

# **Effects of Divorce on Mental Health Through the Life Course**

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Hopkins Population Center Papers on Population  
**WP 97-1**  
(February, 1997)

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This project was supported by research grant R37 HD25936 and population center grant P30 HD06268 from the National Institute of Child Health and Human Development. We thank Paul Allison, Mark Appelbaum, Lingxin Hao, and Stephen Raudenbush for their comments on earlier versions of this paper.

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## Effects of Divorce on Mental Health Through the Life Course

### Abstract

The long-term effects of divorce on individuals after the transition to adulthood are examined using information from a British birth cohort that has been followed from birth to age 33. Growth-curve models and fixed-effects models are estimated. The results suggest that part of the seeming effect of parental divorce on adults is a result of factors that were present before the parents' marriages dissolved. But in addition, the results also suggest that there is an effect of the divorce and its aftermath on adult mental health. Moreover, a parental divorce during childhood or adolescence appears to continue to have a negative effect when a person is in his or her twenties and early thirties.

## **Effects of Divorce on Mental Health Through the Life Course**

Although there is now a substantial literature on the effects of divorce on children and young adults, information about the long-term effects of divorce after the transition to adulthood is less comprehensive. We lack adequate knowledge of the patterns of continuing effects, if any, of a parental divorce during the adult life course. To be sure, there are studies that extend to older ages (Amato and Keith, 1991a); and these show that parental divorce appears to have an effect on outcomes such as lower psychological well-being (Glenn and Kramer, 1985), more depressive symptoms (Crook and Raskin, 1975; Roy, 1985), having an income below the poverty line (McLanahan, 1985), and a greater likelihood of becoming divorced (Bumpass et al., 1991). But the vast majority of these studies of adults are based on cross-sectional surveys in which individuals retrospectively report their family structure. Moreover, most studies address young adulthood. Consequently, the inferences we can draw about patterns of effects into midlife are limited.

For the purposes of this article, let us consider three time periods in which the effects of a parental divorce during childhood can be manifest. The first period is the time between the individual's birth and the point at which his or her parents separate.

Prospective studies of children (e.g., longitudinal studies that begin before the children's parents divorce---see Block et. al, 1986; Cherlin et al., 1991) suggest that some of the differences between children who would later experience parental divorce versus those who would not were observable before any of the divorces occurred. We will call these differences pre-disruption effects. They cannot be due to the disruption itself because it has not yet happened. Rather, they may be caused by parental conflict prior to the disruption; or they may indicate characteristics of the children or their parents that have influenced both their parents' marriages and their own lives. These pre-disruption differences could still be visible in adulthood. For example, a shared genetic tendency in a family, such as a history of depression, could contribute both to parents' marital distress and divorce as well as to children's depression in young adulthood, thus giving the misleading impression that the parental divorce caused the depression.

The second period is the time between a parental divorce and the end of the transition to adulthood. This period has been the most widely studied (Amato and Keith, 1991b). Many articles and books demonstrate associations between parental divorce and aspects of the transition to adulthood, such as lower educational attainment and early childbearing (McLanahan and Sandefur, 1994), and more premarital cohabitation (Thornton, 1991). The few prospective studies that extend to young adulthood indicate that these associations persist even after taking into account measured pre-disruption differences between the divorced and not-divorced groups

(Cherlin et al. , 1995). Chase-Lansdale et al. (1995) found that experiencing a parental divorce before age 16 was associated with poorer mental health in a large sample of 23-year-olds, even controlling for measured pre-disruption differences.

The third period is the time after the transition to adulthood--after the completion of school, entry into the labor force and, for many, entry into marriage, i.e., midlife in the adult life course. A key question regarding this time period is whether the effects of the disruption, if any, diminish or stabilize after the transition to adulthood or whether the effects of disruption may increase as adults from divorced families enter their 30s. If the former were true, then we might expect that the life course of adults from divorced families who successfully navigate the waters of young adulthood would differ little from adults whose parents' marriages had remained intact. In other words, if the effect of parental divorce on the life course is confined largely to events that occur in childhood and adolescence, then its significance is mainly to alter life course trajectories before adulthood. If, in contrast, a parental marital disruption experienced in childhood continues to exert an effect, it may alter mid-life-course trajectories as well.

The study we report in this article is an attempt to increase our knowledge about the effects of a childhood parental divorce on adults. We use a prospective, longitudinal design to analyze information on measures of mental health--behavior problems and malaise--from a unique, large study of British individuals who have been followed since their birth in 1958 and who were last interviewed at age 33 in

1991). By age 33, virtually all of the sample had completed their full-time education, 83 percent had ever married, and 67 percent had had a child (Ferri, 1993). The data set is of interest because its depth and breadth are unusual for survey research.

Moreover, this data set has also been central to the recent literature, although this is the first U.S. article to report on the age 33 wave. Earlier waves have been used to claim support for pre-disruption effects (Cherlin et al., 1991), in that the behavior problems of 7-year-old boys whose parents would divorce in next four years were observably greater than the behavior problems of those whose parents would remain together.

The surveys also have been used to claim support for the effect of the disruption itself and its aftermath, in that the negative effects of parental divorce before age 16 on mental health and life transitions at age 23 were statistically significant even controlling for pre-disruption child and family characteristics (Chase-Lansdale et al., 1995; Cherlin et al., 1995).

Can predisruption characteristics account for much of the apparent effects of divorce at ages 7 to 11 but not at age 23 in the same individuals? In previous work, (e.g., Chase-Lansdale et al., 1995), it has been argued that these findings are not necessarily contradictory. In the present study, we aim to integrate these disparate findings by tracing the path of the effects of divorce on indicators of mental health during the entire sweep of the British study: from age 7--the first time behavioral information was collected--through assessments at ages 11, 16, 23, and 33. To do so, we present estimates from growth-curve models and fixed-effects models, both of

which can be applied to what is referred to as panel data--longitudinal data in which a cross-section of individuals is repeatedly interviewed. In both models, the outcome variables are measures of emotional problems at each wave.

