Parental Divorce in Childhood and Demographic Outcomes in Young Adulthood*

Andrew J. Cherlin

Johns Hopkins University

Kathleen E. Kiernan

London School of Economics and Political Science

P. Lindsay Chase-Lansdale

Harris Graduate School of Public Policy Studies

University of Chicago

*This study was supported primarily by NICHD Grant HD25936. We also wish to acknowledge the support provided to Kiernan by the Economic and Social Research Council of the United Kingdom, Grant R000 23 2161, and NICHD Population Center Grant HD06268 to Johns Hopkins University. In addition, we wish to thank Michele Trieb for programming assistance.

Abstract

We investigated the long-term effects of parental divorce in childhood on demographic outcomes in young adulthood, using a British longitudinal national survey of children. Our analyses control for predisruption characteristics of the child and the family, including emotional problems, cognitive achievement, and socioeconomic status. The results show that by age 23, those whose parents divorced were more likely to leave home because of friction, to cohabit, and to have a child outside marriage than were those whose parents did not divorce. Young adults whose parents divorced, however, were no more or less likely to marry or to have a child in a marriage. Moreover, even in the divorced group, the great majority did not leave home because of friction or have a child outside marriage.

Accumulating evidence across an array of developed countries shows that children who experience parental divorce differ from those who do not with respect to leaving home, union formation, and childbearing behavior. Those from divorced families, and particularly from stepfamilies, are more likely than children reared by both biological parents to leave home at a young age and to do so for negative reasons, such as conflict and friction (see Goldscheider and Goldscheider 1989, 1993 on the United States; Kiernan 1992 on Great Britain; Mitchell, Wister, and Burch 1989 on Canada; Young 1987 on Australia). Young persons who have experienced parental divorce are more likely than their peers to cohabit (Ghilagaber 1993 on Sweden; Kiernan 1992 on Great Britain; Liefbroer, Corijn, and Gierveld, 1993 on the Netherlands; Thornton 1991; Furstenberg and Teitler 1994 on the United States). The studies by Goldscheider and Goldscheider (1989, 1993), Kiernan (1992), and Thornton (1991) also reveal that when parents remarry their daughters are more prone to enter marriage at younger ages. Young women from families disrupted by divorce are more likely than those without disruption to bear children in their teens and to conceive and bear children outside marriage (Kiernan 1992; McLanahan and Bumpass 1988).

Moreover, a growing body of research also indicates that divorce during childhood has long-term consequences for adult life, particularly with respect to socioeconomic attainment (Manski et al. 1992; McLanahan and Sandefur 1994; Wadsworth et al. 1990) and emotional well-being (Amato and Keith 1991; Aro and Palosaari 1992; Chase-Lansdale, Cherlin, and Kiernan forthcoming; McLeod 1991). These latter findings generally persist after controlling for socioeconomic origins and educational attainment.

Yet recent research on the effects of divorce on children suggests that certain factors begin to influence children long before their parents separate and that these effects that continue through the divorce itself and far beyond (Kiernan and Chase-Lansdale 1993). Thus some of the reported effects of marital disruption on demographic outcomes could be manifestations of predisruption problems or characteristics of the child or the family. Studies based on national longitudinal survey data indicate that the behavior of children whose parents will separate years later is observably different from that of children whose parents will never separate (Cherlin et al. 1991; Elliott and Richards 1991). For example, the study by Cherlin et al. reported that children whose parents were still married but later would divorce were showing more behavior problems and doing less well in school than children whose parents would remain married. These predisruption effects were stronger for boys than for girls.

Prospective, longitudinal studies that follow children from childhood into adulthood are rare. The National Child Development Study (NCDS) of Great Britain, used in this study, has charted the lives of children born in 1958. The mothers of these children were interviewed originally around the time of the birth, and follow-up interviews with the mothers and with the children's teachers occurred at ages 7, 11, and 16. The cohort members themselves were interviewed in 1981, when they were 23. Thus the prospective design of NCDS allows us to examine the extent to which childhood and family characteristics before separation affect the long term impact of divorce on demographic behavior in young adulthood. Controlling for predisruption characteristics, we also will examine in some detail the magnitude of the effect of divorce in two senses: relatively (e.g., the extent to which a parental divorce increases the odds of having a child outside marriage) and absolutely (e.g., the probability that a person whose parents divorce will have a child outside marriage). We also examine the long-term effects of a child's degree of emotional problems at age 7 on his or her demographic outcomes in young adulthood. In addition, we briefly discuss two extensions of these analyses: cohabitation and nonmarital childbearing considered as joint outcomes, and the effects of the timing of the parental divorce.

We believe that lessons from the study of Great Britain can be drawn cautiously for the United States. Demographic trends in the United States over the past few decades are more similar to trends in Great Britain than to any other European country (Cherlin and Furstenberg 1988; Kiernan 1988). As in the United States, the divorce rate in Great Britain increased greatly during the 1960s and 1970s. The divorce rate for England and Wales was 2.1 per 1,000 married women age 16 and older in 1961, 5.9 per 1,000 in 1971, and 11.9 per 1,000 in 1981 (Office of Population Censuses and Surveys 1990). Although these figures are lower than in the United States, where the rates in corresponding years were 9.6, 15.8, and 22.6 for 1,000 married women age 15 and older (U.S. National Center for Health Statistics 1978, 1984), they represent one of the highest rates in Europe. Of course, there may be cultural or social structural differences between the family lives of children in Great Britain and in the United States. Moreover, the 1958 birth cohort experienced less parental divorce than did British (or American) children born a decade later.

A previous article (Kiernan 1992) also used the NCDS to examine the effects of parental divorce on demographic outcomes in young adulthood. That article, however, did not examine whether predisruption childhood and family characteristics account for later demographic outcomes. Also, it did not use the multiple indicators of children's emotional well-being that were collected during the study.

DATA AND METHODS

The National Child Development Study

The NCDS is a longitudinal study of children who were born in Great Britain in the first week of March 1958. It was originally designed to examine the social and obstetric factors associated with stillbirth and death in early infancy. Interviews were conducted with a total of 17,414 mothers, representing 98% of all births that week

(Shepherd 1985). The first three follow-up interviews were conducted when the children were ages 7 (N = 15,468), 11 (N= 15,503), and 16 years (N = 14,761). These three ages represented important turning points in the children's educational experience in the British school system. When the cohort was 23 years old, a survey research organization traced and interviewed 12,537 of the members. The number of years between points of data collection contributed to sample attrition from this data set.[1]

At the first follow-up interview, conducted when the children were 7, the focus of the study was broadened considerably. Detailed measures of the children's personalities were obtained from interviews with each child's mother and teacher (Rutter, Tizard, and Whitmore 1970). Cognitive achievement was measured by administering standardized reading and mathematics tests. Interviewers were asked to note whether the family had required the services of child welfare agencies or professionals. The age-23 interview included a life history calendar yielding the dates of important life transitions, such as leaving home for the first time (and reason for leaving home), entry into first cohabiting union, first marriage, and parenthood.

