After the Shock
Reform, Resilience, and Economic Transformation in MENA
ERF 32nd Annual Conference
June 14-16 | Cairo, Egypt

2026

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The Impact of Maternity Leaves Extensions on Fertility Decisions in Iran

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ECONOMIC RESEARCH FORUM
منتدى البحوث الاقتصادية

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Abstract

In recent decades, Iran has witnessed a significant decline in its Total Fertility Rate, a trend that could pose substantial economic and social challenges in the future. In response, Iran's policymakers have introduced various economic incentives to encourage couples to have more children. One notable initiative was the extension of maternity leave from six to nine months, implemented in 2013. This research examines the impact of this extended maternity leave on fertility decisions among employed women, using data from the Labor Force Surveys (LFS) and a difference-in-difference identification methodology. The findings reveal, while the extension did not significantly affect the childbirth probability among married working women aged 20 to 35, it led to a statistically significant increase in fertility among women aged 35 to 45, who are approaching the end of their childbearing ages. Furthermore, the most pronounced effects were observed among public sector employees in this age group.

Keywords: Total Fertility Rate, Childbearing, Maternity Leave, Population policy
JEL Classification Codes: J1; J4; K2

1. Introduction

The female fertility rate is a key determinant of a society's future age structure. Since the 1990s, Iran has experienced a decline in population growth rates and female fertility rates. By the late 1990s, the fertility rate had fallen below the replacement level. If this trend continues, it will lead to an aging population, resulting in significant economic implications for the labor market, pension systems, and possibly immigration patterns. Consequently, policymakers have recently prioritized childbearing-supportive policies. One such policy involves extending maternity leave for employed women. By reducing the costs associated with childbearing, extended maternity leave can incentivize working women to have more children.

In post-revolutionary Iran, maternity leave durations were extended in 1980, 1995, 2007, 2013, and 2021. Figure 2.2 illustrates the changes in maternity leave durations and fertility rates in Iran from 1975 to 2021. The most recent substantial change occurred in 2013, when maternity leave for singleton births was increased from six months to nine months. In 2021, the Social Security Organization reported that 234,000 mothers benefited from maternity leave [Shadpoor and Mirkarimpour, 2021].

Despite the implementation of this policy, its national impact on childbearing among women in Iran has not been thoroughly examined. This study aims to address this gap by analyzing cross-sectional data from the Iranian Labor Force Surveys to evaluate the effects of the 2013 maternity leave extension on the childbearing behavior of employed women. existing regional-level studies investigating Iran's 2013 legislative increase in the length of maternity leave have been constrained by small sample sizes. This paper significantly expands the scope by utilizing the Iranian Labor Force Survey data to provide the first national-level analysis of the policy's impact on women's childbearing, disaggregated by the age group of employed women. Using the difference-in-differences Identification methodology, the study identifies eligible women as the Treatment and other working women as the Control groups. Furthermore, we leverage the differential implementation of the policy—nearly universal adherence in the public sector versus significant non-compliance in the private sector (a reflection of the Iranian government's challenges in labor law enforcement)—to examine heterogeneous treatment effects on fertility outcomes between public and private sector employees. This comparative approach offers novel insights into the moderating role of institutional capacity and sectoral enforcement on the efficacy of social policy. This research seeks to answer three core questions:

1. Has the 2013 maternity leave extension increased childbearing?
2. What are the policy's effects on childbearing among all employed women, as well as those in the private and public sectors?
3. Which age groups have been most affected by this policy?

The findings reveal that while the policy had no significant impact on childbearing among women aged 20 to 45 and 20 to 34, it did increase the childbearing rate for women aged 35 to 45 by 1.86 percentage points. Notably, the statistically significant and greatest impact was observed among Women aged 35 to 45 working in the public sector, whose childbearing increased by 2.1 percentage points following the policy's implementation.

These results align with economic theory. Given the relatively modest three-month extension of maternity leave, the policy's impact was expected to be concentrated among individuals for whom the marginal benefits and costs of having another child were closely balanced. Women aged 35 to 45, being in their last childbearing years, are more likely to grapple with the decision of whether to have an additional child to achieve their desired family size. For these women, the added three months of fully paid maternity leave, while

not altering the intrinsic benefits of childbearing, might offset some of the associated costs, encouraging eligible women to expand their families.

This paper is structured as follows: Section 1 introduces the topic. Section 2 discusses the background of the policy. Section 3 reviews the literature. Section 4 describes the data. Section 5 outlines the research methodology. Section 6 presents the empirical findings. Section 7 documents the mechanisms underlying the policy's ineffectiveness for younger women and its success for older women, and finally the Section 8 concludes.

2. Policy Implementation Background

The total fertility rate (TFR) is the main driver of the future age and sex structure of a society. The total fertility rate is the average number of children a woman will have during her reproductive years (usually from the ages of 15 to 49). A total fertility rate of 2.1 is defined as the replacement rate. This level of TFR ensures that a society will continue to maintain its population replacement rate.

Population growth in Iran has varied over the decades; the highest percentage of population growth occurred between 1976 and 1986, with an annual average of 3.91 percent. Then, population growth slowed down and reached 1.24 percent in the 2016 census. This rate is similar to the global average, but it has decreased significantly in Iran compared to previous decades. The share of the population over 30 years of age in the country's population age pyramid in recent censuses indicates that the age pyramid of the Iranian population is changing towards aging, and the share of the elderly population is increasing [National Institute of Health Research, 2019].

The cost of childbearing generally includes the financial cost of the child itself and the opportunity cost to parents of raising them, which increases household expenses and can affect mothers' labor supply and household income streams. Therefore, policies such as providing childbearing benefits or paid maternity leave can increase childbearing by reducing the cost of childbearing for the household.

The decline in fertility rates since the 1990s is attributed to changes in the country's population policies, the evolving role of Iranian women in society, and the shifts in women's bargaining power within households [K. Haddad and Kabiri-Renani, 2014]. Although Iran's total fertility rate rose slightly in 2016-reaching 2.1 children per mother-evidence suggests that fertility rates have been declining again since 2017, [Statistical

Center of Iran]. In recent years, a decline in fertility rates and low population growth have led policymakers and planners to voice concerns about the persistence of these trends. As a result, they have introduced policies to support childbearing, one of which involves extending the length of maternity leave for women.

2.1. Fertility Promotion Policies for Working Women in Iran

From 1975 to 1980 in Iran, women were entitled to a total of 84 days of maternity leave, paid at two-thirds of their average salary calculated over the previous 90 days of employment. According to the law, an insured woman who had paid insurance premiums for at least 60 days prior to giving birth could receive pregnancy benefits, provided she did not work during this leave. She also had to remain employed until the start date of her maternity or sick leave [Omran Naeimi, 2010].

Since 1980, under Article 76 of the Labor Code, women have been entitled to 90 days of maternity leave, at least 45 of which must be used after childbirth. In the case of twins, an additional 14 days of leave is granted. After their maternity leave ends, mothers may return to their previous jobs, and this leave period counts toward their service record. During the maternity leave, a mother's salary is paid in accordance with the provisions of the Social Security Law. Furthermore, in workplaces employing female workers, the employer must provide half an hour of breastfeeding time after every three hours of work until the child is two years old [Islamic Parliament Research Center, c].

In 1995, under the Law of Promoting Breastfeeding and Supporting Mothers During Breastfeeding, maternity leave was extended to four months for female workers in both the public and private sectors, covering up to three children. After returning to work, mothers of infants may, if necessary, take one hour of leave per day (without deduction from their paid leave) until the child is 20 months old, and their job security must be ensured after maternity leave and during breastfeeding [Islamic Parliament Research Center, e].

Since 2001, mothers of twins have been entitled to five months (150 days) of maternity leave, while mothers of triplets or more have been entitled to one year (365 days). In both cases, pregnancy benefits are provided throughout this period [Omran Naeimi, 2010]. In 2007, the total maternity leave was extended to six months, with an additional one hour of breastfeeding leave per day until the child reaches 24 months of age [Islamic Parliament Research Center, d].

In July 2013, during the last days of Mahmoud Ahmadinejad's presidential term, and by virtue of the powers granted to the cabinet by the Population and Family Regulation Amendment Law of 2013, the then-President issued a decree setting the maternity leave period (for single and twin births) at nine full months for women employed in both public and non-public sectors, with salaries and related bonuses paid in full. The decree also granted two weeks of compulsory (incentive) leave to the spouses of these women. However, in 1395, following the approval of the Sixth Development Plan, fathers' leave was reduced from two weeks to three days. The decree's provisions were extended to mothers whose children were not yet nine months old, allowing them to remain on leave until the baby reached nine months of age [Islamic Parliament Research Center, a].

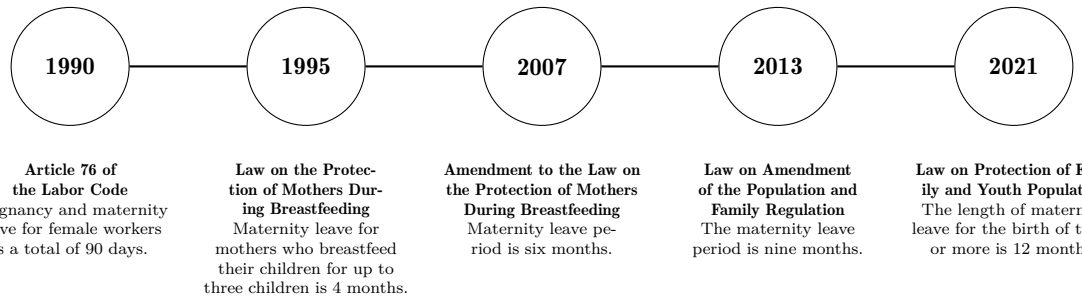

Figure 2.1: Timeline of reforms to maternity leave length in Iran during the post-Islamic Revolution period.

