CAHIER DE RECHERCHE #2008E
DÉPARTEMENT DE SCIENCE ÉCONOMIQUE
FACULTÉ DES SCIENCES SOCIALES
UNIVERSITÉ D'OTTAWA

WORKING PAPER #2008E
DEPARTMENT OF ECONOMICS
FACULTY OF SOCIAL SCIENCES
UNIVERSITY OF OTTAWA

Is excess (fe)male mortality caused by the prenatal environment, child biology, or parental discrimination? New evidence from male-female twins[1]

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December 2020

[1] This paper substantially revises earlier versions that circulated under different titles as Pongou (2009, 2010). I thank Anna Aizer, Jason Boardman, Marie Patience Bayab, Ken Chay, Jakina Debnam Guzman, Esther Duflo, Andrew Foster, Seema Jayachandran, Thomas Lemieux, Emily Oster, Steve Pischke, Mark Pitt, Yona Rubinstein, Olumide Taiwo, and several participants at PAA (2008 and 2009) and IPC (2009) for useful discussions and comments. I am grateful to Mikaila Grace Boyd and Marie Christelle Mabeu for invaluable assistance. Generous financial support was provided by Brown University, the Compton Foundation, The Hewlett Foundation, the University of Ottawa, and Ontario's Ministry of Research, Innovation and Science under ERA grant 400201-190299-2001.
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Abstract

Male-female differences in early age mortality continue to be an important source of child inequality in the world, and are a likely cause of gender disparities in human capital accumulation. Recent literature highlights the important role of the prenatal environment in inducing these differences, in addition to biological influences and gender discrimination in the allocation of resources. However, the distinct roles of these three sets of factors have not been quantified in a unified framework. We propose a new methodology for decomposing male-female differences in mortality into the distinct effects of the prenatal environment, child biology, and parental preferences. We implement this methodology by comparing the mortality sex gap among male-female twins versus all twins in India, a country where daughters are discriminated against, and sub-Saharan Africa, a region where sons and daughters have been found to be valued by their parents about equally. We uncover three main findings: (1) both the prenatal environment and biology increase the mortality risk of boys in these regions; (2) the relative importance of the prenatal environment increases with age, while the effect of biology decreases and even reverses in later childhood; and (3) parental discrimination against girls in India significantly raises their mortality; however, failure to control for the effect of the prenatal environment, biological influences, and the endogeneity of sex determination (due to parental factors and sex-selective abortion) leads traditional methodological approaches to underestimate the effect of discrimination on excess female mortality by 173 percent in the period from birth to 1 year, and by 23 percent between the ages of 1 and 5. Taken together, the findings provide novel quantitative evidence on the relative importance of nature versus nurture in the mortality gap between males and females, and show that the impact of discrimination against girls in certain societies has been underrated.

Key words: Male-female differences in mortality; nature versus nurture; prenatal environment; child biology; discrimination against girls; twins; decomposition methodology

JEL classification: I15, J13, J16, J24, O15

1 Introduction

Male-female differences in early age mortality continue to be an important source of child inequality in many countries of the world. In societies where gender discrimination in the allocation of health and household resources is minimal, female children have a better chance of survival than their male counterparts, an empirical fact long observed by Graunt (1662) and explained in a relatively recent literature by a combination of factors in the prenatal environment and biological influences (Pongou (2013, 2015)), Pongou et al. (2017)).[1] In some countries of South and East Asia, however, the survival advantage of female children diminishes and eventually reverses by age five (see Figure 1).[2] This latter fact has been explained by pro-male bias in investment by parents and other caretakers.[3]

No study has so far quantified the distinct roles of the prenatal environment, child biology, and gender discrimination in male-female differences in early age mortality in a unified framework. Conducting such an analysis is important for several reasons. In addition to contributing to the ongoing debate on the relative importance of nature versus nurture in life expectancy and human capital accumulation, such an analysis is likely to shed light on the real extent of discrimination against girls in certain countries of the world.[4] For example, the observation (in Figure 1) that crude mortality rates are higher for boys compared to girls in the first year after birth in India has led some scholars to claim that discrimination against girls is negligible during this period and is only present between the ages of 1 and 5 (e.g., Sen (2001), Osmani and Sen (2003)). Clearly, such claims are inconsistent with the long-documented existence of infanticide, neglect and

[1] Naeye et al. (1971) and Waldron (1983) argue that male children are biologically weaker and are more susceptible to premature death than female children. Pongou (2013, 2015) propose the preconception origins hypothesis, which is a generalization that accounts for the role of the prenatal environment in addition to biological influences. In a relatively recent study, Biologist William James provides additional evidence in support of this latter hypothesis (James (2015)).

[2] Figure 1 shows that mortality rates are greater for males than females during the first year of life both in sub-Saharan Africa and India; however, between the ages of 1 and 5, while female children continue to die less frequently than their male counterparts in sub-Saharan Africa, the pattern is reversed in India.

[3] There is a large body of literature on the neglect of female children in South and East Asia: see Sen et al. (1988), Sen (1989, 1990b, 1992, 2014), Lin and Luoh (2008), Lin et al. (2014), Ebenstein (2010), Klasen and Wink (2002), Coale and Banister (1994), Behrman (1988), Kynch and Sen (1983), Alderman and Gertler (1997), Basu (1989; 1992), Chen et al. (1981), Borooah (2004), Hazarika (2000), Pande (2003), Oster (2009), Sen and Sengupta (1983), Croll (2000), Preston and Weed (1976), Anderson and Ray (2010), and Jayachandran and Kuziemko (2011), among numerous others. From an economic point of view, it is argued that higher parental investment in sons responds to differential labor market returns by sex (e.g., Rosenzweig and Schultz (1982) Sen (1990a)). This view is consistent with studies showing that female labor market participation and higher female income and education decrease the relative mortality of girls (Rose (1999), Qian (2008), Mason and King (2001), Dreze et al. (1999)). Also, using data from China, Wang (2019) shows that mothers work less, are more likely to be in school, participate more in household decision-making, and have more leisure following the birth of a boy.

[4] Mabeu and Pongou (2020) calculate that among children who died before their fifth birthday between 2000 and 2015 globally, boys outnumbered girls by 34 million. This outcome, however, masks important heterogeneities across regions of the world. In regions where girls are discriminated against in the allocation of household and health resources, they outnumber boys by a significant amount among children who die before reaching the age of five.

abandonment of female newborns in some countries of South and East Asia (e.g., Sudha and Rajan (1999), Coale and Banister (1994)). We argue that such claims follow from the fact that conventional methodological approaches do not account for the effects of prenatal and biological factors when estimating the effect of anti-female discrimination on male-female differences in early age mortality and other health outcomes. traditional approaches also fail to account for the endogeneity of sex determination due to parental factors or sex-selective abortion. In this paper, we introduce a new methodology for estimating the distinct effects of the prenatal environment, child biology, and gender discrimination on differences in mortality between boys and girls. Our methodology addresses the issues mentioned above and leads to new findings.

Our methodological approach exploits variation in the mortality sex gap in male-female twins versus all twins. It addresses two endogeneity issues not accounted for in conventional methodological approaches. First, it accounts for the fact that the biological sex of a fetus is not randomly determined. It is partially determined by prenatal factors that may also affect survival in utero and after birth.[5] Second, our approach accounts for prenatal gender discrimination, including the neglect of female fetuses (Bharadwaj and Lakdawala (2013)) and sex-selective abortion (Sen (1990b), Lin et al. (2014)) in societies where such practices exist. Indeed, by comparing the mortality risk of a female twin to that of her male co-twin, our methodology addresses an endogeneity issue that may arise from the fact that prenatal health care during pregnancy or a parent's choice to give birth to a child may depend on the biological sex of the latter. A female twin is subject to the exact same treatment as her male co-twin during pregnancy. Failure to control for the prenatal environment in conventional methodological approaches may lead to an under-estimation of parental discrimination.[6] We implement this methodology using individual

[5]These factors are numerous (see Pongou (2013)). According to James (1998), the biological sex of a child is determined by parental circumstances and hormonal levels around the time of conception. Levels of hormones are themselves related to such factors as diet, exposure to environmental hazards, occupation, stress, medical conditions, etc. Several empirical studies find results that are consistent with these hypotheses (see Pongou (2013)). For example, Almond and Edlund (2007) and Almond and Mazumder (2011) find that social class and maternal fasting during Ramadan affect sex ratio at birth. Pongou (2013) finds that educated mothers are more likely to bear boys. Garry et al. (2002) found that the sex ratio of children with birth defects was 1.75 males to 1 female if their parents were exposed to certain pesticides in the Red River Valley of Minnesota. Mocarelli et al. (1996) found that parents who were exposed to high concentrations of 2,3,7,8-Tetrachlorodibenzo-pdioxin (TCDD) following a plan explosion in Seveso, Italy, in July 1976 when they were younger than 19 years old were about three times more likely to produce girls than boys, with a male-to-female ratio of 0.38. We argue that such effects are important enough to constitute a source of bias in conventional methodological approaches to studying male-female differences in early age health outcomes.