For this article we restricted our focus to the 10,353 cohort members whose parents were in intact marriages from the child's birth until the age-7 interview and who were interviewed at age 23. This restriction allowed us to obtain baseline measures of child and family characteristics for a group of children from nondisrupted families, some of whom later experienced parental marital disruption and some of whom did not. It also introduced issues of possible sample selection bias, which we address below.

We developed outcome variables from the life transition data in the age 23 interviews. Then we examined the consequences of parental marital disruption (divorce, separation, or death) for age-23 outcomes, controlling for predisruption, age-7 characteristics of the cohort member and family.

Whether a parental divorce, separation, or death occurred can be ascertained from the parent interviews at the age-7 wave and again at the two intervening waves at ages 11 and 16. These interviews with parents included information on the child's relationship to the mother or father figure who lived in the household at the time and, where applicable, on the reasons why the child was not living with his or her biological mother or father. Even so, one can learn only *whether* a disruption occurred in the interval between the interviews (i.e., between birth and 7, 7 and 11, or 11 and 16) but not when during the interval. Of the 10,353 cohort members who still were living with two biological parents at age 7 and who were interviewed at age 23, we could determine whether or not the parents divorced or separated between age 7 and age 16 for 7,966 (77%). Of the remainder, 96% had been lost temporarily through attrition at the age-11 or age-16 interviews, but were found and reinterviewed at age 23.[2] At the age-11 and age-16 interviews, information also was obtained on whether a stepparent was present in the home. However, we could not observe stepparents who may have entered the household after an interview (for example, when the cohort member was 12) and left before the next interview (for example, at age 15). Also, restricting the analysis to children who experienced family disruption after age 7 excluded a significant proportion of the cohort members who had experienced a stepfamily.

Assessments of Children at Age 7

The age-7 interview provided a great deal of information about the child from several sources: parents, teachers, and the interviewers. The latter were Local Authority Health Visitors, trained nurses employed by the local government who normally saw every family before and after the birth of a child and frequently conducted follow-up visits. We discuss these briefly; further detail can be found in Chase-Lansdale et al. (forthcoming).

Emotional indicators. The emotional indicators include separate ratings of behavior problems by parents and by teachers, and information from the Health Visitors on use of child-oriented social services. Parents were asked most of the items from the Rutter Home Behaviour Scale (Rutter et al. 1970). The scale was designed to identify two broad bands of behavior problems in children: externalizing disorders, in which the child exhibits undercontrolled behavior such as aggression or disobedience, and internalizing disorders, in which the child exhibits overcontrolled behavior such as anxiety or depression. We constructed an 18-item summated scale which had an alpha reliability of .72[3] Teachers filled out the Bristol Social Adjustment Guide (Stott 1969), consisting of 146 "items of behaviour."[4] The responses were combined by expert raters into 12 syndrome scores (e.g., hostility towards children, restlessness, depression). From the syndrome scores we constructed a scale that had an alpha reliability of .68.

In addition, we constructed a composite measure that ranged from 0 to 2 (with 2 coded as "2 or more"), based un the Health Visitor's report (from a checklist) regarding whether the family had been in contact with any of five social service agencies or professionals.[5]

Cognitive indicators. The children were given the Southgate Group Reading Test (Southgate 1962), which tests how well children have learned to read in school and involves both word recognition and sentence completion. Thus it probably confounds children's cognitive ability with their mastery of reading. The test was standardized against a large, normal population of British schoolchildren (Pringle 1965). We also created a mathematics score based on the number of correctly answered questions in a 10-question problem arithmetic test (e.g., "A man has eight books and he lost three of them. How many books did he have left?"). This test was developed specifically for the NCDS (Pringle, Butler, and Davie 1966). In addition, a summated scale was created from the teachers' responses to five school performance items: "oral ability," "awareness of the world around him," "reading," "creativity," and "number work." The alpha reliability was .89.

Socioeconomic indicators. The socioeconomic indicators were limited; for example, no information was collected on family income. Nevertheless, we were able to construct six indicators. The first five were binary variables: 1) whether the father's occupation was nonmanual (versus manual); 2) whether the family owned its dwelling unit (versus rented); 3) whether the father had stayed in school past the minimum school-leaving age; 4) whether the mother had stayed in school past the minimum school-leaving age; and 5) whether the Health Visitor had noted economic difficulties by marking financial difficulties, unemployment, or housing difficulties on the checklist. The sixth variable was the mean number of persons per room in the dwelling unit.

Confirmatory factor analysis. We then subjected these emotional, cognitive, and socioeconomic indicators to a confirmatory factor analysis, estimated separately for girls and for boys.[6] Figure 1 displays the standardized results for girls; the results for boys were very similar. The best-fitting model consistent with our theoretical expectations had four latent constructs: one was associated with the emotional indicators, one with the school-achievement indicators, and two with the socioeconomic indicators.[7] Adjusted goodness-of-fit indices were .96 for boys and .95 for girls. No other plausible models improved the fit.

The first of the two socioeconomic constructs was composed of three indicators: owned or rented their dwelling unit, economic difficulties, and number of persons per room. In interpretation, this factor represents the current financial status of the child's family. The second socioeconomic construct consisted of the remaining indicators: father's occupation, father stayed in school, and mother stayed in school. This construct appears to represent the social class background of the child's family rather than the current financial situation.

We then used the weights from the confirmatory factor analyses then were used to create a predicted score on each of the four latent constructs for each cohort member. A higher score indicates greater emotional problems, greater school achievement, higher economic status, and higher social class background. Table 1 presents the means and standard deviations of these variables.

Marital Disruption

Our other independent variables consist of two binary indicators of marital disruption: first, whether the child's parents separated or divorced between the age 7 and the age-16 interview (175 boys and 208 girls in our age-7 subsample experienced these events and were interviewed at age 23); and, second, whether a parent died during this period (182 boys and 185 girls). These variables are coded 1 if the disruption occurred and 0 otherwise. Henceforth we use the term divorce to denote both divorce and marital separation.

