SPOUSAL ALTERNATIVES AND MARITAL DISSOLUTION*

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Using three data sources we explore the effects of the quantity and quality of potential new marital partners available in local marriage markets on the risk of marital dissolution. Data from the National Survey of Families and Households demonstrate that, among recently-divorced men and women, a substantial percentage had been romantically involved with someone other than their spouse prior to divorcing. Merging microlevel data from the National Longitudinal Survey of Youth with aggregated Public Use Microdata from the 1980 U.S. Census, we examine the impact of marriage market characteristics and other contextual variables on the risk of marital dissolution, net of individual-level predictors. Proportional hazards models reveal that, among non-Hispanic Whites, the risk of dissolution is highest where either wives or husbands encounter an abundance of spousal alternatives. Increased labor force participation among unmarried women and high geographic mobility rates in the local area also increase marital instability. Our results suggest that many persons remain open to alternative relationships even while married, and that the supply of spousal alternatives in the local marriage market significantly increases the risk of marital dissolution.

The transformation of the American family is perhaps nowhere more evident than in the rapid rise in divorce since World War II, and few research areas have attracted as much scholarly attention (Cherlin 1992). Trends in divorce and separation have been widely discussed (Martin and Bumpass 1989; Morgan and Rindfuss 1985; Thornton and Rodgers 1987), and the many social and demographic characteristics of persons and couples that influence the probability of marital disruption have been documented extensively (see White 1990, for a recent review). Among the most important individual-level risk factors for divorce are age at marriage (Booth and Edwards 1985; Bumpass, Martin, and Sweet 1991), marital and premarital childbearing (Billy, Landale, and McLaughlin 1986; Waite and Lillard 1991), race (Menken, Trussell, Stempel, and Babakel 1981), wife's education and employment (Greenstein 1990; Spitze and South 1985), home ownership (Becker, Landes, and Michael 1977), and premarital cohabitation (Bennett, Blanc, and Bloom 1988). Marital happiness (Booth, Johnson, White, and Edwards 1986) and conjugal interaction (Booth, Johnson, White, and Edwards 1985) also play important roles and may serve to mediate the influences of various sociodemographic factors on divorce (White 1990).

While studies of risk factors have contributed significantly to our knowledge of divorce differentials, they have focused exclusively on characteristics of couples or the individual marriage partners, paying little or no attention to features of the social context that might influence marital disruption. Thus, possible contextual influences on marital dissolution have gone largely unexplored. Foremost among the contextual characteristics that might influence divorce is the availability of numerous, desirable al-

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ternatives to one's current spouse. Although no study has yet linked the available pool of spousal alternatives to divorce probabilities, suggestive evidence can be found in two areas of research. First, several recent studies have documented an association between marriage opportunities (as reflected in the age/sex composition of local populations) and actual marriage rates (Lichter, LeClere, and Mclaughlin 1991; Lichter, McLaughlin, Kephart, and Landry 1992; South and Lloyd 1992). Given the high remarriage (and cohabitation) rates among divorced persons (Cherlin 1992) and the evidence (presented below) that many divorced persons were romantically involved with someone other than their spouse prior to divorcing, it seems plausible that a surfeit of spousal alternatives may draw people away from their current marriage. The availability of desirable alternatives to one's present spouse may be particularly important for individuals in unhappy marriages. Second, White and Booth (1991) and Udry (1981) have reported that married persons who perceive a high likelihood of remarrying should they divorce are more likely than others to dissolve their marriages. Again, this suggests that the availability of spousal alternatives may attract persons away from their current marriage partners, independent of the many other factors that motivate individuals to divorce. Neither of these two studies, however, attempts to directly ascertain the impact of the supply of spousal alternatives in the local marriage (or remarriage) market on the probability of divorce.

We examine the effect of the *quantity and quality of spousal alternatives*, as well as other characteristics of the local context, on the probability of marital dissolution. In the primary analyses, marital history data from the National Longitudinal Survey of Youth (NLSY) are merged with Census data describing the quantity and quality of alternative spouses in each respondents' local geographic area. We then use proportional hazards regression models with time-varying explanatory variables to determine the impact on marital dissolution of these marriage market characteristics and of other features of the local context, net of the conventional individual-level (and couple-level) predictors of divorce.

THEORY

Although few studies have directly examined the influence of marriage market characteristics on marital dissolution, theories of marriage and divorce have not been completely silent on this issue. Becker's (1981; Becker et al. 1977) analogy of "marital bargains" as resembling international trade implies that a current marriage (or trading relationship) will dissolve when either marriage (or trading) partner is able to form a more profitable relationship with a different party. Presumably, the likelihood of finding an associate more attractive than one's current partner will increase as the volume of potential partners in the population increases. More generally, Becker (1981) argued that the principal causes of divorce are marital searchers' imperfect information about marriage markets and the accumulation of better information about both one's remarriage prospects and one's spouse during the course of marriage.[1] One type of information that is likely to improve during marriage is information about the availability of desirable alternatives to one's current spouse. Of course, as Becker (1981:220) notes, the gains from marriage and the significant costs to marital dissolution make it unwise, while married, to search as fervently for a remarriage partner as for a first spouse. Yet, such searches, however tepid, are more likely to be successful when the supply of alternative, potential spouses is abundant. Moreover, one need not have a specific remarriage partner in mind for remarriage opportunities to influence marital dissolution, since even a vague perception of significant prospects for forming post-divorce relationships may increase the

