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Research Article

The topography of the divorce plateau: Levels and trends in union stability in the United States after 1980

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Research Article

The Topography of the Divorce Plateau: Levels and Trends in Union Stability in the United States after 1980

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Abstract

The probability of divorce in the U.S. has remained constant for the last two decades at about "half of all marriages." While this estimate is well established, and marked differentials in divorce rates are well known, there are no reliable estimates of differences in the lifetime probability of divorce. Using data from the 1990 June CPS, we document very large differentials by race, age at marriage, and education in the probability that recent cohorts of marriage will end in separation or divorce. Then, using data from the 1995 NSFG, we find important increases in differentials in marital dissolution, and especially in all unions, during this period of stable aggregate rates. These results indicate that examining only marital transitions obscures the growth in family instability that has resulted among some groups because an increasing proportion of unions began as cohabitation.

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1. Introduction

A recent analysis of marriage, divorce and remarriage, presented estimates of divorce calculated from vital statistics data. Schoen and Standish's (2001) paper is important because it updates prior analyses of the implications of trends in the timing of marriage, divorce, and remarriage on the U.S. life course. We expand on their analysis in three ways. First, we present differentials in estimates of marital dissolution by three socioeconomic characteristics: race-ethnicity, age at marriage, and education. Studies of the relative risk of disruption are common (e.g. Bramlett & Mosher 2001; Castro Martin & Bumpass 1989; Sweeney & Phillips 2001), and important for our understanding of the processes involved. Nonetheless, the substantive implications of differing rates cannot be appreciated fully without estimates of lifetime disruption for subpopulations. Second, we conduct analyses of recent marriage cohorts to identify whether the large differentials in marital dissolution documented in the first step are growing. In light of the importance of family stability for the lives of both adults and children (Amato 2000), it is surprising that there has been little attention in the recent literature to either subgroup differences in the cumulative probability of divorce or to changes in relative risks over time. One exception is Teachman's (2002) recent analysis of trends in divorce from the 1950s to the early 1980s. Our analysis extends this trend analysis by examining the period between the early 1980s to the early 1990s. Finally, we investigate the impact of cohabitation on the differential in the instability of family life. If we are to understand trends in family stability, we must take into account the dramatic increase in cohabitation. Cohabitation has moved some of the instability of family life out of the statistical accounting system of marriage and divorce. It has thus likely contributed to the plateau by removing an increasing number of disruptions that, in an earlier time, would have occurred after marriage. While some cohabiting unions are casual alliances, the majority of cohabiting persons think that they will marry their partner and many have definite plans to do so (Bumpass, Sweet, & Cherlin 1991). Our understanding of family stability is impaired if we focus only on divorce rates: for example, children's family lives have become increasingly unstable during the plateau in the divorce rate (Bumpass & Lu 2000).

2. Data and methods

2.1 Differentials in marital dissolution

Our first objective is to provide estimates of lifetime marital dissolution by age at marriage, race/ethnicity, and education using data from the 1990 June Current Population Survey

(CPS). The CPS is conducted monthly by the Bureau of the Census for the Bureau of Labor Statistics, primarily to measure characteristics of the labor force. Each month approximately 50,000 households are sampled and interviewed. Every five years between 1975 and 1995, the June survey included a supplement collecting information on the marital, separation, divorce, and birth histories of adult female members of sampled households.

Contrary to some earlier reports (e.g. Preston and McDonald, 1979; McCarthy, Pendleton, and Cherlin, 1989), there is reason to believe that survey data can provide good estimates of divorce. Goldstein (1999) shows that although both marriages and divorces are under-reported in the Current Population Survey, when survey data are used for both the numerators and denominators, the 1990 June CPS closely tracks the crude divorce rates estimated from the NCHS vital statistics. Our own previous analyses also show that survey data can replicate vital statistics and that estimates of cohort disruption by specific marital durations are almost identical in each of the June Current Population Surveys from 1975 to 1990, for a period up to 20 years prior to each survey (Bumpass and Raley 1992).

We would prefer to use the 1995 June CPS to represent as recent a period as possible. (The 2000 June CPS did not collect marital histories.) Unfortunately, the 1995 CPS appears to underestimate levels of marital disruption relative to estimates for the same cohorts from the earlier surveys. Consequently, we use the 1990 CPS to calculate our period estimates of divorce. Given the overall stability in levels of divorce, this should not affect our estimates for the total population of first marriages. However, to the extent that there has been divergence in levels of marital instability for different population groups, our lifetime estimates understate the differentials we would observe if we had more recent data.