The number of parental deaths is comparable to the number of parent divorces for several reasons. First, as stated earlier, we excluded children who experienced parental divorce or death before age 7. More divorces (245) than deaths (129) were excluded because divorces tend to occur earlier in marriages, whereas deaths tend to occur when parents are older. Second, divorce rates in Great Britain, as noted, were lower than in the United States, although they were rising and were high by European standards. Third, the age pattern of childbearing was older in Great Britain than in the United States, this created a higher risk of experiencing parental death. The mean age at birth in Great Britain in 1958 was 27.9 (Office of Population Censuses and Surveys 1990), whereas in the United States it was approximately 25.5(Rindfuss and Sweet 1977).[8]

Outcome Variables and Statistical Models

Table 2 shows the outcome variables obtained from the age-23 questionnaire, their definitions, and the statistical models that were used with them. All analyses were performed separately for women and for men because of the different patterns of life course transitions, such as the timing of union formation and of parenthood, for the two genders in late adolescence and early adulthood. As for the first event listed, leaving home, 79% of the female cohort members and 61% of the male cohort members reported that they had done so. The age-23 interview presented the cohort members with a series of reasons why they might have left home, including "wanted to leave because of friction at home" and "was asked to leave because of friction at home." Preliminary analyses showed that persons in our subsample whose parents' marriages had been disrupted were most likely to have responded that they had left because of friction. Therefore we created a three-category outcome variable: (0) respondent had not yet left home, (1) respondent had left home because of friction (either "wanted to leave" or "was asked to leave"), and (2) respondent had left home for other reasons (e.g., "to get married," "to take up a job"). We then specified a multinomial logistic model of responses to this three-category outcome was then specified. It can be written as

$$\frac{p_1}{p_0} = e^{x\beta_1}, \qquad \frac{p_2}{p_0} = e^{x\beta_2},$$

where p_0, p_1, and p_2 are respectively the probabilities of not leaving home, leaving home because of friction, and leaving home for other reasons; x is a vector of characteristics of the cohort member and his or her family (the divorce and death binary variables and the four age-7 latent variables); and β_1 and β_2 are vectors of parameters to be estimated (Maddala 1983).[9]

All the other outcomes in Table 2 are demographic events pertaining to union formation and first birth. Because most of the events were clustered in the few years preceding the age-23 interview, the time path of occurrence was less salient than in a fuller age distribution. Consequently we confined our attention to whether or not an outcome had occurredby age 23. Therefore we estimated binomial logistic models of whether or not a cohort member had experienced the event by age 23:

$$\frac{p_1}{p_0} = e^{x\beta},$$

where p_1 is the probability of the event's occurring, p_0 is the probability of the event's not occurring, x is as above, and β is the vector of parameters to be estimated. (Nevertheless, when we repeated the analysis using timing information and Cox [1972] proportional hazard models, the results were virtually identical.)

As noted earlier, a potential sample selection problem was created by our research design, which involves analyzing only those cohort members whose parents were in intact marriages from the cohort member's birth until age 7. If unmeasured factors affect both the likelihood of falling into this subsample and the risk that one of the age-23 outcomes will occur, then our coefficient estimates of the latter risk would be biased. To correct for this possible selection bias, we used a two-step procedure (Heckman 1979; Maddala 1983) that yields an inverse Mills ratio (which we identify, following convention, as lambda) for each person in the subsample and added lambda to the right-hand side of each model for the outcomes in Table 2.

The success of the selection-bias correction procedure depends on the inclusion in the first-step probit regression of instrumental variables that are associated with the likelihood of parental divorce but are not associated with the substantive outcome (e.g., cohabitation by age 23). It is difficult to find useful instruments in applications similar to ours. We included potential instruments such as the age difference between the mother and the father of the cohort member and whether or not the cohort member's mother worked for wages during her pregnancy; we hypothesized both of these to relate to the probability of parental marital dissolution (Cherlin 1977) but not to age-23 outcomes. Still, these instruments are less than ideal; and we are aware that this two-step procedure sometimes provides estimates inferior to those generated by OLS (Stolzenberg and Relles 1990).

In Tables 3 and 4, the estimated parameters of all of the logistic models are presented in the exponentiated form e^{β}. These exponentiated parameters are odds ratios: they represent the proportional increase (if greater than 1.0) or decrease (if less than 1.0) in the odds of the event's occurring, for a one unit change in the corresponding independent variable. Based on these results, Table 5 presents the predicted probabilities that selected outcomes would occur, according to the parental marital status of the cohort member. We calculate the predicted probabilities from the estimates in Tables 3 and 4 by assigning to hypothetical individuals the mean values (separately by sex) on the four age-7 latent variables and on lambda. Then we calculated predicted probabilities under three assumptions:

No parental divorce or death between ages 7 and 16;

Parental death between 7 and 16;

Parental divorce between 7 and 16.

For the multinomial logistic model of leaving home, we obtained the predicted probabilities p_1 (left because of friction) and p_2 (left for other reasons) were obtained from the following expression, which can be derived from the two-equation model presented earlier:

$$p_i = \frac{e^{x\beta_i}}{1+e^{x\beta_1}+e^{x\beta_2}}, \ for \ i=1,2.$$

For the binomial logistic models, the predicted probability was obtained from:

$$p = \frac{e^{x\beta}}{1 + e^{x\beta}} .$$

RESULTS

Columns 1 and 2 of Tables 3 and 4 show the multinomial logistic model of leaving home; Columns 3 through 9 show the binomial logistic models of various union formation and first birth outcomes. All of these analyses include measures that attempt to control for age-7, predisruption characteristics of the cohort member and his or her family. The tables show that parental marital disruption between ages 7 and 16 affects the likelihood of some outcomes at age 23, even after age-7 characteristics are included in the model. The introduction of the age-7 characteristics into these equations mainly produced modest, nonsignificant reductions in the coefficients for parental marital disruption. Thus, insofar as these age-7 variables control for personal and family difficulties that existed before the disruption, the significant coefficients for parental marital disruption suggest that the disruption and its aftermath does affect some, although not all, life transitions pertaining to leaving home, union formation, and parenthood.

With respect to leaving home, it is evident that having experienced a parental divorce has a greater effect on leaving home because of friction (versus not leaving home) than on leaving home for other reasons (versus not leaving home). For both women and men, the odds ratios (upper left-hand cells of the tables) are approximately

5.0, indicating that the odds of leaving home because of friction versus not leaving home are five times greater for those whose parents divorced. In comparison, a parental divorce increases by about half the odds of leaving home for other reasons (getting married, taking a job) versus not leaving home. In addition, marital disruption due to the parents's death increases the odds of leaving home because of friction, although not by as much as does divorce. (The difference between the effect of divorce and of death is significant at $p < .01$.) Thus some of the effect of parental marital disruption seems to reflect the aftermath of the loss of a parent by either divorce or death. Nevertheless, loss by divorce produces by far the larger effect, especially as a result of conflict at home. This finding is consistent with the studies of homeleaving cited earlier.

In contrast, men who lost a parent because of death are somewhat *less* likely to leave home for reasons other than friction than are those who did not experience a parental marital disruption, as indicated by an odds ratio of less than 1.0. (The corresponding parameter for women also suggests a lower likelihood of leaving but is not statistically significant.) We can speculate that children may play an adultlike role as helpers to widowed parents more than to divorced parents, and that this role postpones their departures from their households.

