The effect of childhood family size on fertility in adulthood. New evidence from IV estimation*

Sara Cools† Rannveig Kaldager Hart‡

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Abstract

While fertility is positively correlated across generations, the causal effect of children's experience with larger sibships on their own fertility in adulthood is poorly understood. Using the sex composition of the two firstborn children as an instrumental variable, we estimate the effect of sibship size on adult fertility. Estimations are performed using high-quality data from Norwegian administrative registers. Our study sample is all first- or second-borns during the 1960s in Norwegian families with at least two children (approximately 110 000 men and 104 000 women). An additional sibling has a positive effect on male fertility, mainly causing them to have three children themselves, while there is a negative effect for women at the same margin. Investigation into mediators reveals that mothers of girls shift relatively less time from market to family work when an additional child is born. We speculate that this scarcity in parents' time makes girls aware of the strains of life in large families, leading them to limit their own number of children in adulthood.

Keywords: Fertility, Intergenerational transmission, Instrumental Variables, Family Size

JEL codes: C26, J13, J22

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†Institute for Social Research. Email: sara.cools@samfunnsforskning.no.

‡Department of Sociology and Human Geography, University of Oslo and Research Department, Statistics Norway. Email: r.k.hart@sosiologi.uio.no

1 Introduction

Most important life outcomes – such as health, education, and income – are positively correlated across generations. This positive relationship is in part due to potential causation from one generation's achievements in these fields to that of the next generation, and in part due to the shared genetic and social circumstances of parents and children. Studies of the intergenerational correlation in fertility across the developed world consistently find that children also tend to replicate their parents' family size (Murphy 2013).

The fact that this relationship resists the inclusion of detailed controls for socioeconomic status, suggests that the transmission of fertility across generations is not merely a by-product of shared social circumstances (Kolk 2014). The remaining correlation, however, is still somewhat of a black box. The fact that parents' preferences for their children's fertility behavior is positively correlated with the next generation's preferences (Axinn et al. 1994; Starrels and Holm 2000) and behavior (Barber 2000), indicates some transmission of "family culture" across generations. However, the origin of this "family culture" could lie in shared environment (beyond socioeconomic position) as well as in genetic similarity (Kohler et al. 1999). Despite demographers' interest in intergenerational transmission of fertility, surprisingly little attention has been devoted to netting out the similarity across generations and obtaining causal estimates of the effect of an additional sibling on own fertility in adulthood.

In this paper, we seek to answer two closely related research questions: First, what is the causal effect of having an additional sibling on fertility in adulthood? Second, which social mechanisms would mediate such an effect?

Answering these questions requires estimates that can plausibly be interpreted as capturing a causal effect. We therefore use the same sex instrumental variable, which exploits the demographic finding that having children of the same sex increases the probability of further childbearing (Andersson et al. 2006; Ben-Porath and Welch

1976; Gini 1951). This increase in sibship size is uncorrelated with all background factors of parents, such as their (initial) preference for family size (Angrist and Evans 1998).

To the best of our knowledge, the only other study applying IV estimation to this end is Kolk (2015), which finds no clear effects of sibship size on adult fertility using the birth of younger twin siblings as instrumental variable. We use the twin instrument in a robustness check in this paper—though we believe it has some invalidating features when applied to the intergenerational transmission of fertility. By applying what we consider a more credible instrument, and by supplying extensive robustness checks and a thorough investigation into the mechanisms explaining our findings, our paper offers a significant contribution to the causal understanding of fertility transmission across generations.

Using data from Norwegian administrative registers, we study the fertility behavior of Norwegian men and women born in the 1960s. We estimate effects on completed fertility (measured at age 43), as well as on the likelihood of making specific parity transitions. The 1960s cohort balances the need for full background information with that of observing completed fertility (see Section 4.1). In this cohort, the modal number of siblings is one, closely followed by two (Rønsen 2004, p. 276). Our estimates thus capture the effect of growing up in a typical larger family, relative to a typical smaller family in a dual-earner society. With high female labor force participation, but rudimentary public support for working mothers, Norway in the 1960s has striking similarities with today's lowest-low fertility context in Southern and Central Europe (McDonald 2000b).

Our main finding is that the increase in sibship size increases men's fertility in adulthood, while it decreases women's fertility. For both, the effect is concentrated in the decision to have a third child. In order to understand the mechanisms behind these heterogeneous effects, we investigate the role of several potential mediators. Most importantly, we find that the additional sibling causes mothers (in the family of origin) to reduce their labor supply significantly more in the male than in the female sample. As a potential consequence, girls who grow up with more than one sibling are more familiar with the work-load and time-squeeze associated with having more children.

Although sibling sex composition is a much used instrumental variable for sibship size, the possibility that sex composition affects more than just sibship size is of particular concern when the outcome considered is fertility in the next generation. We test extensively for such *direct effects* of sibling sex on all outcomes considered in this paper, and we find no evidence for them. Our results are also corroborated by similar (though insignificant) point estimates when applying the twin instrument.

Our findings cast new light on fertility contagion. Fertility contagion is commonly thought of as an effect multiplier – magnifying small changes in the cost of child-bearing into large fertility responses. Our findings, on the other hand, suggest that large families in one generation may also cause *lower* fertility in the next generation. Whether contagion is positive or negative depends on the children's experience of life in larger families. Policies that make life in large families less straining, particularly for women, may hinder negative contagion across generations and hence contribute to maintaining high birth rates.

The remainder of the paper is organized as follows. Section 2 discusses causal mechanisms that could link an additional sibling to fertility in adulthood. Section 3 presents the instrument, and Section 4 describes data and variables. The main results are presented in Section 5, and robustness checks in Section 6. Section 7 shows effects on potential mediators, and Section 8 concludes.

2 Sibship size and fertility in the next generation

The birth of an additional sibling influences the time and money available to each child – and likely also the preferences and beliefs about life in large families. Moving from a sibship of two to three increases the workload at home, often pushing a household's established work-family balance in the direction of family life. Angrist and Evans (1998) find a 5.3% reduction in US families' total income as a result of the increase in sibship size, and similar findings have been documented in other countries (see e.g. Cruces and Galiani (2007) for Latin America, Daouli et al. (2009) for Greece, Hirvonen (2009) for Sweden, and Cools (2013) for Norway). In Section 7.1, we estimate the effect of sibship size on parents' labor supply.

Similarly, parents may shift time from (pure) leisure, such as time for hobbies and friends, to child rearing upon the birth of an additional child. To the extent that parents of larger sibships place relatively more weight on family life, the value of family life as perceived by older siblings in the household may increase with additional siblings. This could lead to more family-oriented behavior in the next generation, which could be reflected in higher probabilities of marrying and having a first child, more stable marriages and larger families. To test this hypothesis, we use union stability (of both the index generation and their parents) as a proxy for being family-oriented (Section 7.2).

Through adaptive preferences, a third child may give parents a preference for three-child families (see e.g. Hayford (2009) for a more general example). Furthermore, parents' preferences for their children's fertility behavior is positively correlated both with their children's fertility preferences (Axinn et al. 1994; Starrels and Holm 2000) and their fertility behavior (Barber 2000). In a similar vein, the *imitation hypothesis* suggests that children model their fertility behavior upon that of their parents, so that those who grow up with two siblings would be disproportionately more likely to have a completed family size of three (Starrels and Holm 2000).

