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Recent Trends in Marital Disruption

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The post-1980 decline in the crude divorce rate must be interpreted in the context of the long-term trend and in terms of what we know about composition effects on crude measures—particularly given shifts in age at marriage and the age composition effects of the baby boom. Data from the June 1985 Current Population Survey permit more detailed, exposure-specific measurements as well as the use of separation as the event terminating marriage. Estimates from these data suggest a decline followed by a recovery. Taking into account well-known levels of underreporting, we find that recent rates imply that about two-thirds of all first marriages are likely to end in separation or divorce. We examine the persistence of major differences in marital stability and evaluate the comparative stability of first and second marriages.

Family life is changing dramatically across virtually all Western industrial societies (Westoff, 1986). Among the most profound of these changes has been the sharp reduction in marital stability, affecting markedly the life course of individuals, the nature of family life, and the household composition of populations. Hence it is not surprising that the recent decline in the U.S. crude divorce rate generated a collective sigh of relief (e.g., Kantrowitz et al., 1987; Norton and Moorman, 1987; "Study Finds Drop in Divorce Rate," 1987).[1] It is too soon, however, to conclude that the long-term trend has been reversed. We must carefully evaluate this decline in light of the historical context and in light of what we know about composition effects on measures like the crude divorce rate.

Preston and McDonald's (1979) estimates of cohort divorce rates strongly suggest that basic structural changes underlie the trend: Whereas annual rates of divorce have fluctuated with wars and the economy, the likelihood of divorce has increased in a steadily accelerating curve, from 7 percent for first marriages in 1880 to recent levels of more than 50 percent (Weed, 1980). The duration of this trend strongly suggests that the roots of current patterns of marital instability are deep, and not just a response to recent changes in other domains such as fertility, sex-role attitudes, female employment, or divorce laws.

Increases in divorce since 1965, as well as the declines in marriage and fertility rates, are best seen as a continuation of the long-term reduction of family functions that has occurred in conjunction with the transformation of our economy (Ogburn, 1954) and an increasing cultural emphasis on individualism (Lesthaeghe, 1983; Westoff, 1986). When viewed from this perspective, we must ask whether the underlying changes have run their course.

In this article we lay out some reasons why a decline in the crude divorce rate need not signal a reversal of long-term trends and evaluate recent trends and levels by using data that permit more detailed, exposure-specific measurement. We then examine the persistence of major differences in marital stability, and finally, we evaluate the comparative stability of first and second marriages.

Data and Methods

This analysis is based on marital histories from the June 1985 Current Population Survey (CPS; U.S. Bureau of the Census, 1986). This source has a clear advantage over vital statistics data because it permits the use of *separation*, instead of *divorce*, dates to identify the timing of marital disruption and because it allows individual-level analysis of a number of related factors, holding the risk exposure constant. For the ever-divorced or currently separated, separation is a more meaningful definition of disruption than divorce because of the dependence of the latter on variations in the legal process and because subgroup variations in the timing and probability of divorce after separation make it a misleading indicator of marital disruption (McCarthy, 1978; Sweet and Bumpass, 1974).

We limit our analysis to the data for women because of the considerably lower quality of the marriage history data for males (Cherlin and McCarthy, 1984) and to cohorts married since 1970 to minimize recall error. (Over the last 15 years or so, each annual first marriage cohort is represented by approximately 1,000 women.) The variables we can examine are limited by the questions included in the CPS and by the cross-sectional design, so we cannot explore important factors such as whether the respondent's parental family was intact or the income prior to disruption. Nonetheless, we can examine trends and differentials by race, education, age at marriage, prior fertility, marriage order, and region.

Life-table methods are used to deal with the censoring imposed by the interview for marriage cohorts and to produce synthetic estimates of lifetime experience from period rates. Estimates of period rates are computed by using the last marriage cohort to complete a given year of duration since marriage by June 1, 1985. Thus we observe the risk of marital disruption at each marital duration (conditional on being married at the start of that duration) for each year between marriage and the 40th anniversary. These duration-specific rates are then combined using life-table logic to yield the cumulative proportion expected to survive to specific durations.

A proportional-hazard model is used for the multivariate analysis (Menken et al., 1981; Teachman, 1982).[2] For continuous variables, the parameter β denotes the effect of a unit change in the independent covariate on the log of the hazard rate, "holding constant" the other independent variables. For categorical covariates, β represents the deviation of a specified group from the hazard of the reference group. The exponentials of the coefficients, $\exp(\beta)$, allow us to express the hazard of a specific group as a proportion of the baseline hazard. An $\exp(\beta)$ of 1.0 means that the characteristic analyzed has no effect, whereas an $\exp(\beta)$ of 1.50, for example, means that marital disruption is 50 percent higher than in the comparison category.