[1] We use the terms "remarriage prospects" and "remarriage markets" loosely to refer to opportunities to form intimate extramarital and postmarital relationships. To confirm the hypothesis that these opportunities increase the probability of divorce, it is not necessary that divorced persons actually remarry someone from their field of eligibles. Some may establish a relationship, but not remarry. And others may decide to divorce their current spouse based only on their *perception* of substantial opportunities to form postmarital relations, which is likely to be enhanced by the availability of numerous spousal alternatives (Udry 1981).

motivation to divorce. Also, extensive opportunities for extramarital sexual relationships may sow the seeds of marital infidelity, which is frequently cited by divorced persons as a reason for marital dissolution (Kitson, Babri, and Roach 1985).

Nothing in Becker's (1981) economic theory appears to suggest that remarriage opportunities would differentially affect wives' and husbands' propensity to divorce. A local marriage market containing favorable remarriage prospects for wives (and thus generally unfavorable prospects for husbands) should increase the probability of divorce, as should a marriage market having favorable remarriage prospects for husbands (and thus generally unfavorable prospects for wives). Hence, Becker's theory implies a *curvilinear*, U-shaped relationship between any sex-ratio based measure of mate availability and marital dissolution—the highest rates of divorce would occur where either men or women eligible for remarriage outnumber those of the opposite sex, and the lowest rates would occur where the numbers of eligible men and women approach parity.

Of course, using such a sex-ratio based measure of mate availability does assume that the wishes of the spouse enjoying abundant remarriage opportunities override those of the spouse faced with only a few. Theoretically, in marriage markets with relatively high sex ratios (i.e., an abundance of available men), wives would have numerous remarriage opportunities and thus be more motivated to divorce, while husbands would have comparatively few spousal alternatives, and thus be motivated to maintain an existing marriage. In marriage markets with low sex ratios (i.e., an abundance of available women), the reverse situation would occur. Although the wishes of the spouse facing few spousal alternatives may temper the effect of the sex ratio, our argument assumes that the desire of the spouse facing abundant remarriage opportunities takes precedence. That is, we assume it takes only one tempted, motivated partner to dissolve a marriage.

Empirical analyses of the impact of spousal alternatives on divorce, while not specifically testing the hypothesis of a curvilinear association, have generated somewhat equivocal findings. Drawing on Guttentag and Secord's (1983) theory of imbalanced sex ratios, several cross-national studies have explored the relationship between societal sex ratios and rates of marital dissolution. Consistent with economic theory, Guttentag and Secord (1983) argued that a relative oversupply of marriageable women will increase men's desire to divorce (i.e., the availability of numerous alternative partners provides incentives for a husband to leave his wife). But unlike a simple supply and demand explanation, Guttentag and Secord (1983) hypothesized different dynamics for women. When women enjoy favorable remarriage opportunities, men use their superior structural power over them, defined as control over the "political, economic, and legal structures of the society" (p. 26), to limit women's marital and familial options, in particular, their ability to divorce. Consistent with these predictions, two cross-national studies have observed *inverse* correlations between the sex ratio and the divorce rate (South and Trent 1988; Trent and South 1989). However, these studies did not test for the curvilinear relationship implied by Becker's theory. Moreover, seemingly in contradiction to the Guttentag and Secord (1983) thesis, South and Lloyd (1992) reported a *positive* relationship between women's marriage opportunities and the divorce rate across U.S. marriage markets simultaneously defined by age, race, education, and geographic area. The direction of the observed relationship between the supply of spousal alternatives and marital dissolution may differ across countries and across marriage markets in the United States because, in the latter context, men may lack the structural power necessary to diminish women's ability to divorce. In any event, these inconsistent findings warrant reevaluation of the impact of the supply of spousal alternatives on the risk of marital dissolution.

Other discussions of the impact of marriage markets on marital behavior, while not specifically addressing the issue of marital dissolution, have called attention to the quality, rather than the quantity, of potential spouses. Thus, in addition to evaluating the effect of the quantity of spousal alternatives on marital dissolution, it seems important to consider the characteristics of potential remarriage partners. Wilson (1987), for example, argued that the dearth of men with desirable economic characteristics, especially steady em-

ployment, diminishes women's gains from marriage and, thus, reduces female marriage rates. And, Oppenheimer (1988) argued that the secular retreat from marriage is less a function of women's increasing financial independence than of men's deteriorating economic circumstances. Hence, in attempting to relate spousal alternatives to divorce, it is important to consider the relative quality, as well as the relative quantity, of the alternative husbands and wives available to currently married persons. For wives, an attractive pool of eligible husbands is likely to contain a comparatively large percentage of employed and highly educated men (Lichter et al. 1992). Men's marital (or re-marital) preferences are more difficult to anticipate but, like women, they appear to express a slight preference for employed over nonemployed spouses (South 1991).

