Prenatal Health Investment Decisions: Does the Child's Sex Matter?*

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Citation:
Lhila, Aparna and Kosali I. Simon. 2008. "Prenatal Health Investment Decisions: Does the Child's Sex Matter?" *Demography*, 45(4): 885-905

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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 paper.

ABSTRACT

Individuals invest in their own health, but children rely on parents to act on their behalf, especially in the case of prenatal health. This paper asks: Do parents in the U.S. who choose to give birth allocate resources differently in improving 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. We use birth certificate data from 1989-2001 on maternal prenatal behaviors (e.g. smoking, weight gain, and prenatal care use) and child health outcomes at birth (e.g. birthweight). Ultrasound receipt proxies for knowing fetal gender, enabling us to separate out child sex-related biological differences from investment differences in sons' and daughters' health. We find 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 U.S., neither in the population as a whole nor among first-generation 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 U.S. and in other countries have been shown to display son preference in several dimensions, including fertility, marriage, and postnatal investment decisions in child health (e.g. Rosenzweig and Schultz 1982; Lundberg and Rose 2003; Dahl and Moretti 2004; Abrevaya, 2005; Almond and Edlund 2008). In this paper we study the effect of knowing fetal gender on one class of particularly influential health decisions, i.e., prenatal health investments 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. In our work, we first examine the possibility of sex –selective abortion but our main question is on the effect of knowing fetal gender on prenatal health investments *conditional upon having made the abortion decision*. Hence we answer the question: Do parents in the U.S. who choose to give birth allocate resources differently in improving 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 public policy perspective the question raised in this paper is an important one. If knowing the child's sex in advance disadvantages some children's health at birth, then a

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

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 India where knowing the sex of the child in advance 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 are well documented in the development literature, especially with regard to parents in South and Southeast Asia (e.g. Sen 1992, 2003). American parents have preferences regarding the sex mix of their children with regard to fertility stopping rules (Angrist and Evans 1998), and Dahl and Moretti (2004) finds that American parents' decisions regarding marriage and divorce are also guided by their preference for sons.[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 U.S., then is there any reason to expect different parental behaviors based on child gender among pregnancies parent decide to bring to term? We present a conceptual argument why parents who opt not to terminate a pregnancy based on preferred child gender may invest differently in the prenatal health of their sons and daughters. First, assuming a continuum of son-preference or "distaste" for girls[4], and a continuum of individual-specific abortion costs which includes time and psychological costs, it is likely that there remain

[2] Authors' calculations from Natality Detail data 1989-2001
[3] There are other studies suggesting that this phenomenon may be lessening over time (e.g. Pollard and Morgan 2002). Neither that study nor Dahl and Moretti (2004) considers immigrants separately.
[4] The use of the term "distaste for daughter", though unsavory, serves to describe parents' lack of preference for a daughter, i.e. the opposite of son preference.

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 still be ways in which their son preference may manifest itself. Second, since prior work has found that post-natal gender discrimination in investments exists it is logical to test whether gender-biased resource allocations 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 paper 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 use data from the 1989-2001 Natality Detail Files and use the method developed in Dahl and Moretti (2004) -- gender difference in the prenatal behavior of mothers who did not have an ultrasound (and hence did not know the sex of the child in advance) are used to separate out the natural or biological difference from the investment difference that is driven by preferences regarding the sex of the child.

We choose to focus on mothers' prenatal investment decisions because they are potentially modifiable, have utility consequences for mothers,[6] and are important in influencing birth outcomes and shaping the long-term health of their children. These costs and benefits give rise to demand functions for child health inputs which will vary across women, perhaps systematically related to whether the fetus she carries is a boy or a girl. If U.S. parents do in fact favor sons we expect to find that after controlling for any biological effects (like morning sickness) caused by carrying a boy vs. girl fetus, knowing that the unborn child is female may, at

[5] We will use the term 'mother' in the rest of the text, although we will return to discuss the role played by fathers vs. mothers in the decisions we study.
[6] E.g. the time and money costs of prenatal health care visits, the social and psychological costs of tobacco and alcohol abstention, or difficulties of attaining suggested weight gain goals.

