# “The Impact of Time-Varying Exposure to a Two-Parent Household on Black-White Skill Gaps at School Entry”

**By Jill Bowdon, University of Pennsylvania, Graduate School of Education**

**Abstract:**

While there is extensive documentation of the effects of father absence on cognitive ability, aggression, and depression or anxiety, we do not know how father presence affects child attentiveness or academic engagement. Since attentiveness and engagement are associated with higher academic achievement throughout elementary school, it is worth understanding what drives disparities in these types of skills by kindergarten entry. In this study, I employed marginal structural models with inverse probability of treatment weighting to estimate the effects of time-varying exposure to an intact family structure on teacher ratings of child behavior in kindergarten. This method of estimating time-varying treatment effects is an improvement over conventional regression models because it does not “control away” the effect of prior exposure to father presence operating indirectly through the time-varying characteristics of the child or family. I found that father presence in both infancy and the year prior to kindergarten has a positive effect on approaches to learning that is comparable in size to its effect on reading ability and externalizing behavior. Given the fact that only one-third of black children had a biological father present in the household in both infancy and at age four, these estimated effects of father presence can help explain black-white skill gaps at kindergarten entry.

**Keywords:** Black-White Gaps, Social-Behavioral Skills, Family Structure, Kindergarten Readiness

## INTRODUCTION

Kindergarten teachers report that black students are more likely to have problems paying attention, staying focused, and persisting on tasks than whites (Rimm-Kaufman, Pianta, and Cox, 2000; Downey and Pribesh, 2004). In nationally representative data on a kindergarten cohort, blacks scored about three-tenths of a standard deviation below whites on measures of “Approaches to Learning”, a scale that measures attentiveness, task persistence, and engagement (Duncan and Magnuson, 2011). These early black-white skill gaps are concerning because children’s social and behavioral skills at school entry have long-lasting impacts.

In fact, most kindergarten teachers view behavioral skills—such as being able to pay attention, practice self-control, and inhibit impulses-- as more important indicators of kindergarten readiness than knowing letters and numbers (Heaviside and Farris, 1993; Lin, Lawrence, and Gorrell, 2003). When children come to school without the ability to sit still or follow directions, they have trouble absorbing the information necessary to learn because they spend less time on task. Furthermore, children with low self-regulation have more troubled relationships with their teachers and peers because their behaviors elicit more correction than affirmation (Ladd, Birch, and Buhs, 1999). High levels of conflict with teachers can be especially detrimental for kindergartners because they interfere with the transition to schooling. Perhaps for this reason, children with low attention spans learn less in kindergarten and in later grades (Duncan et al., 2007; Claessens, Duncan, and Engle, 2009).

But what causes black-white gaps in noncognitive skills in kindergarten? Prior to school entry, racial disparities in home and neighborhood environments play the most important role in shaping skill development. Past research, which has focused primarily on explaining black-white gaps in reading and math, has shown that black-white gaps in socioeconomic status can account for the majority of differences in kindergarten readiness (Fryer and Levitt, 2004).[1] However, understanding that racial differences in socioeconomic status explain kindergarten readiness gaps does not clarify the mechanisms at work in the intergenerational transmission of disadvantage (Duncan and Magnuson, 2005). This study examined how large the joint causal impact of living with

[1] Using a limited set of covariates, including measures of family socioeconomic status, child age, child birth weight, WIC participation, mother's age at first birth, and number of children's books in the household, Fryer and Levitt (2004) reduced black-white gaps on cognitive measures to the point that they were statistically insignificant.

two biological parents (either cohabiting or married) in infancy and at age four was and whether black-white disparities in father presence could explain racial gaps in kindergarten readiness. In doing so, it makes two contributions.

First, while there is extensive documentation of the effects of father absence on cognitive ability, aggression, and depression or anxiety (Osborne and McLanahan, 2007; Cavanagh and Huston, 2006; Carlson and Corcoran, 2001; McLanahan and Sandefur, 1994), there is little evidence of the effects of father presence on child attentiveness or academic engagement in the classroom. Second, while many prior studies on the effects of family structure or family instability do not address time-dependent confounders, this study used marginal structural models to estimate the effects of time-varying exposure to father presence (e.g. Sharkey and Elwert, 2011). These models offer a significant improvement over conventional regression models because they allow for the estimation of a joint effect but do not "control away effects" of prior treatments.

## BACKGROUND

### *Black-White Differences in Father Presence during Early Childhood*

In the last few decades, there has been a dramatic rise in non-marital birth in the population at large (Ventura and Bachrach, 2000). Now, nearly four in ten babies are born outside of marriage (Ventura, 2009). Though the majority of unmarried parents are romantically involved at the time of the birth, only 44% of these relationships remain intact by the time the child is three year old (Carlson, McLanahan, and Brooks-Gunn, 2008). Together with more traditional forms of family disruption—divorce and remarriage—these trends mean that children of all races are increasingly growing up in unstable family environments. In fact, by one estimate, one-third of all children experience a transition from one family structure type to another by school entry (Cavanagh and Huston, 2006).

Even in this context of rising instability of family structure, however, there remain notable black-white differences. In 1965, Moynihan reported that about one in five black children were born outside of marriage. In 2000, 69% of black children were born to an unmarried mother, compared to about one-quarter of white children (McLanahan, 2009). Racial differences in nonmarital births arise because of black-white differences in the rate and timing of marriage. Black women marry less, even after the birth of a child, because of both structural and cultural conditions that make their lived experiences different from those of white women. Black women are less likely to marry across race (Qian and LIchter, 2007), and because of a shortage of non-incarcerated black men who have demonstrated earnings potential and similar educational attainment, they are also less likely to find suitable marriage partners within their race (Harknett and McLanahan, 2004; Wilson 1987). Given the scarcity of marriageable black men, though black women have higher pro-marriage attitudes than whites, they are also less likely than whites to view marriage as a likely occurrence or a prerequisite for conceiving a child (McLanahan, 2009). Sacrificing their independence and household authority for a marriage that does not promise emotional or financial stability seems irrational to many black women, especially because premarital birth is so normative in their social networks (Edin and Kefalas, 2005).

One implication of this higher incidence of non-marital birth among black women is that black children are less likely to have biological fathers in the household during early childhood. In general, unmarried parents' relationships are less stable than marital bonds (Osborne, Manning, and Smock, 2007; Carlson, McLanahan, and Brooks-Gunn, 2008), and black mothers are especially unlikely to marry the birth father by the time their child is a toddler (Harknett and McLanahan, 2004). As a result, black children are more likely to miss out on many of the protective factors associated with a two-parent household during the first, formative years of life.

Of course, family instability may not have as large of an effect for black children because, regardless of family structure, black children face more economic hardship and family stress than white children (Amato and Keith, 1991; Shaw, Winslow, and Flanagan, 1999). In addition, because black children are more likely to have a larger network of kin to provide support, growing up in a single parent household may create less emotional and economic strain for blacks (McLoyd et al. 2000). Indeed, many researchers have found a weaker association between family structure and child outcomes (e.g. Fomby and Cherlin, 2007).