The assumption of proportionality of hazards is not problematic for most of our covariates (see Morgan and Rindfuss, 1985; Teachman, 1986). Marriage cohort was the only variable that did not meet the proportionality assumption. This problem will arise with a trend variable whenever there has been a period fluctuation in the transitions being analyzed. By definition, such a fluctuation will occur at systematically different durations for cohorts of entry to risk—destroying the assumption of a common duration function. Because of this lack of proportionality, and because of significant interactions between marriage cohort and a number of our predictors, separate models were run for marriage cohorts.

Recent Trends

The U.S. crude divorce rate declined 10 percent between 1980 and 1987. This decline *may* indicate a return to more stable family life; however, it would be foolish to jump to that conclusion from this brief deviation from the trend. The current rate represents a very high level of marital disruption. Moreover, there are a number of reasons why a leveling of divorce rates, or even a slight decline, might occur without any change in values:

1. The acceleration after the mid-1960s cannot continue indefinitely, since an absolute upper limit must soon be reached.
2. Further, it is axiomatic in demography that trends in crude rates may be heavily influenced by shifting composition. Two components are particularly relevant—the increasing age at marriage and the age composition consequences of the baby boom. The marked reduction in marriage rates implies that the composition of recent marriages would tend to make them more stable, all else equal. For example, teenage marriages are much more likely to end in divorce than marriages at older ages, and recent marriages include a lower proportion who married as teenagers. In addition, the lower marriage rate would suggest that at all ages, many of the more tenuous relationships, which previously would have married and divorced, do not now make it to the altar. Independent of the changes in marriage rates, the movement through the population of the large baby boom cohorts first increased, then decreased the proportion of marriages beginning at high-risk young ages. It had a similar net influence on the proportion of marriages at short durations (for which divorce rates are highest).
3. Finally, it is common demographic wisdom that period measures will be inflated to the extent that divorce is occurring progressively earlier in marriage. A downturn in period measures can be expected when timing ceases to change.

In the remainder of this section, we examine the trends in the risk of marital disruption by using the more detailed data available from the June 1985 CPS. Then we reevaluate expected lifetime experience. The trends are examined from two perspectives: disruption within the first 5 years of marriage for successive marriage cohorts, and synthetic estimates of disruption within 40 years of marriage based on duration-specific rates in different periods.

We began by examining life-table estimates of the cumulative proportion of marriage disrupted within five years, by order of marriage, for marriage cohorts since 1970. There has not been a downturn in cumulative cohort experience as far as we can compare it for recent cohorts. The cumulative proportion with a disruption within five years for cohorts marrying in the late 1970s and those marrying in the early 1980s was 0.22 and 0.23 for first marriages, 0.24 and 0.27 for second marriages, and 0.20 and 0.38 for third marriages (though there were only 238 and 451 third marriages in the sample for these two periods).

Period life tables estimate lifetime marital disruption at 57 percent, 51 percent, and 56 percent for 1980, 1982, and 1984, respectively. Hence we find a downturn followed by a recovery to the preceding all-time high. Neither the downturn nor the recovery should be overemphasized. Period fluctuations are to be expected, and it is prudent to evaluate such fluctuations in light of the longer trend. These data suggest, at the most, a plateau rather than a reversal in that trend. Such a plateau (without a recovery) persists in the crude rates through 1987. We have already noted, however, that shifts in age and age at marriage composition may depress the crude measure relative to duration-specific measures such as those used here.[3]

The cohort implications of recent disruption rates were made salient recently by the Lou Harris claim that only about 12 percent of marriages are likely to disrupt.[4] It appears that previous demographic estimates have been *downwardly* biased for two reasons: (1) Those based on the vital statistics have failed to take into account the lower rates in the Divorce Registration Area (DRA) compared with the country as a whole (National Center for Health Statistics, 1987b), and (2) those based on surveys have failed to take into account the underreporting of marital disruptions in surveys relative to those recorded by the national vital statistics system (Preston and McDonald, 1979).

We adjusted for these sources of downward bias in the following manner. First, we recalculated the vital statistics life tables for 1975 marriages (Weed, 1980) after adjusting upward the duration-specific rates by 10 percent to compensate for the difference in divorce rates in the U.S. as a whole compared with that in the DRA. This adjustment resulted in an estimate of 51 percent disrupting after 40 years rather than the published figure of 46 percent estimated on the basis of the DRA. We then compared our estimate based on 1975 rates with this adjusted vital statistics estimate, finding that the CPS estimate was 22 percent too low.

