Children's gender and parents' subsequent fertility and partnerships in Iceland

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MS thesis in Economics

September 2016

UNIVERSITY OF ICELAND

Abstract

Are Icelandic parents influenced by the gender of their children? A number of studies in the United States have found that the parents of boys are more likely to remain and become married than parents of girls. However, the cause of this discrepancy and its external validity remain in question. Using data from the Icelandic Birth Registry we follow 54 689 mothers and 39 941 fathers whose first birth is covered by the data and tested for differences in averages of four measurements conditional on gender composition of parents' first children. Our results suggest that boys in Iceland are also associated, albeit more weakly, with a stabilizing effect on the relationship between their parents. Gender also has some measurable associations with fertility and timing between births. The causal mechanisms remain unclear but could be related to biological factors and/or the father's preference.

Preface

This 30-ECTS thesis was supervised by professor Tinna Laufey Ásgeirsdóttir and is submitted to the Faculty of Economics of the University of Iceland in partial fulfillment of the requirements for the degree of Master of Science in Economics.

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1 Introduction

Differential treatment of children based on their gender has been documented widely in many developing countries. A particularly strong pattern of parents' preferences is present in India, China, and some neighboring countries, where having girls appears, on average, less desired. Evidence includes girl neglect, which has been measured by the difference in health outcomes and child mortality between the genders (Sen, 1990) and more recently sex-selective abortions (Westley, 1995). Both result in a skewed male to female ratio whose consequences have not fully materialized since the most imbalanced generations have yet to reach adulthood (Poston, Conde, & DeSalvo, 2011; Jha et al., 2011), but could theoretically further amplify gender inequality (Edlund, 1999). There are some indications that preferences and thus sex ratios in these countries are affected by changes in relative economic and legal conditions of women (Jensen, 2012; Qian, 2008; Sun & Zhao, 2016).

This research looks at the behavioral response of fathers and mothers in Iceland to the gender of their children using birth registry data. Specifically, we evaluate whether children's gender is associated with their parents' future fertility and marriage and cohabitation status. While any discovered association might hint at parents' preferences, deciphering them is complex since fertility and partnership decisions are the result of negotiations between two individuals who do not necessarily agree. Previous studies on the effects of gender on parents' behavior include effects on labor supply (Lundberg & Rose, 2002), savings (Deolalikar & Rose, 1998), political orientation (Washington, 2008; Oswald & Powdthavee, 2010) and, as the present work, effects of gender on fertility and partnerships.

Relatively few studies on the latter mentioned issues exist in the West, where the main focus has been on the United States. Notably, it has been shown that parents of boys are more likely to remain and become married than parents of girls (Dahl & Moretti, 2008; Katzev, Warner, & Acock, 1994; Lundberg, McLanahan, & Rose, 2007; Mammen, 2008; Morgan, Lye, & Condran, 1988; Mott, 1994; Spanier & Glick, 1981). Another consistent pattern is that parents of same-sex siblings have higher expected future fertility (Angrist & Evans, 1998; Pollard & Morgan, 2002; Yamaguchi & Ferguson, 1995; Ben-Porath & Welch, 1976).

The United States is a highly multi-ethnic society, in which gender roles are relatively traditional in comparison to many other Western countries. Thus, the heavy focus in the literature on the United States raises questions about external validity, especially as the determinants of gender preferences are still not well understood, but likely influenced by social and institutional settings. Andersson, Hank, Rønsen, and Vikat (2006) advocate that to "investigate whether societal gender systems actually influence sex preferences for children, cross-national comparative research is highly desirable" (p. 256).

There has been some exploration based on European data, particularly Scandinavian population registries, which has so far revealed more nuanced gender effects than in the United States. Yet, a clear association between same-sex siblings and their mothers' increased fertility (Andersson et al., 2006; Jacobsen, Møller, & Engholm, 1999; Schullström, 1996), and likelihood of marital dissolution (Andersson & Woldemicael, 2001), has been revealed. However, at least one study (in Germany) found that fathers are less likely to reside with their children if they are female (Choi, Joesch, & Lundberg, 2008) and Andersson and Woldemicael (2001) do find correlation with number of girls and marital dissolution risk at parity three in Sweden.

Although all new analyses contribute to the understanding of parents' gender preferences, Iceland is of particular interest due to its high level of gender equality in international comparison where it has topped the Global Gender Gap Index since 2009 (World Economic Forum, 2015). Furthermore, we analyze behavior of both mothers and fathers, whereas most research on gender focuses on married couples or mothers only.

Potential popularization of sex-selection technology in the near future, allowing parents to choose gender before conception, makes understanding preferences timely. Strong preferences coupled with new technology could lead to demographic shocks, as have taken place previously in Asian countries via abortions. If parents favor one gender over the other or a certain mix, spread of sex-selection technology may lead to imbalanced sex ratio and/or decreased family size as the preferred (possibly heterogeneous) gender mix can be obtained by fewer attempts. As an example, parents' who want to have at least one girl and one boy and would continue trying until parity 5 will on average have 2.9 children with 1 out of 8 families reaching parity 5, instead of 2 if they relied on technology.[1]

[1] We base calculations on a fixed likelihood of a boy birth (p) equal to 51%. Probability

Furthermore, this study has methodological implications as gender effects can be the foundation for studies of causal effects in labor and family economics. An example of such work is by Angrist and Evans (1998) who use effects of gender mix of children as an instrument for fertility when studying parents' labor supply and Bedard and Deschênes (2005) who instrument for marriage dissolution with the gender of the first child.

