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Effects of Childhood Family Structure on the Transition to Marriage

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Increasing rates of marital dissolution mean that many more children than in the past spend part of their childhood in single-parent families. This paper explores the effects of family structure during the teenage years on the likelihood of marriage later for both males and females. Using data from two national longitudinal surveys of young people, the analysis finds that childhood family patterns do influence the later family formation of the children involved. However, the experience of disruption of parental marriage affects sons and daughters and blacks and whites somewhat differently. The paper describes these patterns in detail and discusses their implications.

INTRODUCTION

Since the 1950s the American family has seen a major retreat from the pattern of early, stable, and nearly universal marriage that characterized that period. The most recent cohorts have been marrying slowly, and those marriages have become increasingly likely to end through separation or divorce (Espenshade, 1983). A wide range of factors has been adduced to account for these phenomena, including economic changes, e.g., the more rapid rise of unemployment among young people, (Easterlin, 1978); demographic factors, particularly the "marriage squeeze" (Heer and Grossbard-Schectman, 1981); as well as the great social transformation that has enhanced women's ability to support themselves and any children outside of marriage, both through paid employment and through public assistance (Waite, 1981; Ross and Sawhill, 1975).

These most recent cohorts, however, not only have experienced many social, economic, and demographic changes in their public environment; they also are increasingly the products of these patterns of family change. While growing up they have experienced both the rise in nonmarriage—as more have been born out of wedlock—and the increase in divorce, by spending decreasing proportions of childhood with both natural parents, compared with earlier cohorts. Bumpass and Rindfuss (1979) find that one-third of white and 60% of black children born to ever-married mothers will experience a parental divorce at some point during their childhood if divorce rates remain at about their level during the early 1970s. In addition to the weight of external factors, then, the retreat from marriage may be fueled by inheritance, in which those who experienced less traditional family forms as children are more hesitant about embarking themselves on a traditional family-building schedule.

Findings in the research literature on the heritability of divorce provide one suggestion that this may be so. Those whose parents divorced have been shown to be more prone to divorce

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themselves (Glenn and Shelton, 1983; Pope and Mueller, 1976). Interpretations suggested include acquaintance with the divorce process so that divorce is a familiar, if painful, option; differences in aversion to divorce; and a more general explanation suggesting that children who have not experienced successful families are less able to choose an appropriate partner and maintain an enduring relationship because they lack the necessary social skills. These explanations all imply that not only might children of divorce separate more easily if they do marry; they also might defer marriage, because of less expertise in the marriage market, or avoid it altogether, since the experience of pain might make them more careful not to risk repeating the process.

This prediction about the heritability of nonmarriage, as well as divorce, has already been made, in fact, for one segment of American society. Dubbed a "tangle of pathology," the black family was characterized in the Moynihan report (Rainwater and Yancey, 1968) as having a large proportion of both illegitimate children and children raised in a single-parent household because of marital instability; and Moynihan predicted that they would repeat the cycle themselves when they reached adulthood, further distinguishing the family patterns (and by arguable inference, the life chances) of black and white Americans.

Recent analyses of the transition to marriage have shown that blacks are, in fact, significantly less likely to marry than whites, holding constant other factors influencing the transition to marriage (Waite and Spitze, 1981; Michael and Tuma, 1983; Spanier and Glick, 1980). These lowered marriage rates may, as Moynihan predicted, reflect the greater family instability blacks experienced while growing up. However, although black women have had much higher levels of marital instability than white women, the extent of marital dissolution through separation and divorce experienced by black women in the late 1950s (the period of Moynihan's data) was reached by white females in the early 1970s; and by the late 1970s white rates had increased even more (Espenshade, 1983). Similarly, recent increases in the proportion of white out-of-wedlock births also have brought whites close to the black levels of the 1950s (Hofferth, 1982). It is clear that the phenomenon is a general one, rather than one that characterizes only the black family, and, to the extent the prediction of heritability is reasonable, a phenomenon that could affect white as well as black family formation.

