The Demand for Sons:
Evidence from Divorce, Fertility, and Shotgun Marriage

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1 Introduction

In many developing countries, parents seem to have preferences for having 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). While there are no "missing" girls in the US, preferences for sons may take more subtle forms. In this paper we present several pieces of empirical evidence on the effect of child sex composition on parental marital status and fertility behavior. Specifically, we show that having girls has significant effects on divorce, marriage, shotgun marriage (i.e., when the sex of the child is known before birth), remarriage, fertility stopping rules, and child support payments.

Using a simple model, we show that, taken individually, each piece of evidence does not necessarily imply the existence of parental gender bias. But taken together, our empirical evidence indicates that US parents strongly favor boys over girls. The bias is quantitatively important, but seems to be slowly decreasing over time. When we compare the US with five developing countries, we find that gender bias in the US is generally smaller, and is only a fraction of the bias in China.

We begin by documenting the effect of offspring gender composition on the probability of divorce. We find that mothers with girls are significantly more likely to be divorced than mothers with boys. The effect is quantitatively substantial, explaining a 4% to 8% rise in divorce rates in the U.S.[1] By itself, this effect is 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.[2] It is possible that parents have unbiased gender preferences, but they decide to avoid or delay divorce if they have boys because they realize the harmful effects of raising a son without a father present in the household. We call this the "role model" hypothesis. Alternatively, it is also possible that the monetary, psychological, 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 turn to the effect of child gender composition on marriage. We find that, controlling

[1] While not the focus of this paper, we also find that gender mix reduces the probability of divorce. Other researchers have also documented a demand for variety (Ben-Porath and Welch, 1976, 1980; Rosenzweig and Wolpin, 1980). Angrist and Evans use the demand for a mixed sibling-sex composition to study the effect childbearing on female labor supply (1998).

[2] There is much evidence 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). Longitudinal data on child development show that the absence of a father has more severe and enduring impact on boys than girls. For example, boys are found to suffer more from divorces than girls (Hetereington, Cox and Cox, 1978). In most cases, children are assigned to the mother, irrespective of sex.

for family size, women with only girls are substantially more likely to have never been married than women with only boys. The chance a women will be married decreases by two to seven percent for an all-girl family relative to an all-boy family, depending on family size. For divorcees, a similar pattern emerges. Not only are divorced mothers with all-girl offspring less likely to remarry, when they do remarry they are more likely to get a second divorce.

Perhaps the most striking evidence comes from the analysis of shotgun marriages using Vital Statistics data. 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 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. We interpret this finding as evidence that fathers who find out their child will be a boy are more likely to marry their partner before delivery.[3]

Although the interpretation is not completely unambiguous, these findings on marriage, shotgun marriage, and remarriage are harder to explain under the role model hypothesis compared to the gender bias and differential costs 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 costs hypotheses from the gender bias hypothesis. We find that in families with at least two children, the probability of having another child is higher for all-girl families that all-boy families. The magnitude of the effect increases when we look at families with at least three children. Based on our model, this result would be hard to explain if parents were completely gender unbiased or if the cost of raising girls was higher than the cost of raising boys.[4]

In the final part of the paper, we extend our analysis to five developing countries: Mexico, Columbia, Kenya, Vietnam, and China. For divorce, the negative effect of an all-girl family is substantial, with effects that are two to eight times as large as in the U.S. When we

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[3] We use information on ultrasounds test during labor to perform a specification check. Unlike for ultrasound during pregnancy, the interaction of gender and ultrasound during labor has no effect on marital status at delivery.

[4] The fertility result alone does not necessarily imply that parents favor boys. Daughter preferences could also explain this finding, if parents can not perfectly control the number of children. When additional children provide negative net values, having more boys and fewer girls reduces the marginal value of an additional child, so that the couple will choose a lower probability of birth, even if they prefer girls (Leung, 1991). But together with our previous results on divorces, marriages, shotgun marriages and re-marriages, our finding on fertility stopping rules implies a preference for boys.

compare the effects on fertility across countries, it is largest for China and Vietnam, with more moderate effects for Mexico, Columbia, and Kenya. To add to the international evidence, we find similar negative effects from daughters on the probability a couple is in a consensual union versus a formal marriage in Mexico and Columbia. Our final 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 lead some husbands to marry another woman if his first wife delivers a girl.

Documenting parental sex bias is important for several reasons. First, understanding the magnitude of parental sex bias has important policy implications. Rapid technological progress promises to make it increasing possible for couples to choose the sex of their children. Although these techniques are still used by a negligible number of couples, they are already extremely controversial.[5] While they are legal in the United States, many European countries have outlawed them or are considering outlawing them. The notion of using gender selection technologies to have only boys (or only girls) appears unacceptable to many critics.

Documenting the presence of sex bias in the family is also important because 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 outcomes (Bedard and Deschenes, 2003). Furthermore, even in families where parents are married, parents who prefer boys may devote less attention and nurturing to their daughters than sons. They may also devote less financial resources to their education and health. In this sense, our results are related to the existing literature that documents unequal intrahousehold allocation of resources.[6] Moreover, our findings are related to the many studies that document gender differences 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 intrahousehold allocation of psychological and material resources.

The remainder of the paper is organized as follows. In Section 2, we present evidence for divorces. In Section 3, we present a simple theoretical model which highlights the various

[5] One technique that is already being utilized with mixed success makes use of the fact that 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 (ref xxx). In the future, IVF techniques, which use a DNA detector to separate sperm which is then artificially inseminated into the woman, promises a higher success rate (ref xxx).

[6] Using data for the US, Brazil and Ghana, Thomas (1994) finds that the education of the mother has a bigger effect on her daughter's height, and that the education of the father has a bigger impact on his son's height. Behrman, Pollack and Taubam (1986) and Behrman (1988) investigate intra-family allocation of

testable implications of each hypothesis. In Section 4, we present evidence on marriages, shotgun marriages and re-marriages. In Section 5 we present evidence on fertility stopping rules. In Section 6, we present some additional evidence for the US. In Section 7, we present our international evidence. Conclusions are in Section 8.

2 The Effect of Sex Composition of Children on Divorce

We begin our empirical analysis by investigating the relationship between the sex composition of children and the probability that a woman is divorced. We uncover a surprisingly strong effect of children's gender on divorce. In particular, we document that women who have only girls are significantly more likely to be divorced than women who have only boys.

Because children of divorcees are in most cases assigned to the mother, one possible interpretation of this finding is that fathers prefer sons to daughters. After all, it is the father who, in most case, loses day-to-day access to his children in case of divorce, while the mother is likely to stay with the children, irrespective of marital status. 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. We provide a more precise definition of what we mean for gender bias in the next section. There, we also clarify that father gender bias is not the only possible explanation of the documented relationship between the sex composition of children and divorce, and propose a series of tests that can be used to isolate gender bias. We also clarify how mothers' gender preferences may affect divorce.

For the empirical analysis of divorce, we use data from the 1940 to 2000 Censuses. To minimize the probability that some of the children might have left the household, our sample includes all ever-married mothers between the ages of 18 and 50, with children less than 12 years old.[7] The first column in Table 1A reports the estimates from a regression of a dummy equal to one if the woman is currently divorced on gender of her first child, for women with only one child. The coefficient indicates that women with a girl are .0064 percentage points more likely to be divorced than women with a boy.

To assess the magnitude of the estimated marginal effect, throughout the paper we report the "all-boy baseline", which in Table 1A is the fraction of mothers with only boys who are

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resources in the US. The literature for developing countries is too large to be summarized here. Examples include XXX.

[7] We have experimented with alternative cut-off ages for children (8, 10, 16) and found similar results.

divorced (i.e., the intercept term). We also report the percent effect, which is the increase in the dependent variable for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (in row 1) over the all-boy baseline. This percent effect is simply the odds ratio minus one. In column 1, the percent effect indicates that the probability of divorce when moving from a family with a boy to a family with a girl increases by 4.44%. We view this as a surprisingly large effect.

In column 2, we repeat the analysis including families with one child or more. Unlike the model in column 1, the model in column 2 is not affected by endogeneity of family size. We expect to find a lower coefficient than the one in column 1, because we are now including not only mothers with 1 child, but also mothers with 2 and more children. Reporting the effect without conditioning on the number of children provides the total effect, including any effect on fertility, of having a girl versus a boy for the first child. The effect of the gender of the first child is diluted when women in larger families are included. The coefficient indicates that women whose first child is a girl are .0044 percentage points more likely to be divorced than women whose first child is a boy. Relative to the baseline, the percentage effect is 3.94%.

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. Women with two girls are significantly more likely to be divorced than women with two boys. For families with 2 or more children, the difference is .0044 percentage points, or 5.16% of the baseline all-boy divorce rate. When we turn to families with 3 or more children (column 6), we find a marginal effect of .0071 and a percent effect of 8.60%.

The baseline effect decreases as me move from column 2 to column 4 to column 6, since the probability of divorce is lower for larger families, as one would expect. It is interesting to note that both the marginal effect and the percent effect increase as me move from families with 1+ children to families with 2+ or 3+ children. For example, the percent effect almost double when moving from families with at least 1 child (4.4%) to families with 3+ children (8.60%). As we will see below, the same pattern remains true when we look at other outcomes.

What explains this increase? We don't have a definitive answer. But we speculate there may be two opposite forces at play. First, parental preferences may 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 6 than column 2. On the other hand, family size is endogenous, and this endogeneity should lead us to find a larger effect in column 2 than column 6. The reason is that the sample used in column 6 does not include many couples that divorced because they had girls as first or second child. In this sense, the sample in column 6 has already been purged from some of the couples whose marital status depends on children gender mix. The comparison of the magnitude of the effects in column 2, 4 and 6 may indicate that the first effect—differences in tastes across fertility groups---dominates the second effect---endogeneity of family size.[8]

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 the ones in Table 1A, where we condition on a vector of observable mother 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 slightly smaller. 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.[9]

A limitation of the results in Table 1A is that the dependent variable is current divorce status. Because many divorcees re-marry, the effects in Table 1A are likely to *underestimate* the true effect of having girls on divorce. To see how large this bias is, we use the 1980, 1985, 1990 and 1995, the CPS Fertility Supplements, that reports the complete fertility and marital history of respondents. Row 1 in Table 1A shows that, when the dependent variable is a dummy for current divorce status, we can reproduce the main results of Table. As expected, when the dependent variable is a dummy equal 1 if the first marriage ended in divorce, the estimated marginal effects are larger. For example, the probability that the first marriage ended in divorce for mothers with a girl is more than double than the probability that the same mother is currently divorced (column 1).[10] A second benefit of Table 1B is that is shows that the lack of information on non-resident children does not significantly affect our Census results in Table 1A.[11]

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[8] During the preparation of our manuscript, we became aware of an interesting paper by Bedard and Deschenes (2003) that uses 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. Their findings on the effect of children gender and divorce are generally consistent with our findings in Table 1. The question addressed by our paper and most of our empirical analysis are different from the Bedard and Deschenes paper.