2 Data and Methods

2.1 Data

Our data comes from the Icelandic Birth Registry (i. Fæðingaskrá) and contains information on all children born in Iceland from 1982 to 2014, including date of birth and gender, and also parental information such as cohabitation status, citizenship and health department.[2] In total, the data covers the births of 145 080 individuals attributed to 72 133 mothers and 68 979 fathers. While the dataset is relatively large, allowing for measurements of small effects, it has its downsides. The unit of analysis, the parents, are only observed when their children are born (1) in Iceland and (2) within the time frame of the data set.

Many of the observed parents started their procreation before 1982 and our main independent variable is the gender of the first children of parents. For this reason, we base our analysis on a subset of 54 689 mothers whose first birth is covered by the data, and the 39 941 fathers who were born after the year 1964. That is, they have yet to turn 17 by the end of 1981—and are thus unlikely to have children born before the registry initiation date.

Instead of a cutoff age for mothers we rely on a variable indicating the number of previous children which is unavailable for fathers. Restricting the data in this way generates a bias in regards to the mother's age. As an example, the cohort born in 1957 contains exclusively women who were childless before the age of 24. Since the variable *number of previous children* has some demonstrable inconsistencies, we have thus analyzed in parallel a smaller subset of women using a cutoff age of 16. The lower cutoff age for

of achieving the missing sex at parity $k \geq 2$ is equal to $p^{k-1}(1-p) + p(1-p)^{k-1}$.

[2]The birth registry is maintained by the Icelandic Directorate of Health. The study was approved by the Icelandic Directorate of Health (permission id 1507138), the National Bioethics Committee (project 15-173) and the Data Protection Authority (reference 2015101358ÞS/–) which was a prerequisite for accessing the data.

women is based on the observation that children in our data are (5.4 times) more likely to have been born to a mother between 16 and 17 in age than a father that age, which is 0.35 percent of children.[3]

Unfortunately, the dataset is incomplete; *father's id* is missing for 8 409 children, gender for 62 surviving children and mother's id for 74 children. Some variables we use for regression controls also have missing values. Furthermore, many observed parents are likely to continue their fertility beyond the year 2014, younger mothers are thus overrepresented in the more recent generations.

2.2 Models

To test whether children's gender is associated with their parents' fertility and partnership we focus on average differences in four measurements conditional on gender composition of parents' first children. We use multilevel OLS regression in which the main explanatory variable is gender combination of all previous children and its interaction with self-reported cohabitation status at last birth. When studying parents at higher parities we divide the gender variables into three categories; *all girls*, *all boys*, with mixed sex siblings as a reference category. Controls introduced are parental age and age squared at last birth, health department organized into urban/non-urban, whether the parents are Icelandic citizens and finally year (rounded to 1980, 1990, etc.) and quarter of birth of last child born.

The analysis can be divided into three parts: effects on (1) fertility, (2) time spacing of births, and (3) partnerships. On fertility, our corresponding response variables are total number of children ever born and whether the kth birth takes place (given $k-1$ children), for $k \in \{2, 3, 4\}$. Time spacing is simply days between birth k and $k-1$. In regards to partnership we measure average difference in probability of parent changing partners between birth k and $k-1$ and whether those that do stick together for their respective first two children have upgraded their cohabitation or marital status at birth two. As before, the main explanatory variable is gender combination of all previous children. At parity one the estimated models take the general form for measurement M:

[3] All data manipulation and statistical analysis was done in the R programming language. Source code is available upon request.

$$M = \alpha + \beta_1 \cdot \text{girl} + \beta_2 \cdot \text{cohabiting} + \beta_3 \cdot \text{girl} \cdot \text{cohabiting} + X\gamma + \epsilon.$$

Where X is the matrix of additional controls which we generally estimate regressions both with and without. Our aim is to estimate the average effect of the event of the first child being born a girl instead of a boy on the measurement of interest:

$$\left.\frac{\partial M}{\partial \text{girl}}\right|_{\text{cohabiting}=1} = \beta_1 + \beta_3 \quad \text{and} \quad \left.\frac{\partial M}{\partial \text{girl}}\right|_{\text{cohabiting}=0} = \beta_1.$$

These are the parameters whose estimates are reported in the results sections at parity one. At parity two parameters for *all boys* are added. Regressions are estimated separately for fathers and mothers since first children of a mother are not necessarily the father's first. One exception is in the last case where the unit of analysis are couples who have their first two children together. In all regressions, parents whose first children (explanatory variable) were not singletons or died in their first year are excluded. Non-singletons are defined as children born within 7 days of each other to the same mother, and 280 days when looking at fathers separately. Assuming that if gender affects fathers at all, it cannot do so unless he knows the gender of his previous children.

2.3 Assumptions

Causal interpretation of the models rests on missing data and the mechanism which chooses whether a child is born male or female to be independent of any parents' characteristics that could be related to our response variables. If each birth cannot realistically be considered a Bernoulli trial with fixed probability of a boy being born, or the data is not representative of parents in Iceland by unobserved characteristics, then any association has to be interpreted cautiously.

Much of the literature estimating the effects of children's gender on parents or using the described effects as instruments presupposes exogeneity of gender. As an example, Angrist and Evans (1998) state that "sex mix is virtually randomly assigned" (p. 451) and Bedard and Deschênes (2005) talk about "the fact that the sex of the firstborn child is random" (p. 411).