The availability of employed women in the local marriage market might also increase married men's *access* to spousal alternatives—and thus their motivation to divorce. That is, in marriage or labor markets employing many unmarried women, married men may be more likely in their routine activities to encounter and develop social relationships with desirable alternatives to their current wife. It seems likely that a substantial percentage of extramarital relationships that precipitate divorce evolve between married men and unmarried women in and around the workplace. The extent to which men work with, or otherwise encounter, unmarried women during their workday activities—encounters that are more frequent as more unmarried women join the labor force—should therefore heighten the risk of marital dissolution.

While our primary objective is to assess the influence of the availability of wives' and husbands' spousal alternatives on marital dissolution risks, the dynamic, multilevel model we develop also allows us to address the possible impact of other contextual factors on divorce. Three such factors have been mentioned in the literature on marital dissolution. First, it has been argued that marriages in large, metropolitan areas will be more likely than others to dissolve. Larger areas are likely to contain numerous desirable alternative spouses and, especially for women, to provide economic alternatives to marriage. And, the liberalizing influence of the urban environment on attitudes may decrease the stability of marriage (Elder 1978). The evidence on the relationship between city size and divorce is inconsistent, however; some studies observe the predicted differentials (e.g., South and Spitze 1986), but others do not (e.g., Greenstein 1990).

Second, residential mobility in the local geographic area has been argued to reduce community social integration and thus to destabilize marriages (Glenn and Shelton 1985). According to White (1990), "[I]ntegration increases the likelihood that people will follow social norms in choosing an appropriate spouse and fulfilling their marital roles, and decreases the likelihood that they will court community stigma by divorcing" (p. 905). Consequently, low levels of social integration, as reflected by high mobility rates, should lead to high levels of marital disruption. In aggregate analyses, the rate of geographic mobility is frequently a significant predictor of community divorce rates (Breault and Kposowa 1987; South 1987).

A third community characteristic that might influence the rate of marital dissolution is the cost of housing. Divorce usually requires that one or the other spouse move to a new residence; for many, this entails renting or purchasing a new home. Although the cost of alternative housing is not likely to be the primary concern of persons contemplating divorce, the ease with which a new household can be established following divorce might nonetheless enter into such calculations, especially for low-income individuals. Such reasoning is consistent with the finding from income-maintenance experiments—that the financial ability to form a new household increases the risk of marital dissolution (Hannan, Tuma, and Groeneveld 1978). Thus, we anticipate that, all else being equal, the likelihood of divorce will be higher in geographic areas with moderate housing costs than in those with high housing costs.

DATA AND MEASURES

The empirical analyses draw on three different data sets. First, data from the National Survey of Families and Households (NSFH) describe the extent to which recently di-

vorced and separated persons had been romantically involved with someone other than their spouse prior to their divorce. Of course, it is not necessary that every divorce be precipitated by marital infidelity for the supply and quality of spousal alternatives to influence marital disruption. Some soon-to-be divorced persons may have a particular remarriage partner in mind when they decide to leave their current spouse, but may not have yet established a romantic relationship with that person; others may decide to divorce partly because they perceive favorable opportunities to establish post-marital relationships, even though they are not involved with an alternative partner. Yet, the hypothesis that the volume of spousal alternatives positively affects marital dissolution does imply that a nontrivial proportion of divorces will have been precipitated by one or the other spouse having formed an extramarital relationship with another while the marriage still existed.

The NSFH is a national probability sample of 13,017 adults interviewed between March 1987 and May 1988 (Sweet, Bumpass, and Call 1988). We use selected questions from the self-administered questionnaire completed by the approximately 2,000 respondents who had experienced a marital separation after January 1, 1977. In two items, respondents were asked: "Were you (and was your husband/wife) involved with someone else just before your marriage ended?" For the report of respondent's own involvement, the possible responses are "Yes" and "No"; in reporting the ex-spouse's infidelity, "Don't know" is also a response category.

The bulk of the analyses attempts to directly ascertain the effect of the supply of spousal alternatives on rates of marital dissolution by merging marital history data from the National Longitudinal Survey of Youth (NLSY) with aggregated data from the Public Use Microdata Samples (PUMS-D) from the 1980 Census. The NLSY is a national probability sample of 12,686 civilian and military respondents ages 14 to 22 in 1979 (Center for Human Resource Research [CHRR] 1992). Respondents have been interviewed annually since 1979; the retention rate is over 93 percent for the civilian sample through the 1985 interview. The NLSY is a rich data source for analyzing marital dissolution. It has detailed data on dates of marriages, separations, and divorces, as well as much information on potential explanatory variables. What makes the NLSY especially appropriate for analyzing the contextual influences on divorce is the supplementary geocode files that provide the state and county of residence for each respondent for each annual interview. As described below, these geographic identifiers can be used to match respondents with census-derived measures of the quantity and quality of spousal alternatives in their local marriage market.

