Mothers' Employment, Parental Absence and Children's Educational Gender Gap*

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Abstract

This paper analyzes the connections between three concurrent trends since 1950: (1) the narrowing and reversal of the educational gender gap; (2) the increasing labor force participation rate (LFPR) of married women; (3) the rising incidence of children living with only one parent. We hypothesize that the education production for boys is more adversely affected by a decrease in parental time input as a result of increasing maternal employment or parental absence. Therefore, a pronounced increase in the labor force participation rate of married women as well as the rising incidence of absent fathers may narrow and even reverse the educational gender gap in the child generation. We use micro data from the Norwegian registry to directly show that the parental employment during their children's childhood has an asymmetric effect on the educational achievement of their 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 *mother* generation as well as the incidence of parental absence (mostly absence of fathers) at the U.S. state level. We then propose a model that generates a novel prediction about the implications of these asymmetric effects on parental labor supply decisions and find supporting evidence in both the U.S. and Norwegian data.

Keywords: Female Labor Force Participation; Absent Fathers; Educational Gender Gap.

JEL Classification Codes: I2, J2.

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1 Introduction

This paper analyzes the connection among three concurrent trends since the end of World War II in the United States and many other developed countries: (1). the narrowing and reversal of the educational gender gap; (2). the increased labor force participation rate (LFPR) of married women; and (3). the rising incidence of children living with one parent. The solid line in Figure 1 plots the gender gap in the four-year college completion rate for whites (on the left scale) for cohorts born in 1950 to 1990 in the United States.[1] It shows that the gender gap in college completion for the cohort born in 1950 was negative—female college completion rate was 5.5 percentage points *lower* than males. However, the college gender gap has narrowed rapidly since then, in fact it disappeared within a decade, and reversed for cohorts born after 1960. For cohorts born in late 1980s, female college completion rate is nearly 11.2 percentage points *higher* than male's.[2] The long-dashed line in Figure 1 plots the trend in the labor force participation rates (LFPR) of married white women (on the first right scale) in each cohort's mother generation when the cohort was between 0 and 5 years old (the preschool period). It shows that the LFPR of married white women in their mother generation has more than tripled, rising from below 20% for the 1950 cohort to nearly 70% in late 1980s.[3, 4] The short-dashed line in Figure 1 plots the rising incidence rate of white children living with only one parent (on the second right scale) when the cohort was between 0 and 5 years old. Notably, the rate has more than doubled from 2.5% for the 1950 cohort to around 6.9% for cohorts born in late 1980s.[5]

In this paper, we argue that the latter two trends contribute to the first one via the mechanism that both the increasing LFPR of mothers and the increasing incidence of children living with one parent only have *gender asymmetric* effects on male and female children's educational achievement. The gender asymmetric effects of maternal employment and parental absence can be further decomposed into an effect from the *reduced parental time input* into the children's educational production and a *role model* effect from maternal employment.[6] We argue that both the time input effect and the role model effect are gender asymmetric and are in favor of girls. First, parental time input is an important factor in the children's education production, and it decreases with the increasing LFPR of mothers and the increasing

[1]The college completion rates are calculated from individuals 25 years or older, separately for females and males; gender gap in college completion is equal to the female college completion rate minus the male college completion rate. If we were to alternatively define college achievement as "Some College or College Completion," the trend of the gender gap is similar. In the rest of the paper we use *educational gender gap* to refer to *gender gap in college completion*, or simply, *college gender gap.*

[2]This trend is well documented by many other researchers (**???**). **?** 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 because the educational attainments for cohorts born between 1910 and 1950 are likely affected by, to a significant degree, the World War II and Korean War GI bills (**??**), as well as the Vietnam War (**?**).

[3]In contrast, the labor force participation rates of either unmarried women or men (either married or unmarried) in the cohort's parent generation were relatively constant after the World War II.

[4]Such increasing trend is similar across different education and age groups. This trend has been documented and studied in **?** and **?**, among many others.

[5]The incidence of children living with only one parent is calculated as the unconditional probability of one person having at least one child and not living with spouse. It includes the following cases: single, married but an absent spouse, separated, divorced, and widowed. The majority of one parent families have an absent father. This trend is also documented in **?**.

[6]Parental employment may also have an income effect. Most of the literature, however, finds it to be either insignificant or symmetric across the child's gender.

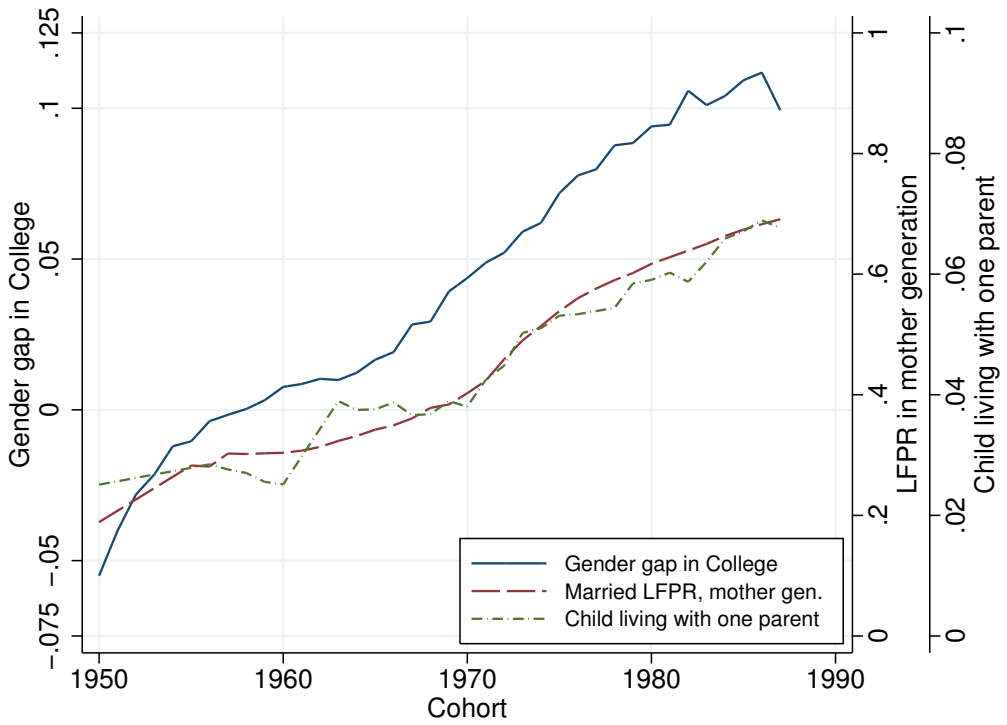

Figure 1: Three Concurrent Trends in The United States Since 1950

Notes: The solid line represents the gender gap in the four-year college completion rate among whites (females versus males) for each birth cohort. College completion rates are calculated from individuals aged 25 or older in the U.S. Census data. The long-dashed line represents the labor force participation rate (LFPR) of married women when the cohort is aged between 0- and 5-year-old (which we refer to as the cohort's mother generation). The dot-dashed line represents the incidence rate of children between 0- and 5-year-old living with only one parent. The LFPR and the incidence rate of children living with one parent since 1962 are calculated from the U.S. CPS data, while those for decennial years in 1950 and 1960 are from the U.S. Census data.

incidence of one absent parent; however, reductions in parental time input has less a detrimental effect on the girls than on the boys.[7] Second, a parent working in the labor market has a role model effect on the children, and to the extent that the role model effect is stronger for the children of the same gender as the working parent, the increasing maternal employment after the World War II provides role models more for the girls than for the boys. Because boys have their own role models from their fathers whose LFPR stayed high throughout the post-World War II period, girls' role model effect from their mothers' employment at most reduces their disadvantages relative to the boys. Thus, the gender asymmetric time input effect is the key to explain the *reversal* of the educational gender gap.

We find empirical evidence supporting our hypothesis in the data from Norway and the United States. Using individual level data from Norway, we find that the parental employment has a gender

[7]The existing evidence is reviewed in Section 2, and our own additional evidence is provided in Section 3.

asymmetric effect on their children's educational attainment in favor of their girls.[8] The gender asymmetric effect is statistically significant, even after controlling for gender-specific cohort fixed effects and family fixed effects. The positive effect of maternal employment on educational gender gap is also confirmed in the analysis using the state level aggregate data in the United States: their educational gender gap of a particular generation is positively correlated with the labor supply of married women of their mothers' generation in their birth state when they were still aged between 0 and 5. We also find that the educational gender gap is positively correlated with the incidence of children living with only one parent (with the absent parent mostly the father). Both positive correlations are statistically significant in the difference-in-differences model controlling for the cohort and state fixed effects.

The above results suggest that, at the aggregate level, the pronounced increase in the LFPR of married women and the rising incidence of children living with only one parent may narrow, and even reverse, the educational gender gap in their children's generation. A back-of-the-envelope calculation based on our empirical results suggests that, in the United States, the increase of the LFPR of married women can account for 17%-18% of the observed narrowing and reversal of the educational gender gap for cohorts born between 1950 and 1986, and the rising incidence of children living with only one parent can account for an additional 4.7%-5.5% of the changes in the educational gender gap. Similarly, our estimates suggest that in Norway the increase of the LFPR of married women can account for 9% of the observed changes in the educational gender gap for cohorts born between 1967 and 1986.

We further provide auxiliary evidence for our mechanism. Using a parsimonious model, we derive a novel prediction on the implication of the gender asymmetric effects, namely, conditional on having the same number of children, maternal labor supply should be higher if they have a higher fraction of girls. We find empirical evidence supporting this prediction in both the U.S. Census data and the Norwegian registry data.

The remainder of the paper is organized as follows. In Section 2, we review the related literature. In Section 3, we present the estimates of the gender asymmetric effects of parental employment or parental absence on their children's educational achievement, both in the Norwegian micro data and the U.S. state level aggregate data. In Section 4, we propose a simple model of parental labor supply featuring the gender asymmetric effect of parental employment on children's educational production. We test the empirical prediction derived from the model using both the U.S. Census data and the Norwegian registry data; we also discuss alternative explanations. Finally, in Section 5, we conclude.

2 Related Literature

First, this paper is related 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

[8] Interestingly, we find that the gender asymmetric effect holds for both mother's and father's employment. The employment for married men did not change much in Norway and the United States during our study period, hence it is unlikely an important driver for the changes of the educational gender gap at the aggregate level, although it is as important as mother's employment at the individual level.

have documented the dynamic changes of the gender gaps in high school graduation and dropout, as well as in college enrollment and graduation (e.g., **????**). 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 **?**, for a comprehensive review).[9] The second strand of the literature argues that the reversal of the college gender gap is due in part to the behavioral and developmental differences between girls and boys—girls have lower costs to prepare and attend colleges than boys (e.g., **?**). Girls are found to have higher non-cognitive skills (**??**), lower rates of Attention Deficit Hyperactivity Disorder (ADHD) (**?**), lower incidence of arrests and school suspension (**?**), or higher elasticity of college attendance (**?**).[10] The 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 in the educational gender gap (e.g., **???**). The mechanism we emphasize in our paper provides a novel and complementary explanation for the narrowing and reversal of the educational gender gap.