Both papers proceed to check for evidence of correlation with available parents' characteristic without finding significant correlations and proceed with causal interpretation.[4]

While absence of evidence (of endogeneity) is not evidence of absence there are some theoretical and empirical indications of presence. As far back as 1835 Quetelet and Heuschling note after examining records from several European countries that: "the excess of male births over female births can be observed in all [14] countries [...] but this excess is more marked for legitimate births than for illegitimate births" (as translated and cited in Brian and Jaisson (2007), p. 73). They theorize that the difference might stem from male fetuses being more sensitive to mothers having less "energy". Darwin (1871) highlights the same "singular fact" (p. 301, part II) using different data.

More recent studies have found lower offspring sex ratio (the proportion that is male) at times of war and famine (Ansari-Lari & Saadat, 2002; Song, 2012; Valente, 2015) and after natural disasters (Fukuda, Fukuda, Shimizu, & Møller, 1998). Helle, Helama, and Jokela (2008) find that higher temperature correlates with sex ratio amongst the Sami people in northern Finland. Fukuda, Fukuda, Shimizu, Andersen, and Byskov (2002) find negative association between parents smoking and sex ratio. Mathews, Johnson, and Neil (2008) find differences in mother's diet are associated with sex ratio but fail to find any effect of smoking. And in line with Darwin, Norberg (2004) reports that children in the United States whose parents were living together before the child's sex was known were 3.2 percent[5] more likely to be male than if their parents were living apart. Almond and Edlund (2007) report similar albeit much smaller increase in sex ratio at birth for married parents and mothers with more years of education, concluding "sex is not random" (p. 2495). Almond, Edlund, Joffe, and Palme (2016) discover strong correlation, in Swedish data, between hospital admissions of pregnant woman due to severe morning sickness (hyperemesis gravidarum) and both education and likelihood of pregnancy resulting in live girl birth. Hamoudi and Nobles

[4]More examples include Oswald and Powdthavee (2010) who state "we think of the gender of a child arriving in the household as a kind of exogenous event" (p. 214), and Pollard and Morgan (2002) who seem confident that "the stochastic process determining sex of previous children assures it is largely unrelated to other individual characteristics".

[5]Norberg reports this number to be 5.4 percent (p. 2407), as an interpretation of 5.4 percent increase in odds instead of probability.

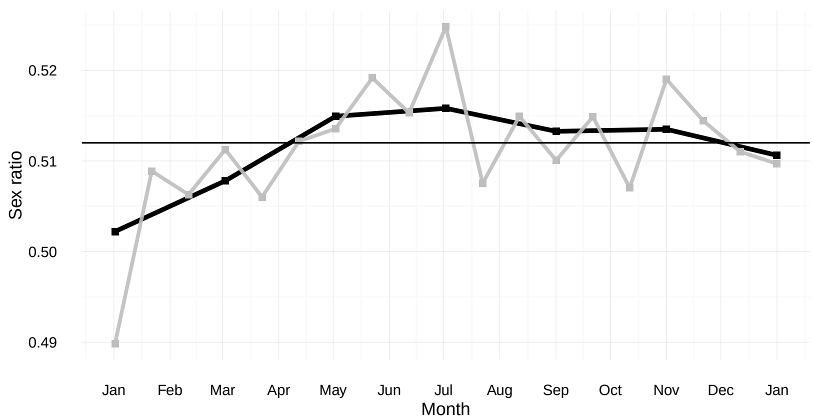
(2014) find a measure of depression has statistically significant correlation with both future divorce and lowered offspring sex ratio. Yet, another group of studies reports negative association between parents age and sex ratio, see Jacobsen, Møller, and Mouritsen (1999) for an overview.

A seminal paper by Trivers and Willard (1973) shows that a mother aiming to maximize her expected number of descendants could be better off varying the likelihood of male births. That is, if males brought up in relatively bad condition have a lower likelihood of reproductive success than females and vice versa, then the hypothesis predicts, via natural selection, that the human sex ratio "at birth correlates with socioeconomic status" (p. 91). Trivers-Willard hypothesis' relevance for mammals is still disputed but has been found to have strong predictive power in some other classes of animals (Cameron, 2004; West & Sheldon, 2002). However, it does not seem far-fetched to relate the hypothesis to the above mentioned findings of reduced sex ratio for unmarried women (being a single mother could be a proxy for poverty in 19th century Europe and USA in the 20th century), psychological depression, nationwide disasters, education, diet, and even old age. Although the apparent share of significant findings might be affected by publication bias, exogeneity of gender can perhaps not be taken as "fact".

Our data provides little information on socioeconomic status, but we make some comparisons of differences in demographic characteristics of people who have boys versus girls. Table 1 provides an overview; only one out of twenty comparisons are significantly different from zero at the five-percent level; *mother's health department* at first birth. One out of twenty is exactly what one would except from randomly generated data. Age is correlated in the direction expected but with high $p-$values. Parents of boys are not clearly more likely to be cohabiting nor married.

Furthermore, we explore associations between sex ratio and season as well as birth weight of older siblings that share the same mother and are close in age. The first exploration is inspired by a study which shows correlation between education (and other socioeconomic indicators) of mothers and season of birth using data from the United States (Buckles & Hungerman, 2013). Average education of mothers is consistently highest in the months around May and hits bottom in January. A study based on data from an Icelandic survey found a correlation between woman's month of birth and her education; those born in July and August had more years of schooling (p

Figure 1: Sex Ratio at Birth by Season in Iceland from 1982 to 2014



Note: The lines have have different degree of smoothing, horizontal line is total average (0.512), dark line two-month average, and gray line approximately three-week average.

