WIDER Working Paper 2021/133

Maternity benefits mandate and women's choice of work in Viet Nam

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August 2021

United Nations University World Institute for Development Economics Research
wider.unu.edu

Abstract: Despite a sizeable literature on the labour market effects of maternity leave regulations on women in developed countries, how these policies affect women's work in developing countries with a large informal sector remains poorly understood. This study examines how extending the maternity leave requirement affects women's decisions to work in the informal or formal sector in Viet Nam. We use a difference-in-differences approach to evaluate the 2012 Amendment to the Viet Nam Labour Law, which imposes a longer maternity leave requirement than before. We find that the law increases formal employment and decreases unpaid work among women in the female labour market. This is driven by women switching from agricultural household work to employment in the public sector. In contrast, we find no effects on formal employment in the private sector. These findings suggest that an increase in the required maternity leave encourages women to switch from informal, unpaid work to working in the formal sector.

Key words: maternity leave regulations, female labour market, informal sector

JEL classification: J08, J2, J3

Acknowledgements: The authors are grateful for comments and feedback from the participants of the UNU-WIDER Workshop on Women's Work and PAA 2021. All errors are the authors'.

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This study has been prepared within the UNU-WIDER project Women's work – routes to economic and social empowerment.

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ISSN 1798-7237 ISBN 978-92-9267-073-3

https://doi.org/10.35188/UNU-WIDER/2021/073-3

Typescript prepared by Gary Smith.

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The Institute is funded through income from an endowment fund with additional contributions to its work programme from Finland, Sweden, and the United Kingdom as well as earmarked contributions for specific projects from a variety of donors.

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1 Introduction

Labour regulations requiring employers to provide maternity benefits are traditionally considered as protection for female workers during the childbearing and child-rearing years, as they provide job security and shield female workers from employer discrimination. In developing countries, many women engage in informal work or care work, where maternity benefits are typically unavailable (ILO 2014). Therefore, when the government imposes requirements on employers to provide maternity benefits for female workers who are pregnant, it may create an incentive for women to switch to a formal job so they can receive the benefits, because such policies often apply only to the formal, regulated sector (Almeida and Carneiro 2012; Pettit and Hook 2005). At the same time, a maternity benefits requirement may impose costs on employers in the formal sector, which, in turn, can discourage the hiring of women in the formal sector (Uribe et al. 2019).

However, the role of maternity benefits in women's decisions on whether to work in the informal or formal sector in developing countries remains understudied. Although there are several studies examining the impacts of general labour regulations on the informal sector of the labour market (e.g. Almeida and Carneiro 2012; Freeman 2010), these studies do not focus on maternity benefits regulations or on women's labour market outcomes. Similarly, although there is a sizeable literature on how mother's labour market outcomes are affected by maternity benefits mandates,[1] they tend to focus exclusively on mothers and, therefore, do not provide any empirical evidence on whether female workers respond to the incentive provided by the maternity benefits. More importantly, most of the existing studies on the employment effects of maternity benefits regulation tend to focus on developed countries, where the labour markets can be very different from those in developing countries. To the best of our knowledge, only Uribe et al. (2019) and Amin and Islam (2019) examine the effect of maternity leave policies on female labour force participation, and only Uribe et al. (2019) examine changes in terms of informal and formal employment among women.

In this paper we assess these important questions in the context of the labour market of Viet Nam during the period between 2010 and 2018. The 2012 Amendment to the Viet Nam Labour Law raised the required maternity leave by two months compared to the original requirement in the law of 1994, providing a unique opportunity to study maternity benefits regulations and female labour market outcomes in developing countries. Viet Nam is particularly interesting to examine because it is known for having a relatively high female labour force participation rate compared to other countries of the same income level (Klasen et al. 2020), and a high share of informal sector workers (ILO 2016). It also imposes relatively generous maternity benefits on employers. Many have argued that Viet Nam's high share of women in work is partly due to the socialist regime's policies to promote gender equality and to draw women into the labour force, such as maternity leave requirements (e.g. Gaddis and Klasen 2014; Klasen 2019; Klasen et al. 2020).

We examine whether the 2012 Amendment has encouraged female workers to transition from informal work to formal employment. We use a difference-in-differences (DiD) approach and compare the labour market outcomes of women of childbearing age and women beyond childbearing age both before and after the law came into effect. Our empirical strategy assumes that longer maternity leave requirements would make formal employment attractive to women of childbearing age, but not to women of older age. We specifically focus on whether women choose to work in the formal sector in response to the maternity benefits they will receive when they become pregnant or when they give birth. Therefore, we do not restrict our study to mothers or those who would become mothers.

[1] See Thomas (2016), Strang and Broeks (2017), and Rossin-Slater (2017) for a summary of this extensive literature.

We find that the law increases formal employment by 3.1 percentage points and decreases agricultural household unpaid work by 4.6 percentage points. We show that these findings are not driven by the cohort effects, pre-treatment trends of differences between the childbearing-age women and older women, or other unrelated factors.

The contributions of this study are twofold. First, despite an extensive body of studies on the effects of maternity leave regulations on women in developed countries (e.g. Akgunduz and Plantenga 2013; Rossin-Slater 2017; Ruhm 1998), there are not many studies on the effects on women in developing countries (Amin and Islam 2019; Uribe et al. 2019). Uribe et al. (2019) find that extending the required maternity leave has a negative effect on women's formal employment in Colombia. In contrast, we find that a similar law has a positive effect on women's formal employment in Viet Nam. This difference means that maternity benefits regulations may have very different effects in different labour market settings.

Second, our study complements a growing literature on women's work in Viet Nam (Feeny et al. 2021; Klasen 2019; Klasen et al. 2020; Kreibaum and Klasen 2015). While the existing studies already explore what might have led to Viet Nam's relatively high female labour force participation, little is known about women's formal employment in this country. Our study shows that extending maternity benefits encourages women to switch from working for family to working for the government. Employment in the public sector has been steadily declining for the last two decades (McCaig and Pavcnik 2013), allowing easier switching to employment in the public sector.

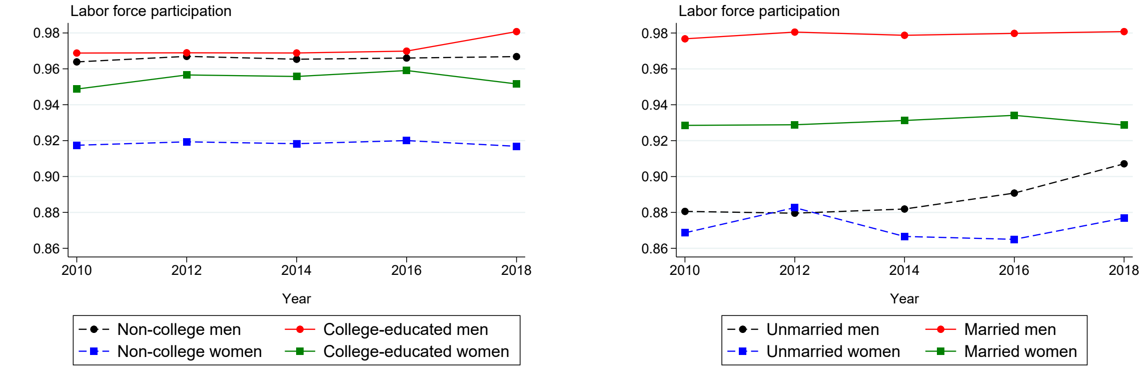
The rest of the paper is organized as follows. Section 2 describes recent trends of women's work in Viet Nam and the context of the 2012 Amendment. Section 3 provides a discussion on our empirical strategy. Section 4 reports on the different data sources used in the study. In Section 5 we report and discuss the findings, and Section 6 concludes the paper.

