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THE DEMAND FOR SONS: EVIDENCE FROM DIVORCE, FERTILITY, AND SHOTGUN MARRIAGE

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ABSTRACT

This paper shows how parental preferences for sons versus daughters affect divorce, child custody, marriage, shotgun marriage when the sex of the child is known before birth, and fertility stopping rules. We document that parents with girls are significantly more likely to be divorced, that divorced fathers are more likely to have custody of their sons, and that women with only girls are substantially more likely to have never been married. Perhaps the most striking evidence comes from the analysis of shotgun marriages. Among those who have an ultrasound test during their pregnancy, mothers carrying a boy are more likely to be married at delivery. When we turn to fertility, we find that in families with at least two children, the probability of having another child is higher for all-girl families than all-boy families. This preference for sons seems to be largely driven by fathers, with men reporting they would rather have a boy by more than a two to one margin. In the final part of the paper, we compare the effects for the U.S. to five developing countries.

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First, methought
I stood not in the smile of heaven, who had
Commanded nature that my lady's womb,
If it conceiv'd a male child by me, should
Do no more offices of life to't than
The grave does to the dead; for her male issue
Or died where they were made, or shortly after
This world had air'd them. Hence I took a thought
This was a judgment on me, that my kingdom,
Well worthy the best heir o' th' world, should not
Be gladded in't by me. Then follows that
I weigh'd the danger which my realms stood in
By this my issue's fail, and that gave to me
Many a groaning throe. Thus hulling in
The wild sea of my conscience, I did steer
Toward this remedy, whereupon we are
Now present here together…

King Henry VIII, by William Shakespeare, Act 2, Scene IV.
[Summary: The King argues to church officials that he should be granted a divorce from Katherine, his wife of 20 years, since she has given him no male heir.]

1 Introduction

In many developing countries, parents seem to have preferences for sons over daughters. In Asia, for example, some researchers have argued that 80 million girls are "missing," perhaps because they have been aborted, neglected, or directly killed (Kevane and Levine, 2001). Because of a rising gender imbalance in India, the use of sonograms for selective abortions has been officially outlawed. While there are no "missing" girls in the U.S., preferences for sons may take more subtle forms. In this paper we present several pieces of empirical evidence on the effect child sex composition has on parental marital status and fertility behavior. Specifically, we show that having girls has significant effects on divorce, child custody, marriage, and shotgun marriage *when the sex of the child is known before birth*, fertility stopping rules, and child support payments. Survey evidence suggests these effects stem largely from the father's preference for a son.

Using a simple model, we show that each piece of evidence taken individually is consistent with parental gender bias, but does not necessarily rule out plausible alternatives. But taken together, our empirical evidence indicates that American parents favor boys over girls. The bias is quantitatively important, and although decreasing over time, it is still significant today. When we compare the U.S. with five developing countries, we find that the effects of gender bias in the U.S. are generally smaller than the effects in Mexico, Columbia, and Kenya, and only a fraction of the effects in Vietnam and China.

We begin by documenting the effect of the gender composition of children on the probability of divorce. We find that parents with girls are significantly more likely to be divorced or separated compared to parents with boys. The effect is quantitatively substantial, accounting for a one to seven percent higher divorce probability for parents with only daughters, depending on family size. This effect is present in every region of the country and occurs across race and education categories. It is strongest in the early years of the data and declines over time, so that by the year 2000, the effect seems to have disappeared. In addition, divorced fathers are much more likely to obtain custody of sons compared to daughters. This custody effect is large, with divorced fathers being 11 to 22 percent more likely to have custody of the children in all-boy versus all-girl families. This effect has become quantitatively more important over time as the number of children living with divorced fathers has increased.

By themselves, these divorce and custody results are not necessarily evidence of parental bias. For example, it is well documented in the child psychology and sociology literature that the presence of the father in the household when kids are growing up is more important for boys than girls.[1] It is possible that parents have unbiased gender preferences, but they decide to avoid or delay divorce if they have boys because they recognize the harmful effects of raising a son without a father present in the household. For the same reason, a divorcing couple may agree that it is in the child's best interest to live with the father if the child is a boy. Alternatively, it is also possible that the monetary or time costs of raising girls are different than the costs of raising boys. A higher cost of raising girls could also explain the documented effect of children's gender on divorce.

We next turn to the effect of child gender composition on marriage. Controlling for family size, women with only girls are substantially more likely to have never been married than women with only boys. The chance a woman will be married decreases by two to seven percent for an all-girl family relative to an all-boy family, depending on family size.

Perhaps the most striking evidence comes from the analysis of shotgun marriages using birth certificate data from California. First we show that, at delivery, gender of the first child is not correlated with marital status for first-time mothers. This is reassuring, because for most parents in the sample, gender of the first child is unknown until birth. We then test whether gender of the child affects marital status at delivery when gender is known in advance (with high probability) because the mother has taken an ultrasound test during pregnancy. Among women

[1] There is a large literature which documents that fathers play a bigger role in the development of their sons than their daughters. Fathers spend more time with their sons (Lamb, 1976; Morgan, Lye, and Condran, 1988) and boys are found to suffer more from divorce than girls (Hetherington, Cox, and Cox, 1978).

who have had an ultrasound test, we find that mothers who have a girl are less likely to be married at delivery than mothers who have a boy. This evidence suggests that fathers who find out their child will be a boy are more likely to marry their partner before delivery.

Although these findings on marriage and shotgun marriage are consistent with the gender bias hypothesis, they do not rule out the possibility of the role model and differential cost hypotheses. We turn to evidence on the effect of sons versus daughters on fertility stopping rules for an additional piece of evidence that can help separate the role model and differential cost hypotheses from the gender bias hypothesis. We find that in families with at least two children, the probability of having another child is significantly higher for all-girl families than for all-boy families. The magnitude of the effect increases when we look at families with at least three children. Based on our simple model, this result would be hard to explain if parents were completely gender unbiased. It is important to realize that the fertility result alone does not necessarily imply that parents favor boys. It could in theory be explained by a higher cost of raising boys relative to girls. But together with results on divorce, marriage, and shotgun marriage, our finding on fertility stopping rules are consistent with a preference for boys.

Is the demand for sons driven by fathers or mothers? In addition to looking at revealed preferences on family size, we also look at stated preferences for boys versus girls. At least since 1941 and continuing to the present, more Americans have stated that they would prefer to have sons over daughters. This gender preference is largely driven by men. While women have only a slight preference for daughters in the population, men say they would rather have a boy by more than a two to one margin.

In the final part of the paper, we extend our analysis to five developing countries: Mexico, Columbia, Kenya, Vietnam, and China. For divorce, the effect of an all-girl family is fairly comparable to the U.S. in percent terms. However, in these countries divorced fathers are substantially more likely to have custody of their sons, so that a larger fraction of girls than boys will grow up without a father in the household. When we compare the effects on fertility across countries, they are substantially higher in China and Vietnam compared to the U.S., with more moderate effects for Mexico, Columbia, and Kenya.

An additional piece of international evidence utilizes the fact that 12 percent of marriages in Kenya are polygamous. Among all married women, we find that those with girls are more likely to be in a polygamous relationship compared to women with boys. We interpret this as evidence that the desire for boys leads some husbands to marry another woman if his first wife delivers a girl. Finally, using a different set of 15 foreign countries, we also document stated preferences for boys around the world. The U.S. has less of a son preference than countries like

Taiwan, Hungary, Guatemala, Singapore, and Thailand, but stronger son preference than countries like Great Britain, Germany, and Iceland.

Documenting parental sex bias is important for several reasons.[2] First, such bias may have direct and indirect consequences on women's socio-economic progress. Growing up in a divorced family, for example, may have long-lasting effects on children's outcomes. Previous studies have argued that children in divorced families are more likely to grow up in poverty, drop out of high school, become teenage parents, and experience unemployment. Furthermore, if there is evidence of parental sex bias in divorce and fertility decisions, it may be indicative of other ways in which parents treat boys and girls unequally. For example, even in families where parents remain married, parents who prefer boys may devote less attention and nurturing to their daughters than sons. They may also devote fewer financial resources to their education and health. In this sense, our results are related to the existing literature that documents an unequal intra-household allocation of resources.[3] Moreover, our findings are related to the many studies that document a persistent gender gap in labor market outcomes, most notably wages. Because parental sex preferences are not easy to control for in wage equations, the finding of lower wages for women may in part reflect parental bias for boys that results in unequal intra-household allocation of psychological and material resources.

Understanding the magnitude of parental sex bias and the way it changes over time also has important policy implications. Rapid progress in sex-selection technologies promises to make it increasingly possible for couples to choose the sex of their children. Although these techniques are still used by a negligible number of couples due to their costs, they are expected to become substantially cheaper and more reliable in the near future.[4] High-tech sex selection poses

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[2] Other studies that have looked at the relationship between gender and divorce are mainly in sociology (Spanier and Glick, 1981; Morgan, Lye, and Condran, 1988; Bracher et al., 1993; Mauldon, 1990; Devine and Forehand, 1996; Wu and Penning, 1997). An exception is Bedard and Deschenes (2003). We are not aware of any study that looks at the effect of child gender on shotgun marriages. Lundberg and Rose document a relationship between child gender and father labor supply (Lundberg and Rose, 2002) and the transition rate into marriage when the child is born nonmaritally (Lundberg and Rose, 2003). Ben-Porath and Welch (1980) have studied the effect of gender on fertility in the 1970s. We discuss the relationship between our study and the previous literature below.

[3] Butcher and Case (1994) look at the effect of sibling sex composition on education and earnings. Thomas (1994), using data for the U.S., Brazil, and Ghana, finds that the education of the mother has a larger effect on her daughter's height, and that the education of the father has a larger impact on his son's height. More generally, Behrman, Pollak and Taubman (1986) and Behrman (1988) investigate the unequal intra-family allocation of resources by gender in the U.S.

[4] Some analysts estimate that the market for sperm sorting in the U.S. will be "between $200 million and $400 million" in the near future (Fortune, 2001). Sperm sorting is already being utilized with mixed success. Because male sperm is lighter, the sperm are rapidly spun to sort sperm into male or female. This technique was pioneered on cattle and other livestock. In the future, IVF techniques, which use a DNA

a range of difficult policy dilemmas, and is currently extremely controversial. While it is legal in the United States, other countries have outlawed it or are considering outlawing it.[5] Obviously, the use of gender selection technologies becomes a more pressing issue if parents have strong preferences for one gender. As these technologies become more widely used, strong preferences for boys could, in the long run, lead to imbalances in the population gender ratio.

The remainder of the paper is organized as follows. In Section 2, we present a simple model which highlights the various testable implications of each hypothesis. In Section 3, we present evidence for divorce and child custody. In Section 4, we present evidence on marriage and shotgun marriages. In Section 5 we present evidence on fertility stopping rules and in Section 6, we present evidence on stated preferences for sons. Section 7 presents some additional evidence for the U.S. In Section 8, we present our international evidence. Conclusions are found in Section 9.

2 A Simple Model

In the empirical analysis of this paper, we investigate the effect of child sex composition on divorce, custody, marriage, and fertility stopping rules. We consider three possible explanations that could give rise to an empirical relationship. First, there is the possibility that parents are *gender biased*, i.e., they prefer having sons to daughters (or vice-versa). This explanation is purely based on tastes: for whatever reason, one of the two parents (or both) derives more utility from having boys than girls. Alternatively, it is possible that parents are unbiased, but realize that the presence of the father in the household has a differential impact on boys versus girls. We call this the *role model* hypothesis.[6] According to this hypothesis, parents care about the well-being of their children, and when deciding whether to marry, divorce, or have more children, take into account the asymmetric effect of a father's presence on boys and girls. A third possible explanation is that the monetary or time cost of raising girls is higher than the cost

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detector to separate sperm which is then artificially inseminated into the woman, promises a higher success rate.

[5]Canada, the U.K., and some Australian states have outlawed sex selection technologies for non-medical reasons. Sex selection is also forbidden by the European Convention on Bioethics. The American Society of Reproductive Medicine argues that sex-selection technologies have "the potential to reinforce gender bias in a society." Under pressure from the industry, ASRM recently gave an ethical go-ahead to sperm sorting, but remains officially opposed to the use of pre-implantation genetic diagnosis, in which embryos are produced outside the body and then screened and selected for genetic characteristics.

[6] There is a consensus in the child psychology literature that the presence of the father in the household when kids are growing up is relatively more important for boys than girls (see Lamb, 1976 and 1987 for a survey of the evidence). This consensus is based on evidence that fathers play a larger role in the development of their sons than their daughters. Longitudinal data on child development shows that the

of raising boys (or vice-versa). We call this the *differential cost* hypothesis.[7] Of course, these three explanations are not mutually exclusive. While there may be other candidate hypotheses which could be important for marriage, divorce, and fertility decisions, we focus on these three throughout the rest of the paper.

In the Appendix, we formally develop a simple two-period model, where parents have transferable utility functions and are forward-looking. The goal of the model is to define precisely what we mean for gender bias, role model, and differential costs and to better understand the implications of each hypothesis. Here we outline the intuition of the model and its predictions for divorce and fertility. Predictions for custody, marriage, shotgun marriage, and other outcomes have a similar flavor as the divorce predictions.

The simple model illustrates that a gender bias for sons, a role model effect for sons, and a higher cost of raising girls all have the same predictions for divorce: parents are more likely to divorce if they have a daughter versus a son. In other words, any of these hypotheses can generate a relationship between having girls and divorce. The intuition is straightforward. If parents have a net bias for boys, the utility loss which occurs with a divorce is larger for boy families compared to girl families. This immediately implies that, ceteris paribus, marriages with a girl are less happy than marriages with a boy, and more likely to result in divorce. If fathers prefer boys, it is easy to see why they might be more reluctant to divorce if they have sons, since mothers are more often granted custody of the children. Alternatively, if parents are unbiased and concerned about the role model effect, having a boy reduces the chance of a divorce since parents recognize the harm done to a boy will be greater. Finally, if parents are unbiased, parents don't care about the role model effect, and girls are more expensive, some fathers may opt out of the marriage. The assumption is that in the divorced state, *for the father*, child support and time costs are likely to be the same regardless of the sex of the child. Indeed, courts are unlikely to order child support payments that are vastly different for girls and boys. Some mothers may not be able to provide a big enough transfer to get a marginal father to remain married if she has a girl. In this corner solution possibility, their husbands are more likely to leave the family since marriage versus divorce is relatively more expensive with a daughter.

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absence of a father has more severe and enduring impacts on boys than girls (Hetherington, Cox. and Cox, 1978). In most cases, children are assigned to the mother after a divorce, irrespective of sex.

[7] There is some evidence that girls are more costly to raise than boys, and that individuals think girls are harder to raise. Olsen (1983) estimates that for one-child families, a girl costs around $900 more each year to raise up to the age of 18 compared to a boy. A December 2000 Gallup Poll of 1,026 American adults indicated that a majority of respondents thought girls were harder to raise. The question asked was "Which do you, yourself, think is easier to raise – a boy or a girl?" Fifty-three percent of respondents answered

However, the three hypotheses have different testable implications for fertility. With only gender bias, parents will be more likely to have an additional child if their first child was a girl. In contrast, with a pure role model story or if the cost of raising girls is higher, the opposite is true. Again, the intuition is simple. If couples prefer sons, one way to think about it is that there is a quality versus quantity tradeoff, so the *effective* number of children is larger in a boy family. Since the marginal utility of an additional child is decreasing in the number of effective children, boy families have a lower marginal utility for an additional child. On the other hand, in the case where parents are concerned about the role model effect and they value additional children equally regardless of sex, fertility is higher for boy families. The reason is that parents will take into account the effect of existing children's gender on future divorce. Couples who already have a boy are more likely to stay together in the future and therefore have another child, since divorce is more costly if you already have a boy. (If parents do not take into account the effect of their children's gender on future divorce, then the role model hypothesis predicts no effect of gender on fertility.) Finally, suppose parents are unbiased, parents don't care about the role model effect, and girls are more "expensive" than boys. In this case, couples whose first child is a girl will be less likely to have a second child. The reason is that, as long as children are normal goods, families with a girl are poorer than families with boys. Due to a pure income effect, couples with girls will have fewer children. Note that the opposite is true if boys are more expensive than girls. If raising boys is more expensive than raising girls, families with boys are "poorer" than families with girls, and therefore will have fewer children. (A priori we don't know whether boys are more expensive than girls or vice versa.)

In our empirical analysis, we find that having girls (relative to having boys) increases the probability of divorce, decreases the probability of paternal custody, decreases the probability of marriage and shotgun marriage, and increases the probability of having another child. The important point to note here is that each piece of evidence does not necessarily imply the existence of parental gender bias. To draw conclusions on the existence of gender bias, we need to consider all the evidence. In particular, our finding that parents with only girls are more likely to divorce and less likely to be married (Sections 3 and 4) can be explained by either gender bias, or the role model story, or a relatively higher cost of raising girls. On the other hand, our finding that families with only girls are more likely to have another child (Section 5) is inconsistent with the role model story, but can be explained by either parental gender bias or by a relatively higher

"boy," 28 percent answered "girl," with the remainder volunteering a response of "no difference" or "no opinion."

cost of raising *boys*. The fertility results alone are not enough to distinguish between the three hypotheses. A test of gender bias requires evidence on fertility as well as divorce and marriage. Although more than one force may be present, taken together, our evidence indicates that U.S. parents do have a preference for boys over girls.[8]

An interesting question that immediately follows from this conclusion is whether the documented bias reflects fathers' preferences or mothers' preferences. While casual observation suggests that, if there is any parental sex bias for boys, it is likely to come from fathers, we cannot draw unqualified conclusions from our evidence on divorce, custody, marital status, and fertility. The model in the Appendix clarifies that, in a transferable utility setting, identifying whether it is a father's preference or mother's preference that drives the results is not possible without some restrictions on utility functions. For example, take the finding that having a girl first increases the probability of having an extra child. If this result is indeed explained by sex bias on the part of one of the parents, it could be coming from either the father or the mother. In the case where an extra child generates enough surplus for the father, he can transfer enough utility to the mother to make her at least as well off with the extra child. Alternatively, the mother could make transfers to the father.

To shed some light on the potential difference in paternal and maternal gender preferences, in Section 6 we turn to data on stated preferences. Survey evidence from Gallup Polls reveals that the bias is especially strong for fathers. Men express an overwhelming preference for sons over daughters, while women's preferences are more evenly split.

3 The Effect of Sex Composition on Divorce and Custody

We begin our empirical analysis by investigating the relationship between the sex composition of children and the probability of divorce. We uncover a surprisingly consistent pattern for the effect of children's gender on divorce. In particular, we document that parents who have only girls are significantly more likely to be divorced than parents who have only boys. The effect occurs across race, education, and geographic location categories. The effect is strongest for older cohorts (parents born in the 1930s) and diminishes for younger cohorts (parents born in the 1970s). The effect declines over time, so that by the year 2000, it is

[8] Alternative explanations are also possible. For example, if daughters are more likely to care for and assist their parents in old age, it would imply a demand for daughters. Alternatively, sons may be more likely to support their parents financially in their old age, which would imply a demand for sons. On a related note, in developing countries, sons may play a more productive role in generating household income, and hence be preferred. Another potential hypothesis is that mothers and their daughters are more vulnerable to physical or emotional abuse from their male partners compared to mothers with sons. Although we recognize these possibilities as alternative explanations, we do not believe these are the main forces at play.

essentially zero.

The existing empirical evidence on divorce is mixed. The possibility that children's gender may be correlated with divorce was originally raised by sociologists in the 1980s (Spanier and Glick, 1981; Morgan, Lye, and Condran, 1988). The evidence in these studies is suggestive, but by no means conclusive. Sample sizes are limited and standard errors are generally not reported, so it is difficult to draw firm conclusions on the statistical significance of the gender effect. Furthermore, because of relatively small samples, these studies pool together families of different size, thus confounding the effect of gender on divorce with the effect of gender on family size. (In Section 5 we document that the gender mix of existing children affects future fertility decisions). Also, these studies are not informative on whether the effect of gender composition on divorce changes over time, or varies across groups. Later sociological studies have typically found no statistically significant correlation between child gender and the probability of divorce or separation.[9] This more recent sociological research questions whether the original finding was due to sampling variability and whether the effect is still present in more recent years. In economics, the only work in this area that we are aware of is a recent paper by Bedard and Deschenes (2003), who find an effect of child gender on divorce in 1980.[10]

Main evidence for divorce. For the main empirical analysis of divorce, we use data from the 1940 to 2000 U.S. Censuses. Specifically, we use the set of all parents who have children living at home with them. The unit of observation is a family: intact families have a mother and a father present, while divorced families have either a mother or a father.[11] To minimize the probability that some of the children might have left the household, our sample includes all

[9] See, for example, Bracher, et al. (1993); Mauldon (1992); Devine and Forehand (1996); Wu and Penning (1997). One possible explanation for the failure to find a statistically significant effect is the use of small samples. Sample sizes in these studies range between 140 and 10,000. In a study based on data from 18 countries, Diekmann and Schmidheiny (2001) find no evidence that child gender affects divorce or separation. Morgan and Pollard (2002) find an effect in some years but not in other years. Other studies by sociologists have argued that gender composition affects the *perception* of marital stability, as measured by responses to questions like "What do you think the chances are that you and your husband will eventually separate or divorce?" See for example Katzev, Warner and Acock (1994) and Heaton and Albrecht (1991).

