

A Study on the Causal Effects of Childbirth on Maternal Labor Supply in Korea

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We estimate the causal effect of additional childbirth on mothers' probability of working by using the two-stage least squares estimation method. Samples of mothers drawn from the 10%–20% Restricted Access Samples of the 2000, 2005, 2010, and 2015 Censuses are used. We use children's gender composition as the instrumental variable for childbirth. We find that additional childbirth had significantly adverse effects on mothers' probability to work up to 2010, but it is no longer likely to be the case. The ordinary least squares estimates still show a strong negative correlation between childbirth and maternal labor supply, which, our results indicate, is driven mainly by endogenous self-selection. We find that the labor supply of less-educated mothers is more adversely affected by childbirths.

JEL Classification: J13, J21, J22

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I. Introduction

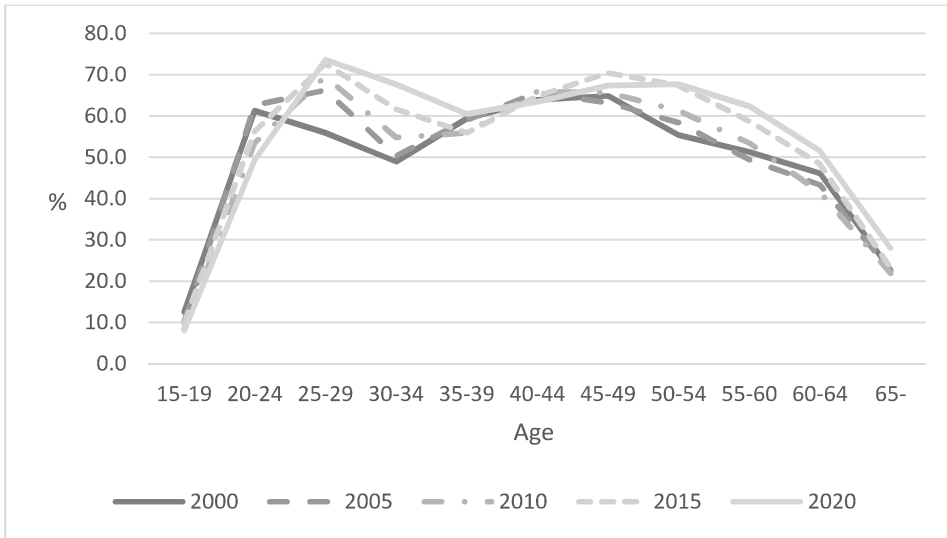
In Korea, childbirths appear to be strongly associated with a drop in the female labor supply. Figure 1 shows the female labor force participation rate (LFPR) by age every five years from 2000 to 2020. Female LFPR by age is M-shaped. Women in their 30s participate less in the labor market than women in their 20s or 40s. The shape has not changed for the last two decades, except for an increase over time of the age at the middle point of M. Figure 1 suggests that female labor force participation in Korea is strongly correlated with marriage, childbirth, and child-rearing. The LFPR curve shifts appear to be associated with the increase of the typical wedding and childbirth age of women.

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[Figure 1] Female labor force participation rate by age and year



Previous studies show that female LFPR in Korea drops after childbirth. Using the first wave of Korean Labor and Income Panel Study data of 1998, Kim (2003) calculates that 48.5% of women employed before childbirth stopped working within a year after childbirth. The trend is particularly strong among 35-to-44-year-old women and those who had been employees. Using the first wave of Korean Longitudinal Survey of Women and Families data of 2007, Kim and Woo (2010) calculate that 26.6% of women employed before childbirth stopped working within half a year after childbirth. They also find that women’s probability of working after giving birth changes little for three to four years but increases steeply after five to six years. The likelihood of dropping out of work is much lower after the second childbirth. Using the same dataset as Kim and Woo (2010) but a different sample, Lee (2014) reports that the probabilities of non-employment after the first and second childbirths are 33.9% and 14.2%, respectively, among 25-to-54-year-old married women employed before giving birth.

However, we should not consider the statistics as evidence that childbirth causes women to drop out of employment. Labor force participation decisions are likely to be made together with choices on having a child. For example, a couple wanting to have a child may plan the timing of birth and, accordingly, the timing of the mother stopping work. A married woman who prefers a higher income and tries to avoid the disadvantages associated with childbirth (e.g., lower earnings) would not have a child and keep working instead. The third factor, such as the preference for child or income, may generate the observed relationship between fertility and female labor supply.

Many studies use instrumental variables (IVs) to deal with the endogeneity of

childbirth. Popular IVs are twinning and gender combinations of children. For example, Rosenzweig and Wolpin (1980), Bronars and Grogger (1994), and Jacobsen, Pearce, and Rosenbloom (1999) use twinning as the IV for the number of children under the assumption that twinning is a random event. Using samples from the U.S. Population Census data of 1970 and 1980, Jacobsen, Pearce, and Rosenbloom (1999) take twinning at the first childbirth as the IV and estimate that an additional child decreases the probability of the mother working by 2 percentage points (%p).