[9] 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 many case almost identical) to the simple average for all-boy families.

[10] Because the baseline effects in the bottom panel are 2 to 3 times larger than the ones in the top panel, the percent effect are larger in the top panel.

[11] As we have mentioned above, in the Census we observe only children living in the household. To minimize the probability that the respondents have other children who live on their own, in Table A1 we have included only women with children younger than 12. Yet, it is in theory still possible that some

Is children gender really random? 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 still not very accurate. Importantly, they were completely unavailable for most of the years under considerations in this paper. Furthermore, recent medical literature shows that natural methods based on timing of intercourse have no significant effect (cite 1995 JAMA paper here XX). In this paper, we therefore assume that child sex is random. This assumption is consistent with the one in Angrist and Evans (1998). As further support of the randomness of child sex, we have calculated the fraction of first-born children who are boys in each Census, from 1940 to 2000, and found that it is constant over time. This suggests that improvements in technology in gender selection have not yet received widespread adoption.[12]

3 A Simple Model

In the empirical analysis of this paper, we investigate the effect of child sex composition on divorce, marriages, 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., that they prefer having sons to daughters (or vice-versa).

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 and girls. There is a consensus in the child psychology literature that the presence of the father in the household when kids are growing up is more important for boys than girls. This consensus is based on abundant evidence that fathers play a larger role in the development of their sons than their daughters. Longitudinal data on child development show that the absence of a father has more severe and enduring impact on boys than girls (Heteringhton, Cox and Cox, 1978). For example, boys are found to suffer more and longer from divorces than girls since, in most cases, children are assigned to the mother, irrespective of sex (see Lamb, 1976 and 1987 for a survey of the evidence). We call this the role model hypothesis. According to this hypothesis, parents care about the well-being of their children, have unbiased gender preferences, 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, time, or psychological cost of raising

children have left, and that males leave earlier than females. Table 1B, based on the entire fertility history of respondents, indicates that this is unlikely to be a major problem.

[12] Add means by year in a footnote here XX.

girls is higher than the cost of raising boys (or vice-versa). We call this the differential costs hypothesis.

(A) Intuition. This section develops a simple model to understand the implications that (i) gender bias, (ii) role model, and (iii) differential costs have for divorce, marriage 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. In other words, any of these hypotheses can generate the relationship between gender and divorce uncovered in Section 2. The intuition is straightforward. If parents are biased in favor of boys, the utility loss of moving from the married state to the divorced state is larger for boy families versus girl families. This immediately implies that marriages with a girl are less happy than marriages with a boy, and more likely to result in divorce. If parents are unbiased and are concerned about the role model effect, having a boy reduces the probability of divorce for marginal marriages since the harm done to a boy as a consequence of divorce is greater. If parents are unbiased, don't care about the role model effect and girls are more expensive, marriage versus divorce is relatively more expensive for a father with a girl because alimony payments and other time costs are similar in the divorced state for the father. This assumption seems realistic, because the time spent by divorced fathers with their children is limited, and because courts are unlikely to order child support payments that are vastly larger for girls than boys.

The predictions for marriage are also similar across the three hypotheses. Parents are less likely to marry if they have a daughter versus a son, irrespective of whether the gender bias, role model or differential cost story is true. The intuition is similar to the one for divorce.

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.[13] Again, the intuition is very simple. If couples prefer sons, the *effective* number of children is larger in the 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

[13] We note that in a model where the decision to have a child is stochastic and can be influenced by the parent's precautionary or proactive measures to have a child, costs cannot always be separately identified from the role model story. See Leung (1991).

reason is that parents will take into account the effect of children gender on future divorce. Having a boy has a lower option value, because it may force some couples to stay married for the sake of the boy, even when they would have divorced in his absence. (If parents do not take into account the effect of children gender on future divorce, then the role model hypothesis predicts no effect of gender on fertility.) Finally, if parents are unbiased, don't care about the role model effect and girls have a higher price than boys, then couples whose first child is a girl will be less likely to have a second child than couples whose first child is a boy. The reason is that families with a girl are poorer than families with boys, because girls are more costly.

In the rest of this section, we 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 the three different assumptions.

(B) 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 (ref xx). 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 know 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:[14]

[14] 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.

$$\varepsilon_t \sim N(0,\sigma^2)\,.$$

This shock is marriage-specific, so if the couple separates the shock is not present.[15] We assume the shock is independent of the gender composition and size of the family. 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. To facilitate discussion, we choose to separately write out the value of this shock when writing the utility function.

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

$$(1) \qquad U^M(B_t, G_t, C_t) + X_t^M + \varepsilon_t$$

if they choose to stay married after the realization of the shock. If they divorce their utility is

$$(2) \qquad U^D(B_t, G_t, C_t) + X_t^D\,.$$

The superscripts M and D in these to equations refer to the utility in the married and divorced states, respectively. The period budget constraint for the married state is

$$(3) \qquad p^M B_t + q^M G_t + r\,C_t^M + X_t^M = Y_t^M$$

where p^M, q^M, and r are the prices of boys, girls, and nontransferable consumption and Y_t^M is combined income for the married couple.[16] The numeraire good is transferable utility. A similar budget constraint holds in the divorced state.

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, so that

$$U^M(B_t, G_t, C_t) = U^M(B_t + \gamma^M G_t, C_t)$$

$$U^D(B_t, G_t, C_t) = U^D(B_t + \gamma^D G_t, C_t)$$

where γ^M and γ^D are positive scalars which weight how much parents value girls compared to boys in the married and divorced states. The sums appearing in the utility functions can be thought of as representing the *effective* number of children in the married and divorced states. We write the effective number of children in the married and divorced states as

$$\widetilde{K}_t^M = B_t + \gamma^M G_t$$

$$\widetilde{K}_t^D = B_t + \gamma^D G_t\,.$$

In this setup, boys and girls are perfect substitutes, with potentially different marginal rates of substitution. It follows that there is a quality-quantity tradeoff between the gender and number of

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[15] A shock which occurs in divorced state for each spouse could be added to the model. However, this complication does not change any of the key predictions.

[16] For simplicity, we assume the budget constraint holds with equality in each period.

children.

We assume that utility in the married state increases at a decreasing rate as the number of effective kids increases. In the divorced state, utility decreases at a decreasing rate as the number of effective kids increases. That is, we assume

$$\frac{\partial U^M}{\partial \tilde{K}^M} > 0, \quad \frac{\partial^2 U^M}{\partial (\tilde{K}^M)^2} < 0$$

$$\frac{\partial U^D}{\partial \tilde{K}^D} < 0, \quad \frac{\partial^2 U^D}{\partial (\tilde{K}^D)^2} > 0 .$$

It is now a simple matter to describe what we mean by a gender bias, a role model, and a cost differential. A gender bias for boys occurs when $\gamma^M<1$, meaning that a girl is worth some fraction of a boy in the married state. In the simplest version of gender bias, boys and girls are equally valuable in the divorced state, so that $\gamma^D=1$. This definition says that holding the number of children fixed, the effective number of children in the married state increases with the number of boys. A role model for boys exists if $\gamma^D<1$, which means that boys cause more disutility compared to girls in the divorced state. The reason is that in most cases children are assigned to mothers, irrespective of their gender. The utility function of altruistic parents takes into account the fact that boys have lower utility in the divorced state relative to girls, since the absence of a father has a more severe impact on boys than girls. Finally, differential costs favoring boys occurs if $p^M<q^M$ so that boys are cheaper than girls in the married state. For simplicity we assume that boys and girls cost the same in the divorced state, i.e., $p^D=q^D$.

(C) 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 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 marriage-specific shock. The divorce decision takes into account the number and sex composition of children, as well as the expected future value of remaining married. If the couple chooses to stay together, they then decide whether to have an additional child. The additional child decision takes into account the probability of divorce in the future, i.e., the expected value of the marriage-specific shock in period two.

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. For example, utility in the divorced state for a family with one boy and one girl is written as

$$U_{BG}^{D} = U^{D}(1+\gamma^{D}1, C_t) + X_t$$

where consumption (transferable and nontransferable) have been chosen optimally for given prices, income, and family composition. From this point on, we omit the time subscript on variables when it is clear which time period is the relevant one.

We begin by writing out the option value of staying in the marriage. It is the expected benefit of optimally chosen period *t*+1 variables if the couple chooses to remain married in period *t*, where the expectation is taken in period *t*. For a family with composition *c* in period *t*+1, the expected benefit is:

$$(4) \qquad \max_{\substack{additional \\ kid}} \left\{ \underset{B,G}{E_t} \left(\underset{\varepsilon}{E_t} \left(\max_{M,D} \begin{Bmatrix} U_c^M + \varepsilon_{t+1}, \\ U_c^D \end{Bmatrix} \right) \right) \right\}$$

where family composition *c* can take on the family types *B, G, BB, BG, GB,* and *GG*. The first max is with respect to the fertility decision, i.e., whether to have an additional child in the next period. The first expectation is taken over the lottery that an additional child (if the couple chooses to have one) will be a boy versus a girl. This simply reflects the fact that couples don't know the sex of an additional child in advance. For simplicity, we assume that there is always a 50-50 chance of a boy versus a girl.[17] The second expectation is taken over the marriage-specific shock. The second max is taken with respect to the divorce decision as a function of the sex and number of children in period *t*+1. Note that while the expectations are taken in period *t*, any additional child is realized in period *t*+1.

The idea behind equation (4) is fairly simple. 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 divorcing and staying married in the next

[17] The actual ratio is closer to 49% girls and 51% boys, but using 50% does not alter the general predictions.

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.

