# Parental Divorce and Child Educational Attainment: A Dynamic Approach

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**Abstract**

Using the Panel Study of Income Dynamics, a model of schooling duration is employed to examine (i) the differential effect of the timing of parental divorce; and controlling for uncorrelated unobserved heterogeneity, (ii) the changes in these effects over time, on children's educational attainment. On average, boys who either first experienced parental divorce during early childhood or during teenage years are at higher risk of dropping out of school; for girls, higher risk of dropping out is associated with parental divorce before the age of 11. Controlling for the degree of marital tension and socioeconomic disadvantages prior to the divorce, the "net" impact of divorce is worse the younger the child was upon the separation. In addition our findings suggest that boys may benefit from their parents staying together as long as possible, with much of the benefit working through the importance of resources in their development. For girls this is not true, possibly because of more weight on non-pecuniary attributes of the household. Furthermore, although the negative effect of divorce initially was more pronounced for boys, they also re-adjust faster. The negative effect lingers over time for girls. This suggests that effect of divorce on girls may be more permanent compared to boys, and girls may never fully recover from the initial shock.

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# 1 Introduction

Over the past thirty years, a series of changes in divorce laws in the United States has dramatically increased access to divorce[1]. This, combined with a simultaneous decline in the social stigma associated with being divorced, resulted in a dramatic rise in divorce rates by over 200% (Kreider and Fields [31, 2002]). Using age-specific divorce rates computed from the June 1985 Current Population Survey, Martin and Bumpass ([39, 1989]) developed life-table estimates predicting that between half to two-thirds of current marriages will end in divorce. During this period, the number of children involved more than doubled, reaching over one million per year.

The startling rise in divorce rates and its implications on children involved have been a topic of heated debate among policy-makers, academics, and the general public. A large body of literature emerged in examining the relationship between family structure and child welfare. Using cross-sectional comparisons of the average outcomes of children from different family structures, most studies found that children from divorced families are linked to a variety of adverse long term outcomes, such as lower academic achievements, premarital childbearing, and poor marital outcomes (Case and Katz [9, 1991], Mayer [40, 1991], Graham et al. [22, 1994]).

Lower educational attainment has been shown to be correlated with welfare recipiency, poverty, and unemployment (Bane and Ellwood [5, 1983]), and these findings have been frequently cited by conservative policy-makers who advocate tougher access to divorce[2]. Legislations such as "covenant marriages"[3], in which couples agree upon marriage to conditions that will make divorce more difficult to obtain, have been passed in the states of Arizona, Arkansas, and Louisiana[4]. Despite their wide use, findings based on static comparisons of children from in-

[1]Most states switched from requiring mutual consent to allowing unilateral or no-fault divorce between 1970 and 1985 ([18, 1998]).

[2]The underlying assumption in restricting access to divorce is that uniltaral divorce increase the incidents of divorce. Using 40 years of census data to explore variations over states and over time, Gruber ([24, 2000]) found evidence to support this line of reasoning.

[3]Covenant Marriage laws generally provide that a couple can choose, at the time of marriage or later, to have slightly more limited grounds for no-fault divorce in their marriage. The couple is required to get marriage education or counselling before their marriage (or the conversion of their marriage to a covenant marriage), and also before divorcing.

[4]Legislation for covenant marriages also has passed one house, but not both, in the states of Oregon, Georgia, Texas and Oklahoma.

tact and divorced families cannot fully capture the dynamic variations in environmental factors surrounding a child while growing up that may be correlated with changes in family structure. Child development is a dynamic process in which the timing and the sequence of events, such as parental divorce, may have differential implications on the long run outcome.

While the debate about policies to affect family structure is taking place, there is no consensus among researchers as to why and how family structure matters for child development. The experiences of children whose parents divorced may be different based on their age at parental divorce, and these effects may also differ by gender. In examining the effect of divorce timing, static models are unable to keep track of the sequence of events, which leads to serious identification problems. Existing evidence on the effect of parental divorce based on child observed characteristics are inconclusive, and often based on selected samples and therefore not generalizable (Shaw [53, 1982], Krein and Beller [32, 1988], Lang and Zagorsky [33, 2000], Painter and Levine [48, 2000], Corak [11, 2001]).

Part of the disadvantages confronted by children from divorced families may be attributed to the loss of parental time and income as a direct result of marital dissolution (Corak [11, 2001]). Parental choice of time and material investments, given available family resources, are correlated with changing family structure over time. Weiss and Willis ([56, 1985]) noted that divorce may lead to an inefficiently low level of expenditures on children, since the custodial parent does not internalize the preferences of the ex-spouse for the expenditures on children. On the other hand, poverty may also be a cause for divorce, given that the stress induced by economic hardship may be a burden on marital relations (Painter and Levine [47, 1999]). Cross sectional comparisons between intact *vs.* divorced families may overstate the effects of family resources on child outcome; and simple "before and after" comparisons are likely to be understated since they do not account for life-cycle earnings growth.

In addition, child long run outcomes may also depend on the child's unobserved ability to cope and re-adjust from parental separation (Jenkins et al. [30, 1989], O'Brien et al. [45, 1995]). The initial reactions to parental divorce may be different for each child, and their ability to cope

with such a shock over time may also differ. Holding everything else constant, given two children from different families whose parents divorced at the same time, their long run outcomes may differ due to their unobserved ability to cope. To our best knowledge, there is no study that explicitly address the potential impacts of state-dependent child unobserved heterogeneity on child long run attainments.

This study re-evaluates the differential impact of the timing of parental divorce on child educational attainment and explores changes in these effects over time using a hazard model of schooling duration with random effects. Given the lack of consensus with respect to the effect of divorce on child educational outcomes based on child observed characteristics, we aim to re-examine these effects in a framework that models child educational attainment as a dynamic process. Furthermore, we aim to explore how the effects of parental marital dissolution on children changes as they grow up. Using annual data on a sample of children since birth until the completion of their schooling, we treat child attainment as a dynamic process which accounts for the time-varying nature of family characteristics that are correlated with changing family structure. Incorporating random effects enables us to control for child unobserved heterogeneity with respect to their ability to re-adjust from the initial shock of divorce.

Given that a large portion of American children may experience parental divorce while growing up, and their experiences may be very different based on the timing and nature of the separation, any simple generalization of the effects of divorce should be made with caution to prevent stigmatizing a large proportion of the population. We attempt to answer the following questions: Are all children hurt by divorce? If not, how are the effects of parental divorce different based on its timing, and across genders? Furthermore, how do the differential impact of divorce on children change over time? Using static models and selected samples, previous findings on the differential effects of parental divorce and its timing on child long run outcomes have been inconclusive, which prevent the design of appropriate policies to help children from divorced families. By using a nationally representative sample of children and observing them from birth until the completion of their schooling, and employing a model which treats child

development as a dynamic process, we will be able to separate out the confounding factors associated with the timing of parental divorce, and pinpoint the differential magnitudes of the timing effects. For example, it is found that the variations in family resources (time or money) is the most important determinant of child outcomes, then policies and social programs that enforce child support payments, increase financial assistance to divorced parents, and child counselling programs would be helpful for these children to adjust. If it is found that the adverse effect of parental divorce is significant only for a sub-group of children based on timing and child gender, then restricting access to divorce in general may not help children. In such a case, limiting access to divorce for parents of children with such characteristics may help. Therefore, allowing for unobserved differences, and testing for the effect of divorce independent of related factors, is critical for informing policy.

# 2 Literature Review

Children of divorced families are found to be more likely to drop out of high school, have children out-of-wedlock, be arrested, less likely to go to college or hold a good job, and more likely to experience failed marriages themselves (Amato et al. [3, 1995], Biblarz et al. [6, 1997], Garasky [20, 1995], McLanahan [41, 1985], McLanahan and Sandefeur [42, 1994], Simons [54, 1996]). Furthermore, female-headed families are found to be economically disadvantaged, invest less time in children, have fewer and weaker male role-models and social networks (Duncan and Hoffman [14, 1985], Manski et al. [38, 1992]). Children of divorced families suffer the separation from one parent, parental hostility, and residential dislocation (Amato [2, 1987]). Furthermore, children in single-parent and stepfather families are more likely to be sexually abused (Russell [50, 1984], Gordon [21, 1989]).

It is well known that children from divorced families have fewer economic resources than their counterparts in intact families. The negative outcomes of children from divorced families may be explained by the reduction of socioeconomic resources associated with parental separation.

McLanahan and Sandefur ([42, 1994]) estimated the effects of differences in assets across family types, and attributed half of the difference in outcomes to differences in family income. However, their studies are based on cross sectional comparisons which tends to overstate the effects of family resources, since the changes in available resources are correlated with divorce. However, there is evidence that suggests that there are also significant recoveries due to subsequent re-marriages (Duncan and Hoffman [14, 1985], Page and Stevens [46, 2002]). Controlling for family resources "before and after" divorce in a static model is also problematic since it does not account for life-cycle earnings growth, and overlooks the probability of dynamic re-adjustments to available resources due to parental remarriages. Therefore, to identify the effect of parental divorce on child long run outcomes, variations in available resources correlated with changing family structure at each point in time may need to be accounted for in a dynamic setting.

The effects of divorce on child educational achievements may differ based on its timing, and these effects on child long run outcomes may also be different by gender. Existing research show conflicting findings on the effect of the timing of divorce on child outcomes, and most results found are not generalizable due to the selected samples used in these studies. Sandefur et al. ([51, 1992]) explore the effect of family type at age 14 on a child's likelihood of high school graduation. They found that disruptions of family structure have persistent negative consequences on high school graduation, and family income accounts for approximately 15% of the negative effects. Focusing on children during their teenage years, Shaw ([53, 1982]) examines the effects of time spent in a single-parent family at different ages and for different durations on child educational attainment. She found that living in a single-parent family for more than two years reduces the likelihood of black girls in completing high school.

A few studies use richer longitudinal dataset to further explore the effects of divorce timing. However, conflicting results remain. Krein and Beller ([32, 1988]) examine the effect of the number of years spent in a single-parent family during preschool, elementary school, and high school on years of schooling completed by a child. They found that the effect is most notable during preschool, increases with the length of exposure, and the effect is larger for boys than for

girls. Using the National Education Longitudinal Studies (1988), Sun ([55, 2001]) finds children of either sex are equally vulnerable to parental separation. Furthermore, it has been suggested that parental marital conflict affects children's emotional development and their coping responses (Jenkins et al. [30, 1989], Fauber et al. [16, 1990], Davies and Cummings [13, 1998]). Using meta analysis, Buehler and Gerard ([7, 2002]) found that martial conflict placed daughters in greater jeopardy between the ages of 2 and 11, given that daughters may be more vulnerable to a stressful family environment because they were socialized to care-take and safeguard family relationships.

The lack of consensus in the effects of the timing of parental divorce may be partially due to the static methodologies used. Previous studies typically explore the differential impacts of the timing of parental divorce in a static framework, which may be unable to properly identify the differential impact of the timing of parental divorce. In examining the effects of divorce on child educational outcomes, these models typically either (i) categorize children into different groups based on the timing of parental divorce (Wojtkiewicz [60, 1993], Krein and Beller [32, 1988], Furstenberg and Kiernan [19, 2001]); or (ii) use controls for whether the divorce took place prior to the completion of schooling (Sandefur et al. [51, 1992], Ribar [49, 1993], Painter and Levine [48, 2000]). Identification problems arise since these models cannot keep track of the sequence of events. Furthermore, the explanatory variables are potentially confounded with the dependent variable (Montgomery [44, 2001]). Hazard models are able to keep track of the sequence of events, and therefore allow us to properly identify the differential effects of the timing of parental divorce. Duration models have been applied to examine the effect of family structure on pre-marital childbearing behavior of youths (Hogan and Kitagawa [29, 1985], Bumpass and McLanahan [8, 1989], Wu and Martinson [61, 1993]), but not on child educational attainment. A hazard model of schooling is a logical application to examine the differential impacts of divorce timing on child educational outcomes.

Existing cross-sectional empirical evidence suggest that there are negative consequences associated with parental divorce. However, there may be unobserved family characteristics that

jointly determines changes in family structure and child long run outcomes (Haveman and Wolfe [27, 1995], Sun [55, 2001]). If the negative child outcomes are caused by these unobserved factors, rather than divorce *per se*, than the estimated effects of parental divorce may be downward biased. Cherlin et al. ([10, 1991]) uses longitudinal data of children from the U.S. and Great Britain, and found that the observed negative effects of divorce can be largely explained by conditions that existed prior to the separation.

Since a natural experiment of how the children would have fared if their parents stay together is impossible, various studies have attempted to test the endogeneity of divorce. Using a variety of datasets and approaches, empirical evidence on whether divorce is endogenous is inconclusive. Grogger and Ronan ([23, 1996]) and Sandefur and Wells ([52, 1997]) exploit the presences of sibling pairs in the National Longitudinal Survey of Youth (NLSY), to separate out the shared unmeasured family characteristics among siblings from the impact of parental loss. Both find that fatherlessness reduces child educational attainment. However, within family comparisons cannot account for changes in intra-household resource allocation as a result of divorce, or the potentially differential impacts of parental loss on older vs. younger siblings (Lang and Zagorsky [33, 2000]). Using an instrumental variables approach and a quasi-experimental setting, respectively, Lang and Zagorsky ([33, 2000]) and Corak ([11, 2001]) both test the endogeneity of divorce using parental bereavement as an exogenous source of parental loss, and both found significant differences between outcomes of children from bereaved vs. divorced families. However, whether this difference is due to unobservables that are endogenous to the divorce decision, cannot be determined. Painter and Levine ([48, 2000]) address the endogeneity of divorce by controlling for characteristics that preceded the birth of the child and when the child was in eighth grade. In contrast to other findings, Painter and Levine found that the effects of divorce on child educational outcomes do not diminish and remain significant, which suggests that parental divorce may have independent effects on children in the long run. In this study, we assume that parental divorce is an exogenous determinant of child educational outcomes. However, this assumption will be tested using an instrumental variables approach to control for

marriage quality[5].