The 2013 ratification by the Council of Ministers was announced without specifying the necessary funding, which led to opposition from the Social Security Organization, responsible for private-sector workers. Meanwhile, women employed in public-sector jobs and covered by workplace insurance had already been eligible for nine months of maternity leave since the ratification was announced. Finally, in 2015, the General Board of the Administrative Court of Justice ruled that the Social Security Organization must also grant nine months of maternity leave [Islamic Parliament Research Center, f]. In 2021, the length of maternity leave for singleton births and for twins or more was set at nine and twelve full months, respectively, with full pay and associated benefits [Law and Regulations Portal of Islamic Republic of Iran]. Figure 2.1 illustrates the trend over time in the duration of maternity leave.

The concurrent enactment of policies aimed at encouraging population growth and increasing fertility rates in the mid-2010s underscores the importance of examining their effects. Among these policies, perhaps the most significant was the extension of maternity leave in 2013. Despite the plan's substantial costs and the debates surrounding its

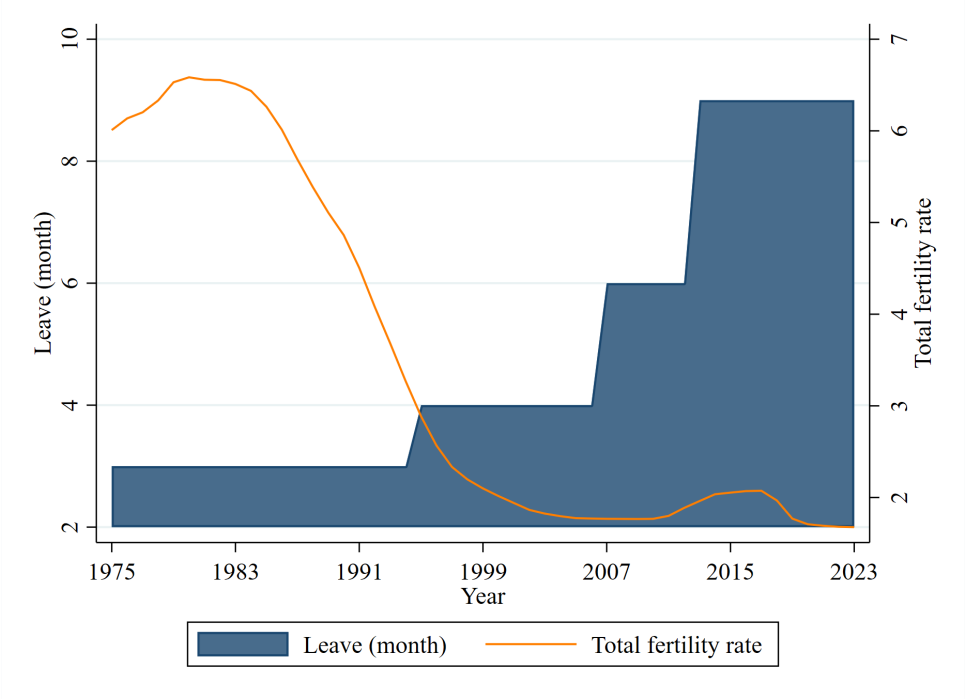

Figure 2.2: Time trend of changes in the length of maternity leave [Nations].

funding, its impact on women's childbearing—the primary goal of the policy—has not yet been investigated. Therefore, in this study, we examine the effect of extending maternity leave on women's childbearing.'

3. Literature Review

The decision to have children is shaped by various economic and social factors, with financial costs being among the most significant. In both developed and developing countries, the economic burden of childbearing—which includes direct expenses such as food, housing, and education, as well as the opportunity cost of time spent on child care—plays a crucial role in shaping fertility decisions. This section draws on Becker's economic theory of childbearing and investment to explore how financial costs affect fertility, particularly in relation to maternity leave policies. Although Becker's framework emphasizes the impact of economic constraints on family decisions, it is evident that government policies—such as maternity leave and income compensation—can alter these dynamics by reducing the financial burden of childbearing. Maternity leave policies, in particular, serve as a key mechanism for supporting working mothers and encouraging higher fertility rates. The literature on maternity leave and fertility offers a range of findings, suggesting that the success of such policies depends on various factors, including

the duration of leave, the level of income compensation, and the broader socioeconomic environment. This section reviews both theoretical perspectives and empirical studies to provide a framework for understanding how maternity leave influences fertility rates and its potential as a policy tool to address demographic challenges.

3.1. Theoretical Backgrounds

Becker's [Becker, 1960] economic theory of childbearing posits that financial costs are the primary factor influencing parents' decisions about the number of children to have. He identifies the costs of food, education, and real income as critical in shaping family decisions. Moreover, Becker suggests that the quality of children is positively correlated with the level of parental investment in a child's life, which may lead families to have fewer children in an effort to maintain higher quality. Although rising real income and declining costs should theoretically encourage childbearing, empirical evidence from developed countries challenges this expectation. Furthermore, Becker argues that childbearing rates should be higher in urban areas, where incomes are typically greater and investment costs are lower, yet evidence often shows the opposite.

In a later work, Becker [Becker, 1992] expanded his theory to include the cost of time spent on childbearing, which he views as a significant economic consideration—particularly in developed countries where women have greater opportunities to participate in the labor market. In less developed nations, where children can contribute to the family income at an earlier age, childbearing is seen more as an investment. Becker also highlights that government policies can influence the costs associated with childbearing. For example, maternity leave policies can reduce these costs by compensating for income lost during childbirth, thus encouraging higher fertility. However, extended maternity leave can also negatively affect women's labor market participation, as prolonged absences may result in lower job positions and wages upon returning, while potentially discouraging employers from hiring women of childbearing age [Schönberg and Ludsteck, 2014].

The relationship between maternity leave policies and childbearing decisions has been examined extensively in the literature. Introducing maternity leave for the first time—especially when it includes a guaranteed return to work or compensation for lost income—has been shown to positively influence women's decisions to have children by reducing the financial burden of childbearing. However, in countries where maternity leave laws are already in place, the effectiveness of extending leave durations or increasing

income compensation largely depends on the magnitude of these changes. Policies offering minimal compensation, such as a minimum wage, tend to benefit low-income women more, as they face greater financial constraints. In contrast, higher-income and more highly educated women are less likely to be influenced by such policies.

Several studies support these arguments. [Olivetti and Petrongolo, 2017] demonstrated that increasing the duration of leave and raising income replacement rates resulted in higher fertility rates, while [Averett and Whittington, 2001] found that maternity leave policies increased the likelihood of women having children by facilitating access to jobs offering leave. [Rossin-Slater et al., 2013] analyzed the impact of California's 2004 maternity leave reforms, which provided 55% salary compensation for six weeks, and found that these reforms increased mothers' weekly working hours and encouraged fertility. [Baum and Ruhm, 2016] confirmed that extended maternity leave led to greater employment retention, with mothers more likely to return to their pre-birth jobs. Additional studies, such as those by [Cannonier, 2014] and [Cygan-Rehm, 2016], provide further evidence of the positive impact of maternity leave policies on fertility, particularly in the private sector and among lower-income women.

In contrast, some studies indicate that maternity leave reforms can have more complex or even negative effects. For example, the 2007 German reforms, which introduced income-proportional compensation but shortened the duration of benefits, led to a decline in second births among low-income women [Cygan-Rehm, 2016]. However, a subsequent study by [Raute, 2019] found that these reforms resulted in increased fertility rates among upper-middle-income women, especially those in older age groups. [Malkova, 2018] examined the Soviet Union's 1981 reforms and found an 8.2% increase in fertility, particularly among rural and less-educated women. Other studies have highlighted the varying effectiveness of such policies across different contexts, such as those by [Björklund, 2006] in Sweden and [Bassford and Fisher, 2020] in Australia, which showed that while paid maternity leave did not significantly affect overall fertility, it did increase the desired number of children among highly educated women. Likewise, Quebec's maternity reforms [Ang, 2015] had a measurable impact on fertility, especially among women aged 25–45, with the 2006 reforms leading to a 23% increase in fertility. While maternity leave generally has a positive effect on fertility, the extent of its impact depends on specific policy characteristics, such as compensation levels and targeted demographics.

In addition to maternity leave, paternity leave policies have been adopted in some countries to encourage fathers' active participation in childcare, potentially reducing both the financial and time burdens on mothers. [Farré and González, 2019] investigated the effects of increased paternity leave and found that, while fathers' involvement in childcare rose, fertility rates actually decreased in the first two years, as fathers became more aware of the costs of raising children. This suggests that although maternity and paternity leave policies can shift family dynamics, they may not invariably lead to higher fertility rates in the long term. Overall, the literature indicates that maternity leave policies can affect fertility rates, but their success in promoting higher birth rates depends on factors such as compensation levels, the duration of leave, and broader social and economic contexts.

