PRENATAL HEALTH INVESTMENT DECISIONS: DOES THE CHILD'S SEX MATTER?*

APARNA LHILA AND KOSALI I. SIMON

Individuals invest in their own health, but children rely on parents to act on their behalf, especially in the case of prenatal health. In this article, we ask, Do parents in the United States who choose to give birth allocate resources differently in the prenatal health of their sons and daughters when the sex of the child is known in advance? We pay special attention to prenatal health behaviors, which can be viewed as investment decisions, of first-generation immigrant parents from India and China, two countries with demonstrated son preference. Ultrasound receipt proxies for knowing fetal gender, enabling us to separate child sex-related biological differences from investment differences in sons' and daughters' health. There is evidence consistent with sex-selective abortions among Indian and Chinese populations, but among parents who choose to carry the pregnancy to term, our findings do not suggest that knowledge of child sex drives prenatal health investments in the United States, neither in the population as a whole nor among Indian and Chinese immigrants.

Individuals combine medical care and other market goods with their own time to invest in their health (Grossman 1972). But for children, parents act as agents who make intra-household resource allocation decisions regarding their health.[1] Parents in the United States have been shown to display son preference in several dimensions, including fertility, marriage, and postnatal investment decisions in child health (e.g., Abrevaya 2005, forthcoming; Almond and Edlund 2008; Dahl and Moretti 2004; Lundberg and Rose 2003; Rosenzweig and Schultz 1982a). In this article, we study the effect of knowing fetal gender on one class of particularly influential health decisions: prenatal health decisions that impact maternal health and the health of the unborn child.

Prior to the prenatal health investment decisions, parents choose to either terminate the pregnancy or carry it to term. Thus, we first examine the possibility of sex-selective abortion, but our main question is about the effect of knowing fetal gender on prenatal health investments *conditional upon having made the abortion decision*. Hence, we try to answer the question, Do parents in the United States who choose to give birth allocate resources differently in the prenatal health of their sons and daughters when the sex of the child is known in advance? We pay special attention to decisions of first-generation immigrant parents who were born in countries with demonstrated son preference.

From a policy perspective, the question raised in this article is an important one. If knowing the child's sex in advance disadvantages some children's health at birth, then a policy that limits access to such information or urges physicians to be more vigilant when conveying this information might be relevant. A precedent has been set in countries like

*Aparna Lhila, Assistant Professor, Department of Public Administration and Policy, University of Georgia, 204 Baldwin Hall, Athens, GA 30602; e-mail: ALHILA@uga.edu. Kosali I. Simon, Associate Professor, Policy Analysis and Management Department, Cornell University, and Faculty Research Fellow, National Bureau of Economic Research. The authors wish to thank Jason Abrevaya, Nan Astone, Marianne Bitler, John Cawley, Arden Handler, George Jakubson, Robert Kaestner, Don Kenkel, Sharon Long, Phil Levine, Phil Morgan, Sean Nicholson, two anonymous referees, seminar participants at Cornell University (Theoretical Works in Progress seminar), Cornell Weil School of Medicine, NY, the University of Georgia's Department of Public Administration and Policy, participants at the 2006 Midwest Economics Association, Population Association of America, and American Society of Health Economists meetings for their comments. We also thank Deborah Rosenberg for providing the SAS code to construct the Kotelchuck Index. The authors are solely responsible for any errors or opinions implied in this article.

1. Currie (2004) provided a review of the economic motivation for studying factors that affect child health and examples of recent research.

India, where knowing fetal gender has led to a skewed population sex distribution; stricter laws govern the use of ultrasounds. It may be especially policy-relevant to test for gender preference among immigrant mothers because they give birth to approximately 20% of U.S. children born every year.[2]

Gender-biased investments in child health among parents in South and Southeast Asia are well documented (e.g., Sen 1992, 2003). American parents have preferences about the sex of their children, which affects their fertility-stopping rule (Angrist and Evans 1998) and their marriage and divorce decisions (Dahl and Moretti 2004).[3] These decisions, like prenatal health investment decisions, are inherently resource allocation decisions that are likely to affect the well-being of girls and boys differently.

Before proceeding with our analysis, we consider the possibility of sex-selective abortion. If this exists among immigrants in the United States, then is there any reason to expect different parental behaviors based on child gender among pregnancies brought to term? We believe so for two reasons. First, assuming a continuum of son preference or "distaste" for daughters,[4] and a continuum of abortion costs that includes time and psychological costs, there may remain some parents for whom the cost of aborting a daughter outweighs the distaste for a daughter. In such cases, parents will choose not to terminate the pregnancy, but there may be ways in which their son preference can manifest itself. Second, since prior work has found that postnatal gender discrimination in investments exists, it is logical to test whether gender-biased investments extend to the womb.

No U.S. or international research, to our knowledge, has studied whether parents are guided by gender preference in their prenatal health investment decisions. In this article, we test whether parents/mothers[5] who are likely to know fetal gender as a result of having an ultrasound invest differentially in the health of boys versus girls by actions such as avoiding prenatal smoking and drinking, prenatal weight gain, and the frequency and adequacy of prenatal care use.

We choose to focus on mothers' prenatal investment decisions because they are potentially modifiable, have utility consequences for mothers (e.g., the time and money costs of prenatal health care visits, the social and psychological costs of tobacco and alcohol abstention, or difficulties with attaining weight gain goals), and are important in influencing child health at birth. These costs and benefits give rise to demand functions for child health inputs that will vary across women, perhaps systematically related to whether the fetus she carries is a boy or a girl. If parents favor sons, we expect to find that after controlling for any biological effects (like morning sickness) caused by carrying a male versus female fetus, knowing that the unborn child is female may, at the margin, negatively impact the demand for prenatal care and infant health outcomes. This effect may differ according to socioeconomic and cultural identity; it may be particularly pronounced among first-generation immigrant mothers who were born in countries with a history of son preference, such as India and China.

This article contributes to the literature in three ways. First, previous studies have focused on parental *post*natal investments in child well-being, whereas we study gender preference with regard to *pre*natal health investments. Second, we add to the relatively small body of literature on gender preference in the United States. Finally, we study the persistence of gender preference among immigrant mothers—a question that, to our knowledge,

2. Authors' calculations from Natality Detail data 1989–2001.

3. Other studies, such as that by Pollard and Morgan (2002), have suggested that this phenomenon may be lessening over time. Neither that study nor Dahl and Moretti's (2004) considered immigrants separately.

4. The use of the phrase "distaste for daughter," though unsavory, serves to describe parents' lack of preference for a daughter, that is, their preference for sons.

5. We will use the term *mother* in the rest of the text, although we will return to discuss mothers' versus fathers' roles in the decisions we study.

has received very little attention despite the preponderance of evidence on son preference in mothers' country of birth.[6]

We next discuss the previous literature relating to gender preference to place our contribution in context. We then present the theoretical framework underlying our analysis, along with a description of the various measures of prenatal investments we examine. The third section describes our method, introduces the data, and discusses limitations. We then present results and conclusions.

PRIOR LITERATURE

Parents' investments in the prenatal health of their children are preceded by parents' decision to terminate unwanted pregnancies. When abortion costs are low, parents choose to terminate unwanted pregnancies that would have otherwise resulted in babies born in the lower tail of the prenatal health investment and birth outcome distribution (Grossman and Joyce 1990; Joyce and Grossman 1990; Joyce, Kaestner, and Korenman 2000). This article addresses a particular form of wantedness, that is, when parents do not desire a child of that sex. Mothers who are unhappy with the pregnancy and whose pregnancies are unwanted are likely to initiate prenatal care later, are less likely to quit smoking (Weller, Eberstein, and Bailey 1987), and are more likely to give birth to low birth weight (LBW) children (Sable et al. 1997; Sable and Wilkinson 2000). Thus, not all pregnancies carried to term display medically optimal prenatal behavior in the United States or elsewhere.

In developing countries, gender preference generally takes the form of son preference. Female fetuses are less likely to be carried to term, and daughters who are born are likely to be in poorer health and face a higher risk of childhood mortality compared with sons. Sex ratios (boy/girl) at birth for most societies lie between 1.03 and 1.06, and ratios of 1.07 and 1.09 are attributed to sex-selective abortions in India (Arnold, Kishor, and Roy 2002) and China (Coale and Banister 1994), respectively. The higher rate of female childhood mortality has also been attributed to parents choosing to invest resources such as food, nutrients, and medical care in sons (Bardhan 1982; D'Souza and Chen 1980; Kynch and Sen 1983; B. Miller 1981; Rosenzweig and Schultz 1982a).