This estimate of the extent of underreporting is similar to Preston and McDonald's. It is possible, however, that underreporting might be less for the period closest to the date of survey. We do not have a comparable vital statistics life table to evaluate this, but we can compare age-specific rates from the June 1985 CPS with adjusted vital statistics rates (for a period centering on 1983). This comparison indicates that the CPS rates are about 19 percent too low. Hence there is little evidence of deterioration in the quality of reporting with time (and no trend in underreporting was evident in the data presented by Preston and McDonald for earlier periods). We then adjust the duration-specific disruption rates for this level of underreporting and recalculate the expected cumulative disruption.

As expected, these adjustments yield a higher proportion likely to experience marital disruption than has previously been estimated. The 56 percent estimate derived directly from the June 1985 CPS implies an adjusted level of expected marital disruption of about two-thirds (64 percent) after 40 years.

Many will find this higher estimate of two-thirds unsettling. In evaluating this estimate it is relevant to note the following:

1. Divorce rates continued to increase after the 1975 level (which implied 51 percent after 40 years, based on vital statistics). For example, the rates for married women 15 and over increased from 20.3 to 22.8 between 1975 and 1979 (National Center for Health Statistics, 1987a).

2. We include separations as well as divorces. Some separations, particularly those to older women who do not wish to remarry, may never be followed by divorce. A comparison of the estimates of expected total disruption to the cumulative proportion expected to *divorce* suggests that perhaps 5 percent of a marriage cohort would separate without ever divorcing.

3. This estimate is based on experience in a recent period. Such a period estimate may overstate (or understate) actual cohort experience because of either timing shifts or changing rates. Note that we are projecting what would follow *if duration-specific rates remained unchanged*, that is, if the long-term trend has in fact plateaued and rates do not increase further. (The disruption experience of the 1970 and 1975 marriage cohorts was higher at each duration than the levels implied by the period rates for the years in which they married because the upward trend continued.)

4. We are addressing the prospects of recent first marriages. The discussion over the Harris estimates appeared at times to confuse this point with the prospects for *existing* marriages.[5] This raises an interesting point that we do not think has been addressed before. Obviously existing marriages include many at longer durations where disruptions are rare and few in the early years of marriage in which the risk is highest. Existing marriages in the June 1985 CPS have an average duration of about 17 years. When we weight the expected cumulative failure after each duration by the duration distribution of existing marriages, we estimate that 21 percent of marriages that were intact in June 1985 are likely to disrupt if rates of that period persist.

Population Differences

The increases in marital disruption of the last several decades have been shared by all groups in American society, as well as by many European countries (Sardon, 1986). Even though this similarity in trends suggests common underlying factors, there are also marked subgroup differences that persist and are important to our understanding of marital disruption.

In the remainder of this article, we examine differentials in marital disruption and then turn to a pooled analysis of first and second marriages to consider the effect of remarriage.[6] Single-variable life tables are used to provide a sense of the scale of the differences observed in the population, whereas a proportional-hazard model is used to estimate the net effects of the variables considered.

First Marriages

Table 1 presents life table estimates of the proportion of first marriages disrupting within five years for three marriage cohorts. The proportional-hazard results are presented in Table 2. As expected, each variable made a significant net contribution to the probability of marital dissolution. However, we found significant interactions between marriage cohort and several of the predictor variables as well as between race and several of the predictor variables.[7] Consequently, the proportional-hazard results are presented separately by marriage cohort and by race. We will consider these two tables jointly, though the focus will be on the multivariate results in Table 2.

Age at Marriage

The inverse relationship between age at marriage and the likelihood of marital disruption is among the strongest and most consistently documented in the literature (Kiernan, 1986; Moore and Waite, 1981; Teachman, 1983, 1986). This association persists net of education and premarital pregnancy (Bumpass and Sweet, 1972) and at longer marital durations (Morgan and Rindfuss, 1985; Teachman, 1986). A number of mechanisms have been suggested

Table 1. Life Table Estimates of the Proportion of First Marriages Disrupted Within Five Years, for First Marriage Cohorts

