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Cohabitation and Marital Stability in the United States*

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Abstract

Recent evidence from Canada and Sweden indicates that cohabitation prior to marriage significantly increases the risk of subsequent marital dissolution. In this article we present results testing the hypothesis that cohabitation increases marital disruption in the United States. We find that premarital cohabitation increases the risk of subsequent marital instability. However, the effect of cohabitation can be attributed to the fact that cohabitants have spent more time in union than noncohabitants. Once total length of union is accounted for, there is no difference in marital disruption between cohabitants and noncohabitants. We argue that subsequent research comparing cohabitants and noncohabitants with respect to marital behaviors that are duration dependent should account for the total amount of time spent in union.

The importance of cohabitation rests not only in its increased prevalence but also in its link to the process of family formation and dissolution. As the marital and fertility careers of young adults become more complex, it is important to understand the multiple determinants and consequences of variations in the timing and sequencing of important life-course events. In various theoretical and empirical reports, cohabitation has been linked to the rising proportion of out-of-wedlock births, delay in marriage, and increases in the divorce rate (see the review in Macklin 1987). In this article, we attempt to shed light on the debate surrounding the consequences of premarital cohabitation in the United States by examining data pertaining to subsequent marital stability.

Cohabitation and Marital Stability

Prior research and theory on the relationship between premarital cohabitation and subsequent marital stability has taken two basic positions. The first position argues that there is something about cohabitation, apart from the characteristics of cohabitants, that influences the success of marriage. As Bennett, Blanc, and

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Bloom (1988) indicate, marriage in the past was considered to be less of a personal bond between men and women than an economic arrangement designed for a gender-based division of labor as couples strove to meet the demands of community and economy. More recently, marriage has become less a social and economic exchange between spouses and more a source of personal gratification (Bellah et al. 1985; Blumstein & Schwartz 1985). There is now a more extended and well-defined period of courtship during which potential mates are sorted (Modell 1980).

In this context, it is possible to see premarital cohabitation as a form of "trial marriage" (Mead 1966; Rapoport 1965; Cherlin 1981). By cohabiting individuals are able to rehearse marital roles, strengthening the future marital bond. Strengthening may occur by developing a mutually satisfactory division of household and market labor, as well as enhancing sexual and interpersonal compatibility. At the same time, if appropriate and satisfactory role relationships cannot be developed, the partners are able to end the relationship before becoming subject to the additional constraints of the legal system. Following the logic of this position, one would hypothesize that cohabitation should lead to more stable marriages.

Available evidence on the nature of cohabiting relationships provides indirect support for this position. Recent evidence suggests that cohabiting unions are less stable than marriages (Bumpass & Sweet 1988; Hofferth & Upchurch 1988), consistent with the notion that only the most compatible unions survive to marriage. Other evidence indicates that most cohabitants organize their unions much like married couples, although a gender-based division of labor is slightly less prominent (Macklin 1987). This suggests that individuals have the opportunity to practice and adapt to marital roles.

The second position is different from the first because it assumes that cohabitants are fundamentally different from noncohabitants. Specifically, it is asserted that cohabitants are a select group of individuals who are less committed to marriage than noncohabitants. It is not the experience of having cohabited itself that affects marital stability. Rather, it is a preexisting disposition toward lesser commitment on the part of cohabitants that influences subsequent marital stability. Thus, cohabitation is not viewed as a stage in the courtship process by which couples become more enmeshed. Instead, it is an alternative, albeit sometimes temporary, to marriage. Among cohabitants, the decision to marry is more likely to be the result of pressures from family and peers. Following the logic of this position, one would hypothesize that cohabitation leads to less stable marriages.

There exists indirect evidence supporting the thesis that cohabitants are less committed to relationships than noncohabitants. The data cited above concerning the higher dissolution rates of cohabiting versus marital unions may be construed as support for the thesis that cohabitants are less committed to stable relationships. Data from both abroad and from the United States also indicates that individuals who cohabit are less likely to hold traditional sex roles, are more likely to value individual freedom, and are less likely to desire children (Carlson 1986; Clayton & Voss 1977; Newcomb & Bentler 1980; Tanfer 1987). On the other hand, recent data from the United States does not indicate cohabitants are any less likely to desire marriage than noncohabitants (Tanfer 1987).

