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Unequal Families, Unequal Effects: How Parental Divorce Differentially Impacts Children's Educational Attainment

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CHILDREN'S EDUCATIONAL ATTAINMENT*

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ABSTRACT

A substantial literature suggests that family disruption leads to lower educational attainment among children. We focus on how the effects of parental divorce on children's education differ across families with varying likelihoods of disruption. Using U.S. panel data, with careful attention to the assumptions and methods needed to estimate total and mediating causal effects, we find a significant effect of parental divorce on educational attainment among children whose parents were unlikely to divorce, for whom divorce was thus a relative shock. We find no effect among children whose parents were likely to divorce and for whom divorce was one of many disadvantages, and thus less economically and socially disruptive. We also find that the observed effect of divorce on children's education is strongly mediated by post-divorce family income. Children's psychosocial skills also explain a portion of the effect among children with a low propensity for parental divorce, while cognitive skills play no role in explaining the negative association between divorce and children's education. Our results suggest that family disruption does not uniformly disrupt children's attainment.

UNEQUAL FAMILIES, UNEQUAL EFFECTS: HOW PARENTAL DIVORCE DIFFERENTIALLY IMPACTS CHILDREN'S EDUCATIONAL ATTAINMENT

Family disruption has long been implicated as a critical event impacting the lives of children. Parental divorce negatively affects a variety of children's outcomes, including psychological wellbeing and academic achievement (see McLanahan, Tach, and Schneider [2013] for a review). Highly-cited research has shown that divorce is negative for children's educational attainment (e.g., McLanahan and Sandefur 1994). However, families differ in their expectation of and ability to adjust and respond to disruption. Families expecting marital stability, who are consequently unprepared for disruption, may experience considerable adjustment difficulties when divorce occurs, leading to adverse outcomes for children. Divorce, however, among families marked by disadvantage and discord, who have come to expect instability, may not incur the same negative consequences. In this study, we consider how the effects of divorce vary by children's likelihood of experiencing a parental divorce. While prior research on the causal effects of parental divorce often addresses the issue of self-selection bias by conditioning on families' likelihood of experiencing divorce, little is known as to whether the consequences of divorce are associated with such predispositions.

Prior research has also largely overlooked how the predisposition to experience parental divorce may condition the mechanisms by which parental divorce impacts children's education. We assess the degree to which decreased educational attainment among children in response to parental divorce is mediated by three factors: post-divorce family instability, family economic resources, and children's skills. First, family transitions (e.g., remarriage, further divorce, cohabitation, union dissolution) occur more frequently following parental divorce, and such instability disrupts children's schooling (Lee and McLanahan 2015; Sweeney 2010). Second,

family income is, expectedly, a central mechanism for the association between parental divorce and children's educational attainment (Thomson, Hanson, and McLanahan 1994; Thomson and McLanahan 2012). With the loss of a parent in the household, typically the father, the mother generally has fewer economic resources. It is well known that such resource reduction negatively impacts children's education, particularly the ability to attend college. Third, a growing body of research suggests that children's psychosocial skills (also termed "non-cognitive skills," "socio-emotional skills," and "personality traits" by scholars in different intellectual traditions) are associated with both family disruption and educational outcomes (e.g., Cunha and Heckman 2009; Duncan and Magnuson 2011; Lleras 2008). We presume that various additional intermediary factors associated with parental divorce, including parents' psychological well-being and parenting style and involvement, will influence children's schooling by impairing their psychosocial skills. Given the considerable role of children's cognitive skills in promoting their educational success, and attention paid to it in the prior literature on divorce (e.g., Kim 2011), we also explore its mediating effect. However, between-person variation in cognitive skills becomes relatively stable by early childhood, and thus ostensibly impervious to disruptive events. We thus expect minimal mediating effect of children's cognitive skills. By contrast, psychosocial skills evolve and change throughout childhood, allowing disruptive family events to play a considerable role in child development (Borghans et al. 2008; Cunha and Heckman 2009; Guo and Harris 2000; Hsin and Xie 2016; Roberts, Wood, and Caspi 2008; Schweinhart et al. 2005).

Our study is thus motivated by two central aims. First, we assess how the effects of parental divorce vary across families with varying likelihoods, or propensities, of disruption. We hypothesize that family "disruption" is not uniformly disruptive. For some, such events require no more than a modicum of social-psychological and behavioral response. For others, such events are

unexpected shocks and necessitate considerable adjustment. The degree of disruption varies by the likelihood and corresponding expectation that such events will occur. Through examining heterogeneity in the effects of divorce by the observed likelihood that children experience a parental divorce, we shed new light on subpopulations for whom the causal effects of divorce may be considerable, modest, or even absent. Second, once we establish how total effects vary across families, we assess several key mediating effects, or mechanisms through which parental divorce affects children's education. Mechanisms explaining the effects of parental divorce have long been conjectured, investigated, and debated. Yet few studies have identified whether, and for whom, these mediation effects are causal. Throughout our analyses, we carefully attend to the assumptions and methods necessary to estimate both total and mediating causal effects across subpopulations of interest. We also offer alternative interpretations of our results based upon observed and unobserved selection into divorce. Our results suggest that a dichotomous distinction between children with divorced and non-divorced parents oversimplifies children's experience. The effect of divorce on educational attainment, and the mechanisms that accounts for it, are stratified by the likelihood that family disruption occurs.

BACKGROUND

Parental Divorce and Children's Educational Attainment

U.S. families have changed dramatically since the mid-twentieth century. Between about 1950 and 1980, divorce rates more than doubled. Only one-quarter of marriages that began in the 1950s ended in divorce, while roughly half of all marital unions beginning in the 1970s eventually dissolved. The increasing incidence of divorce seemingly leveled off after 1980, and possibly even declined (Rotz 2015; Stevenson and Wolfers 2007; yet see Kennedy and Ruggles 2013). Still,

scholars estimate that since the 1980s, roughly half of children experience a parental divorce before they reach adulthood (Amato 2000; Fagan and Rector 2000). Family disruption is more likely to occur among socioeconomically disadvantaged and racial minority families (Amato 2001; Bumpass and Lu 2000). In addition to socioeconomic and demographic factors, marital and fertility history, marital homogamy, relationship quality, traditional family values, and the circumstances surrounding a child's birth play a significant role in marriage survival (Amato, Loomis, and Booth 1995; Frisco, Muller, and Frank 2007; Kim 2011).

As the incidence of parental divorce increased, at least throughout the 1970s, the social stigma associated with such disruption lessened. Nevertheless, the negative consequences for children experiencing family disruption endured (Amato 2001; McLanahan, Tach, and Schneider 2013). A substantial literature links parental divorce to lower levels of children's educational attainment, particularly high school completion (e.g., Amato 2001; Fomby and Cherlin 2007; Lang and Zargorsky 2001; Lee and McLanahan 2015; McLanahan and Percheski 2008; McLanahan, Tach, and Schneider 2013; Seltzer 1994; Sigle-Rushton and McLanahan 2004; Waldfogel, Craigie, and Brooks-Gunn 2010; Wu and Martinson 1993).¹ Scholars studying the causal effects of parental divorce on children have primarily relied on observational data, as divorce is a social phenomenon not subject to experimental manipulation. However, divorced families systematically differ from intact families, ostensibly in both observed and unobserved ways. While selectivity in observed characteristics can be modeled with statistical analyses, selectivity in unobserved characteristics could lead to biases in the estimated effects of parental divorce on children's outcomes. Prior research on parental divorce has adopted a range of methods using observational data in an attempt to address concerns over selection into divorce (e.g., matching models, lagged dependent variable models, individual and sibling fixed effects models, and instrumental variable models), adding

credibility to key findings regarding the negative effects of parental divorce on children's attainment (McLanahan, Tach, and Schneider 2013). This research, however, has not considered how the effects of parental divorce vary according to selection into, or likelihood of, divorce.

Heterogeneous Effects of Parental Divorce on Children's Educational Attainment

Not only do families and children differ, but the effects of parental divorce on children likewise vary. That is, there is no *a priori* reason to expect that children respond uniformly to disruptive family events. Prior research suggests that children with more educated parents experience larger effects of parental divorce than children of less educated parents (Bernardi and Boertien 2016; Bernardi and Radl 2014; Martin 2012), and that parental divorce has stronger effects on white children than on non-white children (Lee and McLanahan 2015; Wu and Thomson 2001). Children's responses to parental divorce also vary by the degree of family well-being: children of married parents with high levels of conflict are no better off, and in fact may fare worse in some respects, than children of single parents (Amato 2000; Jaffee et al. 2003; Musick and Meier 2009; Thomson and McLanahan 2012). Divorce may, in other words, offer some relief from the stress of a high-conflict family environment, particularly if the conflict is visible to children.²

We contend that a shared latent factor underlying the interactions between parental education, race, and family wellbeing and parental divorce is the likelihood, or propensity, of disruption occurring. That is, the negative effects of parental divorce are greater among more advantaged children because they are less accustomed to disruptive socioeconomic events and disadvantaged circumstances than more disadvantaged children (McLoyd et al. 2000). They are consequently unlikely to be embedded in a social network in which family instability is anticipated, or at least less stigmatized (Brand and Simon Thomas 2014; Cherlin 2004; McDermott, Fowler,

and Christakis 2013; Ross 1995).³ By contrast, parental divorce may not further encumber the educational attainment of children who have grown accustomed to disruption in their lives, such as job loss, health shocks, residential mobility, and income decline (Duncan et al. 1998). In this study, we consider how the effects of parental divorce on children's education differ across families with different likelihoods of disruption.

Mediating Effects of Parental Divorce on Children's Outcomes

Attention to mediation is central to our understanding of how parental divorce impacts children's educational outcomes. A mediation analysis provides estimates for the amount and proportion of the effects of parental divorce that are transmitted through various intermediate pathways. For a mechanism to mediate divorce effects on children's education, it must satisfy two conditions: (1) the mechanism must be influenced by parental divorce; and (2) the mechanism must influence children's educational outcomes. The divorce literature has focused on several plausible mechanisms that seemingly satisfy these two conditions.

Family instability in the form of transitions in household composition and family relationships is such a candidate. It is more likely subsequent to a parental divorce (condition 1) and is associated with high levels of parenting stress and lower-quality parent-child relationships leading to lower attainment among children (condition 2) (Beck et al. 2010; Cavanagh, Crissey, and Raley 2008; Cavanagh and Huston 2006; Halpern-Meekin and Turney 2016; Lee and McLanahan 2015; Osborne and McLanahan 2007; Thomson and McLanahan 2012; Waldfogel, Craigie, and Brooks-Gunn 2010; Wu and Martinson 1993; Wu and Thomson 2001). Instability is likely to vary across families. Families with a low likelihood of an initial parental divorce, and thus whose characteristics made family disruption unlikely, will presumably experience fewer

subsequent transitions than families with a high likelihood of parental divorce. In this case, instability will not serve as a crucial explanatory factor. Among families less resistant to disruption, increased instability may play a more consequential role.

Perhaps most prominent among the mechanisms discussed in the literature on parental divorce and child wellbeing is family economic resources. Divorce is associated with a decline in family income (condition 1), and decades of social science research demonstrates that family economic resources play a significant role in children's education (condition 2) (Coleman, Ganong, and Fine 2000; Crosnoe and Cavanagh 2010; Duncan et al. 1998; Lee and McLanahan 2015; McLanahan and Percheski 2008; McLanahan and Sandefur 1994). In addition to the strong impact on home, neighborhood, and school environment, health and emotional well-being, and procuring educational goods and resources, family income facilitates paying the increasingly high price of college (Goldrick-Rab 2016). Prior research suggests that differences in economic resources account for a substantial share (roughly half) of the differences in child outcomes across family types (McLanahan and Sandefur 1994; Thomson, Hanson, and McLanahan 1994; Thomson and McLanahan 2012). The mediating effect of family economic resources may also vary across families. While more advantaged families are likely to have higher levels of economic resources than disadvantaged families, resource loss as a result of divorce may be more pronounced in the former than in the latter (Bernardi and Boertien 2016). For disadvantaged families, income may have already been below the threshold of investment in higher education prior to resource decline due to divorce.

Children's own skills, related to both cognitive and psychosocial abilities, are another mechanism by which family disruption may limit educational attainment. Cognitive ability has historically played a central role in models of status attainment (Sewell, Haller, and Portes 1969)

and human capital development (Becker 1993), and such skills clearly impact educational outcomes and satisfy condition (2). Condition (1), however, that parental divorce impacts cognitive skills, must also be satisfied for such skills to mediate the relationship between divorce and children's education. While some recent research treats children's psychosocial and cognitive skills as parallel or symmetrical outcomes of family disruption (e.g., Kim 2011), we instead maintain that the developmental literature points to important asymmetry in the acquisition of such skills. Cognitive skills undergo rapid development in early childhood, and gradually stabilize thereafter (Borghans et al. 2008; Cunha and Heckman 2009). Hence, if we observe an impact of parental divorce on children's cognitive skills (e.g., Kim 2011), at least beyond the early childhood years, it is likely influenced by the impact of psychosocial mediators, discussed below, on cognitive assessments. The evidence on the effects of parental divorce on cognitive assessments in math, verbal, and general test scores is mixed, with studies adopting more stringent tests for causal associations suggesting little or no effect (Aughinbaugh, Pierret, and Rothstein 2005; Cherlin et al. 1991; Lee and McLanahan 2015; Morrison and Cherlin 1995; Sun 2001; Sun and Li 2002).

Psychosocial skills encompass a broad class of attitudes and behaviors that are correlated with but distinct from cognitive ability, such as emotional stability, self-esteem, mastery, conscientiousness, locus of control, and behavior (Borghans et al. 2008; Claessens, Duncan and Engel 2009; Heckman, Stixrud and Urzua 2006; Jencks et al. 1979; Lleras 2008; Rosenbaum 2001). In contrast to cognitive skills, psychosocial skills evolve and change from early childhood through adulthood. Children's family environments can thus play a significant role in shaping psychosocial skills (Hsin and Xie 2016; Roberts, Walton, and Viechtbauer 2006; Roberts, Wood and Caspi 2008).⁴ We thus expect condition (1) to be satisfied for psychosocial skills. Regarding

condition (2), scholars increasingly recognize the critical role of psychosocial skills in children's academic achievement and educational attainment (Ainsworth-Darnell and Downey 1998; Cunha and Heckman 2009; DiPrete and Jennings 2012; Duckworth and Seligman 2005; Duncan and Magnuson 2011; Farkas et al. 1990; Hsin and Xie 2016; Jackson 2006; Lleras 2008; Rosenbaum 2001; Wolfe and Johnson 1995), even among individuals who share the same family background and cognitive abilities (Heckman and Rubinstein 2001; Heckman, Stixrud, and Urzua 2006). Given evidence in support of both conditions (1) and (2), we hypothesize that psychosocial skills, more than cognitive skills, mediate the effect of parental divorce on children's education.⁵ We also expect that the marginal effect of psychosocial decline may differ across families. If children unaccustomed to socioeconomic disadvantage and disruption experience a greater psychological shock to parental divorce, it may play a more explanatory role.⁶

While our focal mechanisms clearly link family disruption and children's education, we know little regarding their relative mediating influences across families with varying likelihoods of disruption. In this study, we assess the relative impact of family instability, family income, and children's skills in explaining the total effects of divorce on education for children who vary in their propensity for experiencing parental divorce. Also in contrast to most prior work, we carefully attend to the assumptions and methods needed to estimate causal mediating effects. Our analysis cautions against a uniform explanation for both the total and mediating causal effects of parental divorce, calling attention to conditions under which effects vary across families.