< 0.1) than those born in January which were found to have had the least (Ólafsdóttir & Ásgeirsdóttir, 2015). While not proof of, the results could be explained by seasonality in socioeconomic status of parents in Iceland. Correlation between low birth weight (LBW) (< 2.5 kg) and various indicators of socioeconomic status of mothers has also been found in the United States (Parker, Schoendorf, & Kiely, 1994).

If a Trivers-Willard effect (measured by socioeconomic status) is present, sex ratio ought to also hit bottom in January and peak around May, and if LBW is caused by mothers' "condition" then it may help predict the gender of the next child.

Figure 1 shows an overview of sex ratio by season, and table 10 (page 25) shows regression results showing significantly higher sex ratio of around 1.5 percentage points in May and June and surprisingly November compared to January. Table 11 (page 25) presents estimates of difference in means for siblings born to the same mother whose previous child was low birth weight. The correlation fails to reach statistical significance but is consistent with prior expectations. A woman's next child after LBW birth is estimated to be 1.6 percentage points less likely to be male (about 3 percent).

Predicted seasonality in sex ratio and association with birth weight fits the Trivers-Willard hypothesis, and lacks alternative explanations. If the same seasonality that Buckles and Hungerman find in the United States is present in Iceland these informal results inspire further caution.

A final potential source if endogeneity is technology; while gender is generally chosen by nature, competing sex-selection technologies do exist. They are, however, currently, to our best knowledge, unavailable in Iceland and not legally available in most neighboring countries except the United States and Northern Cyprus.[6] We will also assume sex-selective abortion is sufficiently rare enough to not affect our results.

A second concern is missing data, which is rare except for the variable *father's id*, which is missing for 8 409 observed births (5.8 percent). Two thirds of the missing entries come from the years 1987 to 1994 when it is missing in some years for close to 50% of births in some health departments[7], but is otherwise missing for only between 0.5 and 2 percent of observations. Missing values of *father's id* is correlated with all mothers' demographic variables as well as showing a strong negative correlation with child's birth weight and height but statistically insignificant correlation with sex ratio. Attrition due to migration is possibly the source of a small bias whose direction is unknown.

[6]Accurate information on the topic seems scarce but web search for clinics points at these two countries and is confirmed by news reports and an overview by de Saille (2011). The Council of Europe's 1997 Convention on Human Rights and Biomedicine prohibits use of sex-selection technologies "except where serious hereditary sex-related disease is to be avoided." (Article 14).

[7]Capital area (*urban*) is the worst offender as can be seen in in Figure 2 in Appendix 1

Table 1: Means for Demographic Variables by Child's Gender

	Child's gender		
	Girl	Boy	Difference (SE)
Age at first child			
Mothers	25.14	25.12	0.02277 (0.04173)
Fathers	27.19	27.10	0.08789* (0.04978)
Age at second child			
Mothers	28.98	28.94	0.03922 (0.04710)
Fathers	30.79	30.76	0.02373 (0.05855)
Urban at first child			
Mothers	65.74%	64.91%	0.008309** (0.004097)
Fathers	65.1%	64.21%	0.009035* (0.004828)
Urban at second child			
Mothers	62.80%	62.86%	−0.000574 (0.004972)
Fathers	62.72%	62.90%	−0.001799 (0.006133)
Icelandic at first child			
Mothers	93.26%	92.95%	0.003093 (0.002182)
Fathers	90.55%	90.62%	−0.0006978 (0.002027)
Icelandic at second child			
Mothers	96.23%	96.56%	−0.003331* (0.001918)
Fathers	95.69%	95.89%	−0.001964 (0.002548)
Cohabitation at first child			
Mothers	79.69%	79.69%	−0.000001 (0.003463)
Fathers	79.98%	79.84%	0.00134 (0.004046)
Cohabitation at second child			
Mothers	91.44%	91.66%	−0.002269 (0.002861)
Fathers	90.90%	91.34%	−0.004395 (0.00361)
Married at first child			
Mothers	42.73%	43.00%	−0.002769 (0.00426)
Fathers	37.49%	37.85%	−0.003584 (0.004894)
Married at second child			
Mothers	52.71%	52.58%	0.001269 (0.005137)
Fathers	48.15%	47.30%	0.008512 (0.006339)

Notes: Numbers based on singleton births only.$^{*}p<0.01$; $^{**}p<0.05$

3 Results

3.1 Fertility

Tables 2 and 3 report estimates of the associations between gender and fertility. The overall view suggests a higher expected fertility for single parents who have *girls* (versus *boys*) and for cohabiting parents who have *all boys* at parity two and three (versus mixed gender).

Single mothers at parity two are estimated to be at least four percentage points more likely than single mothers of *mixed gender* to have a third child if they have *all girls* ($p < 0.05$). The effect on cohabiting mothers is much smaller and not statistically significant. Instead, cohabiting woman at parity two and three are estimated to be more likely to continue their fertility if they have *all boys*.

Association for fathers are consistent but fail to reach statistical significance with the exception of single fathers at parity three who are estimated to be more than 8 percentage points less likely (compared to fathers of mixed group) to have a fourth child if they have *all boys* ($p < 0.1$) while the association is almost the opposite for cohabiting fathers. There is also a clear contrast ($p < 0.01$) between single and cohabiting fathers in regards to the association (or estimated marginal effects) of a first girl (versus boy) and total number of children.