the margin, have a negative impact on demand for prenatal care and health outcomes at birth. 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 paper 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 U.S. Finally, we study the persistence of gender preference among immigrant mothers, a question that to our knowledge has received very little attention despite the preponderance of evidence on son preference in mothers' country of birth.[7]

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 into 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 which 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 paper addresses a particular form of wantedness, i.e., a pregnancy may be unwanted if parents do not desire a child of that

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

sex. Studies indicate that 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 birthweight (LBW) children (Sable et al. 1997; Sable and Wilkinson 2000). Thus not all pregnancies carried to term display medically optimal prenatal behavior in the U.S. or elsewhere.

In developing countries gender preference generally takes the form of preference for sons. 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 to sons. Sex ratios (boy/girl) at birth for most societies lie between 1.03 and 1.06, thus 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 mortality in the first five years of life has also been attributed to parents choosing to invest resources, such as food, nutrients and medical care in sons (D'Souza and Chen 1980; B. Miller 1981; Bardhan 1982; Rosenzweig and Schultz 1982; Kynch and Sen 1983).

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

A question this literature brings up is the source of the gender bias. In the U.S. fathers have largely been implicated as the source of son preference. Fathers prefer to invest in the health of their sons compared to mothers who have a greater impact on the health of their daughters (Thomas 1994). Fathers spend more time with their sons (Yeung et al. 2001; Bryant and Zick 1996), are more likely to be involved in the caretaking of the son than their daughter (Lundberg, McLanahan and Rose 2005), and spend more time with their children overall when they have a son (Barnett and Baruch 1987; Harris and Morgan 1991). In addition to revealed behavioral differences, an analysis of stated preferences also suggests that fathers favor boys (Dahl and Moretti 2004). Disentangling the role of mothers and fathers in determining prenatal decisions is beyond the scope of this paper, 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 towards examining the role of fathers' preferences by conducting our analysis for married and unmarried mothers separately. While not perfect, it is plausible that spouses exert greater influence on prenatal decisions than unmarried partners.

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 related behaviors of mothers (e.g. use of medical care, smoking decisions) as investments in the production of child health (e.g. Rosensweig znd Shultz 1983, 1982b). We focus on behaviors which are known to impact health in-utero and at birth, which in turn have long-term consequences for health, educational attainment and earning capacity in adulthood. Medically, maternal nutrition and lifestyle and the fetus' exposure to restricted nutrient intake and smoking are likely to case LBW and have long-lasting health effects such as such as hypertension, stroke, type-2 diabetes, and respiratory and heart disease (Barker 1997; Maritz et al. 2005). Improving mothers' prenatal care use along with altering mothers' prenatal health habits such as tobacco, alcohol and drug use, and improving maternal health like adequate weight gain are likely to increase birthweight (Warner 1995; Shiono and Behrman 1995; Evans and Ringel 1999; Boss and Timbrook 2001). Smoking and alcohol use during pregnancy are also associated with low APGAR scores (Okah, Cai and Hoff 2005; Haddow et al. 1988; Streissguth et al. 1981).

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

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. The scores, between 0 and 10, rate infant health based on five criteria – heart rate, respiration, muscle tone, reflex and color, and are measured one and five minutes after birth. 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 adverse infant health outcomes such as lower APGAR scores at both the first and fifth minute after birth (Berle et al. 2005), and since anxiety and depression may accompany unwanted pregnancies, we study APAGAR scores as one of the outcomes.

THEORETICAL FRAMEWORK

We assume that children enter parents' utility functions and that parents are concerned with and derive utility from the general 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 because sons and daughters may enter parents' utility functions in different ways. Consider a utility maximization decision where parents decide the optimal prenatal investment, I, in the health of a child:

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

$$s.t. W = P_X X + P_I I$$

Where I denotes prenatal investments such as use of prenatal care, X stands for all other market goods that provide utility to parents, G stands for the gender of the child, e and c represent economic and cultural conditions, respectively, that pertain to whether a boy or girl would provide more utility to parents, and H denotes child health at birth, which could 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 the composite market good, X, and P_I is the pecuniary and non-pecuniary cost of prenatal investment.