[10] Bedard and Deschenes (2003) use the relationship between gender of the first child and divorce as an instrumental variable to estimate the effect of divorce on women's labor force participation and family income. We became aware of their paper during the preparation of our manuscript. The questions addressed by our paper and most of our empirical analysis are different from the Bedard and Deschenes paper.

[11] We include in the sample both fathers and mothers, to allow for the possibility that children are assigned to either parent. Our analysis implicitly assumes that all of the children from a divorced family live with either their father or their mother. Using administrative data on divorce from 1989 Vital Statistics records, we find that the children are divided between spouses (or other adults) in only 2% of all divorce cases.

families with a mother and/or a father between the ages of 18 and 40, with children less than 12 years old living with them.[12] The dependent variable for the analysis is a dummy equal to one if the children currently reside with a divorced or separated mother or father, and zero if the children live with parents who are currently married.[13] Whether the first child is a boy or a girl can arguably be viewed as random. However, whether the first two or three children are boys versus girls is no longer random since fertility decisions and divorce decisions are both endogenous (see Section 2). For this reason, throughout the paper, we present as much information by sex order and mix as possible.

The first column in Table 1 reports the estimated coefficient when this divorce dummy is regressed on the gender of the first child, for parents with only one child. There is no discernable effect on divorce for these families. In column 2, we repeat the analysis including families with one child or more. Unlike the estimate in column 1, the estimate in column 2 is not affected by the endogeneity of family size. In other words, because fertility is endogenous, the estimate in column 2 has the cleanest interpretation. It provides the total effect on divorce of having a girl versus a boy for the first child, including any indirect effect which operates through differential fertility. The coefficient indicates that parents whose first child is a girl are 0.11 percentage points more likely to be divorced than parents whose first child is a boy.

To help assess the magnitude of the estimated marginal effect, throughout the paper we report the "all-boy baseline," which in Table 1 is the fraction of all-boy families with a currently divorced parent (i.e., the intercept term). Adding the estimated marginal effect to this baseline gives the fraction all-girl families with a currently divorced parent. In each table, we also report the "percent effect," which is the increase in the probability of the outcome of interest for an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family over the all-boy baseline. This percent effect is simply the odds ratio minus one. In column 2, the percent effect indicates that the probability of divorce when moving from a family

[12] The 12-year cutoff is probably conservative, but minimizes the probability that some of the children have left home. If we look at the spacing between the first and the second child for mothers whose first child is less than 8 (and therefore more certain to be at home), we find that for 96% of mothers the spacing between the first and second child is 5 years or less. If most children do not leave home until they are 17 or 18 (or even later), by using our 12-year cutoff we are including virtually all children ever born to a mother. We have experimented with alternative cutoff ages for children (8 and 10 years old) and found similar results. An alternative would be to use children of any age, and restrict the sample to mothers whose "number of children ever born" is equal to the number of children observed in the household. However, this cannot be done consistently, because the variable "number of children ever born" is not available in the 2000 Census. Using this alternative definition for the 1940 to 1990 Census years yields similar results, however. For intact families, we require that the mother be between the ages of 18 and 40 and do not restrict the age of the father. For divorced families, we require the mother or father to be between the ages of 18 and 40.

whose first child was a boy compared to a family whose first child was a girl increases by 0.9%.

As the size of the family increases, the divorce effect becomes larger and more significant. In columns 3 and 4, we repeat the analysis including families with exactly 2 children and with 2 or more children, respectively. The excluded category is all-boy families. A parent with two girls is significantly more likely to be divorced than a parent with two boys. For families with 2 or more children, the difference is 0.25 percentage points, or 2.3% of the baseline all-boy divorce rate. When we turn to families with 3 or more children (column 6) or families with 4 or more children (column 8), we find even larger effects, with marginal effects of 0.53 and 0.72 percentage points and percent effects of 5.4% and 7.4%, respectively. We view these as surprisingly large effects.

As expected, the all-boy baseline decreases when comparing column 2 to column 4 to column 6 to column 8, since the probability of divorce is lower for larger families. It is interesting to note that both the marginal effect and the percent effect increase as we move from families with 1 or more children to families with 2 or more, 3 or more, and 4 or more children. For example, the percent effect increases seven-fold when moving from families with at least 1 child (0.9%) to families with 4 or more children (7.4%). As we will see below, this same pattern usually holds when we look at other outcomes as well.[14]

What explains the dramatic increase in the divorce effect by family size? We don't have a single answer, but we speculate that there may be at least three forces at play. First, fathers are more likely to get custody of their sons than daughters, and this difference in father custody rates is larger for fathers with fewer children. For example, we show below that the difference in father custody between all-boy and all-girl families is almost twice as large in families with one child versus four children (Table 2). If what drives the divorce effect in Table 1 is gender bias in the form of a stronger desire of fathers to live with their sons relative to daughters, and if the probability of paternal custody of boys declines with family size, the measured effect of gender on divorce is likely to be smaller when the number of children is one than when it is four. We elaborate more on this point below. Second, parental preferences may in theory be different for groups with low fertility and groups with high fertility. If groups with high fertility have stronger preferences for sons over daughters, for example, then we should see a larger effect in column 8 than column 2. Third, family size is endogenous. This endogeneity should have the opposite

[13] Note that the dependent variable captures whether the respondent is divorced at the time of the Census. Later in this section, we look at whether a respondent has ever been divorced.

[14] Another interesting pattern to note from Table 1 is that divorce probabilities are generally lower for all-boy families compared to families with a mix of boys and girls. The notable exception is families with exactly one boy and one girl.

effect of leading us to find a larger effect in column 2 than column 8. The reason is that the sample used in column 8 does not include many couples that divorced because they had girls for their first or second child. In this sense, the sample in column 8 has already been purged from some of the couples whose marital status depends on child gender mix.[15] The comparison of the magnitude of the effects by family size seems to indicate that the father custody effects and differences in tastes across fertility groups dominate the endogeneity of family size.

The results presented so far are obtained from models that do not include any controls. In Appendix Table A1 we show results from models similar to those in Table 1, where we condition on a vector of parents' characteristics, including a cubic in age, race, education, region of residence, and decade of birth. Estimates are qualitatively similar to the ones in Table 1, although the estimated effect of the all-girl dummy on divorce is somewhat smaller. For example, the percent effect in column 2 (the easiest to interpret) drops from 0.9 to 0.7. In Table A1, and throughout the paper, the all-boy baseline for models that include covariates is the predicted value of the dependent variable in an all-boy family, using the estimated regression coefficients and the explanatory variables evaluated at their means.[16]

One possible interpretation of the divorce findings in Table 1 is that fathers prefer sons to daughters. In a majority of divorce cases, fathers lose day-to-day access to children, while, irrespective of marital status, mothers stay with the children. According to this interpretation, fathers in marginal marriages who have boys are more likely to stay in the marriage (rather than divorce) than fathers in marginal marriages who have girls, because they like living with sons more than they like living with daughters. However, as argued in the theoretical model discussed in Section 2, gender bias is not the only possible explanation of the documented relationship between the sex composition of children and divorce. It is possible that unbiased parents in marginal marriages decide to stay together if they have boys because they realize the presence of the father at home is more important for boys than for girls (i.e., the role model hypothesis). Alternatively, it is in theory possible that the cost of raising girls is higher than the cost of raising boys (i.e., the differential cost hypothesis). In the next two sections we present additional evidence in order to differentiate between these three hypotheses.

Furthermore, even if the documented correlation between gender and divorce was

[15] Of course, there is also the possibility that some gender-biased couples with girls stay together in the hope of having a boy in the future, but split up over the disappointment of having another girl.

[16] In conditional models, this baseline makes more sense than using the simple unconditional average of the dependent variable for all-boy families. The simple average of the dependent variable for all-boy families does not take into account the fact that the covariates could be different in all-boys families and all-girls families. Of course, our baseline is very close (in most cases almost identical) to the simple average for all-boy families, as one would expect if sex of children is random.

explained by gender bias, it is not obvious that it is father preferences at play. In a transferable utility framework like the one adopted in Section 2, the documented evidence could in theory be explained by either paternal or maternal preferences. However, survey evidence presented in Section 6 indicates that fathers strongly prefer sons and mothers weakly prefer daughters, suggesting it is the father's preference for sons at work.

In interpreting our results, it is important to keep in mind that our estimates are likely to reflect the effect of having daughters on *marginal* marriages. Happy marriages are unlikely to result in a divorce only because the birth of a daughter, even if parents are gender biased. For this reason, even if the documented correlation between gender and divorce is explained by parental gender bias, we can not use our estimates in Table 1 to determine what fraction of couples has a preference for boys.

Main evidence for custody. To better understand the demand for sons and the divorce results, we turn to the effect of child gender on the custody arrangements for children with divorced mothers and fathers. If fathers do indeed have a preference for sons, they may fight harder to obtain custody of their sons than daughters in the case of a divorce. If this is true, we should observe higher paternal custody rates for all-boy families than for all-girl families.[17]

Paternal custody represents a second channel by which gender bias may manifest itself through its effect on family structure. For the purpose of our paper, the important point to note is that this effect, if true, would attenuate the *measured* effect of gender on divorce documented in Table 1. If fathers are biased in favor of sons but can easily get custody of their sons, the measured effect of having girls on divorce would be smaller than in the case where child custody is automatically assigned to the mother, irrespective of gender.

In Table 2, we present estimates of the probability a divorced father has custody based on the gender of the children. Using Census data, we define paternal custody as whether the children are living with the divorced father at the time of the census. The variable we call "custody" is therefore a broad measure of the father's access to and time spent with his children, and is likely to reflect joint custody and visitation rights in addition to sole paternal custody. The sample includes all divorced mothers and fathers, where as before, the sample is restricted to parents between the ages of 18 and 40 with children less than 12 years old. The table shows that children of divorced parents are more often assigned to the mother, but when the father does obtain custody, he is more likely to have custody of his sons. This effect of gender on custody is smaller

[17] Fox and Kelly (1995) use a sample of 509 court cases in Michigan and find that the odds of father custody are enhanced when children are male.

the larger the number of children.

The first column in Table 2 presents results for single-child families. Divorced fathers are 4.5 percentage points less likely to have their daughter living with them; approximately 19.9% of sons live with their divorced dad (see the all-boy baseline near the bottom of the table) compared to only 15.5% of daughters. This amounts to a 22.4 percent effect, which we view as remarkably large. For consistency with Table 1, we also report the custody coefficients for families with 1 or more children in column 2. The coefficient estimate drops to 2.8 percentage points, or a 16.9 percent effect.

As one might expect, when the number of children increases, divorced fathers are much less likely to have custody of the children. This may reflect the fact that fathers could feel less inclined to take on the responsibility of raising a large number of children. For example, fathers have custody of the children in families with 2 children less than 16% of the time. For families with four or more children, fathers have custody less than 11% of the time. More importantly, both the marginal effect and the percent effect of an all-girl family on father custody generally fall as family size increases, although the estimates remain large and statistically significant across all family sizes. For divorced parents with 2 children, the custody effect is a -2.84 percentage point difference, and for 3 and 4 children, the custody effect is -1.50 and -1.24 percentage points respectively. These translate into percent effects which indicate that fathers with all-boy offspring are 18.0, 11.0 and 11.6 percent more likely to have their children living with them compared to fathers with all-girl offspring (see columns 3, 5 and 7).[18] We have also estimated models conditional on parental characteristics. These conditional estimates, shown in Appendix Table A2, are generally similar to the estimates in Table 2.

As for divorce, one possible interpretation of the results in Table 2 is parental gender bias. But the findings in Table 2 are also consistent with the role model story: unbiased parents may decide that it is in the interest of boys to grow up with their father. Furthermore, even if the explanation of the results in Table 2 is gender bias, it is not clear whether the bias is on the father's side or the mother's side. Fathers may prefer sons to daughters, and fight more to retain custody of their sons in case of divorce, or mothers may prefer daughters, and fight more to retain their custody.

For our purposes, the thing to note is that the decline in the percent effect documented in

[18] Another interesting feature of Table 2 is that fathers are also less likely to have custody of mixed sex offspring relative to the all-boy baseline. For example, consider the results in column 3 for families with exactly 2 children. The marginal effects imply that while fathers take custody of 2 boys 16% of the time, they only take custody of a boy and a girl 14% of the time. For comparison, these marginal effects for mixed gender are almost two-thirds of the marginal effect for 2 girls.

Table 2 as we move from column 1 to 3, 5, and 7 mirrors the increase documented for divorce in Table 1. Although we cannot be certain that these patterns are only driven by parental gender bias, we speculate that the decline in Table 2 may explain, at least in part, the increase in Table 1. Taken together, Tables 1 and 2 indicate that fathers are much more likely to be living with their sons versus their daughters. Part of this effect is driven by the finding that couples are less likely to divorce if they have sons. However, it is also driven by the fact that divorced fathers are more likely to have custody of their sons. The custody effect seems to dominate for smaller families, while the divorce effect seems to dominate for larger families. The remarkably large effect uncovered in column 1 of Table 2 may explain why the divorce effect in column 1 of Table 1 is essentially zero. Fathers are more likely to live with their sons, but in column 1 this operates more through the custody effect than the divorce effect. For larger family sizes, on the other hand, this operates more through the divorce effect than the custody effect.

The effect over time and across groups. So far, we have estimated the average effect of having girls on divorce and father custody. This average effect can potentially mask differences across Census years, cohort of birth, racial, and education groups. For example, Tables 1 and 2 take all 60 years of Census data and run combined regressions. The results are therefore a weighted average of the various Census samples, with weights reflecting the number of observations available in each year. Here, we repeat the analysis in Tables 1 and 2 by Census year, cohort of birth, race, and education.

We begin in the top two panels of Table 3 by documenting how the divorce effect has changed over time. Although the models estimated are identical to the ones in Table 1, only the coefficient on all-girl family is reported. Each row is a separate regression. To further save space, we focus on families with 1 or more, 2 or more, and 3 or more children. The first panel shows that, in percent terms, the estimated effect is far lower in more recent Censuses than it was in earlier Censuses, although the decline is not always monotonic. The estimated effects for 1940-1960, a period when divorce was relatively rare, are imprecisely estimated due to smaller sample sizes. Over the 1970 to 2000 period, the percent effect for families with one or more children declines from 3.0% to 0.2%. For larger families, there is a similar decline. Families whose first two children were girls were 5.2% more likely to divorce in 1970 compared to -0.4% in the year 2000 (column 6). The corresponding numbers for families with 3 or more children are 9.7% compared to 0.6%. Unsurprisingly, the all-boy baseline in all three cases increases monotonically, since divorce rates increase over time.[19] The second panel in Table 3 reports results by 10-year birth cohort of parents, and presents a similar picture.[20] The quantitatively large effects found for earlier cohorts largely disappear for the latest observable cohort.

The first two panels in Table 4 present corresponding results for father custody by Census year and birth cohort. In the top panel, we don't find any clear trend across all years for the marginal effect. But when we look at 1980 and later, we find that the percent effects tend to decline (columns 3, 6, and 9). The marginal effects are large and generally remain statistically significant even in the most recent years (columns 1, 4, and 7). This indicates that, in contrast to the divorce effect, the custody effect has not completely disappeared over time.

An important fact to note is that the all-boy baseline--i.e., the fraction of boys living with their fathers after a divorce--has increased dramatically over time. For example, while only 6.5% of fathers with one child had custody of their son in 1970, almost 25% of fathers had custody of their son by the year 2000. Similar increasing patterns are found for larger family sizes or using the birth cohort grouping.[21] This increase in father custody is probably due to a change in cultural and legal norms in favor of paternal custody. This increase more than offsets the decline in the percent effect for more recent years, indicating that a growing number of fathers are taking custody of their sons compared to their daughters over time.

A comparison of the trends over time in Tables 3 and 4 suggests that the divorce effect falls precisely as more and more fathers are taking custody of their children, especially their sons. Results in Tables 1 and 2, which are an average across Census years, mask the fact that the divorce effect was more prevalent in earlier years, but disappeared as the number of fathers taking custody of their sons increased over time.

The bottom three panels of Tables 3 and 4 reveal that the divorce and custody effects are present across race and education categories. The marginal divorce effects are generally larger for blacks than for whites, while the custody effects are generally larger for whites compared to blacks. However, because the divorce probability is higher and the custody probability is lower for blacks, the percent effects are roughly the same. Results for Asians are imprecisely estimated,

[19] Our findings for 1990 are consistent with Morgan and Pollard (2002), who report no statistically significant effect of gender on divorce in recent years. Our custody findings below help explain why this might be case, as fathers are granted custody of their sons much more often in recent years.

[20] In this panel we focus on mothers age 20-30 (instead of mothers 18-40), to accommodate the fact that for the 1970s birth cohort, the oldest mothers we observe in the most recent Census (2000) are only 30 years of age. By focusing on mothers 20-30 we are holding roughly constant the average age across birth cohorts.

[21] The baseline in 1940-60 is surprisingly large. The large figure for those years is entirely driven by 1950 data (In 1940 and 1960 it is below 5%). We don't have a complete explanation for this anomaly in 1950, but we speculate that it may be due to data issues. In 1950, the rate at which mothers and fathers can be linked to their children is much higher compared to 1940 and 1960.

due to a relatively small sample size. When we look at the bottom panel, we find that the divorce and custody effects are present across the three education groups considered. The marginal effect for divorce is generally largest for those with less than a high-school education, but does not follow any other obvious pattern across education categories. Results by region (not shown, but available on request) indicate that the effects are present in every region of the U.S.

Additional results. A limitation of the divorce results presented in Tables 1 and 3 is that, due to the nature of the Census data, the dependent variable is whether a parent is divorced or separated at the time of the Census survey. However, many divorcees remarry. If having daughters versus sons affects the probability of remarriage, this could confound the effect of having girls on divorce. To gauge the potential impact the divorce definition has on our estimates, in Table 5 we look at the probability a respondent's first marriage ended in divorce using two alternative datasets.

First, we use the 1980, 1985, 1990 and 1995 Current Population Survey (CPS) Marriage and Fertility Supplements, which report the complete fertility and marital history of respondents in detail. The top panel of Table 5 reports larger percentage point effects for all-girl families than those in Table 1; these effects are around 1.0 percentage points. The implication is that parents with all girls are approximately 3 percent more likely to have their first marriage end in divorce. The dramatic rise in divorce rates in recent decades explains why these percent effects are small even though the percentage point effect is large. Another important benefit of the CPS data is that it shows that the lack of information on non-resident children does not significantly affect our Census results in Table 1.[22] However, Table 5 also makes clear the limitations of CPS. Even with four years of pooled CPS data, the standard errors on the estimates are large.

Second, we use the fact that in the 1960 to 1980 Censuses, one can back out whether a respondent's first marriage ended in divorce.[23] The bottom panel of Table 5 also yields larger estimates compared to Tables 1 and 3. For example, for families with 1 or more children, the percentage point effect is 0.36, which is larger than the results reported for the corresponding Census years in Table 3. Similar results are found for families with 2 or more and 3 or more

[22] We only observe children living at home in the Census data. To minimize the probability that the respondents have other children who live on their own, in Table 1 we have only included parents below age 40 with children younger than 12 years old. Yet, it is in theory still possible that some children have left, and that males leave earlier than females. The CPS data, based on the entire fertility history of respondents, indicates that this is unlikely to be a major problem.

[23] This is because the 1960, 1970 Form 1, and 1980 Censuses report if an individual has been married more than once. This information is not available for all respondents in the other Census datasets. To conserve on space, only the coefficient on all-girl families is reported in Table 5.

children, although the standard error is large for families with 3 or more children.

4 The Effect on Marriage, Shotgun Marriage, Remarriage, and Second Divorce

In this section we empirically investigate the effect of the sex composition of children on four marriage outcomes. First, we test whether women who have boys are more likely to ever have been married than women who have girls. We find that they are.

Second, we investigate the effect of the sex of the first child on shotgun marriage for first-time mothers. We begin by asking whether marital status is correlated with the sex of the child *at birth*. In general, it is not. This is not surprising, because for most mothers, the sex of the child in unknown before birth. But this is not true for women who take an ultrasound test, since the test often discloses the sex of the child several months before delivery. Surprisingly, we find that the sex of the child affects marital status at delivery for first-time mothers who have taken an ultrasound. These women are less likely to be married at delivery if they have a girl than a boy. We interpret this as evidence that fathers of boys are more willing to marry their partner between conception and delivery than fathers of girls. We are not aware of any existing empirical evidence on the relationship between child gender and shotgun marriages.

Finally, we look at the probability of second marriage and second divorce. We do not find any systematic evidence that divorcees with boys are more likely to remarry. However, if women do remarry, they are more likely to divorce a second time if they have girls. Our results are consistent with findings in Lundberg and Rose (2003) obtained using the PSID.[24]

(A) Marriage. We begin by documenting the relationship between children's gender and the probability a mother has never married. We use all mothers between the age of 18 and 40 with children younger than 12 years old in the 1940-2000 Censuses. In Table 6, we show that among women who have exactly one child, those whose first baby is a girl are 0.18 percentage points more likely to have never been married than women whose first baby is a boy. The corresponding coefficient for women who have 1 or more children is 0.17. The percent effects are non-trivial. Having a girl first reduces the probability of marriage by 2.2%.