2 Background

Compared to other low- to middle-income countries, Viet Nam has a relatively high female labour force participation rate and gender equality (Klasen et al. 2020). The labour force participation rate among women aged 25–54 is slightly lower than that of men in the same age group. The female labour force participation rate of college-educated women is roughly 95 per cent and that of non-college-educated women is roughly 92 per cent; these rates were relatively stable during the 2010–18 period, as indicated in Figure 1. The labour force participation rate among college-educated and non-college-educated men is roughly 97 per cent, but for college-educated men it increased to 98 per cent in 2018. Married individuals are more likely to work than are unmarried individuals, but participation among unmarried men increased from 88.2 per cent to 90.7 per cent during 2014–17.

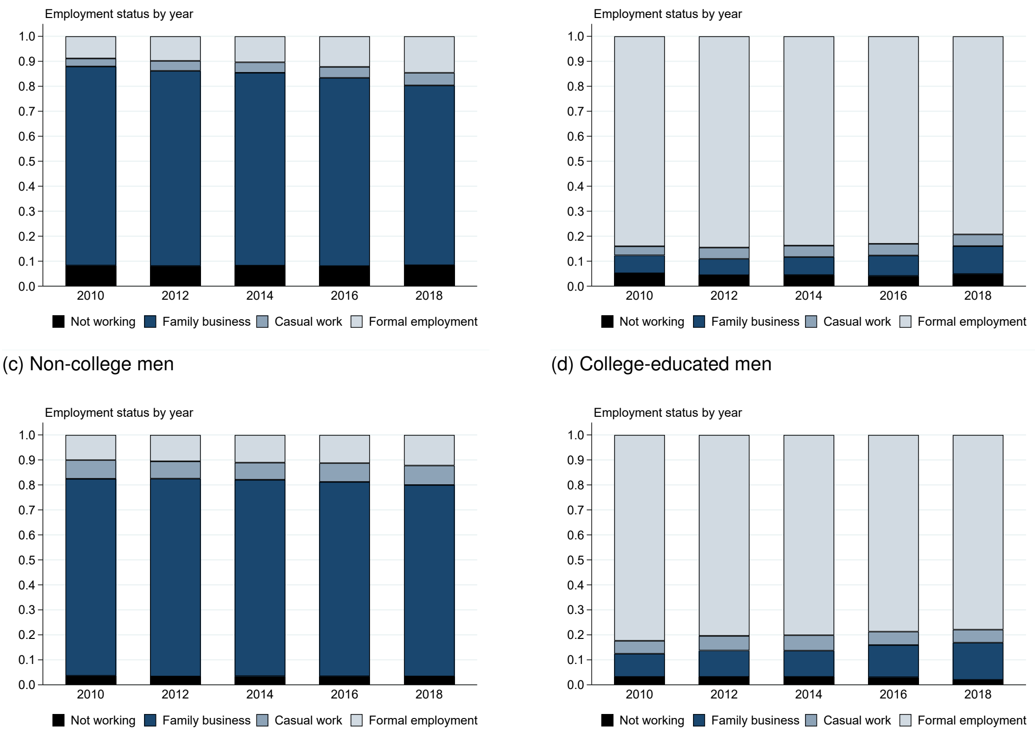
The labour market composition is also remarkably different across gender and college education, as illustrated in Figure 2. Most non-college-educated men and women work for the household business, which is typically unpaid, while college-educated men and women mainly work in the formal sector (defined as wage employment that provides social insurance). Only a small share of non-college-educated women are casual wage workers (waged employment without social insurance) relative to non-college-educated men. The labour market composition is relatively stable over time, although the share of non-college-educated men and women who work in the formal sector appears to increase over time.

Figure 1: Labour force participation of men and women aged 25–54 in 2018 by college education and marital status
(a) By education (b) By marital status



Note: the sample includes all men and women aged 25–54.
Source: authors' calculations based on the Viet Nam Household Living Standard Survey (VHLSS) 2010–18 (see text for description).

Figure 2: Labour market composition among men and women aged 25–54 in 2018 by college education
(a) Non-college women (b) College-educated women
(c) Non-college men (d) College-educated men



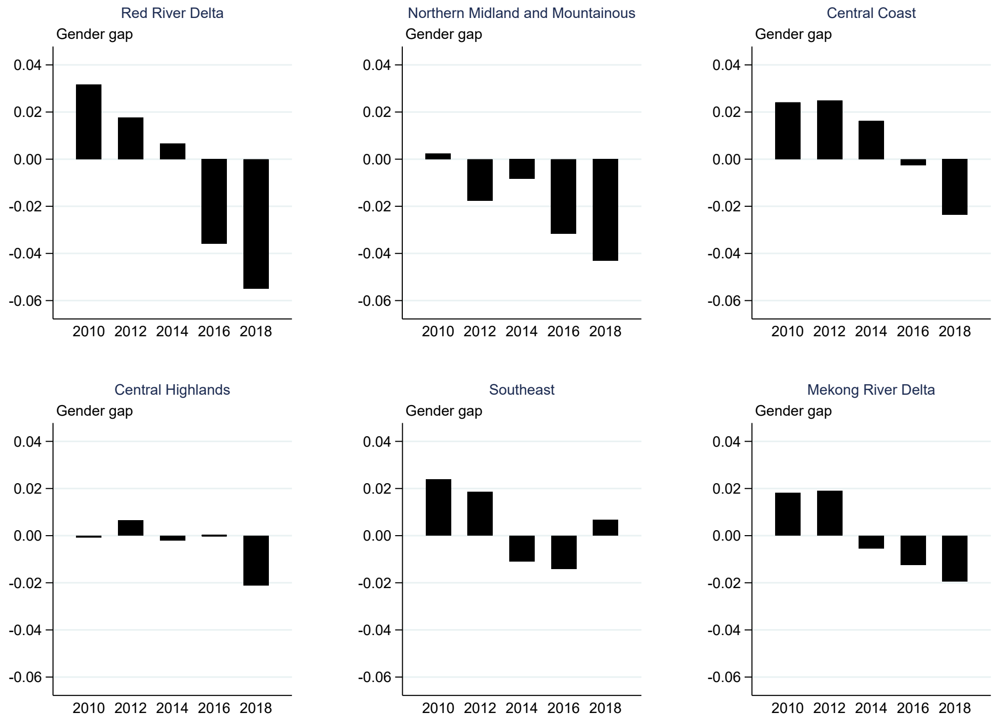
Note: the sample includes all men and women aged 25–54.
Source: authors' calculations based on the VHLSS 2010–18.

The gender gap in formal employment, defined as the difference between the share of men with formal employment and the share of women with formal employment, appears to have declined across the country during this period, as indicated in Figure 3. The Red River Delta has the highest and the Central Highlands has the lowest gender gap in 2010. By 2018, most regions had already reversed the gender gap, except for the Southeast region. This pattern is caused by the share of women working in the formal sector rising faster than the share of men in the formal sector.[2] One potential explanation for this shift

[2] According to the VHLSS data, the share of women with household unpaid work decreases faster than the share of men with household unpaid work. As a result, there is an increase in the gender gap in household unpaid work during 2010–18.

in the gender gap is the fact that the share of women who are college-educated also rises faster than the share of men who are college-educated in these regions (Figure 4).