As information about the consequences of childbearing is imperfect, beliefs about these consequences may significantly influence fertility behavior (Bernardi and Klaerner 2014). Individuals who grow up with an additional sibling may be more familiar with the strains of raising a relatively large family: Children in larger families on average receive less care and attention from their parents, and spend more time taking care of their (younger) siblings (Evertsson 2006). Presumably, such experiences are more pronounced for women than for men: Girls increase their time spent on housework more than boys do when an additional sibling is born (Evertsson 2006; Gager et al. 1999), and are thus made more directly aware of the work required to raise a relatively large family. Additionally, the increase in sibship size mainly impedes women's careers (Cools 2013; Hardoy and Schøne 2008). To the extent that children use the parent of their own sex as a role model, awareness that a large family may limit career opportunities may lead women to limit their family size. Hence, girls from larger sibships may be more aware of the adverse consequences of larger families, both relative to boys from families of the same size, and relative to children with fewer siblings. We describe differences in the housework efforts of teenage boys and girls in Section 7.1 to better understand if they contribute to any effects we see in our sample.

An additional sibling may also affect fertility in adulthood through other causal channels. If less time and money available per child translates into lower human capital investment, or into fewer direct transfers from parents to their adult children, there is scope for a negative effect on fertility in the next generation due to lower overall income. On the other hand, lower human capital implies lower alternative cost in caring for children, which, all else equal, suggests increased fertility in the next generation. However, Waynforth (2011) finds no significant correlation between fertility behavior and economic support from (grand)parents, and empirical studies systematically fail to find a deterring effect on human capital from sibship size at this margin (Black et al. 2005; Mogstad and Wiswall 2009). In Section 7.3, we test for effects of sibship size on educational attainment, assessing if this is a potential mediator of effects on family size.

3 Sibling sex: IV properties and direct effects

As an estimate of the effect of sibship size on fertility in the next generation, the intergenerational correlation in fertility is likely to be severely biased, due to the shared biological, social and economic circumstances of parents and children. Hence, estimating the *effect* of an additional sibling on fertility in the next generation requires a different empirical strategy. We use whether the two firstborn children in the family of origin are of the same sex as an instrumental variable for sibship size. An extensive demographic literature has shown that when the two firstborn children are of the same sex, parents are more likely to have a third child (Andersson et al. 2006; Hank 2007; Kippen et al. 2007). As children's sex composition is uncorrelated with background characteristics of parents (such as fertility preferences), the *same sex* instrument is a much used instrumental variable for sibship or family size (see for example Angrist and Evans (1998); Black et al. (2010); De Haan (2010)).

The validity of using siblings' sex composition as an instrumental variable in our setting hinges on the sex composition having no effects of its own on fertility choices made in adulthood, i.e., that there are no "direct effects". Some studies suggest that family support structures affect fertility decisions (Aassve et al. 2012), and individuals who have a sister will on average receive less practical help from their parents in adulthood, but more help from their sibling (Goodsell et al. 2015; Spitze and Trent 2006). While some qualitative studies suggest that having a sister in itself increases fertility (Bernardi 2003), quantitative studies suggests that brothers influence fertility timing slightly more than sisters (Lyngstad and Prskawetz 2010). As fertility timing is more easily influenced than completed fertility by context (Gauthier 2007), the observed correlations in timing need not imply that siblings affect each other's completed fertility. Importantly, in all these studies, the effect may be channeled exactly through sibship size – in which case it does not pose a problem to our identification strategy.

In order to estimate only the direct effect of sibship sex composition on fertility in adulthood, i.e., net of effects running through sibship size, one needs to look at a situation or sample where sex composition does not influence sibship size (as is done for instance by Angrist et al. (2010); Peter et al. (2014)). In Section 6 we provide an empirical investigation of direct effects of sibling sex composition, utilizing the fact that the third child's sex does not affect the probability of further childbearing in families where the two firstborn children are of opposite sex. We find no evidence of direct effects of sibling sex on any of the outcomes considered in this paper. The empirical distribution across outcomes and sibship size, conditional on instrument status, satisfy the testable implication for instrument validity in Kitagawa (2015). Taken together, these findings inspire confidence in the validity of our IV estimates.

4 Data and study sample

4.1 Study sample

Our point of departure is data from Norwegian administrative registers on all Norwegian residents. Personal identifiers link individuals to their parents and children. For registering to be as complete as possible, we restrict mothers of index persons to be born no earlier than 1935. The need for reliable data on both family background and on own completed fertility makes individuals born during the 1960s particularly suitable.

As the sex composition instrumental variable is defined only for families with at least two children, our sample is limited to families whose first two children were both born between 1960 and 1969. We further exclude families in which the first two children do not share both parents, or where either parent is unknown to the registers. The mechanisms we explore in Section 7 may play out differently in intact and non-intact families. By this restriction, we exclude families that were "complex" before the realization of the instrument. The restriction also strengthens the first stage, improving the precision of the 2SLS estimates. The study sample does not

Table 1: Mean values in family background variables by sibling sex composition

	Same sex		Different sex		Difference	
	Mean	SD	Mean	SD	Est.	SE
Distance two first children (years)	2.45	(1.31)	2.46	(1.33)	-0.01	(0.01)
Mother's						
- year of birth	1941.47	(3.45)	1941.48	(3.47)	-0.01	(0.02)
- age at first birth	22.13	(2.81)	22.16	(2.84)	-0.03†	(0.02)
Father's						
- year of birth	1937.99	(4.95)	1938.02	(4.96)	-0.04	(0.03)
- age at first birth	25.62	(4.38)	25.62	(4.39)	-0.00	(0.03)
N	53431		53813		107244	

Note: The samples are all couples with at least two children, where the two first children are both born in Norway in the period 1960-1969 and are registered with the same mother and father. For the means, standard deviations are reported in parentheses, for the estimated differences, standard errors are in parentheses. † $p < 0.10$.

include individuals who are themselves twins, but they may have twin siblings. Results are not sensitive to these further restrictions.

We study the fertility outcomes in adulthood of both first- and second-borns in the families included in our sample, and estimations are performed separately by index person's sex. In order for the same sex instrument to be internally valid also for the second-borns, sample entry cannot be influenced by the firstborn's sex, and, in addition, there can be no systematic differences within the sample between parents who have firstborns of different sex. We find no effect of firstborn's sex on the probability of entering our sample, nor do we find any statistically significant difference within the sample in parents' background characteristics according to firstborn's sex (results available upon request).

4.2 Family background characteristics

Since the individuals under study are born during the 1960s, background characteristics that are exogenous to the instrument must be observed further back than is recorded in most of the important Norwegian registers. Parents' income could be observed from 1967 onwards, and their education from 1970 onwards, both too late for our purpose. The only available background variables that are realized prior to the instrument for the whole sample are parents' year of birth, their age at first birth and the distance (in years) between the births of the first two children.

The means of these variables are reported in Table 1. We have split the sample into families with two children of the same sex (first column) and of different sex (second column). The last column in Table 1 reports simple t-tests of whether the background characteristics vary with the sex composition of the first children.

When included as controls, background variables enter as a set of dummy variables capturing the distance in years between the birth of the first and the second sibling (censored at six years), and dummies for parents' age at first birth (by age brackets of five years each). All models include birth year and birth order fixed effects (FE), in the form of a set of dummy variables for birth year and birth order. The full set of dummy variables to be used as controls throughout the paper, in addition to t-tests of the difference by instrument status, is given in Supplementary Material (Table S.1). Systematic differences in means by instrument status would indicate that the instrument is not randomly assigned. Some of the estimated differences according to same sex sibship are statistically significant, but they are not significant in size.