With respect to union formation, the general pattern is as follows. Parental marital disruption increases the odds of premarital cohabitation by age 23 far more than it increases the odds of marriage by age 23. (By *premarital* we mean an event that occurred to a person who did not marry by age 23 or who married only after the event in question occurred.) For example, Column 3, Row 1 shows that for both women and men, the odds ratios for divorce are above 2.0, implying that the odds of cohabiting are more than twice as great among those whose parents divorced than among those whose parents remained married. For men, a parental death also increases the odds ratio for premarital cohabitation, although by substantially less than does a parental divorce ($p < .01$). In contrast, Columns 4 and 5 show that the odds ratios for marrying without cohabiting first, and for marrying whether or not cohabitation occurred first, are not affected significantly by marital disruption. Finally, Column 6 estimates the odds of forming any union by age 23, whether marital or cohabiting. Parental marital disruption substantially increases the odds ratio for entering into a union, but the preceding columns make clear that this increase in union formation is due mainly to an increase in premarital cohabitation rather than in marriage. These results are consistent, for instance, with Thornton's (1991) findings from the Detroit-area longitudinal study. We have demonstrated here that these findings still hold after we introduce measures that attempt to control for early, predisruption emotional and cognitive characteristics of the cohort members, which other studies did not measure.

As for first births, Column 7, Row 1 shows that parental divorce increases the odds ratio for having a premarital birth (or becoming a father premaritally). Parental divorce, however, has little effect on the odds ratio for having a birth after marrying, as shown in Column 8, Row 1. The model in the final column ignores whether a first birth occurred before or after a marriage and demonstrates that a parental divorce increases the overall odds ratio for having a birth by age 23. Yet the table shows that this difference is due mainly to the incidence of premarital births. Thus the birth outcome measures show the same pattern that we observed for the union formation outcomes: whereas events occurring outside marriage are stimulated by a parental divorce, the events of marriage itself and of births within marriage hardly are affected. Moreover, for premarital births, the effect of a parental death on women again is similar to the effect of a parental divorce, but smaller in magnitude ($p < .01$). Overall, then, cohort members whose fathers or mothers died have odds of leaving home because of friction, of premarital cohabitation (for men), and of premarital births (for women) that are generally higher than for those whose parents remained married but are lower than for those whose parents are divorced.

All of the estimated coefficients for lambda in the models are positive and sometimes are statistically significant. A positive lambda indicates that the unmeasured attributes that made people more likely to be in our subsample also made them more likely to experience the corresponding demographic event by age 23. In no case, however, does the inclusion of lambda change the other estimated coefficients in the models by more than trivial amounts.

Relative vs. Absolute Effects

So far this discussion of the effects of marital disruption has concentrated on the relative increase or decrease that parental divorce produces in the odds of experiencing a particular transition. Yet in assessing the magnitude of the effect of marital disruption, it also is important to note the percentage of young adults whose parents' marriages are disrupted but who do not experience early transitions. To explore this issue, we present in Table 5 the predicted probabilities for the occurrence of the events that were associated significantly with parental marital disruption in Tables 3 and 4: leaving home, premarital cohabitation, and premarital birth. In calculating each predicted probability, we set all variables equal to their mean values except parental marital disruption, which we varied. Effectively, then, each row of Table 5 presents the predicted probabilities for three hypothetical persons who differ in their experience of parental marital disruption but are otherwise "average." For example, the first row shows that for women, a parental death increases the probability of leaving home because of friction, relative to remaining at home, from .046 to .098; a parental divorce increases it further to .14. Similarly, the table shows that a parental divorce raises from .05 to .17 the predicted probability that a woman will have a premarital birth. Thus our data, if confirmed by other studies, suggest that, in absolute magnitude, a majority of young people whose parents divorce will not experience the potentially serious negative consequences of leaving home because of friction or having a premarital birth. Nevertheless, in relative magnitude, a minority will experience these consequences, and their numbers will cause a substantial increase in the relative size of the population whose transition to adulthood is hindered by parental divorce.

Emotional problems

Because few studies of the transition to adulthood have obtained prospective measures of the respondents' emotional problems in childhood, the effects of our latent construct measuring emotional problems at age 7 are interesting in themselves. Tables 3 and 4 reveal that emotional problems at age 7 displayed a pattern of association with demographic outcomes similar to parental marital disruption. That is, women who had more emotional problems at 7 are more likely to leave home because of friction but not for other reasons, and they are more likely to cohabit before marrying but not to marry. In addition, women who had more emotional problems at 7 are more likely to have a child before marrying and, to a lesser extent, to have a child after marrying. The effects of emotional problems generally are weaker for men than for women. Even so, the estimated parameters for the emotional problems construct cannot be compared directly with the parameters for marital disruption because disruption is measured by dummy variables and emotional problems are measured by a continuous variable.[10] To provide a clearer comparison, we can turn again to predicted probabilities.

Let us consider two hypothetical women who had "average" (mean) values on all independent variables in Table 3 except that Woman A's score for age-7 emotional problems was one standard deviation greater than Woman B's score. Specifically, suppose that Woman A scored one-half standard deviation above the mean for all women, whereas Woman B scored one-half standard deviation below the mean. Woman A would have a .092 predicted probability of leaving home because of friction, more than three times the predicted probability of .028 for Woman B. Woman A also would have about three times as great a probability of bearing a child before marrying as Woman B, .086 compared with .030. In addition, Woman A would have a .24 probability of cohabiting before marrying, compared with a .14 probability for Woman B.

Thus in relative terms, the level of a woman's emotional problems at age 7 emerges as a strong predictor of friction at home and of premarital union formation and childbearing by age 23. This effect persists even after we take into account cognitive achievement and parental socioeconomic status at 7, as well as parental marital disruption between 7 and 16. The similarity in the pattern of effects shown by emotional problems and by parental divorce suggests the possibility of some unmeasured variables, such as parental conflict before the disruption or children's temperamental difficulties or personality characteristics, that may influence both effects (see Caspi, Elder, and Bem 1987).[11]

Additional Analyses

We wish to report briefly on two additional analyses that extend our study. Further information is available from the authors on request.

Cohabitation and nonmarital childbearing as joint outcomes. Cohab-itation and having a child outside marriage are unlikely to be independent outcomes.[12] In our sample, one-third of the women who had a first birth outside marriage by age 23 were cohabiting at the time of the birth. In addition, the odds of having a child within a cohabiting union are 5.5 times greater for women whose parents divorced than for women whose parents did not divorce, when age-7 attributes are controlled. The odds of having a child outside a union are 3.1 times higher for women whose parents were divorced. (Data are not shown.) Thus the effect of parental divorce on the odds of having a child in a cohabiting union is somewhat larger than the odds of having a child outside a union, but the observed differences (i.e., between odds ratios of 5.5 and 3.1) could have arisen by chance. It would appear, at least for this sample, that women who experienced parental divorce have similar propensities to a bear a child either on their own or within a cohabiting union.