Figure 3: Gender gap in formal employment by region and year



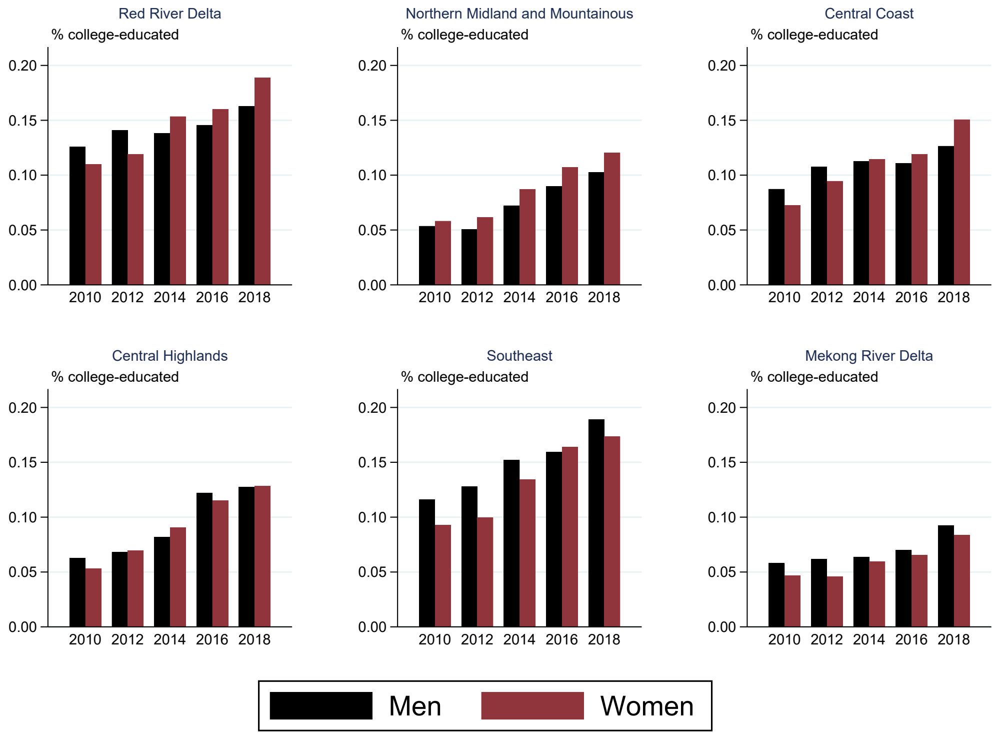
Note: the sample includes all men and women aged 25–54.

Source: authors' calculations based on the VHLSS 2010–18.

The main regulatory framework of the labour market, especially for formally registered businesses, is the Viet Nam Labour Law, which was first written in 1994 (World Bank 1995) and amended in 2002, 2012, and 2019. The original Viet Nam Labour Law of 1994 already provided substantial protections for female workers. Under the original law, female workers were entitled to four months of paid leave during prenatal and postnatal periods, and such leave would be extended by one month for each additional child. During this leave, female workers would still receive full wages from their employer and maternity benefits from the Viet Nam Social Insurance Fund. More importantly, female workers were guaranteed their previous job back, or guaranteed a new job with an equivalent wage. The law also prohibited employers from giving work that might be harmful to the mother or the child. After giving birth, female workers were able to take an additional hour-long break.

The 2012 Amendment to the Viet Nam Labour Law was passed in June 2012 and came into effect on 5 January 2013. The Amendment makes several changes to the original codes. First, it formally defines strikes, including when and how they are allowed. Second, the new law prohibits workers from working more than eight hours/ per day and 48 hours per week; the law also limits overtime work to 30 hours per month and 200 hours per year. Third, the Vietnamese New Year holidays are extended by one additional day, and workers are allowed to take an unpaid day off when a family member passes away. Fourth, the law formally states that: (1) wages shall be paid based on labour productivity and quality of work performed; and (2) the minimum wage is the lowest payment for an employee who performs the simplest work and must ensure his or her minimum living needs. Most importantly, the new law extends the mandated paid maternity for female workers up to six months during the prenatal and postnatal periods. This increases the required maternity leave by two months.

Figure 4: Percentage of men and women who are college-educated by region and year



Note: the sample includes all men and women aged 25–54.
Source: authors' calculations based on the VHLSS 2010–18.

3 Empirical strategy

3.1 Main models

Our empirical strategy relies on the assumption that female workers with higher fertility rates are more likely to switch to the formal sector in response to the maternity benefits incentive from the new law. In the first analysis, we compare employment outcomes between women of childbearing age and women older than childbearing age before and after the law came into effect, using a DiD estimation strategy.

Formally, we estimate the following model for woman i aged 25–54 in year t:

$$Y_{it} = \beta + \delta.(Age\ 25\text{–}44_{it} \times Post_t) + \mathbf{X}'_{\mathbf{it}}\Gamma + \theta_{p,t} + \eta_{p,g} + \varepsilon_{it} \tag{1}$$

where Y_{it} denotes the employment outcomes of woman i in year t, $Age\ 25\text{-}44_{it}$ indicates whether the individual is between age 25 to 44, and $Post_t$ indicates whether year t is after 2013. We control for province-by-year fixed effects ($\theta_{p,t}$), province-by-age-group fixed effects ($\eta_{p,g}$), and other individual characteristics, captured by $\mathbf{X}_{\mathbf{it}}$; ε_{it} denotes the error term. The sample is split into six age groups: 25–29, 30–34, 35–39, 40–44, 45–49, and 50–54. The coefficient of interest is δ. Individual characteristics include urban, ethnicity, household size, number of children under 10 years old in the house, primary education or less, secondary education, and marital status. In all models we cluster standard errors at the commune–year level to account for the survey's sampling design.

The assumptions required to identify the causal effect of the law are that: (1) the employment outcomes of the childbearing age and non-childbearing age groups would have followed the same trends if the required maternity leave length had not been changed; and (2) there are no spillover effects onto the non-childbearing age group. The first assumption would be violated if there were any other factors that affect changes in employment outcomes differently for the childbearing-age and non-childbearing-age women but were unrelated to the new law.

A major concern is that the labour law may be confounded by the rapid structural transformation that Viet Nam has experienced since 2000, as workers moved out of the agricultural sector and into higher-productivity sectors such as manufacturing and services (Liu et al. 2020; McCaig and Pavcnik 2013). A large part of this structural transformation process was due to trade liberalization and enterprise reforms that increased the number of employment opportunities in the manufacturing sector, specifically in the clothing, food products and beverages, furniture, and footwear industries (McCaig and Pavcnik 2013). McCaig and Pavcnik (2015) and Liu et al. (2020) further note that younger and more educated workers are more likely to be affected. Limited employment opportunities in the agricultural sector, along with mechanization and uptake of labour-saving inputs, allow younger workers who are increasingly more educated to transition out of agriculture and into non-farm sectors (Liu et al. 2020).

We address this concern in various ways. First, our estimation strategy in Equation 1 includes province–year fixed effects to account for changes in the local labour market over time that were caused by the structural transformation process. It also includes province–age-group fixed effects and year-specific cohort trends to control for unobserved heterogeneity between age groups within each province and to account for the fact that younger cohorts might be more likely to switch between agriculture and non-farm sectors over time. Lastly, we exclude college-educated women and control for educational attainment in Equation 1, since younger cohorts tend to have higher educational attainment and, hence, are more likely to switch to the non-farm sectors.

We also assess the extent to which our main results are driven by the differential effects of the structural transformation on younger and older workers by replacing women aged 45–54 with men aged 25–44 as the control group for the DiD estimation. Because this alternative control group is as young as the treatment group, the results from this estimation are not driven by the differential effects of the structural transformation process.[3] Therefore, if the estimates using men aged 25–44 as the control group are not different from the main findings (i.e. using women aged 45–54 as the control group), then we may conclude that the structural transformation does not confound the labour law that raises the maternity leave requirement.

3.2 Heterogeneous effects

We assess the dynamics of treatment effects over time by estimating the following event-study specification:

$$Y_{it} = \beta + \sum_{s \neq 2012} \delta_s.(Age\ 25\text{–}44_{it} \times Year_s) + \mathbf{X}'_{\mathbf{it}}\Gamma + \theta_{p,t} + \eta_{p,g} + \varepsilon_{it} \tag{2}$$

where $s = \{2010, 2014, 2016, 2018\}$ and 2012 is the reference year. Similarly, we control for province–year fixed effects and province–age-group fixed effects to absorb any unobserved heterogeneity across age groups at the province level and changes in the labour market across provinces and years. Embedded in this model is a check for parallel pre-treatment trends (or pre-trends): δ_{2010} represents the difference-of-differences of employment outcome between the childbearing group and the non-childbearing group in 2010 and in 2012. Because the law was implemented in 2013, we would expect δ_{2010} to be close to zero and/or statistically insignificant.

