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Does Child Gender Affect Marital Status?

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ABSTRACT

Pooling microdata from five Australian censuses, I explore the relationship between child gender and divorce. By contrast with the United States, 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. Surveys of parental attitudes suggest that this effect is more likely to be driven by fathers than by mothers. This finding is not consistent with theories of preference for sons over daughters, differential costs, role models or complementary costs, but is consistent with a theory of parity preference.

JEL Classification: J12, J13

Keywords: marriage, divorce, daughters, sons, child gender

I. Introduction

Does the gender of their children 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 United States. US 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.

Another possibility is that parents are concerned with the sex combination of their children. A spate of studies have shown the existence of "parity progression", by which parents with two children of the same gender are more likely to have a third child. 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, using a large sample of 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 in using a large number of observations. Second, it takes account not only of the number of boys and girls in a family, but also allow for the possibility of parity preference – that instead of preferring one sex over another, parents might instead prefer to have children of both sexes than children of the same sex.

To preview the results, I do not find evidence that the gender of the first child affects the decision to marry or divorce in Australia. However, there is strong evidence that parity preference affecting 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. Turning to attitudinal data, I find suggestive evidence that most of this effect is driven by fathers, rather than by mothers.

The remainder of this paper is structured as follows. Section II briefly reviews the relevant literature. Section III outlines a simple theoretical model of the effect of child gender on marital status. Section IV presents the census data and empirical results. Section V analyses attitudinal data, and the final section concludes.

II. Previous Research

This analysis is grounded in the theoretical work of Becker (1973, 1974) and Becker, Landes and Michael (1997), 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 US, 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, Lye and Condran (1988) find that parents of daughters were 6 percent more likely to have their first marriage end in divorce. Using the 1980 US Census, which contained information on marital history, Bedard and Deschene (2005) and Ananat and Michaels (2004) find that daughters increased the risk of divorce by 4 percent and 3 percent respectively. Supplementing Morgan, Lye and Condran (1988)

with data from the 1985, 1990 and 1995 CPS, Morgan and Pollard (2002) find no effect of child gender on divorce in the post-1980 CPS samples, and argue that this reflects a change in parental response to child gender.

An alternative approach is that of Dahl and Moretti (2004), who pool the 1940-2000 US Censuses, and look at current marital status instead of marital history. This has the advantage of greatly increasing the sample size, but the disadvantage that in observing only current marital status, divorce effects may by attenuated by remarriage. Dahl and Moretti conclude that daughters are 1 percent more likely to reside with a currently divorced or separated mother or father. Their effects are largest in the 1940-80 census samples, and they do 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 US have generally failed to find any relationship between child gender and divorce. Across eighteen European countries, using data from the 1980s and 1990s, Diekmann and Schmidheiny (2004) find no consistent relationship between child gender and divorce. In Canada, Wu (1995) and Wu and Penning (1997) find 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 2500, it is possible that their study was unable to reject large effects in either direction.

Another relevant literature is that relating to parity progression. A consistent finding in the demographic and economic literature 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; Gray and Evans 2004; Kippen, Gray and Evans 2005), Denmark (Jacobsen, Møller and Englholm 1999), and Sweden (Schullström, 1996). Attitudinal data paint a similar picture – parents consistently say that if they were to have two children, they would prefer to have both a son and a daughter.

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 US (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) report evidence from Gallup polls conducted in the US over the period 1941-2003, and find that US respondents consistently say 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 substantially 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 ask respondents about the factors that were most important to them in deciding whether to have children.[1] Among the possible responses were "have at least one/another boy" and "have at least one/another girl". Australian men are more likely to say that having a boy was important (23 percent) than they are to say that having a girl was important (18 percent). Conversely, women are more likely to say that having

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

a girl was important (16 percent) than to say that having a boy was important (12 percent). Among childless respondents, women are almost indifferent between boys and girls (13 percent boys, 14 percent girls), while men's preferences are strongly in favor of boys (25 percent boys, 19 percent girls).[2]