However, in a recent study of 17 million observations across 72 countries, Bhalotra and Clarke (2019) find that twinning correlates positively with the mother's health and the prenatal environment's quality. They attribute the correlation to selective abortion and raise doubts about the validity of twinning as an IV. Their findings suggest that the IV estimates understate the (negative) effect of childbirth on female labor supply. Farbmacher, Guber, and Vikström (2016) also raise a similar issue, suggesting that high-income earners delay childbirth and increase the risk of having dizygotic twins through fertility treatment. They agree with Bhalotra and Clarke (2019) that IV estimates using twinning underestimate childbirth's effect on labor supply.

The studies that use gender combinations of children as the IV include Angrist and Evans (1998) and Cruces and Galiani (2007). The assumptions are that parental preference over gender combinations of children affects the number of children and that gender combinations (at least for the first two children) are random.

Angrist and Evans (1998) use a sample from the U.S. Population Census data of 1980 and take the first two children of the same gender as the IV. They estimate that having more than two children decreases the mother's probability of working by 12%p compared to having only two children. Their IV estimates are smaller than the ordinary least squares (OLS) estimates by about 6%p. Cruces and Galiani (2007) take a similar approach to data from Argentina and Mexico and estimate that having more than two children decreases the mother's probability of working by 10%p. Their IV estimates are not that different from the OLS estimates. Cho (2006) applies the same method in Korea using the Survey of Family Income and Expenditure data of 2000. She finds that the IV estimates are smaller than the OLS estimates and not statistically significant.

Some studies use other IVs. For example, Lundborg, Plug, and Rasmussen (2017) use the success/failure of in vitro fertilization as the IV. They estimate the effect of having a child on various labor market outcomes of women in Denmark.

The previous studies suggest that the IV estimation method is the leading method to deal with the endogeneity problem of childbirth. However, the availability of a valid IV and the generalizability of the results are the issue (Imbens and Angrist, 1994). The IV estimates of childbirth's effect on female labor supply

are often smaller than the OLS estimates, but it is not always the case.

Considering the strong negative correlation between childbirth and female labor supply and the history of robust policy responses to the issue, it is somewhat surprising that few scholars have tried to use an IV to estimate the *causal* effect of childbirth on female labor supply in Korea. Cho (2006) seems to be the only study that has tried it even in passing.

Our objectives for this study are twofold. First, we estimate the causal effect of additional childbirth on mothers' labor supply in Korea by using IVs. We use samples from the Population Census of 2000, 2005, 2010, and 2015. Our samples are drawn from the Censuses' restricted access data and are five to ten times as large as publicly available samples. We use the children's gender combinations as the IVs.

Second, we compare estimates over time to see whether childbirth has become less of a barrier to female labor supply. Public expenditure on childcare and early education in Korea has considerably expanded over the past two decades. According to official statistics (index.go.kr), the expenditure was negligible in 2000 and has since increased to 0.5% of the GDP in 2010 and 0.9% of the GDP in 2015. The greater availability and affordability of public and private childcare services should motivate (at least some) mothers to stay in or enter the labor market. We examine whether women are more likely to work after childbirth nowadays than, say, a decade ago.

The remainder of this paper is organized as follows. In Section II, we discuss the estimation method and describe the data we use. In Section III, we present the estimation results. Checks on the IVs' randomness, first-stage estimation results, naïve OLS estimation results, and second-stage estimation results are shown and discussed. We also show the results for fathers and compare the estimation results by mothers' education level. In Section IV, we conclude.

II. Estimation Method and Data

We use the two-stage least squares (2SLS) estimation method to estimate the causal effect of childbirth on labor supply. The outcome variable y_i is the dummy variable indicating labor market status, the endogenous explanatory variable n_i is the indicator of fertility, and the IV z_i is the gender mix of the children.

In the first stage, we estimate

$$n_i = \alpha_0 + \alpha_1 z_i + \alpha_2 X_i + \epsilon_i \quad (1)$$

where X_i is the vector of control variables. In the second stage, we estimate the following linear probability model:

$$y_i = \beta_0 + \beta_1 \hat{n}_i + \beta_2 X_i + \varepsilon_i \quad (2)$$

where \hat{n}_i is the predicted value. $\hat{\beta}_1$ is the estimated causal effect of childbirth on labor supply. For comparisons, we also estimate the OLS model by replacing \hat{n}_i with n_i .