We will first present estimates from the growth-curve models, which belong to a larger class of models for panel data called random-effects models; then we will present estimates from the fixed-effects models. Growth curve analysis is presented in texts on hierarchical models that are well-known to sociologists (e.g., Bryk and Raudenbush, 1992); but it has been used primarily by psychologists (Ragosa, Brandt, and Zimowski, 1982; Burchinal and Appelbaum, 1991; Willett, Ayoub, and Robinson, 1991). A recent application is to job-related psychological distress among dual-career married couples (Barnett et al., 1993, 1995). To our knowledge, no articles using growth-curve models have been published in mainstream sociology journals.<sup>1</sup>

Random-effects and fixed-effects models make different assumptions about unobservable sources of variation and typically use the information in the data in somewhat different ways. Both kinds of models correct for the non-independence of the error terms in multiple observations on the same individual. Random-effects models assume that unobserved person-specific characteristics are uncorrelated with observed characteristics, whereas fixed effects models relax that assumption for time-invariant unobserved characteristics (Allison, 1994; Johnson, 1995). In our application unobserved characteristics such as a shared predisposition toward depression could be correlated with observed characteristics such as parental divorce,

which implies that fixed-effects models could be more appropriate. Moreover, fixed-effects models produce parameter estimates that control for unobserved characteristics that do not change over time. On the other hand, fixed effect models cannot separate out the main effects of observed characteristics that do not change over time, such as gender, from the unobserved characteristics.

Substantively, the main difference between the two models in this application is in the treatment of variables measuring parental divorce. In the growth-curve models, a set of dummy variables for age at parental divorce is taken as time-invariant. In other words, at every wave a given individual has the same scores on the set of dummy variables (e.g., a divorce never occurred, a divorce occurred between ages 7 and 10, a divorce occurred between ages of 11 and 15, etc.), no matter whether the wave occurred prior to or after the divorce. This specification allows us to estimate and to graph trajectories of mental health over the entire age range of 7 to 33 for individuals who experienced a divorce during a particular age interval. Crucially, these trajectories, or growth curves, let us examine pre-disruption as well as post-disruption effects. The growth curve specification also allows us examine the main effects of time-invariant variables such as gender and social class.

In the fixed effects models, parental divorce is measured as a single, time-varying dummy variable. At each wave, the variable is coded 1 if a divorce had occurred by that time, and 0 if it had not yet occurred (or never occurred). The estimates from our fixed-effects models pertain only to post-disruption effects and



cannot include the main effects of gender and social class. However, their ability to control for time-invariant unobserved variables suggests that they may provide a stronger test of whether post-disruption effects can be said to exist.

A challenge in any longitudinal study of individual characteristics is to develop comparable measures of the characteristics over time. When the study extends through several decades, as ours does, and the characteristic relates to mental health, the challenge is even greater. One cannot ask identical questions concerning the behavioral problems of seven-year-olds and thirty-three-year olds. We will describe how we addressed this problem. Nevertheless, identical measures were asked at ages 23 and 33. Consequently, we will supplement our analysis with a change-score regression model for ages 23 and 33.

An additional question is whether the effects of a parental divorce differ for individuals of different socioeconomic positions and for women compared to men. If, for example, a substantial portion of the effects of divorce are due to the economic difficulties that single parents and their children often face, then the consequences could be different for children from comfortable homes as opposed to poor homes. If parents shield girls from conflict more than boys, (c.f., Chase-Lansdale and Hetherington, 1990), then the long-term consequences could differ by gender. We examine these questions, although the socioeconomic information in the British study is somewhat limited.<sup>2</sup>

## **PARENTAL DIVORCE AND ADULT CHILDREN'S MENTAL HEALTH INTO MIDLIFE**

Would we expect effects of parental marital disruption on emotional problems to persist at age 33? Even if marital disruption had an effect that persisted beyond the first several years, it could fade afterward. On the other hand, a life-course perspective might lead one to expect that the disruption could trigger intervening events that negatively affected adult mental health. Clinical and developmental psychological studies also suggest that negative trajectories of poor mental health may be lasting for some individuals. Research shows substantial continuity between childhood depression and adult depression (Harrington et al., 1990), with evidence that persons who have first onset of depression before age 20 have a higher likelihood of recurrence than do those whose first episode occurs after age 20 (Giles et al., 1989). And some studies also report that parental divorce in childhood is associated with greater depression in adulthood (Lauer and Lauer, 1991). It is plausible, then, that parental divorce may cause an initial depressive episode in children and adolescents and that depression may reoccur in adulthood. The evidence, however, is inconsistent. Kessler and Magee (1993), in an analysis of a two-wave national survey of adults, report that parental divorce is not associated with early (before age 20) onset of depression but that parental divorce does increase the likelihood that adults who have had one depressive episode will have another one. It is also possible that continuity exists because individuals with symptoms of emotional problems are more likely to

have experienced a parental divorce--either because they came from families with histories of emotional problems or because their own disorders helped precipitate parental divorces. In these cases, a pre-disruption effect should be visible.

## **DATA**

The National Child Development Study (NCDS) is a longitudinal study of children who were born in England, Scotland, and Wales in the first week of March, 1958. Since the NCDS has been described and analyzed elsewhere (Cherlin et al., 1995; Chase-Lansdale et al., 1995), we will not discuss it in full detail here.

Interviews were conducted with 17,414 mothers, who represented 98 percent of all births in that week (Shepard, 1985). Follow-up interviews were conducted with parents and teachers at 7, 11, and 16. At ages 23 and 33, the cohort members were interviewed. We restrict our focus to the 11,759 cohort members whose parents were in an intact marriage at age 7--the first time there is information about the child other than birth weight--and for whom there is subsequent information on parental marital status.<sup>3</sup> This restriction allowed the construction, using confirmatory factor analysis, of three latent-variable measures representing pre-disruption characteristics at age 7:

- Class background: a combination of father's occupation (manual vs. non-manual), whether the father stayed in school past minimum age, and whether the mother stayed in school past minimum age;

- Economic status: a combination of whether the family owned vs. rented its home, the number of persons per room in the household, and an indicator of whether or not the family was experiencing “economic difficulties;” and
- School achievement: a combination of a score on a standardized reading achievement test, a score on a standardized mathematics test, and a score on a 5-item scale of teacher’s assessments of “oral ability,” “awareness of the world around him,” “reading,” “creativity,” and “number work” (alpha reliability .89).

In addition, at the age 7 interview, parents were asked to rate the children’s behavior problems using most of the items from the Rutter Home Behaviour Scale (Rutter et al., 1970). The scale was designed to identify two broad groupings 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. An 18-item summated scale had an alpha reliability of .71.<sup>4</sup> This age-7 behavior problems scale became our initial, pre-disruption measure of emotional problems. At ages 11 and 16, parents were again asked to rate behavior problems using similar but not identical items. A 10-item scale at age 11 was constructed and had an alpha reliability of .68<sup>5</sup>; and a 22-item scale at 16 had an alpha reliability of .75<sup>6</sup>.