Cohabitants may also vary according to other characteristics that affect marital stability. For example, prior research suggests that cohabitants are more likely to have been married before and are more likely to have children at the beginning of marriage (Macklin 1987; Tanfer 1987), both characteristics that may lead to an increased risk of marital disruption. It is not the case, however, that all differences in characteristics would act to increase the likelihood of marital dissolution for cohabitants. For instance, cohabitants are likely to be older at marriage than noncohabitants, and an older age at marriage has been consistently linked to a reduced probability of divorce (Cherlin 1981).

Available empirical evidence linking cohabitation to marital dissolution is scarce. Only recently have appropriate data become available to test the competing hypotheses outlined above. Using data from the Canadian Fertility Survey conducted in 1984, Balakrishnan et al. (1987) find that cohabitation significantly increases the likelihood of subsequent marital disruption. Couples who cohabited before marriage had rates of marital dissolution 50% greater than noncohabitants. Using data from a 1981 Swedish study, Bennett, Blanc, and Bloom (1988) present similar findings. They report that cohabitation increases the rate of subsequent marital disruption by 80%.

Bennett, Blanc, and Bloom (1988:137) state that "the direction of effects found in Sweden in all likelihood holds true in the United States as well." Although the use of data from Sweden has been questioned with respect to its applicability to the United States (Cherlin 1981), similar results from Canada strengthens support for the hypothesis that premarital cohabitation impacts negatively on marital stability. We perform what we believe to be the first empirical test of the effect of premarital cohabitation on marital stability using representative data from the United States.

Data

We take our data from the fifth follow-up to the National Longitudinal Study of the High School Class of 1972 (NLS) conducted in 1986. The NLS is "a stratified two-stage probability sample of persons who were students in U.S. schools that contained 12th graders in the 1971-72 academic year" (Tourangeau et al., 1987:8). The fifth follow-up is a probability sample (N=12,841) of respondents who participated in the 1972 baseline survey or one of the four earlier follow-ups (conducted in 1973, 1974, 1976, and 1979). The fifth follow-up to the NLS contains a relationship history for each respondent with the beginning and ending dates of the first three "intimate," live-in unions with an unrelated adult of the opposite sex.[1] Living arrangements that lasted less than one month were not counted. For each union, it was ascertained whether the couple was married at the time they started living together or if they married prior to the end of the relationship.

Because we are interested in the impact of premarital cohabitation on marital dissolution, we select a subsample of first marriages for ever-married respondents. Using the dates (in terms of day, month, and year) contained in the relationship history, we define cohabitants as respondents who lived with their spouse for more than one month prior to marriage. Consistent with most

TABLE 1: Prevalence and Duration of Premarital Cohabitation in the NLS Sample by Sex

	Women	Men
Cohabited with spouse	20.7%	22.9%
Cohabited more than once	2.6%	2.4%
Duration of cohabitation with spouse:		
0-6 months	29.7%	26.5%
7-12 months	25.0%	23.2%
13-24 months	21.5%	26.5%
> 24 months	23.8%	23.8%

recent literature on marital dissolution, we define the date of marital disruption as the date the couple stopped living together rather than the date of divorce (Becker, Landes & Michael 1977; Menken et al. 1981; Morgan & Rindfuss 1985; Teachman 1982).[2]

Most studies of marital dissolution focus only on women, and the available research on the effects of cohabitation on marital and union disruption is all restricted to reports from women. The NLS contains information on relationship histories ascertained from both men and women. We take advantage of the NLS data to obtain and compare results by sex. Although the NLS contains data on both blacks and whites, we restrict our attention to whites due to the small sample size for blacks.[3] Our final sample size is 4,354 for women and 3,837 for men.