ANALYTICAL APPROACH

Estimating Average Effects of Parental Divorce on Children's Educational Attainment

For a focal child i , the treatment effect (TE) of parental divorce is defined as the difference between the two potential outcomes in the treated (i.e., divorced parents) and untreated (i.e., non-divorced parents) states ($D = 1, 0$):

$$TE_i = Y_i(1) - Y_i(0). \quad (1)$$

That is, we ask whether a child whose parents divorced had different outcomes than he or she otherwise would have had if his or her parents had not divorced. Given the impossibility of observing both treated and untreated outcomes for any individual, the individual-level causal effect, as defined in equation (1), is unidentifiable. The researcher is constrained to estimate subpopulation average treatment effects (Morgan and Winship 2014). Formally, the population average treatment effect (ATE) is defined as the overall average difference in outcomes between children whose parents did and did not divorce:

$$ATE = E(Y(1) - Y(0)). \quad (2)$$

With observational data, the key to our identification strategy is the *ignorability* assumption, i.e. the assumption that parental divorce is uncorrelated with unobserved factors that affect children's outcomes (Heckman 2005). To guard against potential selection bias and improve confidence in the ignorability assumption, we condition the analyses on a rich set of observed characteristics (shown in Table 1), indeed a more extensive set than most prior analyses of family disruption. Still, there is no *a priori* reason why the ignorability assumption holds true, as parents may self-select into divorce due to unobserved factors. To address concerns of selection bias, our analytical approach begins with the estimation of the propensity for parental divorce (P) based on observed covariates (X) (Rosenbaum and Rubin 1983, 1984):

$$P = P(D_i = 1|X_i). \quad (3)$$

Under the ignorability assumption, conditioning on the propensity score is as sufficient as conditioning on the full array of covariates X for the estimation of treatment effects (Morgan and Winship 2014; Rosenbaum and Rubin 1983, 1984; Rubin 1997). Departing from most previous research on parental divorce effects on children, our approach necessitates that we explicitly model parental divorce as a first step.

We next estimate an average treatment effect conditional on the observed propensity for parental divorce:

$$ATE_p = E(Y(1) - Y(0)|P = p). \quad (4)$$

The conditional average treatment effect ATE_p measures the reduced-form, or total, effect of parental divorce operating through all mediating pathways. We estimate a series of linear probability models of the effects of parental divorce on children's high school completion, college attendance, and college completion as follows:⁷

$$Y_i = \alpha + \beta_1 D_i + \beta_2 P_i + \varepsilon_i. \quad (5)$$

For simplicity and ease of interpretation, we include the propensity score as a linear term in equation (5) to reduce confounding biases due to observed covariates and effect heterogeneity (Xie, Brand, and Jann 2012). In subsequent analyses, we explore effect heterogeneity more closely.

Estimating Heterogeneous Effects of Parental Divorce on Children's Educational Attainment

When individuals differ in their response to a treatment, as is generally the case, average treatment effects vary depending on population composition (Xie 2013). It is thus important to understand treatment effect heterogeneity. Toward this goal, we assess whether the effects of divorce vary with the propensity for parental divorce. There are important substantive payoffs to understanding effect heterogeneity by the propensity for treatment, whether or not the ignorability

assumption holds true. If the ignorability assumption is true, effect variability by the propensity score is the only interaction effect consequential for selection bias (Xie, Brand, and Jann 2012). When the ignorability assumption does not hold, we may interpret effect variability by the propensity score as resulting, at least partially, from unobserved selectivity (Brand and Simon Thomas 2013; Brand and Xie 2010; Xie, Brand, and Jann 2012; Zhou and Xie 2016a, 2016b), to be described in more detail below.

We adopt two approaches for estimating effect heterogeneity under the ignorability assumption. First, we use the *matching-smoothing* method consisting of the following steps (Xie, Brand, and Jann 2012): (1) estimate propensity scores for all units; (2) match treated units to untreated units with a matching algorithm; (3) plot the observed difference in a pair between a treated unit and an untreated unit against a continuous representation of the propensity score; and (4) use a local polynomial model to smooth the variation in matched differences to obtain the pattern of treatment effect heterogeneity as a function of the propensity score. Second, if there appears to be effect heterogeneity that could be sufficiently captured by discrete strata of the estimated propensity score, we assess stratum-specific total (and mediating) effects. The number of strata we construct will depend upon the shape of the nonparametric response function. Using this approach, we define the stratum-specific conditional average treatment effect as:

$$ATE_{s,p} = E(Y(1) - Y(0)|S = s, P = p), \quad (6)$$

where $S = \{1, 2, \dots, s\}$ indicates the stratum of the estimated propensity score. We estimate linear probability models of the form described in equation (5) separately by propensity score strata. Admittedly, it is also possible to interact particular covariates with divorce as an alternative way of modeling treatment effect variation. The propensity score, however, provides a parsimonious

measure of an extensive set of observed covariates that indicate the likelihood of divorce and lends itself to a unique interpretation based on observed and unobserved selection.⁸

Modeling Heterogeneous Effects of Parental Divorce by Observed and Unobserved Selection

The previous literature has well established, and we will reaffirm, that a host of observed social background, socioeconomic, and family factors systematically affect the likelihood that a child's parents divorce. As we note above, we label the observed determinants of parental divorce X , and we construct the propensity score P as a summary index of observed X . However, we also acknowledge the presence of unobserved factors that affect the likelihood of divorce. Of these unobserved factors, some are systematic, reflecting parents' unwillingness, or resistance, to divorce. We denote the unobserved resistance to divorce as U . For example, parents' resistance to divorce may be partly affected by their concern that children's future outcomes will be negatively affected by a disruption. We describe the latent divorce function $D^*(\cdot)$ as:

$$D^* = P - U, \tag{7}$$

where P is the propensity of divorce based on observed covariates (Rosenbaum and Rubin 1983, 1984), and U is resistance to divorce, distributed between 0 and 1. Parents divorce when $D^*(\cdot)$ exceeds 0:

$$D = \begin{cases} 1 & \text{if } D^* \geq 0, \\ 0 & \text{otherwise} \end{cases} \tag{8}$$

In this model, we allow for the presence of U that may affect children's attainment subsequent to divorce. In general, the treatment effect varies both by P and by U . The ignorability assumption for estimating heterogeneous treatment effects refers to the special case where the treatment effect varies by P and not by U , i.e., U being ignorable.

Borrowing the notation of Heckman, Urzua, and Vytlacil (2006) and Zhou and Xie (2016a), we propose a simple model for parental divorce that incorporates unobserved response heterogeneity. We define the effect of parental divorce on children's outcomes to be a function of both P and U , called the marginal treatment effect (MTE)⁹:

$$MTE = E[(Y(1) - Y(0))|P = p, U = u]. \quad (9)$$

In Figure 1 (adapted from Zhou and Xie [2016b]), we depict alternative ways in which we can interpret treatment effect heterogeneity. The darker shaded regions indicate a larger treatment effect magnitude (i.e., larger negative effects of parental divorce on children's outcomes). In Figure 1(a) and 1(b), we assume ignorability but allow for divorce effect heterogeneity by P . Under this assumption, equation (9) is reduced to:

$$ATE_p = TT_p = E(Y(1) - Y(0))|P = p), \quad (10)$$

where ATE_p refers to the average treatment effect and TT_p refers to the treatment effect on the treated, both assumed to vary by the propensity for divorce P but not by the unobserved resistance to divorce U .

In Figures 1(c) and 1(d), we consider the general case of equation (9) and allow MTE to be a function of both P and U . Figures 1(a) and 1(c) depict effects for all units. In this case, we cannot estimate ATE_p without the ignorability assumption. However, Zhou and Xie (2016a) show that, if we have treatment heterogeneity bias but not baseline heterogeneity bias, the conventional method for estimating heterogeneous treatment effects by the propensity score under ignorability still yields valid average treatments effect for the treated. That is, we are able to identify TT_p as:

$$TT_p = E(Y(1) - Y(0)|P = p, D = 1) = \frac{1}{p} \int_0^p MTE(P = p, U = u) du, \quad (11)$$

the integration of MTE over $U = u$ given the propensity score $P = p$. Note that the integral of (11) is systematically correlated with the propensity score P : lower observed propensity for divorce is

associated with lower unobserved resistance to divorce, and thus more density in the integral for lower levels of U . Substantively, we take lower resistance to indicate that parents choose divorce despite potential negative effects for children's wellbeing. In Figure 1(d), we illustrate this relationship among treated units, i.e., the subpopulation for which $P > U$. We note the high correlation between P and U among treated units in Figure 1(d): At low values of P , the estimated effect includes proportionally more children whose parents have low values of resistance U ; at high values of P , the estimated effect includes more variation with respect to U , and thus proportionally more children whose parents have high values of U .

– FIGURE 1 ABOUT HERE –

In the subsequent analyses, we present simple results pertaining to the heterogeneous effects of parental divorce on children's outcomes as a function of the estimated propensity of divorce under ignorability. They are informative descriptive results in their own right (Xie, Brand, and Jann 2012). If ignorability is true, we may interpret the pattern in the effect of divorce as a function of the likelihood, or propensity, of disruption. However, if ignorability does not hold, such that we have heterogeneous responses to latent determinants of divorce, the same results are still interpretable because they indicate variation in effects of parental divorce by the latent unobserved parental resistance to divorce.

Estimating Heterogeneous Mediating Effects of Parental Divorce on Children's Educational Attainment

A causal mediation analysis is designed to estimate the role of mechanisms in transmitting the effect of a treatment on an outcome. Mediation methods using a potential outcomes framework have dramatically expanded in recent years (VanderWeele 2016). Compared to traditional path analysis, a potential outcomes approach provides a coherent framework clarifying the assumptions

needed to estimate valid mediation effects (Hicks and Tingley 2011; Imai, Keele, and Tingley 2010; Imai, Keele, and Yamamoto 2010; Keele, Tingley, and Yamamoto 2015; Pearl 2001, 2009; Robins and Greenland 1992; VanderWeele 2015, 2016). The goal in causal mediation analyses is to decompose the total, reduced-form, treatment effect (i.e., the stratum-specific average treatment effect defined in equation [6]) into direct and mediating (or indirect) treatment effects. The mediating effect reflects one potential pathway through which the treatment produces the effect on the outcome of interest. We assess key post-divorce mediating mechanisms that may transmit the effects of parental divorce to children’s educational outcomes: family instability, family income, and children’s cognitive and psychosocial skills. Figure 2 is a directed acyclic graph (DAG) that illustrates the relationship between the propensity for parental divorce (P), parental divorce (D), the mediators (M), and children’s educational attainment (Y).

– FIGURE 2 ABOUT HERE –

Let $M_i(d)$ denote the potential value of the mediator that would be realized under treatment status $D = d$. For example, $M_i(d)$ may indicate child i ’s post-divorce psychosocial skills that would have been observed had the child experienced a parental divorce ($D = 1$) or not ($D = 0$). Only the potential mediator that corresponds to the actual received treatment is observed. Let $Y_i(d, m)$ represent the potential outcome that would result if the treatment and mediating variables equaled respectively d and m for i . For example, $Y_i(1, 0.6)$ represents high school completion status for child i that would be observed if the child had experienced a parental divorce and the psychosocial skills scale equaled 0.6 (the mean value for children of divorced parents). The observed $Y_i(1, 0.6)$ is only one of many potential outcomes of $Y_i(d, M_i(d))$.

Using this notation, we define the total treatment effect for unit i as follows:

$$TE_i = Y_i(1, M_i(1)) - Y_i(0, M_i(0)). \quad (12)$$

This is the same effect described in equation (1), yet equation (12) explicitly expresses the mediating mechanisms. We define the causal mediation effect of the treatment, also known as the natural indirect effect (IE) (Pearl 2009), on the outcome through the mediating variable for unit i as follows:

$$IE_i = Y_i(d, M_i(1)) - Y_i(d, M_i(0)).^{10} \quad (13)$$

The indirect effect shows what change would occur to the outcome if the mediator changed from what would be observed when units are treated ($M_i(1)$) to what would be observed when units are untreated ($M_i(0)$), while holding the treatment status constant at d . This deactivates all pathways except for that operating through the focal mediator. For example, $Y_i(1, M_i(1))$ could represent high school completion status for child i with divorced parents and the level of psychosocial skills after parents divorced and $Y_i(1, M_i(0))$ could represent high school completion status for the same child with divorced parents but with the level of psychosocial skills had parents not divorced. The mediating effect in this example explains the degree to which parental divorce impacts high school completion by decreasing children's psychosocial skills.¹¹ Following our definition of ATE in equation (6), we estimate heterogeneous mediation effects conditional on the estimated propensity score and within each propensity stratum S :

$$IE_{s,p} = E(Y(d, M(1)) - Y(d, M(0)) | S = s, P = p). \quad (14)$$

Mediation analysis estimates the proportion of the total effect that is indirect:

$$PM_S = \frac{IE_{s,p}}{TE_{s,p}}, \quad (15)$$

i.e., the proportion of the causal mediation effect to the total effect. We perform this decomposition exercise for the stratum-specific average treatment effects (i.e., $TE_{s,p}$ is equal to ATE defined in equation [6]).

To identify mediating effects, we must assume *sequential ignorability*. That is, given adjustment for observed pretreatment covariates, the treatment is assumed to be independent of potential outcomes *and* of potential mediators. The first part of the assumption is necessary to identify the total effect (as we described above), while the second part is necessary to identify the mediation effect. That is, we assume no treatment-outcome confounding, treatment-mediator confounding, or mediator-outcome confounding, as well as no variable that is a consequence of the treatment that confounds the mediator-outcome relationship (VanderWeele 2016).

Our mediation analysis proceeds as follows. First, we fit a regression predicting the mediator (M) that includes the treatment (D) and the propensity for treatment (P) within each stratum (S):

$$M_i = \alpha_m + \beta_{m1}D_i + \beta_{m2}P_i + \varepsilon_{im}. \quad (16)$$

Second, we fit a strata-specific regression predicting the outcome that includes the mediator, treatment, and relevant covariates:

$$Y_i = \alpha_y + \beta_{y1}D_i + \beta_{y2}P_i + \beta_{y3}M_i + \varepsilon_{iy}. \quad (17)$$

We simulate model parameters in the mediator and outcome models from their sampling distributions. For each simulation, based on the mediator model, we generate two sets of predicted mediator values for each unit, one when $D = 1$ and one when $D = 0$. We use the outcome model to impute potential outcomes: first, the predicted outcome when $D = 1$ and the predicted mediator value when $D = 1$ (from the previous step); and second, the predicted counterfactual outcome when $D = 1$ and the predicted mediator when $D = 0$. The average causal mediation effect is obtained by averaging the differences between the predicted outcomes under the two values of the mediator across units. For example, we could generate the average difference in children's high school completion across levels of psychosocial skills with and without experiencing a parental divorce.

We repeat the simulation 1,000 times to obtain estimates of uncertainty and statistical significance tests (see Imai et al. [2010] Appendix D for technical details).