Dahl and Moretti (2004) review reasons why gender preference may exist in the U.S. Internationally, parents may prefer sons as the costs associated with raising daughters (c and e in the utility maximization problem) are traditionally higher in countries that practice the custom of transferring a dowry from the bride's parents to the groom's parents at the time of marriage (e.g. Bloch and Rao 2002). Additionally, daughters in many societies leave the parents' home upon marriage, whereas sons bring wives into the parent's house and provide old-age economic security to parents. This is especially important in countries where the government makes no provision for post-retirement income, and is correspondingly less important in a U.S. setting although daughters have been shown to provide more old age assistance than sons in the U.S. (Mellor, 2001). Parents may choose to invest in sons because they expect that returns on human capital investment wil be higher for sons than for daughters. There are also important cultural differentials; for example, 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. On the other hand, parents may prefer to have daughters for equally plausible reasons. For instance, daughters are more likely to provide emotional support to parents and are more likely to be caregivers to elderly parents. 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 differentially 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 as 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 differential relative to a context in which abortions were 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[8] despite a preponderance of information on their health consequences, we consider this evidence that prenatal health investments involve non-trivial costs.

The degree to which we should expect son preference among immigrants from countries with demonstrated son preference is unclear for several reasons. One, there may be selective migration, and immigrant mothers may not be predisposed to son preference even in their country of origin. Alternatively, the new environment they enter may not provide conditions that drive son preference- i.e. resource constraints, policy constraints (such as the one child policy) or cultural considerations, or assimilation into the new environment may change c and e in Equation 1. For instance, Malays of Indian and Chinese descent do not display son preference in their fertility decision, and are instead motivated by the desire for a balanced sex composition (Lhila, 2005).

Prenatal investment differentials 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, alcohol or tobacco use during pregnancy and adversely affect the child's weight and APGAR scores at birth. We present estimates from

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[8] Based on the authors' calculations from the 1989-2001 Natality Detail Files and the 1989 National Maternal and Infant Health Survey,14.7% of mothers report using tobacco during pregnancy, 19.58% initiate prenatal care in the second trimester or later, 28.3% of pregnancies receive a score of inadequate/intermediate on the Kotelchuck adequacy of prenatal care use index, and 58.5% of mothers fail to attain their prenatal weight gain goals.

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. This varies with length of gestation, the presence of medical risk factors. The overall adequacy of prenatal care use is rated as adequate, intermediate or inadequate by the Kessner Index which uses the number of prenatal care visits and the trimester of prenatal care. . Kotelchuck (1994) suggests an alternate measure of adequacy of prenatal care which adjusts for length of gestation and rates adequacy into four categories -- inadequate, intermediate, adequate and adequate plus. It is generally recommended that a woman of normal weight gain 25-30 pounds during pregnancy, overweight women gain 15-20 pounds and underweight women gain 28-40 pounds. As discussed earlier, each of these inputs along with nutrient intake and maternal smoking and alcohol use during pregnancy impact the health status of the child in-utero and at birth.

METHOD

There are four prenatal tests (obstetric ultrasound imaging, amniocentesis, chorionic villus sampling (CVS) and percutaneous umbilical blood sampling (PUBS)) used for diagnosing fetal health that could, as a byproduct, also determine fetal gender. Ultrasounds[9], the most pervasive method, are performed on 68% of mothers typically in the 18th or 20th week of the pregnancy,

[9] 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.

and can reveal the sex of the child with 95-100% accuracy by the 16th week. Amniocentesis[10] is 100% accurate, but this procedure 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[11] carry the greatest risk of miscarriage and have lower useage rates.

After ascertaining the sex of the child and before making prenatal investment decisions parents decide whether or not to abort the pregnancy. Thus we begin our analysis by examining whether the practice of sex-selective abortions exists among immigrants in the U.S. Since we do not observe pregnancies that are terminated we compare the sex ratio (number of boy births, per girl birth) in the U.S. to the sex ratio among first generation 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. T-tests reveal that immigrant mothers are statistically significantly more likely to have sons than daughters relative to all U.S. mothers. Statistically significant differences in sex ratios persist until the fourth parity for immigrants from China; whereas among Indian mothers sex ratios differ from American mothers at second, third, fourth and sixth or greater parities. Concurrent work (Abrevaya 2005; Almond and Edlund 2008) examine there issues in greater detail and also finds evidence consistent with sex-selective abortion.

