1987	1995	2002
19-24	29	38	38	14	15	19
25-29	41	47	58	20	21	26
30-34	40	49	61	17	21	20
35-39	30	48	59	11	17	18
40-44	22	41	54	14	13	14
Total	33	45	54	15	17	19

Notes: Percentages for years 1987 and 1995 are taken from table 1 in Bumpass and Lu (2000). Percentages for 2002 are taken from table 2 in Kennedy and Bumpass (2008)

Table 2: Percentage of Women Aged 19-44 Who Have Ever Cohabited

	Percentage of women who have ever cohabited		
	1987	1995	2002
Education			
Less than high school	43	59	64
High school	32	46	63
Some College	30	39	49
College 4+	31	37	45
Race/Ethnicity			
White non-Hispanic	32	45	54
Black	36	45	57
Hispanic	30	39	52

Notes: Percentages for years 1987 and 1995 are taken from table 2 in Bumpass and Lu (2000). Percentages for 2002 are taken from table 3 in Kennedy and Bumpass.

Table 3: Means of Variables Relative to First Marriage

	Non-cohabitors	Cohabitors
Percentage in population	63.8	36.2
Less than high school (%)	14.9	18.9
High school (%)	35.3	31.9
More than high school (%)	49.8	49.2
Year of birth	1959.5	1962.9
Age at wedding (wife)	21.3	23.3
Age at wedding (husband)	24.1	26.5
Age difference	2.8	3.2
Premarital conception (%)	28.9	51.9
Premarital live-birth (%)	7.7	23.2
Marital live-birth (%)	85.3	75.4

Notes: Sample weights are used.

Table 4: Proportional Hazard Regressions in Pooled Sample.
Dep. Var. Hazard of Separation of First Marriage

Variable	I	II	III	IV
Cohabitation	1.293***	0.747*	1.244***	
	(0.053)	(0.121)	(0.051)	
Cohabitation × age in 2002		1.014***		
		(0.004)		
Cohabitation × year of marriage			0.979***	
			(0.004)	
No HS × cohab				1.158*
				(0.092)
High School × cohab				1.414***
				(0.088)
More than HS × cohab				1.270***
				(0.070)
Education[a]				
No HS	1.008	1.012	1.018	1.046
	(0.051)	(0.052)	(0.052)	(0.065)
High School	1.021	1.025	1.032	0.985
	(0.042)	(0.043)	(0.043)	(0.050)
Fertility Var.				
Premarital conception	1.496***	1.496***	1.500***	1.495***
	(0.064)	(0.064)	(0.064)	(0.064)
Premarital livebirth	1.107*	1.113*	1.115*	1.113*
	(0.064)	(0.065)	(0.065)	(0.064)
Marital live-birth	0.706***	0.709***	0.710***	0.705***
	(0.030)	(0.030)	(0.030)	(0.030)
Religion[b]				
Protestant	0.770***	0.768***	0.764***	0.769***
	(0.043)	(0.043)	(0.043)	(0.043)
Catholic	0.728***	0.725***	0.719***	0.728***
	(0.046)	(0.046)	(0.045)	(0.046)
Other	0.786***	0.785***	0.783***	0.785***
	(0.072)	(0.072)	(0.072)	(0.072)

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Table 4 cont'd

Variable	I	II	III	IV
Race[c]				
White	0.784***	0.785***	0.785***	0.786***
	(0.037)	(0.037)	(0.037)	(0.037)
Other	0.623***	0.620***	0.614***	0.624***
	(0.055)	(0.055)	(0.054)	(0.055)
Family Background				
No int. family	1.276***	1.279***	1.280***	1.277***
	(0.053)	(0.053)	(0.052)	(0.053)
Wife's age at wedding	0.901***	0.901***	0.896***	0.901***
	(0.007)	(0.007)	(0.007)	(0.007)
Husband's age at wedding	1.010**	1.009**	1.010**	1.010**
	(0.004)	(0.004)	(0.004)	(0.004)
Wife's age in 2002	0.992***	0.987***		0.992***
	(0.003)	(0.003)		(0.003)
Year of marriage			1.015***	
			(0.003)	
NSFG 1995	0.980	1.002	0.969	0.979
	(0.040)	(0.042)	(0.041)	(0.040)
NSFG 2002	1.044	1.032	1.047	1.043
	(0.069)	(0.069)	(0.070)	(0.070)
Number of obs.	15849	15849	15849	15849

Notes: Sample weights adjusted for different sample sizes are used. Estimates reported as hazard ratios. A coefficient of greater than one indicates an increase in the hazard of marital dissolution while a coefficient of smaller than one indicates a decrease in the hazard.

a: Omitted category is more than high school.

b: Omitted category is no religion.

c: Omitted category is black.

Standard errors in parentheses. *** 1% ** 5% *10% significant different from 1.

Table 5: Robustness Check in Pooled Sample.Dep. Var. Hazard of Separation of First Marriage

