The Effect Parental Divorce and Its Timing on Child Educational Attainment: A Dynamic Approach

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Abstract

Using 25 years of data on individuals spanning the entire developmental horizon from birth until adulthood drawn from the Panel Study of Income Dynamics (PSID), this study models educational attainment in a duration framework to examine the long-term effects of parental divorce and its timing on child educational attainment. Controlling for both pre- and post-divorce family characteristics and unobserved heterogeneity, divorce involving young children are found to be most detrimental for child long-term educational attainments, with the effects being slightly larger for males than females. The results also suggest that family disruptions affect the educational attainment of male and female children through different channels. While the father's presence and pecuniary resources partially explain the lower educational attainments of male children of divorced parents, these factors cannot explain the differences in attainments among female children, suggesting that potentially more weight are placed on non-pecuniary family attributes in the development of female children.

Keywords: Parental Divorce, Divorce Timing, Gender, Educational Attainment, Hazard Models

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1 Introduction

The startling growth in divorce rates over the past forty years have received widespread attention and spawned voluminous research on the implications of divorce on children involved,[1] in part because single-parent families typically have fewer resources available to invest in their children's human capital than intact families, and the lack of education has been found to be a strong predictor of welfare recipiency, chronic unemployment, and persistent poverty (Bane and Ellwood, 1983). While there is considerable evidence suggesting that parents' divorce has short-run detrimental effects on children, comparatively little evidence exists concerning the long run effects of marital disruption on children's educational attainments. In particular, few studies consider how the long-term effect of divorce may differ by child's age upon divorce occurrence and by child gender. The sequence and timing of life events, such as family disruptions, may be critical in determining the developmental trajectories for children (Chase-Lansdale, 1998). For instance, the socioeconomic disadvantages as a consequence of divorce combined with age- and gender-specific developmental needs may alter the developmental trajectories of children who experienced divorce at different ages, and eventually lead to differences in their adult outcomes.

This study examines, separately for male and female children, the long-term effects of parental divorce by its timing on child educational attainment. Past research indicates that the effect of divorce on child wellbeing may be attributed to both the antecedent family climate (Cherlin et al., 1991; Fronstin et al., 2001), and socioeconomic consequences of divorce (McLanahan and Sandefur, 1994; Duncan and Hoffman, 1985). Using 25 years of data on individuals spanning the

[1] For example: Allison and Furstenberg, (1989); Amato and Keith, (1991); Amato, (1993); Chase Lansdale et al., (1995); Cherlin et al., (1991); Dawson, (1991); Fergusson et al., (1994); Furstenberg, (1990); Hetherington et al., (1985).

entire developmental horizon from birth until adolescence drawn from the Panel Study of Income Dynamics (PSID), this study models educational attainment in a duration framework incorporating random effects to account for both the sequence and timing of events and the dynamic variations in mediating factors potentially correlated with divorce.

Controlling for both pre- and post- divorce family characteristics and child unobserved heterogeneity, we find that children whose parents divorced when they were young attain significantly less education compared to their intact counterparts, however the effect dissipates over time as the longer passage of time since divorce allows for fuller recovery. The results also show that family disruption affects the educational attainments of male and female children through different channels. While the father's presence and pecuniary resources help explain the differences in educational outcomes between male children from intact and divorced families, these factors cannot explain the outcome differences among female children, suggesting that potentially more weight are placed on non-pecuniary family attributes in the development of female children

2 Background

This section provides the conceptual background and discusses related literature on the effect of divorce on children, with special emphasis on how the effects of divorce on child educational attainment may differ by child's age at separation and gender.

Conceptual Background

We draw on the theoretical literature by Becker (1965, 1981) on household production. Within this framework, the child's educational attainment is viewed as a commodity desired by the household, which is produced by a combination of parental time and material inputs. Financial resources allow parents to purchase goods and services important for child development (Blau, 1999), and are complemented by parenting resources — the services provided by the parents using their time and childrearing ability (McLanahan and Sandefur, 1994). For healthy child development, both time and material resources are needed (Coleman, 1988), and they are substitutable to a certain extent as money can buy childcare services and working in the labor market increases available financial resources.

Time and money operate as constraints for all families, but these resources are typically more limited in single-parent than two-parent families. As a direct consequence of divorce, the economic status of the custodial parent (usually the mother) tends to decline (Hoffman, 1977). Limited family income may affect child educational attainment by reducing financial support for further schooling, fewer out-of-school lessons, or living in less desirable neighborhoods with lower quality schools (Fronstin et al., 2001). In addition, both the custodial and the non-resident parent may spend less time with their children after divorce. Furstenberg et al. (1983) found that divorced fathers are likely to have no contact with their children. The father's absence can also reduce the availability of the mother's time for childrearing. Custodial mothers may need to spread their time and energy beyond their pre-divorce tasks: even though the mother no longer needs to devote time to her husband, her childrearing time may be lowered because she must perform tasks done by fathers in two-parent families. Finally, the custodial parent may spend additional time in the labor market. Although working parents are likely to substitute market-produced goods and services for their own time, single working parents typically have little income to make these substitutions.

We expect divorce to have a negative effect on the educational attainment of children because of this reduction in resources available for human capital development. However, the lack of family resources associated with divorce may affect child educational attainment differently by the timing of divorce occurrence. For example, the loss of income and time resources may have larger effects on young children, because (1) young children tend to rely on their parents as their primary caregivers, whereas teachers and peers share much of the parents' role in producing human capital for older children; and (2) the reduction in these inputs occurs over a longer period of time. However, young children may be more likely to experience a recovery of these resources through the remarriage of the custodial parents (Duncan and Hoffman, 1985; Page and Stevens, 2002).[2] Single parents with young children may have better prospects in the marriage market than single parents with older children because (1) parents with young children also tend to be young, and (2) raising a young child may be perceived by prospective mates as being less challenging and/or burdensome.

Divorce may also subject children to emotional stress that may negatively affect their educational attainment in later years. Emotional stress can be caused directly by the divorce itself, or from lifestyle changes that result from, for instance, residential moves and transfers to new schools. It has been argued that marital dissolution might be more emotionally harmful for younger children than for older children, since younger children might be less resilient to the effects of parental separation due to their relative immature social and cognitive development, and their greater dependence on parents (Allison and Furstenburg, 1989). In contrast, Chase-Lansdale et al. (1995) compared the effects of parental separation between ages 7 and 11 and between ages 11 and 16 on child mental health outcomes in adulthood, and found that later separation to be more harmful. While it remains unclear as to how the emotional stress induced by parents' divorce is

[2] Furstenberg and Cherlin (1991) reported that about half of the children whose parents divorce eventually live with a step-parent.

related to the child's age when the event occurs, a longer passage of time between the event and adulthood may allow for a fuller recovery.

Finally, divorce may affect sons and daughters differently. There is evidence that fathers maintain more contact with sons than daughters (Hetherington et al., 1982) and provide more educational financing for sons than for daughters (Wallerstein and Corbin, 1986). The differences in paternal behavior may be motivated by parental beliefs that boys and girls have different developmental needs. For example, parents may perceive that sons need a male-role model in the home; hence, the emotional stress of losing the resident father may be greater for boys. Given that the same-sex parent may have comparative advantages in meeting children's gender specific needs, father's time may be a more valuable input in the care of boys than girls. If sole paternal custody is rare, we may expect that marital disruption and father absence to be more harmful for boys than girls.

Related Literature

As a direct consequence of the rising divorce rates, more and more children spend part of their lives in single-parent families. Early research typically focuses on the implications of single-parenthood on child outcomes. Single-parent families tend to be economically disadvantaged, invest less time in children, and have fewer and weaker social networks (Duncan and Hoffman, 1985; Manski et al., 1992). Consequently, children raised in such conditions tend to enjoy lower investments in human capital (e.g., Blau and Duncan, 1967; McLanahan, 1983).