4.3 Fertility outcome variables

The main outcome variable considered in this paper is the total number of children registered to the individual at the age of 43. The whole sample can be followed until age 43, after which we lose more than 10% of the original sample with each yearly increment in age. We therefore present completed fertility at 43 in our main results. As a robustness check, we have estimated effects on fertility up to age 45 (Supplementary Material Figure S.1). The point estimates are in line with the main findings, though precision decreases with sample size. It does not matter for the difference in estimates across sex whether we use the same age of observation for men and women, or whether we move the female (male) estimates down (up) by a

Table 2: Mean values in outcome variables, by index person's sex

	Men		Women	
	Mean	SD	Mean	SD
N. children at 43	1.75	(1.24)	2.03	(1.14)
Has children at 43	0.78	(0.42)	0.87	(0.33)
Has >1 child at 43	0.63	(0.48)	0.74	(0.44)
Has >2 children at 43	0.27	(0.44)	0.33	(0.47)
Has >3 children at 43	0.06	(0.24)	0.07	(0.26)
N	110225		103760	

Note: The samples consist of all first- and second-born men and women born in Norway between 1960 and 1969 in families with at least two children, where the two first children are registered with the same mother and father. Standard deviations in parentheses.

few years.

We also evaluate parity specific outcomes by considering separately the probability of having more than zero, one, two and three children at this age. Descriptive statistics for these outcomes are given in Table 2.

4.4 Additional outcome variables

In the investigation into mechanisms (Section 7), we study three sets of additional outcome variables: Parents' labor supply, the union stability of index persons and their parents, and the educational attainment of index persons.

The descriptive statistics for these outcomes are given in Table 8 in the Appendix. Information on yearly personal income (consisting of wages, pensions and entrepreneurial income) goes back to 1967 and covers the population residing in Norway each year.

Labor supply at the extensive margin (employment) is defined as having a yearly income above 1 BA. We proxy labor supply at the intensive margin (i.e., working hours conditional on employment) by taking the log of income above 1 BA. This is a valid approximation in our case if a third child has no or only a negligible effect on hourly wages, as documented in Cools (2013). As we find no effects on fathers' labor supply at either margin, these results are not displayed.

Education data come from Statistics Norway's education registers, which record all changes (and their dates) in individuals' highest educational attainment from 1970 onwards. Finally, we have data on parents' marital status from 1992 onwards, that is, from when the youngest individuals in our sample are aged 23 and the oldest 32 years. Parents' marital status when the second child is aged 28 (when about half the sample can be observed) therefore serves as a proxy for their marital status when the children still live at home.

The same registers are used for the index persons' own marital status in adulthood (married, divorced or neither). In order to capture family stability regardless of marital status (more than half of first births to Norwegian coresidental couples are currently to cohabiters), we estimate the effect of an additional sibling on the probability of having (at least) two children with the same partner. Having two children is itself an outcome affected by sibship size, hence we present an unconditional measure, which is equal to one if the individual, at age 43, has at least two children and these two share both parents, and zero otherwise.

5 Effects on fertility in adulthood

IV estimation is performed in two steps, using two stage least squares (2SLS) regression. We first estimate the effect of sibship sex composition on sibship size – captured by the probability of having more than one sibling – giving the *first stage* estimates. IV estimates are then obtained by regressing the index persons' fertility in adulthood on the part of the variation in the sibship size tied to sex composition. The IV estimate captures the average treatment effect among those moved by the instrument, that is, those parents who will have a third child if and only if their two first children are of the same sex (Imbens and Angrist 1994). As the instrumental variable always takes the same value for siblings, treating siblings as independent observations would underestimate standard errors. We avoid this by clustering at the family of origin. We also present *reduced form* estimates of the effect of sibship

Table 3: Effects of sibling sex composition and sibship size on fertility in adulthood

	First stage		*Red. form*		*IV estimate*		*OLS estimate*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
MEN	>1 sibling		N. of children		N. of children		N. of children	
	OLS	OLS	OLS	OLS	2SLS	2SLS	OLS	OLS
Same sex	0.059**	0.057**	0.015*	0.015†				
	(0.003)	(0.003)	(0.008)	(0.008)				
>1 sibling					0.256*	0.258†	0.138**	0.123**
					(0.130)	(0.134)	(0.008)	(0.008)
Birth year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes	No	Yes	No	Yes
Adj. R2	0.018	0.106	0.002	0.005	0.003	0.005	0.005	0.007
N	110225	110225	110225	110225	110225	110225	110225	110225
WOMEN	>1 sibling		N. of children		N. of children		N. of children	
	OLS	OLS	OLS	OLS	2SLS	2SLS	OLS	OLS
Same sex	0.063**	0.061**	-0.013†	-0.014†				
	(0.004)	(0.003)	(0.007)	(0.007)				
>1 sibling					-0.210†	-0.233†	0.205**	0.183**
					(0.119)	(0.121)	(0.008)	(0.008)
Birth year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes	No	Yes	No	Yes
Adj. R2	0.019	0.108	0.002	0.006	.	.	0.009	0.011
N	103760	103760	103760	103760	103760	103760	103760	103760

Note: The sample is first- and second-borns in Norwegian families with at least two children (where the two first children are registered with the same mother and father), who are born between 1960 and 1969. Standard errors (in parentheses) are clustered at the family of origin. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

sex composition on fertility in adulthood (by OLS). The reduced form estimate gives the impact of sibling sex on the outcome in question, making no assumption that the effect is channeled through sibship size. Lastly, the intergenerational correlation in fertility is obtained by OLS regression of fertility in adulthood on sibship size.

5.1 Main results

The main results of this paper are presented in Table 3. Columns (1) and (2) give first stage estimates, columns (3) and (4) the reduced form estimates, and columns (5) and (6) the IV estimates. The OLS estimates of the intergenerational correlation in fertility are given in columns (7) and (8). The upper panel gives estimation results for men, the lower for women. All the specifications in Table 3 include birth year and birth order fixed effects. The even-numbered columns also include a set of exogenous control variables (see Section 4 and in Appendix Table S.1): Parents' year of birth, their age at first birth and the distance in age between the first two siblings.

Columns (1) and (2) in Table 3 give the OLS estimates of how being in a same sex sibship affects the likelihood that individuals in our sample will have an additional sibling. These first stage estimates are slightly larger for women than for men, but they are all very close to 6 percentage points, and comparable in size to other applications of this instrument. With t-statistics above 20, they satisfy the criterion of instrument relevance.

Columns (3) and (4) give the OLS estimates of how being in a same sex sibship affects individuals' own number of children when they are 43 years old. Having a brother causes the men in our sample to have 0.015 more children on average ($p < 0.05$). On the other hand, having a sister causes the women in our sample to have 0.014 *fewer* children on average ($p < 0.10$). The effect on sibship size (columns (1) and (2)) is likely to play a major role in the estimated effect of same sex sibship on individuals' own fertility. Under the assumption that it is in fact the *only* causal channel from sex mix to fertility in adulthood (i.e., the *exclusion restriction* for instrument validity), the 2SLS estimates in columns (5) and (6) are consistent estimates of the causal effect of sibship size on individuals' total number of children at age 43. According to these estimates, having an additional (a second) sibling as a child causes men to have 0.26 more children and women to have 0.23 fewer children on average in adulthood.

