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GENDER DIFFERENCES IN (SOME) FORMATIVE INPUTS TO CHILD DEVELOPMENT

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ABSTRACT

While there is a large literature on gender differences in important childhood developmental inputs in developing countries, the evidence for developed countries is relatively limited. I investigate gender differences in some of these inputs in the US and Canada. In the US very low birthweight males face excess mortality compared to their female counterparts. I provide evidence that the previously documented increase in mortality with the withdrawal of critical care at the Very Low Birth Weight (VLBW) threshold is primarily for boys. The fact that the critical care of both boys and girls changes discretely at this threshold suggests a possible misallocation of scarce hospital resources. In the US first born girls are breastfed longer than first born males, but the difference is so small that it is unlikely to have any consequence. Finally, mothers in the US and Canada are more likely to experience depression post birth when the first born child is a boy. Perhaps related, the parenting of first born boys in Canada in the first years of life is more likely to be confrontational.

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Introduction

It is now well documented that females and males face different socio-economic trajectories over the life cycle. These differences span work and pay, health, living arrangements, and human capital investments. The reasons why may include discrimination, individual choice, environmental factors and biology. While the relationships among explanations and outcomes are in some cases reasonably straight forward—mothers and fathers clearly experience the birth of a child differently in part due to biology—the relative contributions of the various explanations to less deterministic outcomes are typically not as clear.

In sorting among the alternatives researchers typically focus on a part of the life cycle that they think critical to lifetime outcomes. The fetal origins hypothesis (Barker 1990) calls our attention to the early years of life and the prenatal period, and much of social science research on inequality is now focused on this interval.[1] However, for the study of gender differences research also reveals that later life events, for example the birth of a first child (e.g., Angelov et al. 2016, Kleven et al. 2018), can significantly change important economic and social choices.

Much of the previous research on early life environmental factors, thought important to child development and adult outcomes, stratifies the data by socio economic status. In turn this research has revealed that changes in the early life circumstances (i.e., these inputs) can have significant effects on lifetime outcomes, especially for disadvantaged children. Prominent here, for example, is the evaluation of RCTs such as the Perry Preschool Project (e.g., Heckman and Karapakula 2019).

[1] See Almond et al. (2018) for a recent review of the economic literature.

While there is now a sizable literature on how these inputs vary by gender in certain developing countries,[2] there is less direct evidence on any corresponding differences in developed countries.[3] There are previous reports that males receive more total time from their parents as a result of extra inputs from their fathers (e.g., Lundberg et al. 2007 and Yeung et al. 2001), and that family structure might vary by the gender of children (e.g., Lundberg and Rose 2003). Lundberg (2005) and Raley and Bianchi (2006) provide reviews of the literature.

In this paper I investigate differences in a selection of these inputs by gender: medical care after birth, breastfeeding, maternal mental health post birth and parenting. The circumstances of birth, for example low birth weight, are known correlates of adult outcomes. While the claimed benefits of breastfeeding probably exceed those rigorously documented, evidence from RCT investigation indicates there are advantages to children from this source of nutrition. Finally, parenting and parental time inputs are widely cited as an important developmental input. For example, some of the benefits of the Perry Preschool intervention have been attributed to its effects on home environment and parent/child attachments (e.g., Heckman and Karapakula 2019).

The analysis reveals differences in both maternal mental health post birth and early parent/child interactions when the first born child is a boy rather than a girl, which are consistent with boys experiencing a less nurturing and move aversive homelife. There are small gender differences in the periods first born children are breastfed, but previous evidence on the impact of breastfeeding suggests these would have little consequence. Finally, while lower birthweight

[2] Much of this research investigates these differences within populations in which there is preference for male offspring. See, for example, Das Gupta (1987), Pande (2003) and Barcellos, Carvalho, and Lleras-Muney (2014) on nutrition, Ganatra and Hirve (1994) and Borooah (2004) on healthcare and vaccination rates.

[3] There is a distinction here between gender differences in the provision of an input and gender differences in the effect of a given input. A developing literature investigates this latter issue (e.g., Bertrand and Pan 2013).

boys receive more hospital care than lower birthweight girls, the impact of variation in this care around the Very Low Birth Weight (VLBW) threshold by sex suggests that the survival and development of low birthweight children might be improved by reallocating even more care to boys.

Do Females and Males Enter the Labour Market Equally Skilled?

Individuals' lifetime income trajectories depend in part on the skills and aptitudes they bring to the labour market. There is now a large literature, much of it outside economics, that has investigated gender differences in these skills. This research has generated both popular (e.g., Gray 1992, Fine 2010) and academic debate (e.g., Hyde 2005, Spelke 2005) debate.

Economic research has focused on gender differences in competitiveness (e.g., Niederle and Vesterlund 2007), people (versus things) skills (e.g., Lordan and Pischke 2018) or more generally interpersonal skills (e.g., Cortes et al 2018), risk preferences (e.g., Borghans et al. 2009) and willingness to work long hours (e.g., Goldin 2014). These are thought to help account for the considerable gender based occupational segregation in the labour market (e.g., Blau et al. 2013) and gender differences in within occupation success. However, direct evaluations of the relationship between the degree of occupational segregation and gender differences in these or other skills are relatively rare (e.g., Baker and Cornelson 2018).

If the gender differences in these skills are important for labour market success, then it is more than an academic question to discover their source. Often this inquiry is cast as a duel between biology and the environment, although research on genetics tells us that this dichotomy is problematic due to the impact of environment on gene expression. However, it retains its allure likely because it appears to present a division between factors thought to be immutable or hardwired (biology) and those (environmental) that can be remediated if necessary.

The focus here is in differences in environmental factors that girls and boys experience as they grow up. The analysis does not presume that these factors are the most important, nor definitively connects any environmental differences to specific adult outcomes. Rather the objective is to help fill in missing evidence to help target future investigation.

Should Gender Differences in Environmental Inputs Matter

Whether to expect that any found gender difference in an environmental developmental input is consequential is complicated by the fact that males and females simultaneously differ in other ways that may mediate the effect of that input. For example, first born females in Canada, the UK and the US receive more reading time with their parents than first born males starting at very young ages (e.g., Baker and Milligan 2016). This might be consequential, and contribute to the sometimes cited, but also disputed, female advantage in verbal skills (Hyde 2014), if a "unit" of reading has the same developmental effect on boys and girls. But, for example, if boys are more receptive to reading than females, then the gender difference in this input might not matter.

One reason the impact of an input might differ by gender is genetic. For example, some recent research argues that the relationship between stress and consequent depression is in part mediated by genetic vulnerability that in turn appears to differ by gender (Salk and Hyde 2012). Therefore, a given amount of stress might create a different level of depression in females and males, or alternatively different amounts of stress would be required to result in the same level of depression.

As a result, it is not possible to interpret the gender differences documented here as important to gender differences in the skills and aptitudes that males and females develop in their lives without additional assumptions or evidence.

Data and Empirical Framework

The following analysis makes use of a variety of data sources which survey the conditions of young children in the U.S and in Canada. For the investigation of medical care after birth I make use of the US National Centre for Health Statistics (NCHS) public use birth cohort linked birth/infant death files for the years 1983-1991 and 1995-2010.[4] These data, for the population of births, provide a variety of measures of mortality and the cause of death for infants who die within the first year, as well as birth certificate information such as birthweight and characteristics of the mother. I also use hospital discharge data for 1995-2005[5] for Arizona, Maryland, New York and New Jersey distributed by the Heathcare Cost and Utilization Project (HCUP). Information provided for each recorded birth includes birthweight, length of hospital stay and hospital charges.

The analysis of sex differences in breastfeeding makes use of the National Immunization Survey (NIS) for the years 2010 through 2018. The NIS is nationally representative of the American public and is conducted annually, primarily to survey the vaccination of young children and teenagers in the US. I use data from the survey of children aged 19 through 35 months (NIS-Child). Included in the questionnaire are questions about whether, and how long, each child was breastfed. Due to the ages of the surveyed children, breastfeeding may still be ongoing. To address this issue, I examine milestones of breastfeeding in the first year—initiation, breastfed at least 3 months etc.—and the number of days of the first year of life the child was breastfed, which is coded 365 for those breastfed 12 months or longer.

[4] These files are not available for the years 1992-1994.

[5] The data from 1997 for Arizona is omitted from the analysis. Average birthweight calculated from these data for this state displays a sharp downward spike of over 400 grams in this year. This is due to a spike in the proportion of births with recorded birthweight of zero and roughly 200 grams

Finally estimates of sex differences in mothers' mental health post birth and parenting behaviour are based on the Canadian National Longitudinal Survey of Children and Youth (NLSCY) and the US Early Childhood Longitudinal Survey-Birth Cohort (ECLS-B). The NLSCY is a nationally representative survey of Canadian children conducted biennially between 1994/95 and 2008/09. A new cohort of children aged 0 and 1 entered the survey each wave and were followed until age 4/5.[6] The ECLS-B is a nationally representative survey of children born in 2001, and followed subsequently through multiple waves (e.g., age 9 months, 2 years…).

