RECENT TRENDS AND DIFFERENTIALS IN MARITAL DISRUPTION

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INTRODUCTION

Rates of marital instability have more than doubled since 1965, and thus the apparent reversal after 1980 was greeted with a collective sigh of relief (e.g. The New York Times, 1987; Norton and Moorman, 1987). This brief fluctuation, however, is insufficient evidence that the longterm trend has reversed. This paper documents the highest level of marital disruption yet recorded for this society, and examines the persistence of major differences by race, education, and age at marriage. Given the growing prevalence of remarriage, the comparative stability of first and second marriages is also examined. We conclude with a consideration of some linkages between marital disruption and changes in other family related areas.

DATA AND METHODS

This analysis is based on the marital histories from the June 1985 Current Population Survey, a multi-stage stratified sample of the U.S. population.

These data have a clear advantage over vital statistics data because they permit the use of *separation* instead of *divorce* dates to identify the timing of marital disruption, and because they allow the analysis of a number of covariates. 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 of subgroup variations in the timing and probability of divorce given separation can make divorce a very misleading indicator of marital disruption (Sweet and Bumpass, 1974; McCarthy, 1978).

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 1000 women.) Our covariates 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 income prior to disruption. Nonetheless, we can examine trends and differentials by race, education, age at marriage, prior fertility, marriage order and region.

A recent analysis of these data by Norton and Moorman (1987) concludes that divorce rates have declined. That analysis, however, improperly controlled risk exposure because it was based on birth cohorts. At given ages, more recent birth cohorts have had markedly less exposure to the risk of marital disruption because of delayed marriage. While the Norton and Moorman estimates describe the experience of successive birth cohorts by age, no interpretation with respect to differing *rates* can be made from that approach.

In the present analysis, life-table methods[1] are used to deal with the *censoring*

[1] Life-table techniques handle the problem of the censoring of cross-sectional data by incorporating both complete and incomplete segments of marital histories in the analysis, adjusting for differences in the

imposed by the interview. Estimates of period rates are computed using the last marriage cohort to complete a given duration segment of experience by June 1, 1985. These period duration-specific rates are then combined using life-table logic to yield the expected cumulative proportion surviving to specific durations. When marriage cohorts are compared, life-table procedures are again applied, this time to adjust also for truncated experience at each duration.

A proportional hazards model is used for the multivariate analysis. Analysis of the simultaneous effect of several independent variables is not possible with conventional life-tables because, besides being unwieldy, sample size will not permit separate tables for each unique combination. Proportional hazards models permit analysis of the net effect of our covariates on (instantaneous) separation rates.[2] Proportional hazards models, first introduced by Cox (1972), combine the basics of life-table and regression analysis techniques. Like life tables, they adjust for the problem of censoring bias, though a basic distinction is that proportional hazards models are continuous-time models, whereas life tables are based on a set of discrete intervals, such as years or months. Like regression analysis, proportional hazards models allow the formulation of equations relating several

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length of exposure to the risk of separation experienced. Complete or closed intervals refer to those that have ended by disruption, and incomplete or open intervals refer to those in which the event of separation has not occurred, but may eventually occur in the future.

[2]We have also examined the effects of these variables using Logit regression of the probability of disruption by 3 years, for cohorts exposed at least three years. As expected, the results from the two methods are consistent. We present the proportional hazards models because they permit us to use more of the data: i.e., exposure begun within three years of interview and durations beyond three years.

independent variables simultaneously to the hazard function. Previous analyses of marital disruption with this technique include Menken et al. (1981) and Teachman (1982).

The proportional hazards model focuses on the hazard rate $\lambda(t)$, the *instantaneous* rate of separation; that is, $\lambda(t)$ will refer to the probability of marital disruption at exactly duration t, conditional on the fact that no marital dissolution has occurred prior to that time.

The model, which specifies the linear relationship between the hazard function and the independent variables (or covariates) at each duration point, can be written as follows:

$$\lambda(t; z) = \lambda_0(t)\, e^{\beta z} \tag{1}$$

where $\lambda(t; z)$ denotes the hazard or the instantaneous separation rate at marital duration t for an individual (or group of individuals) with a known set of characteristics represented by the vector of covariates z [3]. $\lambda_0(t)$ denotes the unobserved hazard (or risk) at duration t for a *baseline* group of individuals, defined when all the independent variables in the model take the value of 0. β is a set of regression-like coefficients which show to which extent given values of the covariates shift, upward or downward, the underlying baseline hazard function $\lambda_0(t)$. These parameters specify how the covariates act on the hazard in a log-linear way [4]. The exponentials of the coefficients in Equation 2, $\exp(\beta)$, allow us

[3] Those covariates may be either discrete or continuous.

[4] For continuous variables, the parameters β denote the effect of a unit change in the independent covariate on the log of the hazard rate, "holding constant" the rest of the independent variables. For categorical covariates, β represents the deviation of a specified group from the hazard of the reference group.

to express the hazard of a specific group as a proportion of the baseline hazard. If e^{β} equals 1.0, it means that the characteristic analyzed has no effect. If the value of e^{β} is 1.50, for example, that means that marital disruption in that group is 50 percent higher [5] than for the baseline group. Conversely, if e^{β} is less than 1.0, say 0.75, it indicates that the probability of separation is 25 percent lower for that group compared to the reference group.

Summarizing the model, the hazard function is the product of an underlying duration-dependent risk $\lambda_0(t)$, common to all individuals, and another factor $e^{\beta z}$ that depends on the specific characteristics of each individual (Menken et al., 1981).[6]

The assumption of proportionality of hazards is not problematic for most of our covariates (see Morgan and Rindfuss, 1985). This proportionality assumption was checked for all the independent variables by inspecting the plot of the log minus log survival function for strata, as recommended by Menken et al.(1981) and the plot of the hazard rates given by life-table. Marriage cohort membership was the only independent variable that did not meet the assumption of proportionality. (It is worth noting that this problem will arise with a trend variable in proportional hazards models 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

[5] The percentage change associated with an independent variable category relative to the baseline hazard function is equal to $100(e^{\beta} - 1)$.

[6] The Maximum Likelihood Method is employed for estimating the coefficients β and their standard errors. The computer program used is BMDP-P2L (Dixon et al., 1981).

assumption of a common duration function.) Because of this lack of proportionality, and because of significant interactions between marriage cohort and number of our predictors, separate models were run for marriage cohorts.

RECENT TRENDS

The recent decline of about 8 percent in the crude divorce rate can be seen clearly in Figure 1. There are a number of reasons why we might expect a leveling, or even decline, such as that observed since 1980:

1) It is obvious that the acceleration after the mid 1960s cannot continue indefinitely since an absolute upper limit would be reached before too long. This trend *must* plateau sometime soon.