*Theorized Benefits of a Two-Parent Household For Child Development*

Two-parent households offer many advantages to infants and toddlers. In the simplest of terms, having a two-parent household matters because it decreases the chances of exposure to factors known to hinder child development: childhood poverty, maternal stress or depression, and harsh or neglectful parenting. When a mother and father break-up, the household's resources diminish. In the case of divorce, fathers may continue to pay child support, but, still, divorced mothers are more likely to be poor than currently married mothers (Cancian and Reed, 2009). If the parents were not married at the child's birth, there is less likely to be a formal child support arrangement in place, and children may truly suffer from poverty.

Exposure to poverty in early childhood is associated with lower skills at school entry and deficits may be even worse for children who are continuously poor (Duncan, Brooks-Gunn, Klebanov, 1994; Korenman, Miller, and Sjaastad, 1995). Poor parents lack the ability to provide nutritious foods and high-quality neighborhoods that boost brain development, let alone purchase toys, books, or enrichment opportunities for their child. Overall, economic deprivation accounts for half of the achievement differences between children who grow up in a single parent as opposed to a two-parent household (McLanahan and Sandefur 1994).

Beyond affecting household finances, father residence also impacts the quality and quantity of interactions both parents have with the child. Especially in the first few years of life, a father's presence in the family may impact household decisions about whether and when both parents should work outside of the home and the quality of non-parental care to utilize if they do. A single mother has less time for childcare, both because she is more likely to work longer hours outside of the home (Bradbury and Katz, 2002), and because she is now solely responsible for all household duties.

When she does have time to interact with her child, the single mother may be too stressed to do so in a consistent and responsive manner (Cooper, McLanahan, Meadows, and Brooks-Gunn, 2009; McLoyd, 1990). After a relationship ends, many mothers have financial worries and continued conflict with the child's father. Furthermore, just when newly single mothers could use the help of friends or family, their social support systems are disrupted by residential moves and loss of ties (McLanahan and Sandefur, 1994; Coleman, 1988). Family instability is associated not only with increased maternal stress but also harsher parenting practices, like yelling at or threatening the child or using corporal punishment (Beck, Cooper, McLanahan, and Brooks-Gunn, 2010). Since maternal nurturance and stimulation are needed to promote positive cognitive and behavioral skills in childhood, children suffer when maternal stress increases (Knudsen et al., 2006).

A father's involvement with the child often decreases even more than the mother's after a break-up. The quality and quantity of father involvement may not have differed as much by father's residence in years past. But now, as father participation in childrearing has expanded to include greater responsibility for both nurturing and teaching their children (Cabrera et al. 2000; Marsiglio et al, 2000; Sayer, Bianchi, and Robinson, 2004), a father's presence in the household can make a difference in the amount of one-on-one attention a child receives.

A resident father is more likely than a nonresident father to have more frequent contact with a child, though frequency of contact alone is not meaningful for child development (Amato and Gilbreth 1999). Developmental psychologists have long recognized that maternal sensitivity and responsiveness matter for skill development of young children (e.g. Ainsworth 1969). More recent research has confirmed that these types of parenting skills also result in better behavioral and cognitive outcomes for toddlers and preschoolers when executed by fathers (e.g. Cabrera, Shannon, and Tamis-LeMonda, 2007; Tamis-LeMonda, Shannon, Cabrera, and Lamb, 2004), and that children with two supportive parents have better outcomes than children with one or no involved parent (Martin, Ryan, and Brooks-Gunn 2007).

We have long known that children who grow up in a single-parent household have lower academic achievement and more problematic behavior by the end of elementary school and during adolescence (e.g. McLanahan and Sandefur, 1994). Emergent research indicates that family instability and father absence are also negatively associated with child development in toddlerhood and can explain individual variation in kindergarten readiness (Osborne and McLanahan, 2007; Cooper, Osborne, Beck, and McLanahan, 2011). Because a single-parent family is a marker for disadvantages ranging from low education to poverty to prior incarceration (McLanahan, 2009), and because these disadvantages are also related to child skill development independently of family structure, the bivariate association between father presence and child skills may be spurious in part. Still, even in studies that use sophisticated methods to control for selection bias, the negative association between father absence and child skill levels persists (Cherlin, 1999; Sigle-Ruston and McLanahan, 2004). Children who spend any time in single-parent households have a higher probability of poor outcomes than children who live with two parents continuously, regardless of their race or their parents' education level.

*Limitations of Prior Research*

Traditionally, research on the association between parental relationships and child outcomes has used measures of family structure at one point in time to predict either concurrent or future child skill levels. However, these one-time point, static measures of family structure seem unsatisfying in the current era of complex and fluid family relationships. Increasingly, children are born to unmarried parents and experience multiple types of family structure as mothers transition into and out of cohabiting relationships. Static measures of family structure disguise how many transitions a child has undergone, and some theorize that it is these transitions, rather than the presence of one or two parents, that worsen parent-child relationships and harm child development (e.g. Osborne and McLanahan, 2007; Wu and Martinson 1993; Fomby and Cherlin, 2007).

In reaction to these demographic trends and the advancements in our understanding of the effects of stress and instability on children (Shonkoff and Phillips, 2000), more scholars are incorporating longitudinal measures of family structure into their studies of child development. Two approaches are common. In the first, researchers construct a set of dummy indicators to capture the trajectory of parental relationships experienced. For example, in Carlson and Corcoran (2001), the authors used a set of mutually exclusive categories: a two parent household for all years, a single parent household for all years, a single to two parent household, a two to single parent household, and multiple transitions. This type of measure captures both current family status and family structure history but does not include duration of exposure to any family structure. A second common approach in constructing a longitudinal measure of family structure is to use a count of the number of transitions a child has experienced from birth to a specified age and include this count measure in a regression model that also includes additional measures of current family structure, length of exposure to certain types of family structure, and whether there was a transition in the past two years (Osborne and McLanahan, 2007; Cavanagh and Huston, 2008).

These longitudinal measures present a problem when trying to establish a causal effect of family structure on child outcomes because they do not account for time-dependent confounders. Some scholars attempt to address selection bias by controlling for maternal characteristics pre-birth (Fomby and Cherlin, 2007; Osborne and

McLanahan, 2007). This approach relies on the strong assumption that pre-birth measures can account for all factors that might potentially confound the association between a sequence of family statuses and child skill levels.  Other scholars lessen the threat of selection bias by including both mother's characteristics prior to the child's birth and additional controls for maternal characteristics during the baby's first years, such as maternal depression or use of child care (e.g. Cavanagh and Huston, 2008). While this approach incorporates confounding factors that occur after the child's birth, it will also "control away" some of the effects of family structure on child outcomes because it holds constant factors that may result from family instability.

In contrast, the method that I used in this paper, marginal structural models with inverse probability of treatment weighting, allowed me to incorporate family structure at multiple time points during early childhood and still provided a credible way of establishing sequential unconfoundedness. Furthermore, it also allowed me to estimate both the separate and joint effects of family structure at different developmental stages: infancy and age four.

**DATA**

I used data from the Early Childhood Longitudinal Study—Birth Cohort (ECLS-B).  This birth cohort study followed a nationally representative sample of children born in 2001 in the United States from birth through kindergarten entry. Birth certificate records were collected for all sampled children and then additional data were collected when children were, on average, nine months old (wave 1), 2 years old (wave 2), 4 years old (wave 3), and entering kindergarten (waves 4 and 5).