At ages 23 and 33, the cohort members were asked the 24 yes-no questions in

the Malaise Inventory, designed by Rutter et al. (1970). It is a screening instrument that samples a wide range of adult emotional disorders, such as depression, anxiety, phobias, and obsessions. Because depression is more prevalent in adult populations than other problems, the Malaise Inventory overrepresents items related to depression. The Inventory was identical at age 23 and 33, with alpha reliabilities of .78 and .81, respectively.<sup>7</sup>

## **METHODS**

### ***Growth-curve models***

Growth curve analysis provides a way to model the change in an attribute of an individual over time. Examples include the development of speech in young children (Burchinal and Appelbaum, 1991), changes in feelings of isolation from friends (Osgood and Smith, 1995), changes in perceived marital quality (Karney and Bradbury, 1995), and the age-pattern of scores on repeated academic achievement or ability tests (Kerbow, 1992; Hoffer, 1994). Theoretically, the model assumes an underlying path of change in the outcome variable that is generally applicable to all individuals in the study; but it also allows particular characteristics or events (such as whether their parents divorced) to modify that path. In our analyses, the outcome is emotional problems, which were measured at ages 7, 11, 16, 23, and 33.

The process is modeled at two levels. Level 1 consists of repeated observations of individuals over time. Let  $Y_{it}$  be the score of person  $i$  on the outcome variable (in our case, an indicator of emotional problems) at time  $t$ , and let  $a_{it}$  be the age of person  $i$  at time  $t$ . The growth curve for the outcome variable is modeled as follows (using the notation of Bryk and Raudenbush, 1992):

$$Y_{it} = B_{0i} + B_{1i}a_{it} + B_{2i}a_{it}^2 + \dots + B_{pi}a_{it}^p + e_{it} \quad (1a)$$

for  $i = 1, \dots, n$  individuals. The  $B_{pi}$  parameters are associated with a polynomial in age of degree  $p$ ; that is,  $B_{0i}$  is a constant corresponding to an intercept for individual  $i$ ,  $B_{1i}$  is a linear slope associated with age at  $t$  for individual  $i$ ,  $B_{2i}$  is a quadratic slope corresponding to the square of age at  $t$ , and so forth. The error term  $e_{it}$  is usually assumed to be independently and normally distributed with a mean of 0 and a constant variance  $F^2$ . In practice, most models have been either linear,

$$Y_{it} = B_{0i} + B_{1i}a_{it} + e_{it}, \quad (1b)$$

or quadratic,

$$Y_{it} = B_{0i} + B_{1i}a_{it} + B_{2i}a_{it}^2 + e_{it}. \quad (1c)$$

In our case, with a maximum of five observations per individual (at ages 7, 11, 16, 23, and 33), we found estimates of linear models to be acceptable but estimates from quadratic models to be of questionable reliability (see below).

Note that the unit of observation in the Level-1 model is not the individual but rather an observation on an individual at one point in time. Importantly, the method allows the analyst to estimate how characteristics of the individual modify the values

of the  $B_{pi}$ . In other words, the method allows characteristics of the individual to alter the way in which the attribute changes over time. This modeling is done in a Level-2 model, for which the unit of observation is the individual and the dependent variables are the  $B_{pi}$  parameters themselves. In the linear case, there are two such parameters, the Level-1 intercept  $B_{0i}$  and the Level-1 slope  $B_{1i}$ . The Level-2 model is:

$$B_{0i} = \$_{00} + \$_{01}X_{1i} + \$_{02}X_{2i} + \dots + \$_{0q}X_{qi} + r_{0i} \quad (2a)$$

$$B_{1i} = \$_{10} + \$_{11}X_{1i} + \$_{12}X_{2i} + \dots + \$_{1q}X_{qi} + r_{1i}, \quad (2b)$$

where the  $X_{qi}$  are measures of characteristics 1 through  $q$  for individual  $i$ ; the coefficients  $\$_{pq}$  are the effects of the characteristics on the  $B_{pi}$  intercept and slope parameters ( $p=0$  refers to the intercept, and  $p=1$  refers to the slope); and the  $r_{pi}$  are error terms that are assumed to be uncorrelated with the  $X_{qi}$  characteristics and multivariate normally distributed with means of zero and a covariance matrix  $\mathbf{T}$ . The error terms represent unmeasured characteristics of individual  $i$  that do not change over time.

It is the  $\$_{pq}$  coefficients that are of greatest interest in our model. They describe how variations in such characteristics as parental divorce, gender, or social class background alter the growth curve for an individual by changing the values of the  $B_{pi}$  intercept and slope parameters. Note that the subscript  $t$  does not appear in the Level-2 equations; the unit of analysis is the individual  $i$ , and the characteristics  $X_{qi}$  are invariant over time. Our key characteristics will be time-invariant measures of whether an individual ever experienced parental divorce. Estimates of the variance

and covariance components of the growth-curve model were obtained by maximum likelihood methods, and then estimates of the  $\beta_{pq}$  coefficients were obtained by generalized least squares (Bryk and Raudenbush, 1992).<sup>8</sup>

The growth-curve model requires that the repeated measures of the attribute of interest, in this case emotional problems, be measured comparably at each point in time. Otherwise, one cannot discern the underlying path of change. But if the span of time is long, such as the 26-year period between 7 and 33 in our case, comparable measurement may not be possible. Asking about reluctance to go to school is a good way to measure anxiety at age 7 but is inappropriate at 33. Conversely, the Malaise Inventory can ask 23- or 33-year-olds whether they have ever had a nervous breakdown, but no such question can be asked of 7-year-olds.

In order to estimate a growth-curve model of emotional problems from age 7 to 33, consequently, we had to standardize the scale scores at each age to have a mean of 0 and a standard deviation of 1. This insured comparability of measurement, but it did so at a price. Given the standardization, the expected score for the average person at every age is 0; in other words, the typical growth curve is a horizontal line. We cannot observe the actual, unstandardized shape of the curve. Although this is a serious limitation, we can still examine how the growth curve differs according to whether, and when, a parental divorce occurred: in the Level-2 model, we can estimate the effects of divorce on the Level-1 slope and intercept parameters. If parental divorce raises the level of emotional problems at certain ages, we would expect the estimated



growth curve for the divorced group to be above the curve for the not-divorced group. We are thus estimating the effects of parental divorce on the growth curve relative to the path of those whose parents did not divorce.