Whereas the impact of gender-biased investments in developing countries has been studied extensively, studies of gender preference in the United States are relatively sparse. In terms of children's education in the United States, Behrman, Pollak, and Taubman (1986) found that parents exhibit equal concern or slightly favor girls. However, subjective well-being (Kohler, Behrman, and Skytthe 2005), marital stability (Mammen 2003; Morgan, Lye, and Condran 1988; Teachman and Schollaert 1989), and expenditure on housing (Lundberg and Rose 2002, 2004) are likely to be higher among parents who have a son than among those with a daughter. Dahl and Moretti (2004) found that parents' fertility, marriage, and divorce decisions are consistent with son preference. Similar to Lundberg and Rose (2003), they also found evidence that suggests that unmarried parents are more likely to marry prior to the birth if they know in advance that the baby is male.

A question this literature brings up is the source of gender bias. In the United States, fathers have largely been implicated as the source of son preference. Fathers prefer to invest in the health of their sons, while mothers have a greater impact on the health of their daughters (Thomas 1994). Fathers spend more time with their sons (Bryant and Zick 1996; Yeung et al. 2001), are more likely to be involved in the caretaking of a son than a daughter (Lundberg, McLanahan, and Rose 2007), and spend more time with their children overall when they have a son (Barnett and Baruch 1987; Harris and Morgan 1991). Stated preferences also suggest that fathers favor boys (Dahl and Moretti 2004).

6. Two recent exceptions include Abrevaya (2005) and Almond and Edlund (2008), works of which we became aware after conducting our analysis.

Disentangling the role of mothers and fathers in determining prenatal decisions is beyond the scope of this article, but this evidence suggests that mothers might receive less emotional support, experience greater stress, or be subject to domestic violence in extreme cases, which might lead to different behaviors when expecting a son versus a daughter. We take one step toward examining the role of fathers' preferences by conducting separate analyses for married and unmarried mothers. While conducting separate analyses for married and unmarried mothers does not perfectly capture the role of mothers' and fathers' preferences in determining prenatal behavior, it is plausible that spouses exert greater influence on prenatal decisions than unmarried partners. Thus, any prenatal investment differences observed between these two subsets of parents may be attributed, at least partly, to fathers.

The research evidence so far on gender preference in parental behavior suggests the plausibility that knowing the gender of the child in advance may alter mothers' behavior during pregnancy. Following the literature, we consider the health behaviors of mothers (e.g., use of medical care, smoking decisions) as investments in the production of child health (e.g., Rosenzweig and Schultz 1982b, 1983). We focus on behaviors that are known to impact health in utero and at birth. Medically, maternal nutrition and lifestyle and the fetus's exposure to restricted nutrient intake and smoking are likely to cause LBW and have long-lasting health effects, such as such as hypertension, stroke, and type 2 diabetes (Barker 1997; Maritz, Morley, and Harding 2005). Improving mothers' prenatal care use along with altering mothers' detrimental prenatal health habits (such as tobacco, alcohol, and drug use) and improving maternal health (including achieving adequate weight gain) are likely to increase birth weight (Boss and Timbrook 2001; Evans and Ringel 1999; Shiono and Behrman 1995; Warner 1995). Smoking and alcohol use during pregnancy are also associated with low APGAR scores (Haddow et al. 1988; Okah, Cai, and Hoff 2005; Streissguth et al. 1981).

In addition to prenatal health behaviors, we study the impact of knowing fetal gender on birth outcomes because this may capture impacts on child health through avenues on which we do not have data (e.g., stress, domestic violence, secondhand exposure to smoke). Birth weight is an important outcome to study because it has long-term health consequences, such as stunting and underweight, and LBW has been found to lower educational attainment and earnings into adulthood (Behrman and Rosenzweig 2004; Behrman, Rosenzweig, and Taubman 1994; Currie and Hyson 1999; Osmani and Sen 2003).

Although APGAR scores are not significant predictors of long-term health, they are indicative of the prenatal and perinatal experiences of the infant and predict infant mortality. A score below 7 is indicative of problems experienced during labor or delivery, and a score below 4 requires physicians to take immediate steps to stabilize the infant. Higher levels of maternal anxiety and depression during pregnancy have been linked to lower APGAR scores at both the first and fifth minute after birth (Berle et al. 2005); since anxiety and depression may accompany unwanted pregnancies, we study APAGAR scores as one of the outcomes.

THEORETICAL FRAMEWORK

We assume that parents are concerned with and derive utility from the welfare of their children, including their health (Becker 1981). But the marginal utility costs and benefits of investing in the health of the sons may be different when sons and daughters enter parents' utility functions in different ways. Consider a utility maximization problem in which parents choose the optimal prenatal investment in the health of a child:

$$\text{Max}[U\{X,G,H(I) \mid e,c\}]$$

$$s.t. W = P_X X + P_I I, \tag{1}$$

where I denotes prenatal investments, X stands for all other market goods that provide utility to parents, G represents child gender, and e and c represent economic and cultural conditions, respectively, that determine whether a boy or girl would provide more utility to parents. H denotes child health at birth and may enter the utility function interacted with the child's gender. Parents maximize their utility with respect to prenatal investments (I) and are subject to the budget constraint, where W represents full income, P_X is the price of market goods (X), and P_I is the pecuniary and nonpecuniary cost of prenatal investment.

Dahl and Moretti (2004) reviewed reasons why gender preference may exist in the United States. Internationally, parents may prefer sons because the costs associated with raising daughters (due to c and e in the utility maximization problem) are higher in countries that practice the custom of transferring a dowry from the bride's parents to the groom's parents (Bloch and Rao 2002; Gangadharan and Maitra 2003). Additionally, while daughters may leave the parents' home upon marriage, sons bring wives into the household and provide old-age economic security to parents. This is especially important in countries where the government makes no provision for retirement income, and is correspondingly less important in a U.S. setting. Parents may choose to invest in sons because they expect greater returns on human capital investment in sons than in daughters. There are also important cultural differences; in some countries, a son is desirable because only a son can perform the religious rites upon a parent's death. The value of a first-born son also increases if policy restricts the number of children per family—an important possible effect of China's one-child policy. Parents may prefer to have daughters for equally plausible reasons. For instance, daughters are more likely to provide emotional support and old-age assistance to elderly parents (Mellor 2001). A daughter may also be favored in countries that reverse the dowry system and use a "bride price" instead. Finally, parents may prefer to have a balanced sex composition when traditional gender roles are replaced by shared roles and when girls and boys are substitutable (Pollard and Morgan 2002). A combination of these factors may drive parents to invest differently in the health of their sons and daughters in this simple utility maximization framework.

If parents were to perfectly plan the gender of their children and abort pregnancies that result in children of unwanted sex, there would be no reason to expect differential prenatal investment between girls and boys who are born because those girls would be ones who provide the greater return to the parents. To the extent that any selective abortion takes place, we expect narrower prenatal health investment differences relative to a context in which abortions are not possible. However, as long as the costs of abortion are not zero, this remains an important empirical question. Furthermore, given that a substantial fraction of women do not adhere to ideal prenatal care routines[7] despite a preponderance of information on the health consequences of such routines, we consider this to be evidence that prenatal health investments involve nontrivial costs.

The degree to which we should expect son preference among immigrants from countries with demonstrated son preference is unclear for several reasons. First, there may be selective migration, and immigrant mothers may not be predisposed to son preference even in their home country. Alternatively, the new environment they enter may not provide conditions that drive son preference—that is, resource constraints, policy constraints (such as the one-child policy), or assimilation into the new environment may change c and e in Eq. (1). For instance, Malays of Indian and Chinese descent do not display son preference in their fertility decisions (Lhila 2005).

7. Based on the authors' calculations, 14.7% of mothers report using tobacco during pregnancy, 19.6% initiate prenatal care in the second trimester or later, 28.3% of pregnancies receive a score of inadequate/intermediate on the Kotelchuck index (described in Kotelchuck [1994] and also later in this article), and 58.5% of mothers fail to attain their prenatal weight gain goals.

Prenatal investment differences can only occur when parents know the sex of the baby in advance. When parents favor boys, knowing that fetal gender is female could induce parents to invest differently in the pregnancy or cause depression or anxiety, which may lead to fewer prenatal care visits, inadequate weight gain, or alcohol or tobacco use during pregnancy, all of which adversely affect the child's weight and APGAR scores at birth. We present estimates from reduced-form equations of health production and input demand functions. Additionally, we use the Kessner and Kotelchuck indices to address the adequacy of prenatal care utilized by the mother. What follows is a brief description of what constitutes medically satisfactory inputs into the infant health production function.

The number of prenatal care visits is one way of characterizing the continuity of care received by the mother. The American College of Obstetricians and Gynecologists (ACOG) recommends a schedule of doctor visits that is applicable to most normal pregnancies. The overall adequacy of prenatal care use is rated as adequate, intermediate, or inadequate by the Kessner Index based on the number of prenatal care visits and trimester of prenatal care initiation. Kotelchuck (1994) has suggested an alternate measure that adjusts for length of gestation and rates adequacy into four categories: inadequate, intermediate, adequate, and adequate plus. The ACOG generally recommends that a woman of normal weight gain 25–30 pounds, overweight women gain 15–20 pounds, and underweight women gain 28–40 pounds during pregnancy. As discussed earlier, each of these inputs, along with nutrient intake and maternal smoking and alcohol use during pregnancy, impacts the health status of the child in utero and at birth.

