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ORIGINAL PAPER

Does child gender affect marital status? Evidence from Australia

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Abstract Pooling microdata from five Australian censuses, I explore the relationship between child gender and parents' marital status. By contrast with the USA, I find no evidence that the gender of the first child has a significant impact on the decision to marry or divorce. However, among two-child families, parents with two children of the same sex are 1.7 percentage points less likely to be married than parents with a boy and a girl. This finding is unlikely to be consistent with theories of preference for sons over daughters, differential costs, role models, or complementary costs but is consistent with a theory of mixed-gender preference.

Keywords Marriage · Divorce · Daughters · Sons

JEL Classification J12 · J13

1 Introduction

Does children's gender affect parents' decision to marry or divorce? While researchers have long documented a preference for sons in developing countries, more recent studies have shown that the same is true in the USA. In the USA, parents with daughters are more likely to divorce if they are married at the time of the birth and less likely to marry if they are unmarried at the time of the birth.

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This paper uses confidentialized unit record file data from the 1981–2001 Australian Census files. As the data used in this paper are confidential, they cannot be shared with other researchers. However, the Stata do-file is available from the author upon request.

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Another possibility is that parents are concerned with the sex combination of their children. A spate of studies have shown that parents with two children of the same gender are more likely to have a third child than parents with one boy and one girl. Given this pattern, it is also possible that the sex combination of the children affects the decision to marry or divorce.

In this paper, I explore the relationship between children's gender and parents' marital status, pooling data from several censuses. This paper is innovative in two respects. First, it uses data from Australia, a country where the relationship between child gender and marital status has not been explored using large datasets. Second, it takes account not only of the number of boys and girls in a family but also allows for the possibility of mixed-gender preference—that parents with more than one child might prefer a mix of boys and girls to having all boys or all girls.

The notion that parents might prefer a mix of boys and girls is a straightforward implication of parents having convex preferences with respect to boys and girls. Yet while it seems intuitive that parents might prefer a 'diversified child portfolio', previous studies have largely ignored this aspect of parental preferences and its implications for parents' marital status. Australian data present an opportunity to test this aspect of the relationship between child gender and parents' marital status.

To preview the results, I do not find evidence that Australian parents have a higher chance of being married if they have more boys (or more girls). However, there is strong evidence that mixed-gender preference affects marriage patterns. Compared to parents with a boy and a girl, parents with two children of the same sex are 1.7 percentage points less likely to marry if they are unmarried.

The remainder of this paper is structured as follows. Section 2 reviews the relevant literature. Section 3 outlines a simple theoretical model of the effect of child gender on marital status. Section 4 presents the census data and empirical results, and the final section concludes.

2 Previous research

This analysis is grounded in the theoretical work of Becker (1973, 1974) and Becker et al. 1977, which characterizes individuals as maximizing utility by choosing whether to marry and whether to remain married subject to uncertainty. A couple is assumed to terminate a marriage if the expected utility of remaining married falls below the expected utility in the separated state. Children are regarded as providing utility to parents, but that utility need not be the same for boys and girls.

In the USA, several studies have focused on the question of whether parents of daughters are more likely to divorce. Most early studies exploited datasets that provided information on marital history. Using data from the 1980 Current Population Survey (CPS), Morgan et al. (1988) found that parents of daughters were 6% more likely to have their first marriage end in divorce. Using the 1980 US Census, which contained information on marital history, Bedard and Deschenes (2005) and Ananat and Michaels (2007) found that daughters increased the risk of divorce by 4 and 3%, respectively. Supplementing Morgan et al. (1988) with data from the 1985, 1990, and 1995 CPS, Morgan and Pollard (2002) found no effect of child gender on divorce in the post-1980 CPS samples and argued that this reflects a change in parental response to child gender.

An alternative approach is that of Dahl and Moretti (2004), who pooled samples from the 1940–2000 US Censuses and looked at current marital status instead of marital history. This had the advantage of greatly increasing the sample size but the disadvantage that in observing only current marital status, divorce effects may be attenuated by remarriage. Dahl and Moretti concluded that daughters are 1% more likely to reside with a currently divorced or separated mother or father. Their effects are largest in the 1940–1980 census samples, and they did not find a significant impact of the gender of the first girl on marital status in the 1990 and 2000 census samples.