Consistent with previous research, the intergenerational correlations shown in columns (7) and (8) are positive, and slightly stronger for women than for men. Compared to the causal effects documented in columns (5) and (6), the OLS estimates are thus substantially more positive for women, while they are slightly less positive for men. This suggests that the unobserved variables netted out in our IV analysis – such as shared environment, preferences and genetics – transmit fertility

Table 4: Effects of sibling sex composition and sibship size on different parity transitions in adulthood

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
MEN	>0 children		>1 child		>2 children		>3 children	
	OLS	2SLS	OLS	2SLS	OLS	2SLS	OLS	2SLS
Same sex	0.003		0.003		0.006*		0.002	
	(0.003)		(0.003)		(0.003)		(0.001)	
>1 sibling		0.056		0.059		0.109*		0.035
		(0.045)		(0.052)		(0.047)		(0.026)
Birth year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj. R2	0.005	0.002	0.003	0.003	0.003	0.005	0.003	0.004
N	110225	110225	110225	110225	110225	110225	110225	110225
WOMEN	>0 children		>1 child		>2 children		>3 children	
	OLS	2SLS	OLS	2SLS	OLS	2SLS	OLS	2SLS
Same sex	0.001		-0.002		-0.010**		-0.003†	
	(0.002)		(0.003)		(0.003)		(0.002)	
>1 sibling		0.014		-0.025		-0.156**		-0.045
		(0.034)		(0.046)		(0.051)		(0.027)
Birth year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj. R2	0.005	0.006	0.002	.	0.005	.	0.003	.
N	103760	103760	103760	103760	103760	103760	103760	103760

Note: The sample is first- and second-borns in Norwegian families with at least two children (where the two first children are registered with the same mother and father), who are born between 1960 and 1969. Standard errors (in parentheses) are clustered at the family of origin. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

across generations more strongly for women than for men.

We have further split the sample according to birth order. The estimates are statistically significant only in the samples of firstborn men and second-born women (Supplementary Material Table S.2).

5.2 Parity-specific effects

In order to know which fertility margins are affected by sibship size, we evaluate the effects of same sex sibship and sibship size on the likelihood of having more than 0, 1, 2 and 3 children. If the negative effects among women are mediated by belief formation, we expect women from larger sibships to avoid forming large families. The imitation hypothesis suggests that growing up with two siblings leads to a preference for three children, predicting particularly strong positive effects at parity three (Starrels and Holm 2000). If, instead, results were driven by transmission of a more general sense of being family-oriented, we would expect effects on all parities among men.

The reduced form (odd-numbered columns) and IV estimates (even-numbered columns) are given in Table 4. For men, the point estimates are positive at all margins, but only significant for the likelihood of having a third child. However, the parity specific estimates do not differ significantly. An additional sibling makes men 10.9 percentage points ($p < 0.05$) more likely on average to have three or more children. This indicates a male pattern of fertility imitation or adaptive preferences, where growing up with two siblings fosters a preference for a three-child family in adulthood.

For women, there are negative point estimates at all margins above the first child, and also here the only significant estimate regards the likelihood of having a third child, which decreases by 15.6 percentage points ($p < 0.01$) due to the increase in sibship size.

6 Internal and external validity

6.1 Direct effects of sibling sex

Although sibling sex composition is a much-used instrumental variable for sibship size, the potential existence of direct effects, which would compromise the instrument's internal validity, cannot be a priori dismissed. In order to assess the likelihood of bias in the IV estimates presented in Tables 3 and 4, we study how individuals' fertility decisions in adulthood are affected by sibling sex mix in the particular case where sibling sex mix does not affect sibship size. Among the families in our sample with at least three children, where the two first children are of opposite sex, parents are not, on average, influenced by the sex of the third child in their decision to have a fourth child. This sample can therefore be used to investigate direct effects of

Table 5: Testing for direct effects of sibling sex composition on fertility in adulthood

	First stage		*Dir. Effects*	
	(1)	(2)	(3)	(4)
MEN	>2 siblings		N. of children	
	OLS	OLS	OLS	OLS
Same sex	-0.001	0.000	-0.002	-0.002
	(0.005)	(0.005)	(0.014)	(0.014)
Birth year FE	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
Adj. R2	0.015	0.058	0.001	0.003
N	32273	32273	32273	32273
WOMEN	>2 siblings		N. of children	
	OLS	OLS	OLS	OLS
Same sex	0.001	-0.001	0.015	0.015
	(0.005)	(0.005)	(0.013)	(0.013)
Birth year FE	Yes	Yes	Yes	Yes
Birth order FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
Adj. R2	0.015	0.059	0.001	0.003
N	32274	32274	32274	32274

Note: The sample is first- and second-borns in Norwegian families with at least three children, where the two first children are of opposite sex, born between 1960 and 1969. Standard errors (in parentheses) are clustered at the family of origin.

sibship sex composition, rid of any effect going through sibship size.

Columns (1) and (2) in Table 5 show how a second sibling (i.e., the family's third born) being of the same sex as the index person affects parents' further childbearing in this sample. For both men and women, the effect is quite precisely estimated to be zero; the sex of the third child does not influence parents' propensity to have a fourth child. In columns (3) and (4), we estimate whether having a same sex second sibling impacts fertility at age 43. The estimates show no significant effect of having a sibling of the same sex on individuals' own fertility in adulthood, neither for men nor for women. The point estimates are small and go in the opposite direction of the reduced form estimates in Table 3, and we therefore find it unlikely that the IV estimates in Table 3 are severely biased. If anything, the bias indicated by the estimates in Table 5 would push the IV estimates towards zero.

We have also estimated direct effects for each parity transition, as in Table 4, and we find no evidence of direct effects for any of the outcomes, further strengthening

Table 6: First and second stage estimates using twin instrument

MEN	>1 sibling		N. of children	
	(1)	(2)	(3)	(4)
	OLS	OLS	2SLS	2SLS
Twin 2nd	0.363**	0.373**		
	(0.003)	(0.010)		
>1 sibling			0.082	0.085
			(0.206)	(0.201)
Birth year FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
R2	0.005	0.105	0.004	0.007
N	55195	55195	55195	55195
WOMEN	>1 sibling		N. of children	
	(1)	(2)	(3)	(4)
	OLS	OLS	2SLS	2SLS
Twin 2nd	0.366**	0.380**		
	(0.003)	(0.011)		
>1 sibling			-0.210	-0.178
			(0.195)	(0.187)
Birth year FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
R2	0.006	0.109	.	.
N	52049	52049	52049	52049

Note: The sample is firstborns in Norwegian families with at least two children (where the two first children are registered with the same mother and father), who are born between 1960 and 1969. Standard errors (in parentheses) are clustered at the family of origin. ** $p < 0.01$.

the case for internal validity (available upon request).

6.2 Alternative IV: Twins

There is also the possibility of using twinning as an alternative instrument for family size (Angrist and Evans 1998; Black et al. 2005). The twin instrument captures the effect of an unintended third birth, with zero spacing to the second birth, and it might therefore differ from the one captured by the same sex IV.

In Table 6, we show the estimates of the effect of an additional sibling on fertility in adulthood using the twin IV. The first stage is given in columns (1) and (2), and the IV estimates in columns (3) and (4). Again, the estimated effect of sibship size on fertility is positive for men and negative for women, comparable in size to the same sex IV estimates, but not statistically significant at conventional levels. The estimates are also comparable to what Kolk (2015) finds when applying the twin instrument to Swedish data.

Applying the twin instrument to fertility outcomes raises some concerns regarding both internal and external validity. For women, the genetic heritability of (monozygotic) twinning could bias the estimates in either direction: Women whose mothers had twins may have more children due to a twin birth, or fewer children if their heightened risk of twinning keeps them from having additional children. As twinning is a shock not only to the number of siblings, but also to the spacing between them, growing up with younger twin siblings is possibly a different experience from growing up with two younger singleton siblings. Using exogenous variation in spacing from miscarriages, Buckles and Munnich (2012) show that spacing in its own right affects outcomes of both children and parents. While the validity of the same sex instrument is corroborated in our setting by tests for direct effects, we do not have similar tests for the twin instrument.