METHOD

Four prenatal tests—obstetric ultrasound imaging, amniocentesis, chorionic villus sampling (CVS), and percutaneous umbilical blood sampling (PUBS)—used for diagnosing fetal health could, as a by-product, also determine fetal gender. Ultrasounds,[8] the most pervasive method, are performed on 68% of mothers, typically in the 18th or 20th week of the pregnancy, and can reveal the sex of the child with 95%–100% accuracy by the 16th week. Amniocentesis[9] is 100% accurate but carries a risk of miscarriage (0.5%) and has been linked to an increased risk of developing health problems for the baby. CVS and PUBS[10] carry the greatest risk of miscarriage and have lower usage rates.

After ascertaining fetal gender and before making prenatal investment decisions, potential parents make the abortion decision. Thus, we begin our analysis by examining whether the practice of sex-selective abortion exists among immigrants in the United States. Because we do not observe pregnancies that are terminated, we compare the U.S. sex ratio (number of boy births per girl birth) to the sex ratios of Chinese and Indian immigrants. Sex ratios for most societies lie between 1.03 and 1.06, and Table 1 shows that the sex ratio of Indian and Chinese immigrant mothers are 1.07 and 1.08, respectively. As revealed by t tests, immigrant mothers are significantly more likely to have sons than daughters relative to all U.S. mothers. Significant differences in sex ratios persist until the fourth parity for Chinese immigrants; among Indian mothers, sex ratios differ from American mothers at second, third, fourth, and sixth or greater parities. The results reported in Table 1 are consistent with findings reported by Abrevaya (2005, forthcoming) and Almond and Edlund (2008).

These results suggest that girls are underrepresented in our data and point to a selection process that determines the pregnancies that we observe in our data. Since we are unable

8. In an ultrasound test, sound waves are used to view the anatomy and internal organs of the fetus and can determine gestational age, identify a multiple pregnancy, monitor fetal growth, and check for birth defects.

9. Amniocentesis removes a small amount of amniotic fluid to test for genetic abnormalities and fetal health.

10. CVS relies on extracting a sample of the placenta, and PUBS withdraws blood from the baby to test for genetic problems or abnormalities.

Table 1. Comparing Sex Ratios: Nonimmigrant Versus First-Generation Immigrant Populations, by Birth Order

	All Births[a] (1)	Indian Immigrant Births (2)	Difference (1) – (2)	Chinese Immigrant Births (3)	Difference (1) – (3)
Sample Size	40,994,488	155,217		308,846	
All Births	1.05	1.07	–0.02**	1.09	–0.04**
Birth Order = 1	1.05	1.04	0.01	1.08	–0.02**
Birth Order = 2	1.05	1.07	–0.02*	1.08	–0.03**
Birth Order = 3	1.04	1.17	–0.13**	1.13	–0.09**
Birth Order = 4	1.04	1.14	–0.10**	1.19	–0.15**
Birth Order = 5	1.04	1.11	–0.07	1.14	–0.10†
Birth Order ≥ 6	1.03	1.26	–0.23*	1.03	0.01

[a]Births to all nonimmigrant mothers in the United States.

†$p < .10$; *$p < .05$; **$p < .01$

to model the selection process in the absence of data on terminated pregnancies, estimates of the effect of knowing fetal gender on prenatal investments will be biased downward. Thus, the results of our analysis should be interpreted as the partial effect of knowing fetal gender on prenatal health investments—that is, the effect of knowing the gender of the unborn child on prenatal investments, conditional upon making the decision to carry the pregnancy to term.

If there were no biology-related gender differences in outcomes, one way to proceed would be to compare the outcomes of parents who had girls with those of parents who had boys. But biological differences may exist between boy and girl pregnancies, so we separate parents by ultrasound receipt. Gender difference in outcomes among non-ultrasound mothers is attributed to biology; among ultrasound mothers, any gender difference in outcomes that exists over and above the biological difference is interpreted as an investment difference. Mechanically speaking, our method of estimation is to regress each of the outcomes of interest—birth weight, APGAR scores, number of prenatal visits, weight gain, alcohol and tobacco use during pregnancy, and the adequacy of prenatal care based on the Kessner and Kotelchuck indices—on FEMALE, an indicator of child gender; ULTRAS, an indicator of ultrasound receipt; and the interaction of the two. Our key parameter is the coefficient on the interaction term. Evidence in support of girl preferences that translate into differential prenatal investment would be indicated by a positive value on this coefficient. One outcome that we know cannot be affected by gender preference is timing of the first prenatal visit because mothers can't ascertain gender before this first prenatal visit. Thus, we include timing of first prenatal visit as a dependent variable that serves as a specification check. We expect to find no significant relationship between our interaction term and the probability of a first-trimester prenatal visit.

The identification of the effect of knowing fetal gender rests on the assumption that ultrasound receipt is not correlated with unobserved factors, specifically gender preference. There are observable socioeconomic differences between mothers who receive ultrasounds and those who do not (Martin et al. 2002), and our analysis controls for these observable characteristics, such as mothers' age, race, ethnicity, education, and presence of medical risk factors. However, if unobservable characteristics differ, especially if gender preference drives parents to seek ultrasounds, then our estimates may potentially be biased toward finding an effect of son preference.[11]

Our estimation method also assumes that all mothers who receive ultrasounds know fetal gender. However, not all mothers who receive an ultrasound find out the sex of the child. Walker and Conner (1993) found that 80% of mothers in the United States wanted to know the gender of their child. This indicates that there is measurement error due to the misclassification of some women into the treatment group when they actually belong in the control group, which is likely to attenuate the effect of knowing fetal gender on health investments toward zero. Moreover, the extent of measurement error is likely to be greater for mothers who are less likely to want to know the sex of the child. Shipp et al. (2004) found socioeconomic and demographic differences between parents who want to know the sex of the child and those who do not. Mothers and fathers were equally likely (58%) to want to know the sex of the child; however, households with fathers without a full-time job, lower household income, unwed mothers, mothers younger than 22 or older than 40, without a college degree, nonwhite and non-Catholic, and those who had preferences regarding the sex of the child, were more likely to want to know the sex of the child (Shipp et al. 2004). As a robustness check, we repeat our analysis for various subgroups defined by mothers' race (white, black, other), marital status (married, unmarried), and mother's age at the time of the child's birth (younger than 23, 23–29, older than 39), and return to this point when interpreting our results.

In order to test the validity of our method, we execute an additional specification test. Knowing fetal gender should have no impact on the mothers' *post*natal decisions, such as smoking, because all parents know the gender of their child postnatally. Findings to the contrary may cast doubt on our method.

After we present results from our main models applied to the general population, we stratify our analysis in a number of ways to investigate how knowledge of fetal gender impacts prenatal health investments across economic and cultural subgroups as well as by family composition. We test whether son preference persists among first-generation immigrants, and distinguish between first and higher-order births, because gender preference may vary with birth order and sibling composition in the household.

DATA

The primary data for this analysis are the 1989–2001 Natality Detail Files, which contain the universe of live births in the United States between 1989 and 2001. These files are a compilation of birth certificates that provide information on birth outcomes, parental demographics, medical risk factors associated with the pregnancy, prenatal care utilization, and congenital abnormalities. This analysis uses information on only those pregnancies that resulted in singleton births, since the prenatal investments in twins are likely to differ sharply from investments in singletons. Approximately 4 million infants are born in the United States every year, which leads to a sample of over 46 million observations for our analysis. The immigrant analysis uses the universe of mothers who report being foreign born and identify their race as Chinese or Indian. The Chinese and Indian[12] immigrant samples contain 304,530 and 154,492 live births, respectively. Means for variables of interest in each of the samples are presented in Table 2.

11. Another possibility is that boy pregnancies may be less robust, leading to a greater probability of receiving an ultrasound and thus a higher likelihood of gender knowledge among boy pregnancies than girl pregnancies. If we compare all pregnancies that ultimately result in a boy versus a girl, we may find less evidence of distaste for girls than would be present if all pregnancies knew gender prenatally. Since we compare boy/girl pregnancies with ultrasounds, relative to boy/girl pregnancies without ultrasounds, our results are not likely to be subject to this type of bias.

12. Asian Indian was included as a distinct race category in birth certificates beginning in 1992. Thus the Indian immigrant sample spans 1992–2001.

The ultrasound information in the Natality Files does not differentiate between ultrasounds performed during pregnancy versus those performed during labor. To consider whether this affects our results, we draw on the 2001 California Birth Statistical Master File (obtained from the California Department of Health Services), which asks whether an ultrasound was one of the procedures performed during labor and delivery. Typically, an ultrasound is performed during delivery if it is a breech birth. We rerun our main specifications on California data using an indicator of whether the ultrasounds occurred during delivery, a point at which it should not influence prenatal investments.