Second, our paper is also related to the literature investigating the asymmetric effects of parents' or mothers' time input on the development of their children's cognitive and non-cognitive skills. For example, **?** analyses a rather complete literature (68 papers) published between 1960 and 2005 on testing the effects of concurrent maternal employment on children's cognitive or academic achievement. They find that maternal employment has more positive effects for girls. **?** 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. Many other researchers (see e.g., **???????**) also find that maternal employment or the reduction of parental time input (due to separation or divorce) has a more detrimental impact on boy's cognitive and non-cognitive skill development.

Third, our paper is also related to a new but growing literature that suggest a gender asymmetric effect of social and economic disadvantages. **?** document a larger college gender gap in families with either low-educated or absent fathers. **?** discover a large gradient of family disadvantage in the gender gap of the behavioral and academic performance; they find boys who are born to unmarried and less-educated mothers or are raised in families with low socioeconomic status perform much worse than girls with similar backgrounds, even after controlling for family fixed effects. An earlier literature has found that boys are more vulnerable to family disruption such as divorce than girls (e.g. **???**). Most family disruption results in absent fathers. **?** find boys raised by single mothers are at a higher risk of having behavioral problems. They argue that boys are more responsive to parental inputs which tend to decrease dramatically among broken families. In the Swedish registry data, **?** find that boys raised in single-parent families have higher risk of mortality from all causes including suicide, morbidity and injury than girls. The contribution of our empirical analysis to link the increasing LFPR of married

[9] However, **?** finds that after controlling for the top coding bias in the CPS data the difference in the college wage premium between women and men is statistically insignificant from 1990s. **?** also find that the benefits of attending college are not higher for women than for 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 (**?**). See **?** for a comprehensive review of the biological and psychological hypotheses for the sex differences in cognitive abilities.

women and the rising incidence of single parenthood to the narrowing and reversal of the educational gender gap, relying on the mechanism of the gender asymmetric effects of parental time input.

Finally, 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. **?** 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 **?**. 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.

3 Gender Asymmetric Effects of Maternal Employment

3.1 Micro Evidence from Norway

We first present micro evidence that shows a positive correlation between the mother's and father's employment and their own children's educational gender gap in the 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; similar results are found for father's employment.

3.1.1 Norwegian data

We use administrative data for Norway which covers the full Norwegian population. In order to observe parental employment as well as educational outcomes for the children we condition on those being born between 1967 and 1988. Furthermore we restrict the sample to nuclear families where biological parents are still married to each other at the child's 18-year birthday. We also exclude families where at the child's birth the mother's age is below 16 or above 45, or the father's age is below 16 or above 65. In total, the sample consists of 816,617 children from 433,427 families. As a measure of the labor supply or the employment status, 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.[11, 12]

[11] In 2015 this base unit equals 90,068 NOK or 10,860 USD. We use "labor supply" and "employment" interchangeably throughout the paper.

[12] It is worth noting that this construction is primarily a measure of labor supply on the extensive margin. Due to limitation of the data, we do not have measures of labor supply on the intensive margin. Nevertheless, our main results are robust to

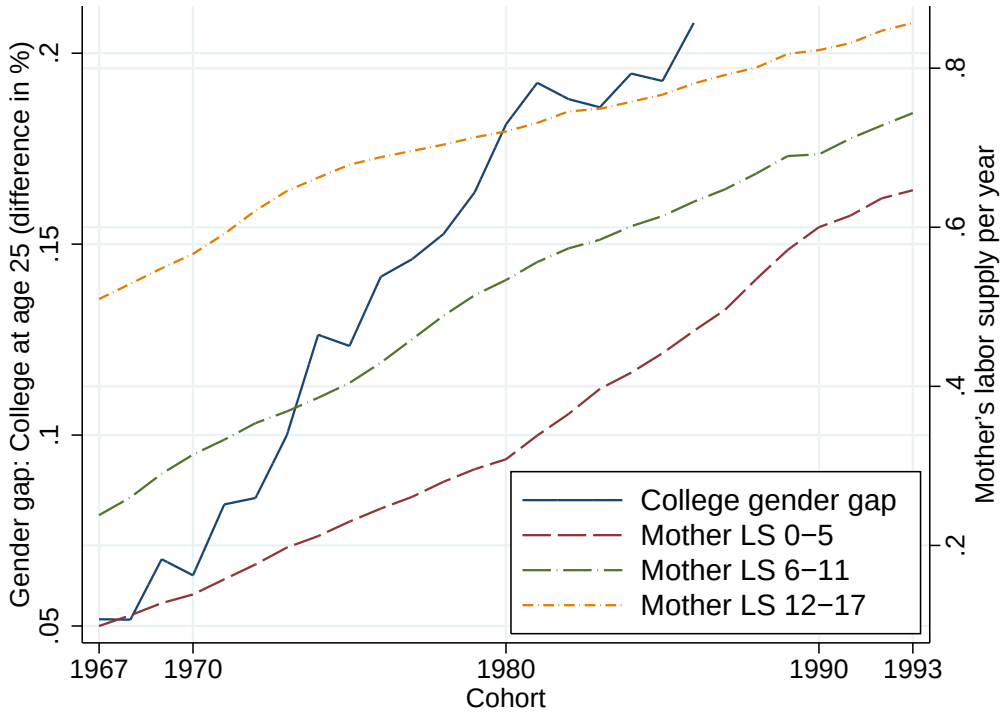

Figure 2: College Gender Gap and Mothers' Employment, Norway 1967-1993

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 (left scale of the y-axis). 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 (right scale of the y-axis).

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 measured as the difference between female and male rates of college graduation. It has been increasing since 1967.[13] There has also been a dramatic increase in mothers' employment, from 10% in 1967 to 65% in 1993 for mothers with young children (0-5 years old). On the other hand, fathers' employment rate has been rather constant, varying between a narrow band of 92% and 94%. Overall, Norwegian data presents similar patterns as the U.S. data shown in Figure 1. Table A10 in Appendix A presents summary statistics of Norwegian registry data used in this paper.

different earnings thresholds. For instance, when we change the threshold of earnings from one base unit to two base units, the results do not change qualitatively, as discussed in Subsection 3.1.2.

[13]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.

3.1.2 Gender Asymmetric Effects of Parental Employment

To ascertain whether parental employment may result in asymmetric effect on the educational achievement of their children depending on the child's gender, we estimate the following equation:

$$y_i = \mathbf{M}_i'\beta + \left(g_i \times \mathbf{M}_i'\right)\gamma + \mathbf{W}_i'\xi + \mathbf{X}_i'\delta + \nu_{g_i,c_i} + \eta_f + \epsilon_i, \tag{1}$$

where i indicates a child, 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 and father's employment at various intervals of i's youth; $\mathbf{W}_i$ is a vector of family earnings; $\mathbf{X}_i$ is a vector of other demographics including the mother's age and the father's age when the child is born and a dummy for the birth order; ν_{g_i,c_i} indicates the gender specific cohort fixed effect; η_f is a family fixed effect; ϵ_i is the i.i.d. unobservable component which is assumed to be orthogonal to all other independent variables.[14]

In order to investigate whether the timing of parental employment during the child's youth matters for the educational achievement, we will construct measures of mother's and father's employment for different age intervals of the child, and run separate regressions. For example, in the first column, the parental employment is measured as the employment rate during their child's entire childhood period, from 0 to 17-year-old.

Discussion

- To control for time-invariant unobservable (of course, also observable) family characteristics that may affect both parent's employment and children's college achievement, we include a family fixed effect, η_f.[15] The inclusion of the family fixed effect η_f implies that the identification of the effects of mother's and father's employment on their children's college gender gap comes from the variation across children of same or opposite genders within the *same* family.[16]

- strictly speaking, without exogenous variation in the parental employment, the results from the DID model are still correlations.[17]

- We do not include g_i as a separate term in (1) because the gender specific cohort fixed effect ν_{g_i,c_i} absorbs the effect of the girl dummy.

- The dummy variable for the birth order is included to control the birth order effect, as the literature has documented that higher birth order has a significantly negative effect on the education (e.g.,

[14] A *family* is defined by a pair of a biological mother and a biological father who stay married to each other until the youngest child's 18-year-old birthday.

[15] For instance, family environment, as well as mother's and father's education and ability (e.g., **?**).

[16] Mother's and father's employment are continuous variables so the difference between different children within the same family is sufficient for the identification of β and γ.

[17] 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 or father's employment and the children's college gender gap, although it is not easy to find such variables.

?).

We include the spousal earnings, instead of family earnings, in the vector $\mathbf{W}_i$ for two reasons. First, we are mainly interested in the *total* effect of mother's or father's employment, which consists of a time effect and an income effect. Since our measure of parental employment is derived from earnings, there is high correlation between employment and earnings. Thus it is difficult to separately identify the time effect and income effect. Second, if including both employment and earnings of mothers, the interpretation for the coefficient of mother's employment is not clear, since varying employment while holding earnings constant for the same individual is difficult in our data. For these reasons, we estimate the total effect of mother's employment using father's employment and earnings as control variables, and estimate the total effect of father's employment using mother's employment and earnings as control variables, in different regressions.

The DID results are presented in Table 1, with Panel A for the total effects of mother's employment and Panel B for father's employment. Each column presents a separate regression. In the first column, the parental employment is measured as the employment rate during their child's entire childhood period, from 0 to 17-year-old. The correlation between the boy's college achievement and the mother's employment is positive but statistically insignificant (Panel A). In contrast, father's employment is negatively correlated with his son's college achievement and such correlation is statistically significant at the 1% level (Panel B). More interestingly, there is an additional positive and significant correlation between the girl's college achievement and either parental employment, measured by the coefficient of the interaction term between the mother's (Panel A) or the father's (Panel B) employment and the girl dummy. This additional correlation is almost 50% higher for the father's employment than the mother's.

We then separate the parental 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 DID regressions. The results reported in Columns 2-4 reveal that the boy's college achievement is negatively but statistically insignificantly correlated with his mother's employment before his high school, after which the correlation is also negative but becomes statistically significant. At the same time, the additional correlation between the mother's employment at all age groups and her own daughter's college achievement is all positive and statistically significant. On the other hand, father's employment, compared to mother's, has more detrimental effect on the son's college achievement, as shown in Panel B. The negative effects are all statistically significant. The additional positive effect of father's employment on the girl's college achievement is also statistically significant and larger than that of mother's employment except during the child's preschool ages, where the effect is positive but statistically insignificant.

