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Maternity benefits mandate and women's choice of work in Vietnam ☆

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ABSTRACT

Despite a sizable literature on the labor market effects of maternity leave regulation 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 decision to work in the informal or formal sector in Vietnam. We use a difference-in-differences approach to evaluate the 2012 Amendments to the Vietnam Labor Law, which imposes a longer maternity leave requirement than before. We find that the law increases formal employment and decreases unpaid work among women. This is driven by women switching from agricultural household work to employment in the private formal sector, especially in the manufacturing industry and among the middle-skilled occupations such as plant and machine workers, craft and related workers, as well as clerks.

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1. Introduction

Labor 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 (Pettit & Hook, 2005; Almeida & Carneiro, 2012). 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, Vargas, & Bustamante, 2019).

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 labor regulations on the informal sector of the labor market [e.g.] (Almeida & Carneiro, 2012; Freeman, 2010), these studies do not focus on maternity benefits regulations or on women's labor market outcomes. Similarly, although there is a sizable literature on how mother's labor 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. Most of the existing studies on the employment effects of maternity benefits regulation also focus on developed countries, where the labor markets can be very different from those in developing countries. To the best of our knowledge, only a few studies examine the association between maternity leave policies and female labor force participation (Besamusca, Tijdens, Keune, & Steinmetz, 2015; Amin, Islam, & Sakhonchik, 2016; Uribe et al., 2019; Amin & Islam, 2019; Erkmen, Panovska, & Anil Taş, 2021), and only Uribe et al. (2019) examine changes in terms of informal and formal employment among women due to changes in maternity leave regulation.

In this paper we assess these important questions in the context of the labor market of Viet Nam during the period between 2010 and 2018. The 2012 Amendment to the Viet Nam Labor 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 labor market outcomes in developing countries. Viet Nam is particularly interesting to examine because it is known for having a relatively high female labor force participation rate compared to other countries at the same income level (Klasen, Le, Pieters, & Santos Silva,

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

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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 labor force, such as maternity leave requirements [e.g.] (Gaddis & Klasen, 2014; Klasen, 2019; Klasen et al., 2020).

We examine whether the maternity leave extension of the 2012 Amendment has encouraged female workers to transition from informal work to formal employment. The basic assumption that we make in this study is that female workers who are more likely to give birth would be more responsive to the incentive created by the maternity leave extension. We use two difference-in-differences (DiD) designs based on this assumption to identify the effect of the maternity leave extension. In the first approach, we compare the labor market outcomes of women of childbearing age with women beyond childbearing age both before and after the new law came into effect. In the second approach, we compare the labor market outcomes of women with different birth rates also before and after the law to quantify such effects. In other words, we leverage the variation in childbearing age in the first DiD approach and the variation in expected birth rates in the second DiD approach. These two different designs allow us to ensure that we are identifying the effect of the maternity leave extension, instead of any other component of the 2012 Amendment. 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.

We find that the maternity leave extension of the new law increases formal employment of women of childbearing age by 2.7 percentage points, and decreases agricultural household work by 3.2 percentage points. The increase in formal employment mostly happens in the private formal sector, especially in the manufacturing industry and among middle-skilled occupations such as plant and machine operators as well as craft workers.

The contributions of this study are threefold. First, despite an extensive body of studies on the effects of maternity leave regulations on women in developed countries [e.g.] (Ruhm, 1998; Akgunduz & Plantenga, 2013; Rossin-Slater, 2017; Hook & Paek, 2020), few studies explore the labor market effects on women in developing countries. Specifically, Besamusca et al. (2015), Amin et al. (2016), Amin and Islam (2019), and Erkmen et al. (2021) document that countries with paid and longer maternity leave regulation are associated with more women choosing to work. 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 using the same empirical strategy as Uribe et al. (2019). This difference suggests that maternity benefits regulations may have very different effects in different labor market settings.[2]

Second, our study complements a small but growing literature on female labor force participation in Viet Nam (Kreibaum & Klasen, 2015; Klasen, 2019; Dang, Hiraga, & Nguyen, 2019; Klasen et al., 2020; Feeny, Mishra, Trinh, Ye, & Zhu, 2021). While the existing studies already explore what might have led to Viet Nam's relatively high female labor force participation, only a few studies examine the causal factors of female labor force participation and especially female formal employment.[3] Our study contributes to this literature by providing direct evidence for the role of maternity leave policies in female workers' decisions on where to work.