2) The marked reduction in marriage rates would imply 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. Further, 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.

3) Independent of changes in marriage rates, the movement through the population of the large baby boom cohorts first increased and then decreased the proportion of marriages in the high-risk young ages at marriage, and it had a similar net influence on the proportion of marriages at short durations (where divorce rates are highest).

4) Period measures such as the crude divorce rate will be inflated if divorce is occurring progressively earlier in marriage, even if the proportion of marriage cohorts that will ever divorce is constant. This means that a downturn in the period measures can be expected when the shift in timing has run its course.

5) Cyclical models based on cohort size predict a downward fluctuation (at least around the trend line) as the smaller birth cohorts of the late 1960s experience comparatively better economic opportunities than the larger cohorts that preceded them (Easterlin, 1978). Further, it is possible that the recessions of the early 1980s reduced the rate of divorce after separation (e.g., South, 1985), or even reduced the rate of separation by decreasing the ability to set up separate households.

What is clear is that the most recent period we can estimate from the June 1985 CPS shows the highest level of marital disruption yet recorded for the U.S. (Figure 2). This sharp increase has not yet appeared in divorce rates because the legal action of divorce lags the actual fact of separation and because the rate of divorce following separation has declined. (We find that the probability of divorcing within 5 years after separation declined by about 4 percent since the late 1970s. Almost all of this decline was concentrated in a 23 percent decline among blacks from 75 to 58 percent–the decline for whites was from 92 to 91 percent.)

Table 1 presents the life-table estimates of the cumulative proportion of marriages disrupted within 5 years, by order of marriage, for 3 marriage cohorts since 1970. While the fluctuation since 1980 has affected these cohorts at different durations, it is clear there has not been a downturn in the cumulative cohort experience–at least as far as we can compare it for the most recent cohorts. [7] Modest increases are observed for second marriages, and very substantial increases for third marriages (though based on a relatively small number).

It is appropriate at this point to examine the cohort implications of the separation rates we observe around 1984.[8] By its design, the total separation rate we have calculated implies that, if a cohort of first marriages were to experience the duration-specific rates observed during this period, 54 percent would have disrupted after 30 years. This is an underestimate of the cohort implications of recent rates for two reasons. First, the relevant period of risk (before widowhood) is longer than 30 years. Separation rates are lower at longer durations, but they are not zero. Secondly, we know that the CPS underreports marital disruptions in comparison to those tabulated by the vital statistics system (Preston and McDonald, 1979).[9]

[7] Though the first 5 years of marriage is a short period from which to draw general conclusions about the trajectory of marital disruption, disruption rates are highest in these early years. Period life tables for 1984 imply that nearly 50 percent of all marital separations over a period of 20 years will occur in the first 5 years.

[8] Based on experience for each one-year duration completed by the end of May, 1985 (cohorts entering each duration of risk June, 1983 to May, 1984) the period spans two years centering on May 1984.

[9] We have calibrated our estimates by two procedures that reach similar results. First, we compared them to the life tables for 1977 marriages (National Center for Health Statistics, 1980). The vital statistics life table was recalculated after adjusting upward the duration-specific rates by 10 percent to compensate for the fact that divorce rates are 10 percent higher for the total U.S. than for the Divorce Registration Area on which this table is based. A comparison of our estimate for this cohort to the adjusted vital

Our resulting estimate is that most recent rates imply that about two-thirds of all first marriages will end in separation within 30 years. This is not a prediction; we know well that such a period estimate can overstate (or understate) actual cohort experience because of either timing shifts or changing rates. Nonetheless, we think this estimate is likely to be realized. Note that this measure estimates what would follow *if duration specific rates remained unchanged*, which is to say, if the longterm trend has in fact plateaued and rates do not increase further. (The separation experience of the 1970 and 1975 marriage cohorts was higher at each age than the levels implied by the period rates for those years because the upward trend continued.)

Children's Single Parent Family Experience

The net stability of the early 1980s compared with the late 1970s would lead us to expect that little has changed from our previous estimates of children's experience based on the June 1980 CPS. Nonetheless, continued replication is important, especially in light of the minor controversy that surrounded the 1980 estimates (Hofferth,1985; Bumpass,1985). As we have repeatedly emphasized, period estimates are a *hypothetical* assertion about what would happen if rates of a particular time were to persist. Obviously, higher estimates

statistics estimate suggests that the proportion disrupting within thirty years is 23 percent too low in the CPS. However, it is possible that underreporting might be substantially less for the period just prior to survey. To evaluate this, we compared age specific rates of divorce calculated from the June 1985 CPS to adjusted vital statistics rates (for a period centering on 1983). This comparison indicates that the CPS estimates of cumulative survival by duration 30 are 21 percent too low. Thus the direct result from the CPS data of 51 percent separated by 30 years must be divided by .77 to .79 to adjust for underreporting.

would be obtained if a projected increase in rates were included in the model. While an increase may well occur in the future, Figure 4 illustrates the trend in our period estimates of children's single-parent experience.[10] The 1985 data closely replicate our estimates for the period previously estimated from the 1980 data, and show (as expected) a fluctuation after 1980 similar to that in Figure 1 for separations from first marriages to women. If rates centering on January, 1984 were to persist over the lifetime of a cohort, the CPS data estimate that about 46 percent of all children would spend some time in a single-parent family. Our previous analysis of children's experience did not take into account the underreporting of marital disruption in the CPS. It is not clear how to do so, since we do not know to what extent such underreporting varies with the presence of children. Nonetheless, if it were independent of parenthood, the adjusted estimate would be 58 percent of all children.

POPULATION DIFFERENCES

Increasing marital disruption has been shared by all groups in American society, as well as by many European countries (Sardon, 1986), suggesting common underlying factors. Nonetheless, marked subgroup differences persist. We examine differentials within first and second marriages separately, and then turn to a pooled analysis of the effect of

[10] As in our previous estimates, births reported by mothers in the June CPS become the units of analysis and premarital births as well as those whose fathers died are included in the definition of single-parent experience. Given the lack of change, the reader is referred to the earlier article for a more detailed treatment.

remarriage. Within each analysis, we consider single variable life-estimates and the net effect of each variable as estimated from the proportional hazards model.

First Marriages

Table 2 presents life-table estimates of the proportion of first marriages disrupting within five years for three marriage cohorts, and the proportional hazards results are presented Table 3. 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 and between race and several of the predictor variables.[11] Consequently, the proportional hazards results are presented separately by marriage cohort and by race. We will consider these two tables jointly because of the consistency of the observed and multivariate effects.