The DID model shows that there appears a positive correlation between both the mother's and father's employment and their own children's gender gap in college achievement. As a back-of-the-envelope calculation, 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 0.0145 in the gender gap. This is about 9% of the increase in the college gender gap in Norway, which increases from 0.0518 in 1967 to 0.2079 in 1986. The father's employment stays relatively flat during the same period.

It is worth noting that the effect of spousal earnings is statistically insignificant in each of the three different age groups, except that it is negative and statistically significant in the pooled case (Column 1) or during the child's high school for the mother's earnings (Column 4 in Panel B). Recall the income effect in this DID model refers to the effect of earnings innovations rather than levels. **?** also find that the effect of changes in income is not significant in Norwegian data.

Our results of parental employment on their children's college gender gap are quite robust to various specifications of the DID model. The estimates of γ in (1) do not change materially in all four variants: when we include the family earnings instead of spousal earnings; when we include an interaction term between earnings and the girl dummy; when we include the logarithm of earnings instead of the earnings level; when we change the labor supply measure from one base unit of earnings to two base units.[18] For example, in the last variant, we assume one individual is working in a certain year if she or he has at least two—instead of one in our baseline model—base units of earnings in the Norwegian social security system. The estimation results of the DID model are presented in Table A11 in Appendix A.

Remark 1 *We would like to emphasize that the estimate of the additional effect of mother's and father's employment on girls relative to boys, namely γ in Eq. (1), 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 parental employment and their children's education but not captured by family fixed effects in the DID model; therefore the estimates of the effects of parental employment on their child's educational achievement would be biased. However, as long as those time-varying factors are symmetric for boys and girls, such bias may be reduced or cancel out in our DID model. In other words, as long as the effects of those omitted time-varying variables are independent of the child's gender, then the bias on the estimate of the effect of mother's and father's employment on children's educational gender gap, namely γ, is likely to be limited.*

It is worth noting that in the regression results in Table 1, the overall effect of mother's employment on her girl's college achievement is positive and statistically significant. Existing literature has found mixed estimates for the effect of mother's employment on her girl's cognitive and non-cognitive outcomes. **?** has a detailed review on this bifurcated effect of mother's employment on boys and girls.[19] In contrast, the overall effect of father's employment on his girl's college achievement is statistically insignificant.

3.2 State Level Evidence from the U.S.

In this subsection we document three positive correlations at the state level in the United States. The first two are the positive correlations between the educational gender gap in one generation and the labor force participation rates (LFPR) of married women in their mothers' generation or the LFPR of married men in their fathers' generation, when the generation was still in their childhood at their birth states. The third one is the positive correlation between the educational gender gap and the incidence of

[18] We try both $\log(\mathbf{W}_i)$ and $\log(1000 + \max\{0, \mathbf{W}_i\})$, where $\mathbf{W}_i$ is the earnings which could be negative (due to business loss, for example). The model with $\log(\mathbf{W}_i)$ has fewer observations.

[19] See the third and fourth paragraphs in page 79 of **?**.

	(1)	(2)	(3)	(4)
Child's Age Interval	$[0, 17]$	$[0, 5]$	$[6, 11]$	$[12, 17]$
Panel A: Total Effects of Mother's Employment				
Mother's Employment	0.006	-0.007	-0.007	-0.012**
	(0.012)	(0.006)	(0.005)	(0.006)
Mother's Employment×Girl	0.042***	0.025***	0.024***	0.037***
	(0.006)	(0.005)	(0.005)	(0.005)
Father's Employment	-0.041	-0.020	-0.026*	-0.015
	(0.026)	(0.015)	(0.015)	(0.013)
Father's Employment×Girl	0.060***	0.019	0.040***	0.045***
	(0.016)	(0.015)	(0.013)	(0.010)
Father's Earnings /1000	-0.349*	-0.150	-0.137	-0.074
	(0.195)	(0.112)	(0.108)	(0.102)
R^2	0.69	0.69	0.69	0.69
Panel B: Total Effects of Father's Employment				
Father's Employment	-0.064***	-0.030**	-0.036***	-0.020*
	(0.022)	(0.013)	(0.014)	(0.011)
Father's Employment×Girl	0.060***	0.019	0.040***	0.045***
	(0.016)	(0.015)	(0.013)	(0.010)
Mother's Employment	0.033**	0.001	-0.004	-0.002
	(0.016)	(0.009)	(0.007)	(0.007)
Mother's Employment×Girl	0.042***	0.025***	0.024***	0.037***
	(0.006)	(0.005)	(0.005)	(0.005)
Mother's Earnings/1000	-0.832**	-0.225	-0.112	-0.373**
	(0.335)	(0.254)	(0.186)	(0.159)
R^2	0.69	0.69	0.69	0.69

Table 1: Regression Results of Children's College Achievement on their Parents' Employment

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses are clustered at the family level. Spousal earnings are measured in thousand USD. There are $816,617$ observations in each regression. The dependent variable is the education achievement in college at age 25. Independent variables include the mother's age and father's age when the child is born, dummy variables for birth orders, the gender specific cohort fixed effect and the family fixed effect.

children living with one parent during the childhood of that generation. In general, these three positive correlation are statistically significant in the difference-in-differences model controlling for the cohort and state fixed effects.

3.2.1 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) (**??**). We limit the sample to white females and males only.

The educational gender gap variables are calculated from the Census data.[20] 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.

From the same Census data we calculate the age distributions of mothers and fathers of each birth year cohort, and define them as the mother generation and the father generation. For each birth cohort at each birth state, during the first six years after birth, we calculate the labor force participation rates (LFPR) of married women among the mother generation and the LFPR of married men among the father generation from both the Census data (1950 and 1960) and the CPS data. For instance, for the cohort born in Wisconsin in 1965, we calculate the LFPR of married women among this cohort's mother generation and the LFPR of married men among this cohort's father generation in Wisconsin between 1965 and 1970.

Table A12 in Appendix A presents summary statistics of the state level data used in this section.[21]

3.2.2 Difference-in-Differences Analysis

We cannot directly apply the DID regression model (1) due to the lack of rich intergenerationally linked data as the Norwegian Registry data. Instead, we construct a pseudo intergenerational panel from the repeated cross-sectional data as in **?**. First we define gsc cohorts, which are groups of individuals with gender g born at state s in year c. Then for the DID regression model (1) we aggregate all individuals to the cohort level for each gender,

$$\bar{y}_{gsc} = \bar{\mathbf{M}}'_{gsc}\beta + g\bar{\mathbf{M}}'_{gsc}\gamma + \bar{\mathbf{W}}'_{gsc}\xi + \bar{\mathbf{X}}'_{gsc}\delta + \bar{\eta}_{gsc} + \bar{\epsilon}_{gsc} \tag{2}$$

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

[21] 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 in their mother generation when that cohort is aged between zero- 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 **?** 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.

where $\bar{y}_{gsc}$ is the average educational achievement of all individuals in cohort gsc; $\bar{\mathbf{M}}_{gsc}$ include the average mother's and father's employment during the childhood of cohort gsc; $\bar{\mathbf{W}}_{gsc}$ include the average family earnings during the childhood of cohort gsc; $\bar{\mathbf{X}}_{gsc}$ include other demographics aggregated at the cohort gsc level, such as the parental ages when the child is born and the birth order; $\bar{\eta}_{gsc}$ is the aggregated summation of the gender specific cohort fixed effect ν_{g_i,c_i} and the family fixed effect η_f; $\bar{\epsilon}_{gsc}$ is the aggregated i.i.d. unobservable component ϵ_i.[22]

At the individual or family level, the parental employment $\mathbf{M}_i'$ are not necessarily same across different genders. Indeed, our theory predicts, confirmed by the data, that they are correlated with gender or gender composition at the individual level, as shown in the next section. But due to lack of genuine intergenerationally linked data, we use the average employment at the state level for gender g born in year c. For each parent the children's gender composition is different: given having same number of children, some have more daughters than others. But when we aggregate them to the state level, the average gender composition, defined by the number of girls divided by the number of children, is almost constant around 0.5 over time. That is, an average family with two children has one daughter and one son. Given the assumption that parental time, income and other resources are public good within the family, we have $\bar{\mathbf{M}}_{gsc}' = \bar{\mathbf{M}}_{sc}'$, $\bar{\mathbf{W}}_{gsc}' = \bar{\mathbf{W}}_{sc}'$ and $\bar{\mathbf{X}}_{gsc}' = \bar{\mathbf{X}}_{sc}'$ Therefore the model (2) becomes

$$\bar{y}_{gsc} = \bar{\mathbf{M}}_{sc}'\beta + g\bar{\mathbf{M}}_{sc}'\gamma + \bar{\mathbf{W}}_{sc}'\xi + \bar{\mathbf{X}}_{sc}'\delta + \bar{\eta}_{gsc} + \bar{\epsilon}_{gsc} \tag{3}$$

Taking the first difference between girls and boys yields

$$\bar{y}_{girl,sc} - \bar{y}_{boy,sc} = \bar{\mathbf{M}}_{sc}'\gamma + \Delta\bar{\eta}_{gsc} + \Delta\bar{\epsilon}_{gsc} \tag{4}$$

We further assume that $\Delta\bar{\eta}_{gsc}$ is composed of some time-varying confounding state-level factors such as income level, $\bar{\mathbf{I}}_{sc}$, a time-invariant state fixed effect, δ_s, and a birth year fixed effect, δ_c,

$$\Delta\bar{\eta}_{gsc} = \bar{\mathbf{I}}_{sc}'\pi + \delta_s + \delta_c \tag{5}$$

Plugging (5) into (4), letting $GAP_{sc} = \bar{y}_{girl,sc} - \bar{y}_{boy,sc}$ denote the college gender gap among cohort sc and $\xi_{sc} = \Delta\bar{\epsilon}_{gsc}$ yields the following difference-in-differences (DID) regression model at the birth state level,

$$GAP_{sc} = \bar{\mathbf{M}}_{sc}'\gamma + \bar{\mathbf{I}}_{sc}'\pi + \delta_s + \delta_c + \xi_{sc}. \tag{6}$$

In this difference-in-differences regression model, the γ and π are identified by the variation of their co-movement across different states. That is, the changes of $\bar{\mathbf{M}}_{sc}$ or $\bar{\mathbf{I}}_{sc}$ and the resulting changes in educational gender gap take place earlier in some states than in other states.[23]

We consider two different measures of $\bar{\mathbf{M}}_{sc}'$ to reflect (the lack of) the average parental time input

[22]The aggregated ν_{g_i,c_i} does not vary across states. The aggregated family fixed effect η_f for cohort gsc varies across states. These two aggregated fixed effects cannot be separated, thus form $\bar{\eta}_{gsc}$ jointly.