In much of the analysis I sample first born children if possible. The assumption that the sex of the child is randomly assigned is most tenable for first born children. The sex of subsequent children may not be random if the probability of higher order births is related to the sex of realized births. Furthermore, the parity of the birth may affect the parental inputs a child receives (e.g., Price 2008), so holding parity constant in any gender comparisons is important. However, the sex of the first born child may be related to some important environmental factors such as whether the father of the child is present in the household.[7] This said, in some analyses I consider the sample of all children or explicitly compare the results for the first born and all children samples, where the behaviour under study is likely to be less sensitive to parity (e.g., hospital provided care directly post birth) or the comparison reveals interesting differences.

Medical Care After Birth

Girls and boys face different medical challenges immediately post birth. I begin by documenting a number of these differences. In figure 1 is evidence of sex differences in one year mortality rates from the NCHS data for the years 1983-1991 and 1995-2010. In the top panel I

[6] An "original" cohort of children surveyed in the first wave and followed throughout the waves.
[7] Both Lundberg and Rose (2002) and Dahl and Moretti (2008) present evidence that a first born female raises the probability of an absent father. Baker and Milligan (2016) find no evidence of this effect in some of the data sets examined here.

present the one year mortality rates for all girls and boys. The rates for both sexes have been trending downwards over time, with more pronounced declines in the 1980s. There is also a persistent gap between the rates for males and females, which has narrowed somewhat over the period. In the bottom panel is the corresponding statistics for the sample of low birth weight (LBW) births defined by birthweight less than 2500 grams. The same patterns are present here with the gender gap declining from over 4 deaths per hundred in the 1980s to under 2 by 2010. In both graphs girls have lower mortality rates than males in all years.[8]

In table 1 I present the causes of death in the first year, by sex, for the full (top panel) and LBW (bottom panel) samples, conditioning on death in this period. The cause of death classifications were changed starting 1990, so here I restrict the sample to 1990-1991 and 1995-2010. In the full sample, the larger gender differences are for congenital malformations (higher for females), Sudden Infant Death Syndrome (higher for males) and deaths associated with the respiratory system (again higher for males). Different in the sample of LBW births is the gender differences in conditions from the perinatal period and deaths associated with the circulatory system, both higher for males.

At least some of these sex differences in morbidities may be responsive to more intensive medical care. A reason why there may be an interaction between sex differences in early life ailments and corresponding differences in medical care is that some of the important indicators for more intensive medical care—low and very low birth weight—do not distinguish by sex. This is perhaps surprising because females are on average of smaller stature than males throughout their lives. They are of lower average birthweight and (in the U.S.) attain an adult

[8] The corresponding graphs for 28 day mortality (not shown) exhibit very similar patterns.

height almost 4 inches lower than males (e.g., Baker and Cornelson 2019). There is, therefore, some intuition for using different cutoffs to indicate birth weights of concern for girls and boys.

Information on gender differences in average birthweights is provided in figure 2. Here I graph the average birthweight by birth percentile for girls and boys. Boys' average birthweight is higher than girls' starting at roughly third percentile. Interestingly, the first and second percentiles of birthweight, and the average birthweights within these percentiles, are almost identical for boys and girls in these data. However, the outcomes of children in these lower percentiles differ markedly by sex. In figure 3 I graph the one year mortality rate of girls and boys in their respective first 10 percentiles of birthweight. The mortality rate for boys is elevated relative to the rate for girls over the first 3-4 percentiles. The fourth percentile of birthweight is 2211 g for boys and 2170 g for girls. Therefore, it is at the percentiles of birthweight in which girls' and boys' average birthweights are roughly similar that we observe the excess mortality for males. The graph for 28 day mortality (not shown) tells a very similar story.

Low birth weight can have both health and socio economic repercussions. While the research on the associations of low birthweight is extensive, there appears to be few studies of either of these classes of outcomes that make a distinction by sex. An exception is Stevenson et al. (2001) who report that boys who are very low birth weight faced higher rates of mortality and a number of morbidities.[9]

The preceding evidence suggests that very low birth weight boys should receive higher levels of critical care than their very low birth weight female counterparts. This is easily checked in the data. The excess mortality of boys at these weights also suggests that this care might be more consequential to their survival. To investigate this possibility I examine how

[9] See also Ernst et al. (2020).

gender differences in mortality in the lower birthweight percentiles interact with the VLBW threshold (birthweight <1500 grams) as an indicator for more intensive medical care. Almond et al. (2010) take this methodological tack to demonstrate the returns to marginal medical care, comparing the mortality of babies with birth weight just above and just below the VLBW cutoff.[10] They find an increase in mortality for babies marginally heavier than the cutoff, which they interpret as a result of specialized care which is withdrawn as birthweight passes through the VLBW cutoff. They also demonstrate that care, captured by data on hospital costs and length of stay in the hospital, changes discontinuously at the VLBW cutoff supporting this interpretation of the results.

The intuition for the investigation here is that because the VLBW cutoff does not vary by sex, but the underlying risk of mortality of low birth weight children does, there should be corresponding differences in the impact of the variation medical care across the VLBW threshold. A difference in the impact of this care by sex might indicate a misallocation of resources through which care is provided children who do not need it at the expense to children who do, or need it more.

I note there has been some debate in this literature over the robustness of the main results in Almond et al. (2010) to the treatment of births exactly at the VLBW cutoff (Barreca et al. 2011, Almond et al. 2011). I abstract from these issues, reporting alternative treatments of these observations, with an eye to whether they affect inference on any gender gap in the effectiveness of this specialized care.

I follow the lead of these articles estimating the equation:

$$y_{it} = \alpha + \beta VLBW_i + \gamma VLBW_i \cdot (1500 - w_i) + \eta(1 - VLBW_i) \cdot (1500 - w_i) + X\mu + \varepsilon_i$$

[10] See also Barreca et al. (2011), Almond et al. (2011), Bharadwaj et al (2013) and Daysal et al. (forthcoming).

by OLS, where y is some outcome for child i born in year t, VLBW is a 0/1 indicator that the child's weight is less than 1500 grams, w_i is the child's weight in grams, and X is a set of control variables, which for the analysis of mortality in the first year include fixed effects for year of birth, mother's and father's five year age groups, race and whether the mother was born out of state, and controls for the child's gestational age and plurality.[11] I restrict the sample to children with weight within 85 grams of the 1500 gram cutoff. Following developments in this research, I estimate models that retain children with weights equal to 1500 grams in the sample, and also so called donut estimators that omit these and surrounding observations. I present both robust standard errors and standard errors clustered at the gram level.

As overview, the graphs of the one year and 28 day mortality rates of boys and girls, by one ounce (28 grams) weight bins in the vicinity of the VLBW threshold are reported in figures 4 and 5. In the top panel of each figure I include 1500 gram birthweights in the first bin above the threshold, while in the lower panel I omit these births to provide some intuition for how the donut estimates might differ from the full sample estimates.

In the top panel there is evidence of elevated mortality rates in the first bin above the threshold, potentially signaling that the withdrawal of care at the threshold is affecting the survival of these marginally heavier births. There is also a distinct difference between boys and girls as the uptick for boys is larger in both absolute and proportional terms, especially for one year mortality. This is consistent with the intuition that the withdrawal of care at the threshold may be more consequential for boys than for girls. In the bottom panel the upticks are smaller for boys and disappear for girls, removing the 1500 gram births from the sample. The diminution of the upticks is expected given the evidence in Barreca et al. (2011) that mortality at

[11] Here I use births of all pluralities to maximize sample size. Also, I am not aware of evidence that hospital personnel administer care differently according to the plurality of the birth.

exactly 1500 grams is elevated compared to mortality at surrounding birthweights. The difference in the effect of this change in sample by sex suggests that the donut estimators are less likely to provide evidence that the cut off of critical care at the 1500 gram threshold has an effect on the mortality of females.

This inference is formalized in table 2, where I present estimates for *VBLW,* of β, from equation (1) for one year mortality rates. In the first row I present the estimate retaining babies weighing 1500 grams in the sample, to match the sample selection in Almond et al. (2010). Note, however, that the sample here is not strictly comparable to the one in this previous study as I use data from an additional 8 years (2003-2010). The pooled estimate in the first column for this longer period is marginally smaller than the estimate in Almond et al (2010).[12] The estimates in the succeeding columns indicate that once the sample is split by gender, it is males that drive the pooled result. The estimate for males is more than twice as large as the estimate for females. It is almost 18 percent of the mean mortality rate for boys in the sample bins above the threshold. The estimate for girls is 8 percent of the mortality rate above the threshold. Also, the estimate for boys is statistically significant at conventional levels, while the estimate for girls is not. Finally, in estimate in the fourth column reveals that this sex difference is marginally statistically significant at conventional levels.[13]

In the subsequent rows I present estimates from the donut estimator following the sample deletions in Barreca et al. (2011). Omitting observations at 1500 grams, the estimates of *VLBW* fall by roughly 50 percent. This is consistent with the findings in Barreca et al. (2011). Only the estimate for males is statistically significant and the estimate for females is now quite small.