Timing of divorce. We also examined whether the timing of divorce (ages 7 to 11 versus ages 11 to 16) was related to demographic outcomes. The child's age at the time of marital disruption did not have such an effect. Thus our study does not suggest that the timing of a parental divorce in a child's life makes a difference for young adult outcomes--at least not for the demographic outcomes studied here. Yet we point out that we excluded children who experienced divorce before age 7. Some studies have reported that preschool children seem more vulnerable to divorce than older children, but studies to date have not adequately disentangled child's age at divorce from duration of time since divorce (Allison and Furstenberg 1989).[13]

SUMMARY AND DISCUSSION

In drawing lessons from this study of the British 1958 birth cohort, we must keep in mind not only the situations in which parental divorce had an effect but also the situations where it did not. For example, parental divorce did not make a person significantly less likely to marry by age 23, nor to have a child within marriage by age 23. (We caution, however, that being married and being a parent at 23 represent early transitions to family roles.) Some observers have argued that the experience of divorce in childhood may lead young adults to be anxious and uncertain about making a long-term commitment to a partner (Wallerstein and Blakeslee 1989). We find no evidence, however, that the 23-year-olds in the British cohort whose parents had divorced were hesitant to marry. Nor were married persons whose parents had divorced significantly less likely to have a child. Parental divorce, our study suggests, does not by itself undermine young adults' desire or ability to undertake long-term commitments to spouses and children.

Rather, a parental divorce seems to have stimulated a pattern of behavior characterized by early homeleaving due to conflict with parents and stepparents and early sexual activity outside marriage--leading, in this cohort, to a greater likelihood of premarital birth and cohabitation. For example, 19% of the women who left home because of friction had given birth to a child outside marriage by age 23, compared with 6% of those who left home for other reasons and 5% of those who did not leave home. (Data are not shown.) Parental death, in contrast, has a substantially weaker effect on leaving home because of friction, on premarital cohabitation, and on premarital birth than does parental divorce. This finding suggests that losing parents for very different reasons leads to different processes (Hetherington 1972).

We would hypothesize that early sexual intercourse may be a central part of the mechanism by which parental divorce produces its particular effects. We think it likely that the age at first sexual intercourse (on which,unfortunately, we had no information) for the British cohort members who had experienced divorce was younger than for those whose parents remained married. Other studies have shown that parental marital disruption indeed is associated with earlier onset of sexual activity (Furstenberg and Teitler 1994; Newcomer and Udry 1987). An important question for future research is whether a pathway can be traced from parental marital disruption to earlier onset of sexual activity and then to early union formation and childbearing.

Early sexual activity may be a key to the effects of divorce on the transition to adulthood for two reasons. First, and more directly, the obvious sexual activity of divorced parents may stimulate nonmarital sexual activity in their children. In other words, the parents' own nonmarital sexual activity may make it difficult for them to discourage their children from following their actions rather than their admonitions (Hetherington and Clingempeel 1992). Second, adolescents' early sexual activity could constitute rebellious or acting-out behavior begun in reaction to an unwelcome parental divorce or to the introduction of an unwelcome new stepparent. That is, early sexual activity among adolescents from families of divorce may be a form of externalizing anger or disappointment at the parent--or it could result from the absence of an appropriate second adult. Such activity obviously places the children at greater risk of bearing or fathering a child outside marriage or of forming a cohabiting union.

Further, the British cohort members were between ages 18 and 23 from 1976 to 1981, a time when cohabitation was beginning to increase noticeably in Britain (Kiernan and Estaugh 1993). Consequently, early cohabitation may have been more of a rebellious, unconventional act in that period than it would be today. Similarly, the children in the Thornton (1991) study were 18 to 23 from 1980 to 1985. It is possible that the connection between parental divorce and cohabitation will become progressively weaker as cohabitation becomes increasingly common and morally accepted. Yet a 1982 survey in Sweden, where cohabitation already was widespread, showed the same pattern (Ghilagaber 1993). Future studies are needed to ascertain whether the relationship between parental divorce and patterns of union formation will persist.

As noted previously, several studies suggest that the impact of divorce on children may begin long before their parents separate, probably as a result of conflict, family dysfunction, and economic difficulties before divorce (Baydar 1988; Block, Block, and Gjerde 1986; Cherlin et al. 1991; Elliott and Richards 1991). Studies such as these, which control for preexisting problems, suggest that difficulties in the predivorce family may be as important for the children involved as those following the divorce. The article by Cherlin et al. was based in part on comparisons of data from the age-7 and age-11 waves of the NCDS. Among boys in the NCDS who were living with two biological parents at age 7, the apparent effect of separation or divorce by age 11 was reduced sharply by considering behavior problems, achievement levels, and family difficulties that already were present at age 7. In the present study, however, the introduction of similar measures of child and family characteristics at age 7, before the divorce, did not significantly weaken the long-term effects of divorce on demographic behavior at age 23.

Possibly the consequences of divorce are different during middle childhood and early adolescence than during young adulthood. The disruption and its aftermath per se may have less effect on 11-year-olds. For example, if we are correct in our hypothesis that early sexual activity is central to the pattern of persistent effects found at 23, then children still in the middle-school period before puberty may show fewer current negative effects. Rather, the negative consequences may become manifest during adolescence, as overall levels of sexual activity rise in the cohort. Even so, we point out, both the Cherlin et al. study and this study suggest that whatever increased risks are caused by a parental divorce, most children whose parents divorce will not show serious negative consequences at age 11 or at age 23. (Also see Chase-Lansdale et al. forthcoming.)

Still, our study has limitations. We could not control directly for several important factors such as family income, parenting practices, and marital conflict. Also, substantial attrition occurred in the cohort from age 7 to 23; although our investigations suggest that attrition did not bias our results, this is not a certainty.

Moreover, we chose to delete cases in which a divorce had occurred before the child's 7th birthday in order to obtain predisruption measures. This deletion could be problematic despite our efforts to correct for sample selection bias. In addition, the age-23 interview did not collect information on whether a parental divorce took place between ages 16 and 23; consequently we could not study these later disruptions. Finally, the childhood of the NCDS cohort of 1958 occurred just as divorce rates began to rise sharply. As a result, the number of divorces is modest in comparison with the size of the full sample, and the number of remarriages is reduced further. The inability to observe stepfamilies of short duration, in combination with our deletion of cases of divorce before age 7, further reduced the number of identifiable remarriages, to the point where multivariate analyses of the role of remarriage in demographic outcomes were not feasible. The consequences of divorce could be different for children growing up in the 1990s, when divorce is even more common, than for children who were born in 1958.

Nevertheless, we believe that the findings presented here expand our understanding of the effects of divorce on young adult transitions. Given the prospective nature of the NCDS, we were able to not only examine the potential role of predisruption factors but also to draw on characteristics that are difficult, if not impossible, to collect retrospectively, such as emotional and cognitive measures. We drew upon multiple reports from parents, teachers, and Health Visitors. Our findings show a pattern of early homeleaving, union formation, and nonmarital childbearing that is associated with a parental divorce. Yet these actions, at least in the British data, have not led to a rejection of marriage at young ages for this cohort. It remains to be seen whether the marriages formed by children of divorce will prove as durable and as satisfactory to the spouses as will the marriages formed by children whose parents remained together. In 1991 the NCDS cohort members were reinterviewed at age 33. The data from this wave, released only recently, will begin to answer this question.

