Are there missing girls in the United States? Evidence on gender preference and gender selection*

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ABSTRACT

Gender selection, manifested by unusually high percentages of male births, has spread in parts of Asia since the introduction of ultrasound technology. This paper provides the first empirical evidence consistent with the occurrence of gender selection within the United States. Based upon fertility-stopping behavior, the aggregate gender preferences among different races in the United States are documented. Analysis of comprehensive birth data shows unusually high boy-birth percentages after 1980 among later children (most notably third and fourth children) born to Chinese and Asian Indian mothers. Moreover, Asian Indian mothers are found to be significantly more likely both to have a terminated pregnancy and to give birth to a son when they have previously only given birth to girls. These findings are consistent with a simple dynamic model of the gender-selection decision in the presence of gender preferences.

*The California natality data used in this paper can not be released due to a confidentiality agreement with the California Department of Health Services (CDHS). The author is grateful to Jan Christensen, Karl Halfman, and Roxana Killian of the CDHS for their assistance during the data-acquisition process. The federal natality data and Census data used in this paper were obtained from the Inter-University Consortium for Political and Social Research (ICPSR). David Hummels provided helpful comments, and Jack Barron provided invaluable computer assistance. Dudley Poston, Jr. kindly provided data on Chinese and South Korean male-to-female birth ratios. This research was partly supported by a University Faculty Scholar grant through Purdue University.

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It is not possible to assess how popular sex-determination tests and gender-selection techniques might be among Indian-Americans or any other group. There are no official statistics, and people who wish to choose the sex of their child do not wish to discuss it publicly...

(*New York Times* article, Aug. 15, 2001)

1 Introduction

Amartya Sen (1990, 1992) coined the term "missing women" to illustrate differential mortality rates experienced by women in several Asian countries. Sen (1990, 1992) has estimated that there are approximately 80–100 "missing women" in Asia, and he has pointed to gender selection as one contributing factor:

> Given a preference for boys over girls that many male-dominated societies have, gender inequality can manifest itself in the form of the parents' wanting the new born to be a boy rather than a girl. There was a time when this could be no more than a wish (a daydream or a nightmare, depending on one's perspective), but with the availability of modern techniques to determine the gender of the fetus, sex-selective abortion has become common in many countries. It is particularly prevalent in East Asia, in China and South Korea in particular, but also in Singapore and Taiwan, and it is beginning to emerge as a statistically significant phenomenon in India and South Asia as well. (Sen (2001))

The existing evidence on gender-selective abortion in Asia is primarily indirect, based upon unusually high percentages of boys being born.[1] In particular, several Asian countries, including China, India, South Korea, and Taiwan, have seen significant increases in the percentages of boys at birth since the 1970's and 1980's, when ultrasound technology (and to a lesser extent amniocentesis technology) became available and affordable to women (see, for example, Poston, Wu, and Kim (2003), Retherford and Roy (2003), and Poston and Glover (forthcoming)). To illustrate these trends, Figure 1 provides a plot of boy-percentages-at-birth for China, South Korea, India, and the United States.[2] Whereas the likelihood of a male birth has remained at just above 51% in the United States since 1980, the percentage of male births has increased to around 53% in China, India, and South Korea.

[1] Direct evidence would require data that could be used to relate voluntary pregnancy terminations to fetus gender.

[2] To smooth the plot somewhat, three-year moving averages are plotted at each year. Sources: China and South Korea data are from Poston and Glover (forthcoming); India data are from Office of the Registrar General of India (2001); United States data are from the federal natality data described more fully in Section 3.

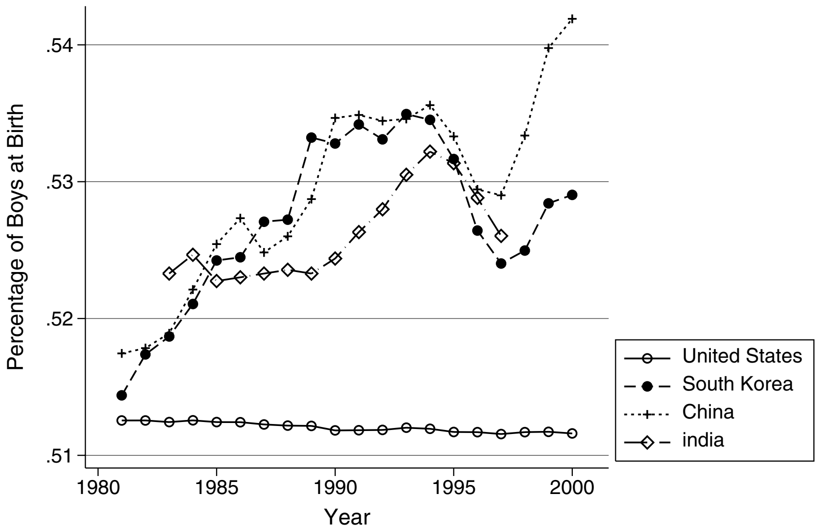

Figure 1: Likelihood of a Male Birth, by Country

Recent research has pointed to more subtle forms of gender bias (specifically, bias favoring sons) in the United States. For instance, Lundberg and Rose (2003) find that single mothers are more likely to marry a child's biological father if the child is a boy. Meanwhile, Dahl and Moretti (2004) find that parents with sons are less likely to be divorced and that divorced fathers are more likely to have custody of their sons. In the conclusion to their study, Dahl and Moretti (2004) point out that gender bias in the United States "does not take the extreme form of 'missing' girls like in some Asian countries..." And, certainly, the boy-percentage trend for the United States in Figure 1 doesn't provide evidence of gender-selective practices in the aggregate.

However, evidence for gender selection may exist at a more disaggregated level. This paper analyzes data on births in the United States, broken down by race, to determine whether evidence for gender selection exists for specific racial groups. One might suspect, for instance, that those races associated with the Asian countries in Figure 1 (Chinese, Indian, Korean) would be more likely to practice gender selection due to cultural biases.[3] This idea has been suggested by others, including Robertson (2001): "Until they are more fully assimilated, immigrant groups in Western countries may retain the same gender preferences that they would have held in their homelands." As anecdotal evidence to this point, a recent *New York Times* article (Sachs (2001)) described efforts by several companies to directly market gender identification and pre-conceptive selection products to Indian expatriates in North America:

"Desire a Son?" asked an advertisement in recent editions of *India Abroad*, a weekly

[3]The term "Indian" will be used to mean "Asian Indian" (rather than "American Indian") throughout this paper.

newspaper for Indian expatriates in the United States and Canada. "Choosing the sex of your baby: new scientific reality," declared another in the same publication. A third ad ran in both *India Abroad* and the North American edition of *The Indian Express*. "Pregnant?" it said. "Wanna know the gender of your baby right now?"

The incentives for gender selection depend not only on gender preferences but also upon family size (i.e., number of children already born). As an extreme example, the "One Child Policy" in China created a strong incentive for *first-birth* gender selection due to a cultural bias toward having sons. Even in the absence of exogenous family-size limits, however, gender-selection incentives (in the presence of gender bias) would likely be stronger as a family approaches its own size limit. For instance, consider a family that has a strong preference for having at least one son and is willing to have at most two children. If the first child is a boy, this family might stop having children; if the first child is a girl, the family would have another child and a greater incentive (than in the first pregnancy) to determine gender and, perhaps, undertake a gender-selective procedure. If there were many such families, the data in the aggregate would indicate a higher percentage of boys among second births (as compared to first births) due to the combination of fertility stopping (by families with first-born sons) and gender determination/selection (by families with first-born daughters).[4] This argument suggests that any evidence of unusual gender percentages at birth is most likely to be found at later births, an issue that is examined empirically in this paper.

Given the possible link between gender-determination incentives and family size, it is important to understand recent fertility trends. At the same time that gender-determination technology has become available, fertility rates have dropped in Southeast Asia over the past few decades (see, e.g., Hirschman (2001)). The average number of children per family has also dropped, mirroring a pattern that has occurred in Europe and the United States over the same time period. For the United States, Figure 2 shows the 1980–2000 trend in the fertility rate among whites, blacks, and Asians.[5] Since 1990, the fertility rates for each of the racial groups has declined, with the fertility rate in 2000 for blacks and Asians significantly lower than it was in 1980. To illustrate trends in family sizes across different races, Figure 3 plots the average birth parity for all births occurring in the United States.[6] (Birth parity refers to the birth-order of the child. First child has a parity of 1, second child has a parity of 2, and so on.) The figure provides a breakdown of Asians into Chinese, Indian, Japanese, and Korean (with the Indian and Korean data available beginning in 1992). For

[4] The fertility stopping by itself has no impact on boy-birth percentages, but the sample of families having second children will be over-represented by those families with first-born daughters. As such, the second-born boy-birth percentage would be even higher than it would have been if all families with first-born sons had also had a second child.

[5] Source: Centers for Disease Control and Prevention (2001, Table 1-7). The fertility rate is defined as the births per 1,000 females between the ages of 15 and 44. Data for Asians is not provided prior to 1980.

[6] Source: United States federal natality data.

all races, the average birth parity has declined since the early 1970's (partly due to the legalization of abortion in 1973). Among Chinese and Japanese mothers, the average birth parity has shown an additional slow decline since the mid-1970's. The data available for Indian and Korean mothers indicates an average birth-parity level very similar to that for Chinese and Japanese mothers.

If parents wish to consider gender selection of their babies, there currently exist three different options in the United States: (1) gender-selective abortion, (2) gender-selective in vitro fertilization (IVF), or (3) sperm sorting. An important distinction is that the latter two options are performed *prior* to pregnancy. Gender-selective IVF is a modified version of the traditional IVF procedure, in which fertilized embryos are transferred into the mother's uterus. For gender-selective IVF, however, embryos are genetically tested ("preimplantation genetic diagnosis") to determine gender and chosen accordingly. Such testing is nearly 100% accurate for gender determination and, when done for gender reasons only (rather than avoiding a genetic disease), has been banned in many countries. Although a very effective means of gender selection, the IVF procedure is very expensive (between $10,000 and $20,000 per implantation cycle). Sperm sorting, on the other hand, is far less expensive (costing a few thousand dollars) but not quite as effective. The procedure involves selecting sperm from a given sperm sample in order to increase the probability of the desired gender when the egg is fertilized. One company that offers sperm sorting in the United States (Microsort) claims a success rate of 91% (295 out of 325) for couples who desired a girl and 73% (39 out of 51) for couples who desired a boy.[7] Although both gender-selective IVF and sperm sorting may be options for gender selection, these two procedures would likely only account for a very small proportion of the gender-selective procedures that might have occurred in the United States in the past few decades. The reasons for this include their recent introduction, their high expense, and the limited number of doctors willing to perform such procedures. As such, both the theoretical model and the empirical investigation of this paper will focus primarily on abortion as the means for gender selection. On the other hand, when thinking about the future of gender selection, these more advanced technologies will no doubt play a larger role.

Turning to gender-selective abortion, the introduction of ultrasound and amniocentesis in the 1970's made such a procedure a possibility. Although neither technology was introduced for the explicit purpose of determining the gender of a fetus, both technologies are capable of this determination during the first half of pregnancy. Amniocentesis, generally performed between the 14th and 18th weeks of pregnancy, is nearly 100% accurate in determining gender but has a small risk (0.5–1.0%) of miscarriage associated with it. (Amniocentesis involves insertion of a needle into the mother's uterus, with a small amount of amniotic fluid removed and then analyzed.) Ultrasound,

[7] These success rates were reported on the company's website (`www.microsort.com`) for pregnancies through January 2004. Scientific evidence of the technology's effectiveness has existed for more than a decade (e.g., Johnson et. al. (1993)).

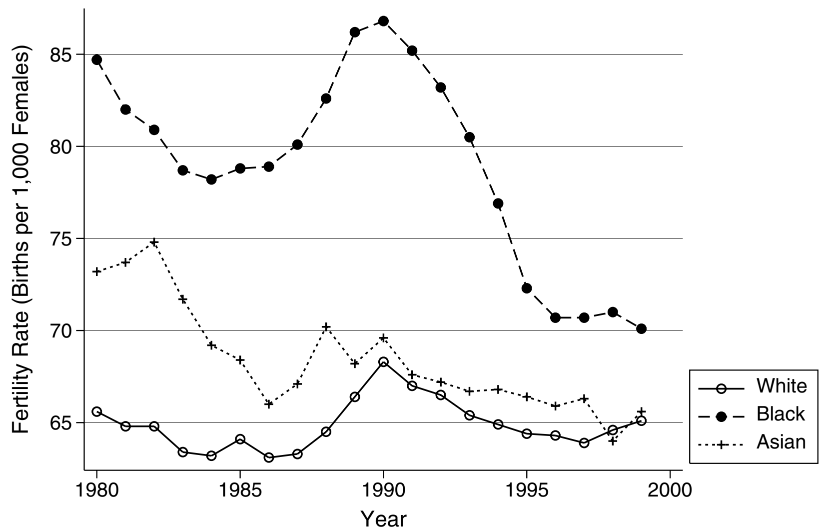

Figure 2: Fertility Rates in the United States, by Race

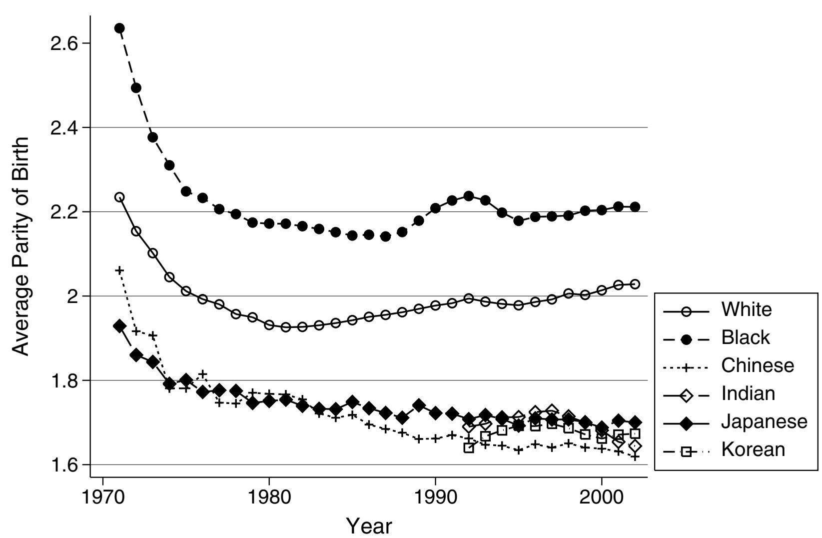

Figure 3: Average Birth Parity, by Race

Table 1: Summary Statistics on Abortion in the United States

	1980	1990	2000
Reported # of legal abortions	1,297,606	1,429,247	857,475
Weeks of gestation:			
8 weeks or less	51.7%	51.6%	58.1%
9–10	26.2%	25.3%	19.8%
11–12	12.2%	11.7%	10.2%
13–15	5.1%	6.4%	6.2%
16–20	3.9%	4.0%	4.3%
21 weeks or more	0.9%	1.0%	1.4%
Previous live births:			
Zero	58.4%	46.2%	40.0%
One	19.4%	25.9%	27.7%
Two or more	22.2%	27.9%	32.3%

which can usually be used to detect gender between the 16th and 20th weeks of pregnancy, is safer than amniocentesis but is somewhat less accurate in gender determination.[8][9] If either ultrasound or amniocentesis were used as a precursor to gender-selective abortion, the abortion would most likely occur during the second trimester of pregnancy. Although most abortions in the United States occur prior to the second trimester, there are a large number of abortions that do occur during the second trimester and later. Table 1 provides some summary statistics on abortions in the United States in 1980, 1990, and 2000, as reported by the Centers for Disease Control and Prevention (2003). Since 1980, roughly 5% of abortions have occurred at 16 weeks or later. These figures, of course, do not represent evidence of gender selection; they merely indicate that a non-negligible fraction of abortions do occur after the point that gender determination is possible. Another interesting fact from Table 1 is that a large percentage of abortions are associated with women who have previously had live births (41.6% in 1980, 54.8% in 1990, and 60.0% in 2000).

The outline of the paper is as follows. Section 2 presents a simple dynamic model of gender selection in the presence of gender preferences (including gender bias or gender-mix preference). Section 3 describes the different data sources (Census data, federal natality data, and California

[8] Chorionic villus sampling (CVS) can also be used for gender determination. CVS is performed at 10–13 weeks and is nearly 100% accurate. However, CVS carries a greater risk of fetal loss than amniocentesis and is rarely performed in the United States. For example, the use of CVS during pregnancy was reported for only 0.1% of births in California between 2000 and 2003.

[9] Non-invasive prenatal DNA testing has also recently been introduced as a method for gender determination. This method requires only a blood sample, has no miscarriage risk, and can be used earlier than amniocentesis. One product, the *Baby Gender Mentor Home DNA Gender Testing Kit*, sells for less than $300 on `www.pregnancystore.com`; the claimed accuracy for this test is 99.9% as early as *five* weeks after conception, but no scientific evidence is yet available to verify this accuracy.

natality data) used in the empirical analysis. Section 4 reports the empirical results. Wherever possible, results are reported separately for the following racial groups: whites, blacks, Chinese, Indian, Japanese, and Korean.[10] First, we examine evidence on aggregate gender (and gender-mix) preferences from Census data, focusing on the decision to have a second and/or third child based upon the gender of previous children. Second, we analyze both federal natality data and California natality data to determine which factors are associated with a baby's gender. Interestingly, the natality data is subject to a "survival bias" since boys have more difficulty surviving pregnancy than girls. As a result, several variables that proxy for the quality of prenatal care and/or difficulty of the pregnancy (including education and month of first prenatal visit) appear as statistically significant determinants of baby gender.[11] The statistical analysis on births after 1980, both with and without controls, indicates that Chinese and Indian mothers are significantly more likely to have sons at higher birth parities (third and fourth children) than for their first child. Third, we analyze a maternally linked version of the California natality data. This version allows us to condition upon the gender of a mother's previous children and to determine whether a variety of variables (current baby's gender, use of ultrasound, use of amniocentesis, and terminated pregnancies) are systematically related to previous children's gender. The analysis suggests that Asian Indian mothers are significantly more likely both to have a terminated pregnancy and to give birth to a son when they have previously only given birth to girls. Fourth, we use a simple framework (similar to the model of Section 2) to infer the prevalence of gender selection from unusual boy-birth percentages. Finally, Section 5 concludes.

2 Model of gender selection

This section presents a simple theoretical model for the gender-determination and gender-selection decisions. The model is intentionally stylized in an effort to focus upon the specific issues of gender determination and gender selection and abstract away from other issues such as unwanted pregnancies, fertility spacing, abortions for non-gender reasons, etc. These latter issues have been considered by other researchers and are nicely exposited in the recent book by Levine (2004). A similar dynamic model to the one presented here has been considered by Fajnzylber, Hotz, and Sanders (2002), who model the optimal (expected-utility-maximizing) use of amniocentesis in a context without gender determination. Dahl and Moretti (2004) provide a model of fertility decisions in the presence of gender preferences but without the possibility of gender selection.

[10] Results for other racial groups (with the largest being American Indian, Vietnamese, and Filipino) are available from the author.