4. Data

The data used in this study come from the Labor Force Surveys (LFS), which have been conducted by the Statistical Center of Iran on a rotating basis since 2005. Each surveyed household is interviewed four times. The Households are interviewed a second time one quarter after the first interview, a third time one year after the first interview, and a fourth time one quarter after the third interview. The LFS collects labor market outcomes and demographic characteristics of household members aged 10 and older, including both Iranian citizens and non-Iranian migrants residing in Iran, typically in the middle month of each quarter. The survey employs a two-stage cluster sampling method, and approximately 25% of the sample is renewed each quarter. It provides detailed information on activity status, occupation, work experience, age, weekly hours worked, education, date of interview, and numerous other variables, all based on international definitions.

For the empirical analysis of this study, we use annual LFS data. Because the LFS records both the month and year of birth for all household members, we can identify households with newborn children under one year old. As a result, we are able to determine the employment status of mothers during the month following the child's birth. This dataset also identifies women on maternity leave, those returning to work after childbirth, and those who were dismissed upon returning to work during their maternity leave period.

A key aspect of this analysis is the difference in the enforcement of maternity leave laws between the private and public sectors after the 2013 reforms. In the public sector,

these policies are generally implemented according to the law. However, in the private sector, the Social Security Organization and employers have not consistently adhered to these regulations [Razeghi Nasrabad and bastani, 2023]. The LFS distinguishes between private-sector and public-sector employment. Table 4.1 presents a statistical summary of employed women aged 20 to 45 in the sample from 2010 to 2017.

Each year, the LFS gathers information from over 100,000 households. While individuals are surveyed for two consecutive quarters, the weights assigned to each individual differ across survey rounds. Each observation in a given quarter is treated as representative of a specific number of citizens, independent from the previous quarter. Therefore, in this study, each observation is considered as a separate individual.

Table 4.1: Descriptive Statistics of Employed Women (Aged 20-45), 2010-2017

		Averages						
Year	**N**	**Age**	**Univ. (%)**	**Married (%)**	**Urban (%)**	**With Kids (%)**	**Insured (%)**	**Public (%)**
2010	20,156	32.598	28.379	66.407	47.614	56.658	28.681	24.400
Std. Dev.	–	7.024	45.1	47.2	49.9	46.0	45.2	43.0
2011	18,393	32.950	31.191	66.041	49.568	55.956	32.295	25.825
Std. Dev.	–	6.916	46.3	47.4	50.0	46.8	46.8	43.8
2012	18,259	33.194	30.971	65.365	49.592	55.458	32.209	24.596
Std. Dev.	–	6.903	46.2	47.6	50.0	47.4	46.7	43.1
2013	22,101	33.177	37.003	68.603	55.952	58.893	37.329	26.745
Std. Dev.	–	6.795	48.3	46.4	49.6	47.2	48.4	44.3
2014	21,216	33.669	39.739	68.392	57.848	58.984	39.659	27.757
Std. Dev.	–	6.699	48.9	46.5	49.4	47.5	48.9	44.8
2015	23,283	33.950	39.449	68.088	57.359	58.588	37.792	25.753
Std. Dev.	–	6.642	48.9	46.6	49.5	47.7	48.5	43.7
2016	24,441	34.138	38.767	67.841	57.187	58.038	36.148	24.209
Std. Dev.	–	6.598	48.7	46.7	49.5	47.9	48.0	42.8
2017	25,993	34.181	38.730	68.468	56.596	59.127	35.206	23.114
Std. Dev.	–	6.546	48.7	46.5	49.6	47.6	47.8	42.2

Note: This statistical summary includes all employed women aged 20 to 45 for the years 2010 to 2017. N = Number of observations.

5. Identification Strategies

This study focuses on child bearing decision of married working women aged 20 to 45 during the period from 2010 to 2017. This time-frame is chosen because the last

amendment to maternity leave duration before 2013 occurred in 2007, and 2008–2009 are considered transitional years that may have been influenced by earlier policy changes. Furthermore, since 2018, the unilateral withdrawal of the United States from the nuclear agreement and the imposition of extensive sanctions on Iran have substantially altered Iran's macroeconomic conditions and labor market dynamics, making it more difficult to isolate the causal effects of the policy on fertility.

According to the law, women eligible for maternity leave must meet the following criteria:

1. They must have worked for at least 60 full days in the year prior to the maternity leave and be insured through their job.
2. They must not be working during the maternity leave.
3. They must be employed up to the day maternity leave or sick leave begins.

This study employs the difference-in-differences (DiD) as an identification method. The treatment group comprises married employed women eligible for maternity leave, while the control group includes married employed women who are ineligible for maternity leave. The main regression model is specified as follows:

$$y_{it} = \beta_0 + \beta_1 \sigma_t + \beta_2 \text{eligible}_i + \beta_3 (\text{eligible}_i \times \text{post}_t) + \beta_4 X_{it} + \varepsilon_{it} \tag{1}$$

Where:

- y_{it}: a dummy variable equal to 1 if the woman has a child under one year old and 0 otherwise
- post: a dummy variable equal to 0 before the 2013 reforms and 1 afterward
- eligible: a dummy variable equal to 1 for women eligible for maternity leave and 0 otherwise
- σ_t: year fixed effects
- X_{it}: a vector of control variables, including individual and workplace characteristics
- ε_{it}: the error term

The sample includes married, employed women aged 20 to 45. Control variables in the regressions include individual and occupational characteristics, such as age, College-educated, occupation, workplace characteristics, number of children, urban or rural location, and provincial fixed effects. Controlling for provincial and urban/rural fixed effects

is crucial due to cultural and economic differences across regions that influence childbearing decisions and costs.

Age plays a pivotal role in fertility decisions. Younger working women often prioritize career progression, higher education, and job stability, which can conflict with childbearing. Conversely, older women face biological constraints. The inclusion of age as a control variable addresses these dynamics.

Occupational factors, including employer cooperation and job type, significantly impact the costs of childbearing. For example, employers in the education sector may provide more support for maternity leave than those in other sectors since in Iran, in the primary education sector and in girls' schools, almost all employees are women. Similarly, replacing a skilled worker is often more challenging than replacing a general laborer, highlighting the importance of occupational controls.

Education is another critical variable. Women with higher education levels tend to prioritize child quality over quantity and are more likely to occupy positions that are harder to replace. Hence, College-educated is included as a dummy variable in the control set.

Since our data are repeated cross-sections, it is crucial to confirm that the conditions necessary for appropriately applying this dataset in a difference-in-differences estimation are met. We find no evidence to suggest that group consistency over time is invalid. The outcomes of these assessments are detailed in Appendix 1.

A potential threat to the internal validity of this study is self-selection, which may occur when women who plan to have children specifically seek employment positions that offer insurance coverage, thereby granting access to maternity leave benefits. If unaddressed, this issue might bias the estimated effect of the policy. However, because the eligibility criteria for maternity leave in Iran remained unchanged during the 2013 reform—which extended leave from six to nine months—any pre-existing self-selection patterns should remain consistent. To verify that these patterns did not change after the reform, we compare eligible women employed for fewer than 12 or 24 months with eligible women employed for longer durations.

The self-selection hypothesis is tested using the following difference-in-differences estimation:

$$\text{leave}_{it} = \alpha_0 + \alpha_1\sigma_t + \alpha_2\text{treatment}_i + \alpha_3(\text{treatment}_i \times \text{post}_t) + \alpha_4 X_{it} + v_{it} \quad (2)$$

Where:

- leave: a dummy variable equal to 1 if the woman took maternity leave and 0 otherwise
- treatment: a dummy variable equal to 1 for women employed for less than a specified period (e.g., 12 or 24 months) and 0 otherwise
- Other variables are defined similarly to the first specification
- v_{it}: the error term

By examining self-selection among women who have been employed for shorter durations, we assess whether patterns of maternity leave uptake changed significantly after the reform.

Another potential source of bias is the growing gender gap in access to jobs with insurance for women. Employers who provide insurance may be reluctant to hire women—particularly those who could have children in the future—because of the additional costs associated with prolonged maternity leave. This reluctance could skew the estimated impact of the policy on women's childbearing decisions. To investigate whether the policy impacted the gender gap in access to insured jobs, we restrict our sample to married working men and women aged 20 to 45 and use specification (3) for the difference-in-differences estimation.

$$y_{it} = \gamma_0 + \gamma_1\sigma_t + \gamma_2\text{treatment}_i + \gamma_3(\text{treatment}_i \times \text{post}_t) + \gamma_4 X_{it} + e_{it} \quad (3)$$

Where:

- y_{it}: a binary variable equal to 1 for insured individuals and 0 for uninsured individuals
- treatment_i: a dummy variable equal to 1 for married working women aged 20 to 45 (treatment group) and 0 for married working men aged 20 to 45 (control group)
- post_t: a dummy variable equal to 1 for the post-policy period and 0 for the pre-policy period
- X_{it}: a vector of control variables, including individual and workplace characteristics
- e_{it}: the random error term

The first attempt to extend the length of maternity leave was initiated by the Health Commission of Parliament in 2012. In March 2012, Parliament approved the amendment to the Population and Family Regulation Law Islamic Parliament Research Center [b]. According to Parliament's resolution, the government was mandated to increase the length of maternity leave from 6 to 10 months. However, the Guardian Council rejected the parliamentary bill due to unspecified funding resources required for the plan and deemed the resolution contrary to Article 75 of the Constitution. Since the government supported the paramilitary bill, a compromise was reached between the parliament and the government. According to which, the phrase *"maternity leave for mothers shall be increased to nine months"* in the bill was revised to *"the government shall be allowed to increase maternity leave for mothers to nine months"*. The revised plan was re-approved by parliament in June 2013 and was sent to the Guardian Council. Shortly after the Guardian Council's approval, the Cabinet of Ministers, in July 2013, citing Note (2) of the Single Article of the Law Amending the Population and Family Regulation Laws, approved and implemented the extension of maternity leave to nine months.