Variable	without imputed			without foreign born			without NSFG 2002		
	I	II	III	IV	V	VI	VII	VIII	IX
Cohabitation	1.302***	0.650**	1.211***	1.275***	0.685**	1.213***	1.359***	0.665**	1.186***
	(0.056)	(0.110)	(0.053)	(0.053)	(0.113)	(0.051)	(0.059)	(0.123)	(0.059)
Cohabitation		1.018***			1.016***			1.018***	
×age in 2002		(0.004)			(0.004)			(0.005)	
Cohabitation			0.971***			0.976***			0.969***
×year of marriage			(0.004)			(0.004)			(0.005)
Education[a]									
No HS	0.965	0.971	0.979	1.047	1.053	1.062	1.035	1.048	1.058
	(0.051)	(0.052)	(0.052)	(0.055)	(0.055)	(0.056)	(0.058)	(0.059)	(0.060)
High School	1.000	1.006	1.014	1.033	1.038	1.046	0.964	0.971	0.979
	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.040)	(0.040)	(0.041)
Fertility Var.									
Premarital	1.507***	1.508***	1.514***	1.481***	1.480***	1.484***	1.418***	1.421***	1.428***
conception	(0.067)	(0.067)	(0.067)	(0.064)	(0.064)	(0.064)	(0.063)	(0.063)	(0.063)
Premarital	1.054	1.060	1.061	1.087	1.093	1.093	1.147**	1.151**	1.147**
live-birth	(0.064)	(0.065)	(0.065)	(0.064)	(0.065)	(0.065)	(0.072)	(0.073)	(0.073)
Marital	0.703***	0.707***	0.709***	0.717***	0.721***	0.722***	0.697***	0.701***	0.704***
live-birth	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)	(0.033)	(0.034)	(0.034)
Religion[b]									
Protestant	0.772***	0.770***	0.765***	0.767***	0.765***	0.760***	0.747***	0.746***	0.744***
	(0.045)	(0.045)	(0.045)	(0.044)	(0.044)	(0.043)	(0.048)	(0.048)	(0.048)
Catholic	0.740***	0.736***	0.728***	0.739***	0.736***	0.730***	0.689***	0.686***	0.681***
	(0.049)	(0.048)	(0.048)	(0.048)	(0.048)	(0.047)	(0.047)	(0.047)	(0.047)
Other	0.783***	0.782***	0.779***	0.825**	0.826***	0.824***	0.737***	0.736***	0.733***
	(0.074)	(0.074)	(0.074)	(0.078)	(0.078)	(0.078)	(0.078)	(0.078)	(0.078)

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Table 5 cont'd

Variable	I	II	III	IV	V	VI	VII	VIII	IX
Race[c]									
White	0.801***	0.802***	0.804***	0.767***	0.768***	0.767***	0.762***	0.763***	0.767***
	(0.039)	(0.039)	(0.039)	(0.036)	(0.036)	(0.036)	(0.039)	(0.039)	(0.039)
Other	0.599***	0.595***	0.588***	0.652***	0.647***	0.640***	0.607***	0.601***	0.593***
	(0.057)	(0.057)	(0.056)	(0.061)	(0.060)	(0.060)	(0.063)	(0.063)	(0.062)
Family									
Background									
No int. family	1.297***	1.301***	1.303***	1.280***	1.284***	1.284***	1.234***	1.236***	1.238***
	(0.056)	(0.056)	(0.056)	(0.053)	(0.053)	(0.053)	(0.051)	(0.051)	(0.051)
Wife's age	0.895***	0.895***	0.893***	0.898***	0.898***	0.892***	0.900***	0.901***	0.892***
at wedding	(0.007)	(0.007)	(0.008)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
Husband's age	1.006	1.005	1.005	1.012***	1.012***	1.012***	1.011**	1.010**	1.010**
at wedding	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
Wife's age in 2002	0.995*	0.989***		0.991***	0.986***	0.892***	0.986***	0.980***	
	(0.003)	(0.003)		(0.003)	(0.003)	(0.007)	(0.003)	(0.004)	
Year of marriage			1.014***			1.017***			1.024***
			(0.003)			(0.003)			(0.004)
NSFG 1995	0.997	1.026	0.989	0.973	0.997	0.960	0.944	0.978	0.939
	(0.042)	(0.043)	(0.042)	(0.040)	(0.042)	(0.040)	(0.040)	(0.042)	(0.041)
NSFG 2002	0.880*	0.871*	0.889	1.098	1.086	1.105			
	0.066	(0.065)	(0.066)	(0.076)	(0.074)	(0.076)			
Number of obs.	15375	15375	15375	15049	15049	15049	11806	11806	11806

Notes: Sample weights adjusted for different sample sizes are used. Estimates reported as hazard ratios. A coefficient of greater than one indicates an increase in the hazard of marital dissolution while a coefficient of smaller than one indicates a decrease in the hazard.

a: Omitted category is more than high school.

b: Omitted category is no religion.

c: Omitted category is black.

Standard errors in parentheses. *** 1% ** 5% *10% significant different from 1.

Table 6: Maximum Likelihood Estimation of Duration Model with and without Bivariate Normal Heterogeneity. Dependent Variable: Hazard of Separation of First Three Marriages

Variable	No Heterogeneity		Bivariate Normal $\rho = 0$		Bivariate Normal no restriction on ρ	
Premarital	0.871**	1.014	0.842***	1.002	0.483***	0.665**
cohabit.	(0.048)	(0.059)	(0.054)	(0.072)	(0.085)	(0.117)
Prem. cohabit.*		0.480***		0.422***		0.443***
2^{nd} marr.		(0.062)		(0.066)		(0.073)
Prem. cohabit.*		0.273***		0.159***		0.176***
3^{rd} marr.		(0.077)		(0.047)		(0.058)
Previous	1.274***	1.232***	1.385***	1.347***	1.454***	1.395***
cohabit.	(0.091)	(0.083)	(0.123)	(0.120)	(0.134)	(0.130)
Time since marriage						
Months						
0-12	1.006	1.008	1.019	1.022	1.020	1.022*
	(0.012)	(0.013)	(0.013)	(0.014)	(0.013)	(0.013)
13-48	1.000	1.001	1.006**	1.007**	1.006**	1.008**
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
49-120	0.996***	0.996***	0.999	1.000	1.000	1.000
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
120	0.997**	0.997*	0.999	1.000	1.000	1.000
	(0.002)	(0.002)	0.002	(0.002)	(0.002)	(0.002)
Time since marital birth						
Months						
0-12	0.804***	0.803***	0.788***	0.785***	0.786***	0.785***
	(0.019)	(0.018)	(0.018)	(0.019)	(0.019)	(0.019)
13-24	1.109***	1.108***	1.107***	1.106***	1.109***	1.107***
	(0.027)	(0.026)	(0.027)	(0.026)	(0.027)	(0.027)
>24	1.000	1.000	0.999	0.999	0.999	0.999
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Constant	0.085***	0.080***	0.090***	0.084***	0.148***	0.120***
	(0.025)	(0.024)	(0.030)	(0.030)	(0.055)	(0.045)
Premarital	1.981***	1.953***	2.167***	2.103***	2.158***	2.095***
conception	(0.118)	(0.117)	(0.156)	(0.156)	(0.157)	(0.155)
Premarital	0.742***	0.667***	0.753***	0.676***	0.726***	0.663***
birth	(0.066)	(0.056)	(0.078)	(0.068)	(0.077)	(0.073)
Education[a]						
No HS	1.052	1.024	1.058	1.038	1.074	1.051
	(0.074)	(0.067)	(0.091)	(0.091)	(0.095)	(0.094)
High school	1.148**	1.132***	1.173**	1.164***	1.155**	1.153***
	(0.066)	(0.083)	(0.081)	(0.085)	(0.082)	(0.083)