[11] For this same reason, the use of ultrasound and amniocentesis are found to have statistically significant associations with baby gender, but these associations should not be taken as evidence on gender selection since the primary uses of these procedures do not have a gender-selective intent. This point is discussed further in Section 4.

2.1 The model

A woman is assumed to have a finite number (T) of fertility periods, indexed $t = 1, \ldots, T$. Figure 4 depicts a single fertility period t. (The word "woman" is used to simplify exposition, though this model could also apply to a "couple.") In a given fertility period t, the woman decides whether or not to become pregnant. If she decides to become pregnant, the woman decides whether or not to determine the gender of the fetus (with the cost of the gender-determination procedure denoted d). If gender is not determined, a boy is born with probability p and a girl with probability $1-p$. If gender is determined, there is a probability q that the gender-determination procedure will result in an involuntary termination of the pregnancy. After gender is revealed (probability of a boy again being p), the woman decides whether or not to terminate the pregnancy. The cost associated with a terminated pregnancy (whether involuntary or voluntary) is denoted c.

Each black dot in the tree from Figure 4 indicates a point at which the woman makes a decision. There are three possible decisions: (1) the pregnancy decision, (2) the gender-determination decision, and (3) the termination decision (after gender is revealed). "Nature" plays a role in two places in the tree: (1) the possibility of an involuntary termination (probability q) resulting from the gender-determination procedure, and (2) the realization of gender (probability p of a boy).

Some additional notation is required to model the utility function. Let the "state variables" n_b and n_g denote the number of sons and daughters, respectively, that a woman already has at a given point in time (e.g., $n_b = n_g = 0$ for no children).[12] Assume that a woman maximizes expected utility, and let $V_t(n_b, n_g)$ denote the expected utility at the outset of fertility period t for a woman with n_b sons and n_g daughters. Let $U_b(n_b, n_g)$ and $U_g(n_b, n_g)$ denote the incremental utilities associated with having a boy and girl, respectively. (Note that these incremental utility functions are assumed to depend only on the existing gender mix and are independent of t.) The incremental utility associated with no pregnancy is normalized to be equal to zero. The discount rate is given by δ, where it is assumed that $\delta < 1$. The finite fertility horizon (T periods) implies that

$$V_t(n_b, n_g) = 0 \text{ for } t \geq T + 1 \tag{1}$$

since no pregnancies occur after period T. Looking again at Figure 4, the realized (expected) utility is given at the end of each possible path in the decision tree. For example, if a boy is born after gender determination, the realized utility is $U_b(n_b, n_g) + \delta V_{t+1}(n_b + 1, n_g) - d$.

[12]The dependence of n_b and n_g on t (i.e., n_{bt} and n_{gt}) is suppressed to simplify notation.

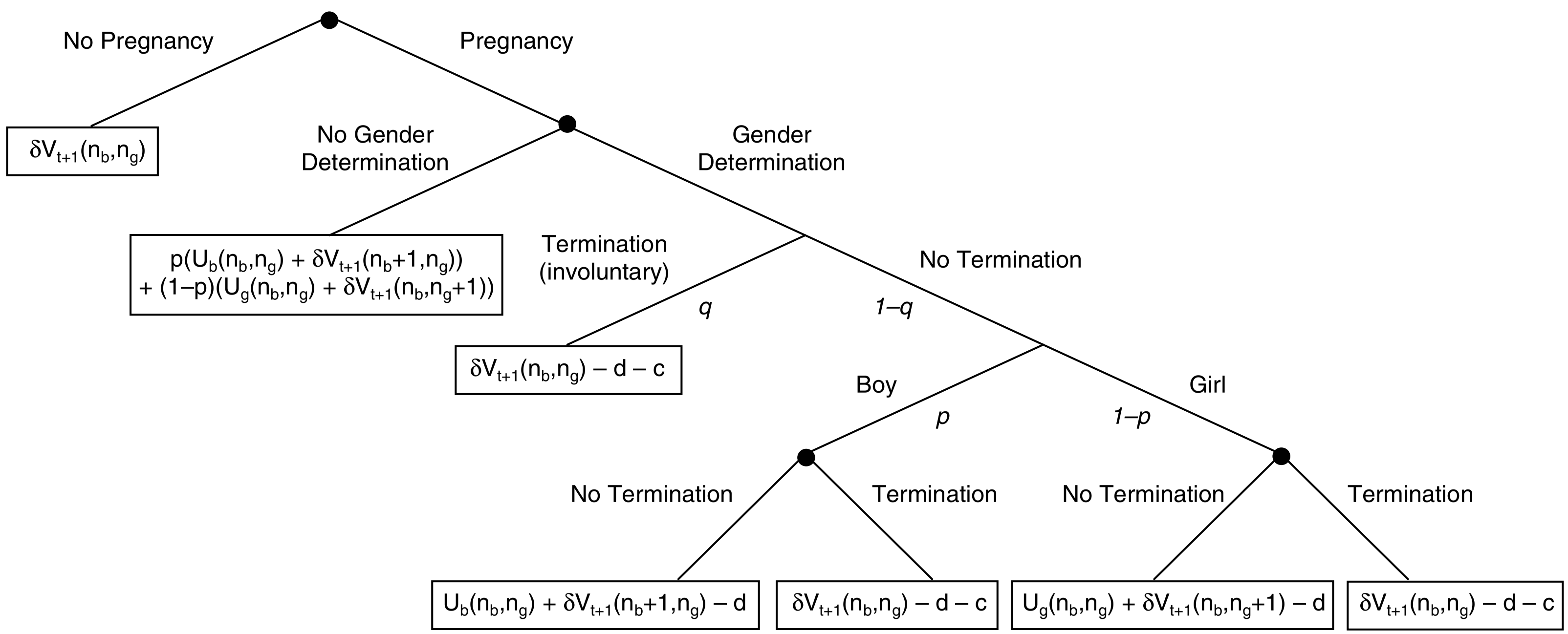

Figure 4: Gender-Selection Decision Tree

The only assumption made with respect to the incremental utility functions is that incremental utility (to either a son or daughter) does not increase as children are added to the family:[13]

$$U_b(n_b, n_g) \geq U_b(n_b + 1, n_g),\ U_b(n_b, n_g) \geq U_b(n_b, n_g + 1),$$

$$U_g(n_b, n_g) \geq U_g(n_b + 1, n_g),\ U_g(n_b, n_g) \geq U_g(n_b, n_g + 1).$$

If an incremental utility $U_b(n_b, n_g)$ ($U_g(n_b, n_g)$) is negative, the woman would prefer no pregnancy to a pregnancy that would result in a boy (girl) baby with certainty. The relative values of $U_b(n_b, n_g)$ and $U_g(n_b, n_g)$ indicate gender preference, with $U_b(n_b, n_g) > U_g(n_b, n_g)$ indicating a boy preference and $U_b(n_b, n_g) < U_g(n_b, n_g)$ indicating a girl preference (conditional on n_b and n_g). The following examples show how the values of the incremental utilities relate to family-size and gender-composition preferences:

Example 1 *(Two-child preference, no gender preferences) $U_b(n_b, n_g) = U_g(n_b, n_g)$ for all n_b and n_g, $U_b(n_b, n_g) < 0$ and $U_g(n_b, n_g) < 0$ if $n_b + n_g \geq 2$*

In this example, the incremental utilities for a boy and girl are equal regardless of the gender composition of existing children. The family-size preference (two children here) is determined by the point at which these incremental utilities become negative.

Example 2 *(Two-child preference, gender-mix bias) $U_b(0,1) > U_g(0,1)$, $U_g(1,0) > U_b(1,0)$, $U_b(n_b, n_g) < 0$ and $U_g(n_b, n_g) < 0$ if $n_b + n_g \geq 2$*

These incremental utilities indicate a preference for at most two children, with a bias toward having a gender mix. In this example, the woman would stop having children after two births even if a gender mix is not achieved.

Example 3 *(Three-child possibility, gender-mix bias) $U_b(0,2) > 0 > U_g(0,2)$, $U_g(2,0) > 0 > U_b(2,0)$, $U_b(1,1) < 0$, $U_g(1,1) < 0$*

These incremental utilities, like those in Example 2, indicate a preference for a gender mix. In this example, however, the woman has positive incremental utility associated with achieving a gender mix on the third child after having two children of the same gender. If a gender mix had been achieved with the first two children, the woman would not become pregnant again. In addition, if the first two children were the same gender, the woman would not become pregnant again if she knew with certainty that the third child would also be the same gender.

[13]The assumption does rule out certain forms of complementarity among children. For instance, after having one son, a mother might have a higher incremental utility for a second son if she values the fact that they will play together as children and grow up to be very close friends.

Example 4 *(One-child preference, boy bias)* $U_b(0,0) > U_g(0,0)$*,* $U_b(n_b, n_g) < 0$ *and* $U_g(n_b, n_g) < 0$ *if* $n_b + n_g \geq 1$

These incremental utilities indicate a preference for at most one child, with a bias toward having a boy. Although our model will take incremental utilities as exogenously given, the distaste for multiple children could arise for a variety of reasons: limited parental resources, the woman's desire to participate in the labor force, externally imposed fertility restrictions (such as China's "One Child Policy"), etc.

Although the model of incremental utilities is quite simple, it is useful to provide a graphical view of different types of gender preferences. Four different situations are depicted in Figure 5, each with incremental utilities plotted (with U_b on the x-axis and U_g on the y-axis) for $n_b + n_g \leq 2$. The arrows indicate the change that occurs after a child is born. A 45-degree line (drawn at $U_b = U_g$) is shown to clarify instances of gender preference. The first three panels in Figure 5 represent specific cases of Examples 1, 2, and 3, respectively. Panel 1 shows gender indifference, with the incremental utility values depending only upon total number of children but not the gender mix. All points lie exactly on the 45-degree line, with the incremental utility values becoming negative after two children. Panels 2 and 3 both show a gender-mix preference, with the primary difference being the number of children desired in the two situations (at most two in Panel 2 and possibly three in Panel 3). To focus on the preference for gender mix, these two situations exhibit a symmetry with respect to gender (note the symmetry about the 45-degree line). There is no intrinsic preference of one gender over the other, but a given gender may eventually be preferred if it would lead to a gender mix. Lastly, Panel 4 shows a son-preference situation. Unlike Example 4, however, this situation does not reflect a one-child fertility limit. There is a strong initial son bias (at $n_b = n_g = 0$), and the girl-birth occurrences do not reduce the incremental utilities of sons at all (whereas boy-birth occurrences do reduce the incremental utilities of daughters).

Obviously, many other different types of gender preferences could be graphically depicted as in Figure 5. The only restriction is given by our assumption that incremental utility values are weakly decreasing. In terms of the incremental-utility plot, this assumption simply implies that, as births occur, the points must (weakly) move downward and leftward. To clearly see the connection between gender preferences and gender selection, the various decision regions for gender determination and/or gender selection will also be shown with the incremental-utility axes (see Figures 6 and 7).

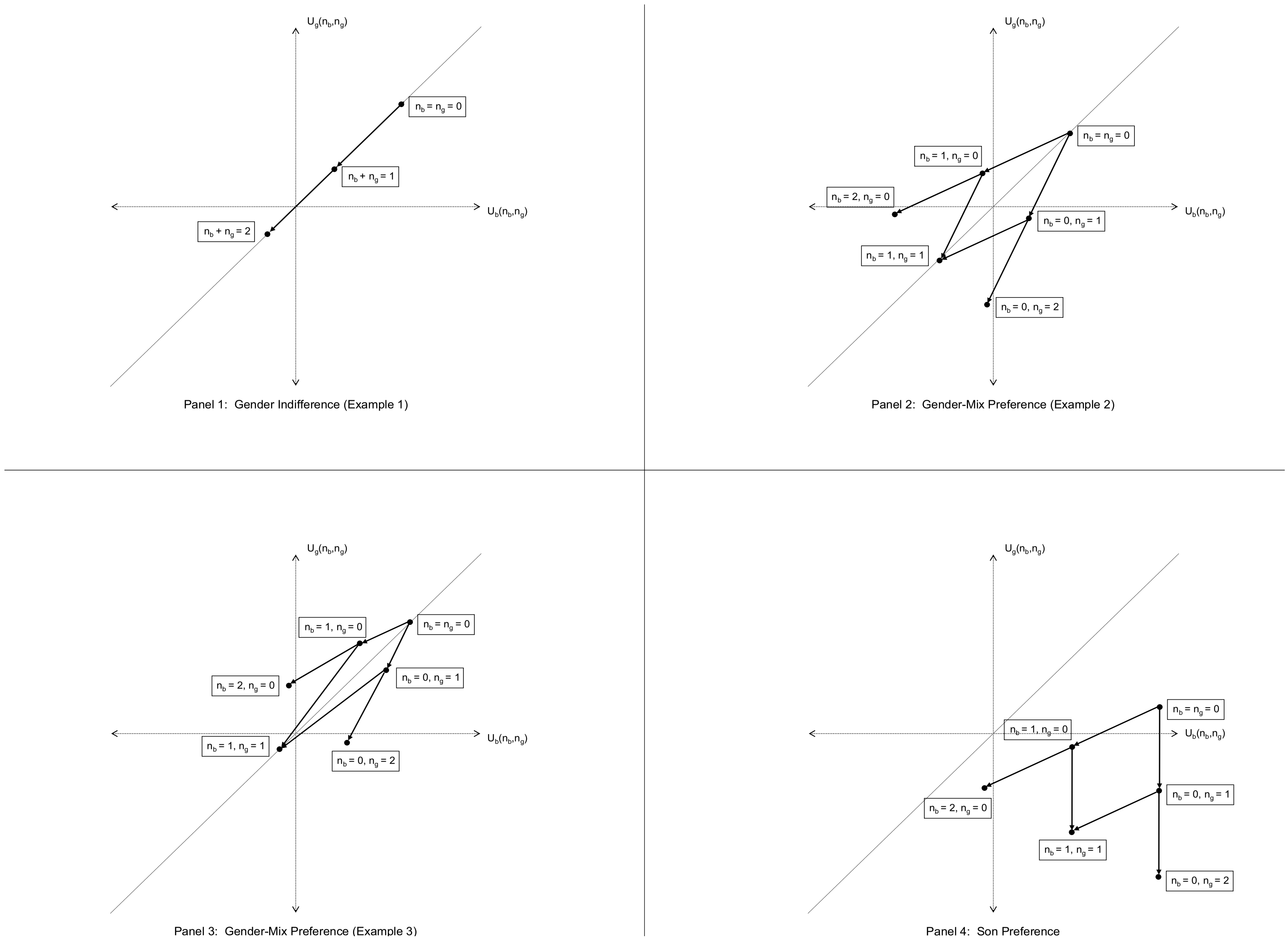

Figure 5: Graphical Representation of Gender Preferences

Before going into the details of a woman's optimal decisions, we briefly note some simplifications implicitly assumed in the model: (1) there are no unintended pregnancies, (2) a woman can become pregnant with certainty, (3) the gender-determination procedure is perfectly accurate, and (4) there is only one type of gender-determination procedure. The first three simplifications could be addressed quite easily by including additional probability parameters at the pregnancy-decision and/or the gender-determination stages. More substantial modification of the model would be required to address the fourth simplification. To fix ideas, assume that a woman can choose between two gender-determination procedures, ultrasound or amniocentesis. Ultrasound involves no risk of involuntary termination ($q = 0$) but has a higher cost associated with voluntary termination (high c) since it is performed later in the pregnancy than amniocentesis. On the other hand, amniocentesis involves a risk of involuntary termination (with q roughly between 0.5% and 1.0%) but has a lower cost associated with voluntary termination.[14] The tradeoff between q and c will determine a woman's choice between the two procedures (if either is considered a better option than no gender determination). Since the primary concern of this paper is to examine the incidence of gender determination (rather than the *form* of gender determination), the model considers only a single form of gender determination. The parameters c, d, and q describe the gender-determination procedure, making the model flexible enough to incorporate either ultrasound or amniocentesis, but the simplification is that the same procedure is considered by the woman in each fertility period.

2.2 Baseline case: gender determination not available

Consider the simpler (baseline) case where a gender-determination procedure is unavailable. In terms of the model above, lack of availability would be equivalent to having the cost $d = \infty$. (Note that $d = \infty$ could also arise if the procedure were available but the woman does not consider it an option because of moral considerations.) In this case, the woman's optimal decisions are straightforward. In each period, she will become pregnant if the expected utility from doing so is positive. The pregnancies will begin immediately since it is costly to wait ($\delta < 1$). The following proposition formally states this result:

Proposition 1 *If $d = \infty$, expected utility is maximized by becoming pregnant in period t (for $t \in \{1, \ldots, T\}$) if and only if $pU_b(n_b, n_g) + (1 - p)U_g(n_b, n_g) > 0$.*

The proof is provided in Appendix A, along with proofs of the other propositions in this section.

[14] Ultrasound may also be cheaper (lower d) and less reliable than amniocentesis.

2.3 The last-period decision with gender determination

To determine a woman's optimal decisions in the last period T, one can work from the bottom of the decision tree (see Figure 4) to the top. First, the termination decision (conditional on having determined gender) is described by the following lemma:

Lemma 1 *(Termination decision in period T) If the gender is known to be male, then the pregnancy will be terminated if and only if $U_b(n_b, n_g) < -c$. If the gender is known to be female, then the pregnancy will be terminated if and only if $U_g(n_b, n_g) < -c$.*

Termination is chosen only when the disutility of having a child (of known gender) outweighs the cost associated with termination. Next, given the termination decision described in Lemma 1, the optimal decision of whether or not to determine gender is given by the following lemma:

Lemma 2 *(Gender-determination decision in period T)*

(i) If $U_b(n_b, n_g) > -c$ and $U_g(n_b, n_g) > -c$, the woman will not determine gender;

(ii) if $U_b(n_b, n_g) > -c$ and $U_g(n_b, n_g) < -c$, the woman will determine gender if and only if

$$pqU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) < -d - qc - (1-q)(1-p)c; \tag{2}$$

(iii) if $U_b(n_b, n_g) < -c$ and $U_g(n_b, n_g) > -c$, the woman will determine gender if and only if

$$pU_b(n_b, n_g) + (1-p)qU_g(n_b, n_g) < -d - qc - (1-q)pc; \tag{3}$$

(iv) if $U_b(n_b, n_g) < -c$ and $U_g(n_b, n_g) < -c$, the woman will determine gender if and only if

$$pU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) < -d - c. \tag{4}$$

Case (i) involves incremental utilities for which termination would not be chosen; as such, gender would not be determined since it is costly ($d > 0$). Case (ii) involves incremental utilities for which a termination would be performed only if the gender were determined to be female. The woman determines gender if the disutility of having a daughter is sufficiently large to outweigh the cost of the procedure (d), the cost of a potential voluntary termination ($(1-q)(1-p)c$), and the cost of a potential involuntary termination (qc plus the lost incremental utility for a male child $-qpU_b(n_b, n_g)$). If there is no chance of involuntary termination ($q = 0$), note that the condition in case (ii) would simplify to $(1-p)U_g(n_b, n_g) < -d - (1-p)c$, in which case gender determination would be more likely than with $q > 0$. Case (iii) is similar to case (ii) with the role of sons and daughters interchanged. Case (iv) turns out to be negligible since a woman with both incremental utilities below $-c$ would not choose to become pregnant.

