Mothers' Employment and Children's Educational Gender Gap*

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Abstract

This paper analyzes the connection between two concurrent trends since 1950: the narrowing and reversal of the educational gender gap and the increased labor force participation rate (LFPR) of married women. We hypothesize that the education production for boys is more adversely affected by a decrease in the mother's time input as a result of increasing employment. Therefore, an increase in the labor force participation rate of married women may narrow and even reverse the educational gender gap in the following generation. We use micro data from the Norwegian registry to directly show that the mother's employment during her children's childhood has an asymmetric effect on the educational achievement of her own sons and daughters. We also document a positive correlation between the educational gender gap in a particular generation and the LFPR of married women in the *previous* generation at the U.S. state level. We then propose a model that generates a novel prediction about the implications of these asymmetric effects on the mothers' labor supply decisions and find supporting evidence in both the U.S. and Norwegian data.

Keywords: Female Labor Force Participation; Educational Gender Gap; Education Production.

JEL Classification Codes: I2, J2.

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1 Introduction

This paper analyzes the connection between two concurrent trends: the narrowing and reversal of the educational gender gap and the increased labor force participation rate (LFPR) of married women since 1950. Figure 1 plots the gender gap in college achievement for different cohorts in the U.S.[1] For the cohort born in 1950 the gender gap is negative—the women's college completion rate is 7.5 percentage points *lower* than the men's. However, the college gender gap narrows rapidly since then, disappears after ten years, and reverses for cohorts born after 1960. For cohorts born in late 1980s the women's college completion rate is 10.5 percentage points *higher* than the men's.[2] Figure 1 also plots trends in the labor force participation rates (LFPR) of both married and unmarried women between the ages of 25 and 30.[3] Since 1950 the LFPR of married women in this age group has more than tripled, rising from 20% to around 70% in 1990. In contrast, the LFPR of unmarried women in the same age group increased only slightly, from 77% to 81%, during the same period. This increasing trend is similar across different education and age groups.[4]

We find that the education production for boys is more adversely affected by a decrease in mothers' time input as a result of their increasing employment. Therefore, an increase in the labor force participation rate of married women may narrow and even reverse the educational gender gap in the following generation. We investigate the direct evidence of the asymmetric effect of mothers' employment on their children's educational achievement as well as its implication on mothers' labor supply decisions. For the direct evidence, we find that in the Norwegian registry data the educational gender gap is positively and significantly correlated with their own mothers' employment during their childhood, even after controlling for the family fixed effect. We also document a positive correlation between the educational gender gap in one generation and the LFPR of married women in the previous generation at the U.S. state level. For the implication, we first propose a model that generates a novel prediction about the implications of these asymmetric effects—conditional on having the same number of children a mother's labor supply should be higher if she has more girls. We then find evidence supporting this prediction in both the U.S. and Norwegian data. Taken together, we argue that the evidence suggests a plausible causal link between mothers' employment and their children's educational gender gap.

We begin with the Norwegian registry data,[5] where we find direct evidence of the gender-asymmetric

[1] The college achievement rate is separately calculated among females and males between age 25 and 64, and the *gender gap in college achievement* is calculated as the rate of females minus the rate of males. The gender gap in Some College and College, calculated among females and males aged between 22 and 64, is very similar. Therefore we do not distinguish the *educational gender gap* and the *college gender gap* in this paper, and use them interchangeably.

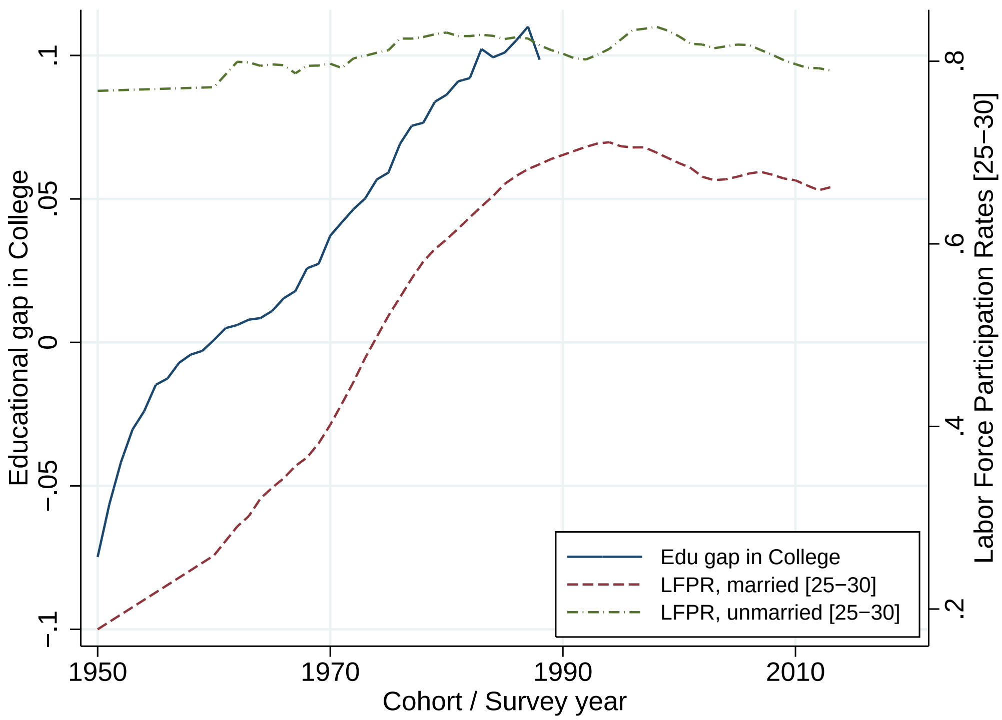
[2] This trend is well documented by many other researchers (Freeman, 2004; Goldin et al., 2006; Becker et al., 2010). Goldin et al. (2006) report that the reversal of the college gender gap is found at all socio-economic status (SES) levels, and in most OECD countries. We focus on the cohorts born after 1950. The widening educational gender gap for cohorts born between 1910 and 1950 are plausibly caused by the World War II and Korean War GI bills (Bound and Turner, 2002; Stanley, 2003), as well as the Vietnam War (Card and Lemieux, 2001). All of them substantially increase the male educational attainment, intentionally or unintentionally.

[3] The LFPR of married and unmarried women in 1950 and 1960 are from the IPUMS Census data, while those since 1962 are from the IPUMS CPS data.

[4] This trend has been documented and studied in Mincer (1962) and Eckstein and Lifshitz (2011), among many others.

[5] It is an administrative data for Norway, covering the full population.

Figure 1: Gender gap in college achievement versus labor force participation rates (LFPR) of married and unmarried women aged 25-30



Notes: The solid line represents the gender gap in college achievement (females versus males) for each cohort while the according x-axis represents the year when the cohort is one-year-old. For example, in year 1980 it represents the gender gap of the cohort who was born in 1979. The college achievement is calculated from individuals aged 25 or older in the U.S. Census data. The dashed and the dot-dashed lines represent the LFPR of married, and respectively, unmarried women aged 25-30 in the following five years. For instance, in year 1980 it represents the average LFPR between 1980 and 1984. The LFPR between 1962 and 2013 are from the U.S. CPS data while those for decennial years in 1950 and 1960 are from the U.S. Census data.

effects of mothers' employment on their own children's educational achievement. In the difference-in-differences regression model controlling for gender specific cohort fixed effects and family fixed effects, we find that the mother's employment has an extra positive effect on her daughter's educational achievement relative to the effect on her son's, and this effect is statistically significant.

We also document a positive correlation between the educational gender gap in a particular generation and their mothers' LFPR *when the generation was still in their childhood* at the state level in the US. We find that the college gender gap in one cohort is positively correlated with the LFPR of married women in the birth state during the cohort's first five years. This positive correlation is statistically significant in the difference-in-differences model controlling for the cohort and state fixed effects. A back-of-the-envelope calculation shows that, in the U.S. (respectively, Norway), about 13% (respectively, 8%) of the total change in the college gender gap can be attributed to the increase in the LFPR of married women since 1950.

What causes this asymmetric effect of mothers' employment on their children's educational achievement? We argue that the asymmetry comes from both the production channel and the role model channel. If the mother's time input has higher marginal productivity in her son's education production, then a reduction in mother's time input as a result of increased employment has more detrimental effect on her son's educational achievement. This is the asymmetric production channel.[6] On the other hand, a working mother provides a positive role model toward working for her daughter directly, which may increase the daughter's perceived expected return of schooling. The asymmetric role model channel has a more positive effect on the daughter's educational achievement. We show that both channels are present in the data and contribute to the narrowing and the reversal of the educational gender gap.

We present further evidence from a different perspective to support our theory. Given the asymmetric effects of maternal employment, a well-informed, rational and altruistic mother who values her children's educational achievement faces the trade-off between participating in the labor market for income and putting her time into her children's education production. Using a parsimonious model we show such a mother would be more likely to participate in the labor market if she has a higher fraction of girls, conditional on having the same number of children, if maternal time input is more productive for her sons' educational production than for her daughters'. We then find empirical evidence supporting this prediction in both the U.S. Census data and Norwegian registry data. Specifically, we find that, among mothers with two or more children, the mother's labor force attachment is positively and significantly correlated with indicators for the first two children both being girls, or the fraction of girls among her children, in most of the U.S. Census years and in the pooled sample, as well as in the Norwegian data.

