Institute for Policy Research
Northwestern University
*Working Paper Series*

**WP-15-16**

# Family Disadvantage and the Gender Gap in Behavioral and Educational Outcomes

**David Autor**
Professor of Economics and Associate Department Head, MIT; and Faculty Research Associate, National Bureau of Economic Research (NBER)

**David Figlio**
Orrington Lunt Professor of Human Development and Social Policy and of Economics; Director and Faculty Fellow, Institute for Policy Research (IPR), Northwestern University; and Faculty Research Associate, NBER

**Krzysztof Karbownik**
IPR Postdoctoral Fellow, Northwestern University

**Jeffrey Roth**
Research Professor of Pediatrics, College of Medicine
University of Florida

**Melanie Wasserman**
Doctoral Student in Economics, MIT

Version: October 21, 2015

2040 Sheridan Rd., Evanston, IL 60208-4100 | Tel: 847-491-3395 Fax: 847-491-9916
www.ipr.northwestern.edu | ipr@northwestern.edu

## Abstract

U.S. women graduate from high school at higher rates than U.S. men, but the female-male educational advantage is larger, and has increased by more, among black students and students of low socioeconomic status (SES) than among white and high-SES students. The authors explore why boys fare worse than girls in low-SES households—both behaviorally and educationally—by exploiting matched birth certificates, health, disciplinary, academic, and high school graduation records for more than 1 million children born in Florida between 1992 and 2002. They account for unobserved family heterogeneity by contrasting outcomes of opposite-sex siblings linked to birth mothers by using administrative records. Relative to their sisters, boys born to low-education and unmarried mothers, raised in low-income neighborhoods, and enrolled at poor-quality public schools have a higher incidence of truancy and behavioral problems throughout elementary and middle school, exhibit higher rates of behavioral and cognitive disability, perform worse on standardized tests, are less likely to graduate high school, and are more likely to commit serious crimes as juveniles. The authors argue that the family disadvantage gradient in the gender gap is a causal effect of the *postnatal* environment: Family disadvantage has no relationship with the sibling gender gap in neonatal health, measured by birth weight, APGAR scores (a test given right after birth to quickly assess a baby’s health), prenatal-care adequacy, congenital anomalies, maternal health, and labor and delivery complications. Although family disadvantage is strongly correlated with schools and neighborhood quality, the SES gradient in the sibling gender gap is almost as large within schools and neighborhoods as it is between them. A surprising implication of these findings is that, relative to white siblings, black boys fare worse than their sisters in significant part because black children—both boys and girls—are raised in more disadvantaged family environments.

**Acknowledgments**: We thank Josh Angrist, Marianne Bertrand, John Ham, Louis Kaplow, Mikael Lindahl, Jeremy Majerovitz, Richard Murnane, Jessica Pan, Kjell Salvanes, Till von Wachter, and numerous seminar participants at the NBER Education Program, CESifo Program on the Economics of Education, MIT, Federal Reserve Bank of Atlanta, Miami University, National University of Singapore, Queens University, Simon Fraser University, University of Quebec at Montreal, University of Toronto, Uppsala University, Washington University in Saint Louis, and the Yrjö Jahnsson Foundation 60th Anniversary Conference on Inequality and Health for valuable suggestions that helped to improve the paper. Autor acknowledges support from the Russell Sage Foundation (Grant #85-12-07). Figlio and Roth acknowledge support from the National Science Foundation (NSF) and the Institute of Education Sciences (CALDER grant), and Figlio acknowledges support from the National Institute of Child Health and Human Development and the Bill and Melinda Gates Foundation. Wasserman acknowledges support from the NSF Graduation Research Fellowship and the National Institute on Aging, Grant Number T32-AG000186. We are grateful to the Florida Departments of Education and Health for providing the de-identified, matched data used in this analysis. The conclusions expressed in this paper are those of the authors and do not represent the positions of the Florida Departments of Education and Health or those of our funders.

## Introduction

The last four decades have witnessed a swift and substantial reversal of the gender gap in educational attainment in the U.S. and much of the developed world. Between 1970 and 2010, the high school graduation rate of U.S. women rose by six percentage points, from 81 percentage points ('points') to 87 points, while the U.S. male high school graduation rate was unchanged (Murnane, 2013).[1] Contemporaneously, women overtook and then surpassed males in college attainment. Amidst this widely remarked rise in female educational attainment hides an equally striking and comparatively unremarked puzzle: the female advantage in high school graduation and college attainment is larger, and has risen by substantially more, among children of minority families. For example, while the overall female advantage in high school completions among U.S. adults ages 20 through 24 was 6.2 percentage points in 2010, it was 4.5 percentage points among whites, 12.2 percentage points among blacks, and 7.8 percentage points among U.S. born Hispanics(Murnane, 2013, Table 3).[2] The contemporaneous race gaps in college attainment among young U.S. adults were equally pronounced: whereas white women ages 25 to 29 were 22 percent more likely to hold a B.A. than white males in 2010, the corresponding gap was 55 percent among both blacks and Hispanics (U.S. National Center for Education Statistics, 2013, Table 104.20).[3]

What accounts for the systematically larger gender gaps in the educational outcomes among minorities? In this paper, we focus on family disadvantage as a potential explanation, where by family disadvantage, we mean cross-group differences in the quality and quantity of available household resources, child-rearing inputs (e.g., nutrition, safety in the home and the neighborhood, stimuli), and parental attention. We hypothesize that family disadvantage may differentially inhibit boys relative to girls, either because behavioral and educational outcomes are more elastic to family circumstances among boys than girls, or because differential parental investment in girls relative to boys varies inversely with household SES. Our goal in this paper is to test whether family disadvantage levies a disproportionate effect on the educational and behavioral outcomes of school-age boys relative to girls, to differentiate this hypothesis from a 'fetal origins' alternative as well as from a neighborhood-and-school-quality explanation, and to utilize the resulting estimates to quantify the degree to which higher rates of family disadvantage among minority populations can partly explain

[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]By comparison, the gap in 1970 was zero overall, $-0.4$ percentage points among whites, 5.1 percentage points among blacks, and $-2.5$ percentage points among U.S. born Hispanics. Thus, the increase in the gap among whites, Hispanics, and blacks in this 40 year period was 4.9 points, 10.3 points, and 7.2 points, respectively

[3]In a similar vein, Goldin et al. (2006) document that 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, moreover, the gender gap in the lower two quartiles had risen by substantially more than among the upper quartiles in the prior twenty years. A challenge in making make meaningful comparisons across race groups in the gender gap in college attainment is that the levels differ so substantially by race: college attainment among 25 to 29 year old U.S. adults in 2010 was 38.6 percent among whites, 19.4 percent among blacks, and 13.5 percent among Hispanics; while, the Hispanic female-female *gap* in college attainment of 6.0 percentage points was smaller than the black female-male gap of 8.3 percentage points, the female to male college attainment *ratio* was 1.55 for both groups (U.S. National Center for Education Statistics, 2013, Table 104.20).

the larger gender gaps in educational outcomes we observe among minorities.[4]

There are two main obstacles to assessing the causal effect of family disadvantage on the gender gap in children's outcomes: suitable data and credible identification. To address the data and measurement challenge, we draw upon a unique matched database of birth certificate, and kindergarten readiness, academic, disciplinary, and high school graduation records for over one million children born in Florida between 1992 and 2002.[5] Florida is particularly well-suited to study the role of SES: it has a large, demographically diverse, and socioeconomically heterogeneous population. Our longitudinal data offer remarkable detail on household characteristics, infant and maternal health at birth, early educational outcomes including pre-school enrollment, assessments of kindergarten readiness at the start of formal schooling, third through eighth grade test scores, truancy, disciplinary outcomes, disability, and high school graduation for the oldest cohorts in our sample. We further supplement these data with information from a large Florida school district that records whether students have been repeat-time felons and/or have been incarcerated as teenagers. Since household disadvantage is imperfectly observable even in this rich database, we apply and combine multiple variables to proxy for this underlying construct, focusing both on maternal education as well as father presence or absence at the time of birth.

The identification challenge for estimating the causal effect of family environment on the gender gap in children's outcomes is that variation in child-rearing environments is likely to be strongly confounded with congenital health, abnormal conditions at birth, maternal health, and hereditary factors that also influence children's outcomes (e.g., parental education affects family environment and is likely correlated with children's latent abilities). Our identification strategy seeks to overcome these confounds by contrasting outcomes of opposite-sex siblings linked by birth and schooling records to the same mother. This strategy provides valid identification of the differential effect of child-rearing environment on boys relative to girls under one stringent condition: the gender gap in potential outcomes between siblings is uncorrelated with our measures of family SES at the time of birth; that is, any intrinsic genetic or biological advantage that girls may possess at birth relative to their male siblings is not systematically larger or smaller in low relative to high-SES families.[6] While

[4]Of course, the erosion of prevailing gender norms that historically inhibited women from attaining education commensurate with their ability certainly has helped to account for the growing female advantage in educational attainment. As documented by Goldin et al. (2006) and Fortin et al. (2015), 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 diminishing 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 of all OECD countries in the same year. But these aggregate patterns do not explain why female overtaking in the U.S. differs substantially and systematically across race and SES groups. Indeed, a simple gender norms explanation would likely predict a *lesser* degree of female overtaking among minority and less affluent families, where children's educational attainment is generally lower and gender norms are arguably stronger. In reality, the opposite has occurred.

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

[6]Extensive evidence indicates that maternal disadvantage leads to worse health at birth through poor health behaviors, more adverse environmental exposures, poorer access to medical care, and worse underlying maternal health (Aizer and Currie, 2014). If these adverse impacts are greater for boys than girls, this could generate an SES gradient in the gender gap that is present at birth. Indeed, animal models (e.g., Giussani et al. 2011; Mueller and Bale 2008) suggest that there could be elevated male stress sensitivity in utero.

this strong unconfoundedness assumption is not fully testable, our detailed and precise measures of infant and maternal health, obtained from birth vital statistics, offer a compelling test of its plausibility.

Under the unconfoundedness assumption, our econometric framework admits the possibility that observed differences in the gender gap across family types may be attributable to two related post-natal causal channels: differential sensitivity of boys versus girls to the same family and environmental inputs; and differential investments made in girls and boys that vary systematically with family disadvantage. As a leading example, consider the impact of father absence on the gender gap in children's outcomes. On the one hand, boys may be differentially sensitive to this family attribute because for boys, father absence also implies the potential lack of a same-sex role model in the home (Autor and Wasserman, 2013). On the other hand, existing literature finds that mothers typically spend more parenting time with daughters and fathers typically spend more parenting time with sons (Lundberg et al., 2007; Baker and Milligan, 2013; Bertrand and Pan, 2013; Gayle et al., 2014). This implies that girls may receive relatively larger parental time investments in father-absent households. As this example suggests, our analysis will not be able to distinguish the differential sensitivity versus differential investment channels. We are, however, able to demonstrate that family family disadvantage influences the gender gap in educational and behavioral outcomes through post-natal exposure.

Family disadvantage could also influence the gender gap via school and neighborhood quality. It stands to reason that neighborhoods might influence boys and girls differently if, for instance, boys are more likely to participate in gang activity or other violent behaviors associated with low-income neighborhoods. School quality is also correlated with family advantage, and it's possible that better schools are more able to accommodate different learning styles, or facilitate a climate where boys are more able to learn. An important contribution of this paper is that we can show the independent effects of family disadvantage on the gender gap above and beyond that contributed by neighborhoods and schools. We demonstrate that post-natal factors relating to family disadvantage, above and beyond the neighborhoods children inhabit and the schools they attend, substantially and differentially influence the likelihood that boys will thrive relative to their sisters.

Our analysis proceeds in three parts. Using the longitudinal Florida birth and school records data, we first explore the foundational assumption of unconfoundedness between SES and the gender gap in birth outcomes. We show that although low-SES newborns have lower birth weight and APGAR scores and shorter gestations, have higher rates of congenital anomalies, complications of labor and delivery, and abnormal conditions at birth, and are delivered by mothers in poorer health who received less adequate care during pregnancy, the brother-sister gap in these birth outcomes is essentially invariant to SES. The finding that the gap in neonatal outcomes is unaffected by SES is particularly important since a wealth of work demonstrates the sensitivity and power of birth weight as a predictor of newborns' subsequent health, and educational outcomes (Black et al., 2007; Aizer and Currie, 2014; Figlio et al., 2014). This evidence suggests that the SES gradient in the gender gap is caused by the postpartum mechanisms discussed above.

Having provisionally accepted the identification assumption of unconfoundedness, we then demonstrate that boys born to low-SES households start school less ready for kindergarten, perform worse on standardized tests throughout elementary and middle school, have higher rates of truancy and behavioral problems and criminal behavior/incarceration, and are less likely to graduate high school than are their sisters.[7] These differentials are economically large and potentially explain a substantial share of cross-group differences in the gender gap. For example, approximately 11 percent of Florida public school children are suspended at least once between third and eighth grade.[8] But suspensions are 6.5 percentage points higher among boys than girls, and an additional 6.3 percentage points higher among black boys relative to black girls. Accounting for the differential effect of family disadvantage on boys relative to girls, however, reduces this gap to 2.5 percentage points—meaning that 4.0 percentage points of the observed black-white gender gap is proximately explained by higher levels of disadvantage to which both black boys and girls are exposed. In households of comparable SES, we would predict this gap to be 60% smaller. Similarly, our estimates of the causal effect of family disadvantage on the gender gap can explain about one-third of the excess high school dropout differential among black males versus black females relative to white males versus white females.

Should the causal effects that we attribute to family environment be interpreted instead as the differential effect of neighborhood and school quality on boys relative to girls, where neighborhood and school quality are correlated with family SES due to the endogenous choice of residence? We empirically differentiate between these alternative interpretations in the third part of the paper. For this analysis, we augment our family SES proxies with measures of neighborhood and environmental quality, including non-parametric indicators of neighborhood, direct measures of school quality (on an A through F scale) produced by the Florida Department of Education, and estimates of the causal effect of neighborhoods on economic mobility produced by Chetty and Hendren (2015). Neighborhood quality makes a very modest contribution to the SES gradient in the gender gap, but school quality matters more: low quality schools are particularly disadvantageous for boys. Nevertheless, accounting non-parametrically for the differential impact of schools and neighborhoods on boys relative to girls reduces the estimated impact of family disadvantage on the sibling gender gap by at most one third. We infer that family disadvantage makes both a substantial direct contribution to the gender gap and an indirect contribution through its influence on schools and neighborhoods.

This paper make three principal contributions. A first is to distinguish between gender differences in outcomes that are caused by family and neighborhood conditions during childhood from those

[7] While the children observed in our sample are not yet old enough for us to observe adult outcomes such as high school dropout (except for the oldest cohorts), 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. 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) reports 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) additionally finds that, holding constant educational attainment, childhood misbehavior is negatively associated with adult earnings at all educational levels, whereas achievement test scores are positively related to earnings only for young men with postsecondary degrees.

[8] We observe disciplinary issues following eighth grade as well, of course, but we limit our analysis to elementary and middle grades to ensure the maximum amount of comparability possible for the largest number of sibling pairs.

gender differences that may already be present at birth. The finding that these neonatal health gaps by gender are essentially invariant across family types serves as a key piece of evidence that our identification strategy for testing the causal effect of postnatal family environment on the gender gap in children's outcomes is credible.

Prior work implicitly draws a contrast between boys and girls born to different mothers to assess the relationship between family disadvantage and the gender gaps in child outcomes. This approach, necessitated by data and 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. Our second contribution is to exploit cross-gender, within-family sibling comparisons (akin to Chetty and Hendren 2015, but adding the gender contrast), to sweep out unmeasured, time-constant family effects that may bias comparisons of children of different genders raised in different families. Additionally, the uniquely detailed 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.[9] As a further robustness test, we are able to contrast outcomes of children born to the same mother with distinct fathers, thus testing whether even within households, paternal status (married, present, or absent) appears to differentially affect boys relative to girls. Across all manner of sibling contrasts, we find unambiguous evidence that behavioral and academic outcomes of boys are differentially affected by family circumstances.

Our third contribution is to assess whether the differential gender gap in childhood outcomes in low-SES families is proximately explained 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), school quality metrics, and school indicator variables with dummy variables for gender to ascertain whether the family environment 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 very little impact on the estimated effect of family disadvantage on the gender gap, though we find that both school quality (measured by Florida school grades) and neighborhood quality (measured by the Chetty and Hendren 2015 mobility measures) exert a meaningful impact on the gender gap in childhood outcomes. We are not aware of prior research that concurrently assesses the impact of family disadvantage, neighborhood environment and school quality on child outcomes at such high resolution.

Our paper builds on a recent literature documenting and examining the determinants of the emerging gender gap in educational and behavioral outcomes.[10] Buchmann and DiPrete (2006), Goldin et al. (2006), DiPrete and Jennings (2012), Autor and Wasserman (2013), Bertrand and Pan (2013), DiPrete and Buchmann (2013), and Fortin et al. (2015) explore the evolution of gender

[9]Unmarried fathers are more likely to establish paternity when the child is male (Almond and Rossin-Slater, 2013). In addition, we have conducted an analysis of longitudinal data from a large Florida school district that shows that unmarried fathers are more likely to leave the household by the time a child is in kindergarten if the child is a daughter rather than a son.

[10]See Lundberg and Pollak, 2007; Cherlin, 2009, 2014, and 2015 for broader discussions of the economics of family structure and its evolution over recent decades.

differences in behavioral and educational outcomes in the United States and cross-nationally. Several papers explore specific hypotheses for this gender gap. Becker et al. (2010) theorize that gender differences in the total costs of education—primarily differences in the distributions of non-cognitive skills—explain the overtaking of men by women in higher education. Papers by Buchmann and DiPrete (2006), DiPrete and Jennings (2012), Baker and Milligan (2013)Bertrand and Pan (2013), Owens (2013), Prevoo and ter Weel (2014), Karbownik and Myck (2015), Riphahn and Schwientek (2015), and Woessmann (2015) focus on the role of family and schooling environment in generating observed gender gaps in behavioral and educational outcomes. Bertrand and Pan (2013) offer the most complete extant analysis of the relationship between family background and gender differences in early childhood outcomes. They 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. Considering 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 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. Consistent with this line of argument, Fan et al. (2015) provide evidence from Norwegian registry data that mother's employment during children's early years has a differentially adverse effect on the educational attainment of sons relative to daughters. They hypothesize that rising female employment may in part explain the reversal of the male-favorable gap in educational attainment.[11]

A complementary literature analyzes the causal effects of neighborhood on developmental, educational, and labor market outcomes, abstracting from the direct effect of family characteristics that are often tightly correlated with neighborhood attributes (e.g., income, race, and educational attainment). A spate of experimental and quasi-experimental studies have found little evidence that the neighborhoods in which children grow up have large or consistent impacts on their educational or labor market outcomes (e.g., Katz et al. (2001); Oreopoulos (2003); Kling et al. (2007); Ludwig et al. (2013)). Two recent studies that pair experimental (Chetty et al., 2015) or quasi-experimental (Chetty and Hendren, 2015) research designs with high resolution tax register data reach a different conclusion: early and prolonged childhood exposure to neighborhoods affect college attendance, labor market earnings, single-parenthood, and neighborhood quality in adulthood. Notably, the quasi-experimental Chetty and Hendren (2015) study, which obtains identification from millions of neighborhood moves finds that boys' outcomes vary more across areas than girls, and that boys' outcomes are more affected by neighborhood exposure than are girls. This work both supports the plausibility of a differential sensitivity channel in driving the gender gap and highlights the value of considering the contributions of family environment and neighborhood simultaneously.

[11] Baker et al. (2008) document that the introduction of highly subsidized, universally accessible child care in Quebec during the late 1990s increased maternal labor supply significantly and, simultaneously, adversely impacted early childhood developmental outcomes among affected cohorts of children. Follow-up work by Baker and Milligan (2013) found that this childcare expansion generated *increased* parental time investment in boys relative to girls.

A final body of work that bears on our identification approach studies the causal effects of environmental and maternal stresses on children's in utero mortality, birth weight, health, cognitive development, and educational and labor market outcomes. Almond and Currie (2011) and Aizer and Currie (2014) review the large literature establishing the causal effect of maternal disadvantage on infant birth weight and infant health, while Almond et al. (2005), Black et al. (2007), and Figlio et al. (2014) document the robust and enduring adverse effect of low birth weight on children's IQ scores, educational attainment, and adult earnings. Several papers confirm the long-standing Trivers-Willard hypothesis (Trivers and Willard, 1973) that in utero stress increases the mortality odds of male relative to female fetuses (Norberg, 2004; Almond and Edlund, 2007). Studies that test for impacts of maternal malnutrition and maternal nutrition supplementation on subsequent academic and labor market outcomes of children who were potentially exposed in utero, find robust, though generally modest, impacts on adult health, skills acquisition, and labor market outcomes (Almond and Mazumder, 2011; Field et al., 2009; Maccini and Yang, 2009; Almond et al., 2014; Hoynes et al., 2015 Forthcoming; Greve et al., 2015). These studies do not, however, consistently find differential impacts by gender on post-natal outcomes. Moreover, where differences are detected, they largely suggest that pre-natal disadvantage has a slightly more adverse impact on girls than boys.[12] Assuming plausibly that pre-natal and post-natal disadvantage are positively correlated, this (modest) pre-natal differential would work *against* our finding that boys are differentially adversely affected by family disadvantage.[13]

The paper proceeds as follows. Section 1 details our data. Section 2 describes our identification strategy using a potential outcomes framework, provides a discussion of potential threats to validity, and specifies our main regression model. Section 3 begins by documenting non-experimental gender differences by family socio-economic status, and shows using sibling fixed effects models that

[12] Greve et al. (2015) find that fetal exposure to Ramadan has a larger negative impact on the achievement scores of Muslim girls than boys in Denmark. Hoynes et al. (2015 Forthcoming) find that improvements in in utero nutrition have a larger positive impact on economic self-sufficiency (an omnibus skills and earnings measure) and overall good health among exposed girls than boys, while impacts for the prevalence of metabolic syndrome show slightly greater benefits for exposed boys than girls. In a developing country context, Maccini and Yang (2009) examine the effect of weather shocks around the time of birth on Indonesian women and men, finding positive effects on health, height, schooling, and household wealth in adulthood among women but not men. Field et al. (2009) find that prenatal iodine supplementation during the first trimester increased completed schooling by as much as half a year, with generally stronger effects for girls than boys. Cullen et al. (2015) review a large body of U.S. cross-national and time series data on the male/female post-natal mortality gap. In wealthy countries such as the U.S., it is unambiguously the case that poorer post-natal conditions (measured by income, education, or prevalence of poverty) are associated with higher differential male mortality. Their paper does not explore mechanisms.

[13] Two recent animal studies test whether environmental and maternal stresses occurring in utero differentially affect health and development of male fetuses. Mueller and Bale (2008) find that male mouse fetuses exposed to high levels of maternal stress in utero exhibit greater adverse behavioral stress responses during adulthood than do similarly exposed female mouse fetuses. Giussani et al. (2011) document that when a pregnant sheep is placed under acute hypoxic stress (i.e., deprived of oxygen), her male fetuses exhibit a substantially elevated cortisol response (a hormonal measure of stress) relative to her female fetuses. It is unknown at present whether these differential gender-specific responses also occur in humans and, if so, how consequential they might be for postpartum outcomes. Aizer et al. (2009) estimate modest negative effects of in-utero exposure to high maternal cortisol levels on the cognition, health and educational attainment of offspring. Persson and Rossin-Slater (2014) shows that prenatal exposure to stress via the death of a maternal relative causes higher take up ADHD medication during childhood, and anti-anxiety and depression medications during adulthood. Neither Aizer et al. (2009) nor Persson and Rossin-Slater (2014) distinguish between impacts on boys and girls.

these differences are *not* present across a wide variety of birth conditions. Section 4 demonstrates that family advantage significantly affects the gender gap in educational and behavioral outcomes, whether measured at the time of kindergarten enrollment, in third through eighth grade, and in eventual rates of high school on-time completion and dropout, and shows that the SES gradient in the gender gap extends to serious criminal behavior and incarceration as well. Benchmarking the impact of family disadvantage on early behavioral and educational outcomes, we find that the correlation between race and family disadvantage can explain about one third of the excess boy-girl gap in on-time high school graduations among black relative to white children. After verifying parametric restrictions imposed by the model in section 5, we examine in section 6 potential pathways by which family structure could affect the gender gap in childhood outcomes, specifically its operation through neighborhood and school quality. Section 7 concludes.