Studies focusing on countries other than the USA have generally failed to find any relationship between child gender and divorce. Across 18 European countries, using data from the 1980s and 1990s, Diekmann and Schmidheiny (2004) found no consistent relationship between child gender and divorce. In Canada, Wu (1995) and Wu and Penning (1997) found no statistically significant impact of child gender on divorce or cohabitation. In Australia, the only relevant study on child gender and marital stability (Bracher et al. 1993) found no significant relationship. The authors did not report the coefficient on child gender or its standard error, but with a sample of only around 2,500, it is possible that their study was unable to reject large effects in either direction.

Another relevant literature is that relating to parity progression and the sex composition of children. A consistent finding in the demographic and economic literatures is that US parents with two same-sex children are more likely to have a third child than parents who have both a boy and a girl (Ben-Porath and Welch 1976; Pebley and Westoff 1982; Angrist and Evans 1998). The same is true in Australia (Young 1977; Kippen et al. 2005; Gray and Evans 2006), Denmark (Jacobsen et al. 1999), and Sweden (Murphy 1992; Schullström 1996). Attitudinal data paint a similar picture; parents in most countries say that if they were to have two children, they would prefer to have a son and a daughter (Sensibaugh and Yarab 1997; Hank and Kohler 2000).

A variety of studies have explored parental preferences for sons versus daughters. When parents are asked whether they would prefer a son or a daughter, there is a clear preference for sons in the USA (Pollard and Morgan 2002), most European countries (Hank and Kohler 2000), and other developed nations (Marleau and Saucier 2002). For example, Dahl and Moretti (2004) reported evidence from Gallup polls conducted in the USA over the period 1941–2003 and find that US respondents consistently said that if they could only have one child, they would prefer a son to a daughter. A 1997 International Gallup poll on the same question found a preference for sons over daughters in 13 of the 16 countries surveyed. In general, men's preference for sons is stronger than women's preference for sons.

The only Australian survey I have been able to find on this issue is a study by Weston et al. (2004), who asked respondents about the factors that were most important to them in deciding whether to have children.[1] Among the possible

[1] I am grateful to Edith Gray for drawing this study to my attention.

responses were "have at least one/another boy" and "have at least one/another girl". Australian men were more likely to say that having a boy was important (23%) than to say that having a girl was important (18%). Conversely, Australian women were more likely to say that having a girl was important (16%) than to say that having a boy was important (12%). Among childless respondents, women were almost indifferent between boys and girls (13% boys, 14% girls), while men's preferences were strongly in favor of boys (25% boys, 19% girls).[2]

Lastly, a series of papers by Shelly Lundberg and coauthors have focused on the effect of child gender on parental time use. In both the USA (Lundberg and Rose 2002; Lundberg 2005a) and Germany (Choi et al. 2005), the birth of a son appeared to lead to a larger increase in the father's labor supply than the birth of a daughter. Exploring time use patterns in the US, Lundberg (2005b) found that highly educated parents devote more childcare time to young sons. Looking only at very young children, Lundberg et al. (2007) found that fathers are more likely to play with, diaper, and feed sons than daughters, while mothers interact similarly with sons and daughters. Having sons also appears to increase parental happiness in the USA. Several studies have found that husbands and wives with sons report higher levels of marital satisfaction than do parents with only daughters (Barnett and Baruch 1987; Katzev et al. 1994; Cox et al. 1999; Mizell and Steelman 2000, all cited in Lundberg 2005c).

3 A simple model of child gender and divorce

The theoretical model adapts that of Dahl and Moretti (2004), who assumed that both husband and wife have transferable utility functions of the form $h(B_t,G_t,C_t)+X_t$, where B_t and G_t denote the number of boys and girls in the household at time t, C_t denotes nontransferable utility, and X_t denotes transferable utility. As they point out, transferable utility functions have the advantage that one can ignore issues of allocation and bargaining power and can simply consider the sum of utility of the two partners.

Where t indexes time, i indexes the married and unmarried states, ε_t is a normally distributed, mean-zero, marriage-specific shock, married and unmarried states are denoted as M and U, respectively, and the utility weights attached to boys and girls are denoted α^i and β^i, respectively, Dahl and Moretti expressed the couple's utility as:

$$U\left(\alpha^i B_t + \beta^i G_t, C_t\right) + X_t + I[i = M] * \varepsilon_t \qquad i = M, U \tag{1}$$

[2] Weston et al. (2004) also found large gaps in boy/girl preferences among respondents aged 20–29 (men favored boys by 30 to 22%, women favored girls by 19 to 12%) and among respondents whose highest level of education was year 12 or less (men favored boys by 31 to 25%, women favored girls by 18 to 15%).