Originally, researchers sampled 14,000 births based on a sampling frame produced by the National Center for Health Statistics of births occurring in 2001. Of the 14,000 families contacted, 10,688 took part in the first wave of data collection, and 10,221 had child assessment data for the first data wave. In each subsequent wave, there was sample attrition. The fourth wave in 2006 followed 7,705 children who completed a parent interview in the previous wave, whether or not they had entered kindergarten at that time. During the fifth data wave in 2007, researchers collected kindergarten measures from 1,759 children who had not started kindergarten in 2006. Because there are two kindergarten waves of data collection, I used data from both waves together to analyze the experiences of all ECLS-B children at kindergarten entry.

I used three components of these data to conduct analyses presented in this paper: birth certificate records, teacher reports, and parent reports. Birth certificate data include the child's date of birth, as well as information on the mother's marital history, age, and education at the child's birth. I also relied on data from Teacher Self-Administered Questionnaires (TSAQ), which were mailed to kindergarten teachers of children enrolled in formal school settings. These questionnaires yielded data on the child's interactions with other children (including their attentiveness and engagement) and demographic and instructional characteristics of the kindergarten program. Finally, parent data were collected in each wave using computer-assisted personal interviewing and self-administered questionnaires. The study design targeted the child's mother as the respondent, and the mother responded in 99% of all cases at wave 1, 98% cases at wave 2, and 97% of all cases at wave 3.

To construct the analytic sample, I started with the children who had complete teacher ratings of their behavior and reading test scores in kindergarten (N=4862). I further restricted the sample to children living with their biological mother in waves 1-3, which resulted in dropping an additional 234 cases. I multiply imputed missing data for other covariates. The final analytic sample consisted of 4,628 children.

*Dependent Variable*

I constructed a factor to measure "Approaches to Learning" comprised of seven items from the Teacher Self-Administered Questionnaires in the child's first kindergarten year. Teachers rated how often the child engaged in certain behaviors on a scale of 1 (never) to 5 (very often). After a factor analysis, I combined measures for the following behaviors: shows eagerness to learn, pays attention well, works/plays independently, keeps working until finished, is overly active, has difficulty concentrating, and is restless/fidgety. The last three items were reverse-coded because they represent negative behaviors. This scale had a Cronbach's Alpha of 0.91. I standardized this measure on the population to have a mean of zero and a standard deviation of one.

*Time-varying Treatment Status*

I measured family structure using binary indicators of whether there were two biological parents in the home at the first wave of data collection, when the average child was nine months old, and at the third wave of data collection, when the average child was four years old. This measure did not depend on the martial status of the biological parents. Because my analytic sample excluded children not living with a biological mother, the reference category of these binary indicators included children living with a single biological mother or a biological mother and her partner but not children living with related or unrelated guardians or adoptive parents. Of the children not in a two-parent household at nine months, 96% lived with a single mother and 4% lived with a single mother and her partner. By age 4, of the children without a two-parent household, 79% lived with a single mother and 21% lived with the mother and her partner. The proportions of children living with a single mother versus the mother and a social father differed for whites and blacks.[2]

Though I estimated the effect of a time-varying exposure to a two-parent family on child outcomes, using just two time points could understate the degree of family instability experienced by children. This is a limitation of my approach. However, in these data, of the children who experienced a two-parent family at both nine months and 4 years, 96% also had both biological parents in the home at birth, 98% had both biological parents in the home at age 2, and 92% had both biological parents in the home at birth, nine months, age two and age four. These proportions are similar for whites and blacks.

*Predictors of The Presence of Two Biological Parents in the Household at Wave 1*

Models accounted for observed child and parental characteristics that could confound the relationship between the time-varying treatment and the outcome. I measured *child race* using a categorical indicator based on parental reports (white, black, Hispanic, Asian, other race). To measure the *parental relationship status at birth*, I included a categorical measure of whether the biological parents were married, cohabitating, or neither at the time of the child's birth. I constructed this measure using an indicator of maternal marital status from the birth records and maternal reports of the

---

[2] Of the whites not in a two-parent household at 9 months, 89% lived with a single biological mother. By age 4, this percentage had dropped to 67%. Of blacks not in two-parent households at 9 months, 98% lived with a single mother; at age 4 90% still lived with a single mother.

dates she started, and in some cases ended, her relationship with the biological father given during the wave one data collection. I also used a categorical measure of *maternal age* at the child's birth (teen, twenties, thirties or higher) and *maternal education* level at the birth (less than high school, high school, some college, or college degree and higher). Both of these measures are from the birth certificate data. In addition, I included a categorical measure of the number of *hours of work* the mother worked in the 12 months preceding the child's birth (none, 20 hours or less, more than 20 hours), as well as a categorical measure of the *number of older siblings at birth* (none, one, two, three or more). Finally, I included an indicator of whether the *mother wanted to be pregnant* with this child, which was reported by the mother in a self-administered questionnaire.

*Predictors of The Presence of Two Biological Parents in the Household at Wave 3*

To predict the treatment status at wave 3, I used all of the above baseline covariates but included measures of *number of siblings* to reflect levels reported during wave 2 instead of at birth. In addition, I included other covariates that may have resulted from family structure in wave 1 but confound the relationship between family structure at wave 3 and the outcome: a binary indicator of whether the family was *above the poverty threshold*, and a binary indicator of whether the child received *parental childcare only*. Because family structure may also have affected parenting behaviors, models also accounted for categorical measures of *times per week the family read to the child at age 2* (none, once or twice, three-six times, and every day) and how many *hours of TV the child watched per weekday at age 2* (none, one, two, three or more). Finally, I included categorical measures of number of siblings of *maternal depression*. Mothers reported their symptoms of depression using a self-administered questionnaire in wave 1, and I used these reports to construct a measure of whether the mother was “not depressed”, “mildly depressed”, “moderately depressed”, or “severely depressed” that corresponds to the cut-points for the Center for Epidemiologic Studies-Depression Scale scores.

**METHODS**

I estimated the effects of time-varying exposure to an intact family structure in early childhood on children’s kindergarten approaches to learning. Specifically, I estimated the joint effect of having two biological parents in the home when the child was both nine months and four years versus not having two biological parents in the home at either age. To identify these joint effects, I utilized marginal structural models with inverse probability of treatment weighting, a method pioneered in the epidemiological literature (Robins, 1998; Robins 1999; Robins et al., 2000) but also utilized in sociology to estimate the effects of time-varying exposure to poor neighborhoods on cognitive ability (Sharkey and Elwert, 2011).

This methodological approach has advantages over conventional regression models for estimating the joint effect of a time-varying treatment in the presence of time-varying confounders. Conventional regression approaches can be used to adjust for observed factors related to both the treatment and the outcome in order to predict the effect of a treatment on the outcome. However, in the case of time-varying treatments and time-varying confounders, taking this conventional regression approach to adjusting for confounders causes two problems: (1) While it is necessary to control for all covariates that influence selection into family structure when the child is four years old, doing so may “control away” part of the estimated effect of family structure at age nine months, (2) Even if it were the case that family structure when the child is nine months has no effects on kindergarten approaches to learning, controlling for factors that influence selection into family structure at age four could induce a non-causal association between family structure when the child is nine months and the outcome if there are unobserved variables that are associated with these covariates and the outcome measure (Pearl 1995, 2000; Greenland, 2003; Hernan, Hernandez-Diaz, and Robins, 2004). Therefore, in cases where there are time-varying confounders affected by past treatments, conventional regression models are not well equipped to estimate the effects of time-varying treatment.