# 1 Data and descriptive statistics

## 1.1 Principal 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 the 1995-96 through the 2012-13 school years from the Florida Department of Education Data Warehouse for children in these birth cohorts who attended the Florida public schools.[14] Birth certificate data were matched to schooling records by the Florida Departments of Education and Health 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 from data from the decennial Census and American Community Survey for years 2000 through 2009.[15]

Florida birth certificates enumerate demographic characteristics of the mother (including education and marital status), health and demographic characteristics of the the newborn, and demographic characteristics of the father if he appears on the birth certificate.[16] We use the birth certificate data to create three proxies of family environment: mother's education, family structure at the time of birth, and average family income in the zip code of the mother's residency at the time of birth. Mother's education is grouped into high school dropout, high school graduate or some

[14] The bulk of our analysis is limited to children born between 1994 and 2002, as we only observe neighborhood measures and sibship status for children in those birth cohorts. We make use of data from 1992 and 1993 in order to study high school graduation, because we observe few children born 1994 onward who had had the opportunity to graduate from high school as our most recent data extract is for the high school class of 2012/2013. We also make use of a distinct dataset provided by a large Florida school district to study the effects of SES on the gender gap in serious criminal activity/juvenile incarceration.

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

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

college, and at least four years of college. Family structure at birth is grouped into three categories: parents married at birth (*married*); unmarried but the father's name appears on the birth certificate (*father present*); and unmarried and the father's name does not appear on the birth certificate (*father absent*).[17]

School records include third through eighth grade Florida Comprehensive Assessment Test (FCAT) math and reading scores, as well as records of absence rates and suspensions school during each academic year.[18] In addition, for some cohorts of children, the Florida Department of Education recorded the results of a universally-administered kindergarten readiness assessment. Each child in the sample can contribute up to six observations for each variable (excluding kindergarten readiness and pre-school enrollment), one per each year observed in each grade. In all regression estimates that follow, we cluster standard errors at the family level to account for the high degree of within-child and within-family serial correlation in these outcomes. We employ direct measures of school quality (official school grades, reported on an A through F scale) produced by the Florida Department of Education.[19] We discuss both birth characteristics and childhood outcomes at length below.

### 1.2 Sample selection and descriptive statistics

Table 1 presents summary statistics for the universe of Florida children born to white non-Hispanic natives, black non-Hispanic natives, Hispanic natives, and immigrants for the years 1994 through 2002 (column 1), which for compactness, we will refer to as white, black, Hispanic, and immigrants for the remainder of the paper.[20] The second column drops the 9% of records that are missing key variables. The third column contains the approximately 81% of column 2 records that were matched to Florida school records; those not matched to school records have either left the state of Florida or attended private school in the state.[21] Column 4 contains the subset of column 3 records with a valid third-grade test score, while column 5 contains the subset of column 4 records containing matched siblings. Due to data availability restrictions we are only able to match siblings born in a

[17]For children born to married couples, information on paternity may be provided by either the mother or father. For children born to unmarried parents, information about the father appears on the birth certificate only if he claims paternity (thus, presence/absence refers to the father's paternity status, not his physical presence/absence at delivery). Laws in Florida that strongly encouraged paternity establishment were not yet in force at the end of our data window, so we feel more comfortable that the voluntary paternity reporting on the birth vital records is a reasonable proxy for paternal presence in a child's life.

[18]Although available for the oldest 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. Furthermore, because of elective subjects and the lack of a standard high school curriculum in mathematics, in particular, the test scores in mathematics are not directly comparable across children.

[19]See `http://schoolgrades.fldoe.org/` for details regarding Florida's school grading process.

[20]We exclude the 2.7 percent of births that fall in neither of these four categories—that is, native births whether the mother is neither white, black nor Hispanic.

[21]Because all disabled children in Florida, regardless of SES, are eligible for the McKay Scholarship Program, a voucher program that provides a subsidy to attend private school for children with disabilities, and the Florida Department of Education maintains utilization records for this program, any disabled child, whether in public or private school, is observed in the school records.

subset of Florida counties and for children born from 1994 forward, representing about 56 percent of the possible set of birth records matched to school records (see Figlio et al. 2014 for details).

Relative to the population of Florida births (column 1), limiting the sample to birth records with complete data (column 2) has almost no effect on birth demographics. Restricting further to births that subsequently appear in Florida public school records (column 3) and eventually obtain a third grade test score (column 4) increases the share of mother who are black, Hispanic, younger, less educated, and unmarried at delivery. These changes in sample composition are consistent with the greater cross-state mobility of high SES adults (Molloy et al., 2011) and higher private school attendance rates of their children. While the matched sample represents a more disadvantaged population than the full sample, the gender composition of those matched to school records does not differ appreciably from the full population of births, and there are virtually no differences in birth weight between the full population and the matched sample.[22] When we further limit the sample to matched sibling births (column 5), we observe relatively fewer white and immigrant mothers, slightly more college educated mothers, and slightly more births with no paternity established. These patterns primarily reflect the greater diversity of the counties where matching is feasible rather than characteristics of households with siblings *per se*.[23]

Tables 2 summarize key demographic characteristics for our main sample observed either at birth or during K-12 schooling according to the race and ethnicity of the mother. These include mother's education and paternal status at birth (married, father absent: paternity claimed, father absent: no paternity claimed), the rated quality of Florida public schools subsequently attended, the median income in the zip code of the mother's residence at the time of birth, and the Chetty-Hendren (2015) measure of county economic mobility associated with that zip code.[24] The pronounced contrasts among race and ethnic groups in each of these measures highlights the degree to which family disadvantage differs systematically across these broad demographic groups. For example, the fraction of children born with no claimed paternity ranges from 7 points among whites to 43 points among blacks. Comparisons of maternal education and income reveal similarly striking contrasts. Notably, infant birth weight mirrors these patterns of family disadvantage: children born to white mothers weigh on average nearly 300 grams more than children born to black mothers, and 50 to 100 grams more than children born to immigrant and Hispanic mothers. These results underscore that comparisons across race and ethnic groups are also implicitly comparisons across education groups, income levels, and household type—meaning that cross-race and cross-ethnic group comparisons are strongly confounded with socioeconomic status. To overcome these confounds, we assess the impact of education, income and household structure on children's outcomes by exploiting within-family, cross-gender contrasts, as explained next.

[22]The slight reduction in the fraction of male births (from 51.2% to 50.6%) as the sample is restricted to more disadvantaged mothers is consistent with the Trivers-Willard (1973) hypothesis.

[23]Immigrant status is substantially more prevalent (27.2%) among births in the counties where sibling matching is feasible than among matched siblings in these counties (20.7%). This pattern arises siblings born outside of Florida (often in the origin country) are not observed in our data.

[24]We detail our procedure for imputing SES from these various variables in section 2 below.

## 2 Empirical framework

Our empirical objective is to isolate the causal effect of family disadvantage on the gender gap in childhood disciplinary and educational outcomes. Since family disadvantage is not randomly assigned to children, we propose an identification approach that does not rely on the assumption that children's potential outcomes are uncorrelated with family disadvantage.

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 gender-discordant siblings, $b$ and $g$, born to mother $j$ whose socioeconomic status at the time of the children's birth is $D_j \in \{0, 1\}$. For expositional simplicity, we treat $D_j$ as discrete (i.e., a family is either disadvantaged or advantaged), and we consider the case where $D_j$ is the same for both births, $D_{jb} = D_{jg}$, though we relax that in the estimation.[25] For sibling pair $j$, we observe only one set of potential outcomes as a function of family advantage, $D_j$:

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

We expect that potential outcomes will differ between brothers and sisters within a family $j$, and further, that these potential outcomes will covary with family disadvantage for both sexes. This means that absent random assignment of disadvantage to households, simple within- or between-family contrasts will not provide a valid causal estimate of the effect of family disadvantage on outcomes $Y$ for either boys or girls. To see why, consider the non-experimental contrast between the outcomes of boys born to disadvantaged versus advantaged families. This contrast is

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

where the first bracketed term on the righthand side is the average causal effect of family disadvantage on outcome $Y$ and the second bracketed expression is a bias term, stemming from differences in potential outcomes between boys born to disadvantaged versus advantaged families. Both intuition and data suggest that this bias term will be non-zero: children born to advantaged families are likely to have genetic and health advantages at birth that may yield more favorable outcomes, holding child-rearing circumstances constant. If so, we can learn little about the causal effect of family disadvantage on children's outcomes simply by contrasting children born to disadvantaged families with those born to advantaged families. By a similar argument, the contrast between brothers and sisters within a family does not identify a parameter of interest since, within families, we expect potential behavioral and educational outcomes to differ systematically between boys and girls (formally, $E\left[Y_{jb}^1\right] \neq E\left[Y_{jg}^1\right]$ and $E\left[Y_{jb}^0\right] \neq E\left[Y_{jg}^0\right]$).

To circumvent this identification challenge, we select as our outcome of interest the difference in the gender *gap* in sibling outcomes, $\hat{Y}_j^1 \equiv Y_{jb}^1 - Y_{jg}^1$ and $\hat{Y}_j^0 \equiv Y_{jb}^0 - Y_{jg}^0$, between advantaged and

[25] We focus on gender-discordant siblings to simplify the formal exposition, but our empirical analysis includes all siblings, regardless of gender discordance. These additional observations help to identify the effects of attributes that vary across individuals within families, such as age and parity.

disadvantaged families. We can identify the causal effect of family disadvantage on this gender gap provided that the gap in potential outcomes between brothers and sisters is independent of family disadvantage. Formally:

**Assumption 1.** *$\hat{Y}_j^1, \hat{Y}_j^0 \perp D_j$. The* latent *gap in childhood outcomes between brothers and sisters is as good as randomly assigned to families.*

Under Assumption (1), any *observed* variation in the gender gap that differs systematically across family types will reflect the causal effect of family type on the gender gap.[26] Invoking this assumption, we can identify the causal effect of family disadvantage on the gender gap by contrasting, within-family cross-sibling differences in outcomes by family type:

$$\begin{aligned} E\left[\hat{Y}_j|D=1\right] - E\left[\hat{Y}_j|D=0\right] &= E\left[Y_{jb}^1 - Y_{jg}^1|D=1\right] - E\left[Y_{jb}^0 - Y_{jg}^0|D=0\right] \\ &= E\left[Y_{jb}^1 - Y_{jg}^1|D=1\right] - E\left[Y_{jb}^0 - Y_{jg}^0|D=1\right], \end{aligned} \quad (3)$$

where the second equality in (3) follows from Assumption (1).

As an example of a setting where this identification approach would be appropriate, define the outcome variables $\tilde{Y}_{jb}^1 \equiv Y_{jb}^1 - \alpha_j$, $\tilde{Y}_{jg}^1 \equiv Y_{jg}^1 - \alpha_j$, $\tilde{Y}_{jb}^0 \equiv Y_{jb}^0 - \alpha_j$, and $\tilde{Y}_{jg}^0 \equiv Y_{jg}^0 - \alpha_j$, where $\alpha_j$ is is a mother-specific random effect that directly affects her children's potential outcomes. The confounds discussed above imply that potential outcomes vary systematically with both gender and with family disadvantage:

$$E\left[\alpha_j D_j\right] \neq 0,\ E\left[\tilde{Y}_{jb}^1\right] \neq E\left[\tilde{Y}_{jg}^1\right],\ E\left[\tilde{Y}_{jb}^0\right] \neq E\left[\tilde{Y}_{jg}^0\right]. \quad (4)$$

However, Assumption (1) implies that:

$$\left\{\tilde{Y}_{jb}^1 - \tilde{Y}_{jb}^0, \tilde{Y}_{jg}^1 - \tilde{Y}_{jg}^0\right\} \perp D_j. \quad (5)$$

That is, the between-sibling gap in potential outcomes is independent of disadvantage. Consequently, contrasting the brother-sister gap in outcomes within families eliminates the cross-family dependency between family disadvantage and potential outcomes (arising through $\alpha_j$), while contrasting the brother-sister gap across advantaged and disadvantaged families eliminates the within-family dependency between gender and potential outcomes (arising from intrinsic gender differences). In this setting, equation (3) provides an unbiased estimate of the causal effect of family disadvantage on boys relative to girls.

Is this strategy for identifying the causal effect of family disadvantage on the behavioral and educational outcomes of boys relative to girls plausible? Assume for example, in violation of our identifying assumption, that family SES differentially affects the fetal development of sons relative to daughters—which could occur if the male fetus is more sensitive than the female fetus to maternal stress levels during pregnancy. In this case, the contrast formed in equation (3) would confound any causal effect of family disadvantage on the gap in outcomes between boys and girls with the

[26] It is actually sufficient for our purposes that the latent gender gap is uncorrelated with our measures of family disadvantage. Thus, Assumption (1) is stronger than we require but nevertheless appear plausible, as we discuss further below.

differential in-utero impact of maternal stress on the subsequent development of children of each sex. Following the fundamental problem of causal inference, this assumption is not directly testable. But we can partially assess its validity by assessing whether gender gaps in neonatal health, measured by birth weight, prenatal care, maternal health, and Apgar scores, differ across families of varying socioeconomic levels. As noted in the Introduction, and documented in section 3, we find no evidence that these gaps vary substantially or systematically with family disadvantage. The multiple studies assessing the impact of *in utero* shocks on post-natal outcomes discussed above generally find that adverse impacts are slightly more pronounced for girls than boys. This would work against a finding that post-natal family disadvantage is differentially detrimental to boys.[27]

## 2.1 Implementation

-We bring this conceptual framework to the data with a set of simple regression models. The key puzzle motivating our inquiry is the substantially greater apparent female advantage in educational attainment in black and Hispanic households relative to white households. This observed gap also serves as a benchmark for assessing the explanatory power of our subsequent models. To provide this benchmark, we estimate the following within-family gender gap model,

$$\begin{aligned} Y_{ij} \;=\; & \gamma_j + \beta_1 \text{Boy}_i + \beta_2(\text{Boy}_i \times \text{Black}_j) + \beta_3(\text{Boy}_i \times \text{Hispanic}_j) + \beta_4(\text{Boy}_i \times \text{Immigrant}_j) \\ & + F_j' \psi + X_i' \lambda + e_{ij}., \end{aligned} \tag{6}$$

where $Y_{ij}$ represents an outcome for child $i$ born to mother $j$, $\text{Boy}_i$ is an indicator variable for whether the child is male, and $\text{Black}_j$, $\text{Hispanic}_j$ and $\text{Immigrant}_j$ are indicators for whether the mother of child $i$ is belongs to one of those race or ethnic categories, with white, non-Hispanic, non-immigrant mothers serving as the reference category. The regression model also includes family fixed effects $\gamma_j$, 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 $F_j$ controls for maternal and family environment characteristics that may vary across births $i$, including mother's education, age, and marital and paternity status. The vector $X_i$ additionally controls for time-invariant child attributes, including birth order and month and year of birth.

In this initial descriptive regression model, the coefficient $\beta_1$ measures the boy-girl difference in outcome $Y_{ij}$ for the omitted category of children of mothers who are white, non-Hispanic natives. The coefficients $\beta_2$, $\beta_3$ and $\beta_4$ represent the demographic differentials of interest, that is, how the boy-girl gap varies with the race-ethnicity-nativity status of the mother. The main effects of family attributes, as parameterized by mother's education and family structure at the time of birth, are constrained to be constant across boys and girls in this baseline model.[28]

[27] 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. We also perform our main analysis using all singletons and siblings and find highly comparable (though less precise) results.

[28] For children of mothers with stable characteristics across births, the effect of family disadvantage are captured by family fixed effects and hence not identified.

The second model implements our proposed identification strategy by permitting the effect of family disadvantage to differ by child gender:

$$\begin{aligned} Y_{ij} \quad = \quad & \gamma_j + \beta_1' \text{Boy}_i + \beta_2' (\text{Boy}_i \times \text{Black}_j) + \beta_3' (\text{Boy}_i \times \text{Hispanic}_j) + \beta_4' (\text{Boy}_i \times \text{Immigrant}_j) \\ & + \beta_5' (\text{Boy}_i \times \text{D}_j) + F_j' \psi + X_i' \delta + e_{ij}. \end{aligned} \quad (7)$$

Here, $\text{D}_j$ is a measure of family disadvantage and the coefficient $\beta_5'$ on interaction term $(\text{Boy}_i \times \text{D}_j)$ captures the differential responsiveness of boys to family disadvantage (and the main effect of $D_j$ is absorbed by the mother fixed effect $\gamma_j$).[29] As detailed above, the interpretation of $\beta_5'$ as the *causal* effect of family disadvantage on the gender gap in cognitive and behavioral outcomes hinges critically on the assumption that the latent gap in outcomes between boys and girls is uncorrelated with family disadvantage at the time of birth. We assess the plausibility of this assumption below by estimating Equation (7) for pregnancy health and at-birth child health outcomes.

Through a comparison of the coefficients from the within-family gap model from Equation (6) and the augmented model from Equation (7), we can calculate what fraction of the race-ethnicity-nativity gradient in the gender gap is explained by the differential effect of family disadvantage on boys. Specifically, we compare $\beta_2$ to $\beta_2'$, $\beta_3$ to $\beta_3'$, and $\beta_4$ to $\beta_4'$. Take, for example, the comparison of the black-white gender gap in the within-family gap model and the augmented model, as estimated by $\beta_2$ and $\beta_2'$, respectively. If we find that the black-white gender gap declines when we permit family disadvantage to differentially affect boys, i.e. $\beta_2'$ is smaller in magnitude than $\beta_2$ , we would attribute this decline to the higher prevalence of disadvantage among black families, and the fact that family disadvantage exerts a disproportionate negative effect on boys relative to girls.

Equation (7) imposes the restriction that the impact of disadvantage on the gender gap in outcomes is constant across race and ethnicity groups—that is, the coefficient on $\text{Boy}_i \times \text{D}_j$ does not require a race-ethnicity-nativity subscript. If this restriction does not hold, the precise contribution of disadvantage to cross-group variation in the gender gap is somewhat ambiguous for the simple reason that the share of the cross-group gap explained by disadvantage will depend upon which group-specific $\text{Boy}_i \times \text{D}_j$ slope is used for the calculation.[30] We test this restriction of constant slopes below and find that it is generally well supported for four of six outcomes (kindergarten

[29] Our analysis permits measures of family disadvantage (mother education, father absent/present/married) to vary across births, but all of the results are robust to either restricting the sample to mothers with stable characteristics across births, or to assigning the characteristics of the first birth (maternal education, marital status) to all subsequent births. Our primary models include main effects for disadvantage to account for this variation and allow $D_j$ to also take an $i$ subscript. We suppress this notation here for simplicity.

[30] Applying the terminology in Fortin et al. (2011)'s Handbook chapter on decomposition methods in economics, the gap between two groups in an outcome of interest can be decomposed into two components: *compositional* differences across groups in the distribution of characteristics; and *wage structure* (or simply 'structural') differences in the returns to those characteristics for each group. In our setting, compositional effects correspond to cross-group differences in disadvantage, while structural effects correspond to cross-group differences in the impact of disadvantage on the gender gap. Our primary hypotheses for this paper focus on compositional differences across groups that augment or diminish the gender gap, holding constant the structural relationship between disadvantage and the gender gap. We do not have specific hypotheses for whether or why the structural relationship between disadvantage and the gender gap would differ across groups.

readiness, absences, suspensions, and high school graduation) but not as well supported for math and reading scores. As it turns out, test scores prove less central to our analysis because (a) in general, they have surprisingly modest explanatory power for our key human capital outcome of high school completion, and (b) our measures of family disadvantage have low explanatory power for these outcomes. We elaborate on these points below.

To estimate these within-family, cross-gender contrasts of educational and behavioral outcomes, our main estimates use a sample of mothers who give birth to two or more children in the observed years, allowing us to achieve identification of the boy main effect (and interactions) while including family fixed effects. To assess whether this identification strategy is likely to limit the generalizability of the findings, we take two further steps: estimating sibling-pair models that exclude family fixed effects; estimating OLS models on an expanded sample that includes all singletons and same-gender siblings and excludes family fixed effects. In practice, these three specifications (siblings with family fixed effects, siblings OLS, and all siblings and singletons OLS) yield largely comparable results. However, the sibling fixed effects models generally find the strongest and most precisely estimated relationship between family disadvantage and the gender gap in outcomes—particularly for the key behavioral measures—suggesting that these models are most successful in sweeping biases stemming from cross-family comparisons of boy versus girl outcomes.

## 3 Testing the unconfoundedness assumption

We begin by assessing the plausibility of the key identifying assumption: whether the latent gender gap in sibling outcomes is independent of family disadvantage. Since the latent gap is by definition unobservable, we analyze the best available proxies for neonatal health and well-being from birth certificate data. Perhaps the most important is birth weight, found by a large medical and economic literature summarized earlier to be robustly predictive of subsequent health, cognitive development, and labor market outcomes. There are substantial birth weight differences across our four main demographic groups, as documented in Table 3 above. At birth, white boys weigh an average of 265 grams (approximately 8 percent) more than black boys, about 100 grams more than Hispanic boys, and about 60 grams more than boys of immigrant mothers, with a similar inter-group pattern evident for girls. There are also substantial differences across the demographic groups in terms of a number of other measures of birth outcomes.

Table 4 estimates equations (6) and (7)in a model with maternal fixed effects to assess whether the sibling gender gaps in birth weight and other neonatal outcomes vary systematically with household demographics, as might occur if family disadvantage differentially impacts *in utero* development of either sex. This table presents results for a range of neonatal outcomes—birth weight; abnormal conditions of the newborn; maternal health; and adequacy of prenatal care. If maternal healthcare utilization or health were affected by (or correlated with) the gender of the fetus—perhaps due to son preference, as in Dahl and Moretti (2008)—the last two measures might detect such a relationship. These results are supplemented by Tables A2and A3, in the Appendix, which reports results for a range of additional birth outcomes, including the clinical estimate of gestation in weeks; the five-minute Apgar score, which is a composite index of neonatal vital sign indicators collected in the minutes following delivery; an indicator of the presence of congenital anomalies; and an indicator of complications of labor and delivery. In addition, since one might expect the boy-girl birth weight differential to be constant in proportional rather than level terms, the appendix table presents a set of estimates for birth weights measured in natural logarithms.[31] Finally, the appendix table presents a set of estimates where the dependent variable is an index constructed from all of the birth outcomes.[32]

For each of the birth outcomes, we report three model specifications. The first set of estimates summarizes mean birth weight (or other birth outcome) differences between boys and girls, overall and by race-ethnicity-nativity, conditional on family fixed effects. Besides the family fixed effects, the only included covariates in this first descriptive model are main effects for boy and race-ethnicity-nativity, as well as interactions between the boy and race-ethnicity-nativity dummies. In the case of birth weight, the coefficient on the boy main effect indicates that white newborn boys weigh on average 125 grams more than their white newborn sisters. Despite the substantially lower average birth weight of black boys and black girls, the black boy-girl birth weight gap within families is essentially identical to that of whites, just 8 grams less than that seen for whites. We also find very modest mean differences in the cross-sibling gender-racial gap in birth weights among Hispanic and immigrant births relative to white births.

