Family Structure and the Gender Gap in Educational and Behavioral Outcomes

David H. Autor
MIT and NBER

David Figlio
Northwestern University and NBER

Krzysztof Karbownik
Northwestern University

Jeffrey Roth
University of Florida

Melanie Wasserman
MIT

June 2014

Abstract

Using birth certificate data for all children born in Florida from 1992 to 2002 linked to detailed schooling outcomes, we investigate the effect of family structure on the gender gap in academic and behavioral outcomes. Specifically, our identification strategy entails contrasting the outcomes of gender-discordant siblings and twins to assess whether the boy-girl gap in standardized test scores and behavioral incidents varies systematically among children born to married and unmarried mothers. Employing sibling and twin fixed effect specifications to finely control for biological and home environment differences, we find robust evidence that children born to unmarried mothers exhibit a larger gender gap in school suspensions and absences than children born to married mothers, with the effects most pronounced among disadvantaged families. Notably, these gender gaps are not attributed to differences in at-birth health or different fathers. We provide additional evidence that the family structure gradient in gender gaps is robust to accounting for differential maternal demographic characteristics across family types. In examining potential mediators of the effects of family structure, namely environmental factors such as neighborhoods and schools, we conclude that the estimated family structure gradient is not mediated by these factors and speculate that it is instead due to differences in home environments and child-rearing strategies.

Introduction

The last four decades have witnessed a swift and substantial reversal of the gender gap in educational attainment in the United States. Following six decades of rapid progress between 1910 and 1970, U.S. high school graduation rates came to a near standstill after 1970 (Goldin and Katz, 2008). As this deceleration took hold, a gap opened between the graduation rates of males and females, and this gap widened over the course of four decades. Between 1970 and 2010, the high school graduation rate of U.S. women rose by an extremely modest 6 percentage points, from 81% to 87% (Murnane, 2013). In the same time interval, U.S. males made no net progress: their high school graduation rate was 81% in both 1970 and 2010. Thus, the ratio of female to male high school graduation rates among U.S. young adults ages 20-24 rose from unity in 1970 to 1.08 by 2010.[1] Because educational attainment is a cumulative process, one would anticipate that the growing female advantage in high school completions would lead to a corresponding gender gap in college graduations. In actuality, something more dramatic occurred. Between 1980 and 2010, the female-male ratio of four-year college attainment among U.S. adults ages 25 to 29 rose from 0.87 to 1.29 (U.S. National Center for Education Statistics, 2013). This swing of 42 percentage points is roughly five times as large as the corresponding gain in the female-male high school graduation ratio.

What accounts for the growing female advantage in educational attainment? Clearly, one central factor is the relaxation of prevailing gender norms that historically inhibited women from attaining the level of education commensurate with their abilities. As documented by Goldin et al. (2006) and Fortin et al. (2014), U.S. girls outranked U.S. boys in high school grade point averages for many decades. And yet, until the early 1980s, girls were substantially less likely than boys with comparable class ranks or IQ scores to enter and complete college. Consistent with the operation of changing gender constraints, women have overtaken men in secondary and tertiary education across the developed world. In 2011, college attainment among young women exceeded that of young men in 28 of 34 OECD member countries. Indeed, the ratio of female to male college attainment in the U.S. lay quite close to the mean all OECD countries in the same year.[2]

Yet, these aggregate facts are potentially less helpful for understanding a second salient feature of the data: the female advantage in high school graduation and college attainment is larger, and has risen by substantially more, among children of minority and low socioeconomic status (SES) families. For example, while the overall female/male high school completion ratio stood at 1.05 among U.S. whites in 2010, it was 1.11 among U.S.

[1] High school graduation rates refer to the status completion rate of U.S. born adults ages 20-24, and they include both traditional high school graduates and GED holders. Thus, 1970 graduation rates refer to cohorts born 1947-50, and 2010 graduation rates refer to cohorts born 1986-1990.

[2] Statistic refers to the fraction of men and women ages 25-34 who had completed at least 3 years of tertiary education as of 2011 in an ISCED 5A level education program (akin to a four year undergraduate program in the U.S.) (Organization for Economic Cooperation and Development, 2013 http://dx.doi.org/10.1787/888932848096, accessed 5/29/2014). Tertiary education rates were higher among young men than young women in Chile, Japan, Luxembourg, Mexico, and Turkey. In South Korea these rates were equal.

born Hispanics, and 1.17 among blacks.[3] The corresponding female/male college completion ratios were even larger: 1.18 among among U.S. whites ages 25-29 in 2010, 1.34 among blacks, and 1.42 among Hispanics (NCES 2013). These differences are also seen across SES groups. Among the high school graduating class of 1992, the female advantage in B.A. attainment was far higher among children of families in the bottom two SES quartiles than among the top two quartiles, and this gender gap had risen by substantially more among the bottom two quartiles during the prior twenty years (Goldin et al., 2006).

This paper empirically assesses one potential contributor to these growing cross-race, cross-SES differences in the gender gap in educational attainment: the changing structure of U.S. families. Specifically, we test whether boys raised in single-headed families are differentially disadvantaged relative to girls in their early educational preparedness, as measured by disciplinary behavioral issues and standardized tests administered in the third through eighth grades. Our focus on changing family structure as a contributor to gender differentials in skills development is motivated by a marked shift in the structure of marriage and child-rearing in U.S. families that has roughly coincided with the dramatic expansion of the gender gap in educational attainment. One aspect of changing family structure is the steep decline in the prevalence of marriage among young adults, and among less educated adults in particular. Between 1979 and 2008, the fraction of U.S. women ages of 25 and 39 who were currently married fell by roughly 10 percentage points among college-educated women, by 15 percentage points among women with some college but no degree, and by fully 20 percentage points among women with high school or lower education (Autor and Wasserman, 2013). These declines reflect a combination of a rise in the age at first marriage, a decline in lifetime marriage rates and, to a lesser extent, a rise in divorce rates among less educated women (Cherlin, 2010). As marriage rates have fallen, the fraction of children born to unmarried mothers and living in single-headed households has grown rapidly. Between 1980 and 2009, the fraction of all U.S. births accounted for by unmarried mothers more than doubled from 18 percent in 1980 to 41 percent (Autor and Wasserman, 2013). The rise in non-marital births was mirrored by an increase in the fraction of children living without biological fathers—and, as with falling marriage rates, this increase was concentrated among children of minority and less-educated individuals.

A corollary to these demographic trends is that U.S. children born into low-education and minority households spend a significant and rising share of their childhoods in single-parent, divorced, and remarried households, are exposed to a disproportionate number of adult partner relationships through cohabitation and remarriage among their primary caregivers, and experience a comparatively smaller number of years in which a second stable parental figure is present in the household (Cherlin, 2005). These differences in home environments may help to explain why children reared in single-headed families appear less successful

[3]Only among blacks was there a gender gap in high school graduation rates as of 1970 (61% among black males versus 66% among black females). Among Hispanics, male graduation rates were slightly higher than female graduation rates (by 2.5 percentage points), while they were identical among white males and females.

across a large set of educational and behavioral metrics, including disciplinary suspensions, involvement in risky behaviors (e.g., smoking and early childbearing), high school dropout (McLanahan and Percheski, 2008; McLanahan et al., 2013), and non-attainment of post-secondary education. Of immediate relevance for this paper, a variety of evidence suggests that boys raised in single-headed families appear to fare particularly poorly on disciplinary and educational outcomes relative to girls raised in similar circumstances (Autor and Wasserman, 2013; Bertrand and Pan, 2013; Buchman and DiPrete, 2006; DiPrete and Buchmann, 2014). In recent work, Bertrand and Pan (2013) document that boys raised in single-parent families exhibit twice the rate of behavioral and disciplinary issues as boys raised in two-parent families and are more than twice as likely to be suspended from school by the eighth grade. Analyzing detailed measures of home environment, school environment, and parent-child interactions, Bertrand and Pan (2013) find only modest evidence that parental and schooling inputs differ systematically between boys and girls across family types. However, their statistical analysis suggests that the quality of parental and child-rearing inputs have larger impacts on the behavioral and disciplinary outcomes of boys than it does of girls. Noting that material resources are fewer, parental attention is scarcer, and the quality of the home environment is generally poorer in single-headed families, Bertrand and Pan (2013) infer that the differential sensitivity of boys to these early inputs can help to explain why the gender gap in child development is larger in single-headed households.[4]

The current paper advances this line of research by providing what we believe is the sharpest feasible test of the relationship between family structure and the gender gap in early development and educational achievement. Using administrative records containing the universe of children born in Florida between 1992 and 2002, matched to their Florida public school records, we test whether the *within-family* gender gap in childhood outcomes differs systematically according to family structure.[5] Concretely, we employ an empirical framework that contrasts the outcomes of gender-discordant siblings (including gender-discordant twins) born to the same biological mother to assess whether, holding constant all common household attributes, the boy-girl (i.e., brother-sister) gap in disciplinary-behavioral incidents and test scores is greater among children born to single-headed families than those born to two-parent families.

The objective of our inquiry is to identify the causal effect of family structure on the gender gap in childhood outcomes. While the causal inference problem is potentially made challenging by the fact that family structure is not randomly assigned to children, our identification strategy does not require that potential outcomes of children are balanced across family types. Rather, it requires that the within-family, cross-sibling

[4] An alternative explanation put forth byBertrand and Pan (2013) addresses differential maternal treatment of boys and girls across family structure types. They provide evidence that single mothers report feeling more emotionally distant from their sons than their daughters, while the gap in emotional warmth is substantially smaller among mothers in families with both biological parents present. It is unclear whether the differential treatment of boys and girls is a cause or a consequence of boys' greater disciplinary issues.

[5] These data are also used by Figlio et al. (2014).

gender gap in potential childhood outcomes is uncorrelated with family structure at the time of birth. Stated concretely, we assume that any intrinsic genetic or biological disadvantage that boys may possess relative to their female siblings at birth are not systematically larger or smaller in single-headed relative to married families. While this assumption is not fully testable, we present evidence below on gender gaps in health at birth that buttresses its plausibility. Under this identifying assumption, we can causally attribute differences in the within-family, cross-sibling gender gap in childhood outcomes to differences in child-rearing environments that siblings experience as a result of family types. These environmental differences do not, of course, exclusively reflect the direct effect of single-headedness on child development; they also encompass other environmental differences that are either caused by, or correlated with, single-headedness, such as parental education, household income, neighborhood and school quality, and characteristics of peers. The unusual richness of our data permit us to explore and, in some cases isolate, these simultaneous channels of causation to potentially isolate the causal effect of family structure on the gender gap in childhood outcomes net of the effects of neighborhood and school quality, parental education, SES, race, ethnicity, and immigrant status.

By exploiting the substantial scale and high resolution of the longitudinal Florida database, our analysis overcomes several limitations of prior literature in identifying the determinants of the gender gap in developmental outcomes. One such limitation concerns the nature of non-experimental comparisons across genders. Prior work implicitly draws a contrast between boys and girls born to different mothers to assess the relationship between family structure and the gender gaps in child outcomes. This approach, necessitated by data and/or sample size constraints, leaves open the possibility that unmeasured differences among families with children of different sexes may in part explain the contrast in developmental outcomes. This threat to validity is of particular concern in light of evidence from Dahl and Moretti (2008) that the gender composition of children affects the evolution of household structure, with fathers relatively more likely to remain in the household following childbirth if the child is a boy. Our empirical approach of exploiting cross-gender, within-family sibling comparisons to identify the effect of family structure on the gender gap sweeps out unmeasured, time-constant family effects that may bias comparisons of children of different genders raised in different families. Additionally, the Florida data, sourced from birth records, enumerate family structure at birth—married, unmarried with father present, or unmarried with father absent—thus abstracting from any changes in family structure that occur postpartum.[6]

A second limitation that we address is that prior work is not able distinguish between gender differences in outcomes that are caused by differences in family conditions during childhood versus those gender differences that may already be present at birth. This is a potential concern because a long line of research, summarized

[6]There is also evidence from Almond and Rossin-Slater (2013) that unmarried fathers are more likely to establish paternity when the child is male.

in Aizer and Currie (2014), documenting that in-utero health differences, some of which are evident at birth, have substantial effects on health and economic outcomes, including education and earnings, over the course of a lifetime. Moreover, as our Florida data reveal, there are sizable and statistically significant differences in infant health at birth across family types, with newborns of married mothers exhibiting greater birthweight and higher Apgar scores than newborns of single mothers. These differences suggest that, even absent any differences in child-rearing environments, children of single-parent households would ultimately be expected to fare less well on behavioral and academic metrics than children born to married households. If the gender gap in neonatal health measures is also larger among siblings of single-parent families than among siblings of two-parent families, this would suggest that any correlation between family structure and gender differences in child outcomes may commence in utero and may therefore not be entirely attributable to differences in family environments. Our linked Florida birth record data allow us to explore this possibility directly by comparing within-household gender gaps in neonatal health across family types. The finding that these neonatal health gaps are largely invariant across family types serves as a key piece of evidence that our identification assumptions are plausible.

To preview the results, we find robust evidence of a causal effect of family structure on the gender gap in childhood behavioral outcomes—school suspensions in particular—though no consistent evidence of a causal effect on educational outcomes as measured by math and reading test scores. Comparing sons and daughters born either sequentially or simultaneously (in the case of gender-discordant twins) to the same mother, we document that boys born to single-headed families have a larger number of disciplinary incidents and suspensions during their primary and middle-school years and have elevated school absence rates. They additionally perform (modestly) less well on reading and math tests, but these patterns appear largely explained by differences in demographics across family types (e.g., mother's race, ethnicity, immigrant status, education and SES).The elevated gender gap in childhood disciplinary outcomes in single-headed relative to married families is found among all demographic subgroups that we examine, according to race, ethnicity, immigrant status, mother's education and household SES. However, the effect of single-headedness on the gender gap in childhood outcomes is substantially greater in families where the mother is from a less advantaged demographic group. Notably, the boy-girl disadvantage in disciplinary outcomes in married relative to non-married families is *as large or larger* in families where siblings possess different fathers as in families where siblings have both biological parents in common. Critical to our identification strategy, we document that although both boys and girls born to single mothers possess lower birthweight and somewhat lower Apgar scores than children born to married mothers, the gender gap in these birth characteristics does not differ by family structure.

We finally exploit the richness of our data to assess whether the differential gender gap in childhood

outcomes in single-headed families is proximately explained (i.e., mediated) by neighborhood of residence and school quality, both of which differ systematically according to family type. We perform this test by interacting a full set of neighborhood (zip code) and school fixed effects with indicator variables for gender to ascertain whether the family structure effect on the gender gap is in part due to the differential effects of neighborhoods and school quality on boys relative to girls. Surprisingly, accounting for these environmental factors has almost no impact on the estimated effect of family structure on the gender gap. In combination, these results suggest that the direct effect of family structure on the home environment may be a primary causal channel for its impact on the gender gap in childhood outcomes.

While the children observed in our sample are not yet old enough for us to observe adult outcomes such as high school dropout, labor force participation or criminal activity, there is strong reason to suspect that observed differences in disciplinary behaviors across genders and family types will later be manifest in differences in academic and labor market performance.[7] This in turn suggests that the steep decline in the prevalence of marriage among young adults, and among less educated adults in particular, is likely to pose an additional source of disadvantage to the educational and labor market achievement of low-SES and minority males.

The paper proceeds as follows. Section 1 details our identification strategy using a potential outcomes framework, provides a discussion of potential threats to validity, and specifies our main regression model. We provide information on our data sources, sample selection and descriptive statistics in Section 2. Section 3 begins by documenting non-experimental gender differences by family structure type and then proceeds to report our main estimates from sibling and twin fixed effect specifications of the effect of family structure on the gender gap in educational and behavioral outcomes. We explore heterogeneity of these results by maternal demographic characteristics in Section 4. Section 5 examines two potential pathways by which family structure could affect the gender gap in childhood outcomes. Section 6 concludes.

1 Empirical Framework

Our empirical objective is to isolate the causal effect of family structure on the gender gap in childhood disciplinary and educational outcomes. While the causal inference problem is potentially made challenging in this setting by the fact that family structure is not randomly assigned to children, we offer an identification strategy that does not rely on this form of random assignment.

[7]For example, Duckworth and Seligman (2005) find that measures of eight grade self-discipline are better predictors of subsequent academic performance than are measures of IQ. Similarly, Segal (2013) finds that eight-grade teacher evaluations of children's misbehavior (arguably akin to our school suspension measure) have important predictive power for adult earnings over and above eighth grade test scores. Segal (2013) also reports that, holding constant educational attainment, childhood misbehavior is associated with adult earnings at all educational levels, whereas achievement test scores are related to earnings only for young men with postsecondary degrees.

Let $Y_{jb} \in \left\{Y_{jb}^0, Y_{jb}^1\right\}$ and $Y_{jg} \in \left\{Y_{jg}^0, Y_{jg}^1\right\}$ equal the potential outcomes of two gender-discordant siblings, b and g, born to mother j whose marital status at the time of her children's birth is $D_j \in \{0, 1\}$, where $D = 1$ indicates the mother is married.[8] For simplicity, and in keeping with our empirical subsequent analysis, we focus on families where the mother's marital status is the same for both births, $D_{jb} = D_{jg}$, a condition that is automatically satisfied for twins. For each pair of siblings j, we observe only one set of potential outcomes as a function of their mother's marital status:

$$Y_{jb} = Y_{jb}^1 \times D_j + Y_{jb}^0 \times (1 - D_j) \text{ and } Y_{jg} = Y_{jg}^1 \times D_j + Y_{jg}^0 \times (1 - D_j). \tag{1}$$

In general, we expect potential outcomes within families j to differ between brothers and sisters and, similarly, for the potential outcome of each sibling to differ as a function of marital status. Consequently, absent random assignment of marital status to families, we cannot obtain a valid causal estimate of the effect of family structure on outcomes Y for either boys or girls. To see why, consider the non-experimental contrast between the outcomes of boys born to married versus single mothers. This contrast is,

$$\begin{aligned} E\left[Y_{jb}|D=1\right] - E\left[Y_{jb}|D=0\right] &= \left\{E\left[Y_{jb}^1|D=1\right] - E\left[Y_{jb}^1|D=0\right]\right\} \\ &\quad + \left\{E\left[Y_{jb}^1|D=0\right] - E\left[Y_{jb}^0|D=0\right]\right\} \end{aligned} \tag{2}$$

where the first bracketed term on the righthand side is the average causal effect of mother's marital status on outcome Y and the second bracketed expression is the well known bias term, stemming from differences in potential outcomes between boys who are born to married versus single mothers. Both intuition and data suggest that this bias term will be non-zero. Children born to married mothers weigh more and have higher Apgar scores at birth than children of single mothers, and the demographic characteristics of single and married mothers differ greatly in terms of race, education and age. Thus, we can learn little about the causal effect of marriage on children's outcomes by contrasting the outcomes of children from married and single families.