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[12] The corresponding estimate in Almond et al (2010) is -0.0072 (0.0022).

[13] The estimate of the difference is from the pooled regression with a full set of interactions between a dummy variable for sex and all other control variables. The reported statistic is for the estimate of the interaction with the *VLBW* dummy variable.

Additional omissions of births within one or two grams of 1500 grams from the sample, lead to the same conclusion. Given the standard errors the sex differences in these samples are no longer statistically significant, although this conclusion is tempered by both the individual estimates by sex and the evidence in figure 4. Note that when births within 3 grams of 1500 grams are omitted the estimate for males is halved again and is no longer statistically significant. As noted by both Barreca et al. (2011) and Almond et al. (2011) this donut omits births at 1503 grams, which is at a spike in reported weights as it corresponds to 53 ounces.

A corresponding analysis of the 28 day mortality rate is presented in table 3. The message here is very similar to that in the preceding table with some important exceptions. First, even when the sample omission is extended to births of weights 1497 grams through 1503 grams, the estimate of *VLBW* for males remains statistically significant and is 10 percent of the sample mean mortality rate. Second, the gender differences are more consistently statistically significant.

This evidence supports the initial intuition to investigate the impact of the VLBW cutoff on mortality by gender. The VLBW cutoff is far more meaningful for males than for females, as judged by their mortality in the first year. While this is interesting in its own right, it does not necessarily imply a misallocation of hospital critical care resources. It may be that medical staff already know the more critical consequences of the VLBW cutoff for males and so the threshold has less impact on the medical care for females. Alternatively, the threshold may lead to a discrete increase in the care of females that consumes resources which might otherwise be devoted to low birthweight males.

In figures 6 and 7 I present the birthweight profiles of two measures of care, total hospital charges and length of stay, from the HCUP data. Again, in the top panel I include 1500 gram births in the first bin above the threshold, while in the lower panel I omit these births from the sample. In both figures there is evidence that boys receive more care than girls, conditional on birthweight, by these measures, likely reflective of the greater risks of mortality boys face at lighter weights.   In the top panels there is a clear, discrete decline in these two measures of care at the threshold, which is modestly larger for males. Omitting the 1500 gram births does not materially affect this inference—the drop-offs in care for both girls and boys remain. The omission of the 1500 gram births has less consequence here because, as noted by Almond et al. (2011), the provision of care to 1500 g and marginally heavier babies does not differ substantially.

In tables 4 and 5 I present the corresponding regression estimates. The standard errors are relatively large for this analysis, reflecting in part the restricted sample of states. For hospital costs (table 4) there is discrete increase in costs for both girls and boys at the VLBW threshold, congruent with the evidence in figure 6. Only the estimate for girls is statistically significant, however, and it is almost 50 percent larger than the estimate for boys.[14] Omitting observations at and around the threshold, the estimate for girls mostly maintains its magnitude, and to a lesser extent its statistical significance, while the estimate for boys attenuates. The corresponding estimates for hospital length of stay are reported in table 5. In the full sample the estimates indicate a discrete increase in stay for both sexes at the threshold, but here it is the estimate for boys that is larger and statistically significant. Again, in the donut estimates the estimates for girls mostly maintain their magnitude while the estimates for boys attenuate. The estimates for

[14] The pooled estimate is smaller than the comparable estimate in Almond et al. (2010), 9065 (2297). Different here is the inclusion for all years between 1995 and 2005 for the four states (except 1997 for AZ), the omission of 2006, and the omission of the 1991-2002 data for California. One consequence of these differences is that the pooled sample used here is roughly one-third smaller than the sample in Almond et al. (2010).

these measures of care appear more sensitive to the sample omissions of the donut estimators than the estimates for mortality.

The preceding evidence supports the intuitive hypothesis that the impact of variation in critical care around the VLBW threshold differs by sex. Boys exhibit higher mortality rates than girls in the first percentiles of birthweight and are more affected by this variation in care than girls. However, the VLBW threshold makes no accommodation by sex, and the HCUP data suggests that critical care varies around the threshold for both girls and boys. While lower birthweight boys receive more care conditional on birthweight than girls, the efficacy of the care provided to girls is challenged, at least by the measures of first year mortality. If hospital care is a scarce resource, these results suggests that there may be overall improvement in child survival and health if a lower VLBW threshold for females freed up care for boys marginally above the current unisex threshold. Furthermore, the evidence in Bharawaj et al. (2013) and Daysal et al. (forthcoming) suggests that a reallocation of this care may have longer term payoffs for the educational attainment of these children, and positive spillover effects for their siblings and their parents. These findings underline gender differences in the contemporaneous consequences of being born with low birthweight, which may in turn lead to corresponding differences in the long run outcomes. As noted above few studies of the long run consequences of LBW have made a distinction by sex.

Breastfeeding

While breastfeeding is widely counselled for newborns (e.g., World Health Organization[15]), the claimed benefits of this practice typically exceed the support of the existing research. Some of the best evidence of positive impacts of breastfeeding come from the

[15] See https://www.who.int/health-topics/breastfeeding#tab=tab_1 accessed March 2, 2020.

Promotion of Breastfeeding Intervention Trial (PROBIT) RCT in Belarus which randomly allocated lactation support across hospitals. This support had a positive impact on breastfeeding incidence and duration, and therefore provides a basis to examine the possible impacts of breastfeeding free of confounding factors. The children who were treated by this RCT have been followed over time (e.g., Yang et al. 2018) and have revealed that the treatment led to lower rates of gastrointestinal infections and eczema (up to the teenage years) and a modest positive impact on IQ at age 6, which appears to have faded out by the teenage years. Many of the other claimed positive impacts of breastfeeding have been harder to substantiate (e.g., Baker and Milligan 2008) outside of correlational analysis.

To investigate any differences in breastfeeding practices by child gender I select cases in the NIS-Child in which the mother is the respondent, and the child is the first born. Summary statistics for this sample are reported in table 5. I also report the statistics for the larger sample of all children, and the marginally smaller sample which has non missing values for the regression analysis, to make clear any impact of sample restrictions. The incidence of breastfeeding is very similar for girls and boys in each sample, at 81 percent. There are sex differences in the unconditional duration of breastfeeding in the first year favoring girls of 1 to 2 days across samples, smallest in the sample for the regression analysis. While girls are more likely to reach each breastfeeding milestone in each sample, it is the difference in the proportions reaching at least 12 months which is largest both in absolute and relative value. It is roughly 1.3 percentage points and statistically significant at the 0.01 level in each sample.

To refine this inference, in table 6 I report estimates for a dummy variable for boys from a linear regression of the various measures of breastfeeding indicated in the top row of the table, on this dummy variable and additional controls for household income below the poverty line, whether the mother is married, whether the household has ever received WIC benefits, and fixed effects for the child's and mother's ages, the mother's education, the child's race and the number of children in the household. The control factors could vary systematically between boys and girls if, for example, the sex of the first born has an impact on family structure.

The estimates are slightly larger than the differences in the table of means, but largely confirm the inference. The difference in incidence between the sexes is very small. Boys are breastfed over the first year less than girls, the difference here approaching 3 days, although not statistically significant at conventional levels. Again, the most substantial gender difference in milestones is at 12 months. It should be noted that the estimates from the larger sample of all children (not reported) are very similar to the ones reported in table 6, except all gender differences in duration are statistically significant at the 5 percent level. For example, the difference in breastfeeding duration in the first year favors girls by 3.34 days (standard error 1.318).

The message of this analysis is that while there are differences in breastfeeding duration among American girls and boys favoring girls, they are quite small. As a point of reference, in the PROBIT study the differences in breastfeeding between the treatment and control groups, which underly the reported positive impacts of this practice, are of an order of magnitude greater than the gender differences reported here. For example (Yang et al. 2018) the treatment/control difference in reporting 6 or more months of breastfeeding is 11.2 percentage points, or 31 percent of the control group mean.

Parenting and Parental Time Inputs

The quality and type of parental/child interactions are thought to be key inputs to child development (e.g., Heckman and Mosso 2014, Kahil 2015). More authoritative (versus authoritarian), sensitive and interactive parenting practices promote development and vary across families stratified by SES. This latter point may account for the evidence that the benefits of ECE appear to accumulate to disadvantaged children, and the universal ECE is sometimes found to have negative impacts on children from more affluent families (e.g., Baker et al. 2008, 2019 and Fort et al. 2020).

Any sex differences in parenting might result from factors on either the child or parent side. For example, boys might be more difficult to parent leading to different discipline and control strategies. However, it is also possible that the birth of a boy or girl differently affects the mental state of a parent, perhaps due to preferences for a child of a particular sex, which in turn affects parent/child interactions.

