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## Abstract

Large racial and ethnic differentials in the risk of marital disruption are observed in the United States, with Blacks exhibiting higher rates of disruption than many other groups. We use data from the 1995 National Survey of Family Growth to investigate whether racial/ethnic differences in exposure to risk factors for disruption can explain variation in levels of marital instability across groups. We consider a wide array of risk factors for disruption and offer one of the few recent analyses of marital instability among Mexican-American women. Our results suggest that, if differences in population composition between groups were removed, the White-Black and Black-Mexican differentials in disruption would be reduced by approximately 30% and 50%, respectively. The story regarding the White-Mexican differential is more complicated, however, and hinges on nativity status of Mexican women. Finally, in light of large differences in marital instability between U.S.-born and foreign-born Mexican women, we also explore the possibility that compositional differences might contribute to differentials in marital instability between these two groups.



**Key Words: Divorce, Race and Ethnicity, United States**

Substantial racial and ethnic variation exists in patterns of marital instability in the United States, with Black Americans experiencing considerably higher levels of disruption than many other groups. For example, recent evidence suggests that almost one third (32%) of marriages among non-Hispanic White women end through separation or divorce within ten years, compared to almost half (49.4%) of marriages among non-Hispanic Black women (Phillips and Sweeney, forthcoming; see also Bramlett and Mosher 2002). Although 27.5% of first marriages among Mexican American women end in disruption within ten years, this overall figure conceals important differences by nativity: only 13.1% of foreign-born Mexican women experience a disruption within the first ten years of marriage, compared to almost 41% of U.S.-born Mexican women (Phillips and Sweeney, forthcoming).

It is important to identify the underlying cause(s) of these large group differences in the experience of marital instability, particularly given evidence of adverse outcomes associated with divorce for adults and children (e.g. Amato, 2000; Waite, 1995). In the current research, we ask the following hypothetical question: How much might racial and ethnic gaps in marital disruption decrease if differences across groups in the distribution of key risk factors for divorce, such as age at marriage or premarital childbearing, were eliminated? Using data from the 1995 National Survey of Family Growth (NSFG), we consider racial differences in exposure to a substantially broader array of risk factors than do prior studies and offer the first formal decomposition of recent differences in levels of disruption between Mexican American and White or Black women. We also provide insight into the potential contribution of group differences in population composition to the large nativity-status gap in marital disruption among Mexican-American women.

**Prior Research on Racial and Ethnic Differences in Marital Disruption**

Social scientists have identified a large array of risk factors for divorce. For example, marrying at an early age, having had a premarital birth or cohabitating relationship, residence in the West or in a metropolitan area, having a low level of education, living apart from a biological parent while growing up, and having a spouse of a different race or who was previously married are all associated with an elevated risk of divorce in the aggregate U.S. population (e.g. Bumpass, Castro Martin, and Sweet 1991; Castro Martin and Bumpass 1989; Phillips and Sweeney, forthcoming; White 1990). Scholars interested in understanding racial and ethnic variation in family patterns have also documented differences across groups in the *effects* of several risk factors for marital instability. Evidence suggests, for example, that the stabilizing effect of an older age at marriage may be somewhat stronger among Whites than among Blacks (Heaton and Jacobson 1994; Castro Martin and Bumpass 1989; Teachman 1983, 1986). On the other hand, premarital childbearing and cohabitation increase the likelihood of disruption among Whites but have weaker effects on marital instability among Blacks (Billy, Landale, and McLaughlin 1986; Castro Martin and Bumpass 1989; Phillips and Sweeney, forthcoming; Teachman 1983, 1986).

However, it is important to keep in mind that racial and ethnic groups also differ in their *exposure* to established risk factors for divorce. For example, some argue that Black men and women are less likely than Whites to marry and to stay married because of a relatively weaker ability to support a family, particularly among African American males (e.g. Wilson 1987). Black married women also tend to be older at first marriage, are more likely to have had a premarital birth, and have accumulated somewhat less education, on average, than White married women (e.g. Greenstein 1990; Teachman 1986; Tzeng and Mare 1995). Mexican Americans are

more likely than non-Hispanic Whites or Blacks to marry at a younger age, to live in the West or in a metropolitan area, and are more likely than Whites to be economically disadvantaged (Bean, Berg, and Van Hook 1996; Oropesa et al.1994; Therrien and Ramirez 2001). Such compositional differences might be expected to contribute to observed racial and ethnic differentials in levels of marital instability, as these characteristics are known to be key risk factors for disruption (Castro Martin and Bumpass 1989).

Although little is known about the contribution of compositional factors to levels of marital instability among Mexican Americans relative to other groups, existing studies indicate that compositional differences can explain only a modest portion of the Black-White gap in disruption. For example, in their analysis of marital disruption from the late 1960s through the mid-1980s among respondents to the National Longitudinal Surveys, Tzeng and Mare (1995) find Black-White differences in socioeconomic and demographic background factors to explain less than 30 percent of the total racial differential in the log odds of marital disruption. Heaton and Jacobson's (1994) analysis of data from the 1982 and 1988 cycles of the NSFG also indicates only limited contribution of compositional differences to the overall Black-White differential in disruption. Results are largely similar for more recent historical periods. Based on analyses of June Current Population Survey data for the 1990-94 period, for example, Sweeney and Phillips (2004) attribute less than 20 percent of the overall Black-White gap in the log odds of disruption to compositional variation between Blacks and Whites in age at marriage, education, premarital fertility, and region of residence. Finally, in one of the few studies to consider recent patterns of marital instability among Mexican Americans, Phillips and Sweeney's (forthcoming) analysis of data from the 1995 NSFG indicates that differences in the experience of premarital cohabitation explain little of the overall racial or ethnic gaps in marital

disruption. This prior study does not provide a formal decomposition of race-ethnic differences in levels of marital disruption, however, and also does not investigate the potential contributions of group differences in exposure to other risk factors for disruption.

It may be that factors other than composition indeed account for much of the variation in patterns of marital stability across racial and ethnic groups. For example, some argue that cultural orientations reduce marital stability among Blacks. Both historical and recent evidence suggest that Black women are more likely than White women to leave a union when faced with infidelity or abuse (Pagnini and Morgan 1996; Patterson 1998), and relatively greater reliance on extended family and kin may well change both the utility and the dynamics of marriage among Blacks (Hays and Mindel 1973). The greater levels of marital stability observed among Mexican Americans have been attributed by some to familism, a term used to describe the overarching importance of family and a belief that the collective needs of the family should prevail over individual needs, which some argue is central to Mexican American culture (Oropesa, Lichter, and Anderson 1994; Vega 1990; but see Raley, Durden, and Wildsmith 2004). To the extent that exposure to relevant risk factors cannot explain the racial and ethnic gaps in disruption, arguments of cultural differences across groups have received indirect support.

On the other hand, the relatively small effects of composition on group differences in levels of marital instability documented by previous studies may be due to the fact that important risk factors for disruption have been overlooked. A recent decomposition of race differences in disruption since 1990, for example, considers only age at marriage, education, premarital fertility, and region of residence (Sweeney and Phillips 2004). Remarkably little work has carefully investigated the role played by racial / ethnic differences in exposure to risk factors such as premarital sexual history, characteristics of husbands, or the resemblance of spouses.

Moreover, no systematic effort has been made to decompose differences in disruption between Mexican Americans and Whites or Blacks. Although patterns of marital disruption also differ considerably between U.S.-born and foreign-born Mexican Americans (e.g. Bean et al. 1996), we know of no recent studies that directly consider the potential contribution of compositional factors to nativity differences in levels of marital instability.

The current research revisits the question of the contribution of compositional differences to gaps in marital disruption, examining Mexican American as well as non-Hispanic White and Black women and incorporating a more complete set of compositional factors than prior studies, including characteristics of the respondent, family background, and spousal attributes. Specifically, we quantify the contributions to racial and ethnic disparities in the risk of disruption made by compositional differences (such as racial/ethnic variation in average levels of educational attainment, premarital sexual histories, or characteristics of husbands). We also investigate whether the effects of various compositional factors on racial and ethnic gaps in disruption may offset one another. For example, Black women's tendency to marry at older ages than White or Mexican American women ought to reduce their risk of disruption relative to these groups, whereas Black women's greater likelihood of premarital childbearing ought to increase their relative risk of disruption. Finally, we explore the potential contribution of compositional differences between subpopulations of married women to large differences in levels of marital instability between U.S.-born and foreign-born Mexican American women.

**Data**

Data for the current analysis are drawn from the 1995 National Survey of Family Growth (NSFG), which contains complete marital and cohabitation histories, as well as detailed

information on childhood living arrangements, employment, education, and pregnancy histories. The survey is based on personal interviews conducted in the homes of a national sample of women 15-44 years of age in the civilian, non-institutionalized population of the United States. Survey response rates were relatively high and approximately equal across White, Black, and Hispanic women (with 79% of eligible women completing interviews), and interviews were conducted both in English and in Spanish (Mosher 1998). A useful feature of the 1995 NSFG is that it over-samples minority groups and therefore contains large numbers of women from different racial and ethnic backgrounds. The 1995 survey includes interviews with 6,841 ever-married women aged 15-44, and includes information on first marriages for 4,452 non-Hispanic white women, 1,144 non-Hispanic black women, and 1,020 Hispanic women (Bramlett and Mosher 2002).