The sample sizes reported in the text and in the tables refer to the actual number of men and women on which various analyses are performed. However, since the NLS is the result of a complex sampling strategy, all descriptive statistics and life table values are based on weighted data. The weights adjust for different probabilities of being selected into the sample and for variations in nonresponse.[4] For the multivariate models, use of weighted data violates the asymptotic theory upon which tests of significance for parameter estimates are based. Fortunately, both weighted and unweighted data provide basically similar parameter estimates. We choose, therefore, to use unweighted data for the multivariate analysis in order to obtain approximate test statistics.

We emphasize that the NLS is representative only of all 12th graders in 1972 (almost all of whom graduated). Earlier and later cohorts are not represented, nor are individuals who do not make it to their senior year in high school. The latter restriction is important due to the extent to which individuals with more education behave differently from individuals with less education with respect to cohabitation and marriage. Because we have no basis on which to evaluate whether selectivity according to education leads to bias, our results must be interpreted as being conditional on having graduated from high school. It remains possible that the relationship between cohabitation and marital stability is different for individuals with less than a high school degree.

TABLE 2: Life Table Estimates of the Cumulative Proportion of Marriages Disrupted by Sex and Premarital Cohabitation Status

Months Since	Women		Men	
Marriage	Cohabit	Not Cohabit	Cohabit	Not Cohabit
12	.028	.031	.041	.020
24	.067	.069	.073	.050
36	.103	.103	.105	.079
48	.139	.133	.149	.122
60	.166	.167	.173	.148
72	.198	.194	.222	.176
84	.262	.225	.258	.197
96	.289	.247	.295	.214
108	.320	.271	.326	.230
120	.351	.290	.379	.243
N[a]	(972)	(3,382)	(895)	(2,942)

[a] Values in parentheses are unweighted sample sizes

Descriptive Results

The proportion of men and women who have cohabited with their spouse prior to marriage is shown in Table 1. Men and women are about equally likely to report having cohabited (21-23%). The values observed for women are roughly congruent with those reported for Canadian women (Balakrishnan et al. 1987). However, the level of cohabitation is much lower than observed for Sweden. Bennett, Blanc, and Bloom (1988) report that almost two-thirds of the ever-married women in their Swedish sample had cohabited. The proportion of cohabitants who cohabited more than 6 months is about the same for both sexes (70-73%). In Sweden, although rates of premarital cohabitation are higher, about the same percent of women (76%) have cohabited for more than 6 months. Only a small fraction (2-3%) of ever-married respondents in the United States have cohabited more than once prior to first marriage.

We use a life table procedure to examine variation in marital disruption, because cohabitants and noncohabitants may have different periods of risk to the event. Life tables yield estimates of the probability that a respondent has ended their marriage at each marital duration (which we calculate in terms of whole months). Individuals who end their marriages contribute exposure at each duration until the time of marital dissolution. Individuals who remain married contribute exposure at each duration until they are truncated by the survey.

The cumulative proportion of respondents ending their marriage, by sex and premarital cohabitation status, is presented in Table 2. The results are consistent with results from Sweden and Canada. Although the differences are small in the first few years following marriage, after 10 years cohabitants are more likely than noncohabitants to have dissolved their marriage. The level of marital dissolution registered for cohabiting women in the United States is only slightly higher than observed for Canadian women (10-year cumulative failure rate of .35 versus .31) but nearly twice that observed for Swedish women (10-year cumulative failure rate of .35 versus .18). The level of marital disruption experienced by non-cohabiting women in the United States is particularly high compared to Canada and Sweden (after 10 years the cumulative failure rates are .29 for the United States, .14 for Canada and .10 for Sweden). Our confidence in these results is strengthened by the fact that the same pattern appears for both men and women.