The analysis begins with an examination of mothers' mental state after the birth of a child. Two studies (de Tychey et al. 2007, Myers and Johns 2019) of small samples of mothers in France and the UK found associations between postpartum depression and the birth of a male child. The suspected mechanisms in these studies are cultural preferences for babies of a specific gender and mothers' greater inflammatory response to the male fetus. Perhaps supportive of the former hypothesis are studies reporting that postpartum depression is related to the birth of a female child in India, Nigeria, Turkey and China, countries where arguably son preference is more common.[16]

Regarding sex preference in North America, it is generally maintained that couples prefer a child of each gender (e.g., Angrist and Evans 1998 for the United States and McDougall, DeWit, and Ebanks 1999 for Canada), although there is some evidence for son preference in certain immigrant communities.[17] Sex preference in this research is typically investigated by

[16] See Patel et al. (2014), Adewuya et al. (2005), Ekuklu et al. (2004) and Xie et al. (2007).
[17] See, for example, Abrevaya (2009) and Almond et al. (2013).

relating fertility decisions to the sex composition of previously born children (e.g., a second child more likely if the first born is male versus female). Baker and Milligan (2016) take a different approach examining parental views at the time of conception of whether a pregnancy was wanted, or at the right time, solicited after the birth has taken place.[18] Interestingly, in the US mothers are more likely to view the pregnancy as unwanted when the child is a male.[19]

Maternal depression is potentially important for the quality of child/parent interactions. Depressed mothers can be either more withdrawn or more hostile and intrusive, either of which can undermine the preferred "serve and return" interaction thought important to brain development. Exposure to depressed caregivers has been associated with higher levels of stress and lower cognition in children (e.g., Liu et al. 2017, Madigan et al. 2018). Finally, there is evidence that maternal stress transmitted to the child post birth through breast milk can negatively impact behavior (Glynn et al. 2008).

I begin by examining the relationship between the incidence of postpartum depression and postpartum problems and the sex of the first born in nationally representative samples of U.S. (the ECLS-B) and Canadian (the NLSCY) children. I estimate linear regressions of indicators of maternal depression on a dummy variable for males, and controls for geography, child's age and mother's age, education, foreign birth and ethnicity, which differ marginally by sample.[20] The estimates of the dummy variable for boys are presented in table 7.[21]

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[18] The respondent is asked to think back to the time just before the pregnancy in answering, but the question is asked after the child has been born.

[19] This result is obtained when the question is asked when the child is age 2 or younger. If the question is asked when the child is ages 3 through 5, the overall incidence of unwanted pregnancy is considerably lower and if anything the pregnancy is more likely viewed as unwanted if the child is female.

[20] Control variables for the ECLS-B regressions are age (single month), birth state, mother's age (single year), education, foreign birth and indicators for whether the mother is black or Hispanic. Control variables for the NLSCY regressions are child's age effects (single month), mother's age effects (single year), mother's education (4 categories), mother's foreign birth, dummy variables for province, urban size (5 categories) and year of birth.

[21] The estimates of these post partum outcomes and the following estimates for parenting practices, first appeared in Baker and Milligan (2011).

The results from the ECLS-B are in the first panel and are from the wave when the child is 9 months old. In the first three rows are the estimates for 0/1 indicators of whether the mothers reported being depressed, sad, blue either moderately or most of the time over the past week.[22] In the fourth row is the result for a 0/1 indicator that the mother had consulted with a health care professional about their emotional or psychological state. For each outcome the mean is higher if the first born child is male. For depression, the difference at 2½ percentage points is statistically significant and just over 40 percent of the mean for a female child. In the last column is the estimate conditional on the controls which make little difference to the estimated gender differences or their statistical significance.

In the second panel are results for mothers in Canada. This switch in country helps calibrate some subsequent analysis of gender differences in parenting, which is possible with the NLSCY. While clearly cross country institutional and cultural differences can confound the US/Canada comparisons, Canada and the US are generally considered reasonably similar for these purposes compared to other possible comparisons.

There is clear evidence that mothers of first born boys in Canada are also more likely to experience postpartum depression and post-partum problems.[23] The differences are quite substantial—two and half percentage points for postpartum depression and over 6 percentage points for postpartum problems. Again, the inference from the conditional differences is very similar to that from the unconditional differences.

The results in table 7 support the previous research, based on more select samples, that mother's face a heightened risk of post birth depression when their child is a boy. I next

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[22] The variables are formed from the survey variables that record responses on a 4 point scale ranging from rarely of never (<1 day per week) to most or all of the time (5-7 days per week).

[23] Post-partum problems include post-partum haemorrhage, post-partum infection, post-partum depression for more than 14 days and post-partum hypertension.

examine how childhood parenting practices vary by gender. While I cannot causally connect any differences in parenting to the preceding results on depression, as noted previous research suggests that maternal depression can spill over into parent/child interactions.

The NLSCY provides parenting scales based on the responses of the person most knowledgeable about the child (in the vast majority of cases the child's mother) to a series of questions about how they relate with their child. The scales attempt to capture positivity, hostility, consistency and adversity in the parent/child relationship. In each case a higher score means more of the indicated parenting dimension. Each scale is built up from the responses to a series of questions tailored to the age of the child.[24] I also examine the responses to a question whether the respondent considers the child "difficult" for his or her age, a higher score indicating more difficult.[25]

The estimates for first born children aged 0-1 in table 8 indicate that boys receive more hostile parenting, and that boys of this age are also more likely to be viewed as difficult by their parent. At ages 3-5, the largest estimates suggest boys receive more hostile and aversive parenting, although only the estimate for aversive parenting is statistically significant. Overall the message here is that the interactions between the parent and the first born boy are more confrontational.

These results suggest that first born males in the Canada spend their first years in a somewhat different parental environment than their female counterparts. The results are consistent with boys simply being harder to parent, and so require different parenting strategies. However, the results are also consistent with boys and girls posing similar challenges, but some

[24] These scales have been used previously in studies of child development (e.g., Baker et al. 2008).
[25] The parents are asked to rate the "difficulty" of child on a scale from very easy to highly difficult to deal with.

mothers meeting these challenges differently according to the sex of their child, due to the association of a male birth with their mental health.

Conclusions

Childhood environmental influences are viewed as an important, although not the sole, determinant of adult social and economic outcomes. Much of the previous research on childhood environment in developed countries has looked for differences in childhood environments by socioeconomic indicators. This is an important distinction in the data as any intergenerational persistence in childhood environments can feed directly into corresponding persistence in adult outcomes, and thereby undermine social mobility and opportunity.

Relatively understudied in this research is another dimension of inequality—gender. While gender differences in socioeconomic outcomes are persistent over time, much of the research on its causes focuses on adulthood factors such as discrimination, family/work balance and employment decisions. However, there is intuition for gender differences in childhood environment to matter to adult outcomes, given that research has documented that corresponding differences in environment by family income or parents' educations have such effects.

Previous research has shown that there are gender differences in the amount of time children spend with their parents. For example, Baker and Milligan (2016) show that parents spend more time in teaching activities with their first born daughters than their first born sons, starting at very young ages, in Canada, the UK and the US. In the paper I extend this evidence to environmental factors including medical care after birth, breastfeeding, mothers' mental health and parenting practices.

Lower birthweight boys in the US receive more care immediately after birth than their female counterparts, as measured by hospital costs and hospital length of stay. This likely reflects the higher risk of mortality these boys face relative to girls. However, variation in this care around the VLBW threshold reveals that while it has a tangible impact on the one year and 28 day mortality of boys, it has a much smaller to no effect on the mortality of girls. This suggests a possible misallocation of resources. Sex specific thresholds for VLBW might result in greater aggregate return to critical care for these children. These findings also suggest that the consequences of low birth weight for longer term outcomes may differ by sex, a topic which has received little attention the literature.

Breastfeeding is widely counselled for children as a source of nutrition in the first year of life. I find sex differences in this nutritional input in the US favour girls. However, the differences are small, and, based on the guidance of previous research on the benefits of breastfeeding, unlikely to have a substantive impact on longer run gender differences in outcomes.

I also examine measures of the interactions between parents and their young children in Canada. Measures of parenting practice reveal that first born boys experience more confrontational, hostile and aversive parenting than first born girls, on average. Perhaps connected is that mothers in the US and Canada are more likely to report postpartum depression after the (first) birth of a boy. This latter result, from nationally representative data, supports previous findings from small selected samples.

The results in the paper paint a picture of male disadvantage in these selected outcomes. They are consistent with a social extension of the "fragile male" hypothesis (e.g., Kraemer 2000), which is more typically posed as a conjecture about males' relative genetic fragility.

This analysis covers a small selection of the wide array of environmental factors which previous research indicates are important to adult outcomes. Clearly taking a gender lens to the study of these inputs is needed to understand if sex differences in childhood environmental inputs are potentially important. However, equally essential is applying this same lens to the causal study of the relationship between these inputs and adult socio-economic status.