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

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

where $F(\cdot)$ is the distribution function of the random variable ε, which by our previous assumption is normally distributed. Let the expected value of the shock conditional on staying in the marriage (i.e., the truncated mean of the shock) be given by

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

and the expected value of the marriage-specific shock conditional on divorce be given by

(7) $$\lambda_c^D = E(\varepsilon_{t+1} \mid \varepsilon_{t+1} < U_c^D - U_c^M)$$

We can now write the expected utility in period *t* for a family with composition *c* in period *t*+1, as:

(8) $$\theta_c = \pi_c(U_c^M - U_c^D + \lambda_c^M) + U_c^D$$

Equation (8) is a function of utility in the married and divorced states, U_c^M and U_c^D, which are in turn functions of the effective number of children in the married and divorced states, $\tilde{K}^M$ and $\tilde{K}^D$.

Equation (8) plays a pivotal role in determining forward-looking divorce and fertility decisions. The appendix shows that θ_c is an increasing, concave function of $\tilde{K}^M$ if

(9) $$\lambda_c^M \left(U_c^M\right)' < -\frac{\left(U_c^M\right)''}{\left(U_c^M\right)'}$$

where the derivatives are taken with respect to $\tilde{K}^M$, holding the number of children fixed. Likewise, θ_c is an decreasing, convex function of $\tilde{K}^D$ if

(10) $$\lambda_c^D \left(U_c^D\right)' > -\frac{\left(U_c^D\right)''}{\left(U_c^D\right)'}$$

where the derivatives are taken with respect to $\tilde{K}^D$, holding the number of children fixed.

(D) Divorce Predictions. We first examine divorce predictions based on child gender in the 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.[18] For example, when considering a gender bias for boys, we assume no role model effect and equal costs.

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

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

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

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

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 (11) and (12) 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 formal proofs can be found in Appendix B and rely on the fact that θ_c is an increasing function of $\widetilde{K}^M$ and a decreasing function of $\widetilde{K}^D$.

The intuition for these results is simple. For example, consider a father with a gender bias for boys and an unbiased mother. There is an asymmetry in the divorced state since custody of the children generally goes to the mother and not to the father. In the extreme case where the father never sees his kids again after a divorce, his utility in the divorced state does not depend on the gender of his children. More generally, one would expect the utility loss of moving from the married state to the divorced state is larger for boy families versus girl families. The benefit of remaining in the marriage for fathers with daughters is smaller, and the marriage is more likely to end in divorce.

The role model story says that parents care about the utility of their children, and in the absence of gender bias they value each child equally. That is, the child's utility enters the parent's utility function. Parents with girls are more likely to divorce since the harm done to a boy as a consequence of divorce is greater. For the differential cost hypothesis, if boys are cheaper, the effect depends on the costs of children in the divorced state. Suppose alimony

[18] Obviously, more than one effect may be present, but the point is to show which effect dominates.

payments and other time costs are similar in the divorced state for the father. This case seems like a realistic approximation, because the time spent by divorced fathers with their children is limited, and because courts are unlikely to order child support payments that are vastly larger for girls than boys. As a consequence, marriage is relatively more costly with a girl for the father and a divorce is more likely to occur.

(E) Fertility Predictions. 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, 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

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

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

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

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 more involved than the divorce decision. Without saying something about the curvature of the utility functions, it is hard to compare the marginal value of an additional child for boy versus girl families. It follows immediately that if θ_c is an increasing, <u>concave</u> function of $\widetilde{K}^M$ and a decreasing, <u>convex</u> function of $\widetilde{K}^D$, the following is true: (i) gender bias predicts higher fertility for girl families, and (ii) role model predicts higher fertility for boy families.[19] If θ_c is an increasing, <u>convex</u> function of $\widetilde{K}^M$ and a decreasing, <u>concave</u> function of $\widetilde{K}^D$, the reverse is true: (i) gender bias predicts higher fertility for boy families, and (ii) role model predicts higher fertility for girl families.

Since the predictions hinge on the convexity or concavity of θ_c, it is important to understand the conditions described in equations (9) and (10). Both equations indicate the

[19] Note that for both the gender bias and role model cases, $|\theta_{GB} - \theta_G| > |\theta_{BB} - \theta_B|$ and $|\theta_{GG} - \theta_G| > |\theta_{BG} - \theta_B|$.

expected future benefit of family composition c, θ_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 (or disutility) of an effective child 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. The left-hand side of equation (9) 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 (9) captures the idea that additional children have potentially rapidly decreasing marginal utility. Since couples prefer sons and because sons and daughters are perfect substitutes, the *effective* number of children is larger in the all-boy family. A boy family has more effective children, 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 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.

The intuition behind equation (10) for the role model hypothesis is similar. The left-hand side of the equation captures the idea that families with a boy are more likely to remain married in period two and therefore have a more negative expected shock, and that the marginal utility of extra children is negative in the divorced state. The right-hand side of equation (10) captures the idea that additional children have potentially rapidly decreasing marginal disutility. If the left-hand side is larger, the boy families are relatively more likely to have an additional child. The intuition behind the result is that the loss in the option value of a divorce is smaller for families with a boy. Since parents do not want their sons to be without a role model, they are more likely to remain in a bad marriage. If families with one girl choose to have another child, they run the risk of having a boy which decreases the option value of a divorce by more than it does for families with one boy. This effect dominates if the utility functions have the right amount of curvature.

If there is sufficient curvature in the married utility function, this indicates a strong preference for the first boy. We refer to strong gender bias as the case when equation (9) holds, and strong role model as the case when equation (10) holds. Under strong gender bias, girl families are more likely to have an additional child and under strong role model, boy families are more likely to have an additional child.

Whether or not equations (9) and (10) hold, differential costs predict no effect on fertility with transferable utility, and in more general models, higher fertility for boy families. The intuition for costs 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.[20] With transferable utility, the income effect on effective children is zero. In more general models, however, couples with boys will be more likely to have an additional child.

(F) Predictions for other Outcomes. In addition to divorce and fertility, in the empirical part of the paper we also look at other outcomes. For the U.S., these outcomes include whether a couple with children born out of wedlock end up marrying, whether couples who find out the sex of their child during pregnancy are more likely to marry before delivery (shotgun marriages), and whether divorced mothers end up remarrying. Predictions for these outcomes are identical to predictions for marriage described above. When we turn to the foreign countries, we also add whether a mother is in a polygamous relationship and whether a couple is in a consensual union versus a marriage.

4 The Effect of Sex Composition on Marriages, Shotgun Marriages, and Remarriages

In the previous section, we have shown what each of our three competing hypothesis---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. 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 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 marriages 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

[20] Note that no price effect arises, because parents gender is revealed *after* parents have made the decision to have another child. Leung (1991) shows that in a more general setting where fertility is not perfectly controlled for by parents, and where consumption and children are perfect substitutes, the opposite result can be true. In Leung's setting, the gender bias and differential cost hypothesis are not separately identifiable.

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.

Finally, we look at the probability of second marriage and second divorce. We find that divorcees with boys are more likely to re-marry and, if they do re-marry, they are less likely to divorce a second time than divorcees with girls.

(A) Marriages. We begin by showing the relationship between children's gender and the probability that a mother has never been married. We use all mothers between 18 and 50 with children younger than 12 in the 1940-2000 Censuses. In Table 2, we show that among women who have exactly one child, those whose first baby is a girl are .0021 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 .0016. The percent effects are large. Having a girl first reduces the probability of marriage in the two samples by 2.1% and 2.5%, respectively.

When we consider women who have at least 2 or 3 children, the estimated effect is even larger. Women with two girls suffer a 5.2% percent decline in the probability of marriage compared with women who have two boys (column 4). Women with three girls suffer a 6.1% percent decline in probability of marriage compared with women who have three boys (column 6). We report estimates conditional on mother characteristics in Appendix Table XX. These conditional estimates are generally very similar to the estimates in Table 2.

(B) Shotgun Marriages. 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 Vital Statistics, for 1989-1994.[21] The first column in Table 3 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

[21] The marital status variable in the Vital Statistics is not ideal. Marital status is imputed on the basis of mothers' last names and babies' last names, 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.

interpretation of our findings in Table 2. When we control for mother characteristics in column 2---including race, education, age Hispanic status, immigrant status, year---the coefficient flips sign, and remains virtually zero.

The most interesting results of the table are in columns 3 and 4. Here we test whether gender of the child matters when the mother has taken an ultrasound test during pregnancy and therefore knows the gender of the baby with high probability in advance. Ultrasound tests are typically able to reveal the sex of the baby by the 16th week of pregnancy. They are considered very accurate (somewhere between 95 and a hundred percent).[22] About 38% of mothers in the sample have taken the ultrasound test during pregnancy.

We regress marital status on a dummy equal one if the child is a girl, a dummy for ultrasound, and the interaction of the female dummy and the ultrasound dummy. The interaction of the female dummy and the ultrasound dummy is negative and statistically significant. The coefficient suggests that women who take the test and have a girl are .0037 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 the probability of ultrasound across mothers.[23] When we also condition on mothers characteristics in column 4, the coefficient is .0030, slightly smaller than the unconditional one, but still significant.[24]

We interpret this finding as evidence that the gender of the child 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. How large is this effect? At the bottom of the table, we report the percent effect, which is the percent change in marital status when moving from a mother with a boy to a mother with a girl. It is the ratio of the coefficient in row 2 over the all boy baseline reported in row 5. The all boy baseline for unconditional models is simply the probability of being married for mothers with a boy who take an ultrasound. That is, the all-boy baseline for models that include covariates is the predicted probability of marriage for a mother who has taken the test and has a boy, using the

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[22] The main reason for taking the test is not disclosure of the child sex, but is diagnostic. The test uses sound waves to view and examine the fetus or and 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. Accuracy of sex prediction increases with stage of pregnancy.

[23] 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 take the test. When we control for mother characteristics in column 4, the coefficient on ultrasound drops to less than half of the unconditional coefficient.

estimated regression coefficients and the explanatory variables evaluated at their means. The percent effect indicates that when moving from a boy family to a girl family, the probability of marriage decline by 0.0045 to 0.0055%.[25]

The California birth certificates also report whether the mother has taken an ultrasound during labor.[26] We use this information to perform a specification check. Unlike for ultrasound during pregnancy, we expect that the interaction of ultrasound during labor and gender not to matter for marital status at birth. There simply is not time between labor and delivery for marital status to be affected. The coefficient on the interaction in column 5 is -.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. In Appendix Table A2, we investigate whether the effect of child gender on marital status for mothers who take ultrasounds varies across racial or educational groups, or by cohort of birth of the mother.