We place several limitations on our analytic sample. For example, we consider disruption only in women's first marriages because stability is known to differ for first and higher order marriages and because sample size restrictions become particularly problematic when disruption patterns are disaggregated both by marital order and by race/ethnicity. Only non-Hispanic White, non-Hispanic Black, and Mexican-American women are included in the analyses, hereafter called Whites, Blacks, and Mexicans, respectively. Other Hispanic groups, such as Puerto Ricans and Cubans, are not considered as the NSFG sample sizes of these groups are prohibitively small. We limit our analysis to marriages formed since 1975 by women between the ages of 15 and 30 to reduce the potential bias during earlier periods toward marriages formed at particularly young ages, which is inherent in the NSFG design. Despite this limitation, acceptably large samples of women by race/ethnicity are achieved, with approximately 3,222, 751 and 471 first marriages among Whites, Blacks, and Mexican Americans, respectively.

We consider a wide array of compositional characteristics that vary across racial and ethnic groups and may affect the likelihood of marital disruption, including respondent characteristics, aspects of a woman's family of origin, and spousal attributes. Among the respondent characteristics we consider are well-known risk factors for disruption such as age at marriage, level of education at marriage, employment status during the year prior to marriage, and having had a premarital birth or conception. Because cohabitation is known to affect marital stability, we include controls for whether the woman lived with a romantic partner (her spouse and/or anyone else) prior to marriage. A recent study by Teachman (2003) suggests that the nature of premarital sexual behavior is associated with patterns of disruption; therefore, indicators of whether a woman had no premarital sex before marriage, premarital sex with only her future husband, or premarital sex with someone other than her future husband are incorporated into the analysis. We also control for both region and metropolitan status of current residence. We do not describe the theoretical rationales for the relationship between these respondent characteristics and the risk of marital disruption here, as their importance is well documented elsewhere (e.g. Castro Martin and Bumpass 1989; White 1990; Teachman 2003).

We also measure certain features of a woman's family of origin, including the educational attainment of her parents, whether she was raised as a Catholic, her place of birth (i.e. whether she is foreign-born), and her childhood family structure (i.e. whether her parents were divorced by age 14). These characteristics may affect the risk of disruption directly, through their influence on attitudes toward marriage and/or the acceptability of divorce (Bumpass et al. 1991). They may also indirectly impact the likelihood of disruption through their influence on other risk factors, such as age at marriage.

Finally, we consider a number of spousal characteristics that may influence marital stability – namely, whether the spouse was previously married, the spouse's age and level of education at the time of marriage, and whether there are differences in terms of age, race, and education between the partners. Differences between partners with respect to these characteristics may contribute to increased conflict, reduced social support for the relationship, and/or indicate differences in values and control between partners, all factors that may lead to marital instability (Bumpass et al. 1991; Kalmijn 1998).

**Method**

We first describe key compositional differences among populations of White, Black, and Mexican American women by comparing the mean values of respondent characteristics, background factors, and spousal attributes for each race and ethnic group at the time of first marriage. To investigate and quantify the contribution of differences in population composition to the overall racial and ethnic gaps in the level of marital disruption, we employ regression decomposition techniques. We examine the role of compositional factors in explaining three racial/ethnic differentials — that between Whites and Blacks, between Whites and Mexican Americans, and between Blacks and Mexican Americans.

We begin by decomposing the differences in group outcomes to identify those risk factors which are the largest contributors to the racial and ethnic gaps in disruption, following the approach of Jones and Kelley (1984). Using this method, the difference in the expected log odds of disruption is expressed in terms of a component due to differences in group means and a component that is unexplained. This approach suggests two possible procedures to decompose

the difference between the marital disruption rates of any two groups, one that uses group 1 as the standard and another that uses group 2 as the standard, as shown below.

$$(1) \qquad \Theta_1 - \Theta_2 = \sum_{k=0}^{k=K} b_{k,1}\left(\bar{x}_{k,1} - \bar{x}_{k,2}\right) + \sum_{k=0}^{k=K} \bar{x}_{k,2}\left(b_{k,1} - b_{k,2}\right),$$

$$(2) \qquad \Theta_1 - \Theta_2 = \sum_{k=0}^{k=K} b_{k,2}\left(\bar{x}_{k,1} - \bar{x}_{k,2}\right) + \sum_{k=0}^{k=K} \bar{x}_{k,1}\left(b_{k,1} - b_{k,2}\right),$$

where $\Theta_1$ is the expected log odds of marital disruption for group 1, $\Theta_2$ is the expected log odds of disruption for group 2, K is the total number of coefficients in the model including the intercept, $b_{k,1}$ is the regression coefficient for variable k in the model for group 1 (where k = 0 indicates the regression intercept and $\overline{x_0} = 1$ for all groups), $b_{k,2}$ is the regression coefficient for variable k in the model for group 2, $\bar{x}_{k,1}$ is the mean value of variable k for group 1 and $\bar{x}_{k,2}$ is the mean value of variable k for group 2. For both equations, the "unexplained" portion of the racial and ethnic gap in disruption is the difference in the effects of covariates weighted by composition. A common approach in decomposition analyses is to adopt the dominant group (e.g. whites or males) as the standard, implicitly assuming that the process that describes the socially dominant group is optimal. However, as there is no clear reason to choose one equation over the other in this application, we present both sets of estimates, viewing the two estimates as approximate upper and lower bounds of the amount of the differential that is explained by compositional differences.

Although the model underlying our decomposition involves a linear relationship between the covariates and the log odds of disruption, we further estimate overall effects of composition by computing average annual probabilities of marital disruption for each racial and ethnic group under three conditions. The average observed annual probability of marital disruption

experienced by a particular group (using that group's own composition) is calculated as well as estimated average probabilities were a group (e.g. Whites) to possess the same set of compositional characteristics as either of the other two groups (e.g. Blacks or Mexicans). In particular, for each race- and ethnic-specific model, we obtain the predicted probability of a marital disruption for all women in the sample and then calculate means of these probabilities for each of the three racial and ethnic groups (for more detail, see Fairlie 1999, 2003).

The regression coefficients applied in the analyses described above are estimated using discrete-time survival models (Allison, 1995). The dependent variable in these models is an indicator of whether disruption occurred in a particular marital duration year. We focus on marital disruption (divorce or separation) rather than divorce alone, as there are well-known differences among racial and ethnic groups in the tendency to divorce after separation (Bean and Tienda, 1987; Bramlett and Mosher, 2002; Sweet and Bumpass, 1987). The process of disruption is examined during the first ten years of marriage only, and women married longer than ten years are censored. Models are estimated separately for White, Black, and Mexican American women to allow the effects of risk factors for divorce to vary across groups. All models include controls for year of marital duration and marriage cohort. The stratified and clustered sampling design of the 1995 NSFG is adjusted using STATA's svylogit routine.

**Results**

*Differences in Risk of Disruption and Composition by Race and Ethnicity.*

Figure 1 displays the annual probabilities of disruption currently experienced by women of different racial and ethnic backgrounds, confirming differentials identified by prior research. Among the three major groups examined, Black women experience the highest annual

probability of disruption, followed by White women, with Mexican women exhibiting the lowest risk of disruption. However, the probability of disruption for Mexican women as a group hides important distinctions by nativity. White women have a substantially higher probability of disruption relative to foreign-born Mexican Americans, but their risk of disruption is *less* than that of U.S.-born Mexican American women. Sample size restrictions preclude a thorough investigation of this important nativity differential in levels of marital instability among Mexican Americans, but we return attention to this issue at the conclusion of our analysis.

[FIGURE 1 ABOUT HERE]

Table 1, which shows the mean values for the compositional measures considered in our analysis, highlights important racial and ethnic differences in the characteristics of women at the time of their first marriage. Consistent with findings from prior work, Black women have an older average age at first marriage in these data, whereas Mexicans tend to marry at much younger ages. Whites are most highly educated, Mexicans the least educated, with Blacks falling in the middle. White and Black women are more likely to be employed full-time than are Mexican women, but Mexican and Black women are more likely to be unemployed than are White women.

Although Black women are far more likely to have experienced a premarital birth than are Mexican or White women, large racial and ethnic differences in the level of premarital conception (that is, having a first birth take place within seven months of the date of first marriage) do not exist. Significant differences do exist, however, in patterns of premarital sexual activity. Approximately one third of women across all three racial and ethnic groups report having had premarital sex only with their future husband, but Black and White women are far more likely than Mexican women to have had sex before marriage with someone other than their

husband. In contrast, Mexican women are more likely to report having been virgins at marriage. In addition, cohabitation before marriage is more common among Whites and Blacks than among Mexicans. The regional concentration of these racial/ethnic groups is apparent, with Mexicans most likely to reside in the West and the majority of Black women dwelling in the South.