By a similar argument, we cannot contrast the outcomes of brothers and sisters within married families (or similarly, within single-headed families) to estimate the causal effect of marriage or single-headedness on boys relative to girls since we would not expect that potential outcomes in categories such as disciplinary incidents or early reading and math achievement would be otherwise comparable between brothers and sisters within any family type (i.e., $E\left[Y_{jb}^1\right] \neq E\left[Y_{jg}^1\right]$ and $E\left[Y_{jb}^o\right] \neq E\left[Y_{jg}^o\right]$.

To potentially surmount this identification problem, we select as our outcome of interest the gender *gap*

[8] For the purposes of exposition, we have limited the discussion here to gender-discordant siblings, since identification of the gender gap comes from these sibling pairs in a sibling or twin fixed effect model. In the empirical analysis below, we include all sibling and twin pairs, regardless of gender discordance. In this discussion, we also collapse the unmarried, dad present and unmarried, dad absent categories into one unmarried category D=0 for simplicity.

in sibling outcomes, $\hat{Y}_j^1 \equiv Y_{jb}^1 - Y_{jg}^1$ and $\hat{Y}_j^0 \equiv Y_{jb}^0 - Y_{jg}^0$, and we invoke the following identifying assumption:

$$\hat{Y}_j^1, \hat{Y}_j^0 \perp D_j, \tag{3}$$

meaning that we assume that the gap in potential outcomes between brothers and sisters is independent of their mother's marital status. Concretely, this assumption says that the *latent* (though not necessarily the observed) gap in childhood outcomes between brothers and sisters is as good as randomly assigned to families. If so, any observed difference in the gender gap across family types will reflect the causal effect of family type on the gender gap. Formally, we can estimate the causal effect of marriage on the gender gap by contrasting the within-family cross-sibling difference in outcomes across family types:

$$\begin{aligned} E\left[\hat{Y}_j | D=1\right] - E\left[\hat{Y}_j | D=0\right] &= E\left[Y_{jb}^1 - Y_{jg}^1 | D=1\right] - E\left[Y_{jb}^0 - Y_{jg}^0 | D=0\right] \\ &= E\left[Y_{jb}^1 - Y_{jg}^1 | D=1\right] - E\left[Y_{jb}^0 - Y_{jg}^0 | D=1\right], \end{aligned} \tag{4}$$

where the second equality in (4) invokes assumption (3).

Is this strategy for identifying the causal effect of family structure on the behavioral and educational outcomes of boys relative to girls plausible? Assume for example, in violation of our identifying assumption, that mothers' marital status during pregnancy differentially affected the fetal development of sons relative to daughters—which could occur if the male fetus is more sensitive than the female fetus to the mother's stress levels during pregnancy. In this case, the contrast formed in equation (4) would confound any causal effect of family structure on the gap in outcomes between boys and girls with the differential in-utero impact of maternal stresses on the subsequent development of children of each sex. Thus, for our estimates to cleanly capture the impact of family structure on child development as it operates through the home, school, and child-rearing environment, it must be the case that there is no differential effect of family structure on the latent gender gap in behavioral and educational outcomes prior to birth. Following the fundamental problem of causal inference, this assumption is not directly testable. But we can partially assess its validity by testing whether gender gaps in neonatal health, measured by birthweight and Apgar scores, differ across family types. As noted in the Introduction, we find no evidence that these gaps vary with family type, which lends credibility to the assumption.[9]

Assuming that our key assumption of non-confoundedness in potential outcomes holds, the econometric

[9] A concern remains that we are introducing selection into our sample by limiting it to mothers who have at least two children. Evidence from Dahl and Moretti (2008) suggests that a first born girl increases women's total fertility, though the effect for the marginal second child is extremely small. If mothers who are induced to have a second child by the gender of their first child have a preference for boys, then estimates from the sample of mothers with at least two children will not be externally valid to the sample of all mothers. Even if the propensity to have a second birth is unaffected by the gender of the first birth, the probability of higher order births is still in question. As shown in Dahl and Moretti (2008), having a female first child positively affects whether a woman has three or more children. If these effects primarily operate in married families, it is possible that the gender differences by family structure estimated in this analysis also include the effect of larger family size for female first births to married mothers. In our analysis, we additionally utilize the sample of twins and twin fixed effect models in order to circumvent these potentially confounding factors.

framework highlights that observed differences in the gender gap across family types may be attributable to three distinct factors: direct causal effects of family structure on child development; indirect environmental factors that are influenced by, but are not a direct consequence of, family structure and exert a causal effect on child development (pathway effects); and (spurious) confounding effects stemming from the (non-causal) association between family structure and underlying causal factors. The direct effect of family structure on the gender gap, in our terminology, refers to any intrinsic differences among family structures that affect the home and child-rearing environment. For example, children of single-headed households will generally experience relatively less contact with the non-resident parent than would occur if the parents were cohabiting, and will typically be exposed to a relatively larger number of adult partner relationships through cohabitation and remarriage of their primary caregivers. We denote these differences in home environments as 'direct' consequences of household structure, though we recognize that there are single-headed families in which both biological parents play primary caregiving roles and, similarly, single-headed families in which the household head's partner relationships are stable over many decades.

We distinguish these direct effects of family structure on the gender gap in childhood outcomes from potential indirect environmental effects that may serve causal as pathways through which family structure affects childhood outcomes but which are not intrinsically tied to family structure. One such pathway is family poverty. In 2010, 41 percent of children in female-headed households without a husband present were living in poverty, relative to 24 percent of children in male-headed households without a wife present, and 8.8 percent of married households with two parents present (U.S. Bureau of the Census, 2011). These sharp differences in poverty across family structures imply that on average children in single-headed households have access to fewer material resources, reside in poorer quality housing and in more dangerous neighborhoods, and likely attend lower quality schools. If boys are more sensitive than girls to these inputs, as suggested by Bertrand and Pan (2013), then these environmental factors may partly or fully account for systematic differences in the gender gap across family types. We refer to these causal effects as indirect because although these cross-family type differences in child-rearing environments may be partly attributable to family type (e.g., single-headed households generally have lower earnings capacity than two-parent households), these differences are not intrinsic to family structure. Our detailed data allow us to directly assesses the role of two such mediating factors—neighborhoods and schools—in inducing the differences in gender gaps in childhood outcomes among family types observed in the data.

The third (confounder) channel of causality refers to other predetermined differences across families that exert a causal effect on the gender gap in childhood outcomes and are correlated with, but not caused by, single-parenthood. Factors in this category will naturally include demographic characteristics of parents that differ across household types. Concretely, single-headedness is more prevalent among among black mothers

than among white or Hispanic mothers, more prevalent among high school dropout mothers than among better educated mothers, and more prevalent among younger mothers than among mothers who are slightly older. If these maternal attributes directly affect the gender gap in child outcomes (e.g., holding constant family structure, boys of high school dropout mothers fare worse than girls of high school dropout mothers) and we fail to control adequately for them, we might incorrectly attribute their causal effects to family structure.[10] We investigate this possibility by using our extensive database to draw sharp comparisons of the gender gap in child outcomes among families with observationally equivalent maternal characteristics.[11]

1.1 Implementation

We bring this conceptual framework to the data by estimating the following model using OLS:

$$Y_{ij} = \beta_0 + \beta_1 \text{Male}_i + \beta_2(\text{Male}_i \times \text{Married}_j) + \beta_3(\text{Male}_i \times \text{DadPresent}_j) + \gamma_j + X'_{ij}\delta + e_{ij}. \quad (5)$$

Here, Y_{ij} represents outcome for child i born to mother j, Male_i is an indicator variable for whether the child is male, Married_j is an indicator for whether the mother of child i is married at birth and DadPresent_j is an indicator for whether child $i's$ mother is unmarried and the father is present at $i's$ birth. The omitted family category is therefore births to unmarried mothers where the father is not present at birth. The regression model also includes mother fixed effects γ_j, which sweep out mother-specific variation that is constant across sibling births, such as shared genetics or common elements of the home environment. The vector X_{ij} additionally controls for mother and child attributes that may vary across births, including month and year of birth, birth order, and the age at birth and the education level of the mother. In our primary analysis, the sample is restricted to siblings with the same marital status at birth (thus, we do not subscript Married_j and DadPresent_j with i since they do not vary across sibling pairs), though we we will subsequently explore (in a later draft) the relationship between family structure and child outcomes for children born to the same mother but with different paternal statuses at birth.[12]

The coefficients of interest are β_2 and β_3, which represent the causal effect of family status at birth (married, unmarried mother with father present) relative to the omitted category of unmarried mother with father not present. One concern with the estimation approach is that we may potentially bias estimates of β_2

[10] Note that even in this case, these predetermined attributes would be legitimate causes of the gender gap under our identification assumption of non-confoundedness between family structure and the gender gap in potential outcomes. But in this case, the causal effects would be attributable to these differences in parent characteristics rather than differences in family structure per se.

[11] More speculatively, our research design allows us to draw within-mother, cross-family structure contrasts between children born to the same mother under different relationship circumstances (married, father present, father absent) to ask both whether, holding constant all fixed mother characteristics, children born to the same mother fare differently if they are born into different family statuses (married, single-headed), and further whether these within-mother, over-time differences in relationship status effect the gender gap.

[12] Due to the sample restriction that both siblings have the same marital status at birth, the main effects Married_j and DadPresent_j are not identified in this model.

and β_3 by restricting the sample to sibling pairs in which the mother's marital status (or dad's presence) did not differ between their births. If the sex of the first child has a causal impact on the probability of a change in family structure (e.g. from married to divorced), then it may be invalid to treat the gender of children as being as good as randomly assigned across family types. We find no evidence in our data, however, that the gender of the first child is correlated with family structure transitions between first and second births.[13] A final concern that our sample is limited to mothers who have at least two children, and thus we cannot estimate the effect of family structure on boys versus girls for children who do not have siblings.

2 Data and Descriptive statistics

2.1 Data Sources

Data for our main analysis are drawn from two sources: the universe of birth certificates for the state of Florida for years 1992 through 2002, sourced from the Florida Bureau of Vital Statistics; and linked school records from 1996 through 2012 from the Florida Department of Education Data Warehouse for children in these birth cohorts who attended the Florida public schools. Birth certificate data were matched to schooling records using first and last names, exact date of birth and social security number. Of the 2,047,663 births recorded by the Florida Bureau of Vital Statistics during 1992-2002, 1,652,333 were observed in Florida public school data, representing an 80.7% match rate. The match rate is almost identical to the percentage of children who are born in Florida, reside there until school-going age, and attend public school, as computed by the Census and American Community Survey for years 2000 through 2009.[14]

Florida birth certificates enumerate demographic characteristics of the mother (including marital status), health and demographic characteristics of the the newborn, and demographic characteristics of the father if he appears on the birth certificate.[15] School records include third grade through eighth grade Florida Comprehensive Assessment Test (FCAT) math and reading scores, as well as records of all suspensions and school absences.[16] We use these data to create three family status categories: parents married at birth (*married*); unmarried but the father's name appears on the birth certificate (*dad present*); and unmarried and the father's name does not appear on the birth certificate (*dad absent*).[17]

[13] Results available upon request.

[14] Figlio et al. (2014) provide many further details on the matching process, as well as extensive validity checks on the matching procedure.

[15] In the very small number of cases (< 100 sibling pairs) where the race or ethnicity on siblings' birth certificates do not match, we assign the race and ethnicity of the firstborn child.

[16] Although available for a subset of the birth cohorts, we exclude high school test scores for two reasons: the subset of children who have reached high school age is still limited leading to small sample sizes and the high school FCAT experienced a major change between birth cohorts, increasing measurement error.

[17] For children born to married couples, information on paternity may be provided by either the mother or father. For children

2.2 Sample Selection and Descriptive Statistics

Table 1 presents summary statistics for the universe of Florida births for years 1992 through 2002 (column 1). Column 2 summarizes the subset of singleton (i.e. non-multiple) births that were matched to school records. Due to data availability restrictions we are only able to match siblings for a subset of Florida counties (containing 56 percent of the population), and only for birth years 1994 through 2002. Column 3 reports statistics for the subset of singleton births with outcomes from the counties where sibling matches are possible. Column 4 limits the sample further to all sibling pairs excluding twins (tabulated separately).[18] The final two columns of the table further restrict the sample to siblings pairs where the mother's marital status (married, father present, father absent) does not change between births. We label such pairs as stable siblings. Column 5 contains all such stable sibling pairs except for twin births while, conversely, column 6 includes twins exclusively. By definition, twins possess the same family structure at the time of birth.[19]

Relative to the population of Florida births (column 1), the matched singleton sample of children (column 2) is more likely to be born to a mother who is black or Hispanic, younger and less educated. Additionally, the matched sample has a higher incidence of unmarried parents. These differences in socio-economic status are consistent with both the higher cross-state mobility of more educated and higher SES adults,and the omission of children born in Florida who subsequently attend private school. While the matched sample represents a more disadvantaged population than the full sample, there are virtually no differences in at-birth characteristics between the full population and the matched sample.

When we limit the sample to matched singleton births with outcomes in the subset of counties where siblings matches are possible (column 3), we observe that the subset of counties where matching siblings is feasible appear quite similar to the overall matched singleton sample (column 2), with the exception that column 3 is comprised of children born to mothers who are more likely to be black, Hispanic, and slightly older. In further restricting the sample to matched siblings in column 4, we observe that mothers in the matched sibling sample are somewhat younger, more likely to be married, and more likely to have a college degree than the full singletons sample. This likely reflects the fact that married mothers have somewhat higher fertility and are better educated than unmarried mothers, and hence are overrepresented in the siblings sample.

Imposing the additional restriction to the column 4 sample that both children in a sibling pair possess the same family structure at birth yields our analytic sample of non-twins in column 5. Although family structure at birth is the same within these sibling pairs, it bears note that siblings may have different fathers and,

born to unmarried parents, information about the father appears on the birth certificate only if he claims paternity. Thus, fathers' presence or absence refers to their legal paternity status, not necessarily their presence or absence at the delivery.

[18] More precisely, this column excludes births that do not meet the following criteria: a sibling's birth is observed in the Florida birth record data; the two siblings correspond to the first and second births observed for this mother between 1994 and 2002; both siblings are matched to Florida school records.

[19] There are a small number of instances (<100) in which twins have different family structures reported on their birth certificate (erroneously, we presume). We drop these cases.

similarly, married mothers may have different spouses at consecutive births. In section 3.2 Family Structure and Paternity, we present results that further subdivide the sample into siblings with and without a father in common. Since women who are married at their first delivery are relatively more likely to maintain their initial family status at the second birth, the column 5 (stable siblings) sample contains a larger fraction of married families relative to father-present and father-absent families. Similarly, mothers in the stable siblings sample possess higher education on average, are slightly older, and are more likely to be white than are mothers in the all siblings sample (column 4).

In comparison to the stable siblings sample (column 5), mothers in the twins sample (column 6) are considerably less likely to be Hispanic or foreign-born, and more likely to be black and older. Despite the demographic differences between the twins and non-twins sample, we use twins as a second main analytic sample because within-family, cross-twin comparisons have high internal validity since twins are necessarily born into the same family circumstances, and are in many cases exposed to very similar household, neighborhood and school environments throughout childhood. Notably, our primary causal estimates for twins and non-twins are largely comparable, with a few exceptions where noted, though the twins analysis is often imprecisely estimated due to the smaller sample size.

Panel B of the table summarizes data on infant health at birth, including birthweight, birth complications, and Apgar scores taken five minutes after delivery[20], and panel C summarizes our outcome measures for childhood disciplinary behaviors and skills development, sourced from Florida school records for grades three through eight for our matched sample. The school-based outcome measures include standardized scores on math and reading tests administered in each grade, days suspended from school in each academic year, and absence rate from school in each academic year. Each child in the sample can contribute up to six observations for each variable, one per each year observed in each grade. In all regression estimates that follow, we cluster standard errors at the level of mothers to account for the high degree of within-child and within-family serial correlation in these outcomes. We discuss these birth characteristics and childhood outcomes below at length. We note here that, consistent with extensive prior evidence, twins have poorer health at birth, including lower average birth weights, more complications and lower five-minute Apgar scores (panel B). These differences in at-birth health persist and manifest in subsequent school outcomes. As is visible in panel C, children in the twins sample have lower reading and math scores and a higher incidence of disciplinary infractions during elementary and middle school grades.

[20]The Apgar score is a summary measure of newborn health. Scores range from 0 to 10, with scores of 7 and above considered normal.

3 Main results

Figures 1 and 2 provide an initial overview of the four educational and behavioral outcome variables that we study using the Florida public schools data: reading test scores, math test scores, suspension days, and absence rates. We plot means of these variables by family structure at birth for our two main samples (non-twins and twins), where the bars in each panel correspond to outcomes for boys (blue), girls (red), and the mean boy-girl difference in outcomes (orange). The whiskers on each bar correspond to the 95 percent confidence intervals for each estimate. Table A1 summarizes these measures by family types. In Panel A of Figure 1, we observe a clear rank ordering of math test scores by the three family structure categories: boys and girls born to married mothers perform best; those born to unmarried mothers with father present perform less well; those born to unmarried mothers with father absent perform least well. The differences in test scores across family types are sizable, particularly for boys. Boys in father-absent families score approximately 0.55 standard deviations below the Florida average on the math test, those in father-present families score approximately 0.15 standard deviations below the average, and those in married families score approximately 0.50 standard deviations above the average. For females, the corresponding gaps are approximately -0.45, -0.15, and $+0.45$. The fact that gradient is shallower for females than males implies that the gender gap in math test scores differs across family types.