By contrast, inverse probability of treatment weighting allowed me to construct a weighted pseudo-population in which the treatment at each wave was no longer confounded by the measured covariates (Hernan and Robins, 2006). In practice, this technique involved multiple steps. First, I constructed this pseudo-population by assigning each child in my sample a weight proportionate to his/her probability of receiving his/her own exposure history. To do so, I estimated a logistic regression model of kindergartners' approaches to learning in a two biological parent family separately for each time point, W, as a function of baseline and time-varying covariates:

$$\frac{P(T_W)}{1-P(T_W)} = \exp[\ \alpha_W + C_W\beta_W], \text{for W} \in \{N, F\} \qquad (1)$$

where T represented the presence of two biological parents in the household, $\alpha_W$ is the intercept, $C_F$ included both $C_N$ and exposure to a two-biological-parent household at nine months. I used these models to predict the actual treatment status of each child—or the probability of being in the family structure actually experienced—separately at nine months and four years of age. I multiplied these two probabilities to get the probability of exposure history:

$$W=P\ (T_N C_N)\ x\ P\ (T_F | T_N,\ C_N,\ C_F) \quad (2)$$

Second, I weighted each case by the inverse of W[3]. I used the weighted pseudo-population, I regressed approaches to learning in kindergarten on exposure to family structure at two time points in early childhood in a model that did not include measured confounders as covariates:

$$E_{weighted}\ [A|T_N,\ T_F]=\alpha + T_N\beta_1 + T_F\beta_2 \quad (3)$$

where A is approaches to learning in kindergarten. Because the final model did not include measured confounders, I avoided the two pitfalls of conventional regression approaches listed above.

In order to produce unbiased estimates using marginal structural models with IPTW weighting, I had to make three of the same strong assumptions inherent in conventional regression approaches. Perhaps, most obviously, I had to assume exchangeability, or the absence of unmeasured confounders. In practice, it is difficult to confirm that this assumption is met. And while it is tempting to include every measure in the data that could be a confounder to minimize this threat, doing so could actually introduce bias due to collider stratification (Greenland, 2003; Hernan, Hernandez-Diaz, and Robins, 2004). Furthermore, if the covariate is not a confounder, its presence in the model estimating the weights could reduce the statistical efficiency of the effect estimate

[3] These weights produce large standard errors. Therefore, it is common practice to use stabilized weights (Robins et al. 2000), constructed from the following formula:
SW=[P($T_N$) x P($T_F$| $T_N$)] x [P($T_N$|$C_N$) x P($T_F$| $T_N$, $C_N$, $C_F$)$^{-1}$.

(Robins and Greenland, 1986). Therefore, I relied on a set of covariates that were both theoretically and empirically related to both family structure and kindergarten readiness.

In addition to exchangeability, I also made two additional assumptions. I assumed that there were exposed and unexposed individuals at every level of the confounders. This assumption, positivity, was necessary in order to estimate effects in all subsets of the population. Finally, I assumed that models estimating the weights were specified correctly. I tested these models under different specifications and used goodness-of-fit tests to guide my model selection.

## RESULTS

### *Fraction of Children Exposed to Two Biological Parents By Age*

Table 1 shows the fractions of children with two biological parents who were living in the same household. In the population, 81% had two parents at home in infancy, 75% had two parents at home when they were four years old, and 71% had two parents at home at both time points. These proportions were even higher for whites: 91% had two parent homes at nine months, 83% had two parent homes at four years, and 81% had two parent homes at both time points. Relative to both the population and to whites, the proportions were much lower for blacks. Only 42% of black children had two parents in the home in infancy, and only 40% had two parents at home in the year prior to kindergarten. Only about one-third of black children had two biological parents in the household at both time points.

### *Description of the Sample*

Table 2 contains the sample means for the observed confounders in the association between the presence of two biological parents in the household and kindergarten readiness. Children who had two biological parents in the home at both nine months and four years (see column 5) were more advantaged than those who had an absent father at either wave (columns 3 and 4) or in both waves (column 2). At birth, their parents were more likely to be married, and their mothers were older and more educated. Fewer children with two present biological parents at nine months and four years of age had severely depressed mothers and were below the poverty line and more were read to daily. Child race was also associated with family structure in early childhood: black children were more likely to have absent fathers in both infancy and at age four than white children. These descriptive statistics confirm what prior literature has already established: selection into family structure is not random, but, rather, associated with other types of structural disadvantage. For this reason, estimates of the association between family structure at nine months and four years and kindergarten readiness cannot be considered causal without establishing sequential unconfoundness.

*Constructing Weights*

I applied the stabilized weights, shown in Table 3, in models predicting the effect of family structure on kindergarten readiness. These stabilized weights accounted for selection into family structure at nine months and four years. To improve efficiency of the weights, I truncated them at the 99$^{th}$ and 1$^{st}$ percentiles (Cole and Hernán, 2008). I estimated these weights using the predictors listed in Table 2 and an interaction term between child race and maternal education. Theoretically, we should expect weights to have a mean of one. The stabilized weights estimated on the basis of these observed traits had means of 0.93 for all children, 0.95 for blacks, and 0.98 for whites. Though all three distributions of weights were skewed to the right, standard deviations were small (0.96 for all children, 1.21 for blacks, and 0.78 for whites).

*Estimates from Conventional Regression Versus Marginal Structural Models*

Table 4 displays bivariate associations between the presence of two biological parents in the household and Approaches to Learning (column 1), the association net of predictors in Table 2 estimated in a conventional regression model (column 2), and the casual effect procured with the use of marginal structural models (3). Having a mother and father in the home is positively associated with Approaches to Learning. The bivariate associations between a two-parent home and the outcome at both nine months and four years were four-tenths of a standard deviation ($\beta$ =0.43, p-value<0.000 and $\beta$=0.43 p-value<0.000). After accounting for confounding factors using a conventional regression model, the estimated association between two biological parents in the household at nine months and Approaches to Learning was not statistically different from zero ($\beta$ <0.00, p-value>0.10) and the association between two biological parents in the household at four years and Approaches to Learning was reduced by more than half ($\beta$=0.16 p-value<0.000).

Because the conventional regression model accounted for all of the potential confounders listed in Table 2, including those that may have resulted from family structure when the child is nine months old, the estimated association between the nine months family structure and Approaches to Learning from the conventional regression model “controlled away” the effect of family structure on Approaches to Learning. This estimated association is downwardly biased and should not be interpreted causally.

By contrast, estimated effects from the marginal structural models with inverse probability of the treatment weighting identified a causal impact, conditional on the assumption that the model accounted for all potential confounders. Using these models, the direct effect of having two-parent household at nine months, holding family structure at four years constant, was two-tenths of a standard deviation, which is approximately half of the bivariate association ($\beta$=0.19, p-value<0.06). The causal effect of a two-parent household at four years, fixing family structure at nine months, was of similar magnitude ($\beta$=0.16, p-value<0.000). And, finally, the joint effect of living in a two biological parent household was 0.34 of a standard deviation ($\beta_1 + \beta_2$=0.34, p-value<0.000).