	Marriage cohort			*N*		
Variable	1970–1974	1975–1979	1980–1985	1970–1974	1975–1979	1980–1985
Age at marriage						
14–19	0.23	0.30	0.31	2,266	1,849	1,477
20–22	0.14	0.18	0.26	1,733	1,641	1,680
23–29	0.11	0.15	0.15	1,126	1,282	1,978
30+	0.14	0.16	0.14	220	256	421
Education						
0–11	0.21	0.29	0.33	709	707	765
12	0.18	0.23	0.26	2,408	2,251	2,421
13+	0.16	0.18	0.16	2,228	2,070	2,370
Children before marriage						
0	0.17	0.21	0.21	4,799	4,411	4,799
1+	0.22	0.31	0.36	546	617	757
Region						
Northeast	0.13	0.17	0.21	1,179	1,064	1,159
North Central	0.16	0.20	0.23	1,307	1,226	1,298
South	0.18	0.25	0.25	1,694	1,528	1,766
West	0.23	0.24	0.22	1,165	1,210	1,333
Race/ethnicity						
White	0.17	0.21	0.22	4,290	4,058	4,403
Black	0.24	0.30	0.36	490	431	472
Hispanic	0.15	0.21	0.24	337	341	447
Total	0.18	0.22	0.23	5,345	5,028	5,556

Table 2. Proportional-Hazard Estimates [exp(β)] of Differentials in First Marriage Dissolution, by Marriage Cohort and Race

	White women, first married in				Black women, first married in			
Variable[a]	1970–1974	1975–1979	1980–1985	1970–1985	1970–1974	1975–1979	1980–1985	1970–1985
Age at marriage								
(14–19)								
20–22	0.60	0.61	0.81	0.63	0.50	0.76[b]	1.12[b]	0.66
22–29	0.42	0.48	0.50	0.44	0.50	0.89[b]	0.78[b]	0.62
30+	0.34	0.52	0.46	0.42	0.53	0.80[b]	1.12[b]	0.67
Education								
(0–11)								
12	0.84	0.81	0.71	0.80	0.87[b]	0.95[b]	0.58	0.84[b]
13+	0.86[b]	0.68	0.49	0.73	0.74[b]	0.97[b]	0.40	0.75
Children before marriage								
(0)								
1+	1.75	1.57	1.80	1.71	1.04[b]	1.43	1.00[b]	1.16
Region								
(North)								
North Central	1.03[b]	1.06[b]	1.00[b]	1.04[b]	1.49[b]	1.12[b]	1.22[b]	1.30[b]
South	1.07[b]	1.34	1.18[b]	1.18	1.00[b]	0.92[b]	0.80[b]	0.94[b]
West	1.52	1.52	1.12[b]	1.45	1.07[b]	1.16[b]	1.07[b]	1.10[b]
N	4,290	4,058	4,403	12,751	490	431	472	1,393

[a] Parenthetical items are the reference categories.
[b] β is less than twice its standard error.

to explain this relationship, including the degree of maturity and competence for marital roles (Booth and Edwards, 1985), search time for a marriage partner (Becker, Landes, and Michael, 1977), and emotional, educational, and economic resources available.

The high disruption rate of teenage marriages is apparent in Tables 1 and 2. Women who married as teenagers are twice as likely to separate as those who married after the age of 22. Consistent with Sweet and Bumpass (1988), we do not find evidence of the higher risk among women marrying over the age of 30 that Glick and Norton (1977) suggested. The effect of age at marriage is little altered by controls for the other variables considered. Among white women marrying in the early 1980s, marrying after the teenage years was associated with a 37 percent lower disruption risk for marriages at the ages of 20–22 and with an additional reduction in risk of more than 20 percent for those marrying after the age of 22.

In Table 1 we see an increase over time in marital disruption for women first marrying at the ages of 20–22: the life tables show nearly a 40 percent increase in the probability of separation between the late 1970s and early 1980s—a period when rates did not increase for other ages at marriage. One possible interpretation for this is that the definition of "young" marriage is changing as the average age at marriage has become progressively older. What once characterized teenage marriages may now extend to marriages in the early 20s.

Among blacks the major age at marriage effect is the contrast between teenage marriages and those at age 20 and older. (Differences for the later cohorts are significant when collapsed to this contrast.) There is some indication of an increase in the size of this effect.

Education

Education is related inversely to marital disruption, though some studies find that the relationship is not entirely linear: women with graduate degrees appear to have higher rates of separation than women with college education (Houseknecht and Spanier, 1980). In the absence of other relevant variables, such as income histories, education may stand as a surrogate for an array of factors ranging from differing levels of economic strain to human capital differences affecting interpersonal skills.