We delimit the NLSY sample in the following ways. First, because of considerable missing geocode data for the military sample, we use only the civilian sample. Second, because very few of the respondents were in second marriages, we focus only on the dissolution of first marriages. Third, because the high levels of Census undercount and the relatively high rates of exogamy among African Americans and Hispanics make it difficult to specify the pool of eligible mates for these groups (Oropesa, Lichter, and Anderson 1993; South and Lloyd 1992), we focus only on the marital experiences of non-Hispanic Whites (hereafter referred to simply as Whites). And fourth, because of the need to match geocode data from each annual interview with 1980 Census data, we include only respondents whose marriages existed sometime between 1979 and 1984. That is, marriages that dissolved prior to the initial 1979 interview are excluded from the analysis because no geographic data are available on the marriage market in which those divorces occurred. Fortunately, there are very few of these left-censored observations; less than 2 percent of respondents were married and divorced prior to 1979.[2] We exclude marriages begun after 1984 so as to increase the accuracy of 1980 Census data as measures of *current* marriage market conditions, that is, of conditions in years subsequent to the 1980 Census date.

By comparing the recorded dates of first marriages and the dissolution of first marriages, through either separation or divorce, we construct marital event histories for the NLSY respondents. In the proportional hazards models that follow, the log of the haz-

[2] This figure represents about 7 percent of all marriages that existed prior to 1985.

ard of dissolution at each duration of marriage, measured in years, constitutes the dependent variable. Respondents who were widowed, who left the sample prior to separating or divorcing, or whose marriage survived to 1985 are treated as censored observations (Allison 1984). Given the above restrictions, 2,592 of the White NLSY respondents are eligible to be included in the analysis, about 22 percent of whom separated or divorced prior to 1985 or before dropping out of the survey.

Of course, by focusing on a relatively youthful sample, we analyze marriages of relatively short duration. The observed duration from first marriage to either dissolution or censoring ranges from less than one year to twelve years, with a mean of 3.43 years and a standard deviation of 2.32 years. We do not view this as a major drawback, however, since the early years of marriage are the least stable (Thornton and Rodgers 1987). Moreover, young and/or recently married persons are probably more likely than others to continue their marital (or remarital) searches even after marriage, and thus marriages of short duration may be most at risk of dissolution because one partner has encountered an attractive spousal alternative.

Measuring Spousal Alternatives

Measuring the supply of alternative spouses in the local marriage market requires that two issues be addressed. First, we must decide how to geographically circumscribe the local marriage market. Although various alternatives exist, our choice here is to use the Labor Market Area (LMA) as defined in the 1 percent PUMS-D file (Tolbert and Killian 1987), a spatial unit used successfully in prior research linking marriage market characteristics to marriage rates (Lichter et al. 1991). LMAs are believed to circumscribe the spatial boundaries within which daily social action occurs and thus from which most potential spouses are chosen. The 382 LMAs are delimited primarily by journey to work patterns and average over 500,000 in population. As Lichter et al. (1992) have noted, LMAs have several important advantages over rival operationalizations of marriage markets: Unlike counties, LMAs can cross state boundaries and, unlike Metropolitan Statistical Areas (MSAs), they encompass the entire U.S. population. Moreover, the microlevel PUMS-D file provides the necessary detail with which to compute appropriate measures of the quantity and quality of spousal alternatives.

The second important issue in measuring the supply of spousal alternatives involves estimating the number of "suitable" potential remarriage partners available to each NLSY respondent. As in prior analyses of marriage market effects (Lichter et al. 1992; South and Lloyd 1992), we require that suitable partners be unmarried, not be institutionalized, and be of the same race and ethnicity as the respondent (i.e., non-Hispanic White). Placing age restrictions on the field of eligibles is somewhat more problematic. Our approach is to compute sex ratios (i.e., number of men per 100 women) for non-Hispanic Whites for the ages of the vast majority of the married NLSY respondents throughout the survey period—ages 20 through 30. While we could attempt to circumscribe more sharply the ages defining the field of eligibles, all age restrictions are, to some extent, arbitrary (Lichter et al. 1992). Moreover, estimating detailed race- and age-specific population counts for years subsequent to 1980, which would be necessary for explaining marriages dissolving in these later years, would involve numerous assumptions we cannot test. Fortunately, Fossett and Kiecolt (1991) demonstrate that detailed age-specific sex ratios are highly intercorrelated, especially at the young ages we examine here. This implies little or no gain by using more detailed age categories. Using age-staggered sex ratios, which allow women to prefer men several years older and men to prefer women several years younger, would also make little or no difference in our results (Fossett and Kiecolt 1991).

The sex ratio is treated as a time-varying explanatory variable, measured (in 1980) for the LMA of residence for each respondent at each annual interview. Thus, the supply of spousal alternatives available to each respondent is sensitive to changes in residence. By construction, the sex ratio assigned to each NLSY respondent includes the number of men and women in the LMA available to each respondent as well as the number of men and women who might be "competing"

with the respondent for available, opposite-sex mates.[3] In this context, low sex ratios indicate relatively numerous spousal alternatives for husbands and relatively few for wives, while high sex ratios indicate the reverse.

To capture the hypothesized curvilinear relationship between the sex ratio and marital instability, we model the effects of the sex ratio as a second order polynomial, including as independent variables the sex ratio in its original metric and the sex ratio squared. We prefer this approach to modeling curvilinearity to one that examines the absolute value of deviations from complete parity in the numbers of eligible women and men (i.e., a sex ratio of 100) because it is likely that fewer unmarried men than unmarried women ages 20 to 30 are actively seeking spouses. On average, men marry later than women, and at these ages fewer men than women express a strong desire to marry (South 1993). Hence, we anticipate that the nadir of the curve describing the relationship between the sex ratio and the risk of marital dissolution falls above a sex ratio value of 100. If, as hypothesized, the risk of dissolution is highest where the volume of spousal alternatives available to either husbands or wives is large, then the coefficient for the untransformed sex ratio should be negative, and the coefficient for the squared term should be positive.