The Moynihan prediction was never tested systematically, either for marital instability or for nonmarriage, at the time it was raised. Few data sources were available that would allow such a test. Individual-level data on young people that include some information on their parental household are needed to test for the presence of a true intergenerational effect. Furthermore, the dynamics of the timing of the process of entry into marriage and its subsequent dissolution require either longitudinal or retrospective data that follow individuals through an extensive segment of their adult years. In addition analyses of data that could test the hypothesis have been generally restricted to women; yet the thrust of Moynihan's argument was on the family/stability/poverty nexus, and most of his images were of the fate of black adolescent boys.

This paper is able to meet the criteria by using data from the National Longitudinal Surveys of the Education and Labor Market Experiences of Young Men and Young Women (NLS). We use these data to estimate the effect of variation in parental household structure on the timing of entry into marriage for both males and females, covering the age range of 17 to 29 for women, and 18 to 29 for men.

HYPOTHESES

This research focuses on a series of questions relating childhood family structure to later marriage for blacks and whites, males and females, and the presentation reflects this focus. In addition, the analysis controls for other factors that have been found in related research to affect the timing of entry into marriage. This has not been done previously in parallel analyses spanning the full period of entry into marriage for both sexes.

Here we ask two central questions:

1. What is the impact of childhood family structure on later family formation? We consider two dimensions of childrens' experiences of family disruption: living in a female-headed household at age 14, and living in some other family type at that age—not including two natural parents—the largest single category of which comprises natural mother-stepfather families. This distinction allows us to assess whether the effects we observe are due to disruption per se, even if remarriage occurs to recreate a nuclear family structure, or whether these effects relate more to continuing differences in household structure. Disrupted families that reformed through remarriage presumably would have different effects on a young person's attractiveness to a prospective mate or willingness to leave home, compared with female-headed families, and these effects might differ for sons and daughters.

2. Does childhood family experience affect later family formation similarly for blacks and whites? If so, the greater experience of blacks of childhood family instability could account in part for the differences observed between the two racial groups in rates of first marriage. Several considerations lead us to expect different effects of childhood family structure for blacks than for whites. First, the much higher rates of out-of-wedlock childbearing and divorce for blacks than for whites mean that black children more often have close contact with family types besides two-parent families, through friends, neighbors, relatives or classmates. Furthermore, the much greater frequency of family types besides the stable two-parent family among blacks may mean the social significance of marriage differs substantially for the two races.

Throughout, we consider how these effects might differ for men and for women.

Data and Methods

Data for this analysis come from the National Longitudinal Survey of Young Women and from the National Longitudinal Survey of Young Men. Conducted by the Ohio State University Center for Human Resource Research, these surveys include information over a recent 15-year period on more than 5,000 young women and more than 5,000 young men. These surveys conducted personal interviews with national probability samples of the noninstitutionalized population of females 14 to 24 in 1968 and males 14 to 24 in 1966. The "Young Women" and "Young Men," as the respondents to these surveys have been called, responded to lengthy interviews in many of the succeeding years through 1980.[1] Attrition from the sample over the panel period has been low; three-fourths of the original Young Women's sample was reinterviewed in 1978 and two-thirds of the original Young Men's sample was reinterviewed in 1980, the last years used for this analysis.

The analysis reported in this paper uses measures of stable respondent characteristics, such as race and family socioeconomic status, measured in the initial survey (1966 for the Young Men, 1968 for the Young Women). Independent variables measuring the respondent's current situation—for example, employment or education—reflect the value at the beginning of the year in question. Table 1 gives definitions, means, and standard deviations of all variables used in this analysis.

Our model of the process of first marriage for males and females builds on previous work on this topic in selection of independent variables (see Waite and Spitze, 1981; MacDonald and Rindfuss, 1981; and Hogan, 1978). In this analysis we focus on the impact of childhood family structure on later family formation and on racial differences in this effect. Thus, the other independent variables in our models act primarily as controls. These variables include parental education, the respondent's educational attainment, school enrollment, employment, region of residence, size of place, year, and whether living away from the parental home; plus, for females, plans for work at age 35 and presence of own children; and, for males, active military service. They have been included in every model presented, and their influence does not vary among the models for a given sex. Some sex differences appear, and these are discussed briefly later in the paper.

We use three measures of structure of the parental family, all referring to the structure when the respondent was age 14. These include whether the family included the respondent's natural mother and father (Intact); whether the family was headed by the respondent's mother or another adult female with no adult male present (Female-headed Household); and other family structures, including stepparent families, families headed by the father only, and families headed by other relatives of the respondent (Other Nonintact). About three-quarters of the respondents lived in intact families during their early teens, and approximately equal proportions lived in female-headed and other family types (10%–12%). As we would expect, however, substantially fewer blacks than whites lived with both natural parents at age 14 (about 55%–60%).