[23]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 **?** and **?**.

and interaction with their children. The first is the LFPR of married women and men among the parent generations. The second is the incidence of children living with one parent. This measure is constructed as the unconditional probability of one adult having at least one child while the spouse is absent in the family.[24] The majority of one parent families is with fathers absent. In $\bar{\mathbf{I}}_{sc}$, we include the median family income among the mother generation and the father generation at the state level.[25]

Table 2 presents the results of the DID regression model (6). In Panel A, the parental employment (LFPR) or absence (Children living with one parent) is measured when the child cohort is aged between 0 and 5. When included separately, the coefficient of the LFPR of married women in the mother generation is positive and statistically significant at the 1% level (Column 1); the coefficient of the LFPR of married men in the father generation is positive and statistically significant at the 10% level (Columns 2); the coefficient of the incidence of children living with one parent is positive but statistically insignificant (Column 3). When LFPR of both married women and men (Column 4) and the rate of living alone with children (Column 5) are included in the same regression, all coefficients only change slightly with the same statistical significance levels. This is also true when the median family income of married women and men are included in the same regression (Column 6), where the coefficients of both parental LFPR increase.

When the regression model includes the parental employment and absence measures during children's different development periods, results differ for different measures. The coefficients of the LFPR of married women are positive although halved and statistically significant when measured during children's primary school ages (Panel B) or high school ages (Panel C). The coefficients of the LFPR of married men become statistically insignificant during either primary or high school ages. The coefficients of the incidence of children living with one parent, on the other hand, remain statistically insignificant when measured during children's primary school ages (Panel B) but are positive and statistically insignificant when measured during the high school ages (Panel C). All results are materially unchanged when the average parental employment is measured by hours worked per week.

The results of the DID regressions in Table 2 imply that the father's generation's employment has much higher positive impact on the educational gender gap of their children generation. Further investigation shows that the positive effect of the parental absence mainly comes from the effect of the incidence of living with the mother alone, while the effect of living with the father alone is either statistically insignificant or negative (Column 7 in Table 2). This indicates that the absence of the father in a family has a more detrimental impact on the boy's educational achievement. However, the LFPR of married men has been mostly flat between 1950 and 1980s, thus at the aggregate national level it has little impact on the children's educational gender gap. These results are consistent with findings from the Norwegian data in the previous subsection.

[24]Spouse absent in the family includes cases of married but spouse absent, separated, divorced, widowed, or single.

[25]We first obtain the cohort distributions of mothers and fathers for each child cohort born in year c. Then for each child cohort sc we calculate the average LFPR, probability of children living with one parent, or the median family income at the state level among married women of the mother generation and married men of the father generation between year c and $c+5$ when the child cohort is aged between 0 and 5, or between year $c+6$ and $c+11$ when the child cohort is aged between 6 and 11, or between year $c+12$ and $c+17$ when the child cohort is aged between 12 and 17.

Coefficient estimates in Table 2 imply that a one-percentage point increase in the LFPR of married women is associated with 0.059-0.072 percentage points increase in the college gender gap of their preschool aged children generation. Note that in our data, the LFPR of married women increased from 18.9% in 1950 to 68.3% in 1986. As a back-of-the-envelope calculation, these coefficient estimates imply that the increase in the LFPR of married women can result in an increase of the college gender gap by about 2.9 to 3.6 percentage points. Since the gender gap in college was about -5.5% for the 1950 cohort and about 11.2% for the 1986 cohort, the effects of the increase in the LFPR of married women can account for 17%-21% of the narrowing and reversal of the college gender gap. This is a bit higher than the 8% in the Norwegian data. On the other hand, the incidence of children living with one parent increases from 2.5% to around 6.9%, which is translated to 3.3%-3.6% of the change in the college gender gap during the same period.[26]

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 (or the incidence of children living with one parent) 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 may 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 between the LFPR of unmarried women and the educational gender gap. We run similar regressions as Table 2, including the LFPR of unmarried women among the mother generation; 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. We find that our results stay unchanged when we exclude these two states from the difference-in-differences analysis.*

3.3 Interpretation

What mechanisms may cause the positive effects of parental employment and absence on their children's educational gender gap? There may be many different explanations from different perspectives. We now discuss three possible channels, namely income effect, time effect, and role model effect.

Extra income from employment may have an time effect. However, previous literature finds either negligible or positive but symmetric income effect. For instance, **?** find that the effect of changes in family income is insignificant in Norwegian data. In **?** the estimated income effects on children's math and reading achievement are positive, statistically significant, and symmetric for boys and girls. In our Norwegian data, the effect of income innovations is found to be either insignificant or negative and significant (Table 1). Thus the income effect is not a likely source for the positive effects of parental employment.[27]

[26]The changes in the incidence of children living with one parent is roughly the same across three age groups.

[27]When we include the gender specific income effect, we find the additional income effect for girls is negative regardless of the source of the income—mother's earnings, father's earnings, or family earnings, have similar negative additional effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: children are aged [0,5]							
LFPR, married women	0.059***			0.060***	0.061***	0.072***	0.072***
	(0.018)			(0.018)	(0.020)	(0.022)	(0.021)
LFPR, married men		0.133*		0.140*	0.152*	0.162*	0.163*
		(0.077)		(0.076)	(0.083)	(0.086)	(0.086)
Children living with one parent			0.084		0.121	0.103	
			(0.079)		(0.086)	(0.090)	
Children living with mother alone							0.064
							(0.049)
Children living with father alone							-0.101
							(0.162)
Observations	1,374	1,369	1,271	1,368	1,220	1,188	1,188
R-squared	0.86	0.86	0.85	0.86	0.85	0.85	0.85
Panel B: children are aged [6,11]							
LFPR, married women	0.033**			0.033**	0.033**	0.038**	0.039**
	(0.016)			(0.016)	(0.017)	(0.018)	(0.018)
LFPR, married men		0.064		0.056	0.037	0.043	0.055
		(0.071)		(0.071)	(0.075)	(0.077)	(0.077)
Children living with one parent			-0.097		-0.084	-0.104	
			(0.070)		(0.072)	(0.074)	
Children living with mother alone							-0.012
							(0.042)
Children living with father alone							-0.386***
							(0.118)
Observations	1,681	1,678	1,574	1,675	1,521	1,491	1,491
R-squared	0.90	0.90	0.89	0.90	0.89	0.89	0.89
Panel B: children are aged [12,17]							
LFPR, married women	0.022*			0.023*	0.030**	0.024*	0.023*
	(0.012)			(0.012)	(0.013)	(0.014)	(0.014)
LFPR, married men		-0.070		-0.068	-0.047	-0.061	-0.061
		(0.043)		(0.043)	(0.044)	(0.045)	(0.045)
Children living with one parent			0.136***		0.126**	0.135**	
			(0.052)		(0.053)	(0.055)	
Children living with mother alone							0.081**
							(0.032)
Children living with father alone							0.018
							(0.078)
Observations	1,703	1,706	1,618	1,701	1,592	1,584	1,584
R-squared	0.92	0.92	0.91	0.92	0.91	0.91	0.91
Median incomes included	N	N	N	N	N	Y	Y

Table 2: Difference-in-differences analysis of college gender gap, U.S data, White only

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Robust standard errors are reported in parentheses. Estimations are weighted by state populations. The dependent variable is the gender gap in college achievement. The age distributions of the mother generation and the father generation for each cohort are calculated from U.S. Census data. Independent variables include birth year fixed effects, birth state fixed effects, and the median income of married women and men (Columns 6-7 only).

Both the time effect and the role model effect may be asymmetric. The time effect refers to the effect of parental time in children's education production, which is one of the most important inputs as previous research discovers (see, e.g., **?**, among many others). The reduction in time input for children as a result of working in the labor market might have asymmetric impact for girls and boys. On the other hand, the role model effect refers to the hypothesis that children might emulate the behavior of their parents (e.g., **?**) or previous generations (e.g., **??**). Therefore, the parent's current employment may increase the child's expected return of education, due to the college wage premium as well as the positive intergenerational correlation in the employment between parents and children. We focus on the role model effect from the same-gender parent, namely the role model effect of mother's employment on girls and the role model effect of father's employment on boys.[28] That is, we assume gender-specific (therefore asymmetric) role model effect. Therefore, if exists, the time effect of parental employment is negative for both the daughter and the son, while the role model effect is positive for the child with the same gender as the parent.

From Table 1, in a nuclear Norwegian family, the negative effect of both mother's and father's employment on the son's educational achievement indicates the presence of the income effect for sons; the insignificant effect of father's employment on daughter's educational achievement implies an insignificant income effect of father's employment for daughters. The positive effect of mother's employment on her daughter's educational achievement indicates the presence of the role model effect of her employment for her daughters. However, we cannot reject the hypothesis of an insignificant role model effect of father's employment on the son.

Now we focus on the time effect of parental employment and argue it may be asymmetric. It is straightforward for father's employment—the total effect of father's employment on the daughter's educational achievement is statistically insignificant (Columns 1-3 in Table 1) or positive and significant (Column 4 in the same table). Thus, the time effect of the father's employment is asymmetric—it has a negligible time effect on the daughter's education but has a large and negative time effect on the son's. This gender-specific asymmetric time effect of father's employment is consistent with findings in literature on boys suffering more from absence of either parent (most likely the father), as well as evidence from the U.S. data.

It is less obvious whether the time effect of mother's employment is asymmetric. In Panel A of Table 1, it appears that the mother's employment on her children's educational gender gap presents a U-shaped pattern—it is larger when the child is either aged $[0, 5]$ or aged $[12, 17]$ than when the child

on girls. As we mentioned earlier in the Norwegian data, it is difficult to isolate the income effect from the time effect due to the high correlation between employment and earnings. This probably explains the negative correlation between the earnings innovation and the college gender gap found in this paper. Furthermore, the estimated income effect on the college gender gap being negative would translate to a negative effect of parental employment on their children's educational gender gap through the extra labor income, thus the income effect is not a likely source for the positive effect of parental employment.