We also examine two plausible alternative explanations for this positive correlation between moth-

[6] Many literature also find that boys are more vulnerable to family disruption such as divorce than girls (e.g. Hetherington et al., 1981; Amato and Keith, 1991; Amato, 2001). Bertrand and Pan (2013) find boys raised by single mothers are at higher risk of having behavioral problems. They argue that boys are more responsive to parental inputs which decrease dramatically among broken families. In the Swedish registry data, Weitoft et al. (2003) find boys raised in single-parent families have higher risk of mortality from all causes including suicide, morbidity and injury than girls. However, our empirical work will not devote more space attempting to investigate the reasons why the production function is asymmetric, but rather show that its presence is consistent with the data.

ers' labor force attachment and the fraction of girls. The first alternative hypothesis is the "differential cost hypothesis"—it may be more expensive to raise girls, thus mothers with more girls work more because of higher financial needs. If so, however, we would also expect that families with first child being girl should have fewer children, to the extent that children are a normal good. This prediction is contrary to what is seen in the U.S. data.[7] The second alternative hypothesis is that mothers with sons may have higher bargaining power in their families, at least in countries with son-biased preferences, and therefore they are less likely to work. This "bargaining power hypothesis" can be distinguished from our gender-asymmetric effect of mothers' employment on children's education production hypothesis in terms of age-dependency. The bargaining power associated with the genders of the children is permanent and should not vary over the children's ages, and therefore it predicts an age-invariant effect of the children's genders on mother's labor supply. On the other hand, our hypothesis of the gender-asymmetric effect of mother's employment implies this effect would vary over time because both the production channel and the role model channel are age-varying. In the pooled U.S. Census data, we find that the positive gradient of the mother's labor force participation rate with respect to the fraction of girls increases with her youngest, oldest, or average child's age. These results do not favor the bargaining power hypothesis but are consistent with our hypothesis of the gender-asymmetric effect of mothers' employment.

The remainder of the paper is organized as follows. In Section 2, we review related literature on educational gender gap, effects of maternal employment and their interaction. In Section 3 we present the estimation of the asymmetric effects of mothers' employment on their children's educational achievement using the administrative Norwegian registry data. In Section 4 we present the relationship between the female LFPR and the educational gender gap at the state level in the U.S. In Section 5 we propose a parsimonious model of mother's labor supply, derive its empirically testable implication, and empirically test it using both the U.S. Census data and Norwegian registry data; we also discuss alternative explanations. Finally, in Section 6 we conclude.

2 Related Literature

By proposing a new and complementary explanation, our paper contributes to the growing literature which documents and explains the narrowing and reversal of the educational gender gap since 1950s in the U.S. and other countries. Researchers have documented the dynamic changes of the gender gaps in high school graduation and dropout (e.g., Murnane, 2013), as well as in college enrollment and graduation (e.g., Goldin et al., 2006 and Diprete and Buchmann, 2006). Various explanations have been explored in the literature. One strand of the literature attributes the narrowing of the educational gender gap to higher returns to college in the labor market for females than for males since at least the 1970s (see Dougherty, 2005, for a comprehensive review).[8] A second strand of the literature argues that

[7]This hypothesis and the evidence are studied clearly in Dahl and Moretti (2008).

[8]However, Hubbard (2011) finds that after controlling the top coding bias in the CPS data the difference in the college wage premium between women and men is statistically insignificant from 1990s. Becker et al. (2010) also find that the benefits of

the reversal of the college gender gap is due in part to the behavioral and developmental differences between girls and boys—girls bear lower costs to prepare and attend colleges than boys (e.g., Goldin et al. (2006)). Girls are found to have higher non-cognitive skills (Jacob, 2002; Bertrand and Pan, 2013), lower rates of Attention Deficit Hyperactivity Disorder (ADHD) (Cuffe et al., 2005), lower incidence of arrests and school suspension (Goldin et al., 2006), or higher elasticity of college supply (Becker et al., 2010).[9,10] A third strand of the literature argues that higher returns to college in the marriage market for women than for men also contribute to the changes of the educational gender gap (e.g., Chiappori et al., 2009; Ge, 2011; Huang, 2013).

Our paper also contributes to a large literature investigating effects of parents' or mothers' time on the development of their children's cognitive and non-cognitive skills. Parental time input is a key ingredient in children's human capital development throughout their childhood and adolescence. Cunha et al. (2010) find that parental investments at younger ages are relatively more important than those at older ages. Del Boca et al. (2014) find that both parents' time investment are important in the production function of children's cognitive skills, and their productivities decline as the children age, while the monetary inputs are not as important. However, Gayle et al. (2012) estimate that the return to mothers' time investment is two thirds higher than that of fathers', and the returns to mothers' time investment are significantly higher for boys than for girls. Table 1 provides a non-exhaustive summary of the studies on the asymmetric (and/or symmetric) effects of maternal employment on children's achievement.[11,12]

The inter-generational connection between the mother's LFPR and her children's educational gender gap presented in this paper is also consistent with the positive correlation between the gender difference in PISA (Programme for International Student Assessment) test scores and the contemporary female LFPR. Guiso et al. (2008) find that the gender gap in the 2003 PISA test scores (both math and reading) is positively and significantly correlated with the contemporary female LFPR in their own countries. They attribute this correlation to the influence of cultures. The result is confirmed in the 2009 PISA data by González de San Román and De La Rica (2012). They further find that having a working mother (in the test year) significantly improves the girl's test scores in both math and reading regardless of the mother's education level, while the effect is insignificant for the boy. Fathers working full time do not have such an asymmetric effect. They call this effect as "the intergenerational transmission of gender role attitudes within the family from mothers to daughters." Both papers use the contemporary female LFPR as a measure of culture or social norm.

attending college are not higher for women than for men.

[9] Becker et al. (2010) argue that women overtake men in higher education because women and men differ in their distributions of the total costs of higher education, primarily due to the difference in their distributions of noncognitive skills. They argue that women's distribution of costs of attending college has a higher mean but a lower variance than those of men's, which leads to a higher elasticity of college supply among women than among men.

[10] These gender differences might come from the fact that women mature earlier and are more patient than men, possibly due to evolutionary selection (Silverman, 2003).

[11] For example, it does not include much research which does not mention gender difference in effects of maternal employment on children's achievements.

[12] See Bernal and Keane (2011) for an extensive list of papers studying the symmetric or asymmetric effects of maternal employment using NLSY data.

Table 1: Effects of mother's employment (ME) on children's achievement.

Studies	Data	Method	Asymmetric effects of maternal employment?
	Generate Positive Difference (Girls vs. Boys)		
Hetherington et al. (1981)	Literature review	-	Divorce hurts boys more often
Belsky and Rovine (1988)	PIFD Project[1]	Comparison	Infants: Boys-[2]
Krein and Beller (1988)	NLS-LME[3]	OLS	Living with mother alone: Girls 0, Boys-
Desai et al. (1989)[4]	NLSY	OLS	Boys of high income family-
Baydar and Brooks-Gunn (1991)	NLSY	OLS	Behavioral: Girls+, Boys-
Mott (1991)	NLSY	OLS	Infants: Healthy girls+, Not healthy Boys-
Brooks-Gunn et al. (2002)	NICHD-SECC[5]	OLS	Boys are more affected negatively
Goldberg et al. (2008)	Meta-analysis	OLS	More positive for girls
Gayle et al. (2012)	PSID	Structural model	Higher return of maternal time for boys
	Generate Negative Difference (Girls vs. Boys)		
Waldfogel et al. (2002)[6]	NLSY	OLS	1st-year ME affects girls more negatively
	Generate No Difference (Girls vs. Boys)		
Blau and Grossberg (1992)	NLSY	OLS	Negative overall
Han et al. (2001)	NLSY	OLS	Negative overall
Baum (2004)	NLSY	OLS	Negative overall
Ruhm (2004)	NLSY	OLS	Negative overall
Lucas-Thompson et al. (2010)	Meta-analysis	OLS	1st-year ME negative overall
	Difference N/A		
Milne et al. (1986)	SESTI & HSB[7]	OLS	Negative overall
Baum (2003)	NLSY	IV	Negative overall
Bettinger et al. (2014)	Norway Registry	Diff-in-Diff	Negative overall

OLS also includes fixed effect regressions.
Note 1: PIFD Project—the Pennsylvania Infant and Family Development Project.
Note 2: "Boys-" means the effect is negative for boys; "Girls+" means the effect is positive for girls; "Girls 0" means the effect is insignificant for girls.
Note 3: NLS-LME: National Longitudinal Survey of Labor Market Experience.
Note 4: Only maternal employment during the boys' infancy had a significant negative effect on PPVT scores at Age 4.
Note 5: NICHD-SECC: National Institute of Child Health and Human Development Study of Early Child Care.
Note 6: They discuss the gender differential effects of the first-year maternal employment only.
Note 7: SESTI: Sustaining Effects of Title I; HSB: High School and Beyond.

3 Asymmetric Effects of Maternal Employment—Micro Evidence from Norway

We first present micro evidence that shows a positive correlation between the mother's employment and her own children's educational gender gap in Norwegian Registry data. Specifically, in both the first-difference regression model controlling for gender specific cohort fixed effects and the difference-in-differences regression model which further controls for family fixed effects, we find that the mother's employment has an additional positive and statistically significant effect on her daughter's educational achievement relative to the effect on her son's.

3.1 Norwegian data

We use the Norwegian Registry data which consists of all children born in Norway between 1967 and 1993, with 724,091 males and 690,888 females. For all these children we identify their parents and then construct variables describing each parent's earnings history in the period when their children are between birth and 17 years of age.