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 (Moore and Waite, 1981; Teachman, 1983; Kiernan, 1986). This association persists net of education and premarital pregnancy (Bumpass and Sweet, 1972) and at higher marital durations (Morgan and Rindfuss, 1985). A number of mechanisms have been suggested

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

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 et al., 1977) and emotional, educational and economic resources available. It has also been argued that marrying late (i.e., outside the normative range of the 20's for women) might also reduce stability, suggesting a nonmonotonic relationship (Glick and Norton, 1977), though this was not found in an analysis of the 1980 census data (Sweet and Bumpass, 1987).

The high disruption rate of teenage marriages is apparent in Tables 2 and 3. Women who married as teenagers are twice as likely to separate as those who married after age 22. We do not find evidence of a higher risk among women marrying over age 30. The effect of age at marriage is little altered by controls for the other variables considered. One interesting change in the age at marriage effect occurs among women first marrying at ages 20 to 22: the life tables show nearly a 40 percent increase in the probability of separation during the period when total rates increased only slightly. One possible interpretation for this is that the definition of "young" marriage is changing as the average age at marriage has become progressively older. In the current setting, what once characterized teenage marriages may now extend to marriages in the early 20s.

An important race difference in Table 3 is the apparent disappearance of an age at marriage effect on marital stability among blacks. Whereas this effect was of a similar order of magnitude among blacks and whites married in the early 1970s, there is no significant effect of age at marriage among black women married after 1975. While it is not immediately apparent how to interpret this change, the longstanding theoretical and empirical centrality of age at marriage suggests that this finding may prove useful for understanding increasing racial differences in marital patterns.

Education

Education is also 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 that may affect interpersonal skills. (In interpreting education "effects" it is also important to keep in mind the selective nature of those who do not complete high school.) Earlier analyses have found that much of the effect of education is attributable to the high correlation between age at marriage and level of educational attainment (Bumpass and Sweet, 1972).

The likelihood of separation increased at lower education level, increasing the size of education differences in marital stability among more recent marriage cohorts. Women who attended college have a 51 percent lower rate of separation than high school dropouts among whites and about a 60 percent lower rate than high school dropouts among blacks. (About half of this effect is attributable to differences by education in age at marriage.)

Children Born Before Marriage

It is not surprising that beginning marriage with a child, especially in circumstances where the husband is not the biological father, would increase the risk of marital disruption (Furstenberg, 1976; Teachman, 1983). In addition to possible increased strains on marriage, women who have already experienced single parenthood may be less reluctant to return to that status (Morgan and Rindfuss, 1985). While there is growing attention to the fact that one of every five children begin life with an unmarried mother (NCHS, 1987), the implications of these patterns for the conditions under which many first marriages begin have been little appreciated. One of every seven recent first marriages involved a woman who already had a child (Sweet and Bumpass, 1987).

In Table 2, we see a substantial increase in marital disruption among women who began marriage with a child. This is of particular interest given the marked increase in rates of unmarried childbearing among white women of all ages (NCHS, 1987) and the correlated increase in the proportion of first marriages that begin with a child. There is no way with current data to determine what proportion of these were cohabiting relationships in which the child was also the husband's, and hence the effect may reflect a greater representation of cohabiting relationships as well as strains introduced by the presence of the child. The effect of a premarital birth on marital disruption is concentrated almost wholly among whites (a 71 percent higher rate than among whites without a premarital birth, averaged over the period since 1970). 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 fact that being an unmarried mother is a much more "deviant" status among whites than among blacks; hence, among whites, it 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 tradition of individualism of some regions (Fenelon, 1971) or of a Durkheimian state of anomie resulting from population growth and migration (Gleen and Supancic, 1984; Gleen and Shelton, 1985).

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, separation rates were 50 percent higher in the West, net of other factors. By the early 1980s, there are no significant regional effects for either whites or blacks (Table 3).

Race/Ethnicity

Because of a number of interactions with race, the preceding discussion has focussed on differing effects by race of other predictor variables; we turn now to the markedly different *levels* of marital disruption between blacks and whites. The higher rates of marital instability among blacks have been the subject of considerable discussion recently, with attention returning to 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, 1987), the average duration in married life has been reduced among black women to about one-third of nonwidowed adult years, compared to about two-thirds among white women (Espenshade, 1985).

It is evident in Table 2 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. (This increase, and the increase among lower education groups noted earlier, may reflect the destabilizing influence of the recent recession.) Despite the increase among Hispanics, the level is only slightly higher than among nonHispanic whites, and is actually lower once composition on the other variables is taken into account (Parker, 1986). The interactions between race and the other predictors preclude a straightforward estimation of the *net* effect of race. 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.[12] The expected rate of disruption within 15

[12]For whites, the low risk estimate is for women who married at age 23 or older between 1970 and 1972

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 and 92 percent respectively. Thus there is much more racial convergence among high risk than among low risk groups: very substantial differences persist between blacks and whites among women our model would classify as having the lowest risk of marital disruption.

Second Marriages

As for first marriages, life-table estimates of the proportion of second marriages disrupting within five years are presented for three marriage cohorts in Table 4, and the proportional hazards results for second marriages are presented in Table 5.[13] These results can be presented succinctly.

Age at Second Marriage

The effect of *age at marriage* continues to be a powerful predictor of disruption for

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 regions 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 who had the opposite characteristics (from the low-risk group) on each of the other variables.

[13]For second marriages, there were no significant interactions between our predictor variables and either marriage cohort or trend. Hence we present only a combined proportional hazards model.

second marriages, though the form of the relationship and its interpretation are quite different. While age may be seen as an indicator of "maturation" relevant for mate selection in first marriage, this should be less relevant in second marriages since the age distribution is quite different. Other factors may be involved, this time perhaps at the older end of the age distribution, such as concern about reduced chances of a third marriage or feeling outside the "normative" ages for initiating new relationships or changing life-style.

However, as noted by Sweet (1977), there are several demographic factors that make the interpretation of age at second marriage difficult. On the one hand, second marriages should be less likely to experience separation, because almost no remarriages occur in the high-risk teenage years. On the other hand, these marriages are selected from among women who were younger on the average when they initiated their first marriage, and this factor may have a persisting negative effect even in second marriage.

The effect of age at remarriage may have increased over the last two cohorts (though the trend interaction is not significant) with increased proportions separated among those married before age 25 and some decline among those married after age 40 (Table 4). Because there are fewer very young marriages, differences are smaller under age 40 (about a 25 percent reduction at ages 25 and over compared to those under 25). However, remarrying after age 40 reduces the risk by 60 percent (Table 5).