Lastly, a series of papers by Shelly Lundberg and coauthors focus on the effect of child gender on parental time use. In both the US (Lundberg and Rose 2002; Lundberg 2005a) and Germany (Choi, Joesch and Lundberg 2005), the birth of a son appears to have a positive impact on fathers' labor supply that is substantially larger than the effect of the birth of a daughter. Exploring time use patterns in the US, Lundberg (2005b) finds that highly educated parents devote more childcare time to young sons. Looking only at very young children, Lundberg, McLanahan and Rose (2005) find 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 US. Lundberg (2006) reports on several studies indicating that husbands and wives with sons report higher levels of marital satisfaction than do parents with only daughters.[3]

III. A Simple Model of Child Gender and Divorce

The theoretical model adapts that of Dahl and Moretti (2004), who assume 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 non-transferable utility, and X_t denotes transferable utility. As they point out,

[2] Weston et al (2004) also find 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%).

[3] Barnett and Baruch, 1986; Katzev et al., 1994; Cox et al, 1999; Mizell and Steelman, 2000.

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

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

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 assume that prices and income are the same in both married and unmarried states, and that the budget constraint holds with equality in each period (ie. 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 γ^i, so the utility function becomes:

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

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 is the additional cost associated with having children of different sexes.

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

(i) The **gender bias hypothesis** suggests that parents prefer one gender over another. For example, if parents have a preference for sons, this implies that $\alpha^M>\beta^M$. In its ordinary form, the gender bias hypothesis implies that $\gamma^M=0$.

(ii) The **role model hypothesis** involves sex-specialization in parenting. For example, that fathers are better at raising boys. Since courts typically grant custody to mothers, altruistic parents will take into account the gender of their parents when deciding whether to separate. Under the example in which parents regard fathers as better at raising boys, this implies that $\alpha^D<\beta^D$. The role model hypothesis implies that $\gamma^D=0$.

(iii) The **differential cost hypothesis** is that the monetary cost of raising boys and girls differs. For example, if girls are more expensive than boys, then $p<q$. This hypothesis implies that $r=0$.[4]

Note that each of the above hypotheses imply that the coefficient on the interaction term equals zero. To take account of the interaction term, I add two hypotheses to those

[4] In the US, Olsen (1983) estimates that for one-child families, a girl costs around $900 each year more to raise up to the age of 18 than a boy. I have been unable to obtain similar evidence for Australia.

listed above.

(iv) The **parity 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 prefer to have at least one boy and one girl, $\gamma^M=0$.

(v) The **complementary cost hypothesis** suggests that the cost of child-rearing is related to whether the family has both boys and girls. For example, if toys and clothes are more easily passed down when children are of the same sex, then for families with two children, one might expect that $r<0$.

IV. 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 five years, and since 1981, the Australian Bureau of Statistics (ABS) has made available 1 percent samples of the full census. I combine the 1981, 1986, 1991, 1996 and 2001 census samples.

The dependent variable is the marital status of the adults in each family. To determine this, I use as the dependent variable the marital status of 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 percent are married, 10 percent are divorced or separated, and 11 percent are never married.[5] If parents with daughters are more likely to divorce,

[5] 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.

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).[6]

The independent variable of interest is the gender of the children in a family.[7] Since the census only provides data on children and adults in the same family, 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 over 40, and those in which the youngest child is aged over 12.[8] The final sample consists of 61,025 families. Of these, 29 percent have only one child, 43 percent have two children, and 28 percent have three or more children. I observe clear evidence of parity progression – among families with three or more children, the first two children are the same sex in 55 percent of cases (in the absence of parity progression, one would expect this probability to be 50 percent).

Since the gender of a child is random, the primary regression specifications are

[6] A preferable approach would have been to construct an "ever divorced" variable. However, I am only able to construct such a variable for the 1981 and 1986 censuses.