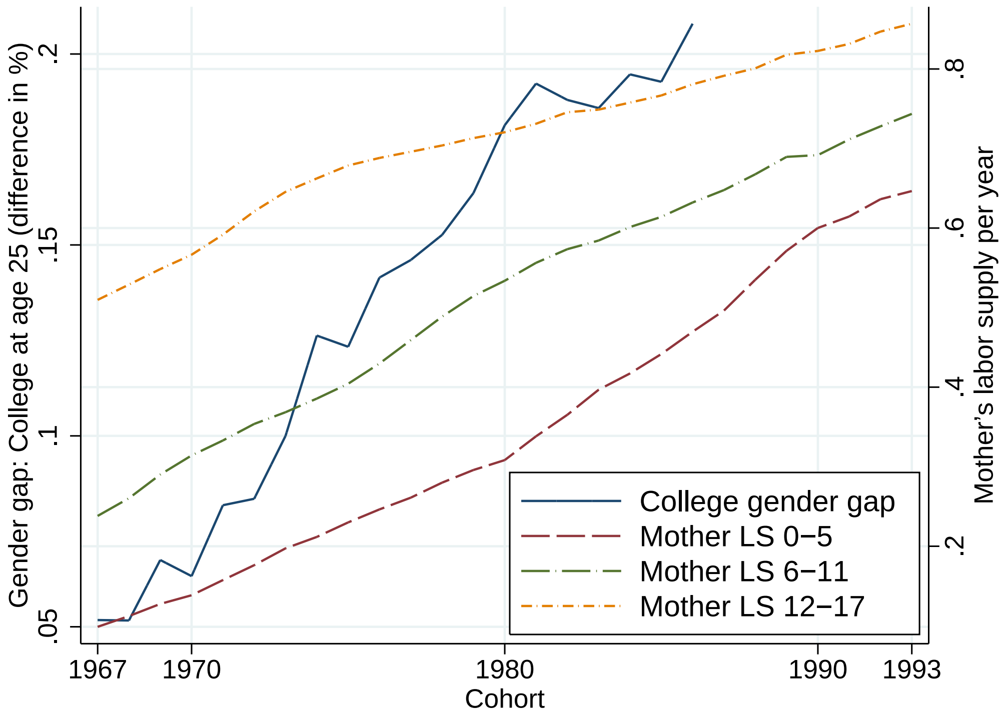
As a measure of labor supply or employment we calculate the number of years where one has at least one base unit of earnings in the Norwegian social security system in a given period, e.g. from when the child is a newborn to 5 years of age, from age 6 to age 11, and from age 12 to 17, respectively.[13] During the period up to 1993, we only observe pension accrual for these parents. Pension accrual is a simple function of earnings such that we can use these to obtain earnings. However, earnings also include other types of income support, such as disability benefits, paid sickness leave and unemployment insurance. In fact part of the earnings increase over time is due to the extension of these income support programs. Therefore in the econometric specifications we include the gender specific cohort fixed effects to control for such policy changes over time.

Figure 2 plots the gender gap in college achievement for cohorts born between 1967 and 1986, as well as their mothers' annual employment rates when they are aged between 0-5, 6-11, or 12-17, respectively. College achievement is measured by whether or not an individual received a college degree by the age of 25. The gender gap is calculated as the difference in the ratio of college graduates between females and males. It has been increasing since 1967.[14] There has also been a dramatic increase in mothers' employment, from 10% in 1967 to 65% in 1993 for mothers with young children. On the other hand, fathers' employment rate has been rather constant, varying between 92% and 94%. Overall, Norwegian data presents similar patterns as the U.S. data in figure 1. Table 2 presents summary statistics of Norwegian registry data used in this paper.

[13] In 2014 this base unit equals 13,300 USD. We use "labor supply" and "employment" interchangeably throughout the paper.

[14] There has been an increase in college achievement for both boys and girls, but much more so for girls. Results are similar if other measures of educational achievement are used, for example whether or not having finished college at the age of 28 or the total years of schooling.

Figure 2: College gender gap and mother's employment, Norway.



Notes: The x-axis represents the cohort's birth year. The solid line represents the educational gender gap in college achievement at age 25 for each cohort (corresponding to the y-axis on the left). All other lines represent the mothers' average labor supply per year (or the labor force participation rate) when their children are aged 0-5, 6-11, or 12-17, respectively (corresponding to the y-axis on the right).

Table 2: Summary statistics of Norwegian registry data.

Variable	Mean	S.D.	Min	Max
College achievement at age 25	0.331	0.471	0	1
Girl	0.488	0.500	0	1
Mother's age at birth	26.889	5.161	8	56
Mother labor supply during $[0, 17]$	0.518	0.341	0	1
Mother labor supply during $[0, 5]$	0.344	0.394	0	1
Mother labor supply during $[6, 11]$	0.505	0.421	0	1
Mother labor supply during $[12, 17]$	0.707	0.389	0	1
Logarithm of family earnings during $[0, 17]$	4.949	0.476	-5.557	9.305
Logarithm of family earnings during $[0, 5]$	3.658	0.526	-6.016	7.617
Logarithm of family earnings during $[6, 11]$	3.840	0.555	-5.972	8.476
Logarithm of family earnings during $[12, 17]$	3.964	0.603	-9.915	9.145

Notes: The final sample has $1,001,157$ observations. Mother's labor supply is measured by the number of years in a given period where one has at least one base unit of earnings in the Norwegian social security system.

3.2 Asymmetric Effects of Mother's Employment

In Norwegian administrative data we regress an individual's educational achievement at the age of 25 on the mother's employment rate at the individual's various age stages. First, we run the following first-difference (FD) linear probability regression model:[15]

$$y_i = \mathbf{M}_i'\beta + g_i\mathbf{M}_i'\gamma + \mathbf{W}_i'\xi + g_i\mathbf{W}_i'\pi + \mathbf{X}_i'\delta + \nu_{g_i,b_i} + \epsilon_i. \tag{FD}$$

The dependent variable y_i is a dummy which takes value 1 if individual i has a college or higher degree by the age of 25, and 0 otherwise. Among the independent variables, g_i is a dummy for girl; $\mathbf{M}_i$ is a vector of variables including the mother's employment at various stages of i's age; $\mathbf{W}_i$ is a vector of the logarithm of family earnings; $\mathbf{X}_i$ is a vector of other demographics including the mother's age when the child is born and dummies for the birth order; ν_{g_i,b_i} indicates the gender specific cohort fixed effect; ϵ_i is the i.i.d. unobservable component which is assumed to be orthogonal to all other independent variables.[16] We include the interaction term between the girl dummy and the mother's employment to allow for gender specific effect of maternal employment. Similarly, the interaction term between the girl dummy and the log of family earnings is included to allow for gender specific income effect.

Regression results from the FD model are presented in Table 3. In the first column, the mother's employment is measured as the employment rate during her child's entire childhood period, from 0 to 17-year-old. The correlation between the boy's college achievement and the mother's employment — the coefficient of the mother's employment — is positive and statistically significant at the 1% level. More interestingly, there is an additional positive and significant correlation between the girl's college achievement and her mother's employment, measured by the coefficient of the interaction term between the mother's employment and the girl dummy. We then separate the mother's employment into three different segments according to the child's age groups, namely when the child was aged 0-5, 6-11, and 12-17 respectively, and run three separate FD regressions. The results reported in Columns 2-4 reveal that the additional correlation between the mother's employment at various age groups and her own daughter's college achievement is all positive and statistically significant at the 1% level. In other words, there is a positive correlation between the mother's employment and her own children's gender gap in college achievement.

When the mother's employments from three different age groups are included in the same regression, Column 5 shows that the signs and magnitudes of the coefficients are similar as those when separately included, except when the child is aged between 6 and 11. The correlation between mother's employment during that age group and her son's college achievement becomes significantly negative while the correlation between the mother's employment and her children's college gender gap becomes statistically insignificant. Table 3 also presents a positive and statistically significant correlation between

[15] We also estimate the logit regression model and the results are very similar to those from the linear model.

[16] ν_{g_i,b_i} absorbs the effect of the girl dummy and therefore the independent variables do not include g_i per se. The dummy variables for the birth orders are included to control the birth order effects. Higher birth order has a significantly negative effect on their education (e.g., Black et al., 2005), and the parents' earning profiles typically increase over time. Therefore it is important to include the birth order dummies.

family earnings and the children's college achievement. The correlation between family earnings and the children's college gender gap is also significantly positive in most cases except when the child is no older than five-year-old where the correlation is positive but statistically insignificant.

However, one limitation with the FD regression model is the omitted variable issue—some unobservable family characteristics might affect both mother's employment and children's college achievement.[17] To control for such unobservable variables, we include a family fixed effect, η_f, and conduct the following difference-in-differences (DID) regression model:[18]

$$y_i = \mathbf{M}_i'\beta + g_i\mathbf{M}_i'\gamma + \mathbf{W}_i'\xi + g_i\mathbf{W}_i'\pi + \mathbf{X}_i'\delta + \nu_{g_i,bi} + \eta_f + \epsilon_i. \qquad \text{(DID)}$$

In this DID model, the identification of the effects of mother's employment and family earnings on her children's college gender gap comes from the variation across children of opposite genders within the same family.[19] Therefore it is the effect of *changes* in the mother's employment or family earnings rather than the effect of employment or income level as identified in the FD model. Of course, strictly speaking, without exogenous variation in the mother's employment and family earnings, the results from the DID model are still correlations.[20]

The results of the DID model are presented in Table 4. After controlling for family fixed effects, the positive correlations between mother's employment and her son's college achievement estimated in the FD model become negative. It is statistically insignificant when mother's employment is pooled across her child's entire childhood (Column 1), but statistically significant when the employment is measured across the child's different age groups (Columns 2-4). This demonstrates the endogeneity issue in the FD model. On the other hand, the estimates of the interaction term between the mother's employment and the girl dummy are still positive and statistically significant at the 1% level; the magnitudes are similar as the FD model as well. This shows that mother's employment has a negative effect on boys, but has a significantly positive effect on her children's college gender gap. Overall, an additional year of mother's employment increases the gender gap in college by 0.002 ($\approx 0.042/18$) in level. The mother's employment increases from 5.079 years in 1967 to 11.296 years in 1986, which is correlated with an increase of 0.0124 in the gender gap. This is about 8% of the increase in the college gender gap in Norway, which increases from 0.0518 in 1967 to 0.2079 in 1986.

The DID regression results show quite similar effects of mother's employment across their children's different age groups (Columns 2-4). The effects for boys are all negative and statistically significant, while the additional effects of mother's employment on girls are all positive and statistically significant in each of three age groups. When mother's employments from all three age groups are included in

[17] For instance, family environment or mother's ability.

[18] A *family* is defined by a pair of a mother and a father. Therefore a mother might be involved into two different families in the case of divorce-and-re-marriage.

[19] Both maternal employment and family earnings are continuous variables so the difference between one boy and one girl is sufficient for the identification. Note that the widely used twin strategy (e.g. Black et al. (2005)) is not applicable in this DID model since there is no variation in $\mathbf{M}_i$ within the twin.

[20] The DID model might still have the endogeneity issue caused by omitted variables. For example, there might be some unobservable variables which are uncorrelated with time or the child's age or family fixed effects but affect both the change in the mother's employment and the children's college gender gap, although it is not easy to find such variables.

Table 3: First difference linear probability regression of children's college achievement on their mothers' employment, Norway.