Where p, q, and s are the prices of boys, girls, and nontransferable consumption, transferable consumption is the numeraire good, and Y_t is combined income, the combined period budget constraint is:

$$pB_t + qG_t + sC_t + X_t = Y_t \tag{2}$$

For simplicity, Dahl and Moretti assumed that prices and income are the same in both married and unmarried states and that the budget constraint holds with equality in each period (i.e., no borrowing or saving).

A key assumption underlying Dahl and Moretti's functional form is that the effect of boys and girls on parents' utility is additively separable. To relax this assumption somewhat, I add the interaction term BG_t and its utility weight weight γ^i, so the utility function becomes:

$$U\left(\alpha^i B_t + \beta^i G_t + \gamma^i BG_t, C_t\right) + X_t + I[i = M] * \varepsilon_t \qquad i = M, U \tag{3}$$

With the interaction term, the combined period budget constraint is:

$$pB_t + qG_t + rBG_t + sC_t + X_t = Y_t \tag{4}$$

where $r>0$ denotes that having both boys and girls is costlier than having only boys or only girls.

Following Dahl and Moretti, I can explore the implications of three hypotheses—gender bias, role model, and differential cost—in this somewhat extended model.

1. The *gender bias hypothesis* suggests that parents prefer one gender over another. For example, if parents with sons have a higher level of utility in marriage than parents with daughters, then $\alpha^M > \beta^M$. In its ordinary form, the gender bias hypothesis implies that parents' utility in marriage does not depend on having both boys and girls (i.e., $\gamma^M = 0$).
2. The *role model hypothesis* posits sex-specialization in parenting. For example, suppose that parents are altruistic, agree that fathers are better than mothers at raising boys, and expect that the mother will have custody of the children in the event of a separation. In this event, parents' utility in the event of a separation will be lower if they have only one son than if they have only one daughter (i.e., $\alpha^U < \beta^U$). The role model hypothesis does not have any implications for the gender mix of the children (i.e., $\gamma^U = 0$).
3. The *differential cost hypothesis* is that the monetary cost of raising boys and girls differs. In Australia and the USA, the evidence suggests that girls are more expensive than boys, implying that $p < q$.[3] This hypothesis does not have any implications for the cost of having both boys and girls (i.e., $r = 0$).

These three hypotheses imply that the coefficient on the interaction term equals zero, in other words, that the gender mix of children has no impact on parents' utility

[3] In the USA, Olson (1983) estimated that for one-child families, a girl costs around US$900 each year more to raise up to the age of 18 than a boy. I have been unable to obtain any published evidence on the cost difference of raising boys and girls in Australia. However, Paul Henman, an Australian social researcher at the University of Queensland who specializes in estimating the cost of children, informs me that his unpublished calculations put the cost of raising teenage girls at around 9–10% (A$800 per year) higher than the cost of raising teenage boys (email correspondence, 8 January 2007).

or the cost of child-rearing. To take account of the interaction term, I add two hypotheses to those listed above.

4. The *mixed-gender preference hypothesis* posits that the utility of parents with two or more children is related to whether they have both boys and girls. For example, if parents have convex preferences over boys and girls, then $\gamma^M>0$.
5. The *complementary cost hypothesis* suggests that (holding constant family size) the cost of child-rearing is lower if all the children in a family are of the same gender. For example, if toys and clothes are more easily passed down to a child of the same sex, then for families with two children, $r>0$.

4 Census data and results

The main empirical analysis focuses on the effect of child gender on parents' marital status in Australia. To maximize statistical power, microdata from all the available census samples are pooled together. The Australian census is conducted every 5 years, and since 1981, the Australian Bureau of Statistics (ABS) has made available 1% samples of the full census. I combine the 1981, 1986, 1991, 1996, and 2001 census samples.

As only two of the five census samples (1981 and 1986) contain information on the respondent's marital history, the dependent variable is the parents' current marital status. So long as child gender does not affect the probability that a divorced respondent remarries, this strategy is likely to cause only attenuation bias in the estimates. (For the 1981 and 1986 censuses, I found qualitatively similar results when using either 'married/currently divorced' or 'married/ever divorced' as the dependent variable.) To partially address this problem, I also show results in which the dependent variable is defined as married/never married.