Since over 2012 and 2013, there was strong support for the policy from both the government and parliament, driven by growing concerns over declining fertility rates. It can be argued that from March 2012 onward, the implementation of the extended maternity leave policy was virtually certain, as the Iranian state was fully committed to enacting the law. Although the policy was officially implemented in July 2013, classifying 2013 as either a pre- or post-policy year introduces ambiguities due to its transitional nature. Consequently, the year 2013 has been excluded from our analysis of the policy's effects.

6. Empirical Findings

This section presents our findings on the impact of the 2013 maternity leave extension on fertility rates among employed women in Iran. Using a difference-in-differences approach, we compare the childbearing outcomes of women who were eligible for maternity leave benefits to those who were not. The analysis highlights key demographic subgroups—including age cohorts and employment sectors—to reveal variations in the policy's effects. We also perform robustness checks to validate the results and explore potential mechanisms underlying the observed outcomes.

6.1. The parallel trends condition

In the first step, the validity of the parallel trends assumption is checked to verify the internal validity of our findings. Since the plan was not approved until the end of 2012 and could not have affected childbearing prior to this, specification (1) is employed to examine the primary assumption. Initially, 2012 is treated as the pre-implementation period and 2013 as the post-implementation period. Subsequently, 2010 and 2011 are defined as pre-implementation periods, while 2012 is considered post-implementation period for conducting a placebo test.

The difference-in-difference estimated coefficients ($\hat{\beta}_3$) are reported in Table 6.1 for all women, women employed in the private sector, and women employed in the public sector. As shown, these estimated values are statistically insignificant, indicating that the parallel trends assumption cannot be rejected.

Table 6.1: Examination of the condition of parallel trends in the period before the increase in the length of maternity leave in 2013

Age (years)	**(1) 20 to 45**	**(2) 20 to 34**	**(3) 35 to 45**
All			
2010 before / 2011 and 2012 after	0.0118 (0.00885)	0.0159 (0.0162)	0.00993 (0.0122)
2010 and 2011 before / 2012 after	0.00982 (0.00615)	0.0156 (0.0109)	0.00893 (0.0113)
2011 before / 2012 after	0.00805 (0.00700)	0.00789 (0.0101)	0.00765 (0.0133)
Private sector			
2010 before / 2011 and 2012 after	0.000777 (0.0160)	0.0133 (0.0256)	0.00640 (0.0156)
2010 and 2011 before / 2012 after	0.00173 (0.0111)	0.00883 (0.0205)	0.00610 (0.00822)
2011 before / 2012 after	0.00132 (0.0158)	-0.00638 (0.0256)	0.0201 (0.0132)
Public sector			
2010 before / 2011 and 2012 after	0.0119 (0.00855)	0.0208 (0.0157)	0.00865 (0.0105)
2010 and 2011 before / 2012 after	0.0119 (0.00949)	0.0172 (0.0162)	0.00988 (0.0121)
2011 before / 2012 after	0.00870 (0.0111)	0.00935 (0.0171)	0.00669 (0.0138)

Note: The difference-in-difference estimate is calculated based on specification (1) and for the years 2010 to 2012. Column (1) reports the estimate for women aged 20 to 45, column (2) the estimate for women aged 20 to 34, and column (3) the estimate for women aged 35 to 45. Standard errors are presented in parentheses. Controls include age, university education, occupation, workplace characteristics, number of children under 18, urban or rural residence, and fixed effects of year and province. * indicates significance at the 90% confidence interval, ** at the 95% confidence interval, and *** at the 99% confidence interval. Standard errors are clustered at the province level.

6.2. Possibility of Self-Selection Presence

According to specification (2), the presence of self-selection cannot be ruled out if the difference-in-difference coefficient ($\hat{\alpha_3}$) is significant. A positive coefficient indicates an increase in self-selection following the policy's implementation, whereas a negative

coefficient signifies a decrease. The policy may incentivize women who are planning to give birth to seek insured employment, which could potentially increase self-selection. On the other hand, stricter hiring practices adopted by employers in response to the policy might reduce self-selection.

Table 6.2 presents the results for the $\hat{\alpha_3}$, reflecting possible presence of self-selection for women employed for less than one or two years before and after the leave plan's implementation. This coefficient is insignificant across all groups, except for public sector employees aged 20–45. Among married women in this category, significant and negative self-selection is observed at the 90% confidence interval. Therefore, there is no strong significant evidence for self-selection among women aged 20 to 45 working in the public sector, however it is significant at 10% critical level. Additionally, the presence of self-selection in other groups is not statistically significant.

Table 6.2: DID Estimates of Self-Selection Bias: Evidence from 1- and 2-Year Employment Thresholds

	(1) 20 to 45	**(2) 20 to 34**	**(3) 35 to 45**
All			
12 months	-0.0153 (0.0140)	-0.00909 (0.0178)	-0.0197 (0.0199)
24 months	-0.0147 (0.00936)	-0.0117 (0.0176)	-0.0196 (0.0152)
Observations	32,675	12,670	20,005
Private sector			
12 months	0.00698 (0.0121)	0.0135 (0.0121)	-0.00936 (0.00776)
24 months	-0.00417 (0.0178)	-0.00159 (0.0242)	-0.0115 (0.00722)
Observations	7,793	4,027	3,766
Public sector			
12 months	-0.0435* (0.0233)	-0.0439 (0.0361)	-0.0164 (0.0275)
24 months	-0.0308* (0.0152)	-0.0335 (0.0240)	-0.0122 (0.0280)
Observations	24,810	8,593	16,217

Note: The difference-in-difference estimate is calculated based on specification (2) and for the years 2010 to 2017. Column (1) reports the estimate for women aged 20 to 45, column (2) the estimate for women aged 20 to 34, and column (3) the estimate for women aged 35 to 45.Standard errors are presented in parentheses. Controls include age, university education, occupation, workplace characteristics, number of children under 18, urban or rural residence, and fixed effects of year and province. * indicates significance at the 90 percent confidence interval, ** indicates significance at the 95 percent confidence interval, and *** indicates significance at the 99 percent confidence interval. The standard errors are clustered at the province level.

6.3. Gender Gap in Access to Insured Jobs

Specification (3) examines the change in the gender gap in accessing jobs with insurance benefit ($\hat{\gamma_3}$) post-implementation compared to pre-implementation. An insignificant estimated value implies that no change in the gender gap can be confirmed. Table 6.3 reports the results for $\hat{\gamma_3}$. It clearly confirms no significant change in the gender gap for access to jobs with insurance benefits following the plan's implementation, indicating that the extension of maternity leave by three months has not had a significant impact.

Table 6.3: Changing gender gap in access to insured jobs (2010 to 2017)

Age (years)	**(1) 20 to 45**	**(2) 20 to 34**	**(3) 35 to 45**
All			
$Treatment \times Post$	-0.0279 (0.0170)	-0.0281 (0.0192)	-0.0295 (0.0176)
Observations	579,900	256,152	323,748
R^2	0.377	0.349	0.395
Private sector			
$Treatment \times Post$	0.00239 (0.0203)	-0.00510 (0.0245)	0.0205 (0.0188)
Observations	468,838	218,268	250,570
R^2	0.248	0.247	0.251
Public sector			
$Treatment \times Post$	0.00341 (0.0109)	0.0310 (0.0201)	-0.0140 (0.0124)
Observations	111,062	37,884	73,178
R^2	0.029	0.038	0.021

Note: The difference-in-difference estimate is calculated based on specification (3) and for the years 2010 to 2017. Column (1) reports the estimate for women aged 20 to 45, column (2) the estimate for women aged 20 to 34, and column (3) the estimate for women aged 35 to 45.Standard errors are presented in parentheses. Controls include age, university education, occupation, workplace characteristics, number of children under 18, urban or rural residence, and fixed effects of year and province. * indicates significance at the 90% confidence level, ** at the 95% confidence level, and *** at the 99% confidence level. Standard errors are clustered at the province level.

6.4. The Effect of Maternity Leave on Childbearing

To examine the effect of maternity leave on women's childbearing, specification (1) is used to conduct a difference-in-difference estimation. The results for the coefficient (β_3) are reported in Table 6.4. As shown in this table, the effect of the policy on women aged 20 to 45 is positive but insignificant across all groups: all employed women, women employed in the public sector, and women employed in the private sector.