[28]The role model effect of mother's employment for boys is not that straightforward. It is possible that competition in the marriage market might generate some indirect effect for boys from maternal employment (**?**), 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.

is aged [6, 11]. To see the pattern more clearly, we include the mother's employment during all three different age groups in the same DID regression, with results presented in Table 3. Interestingly, the total effects of mother's employment during her children's different age groups on the children's educational gender gap present a similar U-shaped pattern—it is positive and statistically significant before school age, and then becomes insignificant when children are in primary school, but turns to be positive and statistically significant again when children go to high school. This U-shaped pattern of the total effect of mother's employment on children's educational gender gap is consistent with a decreasingly asymmetric time effect and an increasingly asymmetric role model effect. That is, if the asymmetry of the time effect shrinks, while the asymmetry of the role model effect widens, with the child's age, then it would generate this U-shaped pattern. The former is supported by other research in finding higher return of mother's time on boys (e.g., **??**), and the literature on education production which suggests that the maternal time input is critical at early stage than at later stages (**???**). For the latter, if the conscious channel of the role model, which increases with the child's age, dominates the subconscious one, which decreases with age, it implies that the asymmetric effect of maternal employment via the role model widens with the child's age. This argument is based on the timing of when the time effect and the role model effect are likely to be the strongest.[29] It is intuitive and straightforward.[30]

The asymmetry of the time effect of father's employment appears to be increasing over his children's age, as shown in Column 2 of Table 3. It is negative at the preschool age and positive at the primary school age, both being statistically insignificant; it is positive and statistically significant at the high school age. Evidence from the U.S. data presents a U-shaped pattern of the asymmetric effect of the father's employment or absence on children's educational gender gap (Table 2). They both points to a large and significant asymmetric effect of father's time with his sons during the high school stage.

In summary, the most direct interpretation of our results is that the asymmetric effect of mother's employment comes from a combination of (decreasingly) asymmetric time effect and (increasingly) asymmetric role model effect, while the asymmetric effect of father's employment comes from a combination of (increasingly or U-shaped) asymmetric time effect and rather weak asymmetric role model effect.[31]

3.4 Racial Differences in Gender Gap

Do we also observe such correlation between parental employment or absence and their children's educational gender gap among other races? In this subsection we investigate the black Americans. The college achievement among black Americans, compared with white Americans, is more balanced between women and men in 1950, with a gender gap of 0.02. However, the college gender gap among

[29] That is, the detrimental time effect of mother's employment 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 in Table 3.

[30] On the other hand, there may be other feasible explanations for asymmetric effect of mother's employment on the children's educational gender gap. That is beyond the scope of this paper and we leave it for future research.

[31] Why the role model effect of father's employment for sons is much weaker than the role model effect of mother's employment for daughters is beyond the scope of our paper. Our speculation is that the historical labor force participation rate is much higher among men than women, thus men's employment status has relatively lower weight on their sons—boys learn that very likely they will work when they grow up regardless.

	(1) **Mother's Employment**	(2) **Father's Employment**
Employment during $[0, 5]$	0.003	-0.020
	(0.007)	(0.014)
Employment during $[0, 5] \times$ Girl	0.015**	-0.001
	(0.007)	(0.017)
Employment during $[6, 11]$	0.008	-0.022
	(0.006)	(0.015)
Employment during $[6, 11] \times$ Girl	-0.002	0.009
	(0.007)	(0.018)
Employment during $[12, 17]$	-0.006	-0.020*
	(0.006)	(0.012)
Employment during $[12, 17] \times$ Girl	0.033***	0.041***
	(0.007)	(0.013)
R-squared	0.69	0.69

Table 3: Difference-in-differences regression of children's college achievement on their mothers' and fathers' employments in different age groups, Norway

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses. Spousal earnings are measured in thousand USD. There are $816,617$ observations in each regression. The dependent variable is the education achievement in college at age 25. Independent variables include the mother's age and the father's age when the child is born, dummy variables for birth orders, spousal employments during three age groups and their interactions with the girl dummy, spousal earnings during three age groups, the gender specific cohort fixed effect and the family fixed effect.

black Americans has increased since then, presenting a similar pattern as white Americans. It peaks in 1985 reaching 0.10, drops to 0.08 a year later, and increases again to 0.11 in 1987, as shown in Figure 3. In late 1980*s*, the college gender gap is much similar among blacks and whites.

We run the same difference-in-differences regressions of 6 for black Americans, with results presented in Table A13. Comparing with whites in Table 2, the coefficient of the LFPR of married women or men among the parent's generation is positive but insignificant regardless of age groups, except when measured during the child's preschool age the latter is negative and insignificant. The coefficient of the incidence of children living with one parent is also positive but statistically insignificant across all three age groups. When we separate it into two different categories, the coefficient of children living with mothers alone is all positive and it is statistically significant when measured after the preschool age. In contrast, the coefficient of children living with fathers alone is negative and it is statistically significant only measured during the child's high school age. It appears that the results with black Americans are similar as those with white Americans, with smaller precision, which is probably due to smaller sample sizes.

Figure 3 shows that the trends of the LFPR of married women and men among blacks assemble the same pattern as whites. On the other hand, the incidence of children living with one parent is much higher and also increases much faster among blacks, almost all of which comes from the prevalence of children living with mothers alone among blacks. As a back-of-the-envelope calculation, such prevalence contributes to 1.8 (measured at primary school) or 0.5 (measured at high school) percentage points of the black college gender gap, which is 20% or 6% of the total change between 1950 and 1987. These figures are higher than those of whites.

4 Children's Gender Effect on Parental Labor Supply

We have documented so far the positive correlation between parental employment and the educational gender gap of their own children in Norway or that of their children generation in the United States. Our interpretation is that this correlation is a result of the asymmetric production effects of mother's and father's time input on boys and girls, and the differential role model effects of their employment on girls. In this section we provide some auxiliary evidence which supports our interpretation from a different angle. Specifically, we investigate a novel prediction of our previous results on the mother's and father's employment decisions. First we present a static model where a parent makes the labor supply decision, aware of the asymmetric effects of her or his time input on sons and daughters. Then we empirically test the model implication regarding a mother's/father's/family's labor supply decision and the gender composition of their children using both the U.S. data and the Norwegian data.

4.1 The Model of Mother's Labor Supply

We illustrate the model using a mother as an example. Same logic applies to a father as well. 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 $y_0 \geq 0$. 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, $T - t \equiv s$.[32] 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:

$$\max_{\{t,c\}} \; u(c) + \lambda \left[n_g e_g (T - t, y_0 + wt - c) + n_b e_b (T - t, y_0 + wt - c) \right] \tag{7}$$

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.[33] 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.[34]

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):

$$\frac{\partial e_g(s, k)}{\partial s} < \frac{\partial e_b(s, k)}{\partial s} \; \forall s, k \tag{8}$$

That is, for any given level of mother's time input and market input, the marginal effect of mother's time is higher for boys. This assumption is derived from previous empirical findings of positive correlation between mother's employment and her children's educational gender gap,

$$\frac{\partial \left[e_g(T - t, k) - e_b(T - t, k) \right]}{\partial t} > 0 \Rightarrow \frac{\partial \left[e_g(s, k) - e_b(s, k) \right]}{\partial s} < 0 \tag{9}$$

This asymmetric effect comes from both the asymmetric time effect and the differential role model effect as discussed earlier.[35]

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

[33] There might be girl-boy difference in direct parental time investment. For example, **?** find parents spend more time reading to girls. However, much parental time investment has spillover effect, as it is observed by all other children in the same family and therefore is likely to become a public good. Disciplining, among others, is another good example.

[34] 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.

[35] 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.

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):

$$\frac{\partial e_g(s,k)}{\partial k} = \frac{\partial e_b(s,k)}{\partial k} \ \forall s,k \tag{10}$$

Note that Assumption 2 immediately implies:

$$\frac{\partial^2 e_g(s,k)}{\partial k \partial s} = \frac{\partial^2 e_b(s,k)}{\partial k \partial s} \ \forall s,k \tag{11}$$

and

$$\frac{\partial^2 e_g(s,k)}{\partial k^2} = \frac{\partial^2 e_b(s,k)}{\partial k^2} \ \forall s,k \tag{12}$$

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):

$$\frac{\partial^2 e_i(s,k)}{\partial s \partial k} \geq 0, \ \forall s,k; \ i \in \{g,b\} \tag{13}$$

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 (7) can be rewritten as:

$$\max_{\{t,c\}} V(t,c;n) \equiv u(c) + \tilde{\lambda}\left[n \cdot e_g(T-t, y_0 + wt - c) + (1-n) \cdot e_b(T-t, y_0 + wt - c)\right] \tag{14}$$

We have the following result:

Proposition 1 *Fixing $\langle \lambda, w, N \rangle$, the optimal labor supply t^* that solves problem (14) is monotone non-decreasing in the fraction of the girls, n.*

Proof. Using Assumption 2, we can immediately derive that

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

where the inequality comes from Assumption 3 and concavity of the production function $e_b(\cdot,\cdot)$. We can also show that[36]

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

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 **?**[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.

4.2 Evidence: Children's Gender Composition and Parental Labor Supply

We now test the prediction from our model in the preceding subsection by examining the relationship between various measures of the parental labor force attachment and the gender composition of the children, in both the U.S. data and the Norwegian data. We run the following econometric specification:

$$y_i = \mathbf{X}_i'\beta + \mathbf{Z}_i'\gamma + \epsilon_i, \tag{15}$$

where y_i is a measure of i's parent's labor force attachment; $\mathbf{X}_i$ is the vector of the parental and family's characteristics including the number of children, age, age squared, age at the first birth, and the year dummies; $\mathbf{Z}_i$ is the vector of variables of interest regarding the children's gender composition. We use two different measures of $\mathbf{Z}_i$: the first child being a girl and the fraction of girls among *all* children in the same family. It is reasonable to assume, at least in developed economies, that both measures are exogenous. The analysis will be first conducted separately on mothers' and fathers' supply, and then on the pooled labor supply of both parents.

We first investigate the U.S. families. The data is constructed from the IPUMS Census data from 1950 to 2000. We select the sample to include only families with both biological white parents present and with at least two children; the mother's age at the first birth has to be between 18 and 40.[37] 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 C14 in Appendix A presents summary statistics of the U.S. data. The fraction of girls has mean 0.485 and standard deviation 0.386, which is not statistically significantly different from 0.5, and is almost constant from 1950 to 2000.

The regression results, presented in Table 4, indicate that the labor force participation status of the mother, the father, or the family average, is positively and statistically significantly correlated with either the first child being a girl or the fraction of girls among all children.[38] Results hold regardless of the labor

[36] Refer to Appendix B for detailed proof.

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

[38] Note this is the effect of the gender of the first child or the fraction of girls among all children, among married men living

force participation status of the spouse. As a back-of-the-envelope calculation, switching the gender of one child would increase the labor force participation rate by 0.2 percentage point or 0.4% of the standard deviation among mothers, and 0.05 percentage point or 0.2% of the standard deviation among fathers.[39] The gender of the first child has a stronger effect on parental labor force participation status, 50% larger for mothers or 100% larger for fathers, than the gender of other children. It also appears that whether the spouse works or not has a positive and significant effect on one's labor force participation decision, and such effect is larger for women than for men. The family average labor force participation status, calculated as the average of the mother and the father, is also positively correlated with the gender of the first child or the fraction of girls among all children.