When we consider women who have at least 2 or 3 children, the estimated effect is even larger. Women whose first two children are girls experience a 4.9% percent decline in the

[24] Lundberg and Rose find that having a boy increases the transition rate into marriage when the child is born non-maritally. Relative to the Census, the PSID has the advantage of following the same mother over time. One disadvantage is its limited sample size, which in the case of Lundberg and Rose is between 400 and 700 observations. Another difference is that Lundberg and Rose use a combined sample of all family sizes which may confound the effect of gender on marriage with its effect on fertility stopping rules.

probability of marriage compared with women who have two boys first (column 4). The percent effect for women with three girls is similar, with a 4.7% lower probability of marriage (column 6). We have also estimated models conditional on mother characteristics. These conditional estimates are generally similar to the estimates in Table 6 and are available on request.

(B) Shotgun Marriage. We now turn to the effect of child sex on marital status at the time of birth of the baby. For this analysis, we use data from all birth certificates of first-time mothers from the California Birth Statistical Master File, for 1989-1994.[25] The first column in Table 7 shows that *at delivery*, gender of the first child is not correlated with marital status. This is reassuring, because for most parents in the sample, gender of the first child is unknown until birth. Finding that gender of the first child is correlated with marital status at delivery would cast doubt on the interpretation of our findings. When we control for mother characteristics in column 2--including race, education, age, Hispanic status, immigrant status, and year--the coefficient flips sign, and remains virtually zero.

The most interesting results of the table are in columns 3 through 6. Here we test whether gender of the child matters when the mother has taken an ultrasound test during pregnancy and therefore knows with high probability the gender of the baby in advance. Ultrasound tests are typically able to reveal the sex of the baby by the 16th week of pregnancy, and are considered very accurate (somewhere between 95 and a hundred percent).[26] About 38% of mothers in the sample have taken the ultrasound test during pregnancy.

We regress marital status on a dummy equal to one if the child is a girl, a dummy for ultrasound, and the interaction of the girl dummy and the ultrasound dummy. The interaction of the girl dummy and the ultrasound dummy is negative and statistically significant. The coefficient in column 3 says that women who take the test and have a girl are 0.37 percentage points less likely to be married at delivery than women who take the test and have a boy. Because we control for the ultrasound main effect, these estimates are not driven by differences in

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[25] Marital status for this dataset is imputed on the basis of the mother's last name and the baby's last name, under the assumption that the baby takes the mother's last name if the parents are unmarried. This imputation is likely to introduce some measurement error in the dependent variable. Measurement error in the dependent variable may increase standard errors. However, it is unlikely to bias our point estimates, since there is no reason to expect that name assignment is systematically correlated with the dependent variable of interest, which is the interaction of ultrasound and child sex, conditional on main effects for child sex and ultrasound.

[26] The main medical reason for taking the test is not disclosure of the child's sex, but is diagnostic. The test uses sound waves to view and examine the fetus and its internal organs. It can also be used to measure bone size (usually femur length and skull diameter) to aid in gauging the gestational age of the fetus. The correct visualization of any fetal part depends of a host of factors such as fetal position, amount of liquor and thickness of the abdominal wall.

the probability of ultrasound across mothers.[27] When we also condition on mothers' characteristics in column 4, the marginal effect is 0.30 percentage points, slightly smaller than the unconditional one, but still significant. As one might expect, the standard error in the conditional model is slightly lower than standard error in the unconditional model.

We interpret these findings as evidence that a child's gender matters for shotgun marriages. Fathers who find out during pregnancy that their child will be a boy are more likely to marry their partner before delivery than fathers who find out that their child will be a girl.[28] How large is the percent effect? Calculating a percent effect for Table 7 is somewhat more difficult, given the available information in our dataset. We are ideally interested in mothers who get pregnant when they are unmarried, get an ultrasound sometime during their pregnancy, and then get married before the baby is born. If we could calculate this fraction for first-time mothers of boys, we would have the all-boy baseline. Unfortunately, we do not know whether the mother was unmarried at the time of conception from birth certificate information. Fortunately, timing of pregnancy and marriage are recorded in the 1980, 1985, 1990, and 1995 CPS Marriage and Fertility Supplements. While not ideal, we use this data to calculate an imputed baseline for the probability an unmarried woman gets pregnant and marries before the birth of her first child.[29] This imputation does not change the percentage point effect, but provides a reference for calculating a percent effect. For the regressions in columns 3 and 4, the percent effect is around 4%.

More importantly, the fact that we do not observe if mothers are married when they get pregnant has implications for our estimates. Ideally, we would like to exclude women who were already married when they got pregnant from the sample, since information from an ultrasound is unlikely to affect their marriage probability at delivery. A priori, we expect women who are already married to automatically have a zero coefficient on the interaction term "girl*ultrasound." In contrast, unmarried women who get pregnant have the potential for this interaction term to influence whether they are married at delivery. By mixing women who are married and unmarried at conception in the sample, we are biasing the coefficient estimate of interest (i.e. the

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[27] The coefficient on the ultrasound main effect is large and positive, indicating that married women are more likely to take the test. In general, women with higher socio-economic status are more likely to have an ultrasound. When we control for mother characteristics in column 4, the coefficient on ultrasound drops to less than half of the unconditional coefficient.

[28] We have also performed the same analysis for women at their second delivery. We found effects that are smaller and not significantly different from zero. This is not particularly surprising, and, in fact, can even be interpreted as a specification check. Knowing the sex of the unborn child is predicted to have more of an impact on first pregnancies than second ones.

[29] The CPS data is not ideal for two reasons. First the timing of pregnancy and marriage is self-reported, and second we do not observe whether a woman had an ultrasound.

coefficient for mothers who are not married at conception) towards zero. Around 66% of women are married at the birth of their baby in this sample, indicating that this concern has the potential to severely bias the coefficient on "girl*ultrasound" towards zero. Viewed in this light, the coefficients in columns 3 and 4 are likely to be downward biased estimates of the effect of child gender for mothers who are not married at conception.

In columns 5 and 6 we attempt to correct for some of this downward bias by weighting observations by the predicted probability the mother is unmarried at conception. The idea is that, although we cannot get rid of always-married mothers in the sample, we can probabilistically identify which mothers are most likely to be unmarried at conception. We use as weights the predicted probability a woman is unmarried at the time of birth using mother's race, age, immigrant status, Hispanic status, and year as well as the interactions of these variables.[30] As expected, the weighted estimates are bigger than the unweighted estimates. In the regression without controls, the effect increases 24% to -0.46 percentage points (column 5), and with controls, the effect increases 30% to -0.39 percentage points (column 6). The corresponding percent effects are -5.5% and -4.6%.

The confidential California Birth Statistical Master File also reports whether the mother has taken an ultrasound during labor.[31] We use this information to perform a specification check. Unlike for ultrasound during pregnancy, we expect that the interaction of ultrasound during labor and child gender not to matter for marital status at birth. The coefficient on the interaction in column 7 is –0.0024, but it is very imprecisely estimated and is not statistically significant from zero. When we control for mother characteristics in column 6, the coefficient on the interaction drops to virtually zero.

(C) Second Marriage and Second Divorce. For completeness, we now turn to the effect of sex composition on the probability of remarriage for divorced mothers and on the probability of a second divorce for remarried mothers. The effect of a child's sex on remarriage

[30] The weights exhibit the expected patterns. For example, younger women and black women have a higher predicted probability of being unmarried. Using the predicted probability that a woman is unmarried at the time of birth (instead of conception) as weights is not ideal. This limitation is likely to lead us to find weighted estimates that are smaller than the weighted estimates that we would find if we could use as weights the predicted probability that a woman is unmarried at the time of conception. However, we suspect that the difference is trivial, as the predicted probability that a woman is unmarried at the time of birth is likely to be highly correlated with the predicted probability that a woman is unmarried at the time of conception.

[31] Sonograms using ultrasound during labor and birth are used for major problems, such as uterine bleeding (to visualize the placenta). Only 2.2% of mothers in our sample have an ultrasound during labor and birth.

is a priori ambiguous, even if it is true that men have gender bias for their natural children.[32] When examining remarriage and second divorce outcomes, data is only available for the 1940-1980 Censuses. Empirically, we find no evidence that divorced women who have only girls are less likely to remarry. The top panel of Appendix Table A3 indicates that the difference in remarriage probabilities for women with all daughters versus all sons is statistically insignificant. The percent effects are clustered around zero.[33] The bottom panel of Appendix Table A3 shows that, among women whose first marriage ended in divorce and who have remarried, those with only girls are more likely to divorce again. The percent effects indicate that having all boys versus all girls increases the probability of a second divorce by 3.5%, 5.2% and 4.3% for families with at least 1, 2, and 3 children, respectively. It appears that the divorce effect is not confined to first marriages only, but continues to affect mothers who remarry as well.[34]

5 The Effect on Fertility

Taken together, the evidence on marriage and shotgun marriage presented in the previous section is suggestive of a gender bias on the part of parents. However, these results alone do not unambiguously rule out the role model hypothesis or the differential cost hypothesis. Take the results for marriage and shotgun marriage, for example. It is possible that unbiased fathers decide to marry their partner if she has boys, not because they prefer boys over girls, but because they want to provide a role model for their sons. Alternatively, it is possible that unbiased fathers decide to marry their partner if she has a boy because the costs of raising boys are lower than the costs of raising girls.[35]

To make further progress in differentiating between the three alternative hypotheses, in this section we focus on fertility decisions. As the model in Section 2 makes clear, if parents are

[32] If potential new husbands have a preference for male step-children over female step-children, one should observe that divorced women with boys have higher remarriage rates than divorced women with girls. On the other hand, it is also possible that men have a preference for the gender of their natural children, but they have no strong preference for the gender of step-children.

[33] One caveat in interpreting these results is that we do not know which children belong to the first marriage versus a subsequent marriage. The same holds true for the second divorce results as well. While the CPS Marriage and Fertility supplements contain this information, unfortunately, the sample sizes are too small.

[34] Note that here we use mothers as the unit of observation, and do not include divorced fathers in the analysis. This is because it is not obvious which fathers to include in the sample. Some women who divorce a second time may have married a man who had never been previously divorced. Therefore, the results in Table A3 are not directly comparable to Tables 1 and 3.

[35] Of course, alternative explanations are also possible. One interpretation of our results on marriage could be that daughters are more likely to care for and assist their parents in old age. A potential explanation for the second divorce result is that mothers with girls may be worried about the possibility of their new husband sexually molesting their daughters. We cannot completely rule out these possibilities, although it seems unlikely these are the only factors at play. Furthermore, these two explanations are harder to reconcile with our findings on fertility for married couples in this section.

biased toward boys, we should see that the probability of having another child is *higher* for families that have all girls than for families that have all boys. On the other hand, if parental preferences are unbiased, and the role model hypothesis is true, we should see that the probability of having another child is *lower or equal* for families that have all girls compared to families that have all boys. Similarly, if parents are unbiased, there is no role model effect, but girls are more expensive than boys, we should see that the probability of having another child is *lower* for families that have all girls compared to families that have all boys.

Main evidence. Table 8 presents fertility results using the sample of married women from the 1940-2000 Censuses used previously. The first column suggests that among families with 1 or more children, the probability of having another child is lower if the first child is a girl. The effect is rather small--only -0.2% of the baseline--but is statistically significant.

Is this a rejection of the gender bias hypothesis? Not necessarily. This small negative effect is partly explained by the fact that women whose first child is a girl are more likely to divorce, as documented in Table 1, and that divorcees have fewer children. Although our sample includes only *currently* married women, a fraction of these women have been previously divorced, so it is not surprising to find that they have slightly fewer children. In this sense, our estimates of the relationship between child gender and fertility in Table 8 are biased toward finding a *negative* relationship between the all-girl dummies and fertility. If we could observe the entire marital history of respondents, we could account for this bias. Although this is not possible for the entire sample, the Censuses for 1980 and earlier years do report if a woman has been married more than once. When we use the 1940-1980 Censuses and restrict our sample to women in their first marriage, the coefficient on the girl dummy from a model like the one in column 1 drops to virtually zero and is no longer statistically significant (the coefficient is -.0007, with a standard error of .0008). This is consistent with the notion that the negative coefficient in column 1 is mostly driven by the combination of a higher probability of divorce for mothers whose first child is a girl and the fact that mothers who experience a divorce spell have fewer children.

The next relevant question is how a zero effect should be interpreted for families whose first child is a girl. One explanation is that most couples plan on having more than one child, *regardless* of the sex of their first child. In this case, sex of the first child should have no effect on the probability of having another child. In the CPS fertility supplements, where the entire fertility history is reported, 85% of mothers with children who are 40 or older (and therefore should have completed fertility) have 2 children or more. If most couples' ex-ante desired number of children is 2 or more, it is not surprising to find a negligible effect for small families. This explanation is also consistent with the evidence for other countries that we report in Section 8. In the 5 developing countries that we look at, where family size tends to be large, there is almost no effect on having an additional child as a function of the sex of the first child, even though there are very large effects as a function of the sex of the first two or three children.

Turning to column 2, which looks at families with two or more children, we find evidence which supports the gender bias hypothesis. The estimates reveal the probability of having another child is 0.89 percentage points *higher* when the first two children are girls than when the first two children are boys. This is 2.3% higher than the baseline probability for all-boy families. In other words, moving from an all-boy family to an all-girl family increases the probability of having a third child by 2.3%. Estimates in columns 3 and 4, based on all families with at least 3 and 4 children, have even larger percent effects.[36] Families with 4 girls versus 4 boys are almost 5% more likely to have a fifth child. In Appendix Table A4, we show that estimates which condition on a mother's characteristics are slightly lower, but generally consistent with the unconditional estimates in Table 8.

To help in interpreting the magnitude of the fertility effect, consider that in 2000, there are about 7.9 million girls younger than age 12 living in all-girl families with married mothers between the age of 18 and 40. Using the distribution across family sizes and the estimates in Table 8 we can calculate how many of these girls would not have been born if parents could perfectly choose the gender of their children with zero cost. We estimate that at least 12,100 girls would not have been born. This estimate is likely to be a conservative lower bound since only a fraction of the parents in our sample who would like to have a boy do go ahead and actually have another child. Many parents in our sample who would have another child if they could determine the gender with certainty choose not to have another child, because the probability of ending up with a boy is around 50%.[37] For this reason, the true number is likely to be substantially larger.

Another feature of Table 8 is that having children of mixed gender has a negative effect on fertility. This is consistent with findings in the previous literature.[38] In column 2, the fertility

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[36] These results are consistent with findings in Ben-Porath and Welch (1980) based on 1970 Census data. On the other hand, more recent work by Teachman and Schollaert (1989) based on 1973-1982 data find that having boys or girls has no differential effect on how long parents wait before having an extra child. We suspect that Teachman and Schollaert's failure to find a statistically significant effect may at least in part be explained by their small sample size (less than 10,000 observations).

[37] On the other hand, it is also possible that some parents who would have liked to have another child tried and failed, because of fertility problems.

[38] We don't focus on mixed gender because other researchers have documented a demand for variety as evidenced in fertility (Ben-Porath and Welch, 1976, 1980; Rosenzweig and Wolpin, 1980). Angrist and

effect of having only girls (relative to only boys) is smaller in absolute value than the fertility effect of mixed gender (relative to only boys). For families with 2 or more children, the former is approximately 18 to 19% of the latter. The importance of mixed gender is also evident in column 3. For families with at least 3 children, the two largest negative coefficients are the ones on combinations where the first two children have the same sex, and the third has the opposite sex: Boy, Boy, Girl and Girl, Girl, Boy. This finding is consistent with the notion that families like gender mix, and if they achieve it with their third child, they are less likely to have a fourth child. The coefficients for combinations where the first two children have opposite sex (BGB, GBG, GBB, and BGG) are all negative, but much smaller, suggesting that parents do like variety in offspring. However, for families with 4 or more children, the demand for sons seems to swamp the demand for variety; families with 4 boys are the least likely to have a fifth child.

One additional feature of Table 8 is important to note. Whether the first child is a boy or a girl can arguably be viewed as random. However, whether the first two or three children are boys versus girls is no longer random since fertility decisions and divorce decisions are both endogenous. This is one reason why we present as much information on the sex order and mix in the tables as possible. Conditioning on the sex of previous children, however, the sex of the last child can still be considered random. Consider couples who have at least three children and whose first two children are a girl and a boy, in that order. Some of these couples have a girl as their third child (i.e., GBG) while other couples have a boy as their third child (i.e., GBB). The sex of this third child can be viewed as random conditional on the sex order of the first two children. Comparing the coefficients on GBG versus GBB from column three in Table 8, we see further evidence of the demand for sons. For this example, the difference in the probability of another child is 0.69 percentage points (-0.22 + 0.91). In other words, for this conditional sample, couples who had a girl as their third child instead of a boy were 2.3% (or in other words, 0.69 percentage points) more likely to have a fourth child.

Additional evidence. One limitation of the Census data used in Table 8 is that we observe family size at the time of the Census, and cannot distinguish between completed and uncompleted fertility. This is not necessarily a source of bias. However, because some mothers have not finished bearing children, part of the effect of gender on fertility may not manifest itself. Estimates in Table 8 are a weighted average of the effect for mothers with completed fertility and the effect for mothers with uncompleted fertility. They are therefore likely to *underestimate* the

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Evans use the demand for a mixed sibling-sex composition to study the effect of children on female labor supply (1998).

effect of child gender composition on lifetime fertility.

To address this problem, we now repeat the analysis using a unique panel of California mothers obtained by longitudinally linking birth certificates for the years 1989-2001. The longitudinal link was obtained by using a confidential version of the California Birth Statistical Master File dataset that provides each mother's name and date of birth.[39] The sample includes all married mothers who are first time mothers in 1989 or 1990, and follows them between 1989 and 2001. One advantage of this sample compared with the Census sample used in Table 8 is that most mothers in the longitudinal sample have completed fertility by 2001. Only 4.3% of mothers are still having babies in 2000 or 2001, suggesting that women concentrate most of their childbearing in the first few years.[40] By 2001, it appears that most mothers in the sample have stopped having children.

Because the longitudinal sample includes women who are more likely to have completed fertility, we expect to find larger effects in this sample compared to the Census. Table 9 shows estimates for the effect of gender composition on the probability of having another child, based on the longitudinal California sample. There is no effect of the sex of the first child on the probability of having two or more children (column 1). Consistent with the Census results, there is a large effect of having two girls and three girls on the probability of having 3 or more (column 2) or 4 or more (column 3) children. The percent effects are 4.0% and 5.5%, respectively.

As expected, the estimates of the all-girl effect from the longitudinal sample in Table 9 appear qualitatively consistent with, but quantitatively larger than, the corresponding estimates using the Census in Table 8.[41] It is also worth mentioning that when we re-estimate our models including all mothers in the Vital Statistics sample--not only those mothers who give birth in 1989 or 1990--our estimates are remarkably close to the Census estimates reported in Table 8. This confirms that including mothers with uncompleted fertility tends to bias downward the estimated effects. It also indicates that the lack of information on non-resident children does not

[39] We matched mothers at first delivery with mothers at second delivery using the mother's name, mother's date of birth (day, month, and year), and the first and second child's date of birth (month and year). We used the same procedure to match mothers at later deliveries. Because there are very few mothers who are born the same day and give birth in the same day and have the same name, the matching is likely to provide an accurate and complete longitudinal sample of California mothers.

[40] For each year between 1989 and 2001, we have calculated the fraction of mothers whose last observed delivery occurs in the specified year: 1989: 21.9%; 1990: 25.9%; 1991: 6.6%; 1992: 9.0%; 1993: 7.8%; 1994: 6.0%; 1995: 5.0%; 1996: 4.0%; 1997: 3.4%; 1998: 3.3%; 1999 2.5%; 2000: 2.3%; 2001: 1.9%.

[41] We find only a weak relationship between the number of girls and magnitude of the coefficient. For families with 3 or more children in column 3, the coefficient for families with 2 boys (BBG, BGB and GBB) appears to be slightly more negative than the coefficients for families with 2 girls (BGG, GBG, GGB), but this pattern is quite weak. We also do not find a strong effect by the gender of the last child, conditional on the gender mix of previous children, as we do in Table 8.

appear to significantly affect our Census results.

The effect over time and across groups. In Table 8, we have estimated the average effect of having girls on fertility. In Table 10, we repeat the analysis for specific sub-groups. We focus on families with 2 or more children. (For families with 1 or more children the effect is generally zero as explained above, and therefore not particularly interesting.) Each row is a separate regression. The top two panels show how the effect of gender on fertility has changed across Census years (panel 1) and decade of birth (panel 2). The effects are not monotonic. In panel 1, the percent effect is low before 1970, it is large in 1980, and declines after that. In panel 2, the percent effect is low for cohorts born before the 1940s, large for the cohort born in the 1950s, and declines after that. When we look across races in panel 3, the largest percent effect is for Asians. This finding stands in contrast with results in Table 3 (panel 3), where Asians display the smallest effect of gender on divorce. We don't have a good explanation for this discrepancy, besides the fact that the Asian sample is relatively small. Finally, panel 4 shows that both the marginal effect and the percent effect are smaller for college-educated women than for high school dropouts and high school graduates.

Interpretation. Overall, we interpret our fertility results in this section as generally supportive of the gender bias hypothesis. In both datasets, families where the first two children are girls have higher fertility than families where the first two children are boys. We find an even stronger effect when comparing families where the first three children are all girls with families where the first three children are all boys. The magnitude appears to be non-negligible. Depending on the sample, an all-girl family is roughly 2 to 5 percent more likely to have another child.