[6]Indeed, in an environment where a parent may choose to give birth to a child based on the biological sex of the latter, a child whose pregnancy is not terminated may be considered a "desired" child, and therefore a child who will receive a better post-natal treatment than a child a parent could not get rid of. Consistent with this hypothesis, Lin et al. (2014) find that access to sex-selective abortion in Taiwan increased the proportion of males born, but reduced relative neonatal female mortality for higher-parity births. Our approach compares a male and a female in a situation in which the parents could not choose to abort the female child without losing her male counterpart. Indeed, while the selective reduction of multiple fetuses exists, its practice is relatively recent, especially in developing countries.

level data on twins from India, a country where parents have been found to discriminate against daughters, and sub-Saharan Africa, a region where sons and daughters have been found to be valued by their parents about equally.[7] Sub-Saharan Africa is also a useful counterfactual that allows to control for the effect of child biology when estimating the effect of discrimination against girls in India (Sen (1990b), Oster (2009)). This is because both regions have comparable epidemiological profiles with children dying from similar diseases (R. E. Black et al. (2003)). We uncover novel findings about the distinct roles of the prenatal environment, child biology and anti-female discrimination in male-female differences in mortality, and find that these effects vary with child age. Below, we provide a concise summary of our methodology and state our main findings.

1.1 An overview of the methodology and findings

Our identification strategy relies on comparing the sex difference in mortality rates in "male-female twin pairs" versus "all twins".[8] The sex difference in mortality rates in "all twins" (denoted A) is the additive effects of child biology, the prenatal environment, and parental discrimination.[9] However, the sex difference in mortality rates in "male-female twin pairs" (denoted B) only measures the additive effects of biology and parental discrimination.[10] Subtracting B from A thus yields the effects of the prenatal environment.

To estimate the distinct effects of child biology and parental discrimination, assume that there are two types of societies: one that discriminates against female children (called D), and one that does not discriminate based on gender (called ND). In ND societies, sons and daughters are valued by their parents and other caretakers about equally. It follows that, while in D societies, B represents the additive effects of biology and parental discrimination, in ND societies, B only represents the effects of biology. At this point, our methodology shows how the distinct effects of biology and the prenatal environment are estimated in ND societies, because the effect of discrimination in such societies is

Our methodology therefore allows to capture the full extent of anti-female discrimination. We view it as complementing conventional approaches that estimate male-female differences in mortality using singletons. However, we argue that conventional approaches are likely to understate the full extent of anti-female discrimination because they do not capture the fact that discrimination often manifests itself even before birth in the form of the abortion of female fetuses.

[7]Several studies find no gender discrimination in the allocation of health resources among children in sub-Saharan Africa (Deaton (1989), Garenne (2003), Mabeu and Pongou (2020)).

[8]All twins include same-sex twins (male-male twins and female-female twins) and mixed-sex twins (male-female twins).

[9]This assumption is consistent with models generally used by biologists and geneticists to disentangle the effects of genetic and environmental factors on health outcomes (e.g., Evans et al. (2002), Neale and Cardon (1992)); see also Fowler et al. (2008) for a study on the role of genetic factors in political participation.

[10]B is estimated (from the same sample of twins as A) using a twin fixed effect (FE) linear probability model (LPM). This latter model automatically controls for the effect of the prenatal environment. Also, it is easy to notice that estimating a twin FE LPM in the sample of "all twins" is equivalent to estimating an LPM in the subsample of "male-female twin pairs."

zero by definition.

To estimate the distinct effects of biology and parental discrimination in D societies, we assume that biology has the same effect in these societies as in ND societies.[11] We therefore derive the effects of discrimination in D societies by subtracting the effect of biology in ND societies (which is B for ND societies) from the combined effects of biology and parental discrimination in D societies (which is B for D societies). In other words, the effect of parental discrimination is obtained by comparing the sex difference in mortality risk (estimated from the sample of male-female twin pairs) in D versus ND societies, which completes our decomposition analysis.

We apply this decomposition methodology to large samples of twins extracted from Demographic and Health Surveys. These surveys are cross-sectional and representative at the national and subnational level, and have been conducted in most developing countries since the late eighties and early nineties. For our purpose, we use data from sub-Saharan Africa and India, with the former region being a relevant counterfactual for the latter (Sen (1990b), Oster (2009), Jayachandran and Kuziemko (2011)) for reasons previously mentioned. For example, comparing India to European countries or other regions would not be a pertinent comparison due to the differences in epidemiological profiles.

In decomposing sex differences in mortality, we distinguish two main periods of child development: infancy, which is the period from birth to 12 months,and childhood, which is the period between the ages of 1 and 5 (12-60 months). Infancy is further divided up into two periods: the neonatal period (from birth to 1 month), and the postneonatal period (1-12 months). We find that the male-female difference in infant mortality estimated from the sample of all twins (that is, the combined effects of the prenatal environment, biology, and parental discrimination (A)) is 45 per thousand points in sub-Saharan Africa and 27 per thousand points in India. The male-female difference in infant mortality estimated from the sample of male-female twin pairs (that is, the combined effects of biology and parental discrimination (B)) is 27 per thousand points in sub-Saharan Africa and -10 per thousand points in India.[12] It follows from our decomposition methodology that in sub-Saharan Africa, the prenatal environment, male biology and parental discrimination raise the mortality of male twins by 18, 27 and 0 per thousand points, respectively. In India, the prenatal environment and male biology raise the mortality of male twins by 37 and 27 per thousand points, while discrimination against daughters raises their mortality by 37 per thousand points. We replicate this analysis in the neonatal and postneonatal periods to take into account the fact that factors contributing to male-female differential

[11] In studies comparing sub-Saharan Africa to India or other countries of South and East Asia, Sen (1990b, 1992) and Oster (2009) make similar assumptions. As in earlier versions of the current paper (Pongou (2009, 2010)), we argue that these comparisons are appropriate because the epidemiological profiles of these regions are similar for children (R. E. Black et al. (2003)).

[12] The male-female difference in mortality being -10 per thousand points in India simply means that the mortality rate is on average 10 per thousand points lower for a male twin compared to his female co-twin.

mortality rates might differ across ages. We find the effect of discrimination against girls in both periods to be substantial.

Applying our decomposition methodology to child mortality, we find that the male-female difference in this outcome estimated from the sample of all twins is 4 and -16 per thousand points in sub-Saharan Africa and India, respectively. In addition, the male-female difference in child mortality (mortality occurring between the ages of 1 and 5) estimated from the sample of male-female twin pairs is respectively -8 and -31 per thousand points in these regions. These results imply that prenatal environmental factors raise the mortality rate of boys by 12 and 15 per thousand points in sub-Saharan Africa and India, respectively; the biological make-up of boys lowers their mortality rate by 8 per thousand points in both regions. Parental discrimination against daughters in India increases their mortality rate by 23 per thousand points.

The findings of this study lead to three main conclusions: (1) both the prenatal environment and biology increase the mortality risk of boys both in India and sub-Saharan Africa; (2) the relative importance of the prenatal environment increases with age, while the effect of biology decreases and even reverses between the ages of 1 and 5; and (3) parental discrimination against girls in India significantly increases their mortality; however, failure to control for the effect of the prenatal environment, biological influences, and the endogeneity of sex determination (due to parental factors and sex-selective abortion) leads conventional methodological approaches to underestimate the effect of anti-female discrimination by 173 percent during infancy, and 23 percent during childhood.[13]

1.2 Plan of the paper

The remainder of this paper is organized as follows. Section 2 situates our paper in the closely related literature. Section 3 presents our methodology. In Section 4, we describe the data used for the analysis. Section 5 presents our findings. Section 6 discusses additional results. Section 7 summarizes the key findings and concludes the study.

2 Closely related literature

The present paper contributes novel quantitative evidence on the relative importance of nature versus nurture in male-female differences in life expectancy and human capital accumulation among children. Its methodology extends that in Pongou (2013) and Pongou et al. (2017) to societies that discriminate against girls in the allocation of health and household resources. Therefore, the proposed approach for decomposing male-female mortality differences into the distinct effects of the prenatal environment, child biology

[13] In particular, conventional approaches underestimate the effect of discrimination on excess female mortality by 446 percent in the neonatal period and 65 percent in the postneonatal period.

and parental discrimination against girls is more general and can be applied to a much greater range of contexts. The implementation of this methodology using data from India is also a new approach. In particular, we find that when parents are unable to discriminate against girls in the prenatal period, they discriminate even more during the post-natal period. This implies that conventional methodological approaches that control neither the effect of the prenatal environment and behavior nor the effect of biology grossly underestimate the effect of parental discrimination against excess female mortality, which is exactly what we find.