Although the Natality Detail Files are a rich source of information about prenatal care utilization, provide large sample sizes to allow intricate subgroup analyses, and lend a great deal of statistical power to our analysis, they provide terse information about some birth outcomes. For instance, mothers report tobacco use during pregnancy, but this question makes no distinction between tobacco use before versus after pregnancy was confirmed. Alcohol consumption is dealt with in a similar manner and is generally considered highly underreported. Furthermore, judging the adequacy of weight gain depends on mothers' prepregnancy weight, which is unavailable in the Natality Files. To address these limitations, we augment our gender preference analysis with data from the 1988 National Maternal and Infant Health Survey (NMIHS). The NMIHS is a survey of mothers who had a pregnancy in 1988 and is designed to study the factors associated with poor pregnancy outcomes. The NMIHS yields a sample of 9,953 live births, and although it is inadequate to study immigrants, NMIHS overcomes the data limitations described above and provides information on household composition, income, and source of payment for prenatal care visits. These may be important controls because gender preference with respect to the current pregnancy may depend on the gender composition in the household, and income and health insurance coverage are two factors that are likely to impact demand for prenatal care. Finally, the NMIHS asks mothers of their decision to smoke and consume alcohol at the time of the survey. Information on mothers' decision to smoke after the birth of the child is used to conduct the second specification test. Thus, data from the Natality Files, CA Statistical Master File, and NMIHS together provide a clearer picture of prenatal investments and birth outcomes in the United States, conditional upon parents' decision to carry the pregnancy to term.

RESULTS

We begin by presenting the unconditional difference-in-differences results in Table 2 for the full sample, and for the Indian and Chinese immigrant samples separately. In each sample, boys tend to weigh more than girls at birth. For example, in the U.S. sample, the boy-girl difference in birth weight tends to be 117.70 g for mothers who received an ultrasound and 114.28 g for mothers who did not receive an ultrasound. This is because boys naturally weigh more than girls at birth. Subtracting this natural difference from the boy-girl difference in birth weight among mothers who received an ultrasound reveals that knowing child's sex is associated with a 3.42 g difference in the birth weight of boys and girls. This difference-in-differences statistic shown in the last column of Table 2 demonstrates that although there is some evidence of gender-biased investments that favor boys, the difference is negligible in substantive terms. The results in this column show that among parents who choose to carry their pregnancy to term, knowing the sex of the child in advance is often statistically significantly correlated with gender differences in health investments and outcomes, but the magnitude of the correlation is always minuscule. Although these results are from simple correlations, the remainder of this section reveals that this relationship holds qualitatively for all samples, irrespective of educational attainment, race, marital status, and age, and even after adding relevant controls to the model. This is reassuring since gender is relatively randomly assigned and the effect of gender in the regression is not expected to be subject to omitted variables bias in a model without controls.

Table 2. Unconditional Means for All U.S. Mothers and First-Generation Indian and Chinese Immigrant Mothers: 1989–2001 Natality Detail Files

	Ultrasound			No Ultrasound			Difference-in-Differences
	Girl	Boy	Difference	Girl	Boy	Difference	
All[a]	48,792,749						
Sample size	14,504,968	15,232,693		9,316,750	9,738,338		
Prenatal care initiated in first trimester	82.44	82.22	0.22**	77.62	77.46	0.17**	0.06*
Birth weight (g)	3,270.30	3,388.00	–117.70**	3,269.65	3,383.93	–114.28**	–3.42**
Probability of low birth weight (< 2,500 g)	8.03	6.90	1.13**	7.55	6.50	1.05**	0.08**
Probability 1 min. APGAR score < 7	8.19	9.09	–0.90**	7.91	8.80	–0.90**	–0.01
Probability 1 min. APGAR score < 4	2.06	2.45	–0.39**	2.01	2.39	–0.38**	–0.01
Probability 5 min. APGAR score < 7	1.29	1.53	–0.24**	1.33	1.57	–0.24**	0.00
Probability 5 min. APGAR score < 4	0.38	0.44	–0.06**	0.44	0.50	–0.06**	0.001
Number of prenatal care visits	11.65	11.61	0.04**	11.00	10.97	0.03**	0.01**
Prenatal weight gain (lb.)	30.52	31.22	–0.70**	29.96	30.60	–0.64**	–0.06**
Probability of inadequate weight gain (< 15 lb. or > 40 lb.)	38.46	39.16	–0.69**	47.55	48.06	–0.51**	–0.19**
Probability of prenatal tobacco use	15.30	15.32	–0.03†	13.46	13.47	–0.01	–0.02
Probability of inadequate/intermediate prenatal care (Kessner Index)	24.27	24.59	–0.32**	30.21	30.45	–0.25**	–0.07**
Probability of inadequate/intermediate prenatal care (Kotelchuck Index)	25.89	25.48	0.42**	32.46	32.00	0.46**	–0.04
Indian Immigrants	154,492						
Sample size	46,845	49,709		27,871	30,067		
Prenatal care initiated in first trimester	83.49	82.73	0.76**	82.03	81.26	0.77*	–0.01
Birth weight (g)	3,104.11	3,189.74	–85.63**	3,104.98	3,191.34	–86.36**	0.73
Probability of low birth weight (< 2,500 g)	10.05	9.11	0.94**	9.28	8.48	0.81**	0.14
Probability 1 min. APGAR score < 7	5.01	5.89	–0.88	4.05	4.38	–0.33	–0.55
Probability 1 min. APGAR score < 4	1.23	1.37	–0.14	1.13	0.83	0.30	–0.44
Probability 5 min. APGAR score < 7	0.72	0.84	–0.12†	0.72	0.85	–0.12	–0.001
Probability 5 min. APGAR score < 4	0.22	0.23	–0.01	0.23	0.24	–0.01	–0.0002

Number of prenatal care visits	11.26	11.19	0.07**	11.09	11.01	0.08*	–0.02
Prenatal weight gain (lb.)	28.52	28.76	–0.24**	27.89	28.29	–0.40**	0.16
Probability of inadequate weight gain (< 15 lb. or > 40 lb.)	46.29	46.88	–0.59†	54.96	55.30	–0.34	–0.25
Probability of prenatal tobacco use	0.31	0.26	0.04	0.25	0.28	–0.03	0.08
Probability of inadequate/intermediate prenatal care (Kessner Index)	25.57	26.48	–0.91**	26.35	27.45	–1.10**	0.18
Probability of inadequate/intermediate prenatal care (Kotelchuck Index)	27.54	27.28	0.26	29.10	28.99	0.11	0.15
Chinese Immigrants	304,530						
Sample size	83,647	91,025		62,421	67,437		
Prenatal care initiated in first trimester	84.90	84.38	0.52**	86.66	86.12	0.55**	–0.02
Birth weight (g)	3,243.65	3,349.49	–105.84**	3,249.16	3,349.19	–100.03**	–5.81
Probability of low birth weight (< 2,500 g)	5.49	4.50	0.99**	5.00	4.31	0.70**	0.29†
Probability 1 min. APGAR score < 7	4.51	5.09	–0.58*	4.27	4.87	–0.60*	0.01
Probability 1 min. APGAR score < 4	1.03	1.36	–0.34*	1.25	1.10	0.14	–0.48*
Probability 5 min. APGAR score < 7	0.63	0.76	–0.13*	0.72	0.67	0.04	–0.17†
Probability 5 min. APGAR score < 4	0.20	0.20	0.00	0.30	0.19	0.11*	–0.11*
Number of prenatal care visits	11.48	11.43	0.04**	11.49	11.45	0.04*	0.00
Prenatal weight gain (lb.)	29.51	29.78	–0.27**	29.23	29.51	–0.28**	0.01
Probability of inadequate weight gain (< 15 lb. or > 40 lb.)	49.46	49.38	0.07	66.35	66.61	–0.26	0.33
Probability of prenatal tobacco use	0.55	0.54	0.01	0.71	0.73	–0.02	0.03
Probability of inadequate prenatal care (Kessner Index)	23.39	23.98	–0.59**	20.81	21.28	–0.47*	–0.13
Probability of inadequate prenatal care (Kotelchuck Index)	26.31	25.56	0.75**	25.21	24.55	0.66**	0.09

Note: Probabilities have been multiplied by 100.

[a]All observations in the 1989–2001 Natality Detail Files.