Hofferth (1982) pointed out that the variability in age and length of time children spent in single-parent families may be more important than whether the child ever lived in a single-parent family, as children may differ in their developmental needs by age and gender (Jenkins et al., 1989; Fauber et al., 1990; Davies and Cummings, 1998). Krein and Beller (1988) examined the effects of years spent with a single mother during preschool, elementary school, and high school on the child's final years of schooling completed. They found that the effect is the greatest during the preschool years, and is larger for boys than girls. Garasky (1995) extended the study by Krein and Beller by examining the effects of living in different types of non-intact families at four different stages during childhood on the child's propensity to graduate from high school. He found that compared to children living in intact families, children who lived in a father-only or father-stepmother family before age 3 have significantly lower probability of high school completion, while no significant differences are found for children who lived in other family types during this period and thereafter.[3]

A drawback of the studies by Krein and Beller (1988) and Garasky (1995) is that they do not differentiate between the causes of single-parenthood. A child can live in a single-parent (or step-parent) family because of a divorce, death, or a non-marital birth (or a combination of these events). Some studies show that children from single-parent families who experienced their parents' divorce have lower educational attainments than children who suffered the death of a parent (Kiernan, 1992; Corak, 2001; Lang and Zagorsky, 2001; Biblarz and Gottainer, 2000). The differences in child educational outcomes may reflect differences in socioeconomic support systems and parenting behavior between these two experiences, and potential selection effects as single divorced parents may have substantially different characteristics than single-parent widows.

While there is ample evidence showing that divorce has short-run detrimental effects on children during all developmental stages,[4] research examining the effect

[3]A recent study by Ermisch and Francesconi (2001) used the British Household Panel Survey to examined the effect of living in a single-parent family between ages 0 to 5, 6 to 10, and 11 to 16 on the probability that a child completed high school (or above). They found that living in a single-parent family between ages 0 and 5 is most detrimental; however, potential gender differences were not considered.

[4]For instance, Wallerstein (1991) and Cherlin et al. (1991) found that parental divorce during elementary school is associated with poorer test scores and more behavioral problems among

of divorce on child long-term educational outcomes remain scarce. In particular, evidence on how divorce may affect child long-term educational attainments differently by the timing of divorce occurrence and by child gender remain scarce and based non-U.S. data. Fronstin et al. (2001) used longitudinal data on a British cohort to examine the effect of family disruptions on child long-run economic wellbeing by timing of disruption occurrence.[5] They found that a parental divorce occurring before age 16 caused substantial reductions in educational attainment for both boys and girls. While they could not reject that the negative divorce effects are constant across all developmental stages for boys, earlier divorce (between ages 0 and 6) are found to have larger effects on the educational attainment than later divorces among females.

The estimated effect of parental divorce on child educational outcomes may be spurious (Manski et al., 1992; Mayer, 1997), as divorce and children's outcomes may share some unmeasured true causal factor. For instance, parents who are less committed to their family may be more likely to divorce and provide less support for their children: parents dealing with medical and/or behavioral problems may be more likely to divorce and less effective as caretakers (Manski et al., 1992). Evidence that children in to-be-divorced families exhibit lower cognitive achievements and behavioral problems prior to divorce lend some support to this proposition (e.g., Cherlin et al., 1991). However, recent research have found

young children, in particular among males (Doherty and Needle, 1991; Block et al., 1986); however Sun and Li (2001) report no gender differences in the test scores among children from divorced families. Children who experienced divorce during adolescence are often found to have lower educational attainments in adulthood (Corak, 2001; Kiernan, 1992; Painter and Levine, 2000), with larger effects found among female children (Kiernan, 1992; Sun, 2001).

[5] A study using U.S. data is worth noting. Loh (1996) used the NLSY to examine the implications of types of single-parenthood and child's age at disruption on years of schooling completed by age 24. While the findings showed that the effect of divorce is larger for women than men, they do not readily speak of differences in the effects by timing of divorce occurrence separately for men and women. This is because the study estimated the effects of "child's age at disruption" and the effects of "types of family disruption (divorce or death)" on educational attainment in separate regressions, rather than interacting these explanatory variables within a single regression.

that the adverse divorce effects persist even after controlling for pre-divorce conditions and shared family unobservable characteristics between siblings (Painter and Levine, 2000; Ermisch and Francesconi, 2001; Fronstin et al., 2001).

Finally, as discussed previously, divorce timing is closely associated with the timing of family socioeconomic status changes and the length of time children are exposed to potential disadvantages. This combined with variations in child developmental needs over time may imply that the amount of resources available at certain developmental stages may be more important than its total amount over time. However, most studies modeled the relationship between the timing of family structure changes and child educational outcomes in static frameworks. Static models are not able to keep track and control for the sequence and timing of family structure and socioeconomic status changes, and may introduce bias in the effects of divorce timing on child educational attainments.

The present study transcends some of the limitations of previous research by focusing on children who experienced parental divorce while growing up to examine the implications of child's age upon divorce occurrence on the long-term educational achievements of male and female children. Using annual data that began at birth on a cohort of American children, profiles of family structure and socioeconomic status changes over time are constructed to account for both pre- and post-divorce family characteristics potentially confounded with the effect of divorce on child long-term educational outcomes. Child human capital accumulation is modeled as a duration process, in which the propensity to accumulate more education in each period depends on choices made in all previous periods. Within this framework, we are able to keep track of the precise sequence and timing of family structure changes and child schooling completion, and control for dynamic variations in family circumstances potentially correlated with the timing of family structure changes.

3 Model

We formulate a simplified model of child human capital accumulation, in which child quality is measured by his/er educational attainment and produced using parents' choices of inputs. Let $T_{i,t}$ be the years of schooling attained for child i at age t. We assume that parents' investment decisions in a given period are contingent upon all of their past investment and marital decisions. Hence, we conceived of knowledge acquisition as a production process in which current and past inputs ($F_i(t)$) are combined with an individual's genetic endowment and capacity (μ_i), conditional on parents' marital status and marital histories ($M_i(t)$) and subject to idiosyncratic shocks ($\varepsilon_{i,t}$):

$$T_{i,t} = T_t[F_i(t), \mu_i | M_i(t), \varepsilon_{i,t}] \tag{1}$$

The duration framework is well suited to estimate this model as it allows for dynamic variations in input choices and is able to keep track of the events as they unfold. Adopting the terminology of duration models, let the completion of formal schooling be the event "failure" of interest. In this context, T can be interpreted as the "survival" time until schooling completion, or years of schooling attained. The (ln) survival time of schooling completion (or "exit") can be modeled as:

$$\ln T_i = \alpha_0 + \alpha_1 X_i(t) + \alpha_2 M_i + \alpha_3 D_i(t) + \mu_i + \sigma\varepsilon_i \tag{2}$$

The quality and quantity of inputs depend on both the availability of time and material resources, and economies of scale in input production within the family. Since direct inputs choices are unobserved, $X_i(t)$ includes both time-invariant and time-varying characteristics of the family and of each parent to proxy for child investment inputs: parents' education, racial background, household income, child parity, mother's age at marriage and at childbirth, and order of the marriage. We also include controls for child's birth cohort and region of residence (in each period) to account for regional and cohort differences in educational and legal environments.[6]

To capture differences in the effect of divorce timing on child educational outcomes, children from divorced families are stratified into five groups by their age upon divorce occurrence using a vector of indicators M_i: less than 5 years old (early childhood), ages 5 to 11 (late childhood), ages 11 to 13 (pre-teens), ages 13 to 18 (teens), and more than 18 years of age (adulthood) Each coefficient represents the underlying baseline schooling exit hazard of the group relative to that of children in intact families.

The most troubled marriages tend to dissolve sooner, and stable marriages are likely to last longer (Ribar, 2006). Even among children whose parents divorced when they were of a similar age, they may differ in terms of their parents' marriage duration. Furthermore, age at parents' divorce is confounded by the length of time a child is exposed to parental absence: children who are younger upon divorce may be exposed to parental absence for a longer period compared to children who experienced divorce when they are older. Conversely, if divorce is most damaging when the child is young, earlier divorce may also be associated with a longer adjustment period. Hence, to account for these potentially confounding effects, the vector $D_i(t)$ includes controls for parents' marriage duration and a time-varying covariate measuring the length of time since divorce (in years) at each period t.

The parameter, μ_i, captures unobserved child heterogeneity (i.e., unobserved

[6]The study grew up during a period of significant legislative changes in state divorce policies. The rise in divorce rates during the 1970s coincided with a sweeping reform in divorce laws that dramatically increased the access to divorce across the U.S. Controlling for region and cohort differences may account for legislative factors that reduce potential peer pressure and stigma associated with having divorced parents.

genetic endowments and capacity), which is treated as a random effect in the model. Idiosyncratic shocks are captured by $\varepsilon_{i,t}$. The distributional assumptions on the idiosyncratic shocks and details on incorporating random effects (μ_i) into our model are discussed in the following sections.