We use the 10% samples of 2000, 2005, and 2010 Population Censuses and the 20% sample of the 2015 Census. The data are available only by restricted access.¹ The publicly available ones are 2% samples. We select for analysis the observations of 31-to-40-year-old women with a spouse who are household heads themselves or spouses of household heads and co-reside with all children to whom they gave birth. We estimate the effects of childbirth on work using a sample of women with two or more children and one or more children.

We focus on women in their 30s because the statistics show that they are more likely to drop out of the labor force than the other age groups. If childbirth has causal effects on work, then we expect the effect to be more prominent among this age group than the others.

[Table 1] Sample sizes

Year	Sampling percentage	Size of the whole sample: Persons (households)	Size of the sample analyzed: Persons (fraction of the whole)	
			Two or more children	One or more children
2000	10%	4,845,543 (1,433,422)	260,177 (5.4%)	314,485 (6.5%)
2005	10%	5,042,490 (1,591,631)	214,510 (4.3%)	272,390 (5.4%)
2010	10%	5,457,530 (1,801,705)	166,845 (3.1%)	232,290 (4.3%)
2015	20%	9,538,188 (3,588,655)	225,136 (2.4%)	342,931 (3.4%)

Table 1 shows the sample sizes. Note that the fraction of women selected for analysis decreases over the years, roughly by about 1%p every five years. This decrease reflects the demographic trend of ages at marriage and childbirth increasing over time and implies that our sample has become more selective as years pass. Thus, the recent sample is more likely to consist of women who have a stronger preference for having (more) children than the others. However, even though the sample has a wider age range, the selection would have become stronger over the years anyway if women's preference shifted in general to having fewer children.

The outcome variable ("working") is equal to 1 if the respondent has either worked or been on temporary leave from her job for the week and 0 otherwise. We use two endogenous explanatory variables, namely, the number of children and the

¹ For information on the restricted data, please visit the website https://mdis.kostat.go.kr/remote/remoteRequestList.do?curMenuNo=UI_POR_P9014

dummy variable indicating whether a mother has given birth to more than two children (compared to only two) or more than one (compared to only one). With the sample of mothers of two children or more, the IVs are dummy variables indicating whether the first two children are all boys, all girls, or a boy and a girl (omitted category). With the sample of mothers of one child or more, the IV is the dummy variable indicating the gender of the first child.

We control for the mother’s age and age squared, age at the first childbirth, and education. We also control for the husband’s age and age squared, husband’s education, and the family’s region of residence.

[Table 2] Distribution of the number of children in the sample of mothers by year

	One child	Two children	Three children	Four children	> 4 children	Total
2000	54,308 (17.3)	223,279 (71.0)	34,265 (10.9)	2,326 (0.7)	307 (0.1)	314,485 (100.0)
2005	57,880 (21.2)	183,016 (67.2)	29,563 (10.9)	1,731 (0.6)	200 (0.1)	272,390 (100.0)
2010	65,445 (28.2)	141,059 (60.7)	23,925 (10.3)	1,643 (0.7)	218 (0.1)	232,290 (100.0)
2015	117,795 (34.3)	187,193 (54.6)	34,797 (10.1)	2,767 (0.8)	379 (0.1)	342,931 (100.0)

Note: Shown are the unweighted numbers of observations in the sample of mothers of one child or more. The numbers in parentheses are the fractions in percentages.

Table 2 shows the distribution of the number of children in the sample of mothers by year. We can see that the fraction of mothers with only one child has increased over time while that of mothers with two children has decreased. The fraction of mothers with three or more children has reduced only slightly. Between 2000 and 2015, the fraction of mothers of one child has increased from 17% to 34% and that of mothers of two children has decreased from 71% to 55%.

Table 3 shows the weighted means of the variables for the two samples. The percentages of working mothers are almost identical between the two samples and have increased by about 3%p between 2000 and 2015. The average number of children has increased slightly among mothers of two or more children (panel A) and decreased by about 0.2 among mothers of one or more children (panel B). The decreasing proportion of mothers of two children, coupled with the stable ratio of mothers of three or more children, explains this trend.

In panel A, the gender composition of the first two children seems to have changed to be more likely to be all girls. The proportion of all girls has increased between 2000 and 2010 and has not changed since then. On the other hand, in panel B, the “first girl” share has been constant and consistent with the known natural sex ratio at birth, 1.05 males per 1 female. We will examine the issue of the

randomness of children's gender in a late section.