In the second column of each panel we test for a relationship between family disadvantage and the gender gap in birth outcomes by adding a set of interactions between child sex, mother's education (high school graduate, college graduate, with high school dropout serving as the reference group), and mother's marital status at childbirth (married, father present, with father absent serving as the reference group). This model also controls for main effects of mother's education, marital status at birth, mother's age at childbirth, and child's birth order, birth month, and birth year.[33] The boy-girl difference in birth weights is trivially (and not statistically significantly) larger—on the order of 8 to 11 grams—for births where the mother has at least a high school education, and is completely unrelated to family structure. A similar lack of a relationship between measures of family advantage and gender are observed across all of the other birth outcomes presented in either the main table or the appendix tables.

The third column of each panel subsumes these four interaction terms into a single composite SES measure based on a principal components analysis of family structure/marital status and maternal education (in years). Construction of this measure is documented in Table A1 We again estimate a tiny and statistically insignificant relationship between family advantage and the gender gap in birth weight: a one standard deviation increase in the SES index (an increment of 1.44) predicts less than a two gram rise in the boy-girl differential in birth weight, an order of magnitude that is of negligible

[31] This distinction could matter since boys weigh more than girls on average within every demographic subgroup.

[32] Construction of each of these outcome variables is detailed in the corresponding table notes.

[33] These main effect variables, either within or between families, are only minimally correlated with newborns' gender and hence have essentially no impact on the estimated gender gaps in birth outcomes reported in the first column of each panel.

educational and health relevance.[34] In all of the other outcomes, the estimated relationship between this SES index and gender is equally inconsequential. In net, comparing sibling pairs, there is no relationship between family advantage and the gender gap in birth weight.[35]

Figure 1 graphically summarizes these patterns by presenting scatter plots of the relationship between maternal SES and the gender gap in neonatal health. Across multiple outcome measures, the relationship between family advantage and the gender gap in infant health is of trivial economic magnitude and is never statistically significant in within-family models. While neonatal outcomes such as the Apgar, APCU, and maternal health indicators should not be considered to be as sensitive or reliable as birth weight as measures of neonatal health, these multiple outcomes nonetheless paint a consistent picture: there is no evidence that the gender gap in infant health at birth is correlated with family advantage. We therefore conclude that any SES gradient in gender gaps in educational outcomes are due to post-natal factors.

# 4 Main results: Gender gaps in behavioral and academic outcomes during K-12

Table 5 and 6 summarize the pronounced gender-contrasts in siblings' educational and behavioral outcomes according to maternal race, ethnicity and nativity. We focus our discussion on a corresponding set of plots (Figure 2) depicting gender gaps in outcomes along four dimensions: kindergarten readiness (a measure of early literacy and/or numeracy skills); school absence rates (a behavioral measure); standardized math scores (an academic measure); and on-time high school graduation (an achievement measure). Kindergarten entry is observed for a subset of cohorts. Absence rates and math scores are observed during grades three through eight for 1994 to 2002 cohorts, while high school completions are observed for the 1992 and 1993 cohorts, which had reached the age of on-time high school completion by the end of our sample.[36]

The top left panel of Figure 2 highlights the striking cross-race and cross-ethnicity differences in the gender gap in academic and behavioral outcomes. Among children born to white mothers, the boy-girl gap in the absence rate is 0.09 percentage points, the boy-girl gap in middle school math scores is 0.06 standard deviations (SDs), and the boy-girl gap in both kindergarten readiness as well as on-time high school graduation is $-5.7$ percentage points (i.e., girls start kindergarten more ready to start school, and also have higher graduation rates). With the exception of KG readiness, the gap in each of these outcomes is monotonically widening (becoming less favorable to boys) as we move the focus of comparison from whites to Hispanics to immigrants to blacks. Among children of

[34] Figlio et al. (2014) estimate that a $1,000$g increase in birth weight is associated with a 0.19 standard deviation increase in test scores in grades three through eight. The 2 gram differential we estimate in Table 4 are one-half to one percent as large, implying an impact on test scores on the order of one one-hundredth of a standard deviation.

[35] When we estimate the same models with singleton comparisons, rather than sibling comparisons, the coefficients on interactions between SES measures and gender are sometimes statistically significantly distinct from zero, but are the same trivial magnitudes as those observed with the family fixed effects models.

[36] We currently observe data through the 2012-13 academic year, while administrative data on high school graduation comes at a one-year lag.

black mothers, the boy-girl gaps in absences, math scores, and on-time graduations are, respectively, 0.48 percentage points, $-0.06$ SDs, and $-12.8$ points—in each case, two to five times as large as among whites. The boy-girl gap in kindergarten readiness among blacks is $-7.9$ points, about one and one-third the size of the boy-girl gap among white children. The gender gap in outcomes among children of immigrant and Hispanics mothers fall roughly in between those of whites and blacks on these three measures.

Panels B through D of Figure 2 plots analogous contrasts by maternal education (high school dropout, high school graduate or some college, four year college degree or higher), and family structure at birth (father absent, father present, married), and quartile of the SES appendix (see tab:app-pca-SES).[37] The boy-girl deficits in behavioral and educational outcomes are robustly larger in families with greater disadvantage, whether measured by maternal education, family structure, or the SES composite.

## 4.1 Primary estimates: Family disadvantage and the gender gap in early outcomes

Having established the veracity of Assumption (1) to the extent permitted by the data, we proceed first to assess the causal effect of family disadvantage on the gender gap in kindergarten readiness and in behavioral and educational outcomes in grades three through eight, followed by disability, then high school completions for the oldest two cohorts in the sample. We then present evidenceon juvenile criminal behavior using a distinct data source.

### Kindergarten readiness

The Florida Department of Education recorded kindergarten readiness measures for entering kindergarteners in two sets of kindergarten cohorts—those who entered kindergarten in fall 2001 and before, and those who entered kindergarten in fall 2006 or later. In the early round of kindergarten readiness assessments, teachers administered a readiness checklist of academic and behavioral skills designed by the state Department of Education with a dichotomous ready/not-ready measure recorded in state records. In the later round of kindergarten readiness, the state universally implemented the DIBELS assessment aimed at measuring early pre-literacy skills; this is a discrete measure that we dichotomize using the approach described in Figlio et al. (2013) so that the percent identified as kindergarten ready corresponds to the percentage in the later assessment. The birth cohorts between 1994 and 2002 who took the kindergarten readiness assessment, therefore, are those born between 1994 and 1996 and those born between 2000 and 2002.

While our preferred specification involves controlling for family fixed effects, it is useful to observe the degree to which our estimated relationships change depending on our inclusion versus exclusion of these fixed effects. We therefore report OLS estimates as well as family fixed effects estimates

[37] High school graduations are not plotted by SES and household structure because these data on SES and household structure are are not available for the 1992 and 1993 birth cohorts, i.e., the cohorts for whom we currently observe high school graduation. .

for kindergarten readiness and the first of our primary outcomes—absences. The first column of Panel A in Table 7 reports conditional mean gaps in kindergarten readiness rates between boys and girls according to mother's race-ethnicity-nativity, where the conditioning variables include mother's education, marital status at birth (married, father present, father absent), zip code income tercile, mother's age at birth, and child's birth order, birth month and birth year. White, non-Hispanic, non-immigrant boys are 5.8 percentage points less likely to start school ready for kindergarten than are demographically similar girls, and the gender gap for blacks is 2.1 percentage points larger.

When we augment this basic model with interactions between the boy indicator and variables that proxy for family advantage/disadvantage at birth (parental status and maternal education) in column 2, we note two fundamental points: first, the interaction terms between boy and various measures of advantage are uniformly positive, indicating that the boy-girl gap in kindergarten readiness is larger in disadvantaged families. Second, we observe that the estimated interaction between boy and black falls by two-thirds, from $-2.09$ to $-0.70$ percentage points, indicating that the bulk of the black-white difference in the gender gap is explained (in the sense of our decomposition above) by differences in family advantage. Observe further that the column 2 model allows race, ethnicity, and immigrant status to exert an independent effect on the gender gap in kindergarten readiness; thus, the interaction terms capture the *differential* effects of parental status and maternal education on boys relative to girls *within* white, black, Hispanic and immigrant families.[38] To simplify exposition, column 3 combines the education and family structure/marital status variables into a single SES index as above; the coefficient on boy $\times$ black remains substantially below the $-2.09$ observed in Column 1, and the interaction between boy and the SES index suggests that SES plays a large role in the boy-girl gap in kindergarten readiness: a one-standard deviation increase in the SES measure ($\sigma = 1.44$) predicts a 1.3 percentage point reduction in the boy-girl kindergarten readiness gap. This patterns of results stands in marked contrast to the *absence* of an SES gradient in the gender gap in birth outcomes. This presents the first piece of evidence that the SES gradient in the gender gap in outcomes is attributable to *post-natal* factors.

The next three columns in Table 7 repeat these estimates on the matched siblings subsample, while the final three columns augment the matched siblings models with family fixed effects. Because of the timing of when kindergarten readiness assessments were administered, we have a limited set of siblings for this analysis.[39] Nonetheless, our basic findings hold in the population for which we can identify siblings and measure kindergarten readiness. With and without family fixed effects, the matched-siblings estimates reinforce the prior conclusions: the boy-girl gap in kindergarten readiness is larger among black and Hispanic families than among white families; conditional on race and ethnicity, the boy-girl gap is larger in father-absent and low maternal education families; and the composite SES index is a powerful predictor of the gap. Figure 3 visually depicts the SES gradient in the gender gap by plotting a bin-scatter of the boy-girl gap in kindergarten readiness

[38]The model permits these factors to have differential impacts by gender but not by race-ethnic group. We relax that assumption in section (5).

[39]Since kindergarten readiness is only available for three adjacent years of birth, we can only compare closely-spaced siblings, as well as a limited set of siblings born numerous years apart.

against the composite SES index while conditioning on all of the covariates used in the corresponding estimate in the table (column 3 of panel C), including, most importantly, race and ethnicity. This figure makes plain the beneficial effect of family advantage on the kindergarten readiness of boys relative to girls. The readiness gap is close to 8 percentage points in the lowest SES bin versus 2 points in the highest bin.[40]

Following our discussion in section 2.1, one benchmark for interpreting the economic magnitudes of these gradients is to scale them relative to the observed gender gaps among blacks, whites and Hispanics. Using our preferred mother-FE models, we find in column 1 of Panel C that the black and Hispanic boy-girl disadvantage in kindergarten readiness is 2.93 and 3.55 percentage points *larger*, respectively, than the comparable boy-girl gap of 5.47 among whites.[41] When in columns 2 and 3 we account for interactions between boy and family disadvantage (as a set of discrete measures or using the composite SES index), the differential disadvantage of boys relative to girls in black households is reduced by nearly one-half to two-thirds. For children of black mothers, this differential disadvantage falls from 2.93 in the non-interacted model to 1.56 in the model that includes discrete boy times family interaction and to 1.52 in the composite model. Under Assumption 1, we can infer that half or more of the additional male disadvantage in kindergarten readiness in black families is due to the differential adverse effect of disadvantage on boys relative to girls rather than on factors specific to black families per se. The apparent role that relative advantage plays for Hispanic families is smaller, though not trivial.

In summary, by age five a boy-girl gap has opened up in kindergarten readiness. The gap is larger for black and Hispanic families than it is for white families, and a large fraction of the racial differences in the gap are accounted for by differences in family advantage. We next turn to the set of outcomes for which we observe the full set of birth cohorts between 1994 and 2002, and therefore, a larger and much more complete set of siblings.

**Absences, suspensions, and test scores in grades 3 through 8**

Table 8 presents parallel findings for the gender gap in absence rates. We focus our discussion on the preferred matched-sibling family fixed effects models, though Table 8 also reports models for singletons, and for matched-siblings without family fixed effects. Column 1 in panel C summarizes conditional gaps in absence rates between boys and girls according to mother's race-ethnicity-nativity, where we condition on the same variables used before. The estimated main effect for the boy indicator variable in this model indicates that white, non-Hispanic, native boys have grade three through eight absence rates that are, on average, 0.21 percentage points higher than those of their sisters. The interaction terms between the race-ethnicity-nativity dummy variables and the boy indicator reveal that the boy-girl gap in absence rates is larger again by 0.30, 0.10 and 0.12 percentage points for black, Hispanic, and immigrant siblings.

[40] The figure also reports the slope of an OLS regression fit to the points in the bin-scatter. While this slope is close to the corresponding estimate in Table 7, it will not match exactly given the differences in specification.

[41] The boy-girl gap is actually slightly smaller for immigrants, but the difference is far from statistical significance.

Following the analysis of kindergarten readiness, Column 2 augments the model with interactions between the boy indicator and variables that proxy for family advantage/disadvantage at birth. Consistent with our findings for kindergarten readiness, each measure that predicts greater family advantage predicts a reduction in the gender gap in school absences. Relative to children born to absent fathers, the boy-girl gap in absences is 0.17 percentage points lower where the father claimed paternity, and 0.31 percentage points lower where the parents were married at birth. Higher levels of maternal education predict additional reductions in this gap in the range of 0.03 to 0.16 percentage points. Because these effects are additive, they make a large net contribution. The model implies that the boy-girl gap in school absences is approximately half a percentage point lower among children born to married parents with college-educated mothers than born to father-absent families with non-high school mothers. Conditional on these measures of education and parental status, the excess boy disadvantage among black boys relative to black girls is reduced from 0.30 to 0.12 percentage points.

Column 3 replaces the education and parental status variables with the composite SES measure, which has a substantial and well estimated relationship to the gender gap. A one standard deviation increase in this index predicts a 0.13 percentage point reduction in the boy-girl absence gap. The upper left panel of Figure 4 illustrates the robustness of this relationship through a bin-scatter plot of the boy-girl gap in absences against the composite SES index (following the format of Figure 3). The SES gradient in the gender gap in absences is a pronounced feature of the data.

A comparison of the three panels of Table 8 also confirms the value-added of the sibling contrast approach used in our preferred specification. When we draw identification of boy-control contrasts exclusively from siblings, we find that the gender gradient in family advantage is typically larger than in the singleton OLS or sibling OLS models. Absent any confounding biases, the addition of thousands of family fixed effects to the sibling-pair models (comparing panels B and C) would be expected to attenuate any estimated relationship between family characteristics and the gender gap. That the coefficients of interest increase in magnitude when fixed effects are added suggests that these fixed effects purge confounding cross-family variation.

We repeat this exercise for suspensions, standardized math scores, and standardized reading scores Table 9.[42] Approximately 11 percent of Florida public school children are suspended for at least one day per school year during grades 3 through 8 (Table 5). But suspension rates are more than twice as high for boys as for girls (15.0 versus 7.3 percent), and the boy-girl differential is twice as large among blacks as whites (12.5 vs. 6.3 percent). The regression estimates in panel A of Table 9 confirm that this pattern is replicated within families: white boys are 6.5 percentage points more likely to be suspended than their sisters, while black boys are 12.8 percentage points more likely to be suspended than their sisters.

The second column of panel A shows that, conditional on race-ethnicity-nativity, the boy-girl gap

[42] Estimates with ancillary estimates for the singletons super-sample and for matched-siblings without family fixed effects are available in a separate reference appendix. As is with the estimates in Table 8, we find qualitatively comparable results in this expanded set of estimates, with highest precision for the main coefficients of interest found in the fixed effects models.

in suspensions is far smaller in families where children are born to married parents, where mothers are better educated, and where household income is higher. Accounting for race and ethnicity, the column 2 estimate implies that the brother-sister gap in suspensions is nearly 10 percentage points *lower* in the most relative to the least advantaged families (i.e. college educated mother married at childbirth versus high school dropout mother and unclaimed paternity). Collapsing our multiple indicators of family disadvantage into the composite SES index in column 3 yields strong evidence of an SES gradient in the gender gap: each standard deviation increase in disadvantage (a reduction in the SES index of 1.44) increases the within-family boy-girl gap in suspensions by 2.8 percentage points. Panel B of Figure 4 depicts this striking relationship using a bin scatter.

What is the economic magnitude of this relationship? Comparing the Boy × Black coefficients across the three columns of panel A indicates that accounting for the differential effect of disadvantage on behavioral outcomes of boys goes a long way towards explaining the differential gender gap in suspensions among blacks relative to whites. The differential black gender gap in suspensions falls by at least 50 percent—from 6.5 percentage points in the baseline model to 3.0 in both the multiple-indicator and composite SES models—when accounting for the differential impact of family disadvantage on boys.

Panels B and C of Table 9 perform the same analysis for standardized reading and mathematics tests administered annually during elementary and middle school (with corresponding bin scatters in panels C and D of Figure 4). Relative to siblings in white families, black, Hispanic and immigrant boys perform less well than their sisters in mathematics and reading. Distinct from the three outcomes analyzed above (kindergarten readiness, absences and suspensions), we find that family disadvantage contributes only modestly to the cross-race and cross-ethnic group variation in the gender gap in these educational measures. The various measures of family advantage generally have the expected sign (reducing relative boy-girl disadvantage in math and reading), but these measures are not consistently statistically significant. When combined into a composite SES measure, the composite measure is also statistically significant but of modest economic magnitude. Accounting for SES differences explains no more than 20 percent of the excess black-white gender gap in test scores.

Causal interpretation of these estimates hinges on the assumption that the family disadvantage gradient in children's outcomes detected in elementary and middle school outcomes is not *already* present at birth. We established above that birth weight—our strongest measure of latent child potential—varies systematically with race and ethnicity, while the gender gap in birth weight is almost entirely uncorrelated with family disadvantage. As a further stress test for this finding, we estimate in Table A4 a set of augmented estimates for the relationship between SES and the sibling gap in kindergarten readiness, and behavioral and educational outcomes that, in addition to prior covariates, control for the log of birth weight and its interaction with gender. For comparison, the bottom panel of the table also reports the coefficient on SES times gender from the corresponding model excluding birth weight. Consistent with expectations, children with higher birth weight attain lower absence rates and higher math and reading scores during third through eight grades (though also higher suspension rates). However, inclusion of infant birth weight has essentially no discernible impact on the estimated SES gradient in the gender gap in behavioral and educational outcomes.[43]

We have performed numerous extensions that elaborate and corroborate these basic findings, results of which are available as reference tables. To analyze the impact of family structure on the gender gap among siblings that experience arguably comparable family circumstances, we focused on both closely-spaced siblings and on the first two siblings observed in larger families. We find comparable impacts of family disadvantage in these subsamples. We also find highly comparable results when we assign each child born to a given mother the marital/paternity status and maternal education observed at the first birth. To further unpack the complex role played by parental structure, we present a set of estimates for behavioral outcomes (i.e., absences and suspensions) in Table A5 where we limit the sample to siblings born to the same mother with distinct fathers, then subdivide further into cases where family structure differs across births (married, father present, father absent) and cases where it does not.[44] Among siblings born to the same mother with distinct fathers, these results confirm that marital status and, to a lesser degree, father presence, differentially reduce absences and suspensions among brothers relative to sisters. These impacts are generally larger in same-status relative to divergent-status families, plausibly because divergent paternal status indicates higher household instability overall. Thus, even in complex family environments where paternity differs across siblings, boys are found to differentially benefit from higher paternal commitment, indicated by marriage or paternity acknowledgment.

We also tested how the SES gradient in the gender gap in early outcomes evolves between grades three and eight, using successive observations across grades (again, contrasting within families). We find that both the SES gap and the gender gap in absences, suspensions, and test scores increases year-over-year across grades within sibling pairs—that is, low SES children fare increasingly poorly relative to high SES children, and boys fare increasingly poorly relative to girls. We do not, however, find a consistent triple-interaction between these forces. While low SES children diverge from high SES children, and boys diverge from girls, the divergence is not consistently greater again among low-SES boys relative to high-SES boys as compared to low-SES girls relative to high-SES girls.[45]

### Physical and cognitive/behavioral disabilities

The Florida Department of Education records whether children have any disability that could influence their educational attainment and life chances. We analyze the SES gradient in the gender gap in disability as a complement to the behavioral and disciplinary outcomes above. To the extent that cognitive and behavioral disabilities are more sensitive to environmental factors than are physical

[43]We have also estimated models that control for a wider variety of birth outcome measures and their interactions with infant gender, and the results are fundamentally unchanged. Models with second and third level interactions between low birth weight (a dummy variable), SES, and gender find no consistent pattern of effects in sibling fixed effects models. While low birth weight is generally predictive of less advantageous behavioral and academic outcomes, it does not have a strong or consistently interaction with either SES, gender, or SES $\times$ gender. These reference tables are available in a separate document.

[44]In each instance where no paternity is claimed on a birth certificate, we treat the father as distinct.

[45]We report this result for completeness, but we did not have a strong prior that this pattern should be present, particularly given that the third-level, within-family comparison makes substantial demands on the data.

disabilities, we expect to see an SES gradient in the gender gap in the former category of disability than in the latter. An unusual feature of the disability data is that, due to Florida's McKay Scholarship Program, *every child* with a disability who is living in the state of Florida during school-going age is recorded in the school records. Thus, if a family sends a disabled child to private school and a non-disabled child to public school, both children are observed in our data.[46]

Table 10 presents family fixed effects models for all disabilities (panel A), cognitive/behavioral disabilities (panel B), and physical disabilities (panel C). We characterize disabilities according to the definitions identified under Florida statute.Cognitive and behavioral disabilities are language impaired; intellectual disability (subdivided in the early years of the data into educable mentally handicapped, trainable mentally handicapped, and profoundly mentally handicapped); developmental delay; specific learning disability; emotionally handicapped; severely emotionally disturbed; and autism spectrum disorders. Physical disabilities are orthopedically impaired; speech impaired; deaf or hard of hearing; visually impaired; hospital/homebound; dual-sensory impaired; traumatic brain injury; and established conditions.

The prevalence of disability in our matched siblings sample is 23 percent, with 9 percent of children possessing an identified physical disability and 14 percent with a cognitive/behavioral disability.[47] Within families, boys are substantially more likely to be classified as physically or cognitively/behaviorally disabled than are their sisters. These gaps also differ across race and ethnic groups. The boy-girl gap in cognitive/behavioral disability is almost twice as large among black as white siblings: 11.5 percentage versus 7.3 percentage points. However, the boy-girl gap in physical disability is slightly smaller among blacks than white siblings: 4.5 versus 5.7 percentage points.

The regression estimates in the second and third columns of each panel detect a strong SES gradient in the prevalence of disability (panel A), which is driven almost entirely by cognitive/behavioral disabilities (panel B). Boys in families with married mothers are significantly less likely to be classified as cognitively or behaviorally disabled. Similarly, there is a strong negative relationship between maternal education and the female-favorable gap in cognitive and behavioral disabilities. The interaction between boy and the composite SES index confirms this strong relationship: a one standard deviation increase in the SES index predicts a 1.4 percentage point reduction in the excess probability that a boy is classified as cognitively or behaviorally disabled (a 10 percent reduction on the base of 14 percent). This SES gradient explains a large fraction of the excess gender gap in cognitive and behavioral disabilities among blacks relative to whites: accounting for SES reduces this excess gap from 4.2 to 2.6 percentage points—a 38 percent reduction—with a similar reduction in the excess prevalence of any disability.

Conversely, there is no evidence of an SES gradient in the gender gap in physical disability.

[46]The McKay Scholarship Program provides a taxpayer-supported school voucher to attend private school to every child with a disability whose family considers the public options to be unacceptable, regardless of family resources or need. In practice, well over 90 percent of all children with disabilities in Florida attend public schools.

[47]Children can potentially have both physical and cognitive/behavioral disabilities. We assign each child with a disability to the major category (physical vs. cognitive/behavioral) first observed (temporally) in their school record. In practice, this breakdown is nearly identical to that which is seen if we instead assign each child to their last-observed major category of disability.