The life table results are consistent with the hypothesis that cohabitants are less committed to marriage. These results, however, do not control for variations between cohabitants and noncohabitants on variables known to affect the risk of marital disruption (e.g., age at marriage, education). The values shown in Table 3 indicate that cohabitants possess characteristics that act to both increase and decrease the likelihood of marital dissolution. Several researchers have found that having a premarital birth (but not a legitimated birth) increases the likelihood of subsequent marital dissolution (Billy, Landale & McLaughlin 1986; Menken et al. 1981; Morgan & Rindfuss 1985; Teachman 1982, 1983). Cohabitants are more likely than noncohabitants to have a premarital birth, while there is little difference between the two groups with respect to legitimated births. Maritally conceived births have been linked consistently to a reduced likelihood of marital disruption (Billy, Landale & McLaughlin 1986; Morgan & Rindfuss 1985; Waite & Kanouse 1985). Cohabitants are less likely than noncohabitants to have a maritally conceived birth. The fertility careers of cohabitants therefore indicate greater risk of marital dissolution.

The marital structures of cohabitants are more complex than those of noncohabitants in that: a) their spouses are more likely to have been married before (unfortunately, the NLS does not contain information on prior cohabitation for spouses), and b) they are more likely to have step-children living in the home (the greater percent in this category registered for men reflects the cultural bias toward giving women custody of children). Several authors have argued that more complex marital structures are more prone to marital instability, either because role performance is less institutionalized in such unions (Cherlin 1978), or because these unions are selective of individuals less committed to marriage in general (Halliday 1980). The empirical evidence shows that complex marital structures are associated with greater marital instability (McCarthy 1978; White & Booth 1985). As was the case for fertility careers, the marital structures of cohabitants indicate a greater risk of experiencing marital disruption.

The most consistent and strongest predictor of marital dissolution is age at marriage (Billy, Landale & McLaughlin 1986; Menken et al. 1981; Morgan & Rindfuss 1985; Teachman 1982, 1983). Individuals who marry early are much more likely to end their marriages than individuals who marry later. The data in Table 3 indicate that cohabitants marry later than noncohabitants. Cohabitants are also more highly educated than noncohabitants and come from higher SES backgrounds. Both factors have been found to decrease the likelihood of marital

TABLE 3: Sample Means of Predictor Variables[a]

	Women		Men	
Variable	Cohabit	Not Cohabit	Cohabit	Not Cohabit
Maritally conceived birth (%)	68	83	66	78
Premarital birth (%)	8	4	13	4
Legitimated birth (%)	10	11	10	9
Spouse married before (%)	29	12	26	7
Step-child in house (%)	5	2	14	4
Age at marriage	24.1	21.5	25.6	23.8
Education (%):				
Less than high school	4	4	2	2
High school	32	48	35	43
Some college	37	28	37	31
College degree	24	18	21	20
More than college	4	2	5	4
SES(%):				
Low	18	22	13	19
Medium	50	58	52	56
High	32	20	35	25

[a] Maritally conceived birth is measured as a time-dependent variable with 1=birth occurring after the first seven months of marriage, 0=otherwise; Premarital birth is measured as 1=birth occurring before marriage and living in the house, 0=otherwise; Legitimated birth is measured as 1=birth occurring in the first seven months of marriage, 0=otherwise; Spouse married before and step-child in the house are dummy variables where 1=spouse previously married, 0=otherwise, and 1=spouse has a child from another relationship living in the household, 0=otherwise; Age at marriage is measured in years; Education refers to level of education measured at the time of marriage where 1=less than high school, 2=high school, 3=some college, 4=college graduate, and 5=more than college; SES is a composite variable based on characteristics of the parental household (father's education, mother's education, father's occupation) and is divided into three approximately equal groups with 1=low, 2=medium, and 3=high.

dissolution (Billy, Landale & McLaughlin 1986; Menken et al. 1981; Morgan & Rindfuss 1985; Teachman 1982, 1983). Thus, variations on age at marriage, education, and SES, contrary to other characteristics distinguishing the two groups, indicate a reduced risk of marital instability for cohabitants versus noncohabitants.

TABLE 4: Effects of the Predictor Variables on Marital Dissolution, Measured from Beginning of Marriage, by Sex[a]

	Women	Men
Cohabited	.187*	.139*
Cohabited more than once	.023	.366*
Maritally conceived birth	-.277*	-.298*
Premarital birth	.065	-.042
Legitimated birth	-.199*	-.203*
Age at marriage	-.208*	-.146*
Education at marriage	-.187*	-.298*
Spouse married before	.327*	.347*
Step-child in house	.119	.007
SES	.055*	.048
Intercept	-2.029*	-1.601*
Model X^2	26476	23184
df	14	14

*The coefficient is at least twice its standard error.

