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Changes in Korean Maternity Protection Law and Labor Market Outcomes for Young Women*

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Focusing on the Korean experience, particularly a recent amendment which extends maternity leave and increases financial benefits during maternity and childcare leave, this paper evaluates how such an expansion of benefits affects the employment and the hourly wages of young women of childbearing age. Empirical results from a difference-in-difference-in-differences model having older women, older men, and young men simultaneously as the control group suggest that neither the employment nor the hourly wages of young women are affected. This implies that the law change does not cause shifts in the labor supply curve and the labor demand curve for young women.

- Key Words: Maternity Protection Law, Maternity Leave, Childcare Leave, Women of Childbearing age, Employment Insurance

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I. Introduction

In order to increase young women's labor market attachment and fertility rate, Korea passed a revised maternity protection law in November 2001. From having been ensured 60 days of full-paid maternity leave and one year of unpaid childcare leave, now all female employees have come to enjoy an additional 30 days of maternity leave. Moreover, if qualified, one receives their usual pay up to a considerably high ceiling during this newly-added third month of maternity leave and modest financial support while taking childcare leave. This study attempts to estimate the impact of this change on the employment and the hourly wages of young women of childbearing age.

Economic theory says that the labor supply will increase in these conditions: childbearing age women who might quit their jobs due to inadequate provision of maternity leave now remain attached to their pre-birth employers, and women who otherwise would not enter the labor market before having a baby now participate to enjoy the lengthened maternity leave and the financial support. However, firms will in turn reduce their demand for young female workers because they bear extra costs, including the need to hire substitute workers and to increase the workload of other remaining workers due to the additional maternity benefits. Thus, the wages of young women should decrease, but how their employment will change is not clear a priori.

The effect of increasing maternity leave on young women's labor market outcomes has been studied in different institutional contexts. Among others, Waldfogel (1999) reports that the enactment of the U.S. federal law guaranteeing job protective leave affects neither the employment nor the wages of women with children. However, Ruhm (1998) uses European aggregate data and finds that slightly extending maternity leave increases the employment of young women, and especially a long spell of maternity leave also lowers their hourly wages. More recently, Lai and Masters (2005) study the effect of employer-provided maternity leave benefits in Taiwan and show that both the employment and the wages of young women are negatively influenced by extending benefits. This result is in contrast with Zveglich and Rodgers (2003) who use the same data but with different model specifications and conclude young women's employment increases, but their wages are unaffected.

Following empirical methodologies of these studies, I employ a difference-in-differences (DD), and further a difference-in-difference-in-differences (DDD) approach. The aforementioned studies have taken differences between gender, time, and as a last dimension, state, country, or industry. However, because Korean maternity protection law has been applied to almost all industries nationwide, I instead condition age range as the third difference by dividing the sample into the young and the old (Puhani & Sonderhof 2008). To measure the changes observed for the treatment group (young women), first, either older women or young men is considered as the control group in a DD model, and second, older women, young men, and older men altogether are used as the control group in a DDD model. Particularly, a DDD specification allows me to control for changes in gender discrimination and young cohort specific trend in employment and wage growth. Further, to account for the effect of unobservable heterogeneity such as motivation, innate ability, or persistent personal environment on coefficient estimates of covariates, in contrast to previous studies using time series cross sectional or aggregate national level data, I use longitudinal microdata and compare the estimated effects of the law change from ordinary least squares (OLS) with those from fixed effects regressions.

Conditional on usual human capital and demographic variables, all fixed effects regressions suggest that the law change did not affect young women's employment. Even when I further control for childbirth between the current and the following year's survey and, additionally, previous year employment status in the DDD model, the result still holds. On the contrary, the effect on the hourly wages of young women obtained from DD specifications varies remarkably by the control group. Whereas OLS and fixed effects estimates of the effect are positive and statistically significant when having older women as the control group, those estimates are not statistically different from zero when having young men as the control group. However, regardless of the estimation methods, DDD estimates show that the effect is positive, and the estimated effect is even statistically significant at the 5% level in the fixed effects regression. Taking into account the possibility that the correlation between employment status and hourly wages causes an upward bias on the estimated effect, I further employ a sample selection model suitable for panel data. Finally, the effect of the law change on hourly wages turns into negative, as the theory anticipates. Taking these together, I conclude that neither the employment nor the hourly wages of young women respond to the law change. This suggests that the costs to employers and the value to young female employees of the law amendment were too small to generate any shifts in the labor supply curve and demand curve for young women.

The next section discusses the background that provoked the changes in Korean maternity protection law and introduces the law change in detail. Section 3 presents the economic theory to explain its impact on the employment and the hourly wages of young women. Section 4 introduces data from the Korean Labor and Income Panel Study and an econometric methodology. Section 5 reports regression results, and section 6 concludes.

Ⅱ. Changes in Korean Maternity Protection Law

1. Background

Korea has continuously had a low female labor force participation rate. Although it has increased during the past 25 years from 43% in 1980 to 50% in 2007, it has been steadily lower than that of the U.S and Canada. Especially, in contrast to Koreans' renowned passion for higher education, the labor force participation rate of women with more than a college degree has been below 65% since 1990[1]. A couple of explanations of these low female labor participation rates have been presented, such as social bias towards women or gender discrimination in work places. However, according to the

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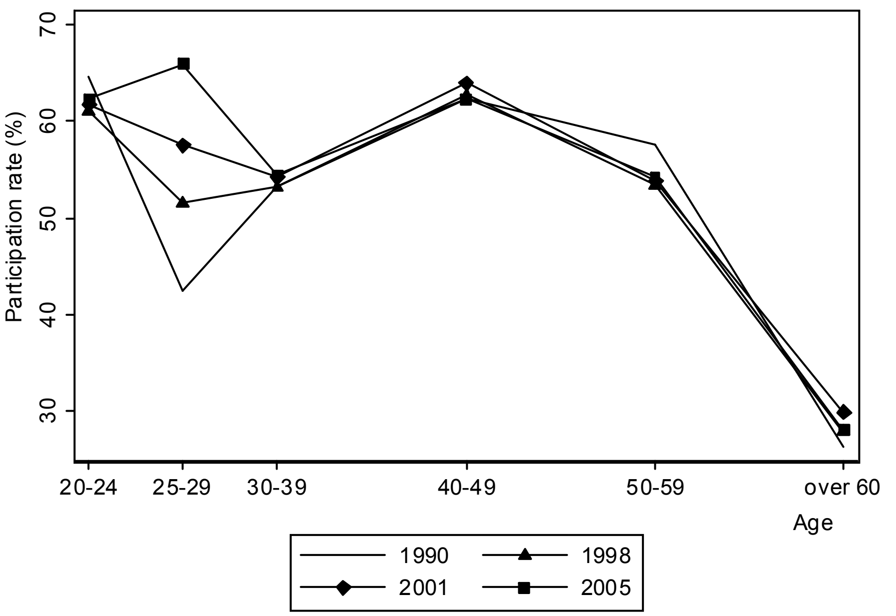
1) International Comparisons of Annual Labor Force Statistics, 10 Countries, 1960-2007, U.S Bureau of Labor Statistics; 2007 Labor Statistics, Korea Labor Institute.

1998 Social Survey, the burdens of childcare and housework were ranked as the main obstacles that keep women from having jobs.

Figure 1 displays Korean women's labor force participation rates in more detail by age group. Even though the lowest point has moved to the right due to delayed marriage and childbirth, it still shows the so-called M shape for 2005; the labor force participation rate is high in the 20's, decreases in the 30's, increases in the 40's to its level of the 20's, and starts decreasing again after the 50's. Kim (2003) and Chang et al. (2004) claim that a woman in her 30's is forced to leave the labor force until her children reach school age due to the burdens of motherhood and housework.