A comparison of the boy-girl gap in math scores across family types, measured by the third bar in each set of columns in the figure, demonstrates this gradient. As detailed in Table A2, the boy-girl gap in math test scores is negative 0.056 standard deviations and highly significant in father-absent families, positive 0.006 standard deviations and insignificant in father-present families, and positive 0.051 standard deviations and highly significant in married families. Thus, the gap flips sign across family types: it favors boys in married households and girls in father-absent households. As shown in the righthand pane of Panel A, the pattern of math test scores for twins is extremely comparable to that of non-twin siblings. Panel B of Figure 1 reveals analogous, though somewhat less pronounced, patterns for reading scores: average reading scores for both sexes are lowest in father-absent families and highest in married families; the female advantage in reading scores is greatest in father-absent families, somewhat smaller in father-present families, and smaller still in married families. This pattern is also observed for twins (righthand pane).

Figure 2 presents levels and boy-girl gaps in behavioral outcomes by family type. As shown in Panel A, the mean number of days that a child is suspended (per grade), a measure of disciplinary infractions, appears highly sensitive to family structure type. Children born to married families experience approximately one-tenth the number of disciplinary infractions of children born to father-absent families. Among children born to married, father-present and father-absent families, the mean number of days suspended per grade is 0.17,

0.66, and 1.43, respectively. Variation in the gender gap in suspensions across family types is equally stark, with the boy-girl suspension gap equal to 0.19 in married families, 0.55 in father-present families, and 1.10 in father-absent families (all contrasts are highly significant).[21] The righthand pane of Panel A displays a highly comparable pattern of disciplinary infractions for twins, while Panel B provides the analogous comparison for school absence rates (measured in percentage points). Absence rates do not differ as substantially across family types as do test scores or disciplinary incidents, but the qualitative pattern is similar: absences are highest in father-absent families and lowest in married families; absence rates of boys exceed those of girls overall; the boy-girl gap is modest in size but highly significant in father-absent and father-present families, and it is essentially non-existent in married families; and the pattern for twins is similar except that there is no boy-girl gap in absences in father-present families.

In summary, these figures provide strong suggestive evidence of a gender gap in educational and behavioral outcomes that varies systematically with family structure. We now provide a rigorous and detailed analysis of these relationships.

3.1 Between and Within-Family Comparisons

We begin by estimating a regression version of the comparisons in Figures 1 and 2 above. We then refine the comparison by contrasting outcomes between brothers and sisters born to the same mother under the same family structure. The first column of Table 2 estimates a simplified version of our primary estimation equation (5) that excludes mother fixed effects and is fit to the primary non-twin sample. Following the ordering of Figures 1 and 2, we present estimates in sequence for math scores, reading scores, suspension days, and school absence rates. In this estimates, and in all that follow, standard errors are clustered on unique mother identifiers.

Table 2 column 1 documents significant gaps in math test scores between boys and girls and among family types, as well as significant differences in the boy-girl gap across family types. Because the column 1 estimates contain no other covariates, the point estimates correspond precisely to the contrasts depicted in Figure 1 above. Girls in father-present families and married families score, respectively, 0.29 and 0.92 standard deviations above girls in father-absent families. The boy-girl gap in math test scores is -0.06 standard deviations in father-absent families. In father-present families, this gap is reduced by 0.06 standard deviations (a net gap of zero), and is reduced by 0.11 standard deviations in married families (a net gap of $+0.05$).

As shown in Table A1, demographic characteristics of parents and children vary substantially across

[21] Since boys have substantially higher levels of suspensions than girls across all family types, it might instead be appropriate to compare the ratio of suspensions by sex across family types rather than the difference in levels. The boy/girl ratio in days suspended is approximately 2.8 in married families, 3.6 in father-present families, and 4.7 in father-absent families.

family types, with father-absent and father-present families appearing disadvantaged relative to married families along a number of dimensions. To account for these demographic differences across family types, we add in column 2 a vector of covariates that includes the mother's age and education at the time of the child's birth, and a set of variables coding the child's birth order (1st or 2nd) and month and year of birth (dropped for twins). Inclusion of these covariates substantially shrinks the family structure gradient in test scores. For girls, the married versus dad-absent gap shrinks from 0.92 to 0.24 standard deviations, and the father-present versus father-absent gap shrinks from 0.29 to 0.09 standard deviations. The gender gap in math test scores is almost entirely unaffected by inclusion of these covariates, however.[22]

Column 3 provides a more rigorous test of the relationship between household structure and the gender gap in childhood outcomes by including a full set of mother fixed effects. The key coefficients in this model are therefore identified exclusively by within family contrasts in gender-discordant sibling outcomes. Naturally, all time-constant mother and family characteristics—including marital status—drop out of the regression model with mother fixed effects included. Strikingly, the estimated gender gap in test scores is essentially unchanged when using within-family, cross-sibling contrasts for identification. Column 4 finally estimates these same relationships using the twins sample, which arguably provides still tighter quasi-experimental control. Though precision is substantially lower in the twins sample (reflecting the 80 percent reduction in sample size), the qualitative pattern of results is clearly maintained.

These initial estimates are of signal importance for our analysis. For our estimation approach to recover the causal effect of family structure on the gender gap in childhood outcomes, it must be the case that the latent gender gap in sibling outcomes is uncorrelated with family type. While we cannot test this assumption directly, the estimates in columns 1 and 2 make clear that key determinants of children's outcomes vary substantially across family types—which explains why including family demographic measures in column 2 greatly reduces the main effect of family type on children's outcomes. This pattern might also suggest that other unmeasured aspects of parental quality could explain the variation in the gender gap across family types. What the column 3 estimates indicate is that if these measured or unmeasured attributes are present and determinative, they are uncorrelated with the gender mix of siblings or with family type, or with both. We explore many such candidate explanatory variables below and allow them to non-parametrically impact the gender gap, something that is not permitted by our specification in Table 2. Before performing these extensions, we summarize additional outcome measures and provide supportive evidence that the main identifying assumption is valid.

We assess the relationship between family structure and the gender gap in reading scores in the second

[22]This pattern suggests that the gender of children is almost uncorrelated with family demographics, a pattern consistent with the means displayed in Table A1.

four columns of Panel A. As suggested by above Figure 1 above, the gender gap in reading appears less sensitive to family structure than does the gender gap in math scores. Nevertheless, the female advantage in reading scores is significantly greater in married families than in father-absent families in all specifications that use the non-twin siblings. When we focus on twins in column 8, we find a similar point estimate to the sibling fixed effects model, but precision is substantially lower and the estimate is insignificant.

Turning to disciplinary outcomes in Panel B, we find large and robust effects of family structure on the number of days suspended of boys relative to girls. The column 2 estimate indicates that boys in father-absent families accumulate 1.1 more suspension days per year than do girls in father-absent families. Father-presence is estimated to reduce this gap by 0.55 days (50%) and marriage is estimated to reduce it by 0.90 days (82%). We detect an almost identical pattern when we contrast outcomes between siblings. Interestingly, the gender gap gradient in disciplinary incidents across family types among twins is somewhat smaller (by 30 to 50 percent) than among non-twin siblings. This could potentially indicate that the between-twin contrast does a better job of holding constant unmeasured family and environmental factors that contribute to the gender gap in non-twin families. Or, alternatively, family structure may have a smaller causal effect on the gender gap among twins than non-twins, perhaps because the challenges that a family faces when raising twins may partly offset the protective effect that family structure (e.g., marriage, father-presence) has on boys.

The final four columns of Panel B find comparable relationships between family structure and the boy-girl gap in a second behavioral outcome, school absences. Relative to the OLS estimates in column 6, the inclusion of mother fixed effects in column 7 slightly accentuates the contrast in gender gaps in absences across family types. Conversely, the twins-only specification suggests a small contrast in the gender gap across family types, particular for father-present families.

In net, the Table 2 estimates offer clear evidence that male disadvantage in educational and behavioral outcomes is larger in unmarried than married families, particularly in families where the father is not named on the birth certificate. Under our identifying assumption of unconfoundedness between family structure and the latent gender gap in sibling outcomes, these estimates may be interpreted as causal effects of family structure on the gender gap. We next provide evidence on the plausibility of this identifying assumption, then subsequently consider a number of channels through which the family structure gradient in the gender gap may operate.

3.2 Family Structure and Paternity

One critical dimension on which family types differ is the stability of adult presence in children's lives. As (Cherlin, 2005) documents, children born to unmarried families in the United States experience a dispropor-

tionate number of adult transitions, reflecting the relative impermanence of non-married family arrangements. One manifestation of these adult transitions is the complexity of paternity relationships within the household, in particular, the extent to which siblings born to the same mother have the same or different fathers. The Florida data permit us to measure the frequency of this phenomenon—i.e., father concordance and discordance—by comparing the exact dates of births (recorded on birth certificates) of siblings' fathers.[23] Father concordance is substantially higher among siblings born to married mothers (94.6% of births) in our Florida data than among siblings more to non-married, father-present households (69.0% of births). While we cannot measure the frequency of father concordance among siblings born to father-absent families—since father's birthdate is not recorded in those cases—we presume that it is considerably lower in these families.[24]

How much does the gradient in the gender gap in child outcomes across family types stem from differences not only in marital arrangements but also differences in patterns of paternity? Table 3 explores this question by splitting our non-twins sample into father-concordant and father-discordant subsamples. The odd numbered columns assess the gender gap in sibling outcomes in father-concordant families by contrasting married and non-married father-concordant families. The even numbered columns of the table use the complementary sample of married and non-married father-discordant families. We exclude father-absent households from both sets of columns because since information on siblings' paternity is not available in these cases.

The Table 3 estimates find that the effect of marriage on the gender gap in childhood outcomes is surprisingly similar between father-concordant and father-discordant families. Among father-concordant siblings, the boy-girl gap in suspensions and absences are 0.32 days and 0.18 percentage points smaller, respectively, in married than in father-present families. Similarly, the boy-girl relative advantage in math and reading scores is approximately 0.03 standard deviations greater in married relative to father-present families. One might have anticipated that these relative gains in boy's outcomes would be smaller in married relative to unmarried father-discordant families since, presumably, family stability is potentially lesser in father-discordant households, even conditional on marriage. The data do not support this intuition, however. The differential gains for boys in married relative to unmarried families appear if anything slightly larger in father-discordant than father-concordant households. However, these slightly larger relative gains for boys in married, father-discordant families should be compared to a baseline of larger boy relative *disadvantage* in father-discordant unmarried families. Indeed, on three of four outcome measures (math, reading, and suspensions), boys in unmarried father-discordant families fare worse than boys in unmarried father-concordant families. On these

[23]Matching only on father's birth date likely results in a small number of false matches—a mother might conceive children with different men who were born on the same date—but these cases are likely very rare (since both day, month and year must match). There seems a greater likelihood of false negatives due to recording errors on birth certificates. But the fact that the father concordance rate is 94.6% for births to married mothers suggests that these recording errors are also rare.

[24]Note that our non-twin samples are limited to families that are type-stable across births (e.g., married, father-present, or father-absent at both births). This requirement likely increases the rate of father-concordance in our married and father-present samples relative to a sample where we did not impose type-stability across births.

same three outcome measures, the boy-girls relative gain to marriage in father-discordant families is larger than in father-concordant families. Thus, the seemingly large benefit that boys obtain relative to girls in married father-discordant families should be understood as an improvement only *relative to* the comparison group of boys in unmarried father-discordant families, a set of boys who appear to fare particularly poorly.

In net, these estimates indicate that the family status gradient in the sibling gender gap does not stem primarily from differences in paternity patterns across family types: whether or not siblings have the same father, the boy-girl disadvantage in educational and behavioral outcomes appears to be reduced by marriage.

3.3 Assessing the Unconfoundedness Assumption Using Balance at Birth

Our identification framework entails orthogonality between the gender gap in potential outcomes and the family structure into which children are born. Though we cannot directly test the validity this assumption, we can assess supporting evidence by comparing the relationship between gender differences in at-birth outcomes and family structure at birth. Specifically, we examine infant health as measured by infant weight, complications at birth and five-minute Apgar scores.[25] Following the format of the prior table, Table 4 present results for the OLS specification without mother fixed effects that are estimated on the sample of siblings with the same marital status both with and without covariates (*OLS* and *OLS with controls*), followed by the sibling fixed effect specification (*FE*) and the twin pair fixed effect specification (*Twins*).

The first row of the three panels of Table 4 indicates that boys are heavier at birth but do not have a higher rate of birth complications or higher five-minute Apgar scores. There are significant differences in measures of newborn health across family types. The two OLS columns reveal that, even while accounting for differences in parent demographics, children born to married and father-present households appear somewhat healthier than children born to father-absent households. Specifically, these children have significantly higher birthweight and higher five-minute Apgar scores. Interestingly, they also appear slightly *more* likely to have birth complications, which may reflect their greater birthweight.

Of critical importance to our analytic framework, the Table 4 estimates provide little evidence that these cross family-type contrasts in newborn health differ by gender. Gender gaps in birthweight and in five-minute Apgar scores are essentially invariant across family-types; in no case is the interaction between male and family-type sizable or significant. Interestingly, birth complications appear to be very slightly higher among boys than girls in married families (FE specification) or in father-present families (twins specifications).[26]

The fact that we find no family type gradient in the gender gap in health at birth and yet find a significant

[25] We have also examined one-minute Apgar scores, gestational length and congenital disorders with the same conclusions. These results are available upon request.

[26] Since the higher incidence of complications is not driven by greater birthweight in these contrasts (as shown in the top panel), we conclude that, if anything, this pattern works (slightly) against a finding of differential male advantage in later outcomes in married and father-present families.

family type gradient in the gender gap in educational and disciplinary outcomes during elementary and middle school years is supportive of a causal effect of family structure on child development. One may legitimately ask, however, if this test of our identifying assumption has any power—that is, should we take the absence of evidence of a gender gap in neonatal health according to family type as evidence of the absence of such an effect? Our answer is yes. As Table 4 demonstrates, the *main effects* of family structure on measures of at-birth health—birthweight in particular—are large and highly significant, indicating that children born to different family types experience different in-utero conditions that are already manifest at birth. As demonstrated by Figlio et al. (2014), these neonatal health differences affect cognitive development in childhood, and have remarkably consistent causal effects across subgroups from a wide range of family socio-economic statuses. This set of observations suggests that there is considerable scope for family structure to influence the gender gap in neonatal health—since in utero conditions clearly vary systematically across family types—and by implication, to cause a gender gap in childhood development. That we find no such relationships among neonates supports our identifying assumption that the latent developmental gap between gender-discordant siblings is uncorrelated with family structure at birth; if such a correlation were present, it would plausibly be observable in a gender gap in neonatal health.

4 Variation in the Family Gradient Across Demographic Groups

The analysis so far estimates an average causal effect of family structure on the gender gap in educational and behavioral outcomes by interacting family structure with indicators for children's sex.[27] So far, we have not allowed this relationship to vary with other family demographics. We explore these interactions here by estimating the gender gap by family structure for subgroups defined by mother's race, ethnicity, immigrant status, educational attainment, SES, and age at her children's births. This exercise serves two goals. One is to characterize heterogeneity in the causal effect of family structure on the gender gap in developmental outcomes. The second is to test for spurious inference. Since, as documented above, parental demographics differ greatly across family types, it is plausible that the family structure gradient in the gender gap measured above may be partly attributable to these demographic differences rather than to family structure per se (e.g., holding constant family structure, boys of high school dropout mothers fare worse than girls of high school dropout mothers). We explore this channel of causation by estimating the family type gradient in the gender gap separately across a large set of demographic characteristics in Tables A3 through A7, which are succinctly summarized in Figures 3 through 7. We subsequently provide two omnibus tests that pool across demographic

[27] This average is specific to the Florida birth meeting our sample criteria, as described in section 2.2. Since our results below suggest that this effect is heterogeneous across demographic groups, this average effect will differ across populations with different demographic compositions.

characteristics.

4.1 Estimates by Demographic Group

Figure 3 and Table A3 report estimates from our primary sibling fixed-effects and twins models separately for children born to white and black mothers. Sample sizes are considerably reduced by these splits and precision falls accordingly. In general, we do not see a robust family gradient in the gender gap in reading in math scores. This is not the case for behavioral outcomes, however. Among blacks and whites, the gender gap in suspensions is substantially smaller among children born to married families as compared to children born to father-absent families. Results for absences are qualitatively comparable to those for suspensions, but precision is lower.

A central two-part pattern of results, hinted at by the absence rate results, and corroborated by subsequent tables, is: (1) the gender gap in both educational and behavioral outcomes is generally smaller among children from more advantaged families; and (2) among demographic groups where these gaps are smaller, the family structure gradient in these gaps is generally shallower. Thus, for example, the boy-girl gap in school absence rates among father-absent families is 0.62 percentage points among blacks versus 0.30 percentage points among whites (that is, only half as large). Similarly, the reduction in this gap among married relative to father-absent families is 0.37 percentage points among blacks versus 0.16 among whites, which is again half as large (and, in this case, insignificant). This pattern suggests, plausibly, that the protective effect of family marital status on boys' relative outcomes is larger in settings where boys are at a greater overall disadvantage relative to girls.

Figure 4 and Table A4 present estimates separately for Hispanics and Non-Hispanics. Here, we see stronger evidence of a family structure gradient in test scores (math scores in particular). For behavioral outcomes—both suspensions and absences—the gradient is quite pronounced. We find a significant protective effect of marriage on boys relative to girls in six of eight specifications; only among Hispanic twins is this effect statistically insignificant, though the point estimates are nevertheless sizable. Estimates that split on immigrant status rather than ethnicity (Figure 5) are highly comparable to the split on Hispanic versus Non-Hispanic, which is logical given that many Hispanics are immigrants and vice versa.