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

[8] These age restrictions are the same as 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 less 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, since this would involve dropping all single parent father-headed households, a characteristic strongly related to the gender of the children. One alternative approach would have been to apply a more stringent age restriction on the reference person. For example, in place of the restriction on the age of the youngest child, I could restrict the reference person to 33 years old or younger. When I do so, the results are qualitatively similar, but this approach reduces the sample size by about one-third. 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 total issue has a low top-code in the 1986 and 1996 census samples, and is missing in the 1991 and 2001 census samples.

straightforward.[9] For the most part, the specifications take the form of estimating the effect of some children's gender on the reference person's marital status. Following the model in Section III, 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):

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

As above, M denotes married and U denotes unmarried (either divorced/separated or never married). D denotes divorced or separated (in this specification, those who are never married are excluded). N denotes never married (in this specification, those who are divorced or separated are excluded). In each case, I estimate the regression 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, {0,½,1} 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.

Table 1 shows the results of this regression, for families with any number of children, one-child, two-children and three-children (columns 1-4), and for unmarried versus

[9] For a detailed discussion of sex-selection technology and the laws governing its use in Australia, see Kippen, Gray and Evans (2005). They point out that sex-selective abortion is likely to be rare in Australia, since 99% abortions/assisted miscarriages are carried out within the first trimester of pregnancy, prior to the point at which fetal gender can be determined. For parents using in vitro fertilization, Preimplantation Genetic Diagnosis can facilitate sex-selection, but its use for non-medical 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 non-medical reasons. Prior to the ruling, it is estimated that around 250 couples used the technology for non-medial reasons (Robotham 2005).

married (Panel A), divorced versus married (Panel B), and never married versus married (Panel C). 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 percent level of significance) effects of more than plus or minus one percentage point.

Table 1: Proportion of Girls and Parents' Marital Status

Panel A: Dependent variable is unmarried (1) or married (0)				
Number of Children	**Any**	**1**	**2**	**3**
Proportion Girls	0.0015	-0.0012	-0.0008	0.0131
	[0.0042]	[0.0070]	[0.0068]	[0.0092]
Observations	61069	17737	26010	17322
Pseudo R^2	0.0000	0.0000	0.0000	0.0001
Observed Prob.	0.2104	0.3268	0.1675	0.1555
Panel B: Dependent variable is divorced/separated (1) or married (0)				
Number of Children	**Any**	**1**	**2**	**3**
Proportion Girls	0.001	0.0011	-0.0028	0.0078
	[0.0036]	[0.0060]	[0.0057]	[0.0080]
Observations	54457	13971	24147	16339
Pseudo R^2	0.0000	0.0000	0.0000	0.0001
Observed Prob.	0.1145	0.1454	0.1032	0.1047
Panel C: Dependent variable is never married (1) or married (0)				
Number of Children	**Any**	**1**	**2**	**3**
Proportion Girls	0.0009	-0.0024	0.002	0.0075
	[0.0035]	[0.0068]	[0.0051]	[0.0065]
Observations	54834	15706	23517	15611
Pseudo R^2	0.0000	0.0000	0.0000	0.0002
Observed Prob.	0.1206	0.2398	0.0792	0.063

Notes: Standard errors in brackets. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels respectively. Panel B excludes never married. Panel C excluded divorced or separated.

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. Since this

variable is always unity for one-child families, Table 2 shows regression results only for two-child and three-child families. Panel A 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. Since the baseline probability is 16.75 percent, this indicates that children's gender can account for around 10 percent of the variation in marital status among two-child families – a surprisingly large effect.

The next two panels of Table 2 break down the effect into divorce (Panel B) and failing to marry (Panel C). The *Children Same Sex* coefficient in Panel B is insignificant, while the *Children Same Sex* coefficient in Panel C 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. But three-child families in which all children are of the same sex were at one point 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.