[3] Liu et al. (2020) find that the gender composition of the agricultural labour force was stable during the 1992–2016 period, which suggests that the structural transformation process may not affect men and women differently.

We also assess the treatment effects by age groups by splitting the treatment group into five-year age groups: 25–29, 30–34, 35–39, and 40–44. The control group is women aged 45–54. We estimate the following model:

$$Y_{it} = \beta_0 + \sum \gamma_g (Post_t \times AgeGroup_{gt}) + \theta_{p,t} + \eta_{p,g} + \mathbf{X}_{\mathbf{it}}'\mathbf{\Gamma} + \varepsilon_{it} \quad (3)$$

where γ_g denotes the treatment effect on women of age group g. We also control for province–year fixed effects and province–age-group fixed effects and individual characteristics, as before.

3.3 Comparison between low-birth-rate and high-birth-rate groups

One concern of using younger women as the treatment group and older women as the control group is that other factors unrelated to the law might have affected young women more than older women. Therefore, we estimate the effects by women's expected birth rates instead of by age. Specifically, we estimate the birth rates for each district-by-age group for ages between 25 and 45, denoted as $BirthRate_{g,d}$, using the 2009 Population and Housing Census. We split this variable into five equal-sized bins, then estimate the following model:

$$Y_{it} = \beta + \sum_b \delta_b (W_b \times Post_t) + \theta_{d,t} + \eta_{d,g} + \gamma_{t,g} + \mathbf{X}_{\mathbf{it}}'\mathbf{\Gamma} + \varepsilon_{it} \quad (4)$$

where W_b denotes whether the birth rate of the individual falls into bin b, $\theta_{d,t}$ are district–year fixed effects, $\eta_{d,g}$ are district–age-group fixed effects, and $\gamma_{t,y}$ are year–age-group fixed effects. These extensive fixed effects account for any unobserved heterogeneity across age groups within districts and any secular shocks across districts and age groups over time. This specification also allows the effects to be non-monotonic across different birth-rate groups. Note that the five birth-rate groups are for women aged 25–45, and the lowest birth-rate group will be used as the reference group. One may add a sixth group, group zero, for women aged 45–54; since this group is unlikely to respond to the maternity benefits, one would expect the effect on this group to be zero and statistically insignificant.

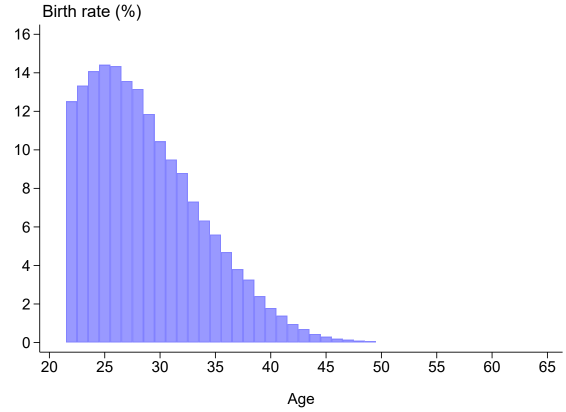
4 Data

Our primary data source is the Viet Nam Household Living Standard Survey (VHLSS), a nationally representative, biennial survey conducted by the General Statistics Office of Viet Nam and the World Bank. The survey sample, which was based on the Population and Housing Census in 1999 and in 2009, consists of roughly 9,000 households from all over the country. The survey collects data from all members in the household, so the sample includes over 36,000 individuals. The data provide rich information about labour force participation, as well as individual and household demographics, economics, and education. For this study, we use the five most recent waves of the data, covering the period 2010–18, to construct a repeated cross-sectional sample.

We restrict the study population to individuals aged 25–54 to exclude those who are still in school or are already retired.[4] We focus our analysis on non-college-educated women because 80 per cent of this group have family unpaid work and, hence, are more likely to respond to the maternity leave increase. We set the childbearing age to be between 25 and 44 because birth rates fall below 0.1 per cent for women older than 45 (Figure 5) and because our sample covers women aged 25–54.

[4] The official retirement ages in Viet Nam are 60 for men and 55 for women.

Figure 5: Birth rates by age



Source: authors' calculations using the most recent live birth information from the 2009 Population and Housing Census.

In order to measure changes in formal and informal employment, we focus on three main employment outcomes, defined as follows, along with individuals' monthly incomes:

- Formal employment: holding a job that (1) comes with a wage or salary, and (2) provides social insurance (ILO 2016). This category mainly includes workers in the public or private sectors (domestic and foreign firms).

- Not working: not holding any job. The reasons for not working include studying, housework, old or retired, disabled or chronic illness, or cannot find a job. The main reason of not working for men is disability, while the main reason for women is housework (Figure 6).

- Agricultural and non-agricultural household work: workers who contribute to a family business in agriculture or non-farm sectors. The ILO classified this work as informal employment (ILO 2016).

Figure 6: Reasons for not working, 2014–18

Source: authors' calculations based on the VHLSS 2014–18 for individuals aged 22–65.

Table 1 provides a summary of key statistics for women aged 25–44 and 45–54 for before and after the new law took effect in 2013. In Figure 7 we also plot the labour market outcomes of interest for these groups by year. In Figure 7(a) we note that the share of women with formal employment in the 25–44 age group is higher than that of the 45–54 age group across all years in the sample. During the 2010–12 period, the formal employment of the younger group increases slightly while that of the older group does not. After 2013 the formal employment share among the childbearing age group increases faster—especially after 2014—while the share among the older group does not.[5] The difference between the two groups widens over time, suggesting that treatment effects may vary over time.