In Figure 6, we divide families according to mother's education recorded on the birth certificate: fewer than 12 years (dropouts); 12 to 15 years (high school graduates); and 16+ years (college graduates) greater.[28] Consistent with the pattern detected above, gender gaps in behavioral and educational outcomes are substantially larger in less advantaged families. Comparing gender-discordant siblings, we detect a male disadvantage

[28]Since mother's education can change across births, this splitting families based upon education (or, similarly, mother's age at birth) may divide non-twin siblings across bins. A mother fixed effect is included in all models, however, meaning that solo-siblings (that is, siblings who are split across panels) do not contribute to identification.

in test scores in high school dropout and high school graduate families but not in college graduate families. Similarly, we see sizable gender gaps in behavioral outcomes across all family types, but these gaps are smallest in college graduate families and largest in high school dropout families (and never significant for college graduate families). Consistent with the hypothesis that the protective effect of family structure on boys is larger among demographic groups where the boy disadvantage is greater, we estimate that father-presence and marriage reduce the boy-girl gap in suspensions by approximately 0.8 days per year in high school dropout families, approximately 0.6 days per year in high school graduate families, and by approximately 0.38 days per year in college graduate families (the latter of which is insignificant). We see a qualitatively similar pattern for absences, while the evidence for a family structure effect on the gender gap in test scores is weaker. In general, estimates of an effect of family structure on the gender gap are insignificant in the twins sample. Note however that the twins samples are typically only one fourth the size of the siblings samples, and hence when we further subdivide the twins sample on mother's education, power is limited—particularly for high school dropout families, which is both the least common educational group in the data and also the group for which we detect the largest family structure effects.

Assessing the evidence for family structure effects on the gender gap among children born to mothers within different age brackets (ages 21 and below, 22 through 29, 30 through 35, and over 35) in Figure 7, we obtain results that compare closely to those for mother's education: the largest and most robust effects are found for disciplinary outcomes, particularly suspensions; they are much greater for children born to younger mothers, for whom the male disadvantage appears largest; these effects are rarely significant in the twins sample, though the sign pattern of the point estimates is in almost all cases consistent with the hypothesis that father-presence and marriage reduce the male disadvantage in childhood outcomes.

4.2 Accounting for all Demographic Effects Simultaneously

A challenge in interpreting the results in Figures 3 through 7 is that many of the demographic categories along which we split the data are highly correlated (e.g., race, education, maternal age). Hence, it is not obvious how much new information is provided by each split of the data. Moreover, no single split accounts for all observable demographic factors simultaneously. We now provide two omnibus comparisons that resolve these questions.

In the first comparison, we follow Figlio et al. (2014) to construct a single index of mother's SES, which allows us to bin the data and then estimate family structure effects within these bins. The socioeconomic index is constructed by regressing pooled test scores (the mean of math and reading) on mother's race, ethnicity, education level, immigrant status, age at birth and the median income in her ZIP code of residence

recorded on the birth certificate.[29] We divide the predicted values from this regression into terciles and, in Table 5, contrast the gender gap within each tercile. These within-SES estimates prove highly consonant with the findings above.

Table 6 provides a second omnibus test. Here, we augment our main regression specification for estimating gender gap gradient across family structures with a set of interactions between child's gender and the demographic variables: maternal race, ethnicity, immigrant status, and age and education at child's birth. Thus, each of these variables enters the specification in parallel to the family structure interaction terms on which we have been focused. Reassuringly, this stringent test of the main results, which are plotted in Figure 8, produces findings that corroborate all prior results in this section. Relative to their female siblings, boys have fewer suspensions and lower school absence rates in father-present and married households than in father-absent households. These contrasts are robustly significant in the non-twin siblings sample. They are also economically sizable and about half as large in the twins sample, but do not attain statistical significance. After accounting for all demographic by gender interactions, we do not find evidence that family structure exerts a direct effect on the gender gap in reading or math test scores.

5 Pathways of Causation: Neighborhoods and Schools

We have so far been agnostic on the mechanism by which family structure affects the gender gap in childhood outcomes. Our results so far establish that this effect is *not* entirely due to demographic differences across family types, nor is it due to the (large) gap in father-concordance rates between married and unmarried families. These results suggest that causal effect is likely to run through a combination of direct effects of family structure on child-rearing conditions (e.g., availability of parental attention, degree of conflict among family members, etc.) and indirect effects that may stem from the impact of family structure on living arrangements. Indeed, a mixture of quantitative and ethnographic evidence from the landmark Moving to Opportunity study highlights that differences in neighborhood environments may in part explain why inner city boys fared worse when moved to low poverty neighborhoods than did boys in the control (non-movers) group while girls moved to low poverty neighborhoods fared better than controls (Clampet-Lundquist et al., 2011). While it is challenging to entirely pin down these direct versus indirect channels, or even to fully conceptually delineate them, we can nevertheless make some progress.[30]

In this final section, we attempt to assess the importance of two indirect channels of causation: neighborhoods and schools. Exploiting the details available from the Florida data, we ask whether, after accounting

[29] Zip code median incomes are obtained from the 2000 U.S. Census Summary Tabulation Files.

[30] For example, if conflict is higher in single-parent families in part because resources are scarcer, it would be difficult to ascertain whether this a direct or indirect effect of family structure on the child-reading environment.

for differences in the neighborhoods and schools in which children of different family types are raised and educated, the estimated effect of family structure on the gender gap remains. Specifically, we perform an exercise akin to that in Table 6 above, except rather than saturating the model with demographic group by child gender interactions, we saturate it with either neighborhood by child gender or school by child gender interactions. Thus, we allow each Florida neighborhood or public school to have an independent effect on the gender gap. To proxy for neighborhoods, we use indicators for the reported zip code of residence of each mother at the time of each of her children's birth. For schools, we include separate indicators for each school that each child attends in the year corresponding to the test score or behavioral measure used as the outcome variable. This yields a total of $1,418$ unique zip code effects and $3,797$ unique school effects, each interacted with child gender. Tables 7 and 8 present the results. It bears note that these specifications are identical to the sibling fixed-effects and twin estimates contained in our initial estimates in Table 2, except that here we include a full set of neighborhood by gender or school by gender interactions. Thus, the contrast between the estimates in Tables 7 and 8 and those in Table 2 answers the question of how large a mediating role school and neighborhood play in the family structure gradient in the gender gap.

The estimates in Table 7 imply a surprisingly modest role for neighborhood in mediating the family structure effect on the gender gap. In all within-sibling comparisons (the odd-numbered columns), the family structure gradient in the gender gap remains robust to inclusion of neighborhood by sex interactions. Comparing these point estimates to the analogous columns in Table 2 (columns 3 and 6) reveals that the neighborhood effects reduce the impact of marriage or father-presence on the gender gap by approximately ten percent across the board. As established in prior tables, the estimates for twins are less robust than those for non-twin siblings, but accounting for a potential neighborhood by gender channel of causation typically only attenuates the direct effect of family structure on the gender gap by 10 to 20 percent.

Table 8 finds a similarly modest role for schools in mediating the effect of family structure on the gender gap. Indeed, the point estimates in this table barely differ from those in 7. This appears logical since which public schools children attend are largely determined by their residential locations, meaning that both zip-code and school fixed effects are likely to absorb the same sources of intermediating variation.

While neighborhood and school are not the only plausible intermediating channels of causation by which family structure may affect the gender gap, they appear two of the most likely, and also the two most often considered in prior literature.[31] In combination with our earlier estimates for paternity, which also appears to play a minor role, we tentatively conclude that aspects of the household child-rearing environment that differ according to family's marital status and father's engagement with children (proxied by paternity acknowledgment of on the birth certificate) appear the most likely causal channels by which family structure

[31] See for example Chetty et al. (2014).

affects the gender gap in children's developmental outcomes.

6 Conclusion

While female-favorable gender gaps in educational outcomes have been widely documented, the determinants remain the subject of empirical inquiry. The emergence of the gender gap in high school graduation and college attendance has not affected demographic groups in the U.S. uniformly, but rather the gender gap among children of minority and less-educated individuals is substantially larger and has grown more quickly than among children from more advantaged backgrounds. Prompted by the marked shift in U.S. marital patterns and the rising incidence of single-parent headed households – particularly among the same demographic groups that have experienced the largest increases in the gender gap in educational attainment – this paper empirically investigates family structure as a potential explanation for the gender gap in educational outcomes. Specifically, our objective is to isolate whether boys born to unmarried mothers, with either the dad present or absent at the time of birth, are differentially disadvantaged relative to girls in their early- to mid-childhood educational preparedness, as measured by standardized test scores and disciplinary outcomes. We provide a test of the effect of family structure on the gender gap in children's academic and disciplinary outcomes with use of birth certificate data for all children born in Florida from 1992 to 2002 linked to detailed schooling outcomes. Our empirical strategy employs sibling and twin fixed effect models to contrast the outcomes of gender-discordant sibling pairs, while holding constant the biological and home environment differences across families that confound prior analyses, and controlling for a host of other time-varying maternal attributes.

We find robust evidence of a causal effect of family structure on the gender gap in behavioral outcomes, as measured by school suspensions and absences, and weaker evidence for the gender gap in math and reading standardized test scores. In particular, there is a clear rank ordering of the gender gap in disciplinary outcomes by the three family structure categories: the gap is the largest among boy-girl siblings born to an unmarried mother with the father absent at birth, it is attenuated among sibling pairs born to an unmarried mother with the father present at birth, and is even smaller among pairs born to a married mother. The systematic variation of the gender gap in behavioral outcomes with family structure type is present among all demographic groups, but is most pronounced among children born to minority, immigrant and less educated mothers.

From our most stringent regression specification in which we allow maternal characteristics to flexibly vary by child gender, our estimates suggest that the gender gap in days suspended per year between boy-girl siblings born to unmarried mothers without the father present at birth is 1.02 days, while the gap shrinks to 0.61 days when the siblings are born to an unmarried mother with the father present at birth, and is

further attenuated to 0.51 days when the children are born to a married mother. Finally, we explore whether the relationship between the gender gap and family structure is mediated by key environmental factors that are systematically correlated with family structure. Notably, our main conclusions are unaffected by either school or neighborhood quality, lending validity to our interpretation of the estimates as direct effects of family structure through the home environment.

Our robust evidence of the effect of family structure on the gender gap in early- and mid- childhood outcomes, particularly disciplinary infractions, sheds light on the consequences of major demographic transitions in the U.S. which have substantially altered the circumstances into which children are born. Our results suggest that common policy levers, such as improving school and neighborhood quality, may not be efficacious for remediating disparities that largely accrue due to differences in home environments and child-rearing strategies.

References

Aizer, Anna and Janet Currie, "The intergenerational transmission of inequality: Maternal disadvantage and health at birth," *Science*, May 2014, *344* (6186), 856–861.

Autor, David and Melanie Wasserman, "Wayward Sons: The Emerging Gender Gap in Education and Labor Markets," Technical Report, Third Way, Washington, DC 2013.

Bertrand, Marianne and Jessica Pan, "The Trouble with Boys: Social Influences and the Gender Gap in Disruptive Behavior," *American Economic Journal: Applied Economics*, January 2013, *5* (1), 32–64.

Buchman, Claudia and Thomas A DiPrete, "The Growing Female Advantage in College Completion: The Role of Family Background and Academic Achievement," *American Sociological Review*, 2006, *71* (4), 515–541.

Cherlin, Andrew, "Demographic Trends in the United States: A Review of Research in the 2000s.," *Journal of marriage and the family*, June 2010, *72* (3), 403–419.

Cherlin, Andrew J, "American marriage in the early twenty-first century.," *The Future of children / Center for the Future of Children, the David and Lucile Packard Foundation*, January 2005, *15* (2), 33–55.

Chetty, Raj, Nathaniel Hendren, Patrick Kline, and Emmanuel Saez, "Where is the Land of Opportunity? The Geography of Economics Mobility in the United States," 2014.

Clampet-Lundquist, Susan, Jeffrey R Kling, Kathryn Edin, and Greg J Duncan, "Moving teenagers out of high-risk neighborhoods: how girls fare better than boys.," *AJS; American journal of sociology*, January 2011, *116* (4), 1154–89.

Dahl, Gordon B. and Enrico Moretti, "The Demand for Sons," *Review of Economic Studies*, October 2008, *75* (4), 1085–1120.

DiPrete, Thomas A and Claudia Buchmann, "The Secret Behind College Completion," Technical Report, Third Way 2014.

Duckworth, Angela L and Martin E P Seligman, "Self-discipline outdoes IQ in predicting academic performance of adolescents.," *Psychological science*, December 2005, *16* (12), 939–44.

Figlio, David, Jonathan Guryan, Krzysztof Karbownik, and Jeffrey Roth, "The Effects of Poor Neonatal Health on Children's Cognitive Development," 2014.

Fortin, Nicole, Philip Oreopoulos, and Shelley Phipps, "Leaving Boys Behind: Gender Disparities in High Academic Achievement," 2014.

Goldin, Claudia and Lawrence F. Katz, *The Race Between Education and Technology*, Cambridge: Harvard University Press, 2008.

_ , Lawrence F Katz, and Ilyana Kuziemko, "The Homecoming of American College Women: The Reversal of the College Gender Gap," *Journal of Economic Perspectives*, September 2006, *20* (4), 133–156.

McLanahan, Sara and Christine Percheski, "Family Structure and the Reproduction of Inequalities," *Annual Review of Sociology*, August 2008, *34* (1), 257–276.

_ , Laura Tach, and Daniel Schneider, "The Causal Effects of Father Absence," *Annual Review of Sociology*, July 2013, *39* (1), 399–427.

Organization for Economic Cooperation and Development, *Education at a Glance 2013* Education at a Glance, OECD Publishing, June 2013.

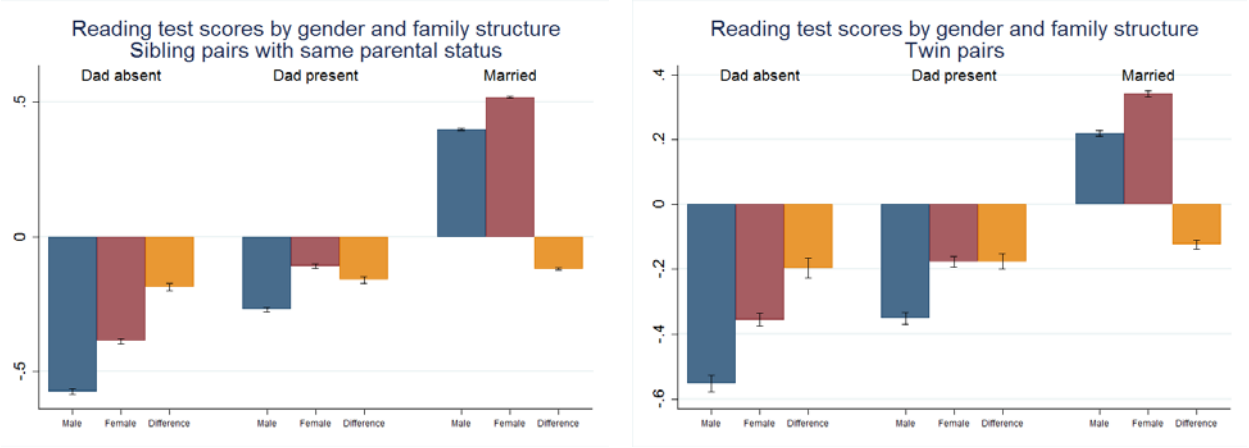
Segal, Carmit, "Misbehavior, Education, and Labor Market Outcomes," *Journal of the European Economic Association*, August 2013, *11* (4), 743–779.

U.S. Bureau of the Census, "Income, Poverty, and Health Insurance Coverage in the United States: 2010 (Report P60-239)," Technical Report 2011.

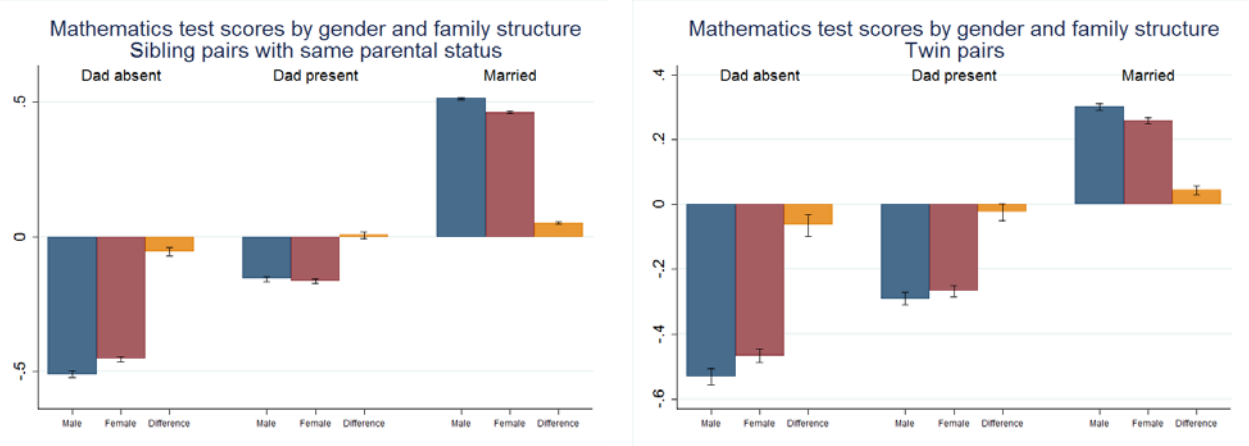
U.S. National Center for Education Statistics, "Digest of Education Statistics 2013," Technical Report, Washington, DC 2013.