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Table 1: Causes of death in the First Year of Life, 1990-1991 and 1995-2010, US

	All Births (N=575173)			LBW Births (N=383445)		
	Girls	Boys	Difference	Girls	Boys	Difference
Infectious and Parasitic	0.0259	0.026	-0.0001 (0.0004)	0.022	0.0221	-0.0001 (0.0005)
Neoplasms	0.0082	0.0067	0.0016*** (0.0002)	0.0058	0.0041	0.0016*** (0.0002)
Diseases of the blood	0.0162	0.0169	-0.0007** (0.0003)	0.0099	0.0102	-0.0003 (0.0003)
Endocrine	0.0149	0.0144	0.0005 (0.0003)	0.013	0.0123	0.0007** (0.0004)
Nervous system	0.0323	0.0318	0.0005 (0.0005)	0.0367	0.0368	-0.0000 (0.0006)
Ear and mastoid process	0.0001	0.0001	-0.0000 (0.0000)	0.0000	0.0000	-0.0000 (0.0000)
Circulatory system	0.0383	0.0401	-0.0018*** (0.0005)	0.0397	0.0442	-0.0046*** (0.0007)
Respiratory system	0.0433	0.0492	-0.0058*** (0.0006)	0.0301	0.0319	-0.0017*** (0.0006)
Digestive system	0.0136	0.0148	-0.0012*** (0.0003)	0.0122	0.014	-0.0018*** (0.0004)
Genitourinary system	0.0099	0.0102	-0.0003 (0.0003)	0.0087	0.009	-0.0003 (0.0003)
Conditions from perinatal period	0.3432	0.3447	-0.0015 (0.0013)	0.4619	0.4818	-0.0199*** (0.0016)
Congenital malformations	0.1525	0.1327	0.0197*** (0.0009)	0.1405	0.1155	0.0250*** (0.0011)
SIDS	0.0797	0.0887	-0.0091*** (0.0007)	0.028	0.0283	-0.0002 (0.00005)
All other diseases	0.0004	0.0006	-0.0001** (0.0001)	0.0003	0.0005	-0.0002*** (0.0001)
External Causes	0.2216	0.2233	-0.0017 (0.0011)	0.1912	0.1894	0.0018 (0.0013)

Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files. Reported statistics are conditional on death in the first year for the indicated samples. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent.

Table 2: Changes in One Year Mortality at the VLBW Threshold, by Sex, 1983-2010, US

Sample Omission	Pooled	Girls	Boys	Difference
None	-0.0057	-0.0032	-0.0081	0.0049
	(0.0015)***	(0.0020)	(0.0023)***	(0.0030)
	[0.0033]*	[0.0031]	[0.0040]**	(0.0025)*
1500g	-0.0026	-0.0007	-0.0044	0.0037
	(0.0015)*	(0.0020)	(0.0023)*	(0.0031)
	[0.0013]**	[0.0018]	[0.0017]**	(0.0024)
1500g +/-1g	-0.0028	-0.0009	-0.0047	0.0038
	(0.0015)*	(0.0020)	(0.0023)**	(0.0031)
	[0.0013]**	[0.0018]	[0.0017]***	(0.0024)
1500g +/-2g	-0.0021	-0.0002	-0.0040	0.0038
	(0.0016)	(0.0021)	(0.0023)*	(0.0031)
	[0.0013]	[0.0017]	[0.0018]**	(0.0024)
1500g +/-3g	-0.0008	0.0006	-0.0022	0.0028
	(0.0018)	(0.0025)	(0.0027)	(0.0037)
	[0.0015]	[0.0025]	[0.0019]	(0.0033)

Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files. In the first row that sample sizes are 344793 for the pooled sample, 172103 for the sample of girls and 172690 for the sample of boys. In subsequent row births with the indicated weight are omitted from the sample. The reported statistics in the first three rows are estimates of a dummy variable for birthweight below the VLBW threshold following (1). The statistic reported in the fourth row is the estimate of the interaction between sex and VLBW from a pooled regression with a full set of interactions between the control variables and sex. Controls variables include fixed effects for year of birth, mother's and father's five year age groups, race and whether the mother was born out of state, and controls for the child's gestational age and plurality, as well as separate linear trends in birthweight above and below the threshold. Robust standard errors in parenthesis. Standard errors clustered at the gram level in square brackets. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent.

Table 3: Changes in 28 Day Mortality at the VLBW Threshold, by Sex, 1983-2010, US

Sample Omission	Pooled	Girls	Boys	Difference
None	-0.0057	-0.0036	-0.0078	0.0042
	(0.0013)***	(0.0017)**	(0.0019)***	(0.0026)*
	[0.0027]**	[0.0023]	[0.0035]**	(0.0021)**
1500g	-0.0033	-0.0017	-0.0047	0.0030
	(0.0013)**	(0.0017)	(0.0019)**	(0.0026)
	[0.0011]***	[0.0014]	[0.0013]***	(0.0015)*
1500g +/-1g	-0.0034	-0.0018	-0.0049	0.0032
	(0.0013)***	(0.0017)	(0.0019)**	(0.0026)
	[0.0011]***	[0.0014]	[0.0013]***	(0.0016)**
1500g +/-2g	-0.0028	-0.0013	-0.0043	0.0030
	(0.0013)**	(0.0017)	(0.0020)**	(0.0026)
	[0.0010]***	[0.0014]	[0.0012]***	(0.0016)*
1500g +/-3g	-0.0021	-0.0003	-0.0039	0.0036
	(0.0016)	(0.0021)	(0.0023)*	(0.0031)
	[0.0013]	[0.0021]	[0.0016]**	(0.0026)

Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files. In the first row that sample sizes are 344793 for the pooled sample, 172103 for the girls sample and 172690 for the boys sample. In subsequent row births with the indicated weight are omitted from the sample. The reported statistics are estimates of a dummy variable for birthweight below the VLBW threshold following (1). The statistic reported in the fourth row is the estimate of the interaction between sex and VLBW from a pooled regression with a full set of interactions between the control variables and sex. Controls variables include fixed effects for year of birth, mother's and father's five year age groups, race and whether the mother was born out of state, and controls for the child's gestational age and plurality, as well as separate linear trends in birthweight above and below the threshold. Robust standard errors in parenthesis. Standard errors clustered at the gram level in square brackets. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent.

Table 4: Changes in Hospital Costs at the VLBW Threshold, by Sex, 1995-2005, US

Sample Omission	Pooled	Girls	Boys	Difference
None	5353	6224	4258	1966
	(2114)**	(2769)**	(3204)	(4230)
	[3527]	[3919]	[4578]	[4829]
1500g	4736	7161	2044	5117
	(2185)**	(2805)**	(3367)	(4377)
	[3930]	[4271]*	[4833]	[4683]
1500g +/-1g	4243	6730	1455	5275
	(2224)*	(2859)**	(3421)	(4453)
	[3996]	[4333]	[4923]	[4782]
1500g +/-2g	3206	6824	-807.3	7631
	(2217)	(2927)**	(3338)	(4435)*
	[4088]	[4480]	[4871]	[4673]
1500g +/-3g	611	4364	-3428	7792
	(2326)	(3060)	(3510)	(4651)*
	[3497]	[4036]	[4588]	[5090]

Notes: Author's calculations from HCUP data. In the first row that sample sizes are 20423 for the pooled sample, 10199 for the sample of girls and 10224 for the sample of boys. In subsequent row births with the indicated weight are omitted from the sample. The reported statistics are estimates of a dummy variable for birthweight below the VLBW threshold following (1). The statistic reported in the fourth row is the estimate of the interaction between sex and VLBW from a pooled regression with a full set of interactions between the control variables and sex. Controls variables include fixed effects for year of birth, state, controls for mothers' race, twin and multiple births, c-sections, and preterm births, as well as separate linear trends in birthweight above and below the threshold. Robust standard errors in parenthesis. Standard errors clustered at the gram level in square brackets. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent.

Table 5: Changes in Hospital Length of Stay at the VLBW Threshold, by Sex, 1995-2005, US

Sample Omission	Pooled	Girls	Boys	Difference
None	1.083	0.772	1.345	-0.572
	(0.506)**	(0.688)	(0.744)*	(1.012)
	[0.747]	[0.788]	[1.047]	[1.119]
1500g	0.979	1.098	0.827	0.271
	(0.524)*	(0.696)	(0.787)	(1.049)
	[0.840]	[0.825]	[1.078]	[0.951]
1500g +/-1g	0.703	0.829	0.536	0.293
	(0.553)	(0.709)	(0.798)	(1.066)
	[0.811]	[0.771]	[1.074]	[0.956]
1500g +/-2g	0.434	0.804	0.003	0.801
	(0.533)	(0.722)	(0.785)	(1.066)
	[0.825]	[0.795]	[1.048]	[0.904]
1500g +/-3g	-0.085	0.397	-0.601	0.997
	(0.556)	(0.753)	(0.818)	(1.110)
	[0.723]	[0.755]	[0.943]	[0.929]

Notes: Author's calculations from HCUP data. In the first row that sample sizes are 20502 for the pooled sample, 10234 for the sample of girls and 10268 for the sample of boys. In subsequent row births with the indicated weight are omitted from the sample. The reported statistics are estimates of a dummy variable for birthweight below the VLBW threshold following (1). The statistic reported in the fourth row is the estimate of the interaction between sex and VLBW from a pooled regression with a full set of interactions between the control variables and sex. Controls variables include fixed effects for year of birth, state, controls for mothers' race, twin and multiple births, c-sections, and preterm births, as well as separate linear trends in birthweight above and below the threshold. Robust standard errors in parenthesis. Standard errors clustered at the gram level in square brackets. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent.