Table 1: Summary statistics

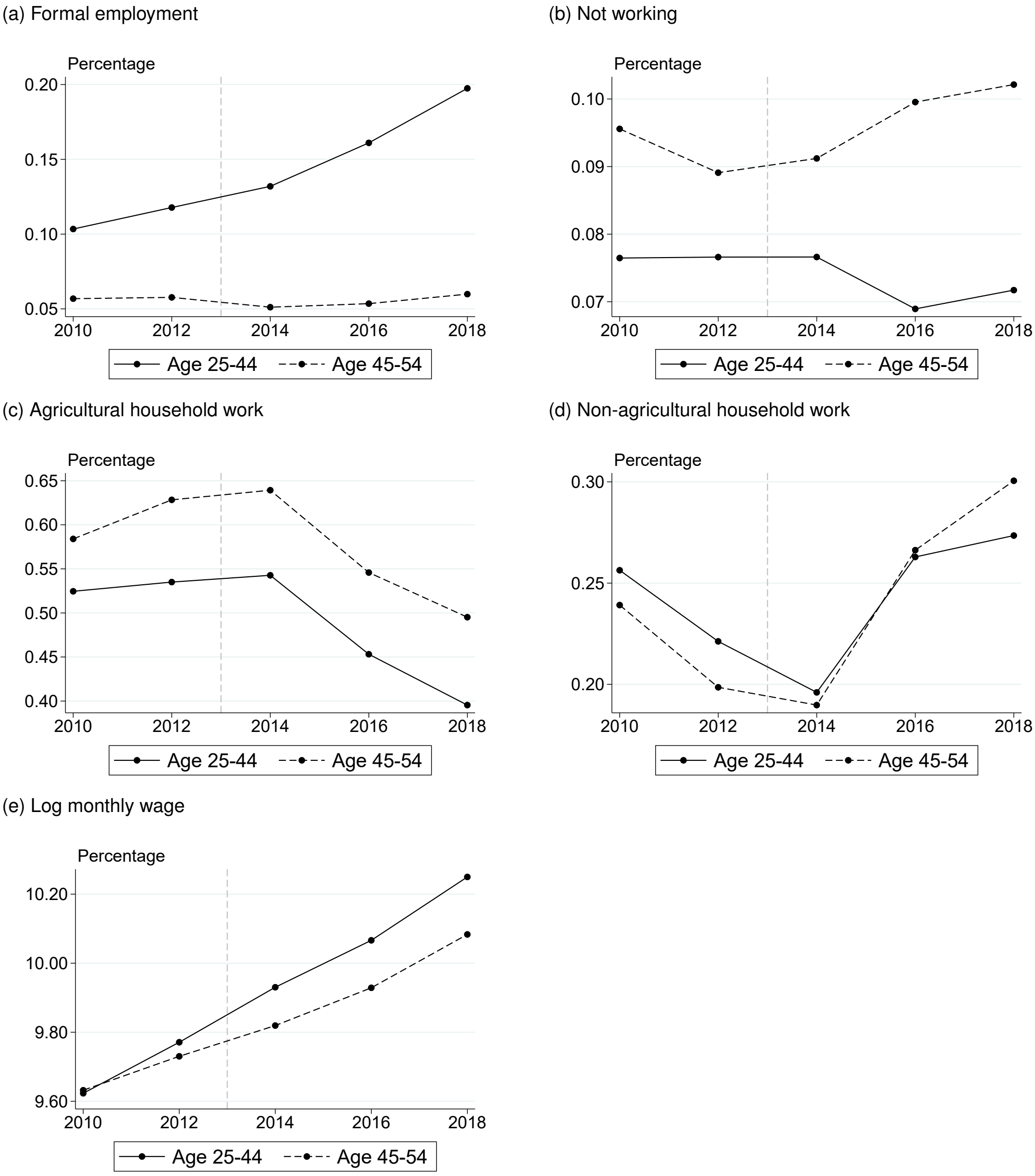
	2010–12		2014–18	
	Age 25–44	Age 45–54	Age 25–44	Age 45–54
Demographics and education level				
Age	34.60	49.36	35.14	49.48
	(5.76)	(2.83)	(5.70)	(2.88)
Urban	0.26	0.29	0.26	0.29
	(0.44)	(0.45)	(0.44)	(0.45)
Married	0.87	0.83	0.88	0.84
	(0.33)	(0.38)	(0.32)	(0.36)
Children in household	1.02	0.37	1.04	0.44
	(0.93)	(0.66)	(0.94)	(0.72)
Primary education or none	0.55	0.52	0.50	0.50
	(0.50)	(0.50)	(0.50)	(0.50)
Lower or upper secondary education	0.45	0.48	0.50	0.50
	(0.50)	(0.50)	(0.50)	(0.50)
Labour market outcome				
Formal employment	0.11	0.06	0.16	0.05
	(0.31)	(0.23)	(0.37)	(0.23)
Not working	0.08	0.09	0.07	0.10
	(0.27)	(0.29)	(0.26)	(0.30)
Unpaid work	0.66	0.74	0.59	0.71
	(0.47)	(0.44)	(0.49)	(0.45)
Agricultural household work	0.53	0.61	0.47	0.56
	(0.50)	(0.49)	(0.50)	(0.50)
Non-agricultural household work	0.24	0.22	0.24	0.25
	(0.43)	(0.41)	(0.43)	(0.43)
Log earning	9.70	9.68	10.09	9.96
	(0.75)	(0.87)	(0.69)	(0.78)
N	19,592	8,972	25,702	14,041

Source: authors' estimation based on the VHLSS 2010–18.

The share of women not working in the 45–54 age group is higher than that of the younger group; after 2013, the share among the younger age group also declines. The share of women doing agricultural household work is also higher for the 45–54 age group than that of the 25–44 age group, and these shares follow similar trends before and after 2013. In contrast, the shares of women doing non-agricultural household work are very similar among the two age groups, and the shares of both groups start to increase after 2014.

[5] The law was passed in 2012, but came into effect in January 2013.

Figure 7: Employment outcome for women aged 25–54 by year



Source: authors' calculation based on the VHLSS 2010–18.

5 Results

5.1 Main findings

We report the main results from estimating Equation 1—that is, a DiD model with women aged 25–44 as the treatment group and women aged 45–54 as the control group—in Table 2. The results for different outcomes are presented in panels A–E. In panel A the outcome is formal employment, defined as holding a job that provides wage/salary and social insurance. In panel B the outcome is not holding a job. In panel C and D the outcome variables are whether individuals work for an agricultural or non-agricultural household, respectively. In panel E is the outcome monthly wage in log from the current job among those who earn a monthly wage. In other words, those who are self-employed or do not work are not included in this regression.

Table 2: DiD estimates for effects on employment outcome

Specification	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Formal employment						
	0.046***	0.045***	0.039***	0.043***	0.038***	0.031**
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel B: Not holding a job						
	–0.002	–0.001	–0.006	–0.001	–0.001	0.002
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel C: Agricultural household work						
	–0.037**	–0.040**	–0.043**	–0.035*	–0.031*	–0.046**
	(0.019)	(0.019)	(0.019)	(0.018)	(0.018)	(0.019)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel D: Non-agricultural household work						
	–0.014	–0.010	0.002	–0.016	–0.014	0.007
	(0.019)	(0.019)	(0.019)	(0.019)	(0.019)	(0.021)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel E: Log monthly income						
	0.080	0.081	0.072	0.110*	0.026	–0.024
	(0.060)	(0.059)	(0.060)	(0.057)	(0.066)	(0.085)
N	9,129	9,129	9,129	9,095	8,432	7,507
Additional controls						
Province FE and year FE	✓					
Province × year FE		✓	✓			
Province × age group FE			✓			
District FE				✓	✓	✓
District × year FE					✓	✓
District × age group FE						✓

Note: this table reports the results from estimating the main DiD model. The interaction term being reported here is *Post-2013 × Age 25–44*. Standard errors are clustered at the commune–year level and reported in parentheses and *p*-values are reported in brackets; sampling weights are applied. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, marital status, and year-specific cohort linear trends. The sample includes non-college-educated women aged 25–54 in the VHLSS 2010–18 sample.

Source: authors' calculations.

We consider six different specifications that control for different levels of fixed effects to absorb any unobserved heterogeneity at the local labour market level over time. The first specification, in column (1), includes fixed effects for province and for year. The second specification, in column (2), includes province–year fixed effects. The third specification, in column (3), additionally controls for province–age-group fixed effects. Since it is possible that the relevant labour market is at the district level instead of the province level, columns (4)–(6) repeat the same specifications in columns (1)–(3) but replace province-level fixed effects with district-level fixed effects.

Formal employment: the estimation in panel A of Table 2 shows that the increase in maternity leave length leads to a statistically significant increase in the probability of formal employment; this result is robust across all specifications. The magnitude of the estimates ranges from 3.9 percentage points when controlling for province–year fixed effects and province–age group fixed effects to 4.6 percentage points when controlling for province fixed effects and year fixed effects.

Not holding a job: in contrast to the estimates for formal employment, we find no significant effects of the increase in maternity leave length on the probability of not holding a job. The estimates are close to zero and insignificant for all specifications, which strongly suggests that the law has no effect on the probability that women do not work.

Agricultural household work: we find that the law is associated with a statistically significant decrease in the probability of agricultural household work. In the least parsimonious model (controlling for province FE and year FE), the estimate is –0.037, while in the most parsimonious model (controlling for district–year FE and district–age-group FE), the estimate is –0.046.

Non-agricultural household work: we find that the law is not associated with any change in non-agricultural household unpaid work. The effect size is small and the estimates are insignificant across all models.

Log monthly income: we also find no association between the new law and log monthly income. All estimates are small and statistically significant (except for column (4)). The sample sizes of these regressions are small because they only include individuals who have wage income.

Taken together, the DiD estimates from this analysis suggest that the new law has a positive impact on formal employment and negative impact on agricultural household work. The law has no effect on the probability of not holding any job and the probability of working for a non-agricultural household business. These findings suggest that female workers respond to the increased maternity leave mandate by switching from family work, especially agricultural household work, to formal employment.