(C) Second Marriages and Second Divorces. 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 is a priori ambiguous, even if it was true that men have gender bias for their natural children. 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 re-marriage rates than divorced women with girls. On the other hand, it is also possible that men have preferences for the gender of their natural children, but they have no strong preferences for the gender of step-children. In this case we should observe no relationship between women re-marriage rates and children sex. It is even possible that having boys reduces the probability of re-marriage, if potential husbands are more concerned about the potential of conflicts with male step-children than female step-children.

Empirically, we find evidence that divorced women who have only boys are more likely to re-marry than divorced women who have only girls. Table 4 indicates that the difference in the remarriage probability between women with only females and women with only males is

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[24] As one would expect, standard errors in conditional models are slightly lower than standard errors in unconditional models

[25] 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.

[26] Ultrasound during labor and birth, also called electronic fetal monitoring, is used to record the baby's heart rate and sometimes the mother's contractions. Only 2.2% of mothers in our sample take ultrasound during labor and birth.

statistically significant. The percent effects indicate that moving from having all boys to having all girls, reduces the probability of re-marriage by 0.97%, 1.78% and 3.6% for families with at least 1, 2 and 3 children, respectively.

Table 5 shows that among women whose first marriage ended in divorce and who have re-married, 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 4.6%, 7.4% and 8.2% for families with at least 1, 2 and 3 children, respectively.

5 The Effect of Sex Composition on the Probability of Having Another Child

Taken together, the evidence on marriages, shot-gun marriages, re-marriages and second divorces presented in the previous section is suggestive of gender bias on the part of parents. However, it is important to note that, like for divorce, these four results alone can not unambiguously rule out the role model hypothesis and 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 boys because the psychological or monetary costs of raising boys are lower than the costs of raising girls. Results for remarriages are a slightly harder to interpret as evidence of role model hypothesis, because potential husband are less likely than natural fathers to feel the need to be a role model to boys who are not their natural sons. Yet, it is in theory still possible that potential husbands decide to marry a divorcee with boys, not because they prefer having male step-children, but because they want to provide a role model for them.

In this section we turn to another test of the role model hypothesis based on fertility decisions. As the model in Section 3 makes clear, if parents are 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 parental preferences 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 or equal* for families that have all girls compared to families that have all boys.

We begin by presenting results based on a sample of all married women with children younger than 12 from the 1940-2000 Censuses. Column 1 in Table 6 suggests that among all 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 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 include 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 children gender and fertility in table 6 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 whether a woman has already been married. 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.

Estimates in column 2 are consistent with the gender bias hypothesis. They suggest that among all families with 2 or more children, the probability of having another child is .0085 percentage points *higher* when the first two children are girls than when the first two children are boys. This is 2.1% of 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 2.1%. Estimates in column 3, based on all families with at least 3 children, are even larger.[27] In Appendix Table A3, we show that estimates conditional on mother characteristics are slightly lower, but generally consistent with unconditional estimates in Table 6.

Another feature of Table 6 is that having children with mixed gender has a negative effect on fertility (see also Angrist and Evans, 1998). In column 2, the effect of having only girls (relative to only boys) is smaller in absolute value than the effect of mixed gender (relative to only boys). For families with 2 or more children, the former is only 18-19% 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

[27] These results are consistent with findings in Ben-Porath and Welch (1980) for the US in 1970, and with a large literature on the effect of child sex on outcomes in developing countries.

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, BGG) are all negative, but much smaller.

Using the Census data, we can also determine if spacing between children depends on the sex composition of the family. In the Census, age of children is measured in years, and this measure can be used to estimate average time intervals between births. For couples with exactly 2 children, those with a girl as the first child have a second child an average of 6.2 days sooner. This effect is precisely estimated, with a t-statistic of 5.1. We interpret this as additional evidence that couples are anxious to have a male child due to gender bias.

One limitation of the Census data used in Table 6 is that we observe family size at the time of the Census, and can not distinguish between completed and uncompleted fertility. Because children sex is random, this is not in itself a source of bias. However, because some mothers have not finished bearing children, part of the effect of gender on fertility does not manifest itself in Table 6. Estimates in Table 6 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 effect of children gender composition on fertility decision when fertility is completed.

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 Vital Statistics dataset that provides mother name and date of birth.[28] 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 is that most mothers in the longitudinal sample have completed fertility. This is shown in Appendix Table A4, which reports, for each year between 1989 and 2001, the fraction of mothers whose last observed delivery occurs in the specified year. Only 4.3% of mothers are still having babies in 2000 or 2001, suggesting that women concentrate their childbearing in the first few years, and by 2001, most mothers in the sample have stopped having children.

Because the Vital Statistics sample includes women with completed fertility, we expect to find larger effects in this sample than in the Census. Table 7 shows estimates of the effect of gender composition of the probability of having another child, based on the longitudinal

[28] We matched mothers at first delivery with mothers at second delivery using the mother name, mother date of birth (day, month and year), and first and second child date of birth (day, 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.

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% and 5.5%, respectively.

As expected, the estimates of the all-girl effect from the longitudinal sample in Table 7 appear qualitatively consistent with but quantitatively larger than the corresponding estimates from the Census in Table 6.[29] 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 6. 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 significantly affect our Census results.

Overall, we interpret our fertility results as generally supportive of the gender bias hypothesis. In both datasets, families where the first two children are both girls have more children than families where the first two children are both 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 of the estimated effects appears to be non-negligible. Depending on the sample, an all-girl family is 2.1 to 5.5% more likely to have another child. As argued in the theoretical model, it would be difficult to explain these findings if there was no parental gender preference and only the role model hypothesis was true.[30]

There is still the question why gender of the first child does not affect fertility, while gender of the first two and three children does affect fertility. We don't have a definitive explanation for this difference. We speculate that it may be due to differences in gender bias for parents with high and low fertility. Consider for example the case where most parents who have strong gender bias want to have at least 2 children--irrespective of children sex---and possibly more, if they are unhappy with the first two draws. In this case, we would see *no effect* of gender of the first child on fertility decisions. But we would see an effect of the gender of the first two

[29] We find only a weak relationship between 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) appear to be slightly more negative than the coefficients for families with 2 girls (BGG, GBG, GGB), but this pattern is quite weak.

[30] The negative effect of having a girl uncovered in column 1 of Table 6 is potentially problematic. However, we have shown that when we include only women in their first marriage, the effect of having a girl drops to zero. Furthermore, the effect is virtually zero in the Vital Statistics sample.

children on the probability of having a third child. Off course this is just a hypothesis, and cannot be empirically tested.

Finally, we want to make clear that failure to reject the gender bias hypothesis does not necessarily imply that the role model hypothesis or the differential cost hypothesis are false. The theoretical model in Section 3 indicates that the effect of children's sex on fertility under the role model hypothesis should be opposite the effect under gender bias. 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 role model on fertility decisions.

6 Additional Evidence for the US: 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 families headed by a woman with children 12 year or younger in the 1995-2000 March CPS. Table 8 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 not statistically significant. The effect among mothers with one or more children is virtually zero. All models control for mother age, mother race, and year.[31] Because we rely on a sample that is substantially smaller than the Censuses and the Vital Statistics, our results are necessarily less precise. For this reason, evidence in Table 8 section 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 NLSY sample appear to be too small to draw firm conclusions.[32]

7 Evidence From Developing Countries

[31] Unconditional models have larger standard errors. None of the relevant coefficients is significant for unconditional models.

[32] 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

We now turn to evidence of gender bias in five developing countries. We were able to obtain large scale samples for China (1982), Vietnam (1989 and 1999), Mexico (1990 and 2000), Colombia (1973, 1985 and 1993), and Kenya (1989 and 1999).[33] To the extent possible, variable definitions were made consistent with definitions in the US Census by IPUMS researchers at the University of Minnesota.

For these countries, we show four pieces of evidence. First, we re-estimate our models for the effect of offspring gender on probability of divorce. Second, we re-estimate our models for the effect on fertility decisions. We can then compare our estimates of the divorce and fertility effects based on US data with estimates based on data from these five developing countries. We find that the divorce and fertility effects are generally larger for developing countries than for the US, with China and Vietnam showing by far the largest percent effects.

We then turn to consensual unions in Mexico and Colombia, which accounts for over 20 and 30 percent, respectively, of all couples with children in the countries. We find that couples are more likely to be in an informal relationship (i.e., a consensual union) compared to a legally recognized marriage if they have girls. Finally, we examine the relationship between polygamy and sex composition. In Kenya polygamy is not uncommon, accounting for 12 percent of all marriages. Surprisingly, mothers with girls are more likely to be in polygamous families than mothers who have boys.

(A) Divorce. We begin by looking at the effect of children gender mix on divorce or separation. The samples and the models are identical to the ones used for the U.S. in Table 1. Because there are virtually no divorcees in the China sample (fewer than .01 percent of women with children report being divorced), we exclude China from this analysis. Table 9 shows that for the 4 countries for which data is available, child gender mix significantly affects the probability of divorce. Column 2 is the easiest to interpret, because it provides a summary measure of having a girl as a first child on fertility, irrespective of the (potentially endogenous) family size. Among women with one or more children, women whose first child is a girl have significantly higher probabilities of divorce in all four countries. The marginal effects for Mexico, Columbia, Kenya, Vietnam are, respectively, .0029, .0040, .0027, .0020.

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that result, we discovered that it is crucially 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.

[33] Specifically, we use the 1990 (1 percent sample) and 2000 (10.6 percent sample) Mexican Censuses; the 1973, 1985, and 1993 Columbian Censuses (all 10 percent samples); the 1989 and 1999 Kenyan Censuses (both 5 percent samples); the 1989 (5 percent sample) and 1999 (3 percent sample) Vietnam Censuses; and the 1982 (0.1 percent sample) Chinese Census.

Because the average divorce rate varies widely across countries, percentage effects at the bottom of the table are easier to compare. In Figure 1, graph these percentage effects, including also the corresponding effect for the US from Table 1. The percentage effect is largest in Vietnam and Kenya, and smallest in the U.S. The effects in Vietnam and Kenya are substantial, with percent effects that are two to three times as large as in the U.S.

(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 identical to the ones used for the U.S. in Table 6. 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 were collected. Because the policy was not retroactive, it should have a negligible effect on our sample.

Estimates have patterns that are qualitatively similar to ones uncovered for the US. In families with one or more children, we find insignificant effects of gender in Mexico and Colombia, and significant effects in Kenya, 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 higher probabilities to have another child than all-boys families.[34] For 2+ families, the marginal effects for Mexico, Columbia, Kenya, Vietnam and China are, respectively, .0246, .0048, .0199, .0874, .1976.