[Table 3] Weighted means of the variables

	Mothers of two or more children				Mothers of one or more children			
	Year				Year			
	2000	2005	2010	2015	2000	2005	2010	2015
Working	0.39	0.38	0.43	0.42	0.39	0.38	0.43	0.42
More than two children	0.14	0.14	0.15	0.16	-	-	-	-
More than one child	-	-	-	-	0.83	0.79	0.71	0.65
Number of children	2.15	2.15	2.16	2.17	1.95	1.90	1.83	1.76
All girls	0.232	0.236	0.244	0.244	-	-	-	-
All boys	0.254	0.255	0.259	0.259	-	-	-	-
First girl	-	-	-	-	0.488	0.488	0.489	0.488
Age	35.7	35.9	36.4	36.4	35.6	35.6	36.1	36.0
Age at first childbirth	25.5	26.1	26.8	27.9	26.0	26.7	27.7	29.0
Spousal age	38.9	39.1	39.5	39.3	38.7	38.7	38.9	38.7
College	0.24	0.36	0.47	0.64	0.26	0.40	0.51	0.67
College, spouse	0.40	0.50	0.56	0.67	0.41	0.53	0.59	0.69

Note: Shown are the weighted means calculated using the population weights.

The average ages of the mothers and spouses in the samples are around 36 and 39, respectively. These change little because we restrict the age range. The average age at the first childbirth, meanwhile, has increased over time, implying that the children are younger in more recent samples. The education level of the mothers and the spouses has increased substantially over time. The increase among mothers is far steeper, consistent with the national trend.

III. Estimation Results

3.1. Randomness of IV

To examine the randomness of the gender composition of the first two children, we estimate a multinomial logit model of three categories (all boys, all girls, and mixed) of gender composition on the mother's characteristics. Table 4 shows the p -values for the test of overall significance of the coefficients. For brevity, the estimated coefficients are omitted. The results indicate that gender composition correlates strongly with the mother's characteristics in 2000 and 2005 but not any

more in 2010 and 2015. In 2000 and 2005, maternal age, age at first childbirth, and spousal age are significantly correlated with gender composition.

[Table 4] Test of the randomness of gender composition of the first two children

	2000	2005	2010	2015
Number of observations	260,177	214,510	166,845	225,054
Test of overall significance ^a	6.66e-09	1.36e-05	0.379	0.161

Note: The multinomial logit model of gender composition is estimated using the sample of mothers with two children or more. The population weights are applied. The estimated coefficients are omitted for brevity. ^aThe *p*-values are shown.

[Table 5] Estimation results of the logit regression of the first child's gender (1 if female, 0 otherwise): Mothers of one child or more

	2000	2005	2010	2015
Age	.0063	-.0330	.0018	-.0766**
Age squared	-.0001	.0005	.0000	.0011**
Age at first birth	.0213**	-.0329**	.0086	.0194*
Age at first birth squared	-.0004**	.0006**	-.0002	-.0004*
Mother's education (omitted: high school)				
None	.0510	.0594	.0363	.3788
Primary	.0673***	-.0726	-.1070	.2194*
Middle	.0438***	.0563**	.0352	.1248***
Vocational College	-.0204	.0048	-.0083	.0013
University	-.0172	-.0073	-.0137	-.0043
Master	.0108	-.0055	-.0302	-.0068
Doctoral	-.0132	-.0488	.0370	N.A.
Spousal education (omitted: high school)				
None	-.0111	.0300	.2593*	-.0104
Primary	-.0963***	.0665	-.0164	.0073
Middle	-.0062	-.0069	-.0500	-.0372
Vocational College	-.0213*	-.0134	-.0112	-.0063
University	.0018	.0008	.0003	-.0051
Master	-.0241	-.0293	.0077	-.0079
Doctoral	.0379	-.0419	.0026	
Spousal age	-.0031	-.0122	-.0096	.0008
Spousal age squared	.00004	.0001	.0001	.00002
Number of observations	314,485	272,390	232,290	342,931
Test of overall significance ^a	3.89e-05	0.034	0.164	0.019

Note: The sample of mothers of one child or more is used. The population weights are applied. Estimated coefficients of the region of residence dummies and the intercept term are not shown. ^aThe *p*-values are shown. * $p < .10$ ** $p < .05$ *** $p < .01$.

Table 5 shows the estimation results of the logit regression of the first child's gender (1 if female, 0 otherwise) on the mother's characteristics using the sample of

mothers of one child or more. It is somewhat surprising that, according to the last row, the gender of the first child appears to be correlated with some aspects of the mother except in 2010. Still, we cannot reject the null hypothesis of zero correlation at the 1% level, which some would consider appropriate given the large sample size, in 2005 and onwards.

The education level and age at childbirth of the mother seem to be correlated with the first child's gender. Mothers with lower education appear to be more likely to give birth to a girl than those with high school education or higher. However, the mother's age does not seem to have a consistent relationship with the first child's gender. The signs of the coefficients seem to change year to year and the magnitudes are small. Although not shown in the table, five coefficients of region dummies are statistically significant at the 5% level in 2000, and a couple of coefficients are statistically significant since 2005. However, they are not consistent over the years.