Table 5: Sample Characteristics of the NIS-Child Breastfeeding Survey, 2010-2018, US

	Mother respondent		Mother respondent-first born		Mother respondent-first born-regression sample	
	Girls	Boys	Girls	Boys	Girls	Boys
Ever breastfed	0.811	0.814	0.818	0.817	0.818	0.819
Duration in first year	180.13	178.11	172.70	170.55	172.05	170.83
Breastfed >3 months	0.656	0.652	0.636	0.632	0.636	0.632
Breastfed >6 months	0.430	0.423	0.402	0.401	0.403	0.402
Breastfed >12 months	0.292	0.279	0.267	0.254	0.267	0.254
N	159,169		60,308		56,574	

Author's calculations from NHIS-Child. All statistics calculated using sample weights.

Table 6 Sex Differences in Breastfeeding Milestones, First Born Children, 2010-2018, US

	Ever Breastfed	Duration first year	Breastfed >3 months	Breastfed >6 months	Breastfed >12 months
Boy	-0.003	-2.908	-0.009	-0.006	-0.017**
	(0.006)	(2.104)	(0.007)	(0.007)	(0.007)
Poverty Cutoff	-0.036***	-2.067	-0.017	-0.004	-0.016*
	(0.009)	(3.190)	(0.011)	(0.011)	(0.010)
Married	0.070***	35.580***	0.106***	0.102***	0.084***
	(0.008)	(3.886)	(0.010)	(0.010)	(0.009)
WIC Benefits	-0.056***	-39.440***	-0.101***	-0.129***	-0.087***
	(0.008)	(3.088)	(0.010)	(0.011)	(0.010)

Author's calculation from NHIS-Child using sample weights. N=56,574. Each column is from a separate linear regression of the indicated dependent variable, on the reported independent variables plus controls for mother's and child's age, mother's education, child's race and number of children in the household. Robust standard errors in parentheses. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent

Table 7: Child Mother's Health Post Birth by the Sex of their First Born Child, US and Canada

	N	Males	Females	Difference	Conditional Difference
			ECLS-B: 9 months		
Depressed	3450	0.087	0.062	0.025*** (0.009)	0.024** (0.011)
Sad	3450	0.083	0.072	0.011 (0.009)	0.008 (0.012)
Blue	3450	0.069	0.063	0.007 (0.008)	0.007 (0.011)
Talked to Doctor about Emotional/Psychological State	3500	0.115	0.107	0.008 (0.011)	0.007 (0.014)
			NLSCY Ages 0 and 1		
Postpartum depression	4754	0.139	0.115	0.024* (0.013)	0.028** (0.013)
Postpartum problems	4723	0.280	0.216	0.064*** (0.022)	0.067*** (0.021)
Mother's self reported depression	7016	0.180	0.165	0.015 (0.012)	0.016 (0.012)

Notes: Author's calculations from ECLS-B and NLSCY. Sample sizes from the ECLS-B are rounded to the nearest 50 observations. The reported regression estimates are the estimated parameter on a 0/1 indicator that the child is male. Control variables for the ECLS-B regressions are age (single month), birth state, mother's age (single year), education, foreign birth and indicators for whether the mother is black or Hispanic. Control variables for the NLSCY regressions are child's age effects (single month), mother's age effects (single year), mother's education (4 categories), mother's foreign birth, dummy variables for province, urban size (5 categories) and year of birth. Robust standard errors in parentheses. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent

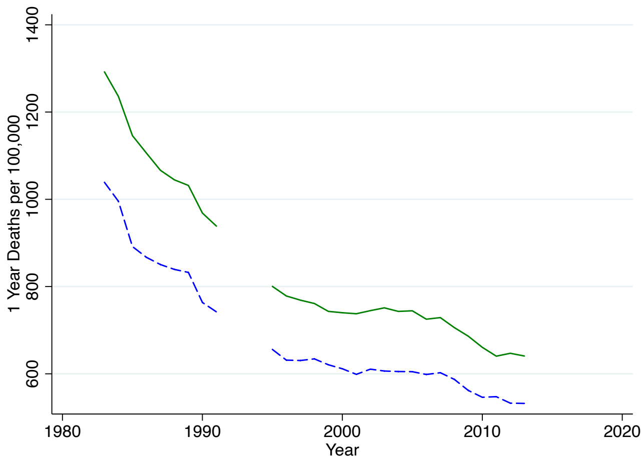
Table 8: Parenting/Child interactions by the Sex of the First Born Child, Canada

	N	Males	Females	Difference	Conditional Difference
Positive: Age 0-1	4900	18.072	18.123	-0.050 (0.090)	-0.056 (0.084)
Hostile: Age 0-1	4904	2.177	2.011	0.166** (0.068)	0.188*** (0.060)
Child is Difficult for age: Age 0-2	6024	2.226	2.080	0.146*** (0.049)	0.140*** (0.046)
Positive: Age 3-5	6847	16.425	16.355	0.070 (0.093)	0.062 (0.082)
Hostile: Age 3-5	6692	8.706	8.500	0.206 (0.141)	0.217 (0.135)
Consistent: Age 3-5	6565	15.502	15.397	0.105 (0.122)	0.080 (0.113)
Averse: Age: 3-5	6812	4.062	3.912	0.150* (0.083)	0.162** (0.080)

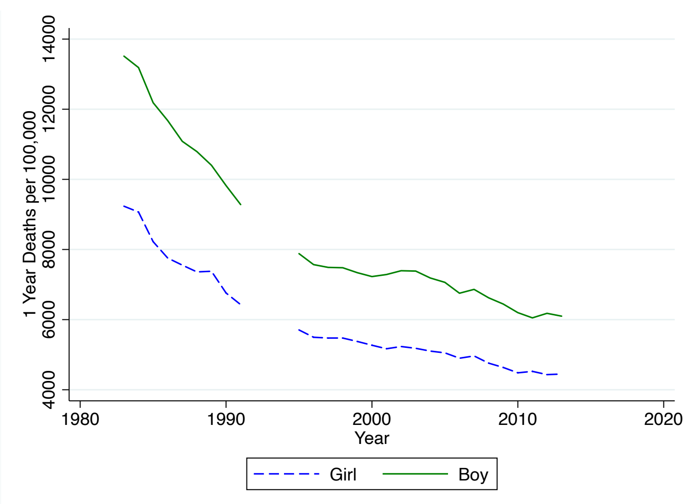
Notes: Author's calculations from NLSCY data. In the regression results the reported statistics are the estimated parameter on a 0/1 indicator that the child is male. Control variables for the Conditional Difference are child's age effects (single month), mother's age effects (single year), mother's education (4 categories), mother's foreign birth, dummy variables for province, urban size (5 categories) and year of birth. Robust standard errors in parentheses. One star indicates significance at the 10 percent level; two stars for 5 percent; three stars for 1 percent