[TABLE 1 ABOUT HERE]

Differences by race and ethnicity in family background and spousal attributes are also evident. Almost half of the Mexican women in our sample were born outside the United States, compared to just 3.6% of Whites and 5.2% of Blacks. Blacks are more likely to report that their parents were divorced by age 14 than are Whites or Mexicans. The educational attainment of White women's mothers exceeds that of the other two groups, and Mexicans are far more likely to have been raised as a Catholic than are Whites or Blacks. Black men, as is the case for Black women, tend to be older at the time of their first marriage while Mexican men are considerably younger. However, there are not substantial differences by race or ethnicity in the measures of age heterogamy. The vast majority of women across all groups marry men who are no more than two years younger or five years older than they are, but Black and Mexican women are more likely than White women to report marrying a man more than two years younger than themselves. Mexican women are least likely to have a husband who was married before and of the same race compared to Whites and Blacks.

[TABLE 2 ABOUT HERE]

As noted earlier, many of these compositional factors are known to be important determinants of marital disruption (e.g. Castro Martin and Bumpass 1989). Indeed, the discrete-time logistic regression models of these compositional factors on the risk of marital disruption

confirm this point, although the results also suggest that the relative strength and statistical significance of the associations vary across racial and ethnic groups (see Table 2). For example, marrying at an older age is associated with protective effects against marital disruption for both Whites and Blacks, but not for Mexicans. Greater levels of educational attainment are associated with more stable marriages, although significantly so only for White women. Premarital cohabitation with either someone other than one's spouse or with both one's spouse and someone else is associated with a significant destabilizing effect on marriage for Whites, but not for Blacks or Mexicans, perhaps because of differences in the meaning and function of cohabitation across race and ethnicity (Phillips and Sweeney, forthcoming). With regard to family background characteristics, being foreign-born is associated with a substantial reduction in the odds of disruption among Mexican women (by almost 70%) and Black women (by about 67%), but not among White women. Although experiencing the divorce of one's parents by age 14 is associated with an elevated risk of disruption for White and Mexican women, no significant effect is found among Blacks. Marrying an older man (older than 27) is associated with a reduced risk of disruption for White and Mexican women, but not for Black women. Marrying a man of the same race is associated with lower rates of disruption, but again, significantly so only for White and Mexican women.[2]

Given the association between many of these characteristics and marital disruption, as well as the striking differences among women of different racial and ethnic backgrounds with respect to these factors at the time of marriage, we expect that racial and ethnic variation in population composition will contribute to the observed differences in the risk of disruption across groups. Yet it is important to keep in mind that the effects of some of the various

[2] See Phillips and Sweeney, forthcoming, for a more detailed discussion of racial and ethnic variation in the process of marital disruption.

compositional factors on racial and ethnic gaps in disruption may offset one another. For example, that Black women tend to marry at older ages than do White and Mexican women should reduce their risk of disruption relative to these groups, while their greater likelihood of premarital childbearing ought to increase their relative risk of disruption. In a similar fashion, Mexican women's lower levels of education should raise their relative risk of disruption compared to White and Black women, but their lower likelihood of having sexual intercourse before marriage ought to reduce their relative risk.[3]

*Contribution of Individual Risk Factors to Racial and Ethnic Differentials*

To quantify the contribution that differences in particular compositional factors make in explaining observed disruption differentials, a regression decomposition approach is applied. The various racial and ethnic differentials are decomposed into the percentages explained (i.e. due to compositional differences) and unexplained. These decomposition results are displayed in Table 3 and are based on the estimated parameters displayed in Table 2 and weighted means for person-years at risk from 1975 to 1994 (not shown). For those characteristics with a positive percentage contribution reported, the differential would be *reduced* by that percentage if racial and ethnic compositional differences in that risk factor were eliminated. For risk factors with a negative percentage contribution, the gap in disruption is estimated to *increase* if racial and ethnic variation in the composition of that characteristic were removed.

[TABLE 3 ABOUT HERE]

[3] These predictions are based on models of disruption for the U.S. population as a whole. To the extent that the effects of these risk factors vary by race and ethnicity, we will see variation in the ways factors counterbalance each other and the size of the offsetting effects.

*White-Black Differential.* The decomposition of the White-Black difference in the log odds of disruption is displayed in the first two columns of Table 3. We find that differences in composition with respect to all factors combined can account for 37% of the differential when White women are used as the standard (column 1) and 18% when Black women are used as the standard (column 2). Taking a simple average of these two estimates (following Oaxaca 1973), we can conclude that about 28% of the White-Black differential in disruption is attributable to compositional differences. Most of the differential explained by variation in composition is attributable to differences in the respondent characteristics of Black and White women, which account for just over 31% of the gap when Whites are used as the standard and about 13% when Black women are used as the standard. Differences in some of these compositional characteristics, such as higher educational attainment and lower likelihood of having a sexual history before marriage among White women, increase the size of the racial gap in disruption. However, these effects are counterbalanced by other compositional differences, such as Black women's older average age at first marriage, that reduce the racial differential.

Overall, racial differences in family background explain virtually none of the total differential in disruption between Whites and Blacks, although the aggregate figure conceals some important points when black women are used as the standard. According to these estimates (column 2), the fact that Black women are less likely to be raised as Catholics increases the gap in disruption, but this positive contribution of religious upbringing is counteracted by the negative effect of foreign-born status and mother's education on the racial differential. However, note that since religious upbringing is not statistically significant in the model of disruption for Blacks, the primary effect of differences in family background characteristics between Blacks and Whites is to reduce the differential. Differences between Black and White

women in their spouses' characteristics account for little of the overall racial differential in disruption, about 3% and 5% when White and Black women are used as the standard, respectively.

*White-Mexican Differential.* The decomposition of the White-Mexican differential is shown in columns 3 and 4. The estimates presented pertain to the differential between White women and all Mexican women combined (both U.S.-born and foreign-born).[4] Again, ignoring nativity among Mexican women, White women have a somewhat *higher* risk of marital disruption than do Mexican women. Estimates of the contribution of composition to the overall disruption differential differ drastically depending on whether the White or Mexican process of disruption is assumed. Using Whites as the standard, our estimates reveal that if ethnic differences in composition were eliminated, the White-Mexican disruption gap would increase by slightly more than half (52.7%) (column 3). In other words, if White women had the same compositional characteristics as Mexican women, their risk of disruption would rise, causing the White-Mexican differential to expand. Using Mexicans as the standard (column 4), however, we estimate that in the absence of compositional differences, the disruption gap would be *reduced* by about 66%.

Whether Whites or Mexicans are used as the standard, we find that differences between the two groups in many respondent characteristics increase the White-Mexican disruption differential. Educational attainment, for example, protects against disruption for both groups, and White women have higher levels of educational attainment. Put another way, given the low levels of educational attainment among Mexican women, we would expect disruption rates among this group to be higher than that observed. Interestingly, the effects on the differential of

[4] Due to small sample sizes, we cannot estimate separate models for U.S.-born and foreign-born Mexican American women.

some characteristics, such as premarital cohabitation experience and region of residence, differ depending on which group is adopted as the standard. If White women had the same level of premarital cohabitation as do Mexican women (using Mexicans as the standard population), the differential would be reduced by about 6%; were Mexican women to have the average cohabitation experience of Whites (using Whites as the standard population), the differential would increase by about 13%. This pattern is due to the fact that premarital cohabitation is associated with an increased risk of disruption for White women, but a *decreased* risk for Mexican women. Note that differences in premarital sexual behavior explain a large portion of the White-Mexican differential. If differences between Whites and Mexicans in premarital sexual activity with someone other than a future spouse were eliminated, the differential in disruption rates would decrease by approximately 62% using Whites as the standard and by about 21% using Mexicans as the standard.

Although family background characteristics slightly increase the differential when Whites are used as the standard (by 9%), they explain about 81% of the differential when Mexicans are used as the standard, primarily because of the much lower risk of disruption experienced by Mexicans who are foreign-born. If ethnic differences in the proportion foreign-born were removed, the White-Mexican disruption gap would disappear (reduced by about 110%), holding everything else constant.[5] This pattern explains the vastly different estimates of the overall contribution of composition to the White-Mexican differential discussed above. Ignoring foreign-born status, the results of the two sets of estimates are much more similar – compositional differences between Whites and Mexicans increase the size of the differential so if

[5] The effect of foreign-born explains much less of the differential when whites are used as the standard because there is essentially no effect of foreign-born status on the risk of disruption among white women.