Notes

[1] Comparisons of the subsample interviewed at age 23 showed them to be "somewhat more middle-class and educationally successful than the original population." The magnitude of the differences, however, was quite modest. For example, 35.1% of those interviewed at age 11 reported that their fathers' occupations were "nonmanual," whereas 36% of those interviewed at age 23 had reported having fathers in nonmanual occupations at age 11. At age 11, 10.4% reported that they received free meals at school (indicating low family income), whereas 9.2% of those in the age-23 sample had reported receiving free meals at age 11 (British National Children's Bureau 1987). We believe that these modest differences do not substantially bias our analyses.

[2] We assigned the mean probability of parental divorce to the temporarily attrited cases and merged them with the cases for which we had complete information. In regression analyses not reported here but available from the authors, we included a dummy variable coded 1 if marital status information at age 16 was valid and 0 if it was missing (and if the mean had been imputed instead). The addition of this dummy variable resulted in virtually no changes in the coefficients for effects of divorce, compared with the models we present here. The dummy variable itself was statistically significant only for leaving home beause of friction among boys.

[3]We also constructed separate subscales for internalizing and externalizing disorders. The reliability of the internalizing subscale, however, was considerably lower than that of the externalizing subscale, and a single-factor solution had higher internal consistency than either subscale. The age-7 items in our final scale were temper tantrums, reluctant to go to school, bad dreams, difficulty sleeping, food fads, poor appetite, difficulty concentrating, bullied by other children, destructive, miserable or

tearful, squirmy or fidgety, continually worried, irritable, upset by new situations, twitches or other mannerisms, fights with other children, disobedient at home, and sleepwalking. The age-16 items were temper tantrums, tears on arrival at school, ever steals things, cannot settle to anything for more than a few moments, often destroys own or others' property, miserable or tearful, squirmy or fidgety, irritable, frequently fights, disobedient, very restless and has difficulty staying seated for long, often tells lies, and bullies other children.

[4] For example, classroom behavior: well-behaved / too timid to be naughty / occasionally naughty / has no life in her or him / constantly needs petty correction / very naughty, difficult to discipline / sly, will abuse trust, hard to catch.

[5]The Children's Society (private service providers), the National Society for the Prevention of Cruelty to Children, the Children's Department (government child welfare system), the Handicapped Children's Association, or a school welfare or attendance officer.

[6] We conducted the confirmatory factor analyses separately for girls and for boys to allow for different factor structures; the results were very similar, however. In preparing the correlation matrices for the confirmatory factor analyses, we replaced Pearson correlations with polychoric or polyserial correlations if one or both variables were dichotomous (Bollen 1989). We used logarithmic transformations of the parent- and teacher-rated behavior problems scales because they approximated the normal distribution more closely. We estimated the confirmatory factor analyses using Lisrel 7 (Jöreskog and Sörbom 1989).

[7] In addition, as Figure 1 shows, we allowed correlations between the error terms of indicators that were rated by the same person, such as a teacher's rating of student's behavior problems and of student's

school performance.

[8]Rindfuss and Sweet (1977) report that the mean age of fertility was 25.62 in 1955 and 25.47 in 1960; we have interpolated.

[9]The multinomial logistic model requires that disturbances associated with the two equations be uncorrelated; this is called "the independence of irrelevant alternatives" (Maddala 1983). This property, which is violated in many demographic applications of the model, implies that the odds of choosing Alternative (0) rather than Alternative (1) are not affected by whether it is possible to choose Alternative (2). The property may not hold in our case because unmeasured factors influencing the choice between leaving home because of friction and staying home may also influence the choice between leaving home for other reasons and staying home. To check our results, we attempted to estimate two separate probit models in a bivariate framework that allowed the disturbances of the two equations to be correlated. For boys, the results were very similar to the multinomial logit estimates; for girls, however, the bivariate probit algorithm would not converge.

[10]Thus a one-unit change in the parental divorce variable corresponds to examining the effects of experiencing divorce versus experiencing no disruption--a reasonable comparison. A one-unit change in the emotional problems variable, however, which has a standard deviation of about .11, corresponds to experiencing a nine-standard-deviation change in emotional problems, an unrealistic comparison.

[11]School achievement scores at age 7 also affect a similar set of age-23 outcomes. For example, if two hypothetical women differed only in that one had a cognitive achievement score of one-half standard deviation above the mean, and the other had a score of one-half standard deviation below the mean, the first would have a .23 predicted probability of cohabiting premaritally, compared with .15 for

the second.

Yet the effect of the cognitive achievement measure is hard to interpret unambiguously because it confounds a higher level of school success and intellectual aptitude, on the one hand, with earlier homeleaving in order to attend university, on the other. The former might lead to more conventional behavior, whereas the latter places a person at greater risk of cohabitation or premarital childbearing.

[12]In Britain children are increasingly being born and reared in cohabiting unions; one extramarital birth in two is registered by couples living at the same address, although younger women are less likely than older women to be in this group (Kiernan and Estaugh 1993).

[13]In addition, we examined whether the introduction of variables from the age-16 round (which included emotional, cognitive, and socioeconomic measures that were similar to the age-7 indicators) into our models in Tables 3 and 4 would mediate the effect of parental divorce and therefore would reduce the direct association between parental divorce and demographic outcomes. Chase-Lansdale et al. (forthcoming) found evidence that age-16 characteristics mediated the effect of parental divorce on mental health at age 23 in the NCDS. Across all the demographic outcomes, the addition of age-16 variables tended to produce a decline in the odds ratios for the effect of parental divorce. Yet the greater propensity of children who experienced parental divorce to leave home because of friction, to cohabit, and (for women) to have an extramarital birth remained substantial and significant. Therefore, these age-16 attributes did not mediate most of the effects of parental divorce on demographic outcomes.

REFERENCES

Allison, P.D. and F.F. Furstenberg Jr. 1989. "How Marital Dissolution Affects Children: Variations by Age and Sex." *Developmental Psychology* 25:540-49.

Amato, P. and B. Keith. 1991. "Parental Divorce and Adult Well-Being: A Meta-Analysis." *Journal of Marriage and the Family* 53:43-58.

Aro, H.M. and U.K. Palosaari. 1992. "Parental Divorce, Adolescence, and Transition to Young Adulthood: A Follow-Up Study." *American Journal of Orthopsychiatry* 62:421-29.

Baydar, N. 1988. "Effects of Parental Separation and Re-entry into Union on the Emotional Well-Being of Children." *Journal of Marriage and the Family* 50:967-81.

Block, J.H., J. Block, and P.F. Gjerde. 1986. "The Personality of Children Prior to Divorce: A Prospective Study." *Child Development* 57:827-40.

Bollen, K. 1989. *Structural Equations with Latent Variables*. New York: Wiley.

British National Children's Bureau. 1987. *The Fourth Follow-Up of the National Child Development Study: An Account of the Methodology and Summary of Early Findings*. London: Social Statistics Research Unit, City University.