In order to explore whether the standardization process influenced our findings, we redid our analyses using only the externalizing disorders sub-scales at 7, 11, and 16. We also redid the analyses using only the internalizing disorders sub-scales at 7, 11, and 16. The results were nearly identical to the ones we report below.<sup>9</sup> In addition, we conducted a regression-based analysis using just the age 23 and 33 data points, at which times the emotional problems measures were identical and standardization was not necessary. We estimated regression equations in which the dependent variable was the change in the (logged) raw score of the Malaise Inventory between 23 and 33 and parental divorce was the key independent variable.<sup>10</sup> It can be shown that a raw difference-score regression is equivalent to a growth-curve model with observations at only two time points.<sup>11</sup> We will report the regression results below. Table 1 presents the means and standard deviations of the variables used in the analyses.

TABLE 1 ABOUT HERE

### ***Fixed-Effects Models***

Along with their strengths, random-effects models, including growth-curve

models, have some limitations, as noted earlier. In Equations 2a and 2b it is assumed that the unmeasured characteristics of individual  $i$ , represented by the error terms  $r_{0i}$  and  $r_{1i}$ , are uncorrelated with the measured characteristics,  $X_{1i} \dots X_{qi}$ . Fixed-effects models do not require this assumption; and they also control for unmeasured characteristics that do not change over time. Consider a single-level model of the form:

$$Y_{ti} = \alpha_i + \beta_1 X_{1ti} + \dots + \beta_q X_{qti} + e_{ti}, \quad (3)$$

where  $\alpha_i$  is a fixed constant that differs for each individual  $i$ , representing unmeasured characteristics of that individual that do not change over time. If the number of individuals is small, the model can be estimated by ordinary least squares by including  $n - 1$  dummy variables for the fixed constants. With many individuals, however, this method is impractical. If each individual is observed twice, at times  $t = 1$  and  $t = 2$ , we can write:

$$Y_{1i} = \alpha_i + \beta_1 X_{11i} + \dots + \beta_q X_{q1i} + e_{1i}, \quad \text{and} \quad (4a)$$

$$Y_{2i} = \alpha_i + \beta_1 X_{12i} + \dots + \beta_q X_{q2i} + e_{2i}. \quad (4b)$$

Then, subtracting the first equation from the second, we obtain:

$$Y_{2i} - Y_{1i} = \beta_1 (X_{12i} - X_{11i}) + \dots + \beta_q (X_{q2i} - X_{q1i}) + (e_{2i} - e_{1i}). \quad (4c)$$

Note that the  $\alpha_i$  terms drop out of this equation, which now expresses the change in the outcome variable for individual  $i$  as a function of changes in the measured right-hand-side variables. Because the  $\alpha_i$  terms drop out, the model controls for time-invariant unmeasured characteristics. It is still possible, however, that there are

unmeasured variables that change over time and that affect both emotional problems and whether parents divorce. These are included in the  $e_{it}$  terms, and fixed effects model do not control for them. The parameters in Equation 4c can be estimated by ordinary least squares.

If there are observations at more than two time points, a generalization of Equation 4c is to express, at each time point, an individual's outcome score as a deviation from his or her mean outcome score across all observations of that individual:

$$Y_{it} - \bar{Y}_i = \beta_1 (X_{1it} - \bar{X}_{1i}) + \dots + \beta_q (X_{qit} - \bar{X}_{qi}) + (e_{it} - \bar{e}_i) \quad (5)$$

Measured variables that do not change over time drop out of Equation 5 because the difference terms are always zero, which is why fixed effects models cannot estimate the main effects of time-invariant variables such as gender. It is possible, however, to estimate the interaction effect of a time-varying variable (such as whether a person's parents have ever divorced) and a time-invariant variable (such as gender). In our application, for every observation on an individual, the key right-hand-side variable is a binary variable coded 1 if a parental divorce has occurred to that individual at any time prior to that observation and coded 0 otherwise.

## FINDINGS

### *Growth-curve models*

We estimated a linear growth-curve model in which the outcome variable was a measure of emotional problems for an individual  $i$  at time  $t$ . The right-hand-side, Level-1 variable was the age of individual  $i$  at time  $t$  minus 7. We subtracted 7 from age so that the intercept parameter  $B_{0i}$  would be the expected level of emotional problems at the start of the observation period (age 7).

The Level-1 model can be written as:

$$(\text{Emotional problems})_{it} = B_{0i} + B_{1i}(\text{age} - 7)_{it} + e_{it}.$$

As expected because of the standardization, the average values of the intercept and slope parameters were close to 0. However, a model that allowed the intercept and slope parameters to vary about their average value from individual to individual fit the data better than a model that treated the intercept or slope as the same for all individuals. The individual variation in the intercept and slope was modeled at Level 2 as a function of whether (and when) the individual's parents divorced, gender, economic status at 7, class background at 7, and school achievement at 7.

Table 2 presents the estimated parameters from two specifications of the Level-2 model. Consider specification 1. The upper panel shows the effects of the covariates on the intercept parameter--which is the expected level of emotional problems at age 7. Recall that none of the individuals in this analysis had experienced

a parental divorce at age 7. Nevertheless, a parental divorce between ages 7 and 22 increased the intercept term by .114 standard deviations, compared to no parental divorce (the reference category). This statistically significant pre-disruption effect implies that persons whose parents would later divorce already had elevated levels of emotional problems at age 7, before any of the divorces had occurred.

#### TABLE 2 ABOUT HERE

Women had a lower predicted intercept than men. It is likely that this finding reflects the lower level of externalizing behavior (fighting, disobedience) in girls than in boys. Individuals from families with a higher economic status showed lower levels of emotional problems, as did individuals with greater school achievement at 7. No plausible interaction terms produced statistically significant findings.

The lower panel of Table 2 shows the effects of the covariates on the slope parameter. A parental divorce between 7 and 22 is associated with a statistically-significant increased slope. Thus, the growth curves for the divorced group and the non-divorced group diverge after age 7, with the curve for those whose parents divorced (by 22) rising more rapidly. In addition, the slope rises faster for women than for men. This gender difference may be connected with the content of the adult measure of emotional problems--the Malaise Inventory. As noted, it is weighted toward depression, the most common disorder in adulthood; and it is known that

women report more depressive symptoms than men (Weissman, 1987). Individuals with better school achievement scores at 7 show a lower growth of emotional problems than do those with poorer school achievement at 7. There are no significant effects of economic status or class background at 7.

Figure 1 presents a graph of the predicted growth curves for emotional problems based on specification 1 in Table 2. The predictions are for hypothetical individuals who differ on whether and when their parents divorced but who are otherwise average on all other measured characteristics. For example, for the intercept of the no-divorce group, the mean values of gender, economic status, class background, and school achievement were multiplied by their respective coefficients from the upper panel of Table 2, the resultant products were summed, and the constant was added. The identical procedure using the coefficients in the lower panel produced the predicted slope. Then, using these values of the intercept and the slope, the line for the no-divorce group was plotted. For the two parental divorce groups (age 7 to 22 and 23 to 33), the same procedure was applied, except that the coefficients for the appropriate binary variable for divorce were added to the prediction equation.