[a] Maximum likelihood logistic-regression coefficients, net of marital duration

Multivariate Results

To conduct a multivariate analysis in a fashion that accounts for variation in duration of exposure to marital disruption, we first construct a set of yearly intervals for which respondents are at risk of marital dissolution. We do this separately for men and women. Individuals contribute intervals, starting at marriage, until the marriage is dissolved or the date of the survey, whichever comes first. In each interval, the dependent variable is a dichotomy — dissolution occurs or it does not. We then pool these intervals into a larger sample for each sex.[5] By using pooled intervals, we are in effect estimating a discrete-time approximation to a continuous-time hazard rate model (Allison 1982).[6]

The net effect of cohabitation on marital dissolution by sex is obtained by fitting the following logistic-regression equation to the data in each of the pooled samples:

$$\mathrm{Ln}[P/(1\text{-}P)] = a + b_1X_1 + b_2X_2 + X_3B_3 + X_4B_4 \qquad (1)$$

where P is the probability that a marriage ends, a is the intercept, X_1 is cohabitation status, b_1 is the effect of cohabiting on the likelihood of marital instability; X_2 indicates having cohabited more than once, b_2 is the effect of having cohabited more than once on marital disruption; X_3 is a vector of control variables, B_3 is a vector of coefficients indicating the effects of the control variables on marital dissolution; X_4 is a set of dummy variables indicating elapsed duration from the beginning of marriage, and B_4 is a set of coefficients indicating whether the likelihood of marital disruption shifts at each marital duration. Since Equation 1 is based on individual-level data, we employ a maximum likelihood procedure for estimation (Hanushek & Jackson 1977).

The coefficients obtained from fitting the logistic-regression equations are shown in Table 4. We use the control variables outlined in Table 3.[7] Having a maritally conceived birth is measured as a time-dependent variable, taking the value zero in each interval until a birth occurs and then taking the value one in that and all subsequent intervals. The values of all other independent variables are fixed at the point of marriage. We include a control for whether the respondent had cohabited previously with someone other than their spouse. We assume that respondents who cohabit more than once prior to marriage are more likely to possess unmeasured characteristics (such as lower commitment) that reduce the success of their unions.

Similar to the findings for Canada and Sweden, premarital cohabitation increases the likelihood of subsequent marital instability for both men and women (the logistic regression coefficients imply that the odds of dissolving a marriage in any given interval are 1.21 and 1.15 times greater for women and men, respectively, who cohabited with their spouse).[8] Having cohabited more than once significantly increases the probability of marital disruption for men but not for women. We note, however, that marrying someone who had been married before significantly increases the risk of marital disruption for both sexes.

The coefficients for the remaining control variables are largely in line with expectations based on prior research. Compared to childless couples, having either a legitimated or a maritally conceived birth significantly reduces the chances of marital disruption. Somewhat surprisingly, having a premarital birth does not have a significant effect on marital stability.[9] Respondents who marry older and who have more education at marriage are less likely to experience marital disruption. Other than the effect of having cohabited more than once, the only difference between men and women in the models estimated is that coming from a higher SES background increases the likelihood of marital disruption for women but not men.[10] Having a step-child in the home does not affect the likelihood of marital dissolution for either sex.[11]

We tested for time dependence in the effects of both cohabitation and the control variables. We found no evidence that the effects of any of the variables change significantly over the marital durations observed in the NLS sample. We note, however, that the distribution of marital duration available in the NLS is restricted to shorter marriages, and it is possible that time dependence would appear at longer durations.