6.3 External validity

The estimates obtained using the same sex, or twin, instrument capture the effect of an additional sibling for individuals whose parents have a third child if and only if the two first children are of the same sex, or if and only if the second birth is a twin birth (Local Average Treatment Effect or LATE). The similarity of the effects as estimated by the two instruments indicates that the effects are not specific to increases in family size driven by twinning or preferences for sex mix. Starting then from the assumption that we seem to capture the general effect of sibship size on fertility for men and women in our index cohort, we discuss the relevance of our findings to other contexts and birth cohorts in Section 8.

7 Empirical tests of mechanisms

In order to gain insight into social mechanisms that could mediate the different effects of sibship size on men's and women's family formation in adulthood, thus answering our second research question, we investigate how several other outcomes (described in Section 4.4) are affected by sibship size. Any such mediator must itself be a causal effect of sibship size, hence we continue to present IV estimates to handle bias from unobservable confounders.

7.1 Parents' labor supply and children's housework

If the addition to the family reduces parents' total labor supply, this will result in lower family income – and in more time spent by at least one parent at home (see Section 2). The upper panel of Table 7 gives the estimates of how mothers' labor supply was affected by additional children during the childhood years of the index persons in our sample. (Fathers' labor supply is not moved by family size in our sample; results are available upon request.)

Labor supply is measured at the extensive margin (employment) and at the intensive margin (log earnings conditional on employment), with averages taken over the years when the second-born child is aged 6-10 and 11-15 years. The estimates are done by age of the second child, since this measure is defined for the whole sample. The third child, if born, will on average be about three years younger than the second child.

There is a substantial difference in how mothers' labor supply is affected by sibship size in the men's and in the women's sample. When the second child is 6-10 years old (and a third child, on average, 3-7 years old), the effect is a 17.6 percentage point reduction in mothers' employment in the men's sample ($p < 0.01$), while there is no reduction in mothers' employment in the women's sample. When the second child is aged 11-15, mothers' employment is reduced by 33 percentage points in the men's sample ($p < 0.01$), and there is still no evidence of an effect in the women's sample. The differences by gender are highly statistically significant. At the intensive margin, labor supply is significantly reduced when the second child is aged 11-15 in both samples, by about 38% in the men's sample, and 12% the

Table 7: Effects of sibship size on childhood circumstances, union stability and educational achievement

Outcome:	Men (1) IV est.	Women (2) IV est.	Diff (3) IV est.
Mothers' labor supply during childhood			
Employment, 2nd child aged 6-10	-0.176**	-0.008	-0.164*
	(0.057)	(0.056)	(0.064)
Employment, 2nd child aged 11-15	-0.330**	-0.015	-0.300**
	(0.052)	(0.050)	(0.058)
Log earnings, 2nd child aged 6-10	-0.053	-0.065	0.014
	(0.073)	(0.073)	(0.082)
Log earnings, 2nd child aged 11-15	-0.324**	-0.119*	-0.193**
	(0.063)	(0.059)	(0.068)
Parents' marital stability			
Parents married, 2nd child aged 28	0.122*	0.044	0.075
	(0.057)	(0.056)	(0.063)
Index person's union stability			
Married at age 43	0.111*	0.097†	0.012
	(0.054)	(0.052)	(0.071)
Divorced at age 43	0.032	-0.105**	0.130*
	(0.037)	(0.040)	(0.052)
$>$ 1 child same partner at 43	0.125*	0.011	0.108
	(0.054)	(0.050)	(0.069)
Index person's educational achievement			
Secondary educ. at age 19	0.061	-0.015	0.076
	(0.051)	(0.052)	(0.066)
Secondary educ. at age 43	0.009	-0.030	0.042
	(0.063)	(0.061)	(0.081)
Lower tert. educ. at age 43	0.071	0.012	0.058
	(0.062)	(0.062)	(0.080)
Higher tert. educ. at age 43	0.007	-0.020	0.026
	(0.039)	(0.030)	(0.046)

Note: In columns (1) and (2), each cell gives the 2SLS estimate of the effect of sibship size on the outcome given by the row heading. In column (3), each cell gives the corresponding 2SLS estimate of the difference in the effect of sibship size by index person's sex, estimated in the pooled sample. The samples are mothers (upper panel), parental couples (middle panel) and children (lower panel) in Norwegian families with at least two children, where the two first children are registered with the same mother and father and are born between 1960 and 1969. The number of observations for each estimate is given in the corresponding cells in Table S.3 and S.4. Standard errors (in parentheses) are clustered at the family of origin. * $p < 0.05$, ** $p < 0.01$.

women's sample. Again, the estimated difference in the pooled model is substantial and statistically significant ($p < 0.01$).

Tests equivalent to those presented in Table 5 show no evidence of a direct effect of sex mix on mothers' labor supply (Supplementary Material, S.3). Also, the conditional distributions across mothers' income and sibship size satisfy the requirement in Kitagawa (2015) (available upon request). A violation of instrument validity thus seems an unlikely explanation of these findings.

Rather, it seems likely that the effect of having a third child on mothers' labor supply is mediated by whether they have a daughter to help out at home. The largest differences by sex are found when the second child is 11-15 years old, and the oldest about 13-18 years, when both are old enough to pull some weight in household work if required.

A daughter in the household may enable mothers to work longer hours in paid work upon the birth of a third child. We have checked survey data on Norwegian teenagers' time use in our index cohorts, collected in 1980 (see Supplementary Material for more details). A simple OLS regression shows that girls spend 32 minutes more on housework per day than boys do ($p < 0.001$). The difference increases by about ten minutes with an additional sibling (not statistically significant). Due to the sample size (N = 415), we cannot instrument for family size, and the estimates do not have a causal interpretation.

Being more involved in housework, teenage girls may have been more aware of the work required to raise a family than their male peers were. Hence, a more negative effect for women than for men could be linked to the possibility that a second sibling was more of a learning experience for teenage girls than for teenage boys.

7.2 Union stability of parents and index persons

Previous research suggests that larger sibship may give a shift to stronger "family-orientedness", and hence mediate a positive effect of sibship size on fertility in the next generation (Section 2). We explore this mechanism by looking at how sibship size affects union stability – both for index persons in adulthood and for their parents – as a proxy for being family-oriented.

The second panel in Table 7 presents estimates of the effect of sibship size on the marital stability of the parents in the family of origin. For both men and women, the estimated effect of sibship size on their parents' likelihood of remaining married is positive. The estimate is however only statistically significant in the men's sample. Again, there is no evidence of direct effects of sex mix on parents' marital stability (Supplementary Material, Table S.3).

Children from intact homes may have a more positive experience of family life in their childhood, leading to increased fertility in the next generation (Axinn and Thornton 1996). The positive effect on parents' marital stability found in the men's sample could contribute to increased family-orientedness – resulting in higher fertility – among men.

The third panel of Table 7 shows the estimated effect of sibship size on the index person's likelihood of being married and divorced at the age of 43. For men, an additional younger sibling increases the likelihood of being married at age 43 by 11 percentage points, and it does not affect the likelihood of divorce. This indicates that growing up in a relatively large sibship increases men's family-orientedness more generally – shifting some men who would otherwise have remained unmarried into marrying and having more children. For women, the effect on marriage is about the same as for men, and there is a negative effect of 10.5 percentage points on the likelihood of being divorced ($p < 0.01$). Hence, the negative effect on women's ferility is not a result of lower union stability.

This interpretation is corroborated in the last row of Table 7, which gives the estimates of the effect of sibship size on the likelihood of having had two first children with the same partner. Relative to the effect on the overall probability of having two children (Table 4), having an additional sibling makes it relatively (but statistically insignificantly) more likely for both men and women to have two first children with the same partner.