Our study is also related to a different literature on issues about labor rights and labor reforms in Vietnam, which mostly comprises of qualitative studies. Tran (2013) reviews and discusses extensively the history of labor organizing to fight for better working conditions, better pay, and better rights as well as the cultural ties among labor right activists, which are among the main drivers of reforms around labor rights (Tran, 2013), including the 2012 Amendments that we study. Evans (2020) and Evans (2021) further argue that labor reforms in Vietnam are driven by domestic activism alone but are assisted by the interests of the communist government around labor right issues.[4] These labor reform studies highlight the significance of labor rights such as maternity leave from the cultural and political lens of workers and policymakers that are not often discussed in typical economics studies.

The rest of the paper is organized as follows. Section 2 provides an overview of the 2012 Amendment.[5] Section 3 reports different data sources used in the study as well as how we define different labor market outcomes for our analyses. Section 4 provides a discussion of our empirical strategies. In Section 5 we report and discuss the findings, and Section 6 concludes the paper.

2. Background

The main regulatory framework for the labor market, especially for formally registered businesses, is the Viet Nam Labor Law, which was first written in 1994 (World Bank, 1995) and amended in 2002, 2012, and 2019. The original Viet Nam Labor 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 per working day.

The 2012 Amendment to the Viet Nam Labor Law (Law No. 10/2012/QH13) 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. This provision is considered by labor rights scholars as weakening the workers' role in resolving strikes (Tran, 2013). Second, the new law prohibits workers from working more than eight hours per day and 48 h per week; the law also limits overtime work to 30 h per month and 200 h 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 labor productivity and quality of

[2] Our study is also broadly related to a large literature on family, fertility and women's work in developing countries, e.g. Bloom, Canning, Fink, and Finlay (2009), Krafft (2020), Selwaness and Krafft (2021), and Finlay (2021).

[3] Dang et al. (2019) find that childcare increases the probability of formal employment among mothers.

[4] The government has an interest in labor repression to maintain the low labor-cost advantage of the Vietnamese economy to promote exports. At the same time, the government recognizes that trade agreements, e.g., the Trans-Pacific Partnership (TPP) and EU Trade Agreement, can strengthen the political regime, but they also require more progressive reforms around the labor right issues (Evans, 2021). The unauthorized strikes by workers also triggered the government's fear for the regime's legitimacy, putting more pressure on them to pursue pro-labor reforms (Evans, 2020).

[5] See Appendix A.1 for an overview of the recent trends in labor market of Vietnam.

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.[6]

Most importantly, the new law extends the mandated paid maternity for female workers up to six months during the prenatal and postnatal periods. In other words, the new law increases the required maternity leave by two months. The salary during the maternity leave is completely paid for by the social insurance, and employers do not have to pay for this expense. Employers do have to contribute 18% of monthly payroll toward social insurance, while workers contribute 8% of their monthly salary.

3. 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. 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 labor 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.[7]

Since we aim to investigate the effect on individual decisions to work, we want to exclude who are still in school or in retirement. This is because those who are in school also have to face the decision whether to continue to invest in their education, while those who are of retirement age have to consider whether to retire (or continue to work in the informal sector); therefore, their decisions to work are likely very different from those of the working age population.

Given these considerations, we choose to restrict the study population to individuals aged 25–54. The official retirement ages in Viet Nam are 60 for men and 55 for women, while the official age of completing schooling, including university education, is 21. However, many individuals remain in school until later. For example, 23.5% of individuals aged 22 and 12.8% of those aged 23 are still school; by the age of 25, however, only 5.4% remain in school (according to the VHLSS 2010–2018). This means that almost everyone in the 25–54 age range can choose between formal or informal employment, or not working at all, for reasons unrelated to schooling and retirement. These age cutoffs also allow us to split the sample into 6 age groups of 5 years: 25–29, 30–34, 35–39, 40–44, 45–49, and 50–54 for further analysis by age group. 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 (Fig. 1) and because our sample covers women aged 25–54.

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

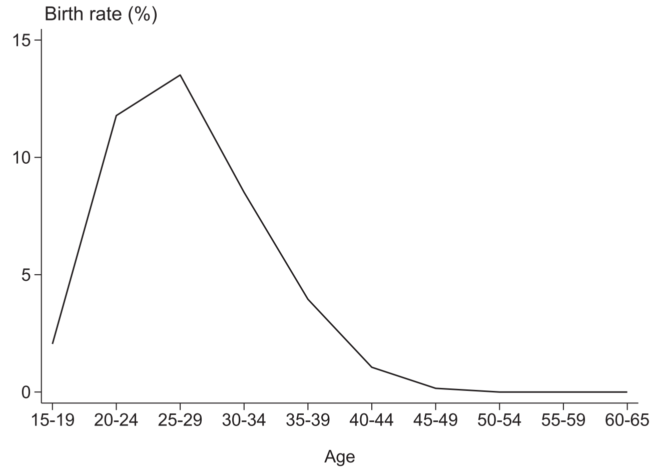

Fig. 1. Birth rates by age. (Source: authors' calculations using the most recent live birth information from the 2009 Population and Housing Census.)