The results shown in Table 4 are consistent with the hypothesis that cohabitants are less committed to marriage than noncohabitants. To explore this proposition in greater depth, we follow Bennett, Blanc, and Bloom (1988) who suggest that long-term cohabitation indicates less commitment to marriage than short-term cohabitation. If this is true, then short-term cohabitants should be less likely to experience marital dissolution. To test this proposition, we reran the logistic-regression models including a control for length of premarital cohabitation (coded as a dichotomy — cohabitation of six months or less versus

TABLE 5: Effects of Premarital Cohabitation and Length of Premarital Cohabitation on Marital Dissolution, Measured from Beginning of Marriage, by Sex[a]

	Women	Men
Cohabited	.287*	.205*
Cohabited more than once	.036	.382*
Cohabited 6 months or less	-.427*	-.294*

*Coefficient is at least twice its standard error.

[a] Maximum likelihood logistic-regression coefficients, net of marital duration and control variables indicated in Table 4

cohabitation of more than six months). The results are shown in Table 5 (since the coefficients for the control variables are virtually unchanged we do not show them).[12] We find that short-term cohabitants are less likely to experience marital dissolution than long-term cohabitants (the logistic regression coefficients imply that the odds of dissolving a marriage for those who cohabited 6 months or less are .65 and .75 times as great for women and men, respectively, in any given interval).

While the length of premarital cohabitation may indicate variation in commitment to marriage, it is also possible that this effect reflects something very different. Specifically, couples who have been together longer, irrespective of marital status, may be more likely to dissolve their unions. For example, respondents who cohabited for two years prior to marriage and have been married for one year may have rates of marital dissolution similar to respondents who have been married for three years. Thus, the negative effect observed for short-term cohabitation may be due to the fact that these unions are still new at the time of marriage.

To test this alternative hypothesis, we estimate the hazard-rate models shown in Equation 1 using total duration in first union. Since premarital cohabitants in our sample cannot, by definition, end their unions before marriage, we censor their relationship histories before the time of marriage. For instance, a respondent who cohabited for 2 years prior to marriage is at risk of marital dissolution in the third year of union and beyond, but not before. The results from conducting this exercise are shown in Table 6. The coefficients indicate no effect of premarital cohabitation on the rate of marital dissolution, although having cohabited more that once is related to an increased likelihood of marital disruption for both men and women. In other words, the increased rate of marital dissolution associated with premarital cohabitation indicated in Tables 4 and 5 can be attributed to the amount of time cohabitants have already spent in union prior to marriage.[13] On the basis of these results, neither of the original hypotheses is supported.

TABLE 6: Effects of Premarital Cohabitation on Marital Dissolution, Measured from Beginning of Union, by Sex

	Women	Men
Cohabited	.032	-.077
Cohabited more than once	.546*	.704*

*Coefficient is at least twice its standard error.

[a] Maximum likelihood logistic-regression coefficients, net of marital duration and control variables indicated in Table 4

Discussion

American men and women are subject to considerable delay in marriage and high divorce rates. Age at first marriage is now higher than at any point since the turn of the century (Espenshade 1987), and estimates indicate that as many as 50% of recently contracted marriages will end in divorce (Weed 1980). In a typical American's life span, marriage has come to occupy an ever smaller proportion of adult years (Watkins, Menken & Bongaarts 1987; Schoen et al. 1985). In such a context, it is likely that cohabitation represents an advanced stage of courtship where marriage is desired but with less commitment to a lifetime partnership (Gwartney-Gibbs 1986; Tanfer 1987). While cohabitation, as a stage in the courtship process, may impact on the nature and timing of marital unions, our results indicate that cohabitation has no direct effect on the stability of marriage.

Our results show that net of important compositional controls, if consideration is given to the amount of time spent in marriage, premarital cohabitation in the United States increases the risk of marital dissolution. However, once allowance is made for the total amount of time spent in union, there is no difference in the rate of marital disruption by cohabitation status. While cohabiting unions remain less stable than marriages (Bumpass & Sweet 1988; Hofferth & Upchurch 1988), the stability of marriages preceded by cohabitation, compared to marriages not preceded by cohabitation, does not indicate less commitment to marriage. Similarly, having cohabited does not confer advantage in terms of marital stability.