Finally, given the results of Lemma 1 and Lemma 2, the optimal decisions regarding both pregnancy and gender determination in period T are characterized as follows:

Proposition 2 *(Pregnancy and gender-determination decisions in period T)*
A woman will become pregnant and not determine gender if

$$pU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) > 0, \tag{5}$$

$$qpU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) > -d - (1-(1-q)p)c, \tag{6}$$

$$\textit{and}\;\; pU_b(n_b, n_g) + q(1-p)U_g(n_b, n_g) > -d - (q+(1-q)p)c. \tag{7}$$

A woman will become pregnant and determine gender if either

$$U_b(n_b, n_g) > \frac{d + (1-(1-q)p)c}{(1-q)p} \tag{8}$$

$$\textit{and}\;\; qpU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) < -d - (1-(1-q)p)c, \tag{9}$$

or

$$U_g(n_b, n_g) > \frac{d + (q+(1-q)p)c}{(1-q)(1-p)} \tag{10}$$

$$\textit{and}\;\; pU_b(n_b, n_g) + q(1-p)U_g(n_b, n_g) < -d - (q+(1-q)p)c. \tag{11}$$

A woman will not become pregnant if

$$pU_b(n_b, n_g) + (1-p)U_g(n_b, n_g) < 0, \tag{12}$$

$$U_b(n_b, n_g) < \frac{d + (1-(1-q)p)c}{(1-q)p}, \tag{13}$$

$$\textit{and}\;\; U_g(n_b, n_g) < \frac{d + (q+(1-q)p)c}{(1-q)(1-p)}. \tag{14}$$

As compared to the baseline case where gender determination is not available (Proposition 1), pregnancy is a more likely outcome in the presence of a gender-determination procedure. When a pregnancy occurs, gender determination is chosen when the incremental utility to having a son is highly positive and the incremental utility to having a daughter is highly negative (or vice versa). Also, note that an immediate implication of Proposition 2 is that a woman with complete gender indifference (that is, $U_b(n_b, n_g) = U_g(n_b, n_g)$ for all values of n_b and n_g) would never determine gender during a pregnancy if either $c > 0$ or $d > 0$.

Figures 6 and 7 provide a graphical representation of the decision regions (as a function of the incremental utilities $U_b(n_b, n_g)$ and $U_g(n_b, n_g)$) described by Proposition 2. For both figures, it is assumed that the probabilities of a boy and girl are the same ($p = 1/2$). Figure 6 considers

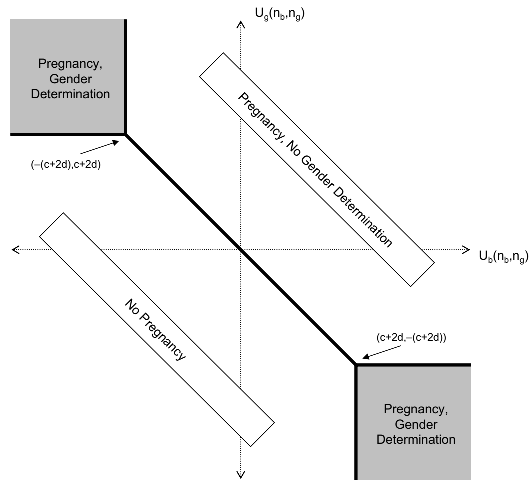

Figure 6: Last-Period Outcome with $p = 1/2$, $q = 0$

the situation where involuntary termination does not occur ($q = 0$), whereas Figure 7 considers the possibility of involuntary termination ($q > 0$). For comparison purposes, the gender-determination region from Figure 6 is shown with a dotted line in Figure 7.[15] As discussed above, the gender-determination region becomes smaller as q rises. In addition, note that the horizontal lines for the gender-determination region become sloped when $q > 0$. Considering the lower-right region in Figure 7, for instance, as the incremental utility of a son becomes larger, the disutility associated with a daughter must also become larger (to offset the potential involuntary termination of a son) for gender determination to occur.

The model has intuitive predictions for how a woman's period-T decisions would be affected by the number of sons and daughters she has at that time:

Proposition 3 *Suppose that* $n'_b \geq n_b$ *and* $n'_g \geq n_g$*. Then, in period* T*,*

(i) if a woman with n_b *boys and* n_g *girls would not become pregnant, she would not become pregnant with* n'_b *boys and* n'_g *girls;*

(ii) if a woman with n_b *boys and* n_g *girls would become pregnant and determine gender, she would either not become pregnant or become pregnant and determine gender with* n'_b *boys and* n'_g *girls;*

(iii) if a woman with n_b *boys and* n_g *girls would become pregnant and not determine gender, she would either not become pregnant, become pregnant and determine gender, or become pregnant and*

[15]This comparison should be considered a comparative-statics exercise. In fact, a higher q procedure such as amniocentesis would arguably have a lower c (as compared to ultrasound) since the termination would most likely occur earlier in the pregnancy.

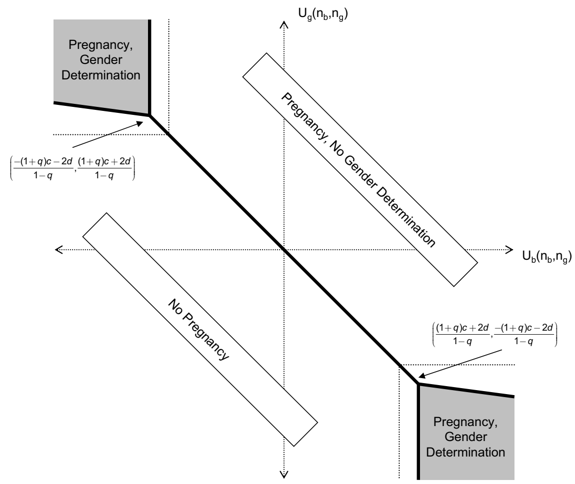

Figure 7: Last-Period Outcome with $p = 1/2$, $q > 0$

not determine gender with n'_b boys and n'_g girls.

In the context of Figures 6 and 7, note that each additional child in the family causes a movement downward and leftward in the diagram (see Figure 5). Proposition 3 simply describes how such movements among the three possible decision regions can occur.

Note that Proposition 3 is a general result that holds regardless of the form of gender preferences. For a woman with strong gender preferences, the predictions become sharper. In particular, for some value of n_b, consider a woman with a strong son bias, in the sense that the incremental utility of having a son remains the same if a daughter is born. (Note that the strong son bias could either result from an overall son bias or from a gender-mix bias.)

Proposition 4 *Suppose that $U_b(n_g, n_b) = U_b(n_g + 1, n_b) > 0$ for given values of n_b and n_g. Then, in period T, if a woman with n_b boys and n_g girls would become pregnant and determine gender, she would become pregnant and determine gender with n_b boys and $n_g + 1$ girls.*

For such women, parts (i) and (iii) of Proposition 3 still hold, but Proposition 4 rules out the possibility of not becoming pregnant for those women in part (ii). An important implication of Proposition 4 is that, *among such women who become pregnant*, the prevalence of gender determination in period T becomes unambiguously higher as daughters are added to the family. (The analogous result would hold for women with a strong daughter bias when sons are added to the family.)

2.4 Decisions in earlier periods with gender determination

The same bottom-up strategy used for period T is used for earlier periods in order to derive a woman's expected-utility-maximizing decisions. To simplify notation within this section, the arguments of the incremental and continuation utilities are suppressed as follows:

$$\begin{aligned} U_b &\equiv U_b(n_b, n_g), \\ U_g &\equiv U_g(n_b, n_g) \\ V_{t+1} &\equiv V_{t+1}(n_b, n_g) \\ V_{t+1}^b &\equiv V_{t+1}(n_b + 1, n_g) \\ V_{t+1}^g &\equiv V_{t+1}(n_b, n_g + 1). \end{aligned}$$

The last three expressions correspond to the continuation utilities associated with having no children in period t, having a son in period t, and having a daughter in period t, respectively.

First, the termination decision (conditional on gender determination) is described by:

Lemma 3 *(Termination decision in period t) If the gender is known to be male, then the pregnancy will be terminated if and only if $U_b + \delta(V_{t+1}^b - V_{t+1}) < -c$. If the gender is known to be female, then the pregnancy will be terminated if and only if $U_g + \delta(V_{t+1}^g - V_{t+1}) < -c$.*

The decisions in periods prior to T are more complex due to the need to consider future fertility periods. If gender is determined to be female, the woman considers not only the incremental utility from a daughter but also the (discounted) difference in continuation utilities between having a daughter and not having a daughter. Note that the difference $V_{t+1}^b - V_{t+1}$ is negative (or zero) under the assumption that incremental utilities are declining. Therefore, a termination is more likely in an earlier period $t < T$ than in the last period T *ceteris paribus.*

As an example, consider a woman in period $T-1$ with no children whose incremental utilities are described by Example 4. Even if the incremental utility of a daughter $U_g(0, 0)$ is *positive*, it is possible that the woman would decide to terminate the pregnancy if the incremental utility of a son is sufficiently high. The reason is that the woman will have at most one child, and the next period (period T) offers an opportunity to have a son (with higher associated incremental utility). This type of argument would also apply to Examples 2 and 3. As a woman approaches the total number of children desired, there will be an additional incentive to terminate a pregnancy if there are future fertility periods and there is a substantial difference between the incremental utilities of a son and daughter.

In the interest of space, the additional results for the pregnancy and gender-determination decisions (analogous to Lemma 2 and Proposition 2 for period T) are explicitly stated in Appendix A. The following proposition provides a simple description of the dynamics of the pregnancy and gender-determination decisions from one period to the next:

Proposition 5 *For given values of n_b and n_g,*

(i) if a woman would not become pregnant in period $t+1$, then she would not become pregnant in period t;

(ii) if a woman would become pregnant and determine gender in period $t+1$, then she would become pregnant and determine gender in period t;

(iii) if a woman would become pregnant and not determine gender in period $t+1$, then she would become pregnant in period t and might determine gender in period t.

This proposition indicates that, for given values of n_b and n_g, the gender-determination region is larger in earlier periods. The possibility of future pregnancies serves to increase the opportunity cost of having a child of the less desired gender.

3 Data Sources

The stylized model of Section 2 makes the link between gender-mix preferences and gender-determination preferences. For parents having a strong bias toward a specific gender, the model predicts that the strongest incentives for gender determination would occur at later births for those parents who previously had children of the less-preferred gender. Unfortunately, existing abortion data in the United States are inadequate for analyzing possible evidence of gender-selective practices. First, gender is not recorded in the two primary abortion surveys conducted in the United States, conducted by the Centers for Disease Control and Prevention (CDC) and the Alan Guttmacher Institute. Second, although information on the number of previous live births is available in these surveys, there is no information on the gender of a mother's existing children. Third, not all states have abortion data available. Joyce et. al. (2004), who have compiled the most comprehensive data on abortions to date, indicate that 19 states (including populous states such as California, Florida, and Illinois) had data unavailable "due to statutory restrictions or inadequate data collection and/or storage." Fourth, when women are asked about the reason(s) for having an abortion, gender preference is rarely mentioned (see, for example, Torres and Forrest (1998)).

In addition to the lack of informative data on abortions, there is no single data source in the United States that records *both* gender preferences and birth outcomes at an individual (mother or family) level. As such, the empirical approach that follows will use Census data in order to document aggregate gender-mix preferences among various different racial groups within the United States. For these same racial groups, birth data will then be used to empirically analyze the link between birth parity and gender likelihood (and for California births, also the link between previous children's gender and gender likelihood). If a given racial group exhibits a strong aggregate preference for sons (daughters), unusually high percentages of boy (girl) births among later children would be consistent with gender-selective practices. On the other hand, if a given racial group exhibits an aggregate preference for a gender mix, seemingly "normal" boy-birth percentages at later births could arise if gender selection for boys and gender selection for girls are both practiced. For this reason, it is useful to have data on the gender of previous births within a family. For instance, if gender determination were practiced to achieve gender mix, families with two daughters (two sons) would have unusually high percentages of male (female) births among their third children.

For this study, three different data sources are utilized: (1) the 5-percent public-use microdata samples (PUMS) of the United States Census (specifically, the 1980, 1990, and 2000 editions); (2) federal natality data (annual files from 1971 to 2002) from the National Center for Health Statistics (NCHS); and, (3) California natality data (annual files from 1982 to 2003) from the California Department of Health Services (CDHS). The Census data and federal natality data are publicly available, whereas the California natality data contain personal identifiers and are subject to confidentiality restrictions.[16] As discussed in more detail below, the personal identifiers were used to maternally link births and identify siblings.

Table 2 provides a summary of the three data sources to clarify the advantages and disadvantages of each. Further details for each of the three data sources are given below:

(1) Census data: The obvious attraction of Census data is that it is representative of the entire United States population. For this study, the additional advantages are the detailed race categories (including Chinese, Indian, Japanese, and Korean) used in the Census survey and the fact that all family members (including their ages and genders) in a household are observed. The latter aspect makes the data suitable for examining how fertility decisions depend on the gender mix of previous children. This idea has been pursued by others (e.g., Dahl and Moretti (2004)) but not at the detailed racial level considered in the next section. Since gender is observed for each child, the Census data also allow one to examine whether a child's gender is related to the gender of previous children. Evidence on this relationship will be also presented in the next section. Unfortunately, due to the 5% sampling, the relatively low sample sizes for most races (except for whites) make the statistical estimates here rather imprecise. The California natality data turn out to be more useful for estimation in this context. Finally, the Census data contain no information about a mother's pregnancies.

[16] A version of the natality data, without personal identifiers, is publicly available from the CDHS.

Table 2: Summary of Data Sources

	Census Data	Federal Natality Data	California Natality Data
Years	1980, 1990, 2000	1971–2002	1982–2003
Sample	5% of U.S. population	1971–1984: 50–100% of births 1985–2002: 100% of births	All California births
Asian race information	Detailed races available	Chinese and Japanese in all years; detailed races available from 1992 on	Detailed races available
Able to link siblings	Yes	No	Yes, using personal identifiers
Prenatal-care data	No	Yes, with ultrasound and amniocentesis usage available 1989 on	Yes, with ultrasound and amniocentesis usage available 1989 on
Data on previous terminated pregnancies	No	Yes, all years	Yes, all years

(2) Federal natality data: These annual data files contain information on births occurring within the United States, obtained from birth certificates filed in individual states. Since 1985, a 100-percent sample of birth certificates has been used to compile these data. In 1971, a 50-percent sample of birth certificates was used. From 1972 to 1984, a 100-percent sample was used for states participating in the Vital Statistics Cooperative Program (with the number of such states increasing from 6 to 46 during the period) and a 50-percent sample for other states. Each record in the federal natality data contains detailed information about the birth (including gender and parity), maternal characteristics (including age, education, and race), and prenatal care (including month of first prenatal visit). Each birth record also indicates the number of previous terminated pregnancies a mother has experienced and, from 1989 on, whether ultrasound and/or amniocentesis were used during pregnancy. The number of terminated pregnancies includes both voluntary and involuntary terminations but does not specify the type(s) of termination(s). Although the detailed information in the federal natality data is very useful for examining the determinants of baby gender, the data has two important limitations. First, detailed Asian races (such as Indian, Korean, and Vietnamese) were only recorded in the data starting in 1992; prior to that, the only specific Asian races recorded were Chinese and Japanese.[17] Second, due to a lack of personal identifiers, there is no way to reliably link births of the same mother together. Therefore, although the birth-order and gender of a given child is observed, there is no way to relate birth outcomes to the gender of a mother's previous children.

[17] Depending on the year, other Asian races were included in a category such as "Other" or "Other Asian or Pacific Islander."

Table 3: Variable Descriptions, Natality Data

Variable	Description
Boy	1 if baby is male
Mother's age	Age (in years) at birth
Birth parity	Number of previous live births reported, plus one
Mother's education	Education in years (note: maximum value of 17)
Amniocentesis	1 if use of amniocentesis during pregnancy is reported
Ultrasound	1 if use of ultrasound during pregnancy is reported
Previous termination	1 if a previous terminated pregnancy (voluntary or involuntary) is reported
Foreign-born mother	1 if mother's birthplace is outside the U.S.
Same-race father	1 if mother and father have same reported race
No prenatal care	1 if mother had no prenatal-care visits
1st-trimester care	1 if first prenatal visit occurred in months 1–3 of pregnancy
2nd-trimester care	1 if first prenatal visit occurred in months 4–6 of pregnancy
3rd-trimester care	1 if first prenatal visit occurred in months 7+ of pregnancy

(3) California natality data: The California natality data contain information on all births that occurred within California between 1982 and 2003 (a total of over 11.6 million births, accounting for roughly 10% of all births in the United States). Each birth record in the California dataset contains essentially the same information (on birth outcomes, mother demographics, and prenatal care) that is available in the federal dataset. The California data, however, overcome the two limitations of the federal data mentioned above. First, detailed Asian race classifications are available from 1982 on. Second, the author was provided with personal identifiers (specifically, mother's full maiden name and mother's birthdate) that enable accurate matching of a given mother's births. This paper will use both an *unlinked* version and a *linked* version of the data. The unlinked version makes no use of the personal identifiers but still serves as a useful complement to the federal data since the detailed Asian race classifications are absent from the federal data between 1982 and 1991. The linked version is used in order to analyze birth and pregnancy outcomes for a mother's second and/or third child, conditioning on gender of previous children. Complete details on the algorithm used to link the California birth records are provided in Appendix B.

Table 3 describes the primary variables from the natality data that are used in the analysis of Section 4. For an overview of the two natality datasets and a comparison of their sample composition, Table 4 reports sample averages of several variables. The results are broken down by mother's race and reported for 1992–2002, the years for which information is available in both datasets for the six races considered. The last row in the table indicates the percentage of U.S. births that occurred in California for each of the racial categories. For the purposes of this study, an appealing feature of the California data is the disproportionately large number of Asian births. The percentage of births occurring in California for the four Asian racial categories ranges from 29.8% for Indian mothers to 46.5% for Korean mothers. Given these large percentages, it is not surprising to find that the descriptive statistics for the California births and U.S. births are very similar for the four Asian races. On average, the California Asian mothers are slightly more educated, slightly less likely to have an ultrasound or amniocentesis, and slightly more likely to have a child with a father of the same race (except for Japanese).

Interestingly, the percentage of foreign-born mothers among Chinese, Indian, and Korean births is extremely high — nearly 90% for Chinese mothers and close to 95% for both Indian and Korean mothers. The percentage of births to fathers of the same race is also very high for these races — between 70% and 80% for Chinese and Korean births and around 90% for Indian births. As a comparison, the percentage of foreign-born Japanese mothers is significantly lower (55.5% for U.S. births), and the percentage of Japanese mothers having a child with a same-race father is also significantly lower (40.8% for U.S. births). The high percentage of foreign-born Asian mothers and same-race fathers suggests that cultural influences could play a role in fertility decisions, a possibility that is examined in further detail in Section 4.