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

[10] Amniocentesis removes a small amount of amniotic fluid to test for genetic abnormalities and fetal health.
[11] CVS relies on extracting a sample of the placenta and PUBS withdraws blood from the baby to test for genetic problems or abnormalities.

downward biased. Thus the results of our analysis should be interpreted as the partial effect of knowing fetal gender on prenatal health investments, i.e. 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 compare the outcomes of parents who had girls to those 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, and among ultrasound mothers any gender difference in outcomes that exists over and above the biological difference is interpreted as an investment differential This method is similar to that used by Dahl and Moretti (2004). Mechanically speaking, our method of estimation is to regress each of the outcomes of interest -- birthweight, APGAR scores, number of prenatal visits, prenatal 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 that indicates that the mother received an ultrasound during her pregnancy, and the interaction of the two. Our key parameter is the coefficient on the interaction between gender and ultrasound. Evidence in support of girl preferences that translate into differential prenatal investment and outcomes at birth 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 1st prenatal visit as the mother can only find out the gender through technology that would have to be used during a prenatal visit. Thus, we include whether or not the mother had her first prenatal visit in the first trimester as a dependent variable which 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 towards finding an effect of son preference.[12]

Our estimation method also assumes that all mothers who receive ultrasounds know fetal gender precisely. Although ultrasounds determine the sex of the child with 95-100% accuracy, not all mothers who receive an ultrasound find out the sex of the child. Eighty percent of mothers in the U.S. stated a desire to find out the gender of their child in a survey conducted at the University of Alabama (Walker and Conner, 1993). Another study of expectant mothers who received an ultrasound in Boston, MA found that parents who want to know the sex of the child and those who do not differ along socioeconomic and demographic characteristics. Mothers and fathers were equally likely (58%) to want to know the sex of the child, however, parents who were planning to move or renovate, not planning to breastfeed, had conceived accidentally, fathers' without a full-time job, lower household income, unwed mothers, mothers less than 22

[12] 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 vs. 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..

or greater than 40, without a college degree, non-White and non-Catholic, and had preferences regarding the sex of the child, were more likely to want to know the sex of the child before delivery (Shipp et al. 2004). 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. This is likely to attenuate the effect of knowing fetal gender on health investments towards zero. Moreover, the extent of measurement error is likely to be greater for groups of mothers who are less likely to want to know the sex of the child. As a robustness check, we conduct our analysis by different sub-groups such as by mothers' race, education, marital status and age, as described in Shipp et al. (2004), and return to this point when interpreting our results.

In order to test the validity of our method we execute an additional specification test. Having an ultrasound and knowing the gender of the child during pregnancy should have no impact on the mothers' postnatal decisions, such as smoking, as 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, as 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 U.S. between 1989 and 2001. These files are a compilation of birth certificate data from all 50 states and the District of Columbia, which provides 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 a singleton birth, since the prenatal investments into twins are likely to sharply differ from investments into singletons. Approximately 4 million infants are born in the U.S. every year, which leads to a sample of over 46 million observations for our analysis. In order to study the prenatal investment patterns among first-generation immigrant mothers we use the universe of mothers who reported being foreign born and identified their race as Chinese or Indian, respectively. The Chinese and Indian[13] immigrant sample contains 304,530 and 154,492 live births respectively. Means for variables of interest in each of the samples are presented in Table 2.

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 re-run 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, provides large sample sizes to allow intricate sub-group analyses, and lends a great deal of statistical power to our analysis, it provides terse information about some birth outcomes.

[13] Asian Indian was included as a distinct race category in birth certificates beginning in 1992. Thus the Indian immigrant sample spans 1992-2001.