#### FIGURE 1 ABOUT HERE

The solid line represents the predicted path of emotional problems for the large group whose parents did not divorce. The path starts near zero and shows no growth;

this is the result of the standardization, which insured that the typical person would have a score near 0 at every time point. We cannot determine the shape of the unstandardized growth curve. Note, however, that the line for the group whose parents divorced when their children were 7-22 begins with a higher level of predicted emotional problems at age 7. The upper panel of Table 2 showed that the initial value for this group at age 7 is significantly higher than for the no-divorce group. Note also that the gap between the no-divorce group and persons whose parents divorced between 7 and 22 is wider at the end of the study than at the beginning because predicted emotional problems have increased for the latter group. Thus, an initial difference in emotional problems is predicted at age 7, before any of the divorces occur; but in addition, the divorced group experienced a further growth in emotional problems relative to persons whose parents never divorced.

As for the group whose parents divorced between 23 and 33, they begin with a level of emotional problems that is slightly above the not-divorced group, although Table 2 showed that the difference is not statistically significant. They show no further increase in emotional problems relative to the non-divorced group. Someone observing these two groups for the first time at age 33 might conclude that a parental divorce after age 23 produces a modest increase in emotional problems. But Figure 1 makes clear that the difference between the two groups is almost entirely a pre-disruption effect that was visible 26 years earlier. There may be some unmeasured characteristics of persons in the divorced-between-23-and-33 group that increased both

their level of emotional problems and the likelihood that their parents would divorce.

To further pursue possible effects due to the timing of divorce, we divided the single binary variable for divorce between 7 and 22 into three binary variables, reflecting the interview dates: divorce between 7 and 10, divorce between 11 and 15, and divorce between 16 and 22. This is specification 2 in Table 2. A likelihood-ratio test indicates that specification 2 fits the data better than specification 1 ( $\chi^2 = 23.5$ , d.f. = 6,  $p < .001$ ). The results allow us to examine the pre-disruption effect more closely. Parental divorces that occurred during adolescence and young adulthood--11 to 22--were associated with significantly higher initial levels of emotional problems at 7. If the pre-disruption effect at age 7 merely reflected parental conflict soon before a divorce, then we would have expected persons in the parental divorce at 7-10 group to have had significantly higher initial levels of emotional problems because the 7-10 interval is closest to age 7; but this expectation was not borne out. It thus appears that the pre-disruption effect of parental divorce at age 7 is not simply a result of the start of the divorce process. Rather, it may also reflect unmeasured characteristics of the individual or the parents that influenced both early emotional problems of the child and a subsequent divorce of the parents.

Figure 2 presents a graph of the predicted growth curves for emotional problems based on specification 2. The highest lines are for divorce between 11 and 15 and divorce between 16 and 22, reflecting their higher initial levels of emotional problems. The largest slopes--representing the linear rate of increase in emotional



problems relative to the no divorce group--are for ages 7-10 and 16-22. Overall, the graph suggests that the effects of divorce are broadly similar for the 7-10, 11-15, and 16-22 group, all of which have elevated levels relative to the divorce at 23-33 and no divorce groups.

### FIGURE 2 ABOUT HERE

A linear model, however, is not fully satisfactory. Because it has an unchanging slope, it cannot answer the question of whether the gap between the divorced group and the not-divorced group continues to diverge or whether it levels off. Put another way, it cannot tell us whether the effects of a divorce in childhood or adolescence continue to raise levels of emotional problems through a person's twenties and early thirties or whether the effects are mainly confined to adolescence and the transition to adulthood. To answer this question, one would need at minimum a quadratic growth curve (containing age squared in the Level-1 model). Our data, however, did not support a quadratic model. When we tried to estimate one, the algorithm converged with difficulty and diagnostics suggested that the results were questionable. (The estimated reliability of the quadratic slope parameter--the proportion of its variance that predicts variation in emotional problems--was 0.0.) It is likely that with only 5 data points per person, it is asking too much of the data to fit a quadratic model that requires three parameters: an intercept, a linear slope, and a

quadratic slope.<sup>12</sup>

Consequently, in order to investigate whether a divorce in childhood or adolescence has a continuing effect on emotional problems after age 23, we turned to a difference-score regression model of change between 23 and 33. As noted earlier, it is equivalent to a two-period growth-curve model. The outcome variable was the difference between the logarithm of the Malaise Inventory score at 33 and the logarithm of the Malaise Inventory score at age 23:

$$\log(\text{malaise at } 33)_i - \log(\text{malaise at } 23)_i = \beta_0 + \beta_1 X_{1i} + \beta_2 X_{2i} + \dots + \beta_q X_{qi} + e_i$$

,

where  $X_{1i} \dots X_{qi}$  are the same set of covariates as in specification 1 of Table 2.

Because the outcome measures were identical at the 2 time points, standardization was not necessary. The results are reported in Table 3.

#### TABLE 3 ABOUT HERE

As can be seen, a parental divorce between 7 and 22 is associated with a significant increase in malaise inventory scores between 23 and 33. So the results indicate that, through their twenties and early thirties, the mental health of adults who experienced parental divorce in childhood or adolescence continued to diverge from the mental health of adults whose parents did not divorce. Whether this enduring effect is due to the disruption or to some unmeasured characteristics that are correlated

with the disruption, we cannot say for sure. The finding did not change when, in an attempt to control for pre-disruption characteristics, we included the age 7 emotional problems scale score as an additional right-hand-side variable. So the finding is not simply a reflection of early, pre-disruption indicators of emotional problems.<sup>13</sup> As before, a parental divorce between 23 and 33 has no significant effect; therefore, the trajectory of mental health of individuals whose parents divorced while they were in their twenties or early thirties did not, on average, differ from the trajectory of mental health of the no divorce group. The NCDS data suggest that a parental divorce that occurs in adulthood does not cause a worsening of mental health.

### *Fixed-Effects Models*

The rising slopes in Figures 1 and 2 are consistent with a post-disruption effect of parental divorce on emotional problems. Yet the rising slopes also could be caused by unmeasured characteristics that are correlated with both parental divorce and levels of emotional problems that rise as a person enters adolescence and adulthood. Using fixed-effects models, we can examine the effects of parental divorce controlling for time-invariant unmeasured characteristics.