To increase confidence in our results, subsequent research should consider more broadly representative samples. As discussed above, we cannot be sure that our results apply to individuals with less than a high school education. Other research indicates that these individuals are more likely to cohabit that individuals with more education (Tanfer 1987) and to experience marital disruption (Morgan & Rindfuss 1985), making it important to determine the impact of cohabitation on their marital stability. It would also be informative to obtain information about the impact of cohabitation on marital stability in different historical contexts by considering earlier and later cohorts of individuals.

Notes

1. The exact question is, "Please start with the first time you got married or lived in an intimate relationship with an unrelated adult of the opposite sex. Do not count any living arrangements which lasted less than one month."

2. Date of separation is used because many marital unions cease to exist as a functioning unit at the time of separation rather than divorce. At separation, individuals often reenter the marriage market, relocate to new communities, and otherwise consider themselves to be "single."

3. There are 566 black women and 316 black men in the NLS data, but only 134 of the women and 78 of the men reported having cohabited.

4. The weights used inflate estimates to the population totals for all high school seniors in 1972.

5. Note that the unit of analysis in this case is the yearly interval. Pooling of intervals leads to the possibility of serially correlated errors, resulting in underestimated standard errors. We tested for serial correlation using ordinary least squares and the Durbin-Watson statistic and found no evidence to suggest a problem. We also estimated our model on each of the intervals separately. While there is some fluctuation in the size of the coefficients and statistical significance across intervals, our substantive conclusions are not changed.

6. We conducted many of our preliminary analyses using a continuous-time hazards model. Our final model takes the form of a discrete-time approximation because of the computational ease with which it can incorporate time-varying predictor variables (see Allison 1982). We note that in all cases where we made a comparison, the results obtained using a continuous-time model were substantively identical to results obtained using a discrete-time model (see also note 8).

7. We control for marital duration by including a set of four dummy variables indicating successive two-year intervals — except the last category, which is open-ended. The assumption is that the likelihood of marital disruption is constant across each two-year period. We estimated a model with a set of dummy variables for yearly intervals and found basically the same results. We do not consider separately marital durations longer than ten years for two reasons. First, since the NLS follows individuals from 1972, their senior year in high school, to 1986, most marriages are less than 10 years in duration. Second, there is an increasing bias toward younger marriages at longer durations.

8. In order to make a more direct comparison between our results and those presented for Canada and Sweden, we reestimated the hazard-rate models using a parametric (Weibull) version of a continuous-time model (excluding marital births, which is a time-varying variable). The resulting coefficients indicate that cohabitants have rates of marital dissolution 35-37% higher (for men and women, respectively) than noncohabitants. While significant, these values are smaller than observed for Canada (50%) and Sweden (80%).

9. The lack of an effect for having a premarital birth may be due to the fact that our sample is more highly educated than the United States population as a whole. Another possibility is that we control for a different set of variables (e.g., cohabitation, previous marital status of spouse) than used in other studies that find an effect of having a premarital birth.

10. This relationship is consistent with a similar finding reported by Bennett, Blanc, and Bloom (1988). They interpret the effect as being due to the greater likelihood of higher SES women having mothers who worked outside the home and who had themselves experienced marital dissolution.

11. Note that step-children refers to the prior children of one's spouse. It does not refer to children (in our sample, necessarily born out of wedlock) born to the respondent in a prior relationship. Such births are included in the premarital birth category. As noted in Table 3, we exclude premarital births not living in the household.

12. As one reader noted, the choice of six months as a cutoff between short- and long-term cohabitation is arbitrary. We chose six months because: a) about one-quarter to one-third of all cohabitations last six months or less, b) Bennett et al. (1988) use six months as a similar cutoff, and c) preliminary results indicated relatively similar results for various cutoffs up through a

full year of cohabitation (beyond which we are hesitant to call cohabitation short-term). Moreover, this analysis is really a precursor to the more detailed analysis that follows.

13. Bennett, Blanc, and Bloom (1988) conduct the same exercise for their Swedish data but still find a significant effect of cohabitation on marital instability.

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