In addition, although the implementation of our methodology uses sub-Saharan Africa only as a "comparison group" or a "counterfactual" to control for the effect of child biology, we obtain new findings for this region too. Indeed, earlier studies testing the preconception origins hypothesis focus only on mortality occurring in the period from birth to 12 months in sub-Saharan Africa (Pongou (2013), Pongou et al. (2017)) and the United States (Pongou (2015)). The present paper adds two new findings. First, we perform a separate decomposition analysis for neonatal and postneonatal mortality. Second, we analyze a new important outcome–child mortality (or mortality occurring between the ages of 1 and 5)–and find, contrary to the long-held biological theory explaining the sex gap in health and mortality, that male biology favors male survival during the childhood period. In other words, excess male mortality during this period in non-discriminatory societies is mostly driven by unfavorable prenatal factors.[14]

By using sub-Saharan Africa as a counterfactual for India, our paper shares similar motivations with prior studies (e.g., Sen (1990b, 1992, 2003), Oster (2009), Jayachandran and Kuziemko (2011)), but we significantly differ in our scopes, methodology, and results. To the best of our knowledge, our study is the first to decompose the sex gap in mortality into the effects of three important factors: the prenatal environment, child biology, and anti-female discrimination. A distinctive feature of our analysis is in explicitly taking into account the fact that offspring sex ratio is not random. In doing so, we find that the prenatal environment increases male mortality both in India and sub-Saharan Africa, and that its effect tends to dominate biological effects as a child grows. In addition, by controlling for both the effects of biology and the prenatal environment using male-female twins, we find that discrimination against girls has a huge effect on their mortality in India, and that this effect is substantial even in the first year of life despite the fact that crude mortality rates tend to be higher for males than for females in this period. In particular, we find that conventional methodological approaches substantially underestimate the effect of discrimination in the neonatal period.[15] To our knowledge, these findings are new

[14]In an earlier version of this paper, we discuss the implications of our results for singleton births (Pongou (2009)).

[15]Importantly, we find that if a method only accounts for the effect of biology but does not account for the effect of the prenatal environment, it still leads to an underestimation of the true effect of anti-female discrimination on girls' mortality. The same is true if a method only accounts for the effect of the

and greatly contribute to our understanding of gender inequality in post-natal mortality trends.

Our finding that biology increases male mortality is consistent with a very large literature that has documented sex differences in immune systems, neurodevelopmental disorders, genetic disorders, unintentional injuries, endocrine response to perinatal stress, and tolerance of prenatal and postnatal malnutrition (e.g., Naeye et al. (1971), Sen (1983, 1985, 1998), Chao (1996)). However, as already noted, this effect is overestimated by conventional methodological approaches and diminishes in importance as a child develops. In particular, our finding that biology favors male survival between the ages of 1 and 5 may be consistent with the fact that females suffer a higher incidence of autoimmune diseases compared to males at certain ages, despite having a stronger immune system (Ahmed et al. (1985), Chao (1996), Bouman et al. (2005)).[16] In addition, the fact that the preconception environment increases the mortality of boys is consistent with other studies showing that other types of negative shocks such as wars affect males more than females (Tapsoba (2020)). Conversely, positive shocks, including institutional (Pongou et al. (2017), Mabeu and Pongou (2020)) and technological (Bharadwaj et al. (2020)), may benefit boys more than girls in terms of survival and other health outcomes.

As already acknowledged, numerous studies have documented the effect of parental bias against daughters. However, by not controlling for prenatal and biological factors, most of them have failed to detect any such effect during in the "very early" period of life following birth.[17] This has led some scholars to claim that female discrimination has damaging effects only during childhood (e.g., Sen (2001), Osmani and Sen (2003)), which is a surprising conclusion in light of the large body of research documenting the existence of infanticide, neglect and abandonment of female newborns in several countries of South and East Asia (e.g., Sudha and Rajan (1999), Coale and Banister (1994)). Our decomposition methodology yields results that are consistent with these realities.

Finally, we view our study as a contribution to the few papers by economists who have used twin samples to study later life outcomes, while at the same time addressing a number of important methodological issues (e.g., Almond et al. (2005), Royer (2009), Behrman and Rosenzweig (2004), Oreopoulos et al. (2006), S. E. Black et al. (2007), Pongou (2013, 2015), Pongou et al. (2017)). While most of these papers analyze data

prenatal environment but does not account for the effect of biology.

[16]Preston and Weed (1976) find excess female mortality from certain diseases such as tuberculosis at ages 5-29, influenza-pneumonia-bronchitis at ages 5-14, and certain infectious and parasitic diseases at ages 1-14. Also, analyzing national data from the World Health Organization, Garenne (1994) finds that mortality from measles is higher for females compared to males.

[17]For example, as mentioned earlier, the probability of death in the period from birth to 12 months is 27 per thousand points higher for male twins than female twins in India. Not accounting for the effects of the prenatal environment and biology would therefore lead one to conclude to the absence of discrimination against daughters in the first year after birth. However, controlling for these factors shows that discrimination substantially increases female mortality by 37 per thousand points during this period.

collected in developed countries, our study is among the very few based on large samples of twins from developing countries.

3 Econometric model

3.1 Estimating sex differences in mortality

We estimate the male-female gap in mortality using the following specification:

$$M_{iht} = \theta Male_i + X_{hyt}\pi + u_h + v_{ht} + w_{iht} \quad (1)$$

where M_{iht} is a dummy variable indicating whether child *i*, born to parents *h*, died at time *t* (M_{iht} is equal to 1 if *i* died at time *t* and 0 otherwise); $Male_i$ is a dummy indicator for whether child *i* is male; X_{hyt} is a vector of observed parental and household characteristics thought to be correlated with sex and mortality; u_h captures parental time-invariant unobserved heterogeneity; v_{ht} captures parental time-variant unobserved heterogeneity; and w_{iht} captures a child random unobserved shock (not correlated with sex). This model differs from conventional estimation approaches by assuming that a child's biological sex is endogenous to prenatal factors that may also affect mortality.[18] These prenatal factors are unobserved in our model. Indeed, this endogeneity is captured in equation (1) by the unobserved variables u_h and v_{ht}. Importantly, the assumption that child sex is randomly determined in studies that follow conventional estimation approaches has led to an ambiguous interpretation of the parameter of interest, θ, which measures male-female difference in the probability of death. In these studies, θ is generally interpreted as measuring the effect of boys' weaker biology when it is positive (that is, when mortality risk is greater for boys than for girls), or the effect of parental discrimination against girls when it is negative. Not surprisingly, as discussed earlier, the observation that *infant mortality* is higher for boys than for girls in both sub-Saharan Africa and India (Figure 1) has led some studies to claim that parental discrimination has a negligible effect on girls' mortality during the first year after birth in India, which is contrary to reality. We will show that once the endogeneity of child sex is corrected, it becomes clearly evident that discrimination against girls in India starts very early. We will decompose θ into the distinct effects of the prenatal environment, child biology, and discrimination against girls.

For our purpose, the unobserved variables u_h and v_{ht} can be interpreted as parental prenatal circumstances and gender bias, respectively. We can explicitly write v_{ht} as the

[18] Pongou (2013) summarizes the literature showing that child sex is not randomly determined. It is partially determined by parental factors that may also affect child mortality in utero or after birth (see also James (1998, 2015)). Moreover, in a context in which sex-selective abortion is prevalent, the sex of a newborn can be thought of as being determined by parents.

sum of time-variant parental prenatal conditions (p_{ht})[19] and parental bias (b_{ht}); that is, $\nu_{ht} = p_{ht} + b_{ht}$. It follows that a cross-sectional linear probability model (LPM) estimate of θ is the additive effects of child biology, prenatal factors, and parental preferences. The effect of parental time-invariant unobserved factors can be removed by comparing the mortality of male-female siblings (children born to the same parents). This can be achieved by estimating a sibling fixed effect regression. For this purpose, assume that i and j are opposite-sex siblings. We rewrite Equation (1) for i and j, respectively, to obtain the following equations:

$$M_{iht} = \theta_{SFE} Male_i + X_{ht}\pi + u_h + p_{ht} + b_{ht} + w_{iht} \quad (2)$$
$$M_{jht'} = \theta_{SFE} Male_j + X_{ht'}\pi + u_h + p_{ht'} + b_{ht'} + w_{jht'} \quad (2)'$$

Therefore, taking (2)-(2)' leads to the following equation:

$$M_{iht} - M_{jht} = \theta_{SFE}(Male_i - Male_j) + (X_{ht} - X_{ht'})\pi + p_{ht} - p_{ht'} + b_{ht} - b_{ht'} + w_{iht} - w_{jht'} \quad (3)$$