When examining age groups, it is observed that for women aged 20 to 34, the coefficients are insignificant in all categories. Specifically, for this age group, the coefficient is

Table 6.4: Results of the difference-in-difference study of the effect of increasing maternity leave on women's childbearing (2010 to 2017)

Age (years)	**(1) 20 to 45**		**(2) 20 to 34**		**(3) 35 to 45**	
	With control	**Without control**	**With control**	**Without control**	**With control**	**Without control**
All						
eligible × *Post*	0.00426	0.00442	-0.00109	-0.000318	0.0186***	0.0106
	(0.00826)	(0.00741)	(0.0142)	(0.0130)	(0.00646)	(0.00721)
Observations	68,267	68,271	32,027	32,028	36,240	36,243
R^2	0.052	0.002	0.065	0.003	0.042	0.003
Private sector						
eligible × *Post*	0.00759	0.00800	0.00484	0.00714	0.0164***	0.0143**
	(0.0105)	(0.0104)	(0.0153)	(0.0148)	(0.00550)	(0.00633)
Observations	43,000	43,003	22,977	22,978	20,023	20,025
R^2	0.051	0.001	0.059	0.001	0.045	0.001
Public sector						
eligible × *Post*	0.00208	0.00476	-0.00584	-0.00235	0.0210**	0.0120
	(0.00831)	(0.00780)	(0.0154)	(0.0144)	(0.00905)	(0.00814)
Observations	60,050	60,052	27,598	27,599	32,452	32,453
R^2	0.053	0.002	0.067	0.006	0.044	0.004

Note: The difference-in-difference estimate is calculated based on specification (1) and for the years 2010 to 2017. Column (1) reports the estimate for women aged 20 to 45, column (2) the estimate for women aged 20 to 34, and column (3) the estimate for women aged 35 to 45. Controls include age, university education, occupation, workplace characteristics, number of children under 18, urban or rural residence, and fixed effects of year and province. * indicates significance at the 90% confidence interval, ** at the 95% confidence interval, and *** at the 99% confidence interval. Standard errors are clustered at the province level.

positive but insignificant for women employed in the private sector, while it is negative and close to zero for all employed women and those in the public sector.

For women aged 35 to 45, the policy's effect is positive and significant across all groups. The largest effect is observed among women employed in the public sector, where the increase in maternity leave resulted in a 2.1 percentage point rise in childbearing, significant at the 95% confidence interval. The smallest effect is found among women in the private sector, where the plan led to a 1.64 percentage point increase in childbearing, significant at the 99% confidence interval. The overall effect of the policy for all women aged 35 to 45 is a 1.86 percentage point increase in childbearing, which is significant at the 99% confidence interval.

These findings suggest that the policy primarily impacted women nearing the end of their childbearing years, encouraging them to have more children. This result aligns with Becker's theory, which posits that households respond positively to incentives that reduce the cost of childbearing. Since a 3-month increase in maternity leave does not substantially lower the overall cost of childbearing, it likely encourages women for whom the costs and benefits of having children are nearly equal, thereby resolving their indecision. In such situations, even small incentives, like extending leave by 3 months, can

reduce the perceived costs of childbearing without altering its benefits, tipping the scale in favor of having more children. Women nearing the end of their childbearing years are often faced with this cost-benefit dilemma, making them particularly responsive to such incentives.

6.5. Robustness Check

To assess the robustness of the results, re-estimations were conducted using specification (1), excluding 2010 as the initial pre-implementation year and 2017 as the final post-implementation year. The outcomes are detailed in Table 6.5. Additionally, as the dependent variable is discrete, specification (1) was re-estimated using the probit method. The probit estimation results, presented in the third column of Table 6.5, closely mirror those obtained via the ordinary least squares (OLS) method, demonstrating the stability of the findings.

Table 6.5: Robustness of Results for Full-Time Married Workers Aged 20 to 45

Sector	**Age (years)**	**(1) 2010 to 2016 (OLS)**	**(2) 2011 to 2017 (OLS)**	**(3) 2010 to 2017 Marginal effect (Probit)**	**(4) 2010 to 2017 (OLS)**
All	20 to 45	0.00279 (0.00871)	0.00197 (0.0111)	0.00531 (0.00794)	0.00426 (0.00826)
	20 to 34	-0.000920 (0.0153)	-0.00663 (0.0180)	-0.00219 (0.0126)	-0.00109 (0.0142)
	35 to 45	0.0147** (0.00685)	0.0160*** (0.00578)	0.0184*** (0.00704)	0.0186*** (0.00646)
Private	20 to 45	0.00892 (0.0106)	0.00719 (0.0136)	0.00851 (0.0115)	0.00759 (0.0105)
	20 to 34	0.00791 (0.0168)	-0.000502 (0.0188)	0.00473 (0.0165)	0.00484 (0.0153)
	35 to 45	0.0149** (0.00566)	0.0210*** (0.00590)	0.0168* (0.00895)	0.0164*** (0.00550)
Public	20 to 45	-0.00256 (0.00847)	-0.00241 (0.0107)	0.00189 (0.00621)	0.00208 (0.00831)
	20 to 34	-0.00872 (0.0166)	-0.0129 (0.0189)	-0.00948 (0.0102)	-0.00584 (0.0154)
	35 to 45	0.0126 (0.00914)	0.0172* (0.00854)	0.0190** (0.00951)	0.0210** (0.00905)

Note: The difference-in-difference estimate is calculated based on specification (1). Column (1) shows the estimation results for the years 2010 to 2016, column (2) shows the estimation results for the years 2011 to 2017, column (3) shows the final effect estimation results using the Probit method for the years 2010 to 2017, and column (4) shows the estimation results for the years 2010 to 2017. Controls include age, university education, occupation, workplace characteristics, number of children under 18, urban or rural location, and fixed effects of year and province. * indicates significance at the 90% confidence interval, ** at the 95% confidence interval, and *** at the 99% confidence interval.

Further robustness checks included re-estimating specification (1) with additional controls, such as work experience, the square of age, the square of the number of children, and replacing university education with a more granular education variable (10 levels). Another estimation excluded individuals aged 20 to 25. These results are summarized in Tables 9.1 through 9.4 in Appendix 2.

7. Potential Mechanisms for the Heterogenous Effects

The divergent impacts of Iran's maternity leave extension can be understood through a life-course and institutional lens, revealing why the policy was ineffective for younger women but successful for an older, specific cohort. For younger working women (aged 20-35), the policy's insignificant effect aligns with theories of fertility postponement Lutz and Skirbekk [2011] and stems from its failure to overcome dominant constraints like high opportunity costs for career-building and anxieties about long-term economic burdens Ebadi [2022]. A narrowly-focused incentive like three additional months of leave is a superficial solution that ignores wider challenges such as job market instability McDonald [2000]. Consequently, for women with a wide biological window, the policy was an insufficient nudge against the powerful tendency to postpone childbirth, failing to shift fertility intentions on the *extensive margin*. In stark contrast, the policy was highly significant for older women (35-45) due to a powerful convergence of biological and institutional factors. The "now or never" pressure of the biological clock meant the extended leave acted as a decisive, last-chance incentive, effectively lowering the immediate cost of a final pregnancy and primarily influencing the *intensive margin* by accelerating the timing of already-planned births Ebadi [2022]. This is consistent with findings that "catch-up" fertility is highly responsive to timely policy interventions Kearney et al. [2022]. Furthermore, this effect was most pronounced in the public sector because, as Ebadi highlights, the institutional context is critical; these jobs provide the job security and benefit enforcement that make the policy a guaranteed, low-risk benefit Pettit and Hook [2005]. Therefore, the policy did not broadly raise fertility but simply accelerated births for a privileged subgroup of older, state-employed women. This analysis underscores that the success of family policies is highly dependent on the specific life-course stage and employment context of the target population. The Iranian case demonstrates that a policy appearing "good from far" can be "far from good" for most of the population if it does not address the foundational economic and institutional barriers that different groups of women face Ebadi [2022].

8. Conclusion

The fertility rate in Iran has declined significantly over the past two decades, prompting policymakers to introduce incentives aimed at increasing fertility due to the severe consequences of continued low fertility for Iranian households. These incentives are designed to raise household fertility rates by reducing the costs associated with having children. In line with this goal, the Iranian government extended maternity leave from 6 months to 9 months in 2013.

Following the Cabinet's resolution to increase maternity leave to 9 months, the public sector swiftly implemented the policy. However, there was resistance in the private sector, where full implementation did not occur until 2015, following the Administrative Court of Justice's decision mandating compliance. This study utilized the difference-in-difference method and Labor Force Survey from the Statistical Center of Iran to examine the policy's impact on women's childbearing.

The results indicate that the policy primarily affected women who were nearing the end of their childbearing years. Specifically, for women aged 35 to 45, the policy led to an increase in childbearing of approximately 1.8 percentage points. Among other age groups, the policy showed no significant effect on childbearing when compared to the pre-implementation period. Furthermore, the largest effect was observed among women employed in the public sector, where the policy was fully implemented. In contrast, the smallest effect was seen among women working in the private sector, where implementation faced resistance from employers and social security systems. Additionally, we found that extending maternity leave by three months does not significantly impact the gender gap in the jobs with insurance coverage.

These findings align with Becker's theory, which suggests that households respond positively to incentives that reduce the costs of childbearing. Given that a 3-month increase in maternity leave is unlikely to substantially reduce the overall costs of childbearing, the policy appears to have influenced women for whom the costs and benefits of childbearing were closely balanced. For such women, even small incentives, like the extended leave, could tip the scales, making the benefits outweigh the costs. Women nearing the end of their childbearing years often face the dilemma of whether having another child will yield greater utility, making them particularly responsive to minimal incentives.