As discussed in Section 3.1, the effects of the labour law may be confounded by the structural transformation that has different effects on young and older workers. McCaig and Pavcnik (2015) and Liu et al. (2020) suggest that the structural transformation has a stronger effect on younger workers because there were more employment opportunities in the non-farm sector that were suitable for young and more educated workers. To check whether the main results were driven by such effects, we replace the women aged 45–54 with men aged 25–44 as the control group in Equation 1. Because this alternative control group is the same age as the treatment group, the results from this estimation would not be affected by the structural transformation. We report the results of these estimations in Table A1 in Appendix A.

We find that using men aged 25–44 as a control group yields slightly more conservative estimates compared to the main findings. The increase in maternity leave length leads to a statistically significant increase in the probability of formal employment; this result is, again, robust across all specifications. The magnitude of the estimates ranges from 2.9 percentage points when controlling for district–year fixed effects and district–age-group fixed effects to 3.6 percentage points when controlling for province fixed effects and year fixed effects.

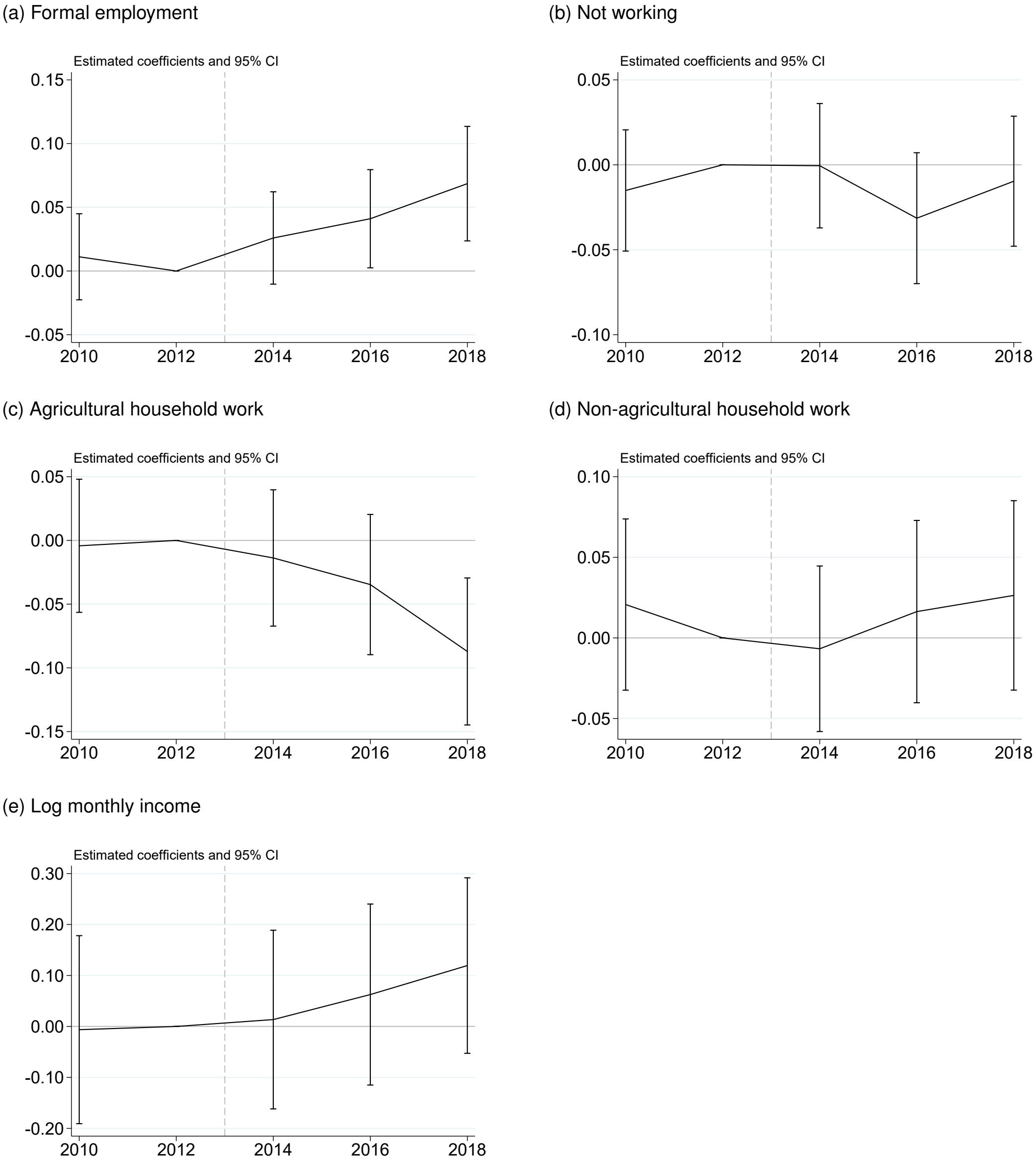
Similarly, we find that the law leads to a statistically significant decrease in the probability of agricultural household unpaid work. In the least parsimonious model (controlling for province fixed effects and year fixed effects), the estimate is –2.8 percentage points, while in the most parsimonious model (controlling for district–year fixed effects and district–age group fixed effects), the estimate is –2.9 percentage points. The estimates are also slightly more conservative than the main findings. Interestingly, while we do not find any effects on log monthly income in the main findings, we do find positive effects using the alternative control group.

To understand this puzzling result, we plot the log monthly income for men and women aged 25–44 over time in Figure A1. The trajectory of women's log monthly income does not appear to have changed during this period, while men's log monthly income appears to have increased at a slower rate after the law was implemented in 2013. Figure A1 indicates that the positive estimates for log monthly income in Table A1 is driven by a change of men's trajectories instead of women's, which is unlikely to be related to the increase in maternity leave requirements.

5.2 Heterogeneous effects

The results from estimating the event-study specification in Equation 2 are largely consistent with the main results, as shown in Figure 8. The estimates for the 2010 coefficients are close to zero and statistically insignificant for all outcome variables. These results suggest that pre-treatment trends are not a major concern in this case.

Figure 8: Event-study estimates for effects on women's labour market outcomes



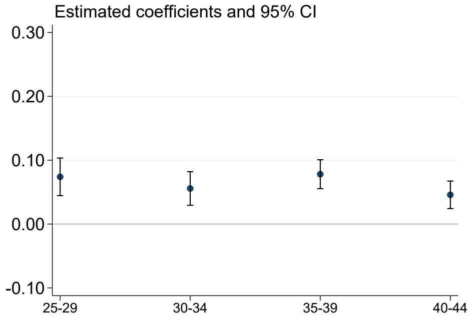
Note: standard errors are clustered at the commune–year level.

Source: authors' estimation using an event study with controls for age variables, individual characteristics, district–year fixed effects, and district–age-group fixed effects.

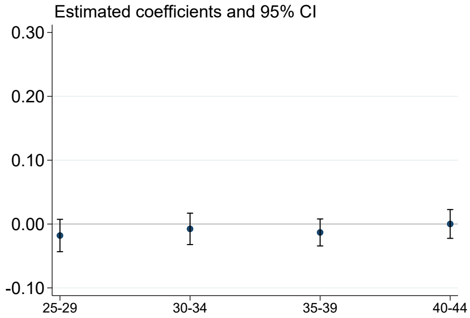
We estimate the heterogeneous effects by age group using Equation 3 and report the results in Figure 9.The effects on formal employment are relatively similar across age groups. The effects on agricultural household work are negative and statistically significant for the 25–29, 35–39, and 40–44 age groups. The estimates for non-agricultural household work are statistically insignificant for all age groups except the 35–39 age group. The effects on the probability of not working at all and on log monthly income are statistically insignificant across all age groups.