Sensitivity Analyses of Heterogeneous Total and Mediating Effects of Parental Divorce on Children's Education Attainment

Sensitivity analyses provide a general framework for investigating the extent to which the estimated treatment effects are sensitive to unobserved confounding covariates by quantifying how the results obtained under the ignorability assumption would change if we relaxed the assumption. A standard approach is the calculation of a bias factor (Arah 2017; Gangl 2013; VanderWeele 2015, 2016; VanderWeele and Arah 2011). The sensitivity of the estimated total effects to unobserved treatment-outcome confounding can be assessed by subtracting the bias factor from the point estimate and confidence interval of the treatment effect obtained under ignorability. The bias term is equal to the product of two (stratum-specific) parameters:

$$B_S = \gamma_S \lambda_S, \quad (18)$$

where

$$\gamma_S = E(Y|U = 1, D = d, S = s, P = p) - E(Y|U = 0, D = d, S = s, P = p) \quad (19)$$

and

$$\lambda_S = P(U = 1|D = 1, S = s, P = p) - P(U = 0|D = 0, S = s, P = p). \quad (20)$$

That is, γ is the mean difference in children's education associated with a unit change in an unobserved binary confounder, U , and λ is the mean difference in the unobserved confounder between the children of divorced and non-divorced parents, both conditional on the estimated propensity for divorce and propensity strata.

Mediation analysis requires an additional sequential ignorability assumption, beyond the ignorability assumption needed to identify the total effect. We assess the sensitivity of the

mediation effects to the simplified assumption of unobserved mediator-outcome confounding with another bias term. The components of the bias term are analogous to those for the total effects, except that they are now conditioned on the mediator. The bias term is in this case equal to the negation of the product of the two parameters, and we subtract this bias term from the mediation effect and the confidence interval.

DATA

National Longitudinal Survey of Youth

The National Longitudinal Survey of Youth (NLSY) is a nationally representative sample of 12,686 respondents who were 14 to 22 years old when first surveyed in 1979. These individuals were interviewed annually through 1994 and biennially thereafter. In 1986, the National Longitudinal Survey began a separate survey of the children of NLSY women, the National Longitudinal Survey's Child-Mother file (NLSCM). Data have been collected every two years since 1986, with new sections added in 1994 as children entered young adulthood. As of 2012, the 6,283 women of the NLSY were 47 to 54 years old and had given birth to about 11,500 children. Several prior studies have used data from the NLSY to investigate the impact of parental divorce (e.g., Aughinbaugh et al. 2005; Lang and Zargorsky 2001; Morrison and Cherlin 1995).

We merged data on women from the NLSY with data on children from the NLSCM ($n = 11,512$ children and $n = 4,931$ mothers) and treat children as our units of analysis. We constructed measures of whether and when a child (0-17 years old) experienced a parental divorce using NLSCM-provided month and year of birth for children and NLSY-provided marriage start and end dates for parents. As our analytical framework requires modeling the probability that children experience divorce over the course of childhood, we begin the period with all children at risk of

experiencing the event. That is, we identified 8,319 children of 3,940 mothers who were born into marriage. This restriction allows the examination of a relatively homogenous population of families. We then identified children who experienced parental divorce at or before age 17, and thus further restricted the sample to those who were at least 18 years old by 2012 ($n = 7,258$ children). About a third of our sample ($n = 2,420$ children) experienced a parental divorce over the course of childhood (the weighted sample proportion is 0.40). The average age of children at the time of divorce is 7 years old.

Descriptive Statistics of Pre-divorce Covariates

Drawing on prior research on the determinants of divorce, we include a rich set of covariates to construct the propensity of parental divorce over childhood: family background factors of mothers (i.e., race, national origin, residential location, religion, family structure and size, household income of mothers during childhood); socioeconomic factors of mothers and households (i.e., education, employment status, job conditions, delinquency, household income, poverty, and welfare status); cognitive and psychosocial factors of mothers (i.e., psychosocial skills [scales for Rotter locus of control, Pearlin mastery, Rosenberg self-esteem], delinquency [based on 16 questions regarding stealing, gambling, fighting, drugs], depressive symptoms [7-item CES-D], body mass index, cognitive ability [ASVAB], high school academic achievement [class rank and college preparatory program]; and family formation and wellbeing factors (i.e., early sexual activity, beliefs about traditional family roles, age at time of child's birth, prior marriages, time between marriage and first birth, desirability of birth of a child, child gender, and child birth weight; whether parents argue about chores, money, cheating, and religion, and whether they match with respect to religion, race, and education).¹² Missing values for the covariates were

imputed based on pre-divorce characteristics. A limitation of these data is that most of our covariates are based on mothers; the strength is that we have a large nationally representative sample with rich longitudinal data on mothers and their children. We also have enough data on fathers and households to construct several key sociodemographic characteristics and indicators of relationship quality. Our sensitivity analyses help us determine how large the influence of an omitted variable need be to invalidate our results. We observe significant differences by parental divorce status for most of the indicators we include, suggesting greater socioeconomic disadvantage and lower family wellbeing among parents who divorce.

– TABLE 1 ABOUT HERE –

Descriptive Statistics of Mediators and Outcomes

Table 2 describes the mediators and outcomes used in the main analyses. Measures of children’s educational attainment include high school completion by age 18, college attendance by age 19, and college completion by age 23. Relationship transitions count the number of times a transition occurs between the states of married, separated, remarried, widowed, and cohabitating, and thus can have positive values for both divorced and non-divorced families. Measures of children’s skills include a scale of cognitive skills and a scale of psychosocial skills. The cognitive skills scale is constructed by averaging three Peabody Individual Achievement Test (PIAT) indicators: Reading Comprehension, Reading Recognition, and Math. Children are between ages 5 and 18 when tested. The psychosocial skills scale is constructed with five indicators: the Pearlin Mastery scale; the Rosenberg Self-Esteem scale; the Ten-Item Personality Inventory (TIPI) Emotional Stability scale; the Behavioral Problem Index (BPI); and the Center for Epidemiologic Studies Depression scale (CESD).¹³ These indicators were all measured when children were age

15 and above, with the exception of behavioral problems, which was measured between ages 4 and 15. For both the cognitive and psychosocial skills scales, the selected items were standardized to have a zero mean and unit variance. We took the mean value of the standardized values to create a composite scale measure and then transformed each scale measure onto the [0,1] interval. We explored each of the indicators separately for both cognitive and psychosocial skills, but ultimately decided that summary scales resulted in little loss of information and greatly increased parsimony.

All mediators were constructed as averages of the measures over the years subsequent to the parental divorce event. For example, if a child's parents were divorced when the child was 7 years old, we averaged the values of the mediator between ages 8 and 17, or when the first measure is available after age 8. Construction of mediators offers inherent measurement challenges due to nonrandom selection into divorce, the timing of divorce, and the expectation that mediators exhibit some degree of age dependency. To assess the degree to which differences in mediators explain effects on children's educational attainment, we need to compare post-divorce mediator estimates for the divorce group to an analogous estimate for children who do not experience parental divorce. Yet while children whose parents divorce have an observable event time, children whose parents do not divorce have no analogous event time.¹⁴ To address this issue, we employ a method that matches children from the divorce group to children in the non-divorce group based on their gender and propensity to experience parental divorce. After the match, we simulate age at divorce for the control group child as the observed age at divorce for the matched child in the treated group. We then average all measures taken after the divorce age (observed or simulated, up to age 17) to create a post-divorce mean.¹⁵ We observe in Table 2 that children whose parents divorced have a greater number of family transitions, substantially lower family income, lower levels of cognitive and psychosocial skills, and lower levels of educational attainment.

ANALYTICAL RESULTS

Propensity Model for Parental Divorce

We model the probability that a child experiences a parental divorce over the course of childhood (age 0-17) as a function of the covariates described in Table 1. Results are presented in Table 3. Allowing our treatment to occur anytime between a child’s birth and age 17 limits our “pretreatment” covariates to those at the time of the child’s birth, which does not allow for the adjustment of time-varying confounders. Still, as the dissolution process is likely to begin well before any formal separation is observed (Furstenberg and Kiernan 2001), too much precision in the window of observation may lead to conditioning on endogenous variables that obscure the effects of an impending divorce. As results from models predicting parental divorce are seldom presented in prior work on divorce effects on children, the literature has not established a widely-accepted prediction model. Our model incorporates a rich set of theoretically informed covariates based on the literature on the determinants of divorce, with attention to limiting covariates that could amplify bias.¹⁶

We show in Table 3 that mothers who themselves were raised in large families with fathers present during childhood are less likely to divorce. Mother’s self-esteem is negatively associated with the odds of divorce, while a high level of depressive symptoms is positively associated with the odds of divorce. High cognitive ability, self-mastery, and academic achievement in high school among mothers appear to be positively associated with divorce. Education and household income generally reduces the odds of divorce, while mothers’ employment, especially employment at a private company without flexible hours, increases odds of divorce. Family formation factors

strongly influence the likelihood of divorce, with women adopting more traditional family practices (e.g., delayed sexual debut and no prior marriages) and attitudes less likely to divorce. Relationship quality measures indicate that arguing about chores is positively associated with divorce, while arguing about money is negatively associated with divorce. Parents who differ in their educational attainment and who are of different races are more likely to divorce. However, those raised in different religions are less likely to divorce, perhaps reflecting strong selection into cross-religion marriages. In sum, with some notable exceptions, the likelihood of divorce generally declines with socioeconomic status and family wellbeing.

– TABLE 3 ABOUT HERE –

Effects of Parental Divorce on Children’s Educational Attainment

We present linear probability model estimates of the effects of parental divorce on children’s educational attainment in Table 4. The first column reports simple zero order estimates. We observe that divorce is associated with an 8 percent lower probability of children’s high school completion, a 12 percent lower probability of college attendance, and an 11 percent lower probability of college completion. We predict that among children whose parents stay married, about 82 percent complete high school, 57 percent attend college, and 24 percent complete college, while among children whose parents divorced, about 75 percent complete high school, 45 percent attend college, and 13 percent complete college. The second column reports estimates adjusted for the propensity of parental divorce and child’s age in 2012. The magnitudes of the coefficients are reduced, but retain significance. We observe that, net of the propensity for parental divorce, divorce is associated with a 4 percent lower probability of children’s high school completion, a 7 percent lower probability of college attendance, and a 7 percent lower probability of college

completion. Holding the propensity for parental divorce at the median, we predict that among children whose parents stay married, about 81 percent complete high school, 56 percent attend college, and 23 percent complete college, while among children whose parents divorced, about 78 percent complete high school, 50 percent attend college, and 17 percent complete college.

– TABLE 4 ABOUT HERE –

Heterogeneous Effects of Parental Divorce on Children’s Educational Attainment

We present local polynomial matching-smoothing (MS) heterogeneity results in Figure 3. The x -axis represents the continuous propensity score and the y -axis represents the observed differences in (a) high school completion, (b) college attendance, and (c) college completion between children whose parents did and did not divorce. We observe a sizable negative effect of divorce on educational attainment among children who had a low likelihood of experiencing a parental divorce, an effect that declines (i.e., becomes less negative) as the propensity increases. The effect nears zero for children with a high propensity for parental divorce. The pattern in effects is curvilinear for high school completion (with little difference between children whose parents had low and moderate likelihoods of divorce), and nearly linear for college attendance and completion. Yet in each case the general trend indicates a reduction in the negative effect of parental divorce on children’s education as the propensity for divorce increases, where the trend is steepest for college completion.

– FIGURE 3 ABOUT HERE –

We next present heterogeneous effect estimates by propensity score strata in Table 5. Given the shape of response functions, and to preserve cases, particularly at the tails of the propensity distribution where bias is most likely to occur, we construct three propensity score strata. Appendix

Table A1 provides descriptive statistics of the covariates and Appendix Table A2 of the outcomes and mediators across strata. Families in which divorce is most likely (stratum 3) generally have the most disadvantaged socioeconomic and family wellbeing attributes.¹⁷ As we expect, given the matching-smoothing results in Figure 3, we find no significant effects for children who have a high propensity for parental divorce (stratum 3). We find significant effects for children who have a low- and mid-propensity for parental divorce (i.e., strata 1 and 2), with the largest effects observed among children with the lowest propensity. We observe a 5 percent lower probability of high school completion (80 percent relative to 85 percent predicted value with the propensity held at the median), a 10 percent lower probability of college attendance (53 relative to 64 percent), and a 12 percent lower probability of college completion (18 relative to 30 percent) among children with a low propensity for parental divorce. We find substantially larger effects for children who have a low propensity for divorce than those we observe for the full sample (reported in Table 4), a consequence of overlooking cross-strata heterogeneity. That is, typically reported average effects under an assumption of effect homogeneity obscure large effects for low propensity children.¹⁸

We underscore that we are comparing the *effects* of parental divorce on children's educational outcomes across strata, not children's *levels* of educational attainment. Children whose parents are unlikely to divorce have advantaged family background characteristics and attain higher levels of education. Between the two dimensions, educational outcomes differ far more by the propensity to divorce, as a summary proxy for family status and wellbeing, than by parental divorce status. As a result, low propensity children with divorced parents outperform high propensity children with married parents. For example, about 53 percent of children attend college who have a low propensity but divorced parents, while about 43 percent of children attend college who have a high propensity but parents who do not divorce (see Appendix A2).¹⁹

Heterogeneous Mediating Effects of Parental Divorce on Children’s Educational Attainment

In Table 6, we report estimates of the mediating effects of family instability, family economic resources, and children’s cognitive and psychosocial skills. As each mediator is assessed in turn, the proportion mediated does not sum to 100 percent for a given stratum and outcome. In general, we find that divorce-induced declines in family income and children’s psychosocial skills account for a large portion of the effect for children with a low propensity, while declines in family income and increases in instability account for a large portion of the negative effect of divorce on children’s education for children with a moderate propensity for parental divorce. The divorce-induced decline in family income accounts for about 40 to (over) 100 percent of the total effect on educational attainment among children who have a low- to mid-propensity for parental divorce.²⁰ Income accounts for more of the association among mid- relative to low propensity children. The larger proportion mediated among children with a mid-propensity for parental divorce relative to children with a low propensity results from the smaller total effects of divorce coupled with larger post-divorce decline in family income among the former than among the latter. Divorce-induced changes in children’s psychosocial skills account for a modest proportion (about 10 to 17 percent) of the total effects of divorce, and only for children with a low propensity. This supports our theory that children who do not expect disruption to occur experience social-psychological distress, which in turn impacts their educational attainment. The same cannot be said for those with a higher likelihood of disruption. In general, the mediators explain more of the divorce effect on high school than on college. As effects on college are further removed from the measurement timing of the mediators than those on high school, this pattern is unsurprising.

Both the total and mediating effects must be significant to obtain meaningful estimates of the proportion mediated. We thus present no estimates of the proportion mediated for children's cognitive skills, as we find no significant mediating effects for cognitive skills among children who have a low- to mid-propensity. We also present no estimates of the proportion mediated for children with a high propensity for parental divorce for any of the mediators, as the total effects of divorce reported in Table 5 are all insignificant. We nevertheless assess the mediating effects for children with a high propensity for parental divorce to aid our understanding of the null total effects. We observe significant negative mediating effects of family instability and income. However, we also find an unexpected positive mediating effect of cognitive skills.²¹ We speculate that the dissolution of parental unions among high propensity, and potentially high-conflict, families results in some improvement in children's performance on achievement tests in school. As we note above, this could reflect measurement error in cognitive skills, such as improvements in concentration or motivation rather than true increases in cognitive functioning. These offsetting influences (i.e., a decline in family stability and income and an improvement in cognitive test performance) potentially explain the null effect on educational attainment among high propensity children.