	(1)	(2)	(3)	(4)	(5)
Mother's employment	0.072***				
	(0.002)				
Mother's employment $\times$ Girl	0.037***				
	(0.003)				
Mother's employment during $[0,5]$		0.120***			0.129***
		(0.002)			(0.003)
Mother's employment $[0,5]$ $\times$ Girl		0.025***			0.016***
		(0.003)			(0.004)
Mother's employment during $[6,11]$			0.037***		-0.038***
			(0.002)		(0.003)
Mother's employment $[6,11]$ $\times$ Girl			0.020***		0.004
			(0.003)		(0.004)
Mother's employment during $[12,17]$				0.025***	0.014***
				(0.002)	(0.002)
Mother's employment $[12,17]$ $\times$ Girl				0.024***	0.019***
				(0.003)	(0.003)
Family earnings	0.194***				
	(0.002)				
Family earnings $\times$ Girl	0.006***				
	(0.002)				
Family earnings during $[0,5]$		0.122***			0.031***
		(0.002)			(0.002)
Family earnings $[0,5]$ $\times$ Girl		0.003			-0.004
		(0.002)			(0.003)
Family earnings during $[6,11]$			0.143***		0.064***
			(0.001)		(0.002)
Family earnings $[6,11]$ $\times$ Girl			0.007***		-0.006*
			(0.002)		(0.003)
Family earnings during $[12,17]$				0.134***	0.084***
				(0.001)	(0.002)
Family earnings $[12,17]$ $\times$ Girl				0.010***	0.012***
				(0.002)	(0.002)
Observations	1,000,931	999,128	999,102	998,202	995,541
R-squared	0.099	0.084	0.086	0.089	0.102

Notes: Standard errors in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
The dependent variable is the education achievement in college at age 25. Independent variables include the mother's age when the child is born, dummy variables for birth orders, and the gender specific cohort fixed effect.

Table 4: Difference-in-differences regression of children's college achievement on their mothers' employment, Norway.

	(1)	(2)	(3)	(4)	(5)
Mother's employment	-0.012				
	(0.008)				
Mother's employment × Girl	0.042***				
	(0.004)				
Mother's employment during $[0, 5]$		-0.011**			-0.006
		(0.004)			(0.005)
Mother's employment $[0, 5]$ × Girl		0.031***			0.024***
		(0.004)			(0.005)
Mother's employment during $[6, 11]$			-0.012***		0.001
			(0.004)		(0.004)
Mother's employment $[6, 11]$ × Girl			0.023***		0.000
			(0.003)		(0.005)
Mother's employment during $[12, 17]$				-0.009**	-0.006
				(0.004)	(0.0014)
Mother's employment $[12, 17]$ × Girl				0.028***	0.022***
				(0.003)	(0.004)
Family earnings	-0.010*				
	(0.006)				
Family earnings × Girl	-0.002				
	(0.003)				
Family earnings during $[0, 5]$		-0.004			-0.003
		(0.003)			(0.004)
Family earnings $[0, 5]$ × Girl		-0.006*			-0.011***
		(0.003)			(0.004)
Family earnings during $[6, 11]$			-0.002		0.000
			(0.003)		(0.004)
Family earnings $[6, 11]$ × Girl			0.001		-0.007
			(0.003)		(0.004)
Family earnings during $[12, 17]$				-0.005*	-0.008***
				(0.003)	(0.003)
Family earnings $[12, 17]$ × Girl				0.003	0.009***
				(0.002)	(0.003)
Observations	1,000,931	999,128	999,102	998,202	995,541
R-squared	0.709	0.710	0.710	0.710	0.710

Notes: Standard errors in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
The dependent variable is the education achievement in college at age 25. Independent variables include the mother's age when the child is born, dummy variables for birth orders, the gender specific cohort fixed effect and the family fixed effect.

the same regression (Column 5), for boys all the effects of mother's employment become insignificant. On the other hand, the additional effect for girls remains to be significantly positive and has similar magnitude as in separate cases when the children are aged 0-5 or 12-17, but it becomes statistically insignificant when the children are aged 6-11.

After controlling for the family fixed effect, the effect of family earnings becomes negative and statistically significant at the 10% level, which is mainly driven by the effect of family earnings when the child is aged 12-17. Recall the income effect in this DID model refers to the effect of changes in family earnings. In other age groups, the income effect is negative but statistically insignificant. The additional income effects for girls are also statistically insignificant except when the child is young. When the child is aged between 0 and 5, the additional income effect for girls is negative and statistically significant at the 10% level. The family earnings are the pooled earnings from both parents. Therefore they overlap with the mother's employment to some extent. When we drop the mother's employment and instead include the father's employment, all estimates of income effects become statistically insignificant. Bettinger et al. (2014) also find that the effect of changes in family income is not significant in Norwegian data.

When we include the father's employment instead of the family earnings, the estimates of all effects of mother's employment either get strengthened or stay unchanged, as shown in Table 5. In particular, the estimate of the pooled mother's employment on boys remains negative but becomes statistically significant at the 1% level (Column 1). The estimates of the effects of mother's employment in the other specifications all have at least the same statistical significance levels and some increase in magnitude. Interestingly, the father's employment also has a negative effect on boys and a positive additional effect on girls. Both effects are statistically significant when individually included.

Remark 1 *We would also like to emphasize that the estimate of the additional effect of mother's employment on girls relative to boys, namely γ in Eq. (DID), is more reliable than those on boys and girls per se, namely β and $\beta + \gamma$ respectively. To the extent that there are other unobserved time-varying factors affecting both mother's employment and her children's education, they are not captured by family fixed effects in the DID model; therefore the estimates of the effects of mother's employment on her child's educational achievement would be biased. However, as long as those time-varying factors are symmetric for boys and girls, such bias would cancel out. In other words, as long as the effects of those omitted variables are independent of the child's gender, then the bias on the estimate of the effect of mother's employment on her children's educational gender gap, namely γ, is likely to be limited.*

It is worth noting that in the regression results in Table 4, the overall effect of mother's employment on her girl's college achievement is positive and statistically significant in the first four specifications. Existing literature has found mixed estimates for the effect of mother's employment on her girl's cognitive and non-cognitive outcomes. Goldberg et al. (2008) has a detailed review on this bifurcated effect of mother's employment on boys and girls.[21]

[21]See the third and fourth paragraphs in page 79 of Goldberg et al. (2008).

Table 5: Difference-in-differences regression of children's college achievement on their mothers' and fathers' employments, Norway.

	(1)	(2)	(3)	(4)	(5)
Mother's employment (LS)	-0.022***				
	(0.008)				
Mother's employment $\times$ Girl	0.056***				
	(0.005)				
Mother's employment during $[0, 5]$		-0.015***			-0.008
		(0.005)			(0.005)
Mother's employment $[0, 5]$ $\times$ Girl		0.036***			0.026***
		(0.004)			(0.005)
Mother's employment during $[6, 11]$			-0.017***		-0.002
			(0.004)		(0.004)
Mother's employment $[6, 11]$ $\times$ Girl			0.031***		0.006
			(0.004)		(0.005)
Mother's employment during $[12, 17]$				-0.013***	-0.009**
				(0.004)	(0.004)
Mother's employment $[12, 17]$ $\times$ Girl				0.036***	0.028***
				(0.003)	(0.004)
Father's employment (LS)	-0.068***				
	(0.014)				
Father's LS $\times$ Girl	0.095***				
	(0.010)				
Father's LS during $[0, 5]$		-0.025***			-0.005
		(0.009)			(0.009)
Father's LS $[0, 5]$ $\times$ Girl		0.029***			-0.005
		(0.010)			(0.011)
Father's LS during $[6, 11]$			-0.034***		-0.022**
			(0.009)		(0.010)
Father's LS $[6, 11]$ $\times$ Girl			0.066***		0.045***
			(0.009)		(0.011)
Father's LS during $[12, 17]$				-0.027***	-0.020**
				(0.008)	(0.008)
Father's LS $[12, 17]$ $\times$ Girl				0.058***	0.044***
				(0.007)	(0.009)
Observations	1,000,931	999,128	999,102	998,202	995,541
R-squared	0.710	0.710	0.710	0.710	0.710

Notes: Standard errors in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
The dependent variable is the education achievement in college at age 25. Independent variables include the mother's age when the child is born, dummy variables for birth orders, the gender specific cohort fixed effect and the family fixed effect.

3.3 Sources of Asymmetry: Education Production and Role Model

There are three possible different channels that the mother's employment might affect her children's educational achievement, namely income, education production, and role model. In this subsection we argue that the income effect is likely to be symmetric, while asymmetry might arise from the education production channel and the role model channel.

As we previously discussed, the effect of additional income is found to be mostly insignificant and symmetric, which is also consistent with previous literature. For example, Dahl and Lochner (2012) find the income effects on children's math and reading achievement are higher for boys but the difference is statistically insignificant. Bettinger et al. (2014) estimate the effect of changes in family income is not significant in Norwegian data.

The education production channel could be asymmetric for boys and girls. If mother's time input has a differential effect in her boy's and her girl's education production processes, then the reduction of the mother's time input due to increasing employment would have differential effects for the boy and for the girl. If the marginal effect of mother's time is higher for boys than for girls, then an increase in mother's employment would have more detrimental effect for boys than for girls.

The role model channel could also be differential for boys and girls. A working mother provides a role model for her children, which may affect the children beliefs about the expected return of their own education. This effect is most direct and significant for girls only, as found in, for example, González de San Román and De La Rica (2012); similar mechanism is also studied in Fernández (2013).[22]

In levels, the results from the Norwegian data support the presence of both the education production channel and the role model channel. The production channel alone indicates negative effect of maternal employment for both sons and daughters, and the role model channel alone indicates positive effect of maternal employment for daughters only. The negative effect of maternal employment on boys can only be generated by the production channel. At the same time, the positive and significant effect of mothers' employment on girls alone indicates that the role model must be in play.