Although it has been established that children from divorced families fare worse than their counterparts in intact families, less is known about how parental divorce affects children as a continual process as they grow up. Children do not react and re-adjust uniformly to parental divorce, but rather respond differently based on age, gender, their temperament, the economic impacts and the potentially lingering conflict between their parents over time (Allison and Furstenberg [1, 1989], Amato and Booth [4, 1997], Emery [15, 1999]). Conflicting results are found among the previous work may be due to: First, static models may not allow for proper identification of the differential effects of divorce timing on child outcomes, given that the effects of family structure and resources may vary at different points in time; second, the lack of adequate data on crucial socioeconomic factors confounded with family structure at each point in time for children from birth to their completion of schooling, without which the estimated effects of divorce may be biased; and finally, child unobserved coping ability have been largely overlooked, and the estimated effects of parental divorce may be understated if it is not explicitly accounted for.

Policies that aim to help children exposed to parental divorce need to be based on generalizable findings that account for these aforementioned issues. The potentially differential effects of parental divorce and its implications on childhood experiences need to be carefully examined in a unified framework and representative dataset. We employ a hazard model of schooling duration with random effects to explore the differential impacts of parental divorce and its timing. This study explores how the effects of parental divorce vary over time and potential differences based on divorce timing. We model child educational attainment as a continuous process. However, we recognize that the effects of parental divorce are potentially over-estimated if part of the negative effects associated with divorce can be explained by unobserved family characteristics prior to divorce. We attempt to address this issue by (i) controlling for potential disadvantages

[5]Conditional on that a marriage eventually end in divorce, we utilize the duration of the marriage as an instrument for the quality of the marriage. We assume that the most troubled marriages tend to end early, and marriage that lasted longer are not the most problematic.

in socioeconomic conditions prior to divorce in a dynamic setting, given that socioeconomic disadvantages may attribute to, and increase the likelihood of divorce; and furthermore, we attempt to control for the unobserved quality of the marriage by (ii) employing the duration of marriage as an instrument for the quality of the marriage, assuming that marriages that are most troubled should dissolve early. However, if the endogeneity of divorce exists and the negative effects are due to factors that are correlated with the observed socioeconomic factors, then the estimated effects of parental divorce found here may be downward-biased[6]. Nonetheless, by modelling child attainments in a hazard framework, we are able to identify the effects of divorce timing as a continuous process by separating out opposing effects due to factors confounded with the timing of parental divorce.

This study examines (i) How do the effect of the initial shock of parental divorce on children differ based on child's age and gender; (ii) How do these effects change over time and its implications on child attainments in the long run. In our hazard model, family resources and family structure are allowed to affect the likelihood of children in obtaining an additional year of schooling from one period to the next. Using data from the PSID, we utilize socioeconomic and demographic information on children that covers the entire period between birth until the completion of their schooling to fully capture all time-varying factors that may impact child development at each point in time. This allows us to identify the effects of differential initial shocks of parental divorce by its timing on child long run outcomes. Our results are generalizable because we consider divorces that take place at many different stages in a child's life, between birth until the completion of their schooling. Due to data limitations, previous studies usually consider only a subset of divorces, namely divorces that take place when children are in high school (Krein and Beller [32, 1988]), therefore the results of which may not be generalizable. Furthermore, incorporating random effects allows us to account for unobserved child abilities to cope with divorce. There may be potentially confounding effects associated with the tim-

[6]To address the endogeneity of divorce, a natural extension to the present work is to model child attainments and divorce timing in a joint decisions framework, which is beyond the scope of this paper. In Liu ([37, 2003]), a simultaneous two-equation hazard model of child attainment and the divorce decision is employed to address the endogeneity issue explicitly.

ing of parental divorce, such as the length of time to re-adjust from the initial shocks. These potentially opposing effects can be separately identified in our model.

# 3 Model

We explore the effects of parental divorce as a continuous process, and examine the changes of these effects over time as children grow up. The underlying conceptual framework for this analysis builds on the model developed by Leibowitz ([34, 1974]). The production of child quality, as measured by the child's educational attainment, is given by:

$$S_t = S(HI_{t-1}|Y_{t-1}, MS_{t-1}, \gamma)$$

In each period, child quality is produced via time and goods home investment inputs chosen by the parents, $(HI)$, chosen optimally by the parents in the previous period. The choice of these inputs is determined conditional on family income $(Y)$, parental marital status $(MS)$, and child innate ability $(\gamma)$. Parental abilities, in various dimensions, are passed onto their children via heredity. Children's innate ability, both intellectual and emotional, together with home investment inputs chosen by the parents, determine the level of schooling attained of the children.

At each point in time, parental choice of investments toward the human capital accumulation of their children are correlated with changes in family structure. Parental choice of the quantity and quality of inputs for their children are constrained by their marital status, given that single-parent families are more limited at the margin in the amount of time and goods available compared to an intact family. In addition, given that divorce can be a stressful event, parental investments both before and after the divorce may reflect the compensating behaviors of parents toward their children. The effect on child development in the long run would depend on the relative magnitude of the negative shock of divorce compared to the benefit of these investments.

The potential effects of divorce and its timing on the welfare of children, both immediately

and in the long run, are important factors parents consider when deciding whether or not to stay married (Weiss and Willis [57, 1997]). Among marriages with children that eventually end in divorce, marriages that dissolve early when children are very young may be intrinsically different than those that stay together. Marriages that end relatively early may be the most troubled ones, and those marriages that lasted longer may be the ones that are fairly stable (Painter and Levine [48, 2000]). Child attainment in the long run may be affected by the effects of parental separation on child cognitive development and their coping abilities at each age. Holding unobserved child coping ability constant, the effects of divorce on child development may be different by their age at parental separation (Corak [11, 2001]) and gender (Krein and Beller [32, 1988]).

In the present context, children's schooling is modelled as a duration process. The higher the level of educational attainment the child ultimately achieves, the longer the duration of her schooling. Let $T_i$ denote the duration of schooling of child $i$. The log-survival time model of schooling is:

$$\ln T_i = \alpha_0 + \alpha_1 HI_i(t) + \alpha_2 PC_i + \alpha_3 Z_i(t) + \alpha_4 D_i(t) + \alpha_5 I_i + \gamma_i + \sigma\epsilon_i \tag{1}$$

As we have mentioned, child attainment depends on the parental allocation of resources at each point in time. $HI_i(t)$ is a vector of covariates which measures the quantity of parental inputs in the production process of child quality for child $i$ at time $t$. Heiland ([28, 2003]) finds that the number of siblings is a proxy for the amount of resources available to a child, given that they compete for the available time and goods resources of the parents. Family income and number of siblings are used as proxies for resources available to the child. Furthermore, family income at each point in time is correlated with changes in family structure such as divorce and remarriages. The estimated effect of divorce may be overstated if we do not account for this dynamic variation.

Parental attributes, $PC_i$, includes time-invariant covariates reflecting observable parental characteristics. Parental educational background have been found to be important determinants

of the quality of parental time investments in children (Leibowitz [34, 1974], Datcher [12, 1982], Haveman et al. [26, 1991], Manski et al. [38, 1992], Ribar [49, 1993]). Additional controls in this category include mother's age at marriage, her age at childbirth, and the order of the marriage. Women who married and have children at younger ages have on average lower educational attainment and weaker attachments to the labor force. These mothers tend to be less financially independent and less likely to be able to support their children on their own once divorced. It has been found that maternal education predicts earlier divorce more strongly than later divorces (Painter and Levine [48, 2000]). In addition, these factors may be proxies for parental preferences with respect to marriage and divorce. Some couples may have a weaker taste for marriage and therefore more likely to divorce (Lillard and Waite [36, 1993]).

The vector $Z_i(t)$ consists of both time varying and time-invariant family and child demographic characteristics. There may be long run implications for children growing up in a unilateral divorce environment[7]. Gruber ([24, 2000]) found that unilateral divorce not only increases the incidences of divorce, but adults who are exposed to unilateral divorce regulations as children are less educated. To control for environmental influences of legislative factors that reduces the stigma of being divorced and its implications on children, we include proxies for the region of residence and the birth cohort. Child ability to cope with family breakup may affect child development independent of other family characteristics. The parameter, $\gamma_i$, captures the unobserved child heterogeneity with respect to their abilities to cope with stressful family events, which is treated as an random effect on child attainments. It should be noted, that the unobserved factors captured by $\gamma_i$ are those uncorrelated with the other observed factors included in our model, which is not to be confused with those unobserved family characteristics that may be endogenous to the divorce decision. Finally, $\epsilon_i$ is a individual-specific error term.

To capture the effect of the timing of parental divorce, we stratified our sample of children from divorced families into five groups based on their age at parental divorce, with the control

[7]Following the passage of unilateral divorce legislation in the state of California in 1969, many states followed suit in a wave of family law reforms (Weitzman [58, 1985]). Unilateral divorce is divorce that does not require the explicit consent of both partners.

group being children from intact families. $I_i$ is a vector of indicators which groups children into separate comparison groups based on their age at parental divorce, with each of the coefficients representing the underlying baseline schooling hazards for each of these group of children compared to those from intact families. This identifies the differential impacts of the timing of divorce. The underlying hypothesis is that marriages that dissolve when their children are of comparable ages are similar in unobserved ways, after controlling for observed differences in family characteristics. Marriages that ended when their children are of similar ages may be alike in unobservable ways, in that they may share similar experiences both before and after the separation. Furthermore, we are assuming that children who were of similar ages when their parents divorced may share similar experiences growing up, after controlling for child unobserved coping abilities. Holding everything else constant, the impact of family dissolution may be similar for a child of a particular age and gender.

To shed light on changes in the effect of parental divorce over time, we further account for two dimensions associated with the timing of divorce. The vector of covariates, $D_i(t)$, includes factors associated with the timing of divorce, such as marriage duration and length of time since divorce. The length of the marriage of those that ended divorce may be an indicator of the quality of the marriage. Marriage duration may be an indicator of the quality of marital relations. It has been argued that parents who are unhappily married but perceive separation as damaging to their children may avoid or postpone the separation (Levinger [35, 1965]). However, if marital relations are sufficiently poor, avoiding divorce may only be delaying it, and subjecting children to an adverse family environment may not be even more harmful for their long term developments. There, we believe that the most troubled marriages should dissolve earlier, even if children are involved. This may explain the existing findings of a strong negative effect of divorce on young children. By controlling for the duration of marriage before it ends in divorce, we attempt to control for the quality of marriage, which may be endogenous to the divorce decision. Furthermore, if early divorce is associated with a larger negative shock, it is also confounded with a longer time to heal. The negative effect of divorce may be most damaging

when the child is young, but it is also associated with the longer adjustment period until the completion of the child's schooling, which may be beneficial. The potential opposing effects with respect to the timing of divorce needs to be accounted for. Without controlling for these two effects, the estimated effects of different timings of divorce may be understated.

## 3.1 Hazard Specifications

In the present framework, child educational attainment is viewed as a duration process. At each point in time, changes in family structure, parental choice of inputs, and child innate abilities interact and influence the likelihood of the child continuing onto the next education level. In estimating the model of schooling duration proposed in the previous section, we need to assume a specific functional form of the hazard function. The choice of an appropriate functional form of the hazard function is very important, since the results may be very sensitive to this decision. In this section, we first discuss the two hazard specifications used in this study, namely the Weibull and the piece-wise linear hazard specifications. Second, we incorporate unobserved heterogeneity, or random effects, into our model.

The first specification assumes that the log of schooling duration, $\ln(T_i)$ as in equation (1), follows a Type II Extreme Value distribution, which implies that 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 Weibull hazard function is:

$$h_i(t) = exp\{-X_i\beta\}\frac{1}{\sigma}t^{\frac{1}{\sigma}-1} \tag{2}$$

Where $\sigma$ and $\{\beta\}$ are parameters to be estimated. Consider the duration of schooling for a given child, it is reasonable to expect that the hazard rate of schooling to exhibit positive duration dependence, given that the longer a child is in school, the more likely that she will finish her schooling in the next period. Therefore, we expect that the shape parameter, $\sigma$, to be in the range of $(0, 1]$, hence the schooling hazard should exhibit positive duration dependence[8]. A

[8]If $0 < \sigma < 0.5$, the hazard increases at an increasing rate; linearly if $\sigma = 0.5$; and at a decreasing rate if 0.5

factor has a negative effect on child educational attainment if it increases the probability of dropping out of school at any stage during the schooling career given the child has achieved a certain level of education by a specific time (i.e. The hazard rate of schooling). For the ease of interpretation, 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 (1), given the fact that: $\widehat{\alpha_k} = -\widehat{\sigma}\widehat{\beta_k}$. This makes intuitive sense since if a factor increases the hazard rate, then the events occur quickly and the survival time (i.e. duration) is short.