In section 2, we have shown what each of our three competing hypotheses--gender bias, role model, and differential costs--would predict regarding the relationship between sex composition of children and the probability of divorce, marriage, and fertility stopping rules. As discussed in Section 2, it would be difficult to explain our findings on fertility if there was no parental gender preference and only the role model hypothesis for boys was true. It would still be possible to explain the fertility findings *in isolation* under a differential cost story. If parents are unbiased, but raising boys is *more* expensive than raising girls, families with boys are "poorer" than families with girls, and therefore have fewer children. However, this explanation is at odds with the divorce results, which require the cost of raising girls to be *higher* than the cost of raising boys if differential costs are the only force in play.

In conclusion, we emphasize that it is difficult to interpret our results in Sections 3, 4, and 5 separately. But taken together, they document a pattern which is most consistent with the presence of a gender bias for sons among U.S. parents. We reiterate that failure to reject the gender bias hypothesis does not necessarily imply that the role model or the differential cost hypotheses are false. For example, it is possible that both gender bias and role model are at play, and that the effect of gender bias is large enough to offset the countervailing effect of a role model concern on fertility decisions. In the next section, we explore using survey data whether men and women have a stated preference for sons versus daughters as an additional piece of evidence on the demand for sons.

Throughout the paper, we have assumed that child gender is random. This assumption is consistent with the one in Angrist and Evans (1998). Is this assumption credible? Recent medical literature documents that natural methods based on timing of intercourse have no significant effect on offspring sex (Wilcox, et al., 1995). Furthermore, although the technology is evolving rapidly, clinical methods to influence the gender of children are currently used by an insignificant fraction of the population and are still not very accurate. Importantly, they were completely unavailable for most of the years under consideration in this paper.[42] It should also be noted that all our results hold true when we condition on a mother's observable characteristics.[43]

6 Stated Preferences for Sons versus Daughters

The previous section documents that parents with two or more children are more likely to have additional children if they have only girls versus only boys. These revealed preferences for fertility are interesting, but do not identify if the preference for sons is due to the father, the mother, or both parents. We turn to stated preferences to help answer this question. Starting in 1941 and continuing to the present, the Gallup organization has collected survey data on individuals' preferences for boys versus girls. While the wording of the question has changed over time, there is a consistent pattern in the data--Americans say they would prefer to have sons

[42] We have calculated the fraction of first-born children who are boys in each Census, from 1940 to 2000, and found that it is roughly constant over time. This is consistent with the notion that improvements in technology in gender selection have not yet received widespread adoption.

[43] One concern is that according to some biological theories, some animal species may be adapted to regulate the sex of their offspring according to environmental conditions (see Charnov, 1982 and Norberg, 2003). If in human populations, births of females become relatively more likely in periods of physical or emotional stress, this theory could explain the divorce results (although it would not explain the fertility results).

over daughters. This trend is plotted in Figure 1.[44] The basic finding is consistent with the observed fertility patterns in Tables 8 and 9, and supports the notion that parents are gender biased. However, it should be noted that a comparison of the magnitude of the effect across most of the years is not possible, since the question varies.

For the results presented in previous sections, we cannot say much about whether the effects are due to a father's versus a mother's gender bias. Recent Gallup Poll surveys conducted in 2000 and 2003 help to pinpoint the source of the bias. The surveys asked both men and women: "Suppose you could only have one child. Would you prefer that it be a boy or a girl?" Figure 2 shows that women seem to have only a slight preference for a daughter (35% say "girl" and 30% say "boy," based on a sample of 1,067 respondents). In sharp contrast, men express an overwhelming preference for a son (19% say "girl" and 48% say "boy," based on a sample of 962 respondents). The fact that men prefer sons by a 2.5 to one margin suggests that the divorce, marriage, shotgun marriage, and fertility findings may largely be due to fathers' gender bias.

Table 11 investigates how the preference for a son versus a daughter depends on parents' characteristics. For parents who stated a preference, we ran a probit regression where the dependent variable was equal to 1 if the respondent preferred a boy.[45] Men are approximately 23 percentage points more likely to indicate a preference for a son, controlling for covariates. Age seems to be a statistically significant factor in preferences. As individuals get older, they prefer boys less, perhaps indicating that life's experiences (including raising girls) help temper any bias. Blacks are about 12 percentage points more likely to prefer a boy compared to whites. However, education level, region, urban status, income, and marital status have no statistically significant impact. Interestingly, those who report that they attend church regularly are 8 percentage points less likely to report a preference for a son. Equally interesting is the finding that respondents who describe their political views as liberal are 15 percentage points more likely to report a preference for a daughter.

7 Additional Evidence for the U.S.: Child Support

Before turning to evidence from other countries, we briefly present an additional result for the U.S. We look at whether the gender composition of children affects the probability that divorced mothers receive child support from their former husbands. We use a sample of all

[44] Surveys were conducted on nationally representative random samples of adults (18 years or older). The surveys typically have 1,000 or more completed interviews per year, which implies a conservative sampling error of plus or minus three percentage points.

families headed by a woman with children 12 years or younger in the 1995-2000 March CPS. Table 12 shows that, among mothers with 2 or more children, the probability of receiving child support is lower for mothers who have two girls compared with mothers who have two boys. The effect among mothers with 3 or more children is also negative, but it is imprecisely estimated. The effect among mothers with one or more children is virtually zero. All models presented in the table control for mother age, mother race, and year. Unconditional models have larger standard errors and none of the relevant coefficients is significant. Because we rely on a sample that is substantially smaller than the Census and Vital Statistics datasets, our results are necessarily less precise. For this reason, evidence in Table 12 should be considered suggestive rather than definitive.

We have also investigated whether parents report being satisfied with their marriage more when they have boys versus girls. For this purpose, we used a question in the NLSY on level of happiness in the marriage. For whites, blacks, and Hispanics, we find that women with all-boy offspring report a higher level of marital satisfaction compared to all-girl offspring, although the difference is not statistically significant. Overall, the combined NLSY samples appear to be too small to draw firm conclusions.[46]

8 Evidence from Developing Countries

In order to put our estimates for the U.S. in a broader context, we now turn to evidence of gender bias in five developing countries. We were able to obtain large scale census samples for China (1982), Vietnam (1989 and 1999), Mexico (1990 and 2000), Colombia (1973, 1985 and 1993), and Kenya (1989 and 1999).[47] For these countries, we show several pieces of evidence. First, we re-estimate our models for the effect of offspring gender on the probability of divorce and father custody. Second, we re-estimate our models for the effect on fertility decisions. We then compare our estimates of the divorce and fertility effects based on U.S. data with estimates

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[45] We also ran multinomial logit regressions for the three-choice model of boy, girl, or no preference on the entire sample. These results are very similar. That is, respondent gender, age, race, church attendance, and liberal ideology are all statistically significant factors as in Table 11.

[46] We were originally inspired to use the NLSY after reading an article by Mizell and Steelman (2000) that claims mothers are significantly happier when they have boys. However, when we tried to reproduce that result, we discovered that it is based on a number of questionable assumptions and on a particular sub-sample. The result is not statistically significant when a more careful analysis is performed.

[47] These data were organized, cleaned up, and documented by IPUMS International. To the extent possible, the international data were converted into a consistent format so that variable definitions are comparable to those in the U.S. Census. Specifically, we use the 1990 and 2000 Mexican Censuses (1 percent and 10.6 percent samples, respectively); the 1973, 1985, and 1993 Columbian Censuses (all 10 percent samples); the 1989 and 1999 Kenyan Censuses (both 5 percent samples); the 1989 and 1999 Vietnam Censuses (5 percent and 3 percent samples, respectively); and the 1982 Chinese Census (0.1 percent sample).

based on data from these five developing countries. We find similar divorce and custody effects, and fertility effects which are generally larger for developing countries than for the U.S. China and Vietnam show by far the largest percent effects.

We next examine the relationship between polygamy and sex composition. In Kenya polygamy is not uncommon. We document that mothers with girls are more likely to be in polygamous families than mothers who have boys. We also discuss evidence on consensual unions in Mexico and Colombia.

Finally, we examine how stated preferences for sons versus daughters compare in the U.S. and abroad using an international Gallup Poll dataset for 15 countries. The U.S. has less of a son preference than countries like Taiwan, Hungary, Guatemala, Singapore, and Thailand, but a stronger son preference compared to countries like Great Britain, Germany, and Iceland.

(A) Divorce and Custody. We begin by looking at the effect of children's gender on divorce or separation. The sample restrictions and the models are similar to the ones used for the U.S. in Table 1. Because there are virtually no divorcees in the Chinese sample (fewer than .01 percent of men and women with children report being divorced), we exclude China from this analysis. To conserve on space in Tables 13 and 14, we do not report all possible child gender orderings as we did for the U.S. Rather we combine families with the same sex mix, regardless of ordering. This more compact grouping does not affect the all-girl percentage point effects or percent effects reported in Tables 13 and 14.

Table 13 reveals that for the four countries for which data is available, the gender mix of children affects the probability of divorce. Consider the column 1, which provides a summary measure of having a girl as a first child on divorce, irrespective of the (potentially endogenous) family size. Among individuals with one or more children, individuals whose first child is a girl have significantly higher probabilities of divorce in both Mexico and Columbia. In Kenya and Vietnam, the effects are not statistically significant. The percentage point increases for Mexico, Columbia, Kenya, and Vietnam are 0.15, 0.46, 0.11, and −0.05, respectively. Because the current divorce or separation rates vary widely across countries, percentage effects are provided in the tables for easier comparison. These magnitudes are similar to those in the U.S. In contrast to the U.S., father custody is much more common after a divorce in these developing countries. For example, in Vietnam, fathers retain custody of their all-boy offspring 48 percent of the time for families with one or more children. The marginal effects for custody are large in these countries. The corresponding percent effects are 16.8%, 10.1%, 19.4%, and 20.0% in Mexico, Columbia, Kenya, and Vietnam, respectively.

(B) Fertility Stopping Rules. Next we turn to the effect of child gender mix on the probability of having an extra child. The models and the sample restrictions are again similar to the ones used for the U.S. in Table 8. It is important to note that the well-known one-child policy in China did not take effect until 1982, the same year that our data was collected. Because the policy was not retroactive, it should have a negligible effect for our sample.

Estimates in Table 14 have patterns that are qualitatively similar to ones uncovered for the U.S. In families with one or more children, we find insignificant effects of the sex of the first child on fertility in Mexico and Kenya, and significant, but very small, effects in Columbia, Vietnam, and China. When we turn to families with at least 2 or 3 children, we find that for all five countries, all-girls families have a much higher probability of adding another child compared to all-boy families. For families with 2 or more children, the percentage point increases associated with all-girl families for Mexico, Columbia, Kenya, Vietnam and China are, respectively, 2.30, 0.50, 2.02, 8.71, and 19.90.

In interpreting these estimates, two points are relevant. First, fertility in developing countries is generally high and occurs at young ages. As discussed previously for the U.S., we expect the percent effect for families with 1 child to be close to zero. In the extreme case where every family plans on having at least two children, we should see no effect based on the sex of the first child. In this case, families with one child would be families that still have uncompleted fertility at the time of the census. Empirically, we find that the effect in column 1 is close to zero, but much bigger for families with 2 or more, 3 or more, and 4 or more children in all countries. Second, as we discuss for the U.S. in Section 4, if girl families have higher divorce rates, our fertility estimates are biased toward finding a negative relationship between all-girl families and fertility. This is because mothers who divorce and then remarry have lower fertility.

We visually compare the fertility effect across countries in Figure 3, where we graph percent effects. The effects for Mexico, Columbia, and Kenya are fairly similar to the U.S., even though the percentage point differences are generally larger than the U.S. The comparison is more dramatic for Vietnam, where all-girl families (with 2, 3, or 4 more children) are between 18% and 29% more likely to have an additional child. In China the effect is so large that we have to use a different scale in the graph for this country. Families with two girls in China are 54% more likely to have a third child compared to families with two boys. For families with three girls, the effect is an astonishing 90% increase. And although few families in China have large families, those with four girls are 128% more likely to have a fifth child.

(C) Polygamy. In Kenya, 12% of married women are in polygamous marriages. If men have a strong bias for boys, they may be more likely to marry a second wife if the first wife has given them a girl. Obviously, child gender is unlikely to be the only factor determining the existence of a polygamous relationship. But it is possible that, for men who are close to indifferent between having one or more wives, a child's gender could be a deciding factor in the decision to take a second wife. This could be either because giving birth to girls lowers the mother's status in the eyes of her husband or society, or because the man believes that having given birth to girls, a woman is more likely to give birth to more girls in the future.[48]

In Table 15 we find that, among all married women age 12 to 40 with children less than 12 years old, women who have girls are more likely to be in polygamous marriages than women who have boys. While we can link mothers to their natural children, we cannot identify which mothers are married to the same man. This is because most women in polygamous relationships in Kenya report living in a separate house, and not with other wives of their husband. Women who live in polygamous relationships are generally less educated, illiterate, older, and live in certain districts in Kenya. Controlling for these variables and census year explains about 10 percent of the variability in polygamy in a linear regression. We report results with and without controlling for these covariates.

The probability a woman with all-girl offspring will be in a polygamous relationship is higher for all family sizes. For families with one or more children, the marginal effect is 0.30 percentage points in models that control for a mother's characteristics, or a 2.8 percent effect. This estimate is significant at the 10% confidence level. For families with 2 or more children, the increase is 0.69 percentage points, so that women with 2 girls versus 2 boys are 6.8% more likely to be in a polygamous relationship. This effect is significant at the 5% confidence level. The all-girl effect for families with three or more children is similar, but the effect in column 6 is only significant at the 10% confidence level.

We interpret this as evidence that the desire for boys leads some husbands to marry another woman if his wife has given him girls. However, it is important to realize that the unconditional estimates are somewhat lower than conditional estimates. While these differences are not inconsistent with the large reported standard errors, we believe the estimates in Table 15 should be considered suggestive rather than definitive.

[48] Ben-Porath and Welch (1980) provide some evidence for the U.S. that the probability of having a second girl conditional on having a girl first is higher than the probability of having a girl as a second child conditional on first having a boy, although the effect is very small.

(D) Consensual Unions. Our next piece of evidence is based on the effect of child gender on "consensual unions" in Mexico and Colombia. Consensual unions are relationships which lack religious or civil recognition and are recorded explicitly in the Mexican and Columbian Censuses. Over 20 percent of couples with children in Mexico and over 30 percent of couples with children in Columbia are in consensual unions. Consensual unions occur primarily for two reasons: either the couple does not want the hassle and expense of a legal marriage, or one of the two partners has not obtained an official divorce from a previous spouse. For both countries, we find that among women who are either married or in a consensual union, women with all girls are more likely to be in a consensual union compared to women with all boys. In Mexico, the probability of a consensual union increases by 2.5 percent for families with one child and 3 percent for families with two children and is statistically significant. The effects are less precise and not significantly different from zero in Colombia. (Table available on request.)

(E) Stated Preferences in 16 Countries. Even more so than the U.S., developing countries appear to have strong observed fertility stopping rules which depend on the sex of previous children. How do stated preferences for sons versus daughters compare internationally? In 1997, the Gallup Organization conducted a survey in sixteen countries on four continents which asked about people's preference for sons versus daughters.[49]

The question asked of respondents was "Suppose you could only have one child. Would you prefer that it be a boy or a girl?" Many adults say that gender would not matter, but for those who do have a preference, boys are generally preferred to girls. Figure 4 graphs stated preference for boys versus girls (for those who state a preference) across 16 countries for which data is available. Traditionally male-centric societies such as Taiwan, Hungary, Guatemala, and Singapore show strong preferences for boys. Taiwan is the extreme example, where boys are preferred by a three-to-one margin. However, the U.S. is not far behind, with preferences closely resembling those in Thailand, Canada, and India. Colombia and Mexico, two countries whose divorce and fertility patterns were examined above, actually show less of a preference for sons compared to the U.S. Developed countries such as Great Britain and Germany have only a slight preference for sons, and in Lithuania, Iceland, and Spain, individuals actually express a preference for daughters.

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[49] Countries were included in the sample if the Gallup Organization had in place a wholly-owned subsidiary or joint venture company, and where an ongoing nationwide public opinion survey was currently being conducted. Surveys were conducted on nationally representative random samples of adults (18 years or older) in all countries except India, Colombia, and Mexico, where interviews were restricted to urban

(F) Interpretation. Overall, our findings from other countries provide a consistent picture. Relative to all-boy families, all-girl families are more likely to experience divorce and to have additional children. Divorced fathers are more likely to have custody of their sons and mothers with daughters are more likely to be in a polygamous relationship. Individuals also generally express a preference for sons over daughters. The interpretation of these findings for developing countries, however, is more difficult than the interpretation of the corresponding findings for the U.S. Unlike in the U.S., children in developing countries often have an important role in generating income for the household. Especially in rural areas, children often help their parents in agricultural production. In countries where social security is not available, children are often expected to provide parents with economic assistance when the parents stop working. Because, in general, males have better economic opportunities than females, results for developing countries do not necessarily reflect only tastes for sons. They are likely to reflect a combination of tastes and differences in the economic productivity between boys and girls.

9 Conclusion

In this paper we show that U.S. parents, most likely fathers, have a strong demand for sons. Although it does not take the extreme form of "missing" girls like in some Asian countries, the demand for sons does affect the marital status and fertility decisions of a significant segment of the U.S. population. The demand for sons seems to be decreasing over time by some measures, but the broad evidence indicates that it is still present in younger generations.

Several pieces of evidence are consistent with a demand for sons. First, parents of girls are significantly more likely to be divorced than parents of boys. The effect accounts for a non-trivial fraction of divorces, with a 1% to 7% higher probability of divorce for families with daughters. After a divorce, fathers are much more likely to live with their sons, possibly because they fight harder to obtain custody if they have sons than daughters. Second, women with only girls are substantially more likely to have never been married than women with only boys. Surprisingly, we also find evidence that the gender of the child affects marital status at delivery *when gender is known in advance* because the mother has taken an ultrasound test during pregnancy. Among women who have taken the ultrasound test, we find that mothers who have a girl are less likely to be married at delivery than mothers who have a boy. Fourth, child gender has a strong effect on fertility stopping rules. For families with at least two children, the

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areas. There are typically 1,000 or more completed interviews per country, which implies a conservative sampling error of plus or minus three percentage points.

probability of having another child is significantly higher for all-girl families compared to all-boy families. Fifth, there is suggestive evidence that child gender affects the probability of receiving child support payments and the probability of second divorce after remarriage.

Taken individually, each piece of evidence does not necessarily indicate parental gender bias. For example, it is possible that parents have unbiased gender preferences, but they decide to avoid divorce if they have boys because they realize that the presence of the father in the family is relatively more beneficial for boys. Alternatively, it is also possible that the monetary or time costs of raising girls are higher than the costs of rising boys. However, a simple model of marriage and fertility indicates that the combination of all this evidence is hard to explain if parents do not have a bias for boys. Of course, the existence of gender bias doesn't rule out the possibility that role model and differential cost concerns also play a part in parents' decisions.

The evidence based on revealed preferences cannot identify whether the documented sex bias reflects fathers' or mothers' preferences. Using stated preferences from survey data, we find that the gender bias is largely driven by men. While women have only a slight preference for daughters in the population, men say they would rather have a boy by more than a two to one margin.

The international evidence presents a similar picture. When we compare the estimated parental sex biases across countries, we find that the estimated bias is largest for China and Vietnam, and smallest for the U.S., with Mexico, Columbia and Kenya in between. Survey evidence indicates that the stated preference for sons in the U.S. is less than in many developing countries, but stronger than most first-world nations.

Taken together, this research suggests that the age-old favoring of boys is not confined to the past or to developing countries. It is subtle and less widespread than it once was in the United States, but it is still significant today. Even though the bias seems to be waning, the consequences of the demand for sons could matter more in the future. Technology already permits parents to choose a baby's sex, but the methods are now costly and unreliable. As the cost of procedures falls and their reliability rises, the sex-ratio in the population may slowly become more male. More importantly, the bias for boys evidenced by our results may lead to worse outcomes for daughters. For example, parental gender bias could potentially explain some of the gender gap in a variety of labor and non-labor market outcomes, since existing evidence shows that children from divorced families are more likely to grow up in poverty, drop out of high school, become parents while teenagers, or be unemployed. More generally, our findings on the demand for sons may be indicative of a more pervasive uneven distribution of economic, psychological, and time resources between brothers and sisters, even in intact families.

References

Angrist, J. D., and W. N. Evans, "Children and their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size," *American Economic Review*, 88(3), 1998.

Becker, G., "A Theory of Marriage: Part I," *Journal of Political Economy*, 1973.

Becker, G., "A Theory of Marriage: Part II," *Journal of Political Economy*, 1974.

Ben-Porath, Y. and F. Welch, "Do Sex Preferences Really Matter?" *Quarterly Journal of Economics*, 285-307, 1978.

Bracher, M., G. Santow, S. P. Morgan, and Tressell, "Marriage Dissolution in Australia – Models and Explanations," *Population Studies*, 47, 403-425, 1993.

Behrman, J. R., R. A. Pollak, and P. Taubman, "Do Parents Favor Boys?" *International Economic Review*, 27(1), 1986.