Table 4 indicates several differences between the Asian mothers and non-Asian mothers. Compared to white and black mothers, Asian mothers are, on average, older when they give birth, more educated, more likely to have first-trimester prenatal care, less likely to have had a previous termination, and more likely to have a boy. Finally, note that the sample composition for white and black births differs somewhat between the federal data and the California data. White mothers in California are far more likely to be classified as Hispanic, and 43.1% of births are to foreign-born white mothers.[18] On average, white mothers in California are less educated, less likely to have an ultrasound, and less likely to have first-trimester prenatal care. Black mothers in California, on the other hand, have a higher average education level and are far more likely to have a child with a same-race father (as compared to the federal sample).

[18] "Hispanic" is not categorized as a race in the natality data, but rather is identified through a separate question. In the interest of space, Hispanic and non-Hispanic births are considered together in all of the results reported here. Breaking the (white) sample into Hispanic and non-Hispanic subsamples yields qualitatively similar results, which are available from the author upon request.

Table 4: Sample Averages for 1992–2002 Births, by Race

	White		Black		Chinese		Indian		Japanese		Korean	
	U.S.	Calif.	U.S.	Calif.	U.S.	Calif.	U.S.	Calif.	U.S.	Calif.	U.S.	Calif.
Boy	0.513	0.511	0.508	0.508	0.519	0.520	0.517	0.521	0.514	0.514	0.520	0.519
Mother's age	27.21	27.21	24.93	26.00	31.28	31.81	28.83	28.81	31.37	32.06	30.19	30.34
Birth parity	1.99	2.15	2.18	2.35	1.63	1.66	1.68	1.65	1.70	1.69	1.67	1.67
Mother's education	12.75	11.67	12.27	12.66	14.14	14.32	14.48	14.78	14.63	14.89	14.60	14.76
Amniocentesis	0.029	0.025	0.016	0.019	0.060	0.053	0.031	0.021	0.081	0.076	0.035	0.023
Ultrasound	0.648	0.518	0.579	0.465	0.624	0.544	0.637	0.583	0.654	0.654	0.550	0.440
Previous termination	0.241	0.171	0.288	0.208	0.228	0.136	0.203	0.141	0.256	0.177	0.231	0.157
Foreign-born mother	0.066	0.431	0.102	0.072	0.899	0.892	0.947	0.940	0.555	0.520	0.948	0.959
Same-race father	0.867	0.958	0.575	0.909	0.780	0.803	0.890	0.929	0.408	0.393	0.727	0.790
No prenatal care	0.010	0.012	0.029	0.019	0.003	0.002	0.010	0.003	0.005	0.004	0.006	0.005
1st-trimester care	0.838	0.787	0.711	0.742	0.867	0.895	0.828	0.874	0.895	0.900	0.836	0.863
2nd-trimester care	0.127	0.166	0.214	0.202	0.108	0.085	0.130	0.104	0.082	0.081	0.125	0.106
3rd-trimester care	0.025	0.035	0.046	0.037	0.022	0.018	0.032	0.019	0.018	0.014	0.033	0.027
# of births	33,592,117	4,845,929	6,593,499	420,839	310,935	130,103	192,241	57,383	96,247	34,509	97,422	45,323
% of U. S. births	—	14.4%	—	6.4%	—	41.8%	—	29.8%	—	35.9%	—	46.5%

4 Empirical Results

This section provides an analysis of the three data sources described in Section 3, examining both gender preferences and boy-birth determinants. Prior to the analysis, however, Section 4.1 briefly discusses the issue of fetal survival and its implications for interpretation of the empirical results. Using Census data, Section 4.2 examines gender preferences (based on fertility stopping) and gender outcomes by race. Section 4.3 provides a detailed analysis, broken down by race, of boy-birth likelihoods. The relationship between gender and birth parity is considered, and several regression analyses are presented for the federal natality data and the (unlinked) California natality data. Section 4.4 focuses on the maternally linked California natality data, which allows the analysis to condition upon the gender composition of a mother's previous children. Finally, Section 4.5 provides a simple framework (similar to the model of Section 2) to infer the prevalence of gender selection from unusual boy-birth percentages. (In this context, the term "unusual boy-birth percentage" simply indicates a percentage that is different from the one that would be expected in the absence of gender selection.)

4.1 Differential fetal survival

Although it is well known that females have a lower mortality rate throughout life, it is perhaps less well known that the mortality differential between males and females begins at conception. The ratio of male to female conceptions is significantly higher than the ratio of male to female births. According to Perls and Fretts (1998), in their review of gender differentials in mortality, there is a "disproportionate rate of spontaneous abortions, stillbirths, and miscarriages of male fetuses." In one study, Mizuno (2000) documents the high ratio of males to females among miscarriages in Japan. Although the data on fetal deaths in the United States is somewhat limited, the existing information indicates that the percentage of male fetal births is significantly higher than the percentage of male live births. For instance, according to data from the NCHS, 53.3% of the 214,043 fetal deaths that occurred after 20 weeks of gestation between 1995 and 2002 were male.[19][20]

For the purposes of this study, the important facts are that males (i) have more difficulty surviving pregnancy (and being born) than females and (ii) are more likely to cause pregnancy complications. The data sources considered have information on live births, so that the possible "survival bias" should be understood before interpreting any empirical results. A bulk of the analysis below considers regressions where the dependent variable is an indicator of a boy birth.

[19]Source: National Center for Health Statistics, Perinatal Mortality Data, 1995–2002. Data was obtained from the National Bureau of Economic Research.

[20]Gender is generally not recorded for fetal deaths prior to 20 weeks of gestation. Among the 21,399 fetal deaths where gender was recorded between 1995 and 2002, 66.9% were identified as male.

Even without gender selection, the pre-birth gender differential in survival leads to several predictions regarding the empirical relationship between gender-at-birth and other observables. These predictions are discussed below and summarized in Table 5:

- *Prenatal care*: If boys have more difficulty surviving pregnancy, one would expect that (all other things equal) mothers who obtain better prenatal care would be more likely to have a boy. Unfortunately, the directly relevant item in the natality data (both federal and California) is the month of first prenatal visit. This variable proxies for both *intended* prenatal care and *unintended* pregnancy problems. For instance, if two mothers have identical intentions (at the beginning of pregnancy) with respect to prenatal-care visits, the mother that experiences problems early in her pregnancy would be more likely to have an earlier first prenatal-care visit. Thinking about the estimated effect of an indicator variable for a first-trimester visit, there would be a positive association with boy-birth likelihood to the extent that the indicator proxies for intended care and a negative association to the extent that it proxies for a problem pregnancy; the estimated association (which turns out to be negative) would be a combination of these two opposite effects. Other variables that might proxy for prenatal care could also have a relationship with boy-birth likelihood. For example, the level of mother's education would not be expected to have a causal effect on baby gender but is probably positively correlated with quality prenatal care, which would lead one to expect a positive relationship between years of education and boy-birth likelihood. In addition, one might expect a negative association between birth parity (i.e., the birth order of the baby) and boy-birth likelihood. This negative association would arise if women who have more children are less likely to obtain good prenatal care, either because they are less privileged or because there is less time (or lower incentives) to obtain prenatal care for later children. If gender selection becomes more prevalent at higher birth parity, as suggested by the model in Section 2, this effect would lead to a more positive association in the case of son bias (more negative in the case of daughter bias).

- *Ultrasound and amniocentesis*: One must be careful in interpreting a relationship between a baby's gender and the decision by the mother to have an ultrasound or amniocentesis. After all, these two procedures are used primarily for prenatal-care purposes. If the use of ultrasound is a proxy for good prenatal care, there would be a positive association between the likelihood of having a boy and the use of ultrasound, in the absence of any gender-selective intent. On the other hand, amniocentesis is a procedure that is generally performed in high-risk situations, such as pregnancies to older mothers or pregnancies to mothers who have experienced problems in their current pregnancy or prior pregnancies. If boys are less likely to survive high-risk pregnancies, there would exist a negative association between the

Table 5: Predicted Associations between Boy-Birth Likelihood and Observables

Observable variable	Expected association (in the absence of gender selection)
First-trimester care	Uncertain
Mother's education	+
Birth parity	−
Ultrasound	+
Amniocentesis	−
Previous termination	−
Previous son(s)	+

likelihood of having a boy and the use of amniocentesis in the absence of gender selection.

- *Previous termination*: A previous involuntary pregnancy termination could proxy for (i) poor prenatal care or (ii) the difficulty that a mother has carrying a pregnancy to full term. In either case, a negative association between previous involuntary terminations and boy-birth likelihood would be expected. On the other hand, voluntary terminations that are not gender-based would not be expected to cause such an association. Gender-based voluntary terminations would cause an association in the direction of the gender bias.

- *Previous child gender*: Mothers who have given birth to boys previously would be expected to be more likely to give birth to another boy. Even without a biological predisposition for having one gender or the other (for which there is little convincing scientific evidence), a previous male birth serves as a (weak) proxy for quality prenatal care and mother's ability to carry a pregnancy to full term. Either of these characteristics leads to a greater chance of a boy birth.

4.2 Census Data: Gender Preferences and Outcomes

This section considers the decision of families to have either a second or third child based on the gender(s) of their previous child(ren). In addition, we examine whether or not the gender of a second or third child is associated with the gender(s) of their previous child(ren). Table 6 provides the results for second-child outcomes, broken down by first-child gender and by mother's race (including four Asian categories: Chinese, Indian, Japanese, and Korean). For every family with at least one child, the table reports (1) the percentage of families that had a second child within 5 years of the birth of the first child and (2) for those families having a second child, the percentage of boy births. Using the 1980, 1990, and 2000 editions of the 5%-sample Census PUMS, results are provided for two time periods (1963–1979 and 1980–1995), with observations categorized by first-child birthyear. Details on constructing the relevant samples are given in Appendix C. For both second-child likelihoods and boy-birth likelihoods, a p-value is reported (in brackets) for the test of equality between the percentages for firstborn-son families and firstborn-daughter families.

Table 6 indicates that the likelihood of having additional children decreases for all races in the more recent time period (1980–1995). The additional-child likelihoods suggest that, in the aggregate, gender of the first child does not play much of a role in determining whether a family has a second child. There is, however, evidence of son preference among Indian and Korean families in 1980–1995 (more likely to have a second child if the first was a girl) and daughter preference among Japanese families in 1980–1995 (more likely to have a second child if the first was a boy). Note that white mothers are slightly more likely to have a second child if their first child was a boy. The slight difference could reflect either (i) a small preference for daughters or (ii) a greater chance of second-child survival given that the first child was a boy (along the lines of the discussion in Section 4.1).

Looking at the boy-likelihood results in Table 6, white and black mothers are more likely to have a second-child boy if their first child was a boy (1.0 and 1.6 percentage points more likely for white mothers and black mothers, respectively, in the later time period (with p-values of 0.000)). As discussed in Section 4.1, this relationship would be expected (even without a biological predisposition) if a first-birth boy proxies for better prenatal care or a greater chance of a pregnancy being carried to term. Unfortunately, the small sample sizes for the Asian races (between 3,000 and 9,000 observations) result in fairly large p-values for the comparison between the two boy-birth likelihoods. The only low p-values are the two corresponding to the 1980–1995 samples of Chinese births (p-value of 0.104) and Indian births (p-value of 0.051). For these two samples, the pattern observed for white and black mothers is reversed, with Chinese and Indian mothers being more likely to give birth to a boy if their first child was a girl. For Chinese mothers, there is a 52.6% chance of having a boy given a first-child girl and 50.8% chance given a first-child boy; for Indian mothers, there is a 52.9% chance of having a boy given a first-child girl and 50.4% chance given a first-child boy.

Analogous to Table 6, third-birth outcomes (likelihood of having a child and likelihood of having a boy) are summarized in Table 7. Observations are categorized by birthyear of the second child, and likelihoods are reported conditional on the number of boys (0, 1, or 2) among the first two children.[21] The third-child results in Table 7 highlight larger gender-preference differences between races. For white, black, and Japanese families, the overall preference is for a gender mix. That is, for these three racial categories, families are most likely to have a third child if the gender

[21] The category "1 boy" could be further broken down by the gender sequence (i.e., boy-then-girl versus girl-then-boy). Since this breakdown offered no additional qualitative insight, the analysis here (and later in Section 4.4) considers only three categories.

Table 6: Second-Child Outcomes, 5-Percent Census Data

	Birthyear of 1st child	Likelihood of having a 2nd child, given gender of 1st child is:		Likelihood that 2nd child is a boy, given gender of 1st child is:		# of families with kids
		Girl	Boy	Girl	Boy	
White	1963-1979	0.719	0.722	0.509	0.516	850,619
		(0.001)	(0.001)	(0.001)	(0.001)	
		[0.001]		[0.000]		
	1980-1995	0.637	0.640	0.507	0.517	826,978
		(0.001)	(0.001)	(0.001)	(0.001)	
		[0.002]		[0.000]		
Black	1963-1979	0.616	0.619	0.494	0.510	117,955
		(0.002)	(0.002)	(0.003)	(0.003)	
		[0.368]		[0.000]		
	1980-1995	0.518	0.522	0.499	0.515	120,889
		(0.002)	(0.002)	(0.003)	(0.003)	
		[0.108]		[0.000]		
Chinese	1963-1979	0.727	0.706	0.502	0.504	5,223
		(0.009)	(0.009)	(0.012)	(0.011)	
		[0.089]		[0.893]		
	1980-1995	0.564	0.550	0.526	0.508	8,759
		(0.008)	(0.007)	(0.010)	(0.010)	
		[0.168]		[0.104]		
Indian	1963-1979	0.714	0.684	0.501	0.496	3,376
		(0.011)	(0.011)	(0.015)	(0.015)	
		[0.065]		[0.767]		
	1980-1995	0.637	0.603	0.529	0.504	5,971
		(0.009)	(0.009)	(0.012)	(0.012)	
		[0.006]		[0.051]		
Japanese	1963-1979	0.688	0.689	0.503	0.509	3,910
		(0.011)	(0.010)	(0.014)	(0.014)	
		[0.957]		[0.684]		
	1980-1995	0.619	0.654	0.500	0.500	3,167
		(0.012)	(0.012)	(0.016)	(0.015)	
		[0.044]		[0.972]		
Korean	1963-1979	0.747	0.737	0.505	0.513	3,566
		(0.010)	(0.010)	(0.014)	(0.014)	
		[0.492]		[0.639]		
	1980-1995	0.637	0.613	0.517	0.508	4,816
		(0.010)	(0.010)	(0.013)	(0.013)	
		[0.090]		[0.499]		

"Having a 2nd child" means that the 2nd child is born within 5 years of the 1st child.
Standard errors are reported in parentheses.
The p-values associated with the two-sided test of equality are reported in brackets.

Table 7: Third-Child Outcomes, 5-Percent Census Data

	Birthyear of 2nd child	Likelihood of having a 3rd child, given # of previous boys:			Likelihood that 3rd child is a boy, given # of previous boys:			# of families with ≥ 2 kids
		0 boys	1 boy	2 boys	0 boys	1 boy	2 boys	
White	1963-1979	**0.444**	**0.365**	**0.426**	**0.506**	0.512	0.514	463,363
		(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	
	1980-1995	**0.367**	**0.298**	**0.359**	**0.499**	**0.514**	**0.521**	606,826
		(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	
Black	1963-1979	0.523	**0.487**	0.523	**0.492**	0.504	0.505	57,140
		(0.004)	(0.003)	(0.004)	(0.006)	(0.004)	(0.006)	
	1980-1995	0.417	**0.379**	0.417	**0.484**	0.509	0.514	82,194
		(0.003)	(0.002)	(0.003)	(0.005)	(0.004)	(0.005)	
Chinese	1963-1979	**0.506**	**0.373**	**0.327**	0.492	0.511	0.547	2,671
		(0.020)	(0.013)	(0.018)	(0.028)	(0.022)	(0.033)	
	1980-1995	**0.317**	**0.188**	**0.216**	0.523	0.487	0.531	5,438
		(0.013)	(0.007)	(0.011)	(0.025)	(0.022)	(0.028)	
Indian	1963-1979	**0.436**	0.269	0.319	0.490	0.542	0.527	1,542
		(0.026)	(0.016)	(0.024)	(0.040)	(0.034)	(0.045)	
	1980-1995	**0.332**	0.190	0.203	0.567	0.542	0.515	4,358
		(0.014)	(0.008)	(0.012)	(0.026)	(0.024)	(0.034)	
Japanese	1963-1979	0.350	**0.264**	0.376	0.532	0.481	0.564	2,007
		(0.021)	(0.014)	(0.021)	(0.038)	(0.031)	(0.035)	
	1980-1995	0.258	0.216	0.271	0.536	0.535	0.537	2,145
		(0.019)	(0.012)	(0.019)	(0.044)	(0.033)	(0.041)	
Korean	1963-1979	**0.448**	0.317	0.299	0.492	0.487	0.498	1,871
		(0.023)	(0.015)	(0.021)	(0.035)	(0.029)	(0.042)	
	1980-1995	**0.268**	**0.130**	**0.160**	0.546	**0.468**	0.569	3,509
		(0.015)	(0.008)	(0.012)	(0.033)	(0.033)	(0.042)	

"Having a 3rd child" means that the 3rd child is born within 5 years of the 2nd child.
Standard errors are reported in parentheses.
Bold indicates an estimate is statistically different (at a 5% level) from *both* of the estimates for the other two categories.

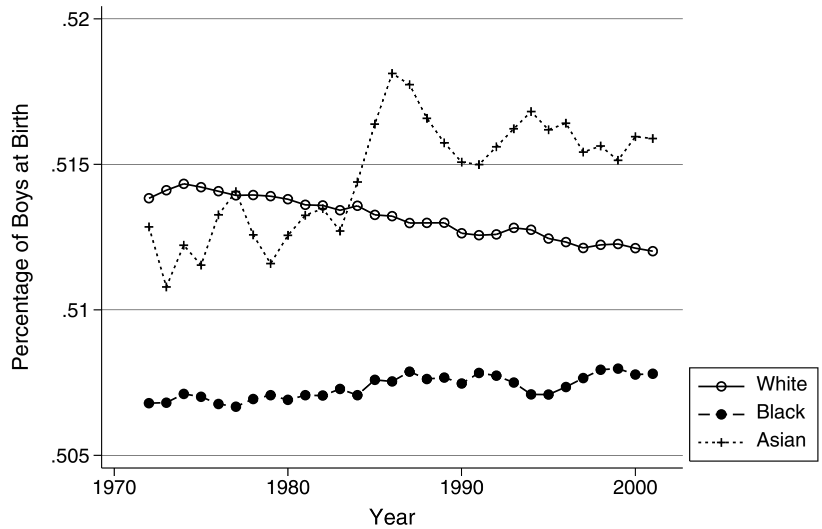

Figure 8: Likelihood of Boy Births in the United States, by Mother's Race

of the previous two was the same (either two sons or two daughters) and least likely to have a third child if they have had a son and a daughter. For the other three Asian races (Chinese, Indian, Korean), there is a definite bias toward having a son: (i) families with two daughters are far more likely to have a third child than families with one or two sons; and, (ii) families with two sons are about equally likely to have a third child as families with a son and a daughter. Although the overall likelihood of having a third child drops in the later time period across all races, the pattern of gender-mix preferences remains very similar across the two time periods for each race.