7.3 Index person's educational attainment

A much hypothesized effect of increased sibship size is that parents will invest less in each child, and that as a result, children from larger sibships will have lower educational attainment (Becker 1960). Educational attainment is again a well-known fertility predictor (Section 2). The lower panel of Table 7 displays no significant effect of sibship size on the likelihood of completing high school by the age of 19. Also when measured at age 43, there is no evidence of effects of sibship size on educational attainment. We find no consistent effects of an additional sibling on the duration of educational enrollment (available upon request). Hence, educational attainment is an unlikely mediator of the effects on fertility behavior.

8 Concluding discussion

While fertility is consistently positively correlated across generations, the findings presented in this paper suggest that the causal effect of sibship size on adult fertility follows a more complex pattern. A second sibling induces some men to have a third child themselves, while it keeps some women from making the same parity transition. Based on the evidence about various mechanisms that could potentially channel the effect of sibship size on adult fertility, there emerges a picture of two processes taking place as a family increases in size.

First, an additional child shifts time and attention to family life from other activities. In the study of mediators, we observe a shift away from mothers' labor supply, easily interpreted as an increase in family time. Evidence of increased union stability both for index persons in adulthood and for their parents further supports a shift towards family values. From this process alone, we would expect a positive impact of sibship size on fertility in adulthood.

Second, the additional child takes up some resources – in terms of time, income, or both. Presumably, other siblings will to some extent receive less time and monetary input from their parents, and be expected to provide some of their own time to the care of their younger sibling. Resources also become scarcer for parents, and mothers' time in particular will be visibly more constrained. As knowledge about the consequences of fertility decisions is often obtained through own experience (Bernardi and Klaerner 2014), an additional sibling might, in this way, make children more conscious of the costs of raising a larger family, potentially causing a negative impact of sibship size on fertility in adulthood.

The findings in Section 7 indicate that the relative impact of these two processes varies with gender. Evidence of the first process is found mainly in the male sample, where mothers' labor supply is much more reduced than in the female sample, and where the additional sibling significantly increases parents' marital stability. One interpretation of these findings is that parents in the male sample adapt their preferences and values to the increase in family size (cf. Hayford (2009)), and then transmit these (adaptive) preferences to their children (Axinn et al. 1994; Barber 2000).

In the female sample, the evidence of such a shift towards family values is weaker, as mothers' labor supply and parents' marital stability is less affected. Insofar as parents' time concerns are felt more keenly by children than their money concerns, the second process will be of relatively greater importance in the female sample. Our female index persons will either have witnessed mothers who were far more time constrained, or they will have had to provide much more for their younger siblings – or both – than their male counterparts. Conley (2004) suggests that families are much more likely to use girls as a "labor reserve" when parental time is scarce. The fact that daughters help out more at home than sons (see Sections 2 and 7.1) could in part explain why mothers of girls and boys respond so differently to the birth of an additional sibling. However, by sharing in the family workload, their daughters may in turn become reluctant to have many children in adulthood. The fact that the women in our study who have an additional sibling are moved to refrain from having a larger family – and not from marriage and parenthood in general – further supports the interpretation that the birth of an additional sibling reveals specific information of the strains associated with life in larger families.

The negative effect of an additional sibling is concentrated among second-born women (Section 5.1). When a third child is born, the second-born is moved to the comparatively less favorable position as middle-born in the sibship (Argys et al. 2007; Kidwell 1982; Salmon et al. 2012). As middle-borns on average receive less time and attention from their parents, they may also be more aware of the disadvantages of a larger sibship (regarding educational attainment, Conley and Glauber (2006) find that second-borns are more negatively affected than firstborns by the birth of an additional sibling). This suggests that our findings likely are driven by an interplay of different social mechanisms, which together create the pattern that emerges.

In the literature on fertility contagion, fertility is expected to be contagious through social networks largely due to imperfect information of its consequences, and individuals draw upon their own experiences and network as information sources (Bernardi and Klaerner 2014). While most studies of fertility contagion consistently find positive effects, our study presents new evidence of belief-mediated negative contagion on female fertility. As controlling for unobserved heterogeneity in the OLS framework is usually only partial, estimates of social contagion are likely to be biased upwards due to similarity within networks and families.

In the male sample, the IV estimates are not significantly different from the OLS estimates, suggesting that the intergenerational correlation in fertility could in large part be driven by the causal effect of an additional sibling. In the female sample, on the other hand, our causal estimates suggest that the intergenerational correlation in fertility is substantially and significantly biased upwards, compared to the effect of an additional sibling. This suggests that exactly those mechanisms that are netted out in our IV design – similarities in both observables and unobservables between parents and children – drive the positive intergenerational correlation in fertility among women. This includes shared social background, heritability of fertility preferences (Rodgers et al. 2001) and transmission of parents' *initial* fertility preferences (Starrels and Holm 2000).

The IV estimates of sibship size presented in this paper pertain, in the strictest sense, to the margin between one and two siblings in the family of origin, and to individuals whose parents are moved to having a third child by the sex mix of the two first children. As the alternative twin instrument also yields a more negative effect of sibship size on women's fertility in adulthood than on men's, we believe the results are not specific to children of parents with exactly these sex mix preferences. Throughout the period of study, the most common number of (maternal) siblings in Norway is one, closely followed by two (Rønsen 2004, p.276). In our index cohorts, our estimates thus capture the effect of moving from a typical small to a typical larger sibship. The margin from two to three children is an important one: The decreasing proportion of families with more than two children has been pinpointed as an important driver of lowest-low fertility (Morgan 2003). One might speculate that having a first sibling on average is perceived as more of a clear gain, potentially giving a more positive effect on fertility in the next generation than indicated by our findings.

Furthermore, the external validity of our findings depend on whether the mechanisms we identify in Section 7 are also plausble in other social contexts than the egalitarian Nordic welfare state. The positive effect found among men seems driven by larger sibships strengthening "family-orientedness" in the next generation. This mechanism is found in other Western contexts in the literature on intergenerational transmission (Axinn et al. 1994; Barber 2000; Starrels and Holm 2000), backing up the external validity of the positive causal effect found for men.

Our analysis of mechanisms suggest that the increase in family size intensifies the conflict between work and family life for mothers, translating into lower fertility for women in the next generation. Norway in the 1960s and 1970s was in transition from a traditional towards a more gender egalitarian society. While female labor supply increased, state provided child care was only partial (Havnes and Mogstad 2011), and fathers' active involvement in child care and housework was marginal (Kitterød and Rønsen 2013; Statistics Norway 1983). Several studies link such partial gender equality – potentially leaving mothers with a "double burden" (Sieber 1974) – to low fertility (Becker 1991; Esping-Andersen and Billari 2015; Goldscheider et al. 2015; McDonald 2000a). This environment resembles that of lowest-low fertility regions in today's Europe, where increased female labor supply has not been paralleled with institutions that support working mothers or increased father involvement (McDonald 2000a; Goldscheider et al. 2015). As of today, our findings for women are of particular relevance for these contexts, and perhaps less important for contexts where high female labor supply and relatively high fertility coexist. Furthermore, our results are of relevance for any parity transition that intensifies the conflict between paid work and motherhood – be it the third or fourth, or maybe fifth, child. At very high parity transitions couples may already practice full gender specialization, and the mechanisms we detect here may be of less relevance. Notably, our results indicate that negative cross-generational effects pertain even when the younger generation faces vastly improved institutional support (Rønsen and Skrede 2010).