For instance, mothers report tobacco use during pregnancy but this question makes no distinction between tobacco use after mothers find out that they are pregnant versus before. 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' pre-pregnancy 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 which is too small to study first generation immigrants from China and India. However, an additional advantage to using the NMIHS sample is that it provides information on household composition, income and source of payment for prenatal care visits. These may be important controls as 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 maternal 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 our 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 U.S., 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 Indian and Chinese immigrant mothers separately, for different outcomes. In each of the three samples, irrespective of whether the mother received an ultrasound, boys tend to weigh more than girls at birth. For instance, in the U.S. sample the boy-girl difference in birthweight tends to be 117.70 grams for mothers who received an ultrasound and 114.28 grams 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 birthweight among mothers who received an ultrasound reveals that knowing child's sex is associated with a 3.42 gram difference in the birthweight of boys and girls. This difference in differences statistic shown in the last column in that table 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 miniscule in substantive or economic terms. Although these results are from simple correlations, the remainder of this section reveals that this relationship holds qualitatively the same for all U.S. mothers, and Indian and Chinese immigrant mothers, irrespective of education 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.

The results of our regression analysis are presented in Table 3, which contains the coefficients and standard errors of the interaction term ULTRASOUND*FEMALE, showing the effect of receiving an ultrasound and having a daughter 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 the method suggested by Ai and Norton (2003)[14]. The regressions control for maternal age, education 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 of knowing fetal gender is female on the prenatal health investments in our total 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 statistically insignificant in almost all cases, suggesting that knowing that fetal gender is female is not a statistically significant determinant of prenatal care use, probability of tobacco use, the 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 Bonferroni adjustment[15] would reduce the statistical significance level of our analysis and 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 lbs lower when expecting a daughter, but this is very little relative to the mean weight gain of 30.7 lbs in our sample. In the birthweight equation the coefficient shows a statistically significant reduction of 3.1 grams, and this applies regardless of

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[14] Ai and Norton (2003) demonstrate 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 which contains the cross-partial derivatives.

[15] This adjustment makes it harder to reject the null hypothesis by making the threshold for statistical significance harder to attain. E.g. instead of using the conventional p-value of 0.05 the Bonferroni adjustment would use a p-value = 0.05/n when one model is estimated repeatedly for n separate outcomes.

marital status, age and race. Substantively the magnitude of the interaction effect on birthweight is small as the average birthweight in the U.S. is 3,329 grams and the standard deviation is 615 grams. Receiving an ultrasound and having a daughter is not a statistically significant determinant 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. A comparison of the unconditional results in Table 2 and the regression coefficients in Table 3 suggests that the results are almost identical-- controls seldom change the magnitude or statistical significance of the results.

Having an ultrasound is not a perfect indicator for knowing fetal gender which 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 23-39 years, 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 are likely to be attenuated towards zero among the groups of mothers who received an ultrasound but do not wish to ascertain fetal gender. Results available upon request showed that the coefficients on the interaction term appear to be the same for all sub-groups, which suggests that our findings are not being driven by measurement error.

Next 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 in Table 4. Overall, the results here too do not point to systematic differences by knowing 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 significantly different from zero. 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 magnitude of these effects are relatively small; although having an ultrasound and eventually having a daughter appears to reduce the probability of quitting smoking by 9.8% among all mothers, and by 14.3% among married mothers. Since stress has been associated with tobacco use these results may be interpreted as suggesting that perhaps mothers who know that they are going to have a daughter are more likely to feel anxious and hence less likely to quit prenatal tobacco use. However, the question about quitting smoking asks respondent if they quit smoking for at least one week after the pregnancy was confirmed. How the respondents interpreted this question is ambiguous as it could be interpreted as smoking behavior the week immediately after 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 they are not present for American mothers in general. The effect 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' education level. There were no notable differences to report, but tables are available upon request.   The lack of results is surprising because India and China have a long history of son preference.

We conduct three specification tests to test the validity of our method. The first specification test considers mothers' post-natal smoking decision.. Results (not included) 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 analysis of our second specification test relating to first trimester prenatal care, it is heartening to note that our analysis passes this test, i.e. no apparent effect of knowing the sex of the child in advance on whether prenatal care was initiated in the first trimester[16]. 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. We estimate coefficients on the interaction of ultrasound receipt during labor/delivery and having a daughter using California data. 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, the likelihood of low birthweight and very low birthweight. The only exception is the effect on birthweight, which implies that determining the sex of the child during labor has a positive and statistically significant impact on birthweight. However, the magnitude of this effect is negligible.