Only one observation per household is used, being the household reference person. The census records the 'registered marital status' of each adult, according to five categories: never married, widowed, divorced, separated, and married. Widows and widowers are excluded, and divorced and separated are coded together. Among those with any children, 79% are married, 10% are divorced or separated, and 11% are never married.[4] If parents with daughters are more likely to divorce, then one should expect to see more girls living in households with a divorced head. If parents with daughters are less likely to marry after their child is born, more girls should be observed living in households with a never-married household head. Both effects would lead one to expect to see more girls living in a household with an unmarried head (where unmarried can be divorced, separated, or never married).

The independent variable of interest is the gender of the children in a family.[5] As the census only provides data on children and adults in the same family, it is not possible to properly take account of instances in children living with neither parent after a

[4] According to the Australian Bureau of Statistics, the proportion of ex-nuptial births was 13.2% in 1981 and 30.7% in 2001: "Population-Births" in *Year Book 2004*, Cat No 1301.0.

[5] Children are identified as family members aged 18 or younger, coded as "dependent child" or "dependent student".

separation (e.g., instances in which children reside with grandparents after a divorce). Neither is it possible to exclude nonbiological children (e.g., adoptees, foster children, and step-children), as the census samples do not code such children in a consistent manner. However, very few children fall into these categories. In the 2001 wave of the Household, Income and Labour Dynamics in Australia survey (HILDA), only 0.3% of resident children were foster children, step-children, or grandchildren living in a household without a natural or adoptive parent. The HILDA survey does not separate adoptive children from natural children, but aggregate national data show that the ratio of adoptions to births was 1 in 80 in 1981 and 1 in 435 in 2001 (Australian Institute of Health and Welfare 2006; Australian Bureau of Statistics 2006).

As the census only contains information on resident children, it is desirable to minimize the possibility that the eldest child has left home. I therefore exclude from the sample families in which the reference person is aged under 18 or more than 40, and those in which the youngest child is aged more than 12. These age restrictions match those used by Dahl and Moretti (2004) and are aimed at maximizing the number of completed fertility spells in the sample. Taking account of the youngest child in the family exploits the fact that the spacing of the first and second children is almost always 5 years or less. Note that the parental age restriction of 40 or younger is based on the age of the family reference person. I cannot exclude on the basis of the mother's age, as this would involve dropping all single-parent father-headed households, a characteristic strongly related to the gender of the children.

An alternative would have been to apply a more stringent age restriction on the reference person. (For example, assuming that parents are at least 18 years old at the time of first birth and children do not leave home before the age of 15, the sample can be restricted to cases in which the reference person is aged 33 years old or below. In this case, the sample is about one-third smaller, but the regression results are qualitatively similar.) Another alternative would have been to exclude families in which the number of "total issue" differs from the number of dependent children in the household. However, this has the drawback that the total issue variable is not available for two of the five census samples (1991 and 2001).

The final sample consists of 61,025 families. Of these, 29% have only one child, 43% have two children, and 28% have three or more children. I observe clear evidence of that the sex composition of children affects parity progression—among families with three or more children, the first two children are the same sex in 55% of cases (if the sex composition of the first two children did not affect parity progression, one would expect this probability to be 50%).

As the gender of a child is random, the primary regression specifications are straightforward.[6] For the most part, the specifications take the form of estimating the

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[6] For a detailed discussion of sex-selection technology and the laws governing its use in Australia, see Kippen et al. (2005). They point out that sex-selective abortion is likely to be extremely rare, as 99% of abortions/assisted miscarriages are carried out within the first trimester of pregnancy, before the point at which fetal gender can be determined. For parents using in vitro fertilization (IVF), Preimplantation Genetic Diagnosis can facilitate sex-selection, but its use for nonmedical purposes is illegal in three Australian states (South Australia, Victoria, and Western Australia). Following a 2005 ruling by the Australian Health Ethics Committee, IVF clinics in other states have agreed not to use the technology for nonmedical reasons. Before the ruling, it is estimated that around 250 couples used the technology for nonmedical reasons (Robotham 2005).