Table 2: Whether Children are the Same Sex and Parents' Marital Status		
Panel A: Dependent variable is unmarried (1) or married (0)		
Number of Children	**2**	**3**
Children Same Sex	0.0167***	-0.0088
	[0.0047]	[0.0071]
Observations	25969	12598
Pseudo R^2	0.0005	0.0001
Observed Prob.	0.1677	0.15
Panel B: Dependent variable is divorced/separated (1) or married (0)		
Number of Children	**2**	**3**
Children Same Sex	0.0054	-0.0097
	[0.0039]	[0.0061]
Observations	24091	11906
Pseudo R^2	0.0001	0.0003
Observed Prob.	0.1028	0.1006
Panel C: Dependent variable is never married (1) or married (0)		
Number of Children	**2**	**3**
Children Same Sex	0.0147***	-0.0002
	[0.0036]	[0.0050]
Observations	23493	11400
Pseudo R^2	0.0013	0
Observed Prob.	0.0799	0.0607

Notes: Standard errors in brackets. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels respectively. Panel B excludes never married. Panel C excluded divorced or separated.

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.[10] If child gender was

[10] One might also be concerned that the results presented here 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 to me.

non-random, 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.

Table 3: Child Gender and Parents' Marital Status Controlling for Parental Demographics		
Panel A: Dependent variable is unmarried (1) or married (0)		
Number of Children	**Any**	**2**
Proportion Girls	0.0025	
	[0.0043]	
Children Same Sex		0.0093**
		[0.0045]
Observations	54528	23210
Pseudo R^2	0.2219	0.2161
Observed Prob.	0.2123	0.1694
Panel B: Dependent variable is divorced/separated (1) or married (0)		
Number of Children	**Any**	**2**
Proportion Girls	0.0018	
	[0.0032]	
Children Same Sex		0.0034
		[0.0034]
Observations	48601	21546
Pseudo R^2	0.1713	0.1834
Observed Prob.	0.1163	0.1053
Panel C: Dependent variable is never married (1) or married (0)		
Number of Children	**Any**	**2**
Proportion Girls	0.0009	
	[0.0026]	
Children Same Sex		0.0059***
		[0.0022]
Observations	48877	20942
Pseudo R^2	0.3087	0.3019
Observed Prob.	0.1213	0.0795

Notes: Standard errors in brackets. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels respectively. Panel B excludes never married. Panel C 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.

Since 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 (since 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.

Table 4: Child Gender and Parents' Marital Status Across Censuses
Dependent variable is unmarried (1) or married (0)

	1981 Census		1986 Census	
Number of Children	**Any**	**2**	**Any**	**2**
Proportion Girls	0.0000		-0.005	
	[0.0079]		[0.0087]	
Children Same Sex		0.0029		0.0204**
		[0.0080]		[0.0093]
Observations	13212	5691	12746	5508
Pseudo R^2	0.0000	0.0000	0.0000	0.0011
Observed Prob.	0.1367	0.1014	0.1703	0.1362
	1991 Census		**1996 Census**	
Number of Children	**Any**	**2**	**Any**	**2**
Proportion Girls	0.0102		-0.0176*	
	[0.0085]		[0.0103]	
Children Same Sex		0.0014		0.0252**
		[0.0090]		[0.0117]
Observations	11269	4787	12324	5129
Pseudo R^2	0.0002	0.0000	0.0002	0.0008
Observed Prob.	0.1448	0.1097	0.2836	0.2254
	2001 Census			
Number of Children	**Any**	**2**		
Proportion Girls	0.0223**			
	[0.0110]			
Children Same Sex		0.0303**		
		[0.0129]		
Observations	11518	4854		
Pseudo R^2	0.0003	0.001		
Observed Prob.	0.325	0.2773		

Notes: Standard errors in brackets. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels respectively.