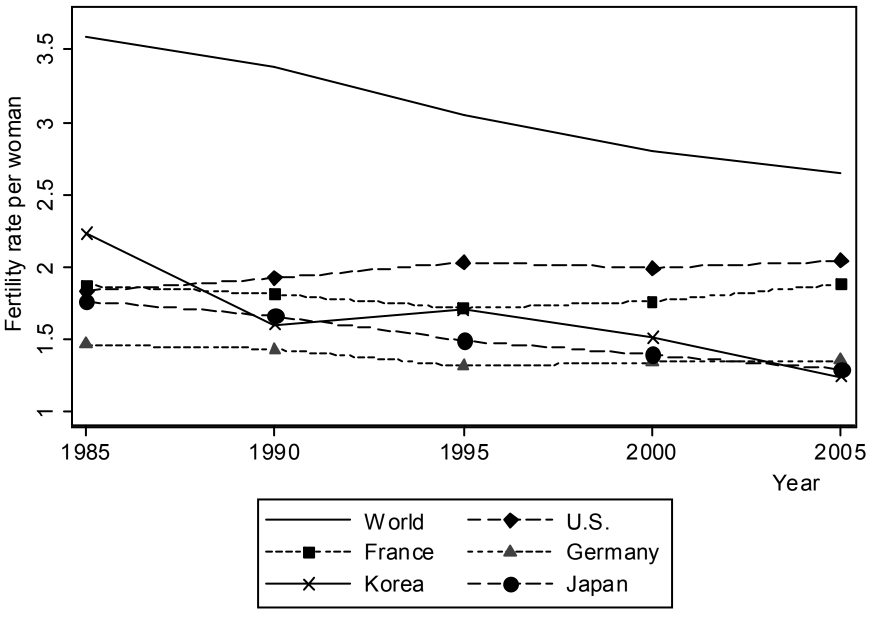
Additionally, low fertility rates have been another social concern. Figure 2 presents the fact that Korean fertility rates dropped remarkably between 1985 and 1990, and this decreasing trend continued beyond 2005. As of 2000, Korean fertility is recorded as the twenty third lowest among 151 countries. Aside from several reasons including high

〔Fig. 1〕 Korean Women's Labor Force Participation Rate by Age Group



Sources : Economically Active Population Survey of Korea National Statistical Office, cited by Korea Labor Institute.

〔Fig. 2〕 Cross Country Comparison of Fertility Rate



Sources : Korean National Statistical Office.

housing prices, excessive educational costs, and sudden cultural change, Chang et al. (2004) claim that young women have felt pressure to choose either family or work due to insufficient social support for raising children while keeping their jobs.

Under these circumstances, the Korean government recognized the need for providing young women with family and work balance to motivate them to join the labor force and to have more children. Thus, in June 2000 a revised maternity protection law was introduced to the National Assembly, and in November 2001 it came into effect.

2. The Law Change in Detail

The Korean maternity protection law consists of maternity leave and childcare leave. Prior to the amendment, a female worker could take 60 days time off at full pay from the pre-birth employer and enjoy up to one year of unpaid childcare leave until the child reaches one year old. After the revision, maternity leave expanded to include an additional 30 days on the condition that the latter half of the leave is assigned after childbirth. Although every female worker now enjoys 90 days of maternity leave and receives the full pay from her pre-birth employer for the first 60 days as before, financial support during the last 30 days depends on the term insured by Employment Insurance; if one has participated in Employment Insurance for more than 180 days before the last day of maternity leave, the insurance offers a third month benefit equal to the usual pay with a cap of 1,350,000 wons (1,047 U.S. dollars in 2001). However, when not qualified, neither the government nor the employer provides any benefit for the third month of leave.

When it comes to childcare leave, the law now guarantees all new mothers a maximum of 10.5 months of leave before a child turns one year old. Thus, the latter half of maternity leave, that is, 45 days, and the maximum period of childcare leave add up to one year of leave after childbirth. Before the law change, no female worker received any benefits during childcare leave; now Employment Insurance provides modest financial support to a new mother only if she was insured more than 180 days before the first day of childcare leave. The monthly childcare benefit began with 200,000 wons (154 U.S. dollars) in 2001, and it has gradually increased to 300,000 wons (240 U.S. dollars) in 2002 and to 400,000 wons (350 U.S. dollars) in 2004. Those not covered under Employment Insurance for at least 6 months are still allowed to enjoy the same length of childcare leave, but neither Employment Insurance nor the employer provides any financial support[2].

Employment Insurance has financed the third month of maternity leave and all periods of childcare leave benefits since November 2001. As one of four types of Korean social security insurance including Pension Service, Health Insurance, and Industrial Accident Insurance, Employment Insurance has provided unemployment insurance, employment stability, and vocational training to the insured since 1995. In its earlier stage, a business hiring more than 10 workers and its employees was required to participate in it.

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2) The maternity protection law has been implemented in practice by imposing a heavy penalty on employers who do not provide maternity or childcare leave to the extent that the law guarantees. As a result, the maternity protection law seems well executed (Ministry of Labor 2008).

In 1998, it expanded to include businesses hiring more than one employee while excluding certain industries such as government, agriculture, forestry, hunting, and private household service. Every insured employee had an obligation to pay 0.3% of one's income between 1995 and 1998, 0.5% between 1999 and 2002, and 0.45% from 2003 onward.

To my knowledge, data showing the total number of women who used maternity or childcare leave before the amendment does not exist. By offering an additional third month of maternity leave benefit, as well as childcare leave benefit to qualified workers through Employment Insurance since 2001, the Ministry of Labor has come to know the exact number of the insured workers who used maternity leave for three consecutive months and any length of childcare leave.

According to the governmental report, the number of women who utilize these benefits has steadily increased. As of 2005, the number of qualified women who received the third month of maternity leave benefit from Employment Insurance increased to 41,104 from 22,711 in 2002. When the number of women who took maternity leave is divided by the imputed number of women aged between 19 and 49 among Employment Insurance participants, who might have delivered a child[3)], 30% of the imputed number of new mothers used three months of maternity leave in 2003, and 37% in 2005. Childcare benefit recipients increased more rapidly. 17% of women who took three months of maternity leave also used childcare leave in 2001, and the rate increased by 9 percentage points in 2005.

The 2003 Maternity Protection Law Enforcement Rate Survey (Kim and Kim 2003) confirms that both the use of leave taking and the length of leave have increased after the maternity protection law change. Before the amendment, 49% of women who gave

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3) Because the number of insured women who bear a child is not available, the imputation is performed. The imputed number of new mothers is obtained by using the Korean age-specific fertility rate from 2002 to 2005 provided by the Korean National Statistical Office. Age-specific fertility rate is defined as 1000 times number of newborn baby in certain age group divided by total population in that age group. Because this rate does not distinguish working women from non-working women, the actual number of working mothers who bear a child should be smaller than the imputed one, and thus a greater percent of working mothers is expected to use three months of maternity leave and any length of childcare leave than that I calculate.

birth took more than 60 days maternity leave, and only 20% of women who took maternity leave further enjoyed childcare leave. However, after the revision, 65% of women reported already took or planned to take more than 90 days maternity leave once they have a child, and 53% of them already enjoyed or planned to enjoy any length of childcare leave.

Ⅲ. Theory

After the law change, all female workers have come to enjoy an additional month of maternity leave and, if qualified, receive financial support during this period and childcare leave from Employment Insurance. Although the benefits for the third month of maternity leave and all periods of childcare leave are paid by Employment Insurance, the subsequent increase in administrative and temporary replacement costs that firms need to bear may have deter them from hiring female workers. This implies that the demand curve for young women shifts downwards. Meanwhile, young female employees now shift their supply curve to the right by the amount that they value the new benefit (Summers 1989); Klerman and Leibowitz (1997) show theoretically that a woman who would otherwise quit a job due to insufficient maternity leave will remain with the pre-birth job when maternity leave is extended, and Baker and Milligan (2008) empirically validate this theoretical prediction by showing that longer maternity leave significantly prolongs time taken off of work after birth when keeping the previous job. In addition, more women could enter the labor market prior to having a child in order to enjoy the financial support during periods surrounding childbirth (Ruhm 1998). As a result, the wages of young women of childbearing age should fall, but the direction of the change in their employment level is not clear unless an empirical analysis is performed. If women do not value the extended benefit as much as the costs that employers bear, then their employment will fall as well. Conversely, when women value it by more than the costs to employers, their employment will increase (Lai and Masters

2005; Zveglich and Rodgers 2003). However, if the law revision had only an insubstantial impact on the labor market, then neither their employment nor wages would respond.

Ⅳ. Data and Methodology

The Korean Labor and Income Panel Study (KLIPS) launched with a representative sample of 5,000 households residing in Korean urban areas in 1998. Since then, it has annually surveyed socioeconomic characteristics of the households and their members aged 15 or older. Similar to the U.S. Panel Study of Income Dynamics, KLIPS allows a new person to join the sample only when he/she joins the household surveyed by marriage and family union, and it continues following the initial interviewees even when family separation occurs. The survey started with 13,321 individual respondents, and as of the 2008 wave, 11,739 individuals still remain.