– TABLE 6 ABOUT HERE –

Sensitivity Analysis for Heterogeneous Total and Mediating Effects of Parental Divorce on Children's Educational Attainment

In our preceding analyses, we invoked the sequential ignorability assumption. Whether this assumption is reasonable is a substantive rather than a methodological issue, which depends upon the quality of the exogenous covariates in capturing potential selection bias. We include an extensive set of covariates to predict divorce. Yet if there remain unobserved confounding

variables that impact parental divorce and the proposed mediators and outcome variables the ignorability assumption would be violated and causal effects unidentified (Keele, Tingley, and Yamamoto 2015). We recognize that even with a rich set of pre-treatment covariates, potential confounders remain (e.g., unobserved paternal characteristics). We address the possibility of unobserved confounding for the total and mediating effects with a series of sensitivity analyses.

In Table 7, we report sensitivity analysis results for total effects. We assume that the unobserved confounder is a binary variable, and assess values of (+/-)5 to (+/-)40 percent for γ with a value of -5 percent for λ . We fix λ at -5 percent because very few characteristics between children of divorced and non-divorced parents differ by more than 5 percent (see Appendix Tables A1 and A2). As the bias factor is the product of γ and λ , the effect reaches non-significance when the unobserved confounder has a strong effect on children's education or a large difference between children of divorced and those of non-divorced parents.²² Suppose, for example, that father's full-time employment status, unobserved in our data, enhances levels of education and is lower among fathers who get divorced (Killewald 2016). When λ equals -5 percent, we assume that the prevalence of fathers having been full-time employed is 5 percent less in the divorced group than in the non-divorced group; when γ equals 5 percent, we assume that children whose fathers are full-time employed have a 5 percent advantage in graduating from high school (or attending or completing college) over children whose fathers are not full-time employed (all else being held equal). From Table 7, we observe that the total effect of divorce on high school completion for children with a low propensity for parental divorce is reduced to non-significance with γ at 40 percent (and remains significant at -40 percent). For college attendance and completion total effects for children with a low propensity remain significant even with γ at 40 percent. Effects of divorce on educational attainment among children with a low propensity for parental divorce

are thus highly robust to unobserved confounding. Effects for children with a moderate propensity for parental divorce are reduced to non-significance with γ at 40 percent for college attendance, but reduced to non-significance with γ at 5 percent for high school and college completion.

– TABLE 7 ABOUT HERE –

We present sensitivity analysis results in Table 8 for mediating effects of family income, where λ remains fixed at -5 percent and γ ranges from (+/-)5 to (+/-)40 percent.²³ Suppose for example that maternal job displacement (Brand and Simon Thomas 2014) is an unobserved confounder in the relationship between family income and children’s education, such that λ and γ are both negative. That is, job displacement is more prevalent among divorced mothers and decreases children’s education attainment. Omitting this confounder, we may overstate the mediating effect of family income. We observe, however, that family income mediating effects remain significant with γ at 40 percent for almost all educational outcomes for children with low and moderate propensities for parental divorce. We present additional sensitivity analyses for mediating effects of family transitions and children’s skills in Appendix Table B1. Mediating effects of children’s psychosocial skills for low propensity children are generally reduced to non-significance with γ at -10 percent (but remain significant with γ at 40 percent). Family instability effects for children with a moderate propensity reach non-significance with γ at -40 percent for high school completion, but reach non-significance with γ at -5 percent for college attendance and completion. Thus, the family income mediating effect is very robust to unobserved confounding, while other mediating effects are more sensitive.

– TABLE 8 ABOUT HERE –

SUMMARY AND DISCUSSION

Children whose parents divorce tend to have lower levels of educational attainment than children whose parents stay together. Influential research, as well as public discourse, suggests that the effects of family disruption on children are uniformly negative. With careful attention to the assumptions needed to estimate total and mediating effects, we assessed whether the impact of parental divorce, and the mechanisms that account for it, vary across families with differing likelihoods of divorce. The results are noteworthy and comprehensible. Effects of parental divorce on children's educational attainment vary inversely with the likelihood of divorce. We find significant effects of divorce on children's education among those with a low-to-moderate likelihood of parental divorce. Conversely, we find no significant effect of divorce on children's education among those who have a high likelihood of parental divorce. Families prone to disruption have high levels of family and socioeconomic hardship and a context in which family shocks and economic distress are normative. That is, for these children, parental divorce is but one of many disadvantaged events faced during childhood. Thus, the effects of any particular adverse event are less disruptive and rendered less severe. Educational attainment rates among children whose parents have a high probability of divorce are relatively low, and these rates are roughly the same whether or not parents divorce. On the other hand, parental divorce may trigger an acute sense of deprivation among children whose peers tend to be advantaged and for whom family instability is uncommon and comes as a shock. In this case, divorce has a more profound impact upon children's educational trajectories. For them, educational attainment rates are generally high, yet significantly differ by parental divorce status.

Given the strong association between cognitive ability and educational outcomes, it is reasonable to consider whether ability mediates the negative effect of parental divorce. However, cognitive ability stabilizes early in childhood, leaving little time for parental divorce to reduce

children's aptitude, and thus to mediate the negative impact of disruptive family events on children's eventual attainment. Children's "non-cognitive" or psychosocial skills, on the other hand, are more malleable throughout childhood and susceptible to the influence of family shocks, and thus are potentially more important in linking disruption to children's education. Skills such as motivation, self-esteem, conscientiousness, and emotional wellbeing are also increasingly implicated as important determinants of educational and socioeconomic success. We find, as expected, that children's cognitive skills do not account for the negative effect of divorce on children's educational attainment. Psychosocial skills, on the other hand, explain a portion of the effect among children with a low propensity for divorce. That is, when family disruption occurs to children with a low expectation of disruption, their social-psychological wellbeing declines and leads to lower educational attainment. While children's psychosocial skills play a surprisingly modest mediating role, family income clearly provides the main explanation as to why divorce negatively impacts children's education among low propensity children. For children with a moderate likelihood of disruption, family income and instability dominate the explanation of divorce-induced declines in education, while psychosocial (and cognitive) skills play no explanatory role.

Our statistical estimation of causal total and mediation effects requires a strong, untestable assumption of sequential ignorability. If the sequential ignorability assumption holds true, we have obtained valid estimates of the total and mediating causal effects of parental divorce. Divorce is a highly selective process; we cannot plausibly account for all the factors that influence both parents' likelihood of divorce and children's educational outcomes and associated mediators. One key advantage and primary motivation for our focus on treatment effect heterogeneity by the propensity score is the heightened recognition of potential violations of the ignorability

assumption. A researcher can begin with the ignorability assumption in order to carry out meaningful analyses without necessarily committing oneself to the validity of the assumption (Xie 2013; Xie, Brand and Jann 2012). Indeed, even when unobserved selectivity is present, it is informative to understand variation in the treatment effect along the propensity score (Zhou and Xie 2016b). This is true because, as we demonstrated in equation (9), it is an average effect over unobserved selection. In this study, while we estimate heterogeneous treatment effects by the propensity for treatment, we allow for interpretation of such effects as revealing variation in treatment effects by unobserved resistance. Our analyses based on the ignorability assumption yields an informative pattern of heterogeneity in the effects of family disruption by the estimated propensity of parental divorce based on observed covariates. If we accept ignorability, the results suggest larger effects among children with a lower likelihood of parental divorce. If ignorability is unaccepted, however, we can also interpret the findings to reflect differential unobserved selectivity of parental divorce: our results then reveal an association between lower resistance to divorce and larger effects of divorce among children whose parents have a lower observed likelihood of divorce (Zhou and Xie 2016b). That is, given an observed low likelihood of divorce, parents must have unobserved characteristics that render them less resistant to divorce for a divorce to occur. Lending confidence to our substantive interpretation based on ignorability, sensitivity analyses indicate that observed total effects are highly robust to confounding.

It is implausible to assume that children respond identically to family disruption. Two questions of sociological importance are whether and why there are systematic variations that can be detected with observed characteristics in the impact of parental divorce on children's educational outcomes. This paper set out to answer these research questions, and has yielded fruitful answers. First, we find important variation in the disruptive effect of parental divorce by

the predicted likelihood of divorce, ranging from significant effects among children whose parents are unlikely to divorce to no effects among children whose parents are likely to divorce. Second, overwhelmingly, this effect is explained by divorce-induced declines in family income; declines in children's psychosocial skills also play a role for children with a low propensity for divorce. While the effect of divorce is seemingly greatest among more advantaged children who do not expect disruption, this is not to say that we should shift attention away from children of disadvantaged families. It is telling that the educational attainment among high propensity children is unaffected by parental divorce, suggesting that social discourse and policy aimed at promoting marital stability among disadvantaged families, without attending to socioeconomic and family conditions in which adverse events are expected, is misguided.

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ENDNOTES

¹ Family disruption includes several possible forms of change in family structure, with the main line of demarcation between adding and losing a partner. The loss of a partner is generally more negatively disruptive than the addition of a partner. Another line of demarcation lies between divorce among married parents versus union dissolution among cohabitating parents. The proportion of children born to cohabiting parents has increased over the last several decades (Bumpass and Lu 2000; Kennedy and Bumpass 2008), yet such children continue to be more disadvantaged relative to children born to marital unions (Osborne and McLanahan 2007). If cohabitating unions are more disadvantaged and unstable from the onset than marital unions, the effects of dissolution on children may be less severe (Brown 2006; McLanahan, Tach, and Schneider 2013). While we considered assessing loss of a cohabitating partner, adding these cases to

form a broader category of family disruption would have increased the heterogeneity of the effect estimates, and there were too few cases of those who began in cohabitating unions at the time of childbirth to allow for meaningful separate estimates (particularly among subgroups).

² Other potentially important sources of variation include gender and child age at parental divorce. Mitchell et al. (2015), for example, find larger effects for boys than for girls on antisocial behavior. They do not find differences by child age when family transitions occur.

³ Given the strong correlation between low income and neighborhood residence with a high proportion of non-traditional households (Cleveland and Gibson 2010), socioeconomically disadvantaged children are more likely to be embedded in a social context with peers who have experienced family instability.

⁴ Results, however, vary according to different measures of psychosocial skills (Aughinbaugh et al. 2005; Cherlin et al. 1991; D’Onofrio et al. 2007; Sun 2001; Sun and Li 2002).

⁵ Lee and McLanahan (2015) find that the effect of family instability is stronger for children’s socioemotional than for their cognitive skills.

⁶ We presume that potential mechanisms such as parental psychological wellbeing, parenting style, family relations, and residential instability (Cheadle and Amato 2010; Meadows, McLanahan, and Brooks-Gunn 2007; Turney 2011) will largely influence children’s educational attainment by way of their impact on children’s psychosocial skills. Another plausible factor is father’s financial contributions to college (Wallerstein and Blakeslee 1989). A portion of this mechanism may be captured by post-divorce family income. However, if fathers contribute directly to college, it would remain unaccounted for. We do not have data to capture this factor, particularly data to capture the intention of fathers to help finance college. We would need to measure the intention to contribute if we were to consider it as a mechanism for the population who do not attend college.

⁷ A critique of the linear probability model (*LPM*) is that it does not estimate the structural parameters of a non-linear model. However, the marginal effects, rather than structural parameters of a binary choice model, are our primary concern. The *LPM* performs reasonably well with respect to estimating the marginal effects. The *LPM* will not give the true marginal effects from the nonlinear model, yet neither will an incorrect

nonlinear model (Angrist and Pischke 2009). The *LPM* is particularly useful when we interpret mediation effects in terms of probabilities rather than odds ratios.

⁸ Breen, Choi, and Holm (2015) offer a critique of the propensity score approach for studying treatment effect heterogeneity. They begin with an assumption of homogeneity, and then apply an assumed distribution of positive selection bias in order to suggest an “erroneous” pattern of heterogeneity as described in Brand and Xie (2010). They do not, however, demonstrate or prove bias. Reconciling divergent findings requires understanding as to how observable and unobservable characteristics influence selection into treatment differently for subpopulations. Reversing the findings from Brand and Xie (2010), for example, depends upon a result in which there is an assumed negative effect of college on wages for low propensity college goers, a substantive result with little if any support from the literature in economics or sociology. Although Breen, Choi, and Holm (2015) take up no substantive issues that provide insight into these processes, Brand and Xie (2010) emphasize *both* selectivity bias and substantive interpretations, as we do here. We cannot, nor can Breen, Choi, and Holm (2015), adjudicate between selection and substantive interpretations. See Zhou and Xie (2016a, 2016b) for additional discussion on this issue. The analyses that form the basis of this paper, nevertheless, include both a nonparametric approach and a simple 3-strata approach (to minimize bias at the tails of the distribution), and not the full *stratification-multilevel* method that forms the basis of their critique.

⁹ The *MTE*, according to Heckman et al. (2006), is defined as the gain from treatment for subjects shifted from the control group into the treatment group by a marginal change in the predicted probability of treatment (namely, D^*). Standard estimands of causal inference, such as the average treatment effect (*ATE*) and the treatment effect of the treated (*TT*), are thus all weighted averages of *MTE* over dimensions of P and U (Zhou and Xie 2016a).

¹⁰ We can also define the direct effect (*DE*) for unit i as follows: $DE_i = Y_i(1, M_i(d)) - Y_i(0, M_i(d))$. The direct effect represents the expected difference in education as a result of all possible mechanisms other than the one under consideration by deactivating the pathway of the mediator under consideration. In the

absence of interactions with the treatment, the total effect is equivalent to the sum of the indirect and direct effects.

¹¹ The causal mediation effect obtained at $D=1$ may be different from that at $D=0$ if there is an interaction between the treatment and mediator variables. We did not find significant interactions for most mediators, and thus do not include them in the outcome models.

¹² As a pseudo measure of whether a mother desired a child's birth, we include a variable indicating the difference between a woman's desired number of children and the birth order position of the current child. "Desired birth" is defined as children for whom the birth order position is less than their mother's 1979 fertility preference. "Undesired births" is defined as children for whom birth order position exceeds fertility preference. For example, if a mother desired three children and the child is the fifth, the indicator for desired births takes a value of zero and the indicator for undesired births takes a value of two.

¹³ The Pearlin Mastery scale measures self-concept and indicates the extent to which individuals perceive themselves in control of forces that significantly impact their lives. The Rosenberg Self-Esteem scale measures the self-evaluation that an individual makes and customarily maintains. The TIPI (Ten Item Personality Index) is a ten-question instrument that yields five scales that measure the "Big Five" personality traits (i.e., extraversion, agreeableness, conscientiousness, emotional stability, and openness to experiences). Here, we focus on the Emotional Stability scale. The BPI measures the frequency, range, and type of childhood behavior problems for children. This includes measures for antisocial behavior, anxious or depressive behavior, conflict behavior, and dependent behavior. Finally, the CESD is a self-report scale that measures the prevalence of depression symptoms. Respondents were asked whether they never/rarely, sometimes, occasionally, or most/all of the time (1) had poor appetite; (2) had trouble keeping their mind on tasks; (3) were depressed; (4) felt that everything took extra effort; (5) had restless sleep; (6) were sad; and (7) could not get going.

¹⁴ If we had instead compared post-divorce mediator measures of the treated group to mediator measures of the control group across childhood, we would expect bias in our estimates due to differences in the child's mean age at the time of mediator measurement between the divorce and non-divorce groups.

¹⁵ Missing mediator values were imputed with a model based on covariates included in Table 1. Models that use imputed and non-imputed mediating variables produce substantively similar results.

¹⁶ In additional analyses (not shown), we use a variety of machine learning techniques (i.e., classification and regression trees (CART), and ensemble methods of random forests and boosted CART) to estimate the likelihood of parental divorce. Although some of these models improve fit, we use the logit specification here for simplicity. We also provide sensitivity analyses to determine how model misspecification may impact our results.