On the other hand, the asymmetry, reflected by the positive effect of mother's employment on her children's college gender gap, could rise from either the asymmetric production channel, or the role model channel, or both. It is in general difficult to separately identify these two channels, but it may be possible to differentiate them based on the timing of when these two channels are likely to be the strongest. The literature on education production suggests that the maternal time input is more critical at early stage than at later stages (Desai et al., 1989; Cunha et al., 2010). On the other hand, the time trend in the role model channel is not clear. The role model might affect children in a subconscious way which decreases with age, and in a conscious manner which increases with age. If the conscious effect of the role model dominates the subconscious one, it would imply that the asymmetric effect of

[22]For boys, it is possible that competition in the marriage market might generate some indirect effect from maternal employment (Fernández et al., 2004), but this effect is indirect and ambiguous in its sign. Sons of working mothers might favor working wives, so more working mothers result in more men who favor working wives, which everything equal could incentivize boys to acquire more schooling; but this effect is dampened by the fact that there are more working women available in the next generation. Thus the equilibrium effect of mothers' employment through role model on sons is unclear.

maternal employment via the role model channel widens with children's ages. In Table 4, the estimate of the additional effect of mother's employment on girls is U-shaped: stronger when children are either young (0-5) or old (12-17) than when they are in between (6-11). When mother's employment during all three age periods (0-5, 6-11, 12-17) are included in the same DID model, the additional effect of mother's employment on girls is significantly positive when children are either young (0-5) or old (12-17), and these effects (both β and γ) are very similar for mother's employment during these two age groups. As long as the effect of role model varies with the child's age, these results indicate the presence of another source of asymmetry. Since the income effect is symmetric, it can only come from the production channel. That is, the detrimental effect of mother's employment in the education production channel is smaller for girls than for boys when they are very young, and this difference decreases with age; the role model channel for girls is small when they are young but increases when they grow up. The combination of these two sources of asymmetry is able to reconcile the U-shaped estimates of γ.

4 Female Labor Supply and Educational Gender Gap in the U.S.

In this section we document a positive correlation between the educational gender gap in one generation and the labor force participation rates of married women in their mothers' generation when the generation was still in their childhood at their birth states in the United States. This positive correlation is statistically significant in the difference-in-differences model controlling for the cohort and state fixed effects.

4.1 Data and Variable Constructions

We use two large national representative data sets in the United States, the March Annual Demographic File and Income Supplement of the Current Population Survey (CPS) data and the Census data, both extracted from the Integrated Public Use Microdata Series (IPUMS) (Ruggles et al., 2010; King et al., 2010). We limit the sample to white females and males only.

The educational gender gap variables are calculated from the Census data.[23] Among all females, the fraction of females having college degrees (aged between 25 and 64) is calculated at the birth state level for each birth year cohort. This is defined as the state level college achievement for females. Similar state level fractions are calculated for males. The educational gender gap is defined as the difference in the college achievement between females and males at the birth state level for each cohort.

The labor force participation rates among all married females aged 25-64 (denoted as married FLFPR), are calculated from the CPS data since 1962 at the state level, averaging over a five-year interval.[24] We also calculate the married FLFPR for three different age groups: between 25 and 30, between 31 and 40, and between 41 and 64. Similar labor force participation rates are calculated for unmarried females (unmarried FLFPR), and for married or unmarried males (respectively, married MLFPR or unmarried

[23] Decennial years from 1950 to 2000 and every year from 2001 to 2012 (ACS).

[24] For example, for the year of 1962 we calculate the average married FLFPR between 1962 and 1966.

Table 6: Summary statistics of the U.S. data.

Variable	Mean	Std. Dev.	Min	Max
College gender gap	0.065	0.049	-0.108	0.367
Birth year	1974.6	8.029	1961	1987
Married FLFPR [25-64]	0.520	0.109	0.154	0.809
Married FLFPR [25-30]	0.523	0.161	0	0.862
Unmarried FLFPR [25-64]	0.730	0.086	0	1
Unmarried FLFPR [25-30]	0.818	0.117	0	1
Married MLFPR [25-64]	0.926	0.030	0.824	1
Married MLFPR [25-30]	0.973	0.022	0.750	1
Unmarried MLFPR [25-64]	0.844	0.062	0.5	1
Unmarried MLFPR [25-30]	0.901	0.071	0	1

Notes: The final sample has 1,225 observations. The college gender gap is the difference in the college achievement between females and males who are aged 25 or older. Married FLFPR: labor force participation rate (LFPR) of married females. Unmarried FLFPR: unmarried female LFPR. Married MLFPR: married male LFPR. Unmarried MLFPR: unmarried male LFPR. These LFPR are calculated at the state level, averaging over the five-year interval.

MLFPR).

We then merge the educational gender gap variables for each birth state and each cohort with the LFPR variables at that state when that cohort is one-year-old. Therefore in the merged data set, for each state and each cohort, we have the educational gender gap between females and males who were born in that state, and the LFPR of married/unmarried females/males in that state when that cohort is aged between one- and five-year-old. Table 6 presents summary statistics of the state level data used in this section.[25]

4.2 Difference-in-Differences Analysis

We run the following difference-in-differences (DID) regression model at the birth state level,

$$GAP_{sb} = \mathbf{LFPR}'_{st}\gamma + \delta_s + d_b + \epsilon_{sb}, \tag{1}$$

[25] At the national level, Figure 1 shows that to some extent the trend in the college gender gap in each cohort "coincides" with the trend in the LFPR of married women aged 25-30 when that cohort is aged between one- and five-year-old. Both of them increased rapidly between 1950 and late 1980s, and both level off after that. The gender gap in 1960 was the same as the gender gap in pre-1910 level, which Goldin et al. (2006) referred to as "homecoming." The educational gender gap for cohorts born in 1989 or later is not available yet, but based on years right before 1989 the trend seems leveling off or even decreasing, in both the Census data and the CPS data. The pattern is similar for the gender gap in More Than High School. Moreover, similar pattern is found in comparing the trend in the educational gender gap of each cohort and the trend in the LFPR of married women aged 31-40 when that cohort is aged between six- and ten-year-old.

where GAP_{sb} denotes the college gender gap among cohort b born in state s; $\mathbf{LFPR}_{st}$ is a vector of the LFPRs of married and unmarried women aged between 25 and 30, and for some specifications also the LFPRs of men in the same age group, in the birth state when cohort b is aged between one- and five-year-old; δ_s and d_b denote state and cohort fixed effects, respectively; ϵ_{sb} is the i.i.d. unobservable component which is assumed to be orthogonal to all other independent variables. In this difference-in-differences regression model, the effect of $\mathbf{LFPR}_{st}$ on GAP_{sb} is identified by the variation of their co-movement across different states. That is, the increases of the LFPR of married women and the resulting changes in educational gender gap take place earlier in some states than in other states.[26]

Panel A in Table 7 presents the DID regression results when we regress the college gender gap on female and male labor force participation rates at the state/cohort level. When included separately, Column 1 shows that the coefficient of the LFPR of married women (married FLFPR [25-30]) is positive and statistically significant at the 1% level; Columns 2-3 show that the coefficient of the LFPR of unmarried women or married men is statistically insignificant; and Column 4 shows that the coefficient of the LFPR of unmarried men is also positive and statistically significant at the 10% level. When LFPRs of both married and unmarried women (Column 5) or all four LFPRs (Column 6) are included in the same regression, all coefficients only change slightly and that of the LFPR of unmarried men becomes statistically insignificant.

Coefficient estimates in Panel A of Table 7 show that a one-percentage point increase in the married FLFPR is associated with 0.053-0.057 percentage points reduction in the college gender gap. Note that in our data, the married FLFPR increased from 20% in 1950 to 65% in 1985. As a back-of-the-envelope calculation, these coefficient estimates imply that the increase in the married FLFPR can result in a reduction of the college gender gap by about 2.4 to 2.6 percentage points. Since the gender gap in college was about -7.5% for the 1950 cohort and about 10.5% for the 1985 cohort, the effects of the increase in the married FLFPR can account for 13% to 14% of the narrowing and reversal of the college gender gap. This is a bit higher than the 8% in the Norwegian data.

If we use the average hours worked per week instead of the labor force participation rate as the measure of labor supply, the results are very similar as shown in Panel B in Table 7. Specifically, only the coefficient of the hours worked per week of married women is statistically significant in all model specifications.

Remark 2 *The identification of this difference-in-differences analysis comes from the fact that some states led others in both the change in the LFPR of married women and the change in the educational gender gap. However, if a third (unobserved) factor caused both changes but is not included in our analysis, then our estimate of the effect of the LFPR of married women on the educational gender gap would be biased, i.e. the omitted variable bias. We cannot completely rule out this possibility, but we argue it is not easy to find such a credible variable for at least two reasons. First, assume that there is such a variable which causes many changes including the two of our interest. Then it should also cause the change in the LFPR of unmarried women. This would lead to a similar correlation*

[26] We take the change in the LFPR of married women as exogenously given. Various explanations for the increase have been explored in the literature, for example Fogli and Veldkamp (2011) and Fernández and Wong (2014).

Table 7: Difference-in-differences regressions of college gender gap on LFPR or hours worked

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Independent Variables—Labor Force Participation Rate (LFPR)						
Married FLFPR [25-30]	0.056***				0.057***	0.053***
	(0.015)				(0.016)	(0.018)
Unmarried FLFPR [25-30]		-0.006			-0.006	-0.005
		(0.010)			(0.008)	(0.009)
Married MLFPR [25-30]			0.104			0.043
			(0.079)			(0.090)
Unmarried MLFPR [25-30]				0.027*		0.025
				(0.015)		(0.017)
Observations	1,171	1,152	1,171	1,144	1,152	1,133
R-squared	0.623	0.611	0.620	0.612	0.615	0.609
Panel B: Independent Variables—Hours Worked per Week						
Married Women [25-30]	0.058***				0.060***	0.057**
	(0.018)				(0.020)	(0.022)
Unmarried Women [25-30]		0.011			0.010	0.010
		(0.011)			(0.011)	(0.012)
Married Men [25-30]			-0.003			-0.024
			(0.025)			(0.023)
Unmarried Men [25-30]				0.009		0.009
				(0.008)		(0.008)
Observations	1,171	1,152	1,171	1,144	1,152	1,133
R-squared	0.622	0.611	0.618	0.611	0.615	0.609

Notes: Robust standard errors (clustered at the birth state level) in parentheses. There are 51 states.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
The dependent variable is the gender gap in college achievement. FLFPR stands for female labor force participation rate (LFPR) and MLFPR stands for male LFPR. The LFPR is calculated for each cohort at each birth state when they are aged 1-6. Independent variables include cohort fixed effects and birth state fixed effects.

between the LFPR of unmarried women and the educational gender gap. As Table 7 shows, this correlation is not statistically significant in all specifications. Second, some states might be likely to lead others in many frontiers, such as the state of California or New York. In unreported regression, we find that our results stay unchanged when we exclude these two states from the difference-in-differences analysis.