The findings also suggest that increasing maternity leave by 3 months has not significantly impacted childbearing among all working women aged 20 to 45. To encourage higher fertility rates more broadly, the government may need to adopt more generous and comprehensive policies that substantially lower the costs of childbearing for women. These findings align with Becker's theory that marginal reductions in childbearing costs can influence fertility decisions when benefits and costs are closely balanced. The policy's limited aggregate impact suggests that broader incentives may be needed to address Iran's demographic challenges. To maximize the efficacy of the maternity leave policy in encouraging childbearing, our findings suggest that the government should pursue three integrated policy adjustments. First, the generosity of benefits should be enhanced—either by extending the duration of paid leave or by providing direct financial assistance—to strengthen the overall incentive for childbearing. Second, the observed heterogeneity in treatment effects necessitates strengthening institutional capacity, particularly by imposing stricter enforcement mechanisms and penalties for non-compliance in the private sector. Such measures are critical to ensuring that the policy's benefits are realized economy-wide. Finally, to address the muted fertility response among younger women—who are more sensitive to career disruption-legislation must go beyond merely guaranteeing a return to the same job and explicitly ensure a return to the exact pre-leave position and status. This amendment is critical because, according to some studies [Zahra Khajeh, 2024], more than one-fifth of women have been forced to continue working in a new or different position after returning, suggesting that the current law fails to adequately address the high perceived opportunity costs that discourage younger, career-oriented women from having children.

Declaration: authors used AI

During the preparation of this work, the authors used ChatGPT in order to improve grammatical precision. After using this service, the authors reviewed and edited the content as needed and take full responsibility for the content of the publication.

Declaration: Conflict of Interest Statement for Manuscript Submission, Funding and Data

1. We confirm that we are residing in Islamic Republic of Iran, the article has been prepared in a personal, academic, or research capacity and not as an official representative or otherwise on behalf of the relevant government.
2. We are writing to disclose any potential conflicts of interest related to the manuscript
3. Our data will be available upon request.
4. There is no funding for this research.

Author Contribution Form

Term	**GholamReza Keshavarz Haddad**	**Mehdi Djalali**
Conceptualization	✓	✓
Methodology / Study design	✓	✓
Software	✓	✓
Validation	✓	✓
Formal analysis	✓	✓
Investigation	✓	✓
Resources	✓	✓
Data curation	–	✓
Writing – original draft	–	✓
Writing – review and editing	✓	–
Visualization	✓	✓
Supervision	✓	–
Project administration	✓	–
Funding acquisition	–	–

(✓ indicates the author who fulfilled the role.)

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Appendix 1

To use cross-sectional data in difference-in-differences (DiD) estimation, three conditions must be satisfied [Botosaru and Gutierrez, 2018], [Sant'Anna and Xu, 2023], [Lee and Kang, 2006]:

1. **Randomization and independence**: The data must be random, and individuals must be assigned to the treatment or control group independently of the characteristics of the control variables or the dependent variable. Furthermore, the assignment of individuals must be independent of time (i.e., before and after the policy implementation).
2. **Group consistency over time**: The characteristics of individuals in the treatment group should remain consistent over time. In other words, individuals in the pre- and post-policy periods should share similar, if not identical, characteristics. This condition applies equally to the control group.
3. **Homogeneous time effects**: The effect of time must be uniform across the treatment and control groups. No external factor, aside from the policy under study, should differentially impact one group relative to the other over time.

In this study, the data is provided by the Statistical Center of Iran and assumed to be random, thereby satisfying the first condition. Furthermore, the parallel trends assumption, as demonstrated in section 6, supports the validity of the third condition.

To satisfy the second condition, individuals in the trearment group must exhibit similar characteristics in both the pre- and post-policy periods, with the same requirement holding for the control group. This necessitates that the sample be independent across all years and drawn from a common distributions. To evaluate fulfillment of the condition, we first conduct a regression analysis using the Probit method, the Propensity Score Matching (PSM). The dependent variable, *eligible*, distinguishes individuals between the trearment and control groups and is regressed on the control variables used in the main DiD analysis.

Subsequently, we calculate the average probability (Propensity Scores) of individuals belonging to the policy group ($D = 1$) and the control group ($D = 0$) during the pre-policy ($T = 0$) and post-policy ($T = 1$) periods. Additionally, these probabilities are calculated without differentiating between pre- and post-policy periods. The results of these calculations are presented in Tables 8.1 to 8.3.

Table 8.1: Probability of being in the Treatment group for all married working women aged 20 to 45

	Mean	Standard Deviation	Minimum	Maximum	Number of Observations
$P(D = 1)$	0.783	0.235	0.002	0.987	35,165
$P(D = 1, T = 0)$	0.802	0.199	0.002	0.986	11,754
$P(D = 1, T = 1)$	0.773	0.251	0.002	0.987	23,411

Table 8.2: Probability of being in the Treatment group for married women working in the private sector aged 20 to 45

	Mean	Standard Deviation	Minimum	Maximum	Number of Observations
$P(D = 1)$	0.394	0.234	0.004	0.947	8,506
$P(D = 1, T = 0)$	0.414	0.233	0.005	0.917	1,980
$P(D = 1, T = 1)$	0.387	0.234	0.004	0.947	6,526

Table 8.3: Probability of being in the Treatment group for married women working in the public sector aged 20 to 45

	Mean	Standard Deviation	Minimum	Maximum	Number of Observations
$P(D = 1)$	0.839	0.164	0.000	0.999	26,659
$P(D = 1, T = 0)$	0.827	0.163	0.000	0.995	9,774
$P(D = 1, T = 1)$	0.846	0.165	0.000	0.999	16,885

In the next step, the matching method was used to evaluate the matching quality for individuals in the policy group during the periods before and after the implementation of the plan. Separate estimates were made for the treatment group and the control group. The dependent variable was defined as having a baby under one year old, while the treatment variable was the dummy variable *Post* (equal to zero for the period before 2013 and one for the period after 2013). The control variables included all the main regression variables except for the industry category.

The results for the treatment group are presented in Figures 8.1, 8.3 and 8.5, and those for the control group are shown in Figures 8.2, 8.4 and 8.6. As demonstrated by the figures, matching quality is consistently high within each group, before and after the plan's rollout. This indicates that individuals with similar characteristics were consistently observed over time across the groups.

Considering the results of the probability calculations and the evaluation of matching quality, there is no evidence to reject the validity of the second condition.

For the third condition to be met, the assumption of parallel trends must hold, as discussed in this article and shown in Table 6.1.

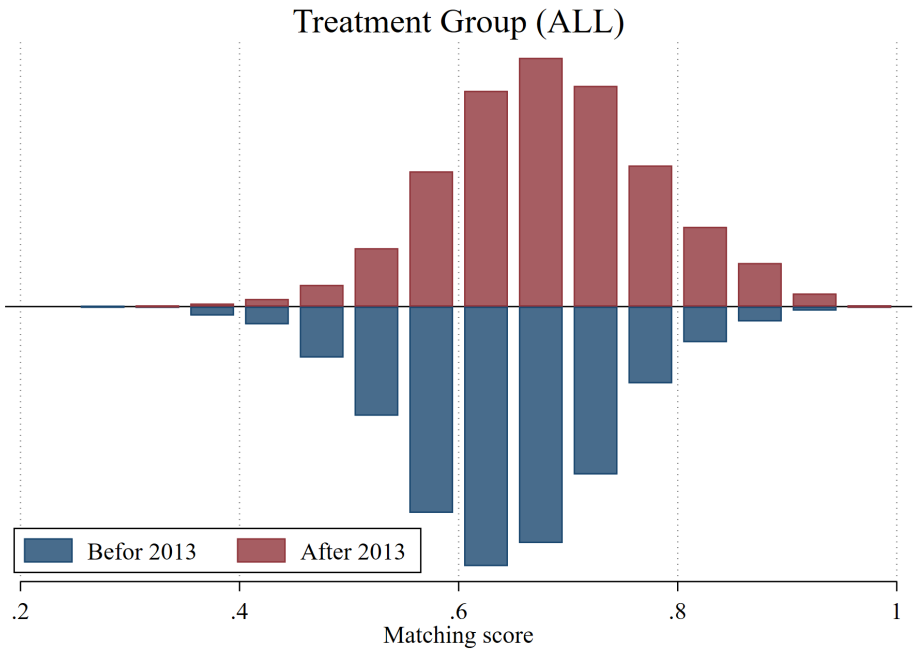

Figure 8.1: Treatment group matching score for all employed women.