Caspi, A., G.H. Elder Jr., and D.J. Bem. 1987. "Moving against the World: Life-Course Patterns of Explosive Children." *Developmental Psychology* 23:308-13.

Chase-Lansdale, P.L., A.J. Cherlin, and K.E. Kiernan. Forthcoming. "The Long-Term

Effects of Parental Divorce on the Mental Health of Young Adults: A Developmental Perspective." *Child Development.*

Cherlin, A. 1977. "The Effect of Children on Marital Dissolution." *Demography* 14:265-72.

Cherlin, A. and F.F. Furstenberg Jr. 1988. "The Changing European Family: Lessons for the American Reader." *Journal of Family Issues* 9:291-97.

Cherlin, A.J., F.F. Furstenberg Jr., P.L. Chase-Lansdale, K.E. Kiernan, P.K. Robins, D.R. Morrison, and J.O. Teitler. 1991. "Longitudinal Studies of Effects of Divorce on Children in Great Britain and the United States." *Science* 252 (June 7):1386-89.

Cox, D.R. 1972. "Regression Models and Life Tables." *Journal of the Royal Statistical Society, Series B* 26:103-10.

Elliott, M. and M.P.M. Richards. 1991. "Children and Divorce: Educational Performance and Behaviour Before and After Parental Separation." *International Journal of Law and the Family* 4:258-76.

Furstenberg, F.F. Jr., and J.O. Teitler. 1994. "Reconsidering the Effects of Marital Disruption: What Happens to Children of Divorce in Early Adulthood?" *Journal of Family Issues* 15:173-90.

Ghilagaber, G. 1993. "Family Initiation among Swedish Males Born 1936-1964: The Choice between Marriage and Cohabitation." *Stockholm Research Reports in Demography* 77. Stockholm: University of Stockholm.

Goldscheider, F.K. and C. Goldscheider. 1989. "Family Structure and Conflict: Nest-Leaving Expectations of Young Adults and Their Parents." *Journal of Marriage and the Family* 51:87-97.

---------------------------. 1993. *Leaving Home before Marriage: Ethnicity, Familism, and Generational Relationships*. Madison: University of Wisconsin Press.

Heckman, J. 1979. "Sample Selection Bias as a Specification Error." *Econometrica* 47:153-61.

Hetherington, E.M. 1972. "Effects of Father Absence on Personality Development in Adolescent Daughters." *Developmental Psychology* 7:313-26.

Hetherington, E. M., and W. G. Clingempeel. 1992. *Coping with Marital Transitions*. Chicago: Society for Research in Child Development.

Jöreskog, K.G. and D. Sörbom. 1989. *Lisrel 7 User's Reference Guide*. Mooresville, IN: Scientific Software, Inc.

Kiernan, K.E. 1988. "The British Family: Contemporary Trends and Issues." *Journal of Family Issues* 9:298-316.

-----------------. 1992. "The Impact of Family Disruption in Childhood on Transitions Made in Young Adult Life." *Population Studies* 46:213-34.

Kiernan, K.E. and P.L. Chase-Lansdale. 1993. "Children and Marital Breakdown: Short- and Long-Term Consequences." Pp. 295-308 in *European Population, vol. II: Demographic Dynamics,* edited by A. Blum and J.-L. Rallu. London: John Libby.

Kiernan, K.E. and V. Estaugh. 1993. *Cohabitation, Extra-Marital Childbearing and Social Policy*. London: Family Policy Studies Centre.

Liefbroer, A.C., M. Corijn, and J. de Jong Gierveld. 1993. "Similarity and Diversity in the Onset of Family Formation in the Low Countries." Presented at the European Science Foundation Conference, Schloss Ringberg, Germany.

Maddala, G.S. 1983. *Limited Dependent and Qualitative Variables in Econometrics*. Cambridge, UK: Cambridge University Press.

Manski, C.F., G.D. Sandefur, S. McLanahan, and D. Powers. 1992. "Alternative Estimates of the Effect of Family Structure during Adolescence on High School Graduation," *Journal of the American Statistical Association* 87:25-37.

McLanahan, S. and L. Bumpass. 1988. "Intergenerational Consequences of Family Disruption." *American Journal of Sociology* 94:130-52.

McLanahan, S., and G. Sandefur. 1994. *Growing Up with a Single Parent: What Hurts, What Helps.* Cambridge, MA: Harvard University Press.

McLeod, J.D. 1991. "Childhood Parental Loss and Adult Depression." *Journal of Health and Social Behavior* 32:205-20.

Mitchell, B.A., A.V. Wister, and T.K. Burch. 1989. "The Family Environment and Leaving the Parental Home." *Journal of Marriage and the Family* 51:605-13.

Newcomer, S. and J.R. Udry. 1987. "Parental Marital Status Effects on Adolescent Sexual Behavior." *Journal of Marriage and the Family* 49:235-40.

Office of Population Censuses and Surveys. 1990. *Marriage and Divorce Statistics,* *1837-1983, England and Wales*. Series FM2, No. 16. London: Her Majesty's Stationery Office.

Pringle, M.F.K. 1965. "Southgate Group Reading Tests." Pp. 812-13 in *The Sixth Mental Measurements Yearbook*, edited by O.K. Buros. Highland Park, NJ: Gryphon Press.

Pringle, M.F.K., N. Butler, and R. Davie. 1966. *11,000 Seven Year Olds*. London: Longman.

Rindfuss, R.R. and J.A. Sweet. 1977. *Postwar Fertility Trends and Differentials in the United States.* New York: Academic Press.

Rutter, M., J. Tizard, and K. Whitmore, eds. 1970. *Education, Health, and Behaviour*. London: Longman.

Shepherd, P.M. 1985. *The National Child Development Study: An Introduction to the Background of the Study and the Methods of Data Collection*. London: Social Statistics Research Unit, City University.

Southgate, V. 1962. *Southgate Group Reading Tests: Manual of Instructions*. London: University of London.

Stolzenberg, R.M. and D. Relles. 1990. "Theory Testing in a World of Constrained Research Design." *Sociological Methods and Research* 18:395-415.

Stott, D.H. 1969. *The Social Adjustment of Children*. London: University of London Press.

Thornton, A. 1991. "Influence of the Marital History of Parents on the Marital and

Cohabitational Experiences of Children." *American Journal of Sociology* 96:868-94.

U.S. National Center for Health Statistics. 1978. *Divorces and Divorce Rates: United States.* Vital and Health Statistics, Series 21, No. 29. Washington, DC: U.S. Government Printing Office.

---------------------. 1984. "Advance Report of Final Divorce Statistics, 1981." *Monthly Vital Statistics Report* 32, No. 9, Supplement (2). Washington, DC: U.S. Government Printing Office.

Wadsworth, M., M. Mclean, D. Kuh, and B. Rodgers. 1990. "Children of Divorced and Separated Parents: Summary and Review of Findings from a Long-Term Follow-up in the U.K." *Family Practice* 7:

Wallerstein, J.S. and S. Blakeslee. 1989. *Second Chances.* New York: Ticknor and Fields.