[16] Results (not reported) reveal that the analyses using the Indian and Chinese immigrant samples also pass the robustness test.

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 as 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 breastfeeding, there are no statistically significant differences in the postnatal investments of mothers who have sons versus those who have daughters. Since these are unconditional means we cannot take this as evidence of daughter preference, as mothers may breastfeed 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 breastfeeding, and insurance take-up are equal for mothers with girls versus boys.

One possible explanation for our finding that child's 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. 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 of gender preference in prenatal investments while it exists for marriage[17]. A third plausible explanation is measurement error. As discussed above, our classification of mothers into control and treatment groups is imperfect and

[17] In order to test the role of fathers' preferences in determining prenatal investments we estimated our models for same race immigrant couples, i.e. prenatal investment decisions when both parents are of Indian or Chinese origin. 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.

this error would bias the coefficients downwards. Although we provide some evidence to the contrary it remains a possibility.

The lack evidence of gender preference in prenatal behavior among immigrant women is worth discussing further because of the home-country findings.[18] 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 U.S. are innately different from their counterparts in India and China. For instance, in 1991 approximately 13 percent of women in India had more than primary education (Velkoff 1998) compared to the college completion rate of 49.6% in the Indian immigrant sample. Education is one dimension along which women who immigrate to the U.S. 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 U.S. 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 to 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 U.S. than in their home country, perhaps because of the economic and policy environment in

[18] Note that for the home countries, evidence of son preference has been found in the context of abortion and post – natal 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.

the U.S. Finally, we thank an anonymous referee for pointing out that the nature of physician intervention in the U.S. 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 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.[19] 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 paper 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, alcohol and tobacco use during pregnancy, i.e., 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.

[19] 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.

Using the 1989-2001 Natality Detail Files and 1988 NMIHS which provide information on live births in the U.S., 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, to 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 it accompanies differences in health investments as we measure it here. Admittedly, measurement error may have biased our results downwards 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.

Other than 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. For one, we study outcomes that are more a result of mother's actions than father's actions, and previous literature has shown that gender preferences are likely to be instigated by fathers. We further posit that the lack of result for Chinese and Indian first-generation women may be because women who immigrate to the U.S. 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 U.S. changes parents' decision-making problem such that gender-biased investments are no longer optimal. Finally, physician intervention in the U.S. might involve closer monitoring of fetal health or greater provision of prenatal health information so that mothers opt to make healthier decisions or 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 these explanations remain plausible since our data and hence our tests are not perfect.

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, a relationship that has not been studied to date. We find some 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. Thus the gender preference effect found in parental marriage, divorce, and fertility behavior in the U.S. does not appear to extend to prenatal health investments among mothers in the U.S.

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Table 1: Comparing Sex Ratios: Non-Immigrant vs. First-Generation Immigrant Populations, By Birth Order

	All Births	Indian Immigrant Births	Difference	Chinese Immigrant Births	Difference
	(1)	(2)	(1)-(2)	(3)	(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

***(**)(*) indicates statistical significance at the 0.01(0.05)(0.1) level
All Births refers to births to all non-immigrant mothers in the US