3.2 Time Spacing of Births

Tables 4 and 5 report differences in average timing between births by gender of previous children. The sample here are mothers and fathers who have had at least two children in columns 1-3 and three children in columns 4-6. The correlations reveal that at parity one a girl can expect their next siblings to come slightly faster than boys if their parents are not cohabiting whereas if her parents are cohabiting the difference is 0 or goes slightly in the opposite direction.

At parity two the same is true (third child comes faster) after same-sex siblings (versus mixed) but irrespective of parents cohabitation status. The estimated effect is between two and three month difference and is significant at the one-percent level when the association is estimated jointly for all mothers/fathers and not separately for those in cohabitation. In fact, the

Table 2: Mother's First Children and the Probability and Expected Count of Subsequent Births.

	Response variable:							
	1 = birth two takes place		1 = birth three takes place		1 = birth four takes place		Total number of children	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All girl(s)	0.014	0.012	0.052**	0.042**	0.006	−0.036	0.021	0.012
	(0.009)	(0.009)	(0.022)	(0.020)	(0.040)	(0.038)	(0.020)	(0.017)
All boys			−0.010	0.009	−0.054	−0.037		
			(0.021)	(0.020)	(0.040)	(0.038)		
All girl(s) × Cohabitation	−0.018*	−0.016*	−0.030	−0.011	0.015	0.062	−0.032	−0.020
	(0.010)	(0.009)	(0.023)	(0.021)	(0.042)	(0.039)	(0.022)	(0.019)
All boys × Cohabitation			0.048**	0.037*	0.105**	0.089**		
			(0.022)	(0.020)	(0.042)	(0.039)		
Controls?		Yes		Yes		Yes		Yes
Observations	53,230	50,333	37,501	35,985	17,201	16,669	53,230	50,333
R^2	0.013	0.109	0.002	0.095	0.005	0.055	0.005	0.150

Notes: OLS estimates with robust standard errors given in parenthesis. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Regressions (7) and (8) are estimated at parity one. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to 1 if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$.

Table 3: Father's First Child(ren) and the Probability and Expected Count of Subsequent Births.

	Response variable:							
	1 = birth two takes place		1 = birth three takes place		1 = birth four takes place		Total number of children	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All girl(s)	0.0003	0.008	0.040	0.022	0.023	0.008	0.025	0.033
	(0.011)	(0.010)	(0.027)	(0.026)	(0.053)	(0.049)	(0.024)	(0.021)
All boys			0.020	0.020	−0.088*	−0.089*		
			(0.026)	(0.025)	(0.049)	(0.053)		
All girl(s) × Cohabitation	−0.017	−0.024**	−0.011	0.007	0.007	0.022	−0.055**	−0.061***
	(0.012)	(0.011)	(0.028)	(0.027)	(0.055)	(0.051)	(0.026)	(0.023)
All boys × Cohabitation			0.015	0.011	0.144***	0.147***		
			(0.027)	(0.026)	(0.051)	(0.055)		
Controls?		Yes		Yes		Yes		Yes
Observations	38,519	36,417	24,484	23,551	10,270	10,029	38,519	36,417
R^2	0.007	0.117	0.001	0.107	0.006	0.118	0.002	0.147

Notes: OLS estimates with robust standard errors given in parenthesis. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to fathers whose first children (explanatory variable) were not singletons or have not survived their first year are excluded. Regressions (7) and (8) are estimated at parity one. Control variables included are father's and mother's age and age squared at last birth, dummies for cohabitation with father of last child, father's citizenship at last birth equal to 1 if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$.

Table 4: Mother's First Child(ren) and Days Until Next Birth

	Response variable:					
	Days between birth 1 and 2			Days between birth 2 and 3		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	1.727	−26.607	−11.644	−83.402***	−102.770	−83.015
	(10.932)	(30.942)	(29.798)	(19.356)	(71.443)	(68.544)
All boys				−71.350***	−72.104	−44.474
				(18.641)	(71.395)	(68.894)
All girl(s) × Cohabitation		32.540	15.886		20.591	20.808
		(32.966)	(31.781)		(74.222)	(71.230)
All boys × Cohabitation					1.110	−1.290
					(73.968)	(71.362)
Controls?			Yes			Yes
Observations	38,294	38,126	36,600	17,602	17,553	17,034
R^2	0.00000	0.024	0.084	0.001	0.002	0.064

Notes: OLS estimates with robust standard errors given in parenthesis. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{***}p$<0.01.

association for mothers is significantly different from 0 at the 0.1-percent level for both mothers of *all boys* and *all girls*.

The effects appear stronger for single fathers than cohabiting ones especially in the case of *all girls* with an estimated 160 days (more than five months) average reduction in wait for third birth for single fathers of *all girls* versus *mixed gender*.

3.3 Partnership

Table 6 and 7 report regression estimates on probability of change in partner between births. The women's sample size for these regressions is smaller than in the time spacing regressions due to missing data in the variable *father's id*.

The results are generally consistent with no association, especially for cohabiting parents. There are however some patterns which take the same form in estimates for fathers and mothers but slightly stronger for mothers.