In interpreting these estimates, two points are relevant. First, fertility in developing countries is generally high. For this reason, we expect the percent effect for families with 1 child to be smaller than the effect of families with two children. (In the extreme, if every family plans on having at least two children, we should see no effect of gender for families with one child. In this case, families with one child would be families that at the time of the Census still have uncompleted fertility.) Empirically, we find--as expected--that the effect in column 2 is much larger than the effect in column 1 for all countries. Second, as we document for the U.S. in Section 4, our fertility estimates are biased toward finding a negative relationship between all-girl dummy and fertility, at least for Mexico, Colombia, Kenya, and Vietnam. For these four countries having girls results in higher divorce rates (Table 9), and therefore in lower fertility.[35]

Because fertility rates vary across countries, in Figure 2 we compare percentage effects. The effects in Columbia and Kenya are fairly similar compared to the U.S., with effects that are

[34] The literature on children sex composition and fertility in developing countries is quite large and can not be summarized here. Results for some countries have already been documented by other authors. See for example Xie (1989) for China.
[35] We speculate that the negative (although not statistically significant) coefficient for Colombia could be due to this bias induced by divorce.

over twice as large for Mexico. The effect is dramatic in Vietnam, where the percent increase for being 18 percent for families with either 2 or 3 girls. 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 50 percent more likely to have a third child compared to families with two boys. For families with three girls, the effect is an astonishing 90 percent increase.

(C) 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. These unions where the couple are living together without officially being married are reported 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. In either case, it is fair to consider consensual union less than marriage. In Table 11 we document that, among women who are either married or in consensual union, women with all-girls are more likely to be in consensual unions that women with all-boys, although the effects not statistically different from zero in Columbia. 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.

(D) 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 polygamous relationship. But it is possible that, for men who are close to indifferent between having one or more wives, the child 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.[36]

In Table 12 we find that, among all married women with children 12 or younger, women who have girls are more likely to be in polygamous marriages than women who have boys. (Where we link women to their *natural* children and not all the children in the household.) For

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

families with one or more children the difference in probability is .0034 percentage points in models that condition on mother characteristics, or a 2.8 percent effect. For families with 2 or more children, the difference in probability twice as big, with a marginal effect of 5.7 percent compared to the baseline probability for all-boy families.

We interpret this as evidence that the desire for boys lead some of the 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 11 should be considered suggestive rather than definitive.

Overall, estimates from five developing countries provide a consistent picture. Relative to all-boy mothers, all-girl mothers are more likely to divorce, to have another child, to be in a consensual union, and to be in a polygamous relationship. The interpretation of these findings, however, is more difficult than the interpretation of the corresponding findings for the US. Unlike in the US, 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 expected to provide parents with economic assistance when the parents stop working. Because in general males have higher economic returns 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.

8 Conclusion

In this paper we show that US parents 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 affects marital status and fertility decisions of a significant portion of U.S. families. The demand for sons seems to be slowly decreasing over time, but is still significant for younger generations.

Several pieces of evidence are consistent with the presence of a demand for sons. First, women with girls are significantly more likely to be divorced than women with boys. The effect accounts for a non-trivial fraction of divorces. 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 probability of having another child is significantly higher for all-girl families compared to all-boy families. Fifth, child gender seems to affect the probability of alimony payments and the probability of second marriages and second divorces.

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 psychological 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 strong bias for boys.

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 US, with Mexico, Columbia and Kenya in between.

Appendix A

Taking the derivatives of θ_c with respect to $\widetilde{K}_t^M$ and $\widetilde{K}_t^D$ while holding the number of children fixed yields:

$$\frac{\partial\theta_c}{\partial\widetilde{K}^M} = F(U_c^M - U_c^D)\frac{\partial U_c^M}{\partial\widetilde{K}^M}$$

$$\frac{\partial^2\theta_c}{\partial(\widetilde{K}^M)^2} = f(U_c^M - U_c^D)\left(\frac{\partial U_c^M}{\partial\widetilde{K}^M}\right)^2 + F(U_c^M - U_c^D)\frac{\partial^2 U_c^M}{\partial(\widetilde{K}^M)^2}$$

and

$$\frac{\partial\theta_c}{\partial\widetilde{K}^D} = (1 - F(U_c^M - U_c^D))\frac{\partial U_c^D}{\partial\widetilde{K}^D}.$$

$$\frac{\partial^2\theta_c}{\partial(\widetilde{K}^M)^2} = f(U_c^M - U_c^D)\left(\frac{\partial U_c^D}{\partial\widetilde{K}^D}\right)^2 + (1 - F(U_c^M - U_c^D))\frac{\partial^2 U_c^D}{\partial(\widetilde{K}^D)^2}$$

These expression follow using properties of normal density functions, normal distribution functions, and truncated normal distributions. The normal distribution is not vital to the results, however. The result is easier to interpret in the normal case, but the general convexity or concavity results can also be written based on other distributions.

Since $F(\cdot)$ is a number between 0 and 1 and utility is increasing in the number of effective kids in the married state, the first derivative is positive with respect to $\widetilde{K}_t^M$. Likewise, since utility is decreasing in the number of effective kids in the divorced state, the first derivative is negative with respect to $\widetilde{K}_t^D$. The second derivative is negative with respect to $\widetilde{K}_t^M$ and positive with respect to $\widetilde{K}_t^D$ if

$$\frac{f(U_c^M - U_c^D)}{F(U_c^M - U_c^D)}\left(U_c^M\right)' < -\frac{\left(U_c^M\right)''}{\left(U_c^M\right)'}$$

and

$$\frac{f(U_c^M - U_c^D)}{1 - F(U_c^M - U_c^D)}\left(U_c^D\right)' > -\frac{\left(U_c^D\right)''}{\left(U_c^D\right)'}.$$

Appendix B

Begin with the gender bias hypothesis, which says that $\gamma^M > 1$ (and for simplicity, assume $\gamma^D = 1$). Consider four possible cases.

Case (i): Both B and G families optimally choose not to have another child. It is easy to verify that $\Pr(D \mid B) < \Pr(D \mid G)$ since U_k^M and θ_k are increasing functions of $\tilde{K}_t^M$ (i.e., $U_B^M > U_G^M$ and $\theta_B > \theta_G$) and since $U_B^D > U_G^D$.[37]

Case (ii): Both B and G families optimally choose to have another child. It is easy to verify that $\Pr(D \mid BB) < \Pr(D \mid GG)$, since U_k^M, and θ_k are increasing functions of the effective number of children (i.e., $U_B^M > U_G^M$, $\theta_{BB} > \theta_{GB}$, and $\theta_{BG} > \theta_{GG}$).

Case (iii): The B family chooses not to have another child, and the G family chooses to have another child. By revealed preference, $\Pr(D \mid B) < \Pr(D \mid BB)$. But in case (ii), we showed $\Pr(D \mid BB) < \Pr(D \mid GG)$.

Case (iv): The B family chooses to have another child, and the G family chooses not to have another child. By revealed preference, $\Pr(D \mid BB) < \Pr(D \mid B)$. But in case (i), we showed $\Pr(D \mid B) < \Pr(D \mid G)$.

Similar reasoning can be used to prove the stated results for the gender bias and differential costs hypotheses.

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[37] Note that we do not actually require that $U_B^D = U_G^D$, but only that there is no role model effect, i.e., $U_B^D \geq U_G^D$.

Table 1A. The Effect of Child Gender on the Probability of Divorce or Separation, U.S. Census Data.

	Marginal Effect on the Probability of 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)
Girl	.0064 (.0006)	.0044 (.0003)	Girl, Girl	.0052 (.0007)	.0047 (.0006)	G, G, G	.0076 (.0015)	.0071 (.0012)
			Boy, Girl	-.0018 (.0007)	.0008 (.0005)	B, B, G	.0008 (.0014)	.0016 (.0012)
			Girl, Boy	-.0004 (.0007)	.0018 (.0005)	B, G, B	.0076 (.0015)	.0055 (.0012)
						G, B, B	.0080 (.0015)	.0067 (.0012)
						B, G, G	.0040 (.0015)	.0044 (.0013)
						G, B, G	.0053 (.0015)	.0043 (.0013)
						G, G, B	.0026 (.0014)	.0029 (.0012)
All-Boy Baseline	.1440	.1118		.0957	.0910		.0811	.0826
Percent Effect	4.44%	3.94%		5.43%	5.16%		9.37%	8.60%
Obs.	1,400,796	3,595,710		1,409,491	2,194,914		550,965	785,423

Notes: Standard errors in parentheses. The dependent is a dummy equal 1 if the respondent is currently divorced. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all ever-married mothers between the ages of 18 and 50 with children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who are divorced, i.e., the intercept term. Percent effect is the increase in the probability of divorce for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline.

Table 1B. The Effect of Child Gender on the Probability of Divorce or Separation, CPS Fertility Supplement.

	Families with ≥ 1 children (1)		Families with ≥ 2 children (2)		Families with ≥ 3 children (3)
		Currently Divorced			
Girl	.0047	Girl, Girl	.0078	G, G, G	.0138
	(.0029)		(.0047)		(.0088)
All-Boy Baseline	.135		.118		.115
Percent Effect	3.48%		6.61%		12.0%
		First Marriage Ended in Divorce			
Girl	.0108	Girl, Girl	.0106	G, G, G	.0098
	(.0030)		(.0050)		(.0095)
All-Boy Baseline	.341		.314		.335
Percent Effect	3.16%		3.37%		2.92%
Obs.	96859		76357		18901

Notes: Standard errors in parentheses. The dependent is the top panel is a dummy equal 1 if the respondent is currently divorced. The dependent is the bottom panel is a dummy equal 1 if the respondent's first marriage ended in divorce. The excluded category is all boys. Models in column 2 include a dummy equal 1 if the gender of the first two children in Girl Boy and a dummy equal 1 if the gender of the first two children in Boy Girl. Models in column 3 include dummies for B, B, G; B, G, B; G, B, B; B, G, G; G, B, G; G, G, B. Data are from the 1980, 1985, 1990 and 1995 CPS Fertility Supplements. The sample includes all ever-married mothers aged 20-70. All-boy baseline is the fraction of mothers in all-boy families who are divorced, i.e., the intercept term. Percent effect is the increase in the probability of divorce for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline.