This analysis uses observations on each respondent over a number of one-year periods. The dependent variable in this analysis is first marriage during the year for those never married at the beginning of the year. Because this process differs substantially by respondent's age—at least for women (Waite and Spitze, 1981)—we estimate all models separately by single years of age. Those who had never married by the interview date corresponding to a particular age—say, age 18—constituted the sample eligible to marry for the first time during the coming year, in this case by age 19. Each respondent supplies annual observations until he or she marries and none thereafter. Because we estimate all models separately by year of age, no respondent ever appears more than once in a single model.

Since the dependent variable in this analysis can have codes of only 0 or 1, we estimated all equations using logit, a maximum-likelihood technique

TABLE 1. VARIABLE DEFINITIONS, MEANS AND STANDARD DEVIATIONS (STANDARD DEVIATIONS IN PARENTHESES)

		Males		Females	
Variable	Coding	Age 18[a]	Age 25	Age 18	Age 25
Wed	Whether marrying for the first time during the year (1 = yes, 0 = no)	.03 (.17)	.11 (.32)	.11 (.31)	.06 (.25)
Parents' education	Average years of schooling completed by parents	10.17 (3.06)	10.69 (3.32)	10.47 (3.17)	10.44 (3.26)
Intact	Whether two natural parents were present when respondent was age 14 (1 = yes, 0 = no)	.80 (.40)	.81 (.39)	.77 (.42)	.79 (.41)
Female-headed household	Whether household head was female when respondent was age 14 (1 = yes, 0 = no)	.13 (.33)	.11 (.32)	.12 (.32)	.16 (.36)
Other intact	Other family type when respondent was age 14 (1 = yes, 0 = no)	.07 (.26)	.07 (.26)	.11 (.31)	.06 (.23)
Black	1 = yes, 0 = no	.29 (.45)	.26 (.44)	.30 (.46)	.42 (.49)
Education	Years of schooling completed	10.46 (1.33)	13.35 (2.64)	10.42 (1.13)	13.15 (2.92)
Enrollment	Enrolled in school full-time (1 = yes, 0 = no)	.67 (.47)	.13 (.37)	.85 (.36)	.07 (.25)
Employment	Current full-time or part-time employment (1 = yes, 0 = no)	.51 (.50)	.63 (.48)	.32 (.47)	.74 (.44)
Year	Year from which observation comes	67.56 (.93)	73.64 (1.76)	69.34 (1.04)	75.68 (1.79)
City size	Coded in categories ranging from 1 = rural, to 8 = urbanized areas of 3 million or more	3.14 (2.38)	3.08 (2.60)	3.13 (2.31)	3.82 (2.42)
South	1 = south, 0 = other	.41 (.49)	.31 (.46)	.41 (.49)	.40 (.49)
Away	Living outside parental home (1 = yes, 0 = no)	.08 (.27)	.52 (.50)	.01 (.11)	.51 (.50)
For females only:					
Kids	Presence of own child (1 = yes, 0 = no)	—[b]	—[b]	.04 (.19)	.21 (.41)
Plans for work at 35	1 = yes, 0 = no	—[b]	—[b]	.46 (.50)	.68 (.47)
For males only:					
Military	Whether currently on active military duty (1 = yes, 0 = no)	.03 (.16)	.15 (.36)	—[b]	—[b]

[a]Age at end of the interval.
[b]This variable not available for this sex.

appropriate for analysis of dichotomous dependent variables (Goodman, 1976). To permit comparison of effects of independent variables across equations, we transformed the logit coefficients to yield measures analogous to unstandardized OLS regression coefficients (Hanushek and Jackson, 1977). The transformed logit coefficients reflect the estimated effect of a unit change in the independent variable on the probability of marriage during a given year, evaluated at the sample means. Our comparisons showed that the results from logit closely approximated those produced by OLS regressions on the same samples and models.

The analysis of the probability of marriage begins with those males (females) never married at age 18 (17) and includes all ages through 29. Since by this age most people have married, the number of those never married becomes too small to support analysis, and results become theoretically less interesting.