Women are estimated more likely to change partners between child two and three if her first two were *mixed gender* (versus same-sex), with a stronger and statistically significant ($p \approx 0.0025$) difference with *all boys*

Table 5: Father's First Child(ren) and Days Until Next Birth

	Response variable:					
	Days between birth 1 and 2			Days between birth 2 and 3		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	11.869	−86.739**	−33.661	−84.580***	−160.068*	−159.237**
	(13.822)	(39.827)	(30.822)	(25.146)	(84.807)	(80.508)
All boys				−87.959***	−116.432	−98.138
				(23.810)	(84.169)	(79.263)
All girl(s) ×		116.807***	45.520		81.096	96.173
Cohabitation		(42.359)	(33.029)		(88.822)	(84.353)
All boys ×					29.605	20.943
Cohabitation					(87.767)	(82.637)
Controls?			Yes			Yes
Observations	25,166	25,008	24,052	10,526	10,477	10,186
R^2	0.00003	0.013	0.339	0.002	0.002	0.063

Notes: OLS estimates with robust standard errors given in parenthesis. All. girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are father's and mother's age and age squared, dummies for cohabitation with mother of last child, father's citizenship at last birth equal to 1 if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. *p<0.1; **p<0.05; ***p<0.01

which is estimated to be 1.8 percentage points when all mothers are pooled together. As with *all boys*, *all girls* are also negatively associated with likelihood partner change compared with *mixed gender* although the parameters fail to reach statistical significance. The association between *all boys* and partner stability is slightly stronger in the alternative women's sample reported in Table 15. Otherwise, the results from the this sample are very consistent to the main one across all measurements. Mother's who reached parity four (reported in Table 12) are significantly more likely to continue with same partner for child three and four, if the first three were *all boys*.

At parity one a first born girl is associated with a slightly higher chance for women to have next child with a new partner (as opposed to the father of her first child), the size of the difference is sizable since 0.7 percentage points represents 4.5% deviation over the baseline of 0.177.

Finally, Table 8 reports estimates association of girls and the likelihood of couples that have their first two children together but are not cohabiting upgrade to cohabitation status (column 1 and 2) or those unmarried upgrade to marriage before child two is born. The estimates suggest that the association between girls and upgrade is positive or nonexistent for parents who

Table 6: Mother's Probability of Change in Partner Between Births

	Response variable:					
	1 = Child 2 with a new partner			1 = Child 3 with a new partner		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	0.007*	−0.007	0.002	−0.008	−0.036	−0.041
	(0.004)	(0.014)	(0.013)	(0.006)	(0.032)	(0.031)
All boys				−0.018***	−0.024	−0.021
Cohabitation				(0.006)	(0.033)	(0.033)
All girl(s) ×		0.014	0.003		0.028	0.039
		(0.014)	(0.014)		(0.033)	(0.032)
All boys ×					0.007	0.009
Cohabitation					(0.034)	(0.033)
Controls?			Yes			Yes
Observations	34,455	34,293	32,986	16,256	16,207	15,750
R^2	0.0001	0.115	0.193	0.001	0.038	0.095

Notes: OLS estimates with robust standard errors given in parenthesis. Response variables are equal to one if father's id is not equal at birth k as it was at last birth, $k-1$. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{*}p<0.1$; $^{***}p<0.01$.

are living apart at parity 1, and absent with regards to marriage.

Table 7: Father's Probability of Change in Partner Between Births

	Response variable:					
	1 = Child 2 with a new partner			1 = Child 3 with a new partner		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	−0.003	−0.023	−0.020	0.0003	−0.009	−0.026
	(0.005)	(0.015)	(0.014)	(0.009)	(0.039)	(0.031)
All boys				−0.013	−0.009	−0.025
				(0.008)	(0.039)	(0.033)
All girl(s) × Cohabitation		0.023	0.020		0.010	0.031
		(0.016)	(0.015)		(0.040)	(0.032)
All boys × Cohabitation					−0.001	0.016
					(0.040)	(0.033)
Controls?			Yes			Yes
Observations	25,166	25,008	24,052	10,526	10,477	10,186
R^2	0.00001	0.089	0.175	0.0003	0.062	0.118

Notes: OLS estimates with robust standard errors given in parenthesis. Response variables are equal to one if mother's id is not equal at birth k as it was at last birth, $k-1$. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are father's and mother's age and age squared, dummies for cohabitation with mother of last child, father's citizenship at last birth equal to 1 if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child.

Table 8: Likelihood of Partnership Upgrades between Child One and Two

	Response variable:			
	1 = Upgrade to cohabitation		1 = Upgrade to matrimony	
	(1)	(2)	(3)	(4)
Girl	0.002	0.008	0.0002	0.001
	(0.017)	(0.017)	(0.006)	(0.006)
Controls?		Yes		Yes
Observations	1,942	1,861	8,996	8,688
R^2	0.00001	0.020	0.00000	0.021

Notes: OLS estimates with robust standard errors given in parenthesis. Samples are restricted to parents who have their respective first two children together and their first child was a singleton and survived its first year. Furthermore all parents in regression (1) and (2) were not in cohabitation at first child, and not married in regression (3) and (4). Control variables included are mother's and father's age and age squared,borth parents citizenship at last birth equal to 1 Icelandic and 0 otherwise and health department at last birth equal equal to 1 if living in capital area, quarter and rounded year of birth of last child.

4 Discussion

Taken together our evidence suggests gender is to some extent associated with parents' fertility and partnerships decisions in Iceland. While some association was expected, the difference between the cohabiting and single parents (representing about a fifth and a tenth at parity one and two respectively) is surprising.