Figure 1: Trends in One Year Mortality of Children , by Sex, US 1983-2010

All Births

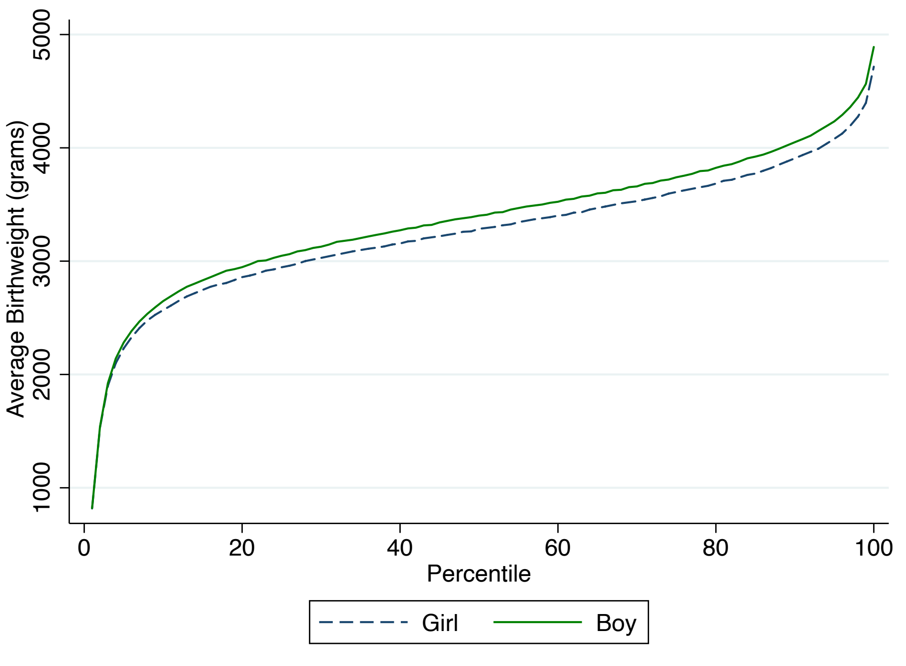


LBW Births



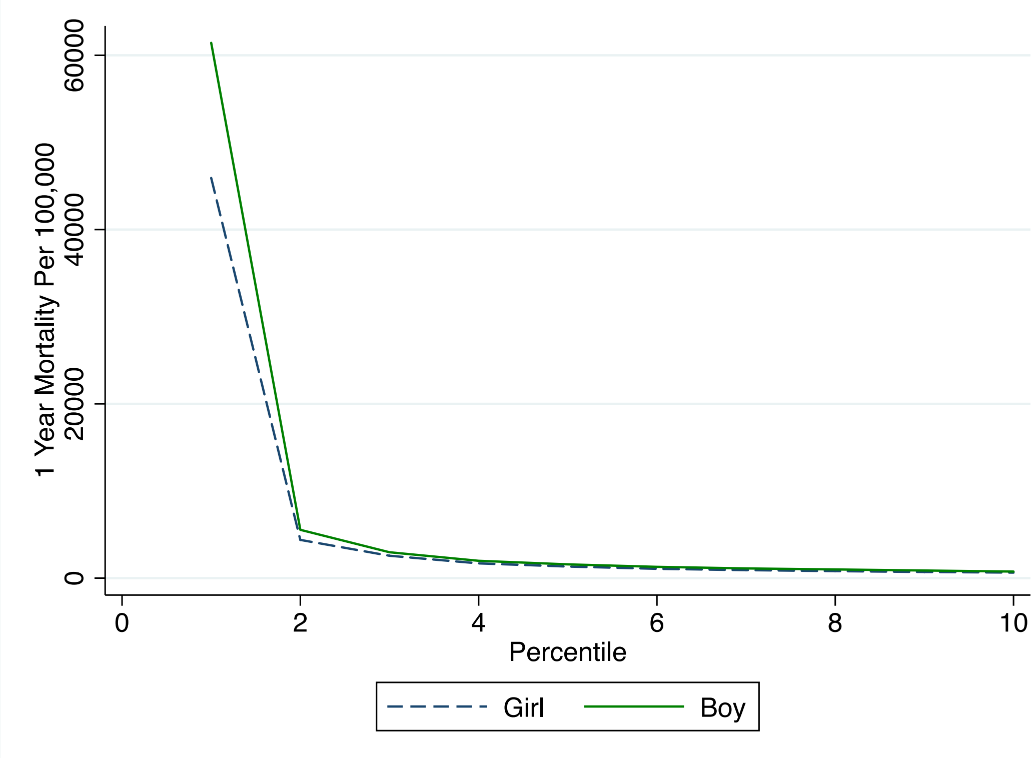
Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files.

Figure 2: Average Birthweight by Sex, by Percentile, US 1983-2010



Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files.

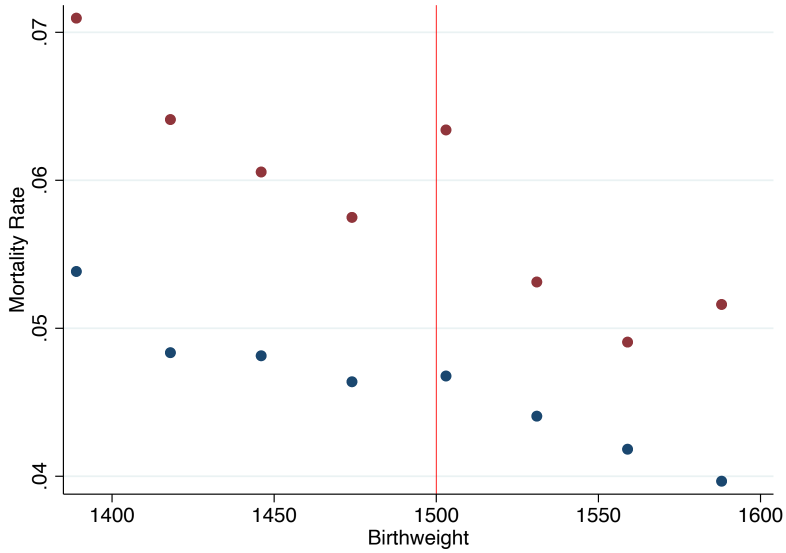
Figure 3: One year Infant Mortality by Sex, by Percentile, Bottom 10 percentiles, US 1983-2010



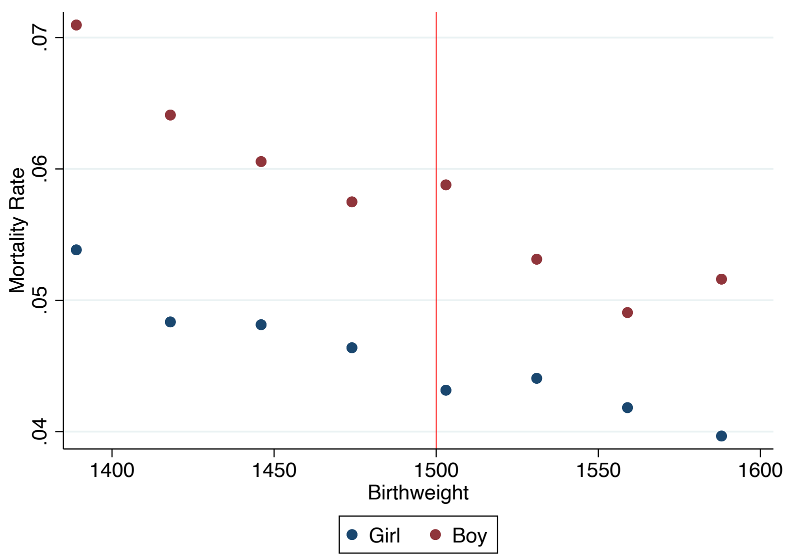
Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files.

Figure 4: One Year Mortality by Sex Around the VLBW Threshold, US 1983-2010

Full Sample



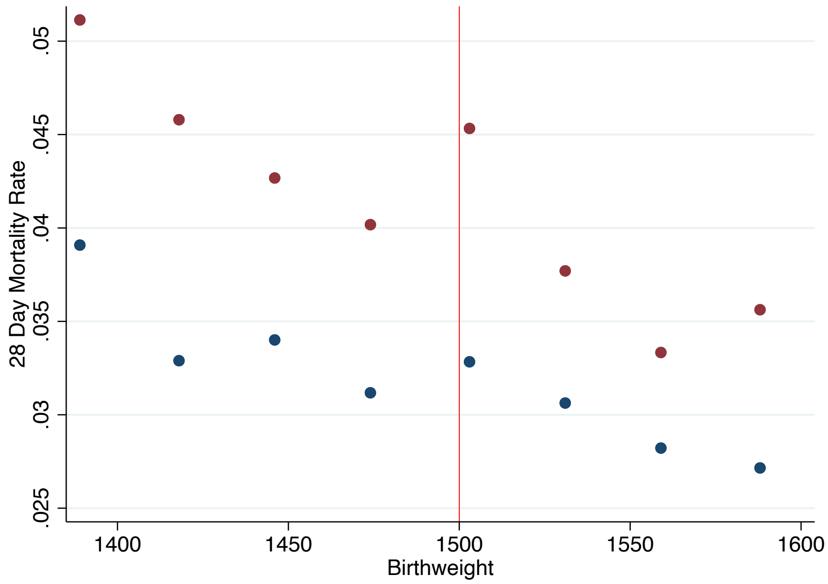
Omit 1500 gram Births



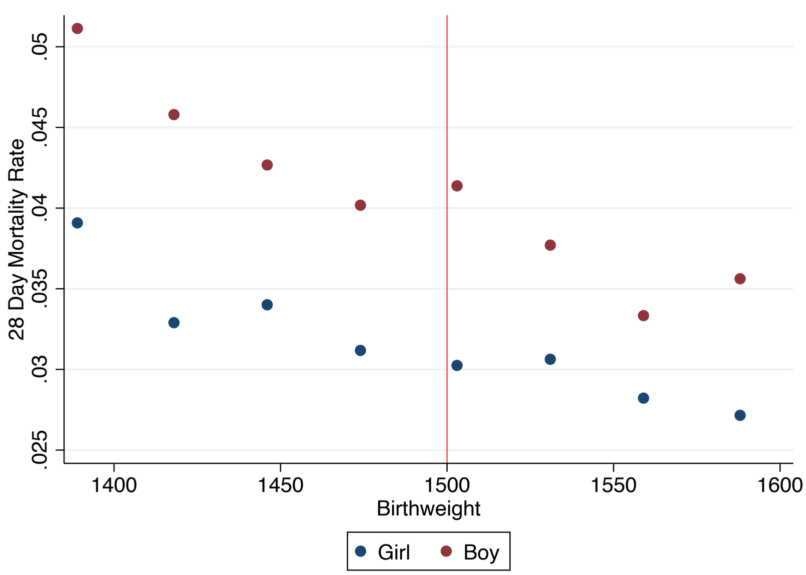
Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files. Median birthweight by one ounce (28g) bins radiating away from the 1500g threshold. Top panel—1500g births included in the first bin above threshold. Bottom panel—1500g births omitted. Plotted at median birthweight in each bin.