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

Means x 100

	Ultrasound			Non-Ultrasound			Difference-
	Girl	Boy	Diff.	Girl	Boy	Diff.	in-
ALL	48,792,749						
Sample size	14,504,968	15,232,693		9,316,750	9,738,338		
Prenatal care initiated in 1st trimester	82.44	82.22	0.22 ***	77.62	77.46	0.17 ***	0.06 **
Birthweight (grams)	3270.30	3388.00	-117.70 ***	3269.65	3383.93	-114.28 ***	-3.42 ***
Probability of low birthweight (< 2500 g)	8.03	6.90	1.13 ***	7.55	6.50	1.05 ***	0.08 ***
Probability 1 minute APGAR score < 7	8.19	9.09	-0.90 ***	7.91	8.80	-0.90 ***	-0.01
Probability 1 minute APGAR score < 4	2.06	2.45	-0.39 ***	2.01	2.39	-0.38 ***	-0.01
Probability 5 minute APGAR score < 7	1.29	1.53	-0.24 ***	1.33	1.57	-0.24 ***	0.00
Probability 5 minute 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 (pounds)	30.52	31.22	-0.70 ***	29.96	30.60	-0.64 ***	-0.06 ***
Pr. 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
Pr of inadequate PNC (Kessner Indx = Inad/Int)	24.27	24.59	-0.32 ***	30.21	30.45	-0.25 ***	-0.07 ***
Pr of inadequate PNC (Kotelchuck Indx=Inad/Int)	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 1st trimester	83.49	82.73	0.76 ***	82.03	81.26	0.77 **	-0.01
Birthweight (grams)	3104.11	3189.74	-85.63 ***	3104.98	3191.34	-86.36 ***	0.73
Probability of low birthweight (< 2500 g)	10.05	9.11	0.94 ***	9.28	8.48	0.81 ***	0.14
Probability 1 minute APGAR score < 7	5.01	5.89	-0.88	4.05	4.38	-0.33	-0.55
Probability 1 minute APGAR score < 4	1.23	1.37	-0.14	1.13	0.83	0.30	-0.44
Probability 5 minute APGAR score < 7	0.72	0.84	-0.12 *	0.72	0.85	-0.12	-0.001
Probability 5 minute 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 (pounds)	28.52	28.76	-0.24 ***	27.89	28.29	-0.40 ***	0.16
Pr. 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
Pr of inadequate PNC (Kessner Indx = Inad/Int)	25.57	26.48	-0.91 ***	26.35	27.45	-1.10 ***	0.18
Pr of inadequate PNC (Kotelchuck Indx=Inad/Int)	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 1st trimester	84.90	84.38	0.52 ***	86.66	86.12	0.55 ***	-0.02
Birthweight (grams)	3243.65	3349.49	-105.84 ***	3249.16	3349.19	-100.03 ***	-5.81
Probability of low birthweight (< 2500 g)	5.49	4.50	0.99 ***	5.00	4.31	0.70 ***	0.29 *
Probability 1 minute APGAR score < 7	4.51	5.09	-0.58 **	4.27	4.87	-0.60 **	0.01
Probability 1 minute APGAR score < 4	1.03	1.36	-0.34 **	1.25	1.10	0.14	-0.48 **
Probability 5 minute APGAR score < 7	0.63	0.76	-0.13 **	0.72	0.67	0.04	-0.17 *
Probability 5 minute 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 (pounds)	29.51	29.78	-0.27 ***	29.23	29.51	-0.28 ***	0.01
Pr. 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
Pr of inadequate PNC (Kessner Indx = Inad/Int)	23.39	23.98	-0.59 ***	20.81	21.28	-0.47 **	-0.13
Pr of inadequate PNC (Kotelchuck Indx=Inad/Int)	26.31	25.56	0.75 ***	25.21	24.55	0.66 ***	0.09

***(**)(*) indicates statistical significance at the 0.01(0.05)(0.1) level

PNC = prenatal care

The All sample represents all observations in the 1989-2001 Natality Detail Files.

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

OLS coefficient or probit interaction effects of ULTRASOUND*FEMALE (standard errors in parentheses)

	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 (pounds)	-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	-0.0019***	-0.0024***	-0.0014***	-0.0015***	-0.0026***
(<15 lb OR >40 lb)	(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/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	-0.0004	-0.0001	-0.0001	-0.0002	-0.0010*
PNC (Kessner Index)	(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	0.0002	0.00001	0.0003	0.00001	0.0004
PNC (Kotelchuck Index)	(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 (grams)	-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:

1. ***(**)(*) indicates statistical significance at 0.01(0.05)(0.1) level
2. 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.
3. The analysis uses a 15% random sample of the 1989-2001 Natality Detail Files
4. 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)

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

OLS coefficient or probit interaction effects of ULTRASOUND*FEMALE (standard errors in parentheses)