†$p < .10$; *$p < .05$; **$p < .01$

The results of our regression analysis are presented in Table 3, which shows the effect of receiving an ultrasound and having a daughter (ULTRASOUND × FEMALE) on prenatal investments and birth outcomes. This is akin to the difference-in-differences result in the last column of Table 2, except that regressions include relevant controls. Regression coefficients on the interaction term derived from ordinary least squares estimation are reported whenever the outcome is continuous. Interaction effects from probit models are reported when the outcome is binary and are calculated according to Ai and Norton (2003).[13] The regressions control for maternal age, educationalå attainment, race and ethnicity, marital status, geographic region and urban/rural residence, mother's foreign-born status, and presence of medical risk factors, and are interpreted as the causal estimates of the effect of knowing fetal gender is female on prenatal health investments and birth outcomes. In Table 3, column 1 indicates the effect for the full sample, and the remaining columns present results separately by birth order and mothers' marital status. The coefficients on the interaction term are close to zero and are statistically insignificant in most cases, suggesting that knowing that fetal gender is female is not a statistically significant determinant of prenatal care use, probability of tobacco use, number of cigarettes smoked during pregnancy, or the adequacy of prenatal care use. In some instances, coefficients are statistically significantly different from zero; however, we interpret the statistical significance of these results with caution because we estimate a number of outcomes, and a statistically significant finding for one of our outcomes could be random. A simple Bonferroni adjustment[14] could render these results statistically insignificant (R. Miller 1981). We also note that the coefficient magnitudes are very small. For example, prenatal weight gain is 0.07 lb lower when expecting a daughter, which is small relative to the mean weight gain of 30.7 lb in our sample. In the birth weight equation, the coefficient shows a statistically significant reduction of 3.1 g, and this applies regardless of marital status, age, or race. Substantively, the magnitude of the interaction effect on birth weight is small; the average birth weight in the United States is 3,329 g, and the standard deviation is 615 g. Receiving an ultrasound and having a daughter are not statistically significant determinants of APGAR scores. In sum, although some coefficients are statistically significantly different from zero, the magnitudes are very close to zero and do not point to any systematic bias in care shown by expected gender of the child. Results were also analyzed by maternal age and race (available upon request) and showed similarly small and generally statistically insignificant effects.

Having an ultrasound is not a perfect indicator for knowing fetal gender; this shortcoming might introduce measurement error in defining the treatment and control groups. Thus, we conduct separate analyses for various subgroups to distinguish between mothers who are less likely to want to know the sex of the child (mothers aged 23–39, white mothers, and married mothers) from those who are more likely to want to know fetal gender. The effect of knowing fetal gender on prenatal health investments is likely to be attenuated toward zero among the groups of mothers who received an ultrasound but who did not wish to ascertain fetal gender. However, results (available upon request) showed that the coefficients on the interaction term appear to be the same for all subgroups, which suggests that our findings are not being driven by measurement error.

Next, in Table 4, we turn to results using the NMIHS data and consider the impact of gender preference on additional measures of prenatal health investments and birth outcomes. Overall, the results here, too, do not point to systematic differences by knowing

13. Ai and Norton (2003) demonstrated that although in a linear model the interaction effect equals the coefficient on the interaction term, in nonlinear models the interaction effect is the sum of the marginal effect on the interaction term and a second term that contains the cross-partial derivatives.

14. This adjustment makes it harder to reject the null hypothesis by making the threshold for statistical significance harder to attain. For example, instead of using the conventional p value of .05, the Bonferroni adjustment would use a p value = 0.05 / n when one model is estimated repeatedly for n separate outcomes.

Table 3. Effect of Knowing Fetal Gender on Prenatal Health Investments and Birth Weight, by Birth Order and Mothers' Marital Status: 1989–2001 Natality Detail Files

	All	First Births	Higher-Order Births	Married Mothers	Unmarried Mothers
Number of Prenatal Care Visits	0.0096**	0.0099**	0.0093**	0.0063*	0.0135**
	(0.0023)	(0.0036)	(0.0031)	(0.0027)	(0.0047)
	[45,172,347]	[18,642,260]	[26,530,087]	[31,192,363]	[13,979,984]
Prenatal Weight Gain (lb.)	–0.0699**	–0.0605**	–0.0756**	–0.0625**	–0.0810**
	(0.0088)	(0.0139)	(0.0112)	(0.0100)	(0.0174)
	[36,173,248]	[15,087,543]	[21,085,705]	[25,224,782]	[10,948,466]
Probability of Inadequate Weight Gain (< 15 lb. or > 40 lb.)	–0.0019**	–0.0024**	–0.0014**	–0.0015**	–0.0026**
	(0.0003)	(0.0004)	(0.0004)	(0.0003)	(0.0005)
	[46,378,124]	[19,126,337]	[27,251,787]	[31,883,393]	[14,494,731]
Probability of Prenatal Tobacco Use	–0.00004	–0.0001	0.0001	0.0001	–0.0005
	(0.0002)	(0.0003)	(0.0003)	(0.0002)	(0.0005)
	[36,809,559]	[15,288,357]	[21,521,202]	[25,429,143]	[11,380,416]
Number of Cigarettes per Day	0.00102	0.00222	0.00004	0.00382	–0.00572
	(0.0033)	(0.0043)	(0.0047)	(0.0036)	(0.0067)
	[36,446,774]	[15,153,978]	[21,292,796]	[25,250,923]	[11,195,851]
Probability of Inadequate/ Intermediate Prenatal Care (Kessner Index)	–0.0004	–0.0001	–0.0001	–0.0002	–0.0010†
	(0.0003)	(0.0004)	(0.0003)	(0.0003)	(0.0005)
	[44,541,536]	[18,415,217]	[26,126,319]	[30,823,881]	[13,717,655]
Probability of Inadequate/ Intermediate Prenatal Care (Kotelchuck Index)	0.0002	0.00001	0.0003	0.00001	0.0004
	(0.0003)	(0.0004)	(0.0004)	(0.0003)	(0.0005)
	[44,831,356]	[18,511,253]	[26,320,103]	[31,000,905]	[13,830,451]
Birth Weight (g)	–3.4461**	–4.0854**	–3.4598**	–3.1176**	–2.8418**
	(0.3412)	(0.5392)	(0.4398)	(0.4028)	(0.6367)
	[46,354,960]	[19,116,638]	[27,238,322]	[31,869,179]	[14,485,781]

Notes: The analysis uses the 1989–2001 Natality Detail Files. OLS coefficients or probit interaction effects of Ultrasound × Female. Standard errors are in parentheses. Sample sizes are in brackets. Controls include birth order, twin status, maternal age, race, ethnicity, and education categories; foreign-born indicator; urban/rural status; region of residence; presence of pregnancy risk factors; and year of birth. Each pair of point estimates and standard errors are obtained from a separate regression; interaction effects are derived using the method described in Ai and Norton (2003).

†$p < .10$; *$p < .05$; **$p < .01$

fetal gender. Expecting a daughter does not appear to be a statistically significant determinant of prenatal care use or the adequacy of weight gain, regardless of the marital status and the gender composition in the household. As before, there are some instances in which the estimates are statistically significant. Knowing fetal gender is female is associated with a statistically significant gain in maternal weight and a decrease in the probability of quitting smoking overall and among married mothers. The magnitudes of these effects are relatively small, although having an ultrasound and eventually having a daughter appear to reduce the probability of quitting smoking by 9.8% among all mothers and by 14.3% among married mothers. Because stress has been associated with tobacco use, these results may be interpreted as follows: mothers who know that they are going to have a daughter are more likely to feel anxious and hence are less likely to quit smoking. However, the question about quitting smoking asks respondents if they quit smoking for at least one week after the pregnancy was confirmed. How the respondents interpreted this question is ambiguous because it could be interpreted as smoking behavior the week immediately after

Table 4. Effect of Knowing Fetal Gender on Prenatal Health Investments and Birth Outcomes, by Mothers' Marital Status and Household Gender Composition: 1988 National Maternal and Infant Health Survey

	All	Married Mothers	Unmarried Mothers	No Sons	No Daughters
Number of Prenatal Care Visits	0.013	–0.035	0.162	0.103	–0.456
	(0.225)	(0.252)	(0.474)	(0.490)	(0.508)
	[7,931]	[4,732]	[3,199]	[1,444]	[1,492]
Prenatal Weight Gain (lb.)	1.576†	1.753†	1.252	–1.760	2.385
	(0.95)	(1.04)	(2.13)	(2.16)	(1.96)
	[9,080]	[5,449]	[3,631]	[1,659]	[1,707]
Probability of Insufficient Weight Gain (< advised)	0.007	0.004	0.009	0.022	–0.057
	(0.034)	(0.039)	(0.064)	(0.080)	(0.077)
	[6,373]	[3,862]	[2,509]	[1,124]	[1,161]
Probability of Inadequate Weight Gain (5 lb. less/more than advised)	0.013	0.008	0.002	–0.011	–0.072
	(0.034)	(0.041)	(0.063)	(0.079)	(0.078)
	[6,373]	[3,862]	[2,509]	[1,124]	[1,161]
Probability of Quitting Tobacco Use	–0.098†	–0.143*	–0.012	–0.044	–0.092
	(0.053)	(0.068)	(0.084)	(0.114)	(0.109)
	[2,684]	[1,429]	[1,251]	[522]	[532]
Birth Weight	25.532	23.861	39.334	89.339	–29.152
	(27.862)	(32.512)	(52.986)	(67.698)	(57.841)
	[9,073]	[5,444]	[3,629]	[1,658]	[1,705]

Notes: OLS coefficient or probit interaction effects of Ultrasound × Female. Standard errors are in parentheses. Sample sizes are in brackets. The analysis is restricted to pregnancies that resulted in a live birth in the 1988 National Maternal and Infant Health Survey. Controls include birth order, twin status, maternal age, race, ethnicity, education, and household income categories; urban/rural status; region of residence; and source of payment for prenatal care. Each pair of point estimates and standard errors is obtained from a separate regression; interaction effects are derived using the method described in Ai and Norton (2003).