While publicly accessible surveys are from 1998 to 2008, I exclude the 2006, 2007, and 2008 samples due to an additional major amendment to the Maternity Protection Law; initially only the third month of maternity leave pay came from Employment Insurance. However, since 2006, women working at small firms (firm size for this qualification differs across industries) have come to receive all three months of maternity leave benefits from Employment Insurance. Moreover, the child age applicable to childcare leave has changed from less than one year old to less than three years old. Nevertheless, I utilize the 2006 wave restrictively to find out whether women gave birth between the 2005 and 2006 waves, and to generate the amount of total family income in 2005 reported as previous total family income in the 2006 wave.

21,199 individuals were interviewed at least once between the 1998 and 2005 surveys. I eliminate 12,848 respondents who are neither *young* nor *old*. *Young* is defined as men and women aged between 20 and 35 in 2001 and *old* as men and women aged between 40 and 60 in 2001. I also discard 1,950 individuals who had not yet started their careers. Whether one started a career is measured through school enrollment status. If individuals who earned a degree, completed coursework, or dropped out do not go back to school through 2005, or stop interviewing at some point after leaving school, I assume they started their careers right after leaving school once they have their first jobs. However, current students and those who return to school after having some work experience are excluded. Also, I drop 622 respondents who report school level or enrollment status unrealistically or inconsistently compared to previous or future surveys are eliminated. Next, 125 employees working in sectors that are not participating in Employment Insurance such as government, education[4], agriculture, forestry, hunting and private household service are excluded[5]. Further, 1,246 individuals who had become self-employed at least one time between 1998 and 2005 are eliminated. Finally, I discard 1,014 respondents who do not provide complete information on the dependent and independent variables. These data restriction rules provide me an unbalanced panel[6] with 3,394 individuals with 21,159 observations for 8 years.

Because the 2001 survey was completed prior to the law change that occurred in November 2001, 1998 through 2001 surveys (inclusive) are considered as the pre-change period and 2002 through 2005 surveys (inclusive) as the post-change period. As a baseline specification, a difference-in-differences (DD) model regards young women as the treatment group and either older women or young men as the comparison group. However, young women and older women are obviously in different stages of their age-wages profiles (Becker 1962; Ben-Porath 1967; Mincer 1974; Heckman 1976). Also, the degree of gender discrimination in Korea against young women relative to young men might have altered within the survey period. Given both these possibilities, none of the control groups in the DD model are likely to face the same time trends with the

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4) Although those working for private educational businesses can participate in Employment Insurance, because educational service is not classified into public and private sectors in KLIPS, I drop the educational sector from the analysis.

5) Unfortunately, sample size in these sectors is too small to be used as the control group for a difference-in-differences specification.

6) Due to the way I restrict the data, the percent of the sample having a certain education level should be stable throughout the survey if attrition occurred at random. Indeed, it does not change substantially enough to generate concerns about sample selection issue.

treatment group even when the maternity law change was not considered. Thus, a difference-in-difference-in-differences (DDD) model having older women, young men, and older men simultaneously as the comparison group is required to control both trends in gender discrimination and the cohort difference in employment and wage growth. Holding controlled variables constant, comparing young women and older women detects the confounding effect of the young specific cohort trend and the law revision, assuming every woman faces the same level of gender discrimination. By taking another comparison between young men and older men, the young cohort effect can be identified. The last difference between the previous two differences finally yield the pure effect of the maternity protection law change.

However, I cannot rule out the possibility that the control group is also affected by the law change. Since hiring young women becomes more expensive than hiring young men, older men, or older women after the law change than before the law change, firms may substitute the control group for young women. Then, the demand curve for the control group will move upward. As a result, the employment and hourly wages of the control group will increase, implying that the DD and DDD models overestimate the effect of the law change on young women. On the contrary, if firms reduce production in response to the increase in the cost of hiring young female workers, the demand curve for the control group will move downward, and the employment and hourly wages of the comparison group will decrease. Consequently, the DD and DDD models will underestimate the effect of the law change on young women.

The DD models are specified as follows when having either older women or young men as the control group, respectively.

$$y_{it} = \beta_0 + X_{it}\beta_1 + young_i\beta_2 + young_i{*}after_{it}\beta_3 + after_{it}\beta_4 + T_t + e_{it}$$

$$\begin{aligned} y_{it} = {} & \beta_0 + X_{it}\beta_1 + X_{it}{*}female_i\beta_2 + female_i\beta_3 + female_i{*}after_{it}\beta_4 \\ & + after_{it}\beta_5 + T_t + e_{it}, \end{aligned}$$

y_{it} represents a dummy variable for $employment\,status$ or $the\,natural$

$logarithm\ of\ hourly\ wages$ deflated in 2005 wons using the Consumer Price Index. Whereas X_{it} is a vector of independent variables such as age, $age\ squared$, $marital\ status$, $the\ highest\ level\ of\ education$, $the\ number\ of\ children\ in\ certain\ age\ ranges$, $monthly\ other\ income$, and $region$ when studying the labor supply equation, it contains $actual\ experience$, $its\ quadratic\ \text{form}$, $actual\ tenure$[7], and dummy variables for $marital\ status$, $educational\ categories$, $firm\ size$, $union\ status$, and $region$ when the dependent variable is $\log - hourly\ wages$. $Young$ is set equal to 1 if an individual is between 18 and 35 years old in 2001, 0 if one is aged between 40 and 60 in 2001, $female$ is a dummy variable for being female, and $after$ represents the post-change period from 2002 to 2005. T_t represents a set of six dummy variables added to capture national year effects, e_{it} is a random error, and β_0 through β_5 are parameters to be estimated. When considering young men as the control group, interactions between most of variables in X_{it} and $female$ are included additionally. Here, the coefficient of $young^*after$ captures the effect of the law change and young cohort specific trend in the dependent variable, and that of $female^*after$ measures the effect of the law change and the change in the level of gender discrimination.

In DDD models, I exploit three dimensions, age range, gender, and time. Thus, not only $female$, $young$, $after$, $young^*after$, $female^*after$ but also $young^*female$ and $young^*female^*after$ should be controlled for. Therefore, the model is specified as follows.

$$\begin{aligned} y_{it} = &\ \beta_0 + X_{it}\beta_1 + X_{it}{}^*female_i\beta_2 + young_i\beta_3 + female_i\beta_4 + young_i{}^*after_{it}\beta_5 \\ &+ female_i{}^*after_{it}\beta_6 + young_i{}^*female_i\beta_7 + young_i{}^*female_i{}^*after_{it}\beta_8 \\ &+ after\beta_9 + T_t + e_{it} \end{aligned}$$

7) Especially, due to its complexity to generate actual experience and tenure using work history information in KLIPS data, almost all papers using this data let potential experience and age replace actual experience and tenure. I believe my effort to gain the exact year of those key variables reduces possible measurement errors in experience that may produce biased estimates, especially for women.

Except that the coefficient of $young*female*after$ now purely represents the effect of the law change, other variables are as above. In order to examine biases resulting from omitting unobservable heterogeneity as a regressor on coefficient estimates of the covariates, I run the DD and DDD models by both OLS and fixed effects regressions. In the fixed effects regressions, because e_{it} is the sum of time-invariant unobservable heterogeneity, α_i, and an idiosyncratic disturbance, ϵ_{it}, time-invariant explanatory variables become redundant.

It is possible that an idiosyncratic disturbance in employment status equation is correlated with that in log-hourly wage equation. If this is the case, the effect of the law change on hourly wages cannot be consistently estimated from Models (1), (2), and (3) by using only individuals whose salary or wages are observed. Also, as the maternity protection law change increases the flow of young female job applicants, firms may raise hiring standards and employ more productive young female workers after the law change than before the law change (Barron, Bishop, and Dunkelberg 1985; Barron and Bishop 1985). As a result, the hourly wages of young women may increase rather than decrease due to the law change, as opposed to the theoretical prediction. Thus, in order to correct for the sample selection bias and find the effect of the law change on the population of young women, I use the following two step procedure proposed by Wooldridge (1995).