Table 1 Proportion of girls and parents' marital status

Number of children	Any	1	2	3
Dependent variable is unmarried (1) or married (0)				
Proportion girls	0.0015	−0.0012	−0.0008	0.0131
	(0.0042)	(0.0070)	(0.0068)	(0.0092)
Observations	61,069	17,737	26,010	17,322
Pseudo R^2	0.0000	0.0000	0.0000	0.0001
Observed probability	0.2104	0.3268	0.1675	0.1555
Dependent variable is divorced/separated (1) or married (0)				
Proportion girls	0.001	0.0011	−0.0028	0.0078
	(0.0036)	(0.0060)	(0.0057)	(0.0080)
Observations	54,457	13,971	24,147	16,339
Pseudo R^2	0.0000	0.0000	0.0000	0.0001
Observed probability	0.1145	0.1454	0.1032	0.1047
Dependent variable is never married (1) or married (0)				
Proportion girls	0.0009	−0.0024	0.002	0.0075
	(0.0035)	(0.0068)	(0.0051)	(0.0065)
Observations	54,834	15,706	23,517	15,611
Pseudo R^2	0.0000	0.0000	0.0000	0.0002
Observed probability	0.1206	0.2398	0.0792	0.063

Results are marginal effects from a probit model. Standard errors in parentheses
The second part (dependent variable is divorced/separated or married) excludes never married. The third part (dependent variable is never married or married) excluded divorced or separated.
*Statistical significance at the 10% level
**Statistical significance at the 5% level
***Statistical significance at the 1% level

effect of child gender on the reference person's marital status. Following the formal model in section 3, I estimate the effect on marital status of the number of boys (B_t), the number of girls (G_t), and the interaction between the two (BG_t) in period t and marital state i:

$$\Pr(i = M) = F(B_t, G_t, BG_t) + \varepsilon \qquad i = U, M;\ i = D,\ M;\ i = N,\ M \tag{5}$$

The effect of child gender on marital status is estimated in three sets of specifications:

– unmarried (U) versus married (M), where unmarried includes divorced, separated, and never married;
– divorced or separated (D) versus married (M), where those who are never married are excluded; and
– never married (N) versus married (M), where those who are divorced or separated are excluded.

In each case, the regressions are estimated using a probit model.

To begin, I estimate the effect of the sex of a child on marital status, using as the dependent variable the proportion of children who are girls. This variable takes the values {0,1} in a one-child family, {0,1/2,1} in a two-child family, and {0,1/3,2/3,1} in a three-child family. To remove the effects of family size, the regressions are estimated separately by family size. In the case of families larger than three children, I focus only on the sex of the first three children. Note that in the one-child case, the variable ‘proportion girls’ is effectively an indicator variable taking the value 1 if the child is a girl and 0 if the child is a boy.

Table 1 shows the results of this regression for families with any number of children, one child, two children, and three children and for unmarried versus married, divorced versus married, and never married versus married. In none of these specifications is the proportion of girls in a family statistically significant at conventional levels. In the specifications with all families, one-child families, and two-child families, the coefficients and standard errors are sufficiently small that I can reject (at the 5% level of significance) effects of more than ±1 percentage point.

I now estimate a second model, replacing the *Proportion Girls* variable with an indicator variable denoting whether all the children are of the same sex. As this variable is always unity for one-child families, Table 2 shows regression results only for two-child and three-child families. The first row indicates that in two-child families where both children are of the same sex, parents are 1.7 percentage points less likely to be married than in two-child families with both a boy and a girl. As the baseline probability is 17%, this indicates that children’s gender can account for around 10% of the variation in marital status among two-child families—a surprisingly large effect.

Table 2 Whether children are the same sex and parents’ marital status

Number of children	2	3
Dependent variable is unmarried (1) or married (0)		
Children same sex	0.0167***	−0.0088
	(0.0047)	(0.0071)
Observations	25,969	12,598
Pseudo R^2	0.0005	0.0001
Observed probability	0.1677	0.15
Dependent variable is divorced/separated (1) or married (0)		
Children same sex	0.0054	−0.0097
	(0.0039)	(0.0061)
Observations	24,091	11,906
Pseudo R^2	0.0001	0.0003
Observed probability	0.1028	0.1006
Dependent variable is never married (1) or married (0)		
Children same sex	0.0147***	−0.0002
	(0.0036)	(0.0050)
Observations	23,493	11,400
Pseudo R^2	0.0013	0
Observed probability	0.0799	0.0607

Standard errors in parentheses

The second part (dependent variable is divorced/separated or married) excludes never married. the third part (dependent variable is never married or married) excluded divorced or separated.