Table 5 report the OLS results when the parent's labor force attachment is measured by whether the parent is employed or whether working full time, instead of the labor force participation status.[40] Full-time working is defined as working more than 1,260 hours per year and is only available since 1980.[41] The effects of the first child being a girl or the fraction of girls on either parent's employment status is positive and statistically significant, and so is for the family average. However, either of them does not have a significant effect on the mother's (Column 4) but has positive and significant effect on father's full time working status (Column 5). We will come back to this comparison in a later subsection. The first child being a girl has a positive and significant effect on the family average full-time status (Column 6).

We conduct a further robustness check relating to the endogeneity issue of the number of children.[42] 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 **?** is not likely to be applicable to our model.[43] 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 (15) where we use the same gender indicator as in **?** as the "instrument" for the number of children. For comparison, we restrict our analysis to families with at least two children in both regressions. Results presented in Table 6 show that they are much similar as those in Table 4. This is not surprising as the

with their spouse. **?** find a *negative* correlation between the father's labor supply and the first child being a girl in a sample of married and unmarried men with or without children, in the Panel Study of Income Dynamics (PSID). Furthermore, their specification does not separate the gender composition effect from the effect of the number of children.

[39]The smaller elasticity of father's labor supply with respect to the fraction of girls than that of the mother's labor supply may come from the fact that in many families men are main breadwinners.

[40]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.

[41]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.

[42]A child's gender is most likely exogenous in the U.S., but the number of children is endogenous.

[43]In **?**, 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 does not necessarily hold according to our model.

variable of interest is likely to be exogenous. It is worth noting that the number of children has a negative effect on the mother's labor force participation status and the effect is smaller in the IV regression. This confirms results in **?**, where they use the Census data of 1980 and 1990 only.[44] the F-statistic on the excluded instrument in the first stage is greater than 10 in all cases.

All above results in Tables 4, 5 and 6 are robust to some model variants. For example, in unreported regressions, we find that the effects of both measures are mostly unchanged when we control for parents' educational attainments. In Probit analysis with or without instrumental variable, the results are also very similar in all cases.

Table 7 presents similar results using Norwegian registry data. Here the dependent variable is the parent's working status inferred from earnings history. Consistent with the U.S. data in Table 6, we control for the parent'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 for the mother's labor supply 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. The father's labor supply is also positively and statistically significantly correlated with the fraction of girls in both the OLS and the IV regressions, but the elasticity is smaller. The differential labor supply elasticity among mothers and fathers is consistent with the previous result from the U.S. data.[45]

	(1)	(2)	(3)	(4)	(5)
	Mothers		**Fathers**		**Family**
Panel A: effects of the gender of the first child					
First child being a girl	0.003***	0.003***	0.001***	0.001**	0.002***
	(0.001)	(0.001)	(0.0002)	(0.0002)	(0.003)
Spousal working		0.106***		0.017***	
		(0.006)		(0.001)	
R-squared	0.14	0.14	0.02	0.02	0.11
Panel B: effects of fraction of girls among all children					
Fraction of girls	0.004***	0.004***	0.001***	0.001**	0.003***
	(0.001)	(0.001)	(0.0003)	(0.0003)	(0.0004)
Spousal working		0.106***		0.017***	
		(0.006)		(0.001)	
R-squared	0.14	0.14	0.02	0.02	0.11

Table 4: Effects of the children's gender composition on parental labor force participation, U.S. Census 1950-2000

Notes: [a]*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors (clustered at the state level) in parentheses. There are 2,667,516 observations in each column. The dependent variable is a dummy variable with 1 indicating participating in the labor force and 0 otherwise. Independent variables include parental age, age squared, age at the first birth, the number of children, and year dummies.

[44]**?** find the number of children has no effect on the father's labor supply while we find a positive, although much smaller, effect. We are using a larger sample covering more Census years.

[45]The number of children has a positive effect on the father's labor supply in Norway. This might indicate that the income effect—a larger family demands more income—dominates.

	(1)	(2)	(3)	(4)	(5)	(6)
	Employment Status			**Full-time Status**		
	Mothers	**Fathers**	**Family**	**Mothers**	**Fathers**	**Family**
Panel A: effects of the gender of the first child						
Gender of 1st child	0.003***	0.001***	0.002**	0.002	0.001**	0.001**
	(0.001)	(0.0003)	(0.0003)	(0.001)	(0.0005)	(0.001)
R-squared	0.14	0.02	0.10	0.04	0.01	0.04
Panel B: effects of fraction of girls among all children						
Fraction of girls	0.004***	0.001***	0.003***	-0.001	0.001**	0.0005
	(0.001)	(0.0003)	(0.0004)	(0.001)	(0.0007)	(0.001)
R-squared	0.14	0.02	0.10	0.04	0.01	0.04

Table 5: OLS results of parental labor force attachment, measured by employment status, U.S. Census 1950-2000

Notes: [a]*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors (clustered at the state level) in parentheses. There are 2,667,516 observations in Columns 1-3 and 746,302 observations in Columns 4-6. Independent variables include parental age, age squared, age at the first birth, the number of children, year dummies, and spousal employment status (Columns 1-2) or full-time status (Columns 4-5).

4.3 Children's Gender Composition and Parental Marital Status

Above we find that children's gender composition affects parental labor supply among married couples, due to the asymmetric effect of parental time input. Given this asymmetric effect, when altruistic parents make other decisions relating to children's education production, they will also take into account their children's gender composition. In this subsection, we investigate whether and how children's gender composition affects parental marital status.

There are other channels through which parental marital status may be affected by children's gender composition. **?** find that a women with the first child being a girl is significantly more likely to be living alone with the children. They argue this is due to the parental biased preference favoring boys over girls. **?** argue that parents may have preference over a variety of their children's genders. To control these channels, we select a sample where each parent, regardless of marital status, has at least one girl and one boy, while otherwise the final sample is constructed in the same way as immediately above form the IPUMS Census data from 1950 to 2000.

We investigate how the children's gender composition affects four different measures of marital status or living arrangement, in the regression model of 15. Similar as before, the variable of interest is the fraction of girls among all children. Table 8 shows that if a parent, mother or father, has more girls among all her or his children, then she or he is more likely to be separated or divorced if ever married, more likely to have an absent spouse if married, or more likely to be living alone with children regardless of the marital status. All coefficients are statistically significant. These results supplement the previous subsection, supporting Proposition 1.

	(1)	(2)	(3)	(4)	(5)	
	Mothers		**Fathers**		**Family**	
	OLS	**IV**	**OLS**	**IV**	**OLS**	**IV**
Panel A: gender of first child						
First child being a girl	0.004***	0.004***	0.001***	0.001**	0.002***	0.002***
	(0.001)	(0.001)	(0.0002)	(0.0002)	(0.0004)	(0.0003)
Number of children	-0.080***	-0.079***	-0.005***	-0.005*	-0.043***	-0.043***
	(0.001)	(0.008)	(0.001)	(0.003)	(0.001)	(0.001)
R-squared	0.15	0.15	0.02	0.02	0.12	0.12
First-stage F-stat		2714		2790		2750
Panel B: fraction of girls among all children						
Fraction of girls	0.005***	0.005***	0.001**	0.001**	0.003***	0.003***
	(0.001)	(0.001)	(0.0003)	(0.0004)	(0.001)	(0.001)
Number of children	-0.080***	-0.079***	-0.005***	-0.005*	-0.043***	-0.043***
	(0.001)	(0.008)	(0.001)	(0.003)	(0.001)	(0.004)
R-squared	0.15	0.15	0.02	0.02	0.12	0.12
First-stage F-stat		2778		2847		2812

Table 6: Effects of the fraction of girls on parental labor force participation among families with 2+ children, U.S. Census 1950-2000

Notes: [a]*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors (clustered at the state level) in parentheses. There are 1,838,743 observations in each column. The dependent variable is a dummy variable with 1 indicating participating in the labor force and 0 otherwise. Independent variables include parental age, age squared, age at the first birth, year dummies, and spousal labor force participation status (Columns 1-4 only). The instrumental variable for the number of children is the same gender indicator as in **?**.

	(1)	(2)	(3)	(4)	(5)	(6)
	Mothers		**Fathers**		**Family**	
	OLS	**IV**	**OLS**	**IV**	**OLS**	**IV**
Number of Children	-0.0352***	-0.0180***	0.0036***	0.0117***	-0.0180***	-0.0026
	(0.0002)	(0.0034)	(0.0001)	(0.0023)	(0.0001)	(0.0022)
Fraction of Girls	0.0009***	0.0011***	0.0005**	0.0005***	0.0013***	0.0015***
	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Spousal working	0.1189***	0.1189***	0.0594***	0.0600***		
	(0.0004)	(0.0004)	(0.0002)	(0.0003)		
Observations	13,532,478	13,532,478	13,735,344	13,735,344	13,532,478	13,532,478
R-squared	0.12	0.12	0.19	0.19	0.16	0.16
First-stage F-stat		1456		1246		1356

Table 7: Effects of the fraction of girls on maternal labor force participation among families with 2+ children, Norway.

Notes: Robust standard errors 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 (Columns 1-2) or the father's (Columns 3-4) or average (Columns 5-6) age, age squared , age at the first birth, and year dummies.

	(1) Separated/divorced	(2) Spouse absent	(3) Living alone
Fraction of girls	0.002**	0.001**	0.003**
	(0.001)	(0.0003)	(0.001)
Number of observations	2,316,459	2,210,519	2,336,641
R-squared	0.04	0.002	0.04

Table 8: Effects of the fraction of girls on parental marriage or living arrangement among families with at least one boy and one girl, U.S. Census 1950-2000.

Notes: [a]*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors (clustered at the state level) in parentheses. The dependent variable is a dummy variable with 1 indicating separated or divorced and 0 married (Column 1), or 1 indicating spouse absent and 0 spouse present (conditional on being married, Column 2), or 1 indicating living alone with children and 0 living with spouse and children (regardless of marital status, Column 3). Independent variables include the number of children, age, age squared, age at the first birth, the female dummy, education dummies, and year dummies.

4.4 Alternative Explanations

It is worth emphasizing that our theory of asymmetric effect of parental time is a sufficient condition for the prediction and findings in this section. It is not necessarily a necessary condition. For this logic, we are not arguing that our theory is the only reason for the discovered positive correlation between the parental labor supply or marital status and the fraction of girls among their own children. Having said that, it is still worth discussing other possible alternative explanations, which is the focus of this subsection.

4.4.1 Differential Costs of Raising Children

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 on parental labor supply, 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 **?** and they find the opposite in the U.S. data. Therefore this hypothesis is unlikely to be the driving force of our findings.