To construct period life-table estimates, we calculate marriage duration specific probabilities of marital dissolution between June 1986 and June 1989. Although the CPS sample is large, we use a three-year window to increase the stability of the estimates for subgroups. This is especially important for longer durations, where we have a small sample of marriages. To construct our lifetable estimates, we calculated the number of people married at each duration from 0 to 30 in January of 1987, 88, and 89. Then we calculated the number of divorces to marriages at each duration between July 86-June 87, July 87-June 88, and July 88-June 89. (We chose these periods so that we would not count dissolutions less than one year old at the time of the survey that may reconcile.) Using this information, we calculated the duration specific divorce rate. We then converted the rate into a probability to calculate the survival function. When duration specific divorce rates are stable, as has been the case, such synthetic period estimates represent well the experience of marriage cohorts.

Separation is used to mark the end of a marriage rather than divorce because the timing of divorce is to some extent an artifact of the legal process and other extraneous factors, and some permanently separated couples never divorce. While separations are sometimes followed by reconciliation, the vast majority of those who have not gotten back together within a year of separation will never do so (Sweet & Bumpass 1974; McCarthy 1978; Bumpass, Castro Martin, and Sweet 1991). More importantly, some population sub groups experience longer durations between separation and formal divorce and are less likely ever to formally divorce. For example, the duration between separation and formal divorce was greater for blacks than whites (McCarthy 1978) during the 1970s and our supplementary analyses of the NSFG data indicate that this continues to be true. Thus, an analysis of divorce would provide distorted estimates of marital dissolution.

2.2. Differentials in trends

The second part of our analysis uses data from the 1995 National Survey of Family Growth (NSFG) to produce cohort estimates of the cumulative probability of disruption within 5 years for first marriages for two recent marriage cohorts. Our goal is to see whether the differentials observed in the first part of the analysis are growing. We also use the NSFG to examine differentials in the stability of cohabiting unions. The NSFG is conducted periodically by the National Center for Health Statistics with the primary goal of providing estimates of factors affecting the U.S. birth rate and the reproductive health of U.S. women 15-44 years of age. Marital and fertility histories have long been a part of the content of this survey, but Cycle 5 provides complete cohabitation histories for the first time. Interviews averaging 105 minutes were conducted with 10,847 respondents over the first ten months of 1995 (Potter et al. 1997).

To examine whether the plateau in marital dissolution was experienced equally for women across all socioeconomic categories, we divide the period since 1980 into two cohorts, 1980-86 and 1987-94. Only two cohorts are used to maximize the size of the samples. To begin with cohorts formed in 1980, we must limit our estimates to first marriages formed at age 30 or younger because of the upper age limit of the sample. Women who were older than age 30 in 1980 are not represented in this sample of women under age 45 in 1995. These estimates are limited to the first five years of marriage because of the second of these two cohorts is necessarily limited to the first seven years of marriage, and the sample size shrinks with each successive year of duration because of censoring by interview (Note 1). Estimates are prepared for first marriages and for all first unions so we can evaluate whether these two perspectives may lead to different conclusions about family experience.

Finally, using data from the NSFG, we construct discrete-time proportional hazard estimates of marital dissolution to examine the statistical significance of any apparent differences in trends across subgroups. That is, whether there is an interaction between union cohort and our variables. Using person-years of exposure, these models begin at union formation and are censored by disruption, 5 years duration, or interview. The covariates examined are marriage cohort, education, race-ethnicity, age at marriage, whether the respondent had a birth prior to the marriage, and whether the respondent cohabited prior to marriage. Parallel models are estimated for divorce from first marriage and for disruption from all first unions (including both marriage and cohabitation). Finally, we include interaction terms in the proportional hazard models to see whether differentials in trends are significant net of controls for other socio-demographic factors. These analyses are weighted to account for the stratified sampling design and differentials in non-response. The weights are scaled so that sample size is the same regardless of whether the data are weighted.

3. Results

Table 1 provides estimates of the cumulative proportions of marriages expected to have ended by 5-year durations up to 30 years. The emergence of differentials can be seen over successive durations, but we focus this discussion on 30 years as an estimate of the proportion of marriages "ever disrupting." (The rate of disruption declines continuously after the first few years of marriage and only 1-2 percent of divorces occur after 30 years of marriage (NCHS 1996)). Similar to other recent estimates of the likelihood that a marriage will end in divorce or separation (e.g. Schoen and Standish 2001), we find that for all first marriages about half will dissolve. As noted in the introduction, our understanding of subgroup differentials in divorce risks is usually set in the context of this overall proportion without a substantive grasp of implications of differential rates for the lifetime experiences. Not surprisingly, the figure for Non Hispanic Whites is very similar to the total. On the other hand, duration-specific rates of the late 1980's imply that 70 percent of all marriages to black women end in separation or divorce.