†$p < .10$; *$p < .05$; **$p < .01$

the pregnancy was confirmed or at any point during the pregnancy, regardless of knowing fetal gender. Since we would ideally like to know mothers' smoking decision after the sex of the child was ascertained, this result should be interpreted cautiously.

Evidence from the development literature leads us to expect son preference among immigrants, even if it is not present for American mothers in general. The effects of knowing fetal gender on prenatal health investments and birth outcomes are presented, overall and by birth order, in Tables 5 and 6 for Indian and Chinese immigrant mothers, respectively. We further broke down our analysis by mothers' educational attainment because degree of son preference may differ with mothers' educational attainment. There were no notable differences to report, but tables are available upon request. The lack of results is surprising because India and China each have a long history of son preference.

We conduct three specification tests to evaluate the validity of our method. The first specification test considers mothers' postnatal smoking decision. Results (not shown) indicate that the gender difference in the probability of smoking after the birth of the child is qualitatively and statistically similar for mothers who had an ultrasound and those who did not. The magnitude of the difference-in-differences estimate is close to zero, as expected, which lends credence to the method used in this analysis. In unreported analyses of our second specification test relating to prenatal care in the first trimester, it is heartening to note that our analysis passes this test—that is, there is no apparent effect of knowing the sex of the child in advance on whether prenatal care was initiated in the first

Table 5. Effect of Knowing Fetal Gender on Prenatal Health Investments and Birth Outcomes, Indian Immigrant Mothers: 1992–2001 Natality Detail Files

	All Births	First Births	Higher-Order Births
Prenatal Investments			
Number of prenatal care visits	0.0197	–0.0237	0.0560
	(0.0387)	(0.0557)	(0.0538)
	[143,278]	[70,705]	[72,573]
Probability of prenatal tobacco use	0.0010	0.0016	0.0006
	(0.0007)	(0.0010)	(0.0010)
Probability of inadequate weight gain (< 15 lb. or > 40 lb.)	0.0012	–0.0061	0.0082
	(0.0042)	(0.0059)	(0.0059)
Probability of inadequate/intermediate prenatal care (Kotelchuck Index)	–0.0022	–0.0002	–0.0031
	(0.0049)	(0.0068)	(0.0070)
Birth Outcomes			
Birth weight (g)	–0.4484	0.5269	–0.0627
	(5.5718)	(7.9395)	(7.8315)
	[148,977]	[73,361]	[75,616]
Probability of low birth weight (< 2,500 g)	0.0003	0.0068	–0.0062
	(0.0029)	(0.0044)	(0.0038)
Probability of very low birth weight (< 1,500 g)	0.0018†	0.0007	0.0027†
	(0.0010)	(0.0015)	(0.0014)

Notes: OLS coefficient or probit interaction effects of Ultrasound × Female. Standard errors are in parentheses. Sample sizes are in brackets. Controls include birth order, twin status, maternal age, and education categories; urban/rural status; region of residence; presence of pregnancy risk factors; and year of birth. The analysis uses the universe of children born to Indian immigrant mothers in the 1992–2001 Natality Detail Files. Each pair of point estimates and standard errors is obtained from a separate regression; interaction effects are derived using the method described in Ai and Norton (2003).

†$p < .10$

trimester.[15] This is reassuring because gender cannot generally be discerned prior to the 16th week.

A further specification check relates to the fact that the Natality Detail Files do not provide information on the timing of ultrasound receipt. Using California data, we estimate coefficients on the interaction of ultrasound receipt during labor/delivery and having a daughter. It is reassuring that the coefficient on the interaction term is essentially zero (table available upon request). These results suggest that receiving an ultrasound during labor and hence determining fetal gender during labor has no impact on the number of prenatal care visits or the likelihood of low birth weight or very low birth weight.

In summary, there is little or no evidence that gender preference plays a role in determining prenatal investments after the decision to carry the pregnancy to term. This is somewhat surprising considering that gender preference has been shown to affect many other facets of family decision-making. In order to investigate whether parents invest differentially in the health of sons *after* birth, we present unconditional means of some measures of postnatal investments that are available in the NMIHS. Table 7 reveals that, with the exception of obtaining food stamps and the length of breast-feeding, there are no statistically significant differences in the postnatal investments of mothers who have sons

15. Results not reported reveal that the analyses using the Indian and Chinese immigrant samples also pass the robustness test.

Table 6. Effect of Knowing Fetal Gender on Prenatal Health Investments and Birth Outcomes, Chinese Immigrant Mothers: 1989–2001 Natality Detail Files

	All Births	First Births	Higher-Order Births
Prenatal Investments			
Number of prenatal care visits	0.0079	–0.0442	0.0614†
	(0.0244)	(0.0347)	(0.0343)
	[288,398]	[146,547]	[141,851]
Probability of prenatal tobacco use	0.0003	0.0005	0.0001
	(0.0009)	(0.0012)	(0.0013)
Probability of inadequate weight gain (< 15 lb. or > 40 lb.)	0.0024	0.0038	0.0010
	(0.0027)	(0.0039)	(0.0038)
Probability of inadequate/intermediate prenatal care (Kotelchuck Index)	0.0006	0.0015	–0.0002
	(0.0032)	(0.0045)	(0.0047)
Birth Outcomes			
Birth weight (g)	–4.9716	–8.7292†	–0.9646
	(3.6293)	(5.1028)	(5.1641)
	[295,724]	[150,435]	[145,289]
Probability of low birth weight (< 2,500 g)	0.0016	0.0030	0.0001
	(0.0014)	(0.0021)	(0.0019)
Probability of very low birth weight (< 1,500 g)	–0.0004	–0.0001	–0.0007
	(0.0005)	(0.0007)	(0.0007)

Notes: OLS coefficient or probit interaction effects of Ultrasound × Female. Standard errors are in parentheses. Sample sizes are in brackets. Controls include birth order, twin status, maternal age, and education categories; urban/rural status; region of residence; presence of pregnancy risk factors; and year of birth. The analysis uses the universe of children born to Chinese immigrant mothers in the 1989–2001 Natality Detail Files. Each pair of point estimates and standard errors is obtained from a separate regression; interaction effects are derived using the method described in Ai and Norton (2003).

†$p < .10$

versus those who have daughters. Since these are unconditional means, we cannot take this as evidence of daughter preference because mothers may breast-feed daughters longer and obtain food stamps simply because girls tend to be lighter at birth. A t test of comparison fails to reject the null hypothesis that the average number of well-child visits, intensity of breast-feeding, and insurance take-up are equal for mothers with girls versus boys.

One possible explanation for our finding that child gender does not impact prenatal health investments is that studies of male-biased investments often implicate fathers as the source of gender bias. Prenatal choices and investment decisions, however, are to a larger extent made by mothers relative to joint decisions like marriage.[16] Prenatal health investments are also likely to involve shorter-term costs than investments in other areas, such as marriage and divorce. This might explain why we fail to see any strong evidence for the role of gender preference in prenatal investments, even though prior literature has established a role for it in marriage decisions. A third plausible explanation is measurement error. As we discussed earlier, our classification of mothers into control and treatment groups is imperfect, and this error would bias the coefficients downward. Although we provide some evidence to the contrary, it remains a possibility.

16. In order to test the role of fathers' preferences in determining prenatal investments, we estimated our models for same-race immigrant couples. We failed to reject the null hypothesis that prenatal investments are equal for sons and daughters. Similarly, the results were virtually identical for married and unmarried parents.