The likelihood of separation increased at lower education levels (from 21 percent to 33 percent by 5 years in Table 1), increasing the size of education differences in marital stability among more recent marriage cohorts. In the multivariate results for the 1980–1984 cohort in Table 2, women who attended college have a 51 percent lower rate of separation than high school dropouts among whites and a 60 percent lower rate than high school dropouts among blacks. Consistent with findings for around 1970 (Bumpass and Sweet, 1972), most of the observed educational differences in marital disruption are explained by other variables for the earliest cohort. This is not true, however, of the most recent marriage cohort. Thus education has become a more important independent factor affecting marital disruption.[8]

Children Born Before Marriage

Marriages that begin with a child are expected to be less stable, especially when the husband is not the child's father (Furstenberg, 1976; Teachman, 1983, 1986). In addition to possible increased strains on the marriage created by the child, women who have already experienced single parenthood may be less reluctant to return to single parenthood (Morgan and Rindfuss, 1985). Whereas there has been increasing attention given to the levels of nonmarital fertility, the implications of its recent levels for the conditions under which marriages begin have been little appreciated. One of every seven recent first marriages involved a bride who was already a mother (Sweet and Bumpass, 1988).

In Table 1, we see a substantial increase in marital disruption among women who began marriage with a child. It is not possible with current data to determine what proportion of these were cohabiting relationships in which the husband was the child's father. Hence the effect may reflect a greater representation of cohabiting relationships as well as strains introduced by the presence of the child. Nonetheless, we find a net effect of premarital birth among whites (a 71 percent higher disruption rate than among whites without a premarital birth, averaged over the period since 1970; see Table 2). The estimates are less stable over cohorts among blacks and average only 16 percent higher over the period since 1970. This difference by race may well reflect the more "deviant" status of an unmarried mother among whites than among blacks; hence among whites, unmarried motherhood may both generate more strain in the marriage and be more selective in terms of the characteristics of the marriage partners involved.

Region

Divorce rates have generally been lower in the East and higher in the West. These differences may reflect demographic composition (Weed, 1974), differences in legal systems, and differences in social tolerance (Sweet, 1973). Others have speculated about the effects of a "frontier atmosphere" or a tradition of individualism of some regions (Fenelon, 1971) or of a Durkheimian state of anomie resulting from population growth and migration (Glenn and Shelton, 1985; Glenn and Supancic, 1984).

It appears that regional differences are narrowing, with particular increases in the Northeast over recent cohorts combined with a slight decline in the West. Among whites in the early 1970s, disruption rates were 50 percent higher in the West, net of other factors. By the early 1980s, there are no significant regional effects among whites (Table 2). Regional differences among blacks were not significant for any cohort or for the period since 1970 combined.

Race/Ethnicity

Because of a number of interactions with race, the preceding discussion has considered predictor variables separately for blacks and whites. We turn now to the much higher rates of marital instability among blacks. This has been the subject of considerable discussion recently, with attention returning to the issues linking single-parent families intergenerationally raised by Moynihan (1965) and to the disadvantaged labor market for black males (Wilson and Neckerman, 1986). In conjunction with markedly lower remarriage rates (Cherlin, 1981; Sweet and Bumpass, 1988), the higher disruption rates among blacks has reduced the average duration in married life among black women to about one-third of nonwidowed adult years, compared with about two-thirds among white women (Espenshade, 1985).

It is evident in Table 1 that differences in marital instability by race increased over the most recent cohorts, with a one-sixth increase among both blacks and Hispanics—again during the period of relative stability. Despite the increase among Hispanics, the level is only slightly higher than among non-Hispanic whites and is actually lower once the composition on the other variables is taken into account (Frisbie, 1986). The interactions between race and the other predictors preclude a straightforward estimation of the net effect of race, though a simple additive model suggests that compositional differences may explain about a quarter of the higher rates among blacks (reducing a 73 percent higher rate to 56 percent).

One way to examine racial differences net of the variables considered here is to estimate the combined risk for women with the lowest and highest risk characteristics in the white and black models separately.[9] The expected rate of disruption within 15 years for the lowest risk group would be 18 percent among whites but 38 percent among blacks; for the highest risk groups the figures are 78 percent and 92 percent, respectively. Thus these variables predict great diversity in the risk of disruption among both blacks and whites. Nevertheless, racial differences persist independent of these variables (e.g., Teachman, 1986), with the risk more than twice as high for blacks as for whites among women who attended college and who neither married early nor had a premarital birth.

Second Marriages

A similar analysis was carried out for second marriages. Life table estimates are in Table 3. There were no significant interactions between our predictor variables and either the marriage cohort or the trend for second marriages. Hence we present only a combined proportional-hazard model in Table 4. For the sake of brevity, we will note only a couple of points specific to second marriages.