We also see very large differences in lifetime marital stability by education and age at marriage. Sixty percent of first marriages disrupt among women who did not complete high school compared to slightly over one-third among college graduates. Similarly over 60 percent of marriages formed before age 20 break up (the proportion is about two-thirds for marriages under age 18) compared to about 40 percent of those begun over age 22. At the same time, marital instability is high relative to most countries in Europe among even those with the lowest rates considered here. For example, 20 percent of college graduates have their first marriage break up within 10 years compared to 20 percent or less for *all* marriages in Sweden, France and the former West Germany (Andersson and Philipov 2001).

Table 1: *Period Estimates of marital dissolution covering 1987 to 1989, using data from the 1990 June CPS*

	% Marriages Dissolved				
Duration	5	10	15	20	30
Education					
Not HS Grad	26	39	47	53	60
High School Grad	23	35	43	47	53
Some College	20	32	38	45	51
College Graduate	11	20	25	28	36
Race					
White	19	30	37	42	47
Black	26	44	53	59	70
Age at Marriage					
13-17	37	50	57	63	68
18,19	30	43	50	55	61
20,21	22	32	39	44	51
22,23	15	23	29	33	40
24-26	14	28	34	41	44
27-29	16	28	31	36	39
30+	13	20	23	27	31
Total	20	31	38	43	50

We turn now to the question of whether the plateau was experienced by each of the subgroups considered here. As described in the methods section, Table 2 presents the proportion disrupting within five years for two marriage and union cohorts formed during the plateau (1980-86 and 1987-94). The plateau is apparent in the top row, as levels of marital instability remained unchanged between 1980 and 1994. It is noteworthy that these estimates replicate almost exactly those from the CPS in Table 1: 21 percent of first marriages end within five years based on the 1987-89 period from the CPS, compared to 22 percent for the 1987-94 marriage cohort based on NSFG. The differentials observed in Table 1 are seen here as well. The first two columns suggest that the plateau was indeed shared by all groups with two exceptions. Marital instability appears to have increased among those without a high school diploma, but not among those with a college degree. There may also be a modest increase in the difference in the probability of disruption for marriages preceded by cohabitation compared to those not. The proportional hazard models discussed below test whether these differences in trends are significant.

Table 2: *Trends in the Proportion of Marriages and Unions Disrupting within 5 years of Formation, Total and by Characteristics, for Women Entering Marriage or Cohabitation under Age 30, 1980-94: Life-table Estimates from the 1995 National Survey of Family Growth*

	First Marriage		**First Union**	
	80-86	**87-94**	**80-86**	**87-94**
Total	22	22	31	34
Education				
Not HS Grad	29	34	39	42
12 Yrs	23	23	31	34
Col 1-3	24	26	34	40
Col 4+	15	13	26	26
No Child Before	18	15	26	24
Child Before	31	32	46	45
No Cohabit Before	20	16		
Cohabit Before Marriage	24	28		
Race/Eth				
Black	33	32	40	55
Non-Hispanic White	21	23	31	33
Hispanic	22	16	28	25
Age at Formation				
15-19	32	33	39	47
20-23	19	21	29	31
24-29	15	17	21	24
Unweighted N	2231	1650	2767	2097

In contrast to marriage, we see was a modest increase in the instability of first unions. This is as we expected given that an increasing proportion of first unions have begun with cohabitation, and that a decreasing proportion of cohabiting unions have lead to marriage (Bumpass and Lu 2000). Most importantly, the plateau was clearly not equally shared when we look at the stability of all unions. In particular, black women experienced a 15 percentage point increase in the likelihood that their first unions dissolve within 5 years, from 40 percent for the 1980-86 cohort to 55 percent for 1987-94 cohort—while disruption rates may have even declined slightly among Hispanic women. Furthermore, the stability of unions formed at younger ages decreased: disruption within five years increased 8 percentage points from 39 percent of unions began 1980-86 to 47 percent of unions began 1987-94. Given that we find no increase in *marital* instability for these groups, an important element of their experience would be missed if cohabitation had not been included in this analysis.