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Table 6 cont'd

Variable	No Heterogeneity		Bivariate Normal $\rho = 0$		Bivariate Normal no restriction on ρ	
Race[b]						
White	1.119	1.162	1.310	1.265	1.313	1.270
	(0.111)	(0.163)	(0.218)	(0.217)	(0.223)	(0.218)
Black	0.993	0.990	1.009	1.015	1.007	1.013
	(0.123)	(0.124)	(0.144)	(0.152)	(0.147)	(0.149)
Religion[c]						
No religion	1.042	1.089	1.027	1.052	1.064	1.076
	(0.143)	(0.151)	(0.144)	(0.190)	(0.178)	(0.183)
Catholic	1.089	1.155	1.100	1.149	1.045	1.100
	(0.139)	(0.151)	(0.167)	(0.197)	(0.163)	(0.175)
Protestant	0.861	0.917	0.843	0.887	0.765*	0.821
	(0.108)	(0.118)	(0.126)	(0.149)	(0.119)	(0.129)
Family Background						
Intact	0.838***	0.714***	0.669***	0.670***	0.633***	0.645***
	(0.035)	(0.036)	(0.043)	(0.046)	(0.043)	(0.044)
Not born in US	0.672***	0.684***	0.642***	0.649***	0.581***	0.602***
	(0.061)	(0.054)	(0.067)	(0.066)	(0.065)	(0.067)
Wife						
Age at wedd.	0.892***	0.889***	0.877***	0.873***	0.876***	0.873***
	(0.006)	(0.006)	(0.008)	(0.007)	(0.008)	(0.008)
Age at interv.	1.013***	1.014***	1.015***	1.015***	1.009	1.011*
	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)	(0.006)
Has kids	0.871	1.037	0.835*	1.030	1.077	1.219
	(0.074)	(0.088)	(0.084)	(0.106)	(0.138)	(0.157)
>1 marriage	1.874***	2.941***	1.534***	2.537***	1.605***	2.563***
	(0.150)	(0.284)	(0.157)	(0.340)	(0.167)	(0.348)
Age diff.	1.011***	1.012***	1.015***	1.017***	1.015***	1.017***
	(0.003)	(0.002)	(0.005)	(0.003)	(0.005)	(0.005)
Husband						
Divorced	1.252***	1.229***	1.329***	1.332***	1.413***	1.383***
	(0.104)	(0.074)	(0.131)	(0.115)	(0.143)	(0.102)
Has kids	1.179***	1.161**	1.224***	1.206***	1.254***	1.230***
	(0.095)	(0.085)	(0.114)	(0.106)	(0.119)	(0.118)

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Table 6 cont'd

	No Heterogeneity		Bivariate Normal $\rho = 0$		Bivariate Normal no restriction on ρ	
Std. dev.	NA	NA	0.749***	0.810***	0.847***	0.857***
			(0.168)	(0.097)	(0.087)	(0.083)
Correlation	NA	NA	NA	NA	0.587***	0.435***
					(0.141)	(0.158)
Log-L	-13593.09	-13574.70	-13547.09	-13522.91	-13541.07	-13519.71
Number of obs.	4021	4021	4021	4021	4021	4021

Notes: Sample weights are used. Estimates reported as hazard ratios. A coefficient greater than one indicates an increase in the hazard of marital dissolution while a coefficient of smaller than one indicates a decrease in the hazard. Standard error based on numerical standard errors in parentheses.

a: Omitted category is more than high school.

b: Omitted category is other race.

c: Omitted category is other religion.

*** 1% ** 5% *** 10% significant different from 1.

Table 7: Maximum Likelihood Estimation of Duration Model with and without Heterogeneity (Finite Discrete Mixture). Dependent Variable: Hazard of Separation of First Three Marriages

Variable	No Heterogeneity		Finite Discrete Mixture $\rho = 0$		Finite Discrete Mixture no restriction on ρ	
Premarital	0.871**	1.014	0.824***	1.009	0.639***	0.902
cohabit.	(0.048)	(0.059)	(0.050)	(0.072)	(0.059)	(0.099)
Prem. cohabit.*		0.480***		0.406***		0.426***
2^{nd} marr.		(0.062)		(0.060)		(0.064)
Prem. cohabit.*		0.273***		0.175***		0.190***
3^{rd} marr.		(0.077)		(0.049)		(0.054)
Previous	1.274***	1.232***	1.511***	1.480***	1.484***	1.485***
cohabit.	(0.091)	(0.083)	(0.135)	(0.136)	(0.134)	(0.135)
Time since marriage						
Months						
0-12	1.006	1.008	1.014	1.016	1.014	1.015
	(0.012)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
13-48	1.000	1.001	1.006**	1.007**	1.005*	1.007**
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
49-120	0.996***	0.996***	1.001	1.002	1.000	1.001
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
120	0.997**	0.997*	1.002	1.002	1.001	1.002
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Time since marital birth						
Months						
0-12	0.804***	0.803***	0.790***	0.788***	0.790***	0.789***
	(0.019)	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)
13-24	1.109***	1.108***	1.109***	1.107***	1.110***	1.107***
	(0.027)	(0.026)	(0.026)	(0.026)	(0.026)	(0.026)
>24	1.000	1.000	0.997***	0.997***	0.998*	0.997**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Constant	0.085***	0.080***				
	(0.025)	(0.024)				
Premarital	1.981***	1.953***	2.092***	2.029***	2.074***	2.011***
conception	(0.118)	(0.117)	(0.146)	(0.146)	(0.145)	(0.144)
Premarital	0.742***	0.667***	0.823***	0.747***	0.831*	0.757***
birth	(0.066)	(0.056)	(0.080)	(0.075)	(0.079)	(0.075)
Education						
No HS	1.052	1.024	1.038	1.008	1.053	1.018
	(0.074)	(0.067)	(0.086)	(0.085)	(0.086)	(0.084)
High school	1.148**	1.132***	1.127*	1.110	1.128*	1.111
	(0.066)	(0.083)	(0.078)	(0.077)	(0.076)	(0.076)