To test our distribution assumption of the Weibull functional form on the hazard function, a more flexible functional form is employed: a piece-wise exponential hazard specification[9]. If we assumes that error term in equation (1), $\epsilon$, follows a Type I Extreme Value distribution, then the duration of schooling, $T_i$, follows an exponential distribution. It is more flexible than the Weibull hazard since it imposes relatively few restrictions on the structure of the hazard function. In this framework, the time span of educational attainment can be arbitrarily divided into intervals, and for the purpose of this study we will divide the time span by grade level[10]. In this specification, the data is used to reveal the baseline hazard. The piece-wise exponential hazard function is given by:

$$h_i(t) = \begin{cases} \lambda_1 e^{X_i(t)\beta}, & \tau_0 \leq t < \tau_1; \\ \lambda_2 e^{X_i(t)\beta}, & \tau_1 \leq t < \tau_2; \\ \vdots & \\ \lambda_M e^{X_i(t)\beta}, & \tau_{M-1} \leq t < \infty \end{cases} \tag{3}$$

Where $\{\tau_m\}$ are $M-1$ known constants, representing the cut-off points for each time interval. The baseline hazard, $\{\lambda_m\}$ are $M$ unknown parameters, together with $\{\beta\}$, are to be estimated. A nice feature of the exponential assumption is that $\widehat{\alpha_k} = -\widehat{\beta_k}$. The results from the piece-wise

$< \sigma < 1$.

[9]For a more detailed discussion, see Meyer ([43, 1990])

[10]Due to data limitations, 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 for every child.

model will be compared to that of the Weibull model, this comparison will determine how well our Weibull model fits our data used in this study[11].

## 3.2 Incorporating Child Unobserved Heterogeneity: Hazard Model with Random Effects

Unobserved child coping abilities may influence child long run outcomes independent of the effects of observed factors. The effects of parental divorce may be understated if child coping ability is not accounted for, therefore by explicitly incorporating the unobserved factors allows us to obtain more precise estimates of the effect of divorce. We employ a hazard model with random effects, or the parametric shared frailty survival model to explicitly account for child unobserved heterogeneity. Following Gutierrez ([25, 2002]), let $h_{ij}(t)$ denote the hazard function without random effects. Child-specific unobserved abilities, as captured by $\gamma_i$, is taken as an random effect that is individual-specific and assumed to be uncorrelated with the other covariates[12]. For the ease of notation, let $\phi_i = e^{\gamma_i}$, which captures the unobserved child heterogeneity. We assume that $\phi_i$ follows a Gamma distribution with mean one and variance $\theta$[13]. 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 the total number of individuals with individual $i$ being observed for $n_i$ periods. The hazard function conditional on the individual-specific random effects 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 $i$ is observed, the standard hazard function is now multiplied by the shared frailty $\phi_i$. Under the Weibull specification,

[11] By incorporating random effects in the hazard model, the Weibull distribution is a good alternative due the limited number of parameters to be estimated. However, it can only be used if it appropriately describes our data.

[12] For a discussion on hazard models with random effects, see Firth and Payne ([17, 1999]).

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

the conditional hazard for individual $i$ at period $j$ is:

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

And the conditional survival function is:

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

To obtain the likelihood function, we calculate the individual-level conditional likelihoods and integrate out the random effect. Suppose $(t_{0ij}, t_{ij}, d_{ij})$ indicate the start time, end time, and whether the observation experiences a failure or censoring for individual $i$ at period $j$. Given $\phi_i$, the contribution to the likelihood is therefore:

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

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 $\phi_i$,

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

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:

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

Given the unconditional individual likelihoods, we can estimate the regression 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 Selection

This study utilizes information on a sample of children from the Panel Study of Income Dynamics (PSID)[14]. The PSID began interviewing a cross-section sample of $5,500$ households in the United States in 1968. Follow-ups are conducted annually, and are still on-going. In each year, the study collected household level information, as well as information on each individual residing in these households. The primary goal of the PSID is to collect information on short-run changes in the economic status of families and individuals. Over time, the study expanded to include detail demographic information of these families and individuals as well. For the purpose of this study, we utilize information from all 25 waves of the PSID, from 1968 to 1992.

We select a sub-sample of children born into the original panel of households interviewed in 1968. Due to sample design, the information on grades attained each year for an individual is not available unless the individual has actually completed his formal education in that year. For the purpose of this study, given that schooling is used to measure child outcomes, it is crucial that we observe the level of schooling attained for each child. Therefore, we select our sample of children who were born between the years of 1968 and 1972, for whom we observe over a span of time long enough, to learn the exact timing of schooling completion and the level of education attained. The information on the timing in which a child actually started schooling (i.e. first grade) is not available, therefore in estimating the hazard of schooling, the duration of schooling

[14]Sponsored and distributed by: Inter-University Consortium for Political and Social research, University of Michigan, Ann Arbor, MI.

is taken as from birth until the completion of schooling. We recognize the possible measurement error in the duration of schooling, particularly in case of left-backs and temporary drop-outs[15]. Therefore we checked the age of each individual at the time of schooling completion with respect to the reported level of schooling attained, and dropped cases that seem very unreasonable[16]

To identify the birth parents of each child included in our sample, we used the *Childbirth and Adoption History Supplementary File* of the PSID. A child is kept in our sample if we are able to identify at least one of his birth parents[17]. If only one birth parent is identified, we then assume that the spouse of the identified parent who resided in the same household the year in which the child is born, is the other birth parent of the child. Furthermore, marriage history information of the birth parents are obtained using the *Marriage History Supplementary File*[18], in order to pinpoint the timing of parental divorce. The final sample used in this study includes 901 individuals. Given the long span of the panel and its annual follow-ups, we have detailed socioeconomic and demographic information on these children and their families since birth until they complete formal schooling.

# 5 Descriptive Statistics

In this section, we first present the descriptive statistics and variable means of the pooled sample of children used in this study. We further compare means by child gender and family structure. In addition, a non-parametric estimator, namely the Nelson-Aelan cumulative hazard estimator, is used to generate graphs to examine the relationship between certain observable characteristics and the probability of dropping out of school.

[15]Given that our focus is the level of schooling attained, and we measure the duration of schooling accordingly, if an individual is left-back or temporarily dropped out for a year, it is still considered a "good" outcome given that it is the ultimate level of education attained by this individual we are interested in. We understand this underestimates bad outcomes.

[16]An individual is included in our sample if he reported to have completed a certain level of education at an age that is at most two years above the expected age at that level of schooling.

[17]The Childbirth and Adoption History File and the Marriage History File include retrospective childbirth, adoption, and marriage histories of individuals in sample households. Both supplementary surveys were conducted in 1985, therefore individuals who attrited prior to 1985, or were institutionalized in 1985, were not included.

[18]Refer to the previous footnote.

Table (2) presents the descriptive statistics of the pooled sample of children used in this study[19], and Table (3) further compares the means by child gender. The average child in our sample is a high school graduate, who has completed 12 years of formal schooling. Approximately 34% of our sample experienced parental divorce while they were growing up[20]. As previously mentioned, to identify the differential effects of the timing of divorce, children from divorced families are categorized into five groups based on their age at parental separation, with each category corresponds to a distinct age range of child development[21]. This rests on two important underlying assumptions: (i) Families that dissolve when their children are of a certain age shared many similar experiences over time; (ii) The initial negative effect of divorce is similar for children of a particular age range. Based on our sample, most divorces take place when the child is fairly young, with the majority of children from divorced families experiencing the separation at elementary school age. Figure (1) illustrate the distribution of divorces by the child's age at parental divorce.

Table (4) compares children from divorce families to those from intact families. Children from divorced families have attain fewer years of schooling, on average, compared to children whose parents stay married. Children from disrupted families are less likely to pursue a secondary education, since the average child from divorced families have just about 12 years of formal education, compared more than 12 years of education for a average child from intact families. The unfavorable outcome observed for children from divorced families may be partially attributable to the economic disadvantage associated with marital separation. Experiencing parental divorce is also associated with economic disadvantages, given that the average annual family income of a divorced family is approximately 83% of those that stayed intact.

In addition, as mentioned previously, the duration of marriage may be an indicator of the

[19]For variable definitions, please refer to Table (1).

[20]Wojtkiewicz ([59, 1992]) estimated that approximately 37.9% and 24.9% of black and white children of the 1960s cohort would experience parental divorce, respectively, which is comparable to the cohorts used in this study.

[21]These categories were chosen based on the age range corresponding to particular education levels: preschool ($0 \leq \text{Age} < 5$), elementary school ($5 \leq \text{Age} < 11$), junior high school ($11 \leq \text{Age} < 13$), high school ($13 \leq \text{Age} < 18$), and beyond high school ($\text{Age} \geq 18$). Each of these age ranges represents a particular developmental stage for children: early childhood, late childhood, pre-teens, teens, and adolescent or young adult.

quality of marital relations. Given two marriages that eventually end in divorce, the one that stay together longer may be relatively more stable. It is worthwhile to note that marriage duration need not be perfectly confounded with the timing of divorce measured by the child's age upon parental divorce. Given two children whose parents divorce when they are of the same age, the one whose parents stayed together longer may have enjoyed relatively secured family environment assuming that the marital relations of the parents are fairly stable. Figure (2) presents the distribution of divorces in our sample by marriage duration. The average marriage duration of our divorced sub-sample is about 14 years.

## 5.1 Nelson-Aelan Cumulative Hazard Estimator

This section uses Nelson-Aelan cumulative hazard estimator to explore the relationship between certain observed characteristics and the probability of dropping out of school. The Nelson-Aelan cumulative hazard estimator (Nelson 1972 and Aelan 1978) is a non-parametric estimator of the cumulative hazard function, which does not require any distributional assumptions. The cumulative hazard function is the sum of the ratio of the number of children who dropped out of school to the number of children who are still in school at each point in time up to time $t$. The Nelson-Aelan estimate of the cumulative hazard rate at time $t$ is:

$$\widehat{H(t)} = \sum_{j|t_i \leq t} \frac{d_j}{n_j}$$

Where $n_t$ defines the population that survived to time $t$; and $d_t$ represents the number of failures at time $t$. Figure (3) shows the cumulative hazard function for children from intact versus divorced families. As we can see, starting at the age of 12, the hazard of schooling, or the likelihood of dropping out, is higher for children from divorced families than children from divorced families. This suggests that children from divorced families are at a higher risk of dropping out of school at each point in time after the age of 12.

The purpose of this study is to examine the differential impacts of divorce timing on children

by gender, and explore the changes in these effects over time. To motivate the discussion, the Nelson-Aelan estimator is used to examine the relationship between the the probability of dropping out of school and parental divorce, with special emphasis on: (*i*) timing of divorce; (*ii*) differences by gender. The findings here motivate the multivariate parametric analysis that follows.

**Children from Divorced Families: Differences by Timing of Divorce**

Previous findings suggest that the effect of parental divorce on child outcomes may differ based on its timing, as measured by the age of the child upon parental divorce. Figure (4) compares the schooling hazard of children from divorced families to those from intact families by the timing of parental divorce. Consistent with previous findings, children whose parents divorced when they are of preschool age or during high school age have significantly higher risk of dropping out of high school compared to children from intact families. On the other hand, the risk of dropping out of school is fairly similar between children whose parents divorced during elementary to junior high school age and those from intact families. This is consistent with previous findings that divorce may be relatively more harmful for very young children and teenagers. Divorce may be detrimental to children at very young ages because children may be most vulnerable and require the most care during early childhood. The teenage years are a period where children learn to form inter-personal relationships and an independent identity. At this age, stress induced by the breakup of a family may harm their development, and shock their expectations.

Children whose parents divorced after they are 18 years of age have lower risk of dropping out of school than children whose parents stay together. There are two potential reasons for this. On the one hand, given that these children are already at an age when they are close to finishing their education, parental marital disruptions may no longer have a strong effect on the children's schooling attainment. On the other hand, marriages that lasted that long, even beyond the completion of the education of their children, may not have been very different to those that stayed intact in terms of the level of stability in the family environment. It may be

that if the marriage is troubled it would have ended earlier.

**Children from Divorced Families: Differences by Gender**

As illustrated in Figure (3), children from divorced families are at a higher risk of dropping out of school, hence have shorter duration of schooling, compared to children from intact families. The effect of divorce may also be different for boys compared to girls, as there may be differences in developmental needs of children by gender. It has been found that divorce may be more harmful for boys, especially if it took place when the child is relatively young (Krein and Beller [32, 1988]). However, the effect of divorce on girls remains controversial. Figure (5) presents the three separate comparisons of the non-parametric cumulative schooling hazard estimates: (*i*) Male children from divorced vs. intact families; (*ii*) Female children from divorced vs. intact families; and (*iii*) Male vs. female children from divorced families.

Consistent with the findings shown in Figure (3), the risk of dropping out of school is higher for children from divorced families for either gender, compared to their counterparts from intact families. Comparing the cumulative hazard rate of male children from divorce families to its female counterpart, we see that female children have a lower risk of dropping out (i.e. longer schooling duration). This seems to suggest that male children from divorced families should fare worse.