To contextualize these results, I also estimated the associations and causal effects of family structure on reading standardized test in the fall of kindergarten and teacher ratings of externalizing behavior. Higher scores on externalizing behavior represent more problem behavior, such as acting out, fighting and arguing. For both of these outcomes, the bivariate associations with family structure at both time points were about four-tenths of a standard deviation, though the associations were positive for reading ability and negative for problem behavior, as we would expect (reading: $\beta$=0.47, p-value<0.000 and $\beta$=0.40, p-value<0.000; externalizing behavior: $\beta$=-0.41, p-value<0.000 and $\beta$=-0.45, p-value<0.000). Conventional regression models that accounted for potential confounders reduced the magnitude of these associations. Neither the association between father presence at nine months with reading ability or externalizing behavior was statistically significant in these conventional regression models ($\beta$=0.06, p-value>0.10; $\beta$=-0.03, p-value>0.10). Marginal structural models recovered the causal impacts of father presence at nine months and four years. Father presence in infancy exerted a stronger effect on kindergarten reading ability than it did in the year prior to school entry ($\beta$=0.26, p-value<0.01 and $\beta$=0.02, p-value>0.10). The joint effect of father presence at the two time points on reading ability was similar to that on approaches to learning in magnitude ($\beta$=0.28, p-value<0.000). Father presence had a significant impact on externalizing behavior at kindergarten entry as well, and the effects of exposure in nine months and at four years were similar ($\beta$=-0.19, p-value<0.01 and $\beta$=-0.22, p-value>0.10). The joint effect on externalizing behavior was -0.40 of a standard deviation (p-value<0.000).

*Differences in Effects Across Race*

Table 5 displays the effects for whites and blacks separately. The estimated effects from the marginal structural models of having a two-parent household in infancy were of identical magnitude for white and black kindergartners, though the confidence interval for black children included zero (whites: $\beta$=0.17, p-value<0.000; blacks: $\beta$=0.17, p-value>0.10). However, because the estimated effect of having a two-parent household at age four were weaker for blacks than whites, the joint effect of family structure at nine months and four years of age was about one-fourth of the size for blacks (Whites: $\beta_1 + \beta_2$=0.46, p-value<0.000; blacks: $\beta_1 + \beta_2$=0.12, p-value>0.10).

## DISCUSSION

This paper investigated the effect of living with two biological parents in infancy and age four on children's approaches to learning at school entry. Unlike prior studies on the longitudinal effects of family structure, I addressed the selection into family structure at multiple points using an approach made popular in the epidemiological literature. The main advantage to this technique-- marginal structural models with inverse probability of treatment weighting-- is that it allows for the estimation of a joint causal effect of family structure at multiple time points without "controlling away" the effect of the first exposure to the treatment.

Using this technique, I found that the presence of two parents in the household at nine months and at four years resulted in better approaches to learning in kindergarten. This study is the first to show that father presence is not only related to reading ability and problem behavior but also to skills such as attentiveness, task persistence, and academic engagement. The magnitude of the effect of family structure in infancy and pre-kindergarten on approaches to learning in kindergarten is comparable to the effects on reading and externalizing behavior: the absolute value of all effects are in the range of 0.3 to 0.4 of a standard deviation unit for the pooled sample of all children. This effect is theoretically meaningful because children with higher levels of academic engagement at school entry have higher reading and math ability later in elementary school (Duncan et al. 2007) and higher quality relationships with teachers and peers (Hughes and Kwok, 2006).

This joint effect of father presence at two points in childhood on kindergarten engagement can help explain the black-white skill gap at school entry in these skills. All children had higher levels of attention and academic engagement in kindergarten when they had a biological father who was present. Notably, the effect of a present biological father during infancy on approaches to learning was similar across race and almost two-tenths of a standard deviation. Since only 42 percent of black infants had two-parent households compared to 91 percent of white infants, this unequal distribution of single-parent homes across race during infancy may explain a sizable part of the three-tenths of a standard deviation black-white gap in approaches to learning in kindergarten. Though the joint effect for blacks of father presence in infancy and age four was one-fourth the size of that for white children, this was entirely due to a the absence of an effect of a present biological father for blacks at age four. One explanation for why a two-parent family mattered less for blacks than whites is that, by age four, blacks faced more economic hardship and family stress, regardless of family structure (Amato and Keith, 1991; Shaw, Winslow, and Flanagan, 1999).

Taken out of context, these analyses relating family structure to kindergarten readiness could be interpreted as situating the blame for black kindergartners' skill gaps with black families. It is important to remember that the many factors that drive family instability for black families—higher rates of black incarceration, joblessness, death, and persistent poverty—are inextricably related to structural racism (Wildeman and Western, 2010; Furstenberg, 2009). If we want to help black children have a more even starting point at school entry, it is not sufficient to promote marriage. Instead, we need to take a systematic approach to supporting families, reducing incarceration, and bolstering employment.

This study has limitations. Just like conventional regression or propensity score models marginal structural models with inverse probability of treatment weighting rest on the strong assumption that all factors that might confound the relationship between family structure at either nine months or four years and approaches to learning in kindergarten are observed. By using maternal characteristics to predict selection into family structure, I relied on the idea that there was strong assortative mating, or that fathers and mothers had similar education, ages, and work histories. This may not be the case, especially for black women who faced a shortage of marriageable men. Another limitation is that because of the need to condition on baseline and time-varying confounders, I could only measure whether a child had a two-parent home at two time points. This time-varying exposure to treatment, while better than a single snapshot measure of family structure, did not register family transitions that happen before or between the two measures. Finally, I did not distinguish between biological mothers who were stably single or partnered and those who had multiple partners in the household during the child's formative years. This distinction would matter in the estimation of the effects if multiple transitions in family structure affect child development more negatively than mere father absence.

## References:


Ainsworth, M. D. S. (1969). Object Relations, Dependency, and Attachment: A Theoretical Review of the Infant-Mother Relationship. *Child Development*, *40*(4), 969–1025.

Amato, P. R., & Gilbreth, J. G. (1999). Nonresident Fathers and Children's Well-Being: A Meta-Analysis. *Journal of Marriage and Family*, *61*(3), 557–573.

Beck, A. N., Cooper, C. E., McLanahan, S., & Brooks-Gunn, J. (2010). Partnership Transitions and Maternal Parenting. *Journal of Marriage and Family*, *72*(2), 219–233.

Bradbury, K., & Katz, J. (2002). Women's labor market involvement and family income mobility when marriages end. *New England Economic Review*, *4*, 41–74.

Cabrera, N. J., Shannon, J. D., & Tamis-LeMonda, C. (2007). Fathers' Influence on Their Children's Cognitive and Emotional Development: From Toddlers to Pre-K. *Applied Developmental Science*, *11*(4), 208–213.

Cabrera, N., Tamis-LeMonda, C. S., Bradley, R. H., Hofferth, S., & Lamb, M. E. (2000). Fatherhood in the Twenty-First Century. *Child Development*, *71*(1), 127–136.

Cancian, M., & Reed, D. (2009). Family structure, childbearing, and parental employment: Implications for the level and trend in poverty. *Focus*, *26*(2), 21–26.

Carlson, M. J., & Corcoran, M. E. (2001). Family Structure and Children's Behavioral and Cognitive Outcomes. *Journal of Marriage and Family*, *63*(3), 779–792.

Carlson, M. J., McLanahan, S. S., & Brooks-Gunn, J. (2008). Coparenting and nonresident fathers' involvement with young children after a nonmarital birth. *Demography*, *45*(2), 461–488.