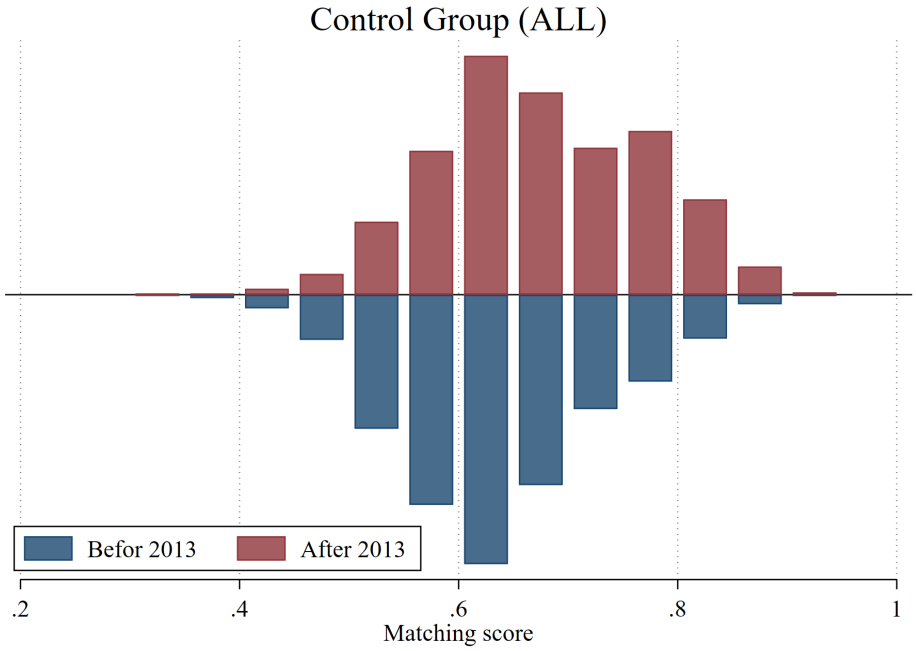

Figure 8.2: Control group matching score for all employed women.

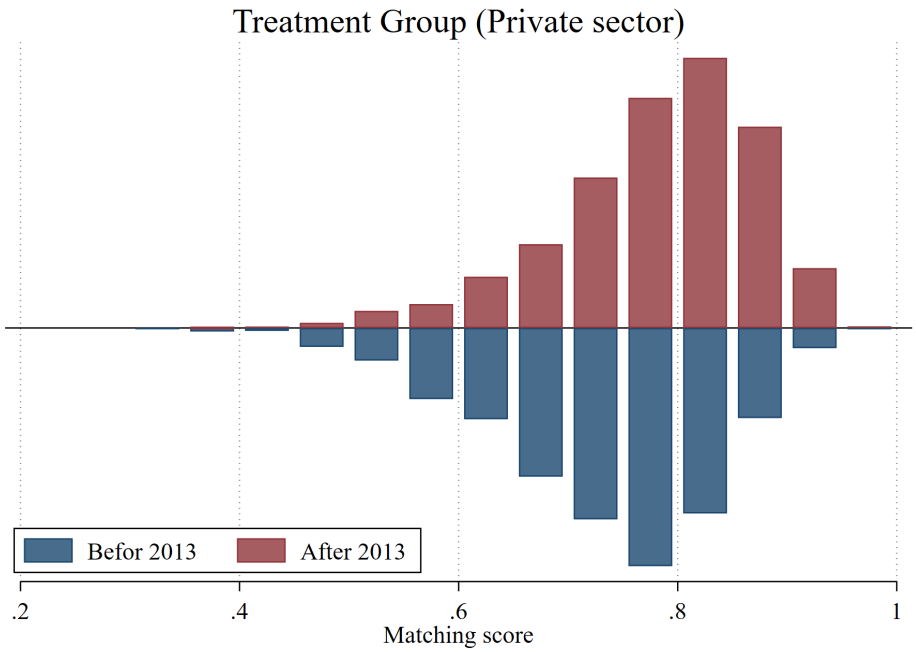

Figure 8.3: Treatment group matching score for women working in the private sector.

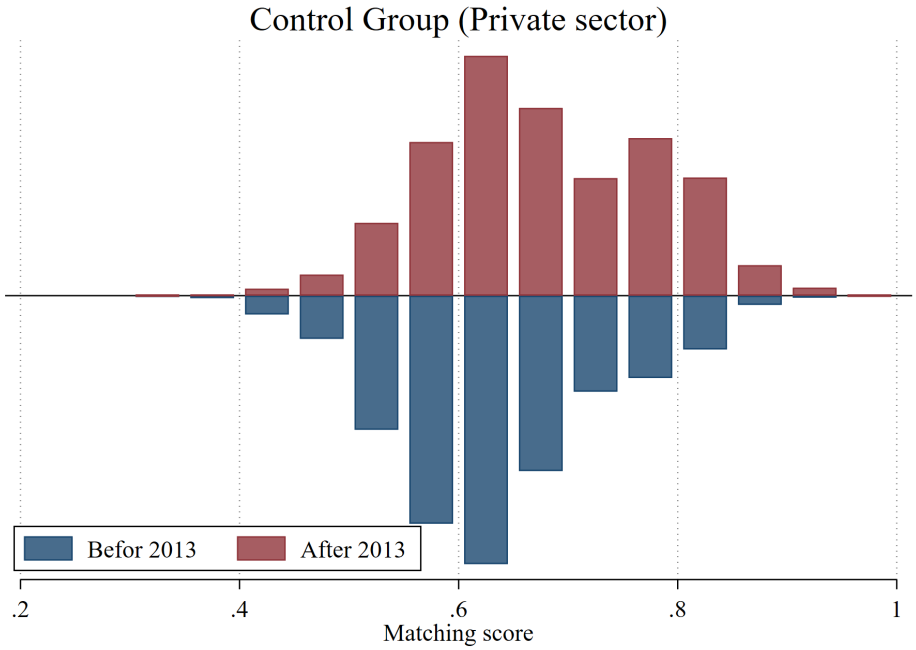

Figure 8.4: Control group matching score for women working in the private sector.

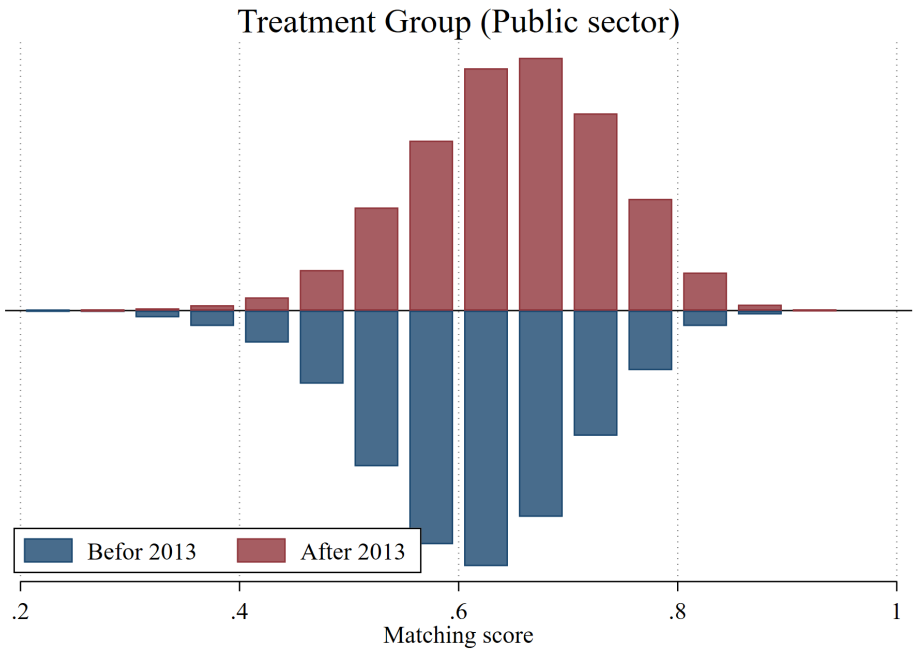

Figure 8.5: Treatment group matching score for women working in the public sector.

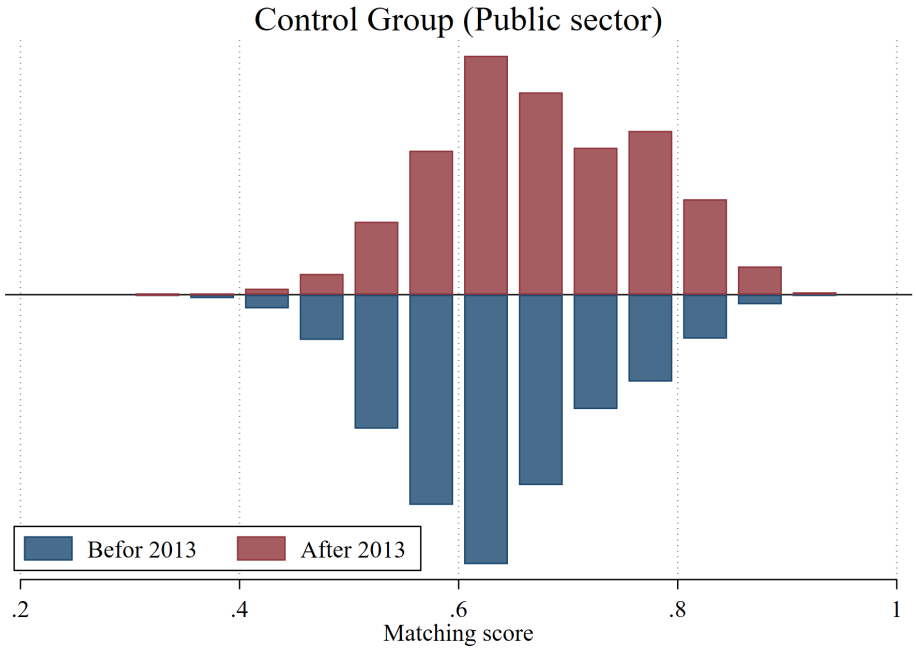

Figure 8.6: Control group matching score for women working in the public sector.