Estimating Equation (3) using a sibling fixed effect regression yields an estimate of θ_{SFE}, which measures the probability of death between male and female siblings. To the extent that parental prenatal circumstances (e.g., health and environmental exposure) vary over time, θ_{SFE} includes the effect of parental time-variant unobserved factors, $p_{ht} - p_{ht'}$. To account for these prenatal circumstances, we resort to a twin fixed effect regression, which compares the mortality risk of a male twin with that of his female co-twin. Assume that $(i, -i)$ is a pair of opposite-sex twins, with i being the male and $-i$ being the female. Rewriting Equation (1) for each leads to the following equations:

$$M_{iht} = Male_i \theta_{TFE} + X_{ht}\pi + u_h + p_{ht} + b_{ht} + w_{iht} \quad (4)$$
$$M_{-iht'} = Male_{-i} \theta_{TFE} + X_{ht}\pi + u_h + p_{ht'} + b_{ht'} + w_{-iht'} \quad (4)'$$

Since the prenatal environment is the same for a pair of twins (that is, $p_{ht} = p_{ht'}$), taking (4)-(4)' yields:

$$M_{iht} - M_{-iht} = \theta_{TFE}(Male_i - Male_{-i}) + b_{ht} - b_{ht'} + w_{iht} - w_{-iht'} \quad (5)$$

The term $b_{ht} - b_{ht'}$ is still correlated with $Male_i - Male_{-i}$, which implies that estimating equation (5) using a male-female twin fixed effect regression (or estimating it from a sample of opposite-sex twins) yields an estimate of θ_{TFE}, which is the additive effects of child biology and parental bias. Note that by subtracting θ_{TFE} from the

[19] The prenatal circumstances of parents that determine offspring sex ratios such as diet, occupation and exposure to environmental hazards generally change over time.

cross-sectional LPM estimate of θ, we obtain an estimate of the effect of the prenatal environment. Also, in societies where parents do not discriminate against children of a given sex in the allocation of household resources (that is, $b_{ht} - b_{ht'} = 0$), θ_{TFE} measures the effect of male biology since $w_{iht} - w_{-iht}$ is uncorrelated with $Male_i - Male_{-i}$ by assumption.

We estimate θ and θ_{TFE} using large samples of twins, which allows us to separate out the effects of the prenatal environment, child biology, and parental discrimination.This decomposition analysis is explained in the next section.

3.2 Decomposing the sex difference in mortality into the effects of the prenatal environment, child biology, and parental discrimination

We posit that θ, which is the male-female difference in mortality estimated from Equation (1), is the additive effects of the prenatal environment, child biology and parental preferences.[20] That is:

$$\theta = \theta_1 + \theta_2 + \theta_3 \quad (6)$$

where θ_1 is the effect of the prenatal environment, θ_2 the effect of child biology, and θ_3 the effect of parental discrimination.

We assume that parental discrimination varies from one society to another. More precisely, we assume two types of societies: non-discriminatory (ND) and discriminatory (D). The effect of parental discrimination in a non-discriminatory society is zero by definition. We also assume that the effect of male biology on sex differences in mortality is the same in discriminatory and non-discriminatory societies. Both assumptions can be formally expressed as follows.

$$\begin{cases} \theta_3^{ND} = 0 \\ \theta_2^{D} = \theta_2^{ND} \end{cases} \quad (7)$$

Plugging the first and the second equation of System (7) into Equation (6), and rewriting Equation (6) for discriminatory and non-discriminatory societies, respectively, yields the following equations:

$$\begin{cases} \theta^{ND} = \theta_1^{ND} + \theta_2^{ND} \\ \theta^{D} = \theta_1^{D} + \theta_2^{D} + \theta_3^{D} \end{cases} \quad (8)$$

[20]As noted earlier, this simplifying additivity assumption largely follows from biological models often used to disentangle the effects of genetic and environmental factors on health outcomes (e.g., Evans et al. (2002), Neale and Cardon (1992)). It is also empirically motivated. Indeed, in Section 6, we discuss some results supporting this assumption. In particular, we did not find any strong interaction between the effects of biology and the prenatal environment.

We then separate out $\theta_1^{ND}$and θ_2^{ND} on one hand, and θ_1^{D}, θ_2^{D} and θ_3^{D} on the other hand. Estimating Equation (1) yields θ^{ND} and θ^{D} for non-discriminatory and discriminatory societies, respectively. Estimating a twin fixed effect regression (Equation (5)) yields θ_{TFE}, which is the additive effects of biology and parental discrimination. Given that the effect of parental discrimination is zero in non-discriminatory societies, θ_{TFE} in these societies solely measures the effect of biology. That is:

$$\begin{cases} \theta_{TFE}^{ND} = \theta_2^{ND} \\ \theta_{TFE}^{D} = \theta_2^{D} + \theta_3^{D} \end{cases} \quad (9)$$

Plugging the first and the second equation of System (9) into the first and the second equation of System (8), respectively, and re-arranging, allows us to extract the effect of the prenatal environment as follows:

$$\begin{cases} \theta_1^{ND} = \theta^{ND} - \theta_{TFE}^{ND} \\ \theta_1^{D} = \theta^{D} - \theta_{TFE}^{D} \end{cases} \quad (10)$$

We have separated out the effects of the prenatal environment and biology in non-discriminatory societies. For discriminatory societies, it remains to separate out the effects of child biology and parental discrimination. Remember that the effect of biology is the same in non-discriminatory and discriminatory societies: $\theta_2^{D} = \theta_2^{ND} = \theta_{TFE}^{ND}$. Plugging this latter equation into the second equation of System (9) and re-arranging yields the effect of parental discrimination:

$$\theta_3^{D} = \theta_{TFE}^{D} - \theta_{TFE}^{ND} \quad (11)$$

We summarize the results obtained from System (7) through Equation (11) below:

$$\begin{cases} \theta_1^{ND} = \theta^{ND} - \theta_{TFE}^{ND} \\ \theta_2^{ND} = \theta_{TFE}^{ND} \\ \theta_3^{ND} = 0 \\ \theta_1^{D} = \theta^{D} - \theta_{TFE}^{D} \\ \theta_2^{D} = \theta_{TFE}^{ND} \\ \theta_3^{D} = \theta_{TFE}^{D} - \theta_{TFE}^{ND} \end{cases} \quad (12)$$

System (12) separates out the roles of the prenatal environment, child biology and parental discrimination in sex differences in mortality in discriminatory and non-discriminatory societies.

3.3 Empirical strategy

3.3.1 Estimating the effects of the prenatal environment, child biology, and parental discrimination by age

It is possible that the distinct effects of the prenatal environment, child biology and parental discrimination against girls on sex differences in mortality vary with age. We distinguish two main periods of child development: infancy (I)–the period from birth to 1 year–, and childhood (CH)–the period between the ages of 1 and 5. Infant mortality is therefore measured as the probability of dying in the first year of life after birth, child mortality is measured as the probability of dying between the first and fifth birthdays, conditional on surviving the infant period.

In most developing countries, mortality occurring during the neonatal period accounts for a large fraction of under-five mortality (R. E. Black et al. (2003)). Therefore, we further divide up the infant period into the neonatal (NN) period (that is, the period from birth to 1 month), and the postneonatal (PNN) period (1-12 months). Neonatal mortality is measured as the probability of dying during the neonatal period, and postneonatal mortality is the probability of dying during the postneonatal period.

We estimate Equations (1) and (5) for each of the periods just defined. Our decomposition of the sex gap in mortality therefore leads to the derivation of the parameters in System (12) for each period (P) as follows:

$$\begin{cases} \theta^{ND}_{1,P} = \theta^{ND}_{P} - \theta^{ND}_{TFE,P} \\ \theta^{ND}_{2,P} = \theta^{ND}_{TFE,P} \\ \theta^{ND}_{3,P} = 0 \\ \theta^{D}_{1,P} = \theta^{D}_{P} - \theta^{D}_{TFE,P} \\ \theta^{D}_{2,P} = \theta^{ND}_{TFE,P} \\ \theta^{D}_{3,P} = \theta^{D}_{TFE,P} - \theta^{ND}_{TFE,P} \end{cases} \quad \text{with } P = I, NN, PNN, \text{ or } CH \qquad (13)$$

3.3.2 Choice of discriminatory and non-discriminatory societies

Our choice of discriminatory and non-discriminatory societies is based on studies conducted in different regions. It is well documented that in most South and East Asian countries, parents have strong pro-male bias, discriminating against female children in the allocation of food, health care, and other resources (e.g., Sen (1990b, 1992, 2003), Jayachandran and Kuziemko (2011)). On the contrary, there is little evidence of such discrimination in sub-Saharan Africa. Based on household data, Garenne (2003) finds that the probability of dying before the fifth birthday is higher for boys than girls, but investment in health care such as immunization does not significantly differ between the two sexes. Further evidence for the symmetrical treatment of boys and girls in a sub-Saharan African country is provided by Deaton (1989). Using household expenditure data from Côte d'Ivoire, Deaton (1989) finds no gender bias in the allocation of goods, while finding a statistically non-significant pro-male bias in Thailand. Mabeu and Pongou (2020) find no gender bias in prenatal and postnatal allocation of health inputs in sub-Saharan African countries. The findings of these studies support the assumption that sub-Saharan Africa is non-discriminatory, as also recognized by Sen (1990b, 1992, 2003). For our analysis, we therefore use data from India, considered a discriminatory society, and sub-Saharan African countries, considered as non-discriminatory.