Butcher, K. F., and A. Case, "The Effect of Sibling Composition on Women's Education and Earnings," *Quarterly Journal of Economics*, 109(3), 1994.

Charnov, E.L.; R. L. Los-den Hartogh, W.T. Jones and J. van den Assem, "Sex Ratio Evolution in a Variable Environment," *Nature*, 289, 27-33, 1981.

Devine, D. and R. Forehand, "Cascading Toward Divorce: the Roles of Marital and Child Factors," *Journal of Consulting and Clinical Psychology*, 64, 424-427, 1996.

Bedard, K., and O. Deschenes, 2003, "Sex Preferences, Marital Dissolution, and the Economic Status of Women," mimeo.

Diekmann, A., and K. Schmidheiny, "Do Parents of Girls Really Have a Higher Risk of Divorce?" mimeo, 2001.

Fox, G. L., and R. F. Kelly, "Determinants of Child Custody Arrangements at Divorce," *Journal of Marriage and the Family*, 57, 693-708, 1995.

Heaton, T. B., and S. L. Albrecht, "Stable Unhappy Marriages," *Journal of Marriage and the Family*, 53, 747-758, 1991.

Hetherington, E., M. Cox, and R. Cox, 1978, *The Aftermath of Divorce in Mother-Child, Father-Child Relationships*, ed. J. Stevens and M. Matthews, National Association for the Education of Young Children.

Katzev, A. R., R. I. Warner, and A. C. Acock, "Girls or Boys? Relationship of Child Gender to Marital Instability," *Journal of Marriage and the Family*, 56, 89-100, 1994.

Kevane, M., and D. Levine, "The Changing Status of Daughters in Indonesia," UC Berkeley, 2001.

Leung, S. F., "A Stochastic Dynamic Analysis of Parental Sex Preferences and Fertility," *Quarterly Journal of Economics*, 1063-1088, 1991.

Lamb, M., "The Role of the Father: An Overview," in *The Role of the Father in Child Development*, ed. M. Lamb, 1-36. New York Wiley, 1976.

Lamb, M., 1987, "The Father's Role: Cross Cultural Perspectives," Lawrence Erlbaum Associates.

Lundberg, S., and E. Rose, "The Effect of Sons and Daughters on Men's Labor Supply and Wages," *Review of Economics and Statistics*, 84(2), 251-268, 2002.

Lundberg, S., and E. Rose, "Child Gender and the Transition to Marriage," *Demography*, 40(2), 333-349, 2003.

Mauldon, J., 1990, "The Effect of Marital Disruption on Children's Health," *Demography*, 27(3).

Mizell, C. A., and L. C. Steelman, "All My Children," *Journal of Family Issues*, 21(7), 2000.

Morgan, S. P., D. N. Lye, and G. A. Condran, "Sons, Daughters and the Risk of Marital Disruption," *American Journal of Sociology*, 94(1), 110-129, 1988.

Morgan, S. P., and M. Pollard, "Do Parents of Girls Really Have a Higher Risk of Divorce?" mimeo, 2002.

Norberg, K., "Dads and Cads: Paternal Investment and Variation in the Human Sex Ratio at Birth," mimeo, 2003.

Olsen, L., *Costs of Children*. Heath and Company, 1983.

Spanier, G. B., and P. C. Glick, "Marital Instability in the United States: Some Correlates and Recent Changes," *Family Relations*, 31, 329-338, 1981.

Rosenzweig, M., "Birth Spacing and Sibling Inequality: Asymmetric Information Within the Family," *International Economic Review*, 27(1), 1986.

Teachman, J. D., and P. T. Schollaert, "Gender of Children and Birth Timing," *Demography*, 26(3), 1989.

Thomas, D., "Like Father, Like Son; Like Mother, Like Daughter," Journal *of Human Resources*, 29(4), 950-986, 1996.

Wilcox A. J., Weinberg C. R., Baird D. D., "Timing of Sexual Intercourse in Relation to Ovulation Effects on the Probability of Conception, Survival of the Pregnancy, and Sex of the Baby," *New England Journal of Medicine*, 333:1517-1521, 1995.

Wu, Z., and Penning M.J., "Marital Instability After Midlife," *Journal of Family Issues*, 18, 459-478, 1997.

Appendix

This appendix develops a simple two-period, forward-looking model for divorce and fertility decisions. In the model, parents have transferable utility functions and experience stochastic shocks which influence the quality of a marriage. This stylized model aids in understanding the implications that (i) gender bias, (ii) role model, and (iii) differential costs have for divorce and fertility stopping rules. The model illustrates that a gender bias for sons, a role model effect for sons, and a higher cost of raising girls all have the same predictions for divorce: parents are more likely to divorce if they have a daughter versus a son. However, the models have different testable implications for fertility under fairly general conditions. With only gender bias, parents will be more likely to have an additional child if their first child was a girl. In contrast, with a pure role model story or if the cost raising girls is higher, the opposite is true.

(A) Utility as a Function of the Sex of Children. Our model for divorce and fertility decisions is grounded in the work of Becker (1973, 1974) which assumes that marriage markets are cleared by transfers between spouses. Specifically, we assume that both the husband and wife have transferable utility functions (i.e., quasi-linear) of the general form $h(B_t, G_t, C_t) + X_t$ where the subscript t denotes time, B_t and G_t are the number of boys and girls in the family, C_t is non-transferable consumption, and X_t is consumption which is transferable. Transferable utility functions have been widely used in bargaining models in addition to the marriage context. The advantage of transferable utility is that when considering divorce, marriage, and fertility decisions, one only needs to compare the sum of the husband and wife's utility. There is no need to consider the allocation of consumption goods in the marriage or determine which spouse has more power in the marriage. Because of this assumption, we do not refer separately to the husband's and wife's utility functions in what follows.

A marriage occurs when there is utility created from the union, although the future value of a marriage is not known with certainty. We model uncertainty in a marriage as follows. In each period, there is a shock to the marriage which is mean zero and independently and identically distributed. We assume the shock is normally distributed: $\varepsilon_t \sim N(0, \sigma^2)$. The shock is marriage-specific, so if the couple separates, the shock is not present.[1] We assume the shock is independent of the gender composition and size of the family. To facilitate discussion, we separate out the value of this shock when writing the utility function.

The couple's aggregated utility in period t can therefore be written as

(1) $$U(B_t, G_t, C_t, i) + X_t + I[i = M] \times \varepsilon_t \qquad i = M, D$$

where I is an indicator equal to 1 if the couple chooses to stay married after the realization of the shock and 0 otherwise. Throughout, we use the letters M and D (usually as superscripts) to refer to the married and divorced states, respectively. The combined period budget constraint is

(2) $$pB_t + qG_t + rC_t + X_t = Y_t$$

where p, q, and r are the prices of boys, girls, and nontransferable consumption and Y_t is combined income for the married couple. We do not allow prices or income to differ in the married and divorced states. While not very realistic, this simplification does not change the main insights of the model. The numeraire good is transferable utility. For simplicity, we assume the budget constraint holds with equality in each period with no borrowing or saving.

To make definitions of the gender bias, role model, and differential costs hypotheses more concrete, let children enter as a single, additive argument in the subutility function, and let

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[1] It is easiest to think of the shock as normalized so that it is measured on the same scale as the linear component of the utility functions. In what follows the normal assumption is not crucial. Other distributions such as a uniform distribution or a logistic distribution yield the same general implications, but do not have as nice an interpretation. A shock which occurs in divorced state for each spouse could also be added to the model, but this complication does not change any of the key predictions.

children be the only arguments whose value interacts with whether the couple is married or divorced, so that

(3) $$U(B_t, G_t, C_t, i) = U(\widetilde{K}_t^i, C_t) \qquad i = M, D$$

where $\widetilde{K}_t^i$ represents the *effective* number of children. If parents value children equally, the effective number of children is an equally weighted sum of the number of boys and girls. However, parents may value boys or girls more, implying unequal weights. They may also have a different value for the effective number of children when married versus divorced. To allow for these possibilities, we write the effective number of children in the married and divorced states as $\widetilde{K}_t^i = \alpha^i B_t + \beta^i G_t,\ \ i = M, D$ where α^i and β^i are positive scalars which weight how much parents value girls compared to boys in the married and divorced states. We assume throughout that $\alpha^M > \alpha^D$ and $\beta^M > \beta^D$, so that children provide more utility in the married state, ceteris paribus. The rationale for this assumption is that one of the parents (most likely the husband) has limited access to the children--and therefore "consumes" less of them--in the divorced state.

In this setup, boys and girls are perfect substitutes. It follows that there is a quality-quantity tradeoff between the gender and number of children. We assume that utility increases at a decreasing rate as the number of effective kids increases. That is,

(4) $$\partial U \big/ \partial \widetilde{K}_t^i > 0, \quad \partial^2 U \big/ \partial (\widetilde{K}_t^i)^2 < 0 \qquad i = M, D.$$

It is now a simple matter to describe what we mean by a gender bias, a role model, and a cost differential. A <u>gender bias</u> for boys occurs when $\alpha^M > \beta^M$, meaning that a girl is valued at some fraction of a boy in the married state. This is because either the husband or the wife derives more utility from living with boys than girls. Holding the number of children fixed, the effective number of children in the married state increases with the number of boys. For simplicity, we also assume that in the divorced state, boys and girls are valued equally: $\alpha^D = \beta^D$. The assumption that $\alpha^D = \beta^D$ simplifies the model, but can easily be relaxed. (One set of assumptions that would imply $\alpha^D = \beta^D$ is that the husband is biased, the wife is unbiased, after a divorce the husband loses access to the children, and that children provide utility only to the parent they live with. In this setup, in the divorced state, the father gets zero utility from his nonresident children and the wife gets equal utility from her children.)

If parents are unbiased, a <u>role model</u> for boys exists if $\alpha^D < \beta^D$ and $\alpha^M = \beta^M$, which implies that boys provide less utility compared to girls in the divorced state. The reason boys provide less utility to parents in the divorced state is that altruistic parents take into account the happiness of their children. In most cases, children are assigned to mothers irrespective of their gender. Under the role model hypothesis, the absence of a father has a larger negative impact on sons, so that boys suffer more after a divorce. The utility functions of altruistic parents take into account this differential utility loss for boys versus girls.[2]

Finally, <u>differential costs</u> favoring boys occurs if $p<q$ so that boys are cheaper than girls. If a father divorces and leaves the family, we assume that boys and girls cost the same for the father since courts are unlikely to order very different child support and alimony payments for fathers with boys and fathers with girls. Any asymmetry in costs in the divorced state is borne by the mother, who retains custody of the children.

(B) Two-Period Model. Using the utility functions described above, we develop a simple two-period model for divorce and fertility decisions. The model is forward-looking with couples making decisions in the first period before knowing the value of the marriage-specific shock in the second period.

To make things simple, in period one we start with couples who are married and already

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[2] We could explicitly model the utility of parents as function of the utility of children, and assume that the utility loss experienced by boys in the divorced state is larger that the utility loss experienced by girls. This would generate the same definition, but would complicate the notation.

have one child. The couple makes two choices in the first period: whether to divorce and whether to have an additional child. The decisions are made sequentially, with the couple first deciding whether to divorce after realizing the value of the first-period marriage-specific shock. If the couple chooses to stay together, they then decide whether to have an additional child, which will be born in the second period. In the second period, the couple observes two new pieces of information: the sex of the additional child (if they chose to have an additional child in the first period), as well as the value of the second-period marriage-specific shock. Given this information, the couple then decides whether to remain married or divorce as before. The second period is the terminal period in the model.

The divorce decision in period one takes into account the number and sex composition of children, as well as the expected future benefit of remaining married. Likewise, the additional child decision in period one takes into account the probability of divorce in the future. The expected benefit of remaining married today for a given family type is a function of the optimal choice of whether to have a child the next period, which in turn is based on the expected benefits of another child (which might be a boy or a girl), given that you have a probability of divorce in the next period (as a function of the sex composition and the realization of the shock). In other words, the expected benefit of staying married today for tomorrow's utility incorporates the possibility of additional children and future divorce.

To figure out divorce and fertility decisions, we need to know the expected utility of various family compositions in the future. To economize on notation, it is helpful to use a shorthand label for various family types. Let *B* stand for boy and *G* stand for girl. In our two-period model, there are six possible family compositions: *B, G, BB, BG, GB,* and *GG*. We refer to utility and other relevant variables which are a function of family type using these abbreviations as subscripts. We also abbreviate the utility function with a superscript to indicate whether the couple is currently married or divorced. For example, utility in the married state for a family with one boy and one girl is written as $U_{BG}^{M} = U(\alpha^{M} + \beta^{M}, C_t) + X_t + \varepsilon_t$. For clarity, we continue to separately write out the shock, ε, and do not include it in the shorthand definition. From this point on, we omit the time subscript on variables when it is clear which time period is the relevant one.

We now define some useful expressions which are a function of utility in the married and divorced states. The probability the marriage will survive in period two is denoted by

(5) $$\pi_c = 1 - F(U_c^D - U_c^M)$$

where the subscript *c* denotes family composition (i.e., *B, G, BB, BG, GB,* or *GG*) and $F(\cdot)$ is the cumulative distribution function of the random variable ε, which by our previous assumption is normally distributed. The expected value of the shock conditional on staying in the marriage (i.e., the truncated mean of the shock) is given by

(6) $$\lambda_c^M = E(\varepsilon_{t+1} \mid \varepsilon_{t+1} > U_c^D - U_c^M).$$

Similarly define λ_c^D as the expected value of the marriage-specific shock conditional on divorce.

For those who remain married in the first period, we can now write the expected utility (evaluated at period *t*) for a family with composition *c* in period *t*+1, as:

(7) $$\theta_c = \pi_c(U_c^M + \lambda_c^M) + (1-\pi_c)U_c^D = \theta(\widetilde{K}^M, \widetilde{K}^D).$$

Equation (7) is a function of U_c^M and U_c^D, which are in turn functions of the effective number of children in the married and divorced states, $\widetilde{K}^M$ and $\widetilde{K}^D$. Equation (7) plays a pivotal role in determining forward-looking divorce and fertility decisions.

(C) The Effect of Sex Composition on Divorce. We first examine divorce predictions based on child gender in a forward-looking context and then turn to the fertility predictions. It is easiest to discuss predictions by comparing a family whose first child is a girl to a family whose first child is a boy. In the discussion, we assume without loss of generality a gender bias for boys, a role model effect for boys, and that boys cost less than girls. (One could easily assume the opposite effects, in which case the predictions would be reversed.) To focus on the effects of each hypothesis, we examine each case separately. For example, when considering a gender bias for boys, we assume no role model effect and equal costs. While more than one effect may be present, the point is to show which effect dominates.

For a family with one boy in the first period, the probability of divorce in period 1 is:

(8) $$\Pr\left(\varepsilon_t < (U_B^D - U_B^M) + \gamma(U_B^D - \max\left\{\tfrac{1}{2}\theta_{BB} + \tfrac{1}{2}\theta_{BG}, \theta_B\right\})\right)$$

where γ is a discount factor. A similar expression for a family with one girl in the first period is:

(9) $$\Pr\left(\varepsilon_t < (U_G^D - U_G^M) + \gamma(U_G^D - \max\left\{\tfrac{1}{2}\theta_{GB} + \tfrac{1}{2}\theta_{GG}, \theta_G\right\})\right).$$

These equations capture the fact that individuals make divorce decisions based on current comparisons of utility as well as the forward-looking option value of remaining married. Comparisons of (8) and (9) reveal whether families with boys or families with girls are more likely to divorce. All three hypotheses (gender bias, role model, and differential costs) have the same prediction: in the first period, families with a girl are more likely to divorce compared to families with a boy. The intuition is given in Section 2.

The proofs for the gender bias and role model results rely on utility being an increasing, concave function in the effective number of kids, properties of the normal distribution, and revealed preference. Detailed proofs are available on request. The differential cost result is proven by first remembering that we assume if a father divorces and leaves the family, boys and girls cost the same for the father since child support payments and other monetary costs are likely to be the same. Any asymmetry in costs in the divorced state is borne by the mother, who retains custody of the children. In a transferable utility setting, when there is an interior solution, there is no effect of differential cost on divorce. However, if the couple reaches a situation where the mother cannot transfer enough utility to her husband, we end up with a corner solution where the father opts out of the marriage.

(D) The Effect of Sex Composition on Fertility. We now consider the fertility predictions of gender bias, role model, and differential costs. As before, we focus the discussion by comparing a family which has one girl to a family which has one boy and assume without loss of generality a gender bias for boys, a role model effect for boys, and that boys cost less than girls. The fertility decision depends on the probability the couple will have a girl versus a boy, the probability the couple will remain married as a function of sex composition, and the expected value of the shock if they remain married.

After choosing whether to divorce in period one, couples who choose to stay together make their fertility decision. For a family with one boy in the first period, the couple will choose to have another child if

(10) $$\tfrac{1}{2}\theta_{BB} + \tfrac{1}{2}\theta_{BG} - \theta_B > 0$$

and similarly, a family with one girl will have another child if

(11) $$\tfrac{1}{2}\theta_{GB} + \tfrac{1}{2}\theta_{GG} - \theta_G > 0.$$

Comparing the left-hand sides of these inequalities reveals whether boy or girl families have higher fertility under the three different hypotheses.

Fertility predictions are slightly more involved than the divorce predictions. Without saying something about the curvature of the utility functions, it is hard to compare the expected value of an additional child for boy versus girl families. It follows immediately that if θ, as described in equation (7), is an increasing, concave function of both $\tilde{K}^M$ and $\tilde{K}^D$, the following is true: (i) gender bias predicts higher fertility for girl families, and (ii) role model predicts higher fertility for boy families. If θ is an increasing, convex function of $\tilde{K}^M$ and $\tilde{K}^D$, the opposite is true.

It is easy to show that θ is an increasing, concave function of both $\tilde{K}^M$ and $\tilde{K}^D$ if

(12) $\lambda_c^i \left(U_c^i\right)' < -\left(U_c^i\right)'' / \left(U_c^i\right)' \qquad i = M, D$

where the derivatives are taken with respect to $\widetilde{K}^i$ for $i = M, D$, holding the number of children fixed. The proof for these concavity results can be seen by taking first and second derivatives and using properties of normal density functions, normal distribution functions, and truncated normal distributions.

Since the predictions hinge on the convexity or concavity of θ, it is important to understand the two conditions described in equation (12). Both conditions indicate the expected future benefit of a change in family composition *c*, depends on the amount of curvature in the utility function. In the current setting, the terms λ_c^M and λ_c^D are simply two different inverse Mill's ratios or hazard functions (i.e., ratios of density functions to survivor functions). This quantity multiplied by the marginal utility of an effective child in the married and divorced states must be less than what is often referred to as the coefficient of absolute risk aversion, an expression which describes the curvature of the utility function.

To make things more concrete, consider a family with one boy versus one girl and let *i=M*. Suppose parents have a gender bias for boys. The left-hand side of equation (12) captures the idea that boy families are more likely to remain married in period two and therefore more likely to enjoy the benefits of an additional child in the married state. The right-hand side of equation (12) captures the idea that additional children have potentially rapidly decreasing marginal utility in the married state. Since couples prefer sons, and because sons and daughters are perfect substitutes, the *effective* number of children is larger in a boy family that remains married. A boy family has more effective children in the married state, so the value of an additional child is lower compared to a girl family by an amount which depends on the curvature of the utility function. For concavity of θ to hold, there must be enough curvature in the utility function so that the benefit due to a lower divorce probability for boy families is smaller than the increase in marginal utility from an additional child for girl versus boy families in the married state. The intuition behind concavity when considering the role model hypothesis and *i=D* follows similar logic. To summarize, if there is sufficient curvature in the utility function so that equation (12) holds, gender bias predicts girl families are relatively more likely to have an additional child. In contrast, the role model hypothesis predicts boy families are relatively more likely to have an additional child.

Regardless of the concavity of θ, the differential cost hypothesis (when girls are more expensive) predicts higher fertility for boy families. The intuition is that having a girl versus a boy can be thought of as a pure income effect. If children are normal goods and girls have a higher price than boys, then couples whose first child is a girl are poorer. The income effect reduces the demand for additional children as well as other consumption goods which are normal goods.[3] Of course, if boys are more expensive the opposite is true, and girl families have higher fertility.

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[3] Note that no price effect arises, because gender is revealed *after* parents have made the decision to have another child. Leung (1991) shows that in a setting where fertility is stochastic and influenced by the parent's precautionary or proactive measures to have a child, and where consumption and children are perfect substitutes, the opposite result can be true. In Leung's setting, the gender bias and differential cost hypotheses are not separately identifiable.

Table 1. Child Gender and the Probability of Current Divorce or Separation, U.S. Census Data.