The major goal of this paper is to test the hypotheses presented earlier about the effects of parental family structure on the transition to marriage for young adult males and females. For this reason, and because of the inherent complexity of our approach, the large number of samples, and the overwhelming number of effects estimated, we present a relatively simple, straightforward analysis. We examine direct effects only; we do not deal with problems of causal direction; we do not explore nonlinearities or nonadditivities (Stolzenberg, 1979).

RESULTS

Childhood Family Structure

Our first concern is to establish the extent to which childhood family experience influences the likelihood of entry into marriage over the critical ages during which first marriages occur. Table 2 presents for males and females the effect of being in a family with two natural parents at age 14, and of two common alternatives to that experience: being in a female-headed household at age 14, and being in a reconstituted family (Cherlin and McCarthy, 1983). Since our measures of family structure and race are all dichotomies, the coefficients in Tables 2 and 3 give the increase (positive sign) or decrease (negative sign) in marriage chances at a particular age for those with a certain parental family structure or race.

TABLE 2. LOGIT COEFFICIENTS FOR EFFECTS OF FAMILY STRUCTURE AT AGE 14 ON WED, WITH CONTROLS FOR OTHER CHARACTERISTICS

	Model 1	Model 2	
Age at End of Interval	Intact	Female-headed Household	Other Nonintact
Males			
18	-.008	.019*	-.011
19	.022*	-.017	-.025
20	-.000	-.002	.001
21	.024	-.026	-.040
22	.032	-.056*	-.015
23	.027	-.055	.003
24	.019	.003	-.049
25	-.018	-.013	.047
26	.047	-.031	-.070
27	-.050	.069*	.024
28	.060	-.343	-.031
29	-.015	-.017	.044
Females			
17	.001	-.008	.008
18	-.034*	.029	.037*
19	.002	-.002	-.005
20	.070**	-.053*	-.089**
21	.021	-.047	.005
22	.008	-.039	.025
23	.011	.004	-.046
24	.012	-.015	-.006
25	.026	-.096	.030
26	.040	-.023	-.101
27	.057	-.061	-.051
28	-.035	.075	-.044
29	.214*	-.245*	-.172

Note: Transformed log-odds coefficients, controlling for Parent's Education, Educational Attainment, Black, School Enrollment, Employment, Region of Residence, Size of Place, Year and Living Away from Home for both sexes. In addition, models for females include Presence of Own Children and Plans for Work at Age 35; models for males include Active Military Duty.

*$.05 < p < .10$.

**$p < .05$.

Overall, it is clear from the predominantly positive effects of Intact on marriage probabilities that children who were not raised in a family with two natural parents still present when they entered their teens were generally somewhat less likely to marry at a given age (Model 1). The results are not terribly strong or fully consistent (although they are as important as most of the other predictors of entry into marriage found in this and related research). For both sexes the pattern of the coefficients suggests that effects are concentrated primarily at the ages when marriage rates are highest—from 20 to 22 for females and from 21 to 24 for males—although similar, if weaker effects characterize both groups at older ages. These results suggest that 3 to 6 percentage point reductions in the likelihood of marriage occur at each age associated with childhood family disruption; and there is little indication that these deficits are made up by increases in the probability of marriage at later ages, suggesting that the overall impact is one of increased nonmarriage.

The general results associated with the single measure appear to be due fairly equally to the effects observed when we consider separately female-headed households and other family forms (Model 2). The coefficients for these two measures tend to have the same size and be of similar magnitude, although this is not always the case. It would seem, then, that the experience of parental disruption deters marriage, whether or not a remarriage occurs to "rebalance" the household. Since those with remarried parents probably experienced divorce at an earlier age than those whose parents have not remarried, this finding suggests that the age at which disruption and/or remarriage occurs is relatively unimportant; but data that can examine these relationships in more detail are needed to confirm the extent to which this is true.

Furthermore, other than the appropriate differences in timing, there are no indications of strong sex differences in the relationship between family structure and later marriage. Whether this relationship is causal, and whether or not the fact of father absence or the arrival of a stepfather influences sons differently from daughters in other ways, the outcome in this analysis is the same—decreased entry into marriage. Since the effects of female-headed family and other nonintact family structure were so similar, we have combined all family forms other than the biological nuclear family into a single specification. This variable, Intact, takes the value 0 for all family forms but two-parent families for whom it takes the value 1.