In the first stage, for each *t,* Model (4) is estimated by the standard probit model.

$$P(employed_{it} = 1) = \Phi(Z_i \gamma_t)$$

Z_i includes all explanatory variables in all time periods except for $after$ and T_t used when the dependent variable is $employment\ status$. I do not control for $after$ *and* T_t in Model (4) because γ_t, the coefficient of Z_i, is allowed to vary by time. By employing $\widehat{\gamma_t}$, the coefficient estimate of γ_t, the inverse mills ratio $\widehat{\lambda_{it}}$ is computed as $\lambda(Z_i \widehat{\gamma_t})$ for each *i* and *t.*

In the second stage, Equation (5) is run by the pooled OLS.

$$\ln(hrwage_{it}) = Z_i\beta + W_{it}\theta + \widehat{\lambda_{i1}}^* T_1\delta_1 + \widehat{\lambda_{i2}}^* T_2\delta_2 + ... + \widehat{\lambda_{it}}^* T_t\delta_t + ... + \widehat{\lambda_{iT}}^* T_T\delta_T + v_{it}$$

W_{it} is a set of all independent variables conditioned on when the dependent variable is log-$hourly\ wages$, β is the coefficient of Z_i, θ is the coefficient of W_{it}, δ_1 through δ_T are the coefficients of the interactions between the inverse mills ratios and time dummy variables, and v_{it} is an random error. I use the nonparametric bootstrap method with 1,000 replicates to obtain the standard errors of the coefficient estimates. The null hypothesis of no selection bias, $H_0 : \delta_1 = \delta_2 = ... = \delta_t = ... = \delta_T = 0$, is tested using the F-statistic.

Using the restricted sample, I present year 1998 descriptive statistics of the variables used in the regressions by group in Table 1. As expected, both the employment rate and hourly wages of men are higher than those of women. While both statistics for young women are higher than those of older women, the opposite is true in men. Differences in years of experience and tenure between older women and older men are quite large, 11.9 years and 5.6 years respectively. This seems relevant to women's labor force participation rate showing the M-shape by age group. Since women leave the labor force due to childbirth, childcare, and housework in their late 20's and early 30's and come back to work in their 40's, years of experience and tenure fall far short of those of men. In contrast with the old, the young tend to have more than high school diplomas. Also, employees cluster at small firms hiring less than 100 employees. This is because as of 1998, 73.3% of total Korean employees are hired by firms that have less than 100 employees, 18.4% at firms with 100 to 999 employees, 8.4% at firms with more than 1000 employees[8]. However, the bigger a firm size is, the greater the number of male employees. Finally, relatively more males are union members than are females. Because employees at bigger firms or with union status are likely to earn higher wages, firm size and union membership will explain part of the gender wage difference.

Before running the aforementioned equations, I present the means for employment rate and hourly wages of young women, older women, young men, and older men,

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8) 1998 Census on Establishments, the Korean National Statistical Office.

〈Table 1〉 Descriptive Statistics

Variables in 1998	Young women	Older women	Young men	Older men
employed	0.35 (0.48)	0.28 (0.45)	0.64 (0.48)	0.70 (0.46)
hourly wages (wons)[a]	4,804.53 (2,403.48)	4,184.33 (2,343.62)	6,235.04 (2,952.85)	7,751.02 (3,964.20)
age	26.29 (3.73)	45.84 (6.08)	26.78 (3.86)	45.61 (6.16)
years of experience	4.03 (3.67)	7.47 (6.98)	3.97 (3.52)	19.33 (8.64)
years of tenure	2.46 (2.53)	3.83 (4.25)	2.75 (2.48)	9.46 (8.27)
married	0.58 (0.49)	0.90 (0.30)	0.33 (0.47)	0.9 (0.25)
≤elementary	0.02 (0.15)	0.34 (0.48)	0.03 (0.17)	0.19 (0.39)
middle	0.05 (0.21)	0.27 (0.45)	0.10 (0.31)	0.22 (0.42)
high	0.67 (0.47)	0.31 (0.46)	0.55 (0.50)	0.43 (0.49)
2 year college	0.14 (0.35)	0.02 (0.13)	0.12 (0.33)	0.04 (0.20)
≥4 year college	0.12 (0.33)	0.05 (0.22)	0.19 (0.39)	0.12 (0.32)
# kids 0-3	0.46 (0.62)	0.02 (0.17)	0.30 (0.54)	0.05 (0.24)
# kids 4-7	0.26 (0.55)	0.06 (0.26)	0.07 (0.29)	0.19 (0.45)
# kids 8-18	0.05 (0.27)	0.87 (0.89)	0.01 (0.09)	1.01 (0.89)
firm size≤100| employed=1	0.52 (0.50)	0.70 (0.46)	0.52 (0.50)	0.51 (0.50)
firm size 100～999| employed=1	0.27 (0.45)	0.23 (0.42)	0.26 (0.44)	0.29 (0.45)
firm size≥1000| employed=1	0.21 (0.41)	0.07 (0.26)	0.22 (0.41)	0.20 (0.40)
monthly other income[a] (1,000,000 wons)	1.66 (1.27)	1.95 (2.13)	1.16 (1.13)	1.14 (2.18)
union status| employed=1	0.16 (0.36)	0.07 (0.25)	0.20 (0.40)	0.29 (0.45)
seoul	0.44 (0.50)	0.45 (0.50)	0.43 (0.50)	0.39 (0.49)
# obs	648	889	524	704
# obs for women and men	1,537		1,228	

[a] 2005 exchange rate : 1,024.13 wons per dollar, The Bank of Korea. Standard deviations in parentheses.

employing only the 1998 and 2005 year data in Table 2. As shown in row (3), whereas young men were 20% more likely to be employed in 2005 compared to 1998, the 2005 employment rates for young women, older women, and older men are similar to their 1998 employment rates. Thus, using the two year data, if employment status is regressed only on $female$, $female*after$, and $after$, the coefficient estimate of $female*after$ is -0.20 (row (5)). However, if employment status is regressed only on $young$, $young*after$, and $after$, the coefficient estimate of $young*after$ is -0.01 (row (4)). To find the DDD estimation result when the demographic variables and year dummy variables are not conditioned on, I further subtract the change in employment rate of older men from that of young men (row (6)), and subtract the result from the difference between the change in employment rate of young women and that of older women (row (4)). As reported in row (7), young women are 19% less likely to be employed due to the maternity protection law change.

I do analogous calculations for hourly wages and present the results from row (8) to row (14) of Table 2. The wages of men increased more than that of women, and the young had higher wage growth than the old between 1998 and 2005. Thus, while row (11) reports that the difference between the change in hourly wages of young women and that of older women is 809.7, row (12) indicates that the difference between the change in the hourly wages of young women and that of young men is -2896.7. To sort out the trends in young-specific wage growth and gender discrimination in finding the effect of the law change on hourly wages of young women, I take another difference between young men's wage growth and older men's wage growth and subtract this difference from row (11). The result suggests that young women earn 453.7 wons less per hour as a result of the maternity protection law change.

Overall, Table 2 shows that the law change contributed to the decrease in both the employment and hourly wages of young women. This implies that young women value the new maternity benefits less than firms perceive them as costs. Now, I control for demographic factors and year effects in the DD and DDD models and run these models by OLS, fixed effects regressions, and sample selection models to see whether this result is stable.

〈Table 2〉 Mean Comparisons between Treatment and Control Group

	Employment rate	Young women	Older women	Young men	Older men
	(1) 2005	0.35	0.29	0.84	0.72
	(2) 1998	0.35	0.28	0.64	0.70
	(3) Row (1) - Row (2)	0	0.01	0.20	0.02
DD1	(4) Young women (3) - Older women (3)	-0.01			
DD2	(5) Young women (3) - Young men (3)	-0.20			
	(6) Young men (3) - Older men (3)			0.18	
DDD	(7) (4) - (6)	-0.19			
	Hourly wages	**Young women**	**Older women**	**Young men**	**Older men**
	(8) 2005	5115.1	4789.3	10425.3	10680.1
	(9) 1998	4790.2	4273.8	6203.7	7731.6
	(10) Row (8) - Row (9)	1324.9	515.5	4221.6	2948.5
DD1	(11) Young women (10) - Older women (10)	809.4			
DD2	(12) Young women (10) - Young men (10)	-2896.7			
	(13) Young men (10) - Older men (10)			1273.1	
DDD	(14) (11) - (13)	-463.7			