4 Data

4.1 Data sources

We use data from Demographic and Health Surveys (DHS) conducted in thirty sub-Saharan African countries, and two National Family Health Surveys (NFHS) conducted in India. Appendix Table A-1 lists all the surveys used for the analysis and provides information on the years in which they were conducted. The DHS and the NFHS surveys are conducted by the same organization (Measure DHS). For most variables, these surveys are comparable across countries and across years. They are also representative at the national and subnational level. They employ a two-stage probabilistic sampling technique to select clusters (or census enumeration zones) at the first stage and households at the second stage. Socioeconomic and demographic information is collected on the characteristics of clusters, households and household members.

In each household, completed information on fertility history is provided by selected women aged between 15 to 49 years old. Information on each live birth includes date of birth; whether the birth is a singleton or a multiple birth; whether the child is still alive; and if the child is dead, when death occurred. For our purpose, the subsample of twins is most relevant. We analyze 75 surveys from sub-Saharan Africa and 2 surveys from India. Merging these surveys yields 1,670,477 observations (or live births) for sub-Saharan Africa, and 543,981 for India. We provide detailed information on the sample size of all births for each country and each survey year in Appendix Table A-1.

Table 1 shows that the sample size of twins is 50,994 for sub-Saharan Africa and 6,920 for India. They represent respectively 3.05 percent and 1.27 percent of the sample of all live births in these settings. Twinning rates vary across sub-Saharan Africa (Table A-1), but the reasons are not entirely known. The proportion of twins for sub-Saharan Africa is a bit smaller compared to that found in the United States by Almond et al. (2005). It is low for India, although the reasons for this are not known. In India, male-male, female-female, and opposite-sex twins represent respectively 35%, 33% and 32% of twins. In sub-Saharan Africa, these figures are respectively 31%, 30% and 39%.

4.2 Summary statistics

We provide some descriptive statistics showing how twins and singletons are comparable along several demographic and socioeconomic factors. The goal of this comparison is to show that twins are not a selected population with respect to these factors.

4.2.1 Sex ratios

The proportion of male births is 0.508 and 0.504 for singletons and twins, respectively, in sub-Saharan Africa, and 0.520 and 0.514 in India. This indicates a slightly lower proportion of male among twins in both settings. However, the relevant comparison of sex ratios should be between singletons and same-sex twins. The proportion of males among same-sex twins is 0.506 and 0.521 in sub-Saharan Africa and India, respectively, figures that are similar to the proportion of males among singletons in these regions (0.508 and 0.520, respectively). This suggests that male-female relative differences in fetal death are similar for twins and singletons, and so are the prenatal environmental factors that determine the sex of a child.

These figures imply a sex ratio at birth (the ratio of males to females at birth) of 1.032 in sub-Saharan Africa and 1.08 in India for the whole sample of twins and singletons. The figure for sub-Saharan Africa is similar to that found by Garenne (2002) based on Demographic and Health Surveys and World Fertility Surveys. The figure for India is the same as that found both in the 2001 Indian Census, as well as earlier work by Rosenzweig and Schultz (1982) that was based on a nationally representative sample of rural households in India. This figure is also consistent with the sex ratio of 1.09 found by Pakrasi and Halder (1971) using the 1961-62 Indian National Sample Survey. High sex ratios at birth in India have been explained by the selective abortion of female fetuses (Sen (1990b, 1992)).

4.2.2 Socioeconomic variables

In Table 2, we show the summary statistics of common demographic and socioeconomic variables for twins and singletons. In sub-Saharan Africa and India, twins and singletons are similar along several characteristics including marital status, mother's age, mother's education, and father's education. Twins tend to be born to slightly older mothers than singletons in both regions. In India, a slightly higher proportion of twins than singletons are born to mothers or fathers with a secondary or higher level education. With respect to household characteristics, we note that twins obviously live in slightly larger households than singletons, but they do not significantly differ along other household characteristics including wealth (e.g., access to or possession of electricity, radio, TV, car). This comparison shows that twins largely mirror the entire population along several important demographic and socioeconomic variables that matter for child health and survival, as found in other studies (e.g., Almond et al. (2005)).

4.2.3 Mortality rates

Mortality rates are much higher for twins than singletons (Table 2). In sub-Saharan Africa, the probability of dying before the fifth birthday is 163 per thousand for singletons and 405 per thousand for twins. In India, these figures are 115 and 329 per thousand, respectively. Twin-singleton differences in mortality rates decline with age (see also Pongou et al. (2019)). The twin-singleton mortality rate ratio in the neonatal period is close to 5 in sub-Saharan Africa (193 vs. 41 per thousand), and 6 in India (287 vs. 50 per thousand). This ratio falls below 2 in both regions by age 5. While mortality rates are higher for twins than singletons in both India and sub-Saharan Africa, in results not shown here, we find that the two groups are comparable in terms of the relative mortality of boys and girls. For example, Pongou (2013) finds that, in sub-Saharan Africa, the probability of death during the first year of life is 1.166 times higher for boys than for girls among twins versus 1.161 among singletons. The fact that twins and singletons are comparable in terms of the sex ratios at birth and male-female relative difference in mortality and along several socioeconomic and demographic variables minimizes threats to external validity. It is also important to mention that the focus on twins is useful because twins constitute an important and fast growing population.

5 Main results

In this section, we present our main findings. We first present a descriptive analysis of male-female differences in mortality. This is followed by the results of the decomposition analysis.

5.1 Sex differences in mortality: A descriptive analysis

In Table 3, we compare sex differences in mortality in all twins versus male-female twins both in India and sub-Saharan Africa. For all twins, infant mortality is higher among males than among females (323 vs. 277 per thousand in sub-Saharan Africa, and 393 vs. 366 per thousand in India). Therefore, the sex gap in infant mortality rate is 46 per thousand points in sub-Saharan Africa and 27 per thousand points in India. In the sample of male-female twins, the mortality sex gap drops to 27 per thousand points in sub-Saharan Africa, and completely reverses in India, where infant mortality is now 10 per thousand points higher among girls than boys. Given that male-female twins have the same exposure to prenatal factors, the smaller sex gap found in the sample of male-female twins in sub-Saharan Africa shows that these factors have an important effect on male mortality, and the reversed gap in India additionally shows the effect of discrimination against female children.

The results are qualitatively similar for neonotal, postneonatal, and child mortality in sub-Saharan Africa, although the male-female mortality gap becomes smaller with age. India, we note that while female children have a survival advantage in the neonatal period, they die at a higher rate in subsequent periods.

5.2 Decomposing sex differences in mortality into the distinct effects of the prenatal environment, child biology and parental discrimination

In this section, we decompose the male-female difference in the probability of death into the distinct effects of the prenatal environment, child biology and parental discrimination. This analysis is performed for each of the four mortality outcomes. This exercise also makes it possible to analyze the extent to which conventional methodological approaches underestimate the effect of parental discrimination on girls' excess mortality in India.

5.2.1 Infant mortality

We estimate the sex difference in mortality in "all twins" (equation (1)) and in "mixed-sex twins" (the latter is equivalent to estimating a twin fixed effect LPM as shown in equation (5)). The dependent variable is a binary indicator equal 1 if the child died before his/her first birthday, and 0 otherwise. Results are presented in Panel A of Table 4. Columns (I)-(III) show the results for sub-Saharan Africa, and Columns (IV)-(VI) show the results for India. In Columns (I) and (IV), the dependent variable is regressed on a binary variable taking on the value 1 if the child is male, and 0 if the child is female. Confirming the descriptive analysis, the probability of dying before the first birthday is 47 and 27 per thousand points higher among male children than their female counterparts in sub-Saharan Africa and India, respectively. In Columns (II) and (V), we control for child, and parental and household characteristics.[21] This changes little from the estimates obtained in Column (I) and (IV). The male-female difference in infant mortality only decreases to 45 per thousand points in sub-Saharan Africa, but remains the same in India.

In order to net out the effect of the prenatal environment from the estimates obtained

[21]These characteristics include child's year of birth, mother's age at survey, education and marital status, husband's education, household size, possession of household assets and facilities, a linear control for year of survey, and country fixed effect.

in Columns (I)-(II) and (IV)-(V), we estimate a twin fixed effect regression in Column (III) for sub-Saharan Africa and Column (VI) for India. Infant mortality is now only 27 per thousand points higher for boys compared to girls in sub-Saharan Africa, but is 10 per thousand points higher for girls compared to boys in India, although the latter is not statistically significant at the 10% level.