Young, C. 1987. *Leaving Home in Australia*. Melbourne: Australian National University and Australian Institute of Family Studies.

Table 1. Means and Standard Deviations for Variables Used in Models

	Men		***Women***	
Variables	***Mean***	***Standard Deviation***	***Mean***	***Standard Deviation***
Parent Divorced	***0.055***	***0.189***	***0.059***	***0.200***
Parent Died	***0.049***	***0.180***	***0.048***	***0.178***
Emotional Problems at 7	***-0.268***	***0.118***	***-0.280***	***0.109***
School Achievement at 7	***21.7***	***5.30***	***20.2***	***4.65***
Economic Status at 7	***0.166***	***0.274***	***0.199***	***0.279***
Class Background at 7	***0.273***	***0.256***	***0.289***	***0.271***
Lambda	***0.419***	***0.134***	***0.403***	***0.144***

Table 2. Age-23 Outcome Variables and Statistical Models.

Outcome Variable	Definition	Statistical Model
1. Leaving home because of friction 2. Leaving home for other reasons	Three-category nominal variable: has left because of friction at home; has left for other reasons; not yet left by age 23.	Multinomial logistic model
3. Cohabitation before first marriage	Whether cohabited before marrying by age 23	Binomial logistic model
4. Marriage before cohabitation	Whether married before cohabiting by age 23	Binomial logistic model
5. First marriage	Whether married by age 23; cohabitation ignored	Binomial logistic model
6. First union	Whether ever in union (cohabiting or marital) by age 23	Binomial logistic model
7. Premarital first birth	Whether gave birth to (or fathered) a child before marriage by age 23	Binomial logistic model
8. Marital first birth	Whether gave birth to (or fathered) a child after marrying by age 23	Binomial logistic model
9. First birth	Whether gave birth to (or fathered) a child by age 23; marital status at time of birth ignored.	Binomial logistic model

Table 3. Estimated Odds Ratios for Models of Effects of Parental Marital Disruption and Age-7 Characteristics on Age-23 Outcomes: Women

	1.00	2.000	3.000	4.000	5.000	6.000	7.000	8.000	9.000
Variables	Left Home Because of Friction vs. Never Left[a]	Left Home for Other Reasons vs. Never Left[a]	Premarital Cohabitation[b]	Marriage, No Cohabitation[b]	Marriage, Whether Cohabited or Not[b]	First Union (Marital or Cohabiting)[b]	Premarital Birth[b]	Postmarital Birth[b]	Birth, Whether Pre- or Postmarital[b]
Parent Divorced	5.07***	1.68**	2.32***	0.789	1.200	1.71***	3.34***	1.030	1.84***
Parent Died	1.76**	0.740	1.030	0.943	0.888	0.951	1.59*	0.786	0.969
Emotional Problems at 7	275.8***	0.666	19.2***	0.198*	1.700	1.540	170.9***	13.74***	155.56***
School Achievement at 7	1.09***	1.04***	1.06***	0.969**	0.994	1.010	1.020	0.990	1.010
Economic Status at 7	0.86	0.619	1.090	0.678	0.795	0.668	1.050	1.130	1.310
Class Background at 7	2.1*	1.99***	2.19***	0.238***	0.283***	0.396***	0.278***	0.203***	0.198***
Lambda	2.41**	0.990	2.51***	0.414***	0.428***	0.788	2.83***	0.739	1.100
Constant	0.147***	1.60**	0.087***	2.69***	4.10***	3.29***	0.134***	1.38*	2.19***
Number of Observations	5,210.00	5,210	5,152.000	5,152	5,209	5,152	5,209	5,209	5209

[a] Columns 1 and 2 constitute a multinomial logistic model of leaving home by age 23 with three categories: never left, left because of friction, left for other reasons.

[b] Binomial logistic regression model.

* p<.10; ** p<.05; *** p<.01

Table 4. Estimated Odds Ratios for Models of Effects of Parental Marital Disruption and Age-7 Characteristics on Age-23 Outcomes: Men

	1	2	3	4	5	6	7	8	9
	Left Home Because of Friction vs. Never Left[a]	Left Home for Other Reasons vs. Never Left[a]	Premarital Cohabitation[b]	Marriage, No Cohabitation[b]	Marriage, Whether Cohabited or Not[b]	First Union (Marital or Cohabiting)[b]	Premarital Birth[b]	Postmarital Birth[b]	Birth, Whether Pre- or Postmarital[b]
Parent Divorced	4.82***	1.42*	2.51***	0.850	1.03	1.60***	1.95**	1.24	1.43*
Parent Died	1.96**	0.732**	1.64***	0.71*	0.842	1.00	0.871	0.921	0.903
Emotional Problems at 7	4.47	0.582	2.85*	0.636	1.37	1.17	2.03	3.79**	3.86***
School Achievement at 7	1.05**	1.06***	1.03***	0.997	1.01	1.01	0.978	0.982	0.980*
Economic Status at 7	0.225***	0.305***	0.624*	0.396***	0.426***	0.338***	0.264**	0.342***	0.289***
Class Background at 7	1.01	3.31***	1.76**	0.412***	0.426***	0.714*	0.604	0.361***	0.399***
Lambda	3.61**	1.96**	1.41	0.775	0.671	0.975	2.87**	0.705	0.982
Constant	0.0423***	0.339***	0.0844***	0.666**	0.955	0.822	0.045***	0.556**	0.637**
Number of Observations	5,183	5,183	5,137	5,137	5,180	5,137	5,180	5,180	5,180

[a] Columns 1 and 2 constitute a multinomial logistic model of leaving home by age 23 with three categories: never left, left becasue of friction, left for other reasons.

[b] Binomial logistic regression model.

* p<.10; ** p<.05; *** p<.01

Table 5. Predicted Probabilities That Selected Outcomes Would Occur, According to Parental Marital Status

Variables	*No Divorce, No Death*	*Death*	*Divorce*
Women			
Left home because of friction	*0.046*	*0.098*	*0.14*
Left home for other reasons	*0.79*	*0.71*	*0.77*
Cohabitation before marriage	*0.22*	*0.23*	*0.51*
First birth before marriage	*0.05*	*0.079*	*0.17*
Men			
Left home because of friction	*0.035*	*0.081*	*0.12*
Left home for other reasons	*0.64*	*0.54*	*0.64*
Cohabitation before marriage	*0.15*	*0.25*	*0.38*
First birth before marriage	*0.025*	*0.022*	*0.048*

FIGURE TITLE

Measurement Model of Age-7 Constructs for Girls.

FIGURE LEGEND

A confirmatory factor analysis of information obtained at age 7 mapped 13 observed indicators onto four latent variables. The figure shows, from top to bottom, the correlations among the latent factors, the loadings of the indicators on the latent factors, and the correlations among the measurement errors associated with the indicators. (The model permitted the correlation of measurement errors of information obtained from a common source, such as all reports from teachers.)