¹⁷ The average age at the time of parental divorce is roughly 7 years for all three propensity score strata.

¹⁸ This pattern of effect heterogeneity may explain results suggesting smaller effects than those we observe here of parental divorce on college attendance and completion using data from the Fragile Families and Child Wellbeing Study, which represents a sample of relatively disadvantaged families.

¹⁹ A limitation of our analyses is that we do not highlight other sources of variation in effects beyond the likelihood of parental divorce, such as gender, race, or timing of divorce, notably important axes of difference. We explored additional sources of variation in preliminary analyses, and findings were generally consistent with the existing literature. Although we considered describing these interactions as well, we decided that focusing on the interaction of parental divorce with the estimated likelihood of divorce advances the existing literature on family disruption in a more theoretically suggestive way.

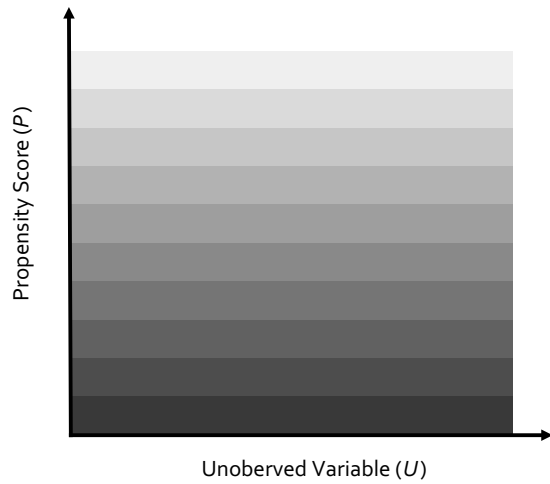
²⁰ When the mediation effect of one variable is over 100 percent, it suggests that the effect is offset by other mediators. For example, although children may suffer loss of family income from their parents' divorces, they may also receive more social support from extended family members, teachers, and friends. These positive and negative factors lead to a total negative effect of divorce smaller than the negative effect caused by the loss of family income alone.

²¹ The positive effect of parental divorce on children's cognitive skills is highly robust to diagnostic tests and different model specifications.

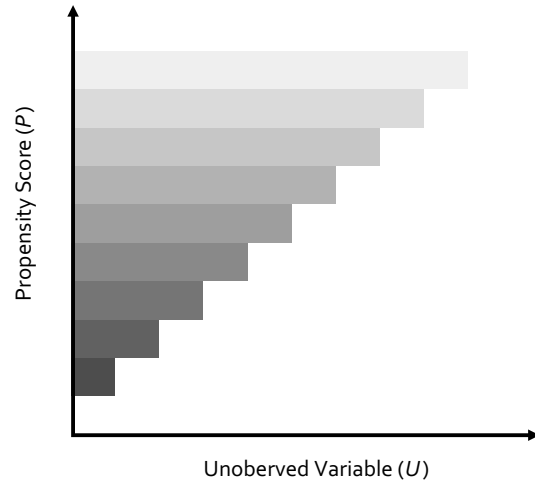
²² In additional sensitivity analyses (not shown), we also explore γ ranging from -5 to -40 and δ ranging from -10 to -20. If γ and δ are both negative, the unobserved confounder can decrease children's education,

and lead to downward bias of our estimates. Our estimates shift no more than about 2 percent even with γ at -40 percent.

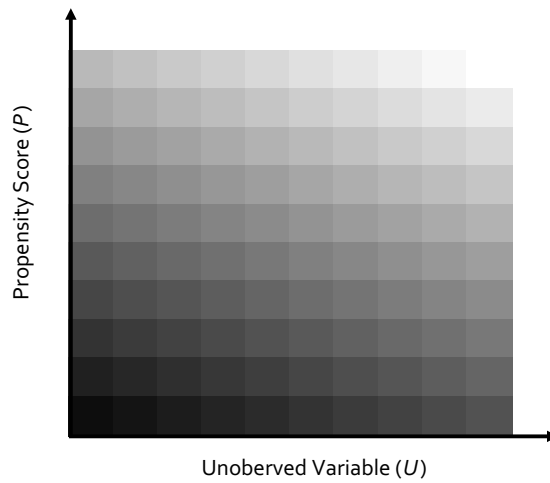
²³ In additional sensitivity analyses (not shown), we also explore γ ranging from 5 to 40 and δ ranging from -10 to -20. Our conclusions remain very similar.



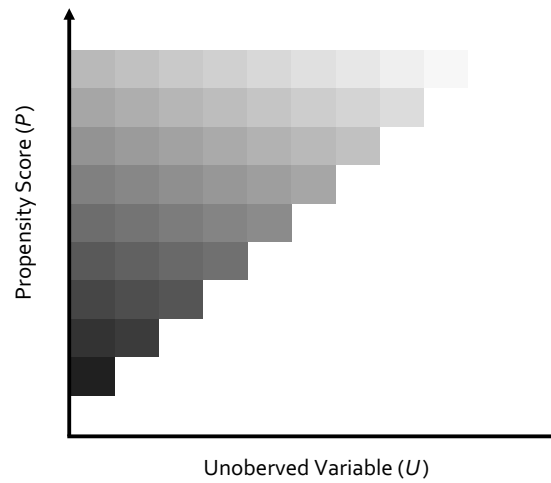
(a) MTE for all units under the ignorability assumption



(b) MTE for treated units under the ignorability assumption



(c) MTE for all units



(d) MTE for treated units

Figure 1. Depiction of Treatment Effect Heterogeneity by the Propensity for Parental Divorce (P) and Unobserved Resistance to Divorce (U)

Note: A darker color indicates a larger treatment effect. Figure 1(c) and 1(d) adapted from Zhou, Xiang and Yu Xie. 2016. "Estimating Heterogeneous Treatment Effects in the Presence of Self-Selection: A Propensity Score Perspective." Princeton University Working Paper.

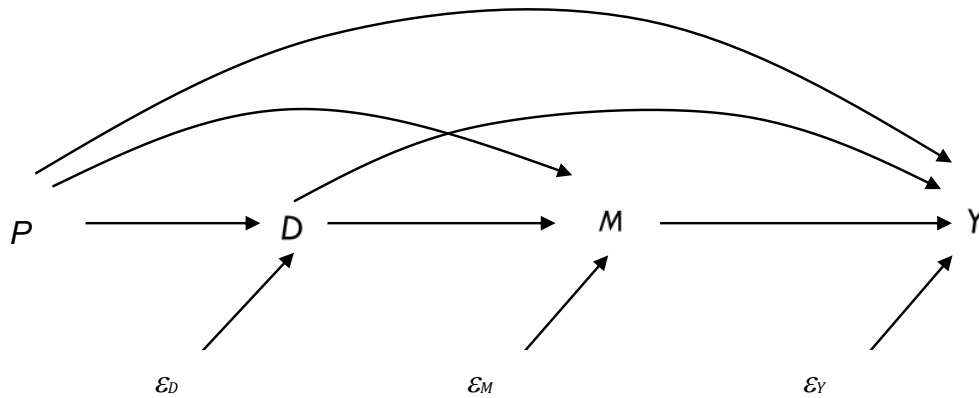
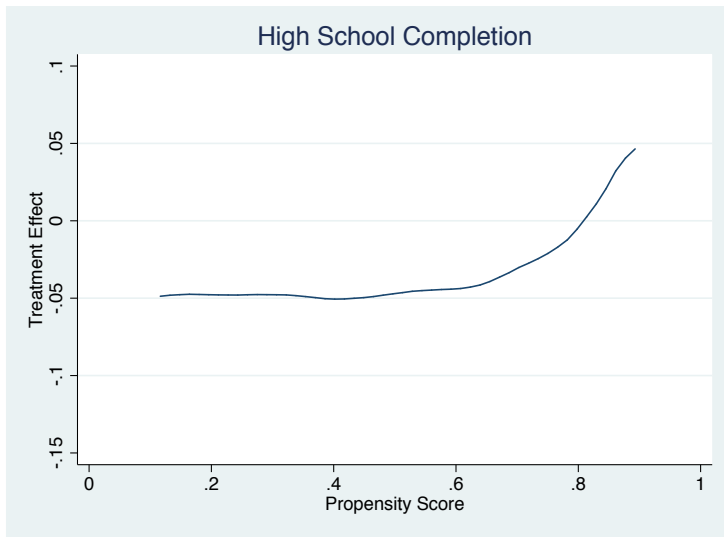
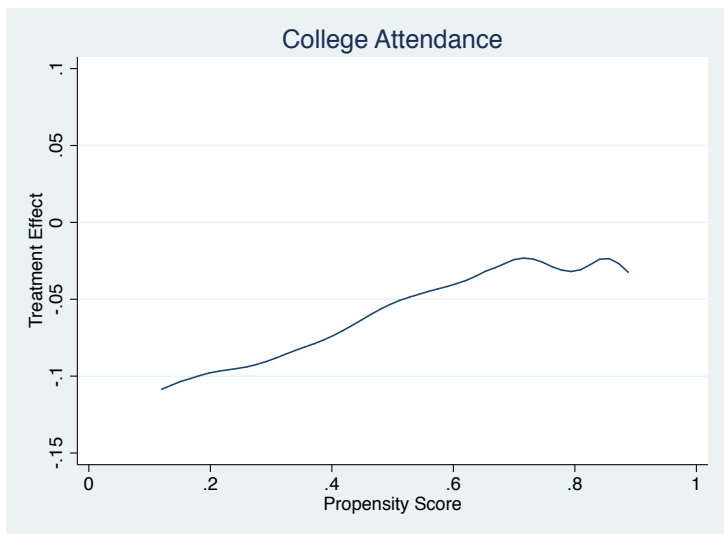


Figure 2. A Causal Framework based on a Directed Acyclic Graph (DAG)

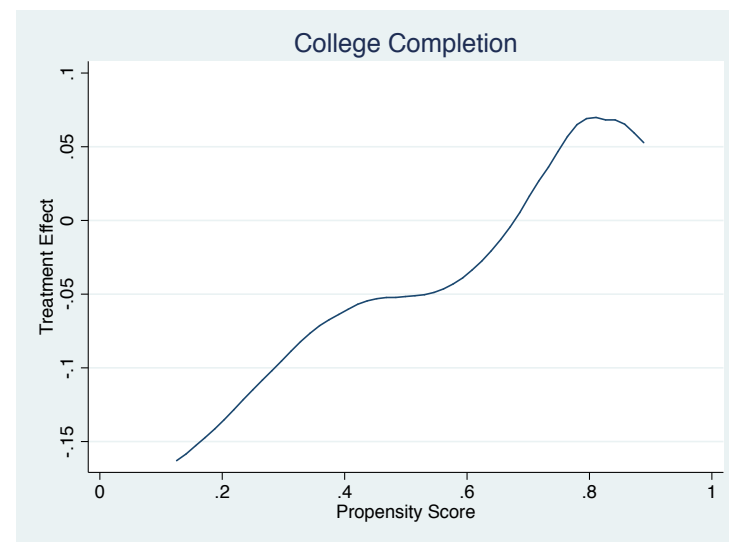
Notes: *P* = Propensity for parental divorce; *D* = Parental divorce; *M* = post-divorce mediators (family instability; family income; children's cognitive skills; children's psychosocial skills); *Y* = Children's educational attainment (high school completion; college attendance; college completion);



(a) MS Effects of Divorce on High School Completion



(b) MS Effects of Divorce on College Attendance



(c) MS Effects of Divorce on College Completion

Figure 3. Matching-Smoothing (MS) Heterogeneous Effects of Parental Divorce on Children's Educational Attainment

TABLE 1
DESCRIPTIVE STATISTICS OF PRE-DIVORCE CHARACTERISTICS (NLSY)

| | Parents not divorced | | Parents divorced | | t-tests |
|--|----------------------|----------|------------------|----------|---------|
| | Mean | (SD) | Mean | (SD) | |
| Family Background Factors | | | | | |
| Black (binary 0/1) | 0.090 | --- | 0.130 | --- | *** |
| Hispanic (binary 0/1) | 0.087 | --- | 0.089 | --- | |
| U.S. born (binary 0/1) | 0.935 | --- | 0.967 | --- | *** |
| Southern residence at age 14 (binary 0/1) | 0.310 | --- | 0.353 | --- | ** |
| Raised no religious preference (binary 0/1) | 0.025 | --- | 0.037 | --- | * |
| Intact family at age 14 (binary 0/1) | 0.783 | --- | 0.668 | --- | *** |
| Absent father before age 14 (binary 0/1) | 0.136 | --- | 0.217 | --- | *** |
| Sibship size (continuous 0-19) | 3.557 | (2.364) | 3.593 | (2.353) | |
| Parents' household income (\$1,000s) (continuous 0-75) | 19.443 | (12.436) | 16.055 | (10.423) | *** |
| Socioeconomic Factors | | | | | |
| Highest education is completed high school (binary 0/1) | 0.564 | --- | 0.585 | --- | |
| Highest education is completed college or more (binary 0/1) | 0.233 | --- | 0.083 | --- | *** |
| Employed (binary 0/1) | 0.538 | --- | 0.496 | --- | ** |
| Employed at a private company (binary 0/1) | 0.023 | --- | 0.010 | --- | * |
| Job offers flexible hours (binary 0/1) | 0.509 | --- | 0.503 | --- | |
| Delinquent activity (binary 0/1) | 0.652 | --- | 0.757 | --- | *** |
| Log household income (continuous 4-14) | 10.254 | (1.095) | 9.823 | (1.198) | *** |
| Household below poverty line (binary 0/1) | 0.132 | --- | 0.182 | --- | *** |
| Household received welfare/TANF (binary 0/1) | 0.096 | --- | 0.205 | --- | *** |
| Cognitive and Psychosocial Factors | | | | | |
| Rotter Locus of Control scale (continuous 4-16) | 8.447 | (2.446) | 8.851 | (2.409) | *** |
| Pearlin Mastery scale (continuous 9-28) | 22.203 | (3.027) | 21.708 | (3.224) | *** |
| Rosenberg Self-Esteem scale (continuous 240-650) | 482.448 | (80.279) | 465.792 | (82.049) | *** |
| Juvenile delinquent activity (binary 0/1) | 0.926 | --- | 0.941 | --- | * |
| CESD score (continuous 0-21) | 3.925 | (3.625) | 5.153 | (4.353) | *** |
| Body mass index (continuous 11-42) | 21.738 | (3.137) | 21.737 | (3.386) | |
| Cognitive ability ASVAB (continuous -3-3) | -0.038 | (0.684) | -0.194 | (0.632) | *** |
| High school class rank percentile (continuous 0-1) | 0.413 | (0.222) | 0.475 | (0.198) | *** |
| High school program was college prep (binary 0/1) | 0.325 | --- | 0.206 | --- | *** |
| Family Formation and Wellbeing Factors | | | | | |
| Sexual debut at age 15 or younger (binary 0/1) | 0.107 | --- | 0.171 | --- | *** |
| "Wife with family has no time for employment" (binary 0/1) | 0.172 | --- | 0.186 | --- | |
| Age at time of child's birth (continuous 13-37) | 26.266 | (4.454) | 24.124 | (4.625) | *** |
| Previously married (binary 0/1) | 0.091 | --- | 0.120 | --- | ** |
| Log months between marriage and first birth (continuous 0-5) | 2.754 | (1.325) | 2.532 | (1.312) | *** |
| Desired birth (continuous 0-13) ¹ | 1.127 | (1.307) | 0.982 | (1.399) | *** |
| Undesired birth (continuous 0-8) ¹ | 0.237 | (0.598) | 0.323 | (0.709) | *** |
| Child male (0/1) | 0.525 | --- | 0.511 | --- | |
| Child birth weight (ounces; continuous 6-268) | 120.150 | (20.091) | 117.667 | (20.125) | *** |
| Mother/father argue about chores often/very often (binary 0/1) | 0.190 | --- | 0.142 | --- | *** |

| | | | | |
|--|-------|-----|-------|---------|
| Mother/father argue about money often/very often (binary 0/1) | 0.207 | --- | 0.091 | --- *** |
| Mother/father argue about cheating often/very often (binary 0/1) | 0.082 | --- | 0.072 | --- |
| Mother/father argue about religion often/very often (binary 0/1) | 0.032 | --- | 0.018 | --- ** |
| Mother/father different race (binary 0/1) | 0.089 | --- | 0.126 | --- *** |
| Mother/father raised different religious preference (binary 0/1) | 0.457 | --- | 0.408 | --- ** |
| Mother/father difference in college completion (binary 0/1) | 0.005 | --- | 0.052 | --- *** |
| <hr/> | | | | |
| Weighted sample proportion | 0.60 | | 0.40 | |
| <i>N</i> | 4,838 | | 2,420 | |

Notes: Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Factors refer to mothers unless otherwise specified. All factors are measured prior to the divorce interval, i.e. at the time of child's birth or earlier. All descriptive statistics are weighted by the NLSY sample weight.