Indeed, while none of the coefficients on the interactions between SES measures and gender are statistically distinct from zero, the evidence suggests that, if anything, physical disabilities are slightly more prevalent among boys than girls in high-SES households.[48]

In summary, we find a large and pronounced SES gradient in the gender gap in cognitive and behavioral disabilities—which are more likely to be influenced by post-natal environmental factors—but no corresponding SES gradient in the gender gap in physical disabilities—which are more likely to be determined by prenatal factors. We view this evidence as supporting the inference that post-natal environmental factors are responsible for the gender gap in educational and behavioral outcomes.

## 4.2 Downstream outcomes: High school completion, criminal behavior, and incarceration

Kindergarten readiness, elementary and middle school disciplinary behaviors, attendance, test scores, and disability are not market outcomes. While these measures are strongly correlated with later labor market success, it is not immediately obvious whether the impact of family disadvantage on these measures is consequential for downstream market outcomes such as educational attainment and earnings. To provide direct evidence this question, we focus on a market outcome that is available in our data: high school graduation.[49] Specifically, we ask whether family disadvantage is predictive of the gender gap in high school graduation, and whether the impact of family disadvantage on kindergarten readiness and third through eight grade behavioral and test scores can account for this relationship. Of necessity, this analysis is limited to the oldest cohorts in our sample, principally those born in 1992 and 1993, for whom on-time graduations are presently available. Our data for these early cohorts do not unfortunately permit sibling linkages, and they lack information on paternity status among those with out of wedlock births. They nevertheless prove quite informative.

Table 11 explores the relationship between family advantage and the gender gap in high school graduations (summarized in Table 6 above) by comparing two OLS models for high school graduations in our Florida sample: one containing the full set of family, mother, and child controls used earlier and a full set of boy × race-ethnicity-nativity dummies; the second containing these covariates plus a set of interactions between gender and family advantage (here, mother's education and her marital status at the time of childbirth). The first column of panel A shows that the conditional mean on-time high school graduation rate of white boys is 6.0 percentage points below that of white girls, and that this gap is 6.7, 2.2, and 3.7 percentage points larger again for black, Hispanic, and immigrant boys, respectively. These very large gender gaps in on-time high school graduation rates—ranging from 6 to 13 percentage points—are about equally composed of higher male dropout rates and higher male grade repetition rates (which may culminate in dropout or high school completion).

The second column of each panel shows that the boy-girl disadvantage in high school comple-

[48] This differential is driven almost entirely by speech impairments. High-SES families appear to be more likely to identify and seek treatment for speech impairments, which are either more prevalent or more often treated in boys.

[49] While high school diplomas are is not allocated by a market mechanism, high school degrees and priced in the labor market (see Autor 2014 for discussion).

tion is substantially smaller in more advantaged families. Conditional on race-ethnicity-nativity, the gender gap in on-time completions is 3.1 percentage smaller for boys born to married mothers, and 3.2 percentage points smaller again for boys with college-educated relative to high school dropout mothers. These are large differentials. Given the substantial differences between whites, blacks, Hispanics, and immigrants in maternal education and in-wedlock childbearing (Table 2), it follows that the pronounced race and ethnic differences in the gender gap in high school graduations are explained in part by the substantial differences in family structure and education across these groups. A comparison of the boy × race/ethnicity dummies in the even and odd-numbered columns confirms this intuition. Accounting for mother's education and marital status at child-birth reduces the black boy-girl disadvantage in high school graduations from 6.7 to 4.6 points (30 percent), reduces the Hispanic boy-girl disadvantage from 2.2 to 1.5 points, and reduces the immigrant boy-girl disadvantage from 3.7 to 3.2 points.[50] This exercise suggests that family disadvantage makes an important contribution to observed cross race-ethnicity-nativity gender gaps in high school completion. We also conjecture that it contributes to gender differentials in downstream market outcomes such as college-going and earnings, though these outcomes are not observable in our data.

How do these findings for the gender gap in high school completions compare to our earlier findings for the SES gradient in the gender gap in kindergarten readiness, third through eight grade disciplinary and educational outcomes. To explore that question, we first estimate the predictive relationship between the K through 8 outcomes analyzed above. We then use our earlier estimates in combination with these simple predictive models to obtain an implied effect of family disadvantaged on high school graduations operating through these channels.[51] Table 12 summarizes this exercise. The first panel presents a linear probability regression of high school on-time graduation on our standard set covariates (child sex, race-ethnicity-nativity interacted with sex, birth order, birth year, and birth month, and mother's education, age and marital status at childbirth), augmented with a vector $T_i$ of elementary and middle school behavioral and educational outcomes. These child-specific outcomes include: kindergarten readiness, absence rates, suspension rates, standardized math scores, and standardized reading scores, each averaged over grades three through eight:

$$\begin{aligned} H_{ij} \quad = \quad & \gamma_j + T_i'\pi + \beta_1 \text{Boy}_i + \beta_2(\text{Boy}_i \times \text{Black}_j) + \beta_3(\text{Boy}_i \times \text{Hispanic}_j) + \beta_4(\text{Boy}_i \times \text{Immigrant}_j) \\ & + F_j'\psi + X_i'\lambda + e_{ij}., \end{aligned} \tag{8}$$

To facilitate comparison across coefficients, we standardize each of the predictive variables in $T$ to have mean zero and unit variance.Thus, regression coefficients correspond to standardized effect sizes.

Estimates of (8) in Table 12 find a highly significant predictive relationship between early behavioral and educational outcomes and subsequent on-time high school completions. Whether entered into the regression individually (columns 1 through 5) or as a group (column 6), the data unambiguously demonstrate that children who have higher absence and suspension rates and lower math

[50] We do not apply the SES index to this exercise because we lack information on paternity status for the the cohorts that have so far reached high school completion age.

[51] Our models for the predictive relationship between K-8 outcomes and subsequent high school graduations should be understood as descriptive rather than causal. We do not wish to claim that any causal effect of family disadvantage on high school outcomes runs entirely through K-8 behavioral and academic outcomes.

and reading scores during grades three through eight, as well as lower kindergarten readiness, are significantly less likely to complete high school within four years. Notably, the standardized effect sizes of early behavioral measures (absences and suspensions) are substantially larger than the effect sizes of math and reading scores, which are in turn larger than the effect sizes of kindergarten readiness and preschool attendance.

To interpret the economic magnitude of these coefficients in relation to our earlier findings, we scale them by the estimated impact of SES on each behavioral and educational outcome, collected from Tables 7 through 9 and reported (in standardized form) in panel B of Table 12). The final panel of Table 12 reports the implied impact of SES on high school graduations scaled by four metrics: a one-standard deviation increment to SES (row 1); the mean white-black SES differential (row 2); the mean white-Hispanic SES differential (row 3); and the mean white-immigrant SES differential (row 4). This set of calculations yields two key findings. First, the implied impact of family disadvantage on the gender gap in high school graduations operating through early behavioral and educational outcomes is large. A one-standard deviation reduction in family SES is predicted to raise the boy-girl high school graduation deficit by approximately 1.65 percentage points. Since cross-group differences in SES are substantial, this mechanism also contributes to observed race and ethnic gaps in on-time high school graduations. The mean black-white, Hispanic-white, and immigrant-white SES differentials of $-1.60\sigma$, $-0.56\sigma$, and $-0.33\sigma$, respectively, can explain 1.69 points of the 6.7 point excess boy-girl deficit in HS graduations among blacks, 0.6 points of the 2.2 point excess boy-girl deficit among Hispanics, and 0.35 points of the 3.7 point excess boy-girl deficit among immigrants.

A second finding is that the primary channel through which these effects appears to operate is behavioral. Variation in the gender gap in reading and math achievement and kindergarten readiness make a negligible contribution to the gender-gap in HS graduations, jointly accounting for only 11 percent of the total explained by the third through eighth grade measure (versus 89 percent explained by absences and suspensions). A plausible interpretation of this finding is that, at least for boys, it is behavioral rather than formal skills deficits that inhibit high school completion, as argued for example by Heckman and Kautz (2012). Indeed, of the four measures considered, suspensions play the largest role in explaining the boy-girl deficit in high school completions (panel C). Moreover, when we estimate models akin to panel A of Table 12 for high school dropout and five-plus years of high school attendance, we find that three-quarters of the estimated impact of behavioral measures on non-completion operates through dropout.

### Juvenile crime and incarceration

As one extreme manifestation of these behavioral gradients, we consider the odds of having "substantial involvement" in criminal activity, leading to multiple felony convictions or incarceration. The Florida Department of Education data do not provide information on participation in the juvenile justice system, and the children born from 1994 onward are typically too young anyway to have many sibling pairs where multiple children were old enough to have had time to engage in serious delinquent behaviors. However, a large unnamed Florida school district has linked their data to juvenile criminal justice data, and recorded an indicator for incarceration or repeat felonies by the child's 16th birthday (the mandatory school attendance age in the state). This school district maintained these data in the 1989-90 school year through the 2004-05 school year, and we match siblings in these data based on children who are consistently co-resident across their years of observation. For most families with a child in the 1989-90 kindergarten cohort in this large county we are able to conduct sibling comparisons of the likelihood of juvenile justice system participation.[52] We observe this information for 8,421 children in 4,056 unique families. Juvenile justice system is on average three times as high among boys as girls, 4.5 versus 1.5 percent, in this Florida county.

Because the rate of juvenile system involvement are so much greater among boys than girls, we consider the odds of participation in juvenile detention among boys relative to girls for matched sibling pairs in our sample.[53] Figure 5 plots these odds ratios according to birth characteristics: race (black or white), marital status (married/unmarried), and maternal education (high school dropout, high school graduate, college graduate). We order these observations on the x-axis of Figure 5 according to mean family SES for each of these twelve demographic groups calculated from our primary sample. For a few groups, we observe no female juvenile justice participation and so cannot compute odds ratios. Among matched sibling paris, the odds ratio for juvenile detention is steeply increasing in our proxies for family disadvantage. For example, among children of white, married, college-educated mothers, this ratio is approximately unity. Among white and black children of unmarried, high school dropout mothers, this ratio is between five and six.[54]

In summary, we observe that the patterns of juvenile justice participation mirror those observed in school outcomes: There is an SES gradient in the gender gap, with boys getting into more trouble relative to their sisters in low-SES households than is observed in high-SES households.

## 5 Testing the Oaxaca restrictions

Our econometric framework imposes the restriction that the differential gender impact of SES on behavioral and educational outcomes is constant across race-ethnicity-nativity groups, as noted in

[52] For the kindergarten cohort of 1989-90, the school district surveyed the universe of parents of incoming kindergarteners, and achieved around a 70 percent response rate. This survey sample frame allows us to assign to each sibling the maternal SES measures (maternal marital status and education) observed at the time of the 1989-90 kindergartener's birth. We limit our analysis to children born on or before August 31, 1989 so that we can observe juvenile justice data through their 16th birthday. A child must have a sibling in the 1989-90 kindergarten class in order to for us to have an SES measure.

[53] Although we cannot compute odds ratios for single sibling pairs, these pairs are balanced within each demographic cell according to race, marital status, and maternal education.

[54] In an available reference table, we perform linear probability estimates with family fixed effects for the relationship between SES measures and the gender gap in juvenile justice system participation. The female-favorable gender gap in juvenile system involvement is slightly smaller among children of married versus non-married mothers, but gap is substantially smaller for children of mothers with a high school or college education. Accounting for the strong SES gradient in the gender gap makes only a modest reduction (on the order of 10 percent) to the excess gender gap in juvenile justice participation among blacks relative to whites. Thus, there are large race difference in the gender gaps (in levels) in juvenile system involvement that are not explained by our SES proxies. Because the odds ratio plot implicitly adjusts for differences in base rates across demographic groups, we see no strong racial pattern in juvenile justice participation in Figure 5 beyond that predicted by the SES composite.

section 2.1. If the impact of SES on the gender gap were to differ substantially by demographic group, this would complicate interpretation of an SES 'effect' on the cross race-ethnicity-nativity gender gap.

We test this restriction here by estimating an augmented version of equation (7) that includes all second and third-level interactions between gender, SES, and race-ethnicity-nativity:

$$\begin{aligned} Y_{ij} &= \gamma_j + \beta_1 \text{Boy}_i + \beta_{j,2}(\text{Boy}_i \times \text{Race}_j) + \beta_{j,3}(\text{Race}_j \times \text{D}_j) \\ &\quad + \beta_{4,j}(\text{Boy}_i \times \text{Race}_j \times \text{D}_j) + F_j'\psi + X_i'\delta + e_{ij}, \end{aligned} \tag{9}$$

where $\text{D}_j$ is the SES (disadvantage) measure and we suppress notation of main effects of race-ethnicity-nativity and disadvantage.[55] Tables A6 and A7 present these augmented specifications.

Focusing first on kindergarten readiness in Table A6, we find no evidence of cross-race-ethnicity-nativity variation in the relationship between SES and the gender gap. A Wald test for cross-race-ethnicity-nativity heterogeneity in the SES gradient in the gender gap accepts the null at $p = 0.92$. Among the educational and behavioral outcomes from third through eighth grade for absences and reading scores, there is again no evidence of differential effects of SES on the gender gap by race-ethnicity-nativity. For suspensions, there is statistically significant heterogeneity in slopes, but its economic magnitude is modest, less than one-third of the size of the main effect. For math scores, we also detect significant heterogeneity: higher SES is particularly advantageous for math performance of boys relative to girls in immigrant families. However, recall from Table 12 that math and reading scores make almost no contribution to the SES gradient in gender gap in downstream outcomes (high school on-time completion, etc). Thus, we believe that this deviation from the constant effects assumption is not substantively significant here.

Table A7 repeats this exercise for on-time high school graduation, high school grade repetition, and high school dropout. Because we do not have all components of the SES variable for the high school age cohorts in our sample, we instead include a full set of main effects and second and third-level interactions for marital status and mother's education (a total of nine third-level interactions among disadvantage, boy, and race-ethnicity-nativity). For on-time high school graduation, we readily accept the null of no heterogeneity in the family advantage gradient in the gender gap ($p \geq 0.68$). Among those not completing high school on time, we find some evidence that Hispanic boys in married families are differentially likely to repeat a grade rather than drop out of high school.[56] Even then, we accept the null at $p \geq 0.07$ that all third-level interactions are jointly zero for grade repetition and high school dropout.

In sum, while there is some evidence of heterogeneity among race-ethnicity-nativity groups in the relationship between family disadvantage and the gender gap in behavioral and educational outcomes, this heterogeneity is quite modest overall. And when we focus on the key market outcomes available to our analysis—high school completion—we find no significant heterogeneity whatsoever.

[55] Regression specifications include main effects for all variables that can differ across siblings, including family disadvantage measures (since, e.g., marital status may change between births).

[56] Murnane (2013) observes that black and Hispanic students are more likely than whites to complete high school in five-plus years. Thus, race gaps in eventual high school completion are smaller than race gaps in on-time graduation.

# 6 Exploring mechanisms: Schools and neighborhoods

Family disadvantage may expand boy-girl disadvantage not because boys are more affected by family environment per se but because the neighborhoods and schools in which disadvantaged children are raised are particularly adverse for boys. For example, boys may be more vulnerable to the risks presented by low-SES neighborhoods and schools (e.g., exposure to gangs and violence); or, boys may be treated relatively harshly by authority figures in these settings (e.g., boys face greater disciplinary and criminal sanctions from teachers and police). Although we cannot assess *why* schools and neighborhoods might matter differently for boys, we can test whether family advantage operates primarily through these channels rather than directly, through family environment per se.

This final empirical section tests whether school and neighborhood can account for the gender-SES gradient in children's outcomes that we have so far attributed to family advantage. For this analysis, we augment our data with measures of school and neighborhood quality. For school quality, we use direct assessments provided by the Florida Department of Education, which assigns each Florida school a letter grade of A, B, C, D or F. We distinguish between schools with a grade of A or B versus the remainder (those rated C, D or F).[57] We use two sets of variables for neighborhood quality: the median income of the zip code of residence observed at birth; and a measure of the causal effects of place of residence on economic mobility by Florida county (Chetty and Hendren, 2015). The Chetty-Hendren mobility measure corresponds to the estimated percentage gain (or loss) in income at age 26 from spending one more year of childhood in each county in the U.S. for a child whose family is at the $25^{th}$ percentile of the national family income distribution.[58]

Table 13 presents estimates of augmented models for the impact of family, school, and neighborhood on gender gaps in behavioral outcomes (absences and suspensions). Columns 1 and 2 replicate our baseline specification for the gender gap in school absences, while subsequent columns augment this specification with school and neighborhood controls, in each case interacted with gender.[59] Column 3 shows that the gender gap in absences is significantly lower in higher income zip codes: an additional $10K in median zip code income predicts a reduction of 0.025 in the gender gap in absences (about 10 percent of the mean boy-girl gap of 0.23 reported in Table 5). The boy-girl differential in absence rates is slightly smaller (less adverse) for siblings raised in more economically mobile neighborhoods (column 4), though this relationship is not statistically significant. The brother-sister gap in absences is substantially smaller, however, for children who attend A or B-rated Florida public schools relative to those attending lower-rated schools (column 5). This reduction of 0.16 points for A schools and 0.08 points for B schools is 35 to 70 percent as large as the overall

[57] Relatively few Florida schools receive grades of D or F.

[58] We thank Jeremy Majerovitz for expert assistance with the Chetty-Hendren data. We use data from their Online Data Table 2: Preferred Estimates of Causal Place Effects by County. We extract the causal exposure effects as percentage gains (or losses) in income at age 26 relative to the national mean for parents at the 25th and 75th percentile of the national household income distribution. We then average the two values to obtain average treatment effect for each county. Due to linearity in ranks this measure is the same as the treatment effect on the median person.

[59] Samples (and hence point estimates) differ very slightly from our main estimates in Tables 8 and 9 because school grades are missing for the small subset of schools observed in their first year of operation, at which point grades have not yet been calculated.

gender gap in absences. The sixth column includes all four school and neighborhood measures simultaneously. School quality remains robustly predictive of the gender gap in absences, while the other variables weaken or remain insignificant.[60]

To interpret the magnitude of these relationships, it is helpful to use two metrics: how much of the SES gradient in that gap do they explain; and how much of the race-ethnicity-nativity gender gap in absences do they explain? The first question can be answered by comparing the coefficients on the Boy × SES interaction term across columns of Table 5 when including versus excluding the neighborhood and school quality interaction terms. The coefficient on Boy × SES of $-0.078$ in the baseline specification falls in magnitude to $-0.062$ in the final specification that accounts for both school and neighborhood quality. Thus, only 20 percent $(1 - 0.062/0.078)$ of the effect of SES on the gender gap in absences is accounted for by measures of school quality and neighborhood. The remainder is implicitly accounted for by family level effects that operate across siblings within schools and neighborhoods.[61]

The answer to the second question—what share of the race-ethnicity-nativity gender gap is explained by family, school, and neighborhood—is summarized in Figure 6. As reported in Table 8 above, the regression adjusted excess gender gap in school absences among blacks relative to whites is 0.29 points. Controlling for the differential impact of family advantage on boys explains 47 percent of this gap. Adding neighborhood income explains an additional 9 points (56 percent in total), while controlling school quality and neighborhood economic mobility increases this explanatory power to 66 percent. In summary, 80 percent of the explained impact of family advantage on the gender gap in absences appears to operate between siblings *within* schools and neighborhoods, and about 70 percent (0.47/0.66) of the explained cross race-ethnicity-nativity variation in the gender gap in absences is accounted for by family-level variation in disadvantage rather than sex differences in the impact of school and neighborhood quality.[62]

Panel B of Table 13 repeats this exercise for grade three through eight suspensions. In this case, we find that both neighborhood income and neighborhood economic mobility are significant predictors of the boy-girl gap in suspensions, with the female-favorable gap smaller in wealthier and more mobile neighborhoods. As with absences, school quality both make a sizable contribution, with suspension rates of boys reduced substantially relative to girls in wealthier zip codes and in higher quality schools. Notably, inclusion of all three neighborhood and school quality measures only minimally affects the family-level impact of SES on the gender gap: the SES × gender interaction term falls by approximately 20 percent in magnitude when conditioning on these detailed measures.

[60] Because neighborhood mobility (based on county) and family SES (based on zip code) differ relatively rarely between siblings, the main effects for these variables are only weakly identified and hence are not included in the table. Main effects for school quality are, however, always large and significant with the expected sign (reducing absences and suspensions, increasing math and reading scores).

[61] This decomposition provides a lower bound on the impact of school and neighborhood since more complete measures of school and neighborhood quality could increase the explanatory power of these constructs. We address this issue non-parametrically in Table A8, summarized below.

[62] Note that neither result is implied by the other. It would be possible, for example, for all of the variation in the gender gap explained by SES to be accounted for by school and neighborhood quality, *and* for SES to account for only a modest subset of the total variation in the gender gap explained by the combination of SES, school quality, and neighborhood.

As reported in panel B of Figure 6, family disadvantage accounts for 50 percent of the excess gender gap among black relative to white siblings, with school and neighborhood quality accounting for an additional 12 points percentage.[63]

Table 14 presents the corresponding analysis for math and reading scores. Relative to the behavioral outcomes, these academic indicators are less central to our analysis since they have comparatively small relationships to high school completion, grade repetition, and dropout. Nevertheless, we find a similar result pattern: greater neighborhood economic mobility and higher school quality significantly reduce the boy-girl disadvantage in early academic performance, though for these outcomes we do not find a significant effect of neighborhood income. Accounting for neighborhood and school quality attenuates the relationship between SES and the gender gap in math and reading. However, *none* of these measures (individually or in combination) accounts for even a third of the excess gender gap in academic outcomes among black relative to white children.

As a final robustness test of these results, we estimate a set of companion models for absences, suspensions, math scores, and reading scores that are saturated with a full set of zip code $\times$ sex and school $\times$ sex interactions. This extension allows the impact of neighborhoods and schools on the gender gap to vary non-parametrically by neighborhood to the level of granularity available in the data. While the inclusion of zip code dummies eliminates both the neighborhood income measure, Chetty-Hendren mobility measure (defined at the county level), and the direct school quality measures, it allows us to retain our explanatory variable of primary interest, family SES (i.e., mother's education and marital/paternity status at birth). These saturated estimates are summarized in Table A8. Comparing across columns of this table documents the robustness of our findings for the contribution of family disadvantage to the gender gap in behavioral outcomes. Controlling non-parametrically for the role of neighborhoods and schools reduces the estimated impact of family disadvantage on the gender gap in absences and suspensions by 30 percent in each case (relative to 20 percent in the Table 13 models that control directly for neighborhood and school quality). For math and reading scores, the attenuation is larger (about two-thirds), and SES is no longer a significant predictor of the gender gap in the saturated models. Family disadvantage appears, however, to account for only a modest component of the gender gap in test scores, as noted above, and test scores in turn play only a modest role in the gender gap in high school graduations.

In net, the results in this section lend substantial support to the hypothesis that early outcomes of boys—particularly, behavioral and disciplinary outcomes—are differentially impacted by disadvantage, whether manifest in family disadvantage, neighborhood quality, or school quality. While our data do not allow us to discern *why* school and neighborhood matter more for boys than for girls, the analysis makes clear that less affluent neighborhoods and lower quality schools generate differentially poor outcomes for boys. Nevertheless, these non-household influences—which are correlated with parental education and marital status—appear far less important for explaining the gender gap, however, than the direct effect of family structure itself.

[63] Our models also tend to over-explain the modest boy-girl gaps in suspensions among Hispanics and immigrants.

# 7 Conclusions

In the U.S. and much of the developed world, boys lag behind girls in numerous measures of educational attainment. The gender gap in U.S. high school and college completion varies also widely by race and ethnicity, with larger gaps among minorities. In this paper, we investigate whether family disadvantage can account in part for the race/ethnicity gradient in these gender gaps in educational outcomes. Utilizing a large data set of Florida births from 1992-2002 linked to schooling records, a parallel analysis of data from a large Florida school district where we can measure juvenile justice participation, and an empirical design that employs within-family brother-sister comparisons, we find that family disadvantage disproportionately negatively affects the academic and behavioral outcomes of school-age boys relative to girls.