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Table 7 cont'd

Variable	No Heterogeneity		Finite Discrete Mixture $\rho = 0$		Finite Discrete Mixture no restriction on ρ	
Race						
White	1.119	1.162	1.447**	1.382*	1.455**	1.388**
	(0.111)	(0.163)	(0.239)	(0.233)	(0.238)	(0.232)
Black	0.993	0.990	1.066	1.087	1.068	1.088
	(0.123)	(0.124)	(0.150)	(0.157)	(0.148)	(0.155)
Religion						
No religion	1.042	1.089	1.025	1.036	0.988	1.022
	(0.143)	(0.151)	(0.177)	(0.185)	(0.170)	(0.180)
Catholic	1.089	1.155	1.094	1.130	1.025	1.107
	(0.139)	(0.151)	(0.178)	(0.190)	(0.167)	(0.185)
Protestant	0.861	0.917	0.823	0.854	0.766*	0.834
	(0.108)	(0.118)	(0.131)	(0.141)	(0.122)	(0.137)
Family Background						
Intact	0.838***	0.714***	0.665***	0.656***	0.660***	0.656***
	(0.035)	(0.036)	(0.042)	(0.042)	(0.041)	(0.042)
Not born in US	0.672***	0.684***	0.663***	0.677***	0.652***	0.672***
	(0.061)	(0.054)	(0.061)	(0.064)	(0.060)	(0.063)
Wife						
Age at wedd.	0.892***	0.889***	0.876***	0.871***	0.876***	0.872***
	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)	(0.007)
Age at interv.	1.013***	1.014***	1.015***	1.016***	1.012**	1.015**
	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)	(0.006)
Has kids	0.871	1.037	0.801***	0.981	0.900	1.022
	(0.074)	(0.088)	(0.075)	0.100	(0.087)	(0.105)
>1 marriage	1.874***	2.941***	1.549***	2.627***	1.621***	2.618***
	(0.150)	(0.284)	(0.148)	(0.341)	(0.154)	(0.340)
Age diff.	1.011***	1.012***	1.018***	1.020***	1.019***	1.020***
	(0.003)	(0.002)	(0.003)	(0.003)	(0.004)	(0.003)
Husband						
Divorced	1.252***	1.229***	1.324***	1.314***	1.356***	1.321***
	(0.104)	(0.074)	(0.112)	(0.112)	(0.114)	(0.112)
Has kids	1.179***	1.161**	1.231***	1.207**	1.252***	1.213**
	(0.095)	(0.085)	(0.104)	(0.103)	(0.104)	(0.103)

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Table 7 cont'd

	No Heterogeneity		Finite Discrete Mixture $\rho = 0$		Finite Discrete Mixture no restriction on ρ	
Rho					0.375***	0.127
					(0.127)	(0.102)
Point1			-3.687***	-3.782***	-3.56***	-3.743***
			(0.399)	(0.392)	(0.399)	(0.396)
Point2			-1.776***	-1.760***	-1.540***	-1.691***
			(0.349)	(0.351)	(0.349)	(0.354)
Weight1[d]			-0.380***	-0.338***	-0.516***	-0.409***
			(0.148)	(0.125)	(0.131)	(0.128)
Log-L	-13593.09	-13574.70	-13537.35	-13511.39	-13532.95	-13510.92
Number of obs.	4021	4021	4021	4021	4021	4021

Notes: Sample weights are used. Estimates reported as hazard ratios. A coefficient greater than one indicates an increase in the hazard of marital dissolution while a coefficient of smaller than one indicates a decrease in the hazard. Standard error based on numerical standard errors in parentheses.

a: Omitted category is more than high school.

b: Omitted category is other race.

c: Omitted category is other religion.

d: Reported as $\Phi^{-1}(w1)$ of weight

*** 1% ** 5% *** 10% significant different from 1.

Figure 1: Survival Functions of First Marriages for Cohabitors and Non-Cohabitors and 95% Confidence Intervals (NSFG 1988)