Figure 1: Gender Gap in Academic Outcomes by Family Structure

Panel A: Reading scores



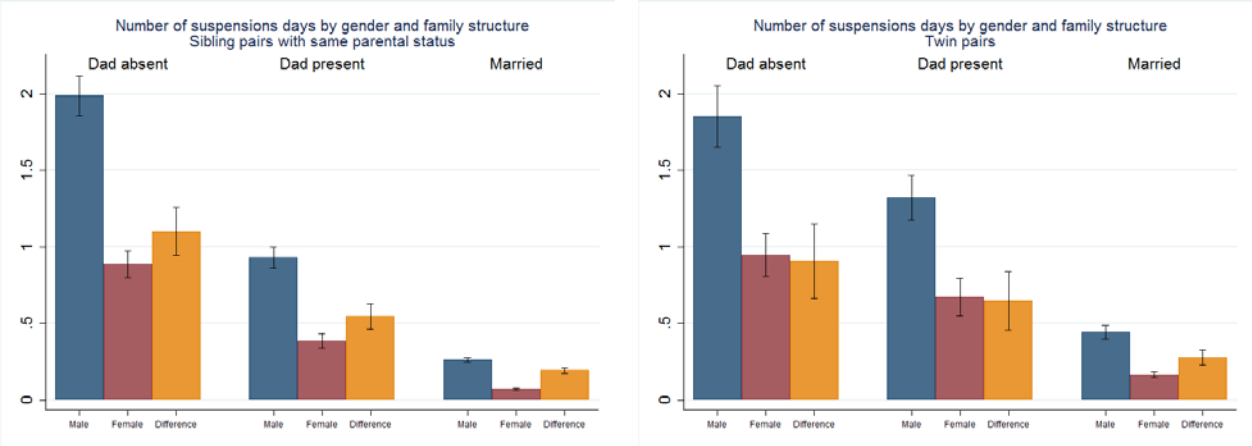
Panel B: Math scores



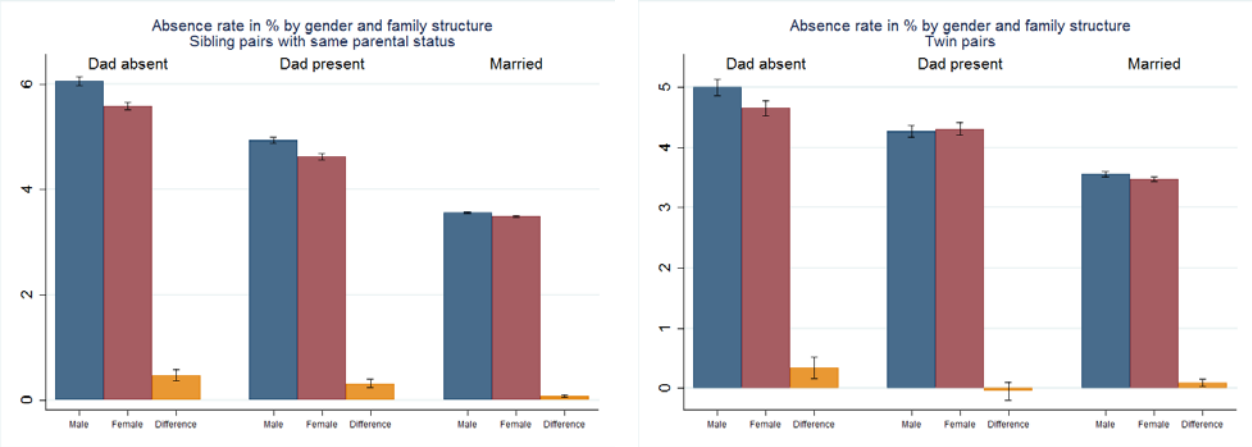
Note: Mean reading and math scores by gender and family structure type are computed from pooled test scores for grades three through eight, with each child contributing up to one observation per grade observed. Figures in the left panel are computed from the sample of non-twin siblings with the same family structure at the time of birth, and figures the right panel are computed from the sample of twins. Confidence intervals are represented by whiskers on each bar.

Figure 2: Gender Gap in Behavioral Outcomes by Family Structure

Panel A: Number of days suspended



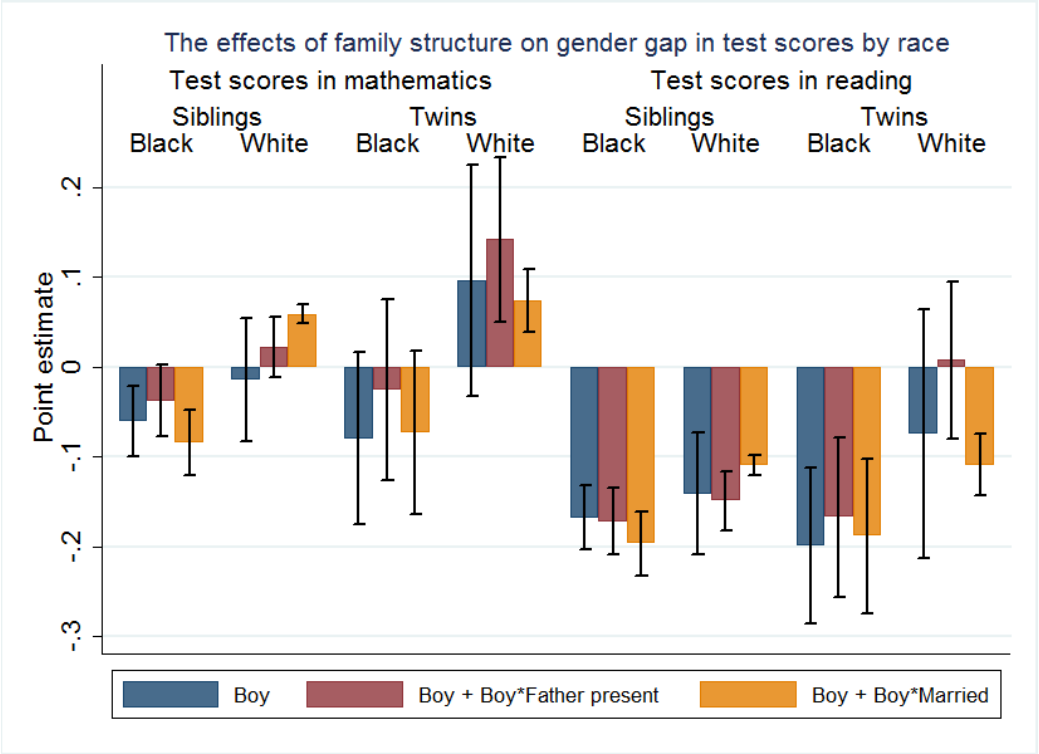
Panel B: Absence rate



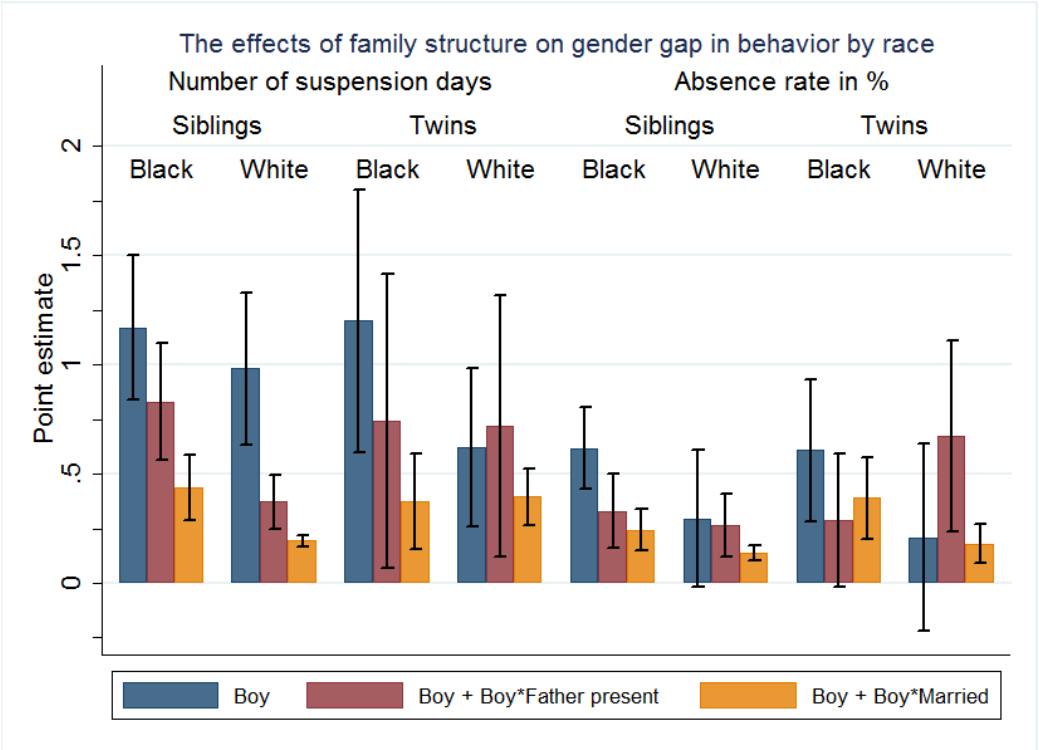
Note: Mean number of days suspended per year and absence rate per year by gender and family structure type are computed from pooled disciplinary data for grades three through eight, with each child contributing up to one observation per grade observed. Confidence intervals are represented by whiskers on each bar.

Figure 3: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, by Race

Panel A: Academic Outcomes



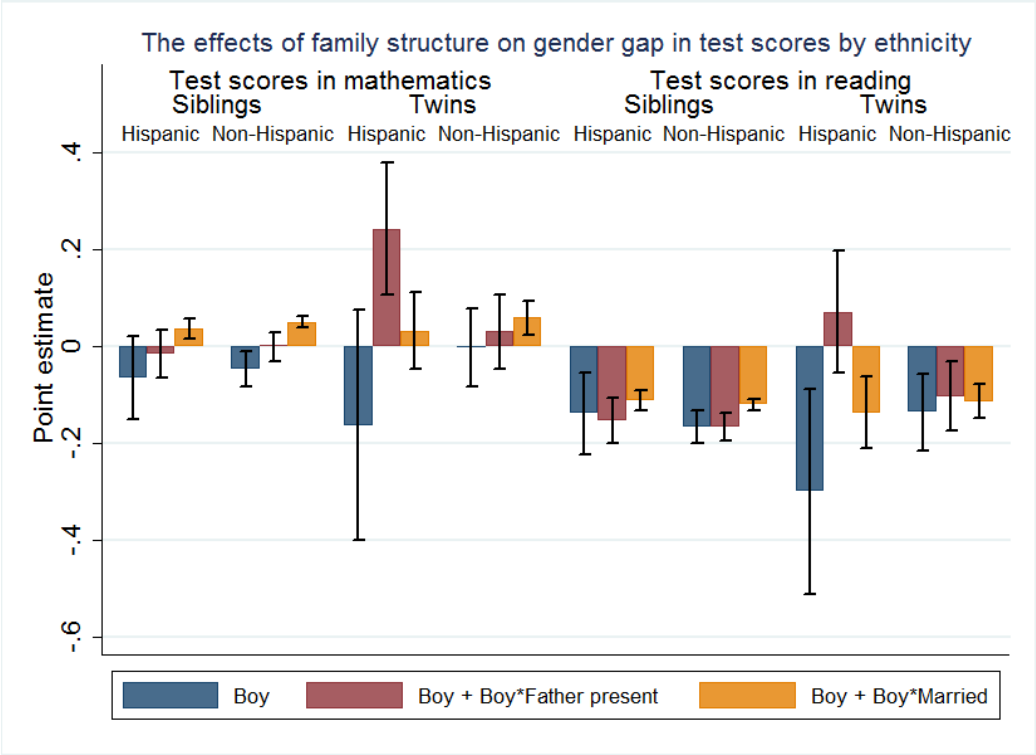
Panel B: Behavioral Outcomes



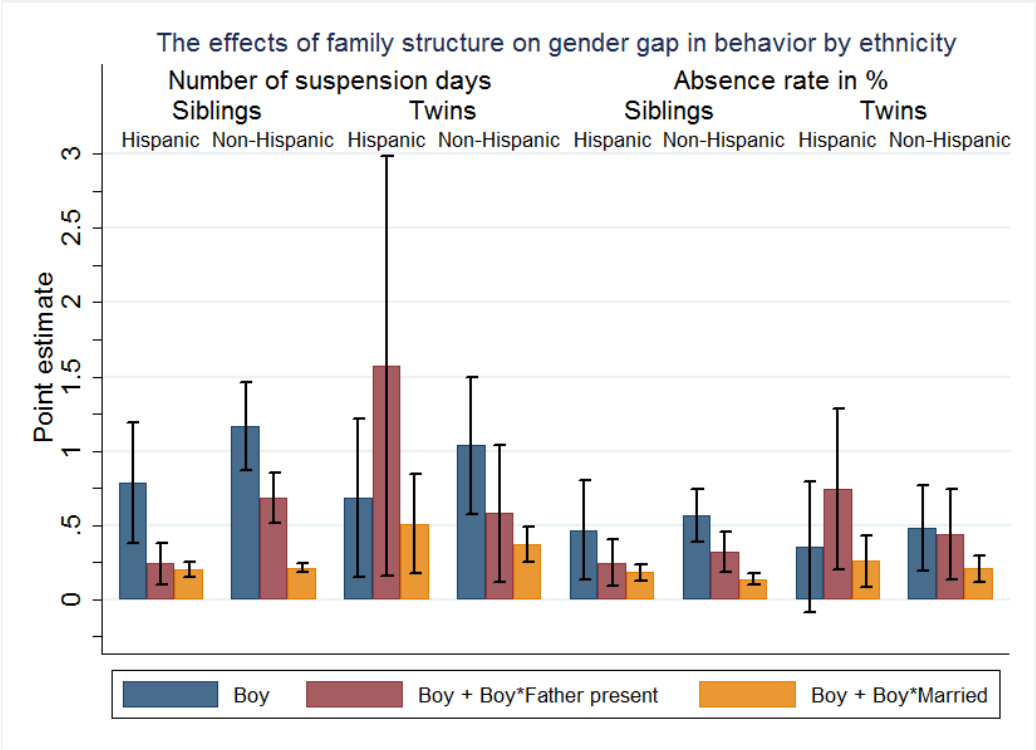
Note: Plot corresponds to estimates in Table A3. Confidence intervals are represented by whiskers on each bar.

Figure 4: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, by Ethnicity

Panel A: Academic Outcomes



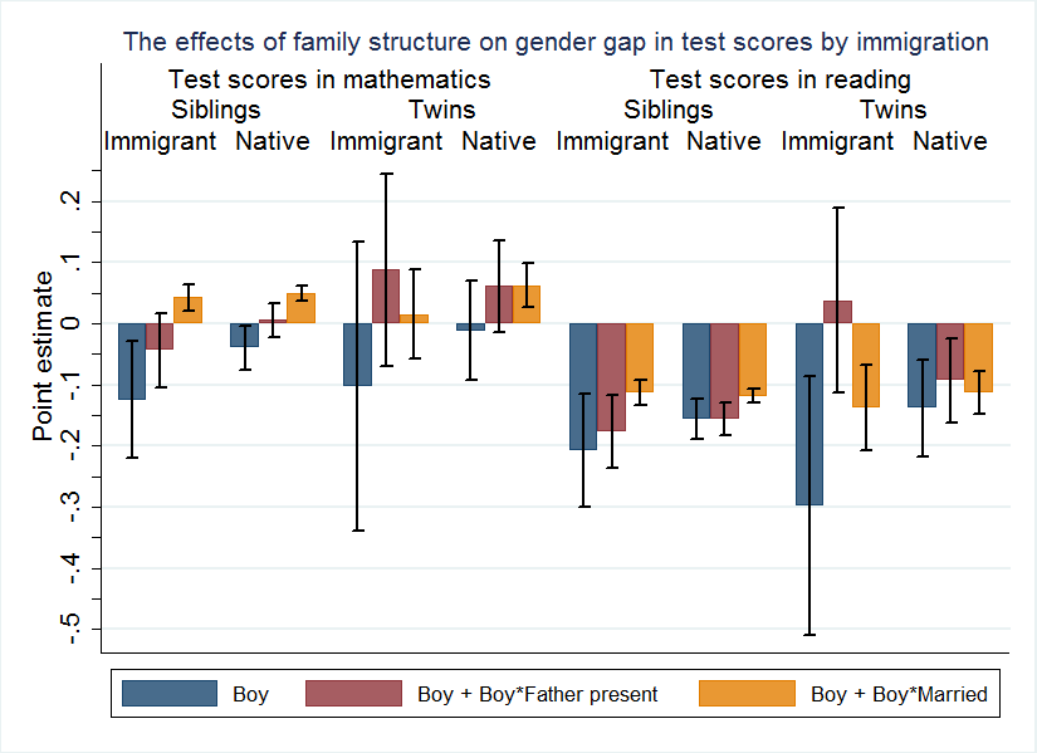
Panel B: Behavioral Outcomes



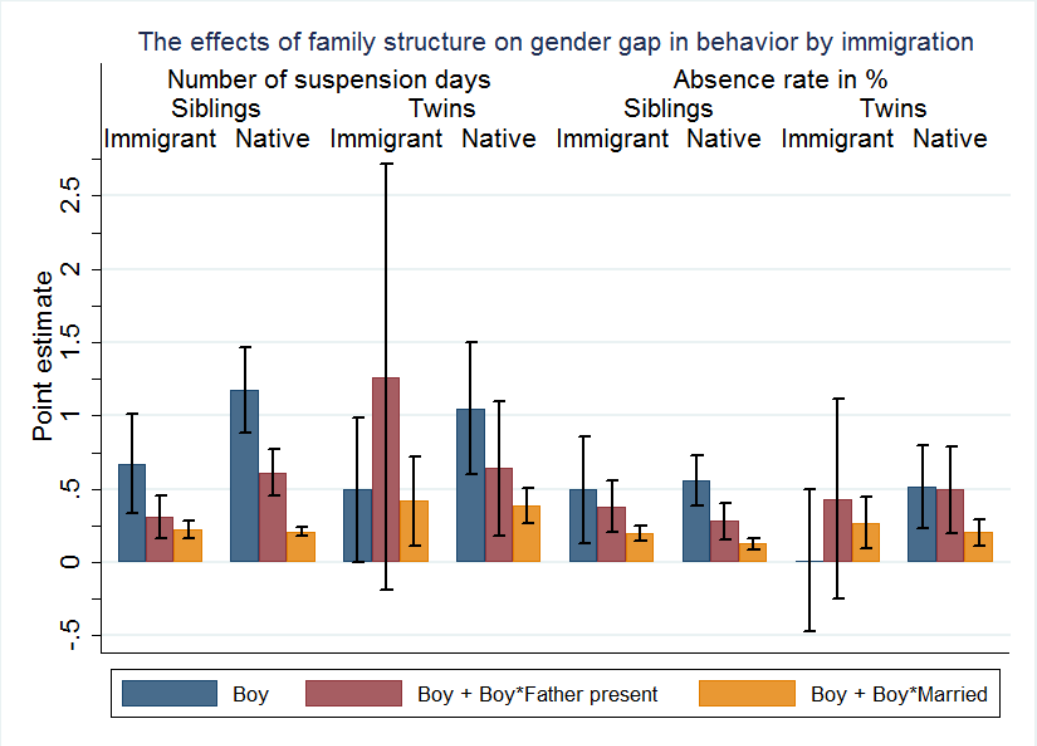
Note: Plot corresponds to estimates in Table A4. Confidence intervals are represented by whiskers on each bar.