Table 7. Gender Difference in Postnatal Investments: 1998 National Maternal and Infant Health Survey

	Girls		Boys		
	Mean	*SE*	Mean	*SE*	Difference
Number of Well-Child Visits	42.572	0.345	43.005	0.348	–0.432
% Mothers Who Got WIC Postnatally, No WIC Before	66.944	0.972	64.118	1.006	2.826*
Number of Days Breast-fed	141.535	3.107	128.856	2.764	12.679**
Intensity of Breast-feeding, First Month	4.782	0.076	4.728	0.077	0.054
Intensity of Breast-feeding, Second Month	3.802	0.072	3.683	0.072	0.119
Intensity of Breast-feeding, Third Month	2.883	0.064	2.813	0.065	0.070
Intensity of Breast-feeding, Fourth Month	2.212	0.056	2.142	0.058	0.070
Intensity of Breast-feeding, Fifth Month	1.803	0.051	1.745	0.053	0.058
Intensity of Breast-feeding, Sixth Month	1.534	0.048	1.478	0.049	0.056
% Mothers With Health Insurance, Postnatal	68.691	0.695	70.004	0.680	–1.313
% Children With Any Immunization	97.183	0.247	97.291	0.240	–0.108

$^{*}p < .05$; $^{**}p < .01$

The lack of evidence of gender preference in prenatal behavior among immigrant women is worth discussing further because of the home-country findings.[17] A compelling explanation for this lack of result is that the preference biases are entirely resolved in apparent gender-selective abortion. Another possible explanation is that women who choose to immigrate to the United States are innately different from their counterparts in India and China. For instance, in 1991, approximately 13% of women in India had more than primary education (Velkoff 1998) compared with the college completion rate of 49.6% in the Indian immigrant sample. Education is one dimension along which immigrant women differ from their counterparts at home, and it is plausible that they behave differently in other aspects as well. A third possibility is that the economic and policy environment in the United States is such that gender-biased prenatal investments are no longer optimal. For example, the one-child policy may be fostering gender preference in the home country, and when such restrictions are relaxed, gender neutrality may be the norm. This could also apply to other policies, such as old-age pensions. Indian immigrant mothers need not worry about the economic ramifications of having girls versus boys, and these factors may help mitigate gender preference among immigrant mothers. In fact, in 2002, Indian immigrants had lower fertility (2.23) compared with women in India (3.07), and Chinese immigrants had higher fertility rates (2.26) relative to women in China (1.70) (Camarota 2005). These results suggest that the fertility decisions of Indian and Chinese mothers are different in the United States than in their home countries, perhaps because of the economic and policy environment in the United States. Finally, we thank an anonymous referee for pointing out that the nature of physician intervention in the United States may be the reason why we don't observe gender-biased prenatal decisions. For instance, the possibility of malpractice lawsuits may lead physicians to monitor mothers' prenatal health more closely so that the negative effects of gender-biased investments are mitigated. Another example is the

17. For the home countries, evidence of son preference has been found in the context of abortion and postnatal investments; to our knowledge, no study has considered the prenatal period, so this sentence is valid under the assumption that in India and China, son preference extends into the prenatal period as well.

prenatal care schedule prescribed by ACOG, which is likely to lead to regular interaction between physicians and patients and hence greater transfer of information regarding healthy behaviors and the negative consequences of inadequate prenatal investments, so that mothers may be less likely to behave in a manner that poses a risk to the unborn child.[18] We cannot prove these claims definitively; nevertheless, the study of immigrants provides an interesting laboratory in which to explore the possible causes of gender bias displayed in home countries and the extent to which it is affected by public policy.

CONCLUSION

Parents invest in their own health and allocate resources that impact child well-being. Parental preferences regarding the gender of the child could impact first their abortion decisions and later their investment decisions. In this article, we study parents' prenatal health investment decisions conditional upon deciding to carry the pregnancy to term. Prenatal health investments are measured by maternal choices during pregnancy, such as number of prenatal care visits, adequacy of weight gain, and alcohol and tobacco use during pregnancy—inputs that are known to affect infant health. Furthermore, we test whether gender preference persists among first-generation immigrants who were born in India and China, two countries with a documented history of son preference.

Using the 1989–2001 Natality Detail Files and 1988 NMIHS, which provide information on live births in the United States, we estimate the effect of gender preference on prenatal health investments. We begin by examining sex ratios among Indian and Chinese immigrants and provide evidence consistent with findings in the literature that the practice of sex-selective abortion exists among these immigrant groups. Since having an ultrasound increases the chances of knowing the gender of the child, we then compare the behavior of women who had ultrasounds and eventually had a girl with women who had an ultrasound and eventually had a boy. Although there are large effects of child gender on parental investment in marriage and time use, we do not find that it accompanies differences in health investments as we measure them here. Admittedly, measurement error may have biased our results downward to the point that we fail to capture the true effect of knowing fetal gender on prenatal investments. This, together with the fact that our results are not precise zeros, means that our failure to find statistically significant evidence of gender bias does not mean it does not exist.

Besides possible measurement error, there are other plausible reasons why we fail to find evidence consistent with gender bias among pregnancies that are brought to term. First, we study outcomes that are more a result of mother's actions than of father's actions, and previous literature has shown that gender preferences are likely to be instigated by fathers. Second, the lack of result for immigrant mothers may be because women who immigrate to the United States are different from their counterparts at home. Third, it is plausible that immigrant mothers do not exhibit gender preference because the economic and policy environment in the United States changes parents' decision-making problem such that gender-biased investments are no longer optimal. Finally, physician intervention in the United States might involve closer monitoring of fetal health or greater provision of prenatal health information so that mothers opt to make healthier decisions or so that the negative effects of gender-biased investments are mitigated. We attempt to test the father preference and physician intervention explanations empirically, and although we don't find much support for these explanations, we believe they remain plausible since our data and hence our tests are not perfect.

18. At one referee's suggestion, we tested whether mothers who had less interaction with their physicians exhibit gender bias in their prenatal investment decisions by estimating our models for the subset of mothers who initiated prenatal care after the first trimester. However, we are unable to reject the null hypothesis that prenatal investments in daughters and sons are the same in any of the three samples.

This analysis is limited by the lack of unambiguous information on parents' knowledge of fetal gender; however, it serves as a starting point for estimating the effect of knowing fetal gender on prenatal health investments. We find evidence consistent with sex-selective abortion among mothers from India and China. However, once the termination decision has been resolved, knowing the sex of the child in advance does not appear to affect prenatal health investments—neither among immigrant nor all U.S. mothers.

REFERENCES

Abrevaya, J. 2005. "Are There Missing Girls in the United States? Evidence on Gender Preference and Gender Selection." Working paper. Department of Economics, Purdue University.

———. Forthcoming. "Are There Missing Girls in the United States? Evidence From Birth Data." *American Economic Journal, Applied Economics*.

Ai, C. and E.C. Norton. 2003. "Interaction Terms in Logit and Probit Models." *Economics Letters* 80(1):123–29.

Almond, D. and L. Edlund. 2008. "Son-Biased Sex Ratios in the 2000 United States Census." *Proceedings of the National Academy of Science* 105:5681–82.

Angrist, J.D. and W.N. Evans. 1998. "Children and Their Parents' Labor Supply: Evidence From Exogenous Variation in Family Size." *American Economic Review* 88:450–77.

Arnold, F., S. Kishor, and T.K. Roy. 2002. "Sex-Selective Abortions in India." *Populations and Development Review* 28:759–85.

Bardhan, P. 1982. "Little Girls and Death in India." *Economic and Political Weekly* 17:1448–50.

Barker, D.J.P. 1997. "Fetal Nutrition and Cardiovascular Disease in Later Life." *British Medical Bulletin* 53(1):96–108.

Barnett, R.C. and G.K. Baruch. 1987. "Determinants of Fathers' Participation in Family Work." *Journal of Marriage and the Family* 49:29–40.

Becker, G. 1981. *A Treatise of the Family*. Cambridge, MA: Harvard University Press.

Behrman, J.R., R.A. Pollak, and P. Taubman. 1986. "Do Parents Favor Boys?" *International Economic Review* 27:33–54.

Behrman, J.R. and M.R. Rosenzweig. 2004. "Returns to Birthweight." *Review of Economics and Statistics* 86:586–601.

Behrman, J.R., M.R. Rosenzweig, and P. Taubman. 1994. "Endowments and the Allocation of Schooling in the Family and in the Marriage Market: The Twins Experiment." *Journal of Political Economy* 102:1131–74.

Berle, J.O., A. Mykletun, A.K. Daltveit, S. Rasmussen, F. Holsten, and A.A. Dahl. 2005. "Neonatal Outcomes in Offspring of Women With Anxiety Disorder and Depression During Pregnancy: A Linkage Study From the Nord-Trondelag Health Study (HUNT) and Medical Birth Registry of Norway." *Archives of Women's Mental Health* 8:181–89.

Bloch, F. and V. Rao. 2002. "Terror as a Bargaining Instrument: A Case Study of Dowry Violence in Rural India." *American Economic Review* 94:1029–43.

Boss, D.J. and R.E. Timbrook. 2001. "Clinical Obstetric Outcomes Related to Continuity in Prenatal Care." *Journal of the American Board of Family Practice* 14:418–23.

Bryant, W.K. and C.D. Zick. 1996. "An Examination of Parent-Child Shared Time." *Journal of Marriage and the Family* 58:227–37.