Hazard Specifications

The hazard of schooling exit, $h_i(t)$, gives the probability that child i will complete t grades of schooling, conditional on s/he remains in school by grade t. To estimate the model as in Eq. 2, we need to impose distributional assumptions on the error term, ε. If we assume that $\ln T_i$ follows a Type II Extreme Value distribution, the duration of schooling T_i would follow the Weibull distribution. For notational convenience, let X_i denote the set of K covariates to be included in our model as discussed in the previous section. The corresponding hazard function is:

$$h_i(t) = exp\{-X_i \frac{\beta}{\sigma}\} \frac{1}{\sigma} t^{\frac{1}{\sigma}-1} \tag{3}$$

where σ and $\{\beta\}$ are parameters to be estimated. We expect that the schooling exit hazard to exhibit positive duration dependence,[7] given that the likelihood of schooling completion (schooling exit) should increase with the grades of education attained. The estimated coefficients of the Weibull hazard function, $\{\beta\}$, can be easily converted back into the coefficients of the covariates in the log survival time model, $\{\alpha\}$ in Eq. 2, since: $\widehat{\alpha_k} = -\widehat{\sigma}\widehat{\beta}_k$. A covariate has a negative effect on child educational attainment if it increases the probability of schooling exit.[8]

To assess the robustness of the Weibull functional form assumption, a more

[7] Positive duration dependence refers to when the hazard increases continuously with time: if $0 < \sigma < 0.5$, the hazard increases at an increasing rate; linear if $\sigma = 0.5$; and at a decreasing rate if $0.5 < \sigma < 1$.

[8] If a factor increases the hazard rate, then the event occurs quickly and the survival time until when failure occurs (i.e., duration) would be short.

flexible distribution is also used: the piece-wise exponential hazard.[9] Assuming that $\ln T_i$ follows a Type I Extreme Value distribution, the duration of schooling, T_i, will follow an exponential distribution. The piece-wise exponential hazard is more flexible than the Weibull hazard since the data is used to reveal the baseline hazard by arbitrarily dividing the span of educational advancement into intervals (e.g., grade levels).[10] The piece-wise exponential hazard function is given by:

$$h_i(t) = \begin{cases} \lambda_1 e^{X_i(t)\beta}, & \tau_0 \leq t < \tau_1; \\ \lambda_2 e^{X_i(t)\beta}, & \tau_1 \leq t < \tau_2; \\ \vdots & \\ \lambda_M e^{X_i(t)\beta}, & \tau_{M-1} \leq t < \infty \end{cases}$$

where $\{\tau_m\}$ are $M-1$ known constants, representing the cut-off points for each interval. The baseline hazard, $\{\lambda_m\}$, are M unknown parameters, together with $\{\beta\}$, are to be estimated. The coefficients of the exponential hazard model can be easily interpreted, since $\widehat{\alpha_k} = -\widehat{\beta_k}$.

Incorporating Random Effects

There may be omitted factors that influence child educational outcomes independent of controls included in the model. For instance, children may differ in their innate cognitive ability, emotional maturity, and genetic endowments (Todd and Wolpin, 2003). Hence, to obtain more precise estimates, we incorporate random effects into our hazard model to account for these unobserved factors.[11] Following Gutierrez (2002), let $h_{ij}(t)$ denote the hazard function without random effects.

[9] For a more detailed discussion, see Meyer (1990).

[10] Although we have information on the timing of schooling completion, we are unable to identify the exact timing of the start of formal schooling. Therefore, we allow the duration process to start at birth.

[11] Hazard models with random effects are commonly known as the parametric shared frailty survival models. For a detailed discussion, see Firth and Payne (1999).

Unobserved factors, denoted by μ_i (as in Eq. 2), is taken as an individual-specific random effect assumed to be uncorrelated with the other covariates. For notational convenience, let $\phi_i = e^{\mu_i}$, and assume ϕ_i follows a Gamma distribution with mean one and variance θ.[12] Children with $\phi_i > 1$ are said to be more frail, or at higher risk of dropping out of school for reasons left unexplained.

Let n denote the total number of individuals with individual i being observed for n_i periods. The hazard function conditional on the individual-specific random effect is:

$$h_{ij}(t|\phi_i) = \phi_i h_{ij}(t) \tag{4}$$

for $j = 1,...,n_i$ with $h_{ij}(t) = h(t|X_{ij})$. For any period in which individual i remains in school, the standard hazard function is now multiplied by ϕ_i. The Weibull hazard function is a good alternative when incorporating random effects, due to the fewer number of parameters to be estimated. However, it can only be used if it appropriately describes our data.[13] Under the Weibull specification, the conditional hazard and survival functions for individual i at period j are given as:

$$\begin{aligned} h_{ij}(t|\phi_i) &= \phi_i \exp\{X_{ij}\frac{\beta}{\sigma}\}\frac{1}{\sigma}t^{\frac{1}{\sigma}-1} \\ S_{ij}(t|\phi_i) &= \{S_{ij}(t)\}^{\phi_i} = \exp\{-\phi_i \exp(X_{ij}\frac{\beta}{\sigma})t^{\frac{1}{\sigma}}\} \end{aligned}$$

Let $(t_{0ij}, t_{ij}, d_{ij})$ indicate the start time, end time, and whether individual i experiences a failure or censoring at period j. Given ϕ_i, the contribution of individual i to the likelihood is:

$$L_{ij}(\phi_i) = \frac{S_{ij}(t_{ij}|\phi_i)}{S_{ij}(t_{0ij}|\phi_i)}\{h_{ij(t_{ij}|\phi_i)}\}^{d_{ij}} = \{\frac{S_{ij}(t_{ij})}{S_{ij}(t_{0ij})}\}^{\phi_i}\{\phi_i h_{ij}(t_{ij})\}^{d_{ij}}$$

[12] The choice of distribution is arbitrary. In theory, any continuous distribution supported on positive numbers that has expectation of one and finite variance θ is allowed.

[13] We access the appropriateness of choosing the Weibull hazard function in the results section and in the Appendix.

and if we define $D_i = \sum_{j=1}^{n_i} d_{ij}$, the conditional likelihood of the i^{th} individual is:

$$L_i(\phi_i) = \phi_i^{D_i} \prod_{j=1}^{n_i} \{\frac{S_{ij}(t_{ij})}{S_{ij}(t_{0ij})}\}^{\phi_i} \{h_{ij}(t_{ij})\}^{d_{ij}}$$

To derive the unconditional likelihood function, we need to integrate out ϕ_i,

$$L_i = \int_0^{\infty} L_i(\phi_i) g(\phi_i) d\phi_i$$

where

$$g(\phi_i) = \frac{\phi_i^{\frac{1}{\theta}-1} \exp(-\frac{\phi_i}{\theta})}{\Gamma(\frac{1}{\theta})\theta^{\frac{1}{\theta}}}$$

The unconditional individual likelihood functions can be expressed as:

$$L_i = [\prod_{j=1}^{n_i} \{h_{ij}(t_{ij})\}^{d_{ij}}] \frac{\Gamma(\frac{1}{\theta} + D_i)}{\Gamma(\frac{1}{\theta})} \theta^{D_i} \{1 - \theta \sum_{j=1}^{n_i} \ln \frac{S_{ij}(t_{ij})}{S_{ij}(t_0 ij)}\}^{\frac{1}{\theta} - D_i} \tag{5}$$

Given the unconditional individual likelihoods, we can estimate the parameters and the variance of the random effect by maximizing the overall log likelihood $\ln L = \sum_{i=1}^{n} \ln L_i$.

4 Data and Sample Descriptives

This study utilizes data from the first 25 waves of the Panel Study of Income Dynamics (PSID), from 1968 to 1992.[14] The PSID began interviewing a cross-sectional sample of $5,500$ households across the United States in 1968. Follow-ups are conducted annually and are still on-going. The primary goal of the PSID is to collect information on short-run economic status changes of families and indi-

[14]The PSID is sponsored and distributed by the Inter-University Consortium for Political and Social research, University of Michigan, Ann Arbor, MI.

viduals. Over time, the study expanded to include detail demographic information as well.

We select a sample of individuals who are (1) part of the original core sample of households interviewed in 1968; (2) born between 1968 and 1972; and (3) have completed all formal education by 1992. Hence, for each individual in our sample, we observe his/er family socioeconomic information annually from birth until the completion of his/er schooling.[15] In 1985, the PSID started collecting retrospective childbirth, adoption, and marriage histories of individuals in the sample households.[16] We use the *Childbirth and Adoption History Supplemental Survey*, to identify the birth parent(s) of each sample individual. Individuals for whom at least one birth parent can be identified are included in our sample. In cases where only one birth parent is identified, we assume that the spouse of the identified parent residing in the same household in the year that the individual was born is the other birth parent. Furthermore, marriage history information of the birth parents are obtained using the *Marriage History Supplementary File* in order to pinpoint the timing of parental divorce. The final sample includes 901 individuals.