1. "Desired birth" is the extent to which mother's 1979 fertility preference meets or exceeds child's birth order.

"Undesired birth" is the extent to which child's birth order exceeds mother's 1979 fertility preference. Each measure equals zero when the measure does not go in the stated direction.

† $p \leq 0.1$ * $p \leq 0.05$ ** $p \leq 0.01$ *** $p \leq 0.001$; two-tailed tests;

TABLE 2
DESCRIPTIVE STATISTICS OF MEDIATORS / OUTCOMES (NLSY)

| | Parents not divorced | | Parents divorced | | t- tests |
|---|----------------------|---------|------------------|---------|-------------|
| | Mean | (SD) | Mean | (SD) | |
| Mediators | | | | | |
| Family Instability | | | | | |
| Relationship transitions (continuous) | 1.289 | (1.075) | 2.596 | (1.615) | *** |
| Family Economic Resources | | | | | |
| Family income (continuous) | \$74,861 | (67673) | \$39,330 | (41156) | *** |
| Children's Skills | | | | | |
| Psychosocial skills scale (continuous) | 0.592 | (0.134) | 0.550 | (0.135) | *** |
| Cognitive skills scale (continuous) | 0.574 | (0.159) | 0.540 | (0.142) | *** |
| Outcomes | | | | | |
| Children's Educational Attainment | | | | | |
| High school completion (by age 18; binary 0/1) | 0.853 | (0.354) | 0.762 | (0.426) | *** |
| College attendance (by age 19; binary 0/1) [^] | 0.624 | (0.485) | 0.463 | (0.499) | *** |
| College completion (by age 23; binary 0/1) [^] | 0.300 | (0.458) | 0.141 | (0.348) | *** |
| <i>N</i> [^] | 3,264 | | 1,912 | | |

Notes: Sample restricted to children whose parents were married at the time of their birth, for children who were at least 18 years old in 2012, and for children with no missing data on educational attainment. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Missing values are imputed for family status and children's cognitive and psychosocial skills. Relationship transitions counts how many changes occurred during childhood between the categories of married, separated, remarried, cohabitating, and widowed for both those married and divorced families. All descriptive statistics are weighted by the NLSY sample weight.

[^] Sample is further restricted to age 19 and above for college attendance ($N = 4,982$), and to age 23 and above for college completion ($N = 3,901$).

† $p \leq 0.1$ * $p \leq 0.05$ ** $p \leq 0.01$ *** $p \leq 0.001$; two-tailed tests;

TABLE 3
REGRESSION ESTIMATES PREDICTING CHILD EXPERIENCING PARENTAL
DIVORCE

| | β /(SE) |
|---|-----------------------|
| Family Background Factors | |
| Black (binary 0/1) | 0.027 (0.087) |
| Hispanic (binary 0/1) | -0.029 (0.083) |
| U.S. born (binary 0/1) | 0.471 *** (0.116) |
| Southern residence at age 14 (binary 0/1) | 0.122 * (0.061) |
| Raised no religious preference (binary 0/1) | -0.064 (0.150) |
| Intact family at age 14 (binary 0/1) | -0.112 (0.073) |
| Absent father before age 14 (binary 0/1) | 0.192 * (0.084) |
| Sibship size (continuous 0-19) | -0.061 *** (0.011) |
| Parents' household income (\$1,000s) (continuous 0-75) | -0.0001 (0.000) |
| Socioeconomic Factors | |
| Highest education is completed high school (binary 0/1) | -0.163 * (0.075) |
| Highest education is completed college or more (binary 0/1) | -0.899 *** (0.140) |
| Employed (binary 0/1) | 0.320 ** (0.101) |
| Employed at a private company (binary 0/1) | 0.375 † (0.207) |
| Job offers flexible hours (binary 0/1) | -0.194 * (0.099) |
| Delinquent activity (binary 0/1) | 0.231 *** (0.061) |
| Log household income (continuous 4-14) | -0.059 * (0.028) |
| Household below poverty line (binary 0/1) | -0.083 (0.072) |
| Household received welfare/TANF (binary 0/1) | 0.412 *** (0.074) |
| Cognitive and Psychosocial Factors | |
| Rotter Locus of Control scale (continuous 4-16) | 0.005 (0.012) |
| Pearlin Mastery scale (continuous 9-28) | 0.026 * (0.011) |

| | |
|--|-----------------------|
| Rosenberg Self-Esteem scale (continuous 240-650) | -0.001 † (0.000) |
| Juvenile delinquent activity (binary 0/1) | 0.058 (0.109) |
| CESD score (continuous 0-21) | 0.050 *** (0.008) |
| Body mass index (continuous 11-42) | -0.034 *** (0.008) |
| Cognitive ability ASVAB (continuous -3-3) | 0.144 ** (0.053) |
| High school class rank percentile (continuous 0-1) | 0.338 * (0.164) |
| High school program was college prep (binary 0/1) | -0.057 (0.072) |
| Family Formation and Wellbeing Factors | |
| Sexual debut at age 15 or younger (binary 0/1) | 0.182 * (0.076) |
| "Wife with family has no time for employment" (binary 0/1) | -0.184 ** (0.070) |
| Age at time of child's birth (continuous 13-37) | -0.092 *** (0.010) |
| Previously married (binary 0/1) | 0.447 *** (0.094) |
| Log months between marriage and first birth (continuous 0-5) | 0.191 *** (0.021) |
| Desired birth (continuous 0-13) | -0.212 *** (0.042) |
| Desired birth (squared) | 0.025 *** (0.006) |
| Undesired birth (continuous 0-8) | 0.101 * (0.043) |
| Child male (0/1) | -0.070 (0.054) |
| Child birth weight (ounces; continuous 6-268) | -0.003 * (0.001) |
| Mother/father argue about chores often/very often (binary 0/1) | 0.215 * (0.110) |
| Mother/father argue about money often/very often (binary 0/1) | -0.419 *** (0.106) |
| Mother/father argue about cheating often/very often (binary 0/1) | 0.101 (0.139) |
| Mother/father argue about religion often/very often (binary 0/1) | -0.139 (0.211) |
| Mother/father different race (binary 0/1) | 0.191 † (0.099) |
| Mother/father raised different religious preference (binary 0/1) | -0.192 ** (0.063) |
| Mother/father difference in college completion (binary 0/1) | 2.279 *** |

| | |
|-----------------------------------|-----------|
| Intercept | (0.253) |
| | 1.994 *** |
| | (0.502) |
| <hr/> | |
| <i>N</i> | 7258 |
| <i>Log Likelihood</i> | -4106.20 |
| <i>P > χ^2</i> | 0 |
| <hr/> | |

Notes: Numbers in parentheses are standard errors. Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Factors refer to mothers unless otherwise specified. All factors are measured prior to the divorce interval, i.e. at the time of child's birth or earlier.

† $p \leq 0.1$ * $p \leq 0.05$ ** $p \leq 0.01$ *** $p \leq 0.001$; two-tailed tests;

TABLE 4
TOTAL EFFECTS OF PARENTAL DIVORCE ON CHILDREN'S
EDUCATIONAL ATTAINMENT

| | Unadjusted total effects | Adjusted total effects |
|--|-----------------------------|---------------------------|
| Educational Attainment Outcomes | | |
| High school completion | -0.078 *** (0.012) | -0.043 *** (0.013) |
| College attendance | -0.122 *** (0.015) | -0.068 *** (0.015) |
| College completion | -0.112 *** (0.012) | -0.065 *** (0.013) |

Notes: Numbers in parentheses are standard errors. Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Estimates are based on linear probability models. Adjusted models control for propensity of parental divorce and children's age in 2012 (estimates not shown). Propensity scores were estimated by a logit regression model of parental divorce on the set of pre-divorce covariates. Analytic sample ($N = 5,176$) is further restricted to age 19 and above for college attendance ($N = 4,982$), and age 23 and above for college completion ($N = 3,901$).

† $p \leq 0.1$ * $p \leq 0.05$ ** $p \leq 0.01$ *** $p \leq 0.001$; two-tailed tests;

TABLE 5
HETEROGENEOUS TOTAL EFFECTS OF PARENTAL DIVORCE ON
CHILDREN'S EDUCATIONAL ATTAINMENT

| | Stratum 1: Divorce least likely Total effects | Stratum 2: Divorce moderately likely Total effects | Stratum 3: Divorce most likely Total effects |
|--|--|--|---|
| Educational Attainment Outcomes | | | |
| High school completion | -0.050 ** (0.018) | -0.039 † (0.022) | -0.036 (0.029) |
| College attendance | -0.101 *** (0.024) | -0.066 * (0.026) | -0.007 (0.032) |
| College completion | -0.122 *** (0.023) | -0.039 † (0.020) | -0.019 (0.022) |

Notes: Numbers in parentheses are standard errors. Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Estimates are based on linear probability models.

Adjusted models control for propensity of parental divorce and children's age in 2012 (estimates not shown). Propensity scores were estimated by a logit regression model of parental divorce on the set of pre-divorce covariates. Analytic sample (N = 5,176) is further restricted to age 19 and above for college attendance (N = 4,982), and age 23 and above for college completion (N = 3,901).

† p ≤ 0.1 * p ≤ 0.05 ** p ≤ 0.01 *** p ≤ 0.001; two-tailed tests;

TABLE 6
HETEROGENEOUS MEDIATION EFFECTS OF PARENTAL DIVORCE ON CHILDREN'S
EDUCATIONAL ATTAINMENT

| | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|--|------------------------------------|---------------|---|---------------|-----------------------------------|---------------|
| | Mediation effects | Prop. med. | Mediation effects | Prop. med. | Mediation effects | Prop. med. |
| <i>Mediators of Effects of Divorce on High School Completion</i> | | | | | | |
| Family Instability | | | | | | |
| Relationship transitions | 0.001 (0.007) | --- | -0.033 *** (0.009) | 82% | -0.024 ** (0.008) | --- |
| Family Economic Resources | | | | | | |
| Family income | -0.035 *** (0.005) | 69% | -0.045 *** (0.007) | 110% | -0.033 *** (0.008) | --- |
| Children's Skills | | | | | | |
| Cognitive skills scale | -0.007 (0.004) | --- | 0.005 (0.004) | --- | 0.033 *** (0.010) | --- |
| Psychosocial skills scale | -0.009 *** (0.002) | 17% | -0.005 (0.004) | --- | -0.010 (0.006) | --- |
| <i>Mediators of Effects of Divorce on College Attendance</i> | | | | | | |
| Family Instability | | | | | | |
| Relationship transitions | -0.010 (0.010) | --- | -0.022 * (0.010) | 33% | -0.016 † (0.009) | --- |
| Family Economic Resources | | | | | | |
| Family income | -0.058 *** (0.007) | 57% | -0.055 *** (0.008) | 81% | -0.025 *** (0.006) | --- |
| Children's Skills | | | | | | |
| Cognitive skills scale | -0.009 (0.008) | --- | 0.011 (0.008) | --- | 0.045 *** (0.012) | --- |
| Psychosocial skills scale | -0.013 ** (0.004) | 13% | -0.004 (0.004) | --- | -0.007 (0.005) | --- |
| <i>Mediators of Effects of Divorce on College Completion</i> | | | | | | |
| Family Instability | | | | | | |
| Relationship transitions | -0.021 † (0.012) | 17% | -0.016 † (0.009) | 39% | -0.016 ** (0.006) | --- |
| Family Economic Resources | | | | | | |
| Family income | -0.050 *** (0.007) | 41% | -0.028 *** (0.005) | 70% | -0.013 *** (0.004) | --- |
| Children's Skills | | | | | | |
| Cognitive skills scale | -0.001 (0.010) | --- | 0.009 (0.006) | --- | 0.026 *** (0.008) | --- |
| Psychosocial skills scale | -0.013 ** | 10% | -0.003 | --- | -0.004 | --- |

(0.005)

(0.004)

(0.003)

Notes: Numbers in parentheses are standard errors. Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Estimates are based on linear probability models. All models control for propensity of parental divorce and children's age in 2012 (estimates not shown). Propensity scores were estimated by a logit regression model of parental divorce on the set of pre-divorce covariates. Proportion mediated is only reported when the total effect (reported in Table 5) and the indirect effect are both significant. $N = 5,176$, and further restricted to age 19 and above for college attendance ($N = 4,982$), and to age 23 and above for college completion ($N = 3,901$).