The results on boy-birth likelihoods from Table 7 are somewhat less informative. For both white and black mothers, the likelihood of having a third-child boy increases with the number of previous boys, which is consistent with the pattern seen among second-child births in Table 6. For the Asian races, the sample-size issues are even more problematic than for the second-child results: the third-child samples are smaller than the second-child samples, and the observations are now broken into three categories rather than two. The standard errors are too large to say much about the statistical differences between the conditional boy-birth likelihoods. The only exception appears to be for Korean mothers in 1980–1995, where the likelihood of a third-child boy is significantly lower when the first two children are a boy-girl mix.

4.3 Boy-Birth Percentages and Regression Analyses

Figure 8, based upon the federal natality data, plots the time series of boy-birth percentages within the United States for three racial categories: white, black, and Asian. The Asian category includes any mother that was classified as Chinese, Indian, Japanese, Korean, Vietnamese, or "other Asian."[22] Over the last three decades, the boy percentage for whites has slowly declined from about 51.4% to about 51.2%, a decline that has been noted in previous research (Davis et. al. (1998), Marcus et. al. (1998)).[23] For blacks, who have lower overall boy percentages (in the 50.7%–50.8% range over the period), there has been a slow increase in the likelihood of boy births, perhaps due to improvements in prenatal care. For Asian mothers, the percentage of boy births experienced an increase of roughly one percentage point during the early 1980's and remained at a higher level through 2000.

Table 8 provides a breakdown of boy-birth likelihoods by time period, by race, and by parity of the child. For this table and the regressions that follow, the samples are restricted to births occurring to parents of the same race. Results for both the federal natality data and California natality data are reported. Each cell in the table has a boy-birth percentage with its associated standard error (in parentheses) and the number of births. The pattern among white births and black births is that higher-parity births are slightly less likely to be boys. The first-child boy percentages among the Asian races are roughly the same as that for whites. For later children in later time periods (since 1980), there appear to be some higher boy percentages for Chinese, Indian, and Korean births.

To determine the statistical significance of these higher percentages and to also control for other factors (such as mother's age, prenatal care, etc.) that might affect the likelihood of having a boy, Tables 9–12 report regression results using these same data sources (Table 9: 1971–1980 federal data, Table 10: 1981–1990 federal data, Table 11: 1991–2002 federal data, Table 12: 1982–2003 California data). The reported results are from linear probability models, where the dependent variable is an indicator variable equal to one for boy births. Heteroskedasticity-robust standard errors are reported.[24] The linear probability model is particularly appropriate for this application since the fitted probabilities are very close to 50%; probit estimation yields nearly identical results in all cases. To make these tables easier to read, all estimates have been scaled up by a factor of 100, so that they can be interpreted as percentage-point effects; for instance, an estimate of 1 would correspond to an increase of one percentage point in the boy-birth probability. In addition, any estimate reported in bold indicates statistical significance at a 5% level.

For each time period and race considered in Tables 9–12, results are reported for two different regressions, one with no other control variables (i.e., including only indicator variables for

[22] As in Figure 1, each point represents a three-year moving average.

[23] This slight increase in boy-birth percentages has also been observed in other countries, including Canada (Allan et. al. (1997)), Denmark (Moller (1996)), and the Netherlands (van der Pal-de Bruin et. al. (1997)).

[24] Since the fitted probabilities are so close to 50%, these standard errors are nearly identical to the unadjusted standard errors.

Table 8: Likelihood of a Boy Birth

	Time Period	1st Child Fraction (s.e.) # births	2nd Child Fraction (s.e.) # births	3rd Child Fraction (s.e.) # births	4th Child Fraction (s.e.) # births
Federal Natality Data					
White	1971-1980	0.515 (0.000)	0.514 (0.000)	0.513 (0.000)	0.512 (0.000)
		7,416,989	6,031,345	2,639,507	1,005,811
	1981-1990	0.515 (0.000)	0.514 (0.000)	0.513 (0.000)	0.514 (0.000)
		10,842,692	9,056,005	4,054,274	1,347,298
	1991-2002	0.514 (0.000)	0.513 (0.000)	0.512 (0.000)	0.511 (0.000)
		12,764,066	10,924,640	5,238,598	1,836,545
Black	1971-1980	0.510 (0.001)	0.508 (0.001)	0.508 (0.001)	0.507 (0.001)
		825,659	727,890	406,290	199,529
	1981-1990	0.511 (0.000)	0.509 (0.000)	0.508 (0.001)	0.508 (0.001)
		1,224,954	1,082,803	624,574	273,886
	1991-2002	0.511 (0.000)	0.510 (0.000)	0.510 (0.001)	0.509 (0.001)
		1,527,312	1,305,316	747,525	320,798
Chinese	1971-1980	0.517 (0.004)	0.515 (0.004)	0.509 (0.007)	0.472 (0.012)
		19,925	15,363	5,672	1,632
	1981-1990	0.518 (0.002)	0.516 (0.002)	0.526 (0.004)	0.526 (0.009)
		63,469	43,604	13,466	3,324
	1991-2002	0.517 (0.001)	0.520 (0.002)	0.532 (0.003)	0.547 (0.008)
		129,302	101,245	23,524	4,170
Indian	1992-2002	0.510 (0.002)	0.518 (0.002)	0.547 (0.004)	0.543 (0.008)
		85,176	62,929	17,153	3,758
Japanese	1971-1980	0.505 (0.005)	0.518 (0.005)	0.522 (0.009)	0.520 (0.021)
		9,954	8,949	3,078	569
	1981-1990	0.508 (0.004)	0.514 (0.004)	0.501 (0.008)	0.492 (0.019)
		16,313	13,460	4,355	721
	1991-2002	0.512 (0.003)	0.516 (0.004)	0.517 (0.007)	0.490 (0.017)
		21,285	16,034	4,748	838
Korean	1992-2002	0.517 (0.003)	0.520 (0.003)	0.531 (0.006)	0.521 (0.016)
		33,886	28,155	7,389	961
California Natality Data					
White	1982-2003	0.514 (0.000)	0.511 (0.000)	0.510 (0.000)	0.509 (0.001)
		3,366,559	2,783,169	1,503,774	636,028
Black	1982-2003	0.509 (0.001)	0.508 (0.001)	0.506 (0.001)	0.509 (0.002)
		284,451	223,380	134,318	64,904
Chinese	1982-2003	0.520 (0.002)	0.518 (0.002)	0.525 (0.004)	0.534 (0.008)
		88,549	69,645	19,505	4,388
Indian	1982-2003	0.509 (0.002)	0.516 (0.003)	0.567 (0.006)	0.594 (0.012)
		40,063	29,527	7,947	1,614
Japanese	1982-2003	0.511 (0.004)	0.516 (0.005)	0.501 (0.009)	0.495 (0.021)
		13,837	10,783	3,402	584
Korean	1982-2003	0.514 (0.003)	0.518 (0.003)	0.529 (0.006)	0.554 (0.017)
		33,967	27,893	6,941	885

birth order) and one with several other control variables. The regressions with no control variables provide the differences (and the associated p-values) between the first-birth boy percentages and later-birth boy percentages that are reported in Table 8. The control variables include year of birth and the variables from Table 3, except that amniocentesis and ultrasound variables are not included for the 1971–1980 and 1981–1990 results. To remain completely flexible about the specification for mother's age, a complete set of age dummies was included in the regressions with control variables. For all regressions reported, the samples were restricted to births of first children through fourth children. As in Table 8, the samples were also restricted to same-race parents. The indicator variable for first-child births is the "omitted category," so that the estimates for the three birth-parity indicators ("2nd child," "3rd child," and "4th child") should be interpreted as a difference in boy likelihood from first-child births. For instance, in the California sample of Chinese births (see Table 12), the regression with control variables indicates that the fourth child is 2.882 percentage points more likely to be male than the first child. For the prenatal-care indicator variables, note that first-trimester care is the omitted category.

For white births, the results from Tables 9–11 indicate that the likelihood of a boy becomes slightly lower at higher parity (consistent with the discussion in Section 4.1), even when other variables are included as controls. This finding holds for 1971–1980, the period in which gender determination would have been either impossible or very unlikely, and then continues in the later periods. The same is true for black births, except that the birth-parity estimates are not statistically significant (at a 5% level) when control variables are included in either the 1991–2002 federal sample or the California sample. Note that the magnitudes of the birth-parity effects are quite low for white births (for example, between 0.077 and 0.310 percentage points in the federal-data regressions with control variables), but the huge sample sizes allow these effects to be precisely estimated.

For Chinese births, statistical evidence of higher boy percentages for third and fourth children is seen in the 1991–2002 federal sample and the 1982–2003 California sample (almost 3 percentage points more likely to have a fourth-child boy than a first-child boy). The evidence of higher boy percentages at later births is even stronger among Indian parents, with larger effects seen for the third child (4.0 percentage points in the federal sample and 6.5 percentage points in the California sample) and the fourth child (over 10.0 percentage points in the California sample). For Indian parents, even the second-child boy percentage is significantly higher (at a 5% significance level, with a magnitude of around one percentage point). For Korean parents, all of the estimates on the higher-order births are positive though only the third-child estimate from the 1991–2002 federal sample is significant at a 5% level. For Japanese births, there appears to be no evidence of unusually high boy percentages after 1980.

Table 9: Boy-Regression Results for Federal Natality Data, 1971–1980

	White		Black		Chinese		Japanese	
2nd child	**-0.083**	**-0.099**	**-0.220**	-0.148	-0.218	0.267	1.330	**2.089**
	(0.027)	(0.034)	(0.080)	(0.101)	(0.537)	(0.744)	(0.728)	(0.940)
3rd child	**-0.197**	**-0.167**	**-0.209**	-0.069	-0.845	-0.746	1.725	2.571
	(0.036)	(0.047)	(0.096)	(0.126)	(0.752)	(1.096)	(1.031)	(1.367)
4th child	**-0.295**	**-0.310**	**-0.321**	-0.202	-4.563	-0.560	1.569	-0.209
	(0.053)	(0.071)	(0.125)	(0.165)	(1.285)	(1.911)	(2.154)	(2.751)
Education		**0.027**		0.024		0.187		-0.058
		(0.008)		(0.024)		(0.110)		(0.214)
Year		-0.007		0.021		-0.022		0.202
		(0.005)		(0.015)		(0.131)		(0.144)
2nd-trimester care		**0.621**		**0.579**		0.388		2.020
		(0.040)		(0.092)		(0.852)		(1.313)
3rd-trimester care		0.087		0.248		2.117		4.975
		(0.091)		(0.175)		(1.697)		(3.609)
No prenatal care		0.145		**0.889**		5.938		-12.024
		(0.199)		(0.301)		(5.312)		(8.590)
Foreign-born		0.048		0.159		-0.835		0.065
		(0.082)		(0.236)		(1.120)		(0.872)
Previous termination		**-0.175**		-0.134		-0.784		-2.727
		(0.040)		(0.106)		(0.950)		(1.112)
Age dummies?	No	Yes	No	Yes	No	Yes	No	Yes
# observations	17,093,652	12,334,129	2,159,368	1,584,020	42,592	24,493	22,550	15,107

Estimates have been multiplied by 100 in order to be interpreted as percentage-point effects.
Standard errors are reported in parentheses.
Bold indicates statistical significance at a 5% level.

Table 10: Boy-Regression Results for Federal Natality Data, 1981–1990

	White		Black		Chinese		Japanese	
2nd child	**-0.101**	**-0.077**	**-0.239**	-0.078	-0.150	-0.526	0.590	0.630
	(0.023)	(0.026)	(0.066)	(0.076)	(0.311)	(0.413)	(0.582)	(0.718)
3rd child	**-0.177**	**-0.132**	**-0.323**	-0.106	0.851	1.242	-0.689	-0.645
	(0.029)	(0.035)	(0.078)	(0.093)	(0.474)	(0.648)	(0.853)	(1.071)
4th child	**-0.186**	**-0.173**	**-0.389**	-0.073	0.795	1.113	-1.532	-0.033
	(0.046)	(0.055)	(0.106)	(0.125)	(0.889)	(1.240)	(1.903)	(2.284)
Education		**0.036**		0.019		0.054		0.314
		(0.006)		(0.017)		(0.059)		(0.168)
Year		**-0.009**		0.015		0.025		-0.027
		(0.004)		(0.010)		(0.063)		(0.108)
2nd-trimester care		**0.541**		**0.610**		**1.396**		0.238
		(0.034)		(0.074)		(0.487)		(1.169)
3rd-trimester care		**0.223**		-0.179		-0.536		-2.162
		(0.074)		(0.142)		(0.966)		(2.450)
No prenatal care		**0.271**		0.074		2.764		1.076
		(0.128)		(0.201)		(2.227)		(6.849)
Foreign-born		-0.004		0.061		0.516		1.220
		(0.055)		(0.111)		(0.819)		(0.662)
Previous termination		**-0.091**		**-0.236**		-0.741		-0.488
		(0.027)		(0.070)		(0.476)		(0.763)
Age dummies?	No	Yes	No	Yes	No	Yes	No	Yes
# observations	25,300,269	20,165,927	3,206,217	2,700,330	123,863	76,317	34,849	24,843

Estimates have been multiplied by 100 in order to be interpreted as percentage-point effects.
Standard errors are reported in parentheses.
Bold indicates statistical significance at a 5% level.

Table 11: Boy-Regression Results for Federal Natality Data, 1991–2002

	White		Black		Chinese		Indian		Japanese		Korean	
2nd child	**-0.123**	**-0.084**	**-0.147**	0.043	0.315	0.294	**0.821**	**0.975**	0.398	0.444	0.273	0.558
	(0.021)	(0.022)	(0.060)	(0.065)	(0.210)	(0.226)	(0.263)	(0.290)	(0.523)	(0.560)	(0.403)	(0.443)
3rd child	**-0.221**	**-0.167**	**-0.139**	0.147	**1.444**	**1.329**	**3.698**	**4.009**	0.480	0.438	**1.450**	**1.803**
	(0.026)	(0.028)	(0.071)	(0.080)	(0.354)	(0.382)	(0.417)	(0.471)	(0.802)	(0.867)	(0.641)	(0.715)
4th child	**-0.334**	**-0.264**	**-0.213**	0.073	**3.021**	**2.689**	**3.260**	**3.488**	-2.160	-2.544	0.436	0.744
	(0.039)	(0.043)	(0.097)	(0.109)	(0.783)	(0.825)	(0.831)	(0.908)	(1.761)	(1.862)	(1.634)	(1.735)
Education		**0.046**		0.010		0.034		-0.060		-0.067		0.002
		(0.004)		(0.015)		(0.035)		(0.053)		(0.143)		(0.101)
Year		**-0.007**		-0.005		-0.011		-0.024		0.000		-0.047
		(0.003)		(0.008)		(0.030)		(0.043)		(0.074)		(0.062)
2nd-trimester care		**0.494**		**0.743**		**1.306**		0.646		-0.502		0.530
		(0.030)		(0.071)		(0.326)		(0.389)		(1.031)		(0.592)
3rd-trimester care		**0.221**		0.152		**1.481**		0.504		1.092		1.834
		(0.067)		(0.151)		(0.688)		(0.747)		(2.129)		(1.066)
No prenatal care		**0.513**		**0.615**		-1.077		-0.002		0.557		-3.994
		(0.120)		(0.230)		(2.118)		(1.265)		(4.920)		(3.082)
Foreign-born		0.031		**0.174**		0.548		-0.249		0.345		0.942
		(0.038)		(0.082)		(0.476)		(0.651)		(0.553)		(1.548)
Previous termination		**-0.145**		**-0.282**		0.052		-0.247		-0.618		0.248
		(0.022)		(0.058)		(0.245)		(0.321)		(0.615)		(0.480)
Amniocentesis		**-0.512**		**-0.567**		-0.696		-0.579		0.306		-1.128
		(0.055)		(0.193)		(0.463)		(0.761)		(0.961)		(1.227)
Ultrasound		**0.087**		0.084		0.121		-0.411		-0.821		-0.260
		(0.020)		(0.054)		(0.213)		(0.267)		(0.538)		(0.480)
Age dummies?	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
# observations	30,763,849	29,498,829	3,900,951	3,659,650	258,241	245,578	169,016	159,163	42,905	40,183	70,391	66,289

Estimates have been multiplied by 100 in order to be interpreted as percentage-point effects.
Standard errors are reported in parentheses.
Bold indicates statistical significance at a 5% level.

Table 12: Boy-Regression Results for California Natality Data (Unlinked), 1982–2003

	White		Black		Chinese		Indian		Japanese		Korean	
	1982–2003	1989 – 2003	1982–2003	1989 – 2003	1982–2003	1989 – 2003	1982–2003	1989 – 2003	1982–2003	1989 – 2003	1982–2003	1989 – 2003
2nd child	**-0.263**	**-0.286**	-0.056	0.162	-0.126	-0.120	0.678	**0.906**	0.480	0.507	0.329	0.459
	(0.040)	(0.051)	(0.141)	(0.189)	(0.253)	(0.306)	(0.383)	(0.450)	(0.642)	(0.824)	(0.404)	(0.516)
3rd child	**-0.355**	**-0.385**	-0.268	0.126	0.542	**1.183**	**5.718**	**6.508**	-0.991	-0.639	**1.462**	1.077
	(0.049)	(0.063)	(0.166)	(0.228)	(0.395)	(0.492)	(0.609)	(0.745)	(0.957)	(1.265)	(0.658)	(0.829)
4th child	**-0.435**	**-0.479**	-0.076	0.452	1.461	**2.882**	**8.485**	**10.003**	-1.652	-2.545	**3.941**	2.969
	(0.068)	(0.087)	(0.217)	(0.293)	(0.772)	(1.009)	(1.248)	(1.473)	(2.112)	(2.716)	(1.693)	(2.052)
Education		**0.037**		0.008		0.043		-0.106		0.197		-0.094
		(0.008)		(0.046)		(0.046)		(0.086)		(0.203)		(0.116)
Year		-0.008		0.021		0.048		0.025		0.146		0.093
		(0.005)		(0.018)		(0.034)		(0.049)		(0.090)		(0.054)
2nd-trimester care		**0.515**		**0.982**		**1.411**		-0.012		1.550		1.046
		(0.061)		(0.201)		(0.507)		(0.665)		(1.609)		(0.790)
3rd-trimester care		**0.489**		0.773		**3.446**		0.511		0.450		2.542
		(0.124)		(0.438)		(1.065)		(1.493)		(4.141)		(1.456)
No prenatal care		**0.594**		**1.366**		-2.911		-0.334		-3.319		-3.763
		(0.227)		(0.666)		(4.176)		(4.584)		(10.378)		(4.237)
Foreign-born		**-0.161**		0.010		0.099		-0.595		0.075		0.247
		(0.050)		(0.295)		(0.585)		(0.945)		(0.822)		(2.122)
Previous termination		**-0.228**		**-0.414**		-0.280		0.539		-1.652		-0.179
		(0.057)		(0.189)		(0.411)		(0.570)		(1.030)		(0.637)
Amniocentesis		**-0.391**		-0.386		0.105		0.353		1.124		-1.460
		(0.136)		(0.557)		(0.688)		(1.455)		(1.538)		(1.955)
Ultrasound		-0.056		0.099		-0.113		-0.062		**-2.231**		-0.699
		(0.043)		(0.156)		(0.289)		(0.397)		(0.795)		(0.482)
Age dummies?	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
# observations	8,289,530	5,773,398	707,053	450,695	182,087	134,349	79,151	66,102	28,606	18,562	69,686	48,885

Estimates have been multiplied by 100 in order to be interpreted as percentage-point effects.
Standard errors are reported in parentheses.
Bold indicates statistical significance at a 5% level.