All in all, possibly except for maternal age and age at first childbirth, parents' characteristics do not appear to have a systematic relationship with the children's gender composition through the years.

3.2. First-stage Results

In the first stage, we regress the endogenous variable on the IV and the control variables. Table 6 shows the key coefficient estimates from the first-stage regressions. Panel A summarizes the results for the sample of mothers of two or more children, and panel B summarizes the results for mothers of one child or more. For brevity, only the coefficient estimates of the IVs are shown. Within each panel, the first subpanel shows the results for the number of children and the second subpanel shows the results for the dummy variables, having more than two children in panel A and having more than one child in panel B. Each subpanel also shows the value of the Kleibergen-Paap rk Wald F statistic to test whether the IVs are weak.

The estimates in subpanel A-1 show that, for mothers of two or more children, the gender composition of the first two children has a strong correlation with the number of children. Except for the All-boys coefficient in 2000, the p-values of the coefficients are practically 0. The results show that in all years, having only two girls has led to more subsequent childbirths. The coefficient size, however, decreased over time. In 2000, mothers whose first two children were girls had, on average, 0.3 more children than those whose first two children were of mixed gender. Fifteen years later, the gap shrank to 0.1. On the other hand, in 2000, having only two boys led to slightly fewer subsequent childbirths than did having a girl and a boy. In 2005 and onwards, however, having only boys led to more subsequent childbirths, and the effect, though much smaller than having only girls, has increased over time.

The results in subpanel A-2 show that the gender composition of the first two

children is also strongly correlated with whether the mother has just two or more than two children. In 2015, mothers with two daughters and mothers with two sons were respectively 10% and 3% more likely to have an additional child than did the mothers of a boy and a girl. The weak IV test F values in subpanels A-1 and A-2 are far above the conventional critical values.

The results in panel B show that having a daughter as the first child leads to more subsequent childbirths. The standard errors are 1/10 or smaller of the coefficient estimates. In 2000, giving birth to a daughter increased the number of children on average by 0.2 and the probability of having more children by 6%. In 2015, the effects were still significant but shrank to 0.05 and 2%, respectively. The weak IV test F values in both subpanels are also far above the conventional critical values.

[Table 6] Summary of the first-stage estimation results

IV	2000	2005	2010	2015
(A) Sample of mothers of two or more children				
(A-1) Dependent variable: Number of children				
All girls	.3290*** (.0024)	.2322*** (.0024)	.1561*** (.0027)	.1082*** (.0023)
All boys	-.0029** (.0013)	.0059*** (.0016)	.0184*** (.0021)	.0309*** (.0020)
Weak IV test F	1.0e+04	4802.4	1719.3	1084.9
(A-2) Dependent variable: More than two children				
All girls	.2951*** (.0020)	.2129*** (.0021)	.1433*** (.0024)	.0967*** (.0020)
All boys	-.0022* (.0012)	.0062*** (.0015)	.0178*** (.0019)	.0298*** (.0018)
Weak IV test F	1.1e+04	5086.5	1841.1	1145.0
Observations	260,178	214,524	166,845	225,136
(B) Sample of mothers of one child or more				
(B-1) Dependent variable: Number of children				
First girl	.1939*** (.0018)	.1247*** (.0020)	.0812*** (.0023)	.0459*** (.0020)
Weak IV test F	1.1e+04	3967.2	1207.2	547.5
(B-2) Dependent variable: More than one child				
First girl	.0576*** (.0012)	.0333*** (.0014)	.0286*** (.0017)	.0180*** (.0015)
Weak IV test F	2155.8	572.1	276.2	160.3
Observations	314,485	272,405	232,290	342,931

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. The weak IV test F is the value of the Kleibergen-Paap rk Wald F statistic. * $p < .10$ ** $p < .05$ *** $p < .01$.

3.3. Main Results

Before we present the main results, let us look at the naïve OLS estimation results. Subpanels A-1 and B-1 of Table 7 show the key OLS estimates. More children are strongly associated with the lower probability of the mother working.

In addition, the association seems to have gotten stronger over the years. According to the results in panel A, a mother of three is less likely to work than a mother of two by about 7%p in 2000. In 2015, the difference grew to 9%p. In panel B, we find a similar trend for the mothers of one child or more. However, we should take the results with caution because they are potentially biased because of the endogeneity of fertility.

The second-stage 2SLS estimation results are shown in subpanels A-2 and B-2 of Table 7. Subpanel A-2 shows that the coefficient estimates are all negative, indicating that having more than two children reduces the mother's probability of working. However, there is a clear trend that the effect decreases over time. The coefficients are statistically significant only up to 2010. In 2000, the birth of the third child lowered the mother's probability of working by about 7%p, and the effect is statistically significant at the 1% level. Fifteen years later, the same event lowered the chance by only about 4%p, and the effect is not statistically significant even at the 10% level.