The Black Family

The results of the analysis to this point, while not powerful, are of a magnitude sufficient to have a considerable effect on racial differences in marriage rates if childhood family structure is not controlled. In general, black males and females have been shown to be about 10% less likely than others to marry, on average (Michael and Tuma, 1983; Waite and Spitze, 1981). Table 3 tests this possibility directly by showing, separately for each sex, the magnitude of the race coefficient with and without childhood family structure controlled (Models 1 and 2, respectively). If a major portion of the previously observed racial effect operates through the differing childhood experiences of the two racial groups, we would expect the race effect to be reduced by controlling for differences in childhood family patterns.

A comparison of columns 2 and 3 indicates that, in fact, virtually none of the race effect operates through childhood family structure, at least as measured in these data. Without such a control, young black men are up to 7% less likely to marry, and young black women are up to 16% less likely to marry than white men and women of the same age who are comparable in other ways. Controlling for childhood family experiences changes the probabilities very little, even at ages when it has a significant effect. There is some suggestion among males that a portion of the negative race coefficients can be attributed to family disruption. For example, controlling family stability reduces the race coefficient among males

TABLE 3. LOGIT COEFFICIENTS FOR EFFECTS OF RACE AND FAMILY STRUCTURE AT AGE 14 ON WED, WITH CONTROLS FOR OTHER CHARACTERISTICS

	Model 1		Model 2	Model 3			
	Intact (1)	Black (2)	Black (3)	Black (4)	Black Intact (5)	Intact (6)	Intact Effect (for Blacks) (5 + 6) (7)
Males							
18	-.008	-.016*	-.015	-.028*	.020	-.016	.004
19	.022*	-.012	-.017	-.029	.023	.012	.035
20	-.000	-.060**	-.060**	-.078**	-.027	-.010	-.037
21	.024	-.068**	-.073**	-.097**	.042	.008	.050
22	.032	-.051**	-.060**	-.073*	.032	.017	.049
23	.027	-.047*	-.055**	-.105**	.083	-.010	.073
24	.019	-.063*	-.068**	-.056	-.010	.023	.013
25	-.018	-.022	-.016	-.047	.039	-.036	.003
26	.047	.004	-.006	-.016	.024	.034	.058
27	-.050	-.026	-.006	-.031	.010	-.054	-.044
28	.060	-.007	-.030	.352	-.408	.384	-.024
29	-.015	-.011	-.006	.009	-.035	-.001	-.036
Females							
17	-.001	-.030*	-.030**	.008	-.066**	.025	-.041
18	-.034*	-.196**	-.092**	-.132**	.060	-.052**	.008
19	.002	-.095**	-.012**	-.105**	.015	-.003	.012
20	.070**	-.074**	-.081**	-.102**	.037	.056**	.093
21	.021	-.074**	-.076**	-.118**	.060	.003	.063
22	.008	-.157**	-.158**	-.014	-.213**	.086*	-.127
23	.011	-.034	-.036	.005	-.054	.036	-.018
24	.012	-.007	-.007	-.054	.060	-.015	.045
25	.026	.036	.035	.058	-.026	.044	.018
26	.040	.047	.045	.191	-.171	.155	-.016
27	.057	-.102*	-.105*	-.070	-.041	.080	.039
28	-.035	-.109	-.098	-.018	-.129	.039	-.909
29	.214**	-.106	-.164*	.545	-.682	.820	.138

Note: Transformed log-odds coefficients, controlling for Parent's Education, Educational Attainment, School Enrollment, Employment, Region of Residence, Size of Place, Year and Living Away from Home for both sexes. In addition, models for females include Presence of Own Children and Plans for Work at Age 35; models for males include Active Military Duty.

*.05 < p < .10.

**p < .05.

at age 21 from −.073 to −.068 and at age 22 from −.060 to −.051; similar reductions occur at several other ages. In contrast, in the equation for females, controlling family disruption rarely makes a difference.