For the other control variables, statistically significant estimates are found mostly for white births and black births, where the sample sizes are far larger than the Asian categories. The direction of the effects for mother's education, previous termination, amniocentesis, and ultrasound are in agreement with the predicted associations from Table 5. To focus the discussion, consider the 1991–2002 federal results in Table 11. For white births, the effect of an additional year of mother's education is a 0.046 percentage point increase in boy-birth likelihood, whereas the marginal effects of a previous termination, amniocentesis, and ultrasound are -0.145, -0.512, and 0.087 percentage points, respectively. Although very significant from a statistical viewpoint, the magnitudes of these estimated effects are far lower than the estimated third-child and fourth-child parity effects for Chinese and Indian births. Interestingly, the likelihood of white male births is lowest for mothers who have first-trimester prenatal care (0.494 percentage points lower than second-trimester care, 0.221 percentage points lower than third-trimester care, and 0.513 percentage points lower than no care), suggesting that its role as a proxy for pregnancy problems is empirically more important than its role as a proxy for quality prenatal care. Even after controlling for a variety of factors, the negative time trend in the white boy-birth probability (seen in the time-series plot of Figure 8) is still statistically significant during both the 1981–1990 period (-0.009 percentage points annually) and the 1991–2002 period (-0.007 percentage points annually). For black births, however, the positive time trend seen in Figure 8 is no longer statistically significant in the regressions with control variables.

We note some other interesting results from the control-variable regressions. The foreign-born indicator variable, which was meant to proxy for potential cultural influences, is not found to have a statistically significant effect on boy-birth likelihood for Asian births. Given the extremely high percentage of foreign-born Asian mothers (see Table 4), the lack of significance of the foreign-born indicator is perhaps not surprising due to the small amount of variation in this variable. For the other control variables, the only statistically significant estimates (at a 5% level) among the Asian births are the prenatal-care variables for Chinese births and the ultrasound indicator for Japanese California births. As is the case for white births, the likelihood of Chinese boy births is higher for second- and third-trimester prenatal care (as compared to first-trimester care); the magnitude of these differences (compared to white births) is, however, larger in the 1991–2002 federal sample and 1989–2003 California sample. For Japanese births in the 1989–2003 California sample, the likelihood of a boy birth is estimated to be 2.231 percentage points lower when the mother had an ultrasound during pregnancy. The direction of this association is contrary to what would be expected if ultrasound proxies for quality prenatal care and, if anything, would be consistent with gender selection in favor of daughters.

In an attempt to gauge the effect of cultural influences on boy likelihoods, Table 13 compares the birth-parity estimates from two different samples: (1) births to mothers of a given race and (2) births to parents of the same race. The same-race samples, where were used for the regressions in Tables 9–12, represent a subset of the samples based only upon mother's race. For the six race categories considered, the table reports the birth-parity estimates from the control-variable regressions for the two samples. (The other coefficient estimates are omitted in the interest of space.) The regression specifications are identical to those in Tables 9–12, meaning that the same-race-sample estimates are also identical. These estimates are provided in the second column of each race panel. The first-column estimates correspond to the samples based upon mother's race only. For white births, the birth-parity effects are extremely similar for the two samples in the 1971–1980 and 1981–1990 federal regressions; for the 1991–2002 federal and 1989–2003 California regressions, there is a slight divergence of the estimates with a slightly higher likelihood of daughters at later births for same-race white parents. The results for Asian births exhibit a divergence in the opposite direction. In every case where birth parity has a statistically significant effect (at a 5% level) for the Chinese, Indian, and Korean samples, the estimated magnitude of the birth-parity effect is larger (more likely to have a boy) for the same-race-parent samples.

4.4 Analysis of Linked California Natality Data

This section utilizes the maternally linked version of the California natality data (described in Section 3 and the Appendix B) in order to determine the relationship, if any, between previous gender(s) of a mother's child(ren) and future birth outcomes. As in the primary analysis of Section 4.3, the analysis of the linked California data will be restricted to those births for which parents are of the same race. The analysis is focused upon second- and third-birth outcomes, as the number of births for most races becomes too small at higher birth parities.

Table 14 provides a brief summary of the gender mix of previous children for second and third births, broken down by race. The results are in agreement with the fertility-stopping results from the Census data previously reported in Tables 6 and 7. There appears to be no strong evidence of gender preferences affecting the decision to have a second child, with the percentages reported in the first row of Table 14 being very close to the overall percentage of firstborn sons. As in the Census data, the gender preferences become evident at the decision to have a third birth. A gender-mix preference across races is evident from the low percentages in the "1 son in first 2 births" category; this percentage would be roughly equal to 50% if there were complete gender indifference. The preference for sons is evident among Chinese, Indian, and Korean parents, where the percentages in the "0 sons in first 2 births" category are quite high (32.72%, 35.66%, and 31.90%, respectively).

Table 13: Comparison of Birth-Parity Effects

Data	Child	White		Black		Chinese		Indian		Japanese		Korean	
		Mother	Mother and Father	Mother	Mother and Father	Mother	Mother and Father	Mother	Mother and Father	Mother	Mother and Father	Mother	Mother and Father
Federal	2nd	**-0.102**	**-0.099**	-0.143	-0.148	-0.184	0.267			**2.349**	**2.089**		
1971–		(0.034)	(0.034)	(0.081)	(0.101)	(0.649)	(0.744)			(0.686)	(0.940)		
1980	3rd	**-0.161**	**-0.167**	-0.065	-0.069	-0.533	-0.746			**2.115**	2.571		
		(0.046)	(0.047)	(0.105)	(0.126)	(0.963)	(1.096)			(0.982)	(1.367)		
	4th	**-0.325**	**-0.310**	-0.113	-0.202	-0.539	-0.560			1.865	-0.209		
		(0.069)	(0.071)	(0.140)	(0.165)	(1.686)	(1.911)			(1.802)	(2.751)		
Federal	2nd	**-0.076**	**-0.077**	-0.114	-0.078	-0.126	-0.526			-0.345	0.630		
1981–		(0.025)	(0.026)	(0.059)	(0.076)	(0.370)	(0.413)			(0.510)	(0.718)		
1990	3rd	**-0.131**	**-0.132**	**-0.151**	-0.106	1.035	1.242			-0.706	-0.645		
		(0.034)	(0.035)	(0.073)	(0.093)	(0.576)	(0.648)			(0.737)	(1.071)		
	4th	**-0.163**	**-0.173**	-0.159	-0.073	0.451	1.113			0.210	-0.033		
		(0.053)	(0.055)	(0.099)	(0.125)	(1.091)	(1.240)			(1.379)	(2.284)		
Federal	2nd	**-0.089**	**-0.123**	-0.093	0.043	0.093	0.294	**0.914**	**0.975**	0.028	0.444	0.317	0.558
1991–		(0.021)	(0.021)	(0.050)	(0.065)	(0.200)	(0.226)	(0.274)	(0.290)	(0.362)	(0.560)	(0.377)	(0.443)
2002	3rd	**-0.181**	**-0.221**	-0.068	0.147	**0.826**	**1.329**	**3.553**	**4.009**	0.802	0.438	**1.228**	**1.803**
		(0.027)	(0.026)	(0.062)	(0.080)	(0.336)	(0.382)	(0.441)	(0.471)	(0.548)	(0.867)	(0.596)	(0.715)
	4th	**-0.261**	**-0.334**	-0.157	0.073	**2.070**	**2.689**	**2.857**	**3.488**	0.359	-2.544	1.079	0.744
		(0.040)	(0.039)	(0.083)	(0.109)	(0.706)	(0.825)	(0.840)	(0.908)	(1.028)	(1.862)	(1.306)	(1.735)
California	2nd	**-0.267**	**-0.286**	0.135	0.162	-0.170	-0.120	**0.870**	**0.906**	-0.240	0.051	0.077	0.459
1989–		(0.049)	(0.051)	(0.174)	(0.189)	(0.271)	(0.306)	(0.430)	(0.450)	(0.519)	(0.824)	(0.455)	(0.516)
2003	3rd	**-0.364**	**-0.385**	0.120	0.126	0.814	**1.183**	**5.729**	**6.508**	-1.201	-0.639	0.835	1.077
		(0.061)	(0.063)	(0.210)	(0.228)	(0.439)	(0.492)	(0.711)	(0.745)	(0.797)	(1.265)	(0.722)	(0.829)
	4th	**-0.439**	**-0.479**	0.337	0.452	**2.416**	**2.882**	**8.554**	**10.003**	-0.986	-2.545	2.158	2.969
		(0.084)	(0.087)	(0.271)	(0.293)	(0.895)	(1.009)	(1.389)	(1.473)	(1.556)	(2.716)	(1.653)	(2.052)

Table 14: Gender Mix for Linked California Natality Data

	White	Black	Chinese	Indian	Japanese	Korean
% of mothers with 2+ births having a firstborn-son	51.48%	50.68%	51.44%	50.39%	50.53%	50.74%
% of mothers with 3+ births with:						
0 sons in first 2 births	25.78%	25.70%	32.72%	35.66%	26.55%	31.90%
1 son in first 2 births	46.04%	47.28%	40.82%	43.54%	44.00%	41.84%
2 sons in first 2 births	28.18%	27.01%	26.46%	20.80%	29.45%	26.27%

Table 15 provides summary statistics for four indicator variables (male-child, ultrasound, amniocentesis, and termination-since-last-birth), with second-child and third-child averages broken down by the number of previous sons that a mother has had. The termination-since-last-birth indicator variable was constructed by comparing the number of previously terminated pregnancies reported in two successive pregnancies; in particular, the variable was set equal to one if the number reported at the later pregnancy was larger than the number reported at the earlier pregnancy, and zero otherwise. To determine the statistical significance of any differences seen in the sample averages of Table 15 and to control for other observables, Tables 16 and 17 report regression results for second-birth and third-birth outcomes, respectively.

Table 16 considers two specifications, one with only an indicator variable for a firstborn-girl child ("no covariates" specification) and one that also includes the full set of control variables considered in the regressions of Section 4.3 ("covariates" specification). Only the coefficient estimate of the firstborn-girl indicator variable is reported for each of the regressions. As discussed in Section 4.1, differential fetal survival would lead one to expect that mothers who have previously given birth to a son (daughter) would be more (less) likely to give birth to another son. This relationship is found to be statistically significant among both white and black births. In the specifications with control variables, white mothers are 0.448 percentage points more likely to give birth to a second-child boy if her first child was a boy, whereas the analogous difference for black mothers is 0.583 percentage points. No such significant effects are found among Chinese, Japanese, or Korean births. However, among Indian births, a mother is 2.089 percentage points more likely to give birth to a second-child son if her first child was a girl.

For the other three dependent variables (ultrasound, amniocentesis, and termination-since-last birth) considered in Table 16, only three estimates (for the firstborn-girl variable) are significant at a 10% level in the control-variable regressions. Japanese mothers were 2.709 percentage points

Table 15: Summary Statistics for Linked California Natality Data

Variable	Race	2nd child average, if 1st child is a:		3rd child average, if # of previous boys is:		
		Girl	Boy	0 boys	1 boy	2 boys
Male child	White	0.508	0.514	0.503	0.511	0.514
	Black	0.503	0.510	0.505	0.510	0.509
	Chinese	0.519	0.517	0.527	0.518	0.515
	Indian	0.529	0.513	0.634	0.525	0.533
	Japanese	0.507	0.523	0.488	0.520	0.509
	Korean	0.518	0.520	0.511	0.517	0.516
Ultrasound	White	0.523	0.524	0.533	0.531	0.536
	Black	0.456	0.456	0.457	0.463	0.467
	Chinese	0.513	0.514	0.539	0.492	0.530
	Indian	0.590	0.583	0.629	0.592	0.609
	Japanese	0.644	0.621	0.685	0.645	0.619
	Korean	0.382	0.377	0.408	0.433	0.409
Amniocentesis	White	0.032	0.033	0.031	0.029	0.032
	Black	0.019	0.019	0.018	0.018	0.019
	Chinese	0.058	0.057	0.097	0.087	0.082
	Indian	0.029	0.023	0.039	0.047	0.030
	Japanese	0.091	0.092	0.162	0.115	0.132
	Korean	0.016	0.012	0.039	0.023	0.032
Termination since last birth	White	0.122	0.122	0.120	0.121	0.122
	Black	0.148	0.148	0.144	0.148	0.149
	Chinese	0.098	0.101	0.104	0.098	0.095
	Indian	0.109	0.100	0.137	0.091	0.087
	Japanese	0.121	0.116	0.116	0.115	0.152
	Korean	0.120	0.116	0.120	0.125	0.122

more likely to have an ultrasound (p-value of 0.047) prior to their second birth if their first child was a girl. Indian mothers were 0.488 percentage points more likely to have an amniocentesis (p-value of 0.063) and 0.921 percentage points more likely to have a terminated pregnancy since their first birth (p-value of 0.078) if their first child was a daughter.

Analogous to Table 16, Table 17 reports two different regression specifications (with and without additional covariates). For both specifications, two indicator variables (a "no-sons indicator" and a "one-son indicator") are included and should be interpreted as differences from mothers with two sons. For white births, the slight gender persistence effect is again estimated to be statistically significant; white mothers with two sons are 0.881 percentage points more likely to give birth to a third-child son than white mothers with two daughters. The only other statistically significant estimate in the male-child regression results is for Indian births, where Indian mothers with no sons are 12.272 percentage points more likely (p-value of 0.000) to have a third-child son than Indian mothers with two sons. Indian mothers with two daughters are also significantly more likely (4.360 percentage points, p-value of 0.022) to have a terminated pregnancy between their second and third births than Indian mothers with two sons. The other effects that are found to be significant at a 5% level are the following: lower ultrasound usage for Chinese mothers with one son, higher ultrasound usage for Japanese mothers with two daughters, and lower amniocentesis usage for white mothers with one son.

Finally, to determine whether there exists more direct evidence of gender-selective practices, Table 18 considers boy regressions on subsamples of births for which there was either a terminated pregnancy since the last birth or an ultrasound performed during the pregnancy. In particular, three different subsamples are considered: (1) mothers who had a terminated pregnancy since last birth, (2) mothers who had an ultrasound during pregnancy, and (3) mothers who had an ultrasound during pregnancy and no ultrasound during their previous pregnancy. Since ultrasound usage may simply proxy for good prenatal care, the third subsample is considered in order to focus upon those mothers for whom the ultrasound usage would be more likely to be for gender-determinative purposes (as compared to the second subsample). For each of the three subsamples (and each of the six racial categories), results are reported for the firstborn-girl indicator variable in the second-child boy regression and the no-sons and one-son indicator variables in the third-child boy regression. The other control variables used in the regressions were birthyear, age, age squared, education, and the prenatal-care indicator variables.

No statistically significant estimates (at a 10% level) are found among black, Chinese, or Japanese births. For white mothers with a terminated pregnancy since last birth, the third-child regression indicates that there is a significantly greater chance (1.398 percentage points) of having a boy when they have a boy-girl mix than when they have two boys. Since boy-birth persistence

Table 16: Second-Child Regressions for Linked California Natality Data

Dependent variable	Race	Coefficient estimate for firstborn-girl indicator variable: No Covariates	Covariates
Male child	White	-0.629 (0.077)**	-0.488 (0.086)**
	Black	-0.691 (0.276)**	-0.583 (0.312)*
	Chinese	0.197 (0.506)	0.324 (0.552)
	Indian	1.632 (0.813)**	2.089 (0.853)**
	Japanese	-1.599 (1.290)	-0.550 (1.474)
	Korean	-0.132 (0.864)	-0.080 (0.959)
Ultrasound	White	-0.103 (0.085)	-0.036 (0.084)
	Black	-0.215 (0.308)	-0.150 (0.301)
	Chinese	-0.103 (0.548)	0.305 (0.529)
	Indian	0.671 (0.833)	0.545 (0.833)
	Japanese	2.290 (1.412)	2.709 (1.366)**
	Korean	0.515 (0.924)	0.532 (0.914)
Amniocentesis	White	-0.053 (0.030)*	-0.015 (0.029)
	Black	0.026 (0.085)	0.037 (0.084)
	Chinese	0.083 (0.256)	0.029 (0.249)
	Indian	0.591 (0.268)**	0.488 (0.263)*
	Japanese	-0.049 (0.845)	0.190 (0.819)
	Korean	0.472 (0.224)**	0.338 (0.222)
Termination since last birth	White	0.049 (0.050)	0.066 (0.057)
	Black	-0.036 (0.196)	-0.046 (0.225)
	Chinese	-0.292 (0.303)	-0.428 (0.332)
	Indian	0.905 (0.499)*	0.921 (0.523)*
	Japanese	0.573 (0.835)	1.005 (0.953)
	Korean	0.437 (0.401)	0.271 (0.628)

**: significant at a 5% level; *: significant at a 10% level.

Estimates have been multiplied by 100, in order to be interpreted as percentage-point effects. The covariates included were birthyear, age, age squared, education, and the prenatal-care indicator variables. The “male child” regression also includes indicators for ultrasound and amniocentesis.