Figure 5: 28 Day Mortality by Sex Around the VLBW Threshold, US 1983-2010

Full Sample



Omit 1500 gram Births

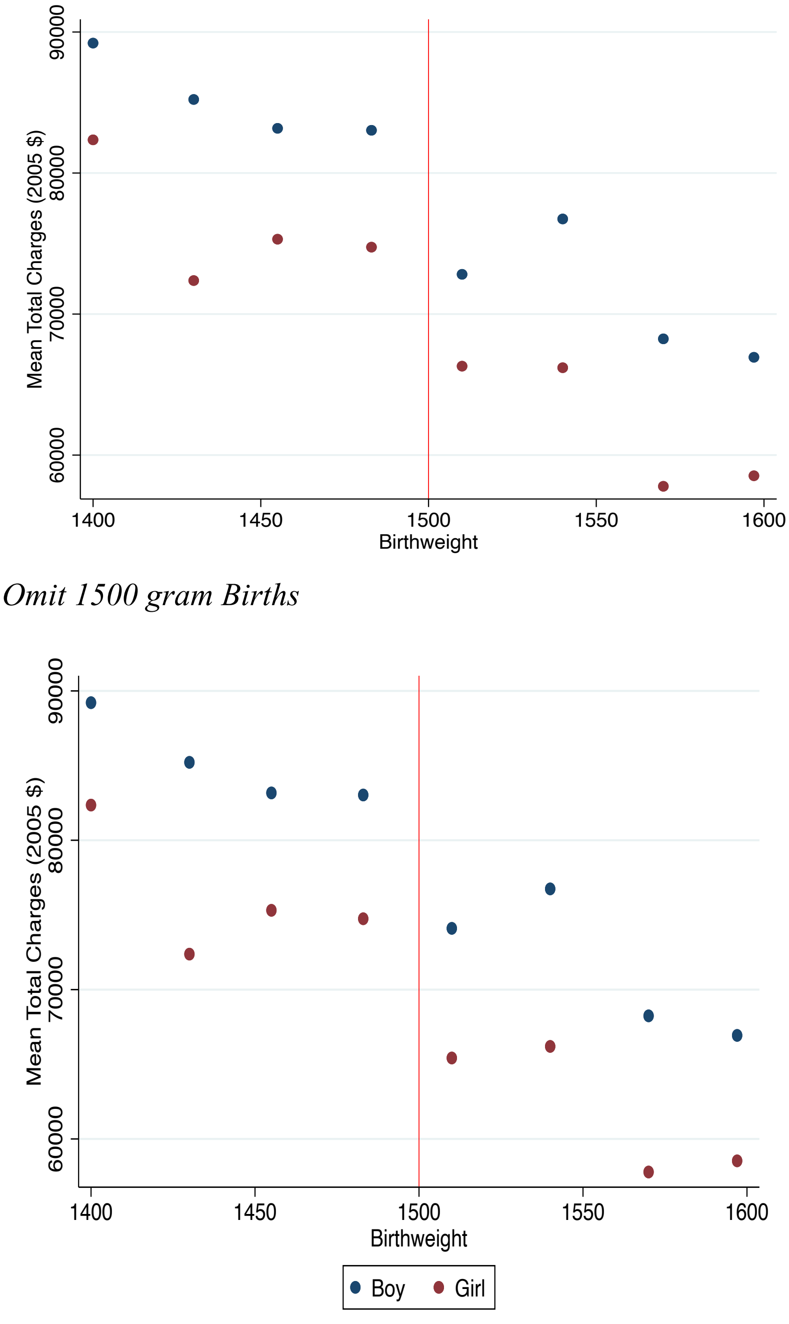


Notes: Author's calculations from NCHS public use birth cohort linked birth/infant death files. Median birthweight by one ounce (28g) bins radiating away from the 1500g threshold. Top panel—1500g births included in the first bin above threshold. Bottom panel—1500g births omitted. Plotted at median birthweight in each bin.

Figure 6: Hospital Charges by Sex Around the VLBW Threshold

Full Sample

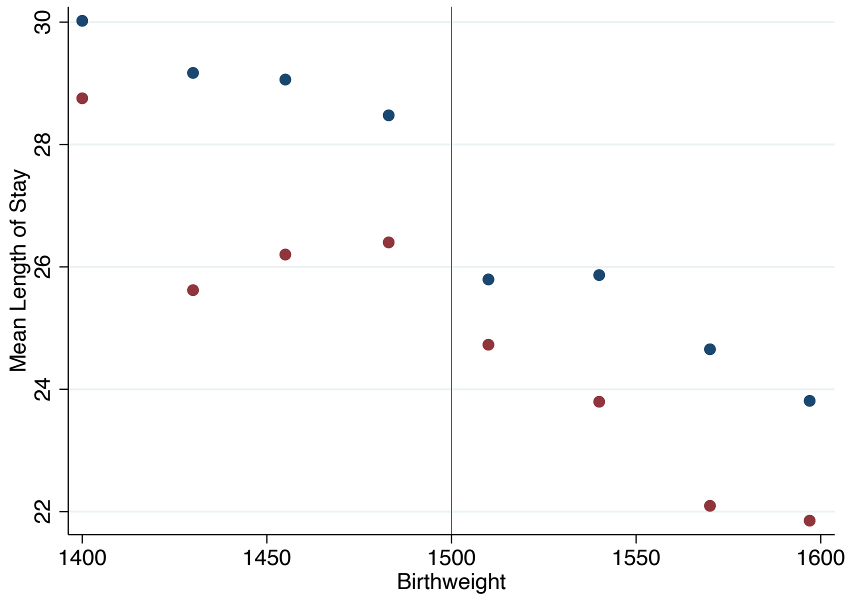
Omit 1500 gram Births



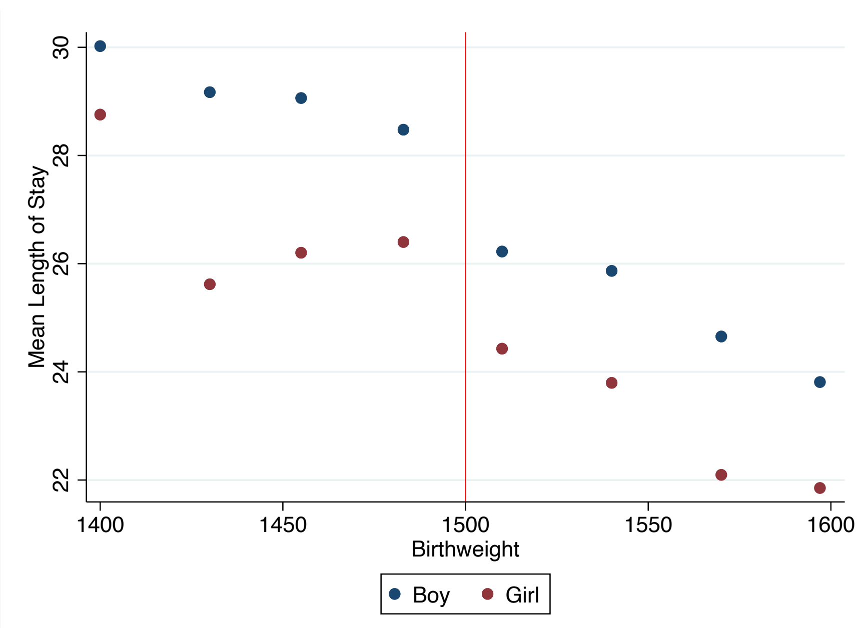
Notes: Author's calculations from HCUP data for births in Arizona, Maryland, New York and New Jersey, 1995-2005. Median birthweight by one ounce (28g) bins radiating away from the 1500g threshold. v Top panel—1500g births included in the first bin above threshold. Bottom panel—1500g births omitted. Plotted at median birthweight in each bin.

Figure 7: Hospital Length of Stay by Sex Around the VLBW Threshold

Full Sample



Omit 1500 gram Births



Notes: Author's calculations from HCUP data for births in Arizona, Maryland, New York and New Jersey, 1995-2005. Birthweight grouped into one ounce (28g) bins radiating away from the 1500g threshold. Top panel—1500g births included in the first bin above threshold. Bottom panel—1500g births omitted. Plotted at median birthweight in each bin.