Description of Study Sample

Table 1 presents the sample means for our pooled study sample.[17] The average child in our sample has completed 12 grades of formal schooling (i.e., a high school graduate). Approximately 34% of our sample experienced parental divorce

[15] A shortcoming in the PSID is that information on "grades of schooling attained" for each individual is not available annually; rather, this information is reported only at the year in which the individual has completed all formal education (i.e., final year in school). As such, we have to abstract away from potential grade retention and assume that the individual has been in school and advanced continuously. If children from divorced families are more likely to be retained, our estimates would be a "lower-bound" of the true effect of divorce.

[16] Individuals who have attrited prior to 1985, or were institutionalized in 1985, were not included.

[17] For details of the variable definitions, see Appendix Table 1.

while growing up.[18] Sub-sample means for children of divorced parents and children of married parents are reported in Table 2. Children whose parents divorced attain fewer years of schooling, on average, compared to their counterparts whose parents stay married (about $1/2$ a grade level).

Among children from divorced families, the average age at divorce is approximately 9, with a majority (66%) experiencing parental divorce before they reached age 11 (see Figure 1). However, as previously discussed, marriage duration may be correlated with marital stability/quality, since the most troubled marriages tend to dissolve quickly and stable marriages are likely to last longer. Note that even among couples whose children are of a similar age upon divorce, the length of time they have been married may vary substantially. Figure 2 shows the distribution of marriage durations of our divorced sub-sample. As shown, marriage duration of varying lengths are represented, with the average marriage among couples who eventually divorced lasting about 14 years.

Finally, consistent with existing evidence, experiencing parental divorce is associated with economic disadvantages, given that the average annual family income of a divorced family is approximately 83% of families that stayed intact (Table 2). However, there are significant variations of family income over time, which appears to be closely linked to the timing of divorce. Figure 3 presents the comparisons of income profiles over-time between divorced and intact families, separately by the child's age upon divorce: less than 5 years of age (top), between 6 and 12 (middle), and between 13 and 18 (bottom). While family income of divorced families appear to be at par with that of intact families prior to divorce, significant reduction in income is observed starting around the time of divorce. At last, there appear to be some evidence of a post-divorce recovery of income among

[18] In line with our estimates, Wojtkiewicz (1992) estimated that approximately 37.9% and 24.9% of black and white children born in the 1960s would experience parental divorce, respectively.

families who divorced when the child is between ages 6 and 11. However, no evidence of income recovery are found for families that divorced when the child is very young (between ages 0 and 5) or when the child is a teenager (between ages 13 and 18).

Kaplan-Meier Survival Estimates

The Kaplan-Meier estimator is a non-parametric maximum likelihood estimate of the underlying survival function (Kaplan and Meier, 1958), which can be used to illustrate the relationship between an individual's grades of schooling attained and the timing of parents' divorce.[19] Figure 4 presents the survival probabilities in schooling attainment by family structure for the male and female sub-samples. Both male and female children from divorced families have significantly lower probability of completing each grade of schooling relative to their counterparts with married parents.[20] However, the difference in the probabilities of attaining each grade of schooling between children of intact and divorced families are more pronounced among males than females. Figure 5 further stratifies the samples by divorce timing.[21] Relative to children of married parents, both male and female children whose parents divorced before they reached age 5 (as oppose to during the teenage years) have significantly lower probabilities of attaining each grade of schooling.[22]

[19] Let the number of individuals who exited after completing t years of schooling be denoted d_t, and n_t denotes the number of individuals who are still in school after completing t years of schooling. The Kaplan-Meier survival function gives the probability that the final years of schooling in our sample is at least t: $\widehat{S(t)} = \prod_{j|t_j \leq t} \frac{n_j - d_j}{n_j}$.

[20] For both the male and the female sub-samples, log rank test of equality of survival functions between children from intact *vs.* divorced families are rejected at the 5% level.

[21] For illustrative clarity, we only present the survival probabilities for children whose parents divorced during preschool (ages 0 $\sim$ 5) and during teenage years (ages 13 $\sim$ 18). Survival probabilities for children whose parents divorced during late childhood and pre-teen years are available upon request.

[22] For both the male and female sub-samples, equality of the survival probabilities between "children of intact families" and "children whose parents divorced before age five" are rejected

5 Main Findings

The non-parametric survival estimates shown in the previous section point to significant variation in the effect of parents' divorce on child educational attainment by child's age upon divorce occurrence and gender. However, the Kaplan-Meier estimator does not accommodate multivariate analysis. In this section, we present estimation results of the duration model of child schooling attainments introduced in Section 3.

Duration Dependence

We begin by examining the underlying duration dependence of the schooling exit hazard. As discussed previously, the Weibull model is a good alternative when incorporating random effects in a hazard model, due to the fewer number of parameters to be estimated. However, it should be used only if it appropriately describes the data. Hence, before incorporating random effects into our model, we first compare the estimation results of the piece-wise exponential hazard model to that of the Weibull model, to determine whether the Weibull model fits the data well.

Table 3 presents the estimation results for the pooled sample (without random effects). The estimated duration dependence parameter for the Weibull model, σ, is 0.072, implying that the schooling exit hazard increases at an increasing rate (i.e., the likelihood of schooling completion/exit increases with the grades of schooling attained). The positive duration dependence is confirmed by the piece-wise exponential model, as revealed by data through a set of indicators representing the educational span. Finally, the estimated log likelihood is higher for

at the 5% level via the log rank test. Equality of survival functions between "children of intact families" and "children whose parents divorced during their teenage years," however, cannot be rejected.

the Weibull model relative to that of the exponential model, suggesting that the Weibull model fits the data better.

Timing of Divorce

The reported coefficients in Table 3 represent the effect of each covariate on the predicted log of schooling duration (i.e., grades of schooling attained). Interpretation of these coefficients deserves some attention. The coefficient on "female" is 0.010 in the Weibull model, implying that the predicted grades of schooling (in logs) for a female is 0.01 longer than that of a male with the same characteristics. Alternatively, time ratios can be calculated by exponentiating the coefficients. In this case $e^{0.010} = 1.010$, implying that females on average will attain 1% more grades of education than their male counterparts. Controlling for gender, parental characteristics and socioeconomic background, we find that children whose parents divorced when they were very young complete significantly fewer grades of education: an individual whose parents divorced before s/he reached age five attains 3.25% fewer grades of schooling relative to individuals with married parents. While lower educational attainments are also found for children whose parents divorced when they were more than five years old, the differences relative to intact children are smaller in magnitude and statistically insignificant.

Gender Differences

The results base on the pooled sample point to potential differences in the effect of divorce by child's age upon divorce. In addition, we find a small but statistically significant difference in educational attainment by child gender. However, the analysis thus far has not considered how divorce timing may affect male and female children's educational attainments differently, and potential unobserved heterogeneity. Table 4 reports the estimated effects of divorce timing on the educational attainments for the male and female sub-samples using the Weibull model with random effects.[23] Three model specifications are adopted: Model 1 controls for basic demographic characteristics; Model 2 adds controls for family resources, proxied by parental characteristics and family income; and Model 3 additionally accounts for factors potentially confounded with the timing of divorce occurrence.

There is evidence of significant unobserved heterogeneity within our sample, as the frailty term (θ) is significant across all model specifications. Controlling for basic demographic characteristics (Model 1), we find that while parental divorce is generally associated with lower educational attainments in children, the association is only statistically significant for children whose parents divorced during childhood, with the effect being slightly larger for males than females. Male children whose parents divorced before they reach age 11 (similarly for female children who experienced parental divorce before age 5) attain 2.2% $\sim$ 3.7% fewer grades of schooling (3.1% for females) relative to their intact counterparts.