Table 17: Third-Child Regressions for Linked California Natality Data

Dependent variable	Race	Regression without covariates		Regression with covariates	
		Estimate for no-sons indicator	Estimate for one-son indicator	Estimate for no-sons indicator	Estimate for one-son indicator
Male child	White	-1.087 (0.187)**	-0.307 (0.164)*	-0.881 (0.197)**	-0.152 (0.172)
	Black	-0.355 (0.624)	0.157 (0.547)	-0.154 (0.662)	0.198 (0.580)
	Chinese	1.146 (1.649)	0.289 (1.575)	1.391 (1.723)	0.294 (1.653)
	Indian	10.130 (2.897)**	-0.820 (2.838)	12.272 (2.992)**	0.348 (2.939)
	Japanese	-2.032 (3.949)	1.170 (3.493)	-2.071 (4.300)	2.209 (3.759)
	Korean	-0.451 (3.085)	0.140 (2.916)	-0.895 (3.238)	0.541 (3.049)
Ultrasound	White	-0.308 (0.195)	-0.477 (0.171)**	-0.067 (0.191)	-0.215 (0.168)
	Black	-0.926 (0.654)	-0.379 (0.574)	-1.175 (0.638)*	-0.404 (0.559)
	Chinese	0.898 (1.703)	-3.854 (1.634)*	1.000 (1.661)	-3.229 (1.589)**
	Indian	2.030 (2.930)	-1.641 (2.855)	1.641 (2.930)	-1.347 (2.845)
	Japanese	6.541 (4.020)	2.567 (3.583)	7.999 (3.933)**	3.090 (3.532)
	Korean	-0.134 (3.157)	2.410 (2.994)	-0.207 (3.137)	1.791 (2.967)
Amniocentesis	White	-0.114 (0.068)*	-0.324 (0.059)**	0.012 (0.066)	-0.123 (0.058)**
	Black	-0.131 (0.176)	-0.137 (0.155)	-0.219 (0.175)	-0.167 (0.154)
	Chinese	1.492 (0.972)	0.488 (0.909)	1.484 (0.951)	0.610 (0.893)
	Indian	0.923 (1.083)	1.665 (1.078)	0.776 (1.086)	1.677 (1.064)
	Japanese	2.909 (3.007)	-1.698 (2.459)	3.754 (2.941)	-1.668 (2.414)
	Korean	0.693 (1.179)	-0.931 (1.006)	0.254 (1.161)	-1.149 (1.023)
Termination since last birth	White	-0.141 (0.122)	-0.065 (0.107)	-0.067 (0.129)	0.025 (0.114)
	Black	-0.471 (0.441)	-0.052 (0.389)	-0.489 (0.473)	-0.029 (0.418)
	Chinese	0.901 (0.985)	0.296 (0.928)	1.121 (1.036)	0.896 (0.985)
	Indian	5.007 (1.804)**	0.357 (1.615)	4.360 (1.897)**	-0.448 (1.686)
	Japanese	-3.577 (2.684)	-3.704 (2.402)	-2.068 (2.995)	-4.599 (2.556)*
	Korean	-0.173 (2.013)	0.324 (1.919)	-0.041 (2.132)	0.586 (2.024)

**: significant at a 5% level; *: significant at a 10% level.

Estimates have been multiplied by 100, in order to be interpreted as percentage-point effects. The covariates included were birthyear, age, age squared, education, and the prenatal-care indicator variables. The "male child" regression also includes indicators for ultrasound and amniocentesis.

would lead one to expect a boy birth to be slightly more likely with two previous sons, this estimate would be consistent with a situation in which a small proportion of white mothers who already have two sons are practicing gender selection in favor of third-child daughters. On the other hand, there is little evidence to suggest gender selection in favor of third-child sons when white mothers already have two daughters. White mothers who have an ultrasound are more likely to give birth to a son when they have previously given birth to sons (0.501 percentage points more likely in the second-child regression, and a 0.869 percentage-point difference between no sons and two sons in the third-child regression). These estimates are extremely similar to those in Tables 16 and 17. Interestingly, the statistical significance of the estimates disappears in the third white subsample.

As in the male-child results of Tables 16 and 17, the estimates on the firstborn-girl and the no-sons indicator variables for the Indian births are statistically significant at a 5% level (except for the firstborn-girl estimate in the third subsample). These estimates are larger in magnitude than the estimates for the full sample. For instance, among Indian mothers who had a terminated pregnancy since last birth, there is a larger chance (6.442 percentage points) of having a boy when the first child was a girl (compared to an effect of 2.089 percentage points in the full sample). Within the analogous third-child subsample, Indian mothers with no sons are 21.340 percentage points more likely to give birth to a son than Indian mothers with two sons (as compared to a 12.272 percentage-point difference in the full sample). The third-child regression estimates for the two ultrasound subsamples are also larger in magnitude than those for the full sample. Unlike the estimates for white births, the statistical significance of the estimates remains in the third Indian subsample, although the standard errors nearly double due to the lower sample size. Finally, note that the estimates on the firstborn-girl and no-sons indicator variables are significant at a 10% level for the third Korean subsample, even though no significant estimates of previous gender had been found in the male-child regressions on the full sample of Korean births in Tables 16 and 17.

4.5 Inferring the Prevalence of Gender Selection from Boy-Birth Percentages

In this section, the following question is considered: If unusual boy-birth percentages are the result of gender-selective abortions, what does the observed boy-birth percentage imply about the prevalence of both gender determination and gender selection? Without loss of generality, consider the case where the gender bias favors sons and gender-selective abortion is only chosen when the female gender is revealed.[25] As in Section 2, let p denote the "natural" probability of a boy birth (in the absence of gender determination and/or selection). Let g denote the probability that a woman has a gender-determinative procedure (meaning that gender-selective abortion would be chosen if

[25]If the reverse is true for a subgroup of the population (daughter bias and gender-selective abortion only for males), the prevalence of gender determination/selection discussed below would be a lower bound on the actual prevalence.

Table 18: Boy Regressions on Subsamples

Subsample	Race	2nd-Child Regression		3rd-Child Regression		
		Firstborn girl	# obs	No-sons indicator	One-son indicator	# obs
Mothers who had a terminated pregnancy since last birth	White	0.223	170,851	0.694	1.398**	59,621
		(0.242)		(0.558)	(0.489)	
	Black	0.084	15,975	0.111	1.941	6,601
		(0.791)		(1.704)	(1.484)	
	Chinese	1.094	3,359	6.952	6.829	586
		(1.731)		(5.417)	(5.237)	
	Indian	6.442**	1,459	21.340**	1.024	225
		(2.623)		(9.405)	(10.027)	
	Japanese	-2.627	557	-11.736	9.324	132
		(4.273)		(11.341)	(10.591)	
	Korean	3.495	1,340	4.722	0.773	213
		(2.733)		(9.305)	(8.764)	
Mothers who had an ultrasound during pregnancy	White	-0.501**	708,471	-0.869**	-0.056	256,914
		(0.119)		(0.269)	(0.236)	
	Black	-0.782	47,224	0.150	-0.203	20,130
		(0.460)		(0.972)	(0.849)	
	Chinese	0.326	16,903	2.022	1.692	2,970
		(0.769)		(2.352)	(2.304)	
	Indian	2.408**	8,089	16.710**	3.339	1,276
		(1.110)		(3.786)	(3.763)	
	Japanese	-0.739	2,919	-2.428	4.038	653
		(1.853)		(5.301)	(4.731)	
	Korean	1.644	4,144	3.356	-4.048	699
		(1.559)		(5.083)	(4.737)	
Mothers who had an ultrasound during pregnancy and no ultrasound during previous pregnancy	White	0.120	197,036	0.088	0.008	72,460
		(0.225)		(0.508)	(0.445)	
	Black	0.021	14,783	1.472	-0.757	6,888
		(0.823)		(1.672)	(1.459)	
	Chinese	-0.112	5,406	4.033	1.194	903
		(1.360)		(4.248)	(4.249)	
	Indian	2.824	2,231	24.039**	15.761**	330
		(2.112)		(7.016)	(7.215)	
	Japanese	3.000	576	-11.697	12.213	121
		(4.204)		(14.181)	(10.860)	
	Korean	4.768*	1,528	14.076*	-7.171	271
		(2.568)		(7.794)	(7.760)	

**: significant at a 5% level; *: significant at a 10% level.

Estimates have been multiplied by 100, in order to be interpreted as percentage-point effects. The additional covariates included were birthyear, age, age squared, education, and prenatal-care indicator variables.

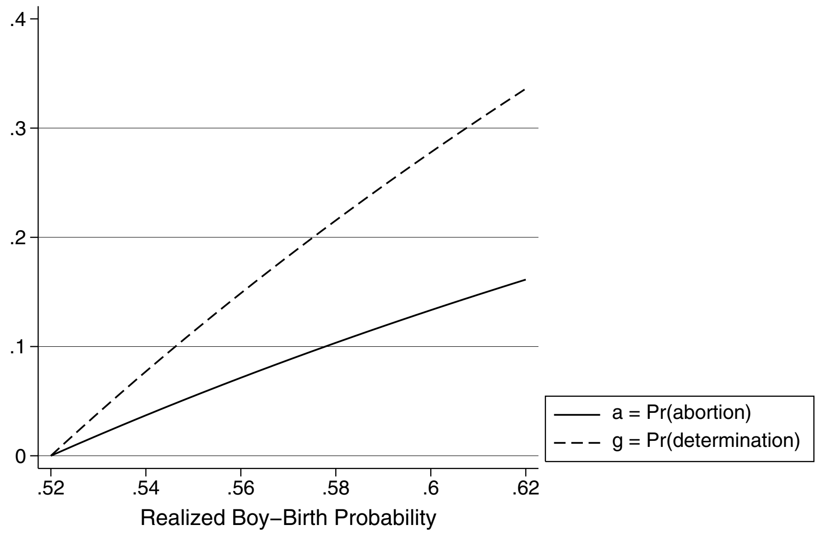

Figure 9: Implied Prevalence of Gender Determination/Selection (for $p = 0.52$)

female gender is revealed).[26] For simplicity, assume that the gender-selective procedure has no associated risk of involuntary termination ($q = 0$ in the notation of Section 2). If a denotes the probability that a woman has a gender-selective abortion, it immediately follows that $a = g(1-p)$. Finally, let $\tilde{p}$ denote the boy-birth probability in the presence of gender-selective practices. The probability $\tilde{p}$ is the quantity corresponding to the boy-birth percentage *observed* in the data. Note that $\tilde{p}$ is related to p and a as follows:

$$\tilde{p} = \frac{\text{Pr(boy birth)}}{\text{Pr(live birth)}} = \frac{p}{1-a}. \tag{15}$$

Equivalently, a and g can be written in terms of the probabilities p and $\tilde{p}$ as follows:

$$a = \frac{\tilde{p} - p}{\tilde{p}} \quad \text{and} \quad g = \frac{\tilde{p} - p}{\tilde{p}(1-p)}. \tag{16}$$

To infer anything about the prevalence of gender determination and gender-selective abortion (g and a, respectively), a value for the "natural" boy-birth probability (p) is needed. A conservative choice of p, based upon the first-birth boy percentages reported in Table 8, is $p = 0.52$.[27] For this value of p and realized boy-birth probabilities ($\tilde{p}$) ranging from 0.52 to 0.62, Figure 9 shows the implied probabilities of gender determination and gender-selective abortion. As an illustration, consider the boy-birth percentages for Indian births reported in Table 8. In the federal natality data, the fraction of boy births among third and fourth children is approximately 0.545. If this

[26] For instance, if *all* pregnant women had a gender-revealing ultrasound performed, g would represent the fraction of women who *would* have a gender-selective abortion if a female is revealed.

[27] An increase in p is related to a decrease in both g and a.

higher percentage is the result of gender selection, Figure 9 indicates that the probabilities of gender determination and gender-selective abortion consistent with this percentage are approximately 9.6% and 4.6%, respectively.[28] For the California estimates (56.7% boy-birth percentage for third children and 59.4% for fourth children), the implied determination and abortion probabilities would be 17.3% and 8.3%, respectively, for third children and 26.0% and 12.5%, respectively, for fourth children.

5 Conclusion

This study has provided empirical evidence consistent with gender selection at later births within the United States. For Chinese and Indian parents, the likelihood of having a son is significantly higher for third-born and fourth-born children as compared to first-born children.[29] Controlling for maternal characteristics, prenatal-care variables, and a time trend, the increase in boy-birth likelihood explained by birth parity is extremely significant and of an order of magnitude larger than other determinants. The birth-parity effects for Chinese and Indian births are found to be larger for same-race couples as compared to samples based on mother's race only. On the other hand, only slight evidence of birth-parity effects is found among Korean births and no evidence is found among Japanese births.

The evidence from the California natality data is particularly striking for Indian births: second-born children are 0.9 percentage points more likely to be boys, third-born children 6.5 percentage points more likely, and fourth-born children 10.0 percentage points more likely. Moreover, Indian parents are significantly more likely to have a boy and a terminated pregnancy (since the last live birth) if they have had only daughters previously. For instance, Indian parents with two daughters are 12.3 percentage points more likely to have a third-born son than Indian parents with two sons and 4.4 percentage points more likely to have a terminated pregnancy since the last birth. The simple framework of Section 4.5 suggests that the unusually high boy percentages among third- and fourth-born Indian children in California would be consistent with gender-selective abortion rates of around 10% (and gender-determination rates of around 20%).

Gender selection seems like a logical candidate for the observed high boy-birth percentages at later births, particularly since the analysis has controlled for maternal and prenatal factors.

[28] If p is taken to be 0.51, which is the observed percentage of first-birth boys for Indian parents in the federal and California samples, the implied probabilities of gender determination and gender-selective abortion would of course be higher — 13.4% and 6.4%, respectively.

[29] Although it is also possible that gender selection occurs among first-born children, the existing data do not support this conclusion. For Chinese births (see Table 8), there has been almost no change since 1971 in the boy-birth percentage among first-born and second-born children. Unfortunately, such a time-series comparison is infeasible for Indian and Korean births since data is not available prior to 1992 at the federal level and 1982 at the California level.

Nevertheless, given that that the empirical evidence provides only a weak link between terminated pregnancies and boy-birth likelihoods, it is important to consider other possible explanations. In particular, a recent study by Oster (2005) suggests that the high incidence of hepatitis B in many Asian countries, particularly China, can partly explain the unusually high observed boy-birth percentages. Oster (2005) estimates that hepatitis B (which is associated with a higher boy-birth likelihood) can account for around 75% of the "missing women" in China as opposed to around 17% of the "missing women" in India. Fortunately, the California natality data allow us to check whether hepatitis B might offer an explanation for our findings, as births between 1989 and 2003 have an indication of whether the mother was either infected with or a carrier of the hepatitis B virus. (No such information for fathers is available.) The percentage of hepatitis B carriers among the six races considered in this study (for 1989–2003 births) are as follows: whites 0.07%, blacks 0.13%, Chinese 1.49%, Indian 0.12%, Japanese 0.15%, and Korean 0.65%.[30] We suspect that the significantly higher incidence of hepatitis B among Chinese and Korean mothers explains at least part of the overall higher boy-birth percentages that are observed for these two races (around 51.7% for first-born children in the United States, as reported in Table 8). On the other hand, hepatitis B does not seem like a plausible explanation for the other findings of this study. First, for hepatitis B to account for the estimated birth-parity effects, it would have to be the case that its incidence becomes significantly higher at later births. The California data, however, suggest no such trends in hepatitis B prevalence at later births for any of the races.[31] In addition, inclusion of an indicator variable for hepatitis B carrying mothers had nearly no quantitative effect on the birth-parity estimates reported in Section 4.[32] Second, the most significant findings of Section 4 involve Indian births, whereas the "hepatitis B effect" of Oster (2005) is found to be relatively smaller in India than in China. The reported incidence of hepatitis B among Indian mothers in California is similar to the reported incidence among black mothers in California.

Overall, the empirical findings are in line with the gender preferences documented with Census data in Section 4.2 and the stronger incentives for gender selection that arise at later births. For Chinese, Indian, and Korean families, the Census data indicate a strong son bias that appears with the decision to have a third child, with a much higher likelihood of having a third child among families with two daughters. In contrast, the third-child outcomes from the Census data indicate a preference for a gender mix among white, black, and Japanese families. Despite the gender-mix preference that appears in the fertility decisions for these races, the empirical results

[30]The incidence of hepatitis B is undoubtedly under-reported in the California data. However, the arguments given below would only become invalid if the level of mis-reporting is systematically related to race and/or birth order.

[31]From a fertility-stopping viewpoint, one would expect hepatitis B carriers (who are more likely to have sons) to have fewer children. Therefore, the percentage of carriers would be lower at higher birth parity in the aggregate.

[32]Interactions of the hepatitis B indicator with the birth-parity indicators were also included, again with no meaningful impact on the results reported.

do not provide much evidence of unusual boy-birth percentages that would suggest that gender selection is being used to achieve a gender mix. In particular, the aggregate birth-parity effects for white parents (estimated in Section 4.3) do not change much from the 1971–1980 time period to later time periods. In addition, the effect of previous gender(s) on gender outcomes for white parents (estimated in Section 4.4) is in agreement with the gender persistence that would be expected from the differential fetal survival rates between males and females. The only evidence consistent with gender selection concerns the third-child outcomes for white mothers in California: among mothers who had a terminated pregnancy between their second and third child, those with two sons had a higher than expected likelihood of having a daughter (1.4 percentage points more likely than mothers with a son and a daughter). Given the small magnitude of this differential and the lack of other systematic evidence, the prevalence of gender selection for the purposes of attaining gender mix appears to be limited. Further research on this point would certainly be interesting.

Although no previous studies have provided empirical evidence concerning the prevalence of gender selection in the United States, there have been a few studies that have examined attitudes toward gender determination and gender selection in the United States. For instance, Wertz and Fletcher (1998) report the following results from a 1996 survey, which included responses from 1083 U. S. geneticists and a random sample of 1000 U. S. citizens:[33]

- 62% of U. S. geneticists had received an outright request for sex selection by prenatal diagnosis. (75% reported that they had received a "suspected request" under alternative pretenses).
- When asked about the hypothetical case of a couple with four daughters who desire a son and will abort a female fetus, 34% of the U. S. geneticists would perform prenatal diagnosis for sex selection and an additional 38% would refer the couple to a geneticist that would.
- 26% of the U. S. public sample respondents would use a "safe and accurate method of preconception sex selection ... such as separation of X and Y bearing sperm." 40% of respondents thought that such a method should be available without restrictions.

In a more recent study, Jain et. al. (2005) questioned infertility patients about sex-selection procedures. Their sample is particularly interesting, given the inherently high cost of pregnancy and lower chance of future pregnancies among infertility patients. Of the 561 respondents, 40.8% indicated that they would want to select the sex of their next baby at no additional cost, with significantly higher percentages among those with either no children or children all of one gender. Of those desiring to gender select, 61.1% wanted to have a daughter and 55.0% would use gender-selective IVF (as opposed to sperm sorting).

[33]The breakdown of the geneticists was as follows: 32% M.D., 18% Ph.D., 34% M.S. ("genetic counselor"), and 16% other.

Several factors could lead to an increase in the prevalence of gender selection within the United States. First, if the declining trend in family size continues, there would be increased incentives (holding gender preferences fixed) for gender selection. Second, the availability of improved preconceptive gender-selective technologies at lower costs will tend to increase the prevalence of gender selection.[34] Most importantly, a preconceptive gender selection method eliminates the need for a gender-based abortion, which involves prohibitive costs (including moral and ethical costs) for most parents.

Although the gender-mix preference among whites and blacks in the United States is not likely to change much in the near future, it is possible that the son bias observed among some of the Asian races (Chinese, Indian, and Korean) could diminish. Such a change could occur for a variety of reasons, including reduced cultural bias toward sons (as has occurred in Japan over the last few decades) and a larger proportion of second- and third-generation Asian mothers in the United States.

What about the possible effects of increased gender selection in the United States? Given that the predominant preference is for a gender mix, gender selection would probably not lead to a gender-imbalance problem in the aggregate. Such a gender imbalance could, however, arise among subpopulations with a bias toward sons or daughters. The effect on family size would be ambiguous, as suggested by the model in Section 2: although families could achieve gender mix with fewer children, some families would be willing to have additional children if they could choose gender. Given that gender-selective procedures are not banned in the United States, the most predictable effect of increased gender selection would be the resulting debate on the surrounding moral and ethical issues and potentially the fight over regulation.[35]

Appendix

Appendix A: Proofs and other material for the theoretical model

For the material in this appendix section, the simplifying notation introduced at the beginning of Section 2.4 is used. Proofs for the propositions and lemmas in Section 2 are provided, as well as the statement of two of the results from Section 2.4.

Proof of Proposition 1: For period T, the expected utilities associated with no pregnancy and pregnancy are 0 and $pU_b + (1-p)U_g$, respectively, so that the result clearly holds. If $pU_b + (1-p)U_g$ is negative, then $V_T = V_T^b = V_T^g = 0$ and pregnancy is not chosen in period $T-1$; this same reasoning holds for all periods

[34] A recent report by the Centers for Disease Control and Prevention (2004) documented the increased use of "assisted reproductive technology" (defined as fertility treatments involving both sperm and eggs, predominantly IVF). The number of live-birth deliveries using this technology increased steadily from 14,507 in 1996 to 33,141 in 2002. Assisted reproductive technology currently accounts for roughly 1% of live births in the United States.