Figure 5: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, by Mother's Immigrant Status

Panel A: Academic Outcomes



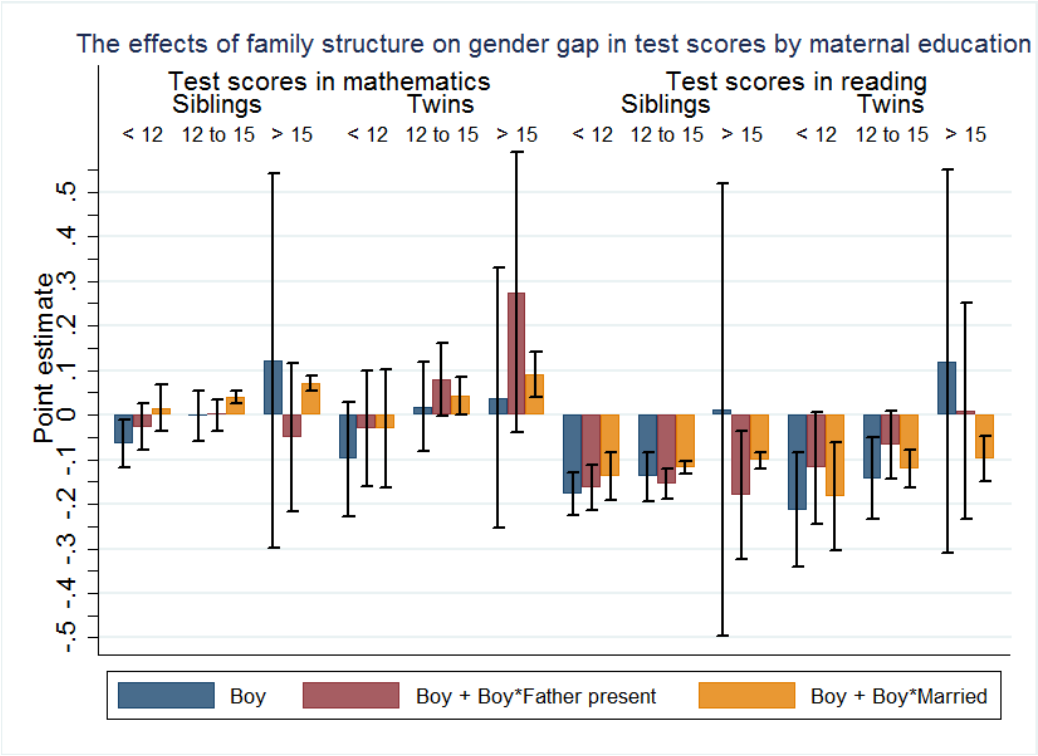
Panel B: Behavioral Outcomes



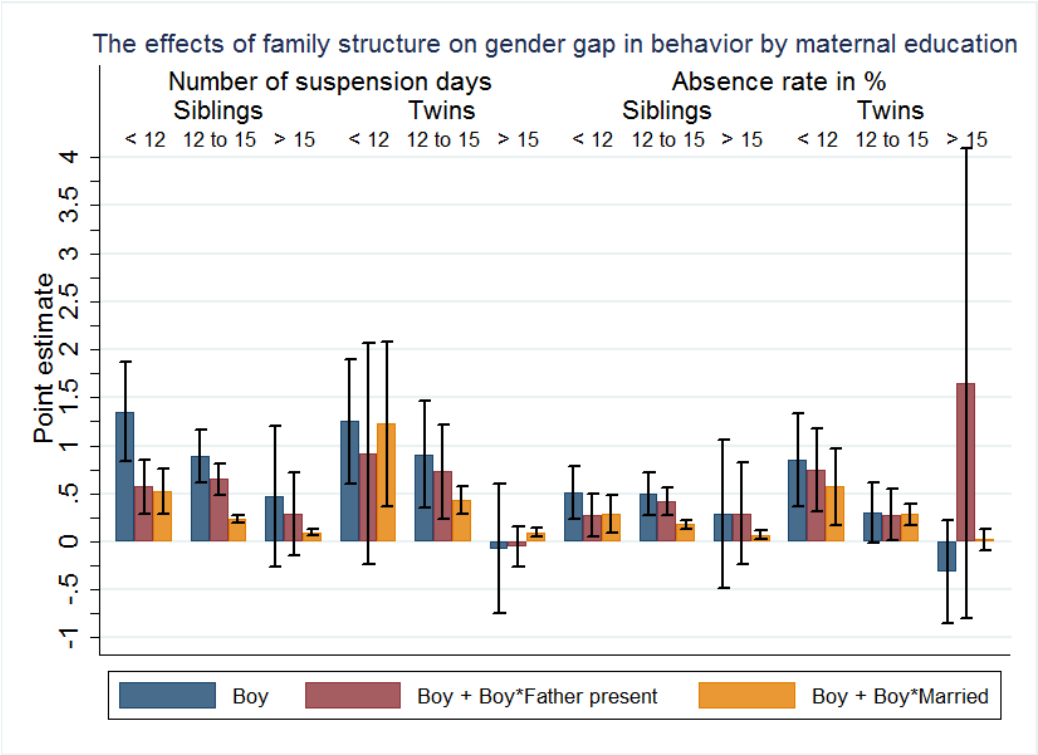
Note: Plot corresponds to estimates in Table A5. Confidence intervals are represented by whiskers on each bar.

Figure 6: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, by Mother's Education Level

Panel A: Academic Outcomes



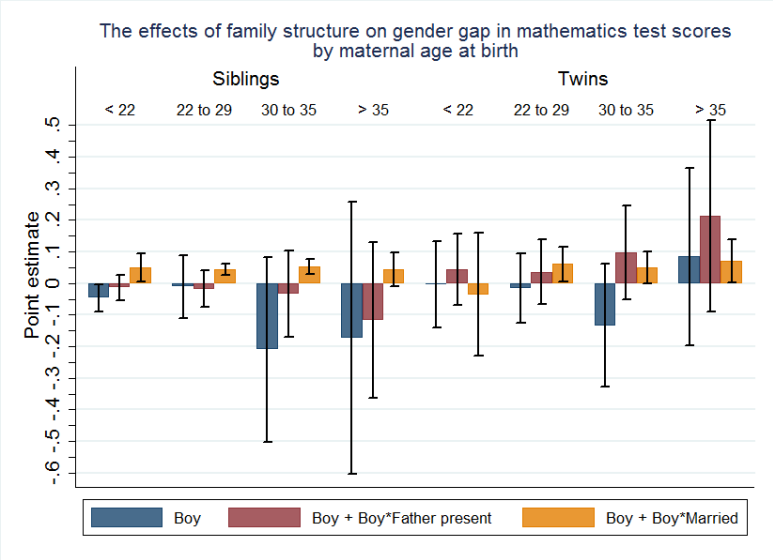
Panel B: Behavioral Outcomes

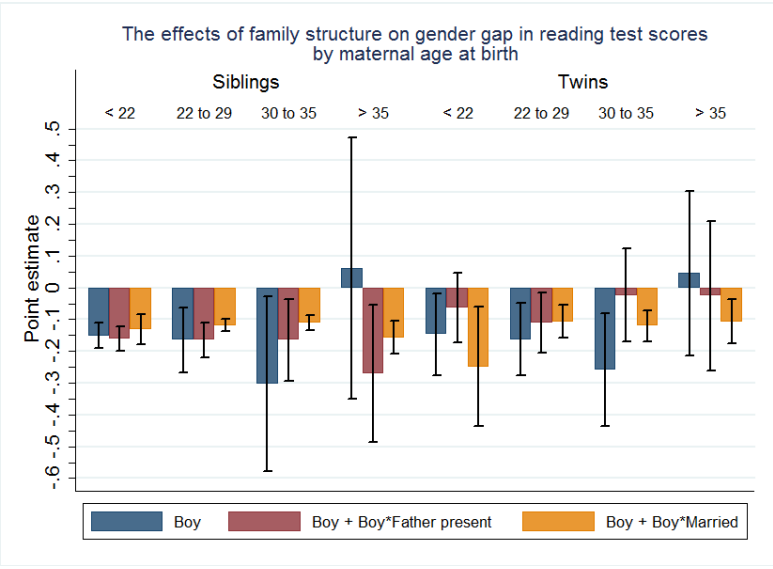


Note: Plot corresponds to estimates in Table A6. Confidence intervals are represented by whiskers on each bar.

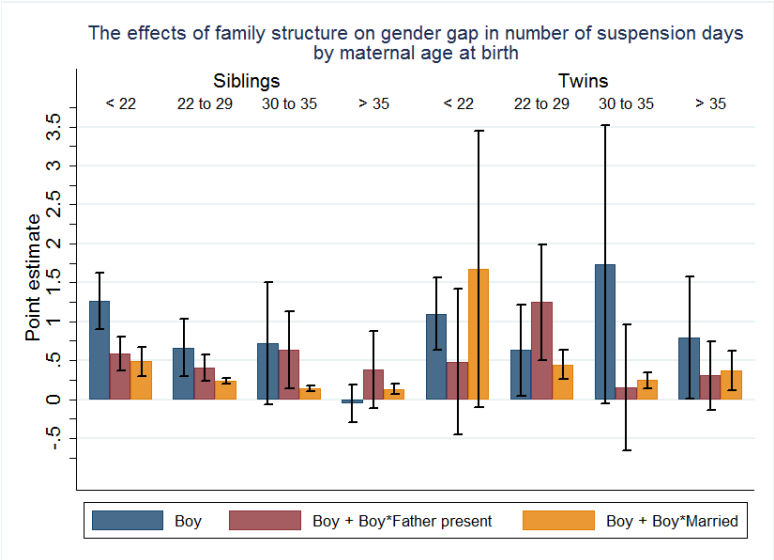
Figure 7: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, by Mother's Age at Birth

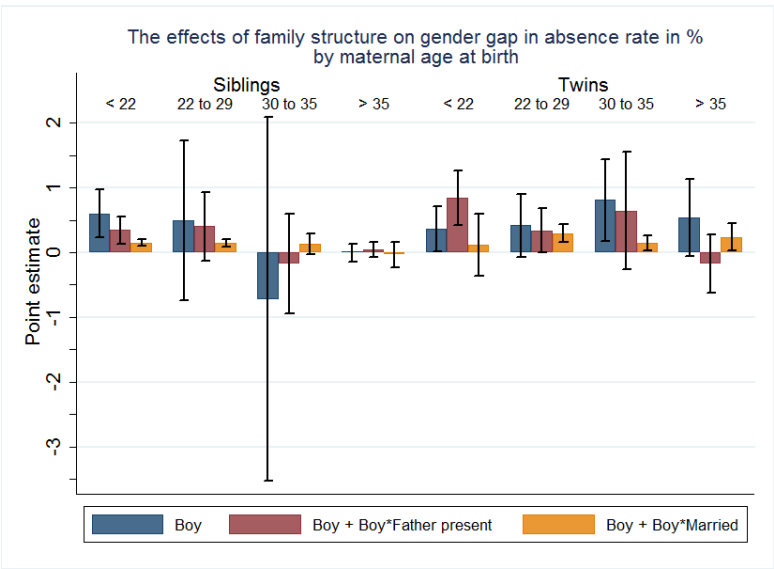
Panel A: Academic Outcomes





Panel B: Behavioral Outcomes

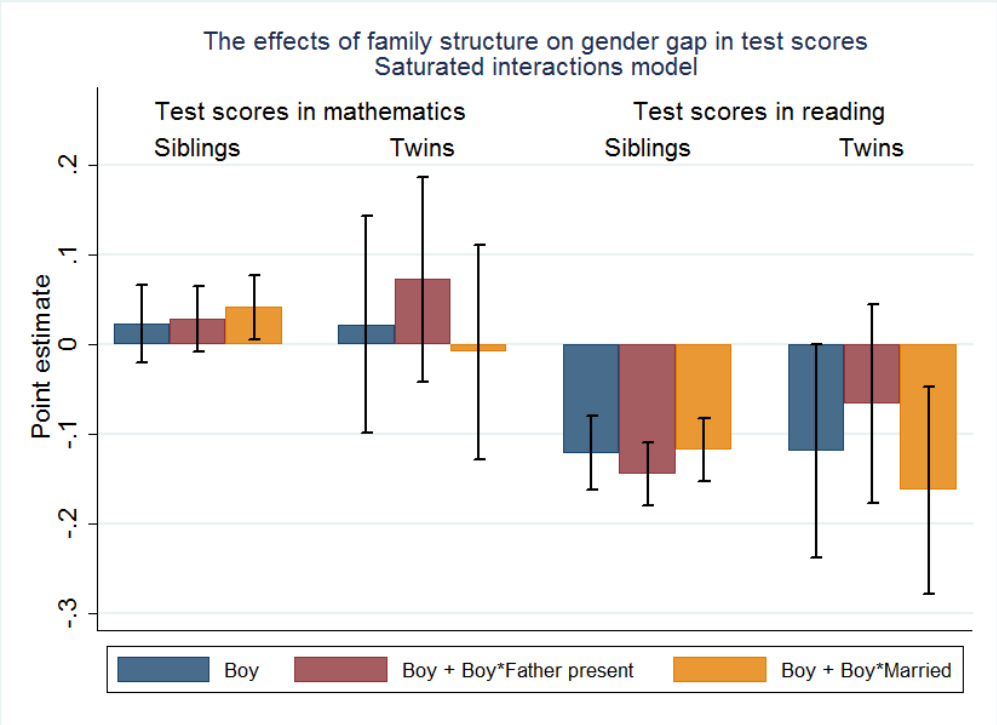




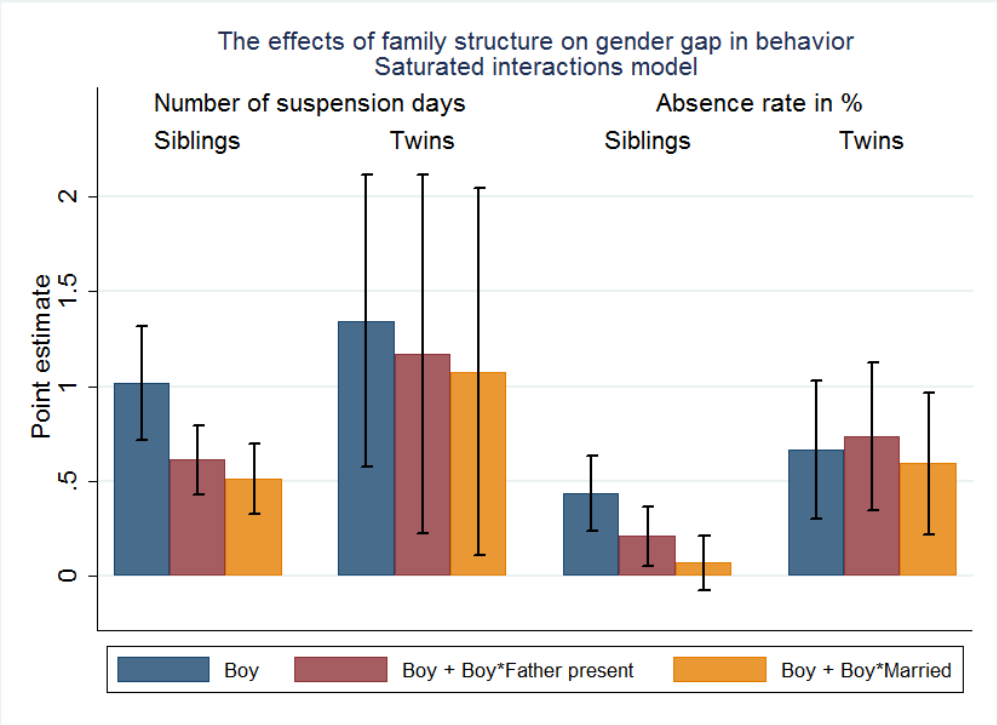
Note: Plot corresponds to estimates in Table A7. Confidence intervals are represented by whiskers on each bar.

Figure 8: Regression Adjusted Gender Gap in Academic and Behavioral Outcomes by Family Structure, Accounting for all Gender by Demographic Group Main Effects and First Order Interactions

Panel A: Academic Outcomes



Panel B: Behavioral Outcomes



Note: Plot corresponds to estimates in Table 6. Estimated within-family, cross-gender contrasts contain a full set of interactions terms between child gender and mother's race, ethnicity, immigrant status, and age and education at child's birth. Confidence intervals are represented by whiskers on each bar.