By the time a child is five years old, a sizable and pronounced SES gradient in the gender gap has opened up: the less advantaged a family is, along a range of dimensions, the less likely it is that boys will be ready to start school relative to their sisters. The impact of family disadvantage on the outcomes of boys relative to girls is already evident by the time of kindergarten entry, is further manifested in behavioral and educational gaps in elementary and middle-school performance, and crystalizes into sharp differences in high school graduations by age 18 and criminal activity by age 16.These SES gradients are especially strong for behavioral outcomes measured during the third through eighth grades, and though more modest for reading and math test scores, the gradient persists to high school graduation. These differences are largely independent of neighborhood quality and school quality, though we also observe that better-quality schools (as defined by the Florida Department of Education) help to mitigate the gender gap in educational and behavioral outcomes. These estimates imply that a sizable portion of the minority-white difference in educational and behavioral gender gaps is attributable to higher degrees of family disadvantage among minority families—which in turn disproportionately impairs the behavioral and educational advancement of minority relative to white boys.

A central finding of our analysis is that, across a large number of measures of neonatal well-being— measured by birth weight, infant and maternal health, and numerous indicators of prenatal care and medical complications before and during delivery—we detect no evidence that the SES gradient in the gender gap in outcomes is already present at birth. Accepting these results to strongly suggest that there is no *latent* gender gap in potential outcomes that varies with SES, we infer that the divergent behavioral and cognitive development of boys relative to girls in low-SES versus high-SES families reflects the causal effect of the post-natal environment on child development.

While there are numerous pathways through which family disadvantage might influence educational and behavioral outcomes, many discussed above, our paper is deliberately agnostic about the specific channels through which post-natal exposure to family disadvantage generates the observed gender gap in outcomes. Our primary contribution, instead, is to rigorously test and, for the most part, reject two natural candidate explanations. One explanation is a "fetal origins" hypothesis for the gender gap, in which the SES gradient in potential outcomes is imparted prior to birth. Our analysis makes clear that, at the time of delivery, boys born to low-education, unmarried, and father-absent mothers are not systematically less fit or less well cared for than their sisters. The second is an environmental factors hypothesis. We document that neighborhood quality makes a small contribution to the SES gradient in the gender gap, while school quality matters more—low quality schools are particularly disadvantageous for boys. Nevertheless, accounting non-parametrically for the differential impact of schools and neighborhoods on boys relative to girls reduces the estimated impact of family disadvantage on the sibling gender gap by less than a third. Rejecting these alternatives implies that family disadvantage makes both a substantial direct contribution to the gender gap *as well as* an indirect contribution through its influence on schools and neighborhoods.

In summary, the impact of family disadvantage on the outcomes of boys relative to girls is already evident by the time of kindergarten entry, is further manifested in behavioral and educational gaps in elementary and middle-school performance, and crystalizes into sharp differences in high school graduations (as well as criminal involvement) by age 18. While our data do not permit us to quantify the causal effect of family disadvantage on the gender gap in subsequent downstream market outcomes—including college attainment, earnings, and labor force participation—we suspect that these early differences in behavioral and educational outcomes continue into adulthood, as boys and girls exit the compulsory school system and matriculate into employment, higher education, and potentially parenthood.

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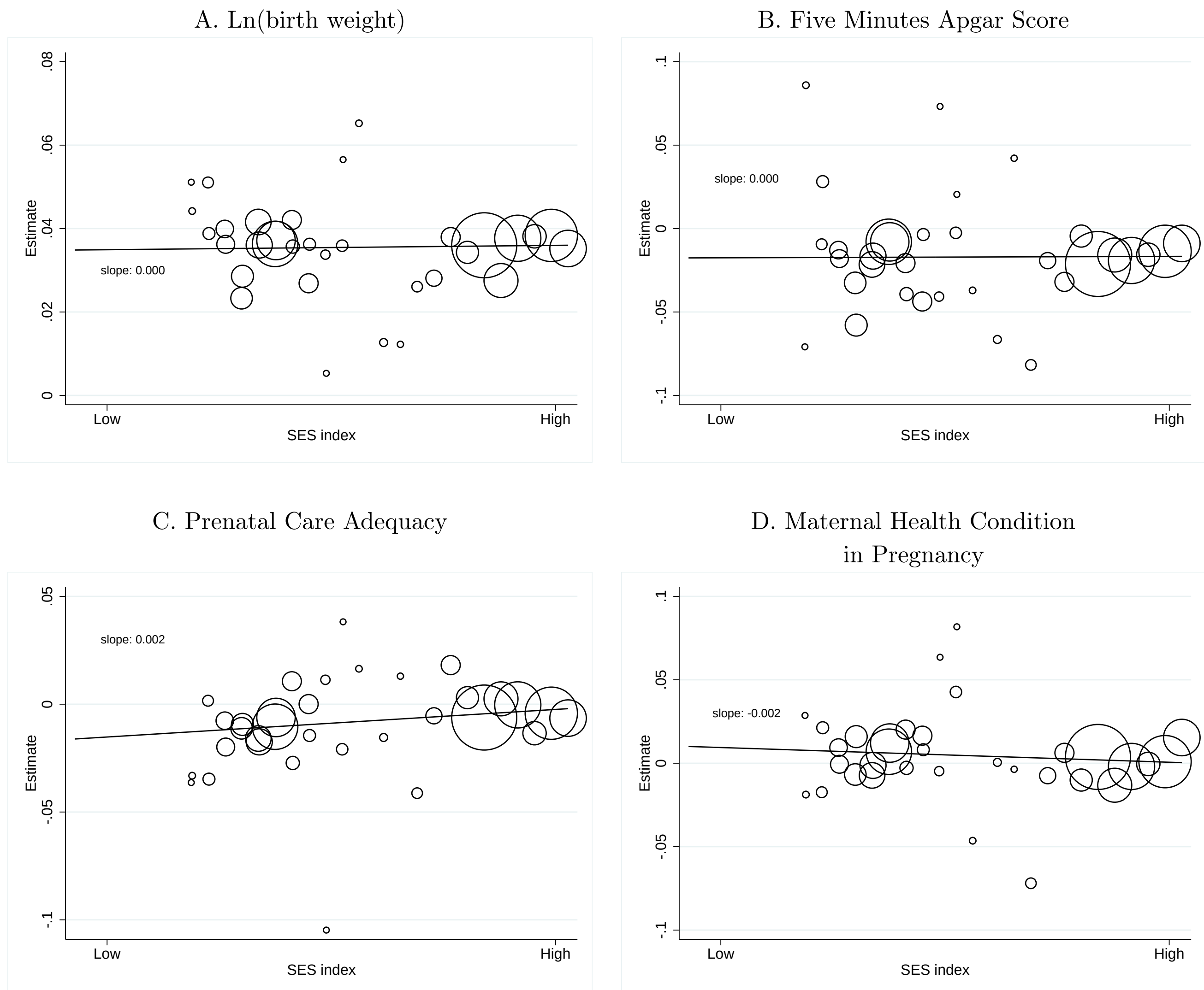
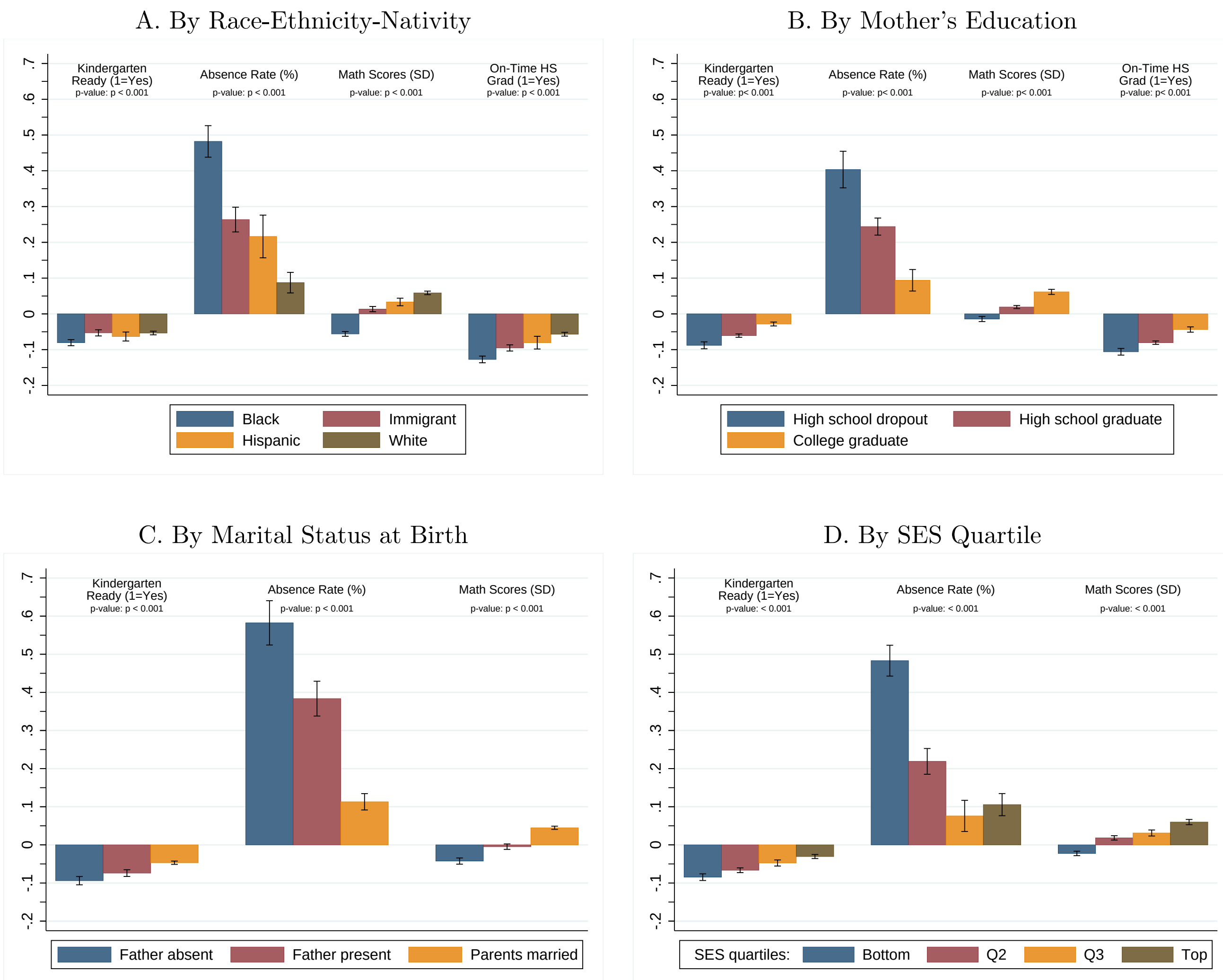
# Figures and Tables

Figure 1: Boy-Girl Gaps in Ln(birth weight), Five Minutes Apgar Score, Prenatal Care Adequacy and Maternal Health Condition in Pregnancy by SES



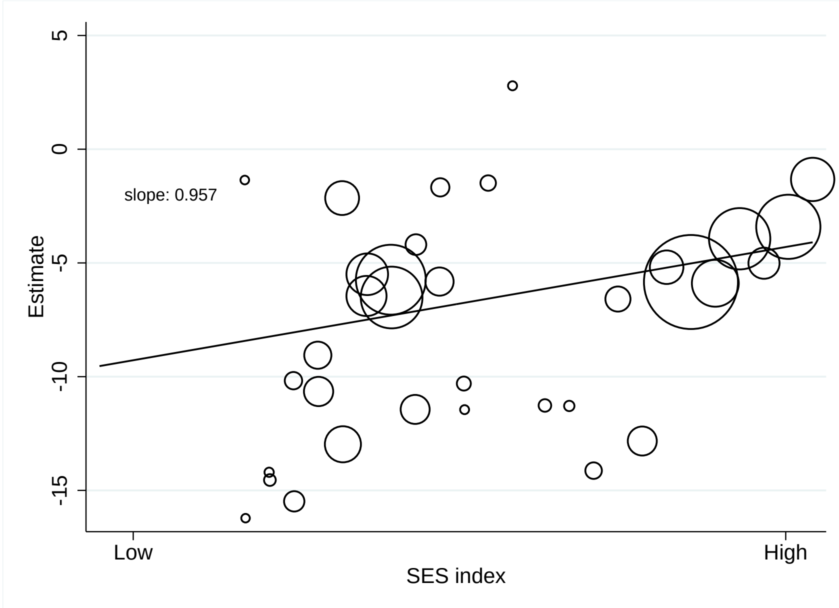

Note: Mean ln(birth weight), five minutes APGAR score, adequacy of prenatal care and maternal health condition in pregnancy by gender and race are computed from the sample of non-twin siblings born between 1994-2002. Prenatal care adequacy is defined according to the Kessner Adequacy of Prenatal Care Utilization index (APCU), which is a binary indicator equal to one if the mother received standard prenatal care services during pregnancy (and zero otherwise). Mother health issues during pregnancy is an indicator variable equal to one if the mother suffered from any of a large set of chronic or pregnancy-related disorders (anemia; cardiac disease; acute or chronic lung disease; diabetes; genital herpes, hydramnios/oligohydramnios; hemoglobinopathy; chronic hypertension; pregnancy-associated hypertension; eclampsia; incompetent cervix; previous infant 4000+ grams; previous preterm or small for gestational age infant; renal disease; RH sensitization; uterine bleeding; other specified health problem) during pregnancy or delivery. Regression-adjusted male-female gaps are plotted for 54 bins of our SES index, plotted against the mean values of the SES index in each of those bins. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth.

Figure 2: Boy-Girl Gaps in Absences, Math Scores, On-Time High School Completion and Kindergarten Readiness by Family Characteristics



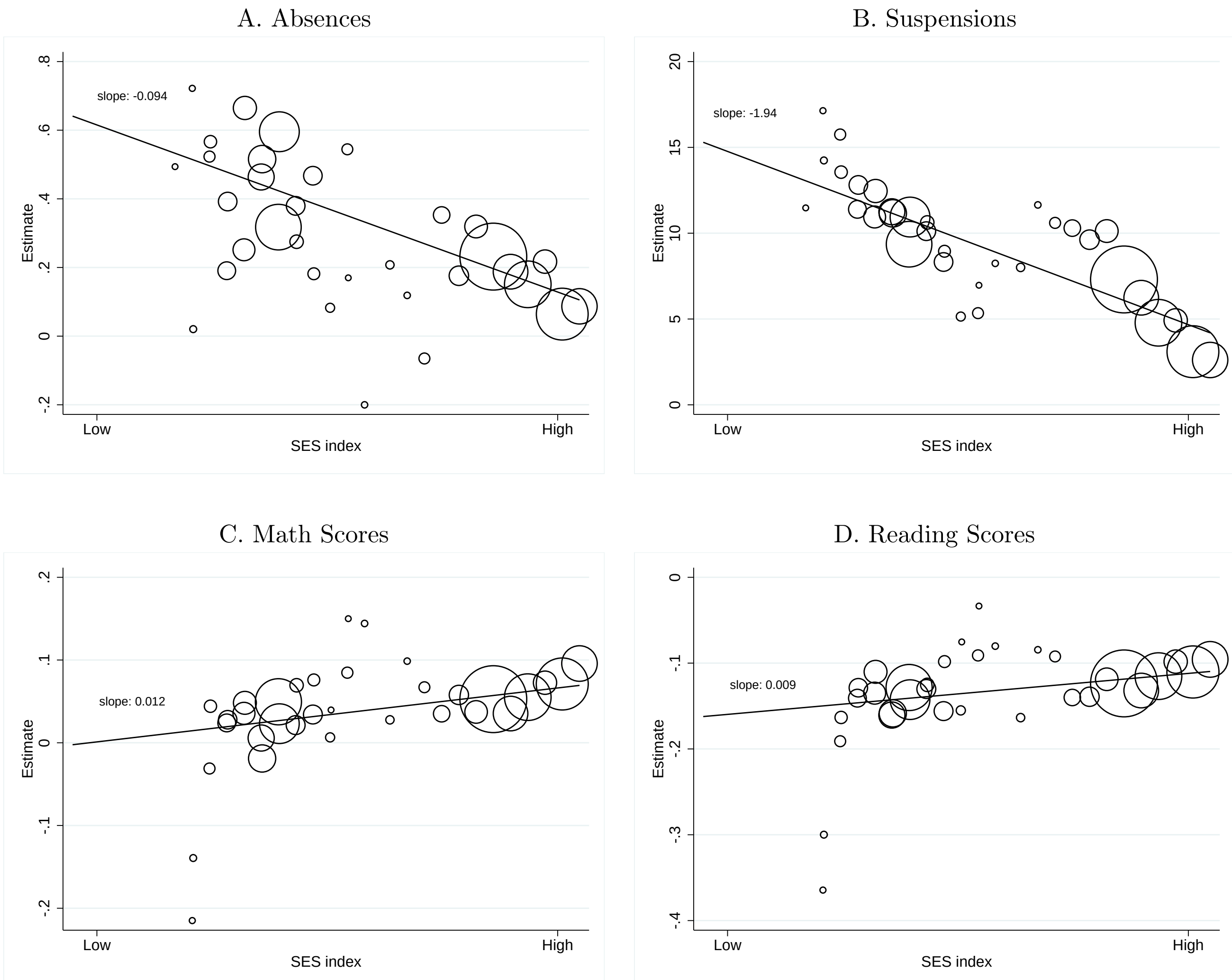

Mean absence rates and math scores by gender and race-ethnicity-nativity are computed from the pooled sample of non-twin siblings spanning grades three through eight, with each child contributing up to one observation per grade observed. High school graduation rates are computed from the two cohorts of non-twin singletons who have reached high school graduation age as of the last year of our data from the Department of Education. Kindergarten readiness rates are computed from the pooled sample of non-twin siblings born in cohorts 1994 to 1996 and 2000 to 2002. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. 95% confidence intervals are represented by whiskers on each bar.

Figure 3: Boy-Girl Gap in Kindergarten Readiness by Family SES at Birth



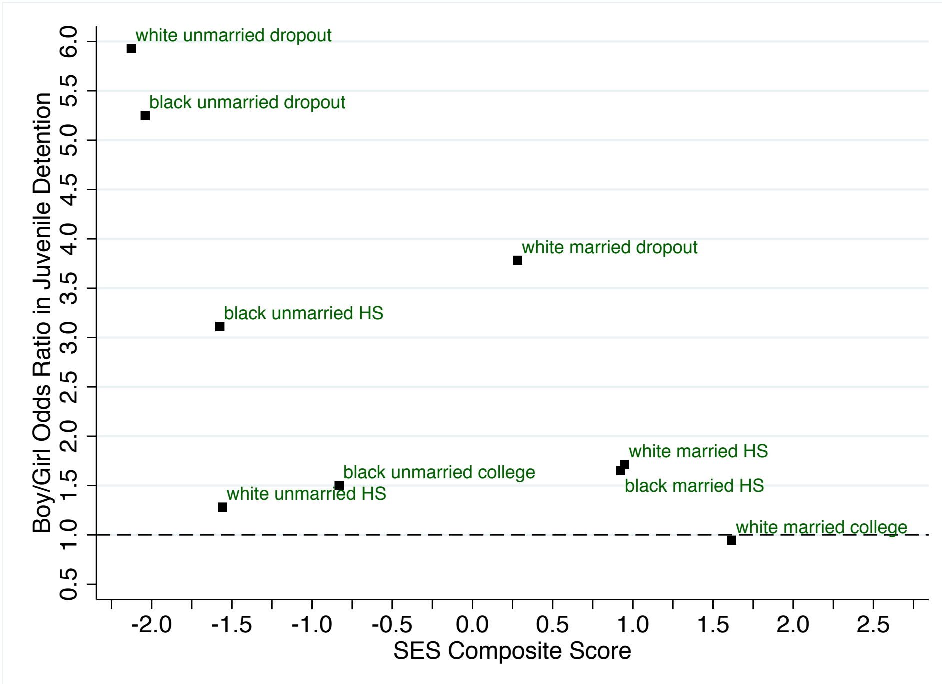

Note: Regression-adjusted mean kindergarten readiness (controlling for child year and month of birth, maternal age at birth, birth order within family/sibling composition, the main effect of the SES index, and the race-ethnicity-nativity of the mother, is plotted for 54 bins of the SES index, against the mean values of our SES index in each of those bins, from the sample of non-twin siblings born between 1994-2002. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth.

Figure 4: Boy-Girl Gap in Behavioral and Academic Outcomes by Family SES at Birth



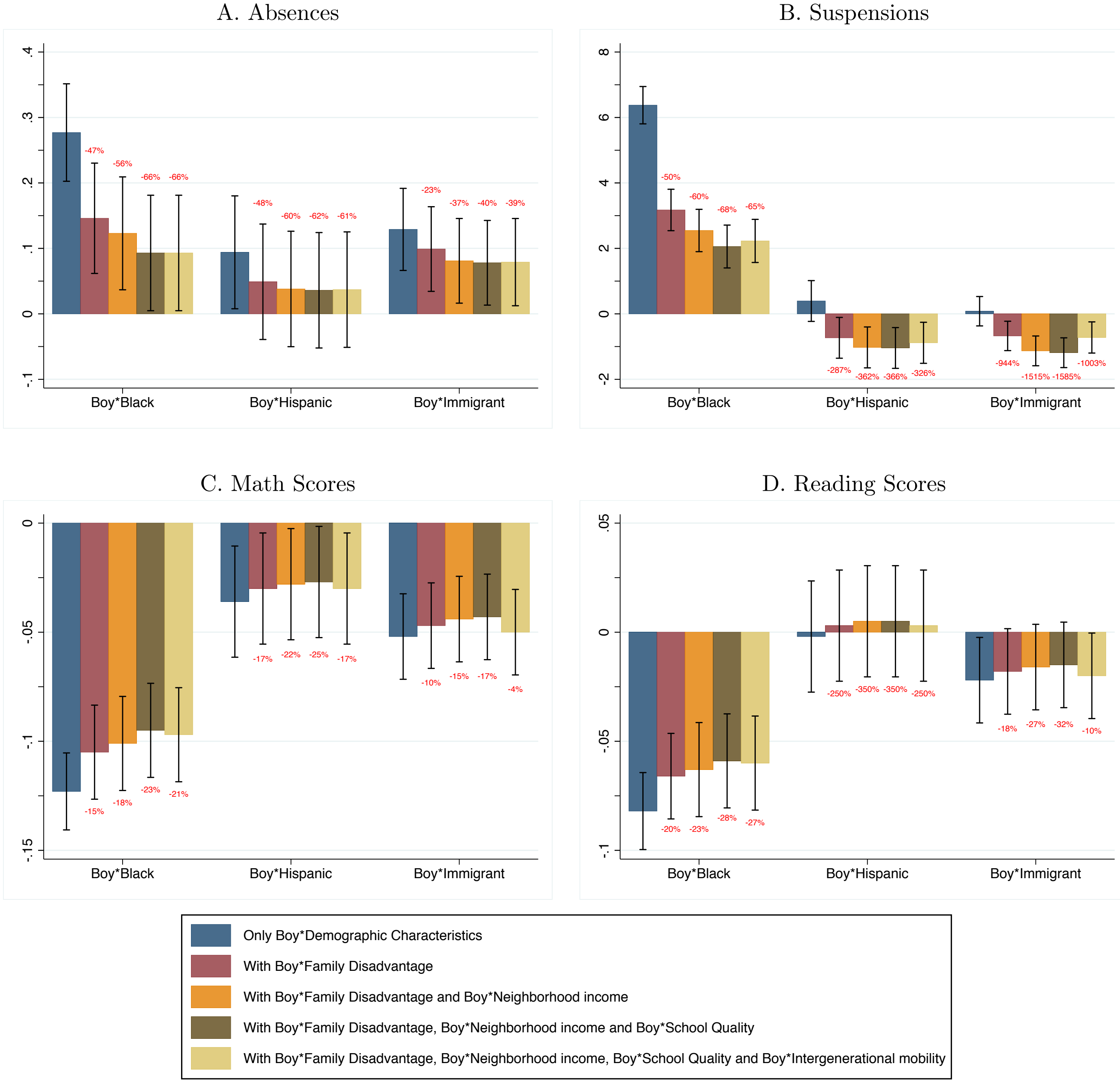

Regression-adjusted mean absence rates, suspension rates, and FCAT math and reading scores (controlling for child year and month of birth, maternal age at birth, birth order within family/sibling composition, the main effect of the SES index, and the race-ethnicity-nativity of the mother, is plotted for 54 bins of the SES index, against the mean values of our SES index in each of those bins, from the sample of non-twin siblings born between 1994-2002. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth.