However, by comparing the cumulative hazard rate of male children from divorced families to male children from intact families, we can see that the two hazard functions eventually converge. The cumulative hazard of schooling duration initially increases at age 14, but once the child reaches 20 years of age, the two hazard rates of male children from intact and divorced families are virtually indistinguishable. This suggests that even though male children from divorced families may be at a higher risk of dropping out during high school, once the child has successfully completed high school, the likelihood that he will continue is not significantly different compared to his counterpart from intact families. On the other hand, the hazard function of female children from divorced families diverges from the hazard rate of female children from intact families. After

the age of 13, a female child from a divorced family is continuously at a higher risk of dropping out of school compared to her counterpart from intact families. The risk of dropping out of school seems to be persistent for girls from divorced families.

To examine the differential effects of divorce on children based on gender, it is important to determine the appropriate comparison groups. By comparing female children from divorced families to their male counterpart, it would seem to suggest that male children from divorced families are more disadvantaged. However, when we compared male children from divorced families to their male counterparts from intact families, it suggests that much of the disadvantage faced by male children from divorced families may diminish over time. There may be some "catch-up" effect for boys from divorced families. However, if we compared female children from divorced families to those from intact families, the disadvantage faced by female children from divorced families seems to actually persist over time.

# 6 Estimation Results

This section presents the empirical results for models developed in section (3). Simple non-parametric hazard estimates presented in the previous section suggest that the effect of parental divorce may be most harmful if it takes place during preschool age and teenage years. In addition, as these children grow up, the effect of divorce may linger for girls but diminish for boys, compared to their counterparts from intact families. However, these findings only show the relationship between one of the observable characteristics and the child's schooling duration, given that the Nelson-Aelan estimator does not accommodate multivariate analysis. As we have mentioned before, child development should be viewed as a dynamic process, and parental divorce and its timing is correlated with a variety of factors that may impact long run child attainment, which we need to account for to properly identify the differential effects of the timing of parental divorce.

As mentioned in previous sections, child unobserved coping abilities may affect child's long

run attainments independently of other unobserved family and child characteristics, which may bias the estimated effects of the timing of divorce upwards if it is not accounted for. Therefore, the estimated effects of divorce and its timing reported here are base on the hazard model with random effects[22]. To extend the analysis in the previous section, we present the estimation results of the hazard model with random effects presented in section (3). The estimation results are presented in the following order. First, to establish a link between parental divorce and the negative child outcomes, the effect of divorce and its timing on child schooling duration is examined for the pooled sample of children. Second, we examine the possible differential impact of divorce on children by gender. There may be gender differences in developmental needs of children and the reactions to parental divorce may be different for boys and girls. Furthermore, the effects of parental divorce on child attainment may change over time. By estimating our hazard model of schooling separately by gender allows us to identify the possible differential effects of timing of divorce and other observed factors on the attainments of children and examine how the effects of divorce timing change as a continuous process over time for children of either sex.

## 6.1 The Effect of Divorce and Its Timing on an "Average Child"

Children from divorced families are found to have lower educational attainments, compared to children who grew up in intact families. In Table (5), estimation results on the pooled sample of children are presented for the two hazard specifications, Weibull and piecewise[23]. The estimated duration dependence parameter of the Weibull distribution, $\sigma$, is compared to the estimated baseline hazard of the piece-wise exponential specification[24], to verify whether the Weibull specification is a reasonable assumption given our data. Furthermore, there is evidence of significant unobserved heterogeneity, since the estimated variance of the random effect, as

[22]For the Weibull specification only.

[23]We expect the estimates of the Weibull specification to be smaller in magnitude compared to that of the piecewise specification due to its more restrictive functional form assumption, therefore the Weibull model does not capture as much variation as the piecewise model.

[24]As illustrated by $\{\lambda_m\}_{m=1}^{M}$ as in equation (3)

given by $\theta$, is significant. This provides justification for incorporating random effects into our hazard model to adjust for the potential upward bias in the estimated effect of parental divorce.

Intuitively, the hazard of schooling should increase with time, given that the longer a child is in school, the closer she is to completing her schooling. Therefore, we expect the hazard rate of schooling to exhibit positive duration dependence. The estimated duration dependence parameter, $\sigma$, is equal to 0.038, in the Weibull model. This indicates that the hazard rate of schooling increases at an increasing rate as a function of time: the longer that a child is in school, she is increasingly more likely to finish her schooling. The second specification, the piecewise model, allows for duration dependence through a set of 22 indicator variables which represents the time intervals. For each time interval, the baseline hazard rate is not allowed to change. These estimates confirm the positive duration dependence of the estimated hazard function suggested by the Weibull model.

The reported coefficients in Table (5) represent the predicted log of schooling duration for each of the observed characteristics. Female children, on average, achieved a higher level of education compared to male children. The interpretation of the reported coefficients deserve some attention. The coefficient on "Female" is 0.010 in the Weibull specification, which means that the predicted log of schooling duration for a female child is 0.010 longer than for another child with the same characteristics, but who is male. Alternatively, the time ratios for each characteristic can also be calculated by exponentiating each of the coefficients. For example, by exponentiating the coefficient on "Female":

$$\exp\{0.014\} = 1.014$$

This means that the schooling duration for a female child is 1.4% longer compared to a male child.

Consistent with previous findings, children from divorced families have lower educational attainments than children from intact families, after controlling for a variety of socioeconomic

background characteristics. Accounting for changes in family resources and a variety of characteristics that may be confounded with divorce, the duration of schooling for an average child from a divorced background is 2.2% shorter than her counterpart from an intact family. To explore the differences in child educational attainment by the timing of parental divorce, we separate the sample of children from divorced families into groups by their age when divorce occurred. Children whose parents divorced when they are in preschool, elementary school, or high school age have a significantly higher risk of dropping out of school. The estimated effect of divorce is most pronounced if it takes place during preschool age: a child whose parents divorce when he is of preschool age have on average 3.2% less schooling than a child raised in an intact family. This is consistent with prior research findings that young children are most vulnerable to family separation. So, by ignoring divorce timing we get an estimated effect of divorce that may be misleading.

We need to keep in mind that the results shown in Table (5) is applicable to our pooled sample of children, i.e. for an average child. As we have mentioned previously, the effect of divorce for children of different ages may be very different for male vs. female children. Investments required for child development, initial response to parental separation, and the ability to cope to stress, may all be gender-biased. The effect of parental separation may have very different implications for children of different ages and gender. We should keep in mind that the estimated effect reported in this section should be interpreted as the "gross" effect of divorce on an average child, which may be confounded with potentially beneficial effects associated with the timing of divorce. Early withdraw of children from stressful family environment, longer period of time to heal, may be potentially beneficial to child development. Self-selection into divorce may yield some benefits greater than would be predicted by the observable characteristics. To separately identify the "net" effect of divorce, we need to account for all the confounding factors correlated with divorce that may explain the observed disadvantages of children from divorced families. In the next section, we will further explore differential effects of the timing of parental divorce by gender. We explore the effect of parental divorce on child development as a continuous process,

identify the "net" effect of divorce, and examine how these effects change over time.

## 6.2 Gender Differences in the Effects of Parental Divorce

Are boys and girls "equally" affected by divorce? Are boys and girls "similarly" affected by divorce, meaning, how do these effects change over time for boys compared to girls? Based on existing research, the answers to these questions are inconclusive. In order to answer these questions, we employ a hazard model with random effects to examine the impact of divorce and its timing on children as a continuous process separately by gender, and explore how these effects evolve over time. All estimation results are based on the hazard models with random effects, which controls for unobserved heterogeneity with respect to child abilities. Table (6) and (7) presents the estimation results for the sub-samples of male and female children, respectively.

Mother's level of education, the number of siblings in the household, and family income are controls for the quantity and quality of home investments available for children. In examining the impact of divorce on children, together they serve as controls for the possible socioeconomic disadvantages correlated with family dissolution at each point in time. Family income is taken as a proxy for the amount of material resources available for the child. For a given level of family income, children compete for the available time and material resources available from the parents. Controlling for family income and number of siblings, mother's education proxies for the quality of parental time investments for the child, assuming that the mother is the primary care-giver. As we can see, having a highly-educated mother, fewer siblings, and higher family income are all associated with higher attainments of children of either sex.

Early divorce may be beneficial if it withdraws a child from an stressful environment that may damage child development, and allows for the longer healing period from the initial shock of parental separation. In addition, unobserved quality of the marriage may influence the likelihood of divorce and have lingering impacts on children. Without controlling for the endogeneity of divorce, the estimated effects of divorce may be overstated. As mentioned before, we attempt to address this bias by using the duration of parents' marriage as an instrument for the quality

of their marriage. Therefore, to identify the "net" effect of divorce timing, we must account for these potentially opposing effects confounded with the timing of separation.

The estimated effects of the timing of divorce reported in the first column of Tables (6) and (7) should be interpreted as the "gross" effects of divorce, since the estimates are confounded with these potentially opposing effects correlated with the timing of divorce. By allowing for these opposing effects to interact, we see that by controlling for unobserved child abilities, parental divorce is harmful for boys during preschool and teenage years, and the effects are almost identical: Based on the estimates of the weibull specification, boys whose parents divorce during preschool have 2.8% less schooling and boys who experienced parental divorce when they are teenagers have 2.9% less schooling, on average, compared to their male counterparts from intact families. For girls, the gross effect of divorce is most pronounced during childhood. Girls who experienced parental divorce during childhood have on average 2.2% to 3.8% less schooling than their female counterparts from intact families, holding all other characteristics constant.

The quality of marital durations may affect the likelihood of divorce and impact child development. For the marriages that eventually end in divorce and conditional on its timing, we use the duration of the marriage as an instrument for the quality of the marriages, assuming that the most troubled marriages would end early. It has been argued that parents may postpone the divorce if they consider the separation to be very harmful to their children. However, I would argue that divorce may not be postponed if the quality of marital relations are sufficiently poor. In addition, we should note that the length of marriage need not be perfectly positively correlated with the timing of divorce, given that a couple may be married for a prolonged period of time before childbearing[25]. Therefore, controlling for the timing of divorce, marriage duration should at least partially control for the endogeneity of divorce.

The second columns of Table (6) and (7) presents the estimated effects of the timing of divorce after controlling for the duration of marriage for the divorced sub-sample. Children

[25]For example, given two children whose parents divorced when they are of the same age, the duration of their parents' marriage need not be the same. In the present context, we assume that the one whose parents have been married longer prior to divorce should be exposed to less stress due to the relative stability of the marital relations.

whose parents divorce when they are more than 18 years of age share similar risks of dropping out of school compared to those from intact families, given that marriages that lasted that much longer should be fairly stable and may not be very different from those that stayed intact. This finding provides support for our assumption that marriage duration proxies for the unobserved quality of the marriage.

Although the effect of marriage duration on child educational attainment is insignificant independently, there is evidence that divorce may be endogenous, at least for girls. The effect associated with marriage duration, albeit insignificant, is negative, which suggests that girls may not benefit from parents staying together if marital relations are fairly poor. For female children, there is evidence that the effect of divorce is downward biased due to endogeneity, given that the estimated effects of divorce are reduced in magnitude and become insignificant once we control for the duration of marriage.

However, the same cannot be concluded for boys. The coefficient on parental marriage duration is positive, albeit insignificant, which may suggest that boys benefit from their parents staying together longer, even if marital relations are very troubled. The reduction in resources as a consequence of divorce may be a relatively more important determinant of their educational outcomes. The development of boys may depend more crucially on the amount of available family resources, given that the estimated effect of family income on child schooling duration is both larger and more significant for boys than for girls. Therefore, this evidence may be suggesting that even if marital relations are poor, boys still yield a net benefit from having their parents stay together.

Based on our estimated results, Figure (6) presents comparisons of the predicted hazard functions for two children from divorced backgrounds based on different parental marriage duration, for males and females, respectively. Consider two children who are of the same age when their parents divorced. Holding everything else constant, we vary only marriage duration for the two children to be 5 years and 10 years respectively. Longer marriage duration is associated with lower hazard rate of dropping out of school for male children; and the opposite is true for

female children.

Finally, we examine the changes in the effects of divorce on child educational attainment over time. Based on the findings of the "gross" effect of divorce, it seems to suggest that the negative effect of parental divorce is the most pronounced when the child is very young. However, earlier divorce is also associated with a longer healing period, which may be beneficial for child development. Given that divorce is a negative shock to the family, the length of time a child has to adjust and cope with the shock may be positively correlated with long-run outcomes. Length of time since parental divorce proxies for the amount of time a child has to readjust after the initial shock of divorce. To control for this potentially opposing effect associated with divorce, we further control for time since parental divorce as a time-varying covariate. After controlling for all these factors potentially correlated with divorce timing, the estimated effects presented in column 3 represent the "net" effect of divorce at each point in time.

The longer the period of time since parental divorce, the longer the healing period. For children of either gender, each additional year since parental divorce is associated with longer schooling duration, hence higher educational attainments. For boys and girls, each additional year that passed since the initial separation is associated with 1.8% and 1.5% longer schooling duration, respectively. Without controlling for this positive effect associated with earlier divorce, the estimated effect of parental divorce may be understated. As we can see, the negative effects of divorce at each point in time is magnified once we control for these benefits associated with time since parental divorce.