9. Appendix 2

Table 9.1: Robustness of results for all full-time married workers aged 20 to 45

	2010-2016	2011-2017	2010-2017							
Age Group	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
20 to 45	0.00279	0.00197	0.00302	0.00130	0.00303	0.00436	0.00424	0.00438	0.00497	0.00426
	(0.00871)	(0.0111)	(0.00805)	(0.00800)	(0.00843)	(0.00830)	(0.00825)	(0.00832)	(0.00831)	(0.00826)
20 to 34	-0.00092	-0.00663	-0.00394	-0.00517	-0.00497	-0.00105	-0.00116	-0.00097	-0.00133	-0.00109
	(0.0153)	(0.0180)	(0.0138)	(0.0142)	(0.0145)	(0.0142)	(0.0143)	(0.0141)	(0.0141)	(0.0142)
35 to 45	0.0147**	0.0160***	0.0196***	0.0174**	0.0197***	0.0170***	0.0166***	0.0191***	0.0196***	0.0186***
	(0.00685)	(0.00578)	(0.00686)	(0.00663)	(0.00659)	(0.00571)	(0.00571)	(0.00647)	(0.00686)	(0.00646)
Excluding 20-25 (2010-2017)										
20 to 45	-	-	-	-	-	-	-	-	-	0.00470
	-	-	-	-	-	-	-	-	-	(0.00835)
20 to 34	-	-	-	-	-	-	-	-	-	-0.00265
	-	-	-	-	-	-	-	-	-	(0.0148)
35 to 45	-	-	-	-	-	-	-	-	-	0.0186***
	-	-	-	-	-	-	-	-	-	(0.00646)
Work Experience	NO	NO	NO	NO	NO	NO	NO	NO	YES	NO
$(Age)^2$	NO	NO	NO	NO	NO	YES	NO	YES	NO	NO
$(Number\ of\ Children)^2$	NO	NO	NO	NO	NO	YES	YES	NO	NO	NO
Education Level	NO	NO	NO	NO	YES	NO	NO	NO	NO	NO
Age Group (10-Year Range)	NO	NO	NO	YES	NO	NO	NO	NO	NO	NO
Age Group (5-Year Range)	NO	NO	YES	NO	NO	NO	NO	NO	NO	NO
Time Constant Effect	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Fixed Effect of Province	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: The difference-in-differences estimates are calculated based on specification (1). Column (1) reports estimates for 2010 to 2016, column (2) for 2011 to 2017, and columns (3) to (10) for 2010 to 2017. Significance levels: * 90%, ** 95%, *** 99%.

Table 9.2: Robustness of results for married women aged 20 to 45 working in the private sector

	2010-2016	2011-2017	2010-2017							
Age Group	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
20 to 45	0.00892	0.00719	0.00663	0.00518	0.00467	0.00767	0.00758	0.00767	0.00715	0.00759
	(0.0106)	(0.0136)	(0.0105)	(0.0106)	(0.0104)	(0.0105)	(0.0105)	(0.0105)	(0.0103)	(0.0105)
20 to 34	0.00791	-0.00050	0.00322	0.00384	0.00199	0.00518	0.00500	0.00505	0.00357	0.00484
	(0.0168)	(0.0188)	(0.0145)	(0.0158)	(0.0156)	(0.0152)	(0.0151)	(0.0154)	(0.0148)	(0.0153)
35 to 45	0.0149**	0.0210***	0.0170***	0.0166***	0.0144**	0.0147***	0.0143***	0.0168***	0.0164***	0.0164***
	(0.00566)	(0.00590)	(0.00561)	(0.00555)	(0.00534)	(0.00448)	(0.00457)	(0.00542)	(0.00578)	(0.00550)
Excluding 20-25 (2010-2017)										
20 to 45	-	-	-	-	-	-	-	-	-	0.00776
	-	-	-	-	-	-	-	-	-	(0.0112)
20 to 34	-	-	-	-	-	-	-	-	-	0.00446
	-	-	-	-	-	-	-	-	-	(0.0173)
35 to 45	-	-	-	-	-	-	-	-	-	0.0164***
	-	-	-	-	-	-	-	-	-	(0.00550)
Work Experience	NO	NO	NO	NO	NO	NO	NO	NO	YES	NO
$(Age)^2$	NO	NO	NO	NO	NO	YES	NO	YES	NO	NO
$(Number\ of\ Children)^2$	NO	NO	NO	NO	NO	YES	YES	NO	NO	NO
Education Level	NO	NO	NO	NO	YES	NO	NO	NO	NO	NO
Age Group (10-Year Range)	NO	NO	NO	YES	NO	NO	NO	NO	NO	NO
Age Group (5-Year Range)	NO	NO	YES	NO	NO	NO	NO	NO	NO	NO
Time Constant Effect	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Fixed Effect of Province	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: The difference-in-differences estimates are calculated based on specification (1). Column (1) reports estimates for 2010 to 2016, column (2) for 2011 to 2017, and columns (3) to (10) for 2010 to 2017. Significance levels: * 90%, ** 95%, *** 99%.

Table 9.3: Robustness of results for all married women aged 20 to 45 working in the public sector

	2010-2016	2011-2017	2010-2017							
Age Group	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
20 to 45	-0.00256	-0.00241	0.00138	-0.00070	0.00123	0.00232	0.00208	0.00230	0.00343	0.00208
	(0.00847)	(0.0107)	(0.00850)	(0.00846)	(0.00872)	(0.00841)	(0.00825)	(0.00846)	(0.00840)	(0.00831)
20 to 34	-0.00872	-0.0129	-0.00948	-0.0129	-0.0117	-0.00684	-0.00674	-0.00595	-0.00594	-0.00584
	(0.0166)	(0.0189)	(0.0151)	(0.0152)	(0.0160)	(0.0156)	(0.0157)	(0.0152)	(0.0154)	(0.0154)
35 to 45	0.0126	0.0172*	0.0220**	0.0188*	0.0242**	0.0197**	0.0193**	0.0214**	0.0231**	0.0210**
	(0.00914)	(0.00854)	(0.00966)	(0.00925)	(0.00928)	(0.00825)	(0.00829)	(0.00903)	(0.00963)	(0.00905)
Excluding 20-25 (2010-2017)										
20 to 45	-	-	-	-	-	-	-	-	-	0.00227
	-	-	-	-	-	-	-	-	-	(0.00799)
20 to 34	-	-	-	-	-	-	-	-	-	-0.0101
	-	-	-	-	-	-	-	-	-	(0.0149)
35 to 45	-	-	-	-	-	-	-	-	-	0.0210**
	-	-	-	-	-	-	-	-	-	(0.00905)
Work Experience	NO	NO	NO	NO	NO	NO	NO	NO	YES	NO
$(\text{Age})^2$	NO	NO	NO	NO	NO	YES	NO	YES	NO	NO
$(\text{Number of Children})^2$	NO	NO	NO	NO	NO	YES	YES	NO	NO	NO
Education Level	NO	NO	NO	NO	YES	NO	NO	NO	NO	NO
Age Group (10-Year Range)	NO	NO	NO	YES	NO	NO	NO	NO	NO	NO
Age Group (5-Year Range)	NO	NO	YES	NO	NO	NO	NO	NO	NO	NO
Time Constant Effect	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Fixed Effect of Province	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: The difference-in-differences estimates are calculated based on specification (1). Column (1) reports estimates for 2010 to 2016, column (2) for 2011 to 2017, and columns (3) to (10) for 2010 to 2017. Significance levels: * 90%, ** 95%, *** 99%.

Table 9.4: Estimation of the final effect using the Probit method to examine the effect of increasing the length of maternity leave on childbearing among working women (2010 to 2017)

Age (years)	**(1) 20 to 45**		**(2) 20 to 34**		**(3) 35 to 45**	
	With control	**Without control**	**With control**	**Without control**	**With control**	**Without control**
All						
eligible × *Post*	0.00531	0.00334	-0.00219	-0.00278	0.0184***	0.0107
(SE)	(0.00794)	(0.00730)	(0.0126)	(0.0116)	(0.00704)	(0.00822)
Observations	68,147	68,271	31,985	32,028	36,076	36,243
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Private sector						
eligible × *Post*	0.00851	0.00858	0.00473	0.00760	0.0168*	0.0194
(SE)	(0.0115)	(0.0124)	(0.0165)	(0.0166)	(0.00895)	(0.0134)
Observations	42,883	43,003	22,935	22,978	19,932	20,025
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Public sector						
eligible × *Post*	0.00189	0.00317	-0.00948	-0.00539	0.0190**	0.0117
(SE)	(0.00621)	(0.00733)	(0.0102)	(0.0109)	(0.00951)	(0.00931)
Observations	59,898	60,052	27,542	27,599	32,299	32,453
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes

Note: The difference-in-difference estimate was calculated based on specification (1) and using the Probit method. Controls include age, academic status, occupation, workplace characteristics, number of children under 18, urban or rural status, and fixed effects of year and province. * indicates significance at the 90% confidence interval, ** indicates significance at the 95% confidence interval, and *** indicates significance at the 99% confidence interval.