Ⅴ. Econometric Results

1. Employment

First, using the DD model, I examine the impact of the maternity protection law change on young women's employment. Here, older women and young men are considered as control groups in turn. The results are presented in Table 3. The first two columns are results obtained from OLS and the remaining two columns are from the fixed effects regressions. The coefficient of $young*after$ corresponds to the effect of the maternity law change plus the young cohort specific trend in employment, and the coefficient of $female*after$ represents the sum of the effect of the law change and the effect of changes in the degree of gender discrimination on young women's employment. When older women are considered as the control group, OLS shows that young women are 3.3% less likely to be employed during years after the law change. As they age, if the number of young women who drop out of the labor force due to motherhood is greater than the number of older women who retire, young women should be less likely to work on average during this period regardless of the law change. Alternatively, unobservable fixed effects may cause a downward bias on the coefficient estimate of $young*after$. To narrow down the possible explanations, I run a fixed effects regression and find that $young*after$ is not statistically significant anymore. This implies that unobserved heterogeneity is responsible for the underestimation of the coefficient of $young*after$, and the cohort difference in employment and the effect of the law change offset each other conditional on the fixed effects. Meanwhile, when young men is the counterpart of young women, no matter the estimation methods, no change in young women's probability to be employed is reported compared to the pre-change period. However, it is possible that gender discrimination has been alleviated to offset the negative effect of the law change, or vice versa (that is, that discrimination has increased to offset the positive effect). Thus, to find the pure effect of the law

〈Table 3〉 DD Analysis - Employment of Young Women

Dependent variable	Employed=1, 0 otherwise			
Estimation methods	OLS		Fixed	
Control group	Older women	Young men	Older women	Young men
age	0.014*** (0.004)	0.114*** (0.017)	0.015 (0.155)	0.138*** (0.357)
age*female		-0.107*** (0.024)		-0.103*** (0.028)
agesq*10^{-2}	-0.035*** (0.005)	-0.180*** (0.030)	-0.046*** (0.009)	-0.185*** (0.035)
agesq*10^{-2}*female		0.151*** (0.041)		0.157*** (0.047)
married	-0.168*** (0.013)	0.174*** (0.022)	-0.229*** (0.021)	0.042 (0.030)
married*female		-0.422*** (0.029)		-0.362*** (0.041)
≤elementary	0.085*** (0.014)	-0.279*** (0.041)		
≤elementary*female		0.021 (0.058)		
middle	0.051*** (0.013)	0.022 (0.022)		
middle*female		0.023 (0.037)		
2 year college	0.020 (0.016)	0.018 (0.020)		
2 year college*female		0.043* (0.026)		
≥4 year college	-0.032*** (0.016)	0.021 (0.017)		
≥4 year college*female		0.015 (0.025)		
# kids 0-3	-0.200*** (0.011)	0.022 (0.017)	-0.131*** (0.011)	-0.010 (0.018)
# kids 0-3*female		-0.183*** (0.021)		-0.096*** (0.023)
# kids 4-7	-0.071*** (0.010)	0.022 (0.016)	-0.042*** (0.011)	-0.007 (0.021)
# kids 4-7*female		-0.057*** (0.020)		-0.014 (0.027)
# kids 8-18	0.002 (0.007)	0.064*** (0.023)	-0.001 (0.009)	0.026 (0.032)
# kids 8-18*female		0.010 (0.027)		0.005 (0.040)

〈Table 3〉 *Continued*

Dependent variable	Employed=1, 0 otherwise			
Estimation methods	OLS		Fixed	
Control group	Older women	Young men	Older women	Young men
monthly other income[a] (1,000,000 wons)	-0.016*** (0.002)	-0.006*** (0.002)	-0.003** (0.002)	-0.004** (0.002)
seoul	0.016* (0.008)	0.044*** (0.009)	-0.007 (0.033)	0.040 (0.029)
young	-0.093*** (0.025)			
female		0.173*** (0.337)		
after	0.142*** (0.019)	0.076*** (0.022)	0.191 (1.085)	-0.056 (2.491)
young*after	-0.033* (0.018)		-0.024 (0.019)	
female*after		0.015 (0.019)		0.004 (0.027)
# individual			1,884	1,588
# obs	11,849	9,358	11,849	9,358
R - squared	0.126	0.323	0.068	0.223

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of *after* represents the 2005 year effect.
[a] 2005 exchange rates : 1024.13 wons per dollar, The Bank of Korea.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

revision on young women's probability to be employed, the DDD model is called for.

Before proceeding to DDD regressions, it is worthwhile to pay attention to other socioeconomic variables. First, note that dummy variables for $the\ level\ of\ education$ and their interactions with $female$ drop out from all fixed effects models that follow, because the sample consists of those who do not receive further education during the survey. Even though their estimated coefficients are likely to be contaminated with biases in OLS due to omitting unobservable characteristics from regressors, I discuss them throughout the paper assuming that the order in the magnitude does not change. When young women are compared to older women, the first column suggests that women who have less than a high school education are more likely to have jobs than women who have more than college degrees. By contrast, when having young men as the control group, the second column reports that young women with an education less than or equal to elementary school are less likely to be employed than young women with more than a middle school education. Thus, depending on age, two women having the same level of education may have different probabilities to be employed. As expected, married women or mothers with young children under age three are less likely to work, but the presence of young children does not affect young men in all columns. Married men are 17% more likely to work in OLS, but this marriage premium disappears in the fixed effects model. This indicates that more productive men are more likely to get married. Interestingly, for women, the coefficient estimate of $monthly\ other\ income$ is smaller in the fixed effects regression than in OLS, implying that unobservable characteristics are positively correlated with $monthly\ other\ income$.

Now, I run the DDD specification by OLS and the fixed effects models using older women, older men, and young men altogether as the comparison group. The coefficient of $young{*}female{*}after$ captures the pure impact of the law change without confounding it with the cohort difference or trend in gender discrimination. The coefficient estimates of this model are presented in Table 4. First, note that while the coefficient estimate of $young{*}female{*}after$ is statistically significant and positive in OLS, it is not statistically significant in the fixed effects regression. Moreover, the standard errors of the coefficient estimates are more or less small enough not to raise concerns about the credibility of the result. Comparing the fixed effects results between the DDD and the DD specification particularly implies that during the sample period, as many older women stop working as young women, and the level of gender discrimination in employment stays unchanged.

The rest of the coefficient estimates of the DDD model shows quite expected results: Whereas men do not differ in the probability to be employed if they have more than an elementary school education, women are more likely to be employed if they are less educated than high school. Thus, the employment pattern of women seems to be determined by older women, who are more likely to be employed if they are less educated in contrast to young women. Other variables discussed using the DD model

〈Table 4〉 DDD Analysis - Employment of Young Women

Dependent variable Estimation methods	Employed=1, 0 otherwise OLS	 Fixed
age	0.057*** (0.004)	0.036 (0.139)
age*female	-0.044*** (0.006)	-0.023** (0.012)
agesq*10^{-2}	-0.077*** (0.005)	-0.065*** (0.010)
agesq*10^{-2}*female	0.042*** (0.007)	0.020 (0.014)
married	0.192*** (0.015)	0.012 (0.023)
married*female	-0.361*** (0.019)	-0.240*** (0.030)
≤elementary	-0.060*** (0.015)	
≤elementary*female	0.147*** (0.020)	
middle	-0.015 (0.013)	
middle*female	0.067*** (0.017)	
2 year college	0.014 (0.017)	
2 year college*female	0.005 (0.022)	
≥4 year college	-0.001 (0.013)	
≥4 year college*female	-0.032* (0.020)	
# kids 0-3	0.012 (0.014)	-0.007 (0.014)
# kids 0-3*female	-0.212*** (0.017)	-0.124*** (0.018)
# kids 4-7	0.008 (0.012)	-0.012 (0.013)
# kids 4-7*female	-0.078*** (0.015)	-0.030* (0.017)
# kids 8-18	0.020*** (0.008)	-0.009 (0.010)
# kids 8-18*female	-0.018* (0.010)	0.011 (0.013)
monthly other income[a] (1,000,000 wons)	-0.015*** (0.001)	-0.005*** (0.001)

〈Table 4〉 *Continued*

Dependent variable Estimation methods	Employed=1, 0 otherwise OLS	 Fixed
seoul	0.018*** (0.006)	0.015 (0.025)
young	0.066** (0.026)	
female	0.976*** (0.135)	
after	0.167*** (0.016)	0.194 (0.969)
young*after	-0.090*** (0.019)	-0.025 (0.020)
female*after	-0.022 (0.016)	0.002 (0.021)
young*female	-0.160*** (0.034)	
young*female*after	0.058** (0.025)	0.004 (0.028)
# individuals		3,394
# obs	21,159	21,159
R - Squared	0.298	0.237

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of *after* represents the 2005 year effect.
[a] 2005 exchange rates : 1024.13 wons per dollar, The Bank of Korea.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

show exactly the same pattern.