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 percent level), and more likely to be unmarried in 2001 (significant at the 5 percent level). In two-child families, *Children Same Sex* is positive and significant in 1986, 1996 and 2001 (all at the 5 percent 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 United States.

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 the choice of specification.

V. Attitudinal Surveys

To further explore the effect of child gender on marital stability, I turn to a different type of evidence – survey data on parental behavior and attitudes towards children. Since the census does not include the relevant questions, the Household, Income and Labour Dynamics in Australia survey (HILDA) is used instead. HILDA is a four-wave panel dataset covering 2001-04, and questions are taken from all available waves (for more information on HILDA, see Watson 2005).[11] As with the results in the previous

[11] With only a four-year panel, I do not have sufficient births to observe the change in parental attitudes before and after childbirth, so I only exploit the cross-sectional nature of the data. This is also more straightforward, since the focus is on the combination of children, rather than the effect of a new child.

section, the fact that child gender is random means that such an approach should not bias the results.

For reasons of comparability, I apply the same sample restrictions as in the previous section – focusing on parents aged between 18 and 40, with at least one child aged 12 or under.[12] The sample is then further restricted to families with two resident children. Although HILDA contains data on non-resident children, the focus here is on the effect of child gender on attitudes and time use. I therefore take account only of resident children. Among those with two children, 74 percent are married, 7 percent are divorced or separated, and 19 percent are never married.

Note that since the HILDA questions for the most part ask about resident children, divorce or separation may lead to attenuation bias. In a dichotomous model, for example if same-sex children either cause parents to angrily separate or happily stay together, I should observe no effects. However, it is more likely that the effect of child gender is on a continuum, in which case it should be possible to learn something about the effect of child gender on marital status from observing parents who are living together.

I estimate the effect of child gender on six outcomes:

- Log usual weekly work hours in all jobs
- Log weekly time (in minutes) spent playing with children
- Satisfaction with partner (scaled from 0-10)
- "I often feel tired, worn out or exhausted from meeting the needs of children" (scaled from 1-7)

[12] Since HILDA contains information on the parentage of each child, I am are able to correctly code children in blended families.

- "I find that taking care of my children is much more work than pleasure" (scaled from 1-7)
- "Do fair share of looking after children" (scaled from 1-5)[13]

Table 5 shows the results from these specifications, with the first two dependent variables estimated using OLS, and the last four using ordered logit.

Table 5: Whether Children Are Same Sex and Parental Behavior

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log work hours	Log time with children	Satisfaction with partner	Children make me tired	Children more work than pleasure	Do fair share of looking after children
Panel A: Mothers						
Children Same Sex	-0.0500	0.1342*	0.1235	0.0522	-0.1390	-0.0112
	[0.0752]	[0.0717]	[0.1302]	[0.1222]	[0.1215]	[0.0789]
Observations	923	1515	1890	2078	2078	2072
R^2 or Pseudo R^2	0.0026	0.0059	0.0026	0.0019	0.0012	0.2912
Mean of dep var:	*3.094*	*3.028*	*8.086*	*4.881*	*2.796*	*2.958*
Panel B: Fathers						
Children Same Sex	0.0350	0.0052	0.0822	0.0647	0.166	0.0795
	[0.0242]	[0.0724]	[0.1460]	[0.1398]	[0.1402]	[0.1084]
Observations	970	948	1355	1327	1328	1322
R^2 or Pseudo R^2	0.0114	0.0012	0.0011	0.003	0.001	0.0166
Mean of dep var:	*3.813*	*2.369*	*8.325*	*4.048*	*2.684*	*2.984*

Note: Sample is parents aged between 18 and 40, with exactly two children, at least one of whom is aged 12 or under. Columns 1 and 2 are estimated using OLS. Columns 3-6 are estimated using ordered logit. All specifications include year fixed effects. Standard errors, clustered at the person level, in brackets. ***, ** and * denote statistical significance at the 1%, 5% and 10% levels respectively.