- 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 for not working for men is disability, while the main reason for women is housework (Fig. A1).
- 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 formal sectors (domestic and foreign firms).
- Informal wage work: Holding a job that comes with a wage or salary, but does not provide social insurance. In other words, these are wage workers who do not receive any social insurance.
- Agricultural and non-agricultural household work: workers who contribute to a family business in agriculture or non-farm sectors. These family-contributing workers are also considered as informal workers (ILO, 2016).[8]

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 Fig. 2 we also plot these labor market outcomes by age groups, gender, and year. In Fig. 2(a) we note that the share of women with formal employment in the 25–44 age group is generally higher than the share of men in the same age group, and also higher than the shares of men and women in the older age group. During 2010–2014, formal employment of women and men age 25–44 follow an upward trend, while formal employment of both gender age 45–54 remain stable. This is consistent with what McCaig Pavcnik (2017) and Liu, Barrett, Pham, and Violette (2020) observe: as more formal business open over time, demands for younger workers in the formal sector also increase as they tend to have better education.

After 2014, formal employment among women between age 25–44 rises more sharply. At the same time, formal employment among men age 25–44 continues to rise at the same pace as before. In contrast, the formal employment among women age 45–54 remains stable throughout the study period, and that among men age 45–54 appears to decrease after 2014. In other words, there is an increase in formal employment among women of childbear-

[6] It is important to note that this provision only put forward the formal definition for minimum wage. The actual regulations on the minimum wage levels were established in 2011 by the Decree 70/2011/ND-CP (Tran, 2013).

[7] Although the VHLSS data has been collected since 2002, only the 2010–2018 surveys ask for information about social insurance, which allows us to construct the main outcome variable for formal employment, as described below.

[8] Since the VHLSS survey does not ask about whether an individual is a self-employed worker, we do not include this labor market outcome. If the self-employed workers earn wage, then they would be under the "Informal wage work" category, but if they work unpaid for their household, then they would be under the "Agricultural or non-agricultural household work" category.

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)
labor 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.

ing age following the 2012 Amendment that we do not observe in men age 25–44 or women age 45–54. We also observe a decrease in formal employment among older men following the new law. The age and gender-specific changes in formal employment suggest that these are more related to the maternity leave extension of the Amendment, instead of the other provisions, which are likely gender-neutral.

The shares of women of both age groups who have informal wage employment increases steadily during the study period. In contrast, the share of informal wage employment among men aged 25–44 stays relatively stable, and that share among men aged 45–54 fluctuates during 2010–2018. The trends in not working are similar across genders and age groups and are also relatively stable over time, suggesting that the new law does not have any effect on this outcome. Agricultural household work, on the other hand, experiences a sharp drop in 2016 across all age groups and genders. This decrease is part of a longer-term decline in agricultural employment in Vietnam. Indeed, McCaig Pavcnik (2017) and Liu et al. (2020) find that agricultural employment has been steadily declining since the early 2000s as trade liberalization and enterprise reforms increase the number of employment opportunities outside the agricultural sector. This structural transformation process, however, has similar effects on both genders in terms of agricultural employment (Liu et al., 2020), which is consistent with what we observe in Fig. 2(d).

We also observe an opposite trend for the non-agricultural household work across all age groups and genders: non-agricultural household work decreases from 2010 to 2014 and reverses, increasing in 2016–18. For both agricultural and non-agricultural household work, these trends represent economy-wide changes as they apply to all age groups and genders.

The log monthly wage also increases for all groups, although it increases faster for women aged 25–44 throughout the study period. Indeed, the wage gap between women of the two age groups was very small in 2010, but has been widening since 2012, i.e., even before the new law was implemented. In contrast, the wage gap between the two groups of men remains stable until it closes in 2018.

These descriptive comparisons illustrate the increase in formal employment among women of childbearing age relative to the other age and gender groups following the implementation of the new law. They also show the overall shift away from the agricultural sector as documented in the previous literature (McCaig Pavcnik, 2017; Liu et al., 2020). In the next section, we present our empirical strategies to identify the effect of the maternity leave extension while accounting for these structural changes of the economy.