Consistent with previous findings, parents' characteristics and family income (both pre- and post-divorce) partially explains the lower educational attainments among males from divorced families, but the same pattern is not found for females (Model 2). Furthermore, while the mother's education is important for the development of both sons and daughters, father's education is found to be significant only for the educational attainment of sons. Taken together, these results are consistent with the idea that divorce affects boys and girls through different channels, in particular girls may be affected more by non-pecuniary attributes of the family (which we cannot account for). In addition, these findings are in line with

[23]The Weibull hazard model with random effects is our preferred specification (relative to the piece-wise exponential) since the Weibull distribution fits our data well (see discussion in the previous section). We also test between alternative distributional assumptions, which reveal that the Weibull distribution is the most preferred and best fitting model. The details of this test are discussed in more detail in the Appendix.

the idea that paternal involvement affects sons and daughters differently. Mammen (2003) and Lundberg et al. (2007) both find that fathers spend significantly more time with sons relative to daughters. If the utility each parent derives from children (and hence their incentive to invest in children) is contingent upon custody/coresidence arrangements (Lundberg, 2005), and that fathers are more likely to obtain shared custody of sons (Dahl and Moretti, 2004; Cancian and Meyer, 1998), fathers may be less involved and hence his inputs less important in the development of his daughters.[24]

Next, we account for parents' marriage duration[25] and a time-varying continuous variable for the number of years since parents' divorce (Model 3). Consistent with evidence that paternal presence benefits male children's development, we find that parents' staying married longer has a positive effect on the long-term educational attainments of sons: holding the timing of divorce constant, each additional year in which the parents stay married increases a male child's education by 0.4%. The benefit of having the parents' stay married longer (or posting divorce), however, is small and insignificant for female children. Amato et al. (1995), Amato and Booth (1997), Hanson (1999), and Jekielek (1998) found that off-springs were better off on a variety of outcomes if parents in high conflict marriages divorced than if they remained married. Given that girls tend to be raised to care take and safeguard family relationships (Buehler and Gerard, 2002), prolonged exposure to parental hostility because of postponing divorce may adversely affect daughters more. Our finding is in line with evidence that divorce have positive consequences for some children, especially daughters. For example, a qualitative study by Arditti (1999) found that many offspring from divorced families,

[24]For a detailed discussion on evidence that daughters receive less paternal time both within and outside of marriage, see Lundberg (2005).

[25]This variable is coded "0" for all children raised in intact families, and "≥ 0" for children of divorced parents.

especially daughters, reported developing especially close relationships with their custodial mothers.[26]

Finally, while children whose parents divorced when they were younger may be exposed to prolonged period of socioeconomic disadvantages associated with divorce, a longer passage of time between the event and adulthood may also allow for a fuller recovery. Controlling for dynamic variations in family income over time (both pre- and post divorce), Model 3 also includes a time-varying control for the "number of years since parents' divorce". While the pattern of the divorce timing effects remains consistent with those shown in Models 1 and 2, the estimates in Model 3 are significantly larger in magnitude. This is as expected as the number of years since parents' divorce is both negatively correlated with child's age at divorce occurrence, and positively associated with child educational attainment. Hence the adverse divorce timing effects would be biased toward zero if we do not control for time since parental divorce. Controlling for dynamic variations in family socioeconomic status and divorce timing, we find that each additional year after divorce is associated with an increase in child educational attainment by 1.8% and 1.5% for males and females, respectively.

Overall, our results show that divorce involving young children is most detrimental for child long-term educational attainments, with the effect being slightly larger for males than females. Although earlier divorce (in terms of child's age at divorce) is most harmful, we find some evidence that the longer passage of time since divorce is potentially beneficial. Taken together, the difference in educational attainments between children from intact and divorced families appears to be modest. Finally, we find some evidence that family disruption affects the educational attainments of male and female children through different channels: while the father's presence and pecuniary resources partially explain the differ-

[26] This is consistent with some qualitative studies, for example, Amato and Booth (1997).

ences in educational outcomes between male children from intact and divorced families; these factors cannot explain the differences in attainments found among female children, which may suggest more weight being placed on non-pecuniary family attributes in the development of female children.

6 Discussion and Conclusions

This study examines the long-term effect of parental divorce on children's educational attainments. While ample evidence exists showing that divorce has short-term adverse effects on children, the long-term implications of divorce on child educational outcomes have been documented only by a few studies using British data. Using longitudinal data on a representative cohort of U.S. children, this study examines, separately for males and females, the differential effects of divorce on child educational outcomes by child's age upon divorce. Unlike many studies, we have been able to construct family socioeconomic profiles over time and control for dynamic variations in socioeconomic characteristics correlated with family structure changes, as we have annual information on these children and their families since childbirth until adulthood.

We estimate a duration model of child schooling attainment to determine how the effect of parental divorce may differ by child's age at family disruption. Consistent with Fronstin et al. (2001), we find that parental divorce has a strong effect on educational attainment of children who experienced divorce during childhood. Controlling for both pre- and post divorce socioeconomic circumstances and child unobserved heterogeneity, male children whose parents divorced before age 11 complete 2.2% to 3.7% fewer grades of schooling, and female children whose parents divorced before age 5 attain 3.1% fewer grades of education, compared to their respective intact counterparts. The long-term educational outcomes of children who experienced divorce at older ages (both males and females), however, do not appear to differ significantly from those raised in intact families. These results are in line with children being needier of parental care and more fragile to disruptions when young (e.g., Allison and Furstenberg, 1989; Haveman et al., 1991). Although divorce when children are young is most harmful, our results also indicate that the longer passage of time since divorce can be beneficial. Taken together, the difference in educational attainments between children whose parents divorced when they were young and children raised in intact families in the long run appears to be modest.

Finally, we find some evidence that divorce affects the educational attainments of male and female children through different channels. We find that the father's presence and pecuniary resources partially explain the differences in educational outcomes between male children from intact and divorced families; these factors cannot explain the differences in attainments found among female children. In addition, sons tend to benefit more from having their parents stay married longer. Our findings are in line with evidence of fathers investing more resources toward their sons (Mammen, 2003; Lundberg et al., 2007) and paternal presence is particularly important for male children's development (Harris and Morgan, 1991); and that daughters may be more affected by non-pecuniary attributes of the family, hence prolonged exposure to potential family conflict by delaying divorce may put daughters in greater jeopardy (Buehler and Gerard, 2002). Given that we cannot account for non-pecuniary family attributes and inputs (for example, time spent with children and marital conflict), and that paternal absence is more likely (since maternal custody is more common), these may help explain the estimated effect of divorce being larger for males than for females.

Our findings have important implications for the growing number of divorces and protecting children involved. As indicated earlier, single-parent families make up an increasing fraction of the poverty population. Educational advancement is one way out of poverty, but children from divorced families achieve less of it. Thus this lack of education may lead to long-term economic hardship for the future families of these children. Policies deigned to break this tendency can help counteract this cycle of poverty. Given that the preschool years seem to matter the most, providing additional resources for children through stronger enforcements of child support and paternal involvement, in particular for divorced families with young male children, may help counteract this tendency towards less education.

We note that the differences in educational attainment found here are robust to controls for dynamic variations in individuals' observable family socioeconomic characteristics from birth to adulthood and omitted variables uncorrelated with these controls. In addition to affecting the availability of economic resources, divorce may influence child scholastic development through other mechanisms not accounted for here, such as the reduction in parental time spent with children and changes in parenting practices after divorce. While differences in socioeconomic characteristics correlated with divorce help explain the lower attainments found for male children, they could not explain the poorer educational outcomes found for female children. Further investigation using better data or a structural framework might be better able to unravel the routes through which parental divorce affect the economic wellbeing of their daughters in adulthood.

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TABLE 1
Sample Means: Pooled Sample

	N	Mean	[S. D.]	Min	Max
Dependent Variable					
Grades of schooling attained	901	12.39	[1.516]	7	17
Family Structure Variables					
Parents divorced	901	0.341	[0.474]	0	1
Age at parental divorce:					
- $0 \leq$ Age < 5	901	0.094	[0.292]	0	1
- $5 \leq$ Age < 11	901	0.131	[0.338]	0	1
- $11 \leq$ Age < 13	901	0.034	[0.182]	0	1
- $13 \leq$ Age < 18	901	0.039	[0.193]	0	1
- Age ≥ 18	901	0.042	[0.201]	0	1
Explanatory Variables					
Female	901	0.471	[0.499]	0	1
Race:					
- White	901	0.697	[0.460]	0	1
- Black	901	0.270	[0.444]	0	1
- Other	901	0.033	[0.180]	0	1
Mother's education	901	12.52	[2.087]	0	18
Father's education[2]	784	12.59	[2.904]	0	19
Father's information missing	901	0.130	[0.336]	0	1
Family income[3]	901	$31,937$	$[20,712]$	$4,624$	$335,352$
Mother's age at childbirth	901	24.98	[5.723]	14	44
Mother's age at marriage	901	20.09	[3.239]	11	38
Order of the marriage[1]	901	1.049	[0.226]	1	3
Parity	901	2.072	[2.058]	0	11
Urban residence[4]	901	0.521	[0.500]	0	1
Region of residence:[5]					
- North East	901	0.165	[0.372]	0	1
- North Central	901	0.280	[0.449]	0	1
- South	901	0.401	[0.490]	0	1
- West	901	0.154	[0.361]	0	1
Birth cohort:[6]					
- 1968	901	0.189	[0.391]	0	1
- 1969	901	0.204	[0.403]	0	1
- 1970	901	0.203	[0.403]	0	1
- 1971	901	0.190	[0.392]	0	1
- 1972	901	0.214	[0.410]	0	1

Notes: 1. The order of the marriage is defined with respect to the mother; 2. Information on the biological father is missing for a sub-sample of children ($n = 117$). The mean here is computed for the group of children for whom we observe father's information; 3. Family income in each year is measured in 1984 constant dollars; 4. The location of residence is defined as ”urban” if the population of the largest city in the county of residence exceeds $100,000$; 5. Constructed using PSID regional codes; 6. Corresponds to the “birth year”.