In addition to considering the quantity of spousal alternatives available to husbands and wives, we also examine the quality of these potential mates. From the women's perspective, one quality of potential husbands—current or potential employment—is believed to be particularly important (Oppenheimer 1988). To measure the desirability of potential husbands on this variable, we compute from the PUMS-D file the percentage of men in each respondent's marriage market (i.e., unmarried non-Hispanic White men in the same LMA, age 20 to 30) who are currently employed or enrolled full-time in school. Analogously, we include the percentage of unmarried women in each respondent's marriage pool who are employed or enrolled full-time in school as a measure of the "quality" of potential wives available to married men. As suggested above, in addition to reflecting the supply of desirable alternative wives available to married men, this variable may also reflect the probability that married men will encounter in their routine, workday activities attractive alternatives to their current wife. As with the sex ratio, current or future employment status as an indicator of the quality of spousal alternatives is treated as a time-varying covariate, measured (in 1980) for each respondent's place of residence at each annual interview.

Other Contextual Characteristics

Three other characteristics of the local geographic area serve as explanatory variables. The variable we call *in MSA* distinguishes respondents who reside in a Metropolitan Statistical Area from those who do not. As argued above, metropolitan areas are likely to contain numerous marital and economic alternatives, and thus may foster higher rates of dissolution. *Geographic mobility*, believed to decrease a community's level of social integration and hence to promote divorce, is measured by the percentage of 1980 residents in a respondent's county who lived in a different county in 1975. The *cost of housing* is measured by the median monthly rent paid by residents in the respondent's county. To adjust for geographic differences in the cost of living, we divide the median monthly rent by annual per capita income (in hundreds of 1980 U.S. dollars) of the county. We use median monthly rent rather than median housing price on the assumption that spouses contemplating leaving the marital household would be more likely to rent than to purchase a new home. Measures of geographic mobility, rental prices, and per capita income are taken from the County and City Data Book tape (U.S. Bureau of the Census 1983) and merged with the individual NLSY records. As with metropolitan residence, both geographic mobility and rental prices are matched with the respondent's area of residence at each interview—thus, these vari-

[3] We observed several extreme observations at both ends of the distribution of the sex ratio; most likely they are consequences of sampling errors in the smaller LMAs. To reduce the impact of these observations, values of the sex ratio below the 5th and above the 95th percentile were recoded to the values at those points in the distribution.

ables can change in value if a respondent changes residence.[4]

Individual-Level Predictors

Prior research on marital dissolution has identified many social and demographic characteristics of individuals and married couples that influence the probability of marital dissolution (White 1990). While not all of these characteristics can be measured by variables in the NLSY, we can include many of the most frequently used predictors. Individual characteristics of the respondent include *age at first marriage* (in years), *years of school completed*, whether or not the respondent *owns her or his home*, *number of children* ages 0 to 17 in the household, *annual income* (in 1,000s of U.S. dollars), and *number of weeks worked* in the year prior to the interview. Income and weeks worked by the respondent's spouse are also included as explanatory variables. Finally, to control for possible gender differences in the reporting of marital dissolutions (Bumpass et al. 1991), we include a dummy variable for respondent's *sex*.[5] All but age at first marriage and respondent's sex are treated as time-varying covariates and are measured at each annual interview.

[4] We also assume that there is considerable stability in the values of the Census-based contextual variables over the study period—that is, that the correlations between the values of the contextual variables in 1980 and the values in 1979 and in subsequent years through 1984 are high. Others have also made this assumption; we, too, find it reasonable (e.g., Lichter et al. 1992).

[5] For the respondent's spouse, we lack data on age at first marriage and the years of school completed. Thus, if the respondent is male, these variables refer to the husband; if the respondent is female, they refer to the wife. In preliminary analyses, we explored possible sex (or spousal) differences in the effects of these variables by including product terms representing the interaction of respondent's sex with both age at marriage and years of school. Neither interaction was statistically significant, suggesting that the effects of these two variables operate similarly for husbands and wives. Because we cannot control for the analogous characteristics of the spouse, however, the effects of these variables should be interpreted cautiously.

METHOD

Proportional hazards regression models are used to assess the impact of marriage market characteristics and the other explanatory variables on the probability of marital dissolution through divorce or separation (Allison 1984; Teachman 1982). As Bumpass et al. (1991) noted, proportional hazards models have become the standard technique for analyzing marital transitions, including marital disruption. Preliminary analyses testing the proportionality assumption (including the examination of log-log survival functions and estimating models that allow the effects of the explanatory variables to vary by marital duration) revealed no obvious violations of this assumption. This finding is consistent with prior research, demonstrating that most predictors of marital dissolution have similar effects regardless of the duration of marriage (Heaton, Albrecht, and Martin 1985; South and Spitze 1986; White and Booth 1991). The models are estimated with the PHREG routine in the SAS package (SAS Institute, Inc. 1985).