4. Empirical strategy

Identifying the effect of the maternity leave extension is not straightforward because the extension was combined with several other changes in the 2012 Amendment, which means we may pick up the effects of those changes as well. Our empirical strategies, therefore, are built on the assumption that female workers who are more likely to give birth are also more likely to respond to the incentive of the maternity leave extension. In other words, we assume that the other components of the Amendment may affect the labor market outcomes of everyone, but the maternity leave extension would differentially affect women who are more likely to give birth.

We consider two difference-in-differences research designs to identify the effect of the extension based on this premise. For the first design, we compare employment outcomes between women of childbearing age and women older than childbearing age before and after the law came into effect. For the second design, we compare employment outcomes between women with different expected birth rates, also before and after the maternity leave extension.

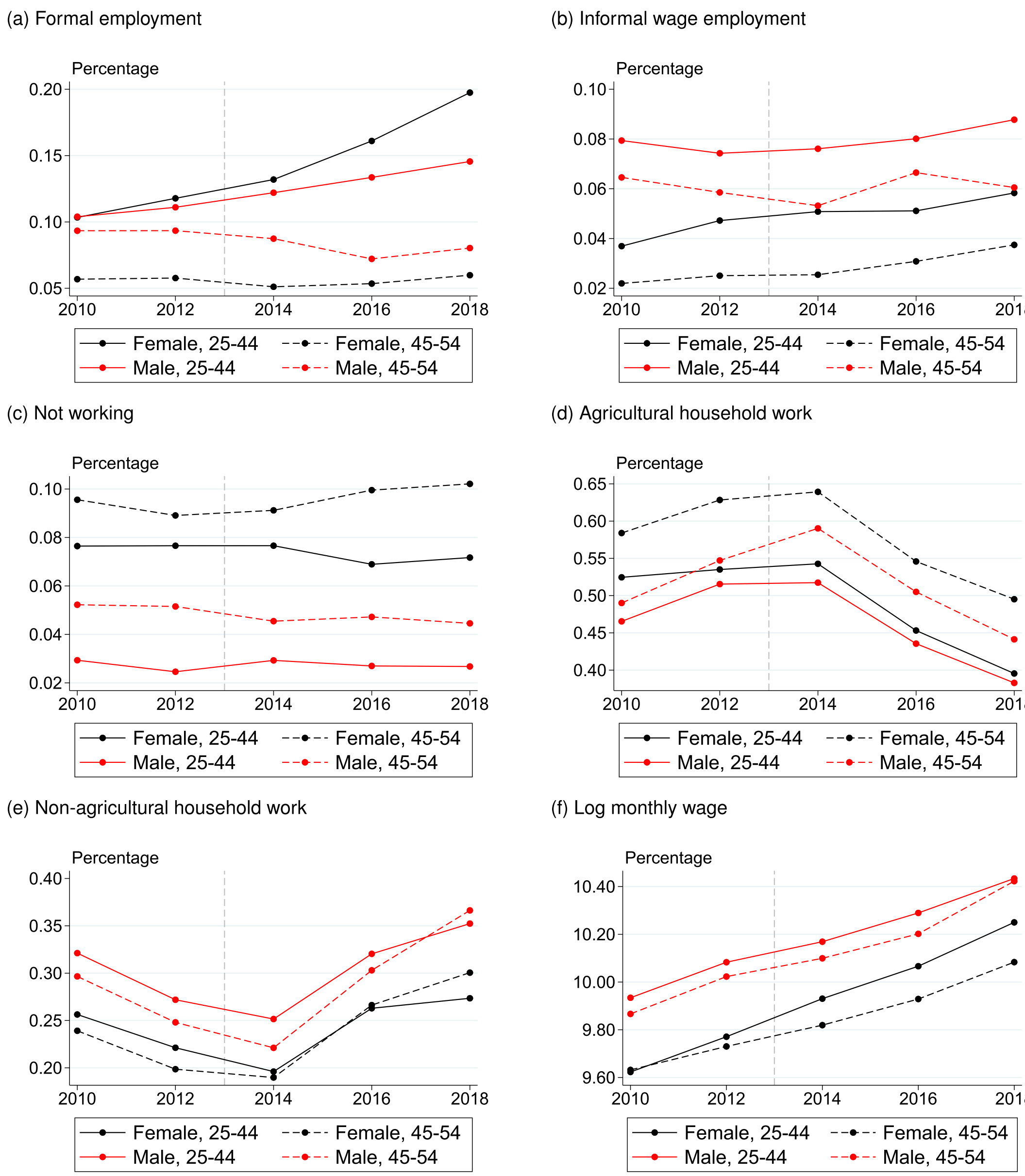

Fig. 2. Employment outcome for women aged 25–54 by year. (Source: authors' calculation based on the VHLSS 2010–18.).