*Statistical significance at the 10% level

**Statistical significance at the 5% level

***Statistical significance at the 1% level

Table 3 Child gender and parents' marital status controlling for parental demographics

Number of Children	Any	2
Dependent variable is unmarried (1) or married (0)		
Proportion girls	0.0025	
	(0.0043)	
Children same sex		0.0093**
		(0.0045)
Observations	54,528	23,210
Pseudo R^2	0.2219	0.2161
Observed probability	0.2123	0.1694
Dependent variable is divorced/separated (1) or married (0)		
Proportion girls	0.0018	
	(0.0032)	
Children same sex		0.0034
		(0.0034)
Observations	48,601	21,546
Pseudo R^2	0.1713	0.1834
Observed probability	0.1163	0.1053
Dependent variable is never married (1) or married (0)		
Proportion girls	0.0009	
	(0.0026)	
Children same sex		0.0059***
		(0.0022)
Observations	48,877	20,942
Pseudo R^2	0.3087	0.3019
Observed probability	0.1213	0.0795

Results are marginal effects from a probit model. Standard errors in parentheses

The second part (dependent variable is divorced/separated or married) excludes never married. The third part (Dependent variable is never married or married) excluded divorced or separated. All regressions control for family income quintile, a quadratic in reference person age, the number of years of education of the reference person, and an indicator for the census year.

*Statistical significance at the 10% level

**Statistical significance at the 5% level

***Statistical significance at the 1% level

The next two parts of Table 2 break down the effect into divorce and failing to marry. The *Children Same Sex* coefficient in the divorced versus married specification is insignificant, while the *Children Same Sex* coefficient in the never married versus married specification is positive and statistically significant. This suggests that the effect is driven primarily by parents with two same-sex children not marrying, rather than by such families divorcing.

For three-child families, the *Children Same Sex* coefficients are insignificant in all specifications, but the standard errors are sufficiently large that I cannot reject effects in either direction. It is probably not surprising that the effects do not persist from two-child to three-child families. Conditional on having two children, both being of the same sex is a random event. However, three-child families in which all children are of the same sex were at one stage a two-child family with both children of the same sex. If parents with two same-sex children have a higher chance of separation (as the results in the first column of Table 2 suggest), then parents who go on to have three same-sex children will have a higher propensity to marry (or stay married).

As a robustness check, Table 3 shows results including controls for family income quintile, a quadratic in reference person age, number of years of education of the reference person, and an indicator for the census year.[7] If child gender was nonrandom, one might expect this to have a significant impact on the coefficients. Instead, adding controls has the effect of inflating the standard errors but does not have a major impact on the coefficients. Across all families, *Proportion Girls* remains positive but statistically insignificant, while in two-child families, the coefficient on *Children Same Sex* in two-child families attenuates slightly but remains statistically significant.

Given that both Morgan and Pollard (2002) and Dahl and Moretti (2004) find that the effect of child gender on marital status is much less pronounced in the 1990s and 2000s than in earlier decades, it is useful to explore whether the same pattern emerges in Australia. In Table 4, I compare the effects across census years. For reasons of space, the dependent variable is unmarried versus married (as this captures both divorce effects and never-married effects). Two specifications are presented—the effect of the proportion of girls for all families and the effect of two same-sex children in two-child families.

For the most part, child gender does not have a statistically significant effect on marital status when considering the census samples individually. Families with more girls are less likely to be unmarried in 1996 (significant at the 10% level) and more likely to be unmarried in 2001 (significant at the 5% level). In two-child families, *Children Same Sex* is positive and significant in 1986, 1996, and 2001 (all at the 5% level). Together, these coefficients suggest that the effect of child gender on marital status has grown stronger in Australia over time—the opposite trend to that observed in the USA.

As a further check, the Appendix shows the results of a more flexible functional form, looking at all combinations of child gender and comparing these results with those of Dahl and Moretti (2004). In general, the results for Australia are not statistically significant. While the coefficient on the sex of the first child is positive, the coefficients on having two girls (against the counterfactual of two boys) and on having three girls (against the counterfactual of three boys) are sensitive to sample selection.