Table 2. The Effect of Child Gender on 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	.0021	.0016	Girl, Girl	.0026	.0020	G, G, G	.0026	.0022
	(.0005)	(.0002)		(.0005)	(.0004)		(.0008)	(.0010)
			Boy, Girl	-.0013	.0004	B, B, G	.0007	-.0006
				(.0004)	(.0004)		(.0008)	(.0009)
			Girl, Boy	-.0015	.0004	B, G, B	.0032	.0024
				(.0004)	(.0004)		(.0008)	(.0010)
						G, B, B	.0034	.0031
							(.0008)	(.0010)
						B, G, G	.0041	.0039
							(.0008)	(.0010)
						G, B, G	.0048	.0042
							(.0008)	(.0010)
						G, G, B	.0002	-.0008
							(.0008)	(.0010)
All-Boy Baseline	.0966	.0618		.0391	.0381		.0361	.0356
Percent Effect	2.17%	2.59%		6.65%	5.25%		7.20%	6.18%
Obs.	1,552,354	3,835,753		1,466,650	2,283,399		816,749	572,265

Notes: Standard errors in parentheses. The dependent variable in columns 1 and 2 is a dummy equal to 1 if the family has 2 or more children; in columns 3 and 4 a dummy equal to 1 if the family has 3 or more children; in columns 5 and 6 a dummy equal to 1 if the family has 4 or more children. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all mothers between the ages of 18 and 50 with family composition calculated using all children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who are never married, i.e. the intercept term. Percent effect is the increase in the probability of never being married for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline.

Table 3. The Effect of Child Gender on the Probability of a Shotgun Marriage, California Vital Statistics Data.

	(1)	(2)	Ultrasound during Pregnancy (3)	Ultrasound during Pregnancy (4)	Ultrasound during Labor (5)	Ultrasound during Labor (6)
Female	-.0003	.0007	.0010	.0019	-.0004	.0007
	(.0008)	(.0007)	(.0010)	(.0009)	(.0008)	(.0007)
Female*Ultrasound			-.0037	-.0030	-.0024	-.0007
			(.0016)	(.0014)	(.0055)	(.0049)
Ultrasound			.0657	.0303	-.0003	-.0057
			(.0011)	(.0010)	(.0038)	(.0034)
Controls?	No	Yes	No	Yes	No	Yes
Baseline for Mothers with a Boy			.666	.642	.628	.618
Percent Effect			-.0055%	-.0045%	-.0038%	-.001%
R-squared	0.00	0.22	0.01	0.23	0.00	0.22
Obs.	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601	1,403,601

Notes: Standard errors in parenthesis. The dependent variable is a dummy equal 1 if the mother is married at delivery. Models in Columns 2,4 and 6 control for mother's race (three groups), mother's education, mother's age, mother's immigrant status, mother's Hispanic status, year. In columns 3 and 5, the baseline is the probability that a woman with a boy who took the ultrasound is married. In columns 4 and 6, the baseline is the predicted probability that a woman with a boy who took an ultrasound is married, using the estimated regression coefficients and the explanatory variables evaluated at their mean. Percent effect is the percent change in the dependent variable when moving from a boy family to a girl family. It is the ratio of the coefficient on the interaction Female*Ultrasound (in row 2) over the baseline. The data are from California Vital Statistics. The sample includes all first time mothers in California, 1989-1994.

Table 4. The Effect of Child Gender on the Probability of Remarriage, U.S. Census Data.

Marginal Effect on the Probability of Remarriage								
Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	-.0056	-.0059	Girl, Girl	-.0076	-.0113	G, G, G	-.0219	-.0237
	(.0021)	(.0013)		(.0032)	(.0023)		(.0060)	(.0047)
			Boy, Girl	-.0034	-.0029	B, B, G	.0020	.0013
				(.0031)	(.0023)		(.0060)	(.0047)
			Girl, Boy	-.0020	-.0031	B, G, B	-.0029	-.0003
				(.0031)	(.0023)		(.0062)	(.0048)
						G, B, B	-.0030	-.0018
							(.0061)	(.0047)
						B, G, G	-.0016	.0031
							(.0062)	(.0048)
						G, B, G	.0068	.0002
							(.0063)	(.0048)
						G, G, B	-.0106	-.0070
							(.0060)	(.0047)
All-Boy Baseline	.5647	.6056		.6124	.6337		.6597	.6584
Percent Effect	-0.99%	-0.97%		-1.24%	-1.78%		-3.32%	-3.60%
Obs.	228,743	577,125		193,493	348,382		93,796	154,889

Notes: Standard errors in parentheses. The dependent variable is a dummy equal 1 if the respondent is re-married. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all mothers between the ages of 18 and 50 whose first marriage ended in divorce with children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who are divorced, i.e. the intercept term. Percent effect is the increase in the probability of never being married for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline.

Table 5. The Effect of Child Gender on the Probability of a Second Divorce or Separation, U.S. Census Data.

	Marginal Effect on the Probability of a Second Divorce or Separation							
Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0062	.0046	Girl, Girl	.0067	.0062	G, G, G	.0038	.0065
	(.0013)	(.0008)		(.0018)	(.0013)		(.0035)	(.0027)
			Boy, Girl	.0002	.0015	B, B, G	-.0027	.0006
				(.0018)	(.0013)		(.0035)	(.0027)
			Girl, Boy	.0012	.0021	B, G, B	-.0002	.0039
				(.0018)	(.0013)		(.0036)	(.0028)
						G, B, B	-.0012	.0036
							(.0036)	(.0027)
						B, G, G	-.0028	.0022
							(.0036)	(.0028)
						G, B, G	-.0001	.0025
							(.0037)	(.0028)
						G, G, B	-.0012	.0057
							(.0035)	(.0027)
All-Boy Baseline	.0970	.0893		.0874	.0835		.0843	.0786
Percent Effect	6.39%	4.68%		7.67%	7.43%		4.51%	8.27%
Obs.	228,743	577,125		193,493	348,382		93,796	154,889

Notes: Standard errors in parentheses. The dependent variable is a dummy equal to 1 if the respondent is at her second divorce or separation. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all mothers between the ages of 18 and 50 whose first marriage ended in divorce and who remarried. Family composition calculated using all children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who are divorced, i.e. the intercept term. Percent effect is the increase in the probability of never being married for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline.

Table 6. The Effect of Child Gender on the Probability of an Additional Child, 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	
Sex of 1st child	At least 1 more child (2+) (1)	Sex and order of 1st two children	At least 1 more child (3+) (2)	Sex and order of 1st three children	At least 1 more child (4+) (3)
Girl	-.0016	Girl, Girl	.0085	Girl, Girl, Girl	.0093
	(.0005)		(.0010)		(.0021)
		Boy, Girl	-.0455	Boy, Boy, Girl	-.0346
			(.0010)		(.0021)
		Girl, Boy	-.0441	Boy, Girl, Boy	-.0092
			(.0010)		(.0021)
				Girl, Boy, Boy	-.0085
					(.0021)
				Boy, Girl, Girl	-.0037
					(.0021)
				Girl, Boy, Girl	-.0003
					(.0022)
				Girl, Girl, Boy	-.0210
					(.0021)
All-Boy Baseline	.6443		.3874		.3107
Percent Effect	-0.25%		2.19%		2.99%
Observations	3,152,194		1,974,664		710,807

Notes: Standard errors in parentheses. The dependent variable in column 1 is a dummy equal to 1 if the family has 2 or more children; in column 2 a dummy equal to 1 if the family has 3 or more children; in column 3 a dummy equal to 1 if the family has 4 or more children. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all currently married mothers between the ages of 18 and 50 children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who have an additional child, i.e. the intercept term. For example, in column 1, all-boy baseline is the fraction of mothers whose first child is a boy who have at least two children. Percent effect is the increase in the probability of another child for the mother of 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.

Table 7. The Effect of Child Gender on the Probability of an Additional Child, California Vital Statistics Longitudinal Sample.

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 and order of 1st two children	At least 1 more child (3+) (2)	Sex and 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	-0.08%		4.08%		5.52%
Obs.	406,412		231,811		83,025

Notes: Standard errors in parenthesis. The dependent variable in column 1 is a dummy equal 1 if the family has 2 or more children; in column 2 is a dummy equal 1 if the family has 3 or more children; in column 3 is a dummy equal 1 if the family has 4 or more children. The excluded category is all boys. All-boy baseline is the fraction of mothers in all-boy families who have an additional child, i.e. the intercept term. For example, in column 1, all-boy baseline is the fraction of mothers whose first child is a boy who have at least two children. Percent effect is the percent change in the dependent variable when moving from an all-boy family to an all-girl family. It is the ratio of the coefficient on all-girl (in row 1) over the all-boy baseline. The sample is a panel of California mothers obtained by linking longitudinally birth certificates for years 1989-2001. In particular, the sample includes all married mothers who are first time mothers in 1980 or 1990, and follows them between 1989 and 2001. Appendix Table A1 shows that by 2001, most mothers in the sample are likely to have completed fertility.

Table 8. The Effect of Child Gender on the Probability of Receiving Child Support, Current Population Survey Data.

	Mothers with 1+ Children (1)	Mothers with 2+ Children (2)	Mothers with 3+ 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 parenthesis. The dependent variable is a dummy equal 1 if the family receives child support. The excluded category is all boys. The 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 mean. Percent effect is the percent change in the dependent variable when moving from an all-boy family to an all-girl family. It is the ratio of the coefficient on girls over the all-boy baseline. Sample includes all families headed by a woman with children 12 year or younger in the 1995-2000 March CPS. All models control for mother age, mother race, year.

Table 9. The Effect of Child Gender on the Probability of Divorce or Separation, International Evidence.

Mexico

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0082	.0029	Girl, Girl	.0054	.0036	G, G, G	.0023	.0022
	(.0027)	(.0005)		(.0011)	(.0007)		(.0017)	(.0013)
			Boy, Girl	.0004	.0024	B, B, G	-.0010	-.0005
				(.0010)	(.0007)		(.0017)	(.0013)
			Girl, Boy	.0025	.0012	B, G, B	.0024	.0021
				(.0010)	(.0007)		(.0017)	(.0013)
						G, B, B	.0015	.0016
							(.0017)	(.0013)
						B, G, G	.0008	.0005
							(.0017)	(.0013)
						G, B, G	.0019	.0016
							(.0017)	(.0013)
						G, G, B	.0021	.0016
							(.0017)	(.0013)
All-Boy Baseline	.1034	.0422		.0413	.0358		.0321	.0299
Percent Effect	7.93%	6.87%		13.08%	10.06%		7.17%	7.36%
Obs.	53,465	626,012		306,943	572,547		169,729	265,604

Columbia

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0083	.0040	Girl, Girl	.0036	.0035	G, G, G	.0044	.0036
	(.0029)	(.0007)		(.0015)	(.0009)		(.0024)	(.0016)
			Boy, Girl	-.0028	-.0002	B, B, G	-.0047	-.0022
				(.0015)	(.0009)		(.0023)	(.0016)
			Girl, Boy	.0012	.0023	B, G, B	-.0006	.0009
				(.0015)	(.0009)		(.0024)	(.0017)
						G, B, B	.0013	.0013
							(.0024)	(.0017)
						B, G, G	.0003	-.0001
							(.0024)	(.0017)
						G, B, G	.0015	.0017
							(.0024)	(.0017)
						G, G, B	.0018	.0014
							(.0024)	(.0017)
All-Boy Baseline	.1201	.0515		.0519	.0423		.0397	.0345
Percent Effect	6.91%	7.77%		6.94%	8.27%		11.08%	10.43%
Obs.	51,398	423,955		182,885	372,557		104,906	189,672

Table 9, continued.