Figure 9: DiD estimates for effects by age

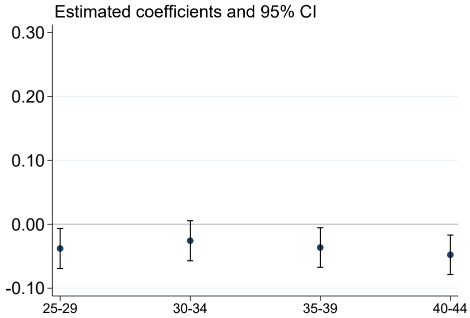
(a) Formal employment



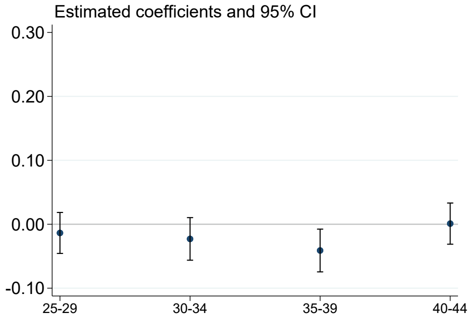
(b) Not working



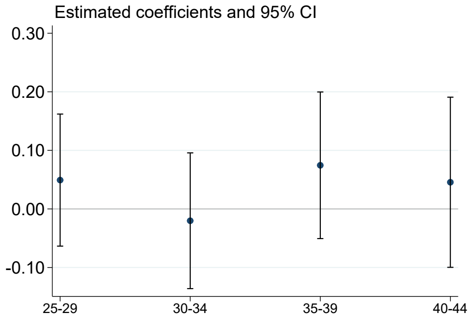
(c) Agricultural household work



(d) Non-agricultural household work



(e) Log monthly income



Note: the plots report the results from estimating the DiD model in Equation 3, in which treatment effects are allowed to vary by age group. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, and marital status. The sample includes women aged 25–54 in the VHLSS 2010–18 sample. Standard errors are clustered at the commune–year level and sampling weights are applied in the regressions.

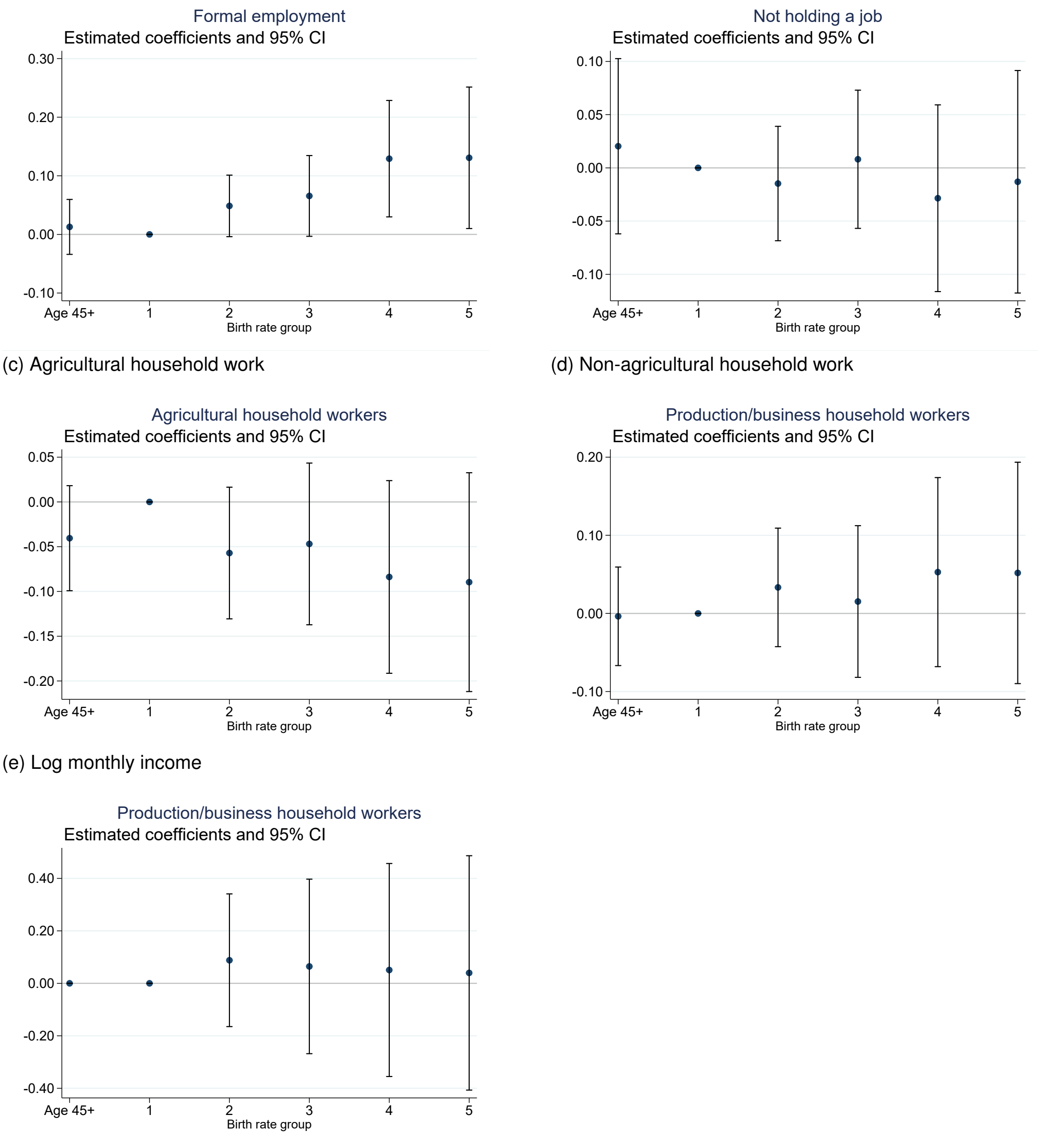
Source: authors' calculations based on the VHLSS 2010–18.

Next, we turn to the results from estimating Equation 4, where we use the district–age-group birth rates from 2009 as a continuous treatment variable instead of childbearing age. To allow for non-monotonicity, we divide the birth rates into five equal-sized bins. The plots in Figure 10 show the effects on the four groups with higher birth rates relative to the group with the lowest birth rates. We also include a group zero, which includes women aged 45–54. Since this group would not respond to the mandated benefits, we would expect the effect on this group to be statistically insignificant.

The estimates for the formal employment outcome are consistent with our main findings: women with higher birth rates experience a larger increase in the probability of formal employment after the new law is enacted. The effect on group zero is statistically insignificant, confirming again that this is the effect of the mandated maternity benefits. Similarly, we find no evidence that the probability of not holding a job is affected in Figure 10(b). Interestingly, we find a negative effect on agricultural household work, but the estimates are insignificant across all birth-rate groups. We find no effects on non-agricultural household work and log monthly income.

Figure 10: DiD estimates for effects by district–age-group birth-rate bin

(a) Formal employment (b) Not working



(c) Agricultural household work

(d) Non-agricultural household work

(e) Log monthly income

Note: the plots report the results from estimating the DiD model in Equation 4, in which treatment effects are allowed to vary by birth-rate groups. Birth rates vary by district and age, and are binned into five equal groups. Individuals age 45+ are included in a separate group. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, and marital status. The sample includes women aged 25–54 in the VHLSS 2010–18 sample. Standard errors are clustered at the commune–year level and sampling weights are applied in the regressions.

Source: authors' calculations based on the VHLSS 2010–18.

5.3 Formal employment in the public and private sectors

The main findings raise a question on why formal employment among childbearing-age women rises in response to the longer maternity leave requirement when the law imposes higher hiring costs on employers. It is possible that the rise in formal employment largely happens in the public sector because the government, as an employer, is less responsive to hiring costs.

To check this, we estimate the DiD model separately for public formal employment and private formal employment and report the results in Table 3. We find that the law increases public formal employment but does not appear to affect private formal employment. We also examine changes in monthly wage by sector, but find no evidence of any effects on these outcomes. These results indicate that the main findings are indeed driven by women moving into employment in the public sector. There are no significant changes to private formal employment, which suggests that the private sector is not responsive to the longer maternity leave requirement.