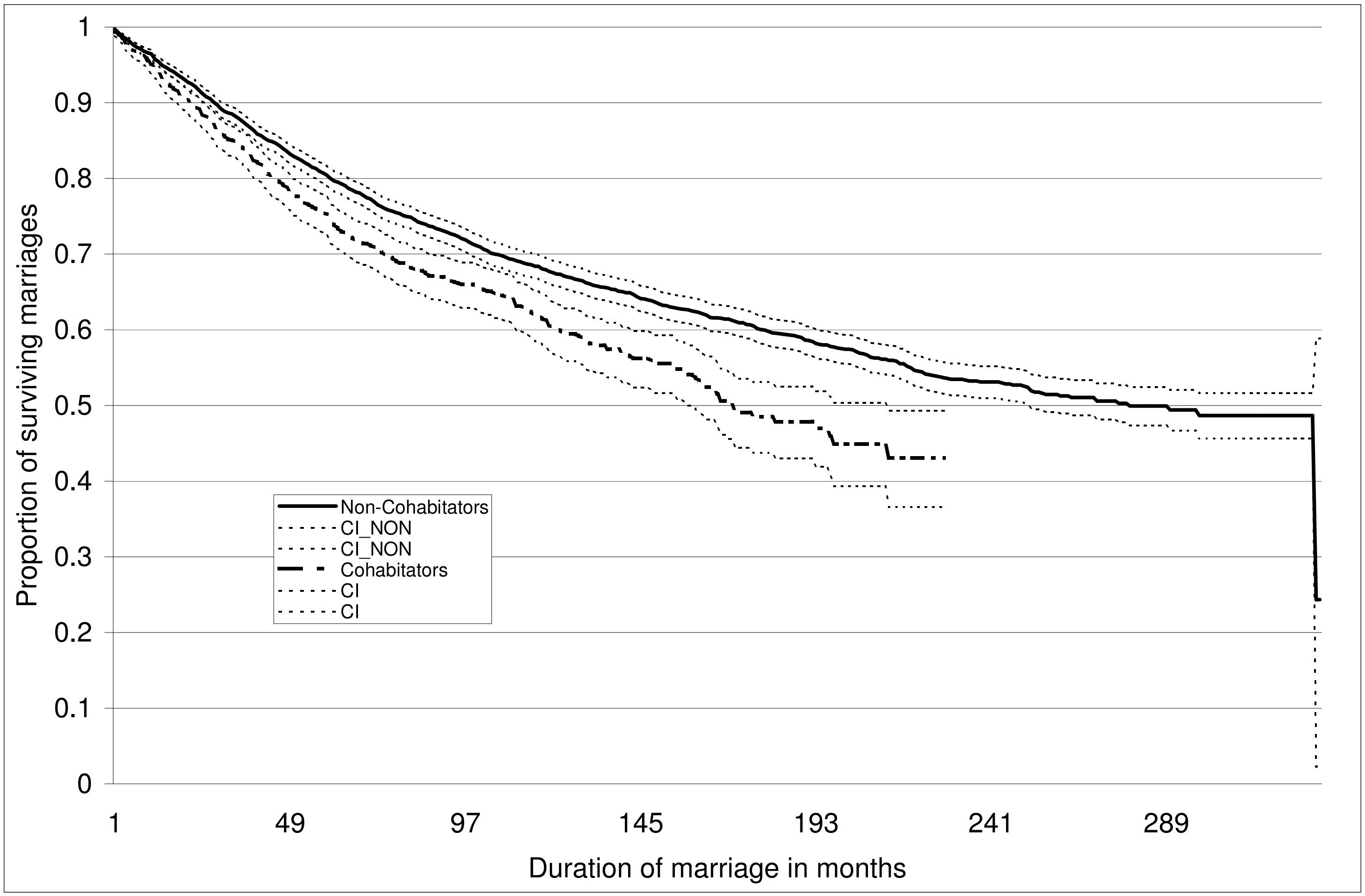


Figure 2: Survival Functions of First Marriages for Cohabitors and Non-Cohabitors and 95% Confidence Intervals (NSFG 1995)

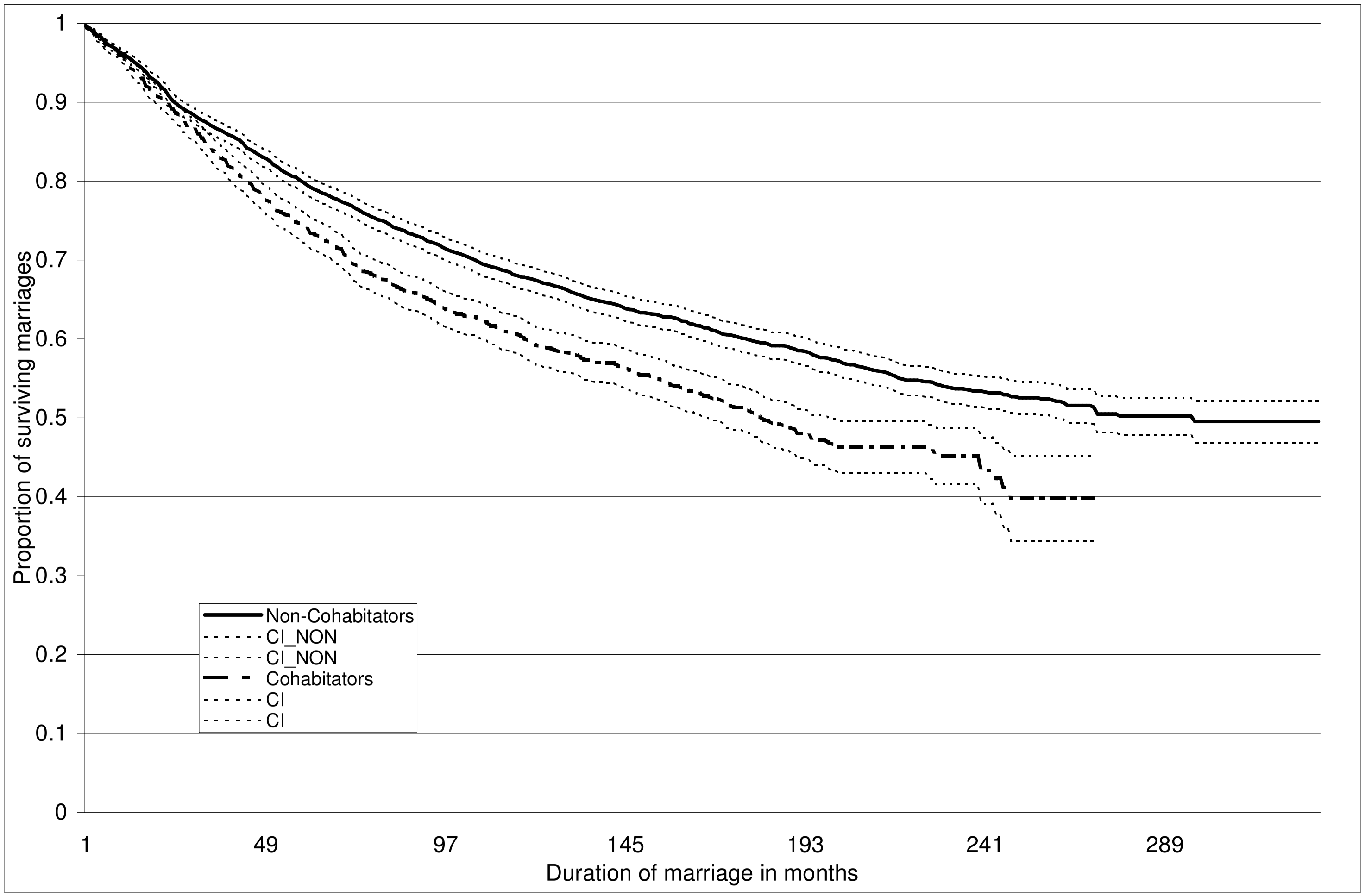


Figure 3: Survival Functions of First Marriages for Cohabitors and Non-Cohabitors and 95% Confidence Intervals (NSFG 2002)