The effect of *education* on the risk of disruption is less for second marriages than for first marriages, primarily because of the role of teenage marriages in the latter. Compared to women not completing high school, having attended college reduces the risk of second marriage disruption by about 18 percent compared to about a 25 percent reduction in the model for first 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 [14] Nonetheless, a number of persons have argued that such prior children are likely to create strains affecting the stability of remarriages. Cherlin (1978) explicitly identifies stepchildren as the principal destabilizing element in remarriages. The literature reports conflicting results regarding this subject. While some of the studies support Cherlin's hypothesis (Becker et al, 1977; White and Booth, 1985), other studies do not (Furstenberg and Spanier, 1984). We find a 14 percent higher net rate.[15] Hence these data suggest that there is very likely a higher disruption rate among remarriages with prior children, but the difference is modest.

We noted the disappearance over time of the effect of *region* on the stability of first marriages; we find no significant regional effect on the stability of remarriages, once the composition on other variables is taken into account.

The effect of *race* is very similar in first and second marriages. This is contrary to the findings of McCarthy (1978), who argues that remarriages are more stable for blacks. In spite of the fact that markedly reduced remarriage rates makes second marriages among blacks more selective, the net rate of marital disruption among blacks in second marriages

[14]About one-third of all nonmarital births, about 7 percent of all births, occur after marital disruption (McLanahan and Bumpass, 1986).

[15]The β is almost twice its standard error.

is of the same magnitude as that we observed in an additive model for first marriages.

The Net Effect of Remarriage on Dissolution Rates

As documented in Table 1, second marriages are less stable than first marriages (McCarthy, 1978), and several explanations have been offered to account for this. Cherlin (1978) argues that the ambiguity of norms and the complexity of the family structure decrease the chances of survival of second marriages. 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); i.e., remarriages may be selective of individuals who are not opposed in principle to divorce, while first marriages include a certain proportion of "stayers" who will never consider divorce (because of religious or other moral and social constraints).

It is evident from Table 6 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 (33 compared to 22). The educational level is somewhat lower for women in their second marriage. Blacks and Hispanics represent a smaller proportion in second marriages (though blacks are more likely to separate from a first marriage, they are less likely to remarry; Hispanics are more likely to remain in their first marriages). Fertility status at the time of the marriage is also very different: 12 percent of women enter first marriage with a child, whereas 76 percent of the women entering second marriages are already mothers.

If we compare the distribution of age at *first* marriage, we observe that a remarkably high proportion of remarried women married for the first time as teenagers (reflecting the higher rate of dissolution of teenage first marriages and the higher remarriage rate associated with re-entering the marriage market at younger ages).

The straightforward way to address the effect of remarriage net of compositional differences would seem to be through pooled hazards models for first and second marriages including remarriage as a covariate. This is complicated, however, by the interactions by race and marriage cohort and by the low overlap between the age distributions of first and second marriages: there are very few teenage second marriages and very few first marriages over age 30. Further, there is a difficulty in interpreting age at marriage effects–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 between 1980 and 1985. Among these recent marriages, remarriages are 25 percent more likely than first marriages to disrupt (first column, Table 7). [16]

When we take into account the lower educational distribution of remarriages, this difference declines to a 15 percent higher rate. When age at *first* marriage is added, we obtain the surprising result that there would be no difference at all in the dissolution risks of first and second marriages if it were not for the differential composition on education

[16] Prior fertility is not included in the pooled models since it has a rather different social meaning for first and second marriages that would make results difficult to interpret. Region is also not included.

and age at first marriage. While it is not clear what factors are involved, this result 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 partially account for the higher disruption rates of remarriages.

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 carefully structured a comparison for white women marrying 1980-84 at ages 25-29 and obtained the result that (net of education) remarriages were 60 percent more likely to disrupt than first marriages. This result is misleading, however, for the simple reason that 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. Ages 25-29 were selected in the example above because these are the ages at which there is the most overlap between the distributions of first and second marriages. Nonetheless, women marrying for the first time at these ages are marrying at "older" ages of first marriage (and thus are among those 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 must conclude that, net of compositional differences with respect to education and age at first marriage, remarriages have no higher risk of disruption than first marriages.

LINKAGES WITH OTHER SOCIAL CHANGES

Decreasing marital stability is just one of many indicators of the shrinking dominance of family roles and obligations. For example, marriage rates in the peak marriage ages dropped by one-half in a decade (Sweet and Bumpass, 1987), and an increasing proportion of young people may never marry; parenthood has been markedly delayed, and perhaps a quarter of young women may never have a child (Bloom and Trussell, 1984); and unmarried young adults and elderly have been increasingly likely to live alone rather than with other family members (Sweet and Bumpass, 1987). There is a growing consensus that competing alternative adult roles are likely to further erode the structural supports of family roles (e.g., Keller, 1971;Bernard, 1972;Westoff and Ryder, 1977;Westoff, 1978;Ryder,1979;Butz and Ward, 1979;Huber, 1980;Davis, 1983;Lesthaeghe, 1983). In the simple demographic metric of years of experience, marriage and family relationships are clearly occupying a shrinking space in our lives (Schoen et al., 1985). Rather than simply an aberration of the late 20th century, these trends are most likely a continuation of the longterm reduction of family functions that has occurred in conjunction with the transformation of our economy (Ogburn, 1954), and with an increasing cultural value on individualism (Lesthaeghe, 1983).

Preston and McDonald's estimates of cohort divorce rates (1979) strongly suggest that basic structural changes are at issue. While annual rates of divorce have fluctuated with wars and the economy, the likelihood of divorce has increased in a steady, accelerating curve, from 7 percent for first marriages in 1860 to the current estimate of perhaps two-thirds of recent marriages. The duration of this trend strongly argues 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. At the same time, many of these changes have been facilitated by delayed marriage and parenthood, as these delays have exposed both men and women to lifestyles competing directly with family roles.

First, it is possible that changing roles within marriage reduce the advantages gained from a domestic division of labor, and hence the net benefit of marriage. The lives of women have been most dramatically altered by the growth in their education and labor market participation, creating in particular a new capacity for independence from marriage (Hannan, Tuma, and Groeneveld, 1978). Men's lives have also been directly altered by these changes. While there has been only a modest increase in shared household tasks (Pleck, 1985), the role of breadwinner is progressively shared. Hence men's claim to home services in exchange for this role is weakened by substitution of purchased services (from cleaning to meals eaten out). These changes in the complementarity of marital roles may well alter how the advantages of marriage are perceived (Becker, 1981) and the qualities sought by men and women in the marriage market.

Second, a critical aspect of the changing meaning of marriage is that marriage and sex have been effectively separated. While marriage is being delayed, sexual relationships begin at increasingly earlier ages (Zelnik and Kantner, 1980), and the norms are increasingly approving (Thornton and Freedman, 1983). This is the culmination of a process that Shorter argued began with industrialization (1973), but the high prevalence and approval of sexual relations among the unmarried makes delayed marriage (or remarriage) less burdensome–also reducing the net benefit of marriage.