This suggests that the effect of family disruption is not the same for the two racial groups, so that controlling for the effect of family history does not achieve the appropriate result. To explore this possibility, we constructed Model 3, which allows different effects of childhood family structure for blacks and whites. This specification means that the race coefficient (column 4) measures differences between blacks and whites who experienced family disruption; the Intact variable (column 6) measures the effect for whites from intact families relative to those from ever-disrupted families. A family-structure effect for blacks is produced by the combined effects of the general Intact coefficient (column 6) and the specific Black Intact coefficient (column 5). This has been computed and presented in column 7.

Thus, more detailed analysis shows that being raised in an intact family primarily affects black males and white females. The Intact coefficient for black males is consistently positive at all but the oldest ages, often with effects of considerable magnitude. For white males, in contrast, coming from an intact family has no consistent impact. The opposite, however, is the case for females. The only exception is the youngest ages, when there is a suggestion that being raised in an intact family protects against very early marriage for women of both races, deterring the high-risk marriages that might "continue the cycle" (Card, 1981). Beyond this exception the effect of Intact which had appeared in the earlier models for females results primarily from its impact on whites. Among black females the coefficients for Intact are unstable and inconsistent.

Determinants of the Transition to Marriage

Although this analysis focuses on racial difference in the effect of childhood family structure on later marriage, our results come from a much more general model of the marriage process. To give a context to the detailed analysis, we summarize briefly here the results for other variables in our model. Table 1 shows the variables included in our models of first marriage and presents definitions of all of them.

The results of the full models indicate that many of the factors that have a strong influence on the marriage probabilities of one sex have a similar influence on the other, although this is not always the case. The most similar effects appear for variables that define marriage markets and, thus, affect everyone within them. Race and size of community constitute two such variables on which homogamy is very high. This paper focuses primarily on the reduced probability of marriage for black men and women. The size of community effects indicate that the larger the community, the lower the marriage rate for both sexes, suggesting that social and economic alternatives to marriage rise with increases in community size and concomitant increases in structural complexity and social heterogeneity (Preston and Richards, 1975; Hannan et al., 1977). Growing up in the South increases the likelihood of marriage at relatively early ages, especially for males.

Measures of family background also had generally similar effects for both sexes. Coming from an intact family generally raises (see the analysis above), and parental educational level consistently lowers the likelihood of marriage at a given age for both sexes. For each sex full-time enrollment in school is associated with decreased likelihood of marriage up to the middle 20s, followed by an increase. The student role competes effectively with the spouse role during these ages. The only variable available for both sexes that consistently differed in its effects on the marriage probabilities of males and females is full-time employment. For females few effects appear. Net of the effects of school enrollment and aspirations for long-term labor-force attachment, women's employment has little implication for their marriage preferences and/or marriageability. For males, however, holding a full-time job is the most important single factor in the likelihood that they will marry in the next year, with strong and consistently positive effects.

Finally, our model included several measures of individual characteristics unique to each sex. Expecting to work at age 35 (measured only for females) had a strong negative effect which was concentrated among the youngest and the oldest of those in this group, suggesting the importance of a nonfamily orientation in marriage decisions. The presence of own children increased the chances of marriage for never-married women but only at the oldest ages; military service had no effect on the matrimonial prospects of the young men.

DISCUSSION

We must make two points about our findings before we summarize and discuss them. First, our hypotheses deal with the "effects" of parental family structure on the transition to first marriage, but we cannot be sure that our findings result solely from these effects. Other unmeasured factors associated with parental family structure (such as personality traits, social skills, or various attitudes and tastes) might account for some of the relationship we observe. Second, many of the individual effects in our models are quite small, even when they achieve statistical significance. However, each of the coefficients reported for a particular age comes from a separate equation which predicts the probability of marriage within a single year; so a number of small effects may cumulate over the ages we consider to equal a sizable impact. For example, the lower probabilities of first marriage at almost all ages for blacks compared with whites means that by the late 20s a much smaller proportion of blacks than whites have entered legal marriage.

Our results suggest that the family patterns children experience when they are growing up continue to have an impact on their own patterns of family formation. However, this effect is a highly qualified one, in that it seems to affect sons and daughters and blacks and whites somewhat differently. The sex-race interaction is a fascinating one and suggests that Moynihan's concerns, while highly partial, were often insightful. His primary concern was that family instability led to "matriarchy" and, thus, to lowered occupational possibilities for black men. What these data show, at least on the much narrower issue of marriage entry, is that a history of family instability does affect black males more fundamentally than black females, in that it reduces the overall chances of marriage. However, Moynihan was only analyzing this issue for blacks. Our results show that the same pattern also characterizes white women relative to white men—experience in a disrupted family lowers the probability of marriage at most ages for females but has much less effect for males.