Previous DDD regression results may be inconsistent and/or insignificant by omitting variables that are correlated with an idiosyncratic error even after conditioning on unobserved heterogeneity. Thus, as a robustness check, I additionally include variables to the DDD model that may have a causal effect on each dependent variable but were not controlled for.

Studying the employment status of young women, I postulate that a woman who expects a newborn baby in the near future may stop working in advance to prepare for delivery and childcare. Also, it is possible that individuals who are employed in the past may be more likely to be employed in this period because job search costs tend to be lower to the employed than to the unemployed (Hyslop 1999). In this sense, I add

$expbaby$, $expbaby*female$, and $employed_{t-1}$ to the DDD employment equation. $Expbaby$ is a dummy variable indicating whether one has a new child between the current and next year's surveys, and $employed_{t-1}$ is a dummy variable for a previous year employment status. $Employed_{t-1}$ is an endogenous variable in a fixed effects model because the time demeaned data on ϵ_{it}, ϵ_{it}-ϵ_{it-1}, where ϵ_{it} is an idiosyncratic error, is correlated with the time demeaned data on $employed_{t-1}$, $employed_{t-1}$-$employed_{t-2}$. Thus, assuming that ϵ_{it}-ϵ_{it-1} is not correlated with $employed_{t-2}$, and using the fact that $employed_{t-1}$-$employed_{t-2}$ and $employed_{t-2}$ are correlated by construction, I employ $employed_{t-2}$ as an instrument variable for $employed_{t-1}$ (Hyslop [1999]). Since these new variables are identified through interviewing the same individuals for every survey, using only fixed effects regressions seems to make sense. I run fixed effect regressions, first including $expbaby$ and its interaction with $female$ and then further adding $previous\ \ employment\ \ status$.

The results are presented in Table 5. The first column shows, as expected, that women who have a baby between this and the following survey are 14% less likely to be employed than women who do not expect a baby within a year. Now, the coefficient estimate of $marital\,status$ interacted with $female$ decreases considerably, compared to that obtained from the fixed effects regression applied to the previous DDD model. However, most importantly, the coefficient estimate of $young*female*after$ does not change noticeably, and it is still statistically insignificant. The second column reports the DDD results after additionally controlling for $employed_{t-1}$. An individual who was employed in the last period is 23% more likely to be employed in this period. When controlling for $employed_{t-1}$ as an independent variable, the negative effect of young children under age seven on the employment of women diminishes markedly compared to that reported in the first column. This implies that having young children and previous employment status are negatively correlated. Although the coefficient estimate of $young*female*after$ changes its sign to negative, it is still statistically insignificant. Therefore, the conclusion that the maternity law change did not cause change in the employment of young women remains sound.

〈Table 5〉 Robustness Check - Employment of Young Women

Dependent variable Instrumental variable	Employed=1, 0 otherwise	
		$employed_{t-2}$
expbaby	0.038* (0.022)	-0.000 (0.029)
expbaby*female	-0.136*** (0.028)	-0.058 (0.037)
$employed_{t-1}$		0.229*** (0.033)
age	0.038 (0.139)	-0.007 (0.362)
age*female	-0.027** (0.012)	0.019 (0.032)
$agesq*10^{-2}$	-0.065*** (0.010)	-0.018 (0.028)
$agesq*10^{-2}*female$	0.023* (0.014)	-0.019 (0.037)
married	0.001 (0.023)	-0.001 (0.040)
married*female	-0.203*** (0.031)	-0.175*** (0.055)
# kids 0-3	0.003 (0.015)	-0.012 (0.029)
# kids 0-3*female	-0.157*** (0.019)	-0.086** (0.038)
# kids 4-7	-0.004 (0.014)	0.010 (0.028)
# kids 4-7*female	-0.056*** (0.018)	-0.039 (0.036)
# kids 8-18	-0.006 (0.010)	0.013 (0.021)
# kids 8-18*female	-0.001 (0.014)	-0.030 (0.028)
monthly other income[a] (1,000,000 won)	-0.005*** (0.001)	-0.001 (0.001)
seoul	0.015 (0.025)	-0.002 (0.044)
after	0.188 (0.968)	0.076** (0.038)
young*after	-0.027 (0.020)	-0.008 (0.027)

〈Table 5〉 *Continued*

Dependent variable Instrumental variable	Employed=1, 0 otherwise	
		$employed_{t-2}$
female*after	0.002 (0.021)	0.009 (0.023)
young*female*after	0.008 (0.028)	-0.023 (0.036)
# individuals	3,394	2,757
# obs	21,159	10,490
R - squared	0.242	0.128

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of after represents the 2005 year effect.
[a] 2005 exchange rates : 1024.13 wons per dollar, The Bank of Korea.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

2. Hourly Wages

Next, I gauge the impact of the maternity law change on young women's hourly wages. I use the Equations (1), (2), and (3) and estimate them by OLS and fixed effects models as before. Compared to the study on employment, now, y_{it} is $the\ natural\ logarithm\ of\ hourly\ wages$, and X_{it} includes variables representing the amount of human capital accumulated at work places and job characteristics, instead of age and family income excluding observed wages. Thus, the key variables, $young{*}after$ and $female{*}after$ in the DD model and $young{*}female{*}after$ in the DDD model are interpreted in terms of hourly wages. Tables 6 and 7 report the results of the DD and the DDD models respectively.

Let's first focus on Table 6. This time, the coefficient estimates of $young{*}after$ and $female{*}after$ differ remarkably depending on the control group. Regarding older women as the comparison group, OLS reports that the mean wage of young women increased by 10.5% during the post-change period. However, the law change should drop the hourly wages of young women if it was substantial enough to bring the changes in the labor supply curve of and the labor demand curve for young women. Therefore, 〈Table 6〉 DD Analysis - Hourly Wages of Young Women

Dependent variable	Ln(Hourly wages)			
Estimation methods	OLS		Fixed	
Control group	Older women	Young men	Older women	Young men
exp	0.015*** (0.003)	0.054*** (0.006)	0.013 (0.013)	0.050*** (0.013)
exp*female		-0.025*** (0.008)		-0.013 (0.016)
expsq*10^{-2}	-0.041*** (0.009)	-0.219*** (0.030)	-0.007 (0.019)	-0.175*** (0.041)
expsq*10^{-2}*female		0.149*** (0.038)		-0.010 (0.080)
tenure	0.027*** (0.002)	0.025*** (0.003)	0.015*** (0.004)	0.016*** (0.004)
tenure*female		0.016*** (0.005)		0.012 (0.008)
married	0.045*** (0.017)	0.121*** (0.016)	0.071* (0.040)	0.033 (0.028)
married*female		-0.102*** (0.026)		0.031 (0.052)
≤elementary	-0.237*** (0.023)	-0.290*** (0.062)		
≤elementary*female		-0.145 (0.111)		
middle	-0.164*** (0.022)	-0.125*** (0.027)		
middle*female		-0.076 (0.051)		
2 year college	0.220*** (0.026)	0.131*** (0.022)		
2 year college*female		0.078** (0.033)		
≥4 year college	0.442*** (0.029)	0.359*** (0.019)		
≥4 year college*female		0.081** (0.035)		
firmsize 100~999	0.048* (0.027)	-0.026 (0.026)	0.065*** (0.028)	0.029 (0.027)
firmsize 100~999*female		0.162*** (0.044)		0.075 (0.047)
firm size ≥ 1000	0.060*** (0.015)	0.069*** (0.017)	0.017 (0.017)	0.037** (0.017)

〈Table 6〉 *Continued*

Dependent variable	Ln(Hourly wages)			
Estimation methods	OLS		Fixed	
Control group	Older women	Young men	Older women	Young men
firm size ≥ 1000*female		-0.008 (0.026)		-0.034 (0.029)
union	0.079*** (0.028)	0.058*** (0.020)	-0.009 (0.034)	0.038 (0.024)
union*female		0.002 (0.039)		-0.032 (0.045)
seoul	0.130*** (0.014)	0.114*** (0.011)	0.203** (0.080)	0.067* (0.040)
young	0.031 (0.025)			
female		-0.188*** (0.033)		
after	0.080** (0.033)	0.250*** (0.027)	-0.117* (0.066)	0.342*** (0.067)
young*after	0.105*** (0.028)		0.155*** (0.027)	
female*after		-0.033 (0.025)		-0.009 (0.035)
# individual			1,031	1,193
# obs	3,989	5,293	3,989	5,293
R - squared	0.300	0.452	0.126	0.272

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of *after* represents the 2005 year effect.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

unobserved fixed effects and the upward sloping age-wages profile should be responsible for this increase in their hourly wages. Nevertheless, the fixed effects regression reports that the increase is even greater by about 50 percent. This implies that young and older women are at different stages on their profiles, which outweighs the probable negative effect of the law change. On the other hand, when having young men as the control group, the mean hourly wages of young women do not increase compared to those before the law change. This result may be due to the law change having no effect on the labor market, and the degree of gender discrimination staying unchanged during the sample period. Conversely, the negative effect of the law change may be offset by the alleviated discrimination. Again, it is not possible to find just the effect of the law change unless I employ the DDD model.