It is in this context that cohabitation has become increasingly common. Disapproval of unmarried couples living together must have rested largely on the way this behavior publicly acknowledged unmarried sexual intimacy. That objection has been effectively destroyed by the knowledge that sexual relationships are common regardless of living arrangements. The decreasing net benefit of marriage associated with changing roles within marriage, and considerable sexual opportunity in the absence of marriage, is reflected in attitudes demonstrating a decreasing sense of the importance of ever marrying (Thornton and Freedman, 1983).

Third, changing marital stability has greatly altered the context of, and expectations about, family life. That the majority of recent marriages are not likely to be permanent may well have feedback effects on marital stability. Everyday observation blatently contradicts norms that marriages are "until death," or that currently married persons are not in the marriage market (permanent availability as noted by Farber, 1973). More than half of recent marriages involved at least one partner who had been previously married, and both partners had been married before in one-quarter of recent marriages (Sweet and Bumpass, 1987).

There may be further feedback, in that the investment persons are willing to make in a relationship is likely to be hedged to the extent that the permanence of the relationship is seen as uncertain (Becker, 1981, p.224). Put simply, insecurity about the viability of relationships may lead to behaviors that, in turn, lower the prospects that the relationship will last.

Other effects of marital uncertainty may occur as young women increasingly believe that it is essential to develop and nurture labor market skills capable of self-support.

Fourth, a similar connection may be involved in the marked increase in *nonmarital* fertility among whites (over a two-thirds increase at all ages since the mid 1970s, NCHS, 1984). Levels of marital stability make it clear that marriage provides little guarantee of a stable parental environment for children. Recognition of this fact may well be loosening the conviction that it is necessary for a woman to be married to have a child. All the problems aside, there is ample evidence that mothers do cope in the absence of a husband;and a single pregnant girl may well question the value of marrying only to divorce a few years later.

Fifth, and finally, there is an ideological component of these changes that is also related to the long term structural transformation of our society. The deep historical roots of increasing individualism have been noted by several authors (Shorter, 1973;Stone, 1982 as cited in Preston, 1984; Lesthaeghe, 1983) Similar themes emerge from a different starting point in the recent book by Bellah and colleagues (1985). When needs and interests conflict, as they do in any collectivity, how much weight must the individual actor give to the interests of other persons in contrast to his or her own, and how large a circle of others must be considered? Over the centuries this circle has narrowed increasingly to the individual. A very important dimension of this trend is the extension to women of the legitimacy of self-interest as a decision criterion that was earlier appropriate only for men (Pankhurst and Houseknecht, 1982).

Divorce is a clear case where the interests of children and adults in the family are often very different (e.g., Wallerstein and Kelly, 1980). The proportion saying a couple should stay together for the sake of the children increased from 51 to 81 percent between 1962 and 1982 (Thornton and Freedman, 1983). Indeed, as we have noted, about half of today's youth will have spent some time in a one-parent family; when combined with their own chances of marital success, only a minority will have stable two-parent families in both childhood and as an adult.

That the family is adaptive, and that family relationships continue to play a very important role in the lives of Americans, are indisputable. At the same time, it seems likely that the relative dominance of these relationships in competition with other adult roles is likely to continue to decline. The forces that have given rise to the long term trends are likely to continue. 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. This suggests that social policies to ameliorate the negative consequences of these patterns–such as the frequent poverty of women and children in single-parent families (Garfinkel and McLanahan, 1986; McLanahan and Bumpass, 1986)–are more apt than concerns with restoring a nostalgic family of the past.

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

	Marriage Cohort			% Change		N		
Marriage Order	70-74	75-79	80-85	70-74 75-79	74-79 80-85	70-74	75-79	80-85
First Marriages	.17	.22	.23	25.1	5.5	5345	5028	5556
Second Marriages	.18	.24	.27	36.0	11.6	1198	1567	2088
Third Marriages	.18	.20	.38	13.5	89.1	165	238	451

Table 2. Life Table Estimates of the Proportion of Marriages Disrupted Within Five Years, for First Marriage Cohorts

	Marriage Cohort			% Change		N		
Marriage Order	70-74	75-79	80-85	70-74 75-79	74-79 80-85	70-74	75-79	80-85
Age at Marriage								
10-19	.23	.30	.31	29.7	4.0	2266	1849	1477
20-22	.14	.18	.26	28.5	38.4	1733	1641	1680
23-29	.11	.15	.15	40.0	0	1126	1282	1978
30+	.14	.16	.14	15.5	-13.4	220	256	421
Education								
0-11	.21	.29	.33	36.8	14.5	709	707	765
12	.18	.23	.26	25.3	13.6	2408	2251	2421
13+	.16	.18	.16	17.3	-9.8	2228	2070	2370
Children Before Marriage								
0	.17	.21	.21	21.2	1.9	4799	4411	4799
1+	.22	.31	.36	43.4	13.4	546	617	757
Region								
Northeast	.13	.17	.21	29.7	24.1	1179	1064	1159
Northcentral	.16	.20	.23	21.5	14.6	1307	1226	1298
South	.18	.25	.25	38.0	2.4	1694	1528	1766
West	.23	.24	.22	5.6	-9.9	1165	1210	1333
Race/Ethnicity								
White	.17	.21	.22	22.5	4.7	4290	4058	4403
Black	.24	.30	.36	25.2	17.2	490	431	472
Hispanic	.15	.21	.24	33.8	16.0	337	341	447
TOTAL	.18	.22	.23	25.1	5.5	5345	5028	5556

Table 3. Proportional Hazards Estimates* of Effects on the Dissolution of First Marriages, by Marriage Cohort and Race

	White Women, First Married:				Black Women, First Married:			
	70-74	75-79	80-85	70-85	70-74	75-79	80-85	70-85
Age at Marriage								
10-19	.60	.61	.81	.63	.50	.76#	1.12#	.66
20-22	.42	.48	.50	.44	.50	.89#	.78#	.62
23-29	.34	.52	.46	.42	.53	.80#	1.12#	.67
30+								
Education								
12	.84	.81	.71	.80	.87#	.95#	.58	.84#
13+	.86#	.68	.49	.73	.74#	.97#	.40	.75
Children Before Marriage								
0								
1+	1.75	1.57	1.80	1.71	1.04#	1.43	1.00#	1.16
Region								
Northcentral	1.03#	1.06#	1.00#	1.04#	1.49#	1.12#	1.22#	1.30#
South	1.07#	1.34	1.18#	1.18	1.00#	.92#	.80#	.94#
West	1.52	1.52	1.12#	1.45	1.07#	1.16#	1.07#	1.10#
N	4290	4058	4403	12751	490	431	472	1393