TABLE 2
Sample Characteristics: Children of Intact vs. Divorced Families

	Intact		**Divorced**	
	Mean	[S.D.]	Mean	[S.D.]
Dependent Variable				
Grades of schooling attained	12.54	[1.471]	12.09	[1.560]
Explanatory Variables				
Female	0.473	[0.500]	0.466	[0.500]
Parity	2.120	[2.156]	1.980	[1.854]
Race:				
- White	0.699	[0.459]	0.694	[0.462]
- Black	0.266	[0.442]	0.277	[0.448]
- Other	0.035	[0.185]	0.029	[0.169]
Mother's education	12.44	[2.046]	12.67	[2.160]
Father's education[2]	12.67	[2.807]	12.32	[3.177]
Father's information missing	0.003	[0.058]	0.375	[0.485]
Family income[3]	$33,921$	[$22,879$]	$28,097$	[$14,995$]
Mother's age at childbirth	25.91	[6.051]	23.19	[4.525]
Mother's age at marriage	20.49	[3.371]	19.32	[2.816]
Order of the Marriage[1]	1.051	[0.234]	1.046	[0.209]
Urban residence[4]	0.525	[0.500]	0.515	[0.501]
Region of residence:[5]				
- North East	0.189	[0.391]	0.121	[0.326]
- North Central	0.268	[0.443]	0.303	[0.460]
- South	0.397	[0.490]	0.407	[0.492]
- West	0.146	[0.354]	0.169	[0.376]
Birth cohort:[6]				
- 1968	0.195	[0.397]	0.176	[0.381]
- 1969	0.192	[0.394]	0.228	[0.420]
- 1970	0.189	[0.391]	0.231	[0.422]
- 1971	0.195	[0.397]	0.179	[0.384]
- 1972	0.229	[0.421]	0.186	[0.389]
Age at parental divorce:				
- $0 \leq$ Age < 5			0.277	[0.448]
- $5 \leq$ Age < 11			0.384	[0.487]
- $11 \leq$ Age < 13			0.101	[0.302]
- $13 \leq$ Age < 18			0.114	[0.318]
- Age ≥ 18			0.124	[0.330]
Age at Parental divorce (mean)			8.954	[5.832]
Duration of parents' marriage			13.83	[7.078]
N	594		307	

Notes: 1. The order of the marriage is defined with respect to the mother; 2. Information on the biological father is missing for a sub-sample of children ($n = 117$). The mean here is computed for the group of children for whom we observe father's information; 3. Family income in each year is measured in 1984 constant dollars; 4. The location of residence is defined as "urban" if the population of the largest city in the county of residence exceeds $100,000$; 5. Constructed using PSID regional codes; 6. Corresponds to the "birth year".

TABLE 3
Effect of Divorce on ln(T = Grades of Education Attained): Pooled Sample

	Weibull	Piece-Wise Exponential
Age at Parental Divorce[a]		
$0 \leq$ Age < 5	-0.032^*	-0.295^*
	[0.012]	[0.106]
$5 \leq$ Age < 11	-0.014	-0.165
	[0.015]	[0.091]
$11 \leq$ Age < 13	-0.012	-0.120
	[0.015]	[0.150]
$13 \leq$ Age < 18	-0.018	-0.180
	[0.015]	[0.123]
Age ≥ 18	-0.017	-0.124
	[0.015]	[0.136]
Duration Dependence		
σ	0.072^*	
	[0.002]	
Birth to Grade 7		-20.92^*
		[0.113]
Grade 8		7.290^*
		[1.018]
Grade 9		5.896^*
		[0.499]
Grade 10		3.975^*
		[0.216]
Grade 11		3.272^*
		[0.165]
Grade 12		2.771^*
		[0.145]
College (Freshman)		0.973^*
		[0.107]
College (Sophomore)		1.099^*
		[0.117]
College (Junior)		0.718^*
		[0.115]
College (Senior)		0.861^*
		[0.155]
Beyond college		0.182^*
		[0.118]
Female	0.010^*	0.125^*
	[0.005]	[0.106]
Constant	2.796^*	-2.001^*
	[0.035]	[0.350]
Log Likelihood	959.86	808.30

Notes: 1. Robust standard errors reported in brackets; 2. Significance level reported: $^* = 5\%$; 3. Additional controls included: ln(Real Family Income), mother and father's levels of education, child parity, mother's age at childbirth, mother's age at marriage, order of the marriage, race, birth cohort, urban residence, and region of residence; *a*. Reference group = "Children in intact families."

TABLE 4

Estimated Effect of Divorce on ln(T = Grades of Education Attained): Male and Female Subsamples

	Male			Female		
	(1)	(2)	(3)	(1)	(2)	(3)
Age at Parents' Divorce[a]						
$0 \leq$ Age < 5	−0.037*	−0.028*	−0.301*	−0.031*	−0.039*	−0.255*
	[0.013]	[0.014]	[0.039]	[0.013]	[0.014]	[0.047]
$5 \leq$ Age < 11	−0.022*	−0.016	−0.226*	−0.017	−0.023*	−0.167*
	[0.010]	[0.011]	[0.033]	[0.012]	[0.012]	[0.038]
$11 \leq$ Age < 13	−0.003	−0.003	−0.158*	−0.027	−0.038	−0.123*
	[0.019]	[0.017]	[0.032]	[0.027]	[0.024]	[0.037]
$13 \leq$ Age < 18	−0.016	−0.029*	−0.147*	−0.014	−0.022	−0.071*
	[0.017]	[0.016]	[0.034]	[0.016]	[0.014]	[0.035]
Age ≥ 18	−0.015	−0.015	−0.015	−0.003	−0.003	−0.007
	[0.015]	[0.014]	[0.013]	[0.026]	[0.023]	[0.020]
Controls:						
Mother's education		0.005*	0.005*		0.004*	0.004*
		[0.002]	[0.002]		[0.002]	[0.002]
Father's education		0.005*	0.004*		0.002	0.002
		[0.001]	[0.001]		[0.002]	[0.002]
ln(Family Income)		0.007*	0.006*		0.004	0.003
		[0.002]	[0.002]		[0.002]	[0.002]
Parity		−0.012*	−0.011*		−0.019*	−0.016*
		[0.003]	[0.002]		[0.003]	[0.003]
Mother's age at marriage		0.001	0.002*		0.002*	0.002*
		[0.001]	[0.001]		[0.001]	[0.001]
Mother's age at childbirth		0.001	−0.001		−0.001	−0.001
		[0.001]	[0.001]		[0.001]	[0.001]
Order of the marriage		−0.031*	−0.029*		−0.007	−0.009
		[0.018]	[0.016]		[0.018]	[0.016]

(Continued)

TABLE 4

Estimated Effect of Divorce on ln(T = Grades of Education Attained): Male and Female Subsamples

	Male			**Female**		
	(1)	(2)	(3)	(1)	(2)	(3)
Duration of parents' marriage			0.004*			0.001
			[0.001]			[0.002]
Years since parents' divorce			0.018*			0.015*
			[0.002]			[0.003]
Constant	2.873*	2.678*	2.704*	2.895*	2.793*	2.795*
	[0.012]	[0.041]	[0.013]	[0.015]	[0.045]	[0.042]
σ	0.042*	0.039*	0.033*	0.043*	0.037*	0.031*
	[0.003]	[0.003]	[0.002]	[0.004]	[0.003]	[0.003]
θ	1.106*	1.057*	1.236*	1.195*	1.266*	1.554*
	[0.167]	[0.174]	[0.198]	[0.207]	[0.251]	[0.301]
Log Likelihood	526.72	568.34	588.16	453.00	488.72	499.00

Notes: 1. Robust standard errors reported in brackets; 2. Significance level reported: * = 5% level; 3. Additional controls included in all models (not reported): region of residence, urban residence, race, birth cohort, and whether father's information is missing; *a*. Reference group = "Children in intact families".