Kenya

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0099	.0027	Girl, Girl	.0042	.0030	G, G, G	.0037	.0025
	(.0034)	(.0009)		(.0022)	(.0012)		(.0032)	(.0020)
			Boy, Girl	.0008	-.0001	B, B, G	-.0048	-.0034
				(.0021)	(.0012)		(.0032)	(.0020)
			Girl, Boy	.0005	.0004	B, G, B	-.0043	-.0021
				(.0021)	(.0012)		(.0032)	(.0021)
						G, B, B	-.0020	-.0017
							(.0031)	(.0020)
						B, G, G	-.0034	-.0028
							(.0032)	(.0021)
						G, B, G	-.0004	-.0007
							(.0032)	(.0021)
						G, G, B	-.0028	-.0006
							(.0032)	(.0020)
All-Boy Baseline	.0506	.0260		.0274	.0224		.0255	.0206
Percent Effect	19.57%	10.38%		15.33%	13.39%		14.51%	12.14%
Obs.	17,883	138,223		47,649	120,340		35,734	72,691

Vietnam

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0193	.0020	Girl, Girl	.0063	.0021	G, G, G	.0027	.0011
	(.0061)	(.0006)		(.0013)	(.0007)		(.0016)	(.0011)
			Boy, Girl	.0007	-.0001	B, B, G	.0012	.0002
				(.0013)	(.0007)		(.0016)	(.0011)
			Girl, Boy	.0026	.0007	B, G, B	-.0007	-.0012
				(.0013)	(.0007)		(.0016)	(.0011)
						G, B, B	-.0015	-.0011
							(.0016)	(.0011)
						B, G, G	.0002	.0001
							(.0016)	(.0011)
						G, B, G	-.0014	-.0010
							(.0016)	(.0011)
						G, G, B	.0005	-.0003
							(.0015)	(.0011)
All-Boy Baseline	.0773	.0149		.0161	.0117		.0083	.0069
Percent Effect	24.97%	13.42%		39.13%	17.95%		32.53%	15.94%
Obs.	8,562	182,514		85,404	173,952		52,669	88,548

Notes: See notes to Table 1 for sample restrictions and definitions. Data sources: 1990 and 2000 Mexican Censuses; 1973, 1985, and 1993 Columbian Censuses; 1989 and 1999 Kenyan Censuses; 1989 and 1999 Vietnam Censuses.

Table 10. The Effect of Child Gender on the Probability of an Additional Child, International Evidence.

Mexico

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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	.0015 (.0008)	Girl, Girl	.0246 (.0022)	Girl, Girl, Girl	.0298 (.0042)
		Boy, Girl	-.0253 (.0021)	Boy, Boy, Girl	-.0112 (.0042)
		Girl, Boy	-.0204 (.0021)	Boy, Girl, Boy	-.0097 (.0043)
				Girl, Boy, Boy	-.0137 (.0042)
				Boy, Girl, Girl	.0045 (.0043)
				Girl, Boy, Girl	.0023 (.0043)
				Girl, Girl, Boy	-.0053 (.0042)
All-Boy Baseline	.9292		.4712		.3567
Percent Effect	0.16%		5.22%		8.35%
Obs.	458,641		426,449		198,669

Columbia

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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	-.0019 (.0011)	Girl, Girl	.0048 (.0029)	Girl, Girl, Girl	.0116 (.0056)
		Boy, Girl	-.0343 (.0029)	Boy, Boy, Girl	-.0174 (.0056)
		Girl, Boy	-.0284 (.0029)	Boy, Girl, Boy	-.0024 (.0058)
				Girl, Boy, Boy	-.0143 (.0057)
				Boy, Girl, Girl	-.0068 (.0057)
				Girl, Boy, Girl	-.0050 (.0058)
				Girl, Girl, Boy	-.0232 (.0057)
All-Boy Baseline	.9149		.5271		.4627
Percent Effect	-0.21%		0.91%		2.51%
Obs.	250,154		228,137		117,008

Table 10, continued.

Kenya

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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	.0038	Girl, Girl	.0199	Girl, Girl, Girl	.0156
	(.0018)		(.0040)		(.0075)
		Boy, Girl	-.0044	Boy, Boy, Girl	-.0112
			(.0041)		(.0075)
		Girl, Boy	.0039	Boy, Girl, Boy	-.0058
			(.0040)		(.0076)
				Girl, Boy, Boy	-.0010
					(.0073)
				Boy, Girl, Girl	-.0025
					(.0076)
				Girl, Boy, Girl	.0117
					(.0076)
				Girl, Girl, Boy	.0029
					(.0074)
All-Boy Baseline	.8773		.6017		.5104
Percent Effect	0.43%		3.31%		3.06%
Obs.	131,046		114,996		69,759

Vietnam

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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	.0044	Girl, Girl	.0874	Girl, Girl, Girl	.0735
	(.0009)		(.0034)		(.0065)
		Boy, Girl	.0092	Boy, Boy, Girl	-.0220
			(.0034)		(.0067)
		Girl, Boy	.0050	Boy, Girl, Boy	-.0401
			(.0034)		(.0067)
				Girl, Boy, Boy	-.0420
					(.0066)
				Boy, Girl, Girl	.0151
					(.0067)
				Girl, Boy, Girl	.0117
					(.0067)
				Girl, Girl, Boy	-.0181
					(.0065)
All-Boy Baseline	.9579		.4887		.4097
Percent Effect	0.46%		17.88%		17.94%
Obs.	176,232		169,191		86,886

Table 10, continued.

China					
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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	.0066	Girl, Girl	.1976	Girl, Girl, Girl	.2110
	(.0016)		(.0061)		(.0115)
		Boy, Girl	.0373	Boy, Boy, Girl	-.0505
			(.0061)		(.0124)
		Girl, Boy	.0123	Boy, Girl, Boy	-.0487
			(.0059)		(.0121)
				Girl, Boy, Boy	-.0520
					(.0119)
				Boy, Girl, Girl	.0336
					(.0122)
				Girl, Boy, Girl	.0571
					(.0122)
				Girl, Girl, Boy	-.0058
					(.0111)
All-Boy Baseline	.9596		.3646		.2343
Percent Effect	0.69%		54.20%		90.06%
Obs.	54,709		52,672		22,394

Notes: Standard errors in parentheses. The dependent variable in columns 1 and 2 is a dummy equal to 1 if the family has 2 or more children; in columns 3 and 4 a dummy equal to 1 if the family has 3 or more children; in columns 5 and 6 a dummy equal to 1 if the family has 4 or more children. The excluded category is all boys. The sample includes all currently married mothers (excluding consensual unions) between the ages of 18 and 50 with family composition calculated using all children less than 12 years old living at home. All-boy baseline is the fraction of mothers in all-boy families who are divorced. i.e. the intercept term. Percent effect is the increase in the probability of divorce for the mother of an all-girl family compared to an all-boy family; that is, it is the ratio of the coefficient for an all-girl family (row 1) over the all-boy baseline. Data sources: 1990 and 2000 Mexican Censuses; 1973, 1985, and 1993 Columbian Censuses; 1989 and 1999 Kenyan Censuses; 1989 and 1999 Vietnam Censuses; 1982 Chinese Census.

Table 11. The Effect of Child Gender on the Probability of a Consensual Union.

Mexico

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0070	.0028	Girl ,Girl	.0067	.0038	G, G, G	.0020	.0004
	(.0043)	(.0011)		(.0022)	(.0016)		(.0040)	(.0032)
			Boy, Girl	-.0056	-.0006	B, B, G	.0002	-.0020
				(.0021)	(.0016)		(.0040)	(.0032)
			Girl, Boy	-.0021	.0009	B, G, B	.0042	.0044
				(.0021)	(.0016)		(.0040)	(.0033)
						G, B, B	.0009	.0026
							(.0040)	(.0032)
						B, G, G	.0055	.0049
							(.0040)	(.0033)
						G, B, G	.0054	.0045
							(.0041)	(.0033)
						G, G, B	-.0009	-.0006
							(.0039)	(.0032)
All-Boy Baseline	.2851	.2219		.2169	.2169		.2106	.2179
Percent Effect	2.46%	1.26%		3.09%	1.75%		0.95%	0.18%
Obs.	45,247	590,516		290,701	545,269		162,398	245,568

Columbia

Sex of 1st child	Families with 1 child	Families with ≥ 1 children	Sex order of 1st two children	Families with 2 children	Families with ≥ 2 children	Sex order of 1st three children	Families with 3 children	Families with ≥ 3 children
Girl	.0057	.0025	Girl ,Girl	.0023	.0029	G, G, G	.0020	.0003
	(.0049)	(.0015)		(.0033)	(.0023)		(.0060)	(.0044)
			Boy, Girl	-.0122	-.0040	B, B, G	-.0120	-.0075
				(.0032)	(.0023)		(.0059)	(.0043)
			Girl, Boy	-.0108	-.0036	B, G, B	-.0002	-.0004
				(.0032)	(.0023)		(.0060)	(.0044)
						G, B, B	-.0046	.0024
							(.0060)	(.0044)
						B, G, G	-.0017	.0010
							(.0060)	(.0044)
						G, B, G	-.0029	-.0032
							(.0061)	(.0045)
						G, G, B	-.0068	-.0006
							(.0059)	(.0044)
All-Boy Baseline	.4774	.3623		.3535	.3506		.3606	.3515
Percent Effect	1.19%	0.69%		0.65%	0.83%		0.55%	0.09%
Obs.	57,798	445,305		191,467	387,507		108,656	196,040

Notes: Standard errors in parentheses. The dependent variable is a dummy equal to 1 if the respondent is in a consensual relationship. Data are from the 1990 and 2000 Mexican Censuses and 1973, 1985, and 1993 Columbian Censuses.