5 Children's Gender Effect on Mother's Labor Supply

We have documented so far the positive correlation between the mother's employment and the educational gender gap of her own children in Norway or the next generation in the United States. Our hypothesis is that this correlation is a result of the asymmetric production effect of mother's time input on boys and girls, and/or the differential role model effect of mother's employment on girls. In this section we first present a static model where a mother makes the labor force participation decision, aware of the asymmetric effects of her time input with her sons and daughters. Then we empirically test the model implication regarding a mother's labor supply decision and the gender composition of her children using both the U.S. data and the Norwegian data.

5.1 The Model of Mother's Labor Supply

Consider the labor supply decision for a mother with $n_g \geq 0$ girls and $n_b \geq 0$ boys. The mother's labor market option is denoted by her wage rate $w > 0$. Her unearned income is $y_0 \geq 0$, which could be her husband's income if she is married. Her time endowment is T, which can be allocated between working in the labor market, $t \in [0, T]$, and maternal time input to her children's education, $T - t \equiv s$.[27] The total income, $y_0 + wt$, is allocated between consumption c and the market-provided education input $k \equiv y_0 + wt - c$ for her children's education production. We assume that the education production functions of girls and boys are $e_g(s,k)$ and $e_b(s,k)$, respectively.

The mother solves the following maximization problem:

$$\max_{\{t,c\}} u(c) + \lambda \left[n_g e_g (T - t, y_0 + wt - c) + n_b e_b (T - t, y_0 + wt - c) \right] \tag{2}$$

where the $u(\cdot)$ is the utility from consumption and the $\lambda > 0$ denotes the weight in the mother's preference on her children's education relative to her own consumption. Note here we assume that the mother's time input is a public good for all of her children's education production. That is, we assume that a mother cannot separately allocate her time to the education production of each of her children.[28] The market-good input to the children's education production is also assumed to be a public good to all children within the same family.[29]

[27] We assume there is no requirement for minimum working hours. Minimum working hours do not change the results.

[28] Disciplining, for example, can be observed by all children in the same family, and therefore is likely to be a public good.

[29] An equivalent assumption is that both maternal time and market-good inputs to the education production of the children are evenly distributed among all children. The results will be identical under this assumption to those under the public good assumption.

The functions—$u(\cdot)$, $e_g(\cdot,\cdot)$ and $e_b(\cdot,\cdot)$—are twice continuously differentiable and concave in each argument. We make three additional assumptions about the education production functions $e_g(s,k)$ and $e_b(s,k)$. First, inspired by the empirical evidence of asymmetric effects of mother's employment on children's educational gender gap, we assume:

Assumption 1. (Marginal Productivity of Maternal Time Input is Higher for a Boy than for a Girl):

$$\frac{\partial e_g(s,k)}{\partial s} < \frac{\partial e_b(s,k)}{\partial s} \ \forall s,k \tag{3}$$

In other words, for any given level of mother's employment and market input, the detrimental effect of mother's employment is higher for boys.[30] Second, consistent with preceding empirical finding that the income effect is either statistically insignificant or gender symmetric, we assume:

Assumption 2. (Gender Symmetric Income Effect in Education Production Functions):

$$\frac{\partial e_g(s,k)}{\partial k} = \frac{\partial e_b(s,k)}{\partial k} \ \forall s,k \tag{4}$$

Note that Assumption 2 immediately implies:

$$\frac{\partial^2 e_g(s,k)}{\partial k \partial s} = \frac{\partial^2 e_b(s,k)}{\partial k \partial s} \ \forall s,k \tag{5}$$

Third, we assume the time input and the market good input in the education production functions are complementary or independent:

Assumption 3. (Maternal Time Input and Market Good Inputs are Complements or Independent):

$$\frac{\partial^2 e_i(s,k)}{\partial s \partial k} \geq 0, \ \forall s,k; \ i \in \{g,b\} \tag{6}$$

An example of the education production function satisfying the above three assumptions is:

$$e_i(s,k) = \alpha\delta_i f_1(s) + \beta f_2(k) + \gamma f_3(s,k), \ i \in \{g,b\}; \delta_g < \delta_b; \alpha,\beta > 0; \gamma \geq 0.$$

where $f_1(s)$ and $f_2(k)$ are increasing and concave functions; $f_3(s,k)$ is increasing and concave in each argument and has positive second-order cross derivative. In this example, time input and market input could be independent $(\gamma = 0)$ or complementary $(\gamma > 0)$.

Let $N \equiv n_g + n_b$, $\tilde{\lambda} \equiv \lambda N$, and let $n \equiv n_g/N$ be the fraction of girls. Then maximization problem (2) can be rewritten as:

$$\max_{\{t,c\}} V(t,c;n) \equiv u(c) + \tilde{\lambda}\left[n \cdot e_g(T-t, y_0 + wt - c) + (1-n)\cdot e_b(T-t, y_0 + wt - c)\right] \tag{7}$$

[30] However, note that we do not impose any restriction on the sign of the marginal productivity of mother's time input in either production function. The sign does not impact our analysis.

We have the following result:

Proposition 1 *Fixing $\langle \lambda, w, N \rangle$, the optimal labor supply t^* that solves problem (7) is monotone non-decreasing in the fraction of the girls, n.*

Proof. Using Assumption 2, we can immediately derive that

$$\frac{\partial^2 V(t, c; n)}{\partial t \partial c} = \tilde{\lambda} \left\{ \frac{\partial^2 e_b(s, k)}{\partial s \partial k} - w \frac{\partial^2 e_b(s, k)}{\partial k^2} \right\} > 0,$$

where the inequality comes from Assumption 3 and concavity of the production function $e_b(\cdot, \cdot)$. We can also show that

$$\frac{\partial^2 V(t, c; n)}{\partial t \partial n} = -\tilde{\lambda} \left[\frac{\partial e_g(s, k)}{\partial s} - \frac{\partial e_b(s, k)}{\partial s} \right] > 0$$

where the inequality follows from Assumption 1. Thus $V(t, c; n)$ is supermodular in (t, c) and has increasing differences in (t, n). The monotone comparative statics result of Milgrom and Shannon (1994, Theorem 5) applies and the optimal labor supply t^* is monotone non-decreasing in n. ■

The intuition of this proposition is straightforward. A mother is more productive in her children's education production if she has a higher fraction of boys among her children, *ceteris paribus.* Thus she would optimally choose to allocate more time at home if she has more boys.

Proposition 1 implies that, *ceteris paribus*, the reservation wage—the wage above which a mother will choose positive amount of labor supply—is a non-increasing function of the fraction of girls. In other words, the model predicts that mothers with more girls are more likely to participate in the labor force. Next we empirically test this model prediction using both the U.S. and Norwegian data.

5.2 Evidence: Children's Gender Composition and Mothers' Labor Supply

We now test the prediction from our model in the preceding subsection by examining the relationship between various measures of a mother's labor force attachment and the gender composition of her children. We run the following econometric specification:

$$y_i = \mathbf{X}_i' \beta + \mathbf{Z}_i' \gamma + \epsilon_i, \tag{8}$$

where y_i is a measure of i's mother's labor force attachment; $\mathbf{X}_i$ is the vector of the mother's and family's characteristics including the number of children, age, age squared, age at the first birth, and the year dummies (in the pooled sample only); $\mathbf{Z}_i$ is the vector of variables of interest regarding the gender of each child, or the gender composition of all children.

It is reasonable to assume, at least in developed economies, that a child's gender is exogenous, but the number of children is likely endogenous. The classic approach of using the gender composition of the first two children as an instrumental variable for the number of children in the family as in Angrist and

Evans (1998) is no longer applicable to our model.[31] However, the variable of our main interest is *not* the number of children, but the gender composition, which is exogenous at least in developed economies such as the United States or Norway. Reflecting on such concerns, we report results from both the standard OLS and the IV regressions for specification (8) where we use the same gender indicator as in Angrist and Evans (1998) as the "instrument" for the number of children. For comparison with the results in Angrist and Evans (1998), we restrict our analysis to families with at least two children in both regressions. For $\mathbf{Z}_i$, we will investigate two different specifications for our main interest. The first specification, reported in Table 9, includes the genders of the first two children, g_1 and g_2, as in Angrist and Evans (1998). The second specification, reported in Table 10, includes the fraction of girls among *all* children in the same family.

5.2.1 Evidence from U.S. Census Data

The data is constructed from the IPUMS Census data from 1950 to 2000. We select the sample to include only families with both biological parents present and with at least two children; we only select white mothers though we put no such restriction on the race of the spouses; the mother's age at the first birth has to be between 18 and 40.[32] Both parents have to be aged between 25 and 65, and all children have to be no older than 18. The children within the same family are ranked in a descending order according to their ages. Thus, the first child is always the oldest, while the second child is the second oldest, etc. Table 8 presents summary statistics of the pooled Census data (1950-2000) used in this subsection. The fraction of girls has mean 0.485 and standard deviation 0.321, which is not statistically significantly different from 0.5, and is almost constant from 1950 to 2000.

Table 9 reports the results for the effects of genders of the first two children on mothers' labor force participation status, from both the OLS regression (Panel A) and the IV regression (Panel B). Columns 1-6 have results separately for each Census year while Column 7 has that for the pooled sample. The dependent variable is a dummy variable indicating the labor force participation status of mothers. The independent variables, besides those reported, include mother's age and age squared, mother's age at the first birth, and the year dummies (Column 7 only). In the IV regressions, the F-statistic on the excluded instrument in the first stage is greater than 10 in all cases.[33]

Results in Table 9 show that, controlling for the gender of the first two children, the number of children has a negative effect on the mother's labor force participation in both regressions. Angrist and Evans (1998) report similar results, but for Census 1980 and Census 1990 only. Our results confirm theirs

[31] In Angrist and Evans (1998), the instrumental variable is a dummy variable indicating whether the first two children are of the same gender, SAMEGENDER= $g_1 g_2 + (1 - g_1)(1 - g_2)$, where $g_j = 1$ if j-th child is female and $g_j = 0$ otherwise. It is a linear combination of g_1, g_2 and $g_1 g_2$. For SAMEGENDER to be a valid instrument for the number of children, it requires that $g_1 g_2$ does not have direct effect on mother's labor supply. This holds only if the total effect of g_1 and g_2 on the mother's labor supply is additively separable, which is not the case in our model.