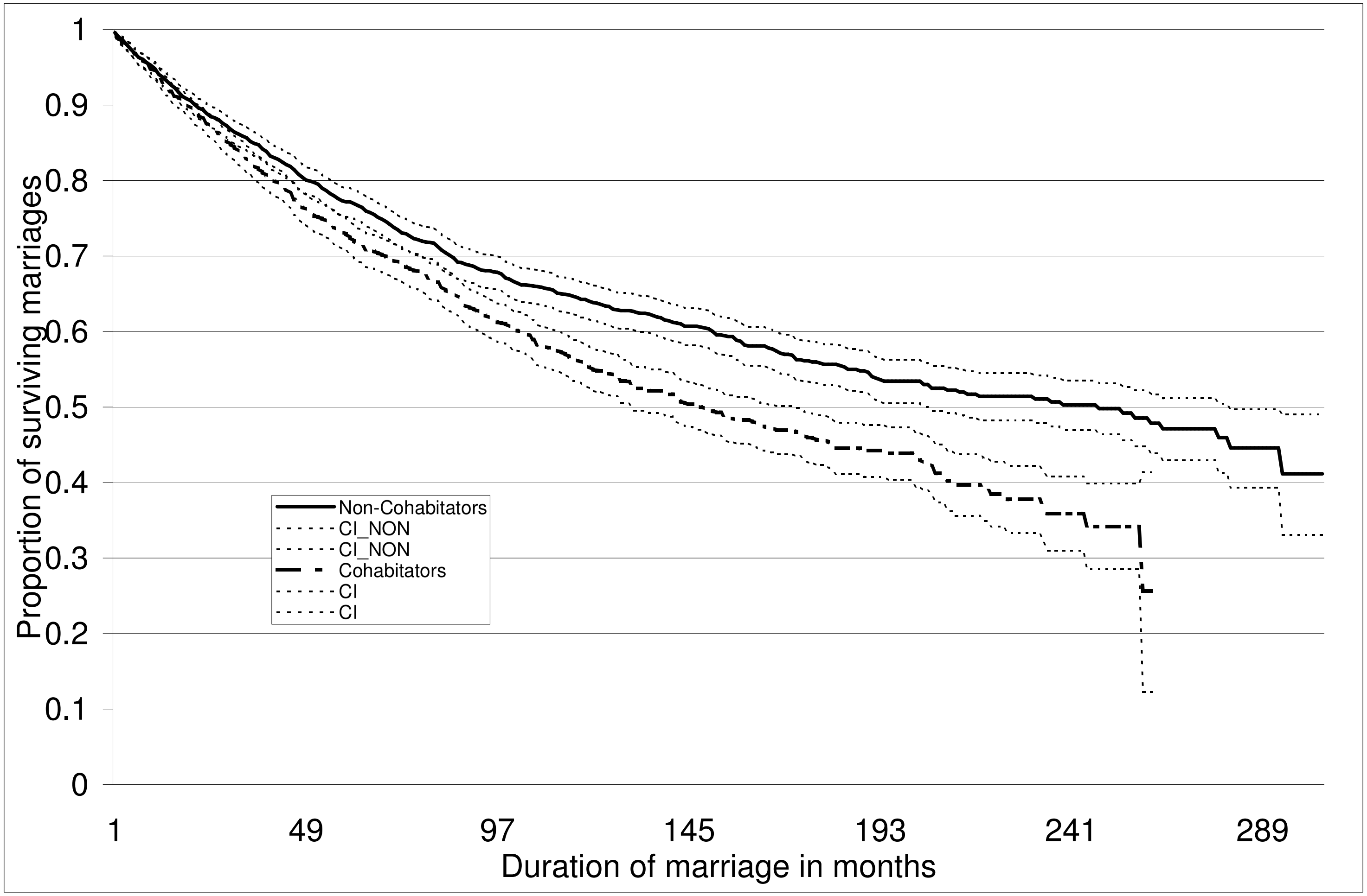


Figure 4: Hazard Ratio for Premarital Cohabitation and Age Cohorts