Given that the length of time since parental divorce is necessarily longer the earlier the divorce takes place, which is measured by the child's age at the time of divorce, the results here may seem contradictory at first. If we assume that the initial shock effect of divorce is constant regardless of its timing, than we would expect that the longer the time since divorce (i.e. the earlier the divorce took place), the more beneficial it would be for the long run outcomes of the child. However, in column 3 of Table (6) and (7), we can see that the initial shock of divorce is not constant regardless of timing. Once we control for the length of time since parental divorce,

the marginal increase in the magnitude of the effect of divorce is larger the earlier the divorce took place. This implies that the initial shock of divorce is greater in magnitude the earlier the divorce took place. Therefore, even if earlier divorce is associated with a longer time to heal, the initial shock is large enough that it never completely diminish as time goes by.

The initial shock of divorce, or the "net" effect of divorce, is found to be negative and significant for both boys and girls during childhood and early adolescence. Although the initial shock of divorce is larger for boys, the findings suggest that boys re-adjust faster from the initial shock of parental separation, which may explain the observed lingering effect of divorce on girls. Comparing the results in column 1 and 3, we see that the "gross" effect of divorce is smaller, but the "net" effect of divorce once we control for marriage duration and time since divorce is larger for boys than for girls. More specifically, there is a larger reduction in the magnitude of the "net" effect of divorce once we allow the benefit of healing time to interact with the divorce timing for boys.

Figure (7) presents the predicted hazard functions for two children of the same age, but one whose parents divorced in when he was 5, and one whose parents divorced when he is 15. The child whose parents divorce when he was 5 necessarily have a longer period to heal retrospectively. As we can see, even though the initial shock of divorce has a larger effect on boys, they also are able to re-adjust faster. Therefore, the hazard of dropping out of school is similar for male children regardless of the timing of divorce. However, even though the initial shock of divorce is not as large for girls, they also heal at a slower pace and therefore never completely recovers. As we can see, the schooling hazard is significantly higher for girls whose parents divorced when they were young.

The order of the marriage of the mother becomes significant for male children, after we control for both the duration of parents' marriage and the length of time since parental divorce. As we have mentioned before, it is more important for boys that their parents stay together, since parental investments are relatively more important for boys development. Couples in which one or both spouse had been married before may be less attached to the marriage and

more likely to divorce again. Higher order marriages are also associated with more step-siblings who compete for the same amount of family resources and parental time investments. After controlling for the beneficial effects of marriage duration and length of time since divorce, we see that male children whose mother is married for the second time when she married his birth father has 3% less educational attainment, compared to a male child whose mother is married for the first time when upon marriage to his birth father.

Due to data limitations, there is a group of children in our sample for whom we do not have any of their fathers' information. However, we believe that father's level of information is an important determinant of child educational attainment. Therefore, in estimating our hazard models, an indicator variable was created to examine whether the group of children with no father's information available is significantly different. In addition, we set all the missing values of the father's schooling to be equal to 0. We expect the coefficient on "father's information missing" to be positive, since even though the father's schooling information is missing, the relationship between father's schooling of these missing fathers and their child's education attainment should be positive. However, the fact that the coefficient is positive and statistically significant is counter-intuitive. Father's information missing is positively correlated with early divorces. The correlation coefficient between "father information missing" and "divorce", and "age at parental divorce" are 0.4786 and $-0.2528$, respectively (Both statistically significant at the 5% level). The fathers of early divorces are more likely to have attrited, and therefore no longer observed in the data set. It is counter-intuitive that the group of children for whom we cannot identify their fathers should have significantly longer schooling duration, given that father's presence is an important determinant of the educational attainment of male children.

There are significant gender differences, both in terms of the effects of timing of parental divorce on child educational outcomes, and also with respect to changes in these effects on child development over time. The finding suggests that it's relatively more important for a male child that his parents stay together, given that parental investments and available resources may be relatively more important for the development of male children. However, parents staying

together may not benefit female children the same way as for male children, and parental investments may not be sufficient enough to compensate for other unobserved factors that are correlated with divorce. By modeling the accumulation of human capital of children and examine the effect of family structure on child attainments in a hazard framework, we are able to explore the changes in the effect of divorce on child development over time. The findings suggest that after controlling for unobserved heterogeneity of child ability, the initial effect of divorce is larger for boys, but boys also re-adjust faster. Although the initial shock of divorce is smaller for girls, they also re-adjust slowly and never fully recovers from the initial shock in the long run. Therefore, it appears that the negative "gross" effect of divorce is larger for girls.

# 7 Summary and Conclusions

This study employs a hazard model of schooling duration to examine the effect of parental divorce on child educational attainment. Child development is viewed as a dynamic process, which is influenced by the choice of parental investments at each point in time. The quantity and quality of available resources that can be devoted to the production of child quality is constrained by family structure. Therefore, changes in family structure will have likely implications on the long term attainments of children. The effects of parental divorce are explored as a continuous process, which may have different implications on child attainments based on its timing and child gender. In addition, there may be significant variations in unobserved child ability, either intellectual or emotional, that can affect the process of child development independently from the observed family and child characteristics. Therefore, we further incorporate random effects in our duration model to address potential child unobserved heterogeneity.

Parental divorce has a significant impact on the educational attainment of children. On average, children from divorced families have 2.2% less schooling than children raised in intact families. However, there are significant gender differences in the effect of divorce based on timing. Male children who experienced parental divorce during preschool and teenage years are

at a higher risk of dropping out of school on average, then male children whose parents stay together; female children, on the other hand, are at risk if their parents divorce during their childhood, or before the age of 11.

Given that the parents' marriage eventually ends in divorce, the findings suggest that male children benefit from their parents staying together longer, while the same cannot be said for girls. This may be due to gender differences in terms of the relative importance of available resources and the impact of a stressful environment for the developmental needs of children. Given poor marital relations, it may be more important for boys that their parents try to stay together, given that available resources may be more important for the attainment of boys and the higher level of resources available associated with being in an intact family. On the other hand, given that marital relations are poor, girls may not benefit from the parents staying together longer, since prolonged exposure to a stressful family environment may be more damaging to the development of female children[26].

There is evidence suggesting that the negative effect of divorce is comparably more permanent for girls than for boys. Although the initial shock of divorce may be larger for boys, boys are also found to re-adjust faster than girls. The magnitude of the negative effects of divorce diminishes over time at a faster rate for male vs. female children. As we have suggested, available family resources are more important for the development of boys. The lost of resources due to the family break-up may be recovered by the subsequent remarriages of the parents, which may explain the faster rate of recovery for boys. On the other hand, girls may be more sensitive to stressful events such as parental separation, and the effects of which may not be fully compensable by time and goods resource. Therefore, even though the initial effect of divorce is smaller for girls, the effect tend to linger over time. This is the first study to draw conclusive evidence of negative outcomes for female children.

Our finding suggest that the long term impact of divorce on children's educational attainment may be different based on age at parental divorce and child gender, and there are significant

[26]Possibly the form of dysfunction is typically for female children.

variation in terms of child unobserved abilities that may independently explain the observed outcome. Furthermore, given that there is weak evidence that the effect of divorce may be causal, and the channels through which they operate is important to design policies to help these children recover from the shock of the separation. Policies that discouraged divorce based on a generalized finding of a negative association between divorce and child outcomes, without consideration for the diversity of child experiences may be more harmful for some groups of children. Discretion is needed designing policies to help children experienced with parental divorce to overcome the observed disadvantages.

Male and female children have very different experiences growing up depending on the timing of the separation. Although initial effect of divorce is found to be negative and significant for children of, much of this impact diminishes over time. However, the process of which differs by child gender. It may be that family resources are more important for the development of male children. Therefore, to help boys to recover from the initial shock of family separation, social assistance programs for divorced parents with male children may be effective. Financial assistance or tax breaks for single divorced parents, mentor programs that provide guidance and counselling may help boys to overcome the observed disadvantages.

Divorce during childhood has an adverse effect on girls. For girls from divorced families, re-adjustment tend to be slow and additional family resources may not compensate for the damage induced by parental divorce. There's weak evidence supporting that delayed parental divorce may not help girls. In families with very high level of tension, such as those with physical abuse, allowing for divorce may improve outcomes. Therefore, the welfare of female children from a substantial proportion of divorced families may be jeopardized if access to divorce is restricted or denied. Our findings suggest that emotional support and guidance may be more important for the development of female children, psychological counselling may be the key to help girls recover from the family breakup.

Our results should be interpreted with caution, given that the effect of divorce may be overestimated since divorce may be endogenous to unobserved family characteristics that affect

child long term attainments and increase the likelihood of divorce. We attempt to control for the endogeneity of divorce by accounting for socioeconomic disadvantages and use marital duration as an instrument for the tension that may have pre-existed in the family. This analysis suggests that the quality of the marriage does affect both the divorce decision and child long run outcomes. As a consequence, there may be other unobserved factors have not been captured by our model. Therefore, we cannot conclude definitively the estimated effects of divorce is *causal*. To properly address the endogeneity of divorce, the production of child attainments and the divorce decision should be modelled simultaneously, which will be explored in future research (See Liu [37, 2003]). To address the endogeneity of divorce, in another paper I model parental divorce and child educational outcomes jointly in a two-equation hazard model, which allows the same unobserved family characteristics to affect both the divorce decision and impact the long run attainment of children.

These results measure the negative associations between parental divorce and observed child educational attainments, but they should not be taken as the counterfactuals, since they cannot predict what the outcomes would have been if the tensions in the family increased and parents are discouraged to divorce. Parents know more about their situation than policy-makers, and self-selection into divorce will always have some unobserved benefits that cannot be captured by the observed factors. Recent trend in legislations have gear towards restricting access to divorce based on the negative associations found in previous research. Establishing the causal link between parental divorce and child outcomes is important, since making divorce more difficult to obtain may not help children if these disadvantages are not actually *caused* by divorce. Nevertheless, based on our findings of the diversity in the experiences and the effects of divorce on children of all ages and gender, policies based on simple generalizations on the implications of divorce on child outcomes are likely to stigmatize a large proportion of the population, given that many American youths will experience parental divorce while growing up.

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Table 1: Variable Definition (By Alphabetical Order)

| 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 $\geq 18$; 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 children from divorced families only) |
| Family Income | Continuous | Yes | Sum of income of all household members measured in 1984 constant dollars |
| Father's Level of 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 Level of 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 |
| Number of Siblings | Continuous | No | Number of Siblings |
| 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 categorized as "North East" according to PSID regional codes;0 otherwise |
| Region of Residence: North Central | Binary | Yes | = 1 if state of residence is categorized as "North Central" according to PSID regional codes;0 otherwise |
| Region of Residence: South | Binary | Yes | = 1 if state of residence is categorized as "South" according to PSID regional codes;0 otherwise |
| Region of Residence: West | Binary | Yes | = 1 if state of residence is categorized as "West" according to PSID regional codes;0 otherwise |
| Urban Residence | Binary | Yes | = 1 if the population of the largest city in the county of residence exceeds $100,000$; 0 otherwise |

Table 2: Descriptive Statistics (Pooled Sample)

| Variable | $N$ | Mean | Std. Dev. | Min | Max |
|---|---|---|---|---|---|
| *Dependent Variable* | | | | | |
| Grades of Schooling Attained | 901 | 12.385 | 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 |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | 901 | 0.131 | 0.338 | 0 | 1 |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | 901 | 0.034 | 0.182 | 0 | 1 |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | 901 | 0.039 | 0.193 | 0 | 1 |
| Age at Parental Divorce (Age $\geq 18$) | 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 |
| Race: Black | 901 | 0.270 | 0.444 | 0 | 1 |
| Race: Other | 901 | 0.033 | 0.180 | 0 | 1 |
| Mother's Level of Education | 901 | 12.522 | 2.087 | 0 | 18 |
| Mother's Age at Childbirth | 901 | 24.984 | 5.723 | 14 | 44 |
| Mother's Age at Marriage | 901 | 20.087 | 3.239 | 11 | 38 |
| Order of the Marriage[1] | 901 | 1.049 | 0.226 | 1 | 3 |
| Father's Level of Education[2] | 784 | 12.585 | 2.904 | 0 | 19 |
| Father's Information Missing | 901 | 0.130 | 0.336 | 0 | 1 |
| Number of Siblings | 901 | 2.072 | 2.058 | 0 | 11 |
| Family Income[3] | 901 | 31,937 | 20,712 | 4,624 | 335,352 |
| Urban Residence[4] | 901 | 0.521 | 0.500 | 0 | 1 |
| Region of Residence[†]: North East | 901 | 0.165 | 0.372 | 0 | 1 |
| Region of Residence[†]: North Central | 901 | 0.280 | 0.449 | 0 | 1 |
| Region of Residence[†]: South | 901 | 0.401 | 0.490 | 0 | 1 |
| Region of Residence[†]: West | 901 | 0.154 | 0.361 | 0 | 1 |
| Birth Cohort (1968) | 901 | 0.189 | 0.391 | 0 | 1 |
| Birth Cohort (1969) | 901 | 0.204 | 0.403 | 0 | 1 |
| Birth Cohort (1970) | 901 | 0.203 | 0.403 | 0 | 1 |
| Birth Cohort (1971) | 901 | 0.190 | 0.392 | 0 | 1 |
| Birth Cohort (1972) | 901 | 0.214 | 0.410 | 0 | 1 |

*Notes:* 1. The order of the marriage defined based on the marriage history of the mother. 2. Information on the biological father is missing for a sub-sample of children ($n = 117$). The mean 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. [†] Corresponding to PSID regional categories for each of the 50 states.