Table 12. The Effect of Child Gender on the Probability of Living in a Polygamous Relationship.

Marginal Effect on the Probability of a Polygamous Relationship								
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)
Girl	.0023	.0034	Girl, Girl	.0041	.0061	G, G, G	.0023	.0059
	(.0018)	(.0018)		(.0026)	(.0025)		(.0047)	(.0047)
			Boy, Girl	-.0015	.0004	B, B, G	-.0030	-.0026
				(.0026)	(.0025)		(.0048)	(.0047)
			Girl, Boy	-.0003	.0014	B, G, B	-.0023	.0006
				(.0026)	(.0025)		(.0048)	(.0047)
						G, B, B	.0004	.0019
							(.0047)	(.0046)
						B, G, G	.0061	.0082
							(.0048)	(.0047)
						G, B, G	.0019	.0041
							(.0048)	(.0047)
						G, G, B	.0034	.0066
							(.0047)	(.0046)
Controls?	No	Yes		No	Yes		No	Yes
All-Boy Baseline	.1226	.1238		.1074	.1079		.1126	.1136
Percent Effect	1.88%	2.75%		3.82%	5.65%		2.04%	5.19%
R-squared	.0001	.042		.0001	.040		.0001	.032
Obs.	131,046	131,046		114,996	114,996		69,759	69,759

Notes: Standard errors in parentheses. The dependent variable in column 1 is a dummy equal to 1 if the family has 2 or more children; in column 2 a dummy equal to 1 if the family has 3 or more children; in column 3 a dummy equal to 1 if the family has 4 or more children. The excluded category is all boys. Data are from the 1989 and 1999 Kenyan Census extracts. The sample includes all married mothers between the ages of 18 and 50 with children less than 12 years old living at home. Regression controls include mother's age (cubic), 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 polygamy 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 (row 1) over the all boy baseline.

Appendix Table A1. The Effect of Child Gender on Probability of Divorce or Separation Controlling for Covariates, U.S. Census Data.

	Marginal Probability of 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)
Girl	.0059	.0039	Girl, Girl	.0042	.0036	G, G, G	.0056	.0049
	(.0006)	(.0003)		(.0007)	(.0005)		(.0014)	(.0012)
			Boy, Girl	-.0020	.0004	B, B, G	.0000	.0006
				(.0007)	(.0005)		(.0014)	(.0012)
			Girl, Boy	-.0007	.0014	B, G, B	.0057	.0034
				(.0007)	(.0005)		(.0014)	(.0012)
						G, B, B	.0060	.0043
							(.0014)	(.0012)
						B, G, G	.0014	.0017
							(.0015)	(.0012)
						G, B, G	.0028	.0015
							(.0015)	(.0012)
						G, G, B	.0016	.0017
							(.0014)	(.0012)
R-squared	.025	.026		.027	.035		.049	.055
Obs.	1,400,796	3,595,710		1,409,491	2,194,914		550,965	785,423

Notes: Standard errors in parentheses. The dependent variable is a dummy equal 1 if the respondent is currently divorced. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all ever-married mothers between the ages of 18 and 50 with children less than 12 years old living at home. 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).

Appendix Table A2: The Effect of Child Gender on the Probability of a Shotgun Marriage, by Race, Education, and Cohort, California Vital Statistics Sample.

	Coefficient on Female*Ultra, by Group	All-Boy Baseline	Percent Effect
	(1)	(2)	(3)
White	-.0036	.667	
	(.0017)		
Black	-.0043	.374	
	(.0057)		
Other Race	-.0010	.820	
	(.0036)		
Hispanic	-.0010	.542	
	(.0026)		
Drop Out	-.0005	.415	
	(.0030)		
High School	-.0063	.624	
	(.0028)		
College	.0008	.828	
	(.0019)		
Born ≤ 1959	.0006	.830	
	(.0036)		
Born 1960s	-.0025	.807	
	(.0019)		
Born 1970s	-.0032	.452	
	(.0024)		
Born 1980s	.0387	.229	
	(.0197)		

Notes: Standard errors in parenthesis. The dependent variable is a dummy equal 1 if the mother is married at delivery. All models include all main effects. Percent effect is the percent change in the dependent variable when moving from a boy family to a girl family. It is the ratio of the coefficient in column XXX over the baseline. The data are from California Vital Statistics. The sample includes all first time mothers in California, 1989-199.

Appendix Table A3. The Effect of Child Gender on the Probability of an Additional Child Controlling for Covariates, U.S. Census Data.

Marginal 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+)	Sex and order of 1st two children	At least 1 more child (3+)	Sex and order of 1st three children	At least 1 more child (4+)
Girl	-.0019	Girl, Girl	.0071	Girl, Girl, Girl	.0073
	(.0005)		(.0009)		(.0020)
		Boy, Girl	-.0459	Boy, Boy, Girl	-.0348
			(.0009)		(.0020)
		Girl, Boy	-.0451	Boy, Girl, Boy	-.0130
			(.0009)		(.0021)
				Girl, Boy, Boy	-.0118
					(.0021)
				Boy, Girl, Girl	-.0078
					(.0021)
				Girl, Boy, Girl	-.0055
					(.0021)
				Girl, Girl, Boy	-.0221
					(.0020)
Obs.	3,152,194		1,974,664		710,807

Notes: Standard errors in parentheses. The dependent variable in column 1 is a dummy equal to 1 if the family has 2 or more children; in column 2 a dummy equal to 1 if the family has 3 or more children; in column 3 a dummy equal to 1 if the family has 4 or more children. The excluded category is all boys. Data are from the 1940 to 2000 U.S. Censuses. The sample includes all currently married mothers between the ages of 18 and 50 with children less than 12 years old living at home. 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).

Appendix Table A4. Year of Last Observed Delivery for First Time Mothers in 1989-1990, Birth Certificates Longitudinal Sample.

Year	Fraction
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	3.0
1999	2.5
2000	2.3
2001	1.9

Notes: Entries are the fraction of mothers whose last observed delivery occurs in the specified year. The sample is a panel of California mothers obtained by linking longitudinally birth certificates for years 1989-2001. In particular, the sample includes all mothers who are first time mothers in 1980 or 1990, observed between 1989 and 2001.

Figure 1. Percent Increase in the Probability of Divorce for an All-Girl Family Relative to an All-Boy Family, by Family Size.

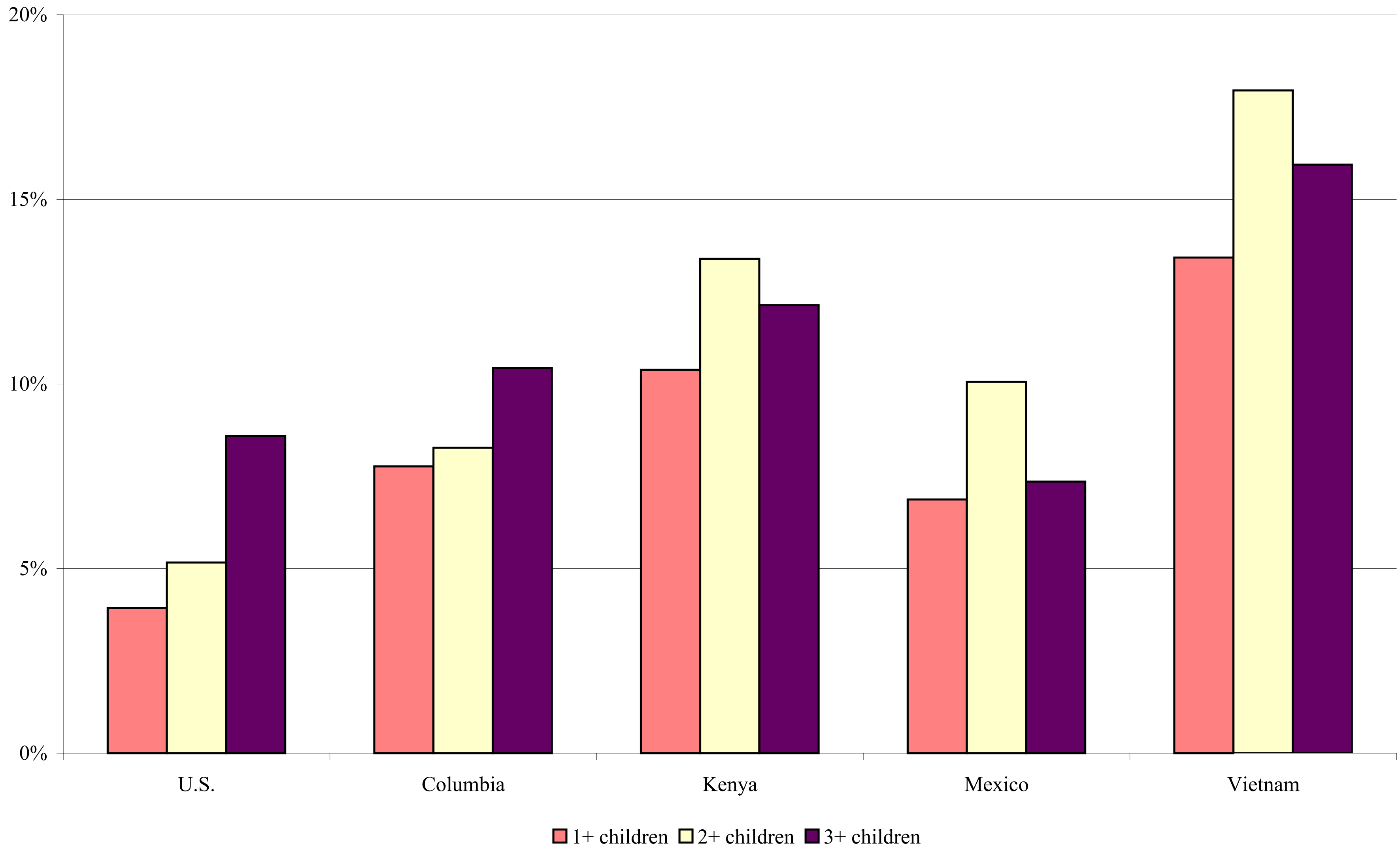

Figure 2. Percent Increase in the Probability of an Extra Child for an All-Girl Family Relative to an All-Boy Family, by Family Size.