The difference might be indicative of gender preferences, at least for fathers. At parity three, single (not cohabiting) fathers are estimated more likely to stop their fertility if they have only boys. On the other hand, cohabiting fathers of three boys are more likely to continue to a fourth birth. This contrast could be due to the mother's influence. Further evidence supporting father's higher demand for sons is the fact that single fathers that have at least three children have the first child considerably faster after *all girls* if they are single than if cohabiting when *same-sex* seems to speed up next birth independently of gender. Consistent with that story is that mothers of two and three children are significantly less likely to have changed partners at next observation (birth of her next child) if she has *all boys*. At the same time, mothers are estimated more likely to have switched partners between child one and two if her first child was a girl.

The general view in the literature is that association of boys with marital stability and decreased fertility are indicative of boys preferences, this view is for example expressed in Andersson et al. (2006) and Dahl and Moretti (2008). However this has not been shown to be an absolute fact and other fitting explanations are not unimaginable Raley and Bianchi (2006) for instance suggest that "parents with two girls may be more likely to have an additional child not because they desire a son but because they so enjoy their girl children that they desire another child." (p. 404).

Generally, gender association with fertility and partnership in Iceland appears to be weaker than what has been found in US data. At parity two the likelihood of progression to parity three is only estimated to be two to five percentage points higher for parents of same-sex siblings, instead of seven reported by Angrist and Evans (1998). Results on partnership show increased stability for mothers of *all boys* at parity two and three and decrease at parity one for mothers of girls. Although not conclusive by itself, a strong prior derived from findings in Sweden, Germany and the United

States makes it likelier than not that boys, somehow, stabilize relationships in Iceland too. However, such interpretation rests on the assumption that nature distributes gender randomly at birth.

A barely acknowledged possibility in the literature so far are biological factors. It could be that the main explanation is that the parents of boys are, on average, marginally different from other parents. According to the Trivers-Willard hypothesis and some, so far, limited empirical evidence suggest they are more likely to be healthier and have higher socioeconomic status which in turn could also cement relationships. One of Dahl and Moretti (2008) impressive findings was that having an ultrasound check reduces the probability of marriage before delivery if the baby is a girl. However, ultrasound checks around the year 1990 seem unlikely to be uncorrelated with mother's characteristics. In our own data we have information on number of mother's medical examinations during pregnancy, which correlates with a number characteristics of the parents and the children, including their gender ($p < 0.1$).

While it is not obvious that comparing average marginal effects in two groups is the most informative type of modeling technique to further understand preferences when responses are likely heterogeneous, there are already some signs of economically significant effects of girls being more likely to grow up relatively absent from their fathers.

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Appendix 1: Support for Data and Methods

Table 9: Means and Standard Deviation of Main Response Variables

Statistic	N	Mean	St. Dev.	Min	Max
Mothers:					
Total number of children	53,808	2.147	0.972	1	10
Proceeds to birth 2 (at parity 1)	53,808	0.711	0.453	0	1
Proceeds to birth 3 (at parity 2)	37,802	0.465	0.499	0	1
Proceeds to birth 4 (at parity 3)	17,307	0.224	0.417	0	1
Changes partners between parity 1 & 2	34,455	0.177	0.382	0	1
Changes partners between parity 2 & 3	16,256	0.113	0.317	0	1
Days time between birth 1 and 2	38,294	1,665	1,070	267	10,311
Days time between birth 2 and 3	17,602	1,747	1,035	262	7,985
Fathers:					
Total number of children	39,073	2.010	0.967	1	9
Proceeds to birth 2 (at parity 1)	39,073	0.643	0.479	0	1
Proceeds to birth 3 (at parity 2)	24,757	0.425	0.494	0	1
Proceeds to birth 4 (at parity 3)	10,364	0.236	0.425	0	1
Changes partners between parity 1 & 2	25,166	0.183	0.387	0	1
Changes partners between parity 2 & 3	10,526	0.141	0.348	0	1
Days time between birth 1 and 2	25,166	1,647	1,096	282	10,158
Days between birth 2 and 3	10,526	1,719	1,038	283	8,587
Couples:					
Cohab. upgrade between birth 1 & and 2	1,943	0.831	0.375	0	1
Married between birth 1 & and 2	8,997	0.101	0.301	0	1

Notes: Samples restrictions are equal to those in corresponding regressions.

Table 10: Porportions of Boys by Month

	Response variable:
	1 = Boy
February	0.007 (0.007)
March	0.002 (0.006)
April	0.007 (0.006)
May	0.014** (0.006)
June	0.016** (0.006)
July	0.012* (0.006)
August	0.008 (0.006)
September	0.010 (0.006)
October	0.004 (0.006)
November	0.017** (0.007)
December	0.007 (0.007)
Constant	0.503 (0.005)
Observations	144,742

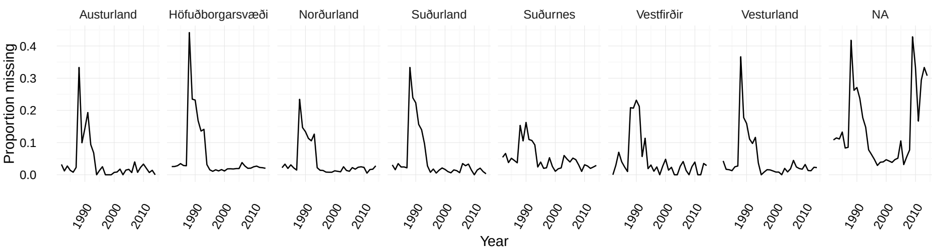
Notes: Based on birth registry full sample except for 338 observations without identification of gender. $^{*}p<0.1$; $^{**}p<0.05$