† $p \leq 0.1$ * $p \leq 0.05$ ** $p \leq 0.01$ *** $p \leq 0.001$; two-tailed tests;

TABLE 7
SENSITIVITY RESULTS FOR HETEROGENEOUS TOTAL EFFECTS OF PARENTAL DIVORCE ON
CHILDREN'S EDUCATIONAL ATTAINMENT

| | Sensitivity parameters | | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|--|------------------------|-------------|---------------------------------|------------------|--------------------------------------|------------------|--------------------------------|-----------------|
| | γ_s | λ_s | Total effects | CI | Total effects | CI | Total effects | CI |
| | | | | | | | | |
| Educational Attainment Outcomes | | | | | | | | |
| High School Completion | -40% | -5% | -0.070 | (-0.106, -0.034) | -0.059 | (-0.102, -0.016) | -0.056 | (-0.113, 0.001) |
| | -20% | -5% | -0.060 | (-0.096, -0.024) | -0.049 | (-0.092, -0.006) | -0.046 | (-0.103, 0.011) |
| | -10% | -5% | -0.048 | (-0.084, -0.012) | -0.036 | (-0.079, 0.007) | -0.033 | (-0.090, 0.024) |
| | -5% | -5% | -0.053 | (-0.089, -0.017) | -0.041 | (-0.084, 0.002) | -0.038 | (-0.095, 0.019) |
| | 5% | -5% | -0.048 | (-0.084, -0.012) | -0.036 | (-0.079, 0.007) | -0.033 | (-0.090, 0.024) |
| | 10% | -5% | -0.045 | (-0.081, -0.009) | -0.034 | (-0.077, 0.009) | -0.031 | (-0.088, 0.026) |
| | 20% | -5% | -0.040 | (-0.076, -0.004) | -0.029 | (-0.072, 0.014) | -0.026 | (-0.083, 0.031) |
| | 40% | -5% | -0.030 | (-0.066, 0.006) | -0.019 | (-0.062, 0.024) | -0.016 | (-0.073, 0.041) |
| College Attendance | -40% | -5% | -0.121 | (-0.168, -0.075) | -0.086 | (-0.137, -0.035) | -0.027 | (-0.090, 0.036) |
| | -20% | -5% | -0.111 | (-0.158, -0.065) | -0.076 | (-0.127, -0.025) | -0.017 | (-0.080, 0.046) |
| | -10% | -5% | -0.106 | (-0.153, -0.060) | -0.071 | (-0.122, -0.020) | -0.012 | (-0.075, 0.051) |
| | -5% | -5% | -0.104 | (-0.150, -0.057) | -0.069 | (-0.120, -0.017) | -0.009 | (-0.073, 0.054) |
| | 5% | -5% | -0.099 | (-0.145, -0.052) | -0.064 | (-0.115, -0.012) | -0.004 | (-0.068, 0.059) |
| | 10% | -5% | -0.096 | (-0.143, -0.050) | -0.061 | (-0.112, -0.010) | -0.002 | (-0.065, 0.061) |
| | 20% | -5% | -0.091 | (-0.138, -0.045) | -0.056 | (-0.107, -0.005) | 0.003 | (-0.060, 0.066) |
| | 40% | -5% | -0.081 | (-0.128, -0.035) | -0.046 | (-0.097, 0.005) | 0.013 | (-0.050, 0.076) |
| College Completion | -40% | -5% | -0.142 | (-0.187, -0.098) | -0.059 | (-0.099, -0.019) | -0.039 | (-0.082, 0.004) |
| | -20% | -5% | -0.132 | (-0.177, -0.088) | -0.049 | (-0.089, -0.009) | -0.029 | (-0.072, 0.014) |
| | -10% | -5% | -0.127 | (-0.172, -0.083) | -0.044 | (-0.084, -0.004) | -0.024 | (-0.067, 0.019) |
| | -5% | -5% | -0.125 | (-0.169, -0.080) | -0.041 | (-0.081, -0.001) | -0.022 | (-0.065, 0.021) |
| | 5% | -5% | -0.120 | (-0.164, -0.075) | -0.036 | (-0.076, 0.004) | -0.017 | (-0.060, 0.026) |
| | 10% | -5% | -0.117 | (-0.162, -0.073) | -0.034 | (-0.074, 0.006) | -0.014 | (-0.057, 0.029) |
| | 20% | -5% | -0.112 | (-0.157, -0.068) | -0.029 | (-0.069, 0.011) | -0.009 | (-0.052, 0.034) |
| | 40% | -5% | -0.102 | (-0.147, -0.058) | -0.019 | (-0.059, 0.021) | 0.001 | (-0.042, 0.044) |

Notes: γ_s refers to the mean difference in children's education associated with a unit difference in the binary unobserved confounder conditional on divorce status and propensity scores. λ_s refers to the prevalence difference of the binary unobserved confounder between the children of divorced and non-divorced parents conditional on propensity scores. The bias factor is equal to the product of the two parameters, and we subtract this bias factor from the total effect and the confidence interval.

TABLE 8

SENSITIVITY RESULTS FOR HETEROGENEOUS MEDIATION EFFECTS OF PARENTAL DIVORCE ON CHILDREN'S EDUCATIONAL ATTAINMENT

| | Sensitivity parameters | | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|--|------------------------|-------------|---------------------------------|------------------|--------------------------------------|------------------|--------------------------------|------------------|
| | γ_s | λ_s | Med. effects | CI | Med. effects | CI | Med. effects | CI |
| <i>Mediators for Effects on High School</i> | | | | | | | | |
| Family Economic Resources | | | | | | | | |
| Family income | -40% | -5% | -0.015 | (-0.024, -0.006) | -0.025 | (-0.038, -0.012) | -0.013 | (-0.028, 0.002) |
| | -20% | -5% | -0.025 | (-0.034, -0.016) | -0.035 | (-0.048, -0.022) | -0.023 | (-0.038, -0.008) |
| | -10% | -5% | -0.030 | (-0.039, -0.021) | -0.040 | (-0.053, -0.027) | -0.028 | (-0.043, -0.013) |
| | -5% | -5% | -0.032 | (-0.041, -0.023) | -0.042 | (-0.055, -0.029) | -0.030 | (-0.046, -0.015) |
| | 5% | -5% | -0.037 | (-0.046, -0.028) | -0.047 | (-0.060, -0.034) | -0.035 | (-0.051, -0.020) |
| | 10% | -5% | -0.040 | (-0.049, -0.031) | -0.050 | (-0.063, -0.037) | -0.038 | (-0.053, -0.023) |
| | 20% | -5% | -0.045 | (-0.054, -0.036) | -0.055 | (-0.068, -0.042) | -0.043 | (-0.058, -0.028) |
| | 40% | -5% | -0.055 | (-0.064, -0.046) | -0.065 | (-0.078, -0.052) | -0.053 | (-0.068, -0.038) |
| <i>Mediators for Effects on College Attendance</i> | | | | | | | | |
| Family Economic Resources | | | | | | | | |
| Family income | -40% | -5% | -0.038 | (-0.052, -0.024) | -0.035 | (-0.050, -0.019) | -0.005 | (-0.017, 0.008) |
| | -20% | -5% | -0.048 | (-0.062, -0.034) | -0.045 | (-0.060, -0.029) | -0.015 | (-0.027, -0.002) |
| | -10% | -5% | -0.053 | (-0.067, -0.039) | -0.050 | (-0.065, -0.034) | -0.020 | (-0.032, -0.007) |
| | -5% | -5% | -0.056 | (-0.069, -0.042) | -0.052 | (-0.068, -0.036) | -0.022 | (-0.035, -0.010) |
| | 5% | -5% | -0.061 | (-0.074, -0.047) | -0.057 | (-0.073, -0.041) | -0.027 | (-0.040, -0.015) |
| | 10% | -5% | -0.063 | (-0.077, -0.049) | -0.060 | (-0.075, -0.044) | -0.030 | (-0.042, -0.017) |
| | 20% | -5% | -0.068 | (-0.082, -0.054) | -0.065 | (-0.080, -0.049) | -0.035 | (-0.047, -0.022) |
| | 40% | -5% | -0.078 | (-0.092, -0.064) | -0.075 | (-0.090, -0.059) | -0.045 | (-0.057, -0.032) |
| <i>Mediators for Effects on College Completion</i> | | | | | | | | |
| Family Economic Resources | | | | | | | | |
| Family income | -40% | -5% | -0.030 | (-0.045, -0.016) | -0.008 | (-0.019, 0.002) | 0.007 | (0.000, 0.015) |
| | -20% | -5% | -0.040 | (-0.055, -0.026) | -0.018 | (-0.029, -0.008) | -0.003 | (-0.010, 0.005) |
| | -10% | -5% | -0.045 | (-0.060, -0.031) | -0.023 | (-0.034, -0.013) | -0.008 | (-0.015, 0.000) |
| | -5% | -5% | -0.048 | (-0.062, -0.033) | -0.026 | (-0.037, -0.015) | -0.010 | (-0.018, -0.003) |
| | 5% | -5% | -0.053 | (-0.067, -0.038) | -0.031 | (-0.042, -0.020) | -0.015 | (-0.023, -0.008) |
| | 10% | -5% | -0.055 | (-0.070, -0.041) | -0.033 | (-0.044, -0.023) | -0.018 | (-0.025, -0.010) |
| | 20% | -5% | -0.060 | (-0.075, -0.046) | -0.038 | (-0.049, -0.028) | -0.023 | (-0.030, -0.015) |
| | 40% | -5% | -0.070 | (-0.085, -0.056) | -0.048 | (-0.059, -0.038) | -0.033 | (-0.040, -0.025) |

Notes: γ_s refers to the mean difference in children's education associated with a unit difference in the binary unobserved confounder conditional on divorce status, propensity scores, and mediator values. λ_s refers to the prevalence difference of the binary unobserved confounder between children of divorced and non-divorced parents conditional on divorce status, propensity scores, and mediator values. The bias factor is equal to the negation of the product of the two parameters, and we subtract this bias factor from the mediation effect and the confidence interval.

APPENDIX TABLE A1

DESCRIPTIVE STATISTICS OF PRE-DIVORCE CHARACTERISTICS BY STRATA (NLSY)

| | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|---|------------------------------------|-----------|---|----------|-----------------------------------|----------|
| | Mean | (SD) | Mean | (SD) | Mean | (SD) |
| | | | | | | |
| Family Background Factors | | | | | | |
| Black (binary 0/1) | 0.065 | (0.247) | 0.130 | (0.336) | 0.130 | (0.336) |
| Hispanic (binary 0/1) | 0.081 | (0.273) | 0.096 | (0.295) | 0.096 | (0.295) |
| U.S. born (binary 0/1) | 0.921 | (0.269) | 0.974 | (0.158) | 0.974 | (0.158) |
| Southern residence at age 14 (binary 0/1) | 0.257 | (0.437) | 0.359 | (0.480) | 0.359 | (0.480) |
| Raised no religious preference (binary 0/1) | 0.024 | (0.154) | 0.035 | (0.183) | 0.035 | (0.183) |
| Intact family at age 14 (binary 0/1) | 0.836 | (0.370) | 0.679 | (0.467) | 0.679 | (0.467) |
| Absent father before age 14 (binary 0/1) | 0.094 | (0.292) | 0.225 | (0.418) | 0.225 | (0.418) |
| Sibship size (continuous 0-19) | 3.646 | (2.432) | 3.547 | (2.317) | 3.547 | (2.317) |
| Parents' household income (\$1,000s) (continuous 0-75) | 206.971 | (127.181) | 159.492 | ##### | 159.492 | ##### |
| Socioeconomic Factors | | | | | | |
| Highest education is completed high school (binary 0/1) | 0.576 | (0.494) | 0.638 | (0.481) | 0.638 | (0.481) |
| Highest education is completed college or more (binary 0/1) | 0.308 | (0.462) | 0.008 | (0.091) | 0.008 | (0.091) |
| Employed (binary 0/1) | 0.561 | (0.496) | 0.496 | (0.500) | 0.496 | (0.500) |
| Employed at a private company (binary 0/1) | 0.036 | (0.187) | 0.016 | (0.126) | 0.016 | (0.126) |
| Job offers flexible hours (binary 0/1) | 0.494 | (0.500) | 0.450 | (0.498) | 0.450 | (0.498) |
| Delinquent activity (binary 0/1) | 0.597 | (0.491) | 0.780 | (0.414) | 0.780 | (0.414) |
| Log household income (continuous 4-14) | 10.501 | (0.894) | 9.795 | (1.038) | 9.795 | (1.038) |
| Household below poverty line (binary 0/1) | 0.110 | (0.313) | 0.187 | (0.390) | 0.187 | (0.390) |
| Household received welfare/TANF (binary 0/1) | 0.044 | (0.205) | 0.153 | (0.360) | 0.153 | (0.360) |
| Cognitive and Psychosocial Factors | | | | | | |
| Rotter Locus of Control scale (continuous 4-16) | 8.256 | (2.427) | 8.895 | (2.400) | 8.895 | (2.400) |
| Pearlin Mastery scale (continuous 9-28) | 22.333 | (3.093) | 21.859 | (3.008) | 21.859 | (3.008) |
| Rosenberg Self-Esteem scale (continuous 240-650) | 490.217 | (79.982) | 461.757 | (80.275) | 461.757 | (80.275) |
| Juvenile delinquent activity (binary 0/1) | 0.922 | (0.267) | 0.949 | (0.220) | 0.949 | (0.220) |
| CESD score (continuous 0-21) | 3.368 | (3.113) | 4.960 | (4.003) | 4.960 | (4.003) |
| Body mass index (continuous 11-42) | 21.740 | (3.206) | 21.877 | (3.219) | 21.877 | (3.219) |
| Cognitive ability ASVAB (continuous -3-3) | 0.026 | (0.691) | -0.230 | (0.606) | -0.230 | (0.606) |
| High school class rank percentile (continuous 0-1) | 0.380 | (0.219) | 0.496 | (0.197) | 0.496 | (0.197) |
| High school program was college prep (binary 0/1) | 0.382 | (0.486) | 0.173 | (0.379) | 0.173 | (0.379) |
| Family Formation and Wellbeing Factors | | | | | | |
| Sexual debut at age 15 or younger (binary 0/1) | 0.058 | (0.235) | 0.164 | (0.370) | 0.164 | (0.370) |
| "Wife with family has no time for employment" (binary 0/1) | 0.175 | (0.380) | 0.175 | (0.380) | 0.175 | (0.380) |
| Age at time of child's birth (continuous 13-37) | 27.520 | (3.915) | 23.685 | (4.013) | 23.685 | (4.013) |
| Previously married (binary 0/1) | 0.064 | (0.245) | 0.131 | (0.337) | 0.131 | (0.337) |
| Log months between marriage - first birth (continuous 0-5) | 2.724 | (1.373) | 2.657 | (1.290) | 2.657 | (1.290) |
| Desired birth (continuous 0-13) ¹ | 1.231 | (1.304) | 0.913 | (1.318) | 0.913 | (1.318) |
| Undesired birth (continuous 0-8) ¹ | 0.208 | (0.564) | 0.321 | (0.680) | 0.321 | (0.680) |
| Child male (0/1) | 0.534 | (0.499) | 0.513 | (0.500) | 0.513 | (0.500) |
| Child birth weight (ounces; continuous 6-268) | 121.959 | (18.936) | 117.468 | (20.831) | 117.468 | (20.831) |

| | | | | | | |
|---|-------|---------|-------|---------|-------|---------|
| Mother/father argue about chores often/very often (binary | 0.218 | (0.413) | 0.123 | (0.329) | 0.123 | (0.329) |
| Mother/father argue about money often/very often (binary | 0.257 | (0.437) | 0.056 | (0.231) | 0.056 | (0.231) |
| Mother/father argue about cheating often/very often | 0.092 | (0.289) | 0.066 | (0.248) | 0.066 | (0.248) |
| Mother/father argue about religion often/very often (binary | 0.040 | (0.195) | 0.011 | (0.105) | 0.011 | (0.105) |
| Mother/father different race (binary 0/1) | 0.062 | (0.240) | 0.144 | (0.352) | 0.144 | (0.352) |
| Mother/father raised different religious preference (binary | 0.480 | (0.500) | 0.390 | (0.488) | 0.390 | (0.488) |
| Mother/father difference in college completion (binary 0/1) | 0.000 | (0.000) | 0.000 | (0.012) | 0.000 | (0.012) |

N

Notes: Sample restricted to children whose parents were married at the time of their birth, and for children who were at least 18 years old in 2012. Parental divorce is measured as divorce that occurred when children were 0-17 years old.

Factors refer to mothers unless otherwise specified. All factors are measured prior to the divorce interval, i.e. at the time of child's birth or earlier.