For the most part, the sex combination of children is not significantly related to parental work patterns, time with children, and attitudes towards children. The only

[13] Coded from 1 "Does much less than fair share" to 5 "Does much more than fair share" (reversed from original).

statistically significant coefficient is on mothers' work time, indicating that mothers with two children of the same sex spend on average 13 percent more time with their children (significant at the 10 percent level). This suggests that mothers are happier with same-sex children, not less happy. The other three coefficients with t-statistics above 1 all point in a similar direction: in families with two same-sex children, fathers tend to work more, fathers are *more* likely to think that children are more work than pleasure, and mothers are *less* likely to think that children are more work than pleasure. Together, these results indicate that if having two children of the same sex reduces the probability of marriage, the impact is more likely to be through the father than the mother.

VI. 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 two 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 percent of the variation in marital status among two-child families. Attitudinal data suggest that most of this impact arises from 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 parity preference hypothesis, and the complementary cost hypothesis. Only the parity preference hypothesis – which posits that parents' utility in marriage is highest when the couple has both a boy and a girl – appears to be consistent with the facts observed here. 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.

References

Ananat, E., and Michaels, G. 2004. "The Effect of Marital Breakup on the Income and Poverty of Women with Children". *mimeo*

Angrist, J. and Evans, W. 1998. "Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size" *American Economic Review*, 88(3): 450-477.

Barnett, R. C., and Baruch, G. K. 1987. "Determinants of Fathers' Participation in Family Work" *Journal of Marriage and the Family*, 49: 29-40

Becker, G.ary S. 1973. "A Theory of Marriage: Part 1." *The Journal of Political Economy*, 81(4): 813-46.

Becker, Gary S. 1974. "A Theory of Marriage: Part 2." *The Journal of Political Economy*, 82(2): S11-S26.

Becker, G..S., Landes, Elisabeth M., and Michael, Robert T. 1977. "An Economic Analysis of Marital Instability." *Journal of Political Economy*, 85(6): 1141-88.

Bedard, K., and Deschenes, O. 2005. "Sex Preferences, Marital Dissolution and the Economic Status of Women" *Journal of Human Resources*. 40(2): 411-434.

Ben-Porath, Y. and Welch, F. 1976. "Do sex preferences really matter?" *Quarterly Journal of Economics,* 90: 285-307.

Bracher, Michael; Santow, Gigi; Morgan, Philip and Trussell, James. 1993. "Marriage Dissolution in Australia: Models and Explanations." *Population Studies*, 47(3): 403-25.

Choi, Hyung-Jai, Joesch, Jutta and Lundberg, Shelly. 2005. "Work and Family: Child Gender, Marriage, and the Work Hours and Earnings of German Men" *mimeo*.

Cox, M., Paley, B., Burchinal, M., and Payne, D. C. 1999. "Marital Perceptions and Interactions Across the Transition to Parenthood", *Journal of Marriage and the Family*, 61(3): 611-625.

Dahl, Gordon B. and Moretti, Enrico. "The Demand for Sons: Evidence from Divorce,

Fertility, and Shotgun Marriage." NBER Working Paper, 2004, 10281.

Diekmann, Andreas and Schmidheiny, Kurt. 2004. "Do Parents of Girls Have a Higher Risk of Divorce? An Eighteen Country Study" *Journal of Marriage and Family* 66(3): 651-660

Gray, Edith and Evans, Ann. 2004. "Parity progression in Australia: What role does sex of existing children play?" Paper prepared for the 12th Biennial Conference of the Australian Population Association, 15-17 September 2004, Canberra.

Hank, K. and Kohler, H-P. 2000. Gender preferences for children in Europe: Empirical results from 17 FFS Countries. *Demographic Research,* 2.

Jacobsen, R. Møller, H., & Engholm, G. 1999. Fertility rates in Denmark in relation to the sexes of preceding children in the family. *Human Reproduction.* 14, 1127-1130.