Figure 5: Boy/Girl Gap in Florida Juvenile Detention Rates by Age 16, by Race, Maternal Education, and Marital Status at Birth




Note: Data are sourced from a large Florida county school district, which recorded whether youths are engaged in the juvenile justice system for "serious involvement" in criminal activity, leading to repeat felony convictions or incarceration, by their 16th birthday, Florida's mandatory school attendance age, in the academic years from 1989-90 through 2004-05. Youths are in the analysis if they were born on or before August 31, 1989 and were either a member of the kindergarten cohort of 1989-90 or had a sibling who was part of that kindergarten cohort, and were in two-plus sibling families represented in these data. Each point represents the ratio of male to female rates of criminal activity by age 16 for the combination of race, maternal education, and maternal marital status measured at the time of the 1989-90 kindergartener's birth.

Figure 6: Sibling Gender Gaps in Behavioral and Academic Outcomes: Observed and Explained




Note: This figure plots the regression-adjusted race-ethnicity-nativity gender gap in absence rates, suspension rates, math scores and reading scores from sibling fixed effects models in Tables 13 and 14. 95% confidence intervals are represented by whiskers on each bar.

Table 1: Sample Selection: Matched Florida Birth and Public School Records

| | White, black, Hispanic, and Immigrant Births (1) | With Complete Data (2) | Matched to Florida School Fecords (3) | Matched to 3rd Grade Test Score (4) | Sibling Sample w/ 3rd Grade Test Score (5) |
|---|---|---|---|---|---|
| White non-Hispanic non-immigrant | 52.0 | 52.4 | 50.8 | 49.7 | 43.0 |
| Black non-Hispanic non-immigrant | 17.1 | 16.8 | 18.7 | 19.4 | 26.1 |
| Hispanic non-immigrant | 8.2 | 8.3 | 8.4 | 8.4 | 10.2 |
| Immigrant | 22.6 | 22.5 | 22.1 | 22.4 | 20.7 |
| High school dropout | 20.6 | 20.1 | 21.8 | 22.6 | 22.9 |
| High school graduate | 58.5 | 59.0 | 60.6 | 61.0 | 58.7 |
| College graduate | 20.5 | 20.9 | 17.6 | 16.4 | 18.4 |
| Age 21 or below | 22.1 | 21.8 | 23.6 | 24.1 | 25.2 |
| Age between 22 and 29 | 41.7 | 41.9 | 42.0 | 42.0 | 44.5 |
| Age between 30 and 35 | 26.0 | 26.2 | 24.9 | 24.5 | 23.6 |
| Age 36 or above | 10.1 | 10.1 | 9.5 | 9.4 | 6.7 |
| Married at time of birth | 63.8 | 64.6 | 61.8 | 60.7 | 60.5 |
| Non-married, father present | 21.0 | 21.6 | 23.3 | 24.0 | 22.2 |
| Non-married, father absent | 13.5 | 13.8 | 14.9 | 15.3 | 17.4 |
| Boy | 51.2 | 51.2 | 51.0 | 50.6 | 50.9 |
| Birth weight of boys (grams) | 3,367 | 3,397 | 3,387 | 3,387 | 3,383 |
| Birth weight of girls (grams) | 3,248 | 3,279 | 3,267 | 3,265 | 3,260 |
| N | 1,636,580 | 1,494,303 | 1,201,745 | 985,084 | 283,597 |

Note: This table reports summary statistics for the statewide data used in this analysis. The first column is the full sample of Florida births from 1994-2002, excluding immigrants who are neither non-white, non-black, nor non-Hispanics; the second column drops the 9% of records that are missing key variables; the third column contains the approximately 81% of column 2 records that were matched to Florida school records; the fourth column is the subset of column 3 for children who remained in Florida public schools through third grade and had at least one test score; and the fifth column is the subset of column 4 in the matched sibling sample – children in families with two or more children born between 1994-2002 and living in counties where siblings are matchable. All demographic characteristics are derived from the birth certificate. White, black, Hispanic and foreign born are mutually exclusive categories.

Table 2: Descriptive Statistics: Parental Characteristics, Neighborhood Characterisitcs, School Quality

| | All | White Non-Hispanic Native | Black Non-Hispanic Native | Hispanic Native | Immigrant |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| Unmarried: Father absent | 0.17 | 0.07 | 0.43 | 0.12 | 0.10 |
| Unmarried: Father present | 0.22 | 0.16 | 0.35 | 0.26 | 0.19 |
| Married | 0.61 | 0.78 | 0.23 | 0.62 | 0.71 |
| Mother HS dropout | 0.23 | 0.15 | 0.35 | 0.28 | 0.22 |
| Mother HS grad | 0.59 | 0.58 | 0.61 | 0.57 | 0.60 |
| Mother college grad | 0.18 | 0.27 | 0.05 | 0.16 | 0.18 |
| Median zip code income | 44,792 | 50,635 | 36,538 | 44,769 | 43,054 |
| | (14,549) | (14,285) | (10,618) | (13,824) | (14,226) |
| N | 283,597 | 122,000 | 73,914 | 28,911 | 58,772 |
| Chetty/Hendren CZ | -0.29 | -0.31 | -0.27 | -0.30 | -0.25 |
| mobility measure | (0.19) | (0.19) | (0.20) | (0.19) | (0.17) |
| N | 283,582 | 121,988 | 73,913 | 28,910 | 58,771 |
| A school quality | 0.62 | 0.74 | 0.35 | 0.66 | 0.65 |
| B school quality | 0.17 | 0.15 | 0.20 | 0.17 | 0.16 |
| C/D/F school quality | 0.22 | 0.11 | 0.45 | 0.17 | 0.19 |
| N | 1,019,582 | 445,550 | 257,421 | 100,651 | 215,960 |

Note: This table reports summary statistics for the sample of non-twin siblings born between 1994-2002, by race-ethnicity- nativity of the mother. All demographic characteristics are derived from the birth certificate. Income is imputed from the 2000 Census based on the child's zip code of birth.

Table 3: Descriptive Statistics: Health Characteristics of Newborns and Mothers: Means and Boy-Girl Contrasts

| | All<br>(1) | White Non-Hispanic Native<br>(2) | Black Non-Hispanic Native<br>(3) | Hispanic Native<br>(4) | Immigrant<br>(5) |
|---|---|---|---|---|---|
| | | | A. Means | | |
| Infant birth weight (g) | 3323 | 3414 | 3148 | 3322 | 3354 |
| | (546) | (523) | (564) | (519) | (532) |
| Apgar 5 score | 8.96 | 8.98 | 8.92 | 8.97 | 8.97 |
| | (0.51) | (0.50) | (0.57) | (0.48) | (0.48) |
| Adequate prenatal care (%) | 85.39 | 90.81 | 73.88 | 87.39 | 87.65 |
| Maternal health issues (%) | 22.40 | 23.13 | 26.59 | 20.68 | 16.43 |
| Abnormal conditions (%) | 4.51 | 4.96 | 4.91 | 3.61 | 3.53 |
| Congenital anomalies (%) | 0.60 | 0.64 | 0.68 | 0.50 | 0.46 |
| Complications of delivery (%) | 26.74 | 27.46 | 26.84 | 26.38 | 25.30 |
| Weeks of gestation | 38.78 | 38.91 | 38.5 | 38.76 | 38.85 |
| | (1.94) | (1.74) | (2.27) | (1.83) | (1.91) |
| | | | B. Boy-Girl Differences | | |
| Infant birth weight (g) | 123 | 121 | 122 | 108 | 118 |
| | (2) | (3) | (4) | (6) | (4) |
| Apgar 5 score | -0.01 | -0.01 | -0.01 | -0.01 | -0.01 |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.00) |
| Adequate prenatal care (%) | 0.09 | -0.23 | 0.01 | -0.17 | 0.01 |
| | (0.13) | (0.17) | (0.32) | (0.39) | (0.27) |
| Maternal health issues (%) | 0.04 | -0.11 | 0.40 | -0.09 | 0.21 |
| | (0.16) | (0.24) | (0.33) | (0.48) | (0.31) |
| Abnormal conditions (%) | 0.58 | 0.87 | 0.30 | 0.75 | 0.26 |
| | (0.08) | (0.12) | (0.16) | (0.22) | (0.15) |
| Congenital anomalies (%) | 0.13 | 0.16 | 0.06 | 0.21 | 0.11 |
| | (0.03) | (0.05) | (0.06) | (0.08) | (0.06) |
| Complications of delivery (%) | 1.14 | 0.85 | 0.97 | 1.99 | 1.53 |
| | (0.17) | (0.26) | (0.33) | (0.52) | (0.36) |
| Weeks of gestation | -0.03 | -0.07 | 0.04 | -0.05 | -0.05 |
| | (0.01) | (0.01) | (0.02) | (0.02) | (0.02) |

Note: See next page.

Table 3 note: Table reports summary statistics for the sample of non-twin siblings born between 1994-2002, by race and ethnicity of the mother and gender of the child. Standard deviations for continuous measures are given in parentheses. Prenatal care adequacy is defined according to the Kessner Adequacy of Prenatal Care Utilization index (APCU), which is a binary indicator equal to one if the mother received standard prenatal care services during pregnancy (and zero otherwise). Mother health issues during pregnancy is an indicator variable equal to one if the mother suffered from any of a large set of chronic or pregnancy-related disorders (anemia; cardiac disease; acute or chronic lung disease; diabetes; genital herpes, hydramnios/oligohydramnios; hemoglobinopathy; chronic hypertension; pregnancy associated hypertension; eclampsia; incompetent cervix; previous infant 4000+ grams; previous preterm or small for gestational age infant; renal disease; RH sensitization; uterine bleeding; other specified health problem) during pregnancy or delivery. Abnormal conditions are coded as an indicator variable equal to one if any of the following conditions are observed: anemia; birth injury; fetal alcohol syndrome; hyaline membrane disease; meconium aspiration syndrome; assisted ventilation $<30$ minutes; assisted ventilation$>30$ minutes; seizure; or other specified abnormal conditions. Congenital anomalies are coded as an indicator variable equal to one if any of the following conditions are observed: anencephalus;spina bifida/meninocele; hydrocephalus; microcephalus; other central nervous system anomalies; heart malformations; other circulatory/respiratory anomalies; rectal atresia/stenosis; trachea esophageal fistula/esophageal atresia; omphalocele/Gastroschisis; other gastrointestinal anomalies; malformed genitalia; renal agenisis; other urogenital anomalies; cleft lip/palate; polydactyly/syndactyly/adactyly; club foot; diaphragmatic hernia; other musculoskeletal/integumental anomalies; Downs Syndrome; other chromosomal anomalies; or other specified congenital anomalies. Complications of labor and delivery are coded as an indicator variable equal to one if any of the following conditions are observed: febrile; moderate/heavy meconium; premature rupture of membranes; abruptio placenta; placenta previa; other excessive bleeding; seizures during labor; precipitous labor; prolonged labor; dysfunctional labor; breech/malpresentation; cephalopelvic disproportion; cord prolapse; anesthetic complications; fetal Distress; or other specified complication. Numbers of observations for infant birth weight: 283,597 for all; 122,000 for white non-Hispanic native; 73,914 for black non-Hispanic native; 28,911 for Hispanic native; 58,772 for immigrant mother. Sample sizes vary slightly (by less that 0.1%) for the other birth outcomes.

Table 4: Testing for a Family Disadvantage Effect on the Gender Gap in Neonatal and Maternal Health

| | A. Infant Birth Weight | | | B. Abnormal Conditions at Birth | | | C. Maternal Health Problems | | | D. Prenatal Care Adequacy | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father | | 1.66 | | | 0.58 | | | -0.37 | | | 0.39 | |
| present | | (9.52) | | | (0.46) | | | (0.87) | | | (0.87) | |
| Boy×Mother | | -2.25 | | | 0.16 | | | -1.70 ~ | | | 0.89 | |
| married | | (9.50) | | | (0.45) | | | (0.87) | | | (0.84) | |
| Boy×High | | 8.47 | | | 0.01 | | | 0.63 | | | 0.22 | |
| school | | (7.32) | | | (0.37) | | | (0.69) | | | (0.67) | |
| Boy×College | | 11.20 | | | -0.06 | | | 1.39 | | | -0.13 | |
| graduate | | (9.55) | | | (0.48) | | | (0.92) | | | (0.73) | |
| Boy×SES | | | 1.17 | | | -0.08 | | | -0.31 | | | 0.21 |
| | | | (2.18) | | | (0.11) | | | (0.21) | | | (0.18) |
| Boy | 125.25 ** | 118.84 ** | 124.73 * | 0.86 ** | 0.66 | 0.90 ** | 0.098 | 0.74 | 0.25 | -0.53 ~ | -1.37 | -0.62 * |
| | (4.06) | (9.91) | (4.25) | (0.22) | (0.49) | (0.23) | (0.406) | (0.92) | (0.42) | (0.28) | (0.89) | (0.31) |
| Boy×Black | -7.78 | -7.03 | -5.97 | -0.41 | -0.45 | -0.54 | 0.181 | -0.40 | -0.33 | 0.15 | 0.54 | 0.48 |
| | (7.04) | (8.22) | (7.98) | (0.36) | (0.41) | (0.40) | (0.671) | (0.77) | (0.75) | (0.60) | (0.69) | (0.68) |
| Boy×Hispanic | -7.99 | -7.09 | -7.45 | 0.04 | -0.01 | -0.02 | -0.377 | -0.45 | -0.56 | -0.21 | -0.12 | -0.10 |
| | (9.21) | (9.32) | (9.31) | (0.44) | (0.45) | (0.45) | (0.915) | (0.92) | (0.92) | (0.70) | (0.71) | (0.71) |
| Boy× | -3.59 | -2.92 | -3.21 | -0.82 * | -0.83 * | -0.85 * | 0.072 | 0.08 | -0.04 | 0.17 | 0.21 | 0.25 |
| Immigrant | (7.37) | (7.42) | (7.41) | (0.35) | (0.35) | (0.35) | (0.659) | (0.67) | (0.67) | (0.54) | (0.54) | (0.54) |
| Mean of Y | | 3,323 | | | 4.51 | | | 22.40 | | | 85.39 | |
| # children | | 283,597 | | | 283,597 | | | 283,595 | | | 283,597 | |

Note: This table reports sibling fixed effects models where the dependent variable is birth weight in grams, abnormal conditions at birth, maternal health problems, or adequacy of prenatal care, where the dependent variables are defined in the note to Table 3. Standard errors (in parentheses) are clustered at the sibling level. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 5: Descriptive Statistics: Behavioral and Academic Outcomes in Grades 3 - 8

| | All (1) | White Non-Hispanic Native (2) | Black Non-Hispanic Native (3) | Hispanic Native (4) | Immigrant (5) |
|---|---|---|---|---|---|
| | | | A. Means | | |
| Kindergarten ready (%) | 83.00 | 88.14 | 77.63 | 80.77 | 82.38 |
| | 51,815 | 19,929 | 15,674 | 5,298 | 10,914 |
| Absence rate (%) | 4.54 | 4.60 | 5.20 | 4.83 | 3.48 |
| | (5.03) | (4.91) | (5.79) | (4.86) | (4.12) |
| | 1,031,904 | 447,481 | 265,560 | 101,377 | 217,927 |
| Suspension rate (%) | 11.20 | 6.70 | 22.70 | 8.50 | 7.50 |
| | (31.50) | (25.00) | (41.90) | (27.80) | (26.30) |
| | 1,036,619 | 449,839 | 266,190 | 101,885 | 218,705 |
| Math score (SD) | 0.07 | 0.37 | -0.47 | 0.03 | 0.14 |
| | (0.98) | (0.92) | (0.92) | (0.93) | (0.93) |
| | 1,188,640 | 514,770 | 304,660 | 118,381 | 250,829 |
| Reading score (SD) | 0.05 | 0.35 | -0.48 | 0.00 | 0.09 |
| | (0.97) | (0.93) | (0.89) | (0.92) | (0.91) |
| | 1,197,661 | 518,460 | 306,110 | 119,377 | 253,714 |
| | | | B. Boy-Girl Differences | | |
| Kindergarten ready (%) | -6.48 | -5.41 | -8.44 | -9.08 | -4.70 |
| | (0.33) | (0.46) | (0.66) | (1.08) | (0.73) |
| | 51,815 | 19,929 | 15,674 | 5,298 | 10,914 |
| Absence rate (%) | 0.23 | 0.09 | 0.48 | 0.22 | 0.26 |
| | (0.01) | (0.01) | (0.02) | (0.03) | (0.02) |
| | 1,031,904 | 447,481 | 265,119 | 101,377 | 217,927 |
| Suspension rate (%) | 7.70 | 6.30 | 12.50 | 6.60 | 6.00 |
| | (0.10) | (0.07) | (0.16) | (0.17) | (0.11) |
| | 1,036,619 | 449,839 | 266,190 | 101,885 | 218,705 |
| Math score (SD) | 0.03 | 0.06 | -0.06 | 0.03 | 0.01 |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.00) |
| | 1,188,640 | 514,770 | 304,660 | 118,381 | 250,829 |
| Reading score (SD) | -0.14 | -0.12 | -0.20 | -0.13 | -0.15 |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.00) |
| | 1,197,661 | 518,460 | 306,110 | 119,377 | 253,714 |

Note: This table reports mean absence rates (multiplied by 100), ever suspended rates (multiplied by 100), and standardized Florida Comprehensive Assessment Test (FCAT) math and reading scores, by race-ethnicity-nativity category, from the pooled sample of non-twin siblings, spanning grades three through eight, with each child contributing up to one observation per grade observed. It additionally reports kindergarten readiness, which equals one hundred if the child is deemed to be ready for kindergarten and zero otherwise. The assessment is performed in the fall of the first kindergarten grade. Kindergarten readiness rates are computed from the pooled sample of non-twin siblings born in cohorts 1994 to 1996 and 2000 to 2002. Standard deviations are reported in parentheses. Numbers of observations are recorded beneath each standard deviation (or mean, in the case of kindergarten readiness).

Table 6: Descriptive Statistics: High School Completion Rates (1992 and 1993 Birth Cohorts)

| | All | White Non-Hispanic Native | Black Non-Hispanic Native | Hispanic Native | Immigrant |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| | | | A. Means (%) | | |
| On-time graduate | 71.45 | 72.70 | 64.63 | 68.00 | 75.94 |
| 5+ years | 12.52 | 10.39 | 19.15 | 13.78 | 11.50 |
| Dropout | 16.03 | 16.91 | 16.22 | 18.22 | 12.56 |
| | | | B. Boy-Girl Differences | | |
| On-time graduate | -7.79 | -5.69 | -12.74 | -8.04 | -9.53 |
| | (0.20) | (0.27) | (0.48) | (0.90) | (0.44) |
| 5+ years | 4.69 | 3.78 | 6.83 | 4.20 | 5.98 |
| | (0.15) | (0.18) | (0.40) | (0.67) | (0.33) |
| Dropout | 3.10 | 1.91 | 5.91 | 3.85 | 3.55 |
| | (0.17) | (0.22) | (0.37) | (0.75) | (0.34) |
| # students | 198,596 | 111,454 | 39,517 | 10,566 | 37,059 |

Note: This table reports rates of on-time graduation, continuation in high school, and dropout by race-ethnicity-nativity category, from the pooled sample of non-twin singletons who were observed in Florida public schools until at least ninth grade. On-time high school completion indicates a high school diploma within four years of entering. 5+ years of high school indicates that at student is enrolled in high school more than four years after entry but had not yet dropped out. High school dropout indicates that an individual does not earn a high school diploma and is no longer enrolled in high school 5+ years after entry. High school graduation rates are computed from the two cohorts of non-twin siblings who have reached high school graduation age. Standard deviations for male-female contrasts are reported in parentheses in panel B. Numbers of observations: 195,596 for all; 111,454 for white non-Hispanic native; 39,517 for black non-Hispanic native; 10,566 for Hispanic native; 37,059 for immigrant mothers.