Specification 1 in Table 4 shows that if, at a given observation point, a parental divorce had already occurred to an individual, his or her level of emotional problems was about 0.1 standard deviations higher than his or her average level across all observation points. The difference is statistically significant at the .001 level. This

result increases our confidence that there is an effect of a parental divorce and its aftermath--that the apparent effect of parental divorce is not simply due to unmeasured time-invariant characteristics of the individual and her or his family. To be sure, it still could be the case that time-varying unmeasured variables are responsible for some part of this effect. But we are on firmer ground in accepting the existence of a true post-disruption effect than was the case with the random-effects models.

#### TABLE 4 ABOUT HERE

The main effect of age on emotional problems is near 0, as would be expected because of the use of standardized scores at each age. Specification 2 adds two interactions: age combined with being female; and age combined with class background at age 7.<sup>14</sup> Age has an upward effect on emotional problems for women, reflecting the greater emphasis of the adult outcome measure, the Malaise Inventory, on depression (which reflects, in turn, the prominence of depression in adult emotional problems). Women, as discussed previously, tend to report more symptoms of anxiety and depression than do men. Persons whose class background at age 7 is higher have a lower increase in emotional problems as they age. The NCDS data do not allow us to determine which of many plausible mechanisms, such as greater material resources or more education, underlies this finding. There were no significant between parental divorce and gender or between parental divorce and the age 7 characteristics.

## DISCUSSION

The models presented in this paper suggest that the difference in mental health at age 33 between persons whose parents had divorced and persons whose parents had not divorced was due in part to pre-disruption differences and in part to the effect of the divorce and its aftermath. Yet it is difficult to be certain about the latter, post-disruption effect. The models also suggest that the mental health of individuals in the two groups continued to diverge in their twenties and early thirties.

As for the pre-disruption effect, growth-curve models indicated that a parental divorce that occurred between age 7 and 22 to a member of the NCDS 1958 birth cohort was related to elevated levels of emotional problems at age 7, even before any of the divorces had occurred. A difference of .11 standard deviations already existed at age 7 between individuals whose parents would later divorce by age 22 compared to individuals whose parents would remain together until their children were age 23. By the time the cohort members were 33 years of age, the difference had expanded to .25 standard deviations.<sup>15</sup> The existence of a statistically significant difference at age 7 indicates pre-existing differences between the families that would later disrupt and those that would remain intact. Although it is possible that the higher level of emotional problems among the 7-year-olds in the divorced group represented their reactions to the beginnings of the process of divorce, we doubt that this is a sufficient explanation because a significant pre-disruption effect at age 7 was observed even for the group whose parents did not divorce until they were 16 to 22. Rather, we think it

may be that some early characteristics of the children or their parents were associated with both a later parental divorce and a later elevated level of emotional problems. The NCDS data, although relatively rich for a large-sample survey, do not have enough information to allow for a determination of what these pre-existing differences were. It is likely that, in many cases, elevated behavior problems at 7 were a reaction to other sources of stress in the family such as continual marital conflict, substance abuse, or violence, of which there are no measure in the NCDS. The sources also could have included financial hardships that were beyond the measurement capability of the somewhat limited economic indicators in the study. It may even be that in some cases a child's difficult temperament was partially responsible for a later divorce.

As for the widening after age 7 of the gap between the no-divorce group and the group whose parents divorced between 7 and 22, it would seem straightforward to interpret it as an effect of the divorce and its aftermath. A large literature documents the diminished parenting, economic declines, and continuing parental conflict that children often experience after a divorce (Hetherington and Clingempeel, 1992). Yet it is possible that some time-varying pre-disruption characteristics that weren't fully measured could have widened the gap after 7. In other words, one's confidence that the results reflect a post-disruption effect depend on one's judgment about how well the NCDS measured relevant pre-disruption factors. Without doubt the NCDS measured them better than retrospective surveys of adults because it included detailed questions on behavior problems, test scores, and other information that cannot be

obtained reliably through retrospective reports. To that extent, it provides a stronger basis than most previous studies for concluding that there is an effect of parental divorce itself on long-term levels of emotional problems. Nevertheless, it is possible that some portion of what appears to be a post-disruption effect could be a result of pre-disruption factors that weren't fully measured.

One might expect the effect of the disruption to vary by the social class background or economic difficulties of the family, but we found no significant interactions in either the growth-curve models or fixed-effects models. As best as can be determined from the NCDS, the relationship of parental divorce to subsequent emotional problems did not vary by social class or economic status. (Nevertheless, the growth-curve models indicated that the level of emotional problems at age 7 was higher for cohort members with economic status or lower school achievement at 7.)

We also found strong effects of gender on levels of emotional problems which were consistent with the literature. Girls showed lower levels of emotional problems at age 7, a time when their tendencies toward internalizing disorders (e.g., anxiety and depression) are harder to measure than boys' tendencies toward externalizing disorders. Conversely, at age 33, women showed higher levels of emotional problems on a scale that was weighted toward internalizing disorders, which are the most common form of adult emotional problems. Many studies have shown that adult women report higher levels of symptoms of mental health problems (Weissman, 1987). Nevertheless, we did not find a significant interaction between gender and the

effects of parental divorce on emotional problems; rather, the process seemed to be similar for women and men.

Finally, let us return to the question of the effect of parental divorce on the adult life course. The NCDS data suggest a continuing effect of parental divorce, or of pre-disruption characteristics associated with it, on adult mental health. The effect size of .25 standard deviations at age 33 appears to be in line with previous studies of divorce (Amato and Keith, 1991a). Studies using the age 23 wave of the NCDS have suggested that divorce raises the risk of negative outcomes such as emotional problems but that most individuals whose parents divorce do not experience these outcomes (Chase-Lansdale et. al., 1995). Our study adds the information that parental divorce in childhood or adolescence, or its correlates, still seems to raise the risk of negative outcomes after age 23: difference-score regressions showed a divergence in the Malaise Inventory scores of the divorce group and the no-divorce group between ages 23 and 33. This continuing effect is not just due to recent divorces; in fact, parental divorces between 23 and 33, after most adult children have left home, were not associated with a divergence in Malaise Inventory scores.

If the continuing effect were truly due to the divorce itself, rather than to unmeasured factors, it would suggest that this childhood event can set in motion a chain of circumstances that are still altering individuals' lives even after most have left home, married, and entered the labor force. The exact nature of these continuing effects cannot be determined from the NCDS. The parental divorce could set in



motion events such as early childbearing or limits in schooling that, in turn, affect adult outcomes. Or parental divorce could be partly a marker for individual characteristics that themselves hinder adult development. In any case, the NCDS data suggest that the life course of individuals whose parents divorced continues to diverge in adulthood from the life course of those whose parents did not.