For countries seeking to maintain both high fertility rates and high female labor force participation, our results underline the significance of facilitating the combination of family life and market work throughout active parenthood. It is well established that the lack of such support structures have an immediate negative effect on fertility. The novelty of our findings lies in that this negative effect on fertility may be even more severe than previously assumed, as it can last across generations.

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Appendix

Table 8: Mean values in additional outcome variables, by index person's sex

	Men			Women		
	Mean	SD	N	Mean	SD	N
Mother's labor supply						
Income, 2nd child aged 6-10	1.47	(1.38)	82809	1.48	(1.37)	78647
Income, 2nd child aged 11-15	2.02	(1.60)	106314	2.05	(1.60)	100291
Employment, 2nd child aged 6-10	0.50	(0.42)	82809	0.50	(0.42)	78647
Employment, 2nd child aged 11-15	0.65	(0.41)	106314	0.66	(0.40)	100291
Father's labor supply						
Income, 2nd child aged 6-10	5.80	(2.08)	109110	5.80	(2.08)	102711
Income, 2nd child aged 11-15	6.10	(2.43)	108466	6.11	(2.44)	102063
Employment, 2nd child aged 6-10	0.98	(0.11)	109110	0.98	(0.11)	102711
Employment, 2nd child aged 11-15	0.97	(0.14)	108466	0.97	(0.14)	102063
Parents' marital stability						
Parents married, 2nd child aged 28	0.74	(0.44)	90080	0.73	(0.44)	84709
Index person's union stability						
Married at age 43	0.52	(0.50)	102376	0.55	(0.50)	98417
Divorced at age 43	0.14	(0.34)	102376	0.18	(0.39)	98417
Has >1 child with one partner at 43	0.55	(0.50)	110225	0.64	(0.48)	103760
Index person's educational attainment						
Secondary educ. at age 19	0.32	(0.47)	108016	0.41	(0.49)	102112
Secondary educ. at age 43	0.69	(0.46)	76463	0.67	(0.47)	72084
Lower tert. educ. at age 43	0.28	(0.45)	76463	0.36	(0.48)	72084
Higher tert. educ. at age 43	0.09	(0.28)	76463	0.06	(0.23)	72084

Note: The samples consist of all first- and second-born men and women born in Norway between 1960 and 1969 in families with at least two children, where the two first children are registered with the same mother and father. Parents' income is measured in base amounts (BA), and employment is defined as having income>1BA. Standard deviations in parentheses.

Table 9: Time spent on housework by child sex and family size. OLS regression results.

	Estimate	(S.E.)
AddSib	0.9	(6.21)
Girl	32.4**	(7.4)
Girl × AddSib	10.9	(12.9)
Adj. R^2	0.107	
N Obs. (Unique ind.)	415	(208)

The sample consists of 415 days of time use entries (208 unique individuals). Standard errors are clustered at the individual. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

Supplementary material

Figure S.1: Effects on index person's number of children measured at ages 25-45

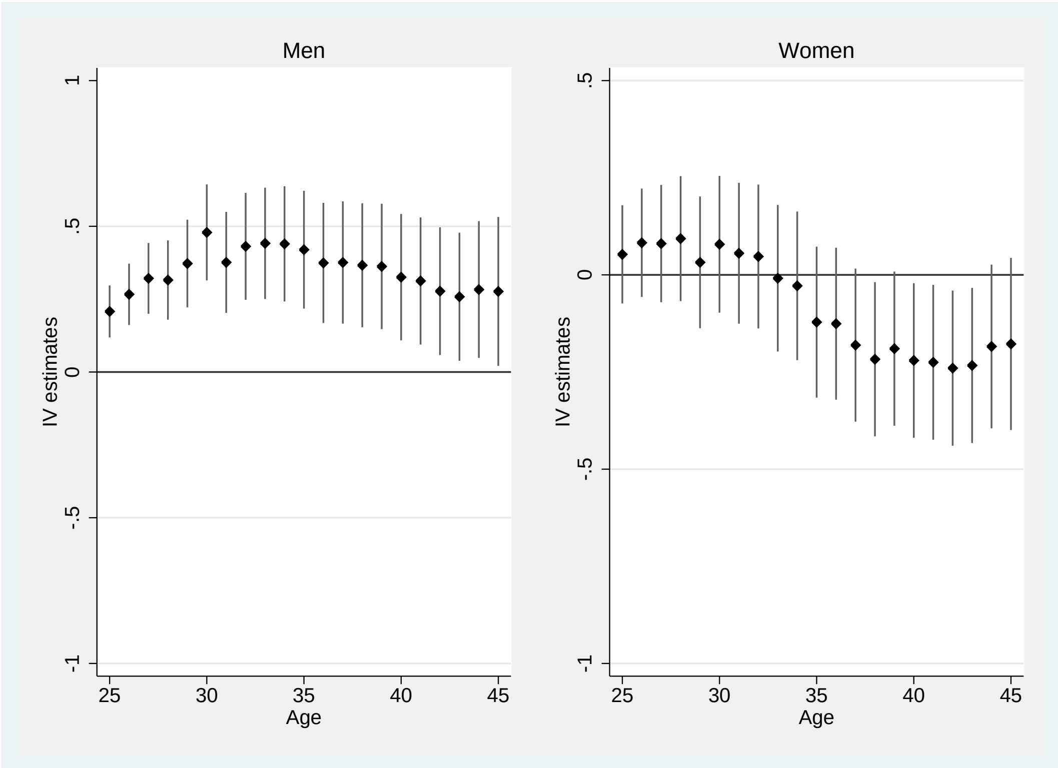

Table S.1: Balancing test of family background variables

First born's birth year	
- 1961	-0.000
	(0.002)
- 1962	-0.002
	(0.002)
- 1963	-0.002
	(0.002)
- 1964	-0.004†
	(0.002)
- 1965	-0.004†
	(0.002)
- 1966	0.004†
	(0.002)
- 1967	0.002
	(0.002)
- 1968	0.005**
	(0.002)
- 1969	0.000
	(0.001)
- 19610	0.000
	(0.000)
Mother's age at first birth	
- $<$20 years	-0.001
	(0.002)
- 20-24 years	-0.005†
	(0.003)
- 25-29 years	0.006*
	(0.002)
- 30-34 years	0.001
	(0.001)
- $\geq$35 years	0.000
	(0.000)
Father's age at first birth	
- $<$20 years	0.001
	(0.001)
- 20-24 years	0.002
	(0.003)
- 25-29 years	-0.005†
	(0.003)
- 30-34 years	0.002
	(0.002)
- $\geq$35 years	0.001
	(0.001)
Distance first and second born	
- $<$1 year	-0.000*
	(0.000)
- 1-2 years	0.002
	(0.003)
- 2-3 years	-0.001
	(0.003)
- 3-4 years	-0.004
	(0.003)
- 4-5 years	0.001
	(0.002)
- 5-6 years	0.000
	(0.001)
- $>$6 years	0.003**
	(0.001)
Observations	107245

Note: The samples are all couples with at least two children, where the two first children are both born in Norway in the period 1960-1969 and are registered with the same mother and father. For the means, standard deviations are reported in parentheses, for the estimated differences, standard errors are in parentheses. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

Table S.2: The effect of sibling sex composition and sibship size on number of children at age 43, effects by index person's birth order

MEN	Firstborns		Second-borns	
	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)
> 1 sibling	0.363*	0.374*	0.095	0.116
	(0.175)	(0.181)	(0.181)	(0.186)
Birth year FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
R2	.	.	0.005	0.007
N	55537	55537	55614	55613

WOMEN	Firstborns		Second-borns	
	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)
> 1 sibling	-0.003	-0.007	-0.459**	-0.471**
	(0.161)	(0.170)	(0.164)	(0.164)
Birth year FE	Yes	Yes	Yes	Yes
Other controls	No	Yes	No	Yes
R2	0.000	0.004	.	.
N	52398	52397	52321	52321