Whites and Mexicans had the same characteristics at the time of first marriage, the White-Mexican differential would be even larger.[6]

Note in addition that compositional differences across groups in the level of mother's education affect the differential differently, depending on whether Whites or Mexicans are adopted as the standard. Assuming the Mexican process of disruption, the gap in disruption is estimated to increase by about 61% were differences in mother's education level eliminated, but the differential is reduced by about 9% using estimates of the White process of disruption. This pattern can be attributed to the fact that greater levels of mother's educational attainment tend to reduce the risk of disruption among Mexicans, and the mothers of White women are far more likely to have a high school degree or more relative to the mothers of Mexican women.

Finally, we find that assortative mating characteristics increase the differential regardless of which group is used as the standard. For example, if White women married men of the same race in the same proportion as do Mexican women, the disruption differential would be increased by 7.8%. If Mexican women had white women's characteristics in this regard, the gap would rise by about 18%.

[6] Additional decompositions (not shown) of the White - U.S-born Mexican differential and the White - foreign-born Mexican differential reinforce this point. Recall that White women have lower rates of disruption than U.S.-born Mexicans but higher rates of disruption compared to foreign-born Mexicans. The White - U.S-born Mexican differential is more than explained by compositional factors; regardless of the standard adopted, about 170% of the differential is explained by composition. In other words, were White women to possess the same compositional characteristics as U.S.-born Mexican women, their rates of marital disruption would substantially exceed those currently exhibited by U.S.-born Mexicans. Estimates of what happens to the White - foreign-born Mexican differential vary depending on whether Whites or Mexicans are adopted as the standard. Assuming the White process of disruption, the gap would increase but assuming the Mexican process, the differential would decline.

*Black-Mexican Differential.* Turning finally to the Black-Mexican disruption differential, we find that 55.7% (using Blacks as the standard - Column 5) and 47.6% (using Mexicans as the standard – column 6) of the differential between these groups can be explained by compositional differences. Although the size of the contribution of particular compositional factors differs depending on whether Blacks or Mexicans are used as the standard, the overall conclusions are similar. For example, both estimates suggest that differences by race and ethnicity in certain respondent characteristics, such as Black women's older average age at marriage and greater levels of education, reduce the gap in disruption. However, that Mexican women are less likely to have a premarital birth and to have had premarital sex than are Black women raises the gap in disruption. Similarly, differences with respect to family background, such as the greater likelihood that Mexican women are foreign-born and raised as Catholics, increase the differential. Taken together, ethnic variation in spousal characteristics increases the differential regardless of which group is adopted as the standard.

[FIGURE 2 ABOUT HERE]

*Overall Contribution of Composition to Gaps in Disruption Probabilities*

To provide perhaps a more intuitive picture of the overall effect of compositional differences on racial and ethnic gaps in disruption, we demonstrate how the average annual probabilities of marital disruption would be expected to change if women possessed varying sets of compositional characteristics.[7] The first three columns of Figure 2 show the estimated annual probability of disruption for White women under three alternate conditions – when they are

[7] Covariates are assumed to have a nonlinear relationship with the probability of disruption in a logit model, and thus the average predicted probability of disruption in the sample will not necessarily equal the estimated probability of disruption for an individual with the sample average values on all covariates (see also Fairlie 1999, 2003).

exposed to their own set of compositional characteristics and when they are exposed to either Black or Mexican women's composition. The average annual probability of marital disruption currently experienced by White women is estimated to be about 0.039, but would increase by about 17% to 0.045 and 0.046 were Whites to possess the same composition as either Blacks or Mexicans, respectively. This rise in the probability of marital disruption for Whites would be expected to cause the White-Black differential to decline, but the White-Mexican differential would actually increase. Black women (second set of columns in Figure 2) currently experience an annual probability of disruption of 0.068, but that risk would be expected to decline by about 2% to 0.067 and by about 35% to 0.044 if Blacks had either the composition of Whites or Mexicans, respectively. Under both of these scenarios, the White-Black and Black-Mexican gaps would be reduced. Referring to the third set of columns, it is evident that Mexican women currently experience the lowest risk of marital disruption (an annual probability of 0.031), but that risk would be expected to *increase* by 29% and 68% if Mexicans had the same characteristics as White or Black women, respectively. In this case where the Mexican process of disruption is assumed, our results suggest that both the White-Mexican and Black-Mexican disruption gaps would diminish were compositional differences across groups to be eliminated.

*Reconsidering Nativity*

As noted above, there is a wide gap in disruption as well as important differences in population composition at marriage between U.S.-born and foreign-born Mexican women (refer to Figure 1 and appendix Table A1). For example, U.S-born Mexican women typically have a higher level of education than do foreign-born Mexican women, and are more likely to be employed, to have cohabited or had a child before marriage, and to have experienced a

disruption in their family of origin by age 14. That is, U.S.-born Mexican women more closely resemble White women in terms of population composition than do Mexican women born outside the United States. Thus, we conclude by reconsidering the issue of nativity. Specifically, we investigate the degree to which the disruption differential between U.S.-born and foreign-born Mexican women can itself be explained by compositional differences across groups. How much might we expect the annual probability of disruption to change for these two groups if U.S.-born Mexican Americans were to possess the same population composition as foreign-born Mexican Americans, and vice versa?[8] Although sample size restrictions preclude a complete formal decomposition of these differences, as we cannot separately estimate the effects of our full set of covariates for U.S.-born and foreign-born Mexican women, we can address the basic issue of how the overall levels of disruption for these two groups might be expected to change, were the compositional characteristics of each group to be altered.

[FIGURE 3 ABOUT HERE]

The results of our exploratory analysis of nativity differences in disruption are shown in Figure 3. In short, we find that the average annual probability of disruption for U.S.-born Mexicans would be expected to decrease by approximately 25% (from .051 to .038), were they to share the same compositional characteristics as foreign-born Mexicans. We would expect the annual probability of disruption among foreign-born Mexicans to increase by about 42% (from .012 to .017), were they to possess the same compositional characteristics as U.S.-born Mexicans. Thus, a large differential in disruption would persist, even if compositional differences between foreign-born and U.S.-born Mexican women were eliminated. This pattern

[8] The regression coefficient estimates obtained from the pooled Mexican American model shown in Table 2 are used in this decomposition since sample sizes are not large enough to estimate separate regression equations for Mexican Americans by nativity status.

reinforces the point that foreign-born status has a strong protective effect against marital disruption among Mexican American women – an effect that cannot be easily explained by compositional differences between these two groups of women.

**Discussion**

Our analysis confirms important variation among non-Hispanic White, non-Hispanic Black and Mexican American women in their characteristics at the time of first marriage. In some respects, White women tend to possess certain characteristics at marriage that ought to lower their risk of disruption relative to other groups; they are more educated and less likely to have had a premarital birth than their Black or Mexican American counterparts, for example. Yet they are also typically younger when they first marry than are Black women and are more likely to marry a man who was previously married – factors that are associated with an increased risk of disruption. Mexican women are less likely to have had premarital sex before marriage and more likely to be born outside the United States, characteristics that are associated with a reduced risk of disruption, but on average they are also less educated and marry at a young age – factors that are associated with increased marital instability.

To investigate the potential contribution of such compositional differences between populations to racial and ethnic variation in levels of disruption, we apply a regression decomposition analysis. With regard to the White-Black differential, our findings reveal that about 28% of the White-Black gap in the log odds of disruption can be explained by differences in the composition of these populations at the time of marriage. Differences between White and Black women in characteristics such as educational attainment, the nature of premarital sexual behavior and premarital cohabitation, which raise the differential, are offset to some extent by

other compositional differences that reduce the gap, such as Black women's older age at marriage. We find that about half of the Black-Mexican gap in the log odds of disruption can be explained by compositional differences between the two groups. Racial and ethnic differences in family background characteristics, such as foreign-born status, are important contributors to the Black-Mexican differential. Likewise, differences between the two groups in premarital sexual activity explain a significant portion of the gap in disruption. On the other hand, variation in spousal attributes across the two groups tends to increase the disruption differential between Blacks and Mexicans. Although differences in the likelihood of a premarital birth and being raised as a Catholic between the two groups appear to explain a fairly large part of the differential, these particular covariates do not have a statistically significant effect on the risk of disruption for either Blacks or Mexicans, and therefore, their contribution in the decomposition analysis should be interpreted cautiously.

The story is more complicated for White and Mexican women, and hinges strongly on nativity status of Mexican women. In the absence of differences in nativity, we conclude that the White-Mexican differential in disruption would actually increase were compositional differences eliminated. These findings are consistent with the purported "paradox of Mexican American nuptiality" (Oropesa, Lichter and Anderson 1994). Mexican Americans are at least as likely to marry as Whites, and tend to do so at relatively younger ages, despite a disadvantaged economic position. As noted earlier, some argue that a familistic cultural heritage may contribute to greater levels of marriage and marital stability, particularly among Mexican immigrant women (Chavez 1991; Fukuyama 1993; Oropesa et al. 1994; Oropesa 1996; Vega 1990; but see Bean et al. 1996; Wildsmith 2002). Certainly, there is evidence that Mexican American women are more likely to report marriage and children as their "major life objective" (Blea 1992), and express

more pro-nuptial attitudes in surveys than do non-Hispanic Whites or Blacks (Oropesa 1996; Oropesa and Gorman 2000; Trent and South 1992).