* exp (β)

\# β less than twice its standard error

Table 4. Life Table Estimates of the Proportion of Marriages Disrupted Within Five Years, for Second Marriage Cohorts

	Marriage Cohort			% Change		N		
Marriage Order	70-74	75-79	80-85	70-74 75-79	74-79 80-85	70-74	75-79	80-85
Age at Marriage								
10-19	.26	.33	.40	24.2	22.0	291	313	374
20-22	.15	.24	.26	60.0	7.5	260	446	538
23-29	.17	.24	.27	42.8	13.3	317	470	748
30+	.13	.17	.14	27.3	-13.7	330	338	428
Education								
0-11	.17	.24	.36	42.8	50.8	329	317	389
12	.17	.25	.26	50.3	1.6	555	743	997
13+	.20	.23	.22	11.3	-0.4	314	507	702
Children Before Marriage								
0	.16	.23	.24	45.8	8.4	238	378	540
1+	.18	.25	.28	35.0	13.0	960	1189	1548
Region								
Northeast	.16	.20	.22	19.6	10.8	187	262	343
Northcentral	.21	.24	.28	15.0	15.5	285	351	489
South	.17	.26	.27	55.9	1.5	402	567	743
West	.16	.23	.30	42.8	28.3	328	387	513
Race/Ethnicity								
White	.18	.23	.26	24.7	12.8	1026	1337	1792
Black	.21	.37	.43	77.5	15.9	90	123	142
Hispanic	.10	.26	.28	168.0	8.1	50	58	87
TOTAL	.18	.24	.27	36.0	11.6	1198	1567	2088

Table 5. Proportional Hazards Estimates of Effects on the Dissolution of Second Marriages. Women. 1970-1985

	β	$z(\beta)$	$\exp(\beta)$
Age at Second Marriage			
(10-24)			
25-29	-0.310	(-3.9)	0.73
30-39	-0.340	(-4.4)	0.71
40+	-1.005	(-10.2)	0.37
Education			
(0-11)			
12	-0.223	(-3.0)	0.80
13+	-0.201	(-2.4)	0.82
Children from Previous Marriage			
(0)			
1+	0.132	(1.8)	1.14
Region			
(Northeast)			
Northcentral	0.151	(1.5)	1.16
South	0.062	(0.7)	1.06
West	0.117	(1.2)	1.12
Race/Ethnicity			
(White)			
Black	0.419	(4.2)	1.52
Hispanic	-0.030	(-0.2)	0.97

Table 6. Demographic Composition of First and Second Marriages

	First Marriages (%)*	Second Marriages (%)*
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	6.1
Age at First Marriage		
10-19	35.4	61.8
20-24	44.9	31.0
25+	19.8	7.3

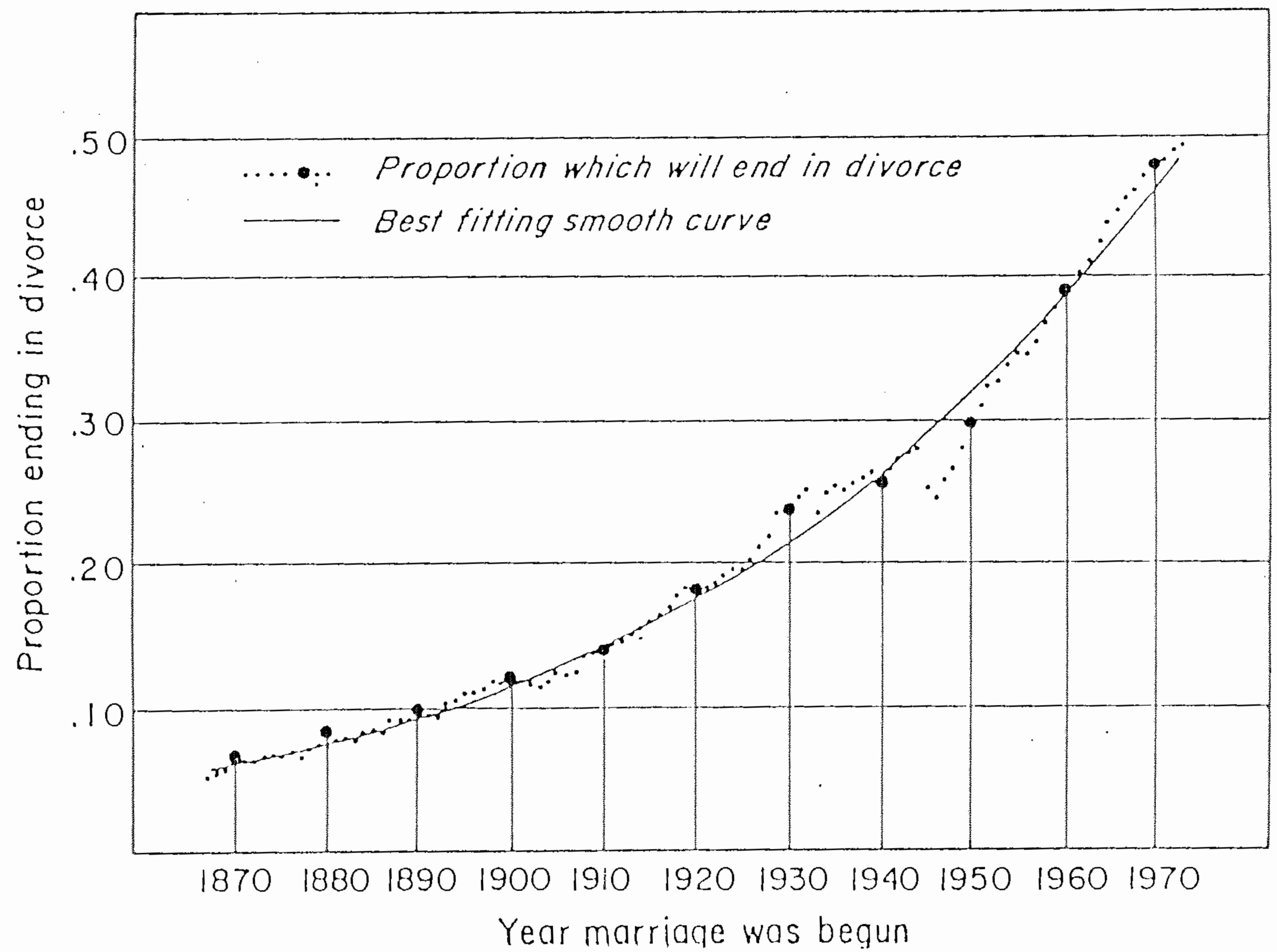
* Weighted percentages

Table 7. Selected Proportional Hazards Models of the Dissolution of First and Second Marriages. White Women, 1980-1985