4.1. First research design: comparison by childbearing age

First, we observe in Fig. 1 that the fertility rates among the 25–44 age group are positive and substantially higher than those among the 45–54 age groups, which are effectively zero. Therefore, this group is unlikely to respond to the maternity leave extension, allowing us to exploit it as the control group.

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

$$Y_{it} = \beta + \delta.(Age25 - -44_{it} \times Post_t) + \mathbf{X}'_{\mathbf{it}}\mathbf{\Gamma} + \sum_s \kappa_s(Year_s \times Age_{it}) + \theta_{p,t} + \eta_{p,g} + \varepsilon_{it} \tag{1}$$

where Y_{it} denotes the employment outcomes of woman i in year t, $Age25 - 44_{it}$ indicates whether the woman 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 $\mathbf{X_{it}}$; 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, which will be used for the age group fixed effects. 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. The term $Year_s \times Age_{it}$ are year dummy variables times age; this allows us to account for changes in the effect of age on labor market outcomes over time that is unrelated to the maternity leave change. In all models we cluster standard errors at the commune–year level to account for the survey's sampling design.

We also extend the main DiD model by estimating the dynamic treatment effects in the following event-study specification:

$$Y_{it} = \beta + \sum_{s \neq 2012} \delta_s.(Age25 - -44_{it} \times Year_s) + \mathbf{X}'_{\mathbf{it}}\mathbf{\Gamma} + \sum_s \kappa_s(Year_s \times Age_{it}) + \theta_{p,t} + \eta_{p,g} + \varepsilon_{it} \tag{2}$$

where $s = \{2010, 2014, 2016, 2018\}$ and 2012 is the reference year. We use the same set of controls, including province–year fixed effects, province–age group fixed effects, and year dummy variables times age. 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.

Based on the different birth rates across age groups, we would expect the effects to be larger among younger age groups if they respond to the incentive from the maternity leave extension. Therefore, we can assess the validity of our approach by allowing the treatment effects to vary by age group. Specifically, we estimate the treatment effects by age group: 25–29, 30–34, 35–39, and 40–44, while the control group is women aged 45–54:

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

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. Importantly, we drop the control variables for year dummy variables times age because there would not be enough variation to differentiate the treatment effects by age groups.

4.2. Threats to validity and robustness checks

The underlying identifying assumption of the first research design is that the employment outcomes of the treatment group and the control group would have followed the same trends if the required maternity leave length had not been changed.[9] This parallel trends assumption would be violated if other factors unrelated to the maternity leave extension also affect employment outcomes differently for the two groups over time. In this section, we discuss different potential threats to this identifying assumption and how we attempt to address them.

The most likely confounders of the maternity leave requirement extension are the other components of the 2012 Amendment. As discussed in Section 2, the new law also defines strike conditions, limits the working hour and overtime work, extends a holiday, and provides a formal definition for the minimum wage. Because our DiD design identifies the effect of the maternity leave extension from variation in whether a women is of childbearing age, our estimation would be biased if these unrelated provisions also affect the two age groups differently.

Another potential confounder is the rapid structural transformation that Vietnam has undergone since 2000, as workers moved out of the agricultural sector and into higher productivity sectors such as manufacturing and services (McCaig Pavcnik, 2017; Liu et al., 2020). 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 Pavcnik, 2017). 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 labor-saving inputs, allow younger workers who are increasingly more educated to transition out of agriculture and into non-farm sectors (Liu et al., 2020). Structural transformation, therefore, can bias upward our estimation as it can increase the number of formal jobs available in the local labor market every year.[10]

Lastly, our results can also be biased by the minimum wage reform in 2011 (Decree 70/2011/ND-CP) which took effect in 2012. Prior to 2011, the government maintained two systems of minimum wage: one was for all domestic employers (including the government), and the other one was for the foreign direct investment sector. As a requirement to join the World Trade Organization (WTO) in 2007, the government equalized the minimum wage levels of the two sectors in 2011 (Tran, 2013; Nguyen, 2017). It is important to note that the minimum wage levels have been increased every year and across every geographic region (and sector), so this is not a shock on the level of the minimum wage, but rather a shock on the difference in minimum wage levels between two sectors.