Camarota, S. 2005. "Birth Rates Among Immigrants in America: Comparing Fertility in the U.S. and Home Countries." Center for Immigration Studies (CIS) Backgrounder #11-05. CIS, Washington, DC.

Coale, A.J. and J. Banister. 1994. "Five Decades of Missing Females in China." *Demography* 31:459–79.

Currie, J. 2004. "Viewpoint: Child Research Comes of Age." *Canadian Journal of Economics* 37:509–27.

Currie, J. and R. Hyson. 1999. "Is the Impact of Health Shocks Cushioned by Socioeconomic Status? The Case of Low Birthweight." *AEA Papers and Proceedings* 89:245–50.

Dahl, G.B. and E. Moretti. 2004. "The Demand for Sons: Evidence From Divorce, Fertility, and Shotgun Marriage." NBER Working Paper #10281. National Bureau of Economic Research, Cambridge, MA.

D'Souza, S. and L. Chen. 1980. "Sex Differentials in Mortality in Rural Bangladesh." *Population and Development Review* 6:257–70.

Evans, W.N. and J.S. Ringel. 1999. "Can Higher Cigarette Taxes Improve Birth Outcomes?" *Journal of Public Economics* 72:135–54.

Gangadharan, L. and P. Maitra. 2003. "Testing for Son Preference in South Africa." *Journal of South African Economies* 12:371–416.

Grossman, M. 1972. "On the Concept of Health Capital and the Demand for Health." *Journal of Political Economy* 80:223–55.

Grossman, M. and T.J. Joyce. 1990. "Unobservables, Pregnancy Resolutions, and Birth Weight Production Functions in New York City." *Journal of Political Economy* 98:983–1007.

Haddow, J.E., G.J. Knight, G.E. Palomaki, and P.K. Haddow. 1988. "Estimating Fetal Morbidity and Mortality Resulting From Cigarette Smoke Exposure by Measuring Cotinine Levels in Maternal Serum." *Progress in Clinical and Biological Research* 281:289–300.

Harris, K.M. and S.P. Morgan. 1991. "Fathers, Sons, and Daughters: Differential Paternal Involvement in Parenting." *Journal of Marriage and the Family* 53:531–44.

Joyce, T.J. and M. Grossman. 1990. "Pregnancy Wantedness and Early Initiation of Prenatal Care." *Demography* 27:1–17.

Joyce, T.J., R. Kaestner, and S. Korenman. 2000. "The Effect of Pregnancy Intention on Child Development." *Demography* 37:83–94.

Kohler, H.-P., J.R. Behrman, and A. Skytthe. 2005. "Partner + Children = Happiness? The Effects of Partnerships and Fertility on Well-being." *Population and Development Review* 31:407–45.

Kotelchuck, M. 1994. "An Evaluation of the Kessner Adequacy of Prenatal Care Index and a Proposed Adequacy of Prenatal Care Utilization Index." *American Journal of Public Health* 84:1414–20.

Kynch, J. and A. Sen. 1983. "Indian Women: Well-being and Survival." *Cambridge Journal of Economics* 7:363–80.

Lhila, A. 2005. "Birth Spacing and Child Health: An Instrumental Variables Approach." Working Paper. Department of Economics, Cornell University.

Lundberg, S., S. McLanahan, E. Rose. 2007. "Child Gender and Father Involvement in Fragile Families." *Demography* 44:79–92.

Lundberg, S. and E. Rose. 2002. "The Effect of Sons and Daughters on Men's Labor Supply and Wages." *Review of Economics and Statistics* 84:251–68.

———. 2003. "Child Gender and the Transition to Marriage." *Demography* 40:333–49.

———. 2004. "Investments in Sons and Daughters: Evidence From the Consumer Expenditure Survey." Pp. 163–80 in *Family Investments in Children: Resources and Behaviors That Promote Success,* edited by A. Kalil and T. DeLeire. Philadelphia: Lawrence Erlbaum.

Mammen, K. 2003. "The Effect of Children's Gender on Divorce and Child Support." Unpublished manuscript. Department of Economics, Columbia University.

Maritz, G.S., C.J. Morley, and R. Harding. 2005. "Early Developmental Origins of Impaired Lung Structure and Function." *Early Human Development* 81:763–71.

Martin, J.A., B.E. Hamilton, S.J. Ventura, F. Menacker, M.M. Park, and P.D. Sutton. 2002. "Births: Final Data for 2001." *National Vital Statistics Reports* 51(2). Hyattsville, MD: National Center for Health Statistics.

Mellor, J.M. 2001. "Long-term Care and Nursing Home Coverage: Are Adult Children Substitutes for Insurance Policies?" *Journal of Health Economics* 20:527–47.

Miller, B.D. 1981. *The Endangered Sex: Neglect of Female Children in Rural North India*. Ithaca, NY: Cornell University Press.

Miller, R.G. 1981. *Simultaneous Statistical Inference*. New York: Springer-Verlag.

Morgan, S.P., D.N. Lye, and G.A. Condran. 1988. "Sons, Daughters, and the Risk of Marital Disruption." *American Journal of Sociology* 94:110–29.

Okah, F.A., J. Cai, and G.L. Hoff. 2005. "Term-Gestation Low Birth Weight and Health-Compromising Behaviors During Pregnancy." *Obstetrics and Gynecology* 105:543–50.

Osmani, S. and A. Sen. 2003. "The Hidden Penalties of Gender Inequality: Fetal Origins of Ill-Health." *Economics and Human Biology* 1:105–21.

Pollard, M.S. and S.P. Morgan. 2002. "Emerging Parental Gender Indifference? Sex Composition of Children and the Third Birth." *American Sociological Review* 67:600–13.

Rosenzweig, M.R. and T.P. Schultz. 1982a. "Market Opportunities, Genetic Endowments, and Intrafamily Distribution of Resources: Child Survival in Rural India." *American Economic Review* 72:803–15.

———. 1982b. "The Behavior of Mothers and Inputs to Child Health: The Determinants of Birth Weight, Gestation, and Rate of Fetal Growth." Pp. 53–92 in *Economic Aspects of Health,* edited by V. Fuchs. Chicago: University of Chicago Press.

———. 1983. "Estimating a Household Production Function: Heterogeneity, the Demand for Health Inputs, and Their Effects on Birth Weight." *Journal of Political Economy* 91:723–46.

Sable, M.R., J.C. Spencer, J.W, Stockbauer, W.F. Schramm, V. Howell, and A.A. Herman. 1997. "Pregnancy Wantedness and Adverse Pregnancy Outcomes: Differences by Race and Medicaid Status." *Family Planning Perspectives* 29(2):76–81.

Sable, M.R. and D.S. Wilkinson. 2000. "Impact of Perceived Stress, Major Life Events and Pregnancy Attitudes on Low Birth Weight." *Family Planning Perspectives* 32:288–94.

Sen, A. 1992. "Missing Women: Social Inequality Outweighs Women's Survival Advantage in Asia and North Africa." *British Medical Journal* 304:587–88.

———. 2003. "Missing Women—Revisited." *British Medical Journal* 327:1297–98.

Shiono, P.H. and R.E. Behrman. 1995. "Low Birth Weight: Analysis and Recommendations." *The Future of Children* 5(1):4–18.

Shipp, T.D., D.Z. Shipp, B. Bromley, R. Sheahan, A. Cohen, E. Lieberman, and B. Benacerraf. 2004. "What Factors Are Associated With Parents' Desire to Know the Sex of Their Unborn Child?" *Birth* 31:272–79.

Streissguth, A.P., D.C. Martin, J.C. Martin, and H.M. Barr. 1981. "The Seattle Longitudinal Prospective Study on Alcohol and Pregnancy." *Neurobehavioral Toxicology and Teratology* 3:223–33.

Teachman, J.D. and P.T. Schollaert. 1989. "Gender of Children and Birth Timing." *Demography* 26:411–23.

Thomas, D. 1994. "Like Father, Like Son; Like Mother, Like Daughter: Parental Resources and Child Height." *Journal of Human Resources* 29:950–88.

Velkoff, V.A. 1998. "Women of the World: Women's Education in India." U.S. Department of Commerce, Bureau of the Census, International Programs Center WID/98-1.

Walker, M.K. and G.K. Connor. 1993. "Fetal Sex Preference of Second-Trimester Gravidas." *Journal of Nurse Midwifery* 38(2):110–13.

Warner, G.L. 1995. "Prenatal Care Demand and Birthweight Production of Black Mothers." *American Economic Review* 85:132–37.

Weller, R.H., I.W. Eberstein, and M. Bailey. 1987. "Pregnancy Wantedness and Maternal Behavior During Pregnancy." *Demography* 24:407–12.

Yeung, W.J., J.F. Sandberg, P.E. Davis-Kean, and S.L. Hofferth. 2001. "Children's Time With Fathers in Intact Families. *Journal of Marriage and Family* 63:136–54.