RESULTS

Descriptive data from the NSFH indicates whether the respondent and/or his or her spouse was involved with someone else prior to the termination of the marriage. Table 1 shows the weighted frequency distribution of responses to NSFH items, separately for recently divorced men and recently divorced women. (We emphasize that these divorced men and women had not been married to each other.) This distribution presents a somewhat ambiguous pattern of the extent of infidelity prior to marital disruption. Approximately 15 percent of the respondents (14.3 percent of the ex-wives and 16.5 percent of the ex-husbands) report that they themselves had been romantically involved with another person prior to their divorce. Yet, over 40 percent of the respondents report that their ex-spouses had been unfaithful, and another 20.6 percent were not sure if their spouses had been involved with someone else. Overall, the distributions are quite similar for divorced men and divorced women.

There appears to be nothing in the survey design that would account for differences be-

Table 1. Frequency Distribution of Respondents' Reports of Marital Infidelity by Self and Spouse by Sex: Recently Divorced Persons from the National Survey of Families and Households, 1987–1988

Item and Responses	Men		Women		Total	
	N	Percent	N	Percent	N	Percent
"Were you involved with someone else just before your marriage ended?"						
Yes	141	16.5	163	14.3	304	15.2
No	714	83.5	980	85.7	1,694	84.8
Total responses	855	100.0	1,143	100.0	1,998	100.0
"Was your (husband/wife) involved with someone else just before your marriage ended?"						
Yes	354	41.4	482	42.2	836	41.8
No	343	40.1	408	35.7	751	37.6
Don't know	158	18.5	253	22.1	411	20.6
Total responses	855	100.0	1,143	100.0	1,998	100.0

tween the respondents' reports of the respondents' own, and of their spouses', marital infidelity. Rather, differences in respondents' reports of adultery by themselves and by their spouses are probably attempts to save face. This implies that the true percentage of divorces preceded by infidelity may fall somewhere between these two reports. On the one hand, respondents may be reluctant to admit their own infidelity to the survey administrators (or perhaps to themselves) to avoid being assigned blame for the failure of the marriage or to escape moral condemnation. On the other hand, consciously or not, respondents may overreport the infidelity of their ex-spouses as a way to justify the dissolution of the marriage (the actual reasons for which may lie with both partners) or as a way to explain being forsaken by one's spouse. Importantly, however, even the most conservative estimate suggests that a substantial percentage of divorces were preceded by marital infidelity. If it is assumed that relatively few divorces were preceded by infidelity on the part of *both* spouses (an assumption supported by the finding that only 11.2 percent of the respondents who reported that one spouse was involved with someone else prior to the divorce reported that the other spouse was also involved), then the percentage of dissolved *marriages* in which one or the other spouse was unfaithful is simply the sum of the percentages of affirmative responses by divorced men and divorced women. Using only the reports of the respondent's own behavior—the most conservative strategy—suggests that 30.8 percent (16.5 + 14.3) of divorces were preceded by infidelity (the lower boundary of a 95 percent confidence interval would be greater than 28 percent). Estimates derived from a respondent's report of infidelity on the part of his/her spouse would be considerably higher.

Of course, these data only suggest extramarital relations as a prominent cause of marital dissolution. Lacking data on the extent of infidelity among couples who do not divorce, we cannot ascertain the strength of the association between extramarital relations and marital disruption. Moreover, marital dissatisfaction may lead to both the decision to divorce and the decision to initiate an extramarital relationship, and thus the inferred relationship between the variables is potentially spurious. Nonetheless, these estimates suggest clearly that at least one-quarter—and perhaps considerably more—of marital disruptions are preceded by the infidelity of one or the other spouse while the marriage still existed. In the rest of our analysis, we attempt to determine the extent to which marital dissolution is influenced by the opportunities available to spouses to cultivate these extramarital relationships.

Table 2. Means and Standard Deviations for Variables in Analyses of Marital Dissolution: Non-Hispanic Whites from the NLSY, 1979–1984

Variable	Mean
Individual Characteristics	
Sex (0 = male; 1 = female)	.62 (.48)
Age at first marriage	20.34 (2.26)
Years of school	12.05 (2.10)
Home ownership (0 = no; 1 = yes)	.33 (.47)
Number of children	.73 (.87)
Wife's weeks worked	26.72 (21.46)
Husband's weeks worked	40.28 (17.56)
Wife's income	4.41 (5.35)
Husband's income	10.79 (9.05)
Marriage Market Characteristics	
Sex ratio (ratio of men to women)	127.48 (22.82)
(Sex ratio)^2	16,772 (6,097)
Percent females employed or in school	84.41 (5.57)
Percent males employed or in school	85.30 (7.12)
Other Contextual Variables	
In MSA (0 = no; 1 = yes)	.61 (.49)
Geographic mobility rate in county	19.47 (8.81)
Median monthly rent in county	2.31 (.35)
Percent divorcing or separating	22.11
Number of cases	2,592

Note: Numbers in parentheses are standard deviations. Descriptive statistics for time-varying covariates are from the interviews immediately preceding the event or censoring.