	Marginal Effect on the Probability of Current Divorce or Separation										
Sex of 1st child	Families with 1 child (1)	Families with ≥ 1 children (2)	Sex order of 1st two children	Families with 2 children (3)	Families with ≥ 2 children (4)	Sex order of 1st three children	Families with 3 children (5)	Families with ≥ 3 children (6)	Sex mix of 1st four children	Families with 4 children (7)	Families with ≥ 4 children (8)
Girl	-0.0004	0.0011	Girl, Girl	0.0020	0.0025	G, G, G	0.0056	0.0053	4 G	0.0067	0.0072
	(0.0006)	(0.0003)		(0.0007)	(0.0005)		(0.0014)	(0.0011)		(0.0029)	(0.0024)
			Boy, Girl	-0.0048	-0.0011	B, B, G	-0.0003	0.0003	2 G, 2 B	0.0018	0.0019
				(0.0006)	(0.0005)		(0.0013)	(0.0011)		(0.0018)	(0.0015)
			Girl, Boy	-0.0028	0.0001	B, G, B	0.0062	0.0045	3 G, 1 B	0.0020	0.0031
				(0.0005)	(0.0005)		(0.0014)	(0.0012)		(0.0021)	(0.0018)
						G, B, B	0.0055	0.0049	1 G, 3 B	0.0008	0.0010
							(0.0014)	(0.0012)		(0.0014)	(0.0018)
						B, G, G	0.0025	0.0029			
							(0.0014)	(0.0012)			
						G, B, G	0.0032	0.0025			
							(0.0014)	(0.0012)			
						G, G, B	0.0014	0.0021			
							(0.0014)	(0.0011)			
All-Boy Baseline	0.1812	0.1360		0.1170	0.1098		0.0980	0.0978		0.0978	0.0977
Percent Effect	-0.2%	0.9%		1.7%	2.3%		5.7%	5.4%		6.8%	7.4%
Obs.	1,554,818	4,169,265		1,679,127	2,614,447		659,523	935,320		195,586	275,797

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all households with parents who are ever married (excluding widows), who are between the ages of 18 and 40, and who have children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the parent is divorced or separated at the time of the survey. The excluded category is all boys. All-boy baseline is the fraction of parents in all-boy families who are divorced or separated, i.e., the intercept term. Percent effect is the increase in the probability of divorce or separation for the parent of an all-girl family compared to an all-boy family.

Table 2. Child Gender and the Probability of Father Custody, U.S. Census Data.

	Marginal Effect on the Probability of Father Custody										
Sex of 1st child	Families with 1 child (1)	Families with ≥ 1 children (2)	Sex order of 1st two children	Families with 2 children (3)	Families with ≥ 2 children (4)	Sex order of 1st three children	Families with 3 children (5)	Families with ≥ 3 children (6)	Sex mix of 1st four children	Families with 4 children (7)	Families with ≥ 4 children (8)
Girl	-0.0447	-0.0281	Girl, Girl	-0.0284	-0.0208	G, G, G	-0.0150	-0.0141	4 G	-0.0124	-0.0154
	(0.0014)	(0.0009)		(0.0022)	(0.0017)		(0.0049)	(0.0040)		(0.0094)	(0.0077)
			Boy, Girl	-0.0175	-0.0128	B, B, G	-0.0123	-0.0087	2 G, 2 B	0.0009	-0.0021
				(0.0022)	(0.0017)		(0.0048)	(0.0040)		(0.0060)	(0.0049)
			Girl, Boy	-0.0180	-0.0147	B, G, B	-0.0133	-0.0078	3 G, 1 B	-0.0079	-0.0076
				(0.0022)	(0.0017)		(0.0050)	(0.0041)		(0.0071)	(0.0058)
						G, B, B	-0.0177	-0.0138	1 G, 3 B	0.0055	0.0006
							(0.0050)	(0.0041)		(0.0070)	(0.0058)
						B, G, G	-0.0098	-0.0110			
							(0.0051)	(0.0041)			
						G, B, G	-0.0198	-0.0157			
							(0.0051)	(0.0041)			
						G, G, B	-0.0069	-0.0048			
							(0.0049)	(0.0042)			
All-Boy Baseline	0.1994	0.1694		0.1578	0.1460		0.1359	0.1268		0.1067	0.1042
Percent Effect	-22.4%	-16.6%		-18.0%	-14.2%		-11.0%	-13.0%		-11.6%	-14.8%
Obs.	281,493	569,601		194,052	288,108		66,574	94,056		19,435	27,482

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all divorced or separated fathers and mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the parent living with the children is the father. The excluded category is all boys. All-boy baseline is the fraction of divorced parents in all-boy families who are fathers, i.e., the intercept term. Percent effect is the decrease in the probability the divorced parent living with the children is the father for an all-girl family compared to an all-boy family.

Table 3. Child Gender and the Probability of Current Divorce or Separation, by Census Year, Decade of Birth, Race, and Education, U.S. Census Data.

	Coeff on Girl (1)	All-Boy Baseline (2)	Percent Effect (3)	Coeff on GG (4)	All-Boy Baseline (5)	Percent Effect (6)	Coeff on GGG (7)	All-Boy Baseline (8)	Percent Effect (9)
Model 1: By Census Year									
1940-60	.0010	.0503	2.03%	.0011	.0416	2.64%	.0025	.0394	6.34%
	(.0009)			(.0015)			(.0030)		
1970	.0025	.0830	3.01%	.0038	.0735	5.17%	.0072	.0741	9.71%
	(.0006)			(.0010)			(.0020)		
1980	.0014	.1419	1.00%	.0053	.1125	4.71%	.0062	.1036	5.98%
	(.0006)			(.0011)			(.0026)		
1990	.0006	.1574	0.46%	.0023	.1253	1.83%	.0084	.1129	7.44%
	(.0006)			(.0011)			(.0026)		
2000	.0004	.1658	0.24%	-.0006	.1365	-0.43%	.0006	.1220	0.57%
	(.0007)			(.0012)			(.0028)		
Model 2: By Decade of Birth (women age 20-30)									
1930s	.0025	.0527	4.74%	.0026	.0459	5.66%	.0130	.0384	33.85%
	(.0016)			(.0025)			(.0051)		
1940s	.0016	.0854	1.87%	.0035	.0791	4.67%	.0079	.0871	9.07%
	(.0007)			(.0013)			(.0028)		
1950s	.0018	.1438	1.25%	.0074	.1222	6.05%	.0092	.1239	7.42%
	(.0008)			(.0015)			(.0039)		
1960s	.0009	.1608	0.56%	.0022	.1449	1.51%	.0038	.1480	2.56%
	(.0009)			(.0017)			(.0044)		
1970s	.0005	.1609	0.31%	.0002	.1488	1.34%	.0003	.1468	0.20%
	(.0011)			(.0021)			(.0048)		
Model 3: By Race									
White	.0006	.1231	0.51%	.0015	.0955	1.57%	.0028	.0793	3.54%
	(.0003)			(.0005)			(.0011)		
Black	.0010	.2834	0.35%	.0042	.2668	1.57%	.0153	.2605	5.88%
	(.0015)			(.0026)			(.0055)		
Asian	-.0011	.0567	-1.94%	-.0049	.0461	-10.63%	-.0044	.0453	-10.30%
	(.0013)			(.0022)			(.0055)		
Other	.0037	.1681	2.20%	.0025	.1441	1.73%	.0044	.1316	3.33%
	(.0021)			(.0035)			(.0073)		
Model 4: By Education									
HS <	.0017	.1651	1.02%	.0043	.1454	2.95%	.0082	.1347	6.11%
	(.0008)			(.0013)			(.0027)		
HS	.0005	.1386	0.36%	.0017	.1102	1.54%	.0017	.0960	1.77%
	(.0005)			(.0008)			(.0018)		
College	.0011	.1207	0.90%	.0014	.0922	1.51%	.0052	.0735	7.02%
	(.0005)			(.0007)			(.0017)		

Notes: Each entry is a separate regression. Models in columns 1, 4, and 7 are identical to models in columns 2, 4, and 6 in Table 1. To conserve on space, only the coefficient for all-girl families is reported. See notes to Table 1.

Table 4. Child Gender and the Probability of Father Custody, by Census Year, Decade of Birth, Race, and Education, U.S. Census Data.

	Coeff on Girl (1)	All-Boy Baseline (2)	Percent Effect (3)	Coeff on GG (4)	All-Boy Baseline (5)	Percent Effect (6)	Coeff on GGG (7)	All-Boy Baseline (8)	Percent Effect (9)
Model 1: By Census Year									
1940-60	-.0271	.1666	-16.26%	-.0258	.1287	-20.00%	-.0280	.1111	-25.22%
	(.0072)			(.0123)			(.0223)		
1970	-.0122	.0654	-18.65%	-.0127	.0579	-21.27%	-.0215	.0595	-36.13%
	(.0018)			(.0029)			(.0053)		
1980	-.0295	.1230	-23.98%	-.0209	.0971	-21.52%	-.0092	.0746	-12.33%
	(.0015)			(.0028)			(.0066)		
1990	-.0346	.1791	-19.31%	-.0205	.1514	-13.54%	-.0114	.1368	-8.32%
	(.0017)			(.0032)			(.0077)		
2000	-.0248	.2474	-10.02%	-.0189	.2291	-8.24%	-.0048	.2145	-2.23%
	(.0021)			(.0041)			(.0101)		
Model 2: By Decade of Birth (women age 20-30)									
1930s	-.0168	.0771	-21.78%	-.0127	.0518	-24.90%	-.0022	.0359	-6.12%
	(.0079)			(.0120)			(.0204)		
1940s	-.0045	.0460	-9.78%	-.0020	.0340	-5.88%	-.0082	.0332	-21.6%
	(.0019)			(.0029)			(.0050)		
1950s	-.0197	.0904	-21.79%	-.0132	.0647	-20.40%	-.0051	.0446	-11.46%
	(.0017)			(.0030)			(.0065)		
1960s	-.0217	.1292	-16.97%	-.0189	.1043	-18.17%	-.0151	.0918	-16.4%
	(.0021)			(.0038)			(.0083)		
1970s	-.0151	.1738	-8.68%	-.0079	.1455	-5.42%	-.0211	.1231	-17.1%
	(.0029)			(.0053)			(.0119)		
Model 3: By Race									
White	-.0308	.1777	-17.33%	-.0219	.1565	-13.99%	-.0111	.1397	-7.94%
	(.0011)			(.0021)			(.0051)		
Black	-.0181	.1195	-15.14%	-.0142	.0932	-15.23%	-.0083	.0777	-10.68%
	(.0019)			(.0031)			(.0059)		
Asian	-.0096	.1774	-5.41%	.0113	.1643	6.88%	.0050	.1200	4.17%
	(.0096)			(.0197)			(.0466)		
Other	-.0186	.2238	-8.31%	-.0269	.2072	-12.98%	-.0644	.2058	-31.29%
	(.0056)			(.0102)			(.0223)		
Model 4: By Education									
HS <	-.0256	.1541	-16.61%	-.0173	.1176	-14.71%	-.0075	.0860	-8.72%
	(.0019)			(.0031)			(.0057)		
HS	-.0300	.1738	-17.26%	-.0238	.1526	-15.59%	-.0225	.1382	-16.28%
	(.0015)			(.0028)			(.0066)		
College	-.0270	.1707	-15.75%	-.0181	.1568	-11.53	-.0054	.1575	-3.42%
	(.0015)			(.0031)			(.0086)		

Notes: Each entry is a separate regression. Models in columns 1, 4, and 7 are identical to models in columns 2, 4, and 6 in Table 2. To conserve on space, only the coefficient for all-girl families is reported. See notes to Table 2.

Table 5. Child Gender and the Probability an Individual's First Marriage Ended in Divorce, Current Population Survey Fertility Supplements and U.S. Census Data.

Marginal Effect on the Probability a First Marriage Ended in Divorce					
Sex of 1st child	Families with ≥ 1 children (1)	Sex of 1st two children	Families with ≥ 2 children (2)	Sex of 1st three children	Families with ≥ 3 children (3)
1980, 1985, 1990, 1995 CPS Fertility Supplements					
Girl	.0108	Girl, Girl	.0106	Girl, Girl, Girl	.0098
	(.0030)		(.0050)		(.0095)
All-Boy Baseline	.341		.314		.335
Percent Effect	3.16%		3.37%		2.92%
Obs.	96,859		76,357		18,901
1960-1980 Census Data					
Girl	.0036	Girl, Girl	.0049	Girl, Girl, Girl	.0038
	(.0006)		(.0010)		(.0022)
All-Boy Baseline	.176		.155		.147
Percent Effect	3.23%		4.02%		2.81%
Obs.	1,621,180		1,025,825		383,442

Notes: Standard errors in parentheses. Data are from the 1980, 1985, 1990, and 1995 CPS Fertility Supplements and the 1960, 1970 Form 1, and 1980 U.S. Censuses. The dependent variable is a dummy equal to one if the respondent's first marriage ended in divorce. The excluded category is all boys. To conserve on space, only the coefficient on all-girl families is reported. All-boy baseline is the fraction of respondents in all-boy families who divorced their first spouse, i.e., the intercept term. Percent effect is the increase in the probability of divorce for the respondent of an all-girl family compared to an all-boy family. The CPS sample includes all-ever married mothers between the ages of 20 and 70 who report having at least 1 child. We focus on mothers in the top panel, because (unlike the Census) the CPS reports full fertility and marital histories. The Census sample includes all ever-married parents (excluding widows) between the ages of 18 and 40 with children living at home between 0 and 12 years old.

Table 6. Child Gender and the Probability a Mother Has Never Married, U.S. Census Data.

	Marginal Effect on the Probability of Never Married							
Sex of 1st child	Families with 1 child (1)	Families with ≥ 1 children (2)	Sex order of 1st two children	Families with 2 children (3)	Families with ≥ 2 children (4)	Sex order of 1st three children	Families with 3 children (5)	Families with ≥ 3 children (6)
Girl	0.0018	0.0017	Girl, Girl	0.0035	0.0024	G, G, G	0.0018	0.0022
	(0.0005)	(0.0003)		(0.0005)	(0.0004)		(0.0010)	(0.0008)
			Boy, Girl	-0.0015	0.0005	B, B, G	-0.0003	0.0012
				(0.0005)	(0.0004)		(0.0010)	(0.0008)
			Girl, Boy	-0.0017	0.0007	B, G, B	0.0028	0.0039
				(0.0005)	(0.0004)		(0.0010)	(0.0009)
						G, B, B	0.0050	0.0053
							(0.0010)	(0.0009)
						B, G, G	0.0052	0.0055
							(0.0010)	(0.0009)
						G, B, G	0.0052	0.0059
							(0.0010)	(0.0009)
						G, G, B	-0.0009	0.0004
							(0.0010)	(0.0008)
All-Boy Baseline	0.1244	0.0786		0.0506	0.0494		0.0461	0.0470
Percent Effect	1.45%	2.18%		6.86%	4.85%		3.91%	4.72%
Obs.	1,734,451	4,463,749		1,749,417	2,729,298		688,766	979,881

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the mother has never married at the time of the survey. The excluded category is all boys. All-boy baseline is the fraction of mothers in all-boy families who never married, i.e., the intercept term. Percent effect is the increase in the probability of never having been married for the mother of an all-girl family compared to an all-boy family.

Table 7. Child Gender and the Probability of a Shotgun Marriage, California Birth Certificate Data.

	Marginal Effect on the Probability of a Shotgun Marriage							
			Ultrasound during Pregnancy				Ultrasound during Labor	
			Unweighted		Weighted			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Girl	-.0003	.0007	.0010	.0019	.0014	.0021	-.0004	.0007
	(.0008)	(.0007)	(.0010)	(.0009)	(.0010)	(.0009)	(.0008)	(.0007)
Girl*Ultrasound			-.0037	-.0030	-.0046	-.0039	-.0024	-.0007
			(.0016)	(.0014)	(.0017)	(.0015)	(.0055)	(.0049)
Ultrasound			.0657	.0303	.0521	.0300	-.0003	-.0057
			(.0011)	(.0010)	(.0012)	(.0011)	(.0038)	(.0034)
Controls?	No	Yes	No	Yes	No	Yes	No	Yes
All-Boy Baseline			.084	.084	.084	.084		
Percent Effect			-4.4%	-3.6%	-5.5%	-4.6%		
R-squared	0.00	0.22	0.01	0.23	0.01	0.19	0.00	0.22
Obs.	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601

Notes: Standard errors in parentheses. Data are from the 1989-1994 California Birth Statistical Master File. The sample includes all first-time mothers. The dependent variable is a dummy equal to one if the mother is married when the baby is born. Models in columns 2, 4, 6, and 8 control for mother's race (three groups), mother's education, mother's age, mother's immigrant status, mother's Hispanic status, and year. The weighted regressions use as weights the predicted probability a woman is unmarried at the time of birth, using mother's race, mother's age, mother's immigrant status, mother's Hispanic status, year and all their interactions. The all-boy baseline is the probability an unmarried woman gets pregnant and marries before the birth of her first child, calculated using the 1980, 1985, 1990, and 1995 CPS Fertility Supplements. Percent effect is the decrease in the probability of being married at the time of birth for the mother of a girl compared to the mother of a boy.

Table 8. Child Gender and Fertility, U.S. Census Data.

Marginal Effect on the Probability of Another Child							
Families with 1 or more children		Families with 2 or more children		Families with 3 or more children		Families with 4 or more children	
Sex of 1st child	At least 1 more child (2+) (1)	Sex order of 1st two children	At least 1 more child (3+) (2)	Sex order of 1st three children	At least 1 more child (4+) (3)	Sex mix of 1st four children	At least 1 more child (5+) (4)
Girl	-0.0018	Girl, Girl	0.0089	G, G, G	0.0079	4 G	0.0139
	(0.0005)		(0.0009)		(0.0019)		(0.0040)
		Boy, Girl	-0.0464	B, B, G	-0.0340	2 G, 2 B	0.0140
			(0.0009)		(0.0019)		(0.0029)
		Girl, Boy	-0.0449	B, G, B	-0.0085	3 G, 1 B	0.0115
			(0.0009)		(0.0020)		(0.0029)
				G, B, B	-0.0091	1 G, 3 B	0.0012
					(0.0019)		(0.0025)
				B, G, G	-0.0044		
					(0.0020)		
				G, B, G	-0.0022		
					(0.0020)		
				G, G, B	-0.0229		
					(0.0019)		
All-Boy Baseline	0.6471		0.3821		0.3045		0.2847
Percent Effect	-0.27%		2.34%		2.61%		4.89%
Obs.	3,599,664		2,326,339		841,264		248,315

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all currently-married mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable in column 1 is a dummy equal to one if the family has 2 or more children; in column 2 a dummy equal to one if the family has 3 or more children; in column 3 a dummy equal to one if the family has 4 or more children; and in column 4 a dummy equal to one if the family has 5 or more children. The excluded category is all boys. All-boy baseline is the fraction of couples with all-boy families who have an additional child, i.e., the intercept term. Percent effect is the increase in the probability of an additional child for the mother of an all-girl family compared to an all-boy family.

Table 9. Child Gender and Fertility, Longitudinal California Birth Certificate Data.

Marginal Effect on the Probability of Another Child					
Families with 1 or more children		Families with 2 or more children		Families with 3 or more children	
Sex of 1st child	At least 1 more child (2+) (1)	Sex order of 1st two children	At least 1 more child (3+) (2)	Sex order of 1st three children	At least 1 more child (4+) (3)
Girl	-.0008	Girl, Girl	.0126	Girl, Girl, Girl	.0113
	(.0015)		(.0026)		(.0052)
		Boy, Girl	-.0437	Boy, Boy, Girl	-.0426
			(.0026)		(.0051)
		Girl, Boy	-.0461	Boy, Girl, Boy	-.0228
			(.0026)		(.0053)
				Girl, Boy, Boy	-.0196
					(.0054)
				Boy, Girl, Girl	-.0173
					(.0053)
				Girl, Boy, Girl	-.0196
					(.0054)
				Girl, Girl, Boy	-.0274
					(.0051)
All-Boy Baseline	.4690		.3082		.2034
Percent Effect	-0.18%		4.08%		5.52%
Obs.	406,412		231,811		83,025

Notes: Standard errors in parentheses. Data are a panel of California mothers obtained by longitudinally linking birth certificates for the years 1989-2001 from the California Birth Statistical Master File. The sample includes all first-time married mothers in 1989 or 1990, and follows them between 1989 and 2001. The dependent variable in column 1 is a dummy equal to one if the family has 2 or more children; in column 2 a dummy equal to one if the family has 3 or more children; in column 3 a dummy equal to one if the family has 4 or more children. The excluded category is all boys. All-boy baseline is the fraction of couples with all-boy families who have an additional child, i.e., the intercept term. Percent effect is the increase in the probability of an additional child for the mother of an all-girl family compared to an all-boy family.

Table 10. Child Gender and Fertility for Families with Two or More Children by Census Year, Decade of Birth, Race, and Education, U.S. Census Data.

	Coeff on Girl, Girl (1)	All-Boy Baseline (2)	Percent Effect (3)
Model 1: By Census Year			
1940-1960	.0005	.4991	0.09%
	(.0040)		
1970	.0051	.4882	1.05%
	(.0020)		
1980	.0154	.3327	4.62%
	(.0017)		
1990	.0073	.3330	2.18%
	(.0017)		
2000	.0071	.3651	1.94%
	(.0019)		
Model 2: By Decade of Birth (women age 20-30)			
1930s	.0033	.5171	0.64%
	(.0062)		
1940s	.0096	.4363	2.20%
	(.0025)		
1950s	.0165	.3097	5.33%
	(.0023)		
1960s	.0095	.3241	2.94%
	(.0025)		
1970s	.0103	.3670	2.85%
	(.0031)		
Model 3: By Race			
White	.0078	.3775	2.07%
	(.0009)		
Black	.0057	.4471	1.26%
	(.0035)		
Asian	.0515	.3041	16.94%
	(.0052)		
Other	.0041	.4399	.93%
	(.0053)		
Model 4: By Education			
HS <	.0125	.4930	2.54%
	(.0021)		
HS	.0096	.3782	2.54%
	(.0014)		
College	.0046	.3361	1.37%
	(.0013)		

Notes: Each panel is a separate regression. Sample includes families with 2 or more children. The models correspond to column 2 in Table 8. To conserve on space, only the coefficient for Girl, Girl families is reported. See notes to Table 8.