Table 3: Descriptive Statistics: Comparing Children by Gender

| | Male | | Female | |
|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. |
| *Dependent Variable* | | | | |
| Grades of Schooling Attained | 12.312 | 1.470 | 12.467 | 1.565 |
| | | | | |
| *Explanatory Variables* | | | | |
| Race: White | 0.675 | 0.469 | 0.722 | 0.449 |
| Race: Black | 0.289 | 0.454 | 0.248 | 0.432 |
| Race: Other | 0.036 | 0.186 | 0.031 | 0.173 |
| Mother's Level of Education | 12.621 | 2.026 | 12.410 | 2.151 |
| Mother's Age at Childbirth | 25.157 | 5.723 | 24.790 | 5.723 |
| Mother's Age at Marriage | 20.151 | 3.169 | 20.014 | 3.319 |
| Order of the Marriage[1] | 1.048 | 0.224 | 1.050 | 0.228 |
| Father's Level of Education[2] | 12.742 | 2.816 | 12.413 | 2.992 |
| Father's Information Missing | 0.138 | 0.346 | 0.120 | 0.326 |
| Number of Siblings | 2.075 | 1.978 | 2.068 | 2.147 |
| Family Income[3] | 31,466 | 15,356 | 32,466 | 25,433 |
| Urban Residence[4] | 0.539 | 0.499 | 0.502 | 0.501 |
| Region of Residence†: North East | 0.168 | 0.374 | 0.163 | 0.370 |
| Region of Residence†: North Central | 0.273 | 0.446 | 0.288 | 0.453 |
| Region of Residence†: South | 0.400 | 0.490 | 0.401 | 0.491 |
| Region of Residence†: West | 0.159 | 0.366 | 0.149 | 0.356 |
| Birth Cohort (1968) | 0.195 | 0.397 | 0.182 | 0.386 |
| Birth Cohort (1969) | 0.214 | 0.410 | 0.193 | 0.395 |
| Birth Cohort (1970) | 0.195 | 0.397 | 0.212 | 0.409 |
| Birth Cohort (1971) | 0.180 | 0.385 | 0.200 | 0.401 |
| Birth Cohort (1972) | 0.216 | 0.412 | 0.212 | 0.409 |
| | | | | |
| Parents Divorced | 0.344 | 0.475 | 0.337 | 0.473 |
| Age at Parental Divorce ($0 \leq$ Age $< 5$) | 0.084 | 0.277 | 0.106 | 0.308 |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | 0.130 | 0.337 | 0.132 | 0.339 |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | 0.042 | 0.201 | 0.026 | 0.159 |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | 0.031 | 0.175 | 0.047 | 0.212 |
| Age at Parental Divorce (Age $\geq 18$) | 0.057 | 0.231 | 0.026 | 0.159 |
| (Mean) Age at Parental Divorce‡ | 9.524 | 5.893 | 8.301 | 5.713 |
| Duration of Parents' Marriage‡ | 14.573 | 7.094 | 12.979 | 6.987 |
| *N* | 477 | | 424 | |

*Notes:* 1. The order of the marriage defined based on the marriage history of the mother. 2. Information on the biological father is missing for 66 male and 51 female children. The mean is computed for the group of children of each sex 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. † Corresponding to PSID regional categories for each of the 50 states. ‡ Mean value calculated based on the sub-sample of children from divorced families.

Table 4: Descriptive Statistics: Children from Intact vs. Divorced Families

| | Intact | | Divorced | |
|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. |
| *Dependent Variable* | | | | |
| Grades of Schooling Attained | 12.539 | 1.471 | 12.088 | 1.560 |
| | | | | |
| *Explanatory Variables* | | | | |
| Female | 0.473 | 0.500 | 0.466 | 0.500 |
| Race: White | 0.699 | 0.459 | 0.694 | 0.462 |
| Race: Black | 0.266 | 0.442 | 0.277 | 0.448 |
| Race: Other | 0.035 | 0.185 | 0.029 | 0.169 |
| Mother's Level of Education | 12.443 | 2.046 | 12.674 | 2.160 |
| Mother's Age at Childbirth | 25.911 | 6.051 | 23.192 | 4.525 |
| Mother's Age at Marriage | 20.485 | 3.371 | 19.316 | 2.816 |
| Order of the Marriage[1] | 1.051 | 0.234 | 1.046 | 0.209 |
| Father's Level of Education[2] | 12.671 | 2.807 | 12.323 | 3.177 |
| Father's Information Missing | 0.003 | 0.058 | 0.375 | 0.485 |
| Number of Siblings | 2.120 | 2.156 | 1.980 | 1.854 |
| Family Income[3] | 33,921 | 22,879 | 28,097 | 14,995 |
| Urban Residence[4] | 0.525 | 0.500 | 0.515 | 0.501 |
| Region of Residence[†]: North East | 0.189 | 0.391 | 0.121 | 0.326 |
| Region of Residence[†]: North Central | 0.268 | 0.443 | 0.303 | 0.460 |
| Region of Residence[†]: South | 0.397 | 0.490 | 0.407 | 0.492 |
| Region of Residence[†]: West | 0.146 | 0.354 | 0.169 | 0.376 |
| Birth Cohort (1968) | 0.195 | 0.397 | 0.176 | 0.381 |
| Birth Cohort (1969) | 0.192 | 0.394 | 0.228 | 0.420 |
| Birth Cohort (1970) | 0.189 | 0.391 | 0.231 | 0.422 |
| Birth Cohort (1971) | 0.195 | 0.397 | 0.179 | 0.384 |
| Birth Cohort (1972) | 0.229 | 0.421 | 0.186 | 0.389 |
| | | | | |
| Age at Parental Divorce ($0 \leq$ Age $< 5$) | | | 0.277 | 0.448 |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | | | 0.384 | 0.487 |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | | | 0.101 | 0.302 |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | | | 0.114 | 0.318 |
| Age at Parental Divorce (Age $\geq 18$) | | | 0.124 | 0.330 |
| (Mean) Age at Parental Divorce | | | 8.954 | 5.832 |
| Duration of Parents' Marriage | | | 13.831 | 7.078 |
| $N$ | 594 | | 307 | |

*Notes:* 1. The order of the marriage defined based on the marriage history of the mother. 2. Information on the biological father is missing for 2 children from intact families, and 115 children from divorced families. The mean 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. [†] Corresponding to PSID regional categories for each of the 50 states.

Table 5: Estimated Effects on the Predicted Log of Schooling Duration: Pooled Sample

| | Weibull | | Piece-wise | |
|---|---|---|---|---|
| Female | 0.010* | 0.010* | 0.124* | 0.125* |
| | (0.005) | (0.005) | (0.050) | (0.050) |
| Parents Divorced | −0.022* | | −0.189* | |
| | (0.006) | | (0.069) | |
| *Timing of Parental Divorce*[1.] | | | | |
| Age at Parental Divorce ($0 \leq$ Age $< 5$) | | −0.032* | | −0.295* |
| | | (0.010) | | (0.106) |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | | −0.019* | | −0.165* |
| | | (0.008) | | (0.091) |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | | −0.020 | | −0.120 |
| | | (0.014) | | (0.150) |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | | −0.023* | | −0.180 |
| | | (0.011) | | (0.123) |
| Age at Parental Divorce (Age $\geq 18$) | | −0.014 | | −0.124 |
| | | (0.012) | | (0.136) |
| *Duration Dependence* | | | | |
| $\sigma$ | 0.038* | 0.038* | | |
| | (0.002) | (0.002) | | |
| Birth to Grade 6 | | | −20.163* | −20.210* |
| | | | (0.356) | (0.358) |
| Grade 7 | | | −5.241* | −5.288* |
| | | | (1.082) | (1.085) |
| Grade 8 | | | −3.847* | −3.894* |
| | | | (0.609) | (0.609) |
| Grade 9 | | | −1.927* | −1.974* |
| | | | (0.409) | (0.410) |
| Grade 10 | | | −1.226* | −1.271* |
| | | | (0.392) | (0.393) |
| Grade 11 | | | −0.724* | −0.769* |
| | | | (0.364) | (0.366) |
| Grade 12 | | | 1.072* | 1.028* |
| | | | (0.353) | (0.355) |
| College (Freshman) | | | 0.946* | 0.903* |
| | | | (0.365) | (0.367) |
| College (Sophomore) | | | 1.322* | 1.284* |
| | | | (0.365) | (0.367) |
| College (Junior) | | | 1.180* | 1.140* |
| | | | (0.381) | (0.383) |
| College (Senior) | | | 1.857* | 1.819* |
| | | | (0.375) | (0.378) |
| Beyond College | | | 2.027* | 2.001* |
| | | | (0.349) | (0.350) |
| $\theta$ | 1.148* | 1.152* | | |
| | (0.138) | (0.140) | | |
| N | 901 | 901 | 901 | 901 |
| Log Likelihood | 1042.59 | 1044.01 | 807.50 | 808.30 |

*Notes:* 1. In reference to "Children from Intact Families"; 2. Robust standard errors reported in parenthesis. 3. Significance level reported by Z - statistics: * – 5% level. 4. Additional controls include: ln(Real Family Income), mother's level of education, father's level of education, number of siblings, mother's age at childbirth, mother's age at marriage, order of the marriage, race (black, other), birth cohort (1969, 1970, 1971, 1972), urban residence, region of residence (south, west, north central).

Table 6: Estimated Effects on the Predicted Log of Schooling Duration: Male Sub-Sample

| | Weibull | | | Piecewise | | |
|---|---|---|---|---|---|---|
| Mother's Years of Schooling | 0.005* | 0.005* | 0.005* | 0.055* | 0.055* | 0.051* |
| | (0.002) | (0.002) | (0.002) | (0.020) | (0.020) | (0.020) |
| Mother's Age at Child Birth | 0.001 | 0.001 | −0.001 | 0.018* | 0.013 | 0.009 |
| | (0.001) | (0.001) | (0.001) | (0.008) | (0.008) | (0.008) |
| Mother's Age at Marriage | 0.001 | 0.002 | 0.002* | 0.002 | 0.010 | 0.016 |
| | (0.001) | (0.002) | (0.001) | (0.017) | (0.016) | (0.016) |
| Father's Years of Schooling | 0.005* | 0.004* | 0.004* | 0.048* | 0.046* | 0.049* |
| | (0.001) | (0.002) | (0.001) | (0.016) | (0.016) | (0.016) |
| Father's Information Missing | 0.059* | 0.056* | 0.048* | 0.681* | 0.666* | 0.680* |
| | (0.022) | (0.022) | (0.019) | (0.241) | (0.239) | (0.234) |
| Order of This Marriage (Mother) | −0.031* | −0.031* | −0.029* | −0.222 | −0.240 | −0.308* |
| | (0.018) | (0.018) | (0.016) | (0.187) | (0.187) | (0.178) |
| Number of Siblings | −0.012* | −0.012* | −0.011* | −0.068* | −0.070* | −0.073* |
| | (0.003) | (0.003) | (0.002) | (0.032) | (0.032) | (0.032) |
| ln (Real Family Income) | 0.007* | 0.007* | 0.006* | 0.116* | 0.118* | 0.110* |
| | (0.002) | (0.002) | (0.002) | (0.038) | (0.039) | (0.040) |
| Duration of Parents' Marriage | | 0.002 | 0.004* | | 0.022 | 0.033* |
| | | (0.002) | (0.001) | | (0.017) | (0.017) |
| Time Since Parental Divorce | | | 0.018* | | | 0.105* |
| | | | (0.002) | | | (0.032) |
| **Timing of Parental Divorce**[a] | | | | | | |
| Age at Parental Divorce ($0 \leq$ Age $< 5$) | −0.028* | −0.043* | −0.301* | −0.284* | −0.482* | −2.061* |
| | (0.014) | (0.020) | (0.039) | (0.149) | (0.227) | (0.552) |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | −0.016 | −0.036 | −0.226* | −0.160 | −0.446* | −1.597* |
| | (0.011) | (0.023) | (0.033) | (0.131) | (0.254) | (0.435) |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | −0.003 | −0.029 | −0.158* | −0.010 | −0.370 | −1.157* |
| | (0.017) | (0.031) | (0.032) | (0.182) | (0.331) | (0.419) |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | −0.029* | −0.062* | −0.147* | −0.202 | −0.636 | −1.155* |
| | (0.016) | (0.037) | (0.034) | (0.207) | (0.422) | (0.442) |
| Age at Parental Divorce (Age $\geq 18$) | −0.015 | −0.015 | −0.015 | −0.215 | −0.221 | −0.248 |
| | (0.014) | (0.014) | (0.013) | (0.155) | (0.153) | (0.152) |
| Constant | 2.678* | 2.681* | 2.704* | | | |
| | (0.041) | (0.041) | (0.037) | | | |
| $\sigma$ | 0.039* | 0.038* | 0.033* | | | |
| | (0.003) | (0.003) | (0.002) | | | |
| $\theta$ | 1.057* | 1.072* | 1.236* | | | |
| | (0.174) | (0.178) | (0.198) | | | |
| Log Likelihood | 568.34 | 568.83 | 588.16 | 523.89 | 525.00 | 538.36 |
| N | 477 | 477 | 477 | 477 | 477 | 477 |

*Notes:* 1. Robust standard errors reported in parenthesis. 2. Significance levels reported by Z statistic: $*$ — 5% level. 3. Additional controls include: (*i*) Geographical region of residence (South, West, North Central); (*ii*) Racial background (Black, Other); (*iii*) Birth cohort indicators (1969, 1970, 1971, 1972); (*iv*) Urban residence indicator. [a] All baseline estimates (by timing of parental divorce) are in reference to children from intact families.