Table 3: DiD estimates for effects on the public and private formal sectors

Specification	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Public sector						
	0.031***	0.030***	0.030***	0.032***	0.027***	0.024***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel B: Private formal sector						
	0.015	0.015	0.009	0.011	0.010	0.006
	(0.011)	(0.010)	(0.011)	(0.011)	(0.011)	(0.012)
N	34,701	34,701	34,701	34,701	34,653	34,422
Panel C: Log monthly income (public)						
	0.038	–0.023	–0.102	0.101	0.252	0.596*
	(0.104)	(0.120)	(0.148)	(0.120)	(0.190)	(0.346)
N	1,388	1,355	1,265	1,231	676	349
Panel D: Log monthly income (private formal)						
	–0.048	–0.063	–0.062	–0.005	0.039	–0.172
	(0.088)	(0.094)	(0.098)	(0.091)	(0.111)	(0.184)
N	2,418	2,366	2,302	2,335	1,877	1,432
Additional controls						
Province FE and year FE	✓					
Province × year FE		✓	✓			
Province × age group FE			✓			
District FE				✓	✓	✓
District × year FE					✓	✓
District × age group FE						✓

Note: this table reports the results from estimating the main DiD model. The interaction term being reported here is *Post-2013* × *Age 25–44*. Standard errors are clustered at the commune–year level and reported in parentheses and *p*-values are reported in brackets; sampling weights are applied. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, marital status, and year-specific cohort linear trends. The sample includes non-college-educated women aged 25–54 in the VHLSS 2010–18 sample.
Source: authors' calculations based on the VHLSS 2010–18.

6 Conclusion

The 2012 Amendment to the Labour Law of Viet Nam effectively extends the mandated maternity leave from four months to six months, which is likely enforceable among firms in the formal sector but not informal businesses. This provides an increase in employment benefits particularly for women of childbearing age, and incentivizes these women to switch from informal work, such as farm or non-farm household work, to formal employment.

We examine this issue by using a DiD empirical strategy—that is, comparing the employment outcomes of women of childbearing age with those beyond childbearing age before and after the implementation of the law. We find robust evidence that the law is associated with an increase of 3.1 percentage points in the probability of formal employment, and a decrease of 4.6 percentage points in the probability of working in a farm or non-farm household. In contrast, we find no evidence that the law affects the probability of not working or the probability of being an unpaid worker for production/household business.

We also find that the positive effects on formal employment happens in the public sector, but not in the private formal sector. This is likely because the government is less sensitive to increases in hiring costs than are private employers. It is also important to note that public sector employment has been declining since the 2000s due to increases in foreign direct investment (FDI) and the number of jobs in the private sector (McCaig and Pavcnik 2013). This means finding jobs in the public sector has likely become easier for those who want to switch jobs.

Why is the private sector not responsive to the longer maternity leave requirement, both in terms of employment and wages? There are several potential explanations for this finding. First, the increase in mandated leave is relatively small, so the maternity leave cost might be small for firms. Second, it is possible that firms find it beneficial to pay for maternity leave to retain workers who give birth or to attract female workers. Therefore, future research should explore the effects of the maternity leave requirement from the employers' perspective.

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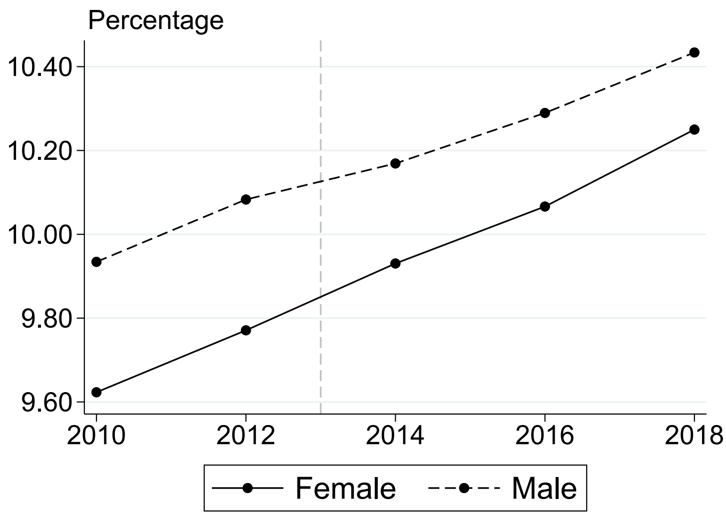
Appendix A: Extra tables and figures

Figure A1: Log monthly income for individuals aged 25–44, by gender



Source: authors' calculations using the 2010–18 VHLSS.

Table A1: DiD estimates for effects on employment outcome using men aged 25–44 as the control group

Specification	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Formal employment						
	0.036***	0.034***	0.034***	0.035***	0.030***	0.029***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
N	45,294	45,294	45,294	45,294	45,263	45,215
Panel B: Not holding a job						
	–0.004	–0.001	–0.001	–0.003	0.000	0.000
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
N	45,294	45,294	45,294	45,294	45,263	45,215
Panel C: Agricultural household work						
	–0.028***	–0.031***	–0.032***	–0.026***	–0.029***	–0.029***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
N	45,294	45,294	45,294	45,294	45,263	45,215
Panel D: Non-agricultural household work						
	–0.010	–0.008	–0.007	–0.011	–0.009	–0.008
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
N	45,294	45,294	45,294	45,294	45,263	45,215
Panel E: Log monthly income						
	0.092***	0.088***	0.087***	0.102***	0.096***	0.091***
	(0.019)	(0.018)	(0.018)	(0.019)	(0.019)	(0.021)
N	17,591	17,591	17,591	17,574	17,145	16,788
Additional controls						
Province FE and year FE	✓					
Province × year FE		✓	✓			
Province × age group FE			✓			
District FE				✓	✓	✓
District × year FE					✓	✓
District × age group FE						✓

Note: this table reports the DiD estimate for all individuals aged 25–44, in which men are the control group. The interaction term being reported here is *Post-2013 × Men*. Standard errors are clustered at the commune–year level and reported in parentheses and *p*-values are reported in brackets; sampling weights are applied. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, marital status, and year-specific cohort linear trends. The sample includes all non-college-educated individuals aged 24–44 in the VHLSS 2010–18 sample.
Source: authors' calculations based on the VHLSS 2010–18.

Table A2: DiD estimates for effects on employment outcome for all women (college-educated and non-college-educated)

Specification	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Formal employment						
	0.074***	0.074***	0.073***	0.069***	0.067***	0.060***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)
N	38,704	38,704	38,704	38,704	38,675	38,486
Panel B: Not holding a job						
	–0.003	–0.004	–0.007	–0.004	–0.005	0.003
	(0.012)	(0.012)	(0.012)	(0.012)	(0.013)	(0.013)
N	38,704	38,704	38,704	38,704	38,675	38,486
Panel C: Agricultural household workers (unpaid)						
	–0.049***	–0.052***	–0.054***	–0.043**	–0.041**	–0.055***
	(0.017)	(0.017)	(0.017)	(0.017)	(0.017)	(0.018)
N	38,704	38,704	38,704	38,704	38,675	38,486
Panel D: Independent production/household business workers (unpaid)						
	–0.020	–0.019	–0.013	–0.024	–0.021	–0.009
	(0.018)	(0.018)	(0.018)	(0.017)	(0.018)	(0.019)
N	38,704	38,704	38,704	38,704	38,675	38,486
Panel E: Log monthly income						
	0.155***	0.165***	0.157***	0.176***	0.137***	0.126**
	(0.050)	(0.050)	(0.050)	(0.047)	(0.052)	(0.063)
N	12,657	12,657	12,657	12,636	12,153	11,243
Additional controls						
Province FE and year FE	✓					
Province × year FE		✓	✓			
Province × age group FE			✓			
District FE				✓	✓	✓
District × year FE					✓	✓
District × age group FE						✓

Note: this table reports the results from estimating the main DiD model. The interaction term being reported here is *Post-2013* × *Age 25–44*. Standard errors are clustered at the commune–year level and reported in parentheses and *p*-values are reported in brackets; sampling weights are applied. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, marital status, and year-specific cohort linear trends. The sample includes non-college-educated and college-educated women aged 25–54 in the VHLSS 2010–18 sample.
Source: authors' calculations based on the VHLSS 2010–18.