4.4.2 Differential Bargaining Power

In countries with son preference, like the U.S. (**?**) and China (**?**), 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 **?** higher bargaining power would imply less labor supply (**?**), thus it generates the results in the previous subsection. This explanation might be true to some extent, but we argue it is

not likely to be the main reason. First, this bargaining power does not explain the positive effect of the fraction of girls on the *father's* labor supply. Second, we show below even for the mother the evidence from the data is not in favor of the differential bargaining power hypothesis.

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. This implies that the effect of the child's gender on mother's employment should not vary with the child's age. 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 have age-dependency, we include an interaction term between the fraction of girls and children's ages in the regression model,

$$y_i = \mathbf{X}_i'\beta + \mathbf{Z}_i'\gamma + \xi t_i + t_i\mathbf{Z}_i'\pi + \epsilon_i, \tag{16}$$

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.[46] In this specification, the coefficient ξ represents the effect of the children's age on the mother's labor supply while the coefficient π represents the age trend in the gradient of the mother's labor supply on the fraction of girls. In this differential but constant bargaining power hypothesis π should be zero.[47] 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 3, the asymmetry is stronger when a child is either young (the asymmetric production channel dominates) or old (the role model channel dominates).

As previous literature finds that the U.S. is likely a country with son preference (**?**), we conduct the econometric specification (16) in the U.S. Census data only. The estimation results from the pooled 1950-2000 U.S. Census data are presented in Table 9. 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. It shows 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.

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 on the fraction of girls increases with her children's age. This is not consistent with the differential bargaining power hypothesis. On

[46] 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.

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

the other hand, 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. Interestingly, similar pattern is found for the father, which helps us evaluate the next alternative explanation.

4.4.3 Family Production

Earlier research find in two-parent families girls, especially adolescent girls, are more likely to engage in family-care household work than boys and they are also more likely to substitute for their mother in housework, in U.S. (**?**), U.K. (**?**), Australia and Canada (**?**), and Sweden (**?**). A mother with higher fraction of girls might find it easier to meet the demands of work and family and therefore participate more in the labor market. This family production hypothesis is able to explain the positive correlation between the fraction of girls and the mother's labor supply as well as its increasing gradient discovered in the previous subsection, thus it is a tempting alternative. What is inconsistent with this hypothesis is that a similar positive correlation and a similar positive gradient hold for the father's labor supply, as presented in Tables 4-5 and Table 9. As most research finds girls might substitute mothers but not fathers in housework.[48] Therefore this theory does not explain the positive correlation and gradient for the father's labor supply in the U.S. data. Furthermore, Panel B of Table 5 reveals that the fraction of girls has an insignificant effect on the mother's full time working status but has a significantly positive effect on the father's. These evidences are harder to be reconciled with the family production hypothesis but are consistent with our asymmetric effect theory.

5 Conclusion

In this paper we argue that the increasing labor force participation rate of married women since 1950 accounts for about 17% (or 9%) of the narrowing and reversal of the educational gender gap among their children's generation in the United States (or Norway); the rising incidence of children raised by one parent contributes to another 3.3% during the same period in the U.S. In particular, we highlight that the connection between the parental employment or absence and their children's educational gender gap is consistent with the asymmetric effect of parental time input in their children's education production as well as the gender-specific role model effect.

In the panel data from Norway, we find direct evidence for the asymmetric effects of parental employment on their own children's educational gender gap. The asymmetry is statistically significant even after controlling for the family fixed effect and is quite robust to different model specifications. In

[48] One exception is **?**, who find that a parental leave reform impacting fathers reduces the likelihood of adolescent girls doing household work in the Norwegian data. This implies that adolescent girls might also substitute for their father in housework in Norway. Therefore we focus on the data from the U.S., not Norway, for the analysis of the gradients of the mother's labor supply on the fraction of girls.

	(1)	(2)	(3)	(4)	(5)
	Mothers		**Fathers**		**Family**
Panel A: Age of youngest child					
Number of children	-0.047***	-0.047***	-0.005***	-0.004***	-0.026***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Fraction of girls	0.003	0.003	-0.001	-0.001	0.001
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Age of youngest child/10	0.185***	0.185***	0.003**	0.001	0.093***
	(0.006)	(0.006)	(0.002)	(0.001)	(0.003)
Fraction of girls × Youngest age/10	0.005**	0.004*	0.003**	0.003**	0.004***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Spousal LM attachment		0.099***		0.014***	
		(0.006)		(0.001)	
R-squared	0.15	0.16	0.02	0.02	0.13
Panel B: Age of oldest child					
Number of children	-0.081***	-0.080***	-0.006***	-0.005***	-0.043***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Fraction of girls	-0.003	-0.003	-0.002*	-0.002*	-0.002
	(0.003)	(0.003)	(0.001)	(0.001)	(0.002)
Age of oldest child/10	0.217***	0.219***	-0.007***	-0.010***	0.104***
	(0.007)	(0.007)	(0.001)	(0.001)	(0.004)
Fraction of girls× Oldest age/10	0.008***	0.007***	0.003***	0.003***	0.005***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Spousal LM attachment		0.099***		0.014***	
		(0.007)		(0.001)	
R-squared	0.15	0.15	0.02	0.02	0.12
Panel C: Average age of all children					
Number of children	-0.050***	-0.050***	-0.005***	-0.004***	-0.027***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Fraction of girls	-0.0004	-0.0003	-0.001	-0.001	-0.001
	(0.0025)	(0.0025)	(0.001)	(0.001)	(0.001)
Average age of all children/10	0.368***	0.368***	0.009***	0.004*	0.186***
	(0.012)	(0.013)	(0.003)	(0.002)	(0.006)
Fraction of girls× Average age/10	0.008***	0.007***	0.003***	0.003**	0.005***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Spousal LM attachment		0.098***		0.014***	
		(0.007)		(0.001)	
R-squared	0.15	0.16	0.02	0.02	0.13

Table 9: OLS results of maternal labor supply, with the fraction of girls interacting with children ages, pooled U.S. Census [1950-2000].

Notes: [a]*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors (clustered at the state level) in parentheses. There are 1,838,743 observations in each column. Independent variables include parent's age, age squared, age at the first birth (Panels A and C only), and the year dummies.

the cross-sectional data from the United States, we find that the college gender gap in one generation is positively correlated with the labor force participation rate (LFPR) of married women in their mother generation, the LFPR of married men in their father generation, as well as the incidence of children living with one parent (mostly mothers), during their childhood in their birth state. We interpret that this correlation is a result of asymmetric effects of parental employment or absence on their children's educational achievement, which benefits daughters more than sons.

We further derive an empirically testable implication of this mechanism on a rational, well-informed and altruistic parent's labor supply. Specifically, in the presence of this gender-asymmetric effect of parental employment, the parental labor supply should be higher if they have 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 three plausible alternative explanations about our findings.

This implication on parental labor supply, however, raises the concern that parental labor supply is endogenous, which might bias the estimation of the effect of parental employment on children's educational gender gap in the first part of the paper. We acknowledge such concern, but argue that the bias, if exists, is minimal. The increases in the LFPR of married women and the divorce rate since 1950 are prevalent across most developed countries and are exogenous to our model. These exogenous increases are the main source for the identification of the effect of parental employment and absence on the children's educational gender gap. The LFPR of married women has increased 114% or 309% of the standard deviation in Norway or the U.S. during the sample periods. On the other hand, the effect of the children's gender composition on parental employment is estimated to be statistically significant but quantitatively small, accounting for less than half percent of the standard deviation of parental employment. For this reason, we believe that, if exists, the bias in our estimation of parental employment and absence on the children's educational gender gap is minimal.

Taken collectively, the weight of our evidence supports the mechanism that parental employment and absence 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 parental employment (mostly mothers) and absence (mostly fathers) 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 parental employment or absence 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.[49] Third, the effect of parental 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 fertility more than in other countries without such preference bias toward sons.

[49]See **?** for a related attempt, though they focus on marriage, and do not consider the impact of mothers' employment on children's education.

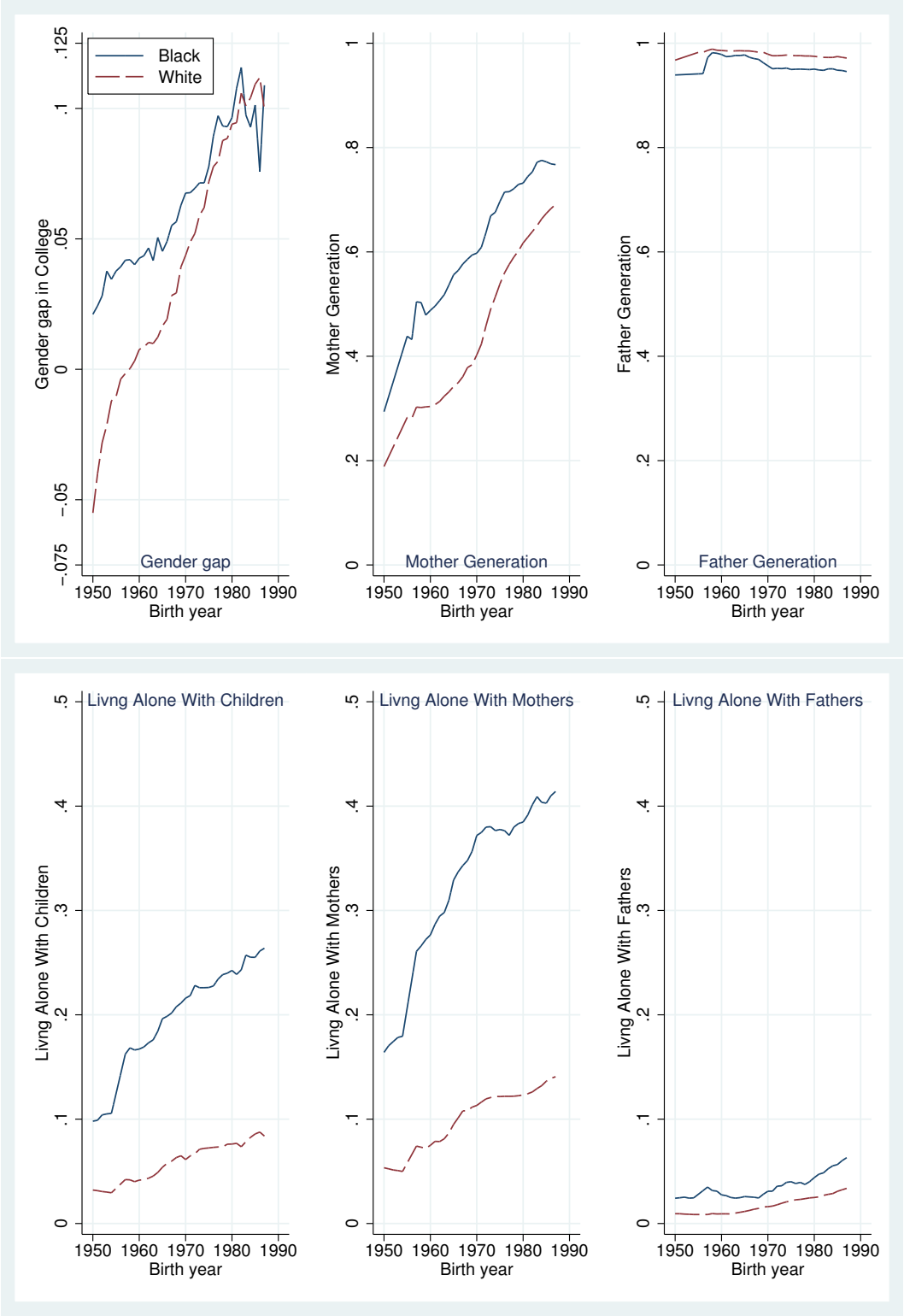

Figure 3: Comparison: black and white Americans

Notes: The solid line represents the college gender gap among black Americans and the long dashed line among white Americans.