The existence of prior children has a different meaning for second marriages because most are likely to have been born in a previous marriage.[10] Nonetheless, a number of persons have argued that prior children are likely to create strains affecting the stability of second marriages. Cherlin (1978) explicitly identified stepchildren as the principal destabilizing element in remarriages. The literature reports conflicting results regarding this subject; some of the studies support Cherlin's hypothesis (Becker, Landes, and Michael, 1977; White and Booth, 1985), and other studies do not (Furstenberg and Spanier, 1984). We find only a 14 percent higher net rate of disruption for those with prior children when age at remarriage is controlled (Table 4), and the rate is actually lower when age at first marriage is taken into account (neither difference is significant). Thus these data provide no support for the seemingly compelling hypothesis that bringing children into a remarriage lowers the odds of marital success. Perhaps the stabilizing effect of the obligations to children offsets the additional strain on the relationship imposed by stepparenting.

The second point that warrants emphasizing with respect to differentials in disruption among second marriages is that the effect of race is very similar in first and second marriages.

Table 3. Life Table Estimates of the Proportion of Second Marriages Disrupted Within Five Years

	Marriage cohort			N		
Variable	1970–1974	1975–1979	1980–1985	1970–1974	1975–1979	1980–1985
Age at first marriage						
14–19	0.20	0.26	0.31	763	972	1,230
20–22	0.16	0.20	0.24	281	375	559
23+	0.11	0.21	0.13	154	220	299
Age at marriage						
14–19	0.26	0.33	0.40	291	313	374
20–22	0.15	0.24	0.26	260	446	538
23–29	0.17	0.24	0.27	317	470	748
30+	0.13	0.17	0.14	330	338	428
Education						
0–11	0.17	0.24	0.36	329	317	389
12	0.17	0.25	0.26	555	743	997
13+	0.20	0.23	0.22	314	507	702
Children before marriage						
0	0.16	0.23	0.24	238	378	540
1+	0.18	0.25	0.28	960	1,189	1,548
Region						
Northeast	0.16	0.20	0.22	187	262	343
North Central	0.21	0.24	0.28	285	351	489
South	0.17	0.26	0.27	402	567	743
West	0.16	0.23	0.30	328	387	513
Race/ethnicity						
White	0.18	0.23	0.26	1,026	1,337	1,792
Black	0.21	0.37	0.43	90	123	142
Hispanic	0.10	0.26	0.28	50	58	87
Total	0.18	0.24	0.27	1,198	1,567	2,088

This is contrary to the findings of McCarthy (1978), who argued that remarriages are more stable for blacks. Even though the markedly lower remarriage rates make second marriages among blacks more selective, the net rate of marital disruption among blacks in second marriages is of the same magnitude as that observed for first marriages.

The Relative Stability of First and Second Marriages

As noted earlier, second marriages are less stable than first marriages (McCarthy, 1978), and several explanations have been offered to account for this. Cherlin (1978) argued that the ambiguity of norms and the complexity of the family structure decrease the chances of remarriages. In addition, persons entering remarriages are, by definition, selected for those willing to consider divorce in the event of marital unhappiness (Bumpass and Sweet, 1972; Halliday, 1980); that is, remarriages may be selective of individuals who are not opposed in principle to divorce, whereas first marriages include a certain proportion of "stayers" who will never consider divorce. Remarriages may also be selective with respect to other characteristics, such as personality or interpersonal styles that are less conducive to marital stability.

Table 4. Proportional-Hazard Estimates [exp(β)] of Differentials in Second Marriage Dissolution, 1970–1985

Variable[a]	(1)	(2)
Age at first marriage		
(14–19)		
20–22	0.85	
23+	0.59	
Age at second marriage		
(14–19)		
25–29		0.73
30–39		0.71
40+		0.37
Education		
(0–11)		
12	0.89[b]	0.80
13+	0.91[b]	0.82
Children from previous marriage		
(0)		
1+	0.96[b]	1.14[b]
Region		
(Northeast)		
North Central	1.20[b]	1.16[b]
South	1.08[b]	1.06[b]
West	1.18[b]	1.12[b]
Race/ethnicity		
(White)		
Black	1.48	1.52
Hispanic	1.06[b]	0.97[b]

[a] Parenthetical items are the reference categories.
[b] β is less than twice its standard error.