[32] Including observations with the mother's age at the first birth beyond this age period does not affect results materially.

[33] In unreported regressions, we found that the effects reported in tables 9, 10 and 11 are mostly robust when we control for mothers' educational attainments, except in 1950 Census year. In 1950, only 20% of mothers in our sample have recorded educational attainment information, and the effect of the fraction of girls is positive but becomes statistically insignificant in all cases when the mothers' educational attainments are included.

Table 8: Summary statistics of the pooled U.S. Census data, 1950-2000.

Variable	Mean	Std. Dev.	Min	Max
Year	1984	13.855	1950	2000
Girl 1st	0.482	0.500	0	1
Girl 2nd	0.488	0.500	0	1
Fraction of girls	0.485	0.321	0	1
Number of children	2.595	0.935	2	21
First two children same gender	0.503	0.500	0	1
Mother in labor force	0.538	0.499	0	1
Mother employed	0.517	0.500	0	1
Mother working full time	0.307	0.461	0	1
Mother age	35.747	5.915	25	58
Mother age at the first birth	24.657	4.381	18	40
Mother education	2.532	0.963	1	4
Age of youngest child	6.133	4.476	0	18
Age of oldest child	11.091	4.550	1	18
Average children age	8.674	4.241	0.25	18

Notes: The final sample has $2,014,236$ observations, among which $1,926,393$ has education information.

and show that this negative effect of the number of children on mothers' labor force participation is also present in the 1960, 1970 and 2000 Census years, as well as the pooled years. In 1950, the effect is not statistically significant in the IV regression. Table 9 also shows, with the exception of 2000 Census year, at least one of the effects, of the first child being a girl or the second child being a girl or both, are positive and statistically significant on mothers' labor force participation. These effects are almost identical in the OLS regression and the IV regression.

In Table 10, we report results using the fraction of girls among all children within the same family as a measure of the gender composition of all children. The effects of the number of children in both regressions stay essentially unchanged when compared with results in Table 9. In both the OLS and IV regressions, the effects of the fraction of girls are positive in all columns, and statistically significant in all Census survey years except 1960 and 2000. When we pool Census years from 1950 to 2000, the effects of the fraction of girls are positive and statistically significant in both the OLS and IV regressions.

In Table 11, we report the OLS results from the second regression specification as in Table 10, but we measure mothers' labor force attachment by whether the mother is employed or whether working full time, instead of her labor force participation status.[34] Full-time working is defined as working more than 1,260 hours per year and is only available since 1980.[35] The effect of the fraction of girls on mother's

[34]If one is employed, then the dependent variable is defined as one; otherwise zero, which includes the cases of unemployed and not in the labor force. Similarly, if one is working full time, then the dependent variable is defined as one; otherwise zero, which includes the cases of working part time, unemployed and not in the labor force.

[35]The hours worked per year is calculated as the product of hours per week and weeks per year. Since these variables refer to last year, in this specification we restrict the sample to those mothers whose youngest child is more than one-year-old at the interview time.

Table 9: Effects of genders of the first two children on maternal labor force participation, U.S. Census 1950-2000.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				Panel A: OLS regression			
Year	**1950**	**1960**	**1970**	**1980**	**1990**	**2000**	**1950-2000**
Number of children	-0.032***	-0.054***	-0.060***	-0.093***	-0.103***	-0.104***	-0.079***
	(0.001)	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)	(0.001)
Girl 1st	0.008***	0.004***	0.007***	0.007***	0.003***	0.001	0.004***
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Girl 2nd	0.004*	0.002	0.001	0.003**	0.002*	-0.000	0.002***
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	104,969	112,013	187,726	536,762	540,702	532,064	2,014,236
R-squared	0.037	0.058	0.058	0.055	0.048	0.050	0.147
				Panel B: IV regression			
Year	**1950**	**1960**	**1970**	**1980**	**1990**	**2000**	**1950-2000**
Number of children	-0.017	-0.070**	-0.101***	-0.084***	-0.074***	-0.066***	-0.074***
	(0.034)	(0.034)	(0.025)	(0.017)	(0.013)	(0.015)	(0.007)
Girl 1st	0.008***	0.004***	0.007***	0.007***	0.003***	0.000	0.004***
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Girl 2nd	0.004*	0.002	0.002	0.003**	0.002*	-0.000	0.002***
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	104,969	112,013	187,726	536,762	540,702	532,064	2,014,236
R-squared	0.034	0.056	0.048	0.055	0.046	0.047	0.147
First-stage F-stat	42.21	59.34	219.9	854.1	1642	1783	3456

Notes: Robust standard errors (clustered at the state level) in parentheses.
*** p<0.01, ** p<0.05, * p<0.1
The dependent variable is a dummy variable with 1 indicating participating in the labor force and 0 otherwise. Independent variables include mother's age, age squared, age at the first birth, and year dummies (column 7 only). Columns 1-6 are for each Census survey year while column 7 pools all years together.

Table 10: Effects of the fraction of girls on maternal labor force participation, U.S. Census 1950-2000.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
			Panel A: OLS regression				
Year	**1950**	**1960**	**1970**	**1980**	**1990**	**2000**	**1950-2000**
Number of children	-0.032***	-0.054***	-0.060***	-0.093***	-0.103***	-0.104***	-0.079***
	(0.001)	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)	(0.001)
Fraction of girls	0.010***	0.006	0.006*	0.010***	0.006***	0.001	0.006***
	(0.003)	(0.004)	(0.003)	(0.002)	(0.002)	(0.001)	(0.001)
Observations	104,969	112,013	187,726	536,762	540,702	532,064	2,014,236
R-squared	0.037	0.058	0.058	0.055	0.048	0.050	0.147
			Panel B: IV regression				
Year	**1950**	**1960**	**1970**	**1980**	**1990**	**2000**	**1950-2000**
Number of children	-0.020	-0.070**	-0.102***	-0.084***	-0.074***	-0.066***	-0.074***
	(0.033)	(0.033)	(0.025)	(0.017)	(0.013)	(0.015)	(0.007)
Fraction of girls	0.010***	0.006	0.007**	0.010***	0.006***	0.000	0.006***
	(0.003)	(0.004)	(0.004)	(0.002)	(0.002)	(0.001)	(0.001)
Observations	104,969	112,013	187,726	536,762	540,702	532,064	2,014,236
R-squared	0.035	0.056	0.048	0.055	0.046	0.047	0.147
First-stage F-stat	43.03	60.93	219.4	866.3	1643	1784	3486

Notes: Robust standard errors (clustered at the state level) in parentheses.
*** p<0.01, ** p<0.05, * p<0.1
The dependent variable is a dummy variable with 1 indicating participating in the labor force and 0 otherwise. Independent variables include mother's age, age squared, age at the first birth, and year dummies (Column 7 only). Columns 1-6 are for each Census survey year while Column 7 pools all years together.

Table 11: OLS results of maternal labor force attachment, measured by being employed or full-time employed.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent Variable in Panel A: 1 - Employed; 0 - O/W						
Year	**1950**	**1960**	**1970**	**1980**	**1990**	**2000**	**1950-2000**
Number of children	-0.031***	-0.053***	-0.059***	-0.092***	-0.102***	-0.104***	-0.078***
	(0.001)	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)	(0.002)
Fraction of girls	0.011***	0.007*	0.007*	0.010***	0.007***	0.001	0.006***
	(0.003)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.001)
Observations	104,969	112,013	187,726	536,762	540,702	532,064	2,014,236
R-squared	0.037	0.056	0.056	0.053	0.046	0.049	0.140
	Dependent Variable in Panel B: 1 - Full Time; 0 - O/W						
Year				**1980**	**1990**	**2000**	**1950-2000**
Number of children				-0.074***	-0.105***	-0.114***	-0.095***
				(0.002)	(0.002)	(0.002)	(0.002)
Fraction of girls				0.007***	0.004**	0.001	0.004***
				(0.002)	(0.002)	(0.002)	(0.001)
Observations				534,639	538,131	529,332	1,602,102
R-squared				0.047	0.043	0.046	0.072

Notes: Robust standard errors (clustered at the state level) in parentheses.
*** p<0.01, ** p<0.05, * p<0.1
Independent variables include mother's age, age squared, age at the first birth, and year dummies (Column 7 only). Columns 1-6 are for each Census survey year while Column 7 pools all years together.

Table 12: Effects of the fraction of girls on maternal labor force participation, Norway.

	(1) OLS	(2) IV
Number of Children	-0.028***	-0.019
	(0.001)	(0.012)
Fraction of Girls	0.002*	0.002*
	(0.001)	(0.001)
Observations	16,429,544	16,429,544
R-squared	0.074	0.073
First-stage F-stat		1456

Notes: Robust standard errors (clustered at the individual level) in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.
The dependent variable is a dummy with 1 indicating working and 0 otherwise. Independent variables include the mother's age, age squared, age at the first birth, and year dummies. Column 1 reports the result from the OLS regression while Column 2 reports the result from the IV regression.

employment status is significantly positive except for Census year 2000 where the effect is positive but statistically insignificantly. The effect on mother's full time working status is significantly positive in 1980 and 1990, and insignificantly positive in 2000. Both effects are positive and statistically significant in the pooled sample. The estimates do not change materially in the IV regression. Also note that in Table 11 the number of children reduces mothers' labor force attachment in all the Census years, with almost identical estimates as in Panel A of Table 10. The results are similar as Panel B of Table 10 in the IV regression. This is again consistent with the findings in Angrist and Evans (1998). If we run Probit analysis with or without instrumental variable for specification (8), the results in this subsection — signs and statistical significance — are very similar in most cases.[36]

5.2.2 Evidence from Norwegian Data

Table 12 presents similar results using Norwegian registry data. Here the dependent variable is mothers' working status inferred from earnings history. Consistent with the U.S. data in Table 10, we control for mother's age, age squared, age at the first birth, and the year dummies in the regressions. The results from both the OLS regression and the IV regression are again consistent with the prediction summarized in proposition 1 in that, *ceteris paribus*, mothers are more likely to be working if they have a higher fraction of girls among their children.