APPENDIX TABLE A2

DESCRIPTIVE STATISTICS OF MEDIATORS / OUTCOMES BY STRATA (NLSY)

| | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|---|------------------------------------|---------------------|---|---------------------|-----------------------------------|---------------------|
| | Parents not divorced | | Parents not divorced | | Parents not divorced | |
| | Parents divorced | Parents divorced | Parents divorced | Parents divorced | Parents divorced | Parents divorced |
| Mediators | | | | | | |
| Family Instability | | | | | | |
| Relationship transitions (continuous) | 1.135 (0.811) | 2.269 (1.478) | 1.484 (1.298) | 2.580 (1.598) | 1.903 (1.660) | 2.975 (1.697) |
| Family Economic Resources | | | | | | |
| Family income (continuous) | \$89,774 (81400) | \$48,883 (47745) | \$50,183 (35984) | \$34,814 (34258) | \$40,132 (30396) | \$33,564 (37950) |
| Children's Skills | | | | | | |
| Cognitive skills scale (continuous) | 0.601 (0.155) | 0.566 (0.141) | 0.535 (0.148) | 0.532 (0.140) | 0.472 (0.153) | 0.518 (0.140) |
| Psychosocial skills scale (continuous) | 0.608 (0.132) | 0.572 (0.137) | 0.567 (0.132) | 0.549 (0.132) | 0.541 (0.134) | 0.524 (0.132) |
| Outcomes | | | | | | |
| Children's Educational Attainment | | | | | | |
| High school completion (by age 18; binary 0/1) | 0.880 (0.325) | 0.817 (0.387) | 0.817 (0.387) | 0.765 (0.425) | 0.748 (0.435) | 0.696 (0.460) |
| College attendance (by age 19; binary 0/1) [^] | 0.686 (0.464) | 0.534 (0.499) | 0.522 (0.500) | 0.426 (0.495) | 0.459 (0.499) | 0.426 (0.495) |
| College completion (by age 23; binary 0/1) [^] | 0.381 (0.486) | 0.172 (0.378) | 0.208 (0.406) | 0.139 (0.346) | 0.149 (0.356) | 0.120 (0.326) |

N[^]

Notes: Numbers in parentheses are standard deviations. Sample restricted to children whose parents were married at the time of their birth, for children who were at least 18 years old in 2012, and for children with no missing data on educational attainment. Parental divorce is measured as divorce that occurred when children were 0-17 years old. Missing values are imputed for family status and children's cognitive and psychosocial skills. Relationship transitions counts how many changes occurred during childhood between the categories of married, separated, remarried, cohabitating, and widowed for both those married and divorced families.

[^] Sample is further restricted to age 19 and above for college attendance (N = 4,982), and to age 23 and above for college completion (N = 3,901).

APPENDIX TABLE B1

ADDITIONAL SENSITIVITY RESULTS FOR HETEROGENEOUS MEDIATION EFFECTS OF PARENTAL DIVORCE ON CHILDREN'S EDUCATIONAL ATTAINMENT

| | Sensitivity parameters | | Stratum 1: Divorce least likely | | Stratum 2: Divorce moderately likely | | Stratum 3: Divorce most likely | |
|--|------------------------|-------------|---------------------------------|------------------|--------------------------------------|------------------|--------------------------------|------------------|
| | γ_s | λ_s | Med. effects | CI | Med. effects | CI | Med. effects | CI |
| | | | | | | | | |
| <i>Mediators for Effects on High School Completion</i> | | | | | | | | |
| Family Instability | | | | | | | | |
| Relationship trans. | -40% | -5% | 0.021 | (0.006, 0.035) | -0.013 | (-0.030, 0.004) | -0.004 | (-0.020, 0.012) |
| | -20% | -5% | 0.011 | (-0.004, 0.025) | -0.023 | (-0.040, -0.006) | -0.014 | (-0.030, 0.002) |
| | -10% | -5% | 0.006 | (-0.009, 0.020) | -0.028 | (-0.045, -0.011) | -0.019 | (-0.035, -0.003) |
| | -5% | -5% | 0.003 | (-0.011, 0.018) | -0.031 | (-0.047, -0.014) | -0.021 | (-0.037, -0.006) |
| | 5% | -5% | -0.002 | (-0.016, 0.013) | -0.036 | (-0.052, -0.019) | -0.026 | (-0.042, -0.011) |
| | 10% | -5% | -0.004 | (-0.019, 0.010) | -0.038 | (-0.055, -0.021) | -0.029 | (-0.045, -0.013) |
| | 20% | -5% | -0.009 | (-0.024, 0.005) | -0.043 | (-0.060, -0.026) | -0.034 | (-0.050, -0.018) |
| | 40% | -5% | -0.019 | (-0.034, -0.005) | -0.053 | (-0.070, -0.036) | -0.044 | (-0.060, -0.028) |
| Children's Skills | | | | | | | | |
| Cognitive skills | -40% | -5% | 0.013 | (0.005, 0.022) | 0.025 | (0.017, 0.034) | 0.053 | (0.034, 0.072) |
| | -20% | -5% | 0.003 | (-0.005, 0.012) | 0.015 | (0.007, 0.024) | 0.043 | (0.024, 0.062) |
| | -10% | -5% | -0.002 | (-0.010, 0.007) | 0.010 | (0.002, 0.019) | 0.038 | (0.019, 0.057) |
| | -5% | -5% | -0.004 | (-0.013, 0.004) | 0.008 | (-0.001, 0.016) | 0.036 | (0.017, 0.054) |
| | 5% | -5% | -0.009 | (-0.018, -0.001) | 0.003 | (-0.006, 0.011) | 0.031 | (0.012, 0.049) |
| | 10% | -5% | -0.012 | (-0.020, -0.003) | 0.000 | (-0.008, 0.009) | 0.028 | (0.009, 0.047) |
| | 20% | -5% | -0.017 | (-0.025, -0.008) | -0.005 | (-0.013, 0.004) | 0.023 | (0.004, 0.042) |
| | 40% | -5% | -0.027 | (-0.035, -0.018) | -0.015 | (-0.023, -0.006) | 0.013 | (-0.006, 0.032) |
| Psychosocial skills | -40% | -5% | 0.011 | (0.006, 0.016) | 0.015 | (0.008, 0.023) | 0.010 | (-0.002, 0.022) |
| | -20% | -5% | 0.001 | (-0.004, 0.006) | 0.005 | (-0.002, 0.013) | 0.000 | (-0.012, 0.012) |
| | -10% | -5% | -0.004 | (-0.009, 0.001) | 0.000 | (-0.007, 0.008) | -0.005 | (-0.017, 0.007) |
| | -5% | -5% | -0.006 | (-0.011, -0.001) | -0.002 | (-0.009, 0.005) | -0.007 | (-0.019, 0.005) |
| | 5% | -5% | -0.011 | (-0.016, -0.006) | -0.007 | (-0.014, 0.000) | -0.012 | (-0.024, 0.000) |
| | 10% | -5% | -0.014 | (-0.019, -0.009) | -0.010 | (-0.017, -0.002) | -0.015 | (-0.027, -0.003) |
| | 20% | -5% | -0.019 | (-0.024, -0.014) | -0.015 | (-0.022, -0.007) | -0.020 | (-0.032, -0.008) |
| | 40% | -5% | -0.029 | (-0.034, -0.024) | -0.025 | (-0.032, -0.017) | -0.030 | (-0.042, -0.018) |

Mediators for Effects on College Attendance

Family Instability

| | | | | | | | | |
|---------------------|------|-----|--------|-----------------|--------|------------------|--------|------------------|
| Relationship trans. | -40% | -5% | 0.010 | (-0.009, 0.030) | -0.002 | (-0.023, 0.018) | 0.004 | (-0.013, 0.021) |
| | -20% | -5% | 0.000 | (-0.019, 0.020) | -0.012 | (-0.033, 0.008) | -0.006 | (-0.023, 0.011) |
| | -10% | -5% | -0.005 | (-0.024, 0.015) | -0.017 | (-0.038, 0.003) | -0.011 | (-0.028, 0.006) |
| | -5% | -5% | -0.007 | (-0.027, 0.012) | -0.020 | (-0.040, 0.001) | -0.014 | (-0.031, 0.004) |
| | 5% | -5% | -0.012 | (-0.032, 0.007) | -0.025 | (-0.045, -0.004) | -0.019 | (-0.036, -0.001) |
| | 10% | -5% | -0.015 | (-0.034, 0.005) | -0.027 | (-0.048, -0.007) | -0.021 | (-0.038, -0.004) |
| | 20% | -5% | -0.020 | (-0.039, 0.000) | -0.032 | (-0.053, -0.012) | -0.026 | (-0.043, -0.009) |

| | | | | | | | | |
|--------------------------|------|-----|--------|------------------|--------|------------------|--------|------------------|
| | 40% | -5% | -0.030 | (-0.049, -0.010) | -0.042 | (-0.063, -0.022) | -0.036 | (-0.053, -0.019) |
| Children's Skills | | | | | | | | |
| Cognitive skills | -40% | -5% | 0.011 | (-0.006, 0.027) | 0.031 | (0.016, 0.047) | 0.065 | (0.041, 0.088) |
| | -20% | -5% | 0.001 | (-0.016, 0.017) | 0.021 | (0.006, 0.037) | 0.055 | (0.031, 0.078) |
| | -10% | -5% | -0.004 | (-0.021, 0.012) | 0.016 | (0.001, 0.032) | 0.050 | (0.026, 0.073) |
| | -5% | -5% | -0.007 | (-0.023, 0.010) | 0.014 | (-0.002, 0.029) | 0.047 | (0.024, 0.070) |
| | 5% | -5% | -0.012 | (-0.028, 0.005) | 0.009 | (-0.007, 0.024) | 0.042 | (0.019, 0.065) |
| | 10% | -5% | -0.014 | (-0.031, 0.002) | 0.006 | (-0.009, 0.022) | 0.040 | (0.016, 0.063) |
| | 20% | -5% | -0.019 | (-0.036, -0.003) | 0.001 | (-0.014, 0.017) | 0.035 | (0.011, 0.058) |
| | 40% | -5% | -0.029 | (-0.046, -0.013) | -0.009 | (-0.024, 0.007) | 0.025 | (0.001, 0.048) |
| Psychosocial skills | -40% | -5% | 0.007 | (-0.001, 0.015) | 0.016 | (0.008, 0.024) | 0.013 | (0.004, 0.022) |
| | -20% | -5% | -0.003 | (-0.011, 0.005) | 0.006 | (-0.002, 0.014) | 0.003 | (-0.006, 0.012) |
| | -10% | -5% | -0.008 | (-0.016, 0.000) | 0.001 | (-0.007, 0.009) | -0.002 | (-0.011, 0.007) |
| | -5% | -5% | -0.011 | (-0.019, -0.003) | -0.002 | (-0.010, 0.006) | -0.004 | (-0.013, 0.005) |
| | 5% | -5% | -0.016 | (-0.024, -0.008) | -0.007 | (-0.015, 0.001) | -0.009 | (-0.018, 0.000) |
| | 10% | -5% | -0.018 | (-0.026, -0.010) | -0.009 | (-0.017, -0.001) | -0.012 | (-0.021, -0.003) |
| | 20% | -5% | -0.023 | (-0.031, -0.015) | -0.014 | (-0.022, -0.006) | -0.017 | (-0.026, -0.008) |
| | 40% | -5% | -0.033 | (-0.041, -0.025) | -0.024 | (-0.032, -0.016) | -0.027 | (-0.036, -0.018) |

*Mediators for Effects on
College Completion*

Family Instability

| | | | | | | | | |
|---------------------|------|-----|--------|------------------|--------|------------------|--------|------------------|
| Relationship trans. | -40% | -5% | -0.001 | (-0.025, 0.022) | 0.004 | (-0.013, 0.022) | 0.004 | (-0.008, 0.016) |
| | -20% | -5% | -0.011 | (-0.035, 0.012) | -0.006 | (-0.023, 0.012) | -0.006 | (-0.018, 0.006) |
| | -10% | -5% | -0.016 | (-0.040, 0.007) | -0.011 | (-0.028, 0.007) | -0.011 | (-0.023, 0.001) |
| | -5% | -5% | -0.019 | (-0.042, 0.005) | -0.013 | (-0.030, 0.004) | -0.014 | (-0.026, -0.001) |
| | 5% | -5% | -0.024 | (-0.047, 0.000) | -0.018 | (-0.035, -0.001) | -0.019 | (-0.031, -0.006) |
| | 10% | -5% | -0.026 | (-0.050, -0.003) | -0.021 | (-0.038, -0.003) | -0.021 | (-0.033, -0.009) |
| | 20% | -5% | -0.031 | (-0.055, -0.008) | -0.026 | (-0.043, -0.008) | -0.026 | (-0.038, -0.014) |
| | 40% | -5% | -0.041 | (-0.065, -0.018) | -0.036 | (-0.053, -0.018) | -0.036 | (-0.048, -0.024) |

Children's Skills

| | | | | | | | | |
|---------------------|------|-----|--------|------------------|--------|------------------|--------|------------------|
| Cognitive skills | -40% | -5% | 0.019 | (0.001, 0.038) | 0.029 | (0.017, 0.041) | 0.046 | (0.032, 0.061) |
| | -20% | -5% | 0.009 | (-0.009, 0.028) | 0.019 | (0.007, 0.031) | 0.036 | (0.022, 0.051) |
| | -10% | -5% | 0.004 | (-0.014, 0.023) | 0.014 | (0.002, 0.026) | 0.031 | (0.017, 0.046) |
| | -5% | -5% | 0.002 | (-0.017, 0.021) | 0.012 | (0.000, 0.024) | 0.029 | (0.014, 0.044) |
| | 5% | -5% | -0.003 | (-0.022, 0.016) | 0.007 | (-0.005, 0.019) | 0.024 | (0.009, 0.039) |
| | 10% | -5% | -0.006 | (-0.024, 0.013) | 0.004 | (-0.008, 0.016) | 0.021 | (0.007, 0.036) |
| | 20% | -5% | -0.011 | (-0.029, 0.008) | -0.001 | (-0.013, 0.011) | 0.016 | (0.002, 0.031) |
| | 40% | -5% | -0.021 | (-0.039, -0.002) | -0.011 | (-0.023, 0.001) | 0.006 | (-0.008, 0.021) |
| Psychosocial skills | -40% | -5% | 0.007 | (-0.002, 0.017) | 0.017 | (0.010, 0.025) | 0.016 | (0.011, 0.021) |
| | -20% | -5% | -0.003 | (-0.012, 0.007) | 0.007 | (0.000, 0.015) | 0.006 | (0.001, 0.011) |
| | -10% | -5% | -0.008 | (-0.017, 0.002) | 0.002 | (-0.005, 0.010) | 0.001 | (-0.004, 0.006) |
| | -5% | -5% | -0.010 | (-0.019, -0.001) | 0.000 | (-0.008, 0.007) | -0.001 | (-0.006, 0.004) |
| | 5% | -5% | -0.015 | (-0.024, -0.006) | -0.005 | (-0.013, 0.002) | -0.006 | (-0.011, -0.001) |
| | 10% | -5% | -0.018 | (-0.027, -0.008) | -0.008 | (-0.015, 0.000) | -0.009 | (-0.014, -0.004) |
| | 20% | -5% | -0.023 | (-0.032, -0.013) | -0.013 | (-0.020, -0.005) | -0.014 | (-0.019, -0.009) |

40% -5% -0.033 (-0.042, -0.023) -0.023 (-0.030, -0.015) -0.024 (-0.029, -0.019)

Notes: γ_s refers to the mean difference in children's education associated with a unit difference in the binary unobserved confounder conditional on divorce status, propensity scores, and mediator values. λ_s refers to the prevalence difference of the binary unobserved confounder between children of divorced and non-divorced parents conditional on divorce status, propensity scores, and mediator values. The bias factor is equal to the negation of the product of the two parameters, and we subtract this bias factor from the mediation effect and the confidence interval.