Both the structural transformation and the minimum wage reform are shocks at the local labor market level. On one hand, the rapid structural transformation means there would be more job opportunities in the non-farm sectors; therefore, it is a positive labor demand shock (McCaig Pavcnik, 2017; Liu et al., 2020). On the other hand, a higher minimum wage in a competitive labor market means that the cost of hiring would increases, so it is a negative labor demand shock (Nguyen, 2017). Our empirical strategy accounts for both of these concerns, to some extent, by controlling for province–year fixed effects in all of models to account for any changes in the local labor market in any given year. We also control for province–age-group fixed effects and year fixed effects times age[11] to absorb any unobserved heterogeneity across age groups over time and across different provinces. Lastly, we exclude college-educated women and control for educational attainment in Eq. 1, since younger cohorts tend to have higher educational attainment and, hence, are more likely to switch to the non-farm sectors.

It is important to note that all of the potential confounding factors discussed above would likely affect both male and female workers. Specifically, the other provisions of the 2012 Amendment to the Labor Code are gender neutral, so we would not expect them to affect the two genders differently. Liu et al. (2020) find that the gender composition of the agricultural labor force was stable during the 1992–2016 period, which suggests that the structural transformation process may not affect men and women differently. Del Carpio, Nguyen, Nguyen, and and Wang (2013) find that the shares of male and female workers who have wages below the minimum wage level are relatively similar (6.1% and 6.5%, respectively). Therefore, we can assess the extent to which our main findings are confounded 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, this estimation should not be affected by any differential effects of the structural transformation process or the other provisions of the 2012 Amendment by age. In other words, 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 can be more assured that our main findings are not driven by the potential confounding factors discussed above.

Alternatively, we can estimate a triple-differences regression in which women of childbearing age are compared to men of the same age and both men and women of age 45–54. The triple-differences approach allows us to further control for labor market

[9] We also assume that the maternity leave extension did not affect our control group, which is women age 45–54. The trends in Fig. 2 suggests that this is not a problem in our data.

[10] Relatedly, the Vietnamese government also passed the 2013 Land Law, which extends the land tenure for farmers who grow annual crop and is found to boost agricultural investment (Bellemare, Chua, Santamaria, & Vu, 2020). The increase in land tenure security may also induce workers to pursue off-farm employment. However, we do not expect this law to have different effects on the childbearing and non-childbearing age groups.

[11] Note that we control for year dummy times age for all specifications except for the specification for treatment effects by age group in Eq. 3 as explained above.

trends across different age groups. Formally, we estimate the following model:

$$Y_{it} = \beta + \delta^{DDD}.(Age25 - -44_{it} \times Post_t \times female_i) + \mathbf{X}'_{\mathbf{it}}\boldsymbol{\Gamma} + \theta_{p,t,f} + \eta_{p,g,f} + \zeta_{p,t,g} + \varepsilon_{it} \tag{4}$$

where $\theta_{p,t,f}$ represents the province–year–female fixed effects, $\eta_{p,g,f}$ represents the province–age group–female fixed effects, and $\zeta_{p,t,g}$ represents the province–year–age group fixed effects. The last term absorbs any differential changes in labor market outcome across different age groups unrelated to gender. This term, therefore, will account for the confounding effects of the structural transformation as well as the other provisions of the 2012 Amendment as they likely affect both genders equally. The triple-difference specification can be easily extended to allow the effects to vary by year or age group as done for the difference-in-differences model.

4.3. Second research design: comparison by expected birth rates

The main DiD approach identifies the effect of the maternity leave extension using variation in childbearing age. These estimates would be biased if other factors unrelated to the maternity leave extension also have differential effects across age groups, as we discussed previously. In the previous section, we extended this design by switching to a different control group and employing a triple-differences model with extensive controls for unrelated labor market shocks by age, gender, and geography to address that concern. However, this approach still relies on the assumption that the effect identified by variation in demographic characteristics related to giving birth such as childbearing age and female would reflect the effect of the maternity leave extension and not other factors.

In this section, we consider an alternative research design to identify the effect of the extension using variation in expected birth rates across districts and age groups. That is, we would expect that women with higher expected birth rates would respond more strongly to the incentive of the maternity leave extension than those with lower expected birth rates. Therefore, we can directly compare the groups with different birth rates over time to identify the effects of the leave extension instead of comparing between childbearing and non-childbearing age or between men and women. We restrict our analysis to the childbearing-age women sample to ensure that the results from this design are driven entirely by differences in expected birth rates.