Table 1: Descriptive Statistics

Variables/Sample	(1) All births	(2) All singletons with outcomes	(3) Column (2) for "sibling" counties	(4) All siblings	(5) Stable siblings	(6) Twins
			Demographic characteristics			
Male	0.512	0.509	0.509	0.512	0.514	0.498
	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Dad absent	0.136	0.152	0.160	0.137	0.093	0.115
	(0.342)	(0.359)	(0.367)	(0.344)	(0.290)	(0.319)
Dad Present	0.212	0.237	0.236	0.208	0.142	0.167
	(0.409)	(0.425)	(0.424)	(0.406)	(0.349)	(0.373)
Married	0.652	0.611	0.604	0.655	0.765	0.718
	(0.476)	(0.488)	(0.489)	(0.475)	(0.424)	(0.450)
Black	0.220	0.252	0.288	0.245	0.197	0.225
	(0.414)	(0.434)	(0.453)	(0.430)	(0.398)	(0.417)
Hispanic	0.228	0.237	0.303	0.248	0.251	0.175
	(0.420)	(0.425)	(0.459)	(0.432)	(0.434)	(0.380)
Immigrant	0.232	0.229	0.291	0.214	0.228	0.181
	(0.422)	(0.420)	(0.454)	(0.410)	(0.419)	(0.385)
Mom high school dropout	0.205	0.232	0.217	0.187	0.145	0.135
	(0.404)	(0.422)	(0.412)	(0.390)	(0.352)	(0.342)
Mom high school grad	0.590	0.609	0.599	0.568	0.554	0.574
	(0.492)	(0.488)	(0.490)	(0.495)	(0.497)	(0.494)
Mom college grad	0.205	0.159	0.184	0.245	0.301	0.291
	(0.404)	(0.366)	(0.387)	(0.430)	(0.459)	(0.454)
Maternal age	27.132	26.69	27.053	25.672	26.734	29.082
	(6.167)	(6.203)	(6.266)	(5.920)	(5.825)	(6.098)
			Birth outcomes			
Infant birth weight	3316.33	3328.376	3316.167	3328.051	3350.286	2402.388
	(580.142)	(559.095)	(559.715)	(539.233)	(532.622)	(572.272)
Complications at birth	0.058	0.056	0.045	0.048	0.046	0.144
	(0.234)	(0.231)	(0.207)	(0.213)	(0.210)	(0.351)
N	1986116	1313428	733423	219974	170346	37738
Five minutes Apgar score	8.963	8.968	8.951	8.958	8.962	8.74
	(0.572)	(0.559)	(0.526)	(0.507)	(0.496)	(0.774)
N	1982897	1311510	732524	219823	170248	37697
			Outcome variables			
Pooled averaged test scores		0.079	0.085	0.196	0.290	0.083
		(0.887)	(0.891)	(0.896)	(0.890)	(0.926)
N		6269850	3407355	762408	593668	113484
Pooled math test scores		0.077	0.089	0.208	0.304	0.079
		(0.956)	(0.963)	(0.966)	(0.958)	(1.000)
N		6219002	3374259	753208	586360	112236
Pooled reading test scores		0.082	0.082	0.185	0.276	0.089
		(0.952)	(0.954)	(0.956)	(0.953)	(0.982)
N		6254355	3399331	759568	591732	113082
Number of days suspended		0.909	0.747	0.429	0.359	0.563
		(6.029)	(4.962)	(3.640)	(3.279)	(4.001)
N		4424873	2264113	380466	301058	73778
Rate of absence (%)		4.908	4.650	4.143	3.924	3.825
		(5.488)	(5.213)	(4.643)	(4.413)	(4.488)
N		5629886	3098324	653716	510574	103242

Note: Means and standard deviations are reported for the samples indicated in the column titles. Column (1) is the full universe of Florida births from the public-use Vital Statistics natality data for years 1992 through 2002; column (2) is all non-multiple births from column (1) that were linked to subsequent schooling outcomes; the column (3) is the subset of children who were born 1994-2002 in counties in which information on siblings is available; column (4) further limits the sample to first and second non-multiple births to the same mother in counties where sibling information is available; column (5) uses a sample further restricted to siblings who have the same family structure at the time of birth (married-married; dad present-dad present; dad absent-dad absent); column (6) is all twins in Florida born between years 1992 and 2002 that were matched to school records. "Father absent" refers to births where the mother is unmarried and the father does not appear on the birth certificate, "Father present" refers to births where the mother is unmarried but the father appears on the birth certificate, "Married" refers to births where the mother is married. "Immigrant" denotes that the mother was born in a country other than the U.S. Infant birth weight is measured in grams; "Complications at birth" is an indicator variable representing whether there are complications at birth; Five Minute Apgar score is measured on a scale ranging from 0 to 10. All school outcome variables are for grades three through eight, with each matched child contributing up to one observation per grade observed. "Pooled average test scores" average math and reading test scores.

Table 2: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Academic Outcomes								
	Mathematics				Reading			
	OLS	OLS with controls	FE	Twins	OLS	OLS with controls	FE	Twins
Male	-0.056***	-0.061***	-0.049***	-0.020	-0.186***	-0.191***	-0.162***	-0.153***
	(0.015)	(0.014)	(0.017)	(0.039)	(0.014)	(0.014)	(0.016)	(0.038)
Male*Father present	0.062***	0.054***	0.044**	0.085	0.026	0.018	0.001	0.079
	(0.019)	(0.018)	(0.022)	(0.052)	(0.018)	(0.017)	(0.020)	(0.050)
Male*Married	0.107***	0.104***	0.096***	0.075*	0.067***	0.064***	0.044***	0.036
	(0.016)	(0.015)	(0.018)	(0.042)	(0.015)	(0.014)	(0.017)	(0.041)
Father present	0.289***	0.093***			0.277***	0.095***		
	(0.014)	(0.014)			(0.013)	(0.013)		
Married	0.915***	0.242***			0.903***	0.266***		
	(0.012)	(0.014)			(0.011)	(0.013)		
N		586,360		112,236		591,732		113,082
Panel B: Disciplinary Outcomes								
	Number of days suspended				Absence rate in %			
	OLS	OLS with controls	FE	Twins	OLS	OLS with controls	FE	Twins
Male	1.102***	1.101***	1.119***	0.993***	0.467***	0.455***	0.551***	0.467***
	(0.097)	(0.095)	(0.133)	(0.209)	(0.087)	(0.085)	(0.081)	(0.132)
Male*Father present	-0.557***	-0.545***	-0.563***	-0.260	-0.155	-0.123	-0.249**	0.019
	(0.111)	(0.109)	(0.149)	(0.309)	(0.107)	(0.104)	(0.099)	(0.190)
Male*Married	-0.908***	-0.901***	-0.907***	-0.601***	-0.397***	-0.381***	-0.404***	-0.252*
	(0.097)	(0.096)	(0.134)	(0.217)	(0.089)	(0.087)	(0.083)	(0.138)
Father present	-0.500***	-0.219***			-0.967***	-0.847***		
	(0.066)	(0.061)			(0.084)	(0.083)		
Married	-0.816***	-0.161***			-2.100***	-1.233***		
	(0.060)	(0.054)			(0.071)	(0.079)		
N		301,058		73,778		510,574		103,242

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Columns "OLS," "OLS with controls," and "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth. Columns titled "OLS" present results from OLS specifications with no control variables. Columns titled "OLS with controls" present results from OLS specifications with the following control variables in addition to the variables listed: mother's age, race, ethnicity, education level, and immigrant status, and indicators for the child's birth order and month and year of birth.. "FE" columns present results that additionally control for mother fixed effects. Columns titled "Twins" present estimates on the twins sample and include mother fixed effects. Standard errors are clustered at mother level. Sibling and twin FE models additionally control for: mother's age at birth, mother's education level, and indicators for the child's birth order and month and year of birth.

Table 3: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, Robustness to Different Fathers

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Mathematics		Reading		Number of days suspended		Absence rate in %	
	Same dad	Different dad	Same dad	Different dad	Same dad	Different dad	Same dad	Different dad
Male	0.013	-0.050*	-0.150***	-0.187***	0.528***	0.627***	0.322***	0.249**
	(0.015)	(0.025)	(0.014)	(0.025)	(0.076)	(0.128)	(0.062)	(0.116)
Male*Married	0.034**	0.097***	0.032**	0.091***	-0.316***	-0.485***	-0.175***	-0.085
	(0.016)	(0.035)	(0.015)	(0.034)	(0.077)	(0.153)	(0.064)	(0.139)
N	477,230	48,596	481,982	48,950	246,288	22,434	415,486	41,372

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. The sample of non-twin siblings with the same family structure at the time of birth is divided into sibling pairs with the same father and sibling pairs with different fathers. All specifications control for mother fixed effects and include the following control variables: mother's age at birth, mother's education level, and indicators for the child's birth order and month and year of birth. Standard errors are clustered at mother level.

Table 4: Relationship between the Gender Gap in At-Birth Health and Family Structure at Birth

	(1) OLS	(2) OLS with controls	(3) FE	(4) Twins
Panel A: Infant birth weight				
Male	105.716***	104.103***	116.867***	102.335***
	(8.801)	(8.739)	(9.506)	(13.244)
Male*Father present	11.831	13.912	1.375	4.167
	(11.261)	(11.127)	(12.059)	(18.264)
Male*Married	10.991	11.932	8.347	20.727
	(9.260)	(9.193)	(9.961)	(14.461)
Father present	93.338***	24.468***		
	(8.644)	(8.785)		
Married	239.140***	74.861***		
	(7.096)	(8.475)		
N		170,346		37,738
Panel B: Complications at birth				
Male	-0.001	-0.000	-0.005	-0.006
	(0.003)	(0.003)	(0.005)	(0.007)
Male*Father present	0.003	0.003	0.008	0.020*
	(0.004)	(0.004)	(0.006)	(0.010)
Male*Married	0.007**	0.007**	0.012**	0.011
	(0.004)	(0.004)	(0.005)	(0.008)
Father present	0.005	0.011***		
	(0.003)	(0.003)		
Married	-0.003	0.007**		
	(0.003)	(0.003)		
N		170,346		37,738
Panel C: Five minutes Apgar score				
Male	-0.008	-0.009	0.001	0.018
	(0.009)	(0.009)	(0.013)	(0.032)
Male*Father present	-0.005	-0.004	-0.024	0.031
	(0.011)	(0.011)	(0.016)	(0.041)
Male*Married	-0.003	-0.003	-0.017	-0.001
	(0.009)	(0.009)	(0.013)	(0.034)
Father present	0.035***	0.016*		
	(0.008)	(0.008)		
Married	0.060***	0.029***		
	(0.007)	(0.008)		
N		170,248		37,697

Note: Each column presents the results from a separate regression. Columns (1), (2) and (3) are estimated on the sample of non-twin siblings with the same family structure at the time of birth. Column (1) presents from OLS specifications with no control variables. Column (2) presents OLS results with the following control variables in addition to the variables listed: Column (3) presents results for a specification that additionally includes mother fixed effects and controls for: mother's age at birth, mother's education level, and indicators for the child's birth order and month and year of birth. Column (4) uses the twins sample, includes mother fixed effects and controls for child birth order. Standard errors are clustered at mother level.

Table 5: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Socioeconomic Status Index

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Mathematics						
		Sibling FE			Twins	
	Bottom	Middle	Top	Bottom	Middle	Top
Male	-0.051***	-0.045	0.321	-0.085*	0.229***	0.078
	(0.019)	(0.056)	(0.305)	(0.044)	(0.089)	(0.175)
Male*Dad present	0.037	0.069	-0.366	0.093	-0.074	0.171
	(0.025)	(0.062)	(0.315)	(0.061)	(0.107)	(0.263)
Male*Married	0.024	0.091	-0.251	0.067	-0.173*	0.016
	(0.026)	(0.057)	(0.305)	(0.060)	(0.092)	(0.177)
N	155,472	226,108	204,780	41,808	42,578	27,850
Panel B: Reading						
		Sibling FE			Twins	
	Bottom	Middle	Top	Bottom	Middle	Top
Male	-0.155***	-0.234***	0.424	-0.208***	0.021	0.100
	(0.017)	(0.057)	(0.335)	(0.041)	(0.092)	(0.270)
Male*Dad present	-0.002	0.081	-0.663*	0.111**	-0.062	-0.079
	(0.024)	(0.063)	(0.342)	(0.056)	(0.110)	(0.317)
Male*Married	-0.014	0.119**	-0.524	0.048	-0.144	-0.188
	(0.025)	(0.058)	(0.335)	(0.055)	(0.095)	(0.271)
N	156,368	228,023	207,341	42,054	42,888	28,140
Panel C: Number of days suspended						
		Sibling FE			Twins	
	Bottom	Middle	Top	Bottom	Middle	Top
Male	1.186***	0.738***	1.061	1.162***	0.439	0.013
	(0.153)	(0.270)	(1.115)	(0.255)	(0.302)	(0.353)
Male*Dad present	-0.486***	-0.296	-0.845	-0.446	0.421	0.009
	(0.184)	(0.279)	(1.130)	(0.419)	(0.393)	(0.360)
Male*Married	-0.750***	-0.478*	-0.967	-0.372	-0.032	0.106
	(0.168)	(0.271)	(1.115)	(0.320)	(0.314)	(0.354)
N	80,235	117,633	103,190	27,476	28,932	17,370
Panel D: Absence rate in %						
		Sibling FE			Twins	
	Bottom	Middle	Top	Bottom	Middle	Top
Male	0.580***	0.332	-0.414	0.615***	0.020	-0.468
	(0.091)	(0.204)	(0.470)	(0.152)	(0.298)	(0.307)
Male*Dad present	-0.269**	0.049	0.736	-0.188	0.439	1.912
	(0.117)	(0.231)	(0.517)	(0.205)	(0.350)	(1.823)
Male*Married	-0.328***	-0.150	0.500	-0.179	0.225	0.514
	(0.107)	(0.206)	(0.471)	(0.182)	(0.305)	(0.313)
N	136,188	197,228	177,158	38,544	39,434	25,264

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for three terciles of a constructed socioeconomic index. The socioeconomic index is constructed by regressing pooled test scores on the following maternal demographic characteristics: race, ethnicity, education level, immigrant status, age at birth and the average income in her ZIP code of residence. The predicted values from this regression are then divided into terciles corresponding to the categories: bottom, middle, and top. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.

Table 6: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, Accounting for All Demographic by Gender Interactions

	(1)	(2)	(3)	(4)
Panel A: Academic Outcomes				
	Mathematics		Reading	
	FE	Twins	FE	Twins
Male	0.023	0.022	-0.121***	-0.119**
	(0.022)	(0.062)	(0.021)	(0.061)
Male*Father present	0.006	0.050	-0.023	0.053
	(0.022)	(0.052)	(0.021)	(0.050)
Male*Married	0.019	-0.030	0.004	-0.044
	(0.023)	(0.050)	(0.022)	(0.049)
N	586,360	112,236	591,732	113,082
Panel B: Disciplinary Outcomes				
	Number of days suspended		Absence rate in %	
	FE	Twins	FE	Twins
Male	1.019***	1.346***	0.435***	0.665***
	(0.153)	(0.392)	(0.101)	(0.187)
Male*Father present	-0.405***	-0.176	-0.223**	0.072
	(0.144)	(0.314)	(0.100)	(0.186)
Male*Married	-0.506***	-0.268	-0.365***	-0.072
	(0.124)	(0.260)	(0.093)	(0.164)
N	301,058	73,778	510,574	103,242

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects, child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level. In all specifications, we further control for interactions of male and the following maternal demographic characteristics: age at birth, education level, race, ethnicity and immigrant status.

Table 7: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, Neighborhood Quality as a Mediator

	(1)	(2)	(3)	(4)
Panel A: Academic Outcomes				
	Mathematics		Reading	
	Siblings	Twins	Siblings	Twins
Male*Dad present	0.043*	0.110**	-0.000	0.071
	(0.022)	(0.054)	(0.021)	(0.051)
Male*Married	0.087***	0.064	0.031*	0.005
	(0.020)	(0.047)	(0.019)	(0.046)
N	585,928	112,114	591,296	112,958
Panel B: Disciplinary Outcomes				
	Number of days suspended		Absence rate in %	
	Siblings	Twins	Siblings	Twins
Male*Dad present	-0.585***	-0.224	-0.249**	0.103
	(0.149)	(0.266)	(0.099)	(0.206)
Male*Married	-0.814***	-0.546***	-0.369***	-0.045
	(0.139)	(0.202)	(0.089)	(0.158)
N	300,864	73,694	510,206	103,124

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. In addition, the child ZIP code of residence fixed effects and the interaction of the male indicator and ZIP code fixed effects are included. Standard errors are clustered at mother level.

Table 8: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, School Quality as a Mediator

	(1)	(2)	(3)	(4)
Panel A: Academic Outcomes				
	Mathematics		Reading	
	Siblings	Twins	Siblings	Twins
Male*Dad present	0.042**	0.087*	0.003	0.059
	(0.021)	(0.051)	(0.020)	(0.049)
Male*Married	0.085***	0.040	0.037**	-0.009
	(0.018)	(0.043)	(0.017)	(0.043)
N	586,360	112,236	591,732	113,082
Panel B: Disciplinary Outcomes				
	Number of days suspended		Absence rate in %	
	Siblings	Twins	Siblings	Twins
Male*Dad present	-0.504***	-0.407	-0.227**	0.076
	(0.138)	(0.275)	(0.095)	(0.181)
Male*Married	-0.756***	-0.709***	-0.347***	-0.010
	(0.127)	(0.234)	(0.082)	(0.158)
N	301,058	73,734	510,574	102,934

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. In addition, the school of attendance fixed effects and the interaction of the male indicator and school of attendance fixed effects are included. Standard errors are clustered at mother level.