Table 7: Estimated Effects on the Predicted Log of Schooling Duration: Female Sub-Sample

| | Weibull | | | Piecewise | | |
|---|---|---|---|---|---|---|
| Mother's Years of Schooling | 0.004* | 0.004* | 0.004* | 0.035* | 0.039* | 0.039* |
| | (0.002) | (0.002) | (0.002) | (0.021) | (0.020) | (0.020) |
| Mother's Age at Child Birth | −0.001 | −0.001 | −0.001 | −0.004 | 0.001 | 0.001 |
| | (0.001) | (0.001) | (0.001) | (0.009) | (0.010) | (0.010) |
| Mother's Age at Marriage | 0.002* | 0.002 | 0.002* | 0.025 | 0.019 | 0.019 |
| | (0.001) | (0.001) | (0.001) | (0.018) | (0.018) | (0.018) |
| Father's Years of Schooling | 0.002 | 0.002 | 0.002 | 0.008 | 0.010 | 0.010 |
| | (0.002) | (0.002) | (0.002) | (0.016) | (0.016) | (0.016) |
| Father's Information Missing | 0.049* | 0.048* | 0.047* | 0.184 | 0.202 | 0.205 |
| | (0.024) | (0.024) | (0.023) | (0.230) | (0.233) | (0.232) |
| Order of This Marriage (Mother) | −0.007 | −0.007 | −0.009 | −0.123 | −0.110 | −0.105 |
| | (0.018) | (0.018) | (0.016) | (0.202) | (0.203) | (0.205) |
| Number of Siblings | −0.019* | −0.019* | −0.016* | −0.121* | −0.121* | −0.121* |
| | (0.003) | (0.003) | (0.003) | (0.036) | (0.036) | (0.036) |
| ln (Real Family Income) | 0.004 | 0.004 | 0.003 | 0.074* | 0.076* | 0.076* |
| | (0.002) | (0.002) | (0.002) | (0.026) | (0.027) | (0.027) |
| Duration of Parents' Marriage | | −0.001 | 0.001 | | −0.026 | −0.024 |
| | | (0.002) | (0.002) | | (0.017) | (0.018) |
| Time Since Parental Divorce | | | 0.015* | | | 0.016 |
| | | | (0.003) | | | (0.039) |
| **Timing of Parental Divorce**[a] | | | | | | |
| Age at Parental Divorce ($0 \leq$ Age $< 5$) | −0.039* | −0.027 | −0.255* | −0.353* | −0.149 | −0.406 |
| | (0.014) | (0.019) | (0.047) | (0.150) | (0.199) | (0.667) |
| Age at Parental Divorce ($5 \leq$ Age $< 11$) | −0.023* | −0.003 | −0.167* | −0.185 | −0.148 | −0.039 |
| | (0.012) | (0.025) | (0.038) | (0.132) | (0.247) | (0.525) |
| Age at Parental Divorce ($11 \leq$ Age $< 13$) | −0.038 | −0.015 | −0.123* | −0.286 | 0.110 | −0.023 |
| | (0.024) | (0.036) | (0.037) | (0.277) | (0.371) | (0.514) |
| Age at Parental Divorce ($13 \leq$ Age $< 18$) | −0.022 | 0.006 | −0.071* | −0.212 | 0.291 | 0.200 |
| | (0.014) | (0.035) | (0.035) | (0.159) | (0.367) | (0.427) |
| Age at Parental Divorce (Age $\geq 18$) | −0.003 | −0.002 | −0.007 | 0.024 | 0.016 | 0.013 |
| | (0.023) | (0.023) | (0.020) | (0.254) | (0.252) | (0.252) |
| Constant | 2.793* | 2.789* | 2.795* | | | |
| | (0.045) | (0.045) | (0.042) | | | |
| $\sigma$ | 0.037* | 0.037* | 0.031* | | | |
| | (0.003) | (0.003) | (0.003) | | | |
| $\theta$ | 1.266* | 1.253* | 1.554* | | | |
| | (0.251) | (0.251) | (0.301) | | | |
| Log Likelihood | 488.72 | 489.11 | 499.00 | 457.70 | 458.81 | 463.37 |
| N | 424 | 424 | 424 | 424 | 424 | 424 |

*Notes:* 1. Robust standard errors reported in parenthesis. 2. Significance levels reported by Z statistic: $*$ — 5% level. 3. Additional controls include: (*i*) Geographical region of residence (South, West, North Central); (*ii*) Racial background (Black, Other); (*iii*) Birth cohort indicators (1969, 1970, 1971, 1972); (*iv*) Urban residence indicator. [a] All baseline estimates (by timing of parental divorce) are in reference to children from intact families.