Appendix

A Summary Statistics

Variable	Mean	S.D.	Min	Max
College achievement at age 25	0.373	0.484	0	1
Girl	0.488	0.500	0	1
Mother's age at birth	27.1	5.1	16	45
Father's age at birth	29.8	5.8	16	65
Mother labor supply during $[0, 17]$	0.581	0.321	0	1
Mother labor supply during $[0, 5]$	0.382	0.395	0	1
Mother labor supply during $[6, 11]$	0.579	0.411	0	1
Mother labor supply during $[12, 17]$	0.780	0.351	0	1
Father labor supply during $[0, 17]$	0.962	0.112	0	1
Father labor supply during $[0, 5]$	0.964	0.127	0	1
Father labor supply during $[6, 11]$	0.967	0.132	0	1
Father labor supply during $[12, 17]$	0.953	0.168	0	1
Mother earnings during $[0, 17]$	22.6	17.6	0	183.3
Mother earnings during $[0, 5]$	13.5	16.6	0	183.3
Mother earnings during $[6, 11]$	21.6	19.9	0	183.3
Mother earnings during $[12, 17]$	32.9	22.0	0	183.3
Father earnings during $[0, 17]$	68.1	25.1	0	183.3
Father earnings during $[0, 5]$	61.6	22.9	0	183.3
Father earnings during $[6, 11]$	69.8	27.8	0	183.3
Father earnings during $[12, 17]$	73.1	32.8	0	183.3

Table A10: Summary statistics of Norwegian registry data

Notes: The final sample has 816,617 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. Earnings are measured in thousand USD.

Child's age group	(1) [0, 17]	(2) [0, 5]	(3) [6, 11]	(4) [12, 17]
Panel A: Total Effects of Mother's Employment				
Mother's employment	-0.006	-0.008	-0.007	-0.011**
	(0.012)	(0.007)	(0.006)	(0.005)
Mother's employment×Girl	0.034***	0.020***	0.021***	0.028***
	(0.006)	(0.006)	(0.005)	(0.005)
Spousal employment	-0.015	0.003	-0.016	-0.010
	(0.024)	(0.012)	(0.013)	(0.012)
Spousal employment×Girl	0.029**	-0.003	0.020*	0.033***
	(0.013)	(0.011)	(0.011)	(0.009)
Spousal earnings /1000	-0.402**	-0.207*	-0.131	-0.080
	(0.203)	(0.117)	(0.112)	(0.106)
R-squared	0.69	0.69	0.69	0.69
Panel B: Total Effects of Father's Employment				
Father's employment	-0.042**	-0.010	-0.025**	-0.015
	(0.019)	(0.010)	(0.011)	(0.010)
Father's employment×Girl	0.029**	-0.003	0.020*	0.033***
	(0.013)	(0.011)	(0.011)	(0.009)
Spousal employment	0.020	-0.008	-0.005	-0.0005
	(0.018)	(0.012)	(0.008)	(0.0075)
Spousal employment×Girl	0.034***	0.020***	0.021***	0.028***
	(0.006)	(0.006)	(0.005)	(0.005)
Spousal earnings/1000	-0.690*	0.015	-0.076	-0.363**
	(0.392)	(0.285)	(0.215)	(0.175)
R-squared	0.69	0.69	0.69	0.69

Table A11: Difference-in-differences regression of children's college achievement on their parents' employment, with two base units of earnings as the threshold for being employed, Norway

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses are clustered at the family level. Spousal earnings are measured in thousand USD. There are 816,617 observations in each regression. The dependent variable is the education achievement in college at age 25. The spousal earnings refer to the father's earnings in Columns 1-4 and the mother's earnings in Columns 5-8. Independent variables include the mother's age and father's age when the child is born, dummy variables for birth orders, the gender specific cohort fixed effect and the family fixed effect.

Variable	Mean	Std. Dev.	Min	Max
College gender gap[b]	0.04	0.055	-0.11	0.367
Birth year	1968.5	10.969	1950	1987
LFPR, married women[c]	0.484	0.16	0.122	0.85
LFPR, married men[c]	0.975	0.014	0.898	1
Hours worked, married women[d]	0.4	0.123	0.092	0.695
Hours worked, married men[d]	1.041	0.075	0.818	1.403
Children living with one parent[e]	0.042	0.016	0.005	0.091
Median income, married women[f]	0.05	0.01	0.023	0.132
Median income, married men[f]	0.051	0.011	0.022	0.13

Table A12: Summary statistics of the U.S. data on college gender gap and labor force participation rates

Notes: [a]The final sample has 1938 observations of state-year. However it is unbalanced. For example, only 1382 observations have non-missing hours worked information for married women. State-year level variables with fewer than 50 observations are dropped.

[b]The college gender gap is the difference in the college achievement between females and males who are aged 25 or older.

[c]LFPR: labor force participation rate. These LFPR are calculated at the state level, averaging over the five-year interval.

[d] Hours worked is measured by hours worked per week divided by 40.

[e]Children living with one parent is measured by the unconditional probability of one adult having at least one child while the spouse is absent in the family.

[f]Median income is in millions of 2004 dollars, deflated by CPI-U.

B Detailed Proof of Proposition 1

$$\frac{\partial V(t,c;n)}{\partial t} = \tilde{\lambda}\left\{n\left[-\frac{\partial e_g(s,k)}{\partial s} + w\frac{\partial e_g(s,k)}{\partial k}\right] + (1-n)\left[-\frac{\partial e_b(s,k)}{\partial s} + w\frac{\partial e_b(s,k)}{\partial k}\right]\right\}$$

$$\begin{aligned}
\frac{\partial^2 V(t,c;n)}{\partial t \partial c} &= \tilde{\lambda}\left\{n\left[\frac{\partial^2 e_g(s,k)}{\partial s \partial k} - w\frac{\partial^2 e_g(s,k)}{\partial k^2}\right] + (1-n)\left[\frac{\partial^2 e_b(s,k)}{\partial s \partial k} - w\frac{\partial^2 e_b(s,k)}{\partial k^2}\right]\right\} \\
&= \tilde{\lambda}\left\{n\left[\frac{\partial^2 e_g(s,k)}{\partial s \partial k} - w\frac{\partial^2 e_g(s,k)}{\partial k^2}\right] + (1-n)\left[\frac{\partial^2 e_g(s,k)}{\partial s \partial k} - w\frac{\partial^2 e_g(s,k)}{\partial k^2}\right]\right\} \\
&= \tilde{\lambda}\left[\frac{\partial^2 e_g(s,k)}{\partial s \partial k} - w\frac{\partial^2 e_g(s,k)}{\partial k^2}\right] > 0
\end{aligned}$$

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: children are aged [0,5]						
LFPR, married women	0.012					
	(0.022)					
LFPR, married men		-0.030				
		(0.073)				
Children living with one parent			0.035			
			(0.041)			
Children living with mother alone				0.011		0.017
				(0.028)		(0.028)
Children living with father alone					-0.054	-0.059
					(0.084)	(0.085)
Observations	578	578	695	640	609	604
R-squared	0.63	0.62	0.57	0.57	0.58	0.58
Panel B: children are aged [6,11]						
LFPR, married women	0.010					
	(0.018)					
LFPR, married men		0.014				
		(0.061)				
Children living with one parent			0.062			
			(0.048)			
Children living with mother alone				0.070**		0.058**
				(0.031)		(0.028)
Children living with father alone					-0.041	-0.057
					(0.059)	(0.058)
Observations	695	701	875	789	741	733
R-squared	0.68	0.68	0.57	0.60	0.65	0.65
Panel B: children are aged [12,17]						
LFPR, married women	0.024					
	(0.017)					
LFPR, married men		0.002				
		(0.032)				
Children living with one parent			0.022			
			(0.039)			
Children living with mother alone				0.053**		0.041*
				(0.022)		(0.023)
Children living with father alone					-0.098**	-0.086*
					(0.050)	(0.050)
Observations	715	725	919	820	768	765
R-squared	0.75	0.73	0.61	0.68	0.69	0.69
Median incomes included	N	N	N	N	N	N

Table A13: Difference-in-differences analysis of college gender gap, U.S data, blacks only

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Robust standard errors are reported in parentheses. Estimations are weighted by state populations. The dependent variable is the gender gap in college achievement. The age distributions of the mother generation and the father generation for each cohort are calculated from U.S. Census data. Independent variables include birth year fixed effects, birth state fixed effects, and the median income of married women and men (Columns 6-7 only).

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

C Effects of Children's Gender Composition on Parental Labor Supply and Marital Status

	Mean	Std. Dev.	Min	Max
Family level				
First child being a girl	0.483	0.5	0	1
Fraction of girls	0.485	0.386	0	1
Number of children	2.093	1.065	1	16
Age of youngest child	6.927	5.24	0	18
Age of oldest child	10.267	5.263	0	18
Average children age	8.642	4.976	0	18
First two same gender[c]	0.505	0.5	0	1

	Mean	Std. Dev.	Min	Max	Mean	Std. Dev.	Min	Max
Parent level		**Mother**				**Father**		
In labor force	0.563	0.496	0	1	0.962	0.19	0	1
Employed	0.543	0.498	0	1	0.939	0.24	0	1
Age	36.19	6.806	25	58	38.696	7.581	25	65
Age at the 1st birth	25.923	4.921	18	40	28.429	5.852	14	65
Education[b]	2.552	0.974	1	4	2.643	1.055	1	4

Table C14: Summary statistics of the pooled U.S. Census data, 1950-2000

Notes: [a]The final sample has $2,667,516$ observations.

[b]$2,560,299$ mothers and $2,558,250$ fathers have education information.

[c]$1,838,743$ families have 2+ children.