It is evident from Table 5 that first and second marriages differ markedly in composition because of the joint consequences of differences in divorce and remarriage rates by these characteristics. Age at second marriage is considerably higher than age at first marriage (a mean of 33 compared with 22). Education is somewhat lower among remarried women. Blacks and Hispanics represent a smaller proportion of second marriages. (Blacks are more likely to separate from a first marriage but are less likely to remarry. Hispanics are more likely to remain in first marriages.) Fertility status at the time of the marriage is also very different: 12 percent of women marrying for the first time since 1970 were mothers when they married compared with 76 percent of remarrying women.

The most striking differences in composition between first and second marriages is the proportion who married for the first time as teenagers. It is critical to recognize that a majority of remarried women were married for the first time as teenagers (62 percent compared with 35 percent among first marriages). Although startling at first, this concentration of young first marriages among remarrying women is readily understandable. It is a consequence of the higher rate of dissolution of teenage first marriages and the higher remarriage rate associated with returning to the marriage market at younger ages.

It would seem straightforward to address the net effect of remarriage by using pooled-hazard models including remarriage as a variable. This is complicated, however, by the interactions of race and marriage cohort that we found among first marriages and by the low overlap between the age distributions of first and second marriages: there are very few

Table 5. Demographic Composition of First and Second Marriages, 1970–1985 (%)

Variable	First marriage	Second marriage
Age at marriage		
10–19	35.4	2.8
20–24	44.9	16.9
25–29	14.0	25.6
30+	5.8	54.7
Education		
0–11	14.5	21.9
12	44.3	46.8
13+	41.2	31.3
Race/ethnicity		
White	78.2	84.8
Black	9.7	7.9
Hispanic	8.6	4.8
Children born before current marriage		
0	87.7	23.9
1+	12.3	76.1
Age at first marriage		
10–19	35.4	61.8
20–24	44.9	31.0
25+	19.8	7.3

teenage second marriages and very few first marriages over the age of 30. Further, there is a difficulty in interpreting age at marriage effects in such a pooled model, to which we will return shortly.

To avoid problems from the race and cohort interactions, we explored pooled models based on marriages to white women married between 1980 and 1985. Among these recent marriages, remarriages are 25 percent more likely than first marriages to disrupt (first panel, Table 6). When we take into account the lower educational distribution of remarriages, this difference declines to 15 percent. When age at first marriage is added, we find that there is no longer any difference in the disruption rates of first and second marriages. Although it is not clear what factors are involved, this is consistent with the Morgan and Rindfuss finding (1985) that effects associated with early marriage persist at longer marital durations; they also appear to persist into second marriage and to account partially for the higher disruption rates of remarriages. This suggests that there may well be factors associated with teenage marriage, perhaps relating to personality or interpersonal style or to uncontrolled social background experiences, that have persisting effects throughout life.

There is a remaining issue that seems important, but for which we have been unable to structure a satisfactory answer: first marriages are handicapped by a much younger age distribution. It seems that this difference in age and maturity might mask higher age-specific dissolution propensities among remarriages. To address this, we examined a pooled model among white women marrying in 1980–1984 at the ages of 25–29; remarriages (net of education) were 60 percent more likely to disrupt than first marriages. This result should be taken as illustrative of the problem, rather than as a finding, because any age-specific (or age-controlled) comparison between first and second marriages contrasts a more stable subset of the former to a less stable subset of the latter. The age range of 25–29 was selected because it contains the most overlap between the age distributions of first and second marriages. Nonetheless, women marrying for the first time at these ages are marrying at

Table 6. Proportional-Hazard Models of the Dissolution of First and Second Marriages: White Women, 1980–1985

Variable*	β	$z(\beta)$	$\exp(\beta)$
	Model A		
Marriage order			
(1)			
2	0.222	(2.9)	1.25
Education			
(0–11)			
12	−0.496	(−5.4)	0.61
13+	−0.977	(−9.5)	0.38
Age at first marriage			
(14–19)			
20–22	−0.322	(−3.9)	0.72
23–29	−0.913	(−8.8)	0.40
30+	−1.056	(−4.5)	0.35
	Model B		
Marriage order			
(1)			
2	0.138	(1.8)	1.15
Education			
(0–11)			
12	−0.483	(−5.2)	0.62
13+	−0.955	(−9.2)	0.34
	Model C		
Marriage order			
(1)			
2	−0.0002	(0.0)	1.00
Education			
(0–11)			
12	−0.421	(−4.5)	0.66
13+	−0.728	(−6.6)	0.48
Age at first marriage			
(14–19)			
20–22	−0.159	(−1.8)	0.85
23–29	−0.701	(−6.2)	0.50
30+	−0.859	(−3.6)	0.53

* Parenthetical items are the reference categories.