	Gross Effects			Model A			Model B		
	β	$z(\beta)$	$\exp(\beta)$	β	$z(\beta)$	$\exp(\beta)$	β	$z(\beta)$	$\exp(\beta)$
Marriage Order									
(1)									
2	0.222	(2.9)	1.25	0.138	(1.8)	1.15	-0.0002	(0.0)	1.00
Education									
(0-11)									
12	-0.496	(-5.4)	0.61	-0.483	(-5.2)	0.62	-0.421	(-4.5)	0.66
13+	-0.977	(-9.5)	0.38	-0.955	(-9.2)	0.34	-0.728	(-6.6)	0.48
Age at First Marriage									
(10-19)									
20-22	-0.322	(-3.9)	0.72				-0.159	(-1.8)	0.85
23-29	-0.913	(-8.8)	0.40				-0.701	(-6.2)	0.50
30+	-1.056	(-4.5)	0.35				-0.859	(-3.6)	0.42

Figure 1

Proportion of marriages begun in each year that will end in divorce, 1867 to 1973.



Source: Cherlin 1981

Figure 2

DIVORCE RATES PER 1000 TOTAL POPULATION. 1920–1984.

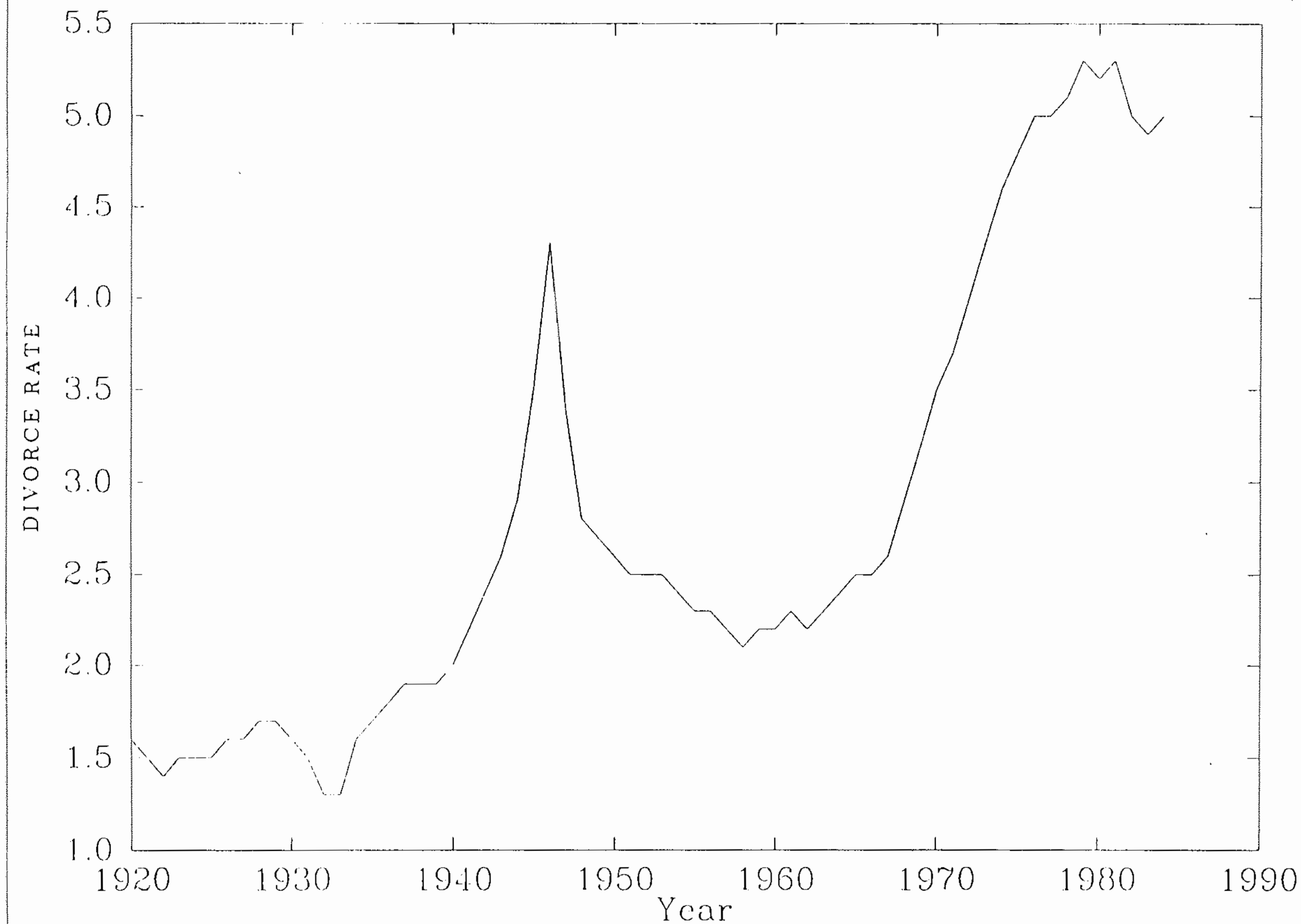


Figure 3

Period estimate of separation by 15 years,two year moving average

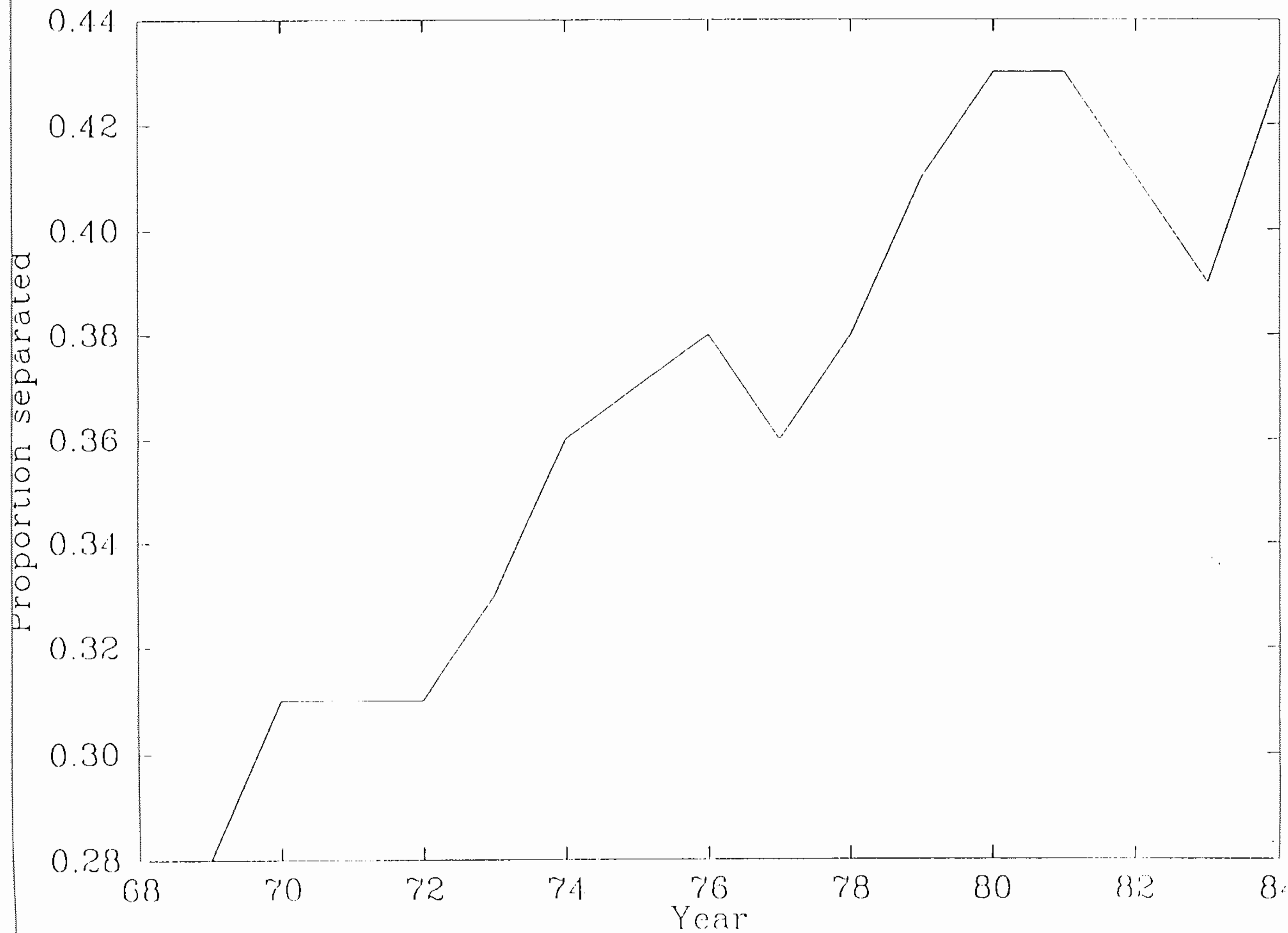


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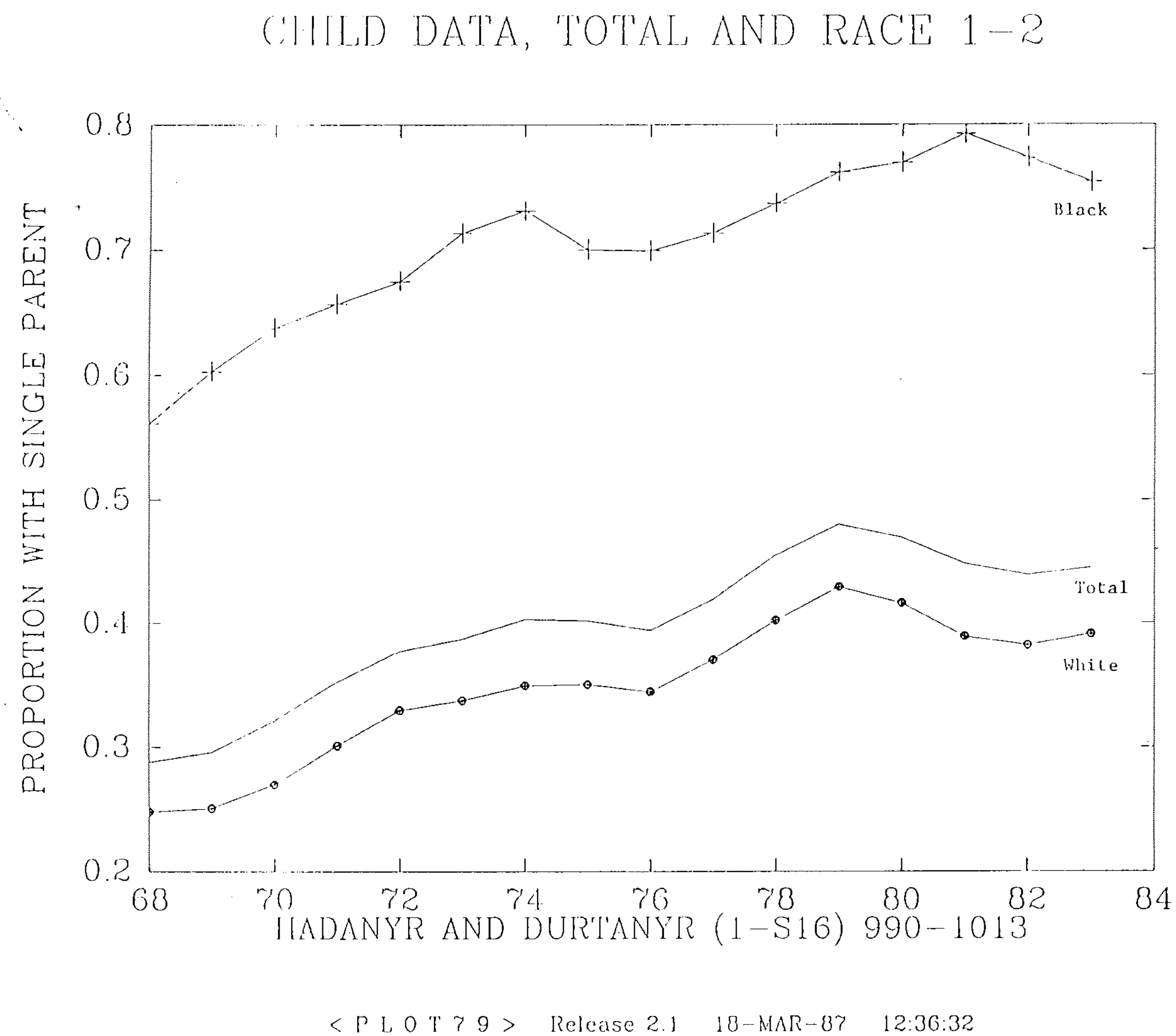
Figure 4


CHILD DATA, TOTAL AND RACE 1-2
PROPORTION WITH SINGLE PARENT
0.8
0.7
0.6
0.5
0.4
0.3
0.2
68
70
72
74
76
78
80
82
84
Black
Total
White
HADANYR AND DURTANYR (1-S16) 990-1013
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