Table 2 presents the means and standard deviations for the variables used in the proportional hazards models. The relatively young age at first marriage ($\bar{x}$ = 20.34) for this sample is, of course, a natural consequence of the youthfulness of the NLSY; marriages contracted at older ages cannot be observed. The typical White respondent resides in a marriage market having 127 unmarried men in their twenties per 100 unmarried women of the same age. This mean sex ratio greater than parity results largely from limiting the ratio to unmarried persons; because men marry later than women, on average, more men than women in this age range will be currently unmarried. Across marriage markets (or LMAs), similar percentages of unmarried men and women in this age range are employed or enrolled in school.

The proportional hazards regression models are shown in Table 3. Model 1 includes as explanatory variables only the individual- (and couple-) level characteristics. With few exceptions, the effects of these variables on the risk of marital dissolution are consistent with prior research. High levels of education, home ownership, and the presence of children all strongly reduce the probability of divorce or separation. Husband's income and weeks worked are also inversely associated with the risk of marital dissolution. However, neither wife's weeks worked nor wife's income significantly influence the risk of divorce, indicating little support for the hypothesized "independence effect." These weak effects of wife's employment on the risk of marital dissolution are consistent with the results of some, but not all, prior research (Greenstein 1990; White 1990). Although age at marriage does not evince a significant net effect on the risk of dissolution, it does have a significant, inverse bivariate effect (equation not shown), implying that its impact may operate through other variables in the model. Finally, the coefficient for sex of respondent fails to attain statistical significance, indicating that neither sex disproportionately reports experiences of marital dissolution.

Model 2 in Table 3 includes as explanatory variables only the marriage market characteristics and other contextual variables. Importantly, the coefficients for the sex ratio and its squared value take on the hypothesized pattern indicative of a U-shaped association.

Table 3. Coefficients from Proportional Hazards Regression Models of Marital Dissolution: Non-Hispanic Whites from the NLSY, 1979–1984

Explanatory Variables	Model 1 b	Model 2 b	Model 3 b
Individual Characteristics			
Sex (0 = male; 1 = female)	.130 (.097)	—	.084 (.099)
Age at first marriage	–.033 (.025)	—	–.029 (.025)
Years of school	–.101** (.024)	—	–.115** (.024)
Home ownership (0 = no; 1 = yes)	–.595** (.115)	—	–.552** (.116)
Number of children	–.319** (.063)	—	–.312** (.063)
Wife's weeks worked	–.002 (.003)	—	–.002 (.003)
Husband's weeks worked	–.005* (.002)	—	–.007** (.002)
Wife's income	–.006 (.011)	—	–.008 (.012)
Husband's income	–.037** (.007)	—	–.037** (.007)
Marriage Market Characteristics			
Sex ratio (ratio of men to women)	—	–.034** (.012)	–.025* (.012)
(Sex ratio)^{2} [a]	—	.013** (.005)	.010* (.005)
Percent females employed or in school	—	.016 (.010)	.028** (.011)
Percent males employed or in school	—	–.008 (.008)	–.009 (.008)
Other Contextual Variables			
In MSA (0 = no; 1 = yes)	—	–.042 (.088)	.068 (.092)
Geographic mobility rate in county	—	.016** (.006)	.018** (.006)
Median monthly rent in county	—	.027 (.140)	–.103 (.142)
–2 log likelihood	7,980.317	8,303.226	7,767.094

*$p \leq .05$ **$p \leq .01$ (two-tailed tests)

[a] Coefficient and standard error are multiplied by 100.

Note: Numbers in parentheses are standard errors. N = 2,592.

Both coefficients are statistically significant. Thus, the risk of marital disruption is greater when either husbands or wives encounter a relatively sizable quantity of alternatives to their current spouse. The only other contextual variable in Model 2 to exert a significant impact on marital dissolution is the county-level rate of geographic mobility. As hypothesized, the risk of dissolution is higher in counties with comparatively high levels of geographic mobility.

Model 3 in Table 3 combines the individual-level and aggregate-level explanatory variables. Most of the coefficients for vari-

Table 4. Predicted Hazard of Marital Dissolution for Selected Sex Ratios

Sex Ratio	Percentile	Predicted Hazard of Dissolution	Ratio of Hazard to the Hazard at the Mean Sex Ratio
105	10th	.041	1.079
115	25th	.039	1.026
127	$= \bar{x}$	.038	1.000
137	75th	.039	1.026
162	90th	.043	1.132

Note: Predicted values are derived from Model 3, Table 3, assuming mean values for other explanatory variables.

ables in the prior two equations retain their sign and statistical significance. That the observed effects of the conventional, individual-level predictors of divorce are not modified by the addition of the contextual explanatory variables suggests that models excluding the latter, as most prior models have done, are not seriously misspecified by their neglect of marriage market characteristics.