In Table 5, we show the results of the decomposition of the sex difference in infant mortality into the effects of the prenatal environment, child biology and parental discrimination against girls. These estimates are computed based on the point estimates obtained in Columns (II) and (III) for sub-Saharan Africa, and Columns (V) and (VI) for India. From this calculation, we find that the prenatal environment plays a significant role in the differential mortality rates of male and female children. It raises boys' infant mortality rate by 45-27=18 per thousand points in sub-Saharan Africa and by 27 -(-10)=37 per thousand points in India. Biology is also an important factor in the sex gap in mortality, but its role is much less important than conventional methodological approaches would suggest. Males' biology increases their mortality rate by 27 per thousand points in sub-Saharan Africa and by the same amount in India as per our assumption.

Finally, discrimination against female children in India significantly increases their mortality risk. It increases female infant mortality rate by 37 per thousand points. This finding contradicts previous conclusions that discrimination against female *infants* in South and East Asia has a negligible effect on their mortality (e.g., Sen (2001), Osmani and Sen (2003)). As noted earlier, such conclusions are based on the fact that infant mortality is higher among boys than girls, as shown in Figure 1 for India. Our analysis shows that once the effects of the prenatal environment and biological differences between the sexes are controlled for, the effect of discrimination against girls becomes more visible, a finding which is consistent with well-documented evidence on the neglect, abandonment, and infanticide of female newborns in most countries of South and East Asia (Sudha and Rajan (1999)).

We note that the underestimation of the effect of discrimination against girls in India by conventional methodological approaches is substantial. Indeed, when the effects of the prenatal environment and child biology are not accounted for, boys are 27 per thousand points more likely to die before their first birthday than girls in India (Table 4, Column (IV)), which would lead conventional approaches to conclude to the absence of discrimination against female children in the period from birth to 1 year. However, once these effects are accounted for, infant mortality is 37 per thousand points higher in girls than in boys, which is the effect of parental preferences reported in Table 5. These estimates imply that conventional methodological approaches underestimate the effect of discrimination on girls' infant mortality by 173 percent. Importantly, the effect of discrimination is underestimated by 100 percent if only the effect of biology is accounted for, and by 73 percent if only the effect of the prenatal environment is accounted for.[22]

5.2.2 Neonatal mortality

The results of the analysis of sex differences in neonatal mortality are reported in Table 4, Panel B. Columns (I) and (IV) show that neonatal mortality is 37 and 43 per thousand points higher among male children compared to their female counterparts in sub-Saharan Africa and India, respectively. After controlling for child, parental and household characteristics in Columns (II) and (V), the coefficient on the male dummy decreases to 36 per thousand points for sub-Saharan Africa, and increases to 45 per thousand points for India. We estimate a twin fixed effect regression in Columns (III) and (VI). We find that the probability of neonatal death is now only 22 and 9 per thousand points higher among males than females in sub-Saharan Africa and India, respectively, but the estimate for India is not statistically different from zero.

In Table 5, we present the decomposition of sex differences in neonatal mortality. Prenatal environmental factors increase the probability of neonatal death among male children by 14 per thousand in sub-Saharan Africa and 36 per thousand points in India. Biology raises male neonatal mortality by 22 per thousand points, and parental discrimination against female children in India increases their mortality by 13 per thousand points.

Again, conventional methodological approaches grossly underestimate the effect of discrimination on girls' neonatal mortality in India. When the prenatal environment and biological influences are not accounted for, boys are 45 per thousand points more likely than girls to die in their first month after birth (Table 4, Column (IV)), and conventional approaches would therefore conclude to the absence of discrimination against girls during the neonatal period. However, accounting for these effects shows that girls are more likely than boys to die by 13 per thousand points due to parental discrimination. These estimates imply that conventional methodological approaches underestimate the effect of parental discrimination on girls' neonatal mortality by 446 percent, which is substantial.

5.2.3 Postneonatal mortality

The results for the analysis of postneonatal mortality are shown in Table 4, Panel C. Columns (I) and (IV) show that the probability of postneonatal death is 18 per thousand points higher for boys than for girls in sub-Saharan Africa, but is 14 per thousand points lower for boys in India, although the estimate for India is not statistically significant at the 10% level. Adding controls changes little to these estimates (see Columns (II) and (V)). We estimate a twin fixed effect regression in Columns (III) and (VI). Postneonatal

[22] Calculations are based on the row of Table 5 reporting the decomposition of infant mortality for India.

mortality is now only 10 per thousand points higher among males than females in sub-Saharan Africa, and is 32 per thousand points higher among females than males in India.

The results of the decomposition analysis, presented in Table 5, show that the prenatal environment increases the mortality of male children by 8 and 18 per thousand points in sub-Saharan Africa and India, respectively. Biology increases male postneonatal mortality by 10 per thousand points, and parental discrimination against girls in India raises their mortality by 42 per thousand points.

Interestingly, the fact that, in India, postneonatal mortality is now higher for girls compared to boys even when no variable is controlled for would lead one to conclude that female children are discriminated against. Note however that traditional estimation approaches would still underestimate the extent of discrimination. Indeed, when the effects of prenatal factors and child biology are accounted for, girls are more likely than boys to die in the postneonatal period by 42 per thousand points, versus 14 per thousand points when these effects are not accounted for. This means that conventional methodological approaches underestimate the effect of discrimination on excess female postneonatal mortality by 65 percent.

5.2.4 Child mortality

The analysis of male-female differences in child mortality is reported in Table 4, Panel D. Columns (I) and (IV) show that child mortality is 4 per thousand points higher among boys in sub-Saharan Africa, but is 17 per thousand points higher among girls in India. These estimates do not change significantly when controls are added (Columns (II) and (V)). The estimation of a twin fixed effect regression in Columns (III) and (VI) shows that a male twin is less likely to die than his female co-twin by 8 per thousand points in sub-Saharan Africa and by 31 per thousand points in India.

Table 5 presents the results of the decomposition analysis of the male-female difference in child mortality. Prenatal factors raise the probability of death among male children by by 12 per thousand points in sub-Saharan Africa and by 15 per thousand points in India. Contrary to the long-held biological theory of sex differences in morbidity and mortality, biology favors the survival of male children during the childhood period, as it reduces their mortality rate by 8 per thousand points. Parental discrimination against female children in India raises their mortality by 23 per thousand points.

As for the other outcomes, we note that conventional methodological approaches underestimate the extent of discrimination on the mortality of girls between the ages of 1 and 5 in India. When the prenatal environment and biological influences are not accounted for, the mortality of girls exceeds that of boys by 17 per thousand points, versus 23 per thousand points when these effects are accounted for. This implies that conventional approaches underestimate the effect of discrimination on girls' mortality by 23 percent.

In summary, our twin-based estimation approach and decomposition technique show that discrimination against girls in India is much greater than traditional methodological approaches would suggest. The extent of the underestimation of the effect of discrimination on girls' mortality under conventional approaches is greatest in the neonatal period, and it decreases with child age. The prenatal environment and child biology increase male mortality both in India and sub-Saharan Africa. But biology negatively affects male survival only in the first year after birth; it tends to favor male survival between the ages of 1 and 5. The findings imply that the relative importance of the prenatal environment in excess male mortality increases with age. Indeed, the prenatal environment is solely responsible for the survival disadvantage of boys between the ages of 1 and 5 in sub-Saharan Africa.

6 Additional findings and discussions

In this section, we discuss additional findings as well as the implications of zygosity and possible biology-environment interactions for our main conclusions. In general, our results suggest that these latter factors are not significant.

Our twin samples include both identical and fraternal twins. However, like in other twin studies (e.g., Almond et al. (2005), Royer (2009), Oreopoulos et al. (2006)), our data do not allow us to distinguish between both types in our analysis. Our estimates rely on comparing "all twins" with "male-female twins." Among all twins, same-sex twins are often, but not always identical, while male-female twins are always fraternal. Identical twins often have perinatal problems (due, for instance, to sharing the same placenta) that fraternal twins do not have. This often results in lower birth weight in the former. Therefore, our estimates of the effect of the prenatal environment would be generalizable only if we assume that perinatal problems due to monozygosity have similar mortality effects on male-male twin pairs and female-female twin pairs. In results not shown here, we uncover two pieces of evidence that support this assumption. First, as already mentioned, the sex ratios at birth among same-sex twins and singletons are very similar in our data, which proves that the two groups do not differ in terms of the sex difference in fetal mortality. Given that singletons, by nature, do not have any perinatal problems (due to monozygosity) that some identical twins have, this result suggests that if such problems exist, they affect male-male twins and female-female twins to the same extent. Second, some of our findings suggest that in reality, perinatal problems due to monozygosity do not significantly affect mortality after birth. In fact, if such problems existed, then mortality would be substantially greater among same-sex twins than opposite-sex twins in each sex category. We find, on the contrary, that the probability of infant mortality is a bit smaller for girls from female-female twin pairs than girls from opposite-sex twin pairs, although the difference is not statistically significant.