"older" ages of first marriage (and thus are among first marriages least likely to separate), whereas women remarrying at this age are selected for women whose first marriage quickly disrupted and who are remarrying at a "young" age relative to other second marriages.

Thus we conclude that net of compositional differences with respect to education and age at first marriage, remarriages have no higher risk of disruption. There seems little basis for focusing on particular adjustment difficulties and incomplete institutionalization as factors decreasing the success rate of remarriages.

Conclusion

Even though the long-term increase in marital instability may have plateaued since 1980, the level remains very high. Using the marriage histories from the June 1985 CPS, life table procedures yielded estimates that 56 percent of recent first marriages would be likely to disrupt within 40 years of marriage. After adjusting, however, for the well-known underreporting of marital disruption in surveys relative to the vital statistics (and for the underrepresentation of divorces in the DRA), we conclude that the best estimate based on these data is that about two-thirds of all first marriages are likely to disrupt.

As high as this estimate may seem, it is based on the assumption that a plateau in the long-term trend has in fact occurred. Concern about the risk of AIDS may have contributed to the duration of the recent plateau, but this plateau had largely occurred before there was widespread information and concern about AIDS. No one knows whether a major social transformation is likely to be brought on by an AIDS epidemic among heterosexuals. In the absence of such an epidemic, it seems most likely that risks of marital disruption will continue to increase. For example, many observers from an array of theoretical perspectives see major links between changing family patterns and increasing equality of opportunity for women (Becker, 1981; Cherlin, 1981). In spite of considerable gains, equality of opportunity for women has a long way to go, and continued change seems likely. As Mason (in press) has recently noted, despite the appearance of a more conservative period, attitudes became increasingly egalitarian over the last decade. Many other relevant factors are also likely to continue to erode the centrality of family roles via-à-vis other adult opportunities (Sweet and Bumpass, 1988).

Surely there will be period variations associated with economic or even attitudinal fluctuations, but we should be slow to interpret plateaus or reversals as turning points in processes with such deep historical roots. The diversity in family life created by patterns of divorce and remarriage are likely an intrinsic feature of modern family life rather than a temporary aberration.

Notes

[1] Indeed, this anxiousness for a return to the "good old days" is illustrated by the *Newsweek* headline announcing that "the age of the disposable marriage is over" (Kantrowitz et al., 1987), as well as by the attention given the recent claim by Lou Harris that disruption rates are really much lower than demographers estimate (Galloway, 1987).

[2] We have also examined the effects of these variables by using logit regression of the probability of disruption by three years, for cohorts exposed at least three years. As expected, the results from the two methods are consistent. We present the proportional-hazard models because they permit us to use more of the data, that is, exposure begun within three years of interview and durations beyond three years.

[3] Vital statistics rates for married women 15 years old and older show a 2 percent increase between 1984 and 1985, the most recent period for which such data are available.

[4] He charged that demographic estimates, such as those cited in this article, are "the most specious pieces of statistical nonsense ever perpetrated in modern times" (Galloway, 1987). The most easily understood answer to whether Harris might be right is the fact that of first marriages occurring in 1975, one-third had already been disrupted by 1985.

[5] As in "54 million marriages just keep flowing along like Old Man River" (Galloway, 1987). We thank Jim Sweet for pointing this out (personal communication).

[6] These analyses to some extent replicate Teachman's (1986) based on data from the National Surveys of Family Growth (NSFGs). The major difference of importance is the attention to changing effects over cohorts reported below.

[7] Tests for interaction were made using the logit model described earlier. There were significant interactions of marriage cohort with education, premarital fertility, and region and of race with premarital fertility, region, and age at marriage.

[8] Teachman's (1986) analysis of pooled NSFGs heavily weighted earlier experience, which explains why he did not find an independent effect for education.

[9] For whites, the low-risk estimate is for women who married at 23 years old or older between 1970 and 1972 without a premarital birth, who had at least attended college, and who lived in the Northeast or North Central regions. The same characteristics were used for blacks, with the exceptions that marriages at 20 and over were included to maximize the number of cases (given the lack of difference beyond that age among blacks) and that residence was based on the Northeast and South, since these had the lowest rates in the equation for blacks. The high-risk estimate for each race is based on women who did not complete high school and were omitted from the low-risk group on each of the other variables.

[10] About one-third of all nonmarital births (about 7 percent of all births) occur after marital disruption (McLanahan and Bumpass, 1988).

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