[36]The only difference is that in the probit analysis (without IV) the effect of the "Girl 2nd" is negative and statistically significant when the dependent variable is mother's labor force participation status. When the SAMEGENDER is used as IV for the number of children, the results of the probit analysis are much similar as panel B in table 9. Probit analysis with "Fraction of girls" yields similar results as tables 10 and 11.

5.3 Alternative Explanations

We have found that mothers with higher fractions of girls among their children have higher labor market participation rates. In this subsection we discuss two plausible alternative explanations for such empirical findings and evaluate their potentials and limitations.

5.3.1 Girls Are More Expensive

If it is more financially expensive to raise girls than boys, then families with higher fractions of girls might be more financially demanding and therefore mothers might be more likely to be forced to work for extra income. Girls could be more expensive for several reasons. For example, traditionally the bride's family pays for the larger share of the wedding costs than the groom's family.

This explanation could potentially generate our findings, but it would also generate another implication. If girls are more expensive and children are a normal good, then families with first child being a girl should have fewer children. This is the *"differential cost hypothesis"* in Dahl and Moretti (2008) and they find the opposite in the U.S. data. Therefore this hypothesis is unlikely to be the driving force of our findings.

5.3.2 Higher Bargaining Power of Mothers with Sons

In countries with son preference, like the U.S. (Dahl and Moretti, 2008) and China (Li and Wu, 2011), mothers with sons would likely have higher bargaining power within families than those with daughters. Then in a standard collective labor supply model as in Chiappori (1992) higher bargaining power would imply less labor supply (Voena, 2015).

In this alternative story, giving birth to a son in a family with son preference would boost a mother's bargaining power. This bargaining power should not fade over time as long as the son is present. In other words, having a daughter would lower a mother's bargaining power for a long period of time. This implies that the effect of the child's gender on mother's employment should not vary with the child's age since the bargaining power is stable over time. On the contrary, in our model, the asymmetric effects of mother's employment on children's education production vary at different age stages of her children, as documented in Section 3.

To see if effects of children's genders on mother's employment vary over time, we include an interaction term between the fraction of girls and children's ages in the regression model,

$$y_i = \mathbf{X}_i'\beta + \mathbf{Z}_i'\gamma + \xi t_i + t_i\mathbf{Z}_i'\pi + \epsilon_i, \tag{9}$$

where t_i measures children's ages in the family. We use three different measures for t_i: the age of the youngest child, the age of the oldest child, and the average age of all children.[37] In this specification, the coefficient ξ represents the effect of the children's age on the mother's labor supply while the coefficient

[37]When the age of the oldest child is included (together with the mother's age), the age at first birth is dropped out of the independent variables due to collinearity.

π represents the age trend in the gradient of the mother's labor supply on the fraction of girls. In the bargaining power hypothesis both ξ and π should be zero. However, the effect of the children's age on the mother's labor supply might vary for other reasons, which would likely generate a positive ξ.[38] But even in this case, the effect of the bargaining power should not vary over the child's age, and therefore π should be zero. On the other hand, in our asymmetric effect hypothesis the π generally should not be zero. The sign of the coefficient π tells us whether the differential effect of children's gender on mother's labor supply widens (if $\pi > 0$) or shrinks (if $\pi < 0$) with the children's age. As discussed in Subsection 3.3 based on results from Norwegian registry data as in Table 4, the asymmetry is stronger when a child is either young (the asymmetric production channel dominates) or old (the role model channel dominates). For this reason, we conduct two different specifications. In the first one, we use the children's age for t_i as described earlier. In the second one, instead of the direct measure of the age, we include two dummy variables indicating whether this age is in the age interval of either [0,5] or [12,17].

As previous literature finds that the U.S. is likely a country with son preference (Dahl and Moretti, 2008), we conduct the econometric specification (9) in the U.S. Census data only. The estimation results from the pooled 1950-2000 U.S. Census data are presented in Table 13. There are several observations worth noticing here. First, we look at the estimates of ξ only, which reflects the effect of the children's ages for those mothers with sons only, i.e., when the fraction of girls is zero. Results in Table 13 show that the effects of the age of the youngest child or the oldest child or the average age of all children on mothers' labor force participation are all positive and statistically significant (Columns 1, 3 and 5). More importantly for our purpose, however, is that the estimates of the interaction terms between "Fraction of girls" and all of the three variables—"Age of the youngest child," "Age of the oldest child" or "Average age of children"—are also positive and statistically significant. The results are similar in the IV regressions and are robust if including the mother's educational attainment. Thus, the evidence suggests that the positive gradient of the mother's labor force participation rate on the fraction of girls increases with her children's age.

When the age interval dummies are used as shown in Columns 2, 4 and 6, the interaction between the fraction of girls and the dummy indicating young age is statistically insignificant in all three cases, while the interaction with the dummy indicating old age is positive and statistically significant. This reflects that this asymmetry stays relatively flat during a child's first eleven years but widens when the child reaches 12- to 17-year-old. This pattern can be explained by a combination of two forces: an asymmetric production channel where the asymmetry does not shrink rapidly with the child's age, and a role model channel where older girls are impacted more significantly than younger ones.

In summary, these evidences are consistent with our asymmetric effect theory and harder to square with the bargaining power hypothesis.

[38] For example, a child attending school would demand less time from the mother.

Table 13: OLS results of maternal labor supply, with the fraction of girls interacting with children ages, pooled U.S. Census [1950-2000].

	(1)	(2)	(3)	(4)	(5)	(6)
Number of children	-0.047***	-0.058***	-0.079***	-0.074***	-0.050***	-0.070***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)	(0.001)
Fraction of girls	0.003*	0.005***	-0.003	0.002	-0.000	0.005***
	(0.002)	(0.001)	(0.003)	(0.002)	(0.002)	(0.001)
Age of youngest child/10	0.179***					
	(0.006)					
Fraction of girls ×	0.005**					
age of youngest child/10	(0.002)					
Fraction of girls ×		-0.000				
youngest child aged [0,5]		(0.002)				
Fraction of girls ×		0.005*				
youngest child aged [12,17]		(0.003)				
Age of oldest child/10			0.216***			
			(0.007)			
Fraction of girls ×			0.008***			
age of oldest child/10			(0.002)			
Fraction of girls ×				0.003		
oldest child aged [0,5]				(0.003)		
Fraction of girls ×				0.006***		
oldest child aged [12,17]				(0.002)		
Average children age/10					0.356***	
					(0.012)	
Fraction of girls ×					0.008***	
average children age/10					(0.002)	
Fraction of girls × average						-0.002
children age is in [0,5]						(0.002)
Fraction of girls × average						0.005***
children age is in [12,17]						(0.002)
R-squared	0.154	0.153	0.147	0.142	0.154	0.149

Notes: Robust standard errors (clustered at the state level) in parentheses. There are 2,014,236 observations in each column.
*** p<0.01, ** p<0.05, * p<0.1
Independent variables include mother's age, age squared, age at the first birth (Columns 1, 2, 5 and 6 only), the year dummies, and the dummies for youngest (Column 2), oldest (Column 4), or average (Column 6) age in [0,5] and [12,17].

6 Conclusion

In this paper we argue that the increasing labor force participation rate of married women since 1950 accounts for about 13% (or 8%) of the narrowing and reversal of the educational gender gap among their children's generation in the U.S. (or Norway). In particular, we highlight that the connection between the mothers' employment and their children's educational gender gap is consistent with the asymmetric effect of mothers' time input in their children's education production and the differential role model effect.

In the panel data from Norway, we find direct evidence for the asymmetric effects of mothers' employment on their own children's educational gender gap. In the cross-sectional data from the U.S., we find that the college gender gap in one generation is positively correlated with the labor force participation rate of married women in their mothers' generation during their childhood. We argue that this correlation is a result of asymmetric effects of mothers' employment on their children's educational achievement, which benefits daughters more than sons.

We then further derive the empirically testable implication of this mechanism on a rational, well-informed and altruistic mother's labor supply. Specifically, in the presence of this gender-asymmetric effect, the mother's labor supply should be higher if she has a higher fraction of girls, after controlling for the total number of children. We find supportive evidence for this implication in both the U.S. data and Norwegian data. We also discuss and rule out alternative explanations about our findings. Taken collectively, the weight of our evidence supports the mechanism that mothers' employment has a gender-specific asymmetric effect on their children's educational achievement. As such, our results highlight a potentially unintended consequence of the increase in mothers' employment on the educational gender gap of their children, which will lead to a potentially important inter-generational "feedback" loop.

We believe that our results regarding mothers' employment and their children's educational gender gap are novel and contribute to a complementary and better understanding of the causes of the narrowing and reversal of the educational gender gap since 1950s. There are several interesting areas for future research as a result of the findings in this paper. First, as we note in Figure 1, the labor force participation rate of married women has stabilized since mid-1990s. Our theory would predict that for cohorts born after mid-1990s the changes in the educational gender gap would slow down as well. This prediction can be checked as new data becomes available. Second, in this paper we have taken the increase in married women's labor force participation rates as given and studied its effects on the educational gender gap of their children's generation. It will be interesting to study an overlapping-generation equilibrium model where mothers make labor supply decisions taking into account such an asymmetric effect on their children's education production, at least when they are young, and their children subsequently make college attendance decisions and labor supply decisions etc.[39] Third, the effect of mothers' employment on their children's educational gender gap may also have implications on the fertility choice. In countries with son preference, the increase in female labor force participation rates might reduce the

[39]See Eckstein, Keane and Lifshitz (2014) for a related attempt, though they focus on marriage, and do not consider the impact of mothers' employment on children's education.

fertility more than in other countries without such preference bias toward sons.

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