Our findings are consistent with the literature that suggests that zygosity may not be so critical for estimating the effect of the prenatal environment on later life outcomes. The medical literature suggests that adult health outcomes among identical twins are similar to those among fraternal twins (e.g., Christensen et al. (1995), Leslie et al. (1993)). Also, S. E. Black et al. (2007) find that the effects of birth weight on outcomes such as education and earnings are similar for identical and fraternal twins. Although the focus in these studies is not on analyzing sex differences in outcomes, they do however suggest that zygosity may not be so important for estimating the effects of intrauterine environment on later life outcomes.[23]

In other results not shown, we provide some evidence that supports our assumption that the effects of the prenatal environment, biology and discrimination are additive. This assumption follows from several studies that analyze twins (e.g., Evans et al. (2002), Neale and Cardon (1992), Fowler et al. (2008)), but it is seldom tested. We estimate the effect of child sex interacted with several proxies for the environment. The interaction between sex and maternal education has no effect on mortality in sub-Saharan Africa and in India.[24] Similarly, interactions between child sex and climate zones (within Africa) in Pongou (2013) show no effect on mortality. These findings are interesting in their own right and support the additivity assumption.

Finally, we control for birth order as a proxy for the intrauterine environment. Several studies have found twin firstborns to be heavier and to have lower mortality rates than twin second-borns (e.g., Smith et al. (2002), Buekens and Wilcox (1993), Pongou (2013)). Consistent with these studies, we find that twin firstborns have a survival advantage over their cotwins. However, controlling for birth order does not change our main findings.

7 Conclusion

Disparities in early age mortality between males and females constitute an important source of child inequality in the world, and are a likely cause of gender differences in human capital accumulation. While traditional explanations for these disparities only focus on biological differences between the sexes and discrimination against females in certain countries, a recent literature points out the important role of the prenatal environment.

[23] Note that monozygous twins account for only 30% of all twins (MacGillivray (1986)). Of all monozygous twins, only 60-70% share the same placenta, and only 1-2% share the same amniotic sac. Therefore, only around 20% of all twins share the same placenta and less than 1% share the same amniotic sac. This suggests that the magnitude of perinatal problems due to monozigisity may be small in theory. Our analysis suggests that monozygosity has little effect, which is consistent with the aforementioned studies.

[24] Interestingly, education is a proxy for socioeconomic status (SES), and SES has been found to determine the sex ratio (e.g., Almond and Edlund (2007), James (1998), Pongou (2013)). Also, in most developing countries, non-educated mothers are more likely to be employed in the agricultural sector (if indeed they are employed), thereby increasing their exposure to certain fertilizers and chemicals likely to affect the sex ratio of their offspring.

However, the distinct roles of these three sets of factors had not been quantified in a unified framework. In this paper, we proposed a new methodology for decomposing male-female differences in mortality into the distinct effects of the prenatal environment, child biology, and parental preferences. We implement this methodology using large samples of twins from India, a country where parents discriminate against daughters in the allocation of health resources, and sub-Saharan Africa, a region where sons and daughters enjoy symmetrical treatment in the household. Sub-Saharan Africa is generally considered a pertinent counterfactual for India in studies of gender discrimination for several reasons, including the fact that children die from similar diseases in both regions.

Our identification strategy and decomposition technique, which rely on comparing male-female differences in mortality in male-female twins versus all twins, show that: (1) both the prenatal environment and biology increase the mortality risk of boys in India and sub-Saharan Africa; (2) the relative importance of the effect of the prenatal environment increases with child age, while that of biology decreases and even reverses between the ages of 1 and 5; and (3) parental discrimination against girls in India significantly increases their mortality; but failure to control for the effects of the prenatal environment, biological influences, and the endogeneity of sex determination (due to parental factors and sex-selective abortion) leads traditional methodological approaches to underestimate the effect of discrimination on girls' mortality by 173 percent in the first year after birth, and by 23 percent between the ages of 1 and 5.

Taken together, our findings provide novel quantitative evidence on the relative importance of nature versus nurture in male-female differences in life expectancy among children. Importantly, our finding that the effects of discrimination against girls on their mortality in India are important even during the neonatal period contradicts previous studies that claim that such effects are only present in later childhood. That these effects have been understated means that new efforts should be undertaken by governments and policymakers to combat this very crucial problem that unjustly denies millions of girls from the right to life in many countries.

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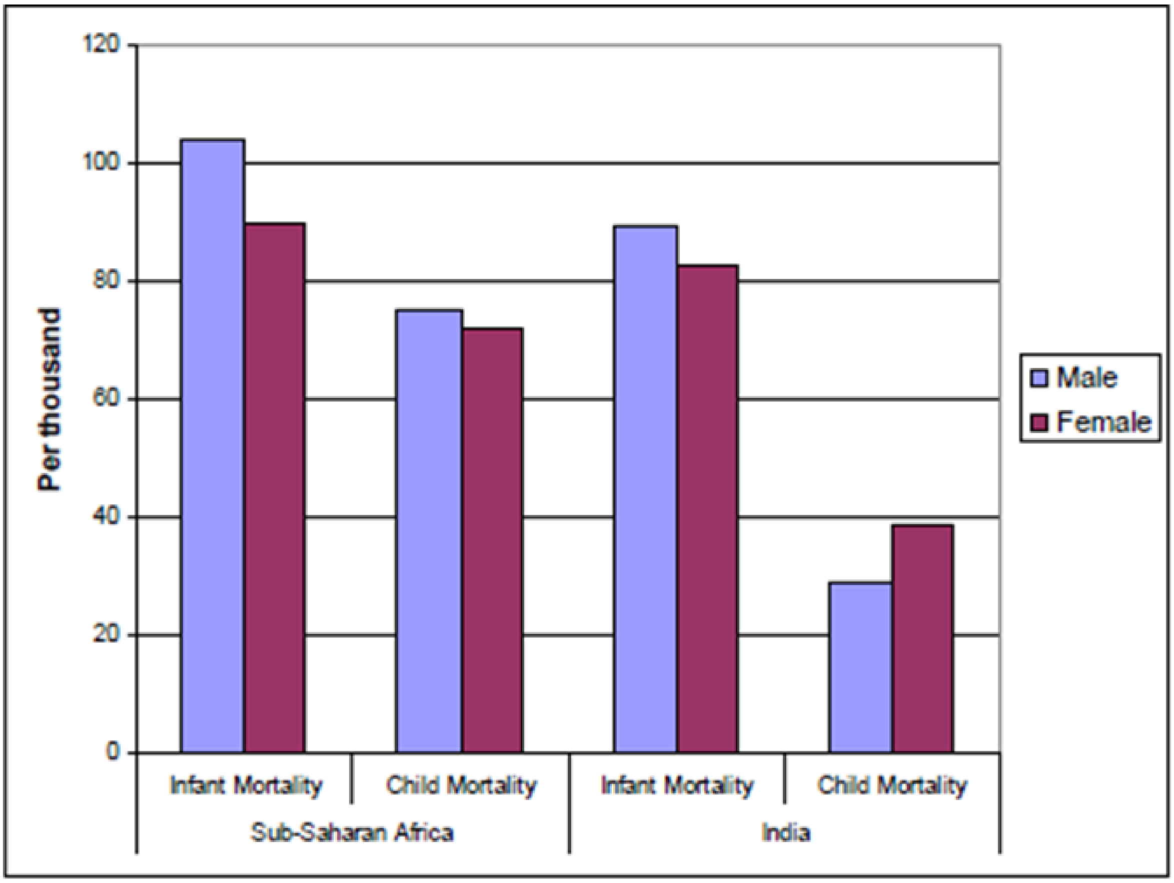
Figure 1: Sex differences in infant and child mortality in sub-Saharan Africa and India



Note: The data are from the Demographic and Health Surveys for sub-Saharan Africa and the National Health Surveys for India. Infant mortality rate is calculated as the probability of dying before the first birthday conditional on being born alive, and child mortality is the probability of dying between the first and fifth birthdays conditional on surviving the first year after birth.