[Table 7] OLS and 2SLS estimation results

Explanatory variable	2000	2005	2010	2015
(A) Sample of mothers of two or more children				
(A-1) OLS estimation results				
More than two children	-.068*** (.003)	-.079*** (.003)	-.083*** (.003)	-.094*** (.003)
Number of children	-.055*** (.002)	-.072*** (.003)	-.077*** (.003)	-.086*** (.002)
(A-2) 2SLS estimation results				
More than two children	-.071*** (.007)	-.059*** (.011)	-.052** (.020)	-.043 (.027)
Number of children	-.064*** (.007)	-.054*** (.010)	-.048** (.019)	-.038 (.024)
Observations	260,178	214,524	166,845	225,136
(B) Sample of mothers of one child or more				
(B-1) OLS estimation results				
More than one child	-.074*** (.002)	-.095*** (.003)	-.098*** (.003)	-.082*** (.002)
Number of children	-.060*** (.002)	-.078*** (.002)	-.082*** (.002)	-.076*** (.001)
(B-2) 2SLS estimation results				
More than one child	-.116*** (.030)	-.081 (.055)	-.063 (.074)	-.007 (.093)
Number of children	-.035*** (.009)	-.022 (.015)	-.022 (.026)	-.003 (.037)
Observations	314,485	272,405	232,290	342,931

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. * $p < .10$ ** $p < .05$ *** $p < .01$.

According to the results in subpanel B-2 with the sample of mothers of one child or more, the birth of the second child has had no statistically significant effect on the mother's working status since 2005. The estimated coefficient size of the number of child variable in subpanel B-2 is less than half of that in subpanel A-2. The results suggest that the birth of the second child has a much smaller—if not zero—effect

on the mother's labor market status than does the birth of the third child.

Note that our main 2SLS results are starkly different from the naïve OLS estimation results. The former shows the decreasing coefficient size over time while the latter shows the opposite. This outcome suggests that self-selection between more children and work among mothers has gotten stronger over the years. As discussed in Section II, women's preference may have shifted in general to having fewer children over the years, and thus the mothers of more than two or one children have increasingly become a more self-selected group.

Our results suggest that changes in the gender norms and the increasing social support for working mothers appear to have lowered the barrier to employment to the point where an exogenous shock to fertility does not affect maternal labor supply. It seems now that, at least among mothers of one or more children, the endogenous choice for (more) children over work based on preference or personal circumstances is the main reason why the number of children and the mother's labor market status are so strongly correlated.

3.4. Results with the Expanded Age Range

One may wonder whether our restriction on women's age for the sample is too strong and how different the results would be with a weaker restriction. To answer the question, Table 8 shows the OLS and 2SLS estimation results with the sample of mothers who are 26–45 years old. The structure of Table 8 is identical to that of Table 7.

Each of the coefficient estimates with the expanded age range is smaller in size than its counterpart in Table 7. This result is in fact expected, as our main sample consists of women in the age range where their labor market status is most likely to be affected by childbirth. However, the differences are not large, ranging usually between 0.02 and 0.03. With the expanded age range, we still find that OLS coefficient estimates get larger in size as years pass. For example, in subpanel A-1, the OLS coefficient estimate of “more than two” is $-.052$ in 2000 and increases to $-.078$ in 2015.

In subpanel A-2 of Table 8, we find that the 2SLS coefficient estimates are all negative and that coefficients are statistically significant at the 5% or smaller level. This is somewhat different from the results in Table 7 where the coefficients in 2015 are not statistically significant at the 10% level. Except for 2005, the coefficient size gets smaller over the years like in Table 7. For example, the 2SLS coefficient estimate of “number of children” is -0.043 in 2000 but shrinks to $-.029$ in 2015.

The 2SLS estimates in subpanel B-2 of Table 8 are far smaller in size than their counterparts in Table 7. The coefficients are statistically significant in 2000 at the 5% level but not significant in later years even at the 10% level. This is also the case in Table 7.

All in all, the results with the expanded sample are qualitatively little different from the main results. Specifically, the effect of childbirth on the working status of the mother is overestimated by OLS and more and more so in recent years. The causal effect of childbirth has become weaker over the years.