Before moving into the DDD analysis, I take a look at coefficient estimates of other variables, reported in Table 6. Interestingly, gender difference in hourly wages depends on $the\ level\ of\ education$. The second column shows that no wage difference exists by gender among young people with less than a high school education, but young women who achieved more than associate degrees have higher wages than comparable young men. The comparison between OLS and fixed effects results suggests that market productivity partly determines the $marital\ status$ of young men. On the contrary, omitting unobserved fixed effects from regressors does not always cause upward biases on coefficient estimates. Among women, the effects of having another $year\ of\ experience$ or $getting\ married$ are greater in the fixed effects regression rather than those in OLS, implying that the innate ability of women may have negative correlation with these variables.

Going back to the discussion of the effect of the law change, Table 7 reports that while the coefficient estimate of $young{*}female{*}after$ is not statistically significant in OLS, the fixed effects regression finds the estimate to be 0.074 and statistically significant at the 5% level. Given that employment has not changed after the law revision, if the law amendment has indeed induced young women to increase their labor supply and firms to reduce hiring young women, then their wages must decrease. At the very least, if this law change did not cause any response from employers and female workers, no wage adjustment should occur. Furthermore, even if OLS reports the coefficient estimate of $young{*}female{*}after$ with an upward bias, the within regression must estimate the effect of the law change as negative to be consistent with the theory prediction. This leads me to set up a more robust model that takes a correlation between employment status and hourly wages into account.

Before studying the sensitivity of these results, for now, let's compare coefficient estimates of other variables between OLS and the fixed effects regression. When not only young women and young men but also older women and older men are analyzed by OLS, no more gender difference in hourly wages by the level of education is

〈Table 7〉 DDD Analysis - Hourly Wages of Young Women

Dependent variable Estimation methods	Ln(Hourly wages)	
	OLS	Fixed
exp	0.026*** (0.002)	0.057*** (0.009)
exp*female	-0.011*** (0.004)	-0.036*** (0.009)
expsq*10^{-2}	-0.063*** (0.005)	-0.079*** (0.013)
expsq*10^{-2}*female	0.022** (0.011)	0.062*** (0.021)
tenure	0.021*** (0.001)	0.013*** (0.002)
tenure*female	0.006*** (0.002)	0.001 (0.004)
married	0.160*** (0.015)	0.048* (0.026)
married*female	-0.117*** (0.022)	0.021 (0.047)
≤elementary	-0.243*** (0.022)	
≤elementary*female	0.001 (0.032)	
middle	-0.187*** (0.016)	
middle*female	0.018 (0.027)	
2 year college	0.200*** (0.019)	
2 year college*female	0.022 (0.032)	
≥4 year college	0.409*** (0.016)	
≥4 year college*female	0.044 (0.033)	
firmsize 100~999	-0.022 (0.019)	0.005 (0.019)
firmsize 100~999*female	0.072** (0.033)	0.060* (0.033)
firm size ≥ 1000	0.068*** (0.012)	0.011 (0.012)
firm size ≥ 1000*female	-0.010 (0.020)	0.005 (0.020)

〈Table 7〉 *Continued*

Dependent variable Estimation methods	Ln(Hourly wages)	
	OLS	Fixed
union	0.058*** (0.015)	0.037** (0.017)
union*female	0.019 (0.032)	-0.049 (0.037)
seoul	0.079*** (0.009)	0.055 (0.040)
young	0.031 (0.021)	
female	-0.156*** (0.040)	
after	0.186*** (0.023)	0.121** (0.049)
young*after	0.068*** (0.022)	0.078*** (0.022)
female*after	-0.054** (0.025)	-0.024 (0.030)
young*female	-0.005 (0.033)	
young*female*after	0.040 (0.036)	0.074** (0.034)
# individual		2,349
# obs	10,923	10,923
R - squared	0.442	0.173

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of *after* represents the 2005 year effect.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

observed. This indicates that gender discrimination by *the level of education* mostly occurs within the young. However, the positive selection into *marriage* for men does not depend on the generation. Also, the effects of all other variables fall in the fixed effects model compared to that in OLS, expect for *experience* and *experience squared* interacted with *female*.

In the previous DDD results, the effect of the law change on hourly wages is positive in both OLS and the fixed effects regression, and it is even greater and statistically significant in the fixed effects regression. Because the theory predicts that the law change should cause a decrease in young women's wages, I presume the coefficient of

$young*female*after$ was inconsistently estimated from Equation (3) due to a correlation between employment status and hourly wages. The effect of the law change can be positive, for example, if more productive women happen to be hired, or if the control group consists of less productive individuals after the law change. I attempt to correct this possible sample selection bias in the DDD model by estimating Equations (4) and (5). The coefficient estimates of the explanatory variables, which are also controlled for in Equation (3), and the inverse mills ratios are reported in the third column of Table 8.

First, according to the F test, the coefficient estimates of the inverse mills ratios are mostly negative, and they are jointly statistically significant at the 1% level. This implies that there is a negative correlation between employment status and the level of hourly wages within the sample period. Thus, the sample selection model should provide the consistent coefficient estimate of $young*female*after$, as opposed to the fixed effects regression. Indeed, the coefficient estimate of $young*female*after$ is negative at -0.011 in accordance with the theoretical prediction, although it is not statistically different from zero. Because the law change has no effect on the employment of young women either, it seems that the law change does not cause any shifts in the supply curve and the demand curve for young women.

Interestingly, the inverse mills ratios for years 2004 and 2005, which belong to the post-change period, are individually statistically significant at the 1% level. This suggests that individuals who earn less are more likely to be employed after the law change. To find out which group contributes the most to the negative correlation between employment status and hourly wages during the survey years and especially in 2004 and 2005, I run the sample selection model separately for four groups: young women, older women, young men, and older men. The results show that the coefficient estimates of the inverse mills ratios are jointly significant and mostly negative only within young men and older men, and particularly they are individually significant and negative for the post-change period (available upon request). Thus, less productive men are more likely to be hired during the post-change period than the pre-change period. However, 〈Table 8〉 Sample Selection Model - Hourly Wages of Young Women

Dependent variable	Ln(Hourly wages)		
Empirical model	DD		DDD
Control Group	Older women	Young men	All
exp	0.017*** (0.003)	0.035*** (0.006)	0.022*** (0.003)
exp*female		0.007 (0.011)	-0.007* (0.004)
expsq*10^{-2}	-0.044*** (0.009)	-0.160*** (0.033)	-0.042*** (0.005)
expsq*10^{-2}*female		0.063 (0.052)	0.002 (0.012)
tenure	0.025*** (0.002)	0.030*** (0.003)	0.020*** (0.001)
tenure*female		0.003 (0.005)	0.005** (0.002)
married	0.091** (0.046)	0.021 (0.032)	0.097*** (0.030)
married*female		0.101* (0.064)	0.015 (0.055)
≤elementary	-0.286* (0.150)	-0.822 (0.653)	0.288 (0.269)
≤elementary*female		0.249 (0.653)	-0.709** (0.320)
middle	0.028 (0.140)	-0.207 (0.146)	-0.037 (0.117)
middle*female		-1.193** (0.685)	-0.102 (0.191)
2 year college	0.062 (0.154)	0.429*** (0.203)	0.344** (0.206)
2 year college*female		-0.476** (0.285)	-0.388 (0.270)
≥4 year college	0.548*** (0.205)	0.570*** (0.126)	0.630*** (0.103)
≥4 year college*female		-0.099 (0.296)	-0.138 (0.227)
firmsize 100~999	0.038 (0.026)	-0.027 (0.025)	-0.028 (0.019)