## NOTES

1. But see two unpublished papers: Kerbow (1992) and Hoffer (1994). A methodological presentation of the use of growth-curve models in studies of marital quality is presented in Karney and Bradbury (1995).
2. Racial and ethnic variation is also of interest, but the 1968 birth cohort in Great Britain was overwhelmingly (96 percent) white.
3. Previous articles that followed the cohort through age 23 attempted to adjust for the possible sample-selection bias inherent in retaining only the subset of children whose parents were still married at age 7 and who were interviewed at age 23. The use of a standard two-step correction procedure (Heckman, 1979; Maddala, 1983) did not alter the results (Cherlin et al., 1995). We have not included a sample-selection correction in the analyses we report in this article.
4. The age-7 items 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.
5. The age-11 items were: difficulty settling in, bullied by other children, destroys others' property, miserable/tearful, squirmy/fidgety, worries, irritable--quick to fly off the handle, upset by new situations, fights with other children, and disobedient.
6. The age-16 items were: stomach ache or vomiting, temper tantrums, tears on arrival at school, steals things, sleeping difficulties, restless, squirmy/fidgety, often destroys property, frequently fights, not much liked by other children, often worries, does thing on own/solitary, irritable--flies off the handle, miserable/tearful, twitches/mannerisms/tics, frequently bites nails of fingers, often disobedient, cannot

settle in, fearful of new situations, fussy/over particular, often tells lies, and bullies other children.

7. The 24 items were: Do you often have back-ache? Do you feel tired most of the time? Do you often feel miserable and depressed? Do you often have bad headaches? Do you often get worried about things? Do you usually have great difficulty in falling or staying asleep? Do you usually wake unnecessarily early in the morning? Do you wear yourself out worrying about your health? Do you often get into a violent rage? Do people often annoy and irritate you? Have you at times had a twitching of the face, head, or shoulders? Do you often suddenly become scared for no good reason? Are you scared to be alone when there are no friends near you? Are you easily upset or irritated? Are you frightened of going out alone or of meeting people? Are you constantly keyed up and jittery? Do you suffer from indigestion? Do you often suffer from an upset stomach? Is your appetite poor? Does every little thing get on your nerves and wear you out? Does your heart often race like mad? Do you often have bad pains in your eyes? Are you troubled with rheumatism or fibrositis? Have you ever had a nervous breakdown?

8. The HLM program was used to generate the estimates (Bryk et al., 1994).

9. Results available upon request. In general, the externalizing and internalizing sub-scales had lower reliabilities than full scales. See Chase-Lansdale et al., 1995.

10. We used the natural logarithms of the malaise scores because the logged scores were more symmetrically distributed than the skewed untransformed scores.

11. To simplify, let us assume that parental divorce for person  $i$ ,  $D_i$ , is the only Level-2 predictor and that, without loss of generality, the age of person  $i$  at time  $t$ ,  $A_{it}$ , is a dichotomous variable coded 0 at the first time point and 1 at the second time point. Then by Equation 1b, the Level-1 model is:

$$Y_{it} = B_{0i} + B_{1i}A_{it} + e_{it}$$

and by Equations 2a and 2b, the Level-2 model is:

$$B_{0i} = \beta_{00} + \beta_{01}D_i + r_{0i}$$

$$B_{1i} = \beta_{10} + \beta_{11}D_i + r_{1i}.$$

Then if  $A_{1i}$  is coded 0 and 1, the Level-1 model for two time points reduces to:

$$Y_{0i} = B_{0i} + 0 + e_{0i}$$

$$Y_{1i} = B_{0i} + B_{1i} + e_{1i}.$$

Therefore,

$$\begin{aligned} Y_{1i} - Y_{0i} &= B_{1i} + (e_{1i} - e_{0i}) \\ &= \beta_{10} + \beta_{11}D_i + r_{1i} + (e_{1i} - e_{0i}) \\ &= \beta_{10} + \beta_{11}D_i + u_i, \end{aligned}$$

where  $u_i (= r_{1i} + [e_{1i} - e_{0i}])$  is normally distributed with mean zero. This last equation is equivalent to the regression of the raw difference score on parental divorce.

12. For some individuals, there were less than 5 time points. One of the advantages of the growth-curve model is that individuals can be included even if they are not observed at all time points. In our sub-sample, 49 percent contributed 5 observations, 35 percent contributed 4, 14 percent contributed 3, and 2 percent contributed two. But even when we re-ran the quadratic model using only individuals who contributed 5 observations, the same problems emerged.

13. In fact, when we (1) restricted the sample to persons whose parents did not divorce before 16 and (2) used the age 16 emotional problems scale score as a control, we still found that a parental divorce between 16 and 22 significantly increased Malaise Inventory scores between 23 and 33.

14. Recall that time-invariant variables such as gender and class background at 7 drop out of the fixed-effects model but can be entered as interactions with time-varying variables such as age.

15. These figures are derived from specification 1 of Table 2. Recall that age minus 7, not age itself was used in these equations. The effect of parental divorce on the intercept of the Level-1 equation is

.114, which yields the gap at age 7. The gap at age 33 equals  $.114 + (.00536 \times (33-7))$ , where .00536 is the effect of divorce on the slope of the Level-1 equation.

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Table 1. Means and Standard Deviations

	<u>Mean</u>	<u>s.d.</u>
Emotional problems at 7	0.0	1.0
Emotional problems at 11	0.0	1.0
Emotional problems at 16	0.0	1.0
Emotional problems at 23	0.0	1.0
Emotional problems at 33	0.0	1.0
Age minus 7	10.4	9.20
Parental divorce between ages 7 and 22	.09	.29
Parental divorce between ages 7 and 10	.02	.15
Parental divorce between ages 11 and 15	.04	.19
Parental divorce between ages 16 and 22	.03	.17
Parental divorce between ages 23 and 33	.02	.14
Gender (1=woman, 0=man)	.49	.50
Economic status at 7	.19	.28
Class background at 7	.28	.27
School achievement at 7	21.1	5.01

n= 11,759

Table 2. Estimated Effects of Individual Characteristics on the Intercept and Slope Parameters of a Linear Growth-curve model of Emotional problems from Age 7 to Age 33