Further indirect evidence is provided in support of the possible role of cultural differences when we look at the decomposition results of the disruption gap between U.S.-born and foreign-born Mexican women. The substantial differences in population composition between these two groups explain only about a third of the total differential and point to the critical role of foreign-born status in promoting marital stability among Mexican Americans. The mechanisms through which nativity produces lower rates of disruption clearly merit further research. Perhaps, as hypothesized above, nativity captures different cultural values between the two groups that promote marital stability, such as a familistic orientation. Recent research attributes other positive outcomes among foreign-born Mexican Americans, such as lower rates of mental illness, to cultural differences, namely close-knit families with many extended family members who provide social and psychological support (Levin, 2005). On the other hand, it may be that foreign-born status contributes to greater levels of stress and uncertainty with regard to U.S. society, factors that may discourage first-generation immigrants from leaving a union. Yet another possibility relates to selection effects. Recent research indicates, for example, that the lower mortality levels of Mexican Americans are due to a "salmon-bias" effect – return migration of foreign-born Mexicans to Mexico when they are ill (Palloni and Arias, 2004). If a similar process is at play with regard to marital disruption – that is, if some foreign-born Mexican women return to their country of origin upon divorce – the effect of nativity may be overstated in our results.

It is also notable that the substantial array of compositional factors considered explains no more than about one quarter of the overall White-Black gap in marital disruption. Heaton and

Jacobson (1994) observe that the types of structural and cultural obstacles faced by Blacks may simply not be captured well by the more basic measures of composition incorporated here. Wilson (1987), for example, has noted the detrimental effects on marriage and marital stability of structural conditions facing the Black population, such as poor labor market opportunities and conditions for Black men. As noted earlier in the paper and may be the case for Mexican Americans, cultural issues may also play an important role in the disruption process among Blacks.

Our findings as a whole suggest that compositional differences between groups do not explain substantial portions of the racial and ethnic gaps in disruption, indicating that a good part of the explanation lies elsewhere. Certainly, arguments citing cultural differences between groups suggest that the compositional factors included in this analysis are not likely to be important determinants of marital stability among Mexican or Black women and thus not a powerful explanation for the racial and ethnic differentials in disruption. Our work provides indirect support for some of these arguments, but future research should work towards incorporating measures of cultural attitudes to test explicitly the notion of cultural values contributing to varying risks of marital disruption among Mexicans and Blacks. Although previous work examines the role of cultural value orientations in explaining racial and ethnic differences in marriage (Oropesa et al. 1994), less is known about their contribution to explaining differentials in marital disruption.

These issues point to several other limitations of the current study. This analysis examines a more complete set of covariates associated with disruption than previous investigations, but there remain potentially important factors that are not included due to data or statistical constraints. For example, potential problems of endogeneity preclude the inclusion of

time-varying measures of employment, although future work should consider such measures of both wife's and husband's employment status over time. In light of research indicating that contextual factors may influence divorce (e.g. South and Lloyd 1995) and as alluded to above with regard to Blacks, future research should incorporate such measures into analyses. Issues related to sample size prevent a thorough analysis of Mexican Americans by generational status, but given the important differences in levels of disruption by nativity, it will clearly be important in future research to examine foreign-born and U.S.-born Mexicans separately. Finally, although the decomposition approach is useful in providing estimates of how much of a given racial/ethnic differential is explained (that is, due to compositional differences), the approach does not distinguish in the unexplained portion between the contribution of differences in process and differences due to the interaction between composition and process. It is possible that as a group's population composition changes, the dynamics of the marriage and disruption process adjust in ways that are difficult to determine. Such changes would alter the estimates of compositional contribution presented here.

Despite these limitations, this study improves our understanding of the determinants of racial and ethnic gaps in marital disruption. Certainly, differences in population composition at the time of marriage do contribute to racial and ethnic differentials in disruption, although their relative importance varies across particular racial and ethnic group comparisons. In no case, however, can varying exposure to risk factors fully explain the racial and ethnic gaps in marital disruption considered here. The important and largely unexplained role of nativity status in promoting marital stability among Mexican Americans is especially noteworthy, and should be a key focus of future research as appropriate data with large sample sizes of Mexican Americans become available.

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Table 1. Mean Values of Covariates in Year of Marriage, by Race/Ethnicity: 1995 NSFG.

| Independent Variable | Non-Hispanic Whites | Non-Hispanic Blacks | Mexican Americans |
|---|---|---|---|
| Age at First Marriage | | | |
| < 20 years | 0.275 | 0.241 | 0.438 [ab] |
| 20-22 years | 0.430 | 0.390 | 0.360 [a] |
| 23-26 years | 0.180 | 0.199 | 0.126 [ab] |
| 27-30 years | 0.115 | 0.169 [a] | 0.077 [ab] |
| Education | | | |
| < 12 years | 0.137 | 0.166 | 0.446 [ab] |
| 12 years | 0.380 | 0.421 | 0.338 [b] |
| 13-15 years | 0.292 | 0.290 | 0.174 [ab] |
| ≥ 16 years | 0.191 | 0.122 [a] | 0.042 [ab] |
| Employment Status | | | |
| Unemployed | 0.179 | 0.258 [a] | 0.381 [ab] |
| Full-Time | 0.635 | 0.596 | 0.488 [ab] |
| Part-Time | 0.133 | 0.116 | 0.111 |
| Both Full- and Part-Time | 0.053 | 0.030 [a] | 0.021 [a] |
| Any Premarital Birth | 0.090 | 0.426 [a] | 0.146 [ab] |
| Any Premarital Conception | 0.111 | 0.115 | 0.158 [a] |
| Premarital Sexual History | | | |
| None | 0.122 | 0.061 [a] | 0.414 [ab] |
| Husband only | 0.303 | 0.312 | 0.349 |
| Someone other than husband | 0.575 | 0.627 [a] | 0.236 [ab] |
| Premarital Cohabitation History | | | |
| None | 0.581 | 0.561 | 0.738 [ab] |
| Husband only | 0.349 | 0.382 | 0.241 [ab] |
| Both husband and someone else | 0.054 | 0.041 | 0.021 [a] |
| Someone other than husband | 0.016 | 0.015 | 0.000 [ab] |
| Region of Residence | | | |
| West | 0.198 | 0.099 [a] | 0.610 [ab] |
| Northeast/Midwest | 0.479 | 0.318 [a] | 0.065 [ab] |
| South | 0.323 | 0.583 [a] | 0.326 [b] |
| Metropolitan Area | 0.748 | 0.857 [a] | 0.901 [a] |
| *Family Background Characteristics* | | | |
| Born Outside the U.S. | 0.036 | 0.052 | 0.467 [ab] |
| Parents Divorced at Age 14 | 0.179 | 0.369 [a] | 0.119 [ab] |

(Continued)

Table 1. (Continued)

| Independent Variable | Non-Hispanic Whites | Non-Hispanic Blacks | Mexican Americans |
|---|---|---|---|
| Mother's Education | | | |
| < 12 years | 0.223 | 0.423 [a] | 0.741 [ab] |
| 12 years | 0.507 | 0.381 [a] | 0.162 [ab] |
| 13-15 years | 0.149 | 0.099 [a] | 0.069 [a] |
| ≥ 16 years | 0.121 | 0.097 | 0.027 [ab] |
| Raised as a Catholic | 0.326 | 0.111 [a] | 0.843 [ab] |
| | | | |
| *Assortative Mating Characteristics* | | | |
| Husband's Age at Marriage | | | |
| < 20 years | 0.112 | 0.093 | 0.220 [ab] |
| 20-23 years | 0.381 | 0.350 | 0.396 |
| 24-26 years | 0.163 | 0.149 | 0.130 |
| ≥ 27 years | 0.345 | 0.408 [a] | 0.253 [ab] |
| Husband's Education at First Marriage | | | |
| < 12 years | 0.128 | 0.177 [a] | 0.467 [ab] |
| 12 years | 0.433 | 0.472 | 0.332 [ab] |
| 13-15 years | 0.220 | 0.219 | 0.133 [ab] |
| ≥ 16 years | 0.218 | 0.131 [a] | 0.068 [ab] |
| Age Heterogamy | | | |
| Husband > 5 Yrs Older than Wife | 0.200 | 0.225 | 0.206 |
| Husband's Age W/in -2 to 5 Yrs of Wife | 0.759 | 0.708 [a] | 0.716 |
| Husband > 2 Yrs Younger than Wife | 0.042 | 0.066 [a] | 0.078 [a] |
| Education Heterogamy | | | |
| Husband More Educated than Wife | 0.243 | 0.259 | 0.211 |
| Husband/Wife same education | 0.514 | 0.424 [a] | 0.552 [b] |
| Husband Less Educated than Wife | 0.243 | 0.317 [a] | 0.237 [b] |
| Husband Married Before | 0.159 | 0.150 | 0.100 [ab] |
| Husband/Wife of Same Race | 0.941 | 0.946 | 0.795 [ab] |
| | | | |
| Marriage Cohort | | | |
| 1975-1979 | 0.257 | 0.206 [a] | 0.215 |
| 1980-1984 | 0.268 | 0.275 | 0.240 |
| 1985-1989 | 0.244 | 0.271 | 0.274 |
| 1990-1994 | 0.230 | 0.247 | 0.271 |
| | | | |
| Number of Observations | 3,222 | 751 | 471 |

*Note.* Means are weighted.