Specifically, we use the 2009 Population and Housing Census to calculate district-level age-specific fertility rates, defined as the number of children born to females of a specific age group in each district, for the women of childbearing age sample. We split the birth rates into five equal-sized bins, then estimate the following model:

$$Y_{it} = \beta + \sum_b \delta_b (W^b_{g,d} \times Post_t) + \theta_{p,t} + \eta_{p,g} + \gamma_{t,g} + \mathbf{X}'_{\mathbf{it}}\boldsymbol{\Gamma} + \varepsilon_{it} \tag{5}$$

where $W^b_{g,d}$ indicates whether the birth rate of the respective district–age-group d and g falls into bin $b, \theta_{p,t}$ are province–year fixed effects, $\eta_{p,g}$ are province–age-group fixed effects, and $\gamma_{t,y}$ are year–age-group fixed effects. The reference group is the group with the lowest birth rates. This model quantifies the effects by birth rates while accounting for any unobserved confounders across age groups within provinces and any secular shocks across provinces and age groups.

Comparing different birth rate groups allows us to assess whether the main results are driven by responses to the maternity leave extension of the 2012 Amendment. If the effects are larger among groups with higher birth rates and smaller among groups with smaller birth rates, it would be consistent with the interpretation that the employment outcomes are responding to the maternity leave extension and not to other unrelated components of the new law. In contrast, if the effects' magnitudes do not correspond to the birth rates, then it is likely that the employment outcomes are driven by other factors.

A useful placebo test for the validity of this approach, therefore, is to estimate the same model on the sample of only men between age 25–54. That is, we merge the expected birth rates of women by district and age groups with a male sample and re-estimate Eq. 5. Since men do not respond to the incentives from the maternity leave extension, we would expect the effects to be small and statistically insignificant.

5. Results

5.1. Main findings

For the main findings, we present three sets of results in Table 2. Columns (1) to (3) report the results from estimating different variants of Eq. 1—that is, a DiD model with women aged 25–44 as the treatment group and women aged 45–54 as the control group. Columns (4) to (6) report the results from estimating the alternative DiD model where men aged 25–44 are the control group. Lastly, column (7) reports the results from estimating the triple-differences model in Eq. 4.

The results for different outcomes are presented in panels A–E of Table 2. 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 informal wage employment, defined as holding a job that provides wages/salary but not social insurance. In Panel C, the outcome is not holding a job. In panels C and D, the outcome variables are whether individuals are working in an agricultural or non-agricultural activity in their own household, respectively. In panel E, the outcome is the log monthly wage 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.

For the main DiD model comparing women of childbearing age (as the treatment group) to older women (as the control group), we consider three different specifications that control for different levels of fixed effects, from least to most parsimonious. All specifications control for urban, ethnicity, household size, number of children under 10 years old in household, educational attainment, marital status, and year fixed effects times age. The baseline specification in column (1) controls only for age group fixed effects and year fixed effects. This accounts for unobserved heterogeneity by age group as well as labor market shocks at the national level. The specification in column (2) instead controls for province–age group and province–year fixed effects; the first fixed effects absorb unobserved heterogeneity by age group and province, while the second fixed effects absorb local labor market shocks at the province level. The specification in column (3), on the other hand, controls for district–age group and district–year fixed effects. This specification accounts for heterogeneity and shocks at the district level, which is more granular than controlling at the province level. The specifications in columns (2) and (3) account for the fact that the local labor market conditions may induce younger workers to move out of agriculture and into formal employment as driven by structural transformation (McCaig Pavcnik, 2017).

To check whether the results from the main DiD model are still confounded by local labor market conditions or other provisions of the 2012 Amendment that are gender-neutral, we contrast the main DiD model with an alternative DiD model comparing women of childbearing age (as the treatment group) with men of the same

Table 2
Difference-in-differences and triple-differences estimates for the effect of the maternity leave extension on labor market outcomes.