Cavanagh, S. E., & Huston, A. C. (2008). The Timing of Family Instability and Children's Social Development. *Journal of Marriage and Family*, *70*(5), 1258–1270.

Cavanagh, S. E., & Huston, A. C. (2006). Family Instability and Children's Early Problem Behavior. *Social Forces*, *85*, 551.

Cheadle, J. E., Amato, P. R., & King, V. (2010). Patterns of nonresident father contact. *Demography*, *47*(1), 205–225.

Cherlin, A. J. (1999). Going to extremes: Family structure, children's well-being, and social science. *Demography*, *36*(4), 421–428.

Claessens, A., Duncan, G., & Engel, M. (2009). Kindergarten skills and fifth-grade achievement: Evidence from the ECLS-K. *Economics of Education Review*, *28*(4), 415–427.

Cole, S. R., & Hernán, M. A. (2008). Constructing inverse probability weights for marginal structural models. *American journal of epidemiology*, *168*(6), 656-664.

Coleman, J. S. (1988). Social Capital in the Creation of Human Capital. *American Journal of Sociology*, *94*, S95–S120.

Cooper, C. E., McLanahan, S. S., Meadows, S. O., & Brooks-Gunn, J. (2009). Family Structure Transitions and Maternal Parenting Stress. *Journal of Marriage and Family*, *71*(3), 558–574.

Cooper, C. E., Osborne, C. A., Beck, A. N., & McLanahan, S. S. (2011). Partnership Instability, School Readiness, and Gender Disparities. *Sociology of Education*, *84*(3), 246–259.

Downey, D. B., & Pribesh, S. (2004). When race matters: Teachers' evaluations of students' classroom behavior. *Sociology of Education*, *77*(4), 267.

Duncan, G. J., Brooks-Gunn, J., & Klebanov, P. K. (1994). Economic deprivation and early childhood development. *Child development*, *65*(2), 296–318.

Duncan, G.J., Dowsett, C. J., Claessens, A., Magnuson, K., Huston, A. C., Klebanov, P., … others. (2007). School readiness and later achievement. *Developmental Psychology*, *43*(6), 1428.

Duncan, G.J. and Magnuson, K.A. (2011). "The nature and impact of early achievement skills, attention skills, and behavior problems." *Whither Opportunity?: Rising Inequality, Schools, and Children's Life Chances*. (Greg J. Duncan & R. J. Murnane, Eds.). Russell Sage Foundation: 47–69.

Duncan, Greg J., & Magnuson, K. A. (2005). Can Family Socioeconomic Resources Account for Racial and Ethnic Test Score Gaps? *The Future of Children*, *15*(1), 35–54.

Edin, K., & Kefalas, M. (2005). *Promises I can keep: Why low-income women put motherhood before marriage*. University of California Press Berkeley.

Fomby, P., & Cherlin, A. J. (2007). Family Instability and Child Well-Being. *American Sociological Review*, *72*(2), 181–204.

Fryer, R. G., & Levitt, S. D. (2004). Understanding the Black-White Test Score Gap in the First Two Years of School. *Review of Economics and Statistics*, *86*(2), 447–464.

Furstenberg, F. F. (2009). If Moynihan Had Only Known: Race, Class, and Family Change in the Late Twentieth Century. *Annals of the American Academy of Political and Social Science*, *621*, 94–110.

Greenland S. (2003). Quantifying biases in causal models: classical confounding vs collider-stratification bias. Epidemiology 2003;14:300–6.

Harknett, K., & McLanahan, S. S. (2004). Racial and Ethnic Differences in Marriage after the Birth of a Child. *American Sociological Review*, *69*(6), 790–811.

Heaviside, S. & E. Farris (1993). Public School Kindergarten Teachers' Views on Children's Readiness for School (NCES No. 93-410). U.S. Department of Education, Office of Educational Research and Improvement, Washington, DC.

Hernán MA, Hernandez-Diaz S, Robins JM. (2004). A structural approach to selection bias. *Epidemiology*, 15:615–25.

Hernán MA, Robins JM. (2006). Estimating causal effects from epidemiological data. *J Epidemiol Community Health*, 60: 578–86.

Hughes, J. N., & Kwok, O. (2006). Classroom engagement mediates the effect of teacher-student support on elementary students' peer acceptance: A prospective analysis. *Journal of School Psychology*, *43*(6), 465–480.

Jackson, A. P., Choi, J.-K., & Franke, T. M. (2009). Poor Single Mothers with Young Children: Mastery, Relations with Nonresident Fathers, and Child Outcomes. *Social Work Research*, *33*(2), 95–106.

Jaffee, S. R., Moffitt, T. E., Caspi, A., & Taylor, A. (2003). Life With (or Without) Father: The Benefits of Living With Two Biological Parents Depend on the Father's Antisocial Behavior. *Child Development*, *74*(1), 109–126.

Knudsen, E. I., Heckman, J. J., Cameron, J. L., & Shonkoff, J. P. (2006). Economic, neurobiological, and behavioral perspectives on building America's future workforce. *Proceedings of the National Academy of Sciences*, *103*(27), 10155 – 10162.

Korenman, S., Miller, J. E., & Sjaastad, J. E. (1995). Long-term poverty and child development in the United States: Results from the NLSY. *Children and Youth Services Review*, *17*(1–2), 127–155.

Ladd, G. W., Birch, S. H., & Buhs, E. S. (1999). Children's social and scholastic lives in kindergarten: Related spheres of influence? *Child development*, *70*(6), 1373–1400.

Lin, H.L., Lawrence, F. R., & Gorrell, J. (2003). Kindergarten teachers' views of children's readiness for school. *Early Childhood Research Quarterly*, *18*(2), 225–237.

Marsiglio, W., Amato, P., Day, R. D., & Lamb, M. E. (2000). Scholarship on Fatherhood in the 1990s and Beyond. *Journal of Marriage and Family*, *62*(4), 1173–1191.

Martin, A., Ryan, R. M., & Brooks-Gunn, J. (2007). The joint influence of mother and father parenting on child cognitive outcomes at age 5. *Early Childhood Research Quarterly*, *22*(4), 423–439.

McLanahan, S., & Percheski, C. (2008). Family structure and the reproduction of inequalities. *Annu. Rev. Sociol*, *34*, 257–276.

McLanahan, S., & Sandefur, G. D. (1994). *Growing up with a single parent: What hurts, what helps*. Harvard University Press.

McLanahan, S. (2009). Fragile Families and the Reproduction of Poverty. *The ANNALS of the American Academy of Political and Social Science*, *621*(1), 111–131.

Osborne, C., & McLanahan, S. (2007). Partnership Instability and Child Well-Being. *Journal of Marriage and Family*, *69*(4), 1065–1083.

Osborne, C., Manning, W. D., & Smock, P. J. (2007). Married and Cohabiting Parents' Relationship Stability: A Focus on Race and Ethnicity. *Journal of Marriage and Family*, *69*(5), 1345–1366.

Pearl, J. (2000). *Causality: models, reasoning and inference* (Vol. 29). Cambridge: MIT press.

Pearl, J. (1995). Causal diagrams for empirical research. *Biometrika*, *82*(4), 669-688.