Table 11: Porportions of Boys and LBW of Previous Sibling

	Response variable:			
	1 = Boy			
	(1)	(2)	(3)	(4)
Low birth weight	−0.011	−0.012	−0.017	−0.016
	(0.030)	(0.017)	(0.011)	(0.011)
Mother's age	−0.023	−0.001	0.0002	
	(0.014)	(0.006)	(0.004)	
(Mother's age)^2	0.0004	0.00001	−0.00001	
	(0.0002)	(0.0001)	(0.0001)	
Constant	0.836	0.536	0.513	0.512
	(0.207)	(0.091)	(0.059)	(0.002)
Observations	4,182	25,096	70,300	70,300

Notes: OLS estimates with robust standard errors given in parenthesis. LBW defined as < 2.5 kg at birth. Column (1) sample includes only children born less than 1.5 year after its older sibling, column (2) is restricted to less than 3 years, column (3) has no restrictions. Total number of children with older sibling with low birth weight was 2210. Regressions are thus necessarily low-powered.

Figure 2: Proportion of *father's id* Missing by Year and Health Department



Appendix 2: Additional Regression Results

Table 12: Mother's Who Reach Parity Four: Partner Change Risk and Time Spacing

	Response variable:	
	1 = Child 4 with a new partner	Days between birth 3 and 4
All girls	0.0003	−71.668
	(0.017)	(52.248)
All boys	−0.032**	39.413
	(0.013)	(45.885)
Constant	0.120	1,617.408
	(0.006)	(20.024)
Observations	3,635	3,893
R^2	0.001	0.001

Notes: OLS estimates with robust standard errors given in parenthesis. All girls/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. No controls included nor estimated for fathers due to small sample size. $^{**}p<0.05$.

Table 13: Mother's First Children and the Probability and Expected Count of Subsequent Births. (Mother's sample 2)

	Response variable:							
	1 = birth two takes place		1 = birth three takes place		1 = birth four takes place			Total number of children
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All girl(s)	0.017	0.013	0.054**	0.046**	0.016	−0.025	0.025	0.015
	(0.010)	(0.009)	(0.023)	(0.022)	(0.044)	(0.042)	(0.021)	(0.018)
All boys			−0.002	0.015	−0.036	−0.020		
			(0.023)	(0.021)	(0.044)	(0.042)		
All girl(s) × Cohabitation	−0.020*	−0.017	−0.037	−0.019	0.005	0.049	−0.036	−0.026
	(0.012)	(0.010)	(0.025)	(0.023)	(0.046)	(0.043)	(0.024)	(0.020)
All boys × Cohabitation			0.043*	0.032	0.084*	0.067		
			(0.024)	(0.022)	(0.046)	(0.043)		
Controls?		Yes		Yes		Yes		Yes
Observations	40,352	38,402	27,365	26,465	12,064	11,809	40,352	38,402
R^2	0.010	0.233	0.001	0.186	0.006	0.138	0.003	0.303

Notes: OLS estimates with robust standard errors given in parenthesis. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Regressions (7) and (8) are estimated at parity one. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to 1 if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{*}p<0.1$; $^{**}p<0.05$

Table 14: Mother's First Children and Days Until Next Birth (Mother's sample 2)

	Response variable:					
	Days between birth 1 and 2			Days between birth 2 and 3		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	6.965	−39.146	−23.894	−87.890***	−109.854	−67.347
	(12.521)	(33.448)	(31.785)	(22.840)	(75.283)	(71.143)
All boys				−95.813***	−67.258	−36.721
				(21.701)	(76.874)	(73.432)
All girl(s) × Cohabitation		56.753	38.507		23.664	3.108
		(35.940)	(34.178)		(79.015)	(74.722)
All boys × Cohabitation					−30.485	−29.841
					(80.142)	(76.550)
Controls?			Yes			Yes
Observations	28,033	27,867	26,857	12,370	12,323	12,004
R^2	0.00001	0.022	0.141	0.002	0.002	0.121

Notes: OLS estimates with robust standard errors given in parenthesis. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{***}p<0.01$.

Table 15: Mother's Probability of Change in Partner Between Births (Mother's sample 2)

	Response variable:					
	1 = Child 2 with a new partner			1 = Child 3 with a new partner		
	(1)	(2)	(3)	(4)	(5)	(6)
All girl(s)	0.008*	−0.005	0.004	−0.008	−0.051	−0.049
	(0.005)	(0.015)	(0.014)	(0.008)	(0.034)	(0.033)
All boys				−0.021***	−0.058*	−0.054
				(0.007)	(0.035)	(0.034)
All girl(s) × Cohabitation		0.013	0.005		0.045	0.048
		(0.016)	(0.015)		(0.035)	(0.034)
All boys × Cohabitation					0.042	0.044
					(0.036)	(0.035)
Controls?			Yes			Yes
Observations	26,043	25,883	24,985	11,911	11,864	11,569
R^2	0.0001	0.105	0.209	0.001	0.042	0.113

Notes: OLS estimates with robust standard errors given in parenthesis. Response variables are equal to one if father's id is not equal at birth k as it was at last birth, $k-1$. All girl(s)/boys means all children up to that point have been girl(s)/boys. Samples are restricted to mothers whose first children (explanatory variable) were singletons and have survived their first year. Control variables included are mother's age and age squared, dummies for cohabitation with father of last child, mother's citizenship at last birth equal to if Icelandic and 0 otherwise, health department at last birth equal to 1 if living in capital area, quarter and rounded year of birth of last child. $^{*}p<0.1$; $^{***}p<0.01$.