	<u>Specification 1</u>		<u>Specification 2</u>	
	\$	s.e	\$	s.e
<u>Intercept</u>				
Parental divorce between 7 and 22	.114***	.0278	--	--
Parental divorce between 7 and 10	--	--	.0761	.0537
Parental divorce between 11 and 15	--	--	.139***	.0416
Parental divorce between 16 and 22	--	--	.113**	.0472
Parental divorce between 23 and 33	.0663	.0574	.0662	.0574
Gender	-.198***	.0162	-.198***	.0162
Economic status at 7	-.268***	.0488	-.269***	.0489
Class background at 7	.0432	.0497	.0435	.0497
School achievement at 7	-.0172***	.00188	-.0172***	.00188
Constant	.485***	.0397	.485***	.0397
<u>Slope</u>				
Parental divorce between 7 and 22	.00536**	.00176	--	--
Parental divorce between 7 and 10	--	--	.00620†	.00357
Parental divorce between 11 and 15	--	--	.00418	.00265
Parental divorce between 16 and 22	--	--	.00607**	.00285
Parental divorce between 23 and 33	.000959	.00347	.000959	.00347
Gender	.0236***	.00103	.0236***	.00103
Economic status at 7	-.00318	.00313	-.00318	.00313
Class background at 7	-.00415	.000316	-.00414	.00316
School achievement at 7	-.000439***	.000121	-.000439***	.000121
Constant	-.000735	.00256	-.000747	.00256
-2 log-likelihood		136,613.5		136,647.0
n = 11,759				
†=p < .10	*=p < .05	**=p < .01	***=p < .001	

Table 3. Change in Malaise Inventory Scores between Ages 23 and 33 as a Function of Parental Divorce and Other Individual Characteristics (OLS Estimates).

	\$	s.e.
Parental divorce between 7 and 22	.114**	.0414
Parental divorce between 23 and 33	.0260	.0782
Gender	-.240***	.0246
Economic status at 7	.151*	.0749
Class background at 7	.0114	.0752
School achievement at 7	-.000656	.00290
Constant	-.0894	.0617
R <sup>2</sup>	.0124	
n = 9,347		

\* = p<.05  
 \*\* = p<.01  
 \*\*\* = p<.001

Table 4. Fixed-Effects Estimates of the Effects of a Parental Divorce and Other Time-Varying Characteristics on Emotional problems.

	<u>Specification 1</u>		<u>Specification 2</u>	
	\$	s.e.	\$	s.e.
Age	.000248	.000448	-.00854***	.000767
Parents have ever divorced	.0983***	.0265	.0834***	.0263
Interaction: Age and female			.0238***	.000852
Interaction: Age and class background at 7			-.0107***	.00159
n = 11,759				

\*\* = p<.01

\*\*\* = p<.001



Figure 1. Linear Growth-curve model of Emotional problems, Age 7 to 33, by Age at Parental Divorce: 7-22 and 23-33.

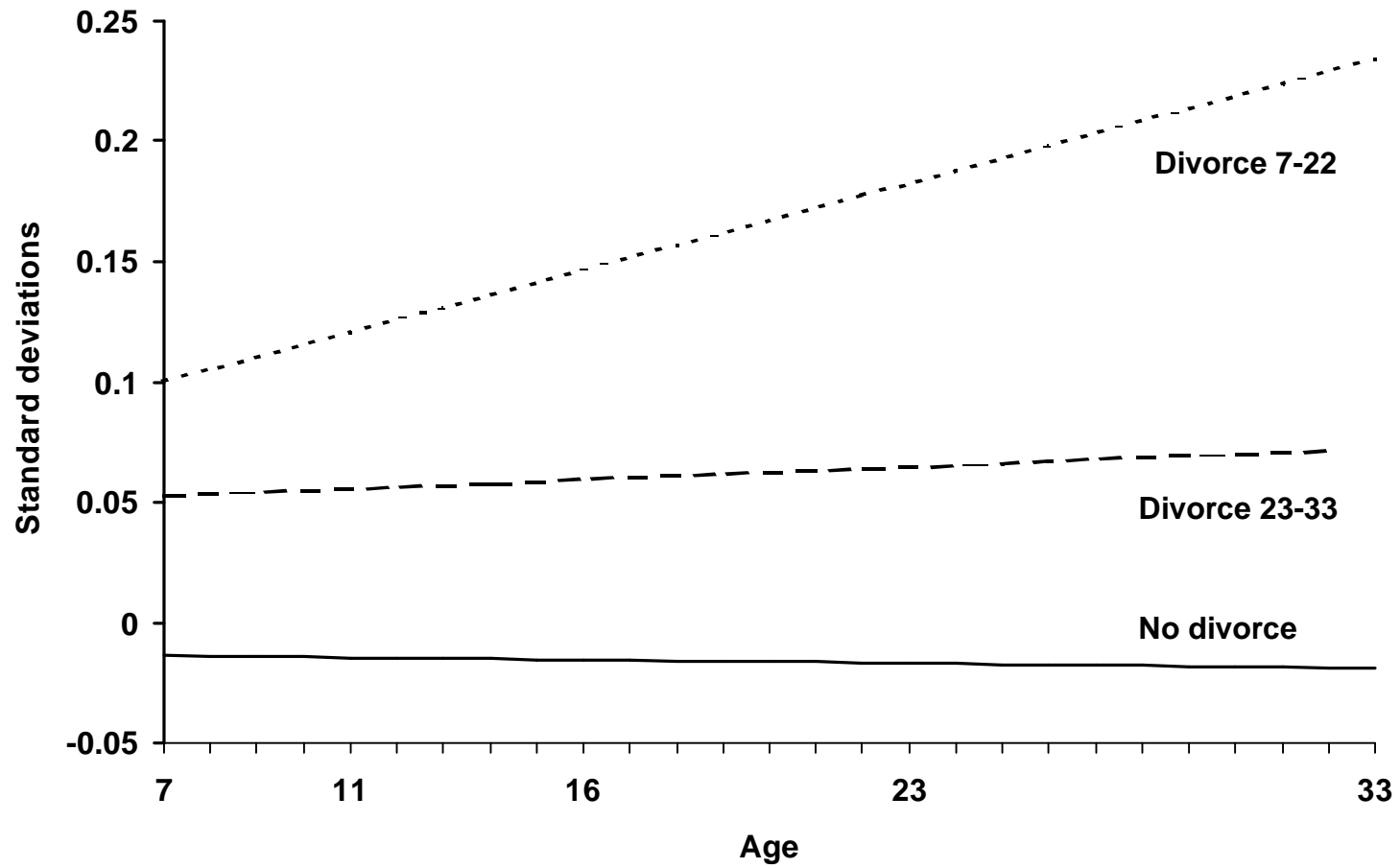


Figure 2. Linear Growth-curve model of Emotional problems, Age 7 to 33, by Age at Parental Divorce: 7-10, 11-15, 16-22, and 23-33.