Qian, Z., & Lichter, D. T. (2007). Social Boundaries and Marital Assimilation: Interpreting Trends in Racial and Ethnic Intermarriage. *American Sociological Review*, *72*(1), 68–94

Rimm-Kaufman, S. E., Pianta, R. C., & Cox, M. J. (2000). Teachers' judgments of problems in the transition to kindergarten. *Early Childhood Research Quarterly*, *15*(2), 147–166.

Robins, J. M. (1999). Association, causation, and marginal structural models. *Synthese*, *121*(1), 151-179.

Robins JM. (1998). Marginal structural models. In: *1997 proceedings of the American Statistical Association, Section on Bayesian Statistical Science*. Alexandria, VA: American Statistical Association, 1998:1–10. (http://biosun1.harvard.edu/~robins/msm-web.pdf).

Robins, J. M., Hernan, M. A., & Brumback, B. (2000). Marginal structural models and causal inference in epidemiology. *Epidemiology*, *11*(5), 550-560.

Robins JM, Greenland S. (1986). The role of model selection in causal inference from nonexperimental data. *Am J Epidemiol*, 123:392–402.

Sayer, L. C., Bianchi, S. M., & Robinson, J. P. (2004). Are Parents Investing Less in Children? Trends in Mothers' and Fathers' Time with Children. *American Journal of Sociology*, *110*(1), 1–43.

Sharkey, P., & Elwert, F. (2011). The legacy of disadvantage: Multigenerational neighborhood effects on cognitive ability. *AJS; American journal of sociology*, *116*(6), 1934.

Shaw, D. S., Winslow, E. B., & Flanagan, C. (1999). A prospective study of the effects of marital status and family relations on young children's adjustment among African American and European American families. *Child Development*, *70*(3), 742-755.

Shonkoff, J. P., & Phillips, D. A. (Eds.). (2000). *From neurons to neighborhoods: The science of early childhood development*. National Academies Press.

Sigle-Rushton, W., & McLanahan, S. (2004). *Father absence and child wellbeing: A critical review*. Russell Sage Foundation.

Tamis-LeMonda, C. S., Shannon, J. D., Cabrera, N. J., & Lamb, M. E. (2004). Fathers and Mothers at Play With Their 2- and 3-Year-Olds: Contributions to Language and Cognitive Development. *Child Development*, *75*(6), 1806–1820.

Teitler, J. O., Reichman, N. E., & Nepomnyaschy, L. (2004). Sources of Support, Child Care, and Hardship among Unwed Mothers, 1999–2001. *Social Service Review*, *78*(1), 125–148.

Ventura, S. J. (2009). Changing patterns of nonmarital childbearing in the United States. *NCHS data brief*, (18), 1-8.

Ventura, S. J., & Bachrach, C. A. (2000). *Nonmarital childbearing in the United States, 1940-99*. US Department of Health and Human Services, Centers for Disease Control and Prevention, National Center for Health Statistics.

Wildeman, C., & Western, B. (2010). Incarceration in Fragile Families. *The Future of Children*, *20*(2), 157–177.

Wilson, W. J. (1987). *The truly disadvantaged: The inner city, the underclass, and public policy*. University of Chicago Press.

Wu, L. L., & Martinson, B. C. (1993). Family structure and the risk of a premarital birth. *American Sociological Review*, 210-232.

TABLE 1: FRACTION OF CHILDREN EXPOSED TO TWO BIOLOGICAL PARENT HOUSEHOLDS, BY WAVE AND RACE

| GROUP | |
|---|---|
| ALL (N=4628) | |
| Nine Months Old | 0.81 |
| Four Years Olds.. | 0.75 |
| Both………………… | 0.72 |
| WHITES (N=1973) | |
| Nine Months Old | 0.91 |
| Four Years Olds.. | 0.83 |
| Both………………… | 0.81 |
| BLACKS (N=673) | |
| Nine Months Old | 0.42 |
| Four Years Olds.. | 0.40 |
| Both………………… | 0.32 |

TABLE 2: SAMPLE MEANS/FRACTIONS FOR THE PREDICTORS OF TREATMENT STATUS AT EACH WAVE BY TREATMENT REGIME

| | All Children | Two Biological Parents in the Home at Neither Wave | Two Biological Parents in the Home at 9 Months But Not 4 Years | Two Biological Parents in the Home at 4 Years But Not 9 Months | Two Biological Parents in the Home at Both 9 Months and 4 Years |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| TREATMENT STATUSES: PRESENCE OF TWO BIOLOGICAL PARENTS IN THE HOUSEHOLD | | | | | |
| Proportion at 9 Months | 0.81 | 0 | 1 | 0 | 1 |
| Proportion at 4 Years | 0.75 | 0 | 0 | 1 | 1 |
| *N* | 4628 | 753 | 398 | 142 | 3335 |
| BASELINE COVARIATES | | | | | |
| *Child Race:* | | | | | |
| White…………………………………………………… | 0.43 | 0.21 | 0.45 | 0.20 | 0.48 |
| Black…………………………………………………… | 0.15 | 0.45 | 0.17 | 0.35 | 0.07 |
| Hispanic……………………………………………… | 0.20 | 0.20 | 0.21 | 0.23 | 0.20 |
| Asian…………………………………………………… | 0.11 | 0.01 | 0.03 | 0.04 | 0.15 |
| Other Race…………………………………………… | 0.12 | 0.12 | 0.15 | 0.17 | 0.11 |
| *Parental Relationship at Birth:* | | | | | |
| Bio Parents Married at Birth…………………… | 0.67 | 0.11 | 0.52 | 0.21 | 0.84 |
| Bio Parents Cohabitating at Birth……………. | 0.15 | 0.16 | 0.37 | 0.20 | 0.12 |
| Neither………………………………………………… | 0.18 | 0.73 | 0.11 | 0.59 | 0.04 |
| *Maternal Age at Birth:* | | | | | |
| Teen……………………………………………………. | 0.10 | 0.26 | 0.17 | 0.25 | 0.04 |
| Twenties……………………………………………… | 0.49 | 0.57 | 0.57 | 0.61 | 0.45 |
| Thirties or Higher…………………………………. | 0.42 | 0.17 | 0.26 | 0.14 | 0.51 |
| *Mother Wanted the Pregnancy*…………………. | 0.50 | 0.16 | 0.37 | 0.16 | 0.61 |