One notable feature of Model 3 is that, once the individual-level explanatory variables are controlled, the percentage of an LMA's unmarried females who are employed or enrolled in school emerges as a significant, positive predictor of marital dissolution. That is, unmarried women's labor force (and school) participation rate in the local marriage market increases the probability of divorce, even net of a wife's own employment situation. Although we cannot dismiss the possibility that this effect occurs because married men prefer employed to nonemployed women as remarriage partners, evidence for such a preference is weak. It is also possible that high levels of employment for unmarried females in a local area signal enhanced economic opportunities for married women, and that both employed and nonemployed married women decide to divorce based on their perceptions of these future opportunities. But the lack of a positive impact of wife's weeks worked or wife's income at the individual level renders this explanation suspect. Rather, as argued above, perhaps the most plausible interpretation of this effect is that high levels of unmarried women's employment increase the probability that married men will encounter, in their routine activities in and around the workplace, desirable alternatives to their current wife. These encounters, in turn, would elevate the risk of marital dissolution.

How large are the effects of the sex ratio on the risk of marital disruption? Table 4 presents the predicted hazard of marital dissolution for selected values of the sex ratio and the ratio of the predicted hazards at these values to the predicted hazard at the mean sex ratio (127). These predicted values are derived from Model 3 of Table 3, holding all other explanatory variables constant at their respective means. Based on the coefficients from this model, the curve describing the relationship between the sex ratio and the (log) hazard of dissolution reaches its nadir at a sex ratio value of 129, approximately the sample mean. Moving out from the mean sex ratio to the sex ratio at the 10th percentile (about 1 deviation below the mean) increases the predicted hazard of dissolution by approximately 8 percent. Moving out from the mean sex ratio to the sex ratio at the 90th percentile (about 1.5 standard deviations above the mean) increases the predicted hazard of dissolution by slightly more than 13 percent. Hence, while the supply of spousal alternatives in the local marriage market is clearly not an overwhelming factor in fostering marital dissolution, neither is its impact trivial.

DISCUSSION AND CONCLUSION

While the theoretical literature on marital dissolution frequently speaks of the supply of marital alternatives as a potential determinant, few studies have attempted to quantify its impact on the probability of divorce. Rather, the bulk of the empirical literature has focused either on sociodemographic differen-

tials in marital disruption (e.g., Bumpass et al. 1991) or on the internal marital dynamics that precipitate a divorce (e.g., Conger et al. 1990). The approach we have adopted here argues for a more "open-systems" perspective on marital dissolution—one that, in addition to the more traditional "push" factors, gives emphasis to the conditions in the marital environment that may serve to "pull" persons away from their current spouses. Specifically, we draw on economic and sociological theories of marriage markets and marital behavior to derive hypotheses relating the quantity and quality of spousal alternatives to marital dissolution. Descriptive evidence showing that a sizable percentage of recently divorced persons had been romantically involved with someone other than their spouse suggests that many married persons continue to "search" for an intimate partner, or at least remain open to the possibility of forming extramarital relationships, even while married. More importantly, multilevel, dynamic models of divorce and separation reveal that, for non-Hispanic Whites, the risk of marital disruption is influenced significantly by the supply of spousal alternatives available to one spouse relative to the supply available to the other. These effects exist, net of the conventional, individual-level predictors. We also find evidence that, net of the wife's own employment, the risk of dissolution is higher for couples who live in marriage markets in which a substantial percentage of unmarried women are employed. The aggregate rate of geographic mobility also reduces marital stability and is likely a consequence of reduced social integration in such a community. However, we find no evidence that either metropolitan residence or local housing costs significantly influence the risk of marital disruption.

Our findings may provide a partial resolution to an apparent paradox regarding the impact of women's labor force participation on rates of marital disruption. While secular trends in rates of divorce and women's labor force participation have generally moved in unison during the postwar period (Cherlin 1992; South 1985), research on the relation between these variables at the individual level has generated inconsistent findings; some studies find the expected "independence effect" of wife's employment; but others, including this analysis, find no effect (Greenstein 1990; White 1990).

Our finding of a positive association between an area's unmarried female labor force participation rate and marital dissolution, net of wives' own employment, suggests that, in the aggregate, women's labor force participation may increase *married men's* propensity to seek a divorce by increasing the probability that men will find a more attractive alternative to their current wife. At least two mechanisms may be operating here. First, increased women's labor force participation is likely to increase rates of social interaction between men and women in contexts—especially those in and around the work place—in which extramarital liaisons are established and remarriage partners selected. As married men work more and more often with women colleagues, it seems likely that interpersonal relationships will develop that may threaten the stability of the men's marriages. Second, increased women's labor force participation (and full-time enrollment in school) may raise the age at which women marry, thereby expanding the pool of unmarried women, and thus the supply of alternative spouses available to married men. In both instances, secular increases in women's labor force participation would be expected to increase the divorce rate quite independently of any effect of wives' employment on marital instability among individual couples.

Most generally, linking marriage market characteristics to divorce anchors the study of marital dissolution in an established conceptualization of social structure (Blau 1977). The distribution of the population along salient structural dimensions—here including sex, age, marital status, and geographic area—presents opportunities to form cross-sex relationships that may decrease marital stability. In a social climate increasingly tolerant of divorce (Thornton 1989) and skeptical of the ideal of marital permanence (Glenn 1991), and in which even married persons are tentatively in the marriage market (Farber 1987), the opportunity to encounter and unite with a more attractive mate will lead many wives and husbands to dissolve their current partnership. In this sense, marital dissolution is, in part, a product of the demographic opportunities embedded in the social structure.

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