Tables 3 and 4 address the question of whether these apparent changes are statistically significant, for marriages and for all unions respectively. In the leftmost columns of Table 3, we see that there was no trend in the likelihood of divorce, but that all of the differentials in the risk of divorce are large and statistically significant—individually, and net of one another in Model 1. During the plateau, the proportion of marriages begun after cohabitation and the proportion begun with children already in the household increased, exerting an upward pressure on the divorce rate. On the other hand, the effect of these changes was moderated somewhat by the increase in age at marriage. Once these factors are controlled, we observe a modest but significant decline in marital disruption. That is, if the characteristics of marriages had not changed between 1980-86 and 1987-94, the marital dissolution rate would likely have decreased (Note2).

Table 3: *Proportional Hazard Models Predicting Dissolution First marriages begun age 30 or younger*

	Differentials				Differentials in Trends Education			
	Zero-Order		Model 1		Model 2		Model 3	
	Relative Risk	B/se	Relative Risk	B/se	Relative Risk	B/se	Relative Risk	B/se
Marriage Cohort (1980-86)								
1987-94	0.93	1.09	0.84	2.32*	0.74	1.98+	0.68	2.46*
Education (College Grad)								
No Diploma	2.52	8.77**	1.53	3.60**	2.09	5.40**	1.24	1.44
High School	1.70	5.69**	1.26	2.31*	1.59	3.96**	1.18	1.40
Some College	1.92	6.52**	1.53	4.10**	1.76	4.40**	1.43	2.74**
Child Before	1.89	9.26**	2.03	9.41**			2.05	9.52**
Cohabited Before	1.40	4.96**	1.38	4.45**			1.37	4.32**
Race-Ethnicity (Non-Hispanic White)								
Black	1.84	6.12**	1.49	3.83**			1.50	3.86**
Hispanic	0.94	0.55	0.82	1.76			0.82	1.68+
Age at Marriage (15-19)								
20-23	0.52	8.35**	0.50	8.30**			0.50	8.34**
24-30	0.41	10.15**	0.37	10.09**			0.37	10.09**
Interaction between Cohort and Education								
No Diploma * 87-94					1.59	2.16*	1.70	2.44*
High School * 87-94					1.21	0.97	1.18	0.84
Some College * 87-94					1.28	1.20	1.21	0.91

Notes
+ Significant at the p < .10.
* Significant at p < .05.
** Significant at p < .01.

Only the interaction terms that were statistically significant—indicating differential trends—are reported in the rightmost columns of Tables 3 and 4. In models 2 and 3 of Table 3, the coefficient for marriage cohort reflects trends for college graduates once an interaction term is included in the model. With all variables controlled, college graduates experienced a substantial and significant decline in marital instability. At the same time, the significance and size of the coefficient for the interaction between marriage cohort and not having completed high school indicates divorce rates increased among those with the least education. Thus the relative disadvantage of women with fewer economic resources was reinforced over this period by increasing marital instability, at the same time that risks declined for women most able to weather the financial consequences of divorce. We also investigated whether the trends in marital instability differed by race or by premarital cohabitation. In neither case was the interaction term significant, indicating that there was no divergence in the trend in divorce by either variable.

A similar analysis of first union dissolution is reported in Table 4. The first row indicates that the modest increase in the proportion of first unions ending in 5 years (seen in Table 2) is marginally significant ($p < .10$). Once we control for the offsetting effects of age at first union and pre-union fertility, there is no trend in the union stability from the early 1980s to the early 1990s. There are again very large and significant differences in stability by socio-demographic characteristics. Once other factors are controlled, as with marriage, unions became more stable during the plateau for those with college degrees, but less stable for those who did not complete high school. Somewhat surprisingly, the apparent increase among those who formed a union at the youngest ages was not supported by a significant interaction term (not shown).

Finally, in contrast to the absence of any change among black women in the risk of marital dissolution, there was a 40 percent increase in the risk of *union disruption*, even though there was no change among white women. This increase is somewhat larger once the other predictors of union stability are controlled, and it is statistically significant in both cases.

Table 4: *Proportional Hazard Models Predicting Dissolution of first unions begun age 30 or younger*