[a] Mean differs significantly from that of non-Hispanic Whites, $p<.05$ level.

[b] Mexican American mean differs significantly from that of non-Hispanic Blacks, $p<.05$ level.

Table 2. Discrete-Time Logistic Regression Estimated Effects of Covariates on the Risk of Marital Disruption within 10 Years, by Race/Ethnicity: 1995 NSFG

| Independent Variable | Non-Hispanic Whites | | Non-Hispanic Blacks | | Mexican Americans | |
|---|---|---|---|---|---|---|
| | Odds Ratio | Coeff./S.E. | Odds Ratio | Coeff./S.E. | Odds Ratio | Coeff./S.E. |
| *Respondent Characteristics* | | | | | | |
| Age at First Marriage *(<20 years)* | | | | | | |
| 20-22 years | 0.615 [a] | -4.548 | 0.731 | -1.330 | 1.159 | 0.464 |
| 23-26 years | 0.547 [a] | -3.350 | 0.386 [a] | -2.536 | 1.226 | 0.390 |
| 27-30 years | 0.383 [a] | -3.633 | 0.377 [a] | -2.446 | 0.379 | -1.123 |
| Education *(< 12 years)* | | | | | | |
| 12 years | 0.672 [a] | -2.007 | 0.713 | -1.037 | 1.106 | 0.181 |
| 13-15 years | 0.556 | -1.831 | 0.567 | -1.036 | 0.685 | -0.341 |
| ≥ 16 years | 0.398 [a] | -2.005 | 0.556 | -0.818 | 1.080 | 0.047 |
| Employment Status *(Not Employed)* | | | | | | |
| Full-Time | 0.932 | -0.636 | 1.110 | 0.566 | 1.093 | 0.302 |
| Part-Time | 0.579 [a] | -3.863 | 1.001 | 0.004 | 1.021 | 0.049 |
| Both Full- and Part-Time | 0.832 | -0.863 | 1.643 | 1.021 | 0.305 | -0.974 |
| Any Premarital Birth | 1.099 | 0.694 | 1.298 | 1.491 | 1.626 | 1.435 |
| Any Premarital Conception | 0.994 | -0.061 | 2.002 [ab] | 2.953 | 1.174 | 0.385 |
| Premarital Sexual History *(None)* | | | | | | |
| Husband only | 1.595 [a] | 2.530 | 1.278 | 0.867 | 1.718 | 1.907 |
| Someone other than husband | 2.901 [a] | 6.015 | 1.837 [a] | 2.051 | 1.415 | 0.817 |
| Any Cohabitation before Marriage *(None)* | | | | | | |
| Husband only | 1.170 | 1.693 | 0.921 | -0.502 | 0.699 | -1.134 |
| Both husband and someone other than husband | 1.535 [a] | 2.795 | 1.198 | 0.473 | 0.468 | -0.721 |
| Region *(West)* | | | | | | |
| Northeast/Midwest | 0.902 | -0.977 | 0.772 | -0.909 | 1.189 | 0.440 |
| South | 1.217 | 1.857 | 0.904 | -0.362 | 0.967 | -0.115 |
| Metropolitan Area | 1.274 [a] | 2.497 | 1.075 | 0.341 | 1.073 | 0.196 |
| *Family Background Characteristics* | | | | | | |
| Born Outside the U.S. | 1.181 | 0.729 | 0.327 [ab] | -2.031 | 0.305 [ab] | -3.153 |
| Parents Divorced by Age 14 | 1.194 [a] | 1.994 | 1.073 | 0.428 | 1.786 [a] | 2.117 |
| Mother's Education *(< 12 years)* | | | | | | |
| 12 years | 1.029 | 0.263 | 1.467 [a] | 1.967 | 0.603 [c] | -1.269 |
| 13-15 years | 1.263 | 1.690 | 1.490 | 1.215 | 0.802 | -0.356 |
| ≥ 16 years | 1.268 | 1.623 | 1.684 | 1.638 | 0.245 [c] | -1.595 |
| Raised as a Catholic | 1.047 | 0.559 | 0.673 | -1.307 | 0.756 | -0.834 |

(Continued)

Table 2. (Continued)

| Independent Variable | Non-Hispanic Whites | | Non-Hispanic Blacks | | Mexican Americans | |
|---|---|---|---|---|---|---|
| | Odds Ratio | Coeff./S.E. | Odds Ratio | Coeff./S.E. | Odds Ratio | Coeff./S.E. |
| *Assortative Mating Characteristics* | | | | | | |
| Husband's Age at First Marriage *(<20 yrs)* | | | | | | |
| 20-23 years | 0.894 | -0.958 | 0.742 | -1.106 | 0.701 | -0.894 |
| 24-26 years | 0.734 | -1.926 | 0.773 | -0.754 | 1.024 | 0.044 |
| ≥ 27 years | 0.656 [a] | -2.154 | 0.926 | -0.173 | 0.227 [a] | -2.149 |
| Husband's Education at First Marriage *(<12 yrs)* | | | | | | |
| 12 years | 1.210 | 0.898 | 1.092 | 0.278 | 1.133 | 0.187 |
| 13-15 years | 0.819 | -0.564 | 0.683 | -0.757 | 0.943 | -0.049 |
| ≥ 16 years | 0.902 | -0.212 | 0.649 | -0.544 | 1.834 | 0.354 |
| Age Heterogamy *(Husband > 5 Yrs Older)* | | | | | | |
| Husband's Age W/in -2 to 5 Yrs of Wife | 0.829 | -1.418 | 1.686 [b] | 1.926 | 0.582 | -1.066 |
| Husband > 2 Yrs Younger than Wife | 0.933 | -0.277 | 2.379 | 1.653 | 0.515 | -0.839 |
| Education Heterogamy *(Husband More Educated)* | | | | | | |
| Husband/Wife same education | 1.166 | 0.734 | 1.450 | 1.037 | 1.012 | 0.017 |
| Husband Less Educated than Wife | 1.450 | 0.985 | 1.723 | 0.878 | 0.858 | -0.120 |
| Husband Married Before | 1.332 [a] | 2.667 | 1.318 | 1.015 | 1.179 | 0.309 |
| Husband/Wife of Same Race | 0.691 [a] | -2.565 | 0.684 | -0.820 | 0.417 [a] | -2.252 |
| Number of Person Years | 22,063 | | 4,626 | | 3,282 | |

*Note.*

[a] Coefficient differs significantly from zero, p <0.05 level.

[b] Coefficient differs significantly from that for Non-Hispanic Whites, p <0.05 level.

[c] Mexican coefficient differs significantly from that for Non-Hispanic Blacks, p <0.05 level.

Table 3. Percentage of Racial Differences in the Expected Log Odds of Marital Disruption Due to Differences in Composition, Based on Person Years at Risk, 1975-1994.