Katzev, A. R, Warner, R. L., and Acock, A. C. 1994. "Girls or Boys - Relationship of Child Gender to Marital Instability" *Journal of Marriage and the Family*, 56(1): 89-100.

Kippen, R., Gray, E. and Evans, A. 2005. "Does parental desire for both a son and a daughter influence the decision to have another child? Evidence from Australian census data". Paper presented at XXV IUSSP International Population Conference, Tours France, 18–23 July

Lundberg, Shelly, 2005a. "Men and islands: dealing with the family in empirical labor economics". *Labour Economics* 12 (4), 591-612.

Lundberg, Shelly, 2005b. "The division of labor by new parents: does child gender matter?" Working Paper, Department of Economics, University of Washington, Seattle, WA.

Lundberg, Shelly. 2006. "Sons, daughters, and parental behavior". *Oxford Review of Economic Policy*. forthcoming

Lundberg, Shelly, McLanahan, Sara and Rose, Elaina. 2005. "Child Gender and Father Involvement in Fragile Families" Working Paper, Department of Economics, University of Washington, Seattle, WA.

Lundberg, Shelly and Rose, Elaina. 2002. "The Effects of Sons and Daughters on Men's Labor Supply and Wages" *Review of Economics and Statistics*, 84: 251-268.

Marleau, J.D., and Saucier, J-F. 2002. Preference for a first-born boy in Western societies. *Journal of Biosocial Science,* 34: 13-27.

Mizell, C. A., and Steelman, L. C. 2000. "All My Children: The Consequences of Sibling Group Characteristics on the Marital Happiness of Young Mothers" *Journal of Family Issues*, 21(7): 858-87.

Morgan, S. Philip; Lye, Diana N. and Condran, Gretchen A. 1988. "Sons, Daughters, and the Risk of Marital Disruption." *American Journal of Sociology*, 94(1): 110-29.

Morgan, S. Philip and Pollard, Michael S. 2002. "Do Parents of Girls Really Have a Higher Risk of Divorce?", *mimeo*, Duke University: Durham, NC

Murphy, M. 1992. The progression to the third birth in Sweden. In J. Trussell, R. Hankinson, and J. Tilton (eds.) *Demographic Applications of Event History Analysis*. Oxford: Clarendon Press, 141-156

Olson, Lawrence. 1982. *Costs of Children*. Lexington, Mass: Lexington Books, 1983.
Pebley, A., and Westoff, W. "Women's sex preferences in the United States: 1970 to 1975." *Demography,* 19: 177-189.

Pollard, M.S., and Morgan, S.P. 2002. "Emerging parental gender indifference? Sex composition of children and the third birth." *American Sociological Review*, 67: 600-613.

Robotham, Julie. 2005. "Banned: choosing a baby's sex" *Sydney Morning Herald*. 25 March

Watson, Nicole. (ed). 2005. *HILDA User Manual – Release 3.0*, Melbourne Institute of Applied Economic and Social Research, University of Melbourne.

Weston, Ruth, Lixia Qu, Robyn Parker and Michael Alexander. 2004. *'It's not for lack of wanting kids...' A report on the Fertility Decision Making Project*. AIFS Research Report 11. Melbourne: Australian Institute of Family Studies

Wu, Z. 1995. “The Stability of Cohabitation Relationships: The Role of Children.” *Journal of Marriage and the Family* 57: 231-236.

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

Young, C. 1977. “Family building differences between same sex and mixed sex families in Australia.” *Australian Journal of Statistics*, 19: 83-95

Appendix Table 1: Comparing Australia and the United States. Dependent variable is divorced/separated (1) or married (0).

Panel A: United States (Source: Dahl and Moretti 2004)

Sex of 1st 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)
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

Appendix Table 1: Comparing Australia and the United States. Dependent variable is divorced/separated (1) or married (0).

Panel B: Australia

Sex of 1st 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
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

Note: "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.