TABLE 2. CONTINUED

| | | | | | |
|---|---|---|---|---|---|
| *Maternal Work in Year Prior to the Birth:* | | | | | |
| No Work…………………………………………… | 0.27 | 0.28 | 0.27 | 0.33 | 0.27 |
| <=20 Hours Per Week………………………….. | 0.09 | 0.08 | 0.08 | 0.04 | 0.09 |
| >20 Hours Per Week……………………………. | 0.64 | 0.64 | 0.65 | 0.63 | 0.64 |
| *Maternal Education at Birth:* | | | | | |
| Less than High School…………………………… | 0.17 | 0.33 | 0.22 | 0.35 | 0.13 |
| High School Degree/GED……………………… | 0.29 | 0.43 | 0.44 | 0.46 | 0.24 |
| Some College…………………………………….. | 0.23 | 0.18 | 0.19 | 0.13 | 0.24 |
| College Degree or Higher……………………… | 0.31 | 0.06 | 0.15 | 0.06 | 0.39 |
| *Number of Siblings at Birth* | | | | | |
| None……………………………………………… | 0.38 | 0.49 | 0.43 | 0.39 | 0.35 |
| One………………………………………………… | 0.34 | 0.23 | 0.30 | 0.30 | 0.37 |
| Two………………………………………………… | 0.18 | 0.17 | 0.16 | 0.16 | 0.19 |
| Three or More………………………………… | 0.10 | 0.11 | 0.11 | 0.14 | 0.10 |
| ADDITIONAL COVARIATES USED TO PREDICT FAMILY STRUCTURE AT 4 YEARS | | | | | |
| *Below Poverty Threshold At Age 2……………* | 0.22 | 0.53 | 0.32 | 0.54 | 0.13 |
| *Child Care at Age 2: Parental Only……………* | 0.50 | 0.36 | 0.41 | 0.47 | 0.54 |
| *Times Per Week Reads to Child at Age 2:* | | | | | |
| Never…………………………………………….. | 0.03 | 0.05 | 0.05 | 0.03 | 0.03 |
| Once or Twice………………………………… | 0.24 | 0.36 | 0.25 | 0.39 | 0.21 |
| 3 to 6 Times………………………………… | 0.27 | 0.26 | 0.29 | 0.33 | 0.27 |
| Every Day……………………………………….. | 0.45 | 0.32 | 0.41 | 0.25 | 0.50 |
| *Number of Siblings At Age 2:* | | | | | |
| None……………………………………………… | 0.28 | 0.41 | 0.33 | 0.35 | 0.25 |
| One………………………………………………… | 0.38 | 0.31 | 0.35 | 0.30 | 0.40 |
| Two………………………………………………… | 0.22 | 0.16 | 0.20 | 0.22 | 0.24 |
| Three or More………………………………… | 0.12 | 0.12 | 0.12 | 0.14 | 0.12 |

TABLE 2 CONTINUED

| | | | | | |
|---|---|---|---|---|---|
| *Maternal Depression at 9 Months:* | | | | | |
| Not Depressed………………………………………. | 0.58 | 0.42 | 0.49 | 0.46 | 0.63 |
| Mildly Depressed……………………………………. | 0.24 | 0.26 | 0.30 | 0.32 | 0.23 |
| Moderately Depressed…………………………… | 0.11 | 0.17 | 0.13 | 0.15 | 0.09 |
| Severely Depressed………………………………… | 0.07 | 0.16 | 0.07 | 0.06 | 0.04 |
| *# Hours of TV/Weekday at Age 2:* | | | | | |
| None………………………………………………………. | 0.10 | 0.09 | 0.11 | 0.08 | 0.09 |
| One………………………………………………………… | 0.29 | 0.26 | 0.31 | 0.25 | 0.30 |
| Two………………………………………………………… | 0.27 | 0.23 | 0.27 | 0.26 | 0.28 |
| Three or More………………………………………… | 0.34 | 0.42 | 0.31 | 0.41 | 0.33 |
| OUTCOME MEASURES | | | | | |
| *Approaches to Learning* | 0.00 | -0.33 | -0.22 | -0.20 | 0.15 |
| *Reading Ability* | 0.00 | -0.33 | -0.10 | -0.36 | 0.17 |
| *Externalizing Behavior* | 0.00 | 0.35 | 0.24 | 0.08 | -0.15 |

TABLE 3:STABILIZED INVERSE PROBABILITY OF TREATMENT WEIGHTS

| | | | Percentiles | | | |
|---|---|---|---|---|---|---|
| | Mean | SD | 1st | 25th | 75th | 99th |
| ALL (N=4628)………………… | 0.93 | 0.96 | 0.18 | 0.73 | 0.87 | 7.63 |
| BLACKS (N=673)……………… | 0.95 | 1.21 | 0.29 | 0.51 | 0.81 | 8.47 |
| WHITES (N=1973)…………… | 0.98 | 0.78 | 0.10 | 0.84 | 0.95 | 6.70 |

TABLE 4: ESTIMATED EFFECTS OF A TWO-BIOLOGICAL-PARENT HOUSEHOLD IN MULTIPLE YEARS ON CHILDREN'S SKILLS IN KINDERGARTEN (N=4628)

| | Approaches to Learning | | | Reading Ability | | | Externalizing Behavior | | |
|---|---|---|---|---|---|---|---|---|---|
| | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model |
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Two Biological Parents in Household at 9 Months Only (Beta 1) | 0.43 | 0.00 | 0.19 | 0.47 | 0.06 | 0.26 | -0.41 | -0.03 | -0.19 |
| | (0.09) | (0.11) | (0.10) | (0.08) | (0.12) | (0.09) | (0.01) | (0.07) | (0.05) |
| Two Biological Parents in the Household at 4 years Only (Beta 2) | 0.43 | 0.16 | 0.16 | 0.40 | -0.01 | 0.02 | -0.45 | -0.25 | -0.22 |
| | (0.04) | (0.02) | (0.04) | (0.03) | (0.01) | (0.04) | (0.02) | (0.03) | (0.03) |
| Multi-Wave exposure (Beta1+Beta2) | | | 0.34 | | | 0.28 | | | -0.40 |
| | | | (0.06) | | | (0.05) | | | (0.08) |

TABLE 5: ESTIMATED EFFECTS OF A TWO-BIOLOGICAL-PARENT HOUSEHOLD ON CHILD SKILLS IN KINDERGARTEN BY RACE

| | Approaches to Learning | | | Reading Ability | | | Externalizing Behavior | | |
|---|---|---|---|---|---|---|---|---|---|
| | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model | Unadjusted Estimates | Regression Adjusted | Marginal Structural Model |
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| WHITES, N=1973 | | | | | | | | | |
| Two Biological Parents in Household at 9 Months Only (Beta 1) | 0.44 | -0.07 | 0.17 | 0.35 | 0.04 | 0.15 | -0.48 | -0.03 | -0.17 |
| | (0.0007) | (0.04) | (0.03) | (0.10) | (0.16) | (0.29) | (0.02) | (0.12) | (0.08) |
| Two Biological Parents in the Household at 4 years Only (Beta 2) | 0.46 | 0.21 | 0.28 | 0.29 | 0.01 | -0.03 | -0.49 | -0.28 | -0.35 |
| | (0.04) | (0.04) | (0.05) | (0.05) | (0.004) | (0.02) | (0.05) | (0.04) | (0.11) |
| Multi-Wave exposure (Beta1+Beta2) | | | 0.46 | | | 0.12 | | | -0.52 |
| | | | (0.02) | | | (0.27) | | | (0.19) |
| BLACKS, N=673 | | | | | | | | | |
| Two Biological Parents in Household at 9 Months Only (Beta 1) | 0.31 | 0.21 | 0.17 | 0.47 | 0.18 | 0.22 | -0.30 | -0.15 | -0.21 |
| | (0.01) | (0.10) | (0.14) | (0.02) | (0.18) | (0.14) | (0.06) | (0.03) | (0.05) |
| Two Biological Parents in the Household at 4 years Only (Beta 2) | 0.25 | 0.07 | -0.05 | 0.32 | -0.11 | -0.10 | -0.35 | -0.26 | -0.09 |
| | (0.01) | (0.06) | (0.07) | (0.04) | (0.05) | (0.06) | (0.004) | (0.04) | (0.21) |
| Multi-Wave exposure (Beta1+Beta2) | | | 0.12 | | | 0.13 | | | -0.30 |
| | | | (0.20) | | | (0.08) | | | (0.26) |