	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)
	7931	4732	3199	1444	1492
Prenatal weight gain (pounds)	1.576*	1.753*	1.252	-1.760	2.385
	(0.95)	(1.04)	(2.13)	(2.16)	(1.96)
	9080	5449	3631	1659	1707
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)
	6373	3862	2509	1124	1161
Probability of inadequate weight gain (5lb less/more than advised)	0.013	0.008	0.002	-0.011	-0.072
	(0.034)	(0.041)	(0.063)	(0.079)	(0.078)
	6373	3862	2509	1124	1161
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)
	2684	1429	1251	522	532
Birthweight	25.532	23.861	39.334	89.339	-29.152
	(27.862)	(32.512)	(52.986)	(67.698)	(57.841)
	9073	5444	3629	1658	1705

Notes:
1. ***(**)(*) indicates statistical significance at 0.01(0.05)(0.1) level
2. 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.

3. The analysis is restricted to pregnancies that resulted in a live birth in the 1988 National Maternal and Infant Health Survey.
4. 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)

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

OLS coefficient or probit interaction effects of ULTRASOUND*FEMALE (standard errors in parentheses)

	All Births	First Births	Higher Order Births
Prenatal Investments			
Number of PNC 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 Care (Kotelchuck Index)	-0.0022	-0.0002	-0.0031
	(0.0049)	(0.0068)	(0.0070)
Birth Outcomes			
Birthweight (grams)	-0.4484	0.5269	-0.0627
	(5.5718)	(7.9395)	(7.8315)
	148,977	73,361	75,616
Probability of low birthweight (<2500 grams)	0.0003	0.0068	-0.0062
	(0.0029)	(0.0044)	(0.0038)
Probability of very low birthweight (<1500 grams)	0.0018*	0.0007	0.0027*
	(0.0010)	(0.0015)	(0.0014)

Notes:

1. ***(**)(*) indicates statistical significance at 0.01(0.05)(0.1) level
2. PNC = prenatal care
3. 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.
4. The analysis uses the universe of children born to Indian immigrant mothers in the 1992-2001 Natality Detail Files
5. 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)

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

OLS coefficient or probit interaction effects of ULTRASOUND*FEMALE (standard errors in parentheses)

	All Births	First Births	Higher Order Births
Prenatal Investments			
Number of PNC 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 Care (Kotelchuck Index)	0.0006	0.0015	-0.0002
	(0.0032)	(0.0045)	(0.0047)
Birth Outcomes			
Birthweight (grams)	-4.9716	-8.7292*	-0.9646
	(3.6293)	(5.1028)	(5.1641)
	295,724	150,435	145,289
Probability of low birthweight (<2500 grams)	0.0016	0.0030	0.0001
	(0.0014)	(0.0021)	(0.0019)
Probability of very low birthweight (<1500 grams)	-0.0004	-0.0001	-0.0007
	(0.0005)	(0.0007)	(0.0007)

Notes:

1. ***(**)(*) indicates statistical significance at 0.01(0.05)(0.1) level
2. PNC = prenatal care
3. 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.
4. The analysis uses the universe of children born to Chinese immigrant mothers in the 1989-2001 Natality Detail Files
5. 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)

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

Means x 100 (standard errors in parentheses)

	Girls		Boys		Difference
	Mean	**(s.e.)**	**Mean**	**(s.e.)**	
Number of Wellchild Visits	42.572	(0.345)	43.005	(0.348)	-0.432
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, 1st month	4.782	(0.076)	4.728	(0.077)	0.054
Intensity of breast feeding, 2nd month	3.802	(0.072)	3.683	(0.072)	0.119
Intensity of breast feeding, 3rd month	2.883	(0.064)	2.813	(0.065)	0.070
Intensity of breast feeding, 4th month	2.212	(0.056)	2.142	(0.058)	0.070
Intensity of breast feeding, 5th month	1.803	(0.051)	1.745	(0.053)	0.058
Intensity of breast feeding, 6th month	1.534	(0.048)	1.478	(0.049)	0.056
Insurance Status, post-natal	68.691	(0.695)	70.004	(0.680)	-1.313
Any Immunization	97.183	(0.247)	97.291	(0.240)	-0.108

***(**)(*) indicates statistical significance at 0.01(0.05)(0.1) level