[35] The President's Council on Bioethics considered some of these issues at its October 2002 meeting. Full transcripts are available at http://www.bioethics.gov/transcripts/oct02/index.html.

back through $t = 1$. If $pU_b + (1-p)U_g$ is positive, then $V_T = pU_b + (1-p)U_g$ and pregnancy is chosen in period $T-1$ (since $V_T^b \geq 0$, $V_T^g \geq 0$, and $\delta < 1$); for earlier periods, note that if pregnancy is chosen in period $t+1$, the expected utility of no pregnancy in period t is $\delta(p(U_b + \delta V_{t+2}^b) + (1-p)(U_g + \delta V_{t+2}^g))$ which is smaller than the expected utility of pregnancy (since $\delta < 1$, $V_{t+2}^b \leq V_{t+1}^b$, and $V_{t+2}^g \leq V_{t+1}^g$).

Proof of Lemma 1: When gender is known, the expected utility of no termination is $U_b - d$ for a boy ($U_g - d$ for a girl) whereas the expected utility of termination is $-d-c$. The results follow immediately.

Proof of Lemma 2: For case (i), the pregnancy would not be terminated if either gender were revealed (by Lemma 1); gender would not be determined since $d > 0$. For case (ii), the pregnancy would only be terminated if female gender is revealed (by Lemma 1). The expected utility from gender determination and no gender determination would be $-d - qc + (1-q)pU_b - (1-q)(1-p)c$ and $pU_b + (1-p)U_g$, respectively, so that the result follows by subtracting $(1-q)pU_b$ from both expressions. Case (iii) is similar. For case (iv), the pregnancy would be terminated for either gender; gender is determined when the expected incremental utility of a birth ($pU_b + (1-p)U_g$) is less than the (dis)utility of determination and termination ($-d-c$).

Proof of Proposition 2: The utility from not becoming pregnant in period T is equal to zero. The expected utility from becoming pregnant is equal to the maximum of the expected utilities from (i) becoming pregnant and determining gender and (ii) becoming pregnant and not determining gender (with the optimal gender-determination decision given by Lemma 2). If $pU_b + (1-p)U_g$ is positive, then the woman becomes pregnant and, by Lemma 2, determines gender if either

$$U_b > -c, U_g < -c, \text{ and } pqU_b + (1-p)U_g < -d - qc - (1-q)(1-p)c,$$

or

$$U_b < -c, U_g > -c, \text{ and } pU_b + (1-p)qU_g < -d - qc - (1-q)pc.$$

In the first case, the first and third inequalities together imply that $U_g < -c - d/(1-p)$, making the second inequality vacuous (i.e., it can't be a binding constraint), and the second inequality (together with $pU_b + (1-p)U_g > 0$) implies $U_b > (1-p)c/p$, making the first inequality vacuous. Similarly, the first two inequalities in the second case are also vacuous. The first region in the proposition (pregnancy and no gender determination) then follows immediately. If $pU_b + (1-p)U_g$ is negative, then the woman only becomes pregnant (and determines gender) if the expected utility associated with gender determination is positive, which occurs if either

$$U_b > -c, U_g < -c, \text{ and } U_b > \frac{d + (1-(1-q)p)c}{(1-q)p},$$

or

$$U_b < -c, U_g > -c, \text{ and } U_g > \frac{d + (q + (1-q)p)c}{(1-q)(1-p)}.$$

Note that the first inequality in the first case and the second inequality in the second case are vacuous. Also, the negative value of $pU_b + (1-p)U_g$ combined with the third inequality in the first case makes the second inequality vacuous and combined with the third inequality in the second case makes the first inequality vacuous. The third region in the proposition (no pregnancy) then follows immediately. For the second region (pregnancy and gender determination), it is useful to consider two cases separately: (i) U_b positive and U_g negative and (ii) U_b negative and U_g positive. (If U_b and U_g are both positive (negative), there would be a pregnancy with no gender determination (no pregnancy).) For case (i), using the results above, a woman becomes pregnant and determines gender if either

$$pU_b + (1-p)U_g > 0 \text{ and } pqU_b + (1-p)U_g < -d - qc - (1-q)(1-p)c$$

or

$$pU_b + (1-p)U_g < 0 \text{ and } U_b > \frac{d + (1-(1-q)p)c}{(1-q)p}.$$

Note that the lines $pqU_b + (1-p)U_g = -d - qc - (1-q)(1-p)c$ and $U_b > \frac{d+(1-(1-q)p)c}{(1-q)p}$ intersect exactly along the line $pU_b + (1-p)U_g = 0$, so that the union of the two regions is simply given by

$$U_b > \frac{d + (1-(1-q)p)c}{(1-q)p} \text{ and } pqU_b + (1-p)U_g < -d - qc - (1-q)(1-p)c.$$

Similarly, for case (ii), the union of the two analogous regions is given by

$$U_g > \frac{d + (q + (1-q)p)c}{(1-q)(1-p)} \text{ and } pU_b + (1-p)qU_g < -d - qc - (1-q)pc.$$

Combining the terms involving c yields the result for the second region (pregnancy and gender determination) of the proposition and completes the proof.

Proof of Proposition 3: Case (i) follows immediately from the assumption that the incremental utilities weakly decrease with more children added to the family. For case (ii), note that the assumption of weakly decreasing incremental utilities implies that either the inequality in (9) or the inequality in (11) will hold for n'_b and n'_g, ruling out the possibility of the woman becoming pregnant and not determining gender; whether the woman becomes pregnant (and determines gender) or does not become pregnant depends upon whether the inequality in (8) or the inequality in (9) continues to hold. Finally, case (iii) vacuously holds.

Proof of Proposition 4: With $U_b(n_g, n_b) = U_b(n_g + 1, n_b) > 0$, it is impossible for the inequalities in (10) and (11) to both hold. Therefore, a woman with n_b sons and n_g daughters becomes pregnant and determines gender since the inequalities in (8) and (9) both hold. With an additional girl ($n_g + 1$ daughters), the inequality in (8) continues to hold by the assumption of strong son bias ($U_b(n_b, n_g) = U_b(n_b, n_g + 1)$) and the inequality in (9) continues to hold by the assumption of weakly decreasing incremental utilities. Thus, a woman with n_b sons and $n_g + 1$ daughters would also become pregnant and determine gender.

Proof of Lemma 3: When gender is known (after gender determination) in period t, the expected utility of no termination is $U_b - d + \delta V^b_{t+1}$ for a boy ($U_g - d + \delta V^g_{t+1}$ for a girl) whereas the expected utility of termination is $-d + \delta V_{t+1}$. The results follow immediately.

The following lemma describes the gender-determination decision in period t:

Lemma 4 *(Gender-determination decision in period t)*
(i) If $U_b + \delta(V^b_{t+1} - V_{t+1}) > -c$ and $U_g + \delta(V^g_{t+1} - V_{t+1}) > -c$, the woman will not determine gender;
(ii) if $U_b + \delta(V^b_{t+1} - V_{t+1}) > -c$ and $U_g + \delta(V^g_{t+1} - V_{t+1}) < -c$, the woman will determine gender if and only if

$$pq(U_b + V^b_{t+1}) + (1-p)(U_g + V^g_{t+1}) < -d + (q + (1-q)(1-p))(-c + \delta V_{t+1}); \tag{17}$$

(iii) if $U_b + \delta(V^b_{t+1} - V_{t+1}) < -c$ and $U_g + \delta(V^g_{t+1} - V_{t+1}) > -c$, the woman will determine gender if and only if

$$p(U_b + V^b_{t+1}) + q(1-p)(U_g + V^g_{t+1}) < -d + (q + (1-q)p)(-c + \delta V_{t+1}); \tag{18}$$

(iv) if $U_b + \delta(V^b_{t+1} - V_{t+1}) < -c$ and $U_g + \delta(V^g_{t+1} - V_{t+1}) < -c$, the woman will determine gender if and only if

$$p(U_b + V^b_{t+1}) + (1-p)(U_g + V^g_{t+1}) < -d - c + \delta V_{t+1}. \tag{19}$$

Note that Lemma 2 (for period T) is actually just a special case of this lemma since the continuation utilities are all equal to zero when $t = T$. As with the termination decision, the continuation utilities play a role for fertility periods prior to T. Consider case (ii), under which a boy would be terminated and a girl would not be terminated. If $q = 0$, the condition in equation (17) for determining gender simplifies to $U_g + \delta(V^g_{t+1} - V_{t+1}) < -\frac{d}{1-p} - c$, meaning that the threshold for gender determination is lower than that for termination (due to the cost d). With $q > 0$, equation (17) appropriately accounts for the possibility of an involuntary termination.

The following lemma characterizes the pregnancy and gender-determination decisions for any fertility period t:

Lemma 5 *(Pregnancy and gender-determination decisions in period t)*
A woman will become pregnant and not determine gender if

$$pU_b + (1-p)U_g > \delta\left[V_{t+1} - pV^b_{t+1} - (1-p)V^g_{t+1}\right], \tag{20}$$

$$qpU_b + (1-p)U_g > -d - (1-(1-q)p)c + \delta\left[(1-(1-q)p)V_{t+1} - qpV^b_{t+1} - (1-p)V^g_{t+1}\right], \tag{21}$$

$$\text{and } \; pU_b + q(1-p)U_g > -d - (q+(1-q)p)c + \delta\left[(q+(1-q)p)V_{t+1} - pV^b_{t+1} - q(1-p)V^g_{t+1}\right]. \tag{22}$$

A woman will become pregnant and determine gender if either

$$U_b > \frac{d + (1-(1-q)p)c}{(1-q)p} + \delta\left[V_{t+1} - V^b_{t+1}\right] \tag{23}$$

$$\text{and } \; qpU_b + (1-p)U_g < -d - (1-(1-q)p)c + \delta\left[(1-(1-q)p)V_{t+1} - qpV^b_{t+1} - (1-p)V^g_{t+1}\right], \tag{24}$$

or

$$U_g > \frac{d + (q+(1-q)p)c}{(1-q)(1-p)} + \delta\left[V_{t+1} - V^g_{t+1}\right] \tag{25}$$

$$\text{and } \; pU_b + q(1-p)U_g < -d - (q+(1-q)p)c + \delta\left[(q+(1-q)p)V_{t+1} - pV^b_{t+1} - q(1-p)V^g_{t+1}\right]. \tag{26}$$

A woman will not become pregnant if

$$pU_b + (1-p)U_g < \delta\left[V_{t+1} - pV^b_{t+1} - (1-p)V^g_{t+1}\right], \tag{27}$$

$$U_b < \frac{d + (1-(1-q)p)c}{(1-q)p} + \delta\left[V_{t+1} - V^b_{t+1}\right], \tag{28}$$

$$\text{and } \; U_g < \frac{d + (q+(1-q)p)c}{(1-q)(1-p)} + \delta\left[V_{t+1} - V^g_{t+1}\right]. \tag{29}$$

The proofs of both Lemma 4 and Lemma 5 follow exactly the same arguments as the proofs for Lemma 2 and Proposition 2, respectively. The only necessary modification to those proofs is the inclusion of the appropriate continuation utilities.

Proof of Proposition 5: Due to the finite number of fertility periods, note that the differences $V_{t+1} - V^b_{t+1}$ and $V_{t+1} - V^g_{t+1}$ weakly decline as t increases (and are equal to zero at $t = T$), a fact that will be utilized for this proof. For case (i) (no pregnancy at $t+1$), the inequalities in (27), (28), and (29) must also hold for period t since $V_{t+1} - V^b_{t+1} \geq V_{t+2} - V^b_{t+2}$ and $V_{t+1} - V^g_{t+1} \geq V_{t+2} - V^g_{t+2}$. Also, since it is costly to wait ($\delta < 1$), note that the converse of (i) also holds: a woman who does not become pregnant at period t would not become pregnant at period $t+1$. Therefore, for case (ii), it must be the case that the woman becomes pregnant at period t and the important analysis concerns the decision to determine gender. Without loss of generality, consider the case of son bias where the inequalities in (23) and (24) hold. (The daughter-bias case follows similar arguments.) For period t, the expected utility associated with gender determination is

$$q(\delta V_{t+1} - d - c) + (1-q)(1-p)(\delta V_{t+1} - d - c) + (1-q)p(U_b + \delta V^b_{t+1} - d),$$

and the expected utility associated with no gender determination is

$$pU_b + (1-p)U_g + pV^b_{t+1} + (1-p)V^g_{t+1}.$$

The difference between these two expected utilities simplifies to

$$(1-p)(V_{t+1} - V_{t+1}^g) + pq(V_{t+1} - V_{t+1}^b).$$

Since both $V_{t+1} - V_{t+1}^g > V_{t+2} - V_{t+2}^g$ and $V_{t+1} - V_{t+1}^b > V_{t+2} - V_{t+2}^b$, the fact that the expected-utility differential is positive for period $t+1$ implies that it must also be positive for period t. Therefore, the woman would become pregnant and determine gender. For case (iii), as discussed above, it must be the case that the woman becomes pregnant in period t (for if she didn't, she wouldn't in period $t+1$). Considering the son-bias case again, the expected-utility differential (between gender determination and no gender determination) must be negative at period $t+1$:

$$(1-p)(V_{t+2} - V_{t+2}^g) + pq(V_{t+2} - V_{t+2}^b) < 0.$$

However, since $V_{t+1} - V_{t+1}^g > V_{t+2} - V_{t+2}^g$ and $V_{t+1} - V_{t+1}^b > V_{t+2} - V_{t+2}^b$, the expected-utility differential at period t could be either positive or negative which would lead to gender determination and no gender determination, respectively. A similar argument would hold for the daughter-bias case, completing the proof.

Appendix B: Construction of the maternally linked California birth data

The CDHS provided data for every birth that occurred in California between 1982 and 2003. The total number of birth records during the 22-year period was 11,657,778. In addition to the publicly available data, the author was provided with (1) each mother's first name, (2) each mother's maiden name, and (3) each mother's date of birth. The birthdate item was available for all births after 1988. A full name for each mother was created by concatenating the first name and maiden name together (with a space in between). Any records that had missing values for mother's name, mother's age, mother's birthdate (for births after 1988), or total number of previous live births were dropped, leaving 11,576,761 observations.

For any two births in the sample, the pair of births is considered a *potential match* if all of the following conditions are met:

- An exact match on mother's full name.
- An exact match between the month and year of the earlier birth and the month-of-last-birth and year-of-last-birth reported at the later birth.
- Consistency of the total-previous-live-births variable (meaning an increase of one from the earlier birth to the later birth).
- Consistency of mother's age information, meaning:
  - if both births occurred after 1988, an exact match on mother's birthdate.
  - if at least one birth occurred between 1982 and 1988, the reported difference between the mother's age at the earlier birth and her age/birthdate at the later birth was possible given the number of months between the two births.

After all potential matches are recorded, a pair of births is then considered an *actual match* if (i) the earlier birth is not a potential match with any other later births and (ii) the later birth is not a potential match with any other earlier births.

To link more than two births for a given mother together, additional linkages are made based upon the actual matches of the birth pairs. For instance, suppose that three births are denoted A, B, and C, in chronological order. If both pairs A-B and B-C represent actual matches, then the birth sequence A-B-C would be linked together. Additional births could be added to this sequence if A is an actual match with an earlier birth or if C is an actual match with a later birth. This process is continued until all matched birth sequences are constructed.

The matching algorithm resulted in a total of 6,800,733 births (58.7% of the total) being part of a matched birth sequence. The remainder of the births consisted of (i) only children, (ii) births that could

Table 19: Matched Birth Sequences, California Data

# of linked children	# of mothers	# of mothers with firstborn observed
2	2,007,361	1,510,597
3	631,291	514,709
4	158,658	131,834
5 or more	47,863	37,982
Total	2,845,173	2,195,122

Table 20: Sample Sizes by Race, California Data

Race	# of mothers with first two births observed	# of mothers with first three births observed
White	1,793,576	569,787
Black	151,031	55,934
Chinese	49.847	8,266
Indian	16,191	2,326
Japanese	16,152	3,304
Korean	16,929	2,517

not be uniquely matched together, (iii) births that could not be matched due to the mother's other births not being in the sample (e.g., because they occurred before 1982 or outside of California), or (iv) births that could not be matched due to coding errors (e.g., misspelled name or incorrect age). Table 19 summarizes the number of birth sequences in the linked data. The last column reports the number of mothers for whom the observed birth sequence begins with her first child. These mothers comprise the samples used for the analysis in Section 4.4. Table 20 provides a racial breakdown of the birth sequences used in the analysis, reporting the number of mothers for whom the first two births and the first three births are observed. The first column of numbers are the relevant sample sizes for analysis that conditions on gender of the first child, whereas the second column of numbers are relevant for analysis that conditions on the gender mix of the first two children.

Appendix C: Details on 5% PUMS Census data analysis

The 1980, 1990, and 2000 editions of the 5% PUMS Census data were used for the analysis in Section 4.2. The racial category was determined by the reported race of the mother. In 2000, the Census questionnaire allowed respondents to also indicate "secondary" racial categories. For the 2000 sample, the categorization was based upon the primary racial category reported for the mother.

In order to condition upon gender of first child or first two children, it is necessary to identify mothers for whom first-child information is available. For the 1990 and 2000 data, only children under the age of 18 who live in the household are recorded in the data. Although the 1980 data contains information on older children, those mothers with children older than 17 were dropped in order to have a comparable sample. In both 1980 and 1990, the data contains an item related to a mother's fertility (specifically, the "number of children ever born"). To identify families for which all of the children are still in the household (so that information on the first child is observed), only those mothers having the same value for number-of-children-born and number-of-children-in-household are kept in the sample. Unfortunately, the number-of-children-born data item is not available in the 2000 data. For these observations, a family was only retained in the sample if the oldest child in the household was 13 years of age or younger. This choice would misclassify birth order for families with older children that have left the household, but the cutoff of 13 was chosen to minimize this possibility. Other cutoff choices yielded extremely similar results, although choosing a lower cutoff obviously reduces the sample size available for analysis.

Each child's age (in years) is reported in the Census data, taking on values between 0 and 17. The birthyear of a child was calculated by subtracting the reported age from the Census year. For example, a 4-year-old in the 1990 Census would have a birthyear of 1986. This birthyear is used to categorize families into the time periods in Tables 6 and 7 (based on first child's birthyear and second child's birthyear, respectively). These two tables report the likelihood of having an additional (second or third) child within five years of the previous child. For the second-child outcomes (Table 6), the families considered are those whose oldest child is at least five years of age. Similarly, for the third-child outcomes (Table 7), the families considered are those whose second-oldest child is at least five years of age. A family is recorded as "having an additional child" if the difference in ages between the previous child and the "additional child" is less than or equal to five years. Finally, note that the earliest birthyear considered is 1963, which corresponds to 17-year-old children from the 1980 sample, and the latest birthyear considered is 1995, which corresponds to 5-year-old children from the 2000 sample.

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