Table 11. Stated Preferences for a Boy or a Girl, Gallup Poll Data.

	Probit (1=Prefer Boy, 0=Prefer Girl)		
Variable	Coefficient (1)	Standard Error (2)	Marginal Effect (3)
Male	0.6128	(0.0752)	0.2335
Age	-0.0335	(0.0129)	-0.0130
Age Squared	0.0003	(0.0001)	0.0001
Race			
White	---	---	---
Black	0.3200	(0.1315)	0.1188
Other	0.0029	(0.1583)	0.0011
Race Unknown	-0.0811	(0.3836)	-0.0317
Hispanic	0.0854	(0.1394)	0.0331
Education			
High School Dropout	0.0142	(0.1108)	0.0055
High School	---	---	---
Some College or More	0.0329	(0.1180)	0.0127
Region			
East	---	---	---
Midwest	-0.0570	(0.1052)	-0.0222
South	-0.1111	(0.0989)	-0.0432
West	0.0656	(0.1094)	0.0253
Urban/Rural Status			
Urban	---	---	---
Suburban	0.0449	(0.0848)	0.0174
Rural	0.1150	(0.1050)	0.0442
Income			
Income < 30k	---	---	---
30k $\leq$ Income $<$ 50k	-0.0973	(0.1850)	-0.0381
50k $\leq$ Income $<$ 75k	0.1193	(0.1011)	0.0459
Income $\geq$ 75k	0.1822	(0.1165)	0.0695
Refused	0.0550	(0.1141)	0.0212
Marital Status			
Divorced or Separated	0.1899	(0.1306)	0.0721
Married	0.1025	(0.1022)	0.0397
Widowed	0.0111	(0.1785)	0.0043
Never Married	---	---	---
Attend Church Often	-0.2098	(0.0767)	-0.0810
Ideology			
Conservative	---	---	---
Moderate	-0.0464	(0.0829)	-0.0180
Liberal	-0.3693	(0.1010)	-0.1452
Refused	-0.0095	(0.2859)	-0.0037
Constant	0.5490	(0.3944)	
Log Likelihood	-837.12		
Observations	1,325		

Notes: Data are from the December 2-4, 2000 and July 18-20, 2003 Gallup Poll Surveys. The question asked of respondents was "Suppose you could only have one child. Would you prefer that it be a boy or a girl?" The survey data is based on telephone interviews of a national random sample of adults age 18 and older. This table includes only individuals with a stated preference for a boy or a girl; multinomial logit results including "no preference" as a third option are summarized in the text and available on request.

Table 12. Child Gender and the Probability of Receiving Child Support, March Current Population Survey Data.

Marginal Effect on the Probability of Receiving Child Support			
	Mothers with 1 or more children (1)	Mothers with 2 or more children (2)	Mothers with 3 or more children (3)
Model 1			
Girl	-.0036		
	(.0065)		
Model 2			
Girl, Girl		-.0290	
		(.0142)	
Mixed Gender		-.0126	
		(.0126)	
Model 3			
Girl, Girl, Girl			-.0550
			(.0338)
Mixed Gender			-.0054
			(.0264)
All-Boy Baseline	.280	.318	.284
Percent Effect	-1.2%	-9.1%	-19.3%
Obs.	17,767	7,706	2,433

Notes: Standard errors in parentheses. Data are from the 1995-2000 March Current Population Surveys. All models control for mother's age, mother's race, and year. The sample includes all single mothers who are the household head between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the family receives child support. The excluded category is all boys. All-boy baseline is the predicted probability that a woman in an all-boy family receives child support, using the estimated regression coefficients and the explanatory variables evaluated at their means. Percent effect is the decrease in the probability of receiving child support for the mother of an all-girl family compared to an all-boy family.

Table 13. Child Gender and the Probability of Divorce or Separation and Father Custody, International Data.

	Marginal Effect on the Probability of Current Divorce or Separation and Father Custody							
	Families with 1 or more children		Sex mix of 1st two children	Families with 2 or more children		Sex mix of 1st three children	Families with 3 or more children	
Sex of 1st child	Divorce (1)	Custody (2)		Divorce (3)	Custody (4)		Divorce (5)	Custody (6)
Mexico								
Girl	0.0015	-0.0385	Girl, Girl	0.0034	-0.0339	G, G, G	0.0016	-0.0403
	(.0007)	(0.0048)		(0.0010)	(0.0067)		(0.0018)	(0.0154)
			1 B, 1 G	0.0010	-0.0257	2 G, 1 B	-0.0001	-0.0448
				(0.0008)	(0.0057)		(0.0015)	(0.0128)
						2 B, 1G	0.0009	-0.0214
							(0.0015)	(0.0127)
All-Boy Baseline	.0639	0.2293		0.0532	0.1735		0.0473	0.2186
Percent Effect	2.35%	-16.78%		6.39%	-19.54%		3.38%	-18.44%
Obs.	453,545	29,332		433,203	23,629		202,411	9,667
Colombia								
Girl	0.0046	-0.0326	Girl, Girl	0.0032	-0.0382	G, G, G	0.0041	-0.0333
	(0.0012)	(0.0055)		(0.0016)	(0.0087)		(0.0030)	(0.0019)
			1B, 1G	0.0008	-0.0123	2 G, 1 B	0.0015	0.0001
				(0.0014)	(0.2924)		(0.0025)	(0.0154)
						2 B, 1G	0.0012	0.0161
							(0.0025)	(0.0153)
All-Boy Baseline	0.1062	0.3221		0.0864	0.2924		0.0800	0.3648
Percent Effect	4.33%	-10.12%		3.70%	-13.06%		5.13%	-9.13%
Obs.	258,932	28,091		238,409	20,878		122,265	9,965
Kenya								
Girl	0.0011	-0.0515	Girl, Girl	0.0015	-0.0708	G, G, G	-0.0007	-0.0801
	(0.0011)	(0.0128)		(0.0016)	(0.0206)		(0.0028)	(0.0402)
			1B, 1G	-0.0002	-0.0131	2 G, 1 B	-0.0037	-0.0287
				(0.0014)	(0.0180)		(0.0023)	(0.0332)
						2 B, 1G	-0.0023	0.0313
							(0.0023)	(0.0328)
All-Boy Baseline	0.0385	0.2658		0.0334	0.2773		0.0351	0.3531
Percent Effect	2.86%	-19.38%		4.49%	25.53%		1.99%	-22.68%
Obs.	114,273	4,459		102,908	3,459		62,099	2,039

Table 13, continued. Child Gender and the Probability of Divorce or Separation and Father Custody, International Data.

	Marginal Effect on the Probability of Current Divorce or Separation and Father Custody							
	Families with 1 or more children		Sex mix of 1st two	Families with 2 or more children		Sex mix of 1st three	Families with 3 or more children	
Sex of 1st child	Divorce (1)	Custody (2)	children	Divorce (3)	Custody (4)	children	Divorce (5)	Custody (6)
Vietnam								
Girl	-0.0005	-0.0960	Girl, Girl	0.0005	-0.1154	G, G, G	-0.0023	-0.1942
	(0.0008)	(0.0146)		(0.0009)	(0.0240)		(0.0016)	(0.0558)
			1B, 1G	-0.0004	-0.0488	2 G, 1 B	-0.0016	-0.0490
				(0.0008)	(0.0208)		(0.0013)	(0.0451)
						2 B, 1G	-0.0019	-0.0104
							(0.0013)	(0.0454)
All-Boy Baseline	0.0265	0.4813		0.0184	0.4020		0.0155	0.5808
Percent Effect	-1.89%	-19.95%		2.72%	-28.71%		-14.84%	-33.44%
Obs.	173,405	4,554		166,660	3,062		84,942	1,176

Notes: Standard errors in parentheses. Data are from the 1990 and 2000 Mexican Censuses; the 1973, 1985, and 1993 Colombian Censuses; the 1989 and 1999 Kenyan Censuses; and the 1989 and 1999 Vietnam Censuses. The divorce sample includes all households with parents who are ever married (excluding widows), who are between the ages of 18 and 40, and who have children living at home between 0 and 12 years old. The dependent variable for the divorce sample is a dummy equal to one if the parent is divorced or separated at the time of the survey. The excluded category is all boys. All-boy baseline is the fraction of parents in all-boy families who are divorced or separated, i.e., the intercept term. Percent effect is the increase in the probability of divorce or separation for the parent of an all-girl family compared to an all-boy family. The custody sample includes all divorced or separated fathers and mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable for the custody sample is a dummy equal to one if the parent living with the children is the father. The excluded category is all boys. All-boy baseline is the fraction of divorced parents in all-boy families who are fathers, i.e., the intercept term. Percent effect is the decrease in the probability the divorced parent living with the children is the father for an all-girl family compared to an all-boy family.

Table 14. Child Gender and Fertility, International Data.

	Marginal Effect on the Probability of Another Child						
	Families with 1 or more children	Families with 2 or more children		Families with 3 or more children		Families with 4 or more children	
Sex of 1st child	At least 1 more child (2+) (1)	Sex mix of 1st two children	At least 1 more child (3+) (2)	Sex mix of 1st three children	At least 1 more child (4+) (3)	Sex mix of 1st four children	At least 1 more child (5+) (4)
Mexico							
Girl	0.0005	Girl, Girl	0.0230	G, G, G	0.0274	G, G, G, G	0.0250
	(0.0006)		(0.0019)		(0.0038)		(0.0069)
		1 G, 1 B	-0.0202	2 G, 1 B	-0.0002	2 G, 2 B	-0.0064
			(0.0017)		(0.0031)		(0.0044)
				2 B, 1 G	-0.0119	3 G, 1 B	0.0146
					(0.0031)		(0.0051)
						1 G, 3 B	-0.0036
							(0.0051)
All-Boy Baseline	0.9556		0.4756		0.3638		0.3309
Percent Effect	0.05%		4.84%		7.52%		7.54%
Obs.	548,885		524,668		247,197		89,688
Colombia							
Girl	-0.0028	Girl, Girl	0.0050	G, G, G	0.0106	G, G, G, G	0.0167
	(0.0009)		(0.0024)		(0.0046)		(0.0079)
		1 G, 1 B	-0.0248	2 G, 1 B	-0.0104	2 G, 2 B	-0.0027
			(0.0021)		(0.0038)		(0.0049)
				2 B, 1 G	-0.0084	3 G, 1 B	0.0076
					(0.0038)		(0.0058)
						1 G, 3 B	0.0051
							(0.0056)
All-Boy Baseline	0.9194		0.5278		0.4554		0.4251
Percent Effect	-0.30%		0.95%		2.32%		3.92%
Obs.	365,423		335,469		173,372		77,997
Kenya							
Girl	0.0021	Girl, Girl	0.0202	G, G, G	0.0169	G, G, G, G	0.0035
	(0.0018)		(0.0041)		(0.0076)		(0.0120)
		1 G, 1 B	-0.0005	2 G, 1 B	0.0047	2 G, 2 B	-0.0056
			(0.0036)		(0.0062)		(0.0074)
				2 B, 1 G	-0.0059	3 G, 1 B	0.0035
					(0.0062)		(0.0086)
						1 G, 3 B	-0.0012
							(0.0085)
All-Boy Baseline	0.8908		0.6039		0.5113		0.4008
Percent Effect	0.24%		3.35%		3.30%		0.87%
Obs.	124,294		110,858		67,477		34,611

Table 14, continued. Child Gender and Fertility, International Data.

	Marginal Effect on the Probability of Another Child						
	Families with 1 or more children	Families with 2 or more children		Families with 3 or more children		Families with 4 or more children	
Sex of 1st child	At least 1 more child (2+) (1)	Sex mix of 1st two children	At least 1 more child (3+) (2)	Sex mix of 1st three children	At least 1 more child (4+) (3)	Sex mix of 1st four children	At least 1 more child (5+) (4)
Vietnam							
Girl	0.0029	Girl, Girl	0.0871	G, G, G	0.0734	G, G, G, G	0.0869
	(0.0008)		(0.0035)		(0.0066)		(0.0105)
		1 G, 1 B	0.0067	2 G, 1 B	-0.0004	2 G, 2 B	-0.0168
			(0.0030)		(0.0055)		(0.0069)
				2 B, 1 G	-0.0365	3 G, 1 B	0.0333
					(0.0055)		(0.0080)
						1 G, 3 B	0.0103
							(0.0081)
All-Boy Baseline	0.9674		0.4875		0.4102		0.3021
Percent Effect	0.30%		17.87%		17.89%		28.77%
Obs.	168,851		163,598		83,766		34,057
China							
Girl	0.0053	Girl, Girl	0.1990	G, G, G	0.2114	G, G, G, G	0.1833
	(0.0012)		(0.0061)		(0.0116)		(0.0181)
		1 G, 1 B	0.0251	2 G, 1 B	0.0231	2 G, 2 B	-0.0133
			(0.0053)		(0.0099)		(0.0130)
				2 B, 1 G	-0.0506	3 G, 1 B	0.0579
					(0.0101)		(0.0150)
						1 G, 3 B	0.0021
							(0.0172)
All-Boy Baseline	0.9777		0.3654		0.2344		0.1429
Percent Effect	0.54%		54.45%		90.20%		128.29%
Obs.	52,788		51,751		22,085		5,677

Notes: Standard errors in parentheses. Data are from the 1990 and 2000 Mexican Censuses; the 1973, 1985, and 1993 Colombian Censuses; the 1989 and 1999 Kenyan Censuses; the 1989 and 1999 Vietnam Censuses; and the 1982 Chinese Census. The sample includes all currently-married mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable in column 1 is a dummy equal to one if the family has 2 or more children; in column 2 a dummy equal to one if the family has 3 or more children; in column 3 a dummy equal to one if the family has 4 or more children; and in column 4 a dummy equal to one if the family has 5 or more children. The excluded category is all boys. All-boy baseline is the fraction of couples with all-boy families who have an additional child, i.e., the intercept term. Percent effect is the increase in the probability of an additional child for the mother of an all-girl family compared to an all-boy family.

Table 15. Child Gender and the Probability of Living in a Polygamous Relationship, Kenyan Census Data.

Sex of 1st child	Families with ≥ 1 children (1)	(2)	Sex order of 1st two children	Families with ≥ 2 children (3)	(4)	Sex order of 1st three children	Families with ≥ 3 children (5)	(6)
	Marginal Effect on the Probability of a Polygamous Relationship							
Girl	0.0030	0.0030	Girl, Girl	0.0052	0.0069	G, G, G	0.0026	0.0074
	(0.0018)	(0.0017)		(0.0026)	(0.0024)		(0.0047)	(0.0045)
			Boy, Girl	-0.0010	0.0012	B, B, G	-0.0030	-0.0055
				(0.0026)	(0.0025)		(0.0048)	(0.0045)
			Girl, Boy	0.0006	0.0013	B, G, B	-0.0036	-0.0018
				(0.0026)	(0.0024)		(0.0048)	(0.0046)
						G, B, B	-0.0002	0.0012
							(0.0047)	(0.0045)
						B, G, G	0.0069	0.0080
							(0.0048)	(0.0046)
						G, B, G	0.0022	0.0015
							(0.0048)	(0.0046)
						G, G, B	0.0039	0.0060
							(0.0047)	(0.0045)
Controls?	No	Yes		No	Yes		No	Yes
All-Boy Baseline	0.1150	.1147		0.1017	.1016		0.1088	0.1089
Percent Effect	2.63%	2.65%		5.12%	6.80%		2.38%	6.79%
R-squared	.0000	.1030		.0001	.0995		.0001	.0973
Obs.	124,294	124,294		110,858	110,858		67,477	67,477

Notes: Standard errors in parentheses. Data are from the 1989 and 1999 Kenyan Censuses. The sample includes all currently-married mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the mother is married to a man with more than one wife, i.e., the dependent variable refers to polygyny and not polyandry. The excluded category is all boys. Controls include mother's age (cubic), district of residence in Kenya, nativity (born in Kenya), census year, literacy (unknown, literate, illiterate), and education (unknown, less than primary, primary, lower secondary, secondary completed, university). In columns 1, 3, and 5 all-boy baseline is the fraction of mothers in all-boy families who are in a polygamous relationship, i.e., the intercept term. In columns 2, 4, and 6 it is the predicted probability a mother in an all-boy family is in a polygamous relationship, using the estimated regression coefficients and the explanatory variables evaluated at their means. Percent effect is the increase in the probability of being in a polygamous relationship for the mother of an all-girl family compared to an all-boy family.

Appendix Table A1. Child Gender and the Probability of Current Divorce or Separation, Controlling for Parents' Characteristics, U.S. Census Data.

	Marginal Effect on the Probability of Current Divorce or Separation										
Sex of 1st child	Families with 1 child (1)	Families with ≥ 1 children (2)	Sex order of 1st two children	Families with 2 children (3)	Families with ≥ 2 children (4)	Sex order of 1st three children	Families with 3 children (5)	Families with ≥ 3 children (6)	Sex mix of 1st four children	Families with 4 children (7)	Families with ≥ 4 children (8)
Girl	-0.0006	0.0008	Girl, Girl	0.0015	0.0017	G, G, G	0.0044	0.0036	4 G	0.0052	0.0055
	(0.0006)	(0.0003)		(0.0007)	(0.0005)		(0.0014)	(0.0011)		(0.0029)	(0.0024)
			Boy, Girl	-0.0050	-0.0014	B, B, G	-0.0006	0.0002	2 G, 2 B	0.0056	0.0053
				(0.0006)	(0.0005)		(0.0013)	(0.0011)		(0.0018)	(0.0015)
			Girl, Boy	-0.0030	0.0001	B, G, B	0.0052	0.0033	3 G, 1 B	0.0024	0.0012
				(0.0005)	(0.0005)		(0.0014)	(0.0012)		(0.0021)	(0.0018)
						G, B, B	0.0046	0.0035	1 G, 3 B	-0.0026	0.0001
							(0.0014)	(0.0012)		(0.0014)	(0.0018)
						B, G, G	0.0010	0.0012			
							(0.0014)	(0.0012)			
						G, B, G	0.0016	0.0007			
							(0.0014)	(0.0012)			
						G, G, B	0.0004	0.0009			
							(0.0014)	(0.0011)			
All-Boy Baseline	0.1813	0.1350		0.1146	0.1098		0.0999	0.0979		0.0985	0.0986
Percent Effect	-0.2%	0.7%		1.3%	1.9%		4.9%	4.1%		5.9%	5.8%
Obs.	1,554,818	4,169,265		1,679,127	2,614,447		659,523	935,320		195,586	275,797

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all households with parents who are ever married (excluding widows), who are between the ages of 18 and 40, and who have children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the parent is divorced or separated at the time of the survey. The excluded category is all boys. All-boy baseline is the predicted probability a parent in an all-boy family is divorced or separated, using the estimated regression coefficients and the explanatory variables evaluated at their means. Percent effect is the increase in the probability of divorce or separation for a parent of an all-girl family compared to an all-boy family. Regressions include controls for parent's cohort (10-year birth cohorts), age (cubic), geographic residence (9 regions), race (black, white, other), and education (less than high school, high school, college).

Appendix Table A2. Child Gender and the Probability of Father Custody, Controlling for Parents' Characteristics, U.S. Census Data.

	Marginal Effect on the Probability of Father Custody										
Sex of 1st child	Families with 1 child (1)	Families with ≥ 1 children (2)	Sex order of 1st two children	Families with 2 children (3)	Families with ≥ 2 children (4)	Sex order of 1st three children	Families with 3 children (5)	Families with ≥ 3 children (6)	Sex mix of 1st four children	Families with 4 children (7)	Families with ≥ 4 children (8)
Girl	-0.0438	-0.0272	Girl, Girl	-0.0258	-0.0182	G, G, G	-0.0119	-0.0104	4 G	-0.0122	-0.0129
	(0.0014)	(0.0009)		(0.0022)	(0.0017)		(0.0049)	(0.0040)		(0.0094)	(0.0077)
			Boy, Girl	-0.0159	-0.0111	B, B, G	-0.0134	-0.0086	2 G, 2 B	0.0018	-0.0003
				(0.0022)	(0.0017)		(0.0048)	(0.0040)		(0.0060)	(0.0049)
			Girl, Boy	-0.0172	-0.0136	B, G, B	-0.0103	-0.0050	3 G, 1 B	-0.0042	-0.0032
				(0.0022)	(0.0017)		(0.0050)	(0.0041)		(0.0071)	(0.0058)
						G, B, B	-0.0144	-0.0101	1 G, 3 B	0.0079	0.0016
							(0.0050)	(0.0041)		(0.0070)	(0.0058)
						B, G, G	-0.0061	-0.0071			
							(0.0051)	(0.0041)			
						G, B, G	-0.0175	-0.0127			
							(0.0051)	(0.0041)			
						G, G, B	-0.0080	-0.0047			
							(0.0049)	(0.0042)			
All-Boy Baseline	0.1995	0.1694		0.1492	0.1398		0.1272	0.1201		0.1075	0.1032
Percent Effect	-22.0%	-16.1%		-18.0%	-13.9%		-10.5%	-12.3%		-11.5%	-14.0%
Obs.	281,493	569,601		194,052	288,108		66,574	94,056		19,435	27,482

Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all divorced or separated fathers and mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable is a dummy equal to one if the parent living with the children is the father. The excluded category is all boys. All-boy baseline is the predicted probability the divorced parent living with an all-boy family is the father, using the estimated regression coefficients and the explanatory variables evaluated at their means. Percent effect is the decrease in the probability the divorced parent living with the children is the father for an all-girl family compared to an all-boy family. Regressions include controls for parent's cohort (10-year birth cohorts), age (cubic), geographic residence (9 regions), race (black, white, other), and education (less than high school, high school, college).