Table 7: Family Disadvantage and the Gender Gap in Kindergarten Readiness

| | A. Singletons OLS | | | B. Siblings OLS | | | C. Siblings FE | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father | | 1.43 ** | | | 1.76 | | | 1.29 | |
| present | | (0.37) | | | (1.12) | | | (2.05) | |
| Boy×Married | | 2.14 ** | | | 1.33 | | | 1.30 | |
| | | (0.36) | | | (1.09) | | | (2.00) | |
| Boy×High | | 1.08 ** | | | 2.45 ** | | | 2.14 | |
| school | | (0.29) | | | (0.89) | | | (1.63) | |
| Boy×College | | 3.91 ** | | | 5.36 ** | | | 5.13 ** | |
| graduate | | (0.33) | | | (1.03) | | | (1.92) | |
| Boy×SES | | | 0.87 ** | | | 0.94 ** | | | 0.95 * |
| index | | | (0.08) | | | (0.25) | | | (0.46) |
| Boy | -5.80 ** | -9.11 ** | -6.11 ** | -5.52 ** | -9.48 ** | -5.92 ** | -5.47 ** | -9.04 ** | -5.84 ** |
| | (0.12) | (0.38) | (0.13) | (0.44) | (1.16) | (0.47) | (0.82) | (2.12) | (0.88) |
| Boy×Black | -2.07 ** | -0.70 * | -0.89 ** | -2.94 ** | -1.54 ~ | -1.49 ~ | -2.93 * | -1.56 | -1.52 |
| | (0.28) | (0.31) | 0.302 | (0.79 | (0.90 | 0.885 | (1.47) | (1.69) | (1.65) |
| Boy×Hispanic | -0.35 | 0.10 | -0.01 | -3.45 ** | -2.79 * | -2.94 ** | -3.55 ~ | -2.88 | -3.02 |
| | (0.38) | (0.38) | (0.38) | (1.11) | (1.12) | (1.13) | (2.09) | (2.11) | (2.11) |
| Boy× | 0.20 | 0.58 * | 0.52 * | 0.77 | 1.20 | 1.11 | 0.79 | 1.20 | 1.07 |
| Immigrant | (0.26) | (0.26) | (0.26) | (0.84) | (0.85) | (0.85) | (1.56) | (1.57) | (1.57) |
| Mean of Y | | 84.74 | | | | 83.00 | | | |
| # children | | 510,729 | | | | 51,815 | | | |

Note: Panel A presents the results for the sample of singleton births whereas panels B and C present results for the sample of siblings. The dependent variable equals one hundred if the child is deemed to be ready for kindergarten and zero otherwise. The assessment is performed in the fall of the first kindergarten grade. We do not observe kindergarten readiness assessment for children born in the years 1997-1999. Panels A and B include main effects for race-ethnicity–nativity of mother. All columns include controls for child year and month of birth as well as maternal age at birth and birth order within family or sibling composition. Columns (1) and (2) further control for maternal education and family structure/marital status at birth. Columns (3) control for the main effect of our SES index. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Standard errors are clustered at the individual level in panel A and at the family (sibling) level in panels B and C. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 8: Family Disadvantage and the Gender Gap in Absence Rates

| | A. Singleton OLS | | | B. Sibling OLS | | | C. Sibling FE | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father present | | -0.124 ** | | | -0.127 * | | | -0.173 ** | |
| | | (0.031) | | | (0.059) | | | (0.057) | |
| Boy×Mother married | | -0.248 ** | | | -0.334 ** | | | -0.307 ** | |
| | | (0.030) | | | (0.058) | | | (0.055) | |
| Boy×High school graduate | | -0.039 | | | -0.027 | | | -0.026 | |
| | | (0.024) | | | (0.048) | | | (0.044) | |
| Boy×College graduate | | -0.131 ** | | | -0.051 | | | -0.158 ** | |
| | | (0.028) | | | (0.054) | | | (0.049) | |
| Boy×SES | | | -0.069 ** | | | -0.092 ** | | | -0.088 ** |
| | | | (0.006) | | | (0.013) | | | (0.012) |
| Boy | 0.171 ** | 0.428 ** | 0.185 ** | 0.117 ** | 0.431 ** | 0.165 ** | 0.209 ** | 0.536 ** | 0.256 ** |
| | (0.012) | (0.033) | (0.012) | (0.023) | (0.065) | (0.025) | (0.021) | (0.062) | (0.023) |
| Boy×Black | 0.268 ** | 0.147 ** | 0.175 ** | 0.369 ** | 0.193 ** | 0.221 ** | 0.298 ** | 0.119 ** | 0.150 ** |
| | (0.023) | (0.025) | (0.025) | (0.042) | (0.049) | (0.048) | (0.039) | (0.045) | (0.045) |
| Boy×Hispanic | 0.068 * | 0.034 | 0.042 | 0.135 ** | 0.088 ~ | 0.085 | 0.100 * | 0.046 | 0.049 |
| | (0.030) | (0.031) | (0.031) | (0.052) | (0.053) | (0.053) | (0.045) | (0.046) | (0.046) |
| Boy× Immigrant | 0.128 ** | 0.104 ** | 0.100 ** | 0.179 ** | 0.153 ** | 0.138 ** | 0.116 ** | 0.081 * | 0.081 * |
| | (0.019) | (0.020) | (0.019) | (0.037) | (0.038) | (0.037) | (0.033) | (0.034) | (0.034) |
| Mean of Y | | 4.80 | | | | | 4.54 | | |
| # children | | 1,011,474 | | | | | 281,150 | | |

Note: Panel A presents the results for the sample of singleton births whereas panels B and C present results for the sample of siblings, grades three through eight. The dependent variable is the child's absence rate in grades three through eight, multiplied by 100. Each child contributes up to six observations. Panels A and B include controls for maternal race-ethnicity– nativity of mother. All columns include controls for child year and month of birth as well as maternal age at birth and birth order within family or sibling composition. Columns (1) and (2) further control for maternal education and family structure/marital status at birth. Columns (3) control for the main effect of our SES index. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Standard errors are clustered at the individual level in panel A and at the family (sibling) level in panels B and C. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 9: Family Disadvantage and the Gender Gap in Suspension Rates, Math Scores, and Reading Scores

| | A. Suspension Rate (%) | | | B. Math Test Scores (SD) | | | C. Reading Test Scores (SD) | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father present | | -0.99 * | | | 0.007 | | | -0.002 | |
| | | (0.40) | | | (0.012) | | | (0.011) | |
| Boy×Mother married | | -3.59 ** | | | 0.020 ~ | | | 0.013 | |
| | | (0.38) | | | (0.012) | | | (0.011) | |
| Boy×High school graduate | | -2.61 ** | | | 0.020 * | | | 0.013 | |
| | | (0.31) | | | (0.009) | | | (0.009) | |
| Boy×College graduate | | -6.16 ** | | | 0.047 ** | | | 0.030 * | |
| | | (0.33) | | | (0.012) | | | (0.012) | |
| Boy×SES | | | -1.93 ** | | | 0.012 ** | | | 0.009 ** |
| | | | (0.08) | | | (0.003) | | | (0.003) |
| Boy | 6.49 ** | 12.69 ** | 7.51 ** | 0.055 ** | 0.013 | 0.048 ** | -0.121 ** | -0.147 ** | -0.126 ** |
| | (0.13) | (0.41) | (0.14) | (0.005) | (0.013) | (0.005) | (0.005) | (0.012) | (0.005) |
| Boy×Black | 6.26 ** | 3.04 ** | 3.03 ** | -0.124 ** | -0.103 ** | -0.104 ** | -0.080 ** | -0.066 ** | -0.064 ** |
| | (0.29) | (0.33) | (0.32) | (0.009) | (0.010) | (0.010) | (0.008) | (0.010) | (0.010) |
| Boy×Hispanic | 0.34 | -0.96 ** | -0.78 * | -0.037 ** | -0.028 * | -0.030 * | -0.004 | 0.003 | 0.002 |
| | (0.32) | (0.32) | (0.32) | (0.012) | (0.012) | (0.012) | (0.012) | (0.012) | (0.012) |
| Boy× Immigrant | -0.09 | -0.94 ** | -0.84 ** | -0.047 ** | -0.041 ** | -0.042 ** | -0.019 * | -0.015 ~ | -0.015 ~ |
| | (0.23) | (0.23) | (0.23) | (0.009) | (0.009) | (0.009) | (0.009) | (0.009) | (0.009) |
| Mean of Y | | 11.16 | | | 0.07 | | | 0.05 | |
| # children | | 281,606 | | | 278,524 | | | 278,705 | |

Note: This table reports the results of sibling fixed effects models where the dependent variable is, in turn, the suspension rate, the standardized reading test score, and the standardized math test score, from grades three through eight. All columns include controls for child year and month of birth, and maternal race-ethnicity- nativity , and maternal age at birth and birth order within family or sibling composition. Columns (1) and (2) further control for maternal education and family structure/marital status at birth. Columns (3) control for the main effect of our SES index. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Standard errors are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 10: Family Disadvantage and the Gender Gap in Physical and Cognitive/Behavioral Disabilities

| | A. Any Disability | | | B. Cognitive/Behavioral Disability | | | C. Physical Disability | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father present | | -0.91 | | | -0.55 | | | -0.36 | |
| | | (0.88) | | | (0.79) | | | (0.56) | |
| Boy×Mother married | | -1.43 | | | -1.42 ~ | | | -0.01 | |
| | | (0.88) | | | (0.78) | | | (0.57) | |
| Boy×High school | | -2.46 ** | | | -2.78 ** | | | 0.32 | |
| | | (0.70) | | | (0.63) | | | (0.45) | |
| Boy×College graduate | | -2.97 ** | | | -3.94 ** | | | 0.97 | |
| | | (0.92) | | | (0.76) | | | (0.67) | |
| Boy×SES | | | -0.74 ** | | | -0.96 ** | | | 0.22 |
| | | | (0.21) | | | (0.17) | | | (0.14) |
| Boy | 13.39 ** | 16.87 ** | 13.75 ** | 7.27 ** | 11.13 ** | 7.74 ** | 6.12 ** | 5.73 ** | 6.01 ** |
| | (0.41) | (0.94) | (0.43) | (0.31) | (0.83) | (0.34) | (0.33) | (0.62) | (0.33) |
| Boy×Black | 2.69 ** | 1.44 ~ | 1.46 ~ | 4.16 ** | 2.65 ** | 2.59 ** | -1.48 ** | -1.20 * | -1.12 * |
| | (0.67) | (0.76) | (0.75) | (0.57) | (0.66) | (0.64) | (0.46) | (0.52) | (0.51) |
| Boy×Hispanic | -0.14 | -0.67 | -0.57 | 1.93 * | 1.23 | 1.37 ~ | -2.06 ** | -1.90 ** | -1.94 ** |
| | (0.92) | (0.93) | (0.93) | (0.79) | (0.80) | (0.80) | (0.62) | (0.62) | (0.62) |
| Boy×Immigrant | -1.13 ~ | -1.43 * | -1.40 * | 1.20 * | 0.80 | 0.86 | -2.33 ** | -2.23 ** | -2.26 ** |
| | (0.68) | (0.69) | (0.68) | (0.57) | (0.57) | (0.57) | (0.47) | (0.47) | (0.47) |
| Mean of Y | | 23.27 | | | 14.05 | | | 9.22 | |
| # children | | | | | 283,597 | | | | |

Note: This table reports the results of sibling fixed effects models where the dependent variable is, in turn, the likelihood that a child has a disability recorded on the school record, a cognitive/behavioral disability (language impaired; intellectual disability – subdivided in the early years of the data into educable mentally handicapped, trainable mentally handicapped, and profoundly mentally handicapped; developmental delay; specific learning disability; emotionally handicapped; severely emotionally disturbed; or autism spectrum disorders), or a physical disability (orthopedically impaired; speech impaired; deaf or hard of hearing; visually impaired; hospital/homebound; dual-sensory impaired; traumatic brain injury; or established conditions). All columns include controls for child year and month of birth, and maternal race-ethnicity- nativity , and maternal age at birth and birth order within family or sibling composition. Columns (1) and (2) further control for maternal education and family structure/marital status at birth. Columns (3) control for the main effect of our SES index. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Standard errors are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 11: Family Disadvantage and the Gender Gap in High School Graduation

| | A. On-time HS grad | | | | B. 5+ years of HS | | | | C. HS dropout | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | | (2) | | (1) | | (2) | | (1) | | (2) | |
| Boy×Mother married | | | 0.031 | ** | | | -0.016 | ** | | | -0.015 | ** |
| | | | (0.005) | | | | (0.004) | | | | (0.004) | |
| Boy×High school graduate | | | 0.008 | | | | -0.006 | | | | -0.001 | |
| | | | (0.006) | | | | (0.004) | | | | (0.005) | |
| Boy×College graduate | | | 0.032 | ** | | | -0.024 | ** | | | -0.007 | |
| | | | (0.007) | | | | (0.005) | | | | (0.006) | |
| Boy | -0.060 | ** | -0.096 | ** | 0.039 | ** | 0.060 | ** | 0.021 | ** | 0.036 | ** |
| | (0.003) | | (0.006) | | (0.002) | | (0.005) | | (0.002) | | (0.005) | |
| Boy×Black | -0.067 | ** | -0.046 | ** | 0.030 | ** | 0.018 | ** | 0.037 | ** | 0.028 | ** |
| | (0.005) | | (0.006) | | (0.004) | | (0.005) | | (0.004) | | (0.005) | |
| Boy×Hispanic | -0.022 | * | -0.015 | ~ | 0.004 | | -0.001 | | 0.018 | * | 0.015 | ~ |
| | (0.009) | | (0.009) | | (0.007) | | (0.007) | | (0.008) | | (0.008) | |
| Boy×Immigrant | -0.037 | ** | -0.032 | ** | 0.022 | ** | 0.018 | ** | 0.016 | ** | 0.014 | ** |
| | (0.005) | | (0.005) | | (0.004) | | (0.004) | | (0.004) | | (0.004) | |
| Mean of Y | 0.715 | | | | 0.125 | | | | 0.16 | | | |
| # children | 198,596 | | | | 198,596 | | | | 198,596 | | | |

This table reports the results of OLS models where the dependent variable is, in turn, on-time graduation, continuation in high school, and dropout, from the pooled sample of non-twin singletons who were observed in Florida public schools until at least ninth grade. Data are limited to the 1992 and 1993 birth cohorts, the only students old enough to be observed in the high school outcomes data. On-time high school completion indicates a high school diploma within four years of entering. 5+ years of high school indicates that at student is enrolled in high school more than four years after entry but had not yet dropped out. High school dropout indicates that an individual does not earn a high school diploma and is no longer enrolled in high school 5+ years after entry. All columns include controls for child year and month of birth as well as maternal age at birth, birth order, maternal education and marital status at birth, and maternal race-ethnicity-nativity. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 12: The Relationship Between Standardized Grade 3 through 8 Behavioral and Academic Measures and On-Time High School Completion

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| | | | | A. OLS Estimates: KG Readiness, Behavioral and Academic Measures, and On-Time HS Graduation | | | |
| Absence rate | -13.30 ** | | | | | | -10.21 ** |
| | (0.15) | | | | | | (0.15) |
| Suspension rate | | -12.22 ** | | | | | -8.03 ** |
| | | (0.16) | | | | | (0.17) |
| Math score | | | 11.35 ** | | | | 5.63 ** |
| | | | (0.16) | | | | (0.25) |
| Reading score | | | | 10.00 ** | | | 2.21 ** |
| | | | | (0.16) | | | (0.25) |
| KG readiness | | | | | 2.99 ** | | 0.45 ** |
| | | | | | (0.15) | | (0.14) |
| Pre-K attendance | | | | | | 0.13 | 0.34 ** |
| | | | | | | (0.14) | (0.13) |
| N | | | | 101,860 | | | |
| | B. 100 × Standardized Coefficients on SES x Boy from Primary Causal Models | | | | | | |
| Absence rate | -2.33 | | | | | | -2.33 |
| Suspension rate | | -8.61 | | | | | -8.61 |
| Math score | | | 1.50 | | | | 1.50 |
| Reading score | | | | 1.17 | | | 1.17 |
| KG readiness | | | | | 2.94 | | 2.94 |
| Pre-K attendance | | | | | | 0.52 | 0.52 |

C. Implied Contribution of Cognitive/Behavioral Gender Gaps in SES to HS Graduation Gender Gaps

| | Absence rate | Suspen-sion rate | Math score | Reading score | KG readiness | Pre-K attend | All |
|---|---|---|---|---|---|---|---|
| One SES σ | 0.342 | 0.994 | 0.121 | 0.037 | 0.019 | 0.003 | 1.52 |
| Black-White | 0.381 | 1.109 | 0.135 | 0.041 | 0.021 | 0.003 | 1.69 |
| Hispanic-White | 0.134 | 0.389 | 0.047 | 0.015 | 0.007 | 0.001 | 0.59 |
| Immigrant-White | 0.078 | 0.228 | 0.028 | 0.009 | 0.004 | 0.001 | 0.35 |

Note: Panel A of this table reports the results from regression of on-time high school graduation on absence rates, suspension rates, reading scores, math scores and kindergarten readiness, respectively, each standardized with mean zero and unit variance. All regressions also include controls for gender, race, ethnicity, immigration, marital status at birth, maternal education, maternal age at birth, child year/month of birth and number of births. Panel B reports the coefficients on the interaction term boy*SES from Tables 7 through 9, from the sibling fixed effects models. Panel C reports the implied contribution of the estimated SES gradient in the gender gaps on the gender gap in high school graduation, by multiplying the coefficients from panels A and B, and scaling the result by the standard deviation in SES, black/white high school graduation gap, Hispanic/white high school graduation gap, and immigrant/white high school graduation gap.

Table 13: Determinants of the Gender Gap in Absences and Suspensions: Neighborhood Income, School Quality and Economic Mobility

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | | A. Absence Rate | | | |
| Boy×SES | | -0.078 ** | -0.070 ** | -0.078 ** | -0.067 ** | -0.062 ** |
| | | (0.012) | (0.012) | (0.012) | (0.012) | (0.012) |
| Boy×Income in 10,000 USD | | | -0.025 ** | | | -0.018 ~ |
| | | | (0.009) | | | (0.010) |
| Boy×Mobility | | | | -0.014 | | -0.048 |
| | | | | (0.073) | | (0.074) |
| Boy×A school | | | | | -0.160 ** | -0.154 ** |
| | | | | | (0.032) | (0.032) |
| Boy×B school | | | | | -0.083 * | -0.082 * |
| | | | | | (0.035) | (0.035) |
| Boy | 0.187 ** | 0.232 ** | 0.356 ** | 0.228 ** | 0.356 ** | 0.425 ** |
| | (0.020) | (0.023) | (0.053) | (0.033) | (0.036) | (0.060) |
| Boy×Black | 0.277 ** | 0.146 ** | 0.123 ** | 0.147 ** | 0.107 ** | 0.093 * |
| | (0.038) | (0.043) | (0.044) | (0.043) | (0.044) | (0.045) |
| Boy×Hispanic | 0.094 * | 0.049 | 0.038 | 0.049 | 0.044 | 0.036 |
| | (0.044) | (0.045) | (0.045) | (0.045) | (0.045) | (0.045) |
| Boy×Immigrant | 0.129 ** | 0.099 ** | 0.081 * | 0.100 ** | 0.089 ** | 0.080 * |
| | (0.032) | (0.033) | (0.033) | (0.033) | (0.033) | (0.033) |
| Mean of Y | | | 4.47 | | | |
| # children | | | 279,361 | | | |
| | | | B. Suspension Rate | | | |
| Boy×SES | | -1.92 ** | -1.72 ** | -1.92 ** | -1.71 ** | -1.55 ** |
| | | (0.08) | (0.09) | (0.08) | (0.08) | (0.09) |
| Boy×Income in 10,000 USD | | | -0.67 ** | | | -0.58 ** |
| | | | (0.07) | | | (0.07) |
| Boy×Mobility | | | | -1.90 ** | | -2.70 ** |
| | | | | (0.55) | | (0.56) |
| Boy×A school | | | | | -2.77 ** | -2.57 ** |
| | | | | | (0.24) | (0.25) |
| Boy×B school | | | | | -0.93 ** | -0.90 ** |
| | | | | | (0.27) | (0.27) |
| Boy | 6.32 ** | 7.42 ** | 10.71 ** | 6.83 ** | 9.48 ** | 11.34 ** |
| | (0.13) | (0.14) | (0.37) | (0.22) | (0.26) | (0.42) |
| Boy×Black | 6.38 ** | 3.17 ** | 2.55 ** | 3.24 ** | 2.49 ** | 2.09 ** |
| | (0.29) | (0.32) | (0.33) | (0.32) | (0.33) | (0.33) |
| Boy×Hispanic | 0.392 | -0.73 * | -1.03 ** | -0.71 * | -0.82 ** | -1.04 ** |
| | (0.32) | (0.32) | (0.32) | (0.32) | (0.32) | (0.32) |
| Boy×Immigrant | 0.080 | -0.68 ** | -1.13 ** | -0.57 * | -0.85 ** | -1.08 ** |
| | (0.23) | (0.23) | (0.23) | (0.23) | (0.23) | (0.23) |
| Mean of Y | | | 10.80 | | | |
| # children | | | 279,808 | | | |

Note: This table presents results from sibling fixed effect models where the dependent variables are, in turn, the absence rate and the suspension rate, as previously defined. All columns include controls for child year and month of birth, maternal race-ethnicity-nativity, maternal age at birth and birth order within family or sibling composition, the SES index, and the grade associated with a school. Column (2) additionally includes the interaction of gender and the SES index. Column (3) further includes the interaction of gender and the Chetty-Hendren mobility measure. Column (4) is the same specification as column (2), but also controls for the interaction of gender and school quality. Column (5) includes the SES index, mobility measure, school quality and the interaction of all these with gender. Sample excludes families for whom zip code is not known at birth. Standard errors are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table 14: Determinants of the Gender Gap in Math and Reading Scores: Neighborhood Income, School Quality and Economic Mobility

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | | A. Math Scores | | | |
| Boy×SES | | 0.011 ** | 0.009 ** | 0.011 ** | 0.009 ** | 0.007 * |
| | | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) |
| Boy×Income in 10000 USD | | | 0.005 | | | 0.004 |
| | | | (0.003) | | | (0.003) |
| Boy×Mobility | | | | 0.053 ** | | 0.060 ** |
| | | | | (0.019) | | (0.020) |
| Boy×A school | | | | | 0.032 ** | 0.031 ** |
| | | | | | (0.006) | (0.006) |
| Boy×B school | | | | | 0.015 * | 0.016 ** |
| | | | | | (0.006) | (0.006) |
| Boy | 0.066 ** | 0.060 ** | 0.037 * | 0.076 ** | 0.036 ** | 0.034 * |
| | (0.005) | (0.006) | (0.015) | (0.008) | (0.008) | (0.016) |
| Boy×Black | -0.123 ** | -0.105 ** | -0.101 ** | -0.107 ** | -0.097 ** | -0.095 ** |
| | (0.009) | (0.011) | (0.011) | (0.011) | (0.011) | (0.011) |
| Boy×Hispanic | -0.036 ** | -0.030 * | -0.028 * | -0.030 * | -0.029 * | -0.027 * |
| | (0.013) | (0.013) | (0.013) | (0.013) | (0.013) | (0.013) |
| Boy×Immigrant | -0.052 ** | -0.047 ** | -0.044 ** | -0.050 ** | -0.045 ** | -0.046 ** |
| | (0.010) | (0.010) | (0.010) | (0.010) | (0.010) | (0.010) |
| Mean of Y | | | 0.10 | | | |
| # children | | | 274,834 | | | |
| | | | B. Reading Scores | | | |
| Boy×SES | | 0.009 ** | 0.008 ** | 0.009 ** | 0.008 ** | 0.007 * |
| | | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) |
| Boy×Income in 10000 USD | | | 0.004 | | | 0.003 |
| | | | (0.003) | | | (0.003) |
| Boy×Mobility | | | | 0.041 * | | 0.046 * |
| | | | | (0.019) | | (0.019) |
| Boy×A school | | | | | 0.022 ** | 0.022 ** |
| | | | | | (0.006) | (0.006) |
| Boy×B school | | | | | 0.005 | 0.005 |
| | | | | | (0.006) | (0.006) |
| Boy | -0.11 ** | -0.116 ** | -0.134 ** | -0.103 ** | -0.132 ** | -0.134 ** |
| | (0.005) | (0.006) | (0.015) | (0.008) | (0.007) | (0.016) |
| Boy×Black | -0.082 ** | -0.066 ** | -0.063 ** | -0.068 ** | -0.061 ** | -0.059 ** |
| | (0.009) | (0.010) | (0.011) | (0.010) | (0.010) | (0.011) |
| Boy×Hispanic | -0.002 | 0.003 | 0.005 | 0.003 | 0.004 | 0.005 |
| | (0.013) | (0.013) | (0.013) | (0.013) | (0.013) | (0.013) |
| Boy×Immigrant | -0.022 * | -0.018 * | -0.016 ~ | -0.020 * | -0.017 ~ | -0.017 ~ |
| | (0.010) | (0.010) | (0.010) | (0.010) | (0.010) | (0.010) |
| Mean of Y | | | 0.07 | | | |
| # children | | | 274,878 | | | |

Note: This table presents results from sibling fixed effect models where the dependent variables are standardized FCAT math and reading scores, as previously defined. All columns include controls for child year and month of birth, maternal race-ethnicity-nativity, maternal age at birth and birth order within family or sibling composition, the SES index, and the grade associated with a school. Column (2) additionally includes the interaction of gender and the SES index. Column (3) further includes the interaction of gender and the Chetty-Hendren mobility measure. Column (4) is the same specification as column (2), but also controls for the interaction of gender and school quality. Column (5) includes the SES index, mobility measure, school quality and the interaction of all these with gender. Sample excludes families for whom zip code is not known at birth. Standard errors are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

## Appendix Tables

Table A1: Construction of Principal Components SES Index for Family Disadvantage

| | First component | Second component |
|---|---|---|
| Mother's education | 0.47 | 0.12 |
| Married | 0.66 | -0.04 |
| Father present | -0.43 | 0.69 |
| Father absent | -0.39 | -0.71 |
| Eigenvalue | 2.07 | 1.24 |

| Summary statistics for the first component | |
|---|---|
| Mean | 0.00 |
| SD | (1.44) |
| Mean White | 0.53 |
| SD White | (1.25) |
| Mean Black | -1.08 |
| SD Black | (1.21) |
| Mean Hispanic | -0.04 |
| SD Hispanics | (1.44) |
| Mean immigrant | 0.20 |
| SD immigrant | (1.37) |

Note: The table reports the results of a principal components analysis of mother's education (in years) and marital status/family structure at birth (married, unmarried-father present, unmarried-father absent). The eigenvectors associated with the first and second components are reported, as well as their associated eigenvalues. The bottom panel reports summary statistics of the SES index, defined as the first component of the principal components analysis, for the overall sample of siblings as well as the white, black, Hispanic and immigrant subsamples.