Another point worth noting about the USA results in Dahl and Moretti (2004) is that they also show some evidence of a mixed-gender preference in two-child families. Relative to families with two boys (BB), parents with two girls (GG) are 0.20 percentage points more likely to be divorced/separated, parents with a boy and then a girl (BG) are 0.48 percentage points less likely to be divorced/separated, and

[7] One might also be concerned that the results presented in this paper do not reflect the effect of the gender of the first two children on marital status but what might be called "sample attrition through parity progression". To take an extreme example, suppose that having a third child required: (a) two same-sex children, and (b) married parents. In this case, families with precisely two same-sex children would be more likely to be unmarried, even if the gender of the first two children had no direct impact on marital status. To test this theory, I estimate the effect of the first two children being of the same sex on marital status, with the sample being those with two *or more* children. The coefficients are smaller than those shown in the first column of Table 2 (0.006 when the dependent variable is 1 unmarried or 0 married, and 0.007 when the dependent variable is 1 never married or 0 married) but still statistically significant. I am grateful to Shelly Lundberg for suggesting this additional robustness check.

Table 4 Child gender and parents' marital status across censuses dependent variable is unmarried (1) or married (0)

	1981 Census		1986 Census		1991 Census		1996 Census		2001 Census	
Number of children	Any	2	Any	2	Any	2	Any	2	Any	2
Proportion girls	0.0000 (0.0079)		−0.005 (0.0087)		0.0102 (0.0085)		−0.0176* (0.0103)		0.0223** (0.0110)	
Children same sex		0.0029 (0.0080)		0.0204** (0.0093)		0.0014 (0.0090)		0.0252** (0.0117)		0.0303** (0.0129)
Observations	13,212	5,691	12,746	5,508	11,269	4,787	12,324	5,129	11,518	4,854
Pseudo R^2	0.0000	0.0000	0.0000	0.0011	0.0002	0.0000	0.0002	0.0008	0.0003	0.001
Observed probability	0.1367	0.1014	0.1703	0.1362	0.1448	0.1097	0.2836	0.2254	0.325	0.2773

Results are marginal effects from a probit model. Standard errors in parentheses
*Statistical significance at the 10% level
**Statistical significance at the 5% level
***Statistical significance at the 1% level

parents with a girl and then a boy (GB) are 0.28 percentage points less likely to be divorced/separated. Assuming the four groups (BB, BG, GB, GG) are equally common, this suggests that USA parents with two children of the same sex (BB or GG) are 0.28 percentage points more likely to be divorced/separated than parents with a boy and a girl (BG or GB).

Recall that in modeling the relationship between child gender and parental marital status, three theories were proposed that predicted a preference for boys over girls (or vice-versa): the gender bias hypothesis, the role model hypothesis, and the differential cost hypothesis. In the Australian context, I do not find evidence that Australian parents have a higher chance of being married if they have more boys (or more girls). One possible conclusion to be drawn from this is that the gender bias hypothesis, the role model hypothesis, and the differential cost hypothesis should be rejected. Another possibility is that two or more of these effects operate, but in opposite directions (for example, it might be the case that Australian parents have a significant gender bias in favor of boys, but this is cancelled out by the cost of girls being significantly higher).

In addition, two theories were proposed that predicted a preference for the sex-mix of children: the mixed-gender hypothesis and the complementary cost hypothesis. The empirical results show that Australian parents with two children are more likely to marry (or stay married) if they have a boy and a girl than if they have two children of the same sex. This is consistent with the mixed-gender hypothesis (which posits that parents' utility in marriage is highest when the couple has both a boy and a girl). It also suggests that the complementary costs hypothesis (which suggests that children are cheaper to raise if they are all of the same sex) is either invalid or else that the magnitude of the cost effect is smaller than the preference effect.

As noted above, the mixed-gender hypothesis is a natural corollary of parents having convex preferences over boys and girls (i.e., a decreasing marginal rate of substitution). However, this also raises the question of why the impact of mixed-gender preferences is not stronger still in the case of families with three boys or three girls. One possible explanation is that the sample of three-child families is too small, or that there are too few three-child families with unmarried parents to detect significant effects. Alternatively, it might be the case that after having two same-sex children, parents choose either to separate/divorce, or to have a third child. In this event, those who have three same-sex children would have already chosen the "marriage track", and their marital status would be unaffected by the gender of the third child.