| Independent Variable | White-Black Differential | | White-Mexican Differential | | Black-Mexican Differential | |
|---|---|---|---|---|---|---|
| | 1<br>White as Standard | 2<br>Black as Standard | 3<br>White as Standard | 4<br>Mexican as Standard | 5<br>Black as Standard | 6<br>Mexican as Standard |
| *Respondent Characteristics* | *31.18%* | *13.47%* | *-24.65%* | *1.47%* | *-13.09%* | *11.34%* |
| Age at First Marriage | -6.35% | -9.20% | -22.14% | -3.12% | -15.04% | -5.45% |
| Education | 7.48% | 4.47% | -46.84% | -5.14% | -12.82% | -2.44% |
| Employment Status | 2.32% | -2.60% | -7.82% | -5.08% | 1.15% | 0.33% |
| Any Premarital Birth | 4.46% | 12.30% | -0.72% | -3.69% | 6.10% | 11.39% |
| Any Premarital Conception | 0.00% | 0.19% | 0.06% | -1.41% | -2.53% | -0.59% |
| Premarital Sexual History | 10.06% | 5.75% | 62.22% | 20.87% | 18.65% | 10.91% |
| Premarital Cohabitation History | -0.18% | -0.83% | 6.46% | -13.17% | -0.67% | -5.77% |
| Region of Residence | 9.63% | 2.29% | -9.12% | 14.17% | -7.69% | 3.19% |
| Metropolitan Area | 3.77% | 1.12% | -6.74% | -1.96% | -0.24% | -0.23% |
| *Family Background Characteristics* | *0.65%* | *-2.69%* | *-9.37%* | *80.89%* | *78.47%* | *55.61%* |
| Born Outside the U.S. | 0.69% | -4.62% | -15.51% | 110.46% | 42.47% | 45.15% |
| Parents Divorced by Age 14 | 4.72% | 1.87% | 1.24% | 4.07% | 1.27% | 10.53% |
| Mother's Education | -3.31% | -12.49% | 9.43% | -61.32% | 10.59% | -17.09% |
| Catholic | -1.45% | 12.55% | -4.54% | 27.68% | 24.13% | 17.02% |
| *Assortative Mating Characteristics* | *2.68%* | *5.39%* | *-18.09%* | *-18.75%* | *-8.96%* | *-20.16%* |
| Husband's Age at First Marriage | -3.20% | 1.68% | -7.83% | -19.94% | -0.25% | -16.46% |
| Husband's Education at First Marriage | 2.37% | 6.27% | -2.64% | 20.93% | -4.33% | 4.47% |
| Age Heterogamy | 1.88% | -3.12% | -1.66% | -2.30% | -1.33% | 1.25% |
| Education Heterogamy | 2.17% | 1.12% | -0.54% | -0.42% | -0.25% | -1.29% |
| Husband Married Before | 0.02% | 0.02% | 2.35% | 1.35% | 0.99% | 0.59% |
| Husband/Wife of Same Race | -0.56% | -0.57% | -7.78% | -18.38% | -3.80% | -8.72% |
| Marriage Cohort | -0.46% | 2.09% | 0.45% | 4.02% | -0.53% | 1.09% |
| Duration of Marriage | 3.29% | 0.03% | -0.99% | -1.52% | -0.20% | -0.27% |
| *% Explained by Differences in Composition* | *37.34%* | *18.28%* | *-52.65%* | *66.11%* | *55.69%* | *47.61%* |

Note. Based on estimated parameters presented in Table 2 and weighted means for person-years at risk from 1975-1994.

**Figure 1: Average Annual Probability of Disruption Within 10 Years of Marriage, by Race and Ethnicity**

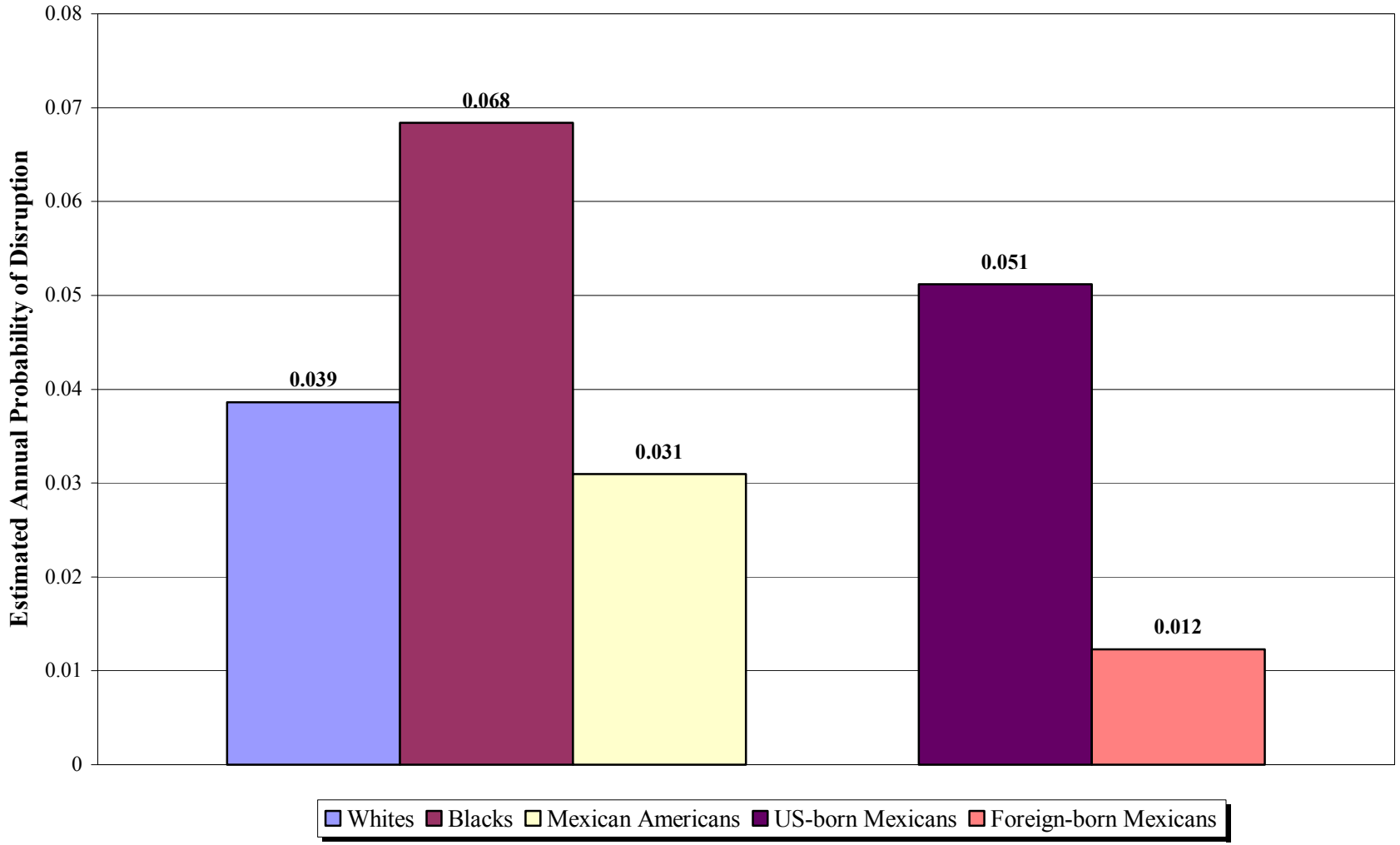

**Figure 2: The Contribution of Compositional Differences to Racial and Ethnic Variation in the Estimated Probability of Marital Disruption**

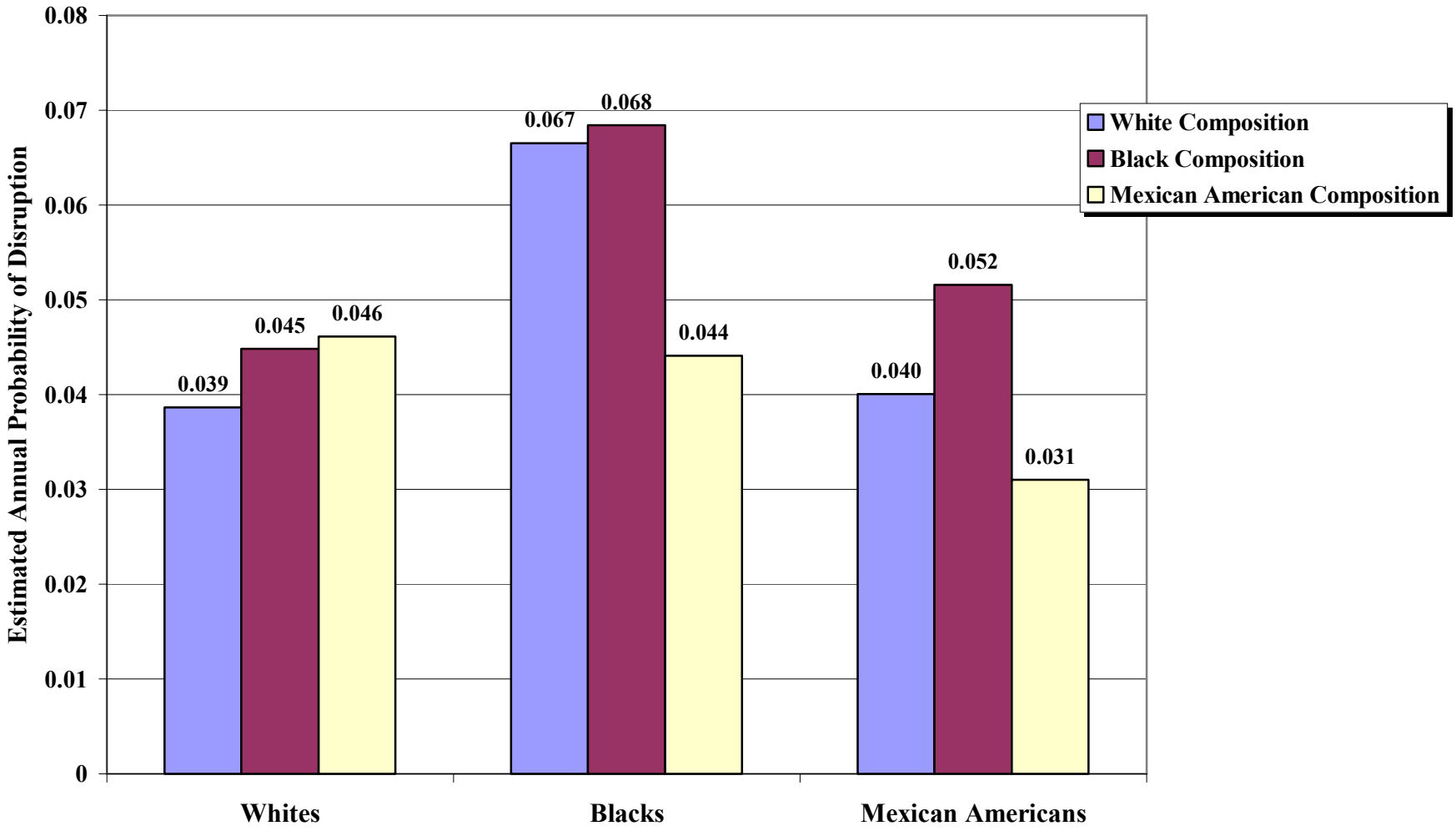

**Figure 3. The Contribution of Compositional Differences to Variation in the Estimated Probability of Disruption among Mexican Americans, by Nativity Status**