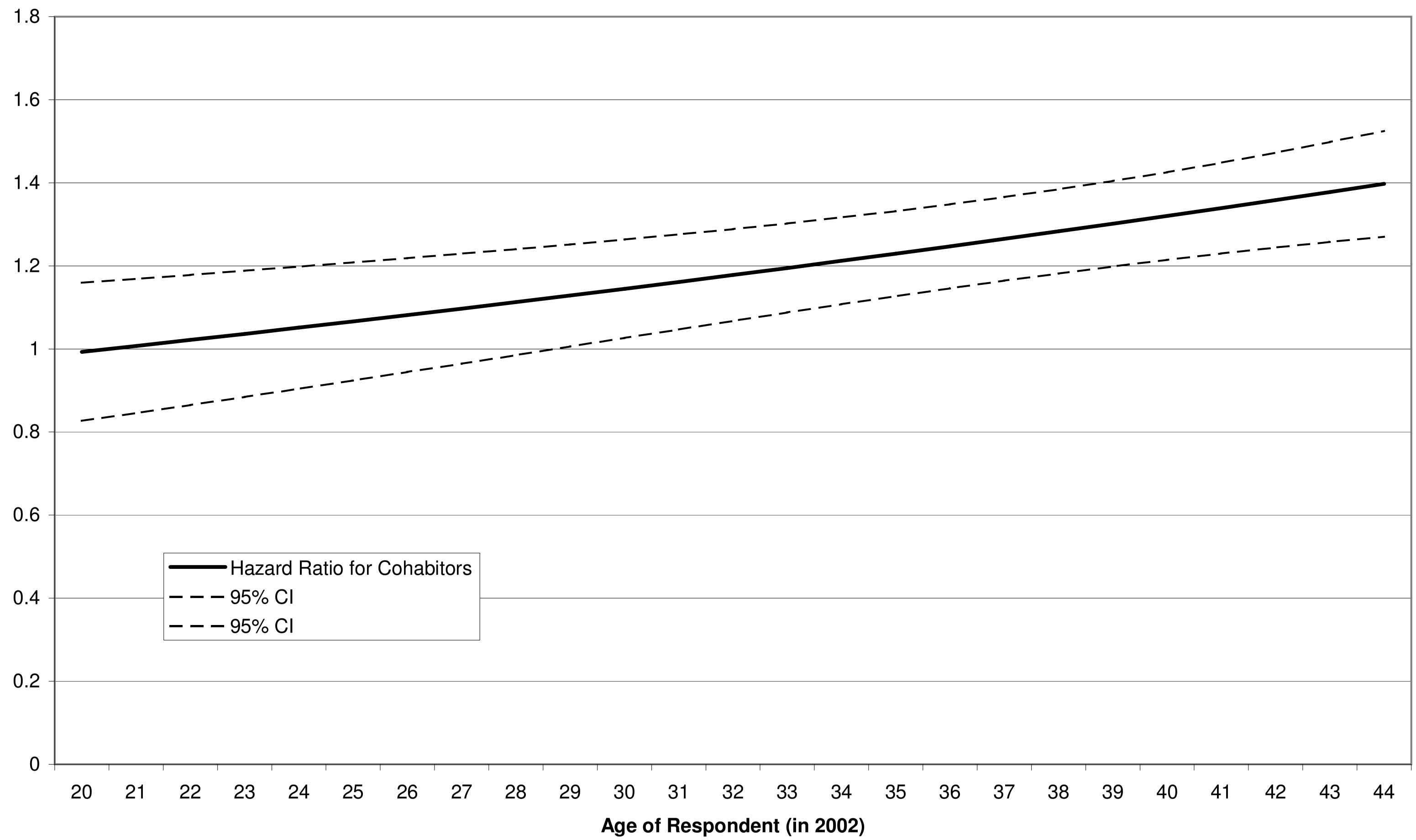
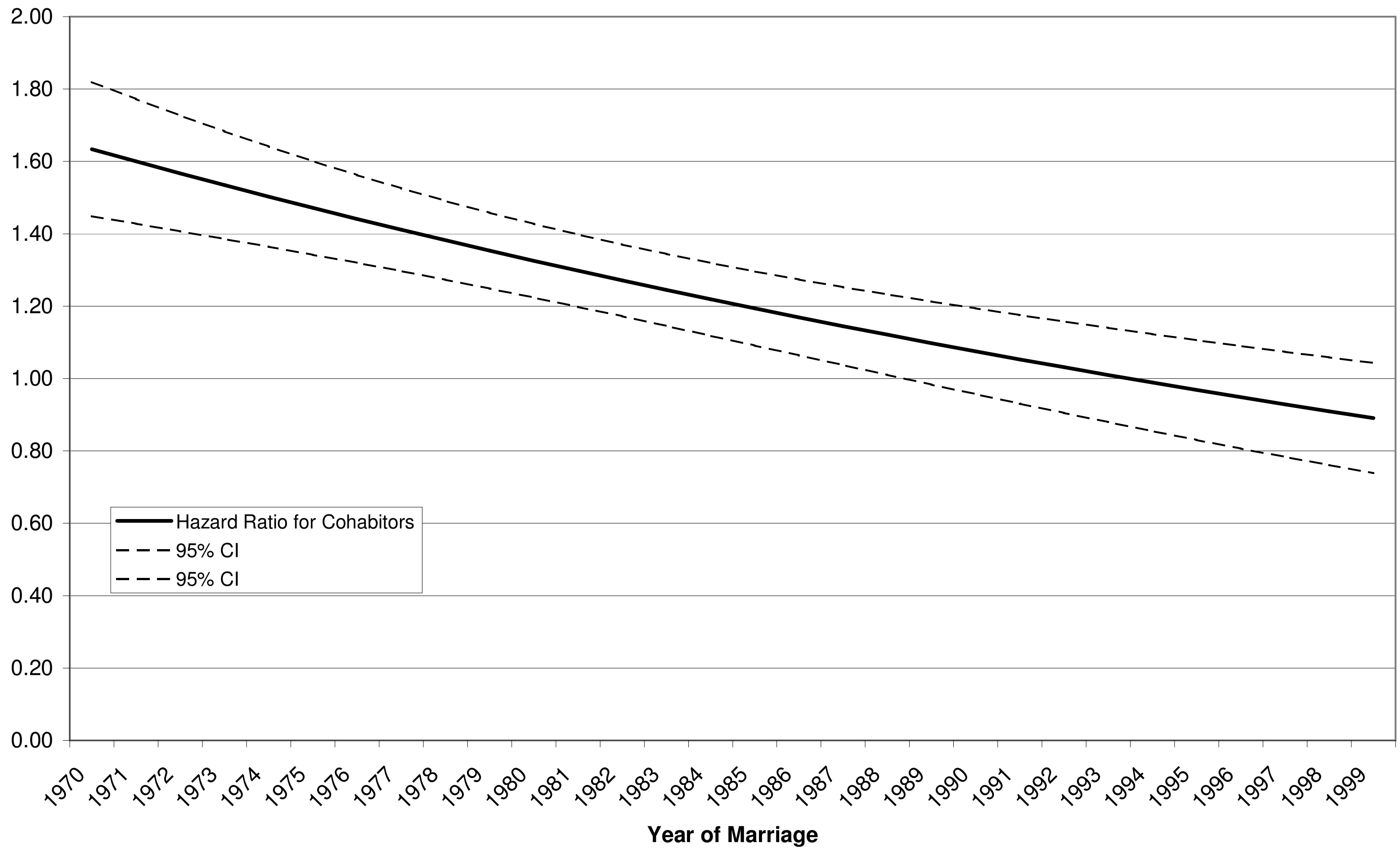