Table A2: Testing for a Household Disadvantage Gender Gap in Neonatal Health: Additional Evidence

| | A. Gestation Weeks | | | B. Five Minute Apgar Score | | | C. Congenital Anomalies | | | D. Labor and delivery Complications | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) | (1) | (2) | (3) |
| Boy×Father present | | -0.002 | | | -0.008 | | | 0.020 | | | 0.180 | |
| | | (0.043) | | | (0.012) | | | (0.179) | | | (0.951) | |
| Boy×Mother married | | -0.027 | | | -0.013 | | | -0.031 | | | 0.761 | |
| | | (0.042) | | | (0.012) | | | (0.178) | | | (0.954) | |
| Boy×High school graduate | | 0.007 | | | 0.004 | | | -0.196 | | | 0.150 | |
| | | (0.033) | | | (0.010) | | | (0.143) | | | (0.753) | |
| Boy×College graduate | | 0.028 | | | 0.019 | | | 0.030 | | | 0.653 | |
| | | (0.040) | | | (0.012) | | | (0.192) | | | (1.030) | |
| Boy×SES | | | -0.005 | | | -0.000 | | | -0.015 | | | 0.325 |
| | | | (0.009) | | | (0.003) | | | (0.042) | | | (0.230) |
| Boy | -0.067 ** | -0.057 | -0.065 ** | -0.015 ** | -0.011 | -0.015 ** | 0.191 * | 0.316 | 0.199 * | 1.594 ** | 0.708 | 1.434 ** |
| | (0.016) | (0.044) | (0.017) | (0.005) | (0.013) | (0.005) | (0.085) | (0.192) | (0.089) | (0.452) | (1.009) | (0.468) |
| Boy×Black | 0.09 ** | 0.081 * | 0.082 * | 0.010 | 0.009 | 0.010 | -0.038 | -0.047 | -0.063 | -0.341 | 0.197 | 0.195 |
| | (0.031) | (0.035) | (0.035) | (0.009) | (0.010) | (0.010) | (0.139) | (0.161) | (0.156) | (0.728) | (0.843) | (0.824) |
| Boy×Hispanic | 0.040 | 0.039 | 0.037 | 0.002 | 0.003 | 0.001 | 0.126 | 0.121 | 0.117 | 0.960 | 1.153 | 1.162 |
| | (0.038) | (0.038) | (0.038) | (0.011) | (0.011) | (0.011) | (0.183) | (0.185) | (0.184) | (1.029) | (1.040) | (1.039) |
| Boy× Immigrant | 0.059 ~ | 0.059 ~ | 0.057 ~ | 0.005 | 0.006 | 0.005 | -0.051 | -0.046 | -0.056 | -0.178 | -0.067 | -0.059 |
| | (0.031) | (0.031) | (0.031) | (0.009) | (0.009) | (0.009) | (0.134) | (0.136) | (0.136) | (0.782) | (0.789) | (0.788) |
| Mean of Y | | 38.78 | | | 8.96 | | | 0.6 | | | 26.74 | |
| # children | | 283,597 | | | 283,212 | | | 283,597 | | | 283,594 | |

Note: This table reports sibling fixed effects models where the dependent variable is weeks of gestation, five minute Apgar score, congenital anomalies, or labor and delivery complications, where the dependent variables are defined in the note to Table 3. Standard errors (in parentheses) are clustered at the sibling level. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A3: Testing for a Household Disadvantage Gender Gap in Neonatal Health: Log Birth Weight and Composite Index

| | A. Log Birth Weight | | | | | | B. Birth Index (PCA) | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | | (2) | | (3) | | (1) | | (2) | | (3) | |
| Boy×Father present | | | -0.001 | | | | | | -0.021 | | | |
| | | | (0.004) | | | | | | (0.032) | | | |
| Boy×Mother married | | | -0.003 | | | | | | -0.020 | | | |
| | | | (0.004) | | | | | | (0.031) | | | |
| Boy×High school graduate | | | 0.003 | | | | | | 0.014 | | | |
| | | | (0.003) | | | | | | (0.024) | | | |
| Boy×College graduate | | | 0.005 | | | | | | 0.032 | | | |
| | | | (0.004) | | | | | | (0.030) | | | |
| Boy×SES | | | | | 0.000 | | | | | | 0.003 | |
| | | | | | (0.001) | | | | | | (0.007) | |
| Boy | 0.036 | ** | 0.035 | ** | 0.036 | ** | 0.020 | | 0.023 | | 0.019 | |
| | (0.001) | | (0.004) | | (0.001) | | (0.013) | | (0.032) | | (0.013) | |
| Boy×Black | 0.001 | | 0.001 | | 0.001 | | 0.050 | * | 0.049 | ~ | 0.054 | * |
| | (0.003) | | (0.003) | | (0.003) | | (0.023) | | (0.027) | | (0.026) | |
| Boy×Hispanic | -0.002 | | -0.002 | | -0.002 | | -0.003 | | -0.000 | | -0.002 | |
| | (0.003) | | (0.003) | | (0.003) | | (0.028) | | (0.028) | | (0.028) | |
| Boy× Immigrant | 0.001 | | 0.001 | | 0.001 | | 0.041 | ~ | 0.043 | ~ | 0.042 | ~ |
| | (0.003) | | (0.003) | | (0.003) | | (0.022) | | (0.023) | | (0.022) | |
| Mean of Y | | | 8.09 | | | | | | 0.00 | | | |
| # children | | | 283,597 | | | | | | 283,164 | | | |

Note: This table reports sibling fixed effects models where the dependent variable is the log off birth weight, or a birth index coming from a principal component analysis of all birth outcomes in Tables 4 and A2. Standard errors (in parentheses) are clustered at the sibling level. The SES measure is constructed as the first component of a principal components analysis of years of maternal education and marital status/family structure (married, father present, father absent) at birth. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A4: Main Estimates for Grade 3 through 8 Outcomes Augmented with Log birth weight

| | A. KG readiness | | B. Absence Rate | | C. Suspension Rate | | D. Math Scores | | E. Reading Scores | |
|---|---|---|---|---|---|---|---|---|---|---|
| | All Singletons | Matched Siblings | All Singletons | Matched Siblings | All Singletons | Matched Siblings | All Singletons | Matched Siblings | All Singletons | Matched Siblings |
| I. Primary Estimates for SES and Gender Gap: Controlling for Child Birthweight | | | | | | | | | | |
| Boy×SES | 0.85 ** | 0.89 ~ | -0.07 ** | -0.09 ** | -1.62 ** | -1.92 ** | 0.005 ** | 0.013 ** | 0.003 * | 0.010 ** |
| | (0.08) | (0.46) | (0.01) | (0.01) | (0.03) | (0.08) | (0.001) | (0.003) | (0.001) | (0.003) |
| Boy | -21.87 ** | -48.94 ~ | 0.26 | 0.70 | 15.70 ** | 13.54 ** | -0.213 ** | 0.233 | -0.338 ** | -0.004 |
| | (4.58) | (26.70) | (0.37) | (0.71) | (1.71) | (4.61) | (0.072) | (0.165) | (0.069) | (0.159) |
| Boy× Black | -0.77 * | -1.135 | 0.18 ** | 0.15 ** | 2.17 ** | 2.97 ** | -0.113 ** | -0.106 ** | -0.067 ** | -0.066 ** |
| | (0.30) | (1.67) | (0.03) | (0.05) | (0.14) | (0.33) | (0.005) | (0.010) | (0.004) | (0.010) |
| Boy× Hispanic | 0.068 | -2.93 | 0.040 | 0.05 | -1.24 ** | -0.79 * | -0.024 ** | -0.029 * | -0.008 | 0.002 |
| | (0.38) | (2.11) | (0.03) | (0.05) | (0.14) | (0.32) | (0.006) | (0.012) | (0.006) | (0.012) |
| Boy× Immigrant | 0.56 * | 1.151 | 0.10 ** | 0.08 * | -1.94 ** | -0.86 ** | -0.039 ** | -0.042 ** | -0.020 ** | -0.015 ~ |
| | (0.26) | (1.57) | (0.02) | (0.03) | (0.09) | (0.23) | (0.004) | (0.009) | (0.004) | (0.009) |
| Boy× ln(Birth-weight) | 1.91 ** | 5.29 | -0.01 | -0.06 | -0.91 ** | -0.75 | 0.029 ** | -0.024 | 0.023 ** | -0.016 |
| | (0.56) | (3.29) | (0.05) | (0.09) | (0.21) | (0.57) | (0.009) | (0.020) | (0.008) | (0.020) |
| ln(Birth-weight) | 6.98 ** | 2.82 | -0.41 ** | 0.081 | 0.72 ** | 2.85 ** | 0.359 ** | 0.313 ** | 0.228 ** | 0.184 ** |
| | (0.38) | (2.67) | (0.03) | (0.07) | (0.12) | (0.44) | (0.006) | (0.017) | (0.006) | (0.016) |
| II. Comparison SES Coefficient from Models Excluding Birthweight | | | | | | | | | | |
| Boy×SES | 0.87 ** | 0.95 * | -0.07 ** | -0.09 ** | -1.63 ** | -1.93 ** | 0.005 ** | 0.012 ** | 0.003 ** | 0.009 ** |
| | (0.078) | (0.461) | (0.006) | (0.012) | (0.031) | (0.081) | (0.001) | (0.003) | (0.001) | (0.003) |
| Mother FE | N | Y | N | Y | N | Y | N | Y | N | Y |
| N | 510,729 | 51,815 | 1,011,474 | 281,150 | 1,012,590 | 281,606 | 1,002,683 | 278,524 | 1,003,578 | 278,705 |

Note: Panel I of this table presents results from OLS regressions on the sample of singleton births, grades three through eight, and results from family fixed effect models for the sample of siblings, grades three through eight. The dependent variables are kindergarten readiness, absence rates, suspension rates, and FCAT math and reading scores, as previously defined. All columns include controls for child year and month of birth as well as maternal age at birth and birth order within family or sibling composition, the grade associated with a school, and main effects of the SES index. The singleton specifications additionally include main effects for mother race-ethnicity-nativity. Standard errors are clustered at the individual level in the singleton specifications and the family level in the matched siblings specifications. Panel II reports the coefficient on the interaction term boy*SES from Tables 8 and 9, from the corresponding singleton or sibling fixed effects models. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A5: The Sibling Gender Gap in Behavioral Outcomes in Complex Families: Distinct Fathers with Same or Different Marital Status at Birth

| | A. Absence Rate (%) | | | | | | B. Suspension Rate (%) | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | All Marital Statuses | | Same Marital Status | | Different Marital Status | | All Marital Statuses | | Same Marital Status | | Different Marital Status | |
| | (1) | | (2) | | (3) | | (1) | | (2) | | (3) | |
| Boy×Father present | -0.12 | ~ | -0.32 | ** | -0.02 | | -0.49 | | -1.44 | | 0.082 | |
| | (0.07) | | (0.12) | | (0.08) | | (0.47) | | (0.86) | | (0.57) | |
| Boy×Married | -0.28 | ** | -0.36 | ** | -0.25 | * | -2.78 | ** | -3.39 | * | -1.56 | * |
| | (0.08) | | (0.12) | | (0.10) | | (0.53) | | (0.84) | | (0.70) | |
| Boy×High school graduate | -0.023 | | 0.032 | | -0.03 | | -2.07 | ** | -1.33 | | -2.36 | ** |
| | (0.06) | | (0.10) | | (0.08) | | (0.44) | | (0.67) | | (0.55) | |
| Boy×College graduate | -0.105 | | -0.08 | | 0.11 | | -5.41 | ** | -5.31 | | -5.03 | ** |
| | (0.12) | | (0.16) | | (0.18) | | (0.80) | | (0.99) | | (1.39) | |
| Boy | 0.53 | ** | 0.52 | ** | 0.52 | ** | 13.38 | ** | 13.03 | ** | 13.52 | ** |
| | (0.09) | | (0.14) | | (0.11) | | (0.54) | | (0.88) | | (0.68) | |
| Boy×Black | 0.11 | | 0.17 | | 0.05 | | 1.74 | ** | 2.22 | | 1.31 | * |
| | (0.08) | | (0.13) | | (0.10) | | (0.52) | | (0.86) | | (0.64) | |
| Boy×Hispanic | 0.06 | | 0.12 | | 0.02 | | -1.31 | ~ | -1.44 | | -1.51 | ~ |
| | (0.11) | | (0.17) | | (0.14) | | (0.73) | | (1.10) | | (0.92) | |
| Boy×Immigrant | 0.112 | | 0.23 | ~ | 0.04 | | -1.76 | ** | -1.02 | | -2.55 | ** |
| | (0.09) | | (0.12) | | (0.12) | | (0.56) | | (0.77) | | (0.74) | |
| Mean of Y | 5.47 | | 5.50 | | 5.45 | | 19.02 | | 19.11 | | 18.61 | |
| # Families | 104,311 | | 62,147 | | 46,392 | | 104,459 | | 62,220 | | 46,476 | |

Note: This table presents sibling fixed effects models in which the sample includes only siblings born to distinct fathers with the same mother. Each birth where no paternity is claimed is coded as having a distinct father. Column 1 includes siblings with concordant and discordant marital statuses at birth. Column 2 limits to siblings with concordant marital statuses at birth, while column 3 limits to siblings with discordant marital statuses. In all cases, the marital status coded for each child is his or her mother's marital status at birth. Standard errors (in parentheses) are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A6: Testing for Heterogeneous Effects of Family Disadvantage on the Gender Gap on Behavioral and Academic Outcomes among Race Groups

| | A. KG Readiness | | B. Absence Rate | | C. Suspension Rate | | D. Math Scores | | E. Reading Scores | |
|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (1) | (2) | (1) | (2) | (1) | (2) | (1) | (2) |
| Boy × SES × Black (A1) | | 0.04 | | 0.02 | | 0.99 ** | | -0.007 | | -0.006 |
| | | (1.18) | | (0.03) | | (0.23) | | (0.007) | | (0.007) |
| Boy × SES × Hispanic (A2) | | -0.54 | | 0.03 | | 0.60 * | | 0.007 | | 0.004 |
| | | (1.53) | | (0.04) | | (0.25) | | (0.009) | | (0.009) |
| Boy × SES × Immigrant (A3) | | -0.24 | | 0.00 | | 0.88 ** | | 0.020 ** | | 0.010 |
| | | (1.19) | | (0.03) | | (0.20) | | (0.007) | | (0.007) |
| Boy × SES | 0.95 * | 1.07 | -0.09 ** | -0.10 ** | -1.93 ** | -2.44 ** | 0.012 ** | 0.008 ~ | 0.009 ** | 0.008 ~ |
| | (0.46) | (0.71) | (0.01) | (0.02) | (0.08) | (0.12) | (0.003) | (0.004) | (0.003) | (0.004) |
| Boy × Black | -1.52 | -1.32 | 0.15 ** | 0.16 ** | 3.03 ** | 3.30 ** | -0.104 ** | -0.119 ** | -0.064 ** | -0.074 ** |
| | (1.65) | (1.75) | (0.05) | (0.05) | (0.32) | (0.36) | (0.010) | (0.011) | (0.010) | (0.011) |
| Boy × Hispanic | -3.02 | -3.05 | 0.05 | 0.05 | -0.78 * | -1.05 ** | -0.030 * | -0.032 * | 0.002 | 0.001 |
| | (2.11) | (2.12) | (0.05) | (0.05) | (0.32) | (0.33) | (0.012) | (0.013) | (0.012) | (0.012) |
| Boy × Immigrant | 1.07 | 1.12 | 0.08 * | 0.08 * | -0.84 ** | -1.17 ** | -0.042 ** | -0.046 ** | -0.015 ~ | -0.017 ~ |
| | (1.57) | (1.65) | (0.03) | (0.04) | (0.23) | (0.26) | (0.009) | (0.010) | (0.009) | (0.009) |
| Boy | -5.84 ** | -5.89 ** | 0.26 ** | 0.26 ** | 7.51 ** | 7.78 ** | 0.048 ** | 0.050 ** | -0.126 ** | -0.125 ** |
| | (0.88) | (0.96) | (0.02) | (0.03) | (0.14) | (0.16) | (0.005) | (0.006) | (0.005) | (0.006) |
| $H_0$(A1=A2=A3=0) | | 0.954 | | 0.752 | | 0.000 | | 0.004 | | 0.148 |
| # children | 51,815 | | 281,150 | | 281,606 | | 278,524 | | 278,705 | |

Note: This table presents results from sibling fixed effect models in which the dependent variables are kindergarten readiness, absence rates, suspension rates, and FCAT math and reading scores, as previously defined. All columns include controls for child year and month of birth as well as maternal age at birth and birth order within family or sibling composition, the grade associated with a school, and main effects of the SES index. Standard errors are clustered at the family level. P-values associated with the Wald tests of coefficient equality are reported at the bottom of the table. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A7: Testing for Heterogeneous Effects of Family Disadvantage on the Gender Gap on High School Outcomes among Race Groups

| | A. On-Time HS Grad | | B. 5+ Years High School | | C. HS Dropout | |
|---|---|---|---|---|---|---|
| Boy × Black × Married (A1) | 0.348 | | -0.417 | | 0.069 | |
| | (1.296) | | (1.039) | | (1.045) | |
| Boy × Hispanic × Married (A2) | -0.257 | | 4.382 | * | -4.125 | * |
| | (2.207) | | (1.711) | | (1.885) | |
| Boy × Immigrant × Married (A3) | -0.301 | | -1.393 | | 1.694 | |
| | (1.289) | | (0.987) | | (1.061) | |
| Boy × Black × HS grad (B1) | -0.996 | | 1.908 | ~ | -0.912 | |
| | (1.378) | | (1.142) | | (1.191) | |
| Boy × Hispanic × HS grad (B2) | -0.897 | | 0.167 | | 0.730 | |
| | (2.379) | | (1.849) | | (2.096) | |
| Boy × Immigrant × HS grad (B3) | 0.798 | | -0.244 | | -0.554 | |
| | (1.379) | | (1.056) | | (1.165) | |
| Boy × Black × College grad (C1) | -0.124 | | 1.637 | | -1.513 | |
| | (2.059) | | (1.636) | | (1.611) | |
| Boy × Hispanic × College grad (C2) | 0.832 | | 0.280 | | -1.112 | |
| | (3.015) | | (2.266) | | (2.523) | |
| Boy × Immigrant × College grad (C3) | 0.190 | | -0.106 | | -0.085 | |
| | (1.661) | | (1.223) | | (1.369) | |
| Boy × Black | -4.038 | ** | 0.636 | | 3.403 | ** |
| | (1.221) | | (1.021) | | (1.091) | |
| Boy × Hispanic | -0.767 | | -3.123 | ~ | 3.891 | * |
| | (2.130) | | (1.685) | | (1.938) | |
| Boy × Immigrant | -3.343 | * | 2.918 | ** | 0.426 | |
| | (1.364) | | (1.073) | | (1.167) | |
| Boy | -9.819 | ** | 6.369 | ** | 3.45 | ** |
| | (0.839) | | (0.643) | | (0.786) | |
| $H_0$(A1=A2=A3=0) | 0.977 | | 0.016 | | 0.025 | |
| $H_0$(B1=B2=B3=0) | 0.681 | | 0.319 | | 0.813 | |
| $H_0$(C1=C2=C3=0) | 0.992 | | 0.775 | | 0.788 | |
| $H_0$(A=B=C=0) | 0.983 | | 0.069 | | 0.130 | |
| # children | | | 198,596 | | | |

Note: This table presents OLS results for the setof singleton births born in 1992-1993. On-time high school completion takes on a value of one if an individual earns a high school diploma within four years of entering, and zero otherwise. 5+ years of high school takes on a value of one if an individual is enrolled in high school more than four years after entry, and zero otherwise. High school dropout takes on a value of one if an individual does not earn a high school diploma and is no longer enrolled in high school. All columns include main effects for race-ethnicity–nativity of mother, child year and month of birth as well as maternal age at birth and birth order within family or sibling composition, maternal education, and marital status at birth. Standard errors are clustered at individual level. P-values associated with the Wald tests of coefficient equality are reported at the bottom of the table. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.

Table A8: Accounting Non-Parametrically for Neighborhood Effects on the Gender Gap in Behavioral and Academic Outcomes

| | A. Absence Rate (%) | | | | B. Suspension Rate (%) | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| Boy×SES | -0.081 ** | -0.061 ** | -0.070 ** | -0.057 ** | -1.95 ** | -1.52 ** | -1.59 ** | -1.39 ** |
| | (0.011) | (0.012) | (0.011) | (0.012) | (0.08) | (0.08) | (0.08) | (0.08) |
| Boy×Black | 0.144 ** | 0.058 | 0.121 ** | 0.072 | 3.26 ** | 3.30 ** | 3.09 ** | 3.00 ** |
| | (0.040) | (0.045) | (0.045) | (0.047) | (0.30) | (0.34) | (0.34) | (0.36) |
| Boy×Hispanic | 0.046 | 0.005 | -0.002 | -0.003 | -0.76 * | -0.002 | -0.168 | -0.118 |
| | (0.042) | (0.045) | (0.045) | (0.046) | (0.29) | (0.32) | (0.32) | (0.32) |
| Boy×Immigrant | 0.103 ** | 0.019 | 0.048 | 0.019 | -0.66 ** | 0.44 ~ | 0.253 | 0.236 |
| | (0.030) | (0.036) | (0.036) | (0.037) | (0.21) | (0.25) | (0.26) | (0.26) |
| Zip & School FE | x | x | x | x | x | x | x | x |
| Boy×School FE | | x | | x | | x | | x |
| Boy×Zip FE | | | x | x | | | x | x |
| Mean of Y | 4.47 | | | | 10.80 | | | |
| # children | 279,361 | | | | 279,808 | | | |

| | C. Math Test Scores (SD) | | | | D. Reading Test Scores (SD) | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| Boy×SES | 0.011 ** | 0.005 ~ | 0.006 * | 0.004 | 0.009 ** | 0.004 | 0.006 * | 0.003 |
| | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) |
| Boy×Black | -0.105 ** | -0.105 ** | -0.103 ** | -0.103 ** | -0.066 ** | -0.066 ** | -0.060 ** | -0.064 ** |
| | (0.010) | (0.011) | (0.011) | (0.012) | (0.009) | (0.010) | (0.011) | (0.011) |
| Boy×Hispanic | -0.029 * | -0.039 ** | -0.038 ** | -0.042 ** | 0.004 | -0.003 | -0.003 | -0.004 |
| | (0.012) | (0.013) | (0.013) | (0.013) | (0.012) | (0.013) | (0.013) | (0.013) |
| Boy×Immigrant | -0.047 ** | -0.054 ** | -0.057 ** | -0.057 ** | -0.018 * | -0.024 * | -0.026 * | -0.027 * |
| | (0.009) | (0.010) | (0.011) | (0.011) | (0.009) | (0.010) | (0.011) | (0.011) |
| Zip & School FE | x | x | x | x | x | x | x | x |
| Boy×School FE | | x | | x | | x | | x |
| Boy×Zip FE | | | x | x | | | x | x |
| Mean of Y | 0.10 | | | | 0.07 | | | |
| # children | 274,834 | | | | 274,878 | | | |

Note: This table presents results from sibling fixed effect models where the dependent variables are absence rates, suspension rates, and FCAT math and reading scores, as previously defined. All columns include controls for child year and month of birth as well as maternal age at birth and birth order within family or sibling composition, the grade associated with a school, and main effects of school quality. All specifications include zip code fixed effects and school fixed effects. Column (2) additionally includes boy x school fixed effects, column (3) includes boy x zip code fixed effects, and column (4) includes both boy x school and boy x zip code fixed effects. Standard errors are clustered at the family level. Point estimates marked **, *, and ~ are statistically significant at the 1, 5, and 10 percent levels, respectively.