5 Conclusion

Combining microdata from five Australian censuses over the period 1981–2001, I estimate the effect of child gender on marital stability. Only in the 2001 census is there clear evidence that daughters are associated with lower marriage rates. However, there is consistent evidence that in two-child families, parents are less likely to be married if the children are of the same sex. Most of this effect appears to be driven by never-married couples failing to marry rather than by married couples divorcing. A two-child couple with same-sex children is 1.7 percentage points less likely to be married than a couple with a boy or a girl. Child gender can therefore explain 10% of the variation in marital status among two-child families. In the working paper version (Leigh 2006), I also analyze survey data and find some suggestive evidence that this impact is driven by paternal rather than maternal attitudes.

In modeling the relationship between child gender and parental marital status, five theories were proposed: the gender bias hypothesis, the role model hypothesis, the differential cost hypothesis, the mixed-gender hypothesis, and the complementary cost hypothesis. The mixed-gender hypothesis—which posits that parents' utility in marriage is highest when the couple has both a boy and a girl—is most consistent with these results. As is well known, couples with two children of the same sex are more likely to have a third child than couples with a boy and a girl. These results also indicate that some such couples may also pursue an alternative strategy: investing less in the relationship by failing to marry or deciding to divorce.

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Appendix

Table A1 Comparing Australia and the USA

Sex of first child	Families with 1 child	Families with ≥1 child	Sex order of first 2 children	Families with 2 children	Families with ≥2 children	Sex order of first 3 children	Families with 3 children	Families with ≥3 children
	(1)	(2)		(3)	(4)		(5)	(6)
USA (Source: Dahl and Moretti 2004)								
Girl	−0.0004	0.0011	Girl, girl	0.0020	0.0025	G, G, G	0.0056	0.0053
	(0.0006)	(0.0003)		(0.0007)	(0.0005)		(0.0014)	(0.0011)
			Boy, girl	−0.0048	−0.0011	B, B, G	−0.0003	0.0003
				(0.0006)	(0.0005)		(0.0013)	(0.0011)
			Girl, boy	−0.0028	0.0001	B, G, B	0.0062	0.0045
				(0.0005)	(0.0005)		(0.0014)	(0.0012)
						G, B, B	0.0055	0.0049
							(0.0014)	(0.0012)
						B, G, G	0.0025	0.0029
							(0.0014)	(0.0012)
						G, B, G	0.0032	0.0025
							(0.0014)	(0.0012)
						G, G, B	0.0014	0.0021
							(0.0014)	(0.0011)
All-Boy Baseline	0.1812	0.1360		0.1170	0.1098		0.0980	0.0978
Percent Effect	−0.2%	0.9%		1.7%	2.3%		5.7%	5.4%
Observations	1,554,818	4,169,265		1,679,127	2,614,447		659,523	935,320

Table A1 (Continued)

Sex of first child	Families with 1 child	Families with ≥1 child	Sex order of first 2 children	Families with 2 children	Families with ≥2 children	Sex order of first 3 children	Families with 3 children	Families with ≥3 children
Australia								
Girl	0.0011	0.0035	Girl, girl	−0.0025	0.0024	G, G, G	−0.0065	0.008
	(0.0060)	(0.0027)		(0.0057)	(0.0043)		(0.0105)	(0.0096)
			Boy, girl	−0.0114	−0.0014	B, B, G	0.0002	0.0065
				(0.0053)	(0.0042)		(0.0103)	(0.0093)
			Girl, boy	−0.0018	0.0048	B, G, B	0.018	0.0247
				(0.0054)	(0.0043)		(0.0118)	(0.0106)
						G, B, B	0.0106	0.0141
							(0.0115)	(0.0101)
						B, G, G	0.0062	0.0109
							(0.0115)	(0.0103)
						G, B, G	0.0052	0.0217
							(0.0114)	(0.0106)
						G, G, B	0.0083	0.0159
							(0.0108)	(0.0098)
All-Boy Baseline	0.1466	0.1123		0.1036	0.1004		0.0987	0.0939
Percent Effect	0.8%	3.1%		−2.4%	2.4%		−6.6%	8.5%
Observations	13,971	54,457		24,147	40,486		11,917	16,339

Dependent variable is divorced/separated (1) or married (0)
“All-Boy Baseline” is the fraction of all-boy families with a currently divorced parent. “Percent effect” is the coefficient in the first row divided by the All-Boy Baseline.

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