Table 1: Sex ratios at birth of singletons and twins in sub-Saharan Africa and India

	Sub-Saharan Africa		India	
	Sample size	% boys (S.D)	Sample size	% boys (S.D)
Singletons	1,619,483	0.508 (0.500)	537,061	0.520 (0.500)
All twins	50,994	0.504 (0.500)	6,920	0.514 (0.500)
Male-female	20,154	0.500 (0.500)	2,232	0.500 (0.500)
Male-male	15,610	1 (0)	2,442	1 (0)
Female-female	15230	0 (0)	2,246	0 (0)
Same-sex twins	30,840	0.506 (0.500)	4,688	0.521 (0.500)

Table 2: Summary statistics

	Sub-Saharan Africa						India					
	Singletons			Twins			Singletons			Twins		
Variables	*N*	Mean	S.D	*N*	Mean	S.D	*N*	Mean	S.D	*N*	Mean	S.D
Child is male	1,619,483	0.508	0.500	50,994	0.504	0.500	537,061	0.520	0.500	6,920	0.514	0.500
Maternal characteristics												
Age	1,619,483	35.104	8.062	50,994	36.343	7.521	537,061	34.772	8.040	6,920	35.225	7.804
Marital status												
Single	1,619,432	0.022	0.148	50,994	0.015	0.122	537,061	0.000	0.000	6,920	0.000	0.000
Married	1,619,432	0.769	0.422	50,994	0.771	0.420	537,061	0.943	0.232	6,920	0.942	0.233
Widowed	1,619,432	0.047	0.211	50,994	0.050	0.218	537,061	0.047	0.212	6,920	0.047	0.211
Living with a partner	1,619,432	0.097	0.296	50,994	0.096	0.295	537,061	0.000	0.000	6,920	0.000	0.000
Not living with a partner	1,619,432	0.034	0.181	50,994	0.037	0.188	537,061	0.008	0.090	6,920	0.008	0.090
Divorced or separated	1,619,432	0.031	0.173	50,994	0.030	0.172	537,061	0.002	0.042	6,920	0.003	0.051
Education												
Not Educated	1,619,404	0.554	0.497	50,990	0.558	0.497	536,070	0.624	0.484	6,908	0.597	0.491
Primary	1,619,404	0.335	0.472	50,990	0.335	0.472	536,070	0.171	0.376	6,908	0.184	0.387
Secondary or higher	1,619,404	0.111	0.314	50,990	0.107	0.309	536070	0.206	0.404	6,908	0.219	0.414
Father's education												
Not Educated	1,548,881	0.579	0.494	49,038	0.580	0.493	536,465	0.623	0.485	6,906	0.597	0.491
Primary	1,540,365	0.352	0.478	48,576	0.351	0.477	535,280	0.171	0.376	6,888	0.184	0.388
Secondary or higher	1,512,371	0.119	0.323	47,740	0.114	0.318	535,115	0.206	0.404	6,886	0.220	0.414
Household characteristics												
Household size	1,619,483	7.993	4.795	50,994	8.447	4.728	537,061	7.229	3.539	6,920	7.427	3.772
Has electricity (0/1)	1,519,492	0.170	0.376	47,648	0.167	0.373	537,061	0.596	0.491	6,920	0.610	0.488
Has radio (0/1)	1,584,591	0.551	0.497	49,820	0.556	0.497	536,867	0.415	0.493	6,920	0.427	0.495
Has TV (0/1)	1,532,985	0.126	0.332	47,972	0.122	0.327	536,918	0.304	0.460	6,920	0.309	0.462
Has car (0/1)	1,527,477	0.042	0.201	47,950	0.039	0.193	536,897	0.016	0.127	6,920	0.019	0.137
Child outcomes												
Infant mortality (0/1)	1,619,483	0.090	0.287	50,994	0.300	0.458	537,061	0.082	0.275	6,920	0.380	0.485
Neonatal mortality (0/1)	1,619,483	0.041	0.199	50,994	0.193	0.394	537,061	0.050	0.218	6,920	0.287	0.453
Postneonatal mortality (0/1)	1,552,795	0.051	0.220	41,175	0.133	0.340	510,302	0.034	0.181	4,932	0.130	0.337
Child mortality (0/1)	1,473,364	0.073	0.260	35,686	0.105	0.306	492,928	0.033	0.180	4,289	0.049	0.217

Table 3: Mortality rates of boys and girls in different age intervals in sub-Saharan Africa and India

	Sub-Saharan Africa				India			
	Boys		Girls		Boys		Girls	
	Mean	S.D	Mean	S.D	Mean	S.D	Mean	S.D
Infant mortality								
All twins	0.323	0.468	0.277	0.447	0.393	0.489	0.366	0.482
Male-female twins	0.307	0.461	0.280	0.449	0.338	0.473	0.348	0.476
Neonatal mortality								
All twins	0.211	0.408	0.174	0.379	0.308	0.462	0.265	0.441
Male-female twins	0.202	0.401	0.180	0.384	0.260	0.439	0.251	0.434
Postneonatal mortality								
All twins	0.143	0.350	0.124	0.330	0.123	0.329	0.138	0.345
Male-female twins	0.132	0.339	0.122	0.328	0.105	0.307	0.129	0.336
Child mortality								
All twins	0.107	0.309	0.103	0.304	0.041	0.198	0.058	0.234
Male-female twins	0.095	0.293	0.101	0.301	0.031	0.174	0.063	0.243

Table 4: Linear probability model estimates of sex differences in mortality based on twins data from sub-Saharan Africa and India

	Sub-Saharan Africa			India		
Panel A: Infant mortality	(I)	(II)	(III)	(IV)	(V)	(VI)
Male	0.047***	0.045***	0.027***	0.027**	0.027**	-0.010
	[0.004]	[0.004]	[0.005]	[0.012]	[0.011]	[0.016]
# Observations	50,994	50,994	50,994	6,920	6,920	6,920
Panel B: Neonatal mortality	(I)	(II)	(III)	(V)	(VI)	(VII)
Male	0.037***	0.036***	0.022***	0.043***	0.045***	0.009
	[0.003]	[0.003]	[0.004]	[0.011]	[0.011]	[0.013]
# Observations	50,994	50,994	50,994	6,920	6,920	6,920
Panel C: Postneonatal mortality	(I)	(II)	(III)	(V)	(VI)	(VII)
Male	0.018***	0.018***	0.010**	-0.014	-0.014	-0.032**
	[0.003]	[0.003]	[0.004]	[0.010]	[0.010]	[0.015]
# Observations	41,175	41,175	37,958	4,932	4,932	4,324
Panel D: Child mortality	(I)	(II)	(III)	(V)	(VI)	(VII)
Male	0.004	0.004	-0.008*	-0.017***	-0.016**	-0.031***
	[0.003]	[0.003]	[0.005]	[0.007]	[0.007]	[0.011]
# Observations	35,686	35,686	30,176	4,289	4,289	3,418
Twins fixed effect	NO	NO	YES	NO	NO	YES
Controls	NO	YES	NO	NO	YES	NO

Controls include child's year of birth; mother's characteristics: age at survey, education, and marital status; husband's education; household's characteristics: household size, possession of assets such as car, television, radio, and electricity; and a linear control for year of survey, and country fixed effect. Standard errors are in brackets, and are corrected for clustering of observations within mothers.

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 5: Decomposition of sex differences in mortality into the effects of prenatal environment, child biology and parental preferences based on twins data

	Sub-Saharan Africa			India		
	Sex differences in mortality attributable to:			Sex differences in mortality attributable to:		
	Prenatal environment	Child biology	Parental preferences	Prenatal environment	Child biology	Parental preferences
Infant mortality	0.018	0.027	0	0.037	0.027	-0.037
Neonatal mortality	0.014	0.022	0	0.036	0.022	-0.013
Postneonatal mortality	0.008	0.010	0	0.018	0.010	-0.042
Child mortality	0.012	-0.008	0	0.015	-0.008	-0.023

Appendix

Table A-1: Sample size by country

Countries	Years of Survey	Total sample size of live births	Sample size of twins	Sample size of singletons
India	1992/93, 1998/99	543,981	6,920	537,061
Sub-Saharan African Countries				
Benin	1996, 2001	38,703	1,880	36,823
Burkina Faso	1992/93, 1998/99, 2003	84,278	2,520	81,758
Burundi	1987	11,880	198	11,682
Central African Republic	1994/95	16,933	444	16,489
Cameroon	1994, 1998, 2004	56,218	2,116	54,102
Chad	1996/97, 2004	47,175	1,350	45,825
Comoros	1996	7,907	294	7,613
Côte d'Ivoire	1994, 1998/99, 2005	45,779	1,486	44,293
Ethiopia	2000, 2005	84,040	1,740	82,300
Gabon	2000	16,862	532	16,330
Ghana	1988, 1993, 1998, 2003	55,743	1,890	53,853
Guinea	1999, 2005	50,021	1,900	48,121
Kenya	1989, 1993, 1998, 2003	94,460	2,572	91,888
Lesotho	2004	14,699	422	14,277
Liberia	1986	17,261	698	16,563
Madagascar	1992, 1997, 2003/04	61,362	1,282	60,080
Malawi	1992, 1996, 2000, 2004	92,571	3,584	88,987
Mali	1987, 1995/96, 2001	98,535	2,788	95,747
Mozambique	1997, 2003	63,157	2,086	61,071
Namibia	1992, 2000	28,309	684	27,625
Niger	1992, 1998	52,702	1,558	51,144
Nigeria	1990, 1999, 2003	74,387	2,628	71,759
Rwanda	1992, 2000, 2005	77,087	1,702	75,385
Senegal	1986, 1992/93, 1997, 1999, 2005	102,487	2,608	99,879
South Africa	1998	22,905	558	22,347
Sudan	1990	25,793	684	25,109
Tanzania	1992, 1996, 2004	96,491	3,228	93,263
Togo	1988, 1998	37,009	1,532	35,477
Uganda	1988, 1995, 2000/01	62,203	1,618	60,585
Zambia	1992, 1996, 2001/02	70,702	2,334	68,368
Zimbabwe	1988, 1994, 1999, 2005/06	62,818	2,078	60,740