Note: The sample is first- and second-borns in Norwegian families with at least two children (where the two first children are registered with the same mother and father), who are born between 1960 and 1969. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

Table S.3: The effect of sibship size on childhood circumstances and educational achievement

	Men			Women			Diff
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Outcome:	Red.form	IV est.	Dir.eff.	Red.form	IV est.	Dir.eff.	IV est.
Mothers' labor supply during childhood							
Employment, 2nd child aged 6-10	-0.011**	-0.176**	-0.006	-0.001	-0.008	0.007	-0.164*
	(0.004)	(0.057)	(0.006)	(0.004)	(0.056)	(0.006)	(0.064)
N	82809	82809	22364	78647	78647	22359	161456
Employment, 2nd child aged 11-15	-0.019**	-0.330**	-0.007	-0.001	-0.015	0.008	-0.300**
	(0.003)	(0.052)	(0.005)	(0.003)	(0.050)	(0.005)	(0.058)
N	106314	106314	30978	100291	100291	30975	206605
Log earnings, 2nd child aged 6-10	-0.003	-0.053	0.010	-0.004	-0.065	-0.011	0.014
	(0.005)	(0.073)	(0.008)	(0.005)	(0.073)	(0.008)	(0.082)
N	53995	53995	12823	51855	51855	12815	105850
Log earnings, 2nd child aged 11-15	-0.019**	-0.324**	-0.004	-0.008*	-0.119*	0.003	-0.193**
	(0.004)	(0.063)	(0.006)	(0.004)	(0.059)	(0.006)	(0.068)
N	82721	82721	22659	78704	78704	22646	161425
Parents' marital stability							
Parents married, 2nd child aged 28	0.008*	0.122*	0.002	0.003	0.044	-0.001	0.075
	(0.004)	(0.057)	(0.005)	(0.004)	(0.056)	(0.005)	(0.063)
N	90754	90754	25421	85404	85404	25421	176158
Index person's educational achievement							
Secondary educ. at age 19	0.004	0.061	-0.004	-0.001	-0.015	0.007	0.076
	(0.003)	(0.051)	(0.005)	(0.003)	(0.052)	(0.005)	(0.066)
N	108016	108016	31471	102112	102112	31658	210128
Secondary educ. at age 43	0.001	0.009	-0.003	-0.002	-0.030	0.007	0.042
	(0.003)	(0.063)	(0.006)	(0.004)	(0.061)	(0.006)	(0.081)
N	76463	76463	23130	72084	72084	23294	148547
Lower tert. educ. at age 43	0.004	0.071	-0.011*	0.001	0.012	-0.002	0.058
	(0.003)	(0.062)	(0.006)	(0.004)	(0.062)	(0.006)	(0.080)
N	76463	76463	23130	72084	72084	23294	148547
Higher tert. educ. at age 43	0.000	0.007	-0.000	-0.001	-0.020	$0.005^{\dagger}$	0.026
	(0.002)	(0.039)	(0.004)	(0.002)	(0.030)	(0.003)	(0.046)
N	76463	76463	23130	72084	72084	23294	148547

Note: In columns (1) and (2), each cell gives the 2SLS estimate of the effect of sibship size on the outcome given by the row heading. In column (3), each cell gives the corresponding 2SLS estimate of the difference in the effect of sibship size by index person's sex, estimated in the pooled sample. The samples are mothers (upper panel), parental couples (middle panel) and children (lower panel) in Norwegian families with at least two children, where the two first children are registered with the same mother and father and are born between 1960 and 1969. Standard errors (in parentheses) are clustered at the family of origin. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

Table S.4: The effect of sibship size on outcomes related to family stability in adulthood

	Men			Women			Diff
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Outcome:	Red.form	IV est.	Dir.eff.	Red.form	IV est.	Dir.eff.	IV est.
Marital status of index person							
Married at age 43	0.007*	0.111*	0.006	0.006†	0.097†	-0.007	0.012
	(0.003)	(0.054)	(0.006)	(0.003)	(0.052)	(0.006)	(0.071)
N	102376	102376	29867	98417	98417	30540	200793
Divorced at age 43	0.002	0.032	-0.004	-0.007**	-0.105**	0.004	0.130*
	(0.002)	(0.037)	(0.004)	(0.002)	(0.040)	(0.004)	(0.052)
N	102376	102376	29867	98417	98417	30540	200793
>1 child same partner at 43	0.007*	0.125*	-0.002	0.001	0.011	0.004	0.108
	(0.003)	(0.054)	(0.006)	(0.003)	(0.050)	(0.005)	(0.069)
N	110225	110225	32273	103760	103760	32274	213985
>1 child at 43	0.003	0.059	-0.003	-0.002	-0.025	0.004	0.080
	(0.003)	(0.052)	(0.005)	(0.003)	(0.046)	(0.005)	(0.066)
N	110225	110225	32273	103760	103760	32274	213985

Note: Each cell gives the coefficient resulting from an estimation where the outcome is given by the row heading. The sample is first- and second-borns in Norwegian families with at least two children (where the two first children are registered with the same mother and father), who are born between 1960 and 1969. Sample sizes correspond to those in Table 7. Standard errors (in parentheses) are clustered at the family of origin. † $p < 0.10$,* $p < 0.05$, ** $p < 0.01$.

Table S.5: Time spent on housework by child sex. Simple means.

	N	Mean	(S.D.)	Min.	Max
Girl	224	(55.5)	69.6	0	405
Boy	191	(18.8)	35.7	0	255

"Housework" refers to minutes spent on house (not care) work as a primary activity during one day. Data are obtained from the Norwegian Time Use Surveys, collected in 1980-1. The samples is girls and boys aged 15-19, who are living with one or two siblings and their mother at the time of the interview. The sample consists of 415 days of time use diaries, reported by 208 unique individuals.

Details on Time Use Data: The time use data used in the estimations in Table 9 come from the 1980 Norwegian Time Use Survey (Statistics Norway 1983), when our index cohort were 11-20 years old, which was conducted on individuals above 15. We rely on the sample of 15-19 years olds (born 1961-1965) to inform us of the time use patterns in our index persons' youth. Our study sample consists of individuals who are unmarried and have no own children, and who live with their mother and one or two siblings (no conditions are imposed regarding the father). The resulting sample includes 415 observations (208 unique individuals reporting time use for (one or) two days each). Information on family structure is self-reported.

The dependent variable is the number of minutes spent by the respondent on housework as the "main activity" (means shown in Table S.5). This is the sum of minutes spent preparing meals and food, cleaning, washing and mending clothes, and fetching water and wood (Statistics Norway 1983, p. 22). While information on child care as the main activity is also available, more than 90% of children and about 40% of mothers with children living at home state that they spend zero minutes on child care as a main activity, which indicates that child care, as found in previous studies, is often classified as "secondary activity" (Kitterød 2001). Minutes spent on housework is therefore likely to be a more reliable indicator of efforts at home. We are grateful to Hege Kitterød for helpful discussions on the interpretation of time use data.

Using time spent on housework as the dependent variable likely gives a conservative estimate of gender differences in contributions at home, as it does not capture babysitting as a "secondary activity"(Kitterød (2001), see also Statistics Norway (1983, p.168-169)), which is disproportionally done by girls (Brannen 1995; Evertsson 2006; White and Brinkerhoff 1981).

Explanatory variables are respondent's sex, number of siblings (two vs. one), and the interaction between respondent sex and number of siblings (controlling for respondent's age and age squared).