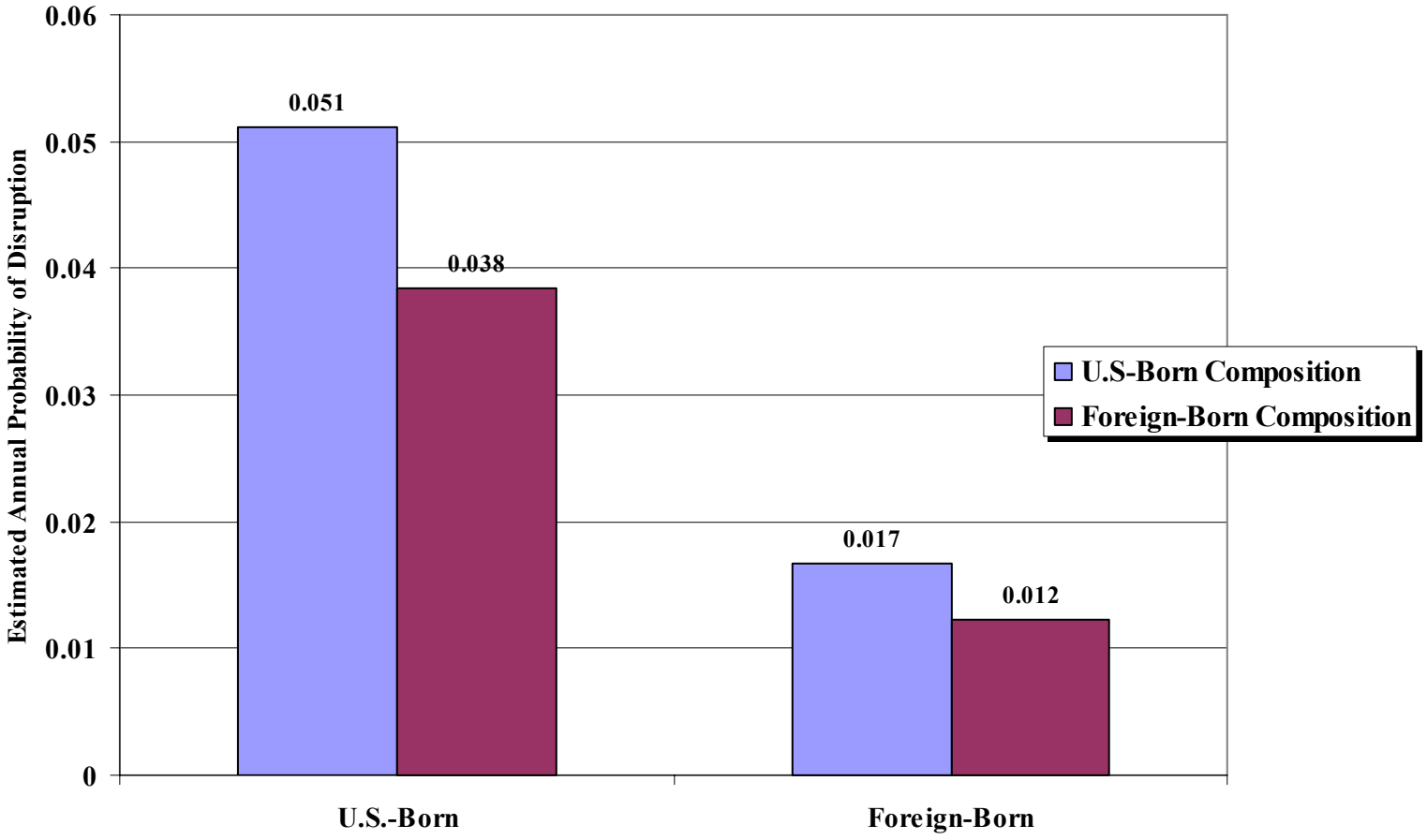

Table A1. Mean Values of Covariates in Year of Marriage, by Nativity: 1995 NSFG.

| Independent Variable | U.S-Born Mexican Americans | Foreign-Born Mexican Americans |
|---|---|---|
| Age at First Marriage | | |
| < 20 years | 0.411 | 0.469 |
| 20-22 years | 0.389 | 0.327 |
| 23-26 years | 0.138 | 0.112 |
| 27-30 years | 0.063 | 0.093 |
| Education | | |
| < 12 years | 0.309 [a] | 0.603 |
| 12 years | 0.402 [a] | 0.264 |
| 13-15 years | 0.231 [a] | 0.108 |
| ≥ 16 years | 0.058 | 0.024 |
| Employment Status | | |
| Unemployed | 0.283 [a] | 0.492 |
| Full-Time | 0.505 | 0.469 |
| Part-Time | 0.177 [a] | 0.035 |
| Both Full- and Part-Time | 0.035 [a] | 0.004 |
| Any Premarital Birth | 0.195 [a] | 0.090 |
| Any Premarital Conception | 0.175 | 0.138 |
| Premarital Sexual History | | |
| None | 0.237 [a] | 0.617 |
| Husband only | 0.376 | 0.319 |
| Someone other than husband | 0.387 [a] | 0.064 |
| Premarital Cohabitation History | | |
| None | 0.664 [a] | 0.822 |
| Husband only | 0.301 [a] | 0.172 |
| Both husband and someone else | 0.035 [a] | 0.006 |
| Someone other than husband | 0.000 | 0.000 |
| Region of Residence | | |
| West | 0.564 | 0.662 |
| Northeast/Midwest | 0.077 | 0.050 |
| South | 0.359 | 0.287 |
| Metropolitan Area | 0.886 | 0.919 |
| | | |
| *Family Background Characteristics* | | |
| Born Outside the U.S. | 0.000 [a] | 1.000 |
| Parents Divorced at Age 14 | 0.142 [a] | 0.093 |

(Continued)

Table A1. (Continued)

| Independent Variable | U.S.-Born Mexican Americans | Foreign-Born Mexican Americans |
|---|---|---|
| Mother's Education | | |
| < 12 years | 0.592 [a] | 0.912 |
| 12 years | 0.258 [a] | 0.053 |
| 13-15 years | 0.108 [a] | 0.025 |
| ≥ 16 years | 0.042 | 0.011 |
| Raised as a Catholic | 0.810 [a] | 0.882 |
| | | |
| *Assortative Mating Characteristics* | | |
| Husband's Age at Marriage | | |
| < 20 years | 0.244 | 0.193 |
| 20-23 years | 0.396 | 0.397 |
| 24-26 years | 0.125 | 0.136 |
| ≥ 27 years | 0.235 | 0.274 |
| Husband's Education at First Marriage | | |
| < 12 years | 0.313 [a] | 0.642 |
| 12 years | 0.434 [a] | 0.215 |
| 13-15 years | 0.182 [a] | 0.078 |
| ≥ 16 years | 0.071 | 0.065 |
| Age Heterogamy | | |
| Husband > 5 Yrs Older than Wife | 0.165 [a] | 0.254 |
| Husband's Age W/in -2 to 5 Yrs of Wife | 0.767 [a] | 0.657 |
| Husband > 2 Yrs Younger than Wife | 0.068 | 0.089 |
| Education Heterogamy | | |
| Husband More Educated than Wife | 0.224 | 0.197 |
| Husband/Wife same education | 0.522 | 0.586 |
| Husband Less Educated than Wife | 0.254 | 0.217 |
| Husband Married Before | 0.116 | 0.081 |
| Husband/Wife of Same Race | 0.661 [a] | 0.947 |
| | | |
| Marriage Cohort | | |
| 1975-1979 | 0.202 | 0.230 |
| 1980-1984 | 0.249 | 0.230 |
| 1985-1989 | 0.215 [a] | 0.341 |
| 1990-1994 | 0.334 [a] | 0.199 |
| | | |
| Number of Observations | 253 | 218 |

*Note.* Means are weighted.
[a] U.S.-born mean differs significantly from that of foreign-born Mexican Americans, $p<.05$ level.