Table A1: Descriptive Statistics, by Family Structure Type

	(1)	(2)	(3)	(4)	(5)	(6)
		Siblings			Twins	
	Married	Dad present	Dad absent	Married	Dad present	Dad absent
		Panel A: Parent demographics				
Black	0.092	0.406	0.740	0.118	0.416	0.616
	(0.289)	(0.491)	(0.439)	(0.322)	(0.493)	(0.486)
Hispanic	0.261	0.274	0.140	0.175	0.192	0.151
	(0.439)	(0.446)	(0.347)	(0.380)	(0.394)	(0.358)
Immigrant	0.251	0.180	0.111	0.192	0.159	0.144
	(0.434)	(0.384)	(0.314)	(0.394)	(0.366)	(0.351)
Mom high school dropout	0.060	0.351	0.534	0.071	0.269	0.343
	(0.237)	(0.477)	(0.499)	(0.256)	(0.444)	(0.475)
Mom high school grad	0.556	0.608	0.454	0.546	0.670	0.611
	(0.497)	(0.488)	(0.498)	(0.498)	(0.470)	(0.488)
Mom college grad	0.384	0.041	0.012	0.383	0.061	0.047
	(0.486)	(0.198)	(0.111)	(0.486)	(0.240)	(0.211)
		Panel B: Birth outcomes				
Male	0.517	0.507	0.505	0.502	0.498	0.473
	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.499)
Infant birth weight	3393.961	3247.388	3147.915	2440.534	2325.461	2275.395
	(518.504)	(548.813)	(553.451)	(559.18)	(590.075)	(597.053)
N	130344	24152	15850	27112	6296	4330
Five minutes Apgar score	8.972	8.945	8.913	8.762	8.698	8.662
	(0.481)	(0.523)	(0.570)	(0.737)	(0.836)	(0.890)
N	130302	24134	15812	27091	6291	4315
		Panel C: Outcome variables				
Pooled averaged test scores	0.472	-0.176	-0.482	0.280	-0.271	-0.476
	(0.826)	(0.805)	(0.835)	(0.876)	(0.855)	(0.908)
N	453778	82878	57012	78296	20716	14472
Pooled math scores	0.488	-0.160	-0.48	0.279	-0.279	-0.497
	(0.893)	(0.886)	(0.935)	(0.944)	(0.938)	(1.015)
N	448126	82004	56230	77464	20496	14276
Pooled reading test scores	0.455	-0.190	-0.478	0.281	-0.261	-0.447
	(0.897)	(0.871)	(0.899)	(0.940)	(0.915)	(0.958)
N	452780	82466	56486	78092	20628	14362
Number of days suspended	0.170	0.663	1.430	0.303	0.981	1.377
	(2.123)	(4.255)	(6.924)	(2.887)	(5.460)	(6.079)
N	232484	38392	30182	51224	12804	9750
Rate of absence (%)	3.527	4.782	5.823	3.518	4.290	4.815
	(3.865)	(5.222)	(6.176)	(4.056)	(5.111)	(5.454)
N	390782	69806	49986	71304	18604	13334

Note: Means and standard deviations are reported for the samples indicated in the column titles, by family structure type. See Table 1 for variable definitions.

Table A2: Boy-Girl Mean Differences, by Family Structure Type

	(1)	(2)	(3)	(4)	(5)	(6)
		Siblings			Twins	
	Married	Dad present	Dad absent	Married	Dad present	Dad absent
			Panel A: Parent demographics			
Black	-0.005	-0.004	-0.013	-0.004	0.011	-0.022
	(0.002)	(0.006)	(0.007)	(0.004)	(0.012)	(0.015)
Hispanic	-0.002	0.006	0.009	0.007	-0.011	0.004
	(0.002)	(0.006)	(0.006)	(0.005)	(0.010)	(0.011)
Immigrant	-0.005	0.007	0.004	0.010	0.001	-0.006
	(0.002)	(0.005)	(0.005)	(0.005)	(0.009)	(0.011)
Mom high school dropout	-0.001	0.000	0.003	0.006	-0.004	0.002
	(0.001)	(0.006)	(0.008)	(0.003)	(0.011)	(0.014)
Mom high school grad	-0.003	0.002	-0.003	-0.001	-0.004	-0.006
	(0.003)	(0.006)	(0.008)	(0.006)	(0.012)	(0.015)
Mom college grad	0.004	-0.001	0.000	-0.005	0.007	0.004
	(0.003)	(0.003)	(0.002)	(0.006)	(0.006)	(0.006)
			Panel B: Birth outcomes			
Infant birth weight	116.707	117.547	105.716	95.712	91.405	89.702
	(2.85)	(7.019)	(8.749)	(6.766)	(14.832)	(18.136)
N	130344	24152	15850	27112	6296	4330
Five minutes Apgar score	-0.012	-0.014	-0.008	-0.005	0.008	-0.010
	(0.003)	(0.007)	(0.009)	(0.009)	(0.021)	(0.027)
N	130302	24134	15812	27091	6291	4315
			Panel C: Outcome variables			
Pooled average test scores	-0.035	-0.077	-0.123	-0.041	-0.101	-0.134
	(0.002)	(0.006)	(0.007)	(0.006)	(0.012)	(0.015)
N	453778	82878	57012	78296	20716	14472
Pooled math score	0.051	0.006	-0.056	0.044	-0.024	-0.065
	(0.003)	(0.006)	(0.008)	(0.007)	(0.013)	(0.017)
N	448126	82004	56230	77464	20496	14276
Pooled reading score	-0.119	-0.16	-0.186	-0.124	-0.176	-0.197
	(0.003)	(0.006)	(0.008)	(0.007)	(0.013)	(0.016)
N	452780	82466	56486	78092	20628	14362
Number of days suspended	0.193	0.545	1.102	0.278	0.649	0.907
	(0.009)	(0.043)	(0.080)	(0.026)	(0.097)	(0.125)
N	232484	38392	30182	51224	12804	9750
Rate of absence (%)	0.07	0.312	0.467	0.089	-0.044	0.345
	(0.012)	(0.039)	(0.055)	(0.030)	(0.075)	(0.095)
N	390782	69806	49986	71304	18604	13334

Note: Means differences between male and female demographic characteristics, at-birth health and school outcomes are presented for the samples indicated in the column titles, by family structure type. Standard errors are in parentheses. See Table 1 for variable definitions.

Table A3: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Race

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Academic Outcomes								
	Mathematics				Reading			
	Black		White		Black		White	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	-0.060***	-0.079	-0.014	0.096	-0.167***	-0.199***	-0.141***	-0.074
	(0.020)	(0.049)	(0.035)	(0.066)	(0.018)	(0.044)	(0.035)	(0.071)
Male*Dad present	0.023	0.054	0.036	0.046	-0.004	0.032	-0.008	0.082
	(0.029)	(0.071)	(0.039)	(0.080)	(0.026)	(0.063)	(0.039)	(0.084)
Male*Married	-0.024	0.007	0.073**	-0.022	-0.028	0.011	0.032	-0.034
	(0.027)	(0.067)	(0.035)	(0.068)	(0.026)	(0.062)	(0.035)	(0.073)
N	114,130	28,162	455,290	81,904	114,816	28,350	459,786	82,550
Panel B: Disciplinary Outcomes								
	Number of days suspended				Absence rate in %			
	Black		White		Black		White	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	1.170***	1.200***	0.981***	0.620***	0.618***	0.609***	0.296*	0.209
	(0.170)	(0.306)	(0.178)	(0.185)	(0.095)	(0.166)	(0.161)	(0.218)
Male*Dad present	-0.340	-0.456	-0.608***	0.098	-0.288**	-0.320	-0.030	0.462
	(0.217)	(0.461)	(0.189)	(0.357)	(0.129)	(0.227)	(0.177)	(0.311)
Male*Married	-0.733***	-0.827**	-0.788***	-0.226	-0.376***	-0.221	-0.159	-0.031
	(0.186)	(0.326)	(0.179)	(0.196)	(0.107)	(0.192)	(0.162)	(0.223)
N	57,990	18,426	234,495	54,004	99,444	25,960	396,477	75,286

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for children born to black mothers and children born white mothers, as indicated by the table labels. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.

Table A4: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Ethnicity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Academic Outcomes								
	Mathematics				Reading			
	Hispanic		Non-Hispanic		Hispanic		Non-Hispanic	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	-0.064	-0.162	-0.047**	-0.001	-0.138***	-0.298***	-0.165***	-0.135***
	(0.044)	(0.121)	(0.019)	(0.041)	(0.043)	(0.108)	(0.017)	(0.040)
Male*Dad present	0.049	0.405***	0.047*	0.032	-0.014	0.369***	0.001	0.032
	(0.050)	(0.140)	(0.024)	(0.057)	(0.049)	(0.125)	(0.023)	(0.054)
Male*Married	0.101**	0.194	0.098***	0.061	0.028	0.162	0.046**	0.021
	(0.045)	(0.128)	(0.020)	(0.045)	(0.045)	(0.114)	(0.018)	(0.044)
N	150,008	20,138	436,352	92,098	151,948	20,336	439,784	92,746
Panel B: Disciplinary Outcomes								
	Number of days suspended				Absence rate in %			
	Hispanic		Non-Hispanic		Hispanic		Non-Hispanic	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	0.784***	0.685**	1.169***	1.038***	0.468***	0.353	0.567***	0.484***
	(0.207)	(0.271)	(0.151)	(0.236)	(0.170)	(0.223)	(0.090)	(0.147)
Male*Dad present	-0.543**	0.891	-0.487***	-0.456	-0.220	0.393	-0.246**	-0.044
	(0.218)	(0.772)	(0.174)	(0.333)	(0.188)	(0.355)	(0.115)	(0.213)
Male*Married	-0.582***	-0.176	-0.954***	-0.668***	-0.285*	-0.096	-0.431***	-0.277*
	(0.207)	(0.319)	(0.152)	(0.243)	(0.172)	(0.240)	(0.092)	(0.154)
N	75,556	12,786	225,502	60,992	130,256	18,398	380,318	84,844

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for children born to Hispanic mothers and children born non-Hispanic mothers, as indicated by the table labels. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.

Table A5: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Immigrant Status

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Academic Outcomes								
	Mathematics				Reading			
	Immigrant		Native		Immigrant		Native	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	-0.125**	-0.103	-0.040**	-0.012	-0.207***	-0.298***	-0.156***	-0.138***
	(0.049)	(0.121)	(0.018)	(0.041)	(0.047)	(0.108)	(0.017)	(0.040)
Male*Dad present	0.081	0.190	0.045*	0.072	0.030	0.336**	0.000	0.045
	(0.058)	(0.145)	(0.023)	(0.056)	(0.056)	(0.133)	(0.022)	(0.053)
Male*Married	0.168***	0.117	0.089***	0.074	0.094*	0.160	0.038**	0.025
	(0.050)	(0.126)	(0.019)	(0.045)	(0.048)	(0.114)	(0.018)	(0.044)
N	135,850	20,446	450,510	91,790	137,780	20,654	453,952	92,428
Panel B: Disciplinary Outcomes								
	Number of days suspended				Absence rate in %			
	Immigrant		Native		Immigrant		Native	
	FE	Twins	FE	Twins	FE	Twins	FE	Twins
Male	0.671***	0.496**	1.173***	1.048***	0.496***	0.013	0.559***	0.517***
	(0.173)	(0.251)	(0.148)	(0.230)	(0.185)	(0.247)	(0.089)	(0.143)
Male*Dad present	-0.361*	0.766	-0.560***	-0.408	-0.116	0.420	-0.278**	-0.022
	(0.187)	(0.784)	(0.168)	(0.328)	(0.205)	(0.426)	(0.110)	(0.207)
Male*Married	-0.449**	-0.075	-0.965***	-0.661***	-0.294	0.258	-0.429***	-0.313**
	(0.175)	(0.295)	(0.149)	(0.238)	(0.187)	(0.263)	(0.091)	(0.150)
N	69,526	13,450	231,532	60,328	118,530	18,908	392,044	84,334

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for children born to mothers native to the U.S. and children born to immigrant mothers, as indicated by the table labels. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.

Table A6: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Mother's Educational Attainment

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Mathematics						
		Sibling FE			Twins	
	< 12	12-15	> 15	< 12	12-15	> 15
Male	-0.065**	-0.003	0.122	-0.100	0.018	0.038
	(0.027)	(0.029)	(0.214)	(0.065)	(0.051)	(0.149)
Male*Dad present	0.037	0.003	-0.172	0.068	0.061	0.236
	(0.038)	(0.034)	(0.231)	(0.093)	(0.066)	(0.219)
Male*Married	0.080**	0.044	-0.052	0.068	0.025	0.053
	(0.038)	(0.030)	(0.215)	(0.094)	(0.056)	(0.151)
N	87,629	330,878	167,853	16,482	68,814	26,940
Panel B: Reading						
		Sibling FE			Twins	
	< 12	12-15	> 15	< 12	12-15	> 15
Male	-0.177***	-0.139***	0.011	-0.213***	-0.142***	0.120
	(0.025)	(0.028)	(0.259)	(0.066)	(0.047)	(0.220)
Male*Dad present	0.014	-0.017	-0.192	0.093	0.076	-0.111
	(0.035)	(0.033)	(0.269)	(0.092)	(0.061)	(0.252)
Male*Married	0.039	0.021	-0.113	0.030	0.022	-0.218
	(0.037)	(0.029)	(0.259)	(0.090)	(0.051)	(0.221)
N	88,096	333,573	170,063	16,564	69,278	27,240
Panel C: Number of days suspended						
		Sibling FE			Twins	
	< 12	12-15	> 15	< 12	12-15	> 15
Male	1.352***	0.885***	0.472	1.254***	0.906***	-0.073
	(0.264)	(0.140)	(0.374)	(0.331)	(0.284)	(0.342)
Male*Dad present	-0.782***	-0.234	-0.184	-0.333	-0.176	0.025
	(0.297)	(0.164)	(0.433)	(0.673)	(0.379)	(0.357)
Male*Married	-0.825***	-0.645***	-0.375	-0.028	-0.469	0.171
	(0.292)	(0.141)	(0.374)	(0.550)	(0.293)	(0.342)
N	46,187	170,767	84,104	11,046	45,966	16,766
Panel D: Absence rate in %						
		Sibling FE			Twins	
	< 12	12-15	> 15	< 12	12-15	> 15
Male	0.506***	0.497***	0.287	0.854***	0.301*	-0.313
	(0.141)	(0.113)	(0.392)	(0.248)	(0.160)	(0.275)
Male*Dad present	-0.227	-0.079	0.005	-0.105	-0.025	1.960
	(0.181)	(0.134)	(0.476)	(0.331)	(0.211)	(1.278)
Male*Married	-0.217	-0.319***	-0.220	-0.282	-0.019	0.334
	(0.174)	(0.115)	(0.392)	(0.321)	(0.169)	(0.281)
N	77,383	288,152	145,039	15,354	63,450	24,438

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for children born to mothers with less than twelve years of education (<12), between twelve and fifteen years of education (12-15) and greater than fifteen years of education (>15), as indicated by the table labels. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.

Table A7: Relationship between the Gender Gap in Academic and Behavioral Outcomes and Family Structure at Birth, by Mother's Age at Birth

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Mathematics								
		Sibling FE				Twins		
	< 21	22 to 29	30 to 35	> 35	< 21	22 to 29	30 to 35	> 35
Male	-0.047**	-0.010	-0.210	-0.174	-0.003	-0.015	-0.134	0.085
	(0.022)	(0.051)	(0.150)	(0.220)	(0.070)	(0.056)	(0.099)	(0.143)
Male*Dad present	0.033	-0.008	0.177	0.058	0.048	0.050	0.230*	0.128
	(0.030)	(0.059)	(0.165)	(0.253)	(0.091)	(0.077)	(0.125)	(0.210)
Male*Married	0.096***	0.053	0.262*	0.218	-0.032	0.075	0.184*	-0.014
	(0.032)	(0.052)	(0.150)	(0.221)	(0.121)	(0.062)	(0.102)	(0.147)
N	134,216	255,338	159,878	36,928	15,896	44,512	36,130	15,698
Panel B: Reading								
		Sibling FE				Twins		
	< 21	22 to 29	30 to 35	> 35	< 21	22 to 29	30 to 35	> 35
Male	-0.151***	-0.164***	-0.303**	0.061	-0.146**	-0.163***	-0.258***	0.045
	(0.020)	(0.052)	(0.140)	(0.210)	(0.066)	(0.058)	(0.091)	(0.132)
Male*Dad present	-0.009	-0.001	0.138	-0.332	0.083	0.052	0.234**	-0.071
	(0.028)	(0.059)	(0.155)	(0.238)	(0.086)	(0.076)	(0.118)	(0.177)
Male*Married	0.020	0.046	0.192	-0.218	-0.102	0.056	0.137	-0.152
	(0.031)	(0.053)	(0.141)	(0.212)	(0.116)	(0.064)	(0.094)	(0.136)
N	135,048	257,632	161,710	37,342	15,982	44,828	36,440	15,832
Panel C: Number of days suspended								
		Sibling FE				Twins		
	< 21	22 to 29	30 to 35	> 35	< 21	22 to 29	30 to 35	> 35
Male	1.262***	0.665***	0.716*	-0.053	1.098***	0.631**	1.738*	0.791**
	(0.186)	(0.188)	(0.399)	(0.125)	(0.239)	(0.298)	(0.911)	(0.398)
Male*Dad present	-0.673***	-0.259	-0.079	0.431	-0.614	0.615	-1.584	-0.487
	(0.217)	(0.206)	(0.474)	(0.271)	(0.530)	(0.481)	(1.002)	(0.457)
Male*Married	-0.778***	-0.429**	-0.577	0.186	0.577	-0.183	-1.493	-0.424
	(0.210)	(0.189)	(0.400)	(0.128)	(0.939)	(0.313)	(0.913)	(0.418)
N	70,533	131,674	81,095	17,756	10,430	29,452	24,022	9,874
Panel D: Absence rate in %								
		Sibling FE				Twins		
	< 21	22 to 29	30 to 35	> 35	< 21	22 to 29	30 to 35	> 35
Male	0.495***	0.603***	0.493	-0.719	0.361**	0.415*	0.809**	0.542*
	(0.105)	(0.192)	(0.627)	(1.436)	(0.179)	(0.249)	(0.324)	(0.304)
Male*Dad present	-0.260*	-0.261	-0.093	0.548	0.483*	-0.075	-0.166	-0.720*
	(0.134)	(0.220)	(0.684)	(1.495)	(0.278)	(0.303)	(0.566)	(0.382)
Male*Married	-0.364***	-0.450**	-0.346	0.854	-0.246	-0.121	-0.662**	-0.303
	(0.133)	(0.194)	(0.628)	(1.438)	(0.304)	(0.258)	(0.328)	(0.322)
N	118,156	222,254	138,509	31,655	14,670	41,040	33,202	14,330

Note: The dependent variables are pooled grade three through eight math and reading test scores, the number of days suspended per year, and the absence rate per year. Each column presents the results from a separate regression. Regressions are estimated separately for children born to mothers of various age groups, as indicated by the table labels. Columns titled "FE" are estimated on the sample of non-twin siblings with the same family structure at the time of birth and present results for specifications that control for mother fixed effects as well as the following set of controls: mother's age at birth, mother's education level, the month and year of the child's birth fixed effects and child birth order. Columns titled "Twins" present estimates using the twins sample, include mother fixed effects and additionally control for child birth order. Standard errors are clustered at mother level.