[Table 8] OLS and 2SLS estimation results for 26-to-45-year-old mothers

Explanatory variable	2000	2005	2010	2015
(A) Sample of mothers of two or more children				
(A-1) OLS estimation results				
More than two children	-.052*** (.002)	-.057*** (.002)	-.067*** (.003)	-.078*** (.002)
Number of children	-.041*** (.002)	-.051*** (.002)	-.062*** (.002)	-.072*** (.002)
(A-2) 2SLS estimation results				
More than two children	-.049*** (.006)	-.028*** (.008)	-.039*** (.013)	-.032** (.016)
Number of children	-.043*** (.005)	-.025*** (.007)	-.035*** (.012)	-.029*** (.014)
Observations	407,321	336,388	272,307	404,713
(B) Sample of mothers of one child or more				
(B-1) OLS estimation results				
More than one child	-.071*** (.002)	-.078*** (.002)	-.081*** (.002)	-.062*** (.002)
Number of children	-.052*** (.001)	-.060*** (.001)	-.067*** (.001)	-.060*** (.001)
(B-2) 2SLS estimation results				
More than one child	-.050** (.023)	.029 (.036)	-.019 (.051)	.004 (.055)
Number of children	-.015** (.007)	.008 (.010)	-.006 (.016)	.002 (.020)
Observations	528,963	441,763	377,856	595,709

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. * $p < .10$ ** $p < .05$ *** $p < .01$.

3.5. Case of Fathers

We also estimate the effect of childbirth on the father's labor market status by using the same 2SLS estimation method, replacing mothers with fathers and spouses with mothers. Table 9 shows the OLS and 2SLS estimation results for fathers with two or more children. The OLS estimation results in panel A suggest that additional childbirth is positively and then negatively correlated with the father's probability of working over the years.

However, the 2SLS estimation results in panel B suggest that the correlation is not likely to be causal. The coefficient size of additional childbirth is practically 0 up to 2010. In 2015, the coefficient size is estimated to be about -0.01, or about a quarter of the estimated coefficient size for mothers, but it is not statistically significant.

[Table 9] OLS and 2SLS estimation results for fathers with two or more children

Explanatory variable	2000	2005	2010	2015
(A) OLS estimation results				
More than two children	.003** (.001)	.001 (.001)	-.013*** (.002)	-.001 (.001)
Number of children	.002* (.001)	.000 (.001)	-.013*** (.002)	-.001 (.001)
Observations	260,178	214,524	166,845	225,136
(B) 2SLS estimation results				
More than two children	-.001 (.003)	.001 (.005)	-.005 (.011)	-.013 (.010)
Number of children	-.001 (.003)	.001 (.005)	-.005 (.010)	-.011 (.009)
Observations	260,178	214,524	166,845	225,136

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. * $p < .10$ ** $p < .05$ *** $p < .01$.

3.6. Heterogeneity of the Effects: Results by Maternal Education

To examine the potential heterogeneous effects of childbirth on maternal labor supply, we divide the sample into two: mothers with post-secondary education and those with up to secondary education. Education is correlated with the quality of jobs one can obtain and possibly with the preference, which can be the source of heterogeneity.

[Table 10] Estimation results by maternal education for mothers of two or more children

Explanatory variable	2000	2005	2010	2015
(A) Mothers with post-secondary education				
(A-1) OLS estimation results				
More than two children	-.082*** (.006)	-.081*** (.006)	-.094*** (.005)	-.101*** (.004)
Number of children	-.075*** (.006)	-.079*** (.005)	-.086*** (.005)	-.093*** (.003)
(A-2) 2SLS estimation results				
More than two children	-.094*** (.024)	-.041 (.027)	-.047 (.040)	-.028 (.040)
Number of children	-.089*** (.022)	-.040 (.026)	-.044 (.038)	-.026 (.037)
Observations	62,451	76,688	77,018	141,923
(B) Mothers with up to secondary education				
(B-1) OLS estimation results				
More than two children	-.064*** (.003)	-.077*** (.003)	-.083*** (.004)	-.100*** (.004)
Number of children	-.052*** (.003)	-.069*** (.003)	-.076*** (.004)	-.090*** (.003)
(B-2) 2SLS estimation results				
More than two children	-.069*** (.008)	-.063*** (.012)	-.054** (.022)	-.058* (.033)
Number of children	-.061*** (.007)	-.057*** (.011)	-.049** (.02)	-.050* (.029)
Observations	197,726	137,822	89,827	83,213

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. * $p < .10$ ** $p < .05$ *** $p < .01$.

Table 10 shows the estimation results for mothers of two or more children. Panel A shows the results for mothers with post-secondary education and panel B shows the results for mothers with less education.

The OLS estimation results in subpanel A-1 indicate that, among the better-educated group, the birth of the third child is associated strongly with a drop in the probability of working by about 10%p in 2015. However, the 2SLS estimation results in subpanel A-2 imply that the causal effect of childbirth is weaker than the OLS estimates and not statistically significant since 2005.

The estimation results of the less-educated group suggest that while the OLS results still overstate the effect of childbirth on maternal labor supply, an additional childbirth has more serious adverse effects. For example, the birth of the third child lowers the chance of the better-educated mother working by 2.6%p but reduces the probability of the less-educated mother working by 5.0%p.