FIGURE 1:
Distribution of Age at Parental Divorce

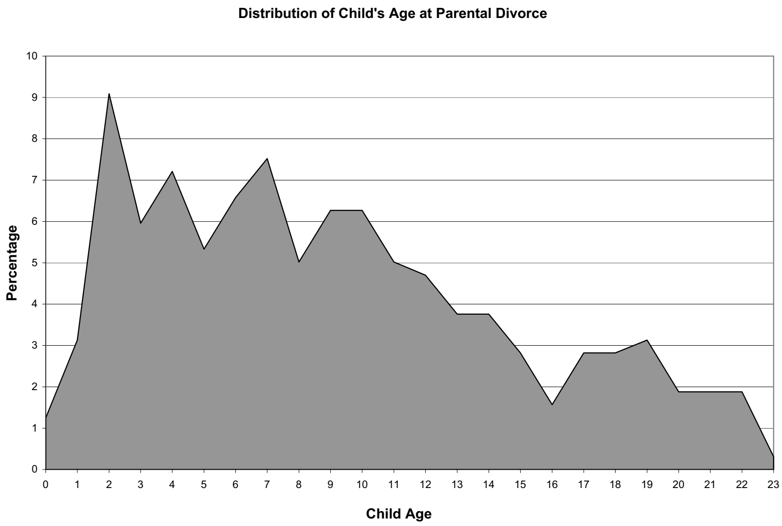


FIGURE 2:
Distribution of Marriage Duration among Divorced Parents

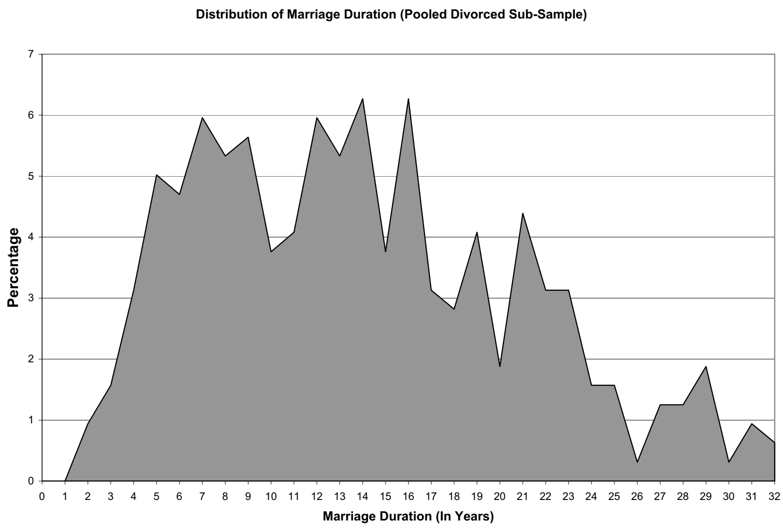
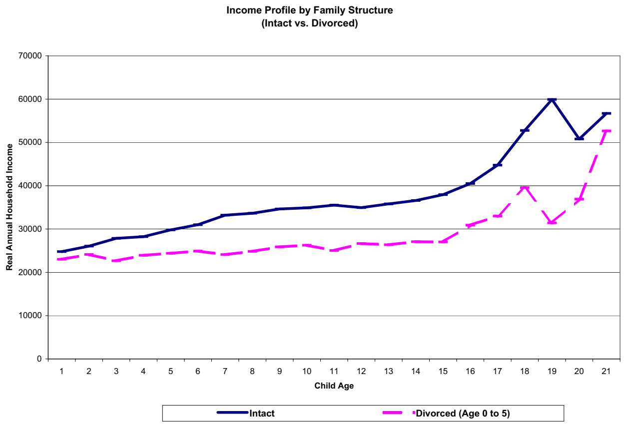
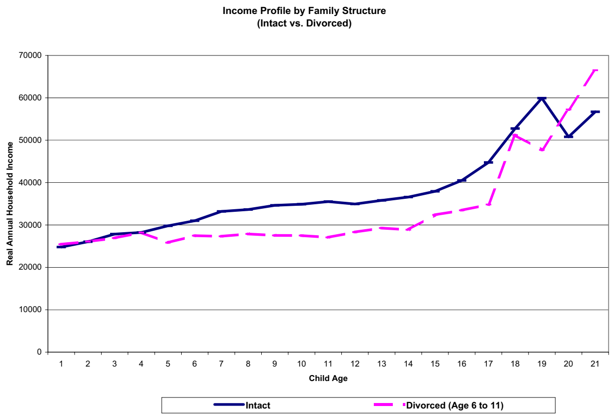


FIGURE 3:
Profile of Family Income by Age at Parents' Divorce





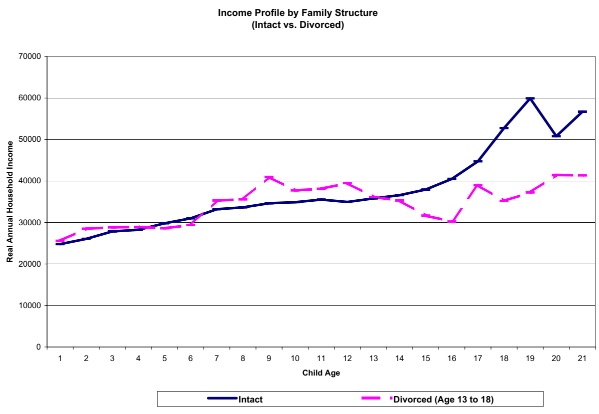
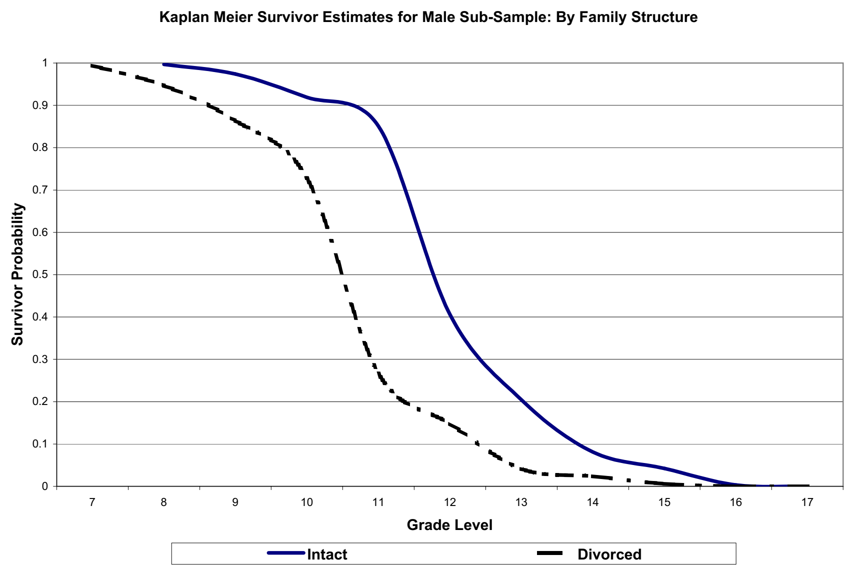


FIGURE 4:
Kaplan-Meier Survival Estimate of Schooling Completion for the Male and Female Sub-Samples