〈Table 8〉 *Continued*

Dependent variable	Ln(Hourly wages)		
Empirical model	DD		DDD
Control Group	Older women	Young men	All
firmsize 100~999*female		0.142*** (0.043)	0.064** (0.030)
firm size ≥ 1000	0.056*** (0.016)	0.055*** (0.016)	0.057*** (0.013)
firm size ≥ 1000*female		-0.001 (0.028)	-0.006 (0.020)
union	0.058** (0.026)	0.041** (0.018)	0.038*** (0.015)
union*female		-0.010 (0.036)	0.015 (0.030)
seoul	0.110* (0.064)	0.080** (0.035)	0.067* (0.033)
young	0.171 (0.111)		0.122 (0.100)
female		-0.107 (0.109)	0.053 (0.157)
after	0.329* (0.189)	0.631*** (0.097)	0.418*** (0.072)
young*after	0.094*** (0.032)		0.104*** (0.023)
female*after		-0.046 (0.037)	0.031 (0.036)
young*female			-0.042 (0.160)
young*female*after			-0.011 (0.040)
inverse mills ratio 1998	0.285 (0.251)	0.161 (0.180)	0.122 (0.149)
inverse mills ratio 1999	0.131 (0.242)	-0.182 (0.187)	-0.072 (0.149)
inverse mills ratio 2000	-0.144 (0.233)	-0.328* (0.196)	-0.187 (0.141)
inverse mills ratio 2001	0.130 (0.257)	-0.029 (0.197)	0.116 (0.146)
inverse mills ratio 2002	0.127 (0.250)	-0.400** (0.189)	-0.103 (0.147)

〈Table 8〉 *Continued*

Dependent variable Empirical model Control Group	Ln(Hourly wages)		
	DD		DDD
	Older women	Young men	All
inverse mills ratio 2003	-0.276 (0.244)	-0.059 (0.191)	-0.208 (0.144)
inverse mills ratio 2004	-0.036 (0.240)	-0.533*** (0.175)	-0.428*** (0.151)
inverse mills ratio 2005	-0.124 (0.257)	-0.635*** (0.171)	-0.494*** (0.140)
# obs	3,989	5,293	10.923
R - squared	0.363	0.519	0.488

Also included in all regressions are dummy variables for years 1999 through 2004. Thus, the coefficient of *after* represents the 2005 year effect.
For each column, the p-value for testing the null hypothesis: inverse mills ratio 1998= inverse mills ratio 1999= …… =inverse mills ratio 2005=0 is 0.652, 0.001, and 0.001, respectively.
Significance levels : * 10%, ** 5%, *** 1% (two tailed test). Standard errors in parentheses.

the coefficient estimates of the inverse mills ratios found exclusively from the young women sample are neither statistically significant nor positive after the law change. Thus, it does not seem that less productive men are more likely to be employed to substitute for less productive young women, or that employers hire more productive young women after the law change in response to an increase of young female job applicants. Therefore, due to unrelated reasons to the maternity protection law change, less productive men are more easily employed during the post-change period than the pre-change period. As a result, having them as the control group overestimates the effect of the law change on young women in the fixed effects regression.

Because the pooled OLS is used in the second step of the sample selection model, the effect of time-invariant variables, such as the dummy variables for the level of education completed, young and female, can be estimated in the sample selection model. However, because these variables are also included in the first stage, and the inverse mills ratios are basically the function of first-stage control variables, they suffer from severe collinearity to have huge standard errors. On the contrary, the rest of the variables are relatively precisely estimated. Compared to the coefficient estimates obtained from the fixed effects model, while the wage gain from each year of experience is much smaller, that from each year of tenure is a bit greater. Also, the effect of marriage is almost double in the sample selection model, and working for a small size firm and a big size firm has a significant wage difference in the sample selection model.

I also estimate the DD models by the sample selection model and report the corresponding results to Table 6 in the first and the second columns of Table 8. Qualitatively, the estimated effects of the law change on hourly wages do not differ from those obtained from the fixed effects regressions.

VI. Conclusion

In November 2001, all Korean female workers were awarded to enjoy one additional month of maternity leave. Moreover, when enrolled in Employment Insurance for more than a certain amount of time, financial support is provided during this additional month and any length of childcare leave, less than 10.5 months. Using data from the Korean Labor and Income Panel Study, this paper studies how the employment and the hourly wages of young women of childbearing age adjusted to the law revison. The theory predicts that the labor demand curve for and the labor supply curve of young women move downwards. Thus, their hourly wages should decrease, but the change in their employment is uncertain because it depends on the relative magnitudes of the shifts in the curves. Empirically, I first use difference-in-differences (DD) models having either older women or young men as the control group with which to compare the changes observed for the treatment group (young women). However, this model cannot distinguish the effect of the maternity law change from the differences in employment and wage growth between the young and the old or change in gender discrimination before and after the law change. Thus, to find the effect of the law change alone, I further employ a difference-in-difference-in-differences (DDD) approach having older women, older men, and young men altogether as the control group. I estimate the DD and DDD models by ordinary least squares (OLS) and fixed effects regressions to provide insight into the role of unobservable attributes such as motivation, innate ability, and persistent family backgrounds in employment and hourly wages. Further, I estimate the DDD model by a sample selection model to incorporate the correlation between employment status and hourly wages when studying the effect of the law change on hourly wages.

In terms of the employment of young women, all fixed effects regressions show that the law revision does not affect young women's employment. Even when expected childbirth within a year and previous employment status are additionally conditioned on in the DDD model, I find that their employment is not affected by the law change. However, when studying hourly wages, the DD model presents completely different results by the comparison group. Setting older women as the comparison group, young women's wages increase substantially in both OLS and the fixed effects regression, but when young men are the comparison group, no change in their wages is reported either in OLS or in the fixed effects regression. Thus, to reconcile these conflicting findings, the DDD model is employed. Although it finds no wage increase for young women after the law change in OLS, the fixed effects regression finds that the hourly wages of young women increase, which is contradictory to the theoretical prediction. Hypothesizing that omitting variables that capture the correlation between employment status and hourly wages causes an upward bias on this estimate, I further employ a sample selection model widely applied to panel data. The estimated effect of the law change turns out to be negative in the sample selection model, although it is not statistically significant at the conventional level.

Overall, neither the employment nor the hourly wages of young women changed following the law revision. Then, the question arises why neither the labor supply nor the labor demand curve adjusts to the law change. Since Employment Insurance defrays the benefit of the third month of maternity leave and up to 10.5 months of childcare leave, which are the major costs incurred by extending maternity protection, firms' costs to hire substitution workers or to reallocate work load among remaining employees seem negligible. Also, this law change may not be able to encourage young women to find a job when they did not have one previously, or who might leave a job due to childcare to not quit.

Until 2005, when a female worker took time off due to childbirth, the pre-birth employer must have provided her with two months of usual pay. Also, she could enjoy childcare leave until her child turned one year old. However, starting from 2006, the government reduced the burden exclusively for small firms by paying all periods of maternity leave benefit from Employment Insurance. Additionally, the child's age applicable to childcare leave extended to until the child reaches three years old. This additional revision diminishes firms' costs, thus the labor demand for young women may increase. Also, more young women may be motivated to participate in the labor market and find jobs. As a result, their employment would increase, but wages would either increase or decrease depending on relative shifts in the labor supply curve of and demand curve for young women.

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abstract

모성보호법 개정과 가임기 여성의 노동시장 성과

김 인 경

정부는 2001년 산전후휴가 기간 확대와 산전후휴가와 육아휴직 기간 동안의 보조금 지급을 골자로 하는 모성보호법 개정을 단행하였다. 본 연구는 삼중차감기법을 통해 가임기 여성을 위한 이러한 추가 혜택이 가임기 여성의 고용과 시간당 임금에 미친 영향을 분석한다. 젊은 남성, 나이든 여성, 나이든 남성을 통제집단으로 간주하였을 때, 모성보호법 개정으로 인한 가임기 여성의 고용과 시간당 임금 변화는 없었다. 이는 모성보호법 개정이 가임기 여성의 노동공급과 기업의 가임기 여성에 대한 노동수요에 어떠한 변화도 초래하지 않았음을 의미한다.

주제어 : 모성보호법, 산전후휴가, 육아휴직, 가임기 여성, 고용보험