We can only speculate on these race/sex differences in the effect of growing up in a disrupted family on young adults' own marriage chances. Black men who grew up in a nonintact family lack models for the father role, especially since stable two-parent families are less common in the black than in the white population (although recent rises in divorce rates are changing this picture). White males without their own fathers in the household may acquire role models more often than blacks from their extended families, neighbors, grandparents, and the like. Some of the differences between effects of parental family structure for black males and females might result from misreporting of legal wedlock, with males omitting and females inventing legal marriage. This revision of their marital history might be especially common among parents—for males if they do not support their children, and for females who think a marital birth more socially acceptable than one out of wedlock.

The lower effect of parental marital disruption for white males than for white females may result from the large relative gains from marriage that this group experiences. White men tend to be substantially more satisfied with their marriages than white women (Ross et al., 1983); and men tend to receive wage bonuses from marriage, whereas women—especially white women—receive significantly lower wages if ever married than if single (Treiman and Roos, 1983; Hudis, 1977). White women whose parents' marriages ended may find marriage less attractive than do women with happier family experiences.

In addition, the social consequences of nonmarriage appear to differ substantially for men and for women. For men singleness is associated with reduced economic success, both prior to marriage and after disruption (Mortimer et al., 1982). For women, on the other hand, nonmarriage seems to be a response to greater access to resources (Preston and Richards, 1975), although White (1981) points out that this relationship does not hold for black females, who do not respond to economic opportunities to the extent white females do. The trend in family instability, then, may have a more problematic effect for black men than for any other race-sex group if marriage improves economic success for them, an effect that we suggest researchers should examine more systematically. At this point, however, differences detected by our model in the pattern of effects linking childhood family disruption to marriage are too weak for this to be more than speculation. These differences in the effects of childhood family structure can do little to account for black-white differences in marriage probabilities or for the rapid declines over time. These two phenomena still need to be explained.

Black-white differences are particularly sharp, with coefficients indicating that, on average, black men and women are more than 10% less likely to marry than are whites. Since the differences are cumulative rather than simply reflecting differences in timing, among those aged 25–29 fully 44% of young black men and women were still never married in 1982, compared with 27% of comparable whites (U.S. Bureau of the Census, 1983). Although younger cohorts of white men and women appear to be following the pattern of increasing nonmarriage that was first analyzed for blacks, much larger increases would have to occur for that gap to close.

Nevertheless, increases in family disruption may have contributed far more both to the declines in marriage and to the racial differences observed than this analysis has been able to demonstrate. One problem is that of measurement. The indicators of family disruption available in the data set do not enable us to distinguish either the timing of disruption in the child's life or the duration of the various family forms that children experienced. Data with more complete parental marital histories, or even histories of the child's living arrangements, are needed to measure variation on these dimensions more precisely. Analyses using mother's marital history to infer children's living arrangements find large differences in experience with family instability between black and white children born to ever-married mothers (Bumpass and Rindfuss, 1979).

A further problem, one which is less tractable, is that a child's personal experience with disruption may not fully indicate children's actual experiences, which are drawn as well from the communities in which they live. Seen in this way, the continuing separation between the white and black communities (U.S. Riot Commission, 1968; Farley, 1977; Bianchi, 1980), plus the earlier increase in family instability experienced by blacks, means that black children have earlier and greater experience with family disruption, even if they did not experience it in their own families. A data set that also includes information on family and community networks is needed to test this possibility.

Our results, although tentative, show children's experiences in their families of orientation to have implications for their later entry into marriage. This conclusion, in conjunction with others' findings that these experiences affect later marital dissolution, suggests that we should look more broadly at the impact of childhood experiences on other aspects of adult lives.

FOOTNOTE

1. The NLS Young Women were interviewed annually from 1968 through 1973 and again in 1975, 1977, 1978, and 1980; the Young Men were interviewed annually from 1966 through 1971 and again in 1973, 1975, 1976, 1978, and 1980.

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