	Differentials						Differentials in Trends											
							Education						Race					
	Zero-Order			Model 1			Model 2			Model 3			Model 4			Model 5		
	Relative Risk	B/se		Relative Risk	B/se		Relative Risk	B/se		Relative Risk	B/se		Relative Risk	B/se		Relative Risk	B/se	
Union Cohort (1980-86)																		
1987-94	1.10	1.84	+	0.99	0.22		0.96	0.37		0.82	1.89	+	1.06	0.91		0.92	1.27	
Education (College Grad)																		
No Diploma	1.79	7.48	**	1.26	2.63	**	1.67	4.98	**	1.09	0.78					1.26	2.63	**
High School	1.33	4.14	**	1.05	0.66		1.26	2.56	*	0.96	0.41					1.05	0.68	
Some College	1.54	5.89	**	1.28	3.14	**	1.38	3.25	**	1.16	1.44					1.27	3.11	**
Child Before	2.08	14.11	**	2.26	14.58	**				2.28	14.69	**				2.28	14.68	**
Race-Ethnicity (Non-Hispanic White)																		
Black	1.72	7.65	**	1.38	4.37	**				1.38	4.32	**	1.45	3.79	**	1.12	1.16	
Hispanic	0.81	2.39	*	0.76	3.06	**				0.76	3.06	**	0.86	1.27		0.79	2.00	*
Age at Union (15-19)																		
20-23	0.63	7.96	**	0.58	8.67	**				0.58	8.68	**				0.58	8.67	**
24-30	0.45	10.93	**	0.40	11.57	**				0.40	11.55	**				0.40	11.55	**
Interaction between Cohort and Education																		
No Diploma * 87-94							1.19	1.11		1.43	2.22	*						
High School * 87-94							1.14	0.96		1.23	1.48							
Some College * 87-94							1.27	1.61	+	1.25	1.49							
Interaction between Cohort and Race																		
Black * 87-94													1.46	2.65	**	1.58	3.16	**
Hispanic * 87-94													0.87	0.78		0.93	0.41	

Notes
+ Significant at the p < .10.
* Significant at p < .05.
** Significant at p < .01.

4. Conclusion

While it is well known that the likelihood of divorce varies by race, age at marriage, and education, no previous study has made clear the implications of these differentials for the lifetime probability of marital dissolution. We estimate that 70 percent of black women's first marriages will end in divorce, as will 47 percent of white women's marriages. There are also substantial variations by education and age at marriage--about 60 percent of the marriages of high school dropouts end in divorce compared to 36 percent among college graduates, and more than three-fifths of teen marriages compared to about 40 percent of marriages begun after age 21.

The very high level of marital dissolution among blacks may provide an important insight for understanding the low, and decreasing, marriage rate for this group (Bumpass and Lu, 2001) as well as the high proportion of the births to unmarried black women (69 percent in 1999, Ventura and Bachrach, 2000). It is likely that the more uncertain the prospects for marital stability, the more the potential gains from marriage are decreased and the potential costs are increased. There may also be a feedback effect from aggregate to individual stability: i.e. the awareness of that marriages are seldom for a lifetime may decrease the investment in a relationship (Becker 1981), and lower the threshold for leaving a relationship (Bumpass 2002).

Not only are differentials large, they appear to be growing. Similar to other studies, our results suggest that the overall rates of marital disruption have remained roughly steady in recent years. At the same time, the educational disparity in divorce has increased, as women with a college degree have experienced a decline in the risk of divorce.

Looking only at marital transitions, however, masks some increase in family instability during the plateau in divorce. Differentials in union stability are increasing considerably. While there was only a modest overall trend, increases in union disruption have been quite substantial for the less educated and for African Americans.

In sum, the crude divorce rate has declined over the last 20 years, leading many to conclude that the common wisdom that half of all marriages will end in divorce is no longer valid (e.g. Popenoe 2001). Some have even suggested that this decline in the (crude) divorce rate may mark the re-emergence of strong, stable families. There are a number of problems with such an interpretation. First, our analysis agrees with Schoen and Standish (2001) that divorce rates over the past 20 years imply that approximately half of all marriages will dissolve, and this has not changed since 1980. Second, there is substantial variation in trends both in marital stability and in the stability of all unions. While the odds of stable relationships are not all that good for even women with a college education, they are clearly lowest for the least advantaged in our society. Further, this disparity has become worse during the plateau in divorce. Unfortunately, those have the least resources to overcome the costs of family dissolution are experiencing the highest levels and the most increase in the risk.

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Notes

1. In the later cohort, those who married earlier contribute more to the estimate than those who marry later. At the extremes, those marrying in 1994 contribute only one year of experience, while those marrying in 1987 to 1989 could contribute up to 5 years of marital experience. This could be problematic if differentials in the first year are more pronounced than differentials in later years. We tested whether the impact of education, premarital fertility, age at marriage, or ethnicity varied across the first 6 years of marriage. They do not, and hence we do not believe our analysis using a censored cohort is problematic.
2. Our estimate of the trend in marital instability includes experience only in the first five years of marriage and may not reflect trends in the dissolution rates at later durations.

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