[Table 11] Estimation results by maternal education for mothers of one or more children

Explanatory variable	2000	2005	2010	2015
(A) Mothers with post-secondary education				
(A-1) OLS estimation results				
More than one child	-.110*** (.004)	-.108*** (.004)	-.105*** (.004)	-.100*** (.003)
Number of children	-.093*** (.003)	-.093*** (.003)	-.093*** (.003)	-.091*** (.002)
(A-2) 2SLS estimation results				
More than one child	-.125 (.080)	-.094 (.096)	-.052 (.100)	-.084 (.119)
Number of children	-.044 (.028)	-.033 (.034)	-.024 (.046)	-.037 (.053)
Observations	81,051	106,546	116,231	226,936
(B) Mothers with up to secondary education				
(B-1) OLS estimation results				
More than one child	-.052*** (.003)	-.078*** (.003)	-.096*** (.004)	-.082*** (.004)
Number of children	-.048*** (.002)	-.068*** (.002)	-.079*** (.002)	-.077*** (.002)
(B-2) 2SLS estimation results				
More than one child	-.114*** (.031)	-.059 (.066)	-.095 (.106)	.132 (.140)
Number of children	-.033*** (.009)	-.014 (.015)	-.026 (.029)	.045 (.047)
Observations	233,434	165,844	116,059	115,995

Note: The population weights are applied. The estimated coefficients of the other control variables are omitted. The numbers in parentheses are robust standard errors. * $p < .10$ ** $p < .05$ *** $p < .01$.

Table 11 shows the estimation results for the mothers of one or more children. The OLS estimation results are similar to those in Table 10. The coefficients of causal effects estimated by 2SLS for the two groups of mothers are not significant in most years—for the better-educated mothers in any year and the less-educated mothers since 2005. The coefficients of the number of children are in general smaller than those of their counterparts in Table 10.

The estimation results in Tables 10 and 11 indicate that the effect of childbirth

on the maternal labor supply has some degree of heterogeneity. Since better education leads to jobs that offer more stability, better maternity leave benefits, and higher earnings, the results are not surprising. The statistics calculated from the August Supplement of the Economically Active Population Survey data indicate that among female employees with up to secondary education, 46% are non-standard workers in 2005, 40% in 2010, and 35% in 2015, while among those with post-secondary education, 36% are non-standard workers in 2005, 31% in 2010, and 28% in 2015.

IV. Conclusion

In this study, we estimate the causal effect of additional childbirth on mothers' probability of working by using the 2SLS estimation method. We estimate these effects using Restricted Access Samples from the 2000, 2005, 2010, and 2015 Censuses, which are five to ten times as large in size as the public samples.

Our estimation results show that additional childbirth used to affect mothers' probability of working adversely, possibly up to 2010, but it is no longer the case. We also find that third or higher-order childbirths affect the labor supply of less-educated mothers more adversely. While the OLS estimates still show a strong negative correlation between additional childbirth and the probability of working, our results indicate that it is driven mainly by endogenous self-selection.

Note that we use samples of mothers who have already given birth to at least one child and estimate the effect of the second and higher-order childbirths on a mother's labor supply. Our estimates cannot say anything about the impact of the first childbirth on female labor supply. It is not possible to estimate the causal effect of the first childbirth owing to the lack of IV.

Our results should also be interpreted as the estimates of the local average effects (Imbens and Angrist, 1994). That is, we estimate the effects of childbirth on the labor supply of mothers whose decision to have another child is affected by the gender composition of their existing children.

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한국에서 출산이 어머니의 노동공급에 미치는 인과효과에 관한 연구

박 철 성* · 손 유 영**

초 록 이 연구에서 우리는 2SLS 추정법을 이용하여 추가 출산이 어머니의 취업 확률에 미치는 인과효과를 추정하였다. 통계청의 인가용 서비스에서 제공하는 2000년, 2005년, 2010년의 인구총조사 10% 표본, 2015년 인구총조사 20% 표본에서 어머니의 표본을 추출하여 분석했다. 기존 자녀의 성별 구성을 도구변수로 이용하였다. 우리는 2010년까지는 추가 출산이 어머니의 취업 확률에 유의하게 부정적인 영향을 주었지만, 그 이후로는 영향이 없었을 가능성이 큼을 발견하였다. OLS 결과는 출산과 어머니의 노동공급 간에 강한 음의 관계가 있음을 보여주는데, 우리 결과는 이것이 내생적인 자기 선택 때문임을 시사한다. 우리는 또한 출산이 학력이 낮은 여성의 노동공급에 더 부정적인 영향을 준다는 것을 발견하였다.

핵심 주제어: 출산, 출산율, 어머니의 노동공급, 성별 구성

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