Figure 5: Hazard Ratio for Premarital Cohabitation and Marriage Cohorts



6 Appendix

This appendix includes several robustness checks for the NSFG 2002. Table 8 presents the results of logistic regressions where a dummy for having an imputed value for the date of marital dissolution is the dependent variable. This analysis shows a strong positive relationship between premarital cohabitation and the probability of having an imputed value. Table 9 shows proportional hazard regressions using the NSFG 2002 separately and including a dummy for having an imputed value. This decreases the coefficient on premarital cohabitation and points to a possible small upward bias when using all observations from the NSFG 2002. Notice that this bias would make it more difficult to find a decline in the relationship between premarital cohabitation and marital instability for more recent age and birth cohorts. For this reason, the estimates using all observations are conservative. Finally, in table 10 a comparison of the NSFG 1995 and 2002 is shown where the same birth cohorts are analyzed and the NSFG 2002 is artificially censored in 1995. One finds large differences in results between the two samples for the coefficients on religious affiliations and race. However, the coefficient on premarital cohabitation is not statistically different between the samples.

Table 8: Logistic Regression. Dep. Var. Dummy for Imputation of Date of Marital Dissolution

Variable	I	II	III
Cohabitation	0.445***	0.601	0.549***
	(0.112)	(0.581)	(0.151)
Cohabitation × age in 2002		-0.004	
		(0.016)	
Cohabitation × year of marriage			-0.015
			(0.014)
Education[a]			
No HS	0.398***	0.398***	0.397***
	(0.133)	(0.133)	(0.133)
High School	0.252**	0.251**	0.257***
	(0.125)	(0.125)	(0.125)
Fertility Var.			
Premarital conception	0.238***	0.238*	0.238*
	(0.131)	(0.131)	(0.131)
Premarital live-birth	0.589***	0.586***	0.596***
	(0.141)	(0.141)	(0.141)
Religion[b]			
Protestant	-0.387***	-0.386***	-0.393***
	(0.147)	(0.147)	(0.147)
Catholic	-0.609***	-0.609***	-0.613***
	(0.165)	(0.165)	(0.165)
Other	-0.541***	-0.541***	-0.541**
	(0.265)	(0.265)	(0.265)
Race[c]			
White	-0.772***	-0.771***	-0.768***
	(0.127)	(0.127)	(0.127)
Other	-0.336*	-0.334*	-0.341*
	(0.198)	(0.198)	(0.199)
Family Background			
No int. family	-0.118	-0.120	-0.116
	(0.110)	(0.111)	(0.111)
Wife's age at wedding	-0.103***	-0.102***	-0.059***
	(0.015)	(0.015)	(0.016)
Husband's age at wedding	0.062***	0.062	0.062***
	(0.009)	(0.009)	(0.009)
Wife's age in 2002	0.044***	0.046***	
	(0.008)	(0.013)	
Year of marriage			-0.036***
			(0.011)
Number of obs.	4038	4038	4038

Table 9: Proportional Hazard Regressions in the NSFG 2002. Dep. Var. Hazard of Separation of First Marriage

Variable	I	II
Cohabitation	1.123	0.982
	(0.096)	(0.084)
Dummy for Imputation		5.778***
		(0.526)
Education[a]		
No HS	0.919	0.839*
	(0.102)	(0.087)
High School	1.236**	1.172
	(0.120)	(0.116)
Fertility Var.		
Premarital	1.716***	1.582***
conception	(0.167)	(0.158)
Premarital	1.024	0.978
live-birth	(0.118)	(0.113)
Marital	0.726***	0.766***
live-birth	(0.065)	(0.068)
Religion[b]		
Protestant	0.824*	0.976
	(0.088)	(0.123)
Catholic	0.825	1.038
	(0.104)	(0.148)
Other	0.933	1.231
	(0.167)	(0.226)
Race[c]		
White	0.820**	1.028
	(0.085)	(0.102)
Other	0.648**	0.747*
	(0.112)	(0.116)
Family Background		
No int. family	1.385***	Dropped because
	(0.126)	of collinearity
Wife's age at wedding	0.906***	0.937***
	(0.013)	(0.013)
Husband's age at wedding	1.005	0.978***
	(0.008)	(0.008)
Number of obs.	4043	4043

Table 10: Proportional Hazard Regressions in the NSFG 1995 and 2002 for the same Birth Cohorts. Dep. Var. Hazard of Separation of First Marriage

Variable	NSFG 1995	NSFG 2002
Cohabitation	1.250*	1.046
	(0.141)	(0.113)
Education[a] No HS	1.052	0.923
	(0.168)	(0.131)
High School	0.980	1.327**
	(0.126)	(0.155)
Fertility Var.		
Premarital	1.401***	1.555***
conception	(0.183)	(0.191)
Premarital	1.047	0.957
live-birth	(0.169)	(0.158)
Marital	0.826	0.708***
live-birth	(0.112)	(0.085)
Religion[b]		
Protestant	0.735**	0.919
	(0.104)	(0.137)
Catholic	0.606***	0.896
	(0.100)	(0.152)
Other	0.587	0.952
	(0.191)	(0.263)
Race[c]		
White	0.646***	0.826
	(0.089)	(0.111)
Other	0.400***	0.725
	(0.124)	(0.172)
Family Background		
No int. family	1.371***	1.578***
	(0.148)	(0.178)
Wife's age at wedding	0.907***	0.926***
	(0.025)	(0.020)
Husband's age at wedding	0.978	0.995
	(0.017)	(0.011)
Number of obs.	1868	2663

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