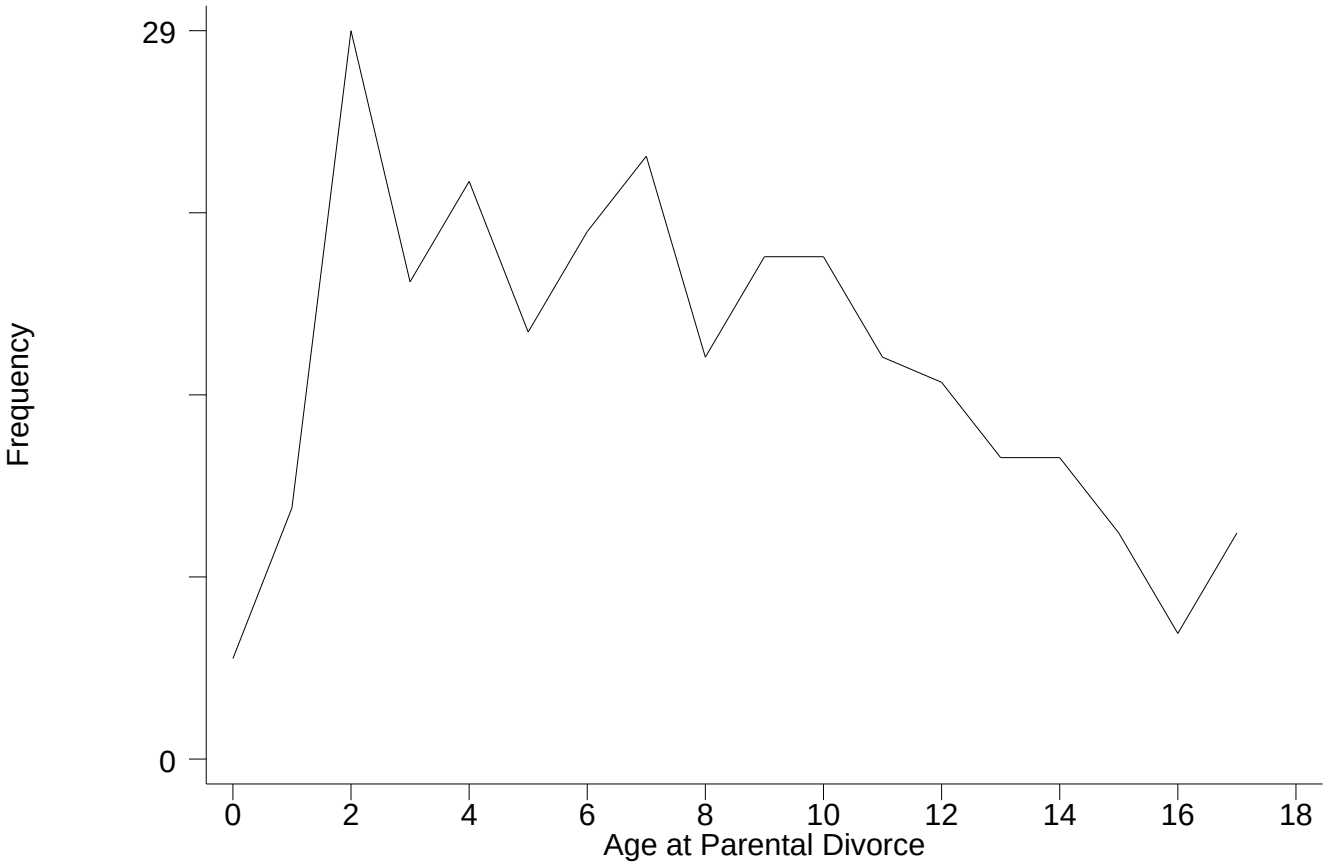


Figure 1: Distribution of Divorces by Child's Age at Parental Divorce

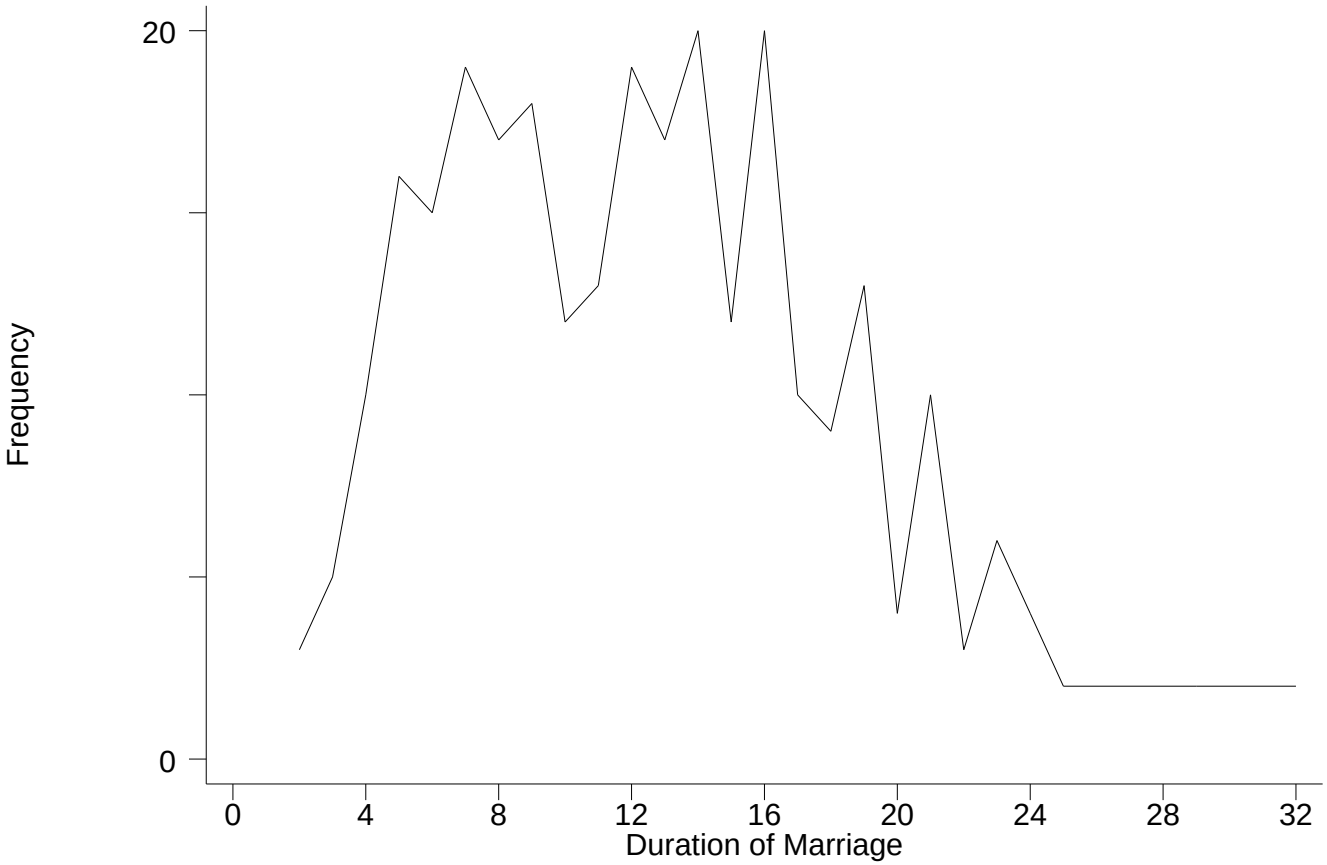


Figure 2: Distribution of Divorces by Marriage Duration

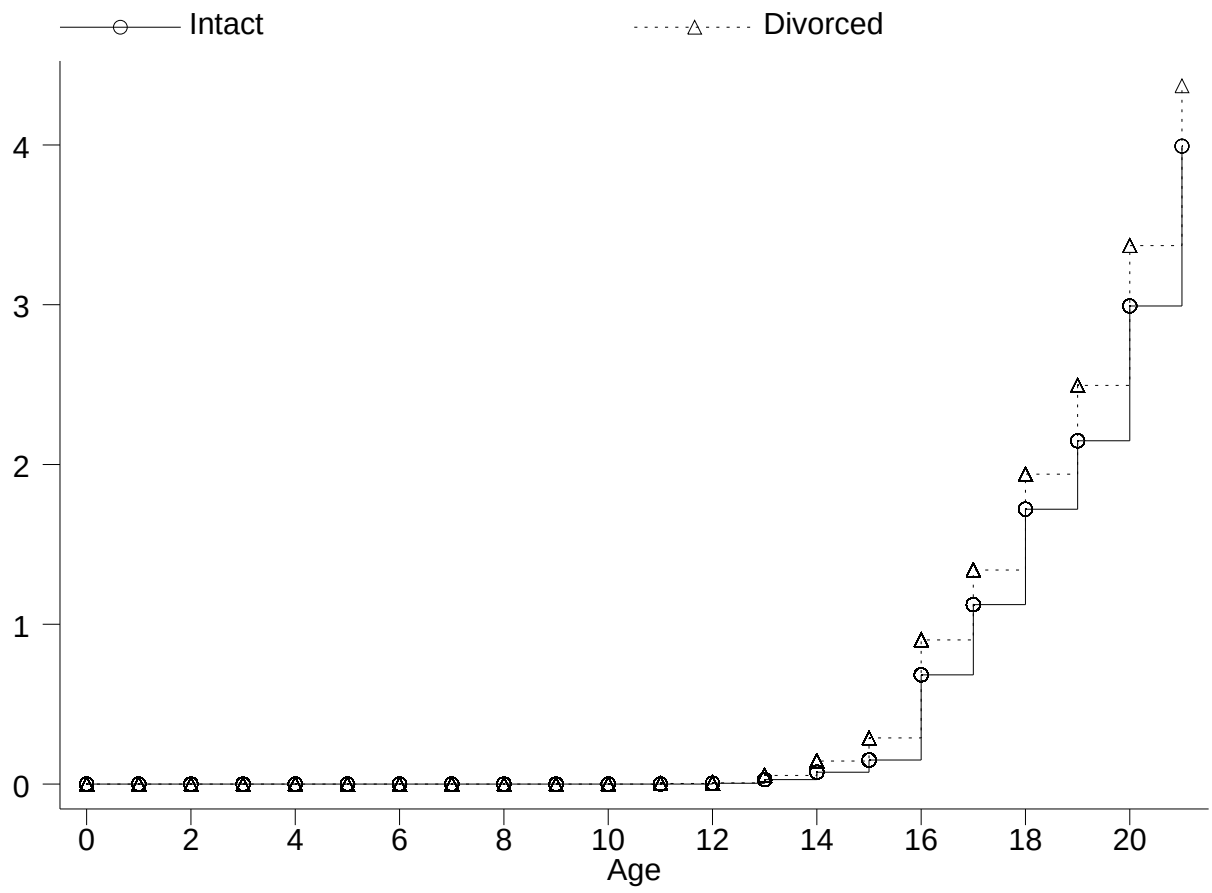


Figure 3: Nelson-Aelan Cumulative Schooling Hazard Estimates: By Family Structure

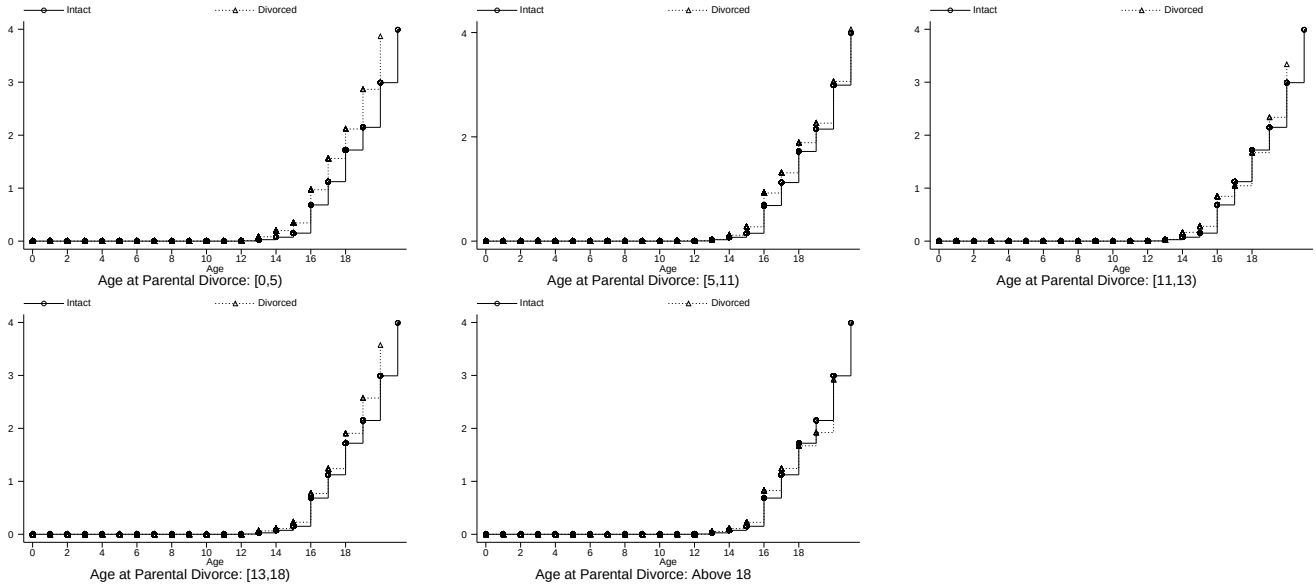


Figure 4: Comparing the Schooling Hazard of Children from Intact versus Divorced Families: By Age at Parental Divorce

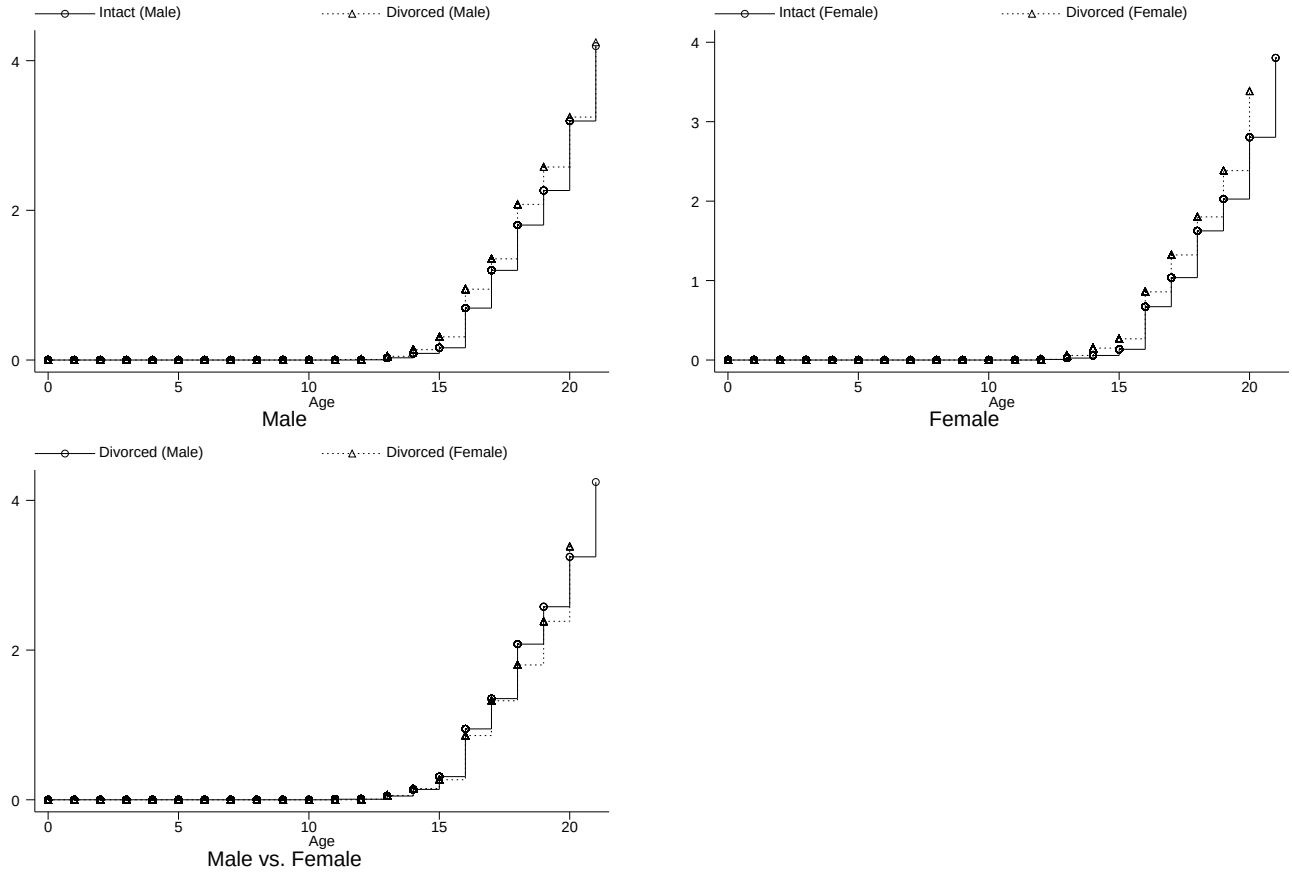


Figure 5: Comparing the Schooling Hazard of Children from Intact versus Divorced Families: By Gender

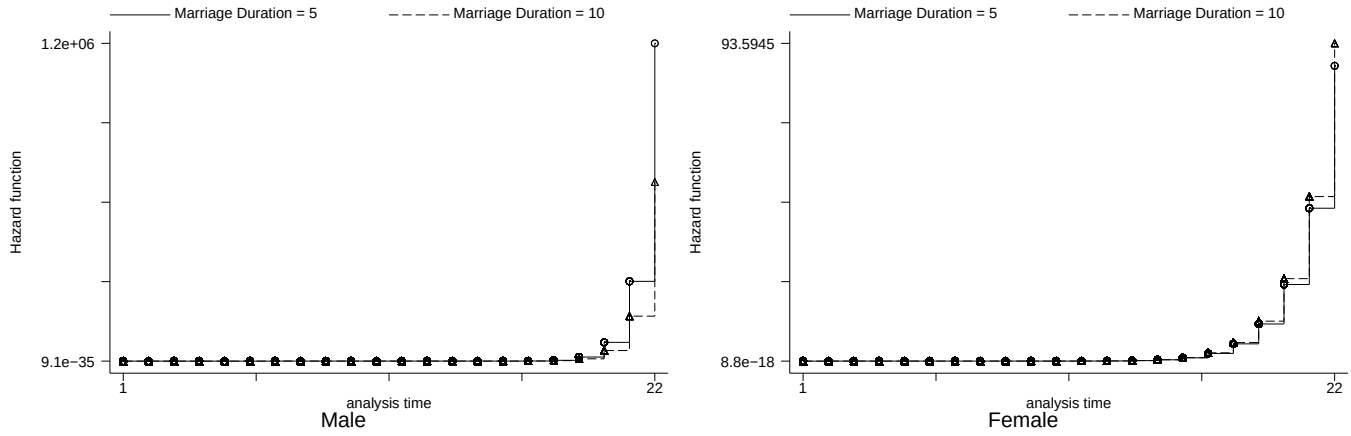


Figure 6: Comparing the Predicted Hazard Functions by Parental Marriage Durations: Male vs. Female

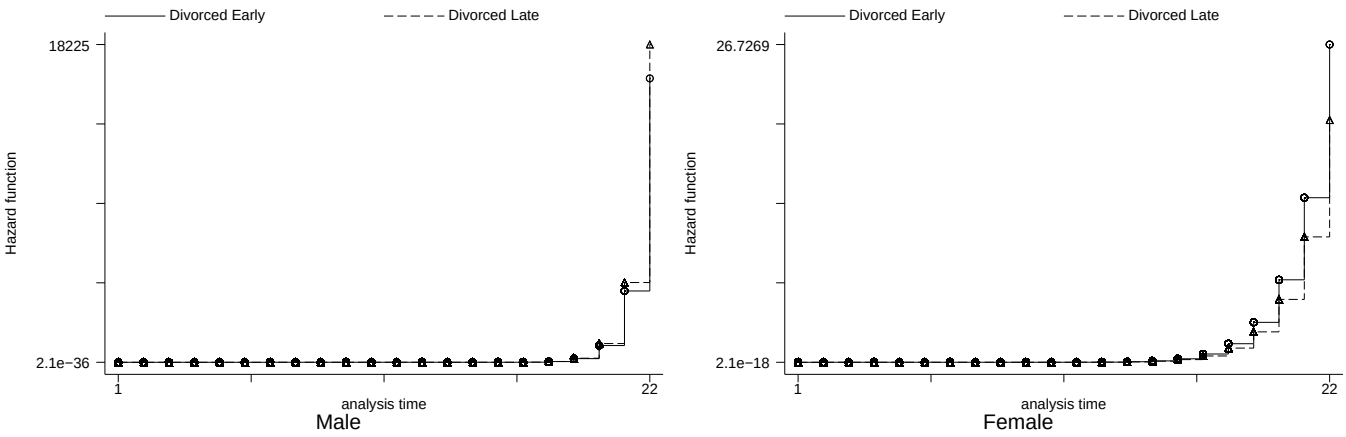


Figure 7: Comparing the Predicted Hazard Functions by Timing of Divorce: Male vs. Female