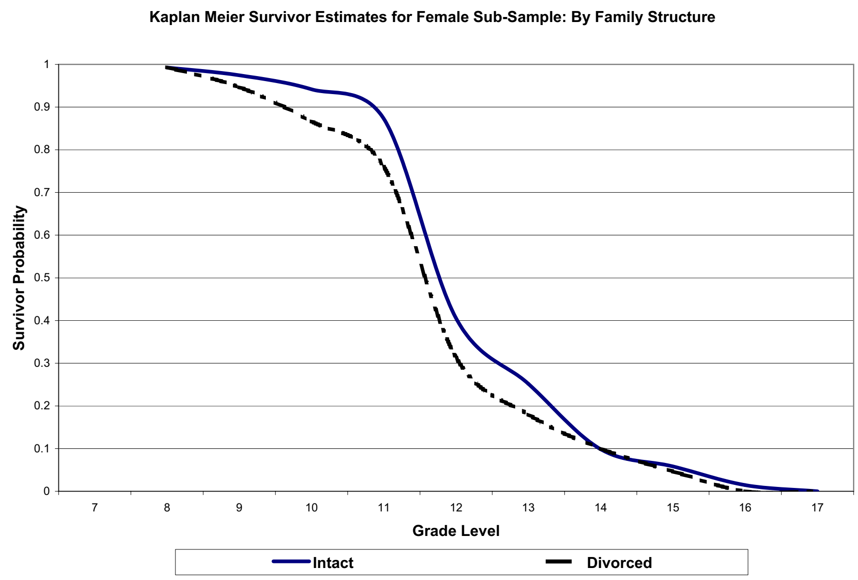
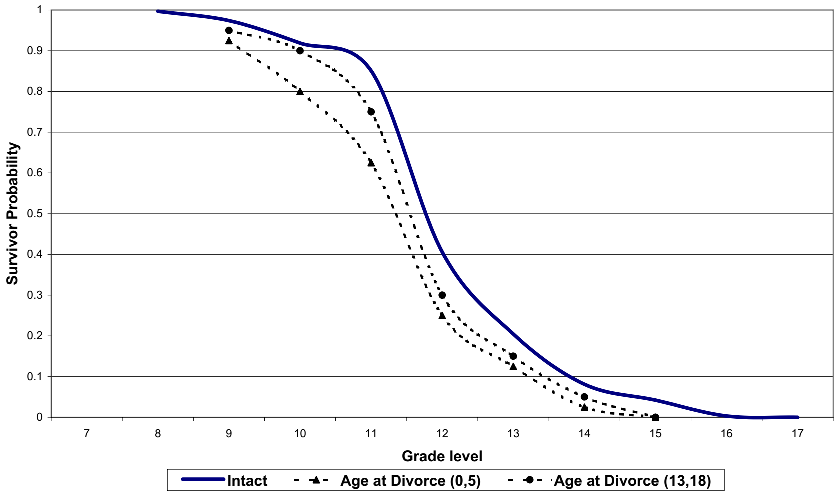
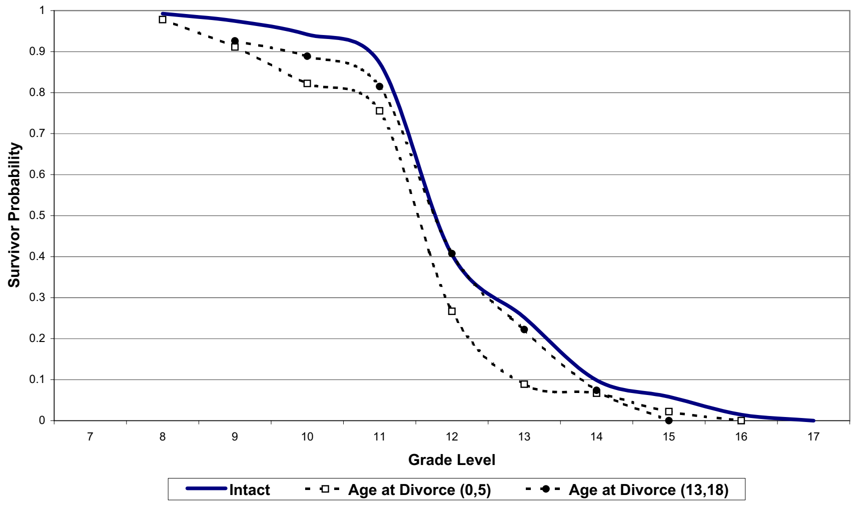


FIGURE 5:
Kaplan-Meier Survival Estimate of Schooling Completion for the Male and Female Sub-Samples
(by Divorce Timing)

Kaplan Meier Survivor Estimates: Comparing Intact and Divorced Families by Timing of Divorce (Males)



Kaplan Meier Survivor Estimates: Comparison between Intact and Divorced Families by Timing of Divorce (Females)



APPENDIX TABLE 1: Variable Definitions

Variable Name	Coding	Time Variant	Definition
Age at parental divorce	Continuous	No	Age when parents divorced
Age at parental divorce ($0 \leq Age < 5$)	Binary	No	$= 1$ if age at parental divorce is [0,5)
Age at parental divorce ($5 \leq Age < 11$)	Binary	No	$= 1$ if age at parental divorce is [5,11)
Age at parental divorce ($11 \leq Age < 13$)	Binary	No	$= 1$ if age at parental divorce is [11,13)
Age at parental divorce ($13 \leq Age < 18$)	Binary	No	$= 1$ if age at parental divorce is [13,18)
Age at parental divorce ($Age \geq 18$)	Binary	No	$= 1$ if age at parental divorce ≥ 18; 0 otherwise
Birth cohort (1968)	Binary	No	$= 1$ if born in year 1968; 0 otherwise
Birth cohort (1969)	Binary	No	$= 1$ if born in year 1969; 0 otherwise
Birth cohort (1970)	Binary	No	$= 1$ if born in year 1970; 0 otherwise
Birth cohort (1971)	Binary	No	$= 1$ if born in year 1971; 0 otherwise
Birth cohort (1972)	Binary	No	$= 1$ if born in year 1972; 0 otherwise
Duration of parents' marriage	Continuous	No	Duration of parents' marriage measured in years (For divorced families only)
Family income	Continuous	Yes	Sum of income of all household members measured in 1984 constant dollars
Father's education	Continuous	No	Highest grade level obtained by father
Father's information missing	Binary	No	$= 1$ If no information on child's father available;0 otherwise
Female	Binary	No	$= 1$ if female; 0 otherwise
Mother's education	Continuous	No	Highest grade level obtained by mother
Mother's age at child birth	Continuous	No	Age of mother when child is born
Mother's age at marriage	Continuous	No	Age of mother upon marriage
Child parity	Continuous	No	Parity with respect to the mother
Order of the marriage	Continuous	No	Order of the marriage defined based on the marriage history of the mother
Parents divorced	Binary	No	$= 1$ if marriage of the parents end in divorce; 0 otherwise
Race: White	Binary	No	$= 1$ if white; 0 otherwise
Race: Black	Binary	No	$= 1$ if Black; 0 otherwise
Race: Other	Binary	No	$= 1$ if racial background is neither white nor black; 0 otherwise
Region of residence: North East	Binary	Yes	$= 1$ if state of residence is ”North East”; 0 otherwise
Region of residence: North Central	Binary	Yes	$= 1$ if state of residence is ”North Central”; 0 otherwise
Region of residence: South	Binary	Yes	$= 1$ if state of residence is ”South”; 0 otherwise
Region of residence: West	Binary	Yes	$= 1$ if state of residence is ”West”; 0 otherwise
Urban residence	Binary	Yes	$= 1$ if the population of the largest city in the county of residence exceeds $100,000$; 0 otherwise
Years since parents' divorce	Continuous	Yes	Years since parents' divorce

Technical Appendix:

Functional Form of the Schooling Exit Hazard

Estimates of hazard models may be sensitive to the choice of functional form (Allison, 1995). We test between four different functional form choices for the hazard function: Weibull, piece-wise exponential, log normal, and log logistic. A common approach to discriminate between models that are not nested is the Akaike information criterion (*AIC*): $AIC =$ -2 (log likelihood) $+ 2(c+p+1)$, where $c =$ number of model covariates, and $p =$ number of model-specific ancillary parameters. Although the model with the highest likelihood is the best-fitting model, the one with the smallest *AIC* value is the most-preferred.[27] We estimated the fully specified model, Model 3 in Table 4, under each of the four distributional assumptions for the hazard function. The corresponding log likelihoods and *AIC* values are reported in Appendix Table 2. Base on these results, the Weibull model is the best fitting and the most preferred model.

APPENDIX TABLE 2
Hypothesis Testing of the Hazard Functional Form Assumptions

	Male		**Female**	
	log likelihood	AIC	log likelihood	AIC
Weibull	588.16	−1122.32	499.00	−944.00
Piece-wise exponential	437.26	−780.52	384.21	−674.23
Log Normal	527.20	−1000.39	460.37	−866.75
Log Logistic	576.59	−1099.18	485.99	−917.98

Note: All statistics based on Model 3 in Table 4.

[27] For a more detailed discussion, see Akaike (1974).