Appendix Table A3. Child Gender and the Probability of Remarriage and Second Divorce, U.S. Census Data.

Sex of 1st child	Families with ≥ 1 children (1)	Sex order of 1st two children	Families with ≥ 2 children (2)	Sex order of 1st three children	Families with ≥ 3 children (3)
Marginal Effect on the Probability of Remarriage after a Divorce					
Girl	0.0007	Girl, Girl	0.0002	Girl, Girl, Girl	-0.0071
	(0.0018)		(0.0032)		(0.0067)
		Boy, Girl	-0.0030	Boy, Boy, Girl	0.0086
			(0.0032)		(0.0067)
		Girl, Boy	0.0011	Boy, Girl, Boy	0.0078
			(0.0032)		(0.0069)
				Girl, Boy, Boy	-0.0011
					(0.0069)
				Boy, Girl, Girl	-0.0020
					(0.0069)
				Girl, Boy, Girl	0.0171
					(0.0070)
				Girl, Girl, Boy	0.0070
					(0.0068)
All-Boy Baseline	0.6592		0.7130		0.7641
Percent Effect	0.11%		0.03%		-0.94%
Obs.	280,598		158,736		58,355
Marginal Effect on the Probability of a Second Divorce or Separation					
Girl	0.0048	Girl, Girl	0.0062	Girl, Girl, Girl	0.0046
	(0.0016)		(0.0027)		(0.0057)
		Boy, Girl	0.0013	Boy, Boy, Girl	0.0020
			(0.0027)		(0.0057)
		Girl, Boy	-0.0017	Boy, Girl, Boy	0.0074
			(0.0027)		(0.0058)
				Girl, Boy, Boy	0.0018
					(0.0058)
				Boy, Girl, Girl	-0.0011
					(0.0059)
				Girl, Boy, Girl	0.0011
					(0.0059)
				Girl, Girl, Boy	0.0077
					(0.0057)
All-Boy Baseline	0.1368		0.1202		0.1064
Percent Effect	3.49%		5.18%		4.33%
Obs.	185,078		113,117		44,801

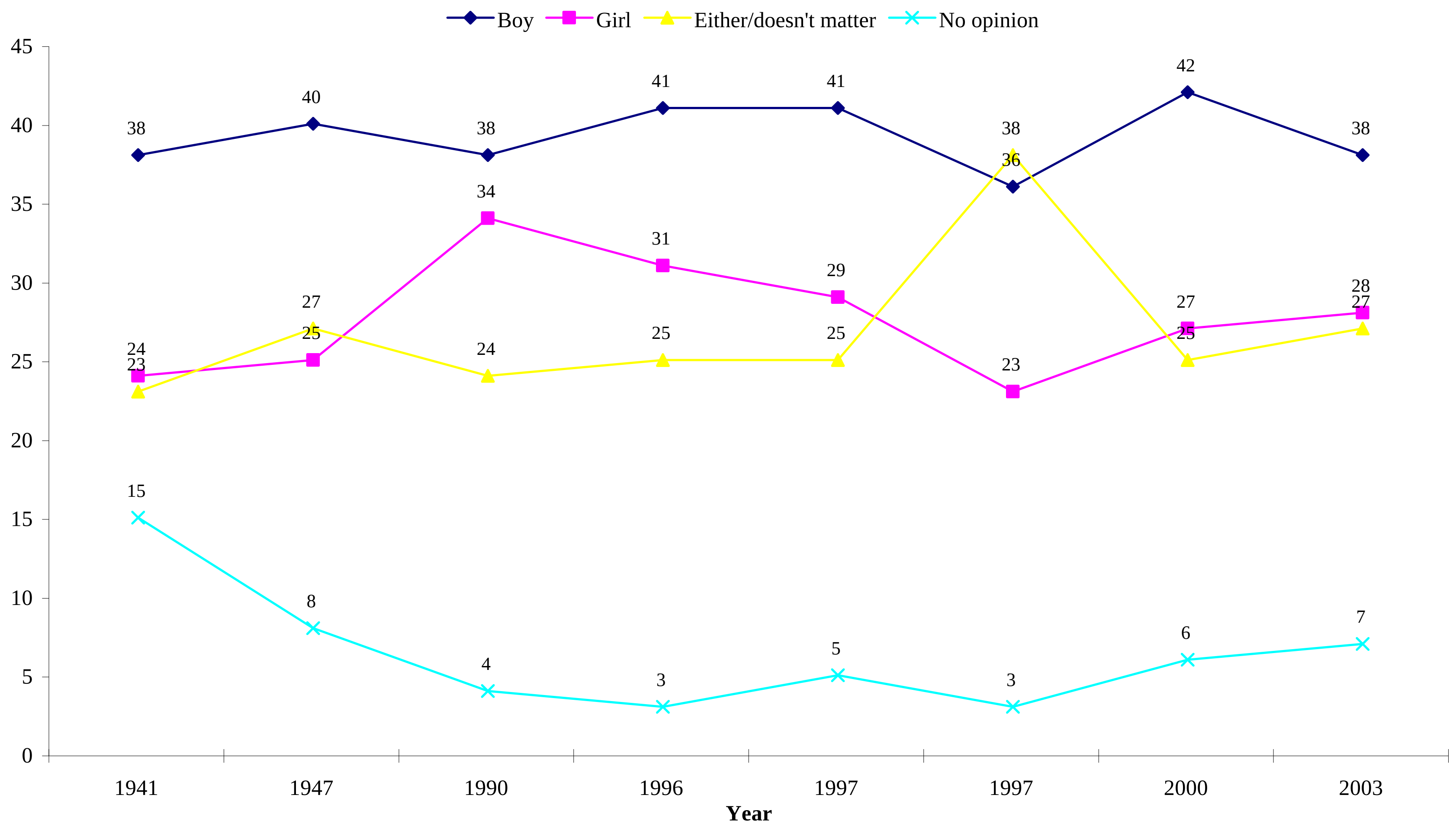
Notes: Standard errors in parentheses. Data are from the 1940 to 1980 U.S. Censuses. The samples include mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. In the top panel, the sample includes all mothers whose first marriage ended in divorce. In the bottom panel, the sample includes all mothers who remarried after a first divorce. The dependent variables are dummies for whether the mother remarried and whether the mother divorced or separated a second time. The excluded category is all boys. All-boy baseline and percent effect are defined as in previous tables.

Appendix Table A4. Child Gender and Fertility, Controlling for Mothers' Characteristics, U.S. Census Data.

Marginal Effect on the Probability of Another Child							
Families with 1 or more children		Families with 2 or more children		Families with 3 or more children		Families with 4 or more children	
Sex of 1st child	At least 1 more child (2+) (1)	Sex order of 1st two children	At least 1 more child (3+) (2)	Sex order of 1st three children	At least 1 more child (4+) (3)	Sex mix of 1st four children	At least 1 more child (5+) (4)
Girl	-0.0020	Girl, Girl	0.0078	G, G, G	0.0063	4 G	0.0125
	(0.0005)		(0.0009)		(0.0019)		(0.0039)
		Boy, Girl	-0.0467	B, B, G	-0.0346	2 G, 2 B	0.0099
			(0.0009)		(0.0018)		(0.0029)
		Girl, Boy	-0.0456	B, G, B	-0.0126	3 G, 1 B	0.0081
			(0.0009)		(0.0019)		(0.0028)
				G, B, B	-0.0131	1 G, 3 B	-0.0023
					(0.0019)		(0.0024)
				B, G, G	-0.0097		
					(0.0019)		
				G, B, G	-0.0075		
					(0.0019)		
				G, G, B	-0.0231		
					(0.0019)		
All-Boy Baseline	0.6471		0.3821		0.3045		0.2845
Percent Effect	-0.31%		2.04%		2.07%		4.40%
Obs.	3,599,664		2,326,339		841,264		248,315

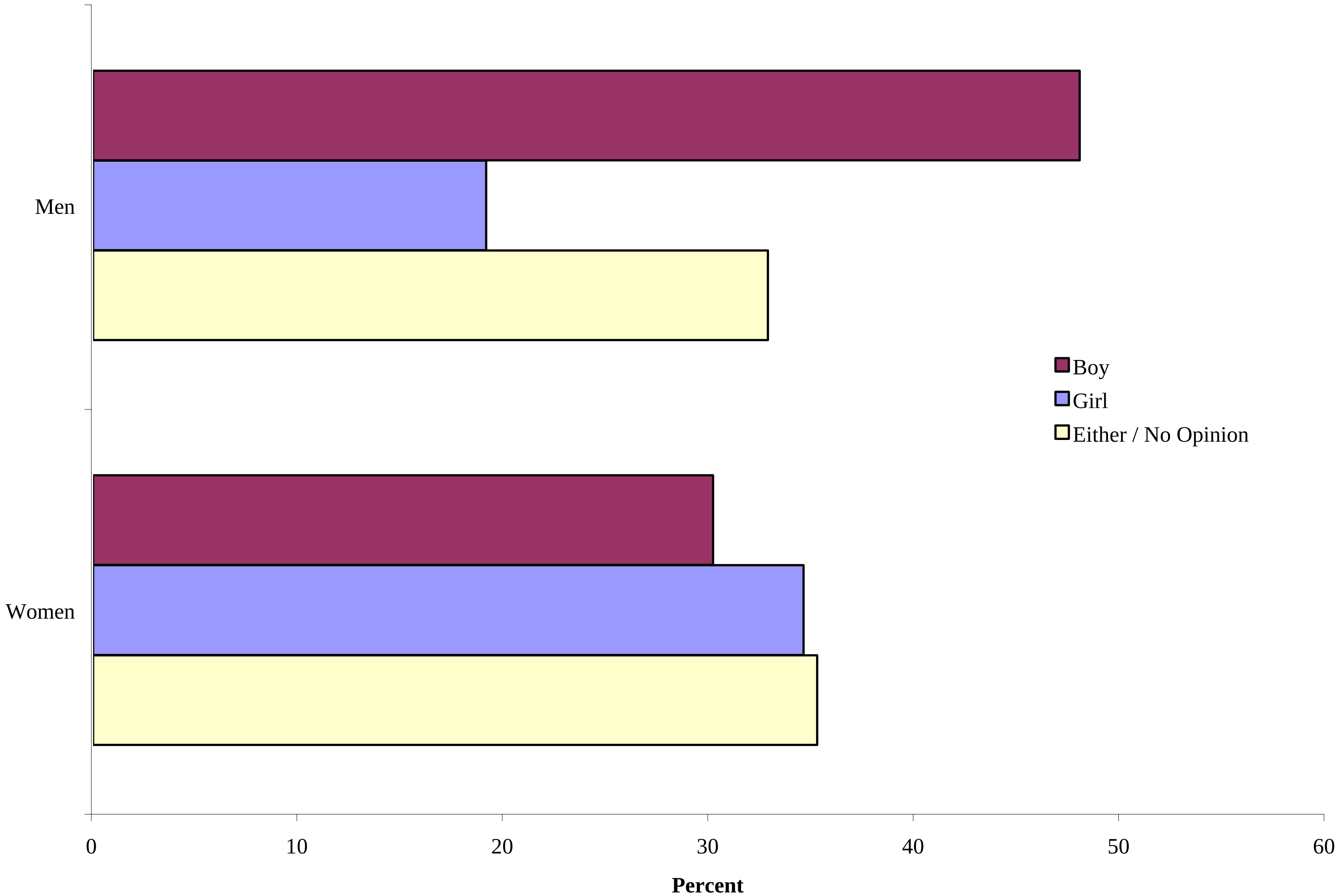
Notes: Standard errors in parentheses. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all currently-married mothers between the ages of 18 and 40 with children living at home between 0 and 12 years old. The dependent variable in column 1 is a dummy equal to one if the family has 2 or more children; in column 2 a dummy equal to one if the family has 3 or more children; in column 3 a dummy equal to one if the family has 4 or more children; and in column 4 a dummy equal to one if the family has 5 or more children. The excluded category is all boys. All-boy baseline is the predicted probability a couple with all boys will have an additional child, using the estimated regression coefficients and the explanatory variables evaluated at their means. Percent effect is the increase in the probability of an additional child for the mother of an all-girl family compared to an all-boy family. Regressions include controls for mother's cohort (10-year birth cohorts), age (cubic), geographic residence (9 regions), race (black, white, other), and education (less than high school, high school, college).

Figure 1. Trend in Preference for Boys versus Girls, Gallup Poll Data.



Questions: *1997, 2000, 2003: Suppose you could only have one child. Would you prefer that it be a boy or a girl?*
1996, 1997: If you were to have a child right now, would you rather have a boy or a girl?
1947, 1990: If you had another child would you rather have a boy or a girl?
1941: If you could have only one (one more) child, which would you prefer to have--a boy, or a girl?

Figure 2. Preference for Boys versus Girls, for Men and Women, 2000 and 2003 Gallup Poll Data.



Question: *Suppose you could only have one child. Would you prefer that it be a boy or a girl?*

Figure 3. Percent Increase in the Probability of an Extra Child for an All-Girl Family Relative to an All-Boy Family, by Family Size.

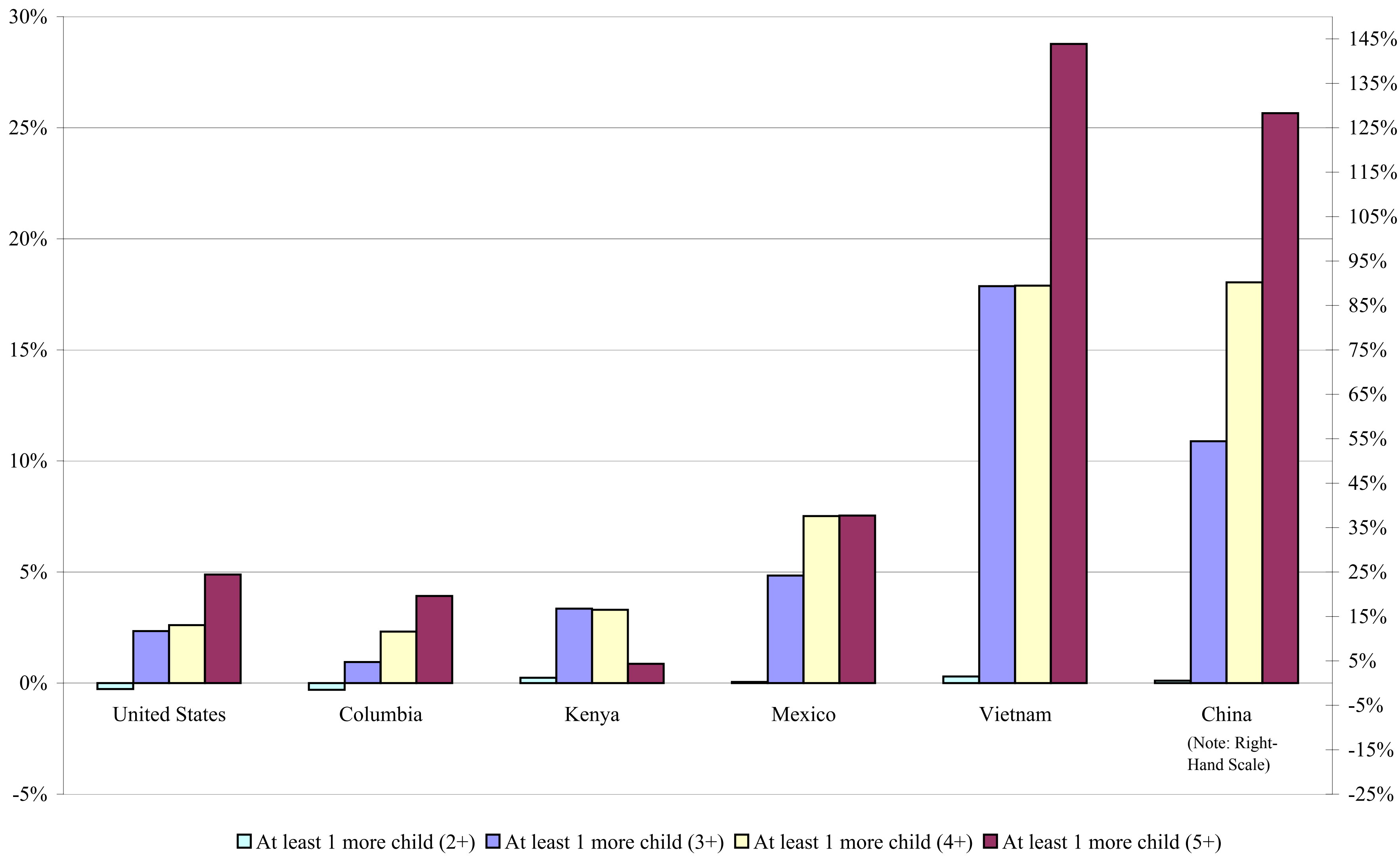
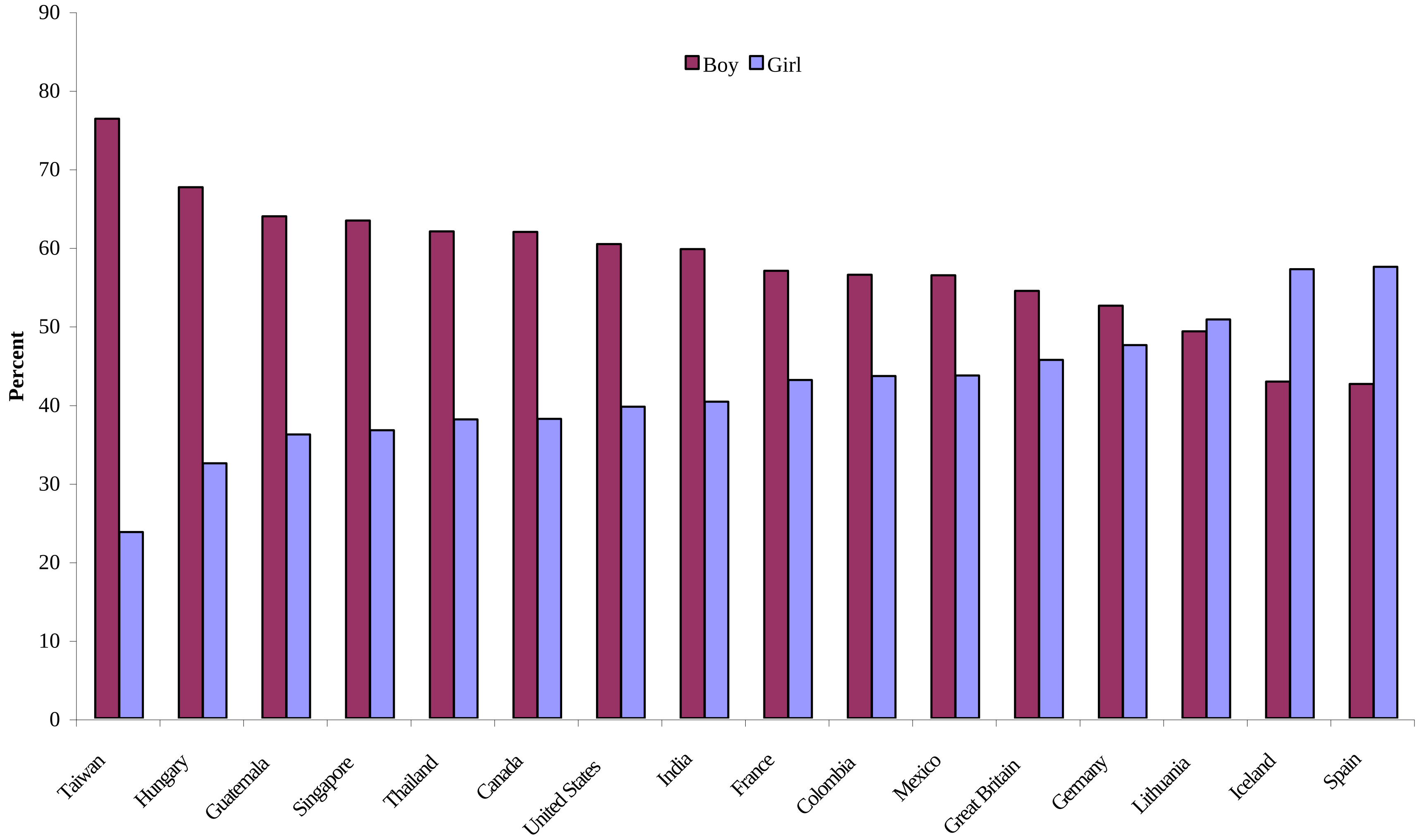


Figure 4. Preference for Boys versus Girls, International Gallup Poll Data.



Question: *Suppose you could only have one child. Would you prefer that it be a boy or a girl?*