	Main DiD model			Alt. DiD model			DDD
	Women age 45–54 as control			Men age 25–44 as control			model
Specification	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Formal employment							
	0.039***	0.039***	0.031**	0.033***	0.034***	0.029***	0.027**
	(0.013)	(0.013)	(0.014)	(0.007)	(0.006)	(0.006)	(0.012)
N	34701	34701	34422	45294	45294	45215	67739
Panel B: Informal wage work							
	0.011	0.011	0.010	0.007	0.008	0.010**	−0.004
	(0.009)	(0.009)	(0.010)	(0.005)	(0.005)	(0.005)	(0.010)
N	34600	34600	34321	45155	45155	45075	67509
Panel C: Not holding a job							
	−0.010	−0.006	0.002	−0.004	−0.001	0.000	−0.009
	(0.013)	(0.013)	(0.014)	(0.005)	(0.005)	(0.005)	(0.010)
N	34701	34701	34422	45294	45294	45215	67739
Panel E: Agricultural household workers (unpaid)							
	−0.032	−0.043**	−0.046**	−0.025***	−0.032***	−0.029***	−0.032**
	(0.020)	(0.019)	(0.019)	(0.008)	(0.008)	(0.008)	(0.016)
N	34701	34701	34422	45294	45294	45215	67739
Panel F: Independent production/business household workers (unpaid)							
	−0.005	0.002	0.007	−0.010	−0.007	−0.008	0.019
	(0.020)	(0.019)	(0.021)	(0.008)	(0.008)	(0.008)	(0.017)
N	34701	34701	34422	45294	45294	45215	67739
Panel G: Log monthly income							
	0.108*	0.072	−0.024	0.084***	0.086***	0.090***	0.019
	(0.064)	(0.060)	(0.085)	(0.020)	(0.018)	(0.021)	(0.069)
N	9129	9129	7507	17591	17591	16788	20684
Additional controls							
Age group FE	✓			✓			
Year FE	✓			✓			
Female FE				✓			
Province × Year FE		✓			✓		
Province × Age group FE		✓			✓		
Province × Female					✓		
District × Year FE			✓			✓	
District × Age group FE			✓			✓	
District × Female						✓	
District × Female × Year FE							✓
District × Age group FE × Female							✓
Province × Age group FE × Year FE							✓

This table reports the difference-in-differences estimate for the effects of the maternal leave extension on labour market outcomes. In columns 1–3, we use women of age 45–54 as the control group, and the results are reported for the interaction term Post-2013 × Age 25–44. The sample for these estimates are non-college females aged 25–54. In column 4–6, we use men of age 25–44 as the control group, and the results are reported for the interaction term Post-2013 × Female. The sample for these estimates includes all non-college individuals aged 24–44. Standard errors are clustered at the commune-year level and reported in parentheses and p-value is reported in brackets; sampling weights are applied. All models control for year FE, age group FE, urban, ethnicity, household size, number of children age under 10 in household, educational attainment, marrital status, and year-specific cohort linear trends. Data is drawn from the VHLSS 2010–2018.

age (as the control group). We estimate three specifications that are analogous to those in the first three columns. Specifically, the specification in column (4) controls for female dummy, year and age group fixed effects. This baseline specification accounts for time-invariant heterogeneity by gender and age group, as well as shocks at the national level. For column (5), we estimate a specification that controls for province–female, province–age group, and female–age group fixed effects. As in the specification in column (2), these controls account for local labor market shocks as well as gender and age differences at the province level. In column (6), we estimate a specification controlling for district–female, district–age group, and district–year fixed effects, which is analogous to that in column (3). Lastly, in column (7), we present the results from estimating the triple-differences in Eq. 4. As described in Eq. 4, we control for province–year–age group, district–female–year, and district–age group–female fixed effects. This model accounts for the gender and age group-specific shocks in any given year.

The estimations in Panel A Table 2 indicates that the maternity leave extension has a positive and statistically significant effect on formal employment of women of childbearing age. In the base specification, the estimate is 0.039. When accounting for age group heterogeneity and local labor market conditions at the province level, the estimate is still 0.039. However, the estimate is reduced to 0.031 when controlling for district–age group and district–year fixed effects. These results indicate that the baseline estimate is partly driven upward by the local labor market shocks at the district level.[12]

Turning to the estimates in columns (4) to (6), the estimate in the baseline specification of the alternative DiD model is 0.033, which is slightly lower than that of the baseline specification of the main DiD model in column (1). Since the two specifications are analogous except for the control group, the difference between the results in column (1) and (4) suggests that the baseline estimate from the main DiD model may be biased to some extent, although the difference is quite small. When controlling for pro-

[12] Note that the sample size is reduced slightly from 34,701 to 34,422 in the third specification. As we control for more granular geographic level, some district–age group cells have one observation and, hence, are dropped from the regression as these are known to introduce bias in high-dimensional fixed effects models (Correia, 2015).

vince–year, province–female, and province–age group fixed effects, the estimate changes only slightly to 0.034. However, when replacing the province-level controls with district-level controls, the estimate is again reduced to 0.029. This suggests that the local labor market conditions also influence the estimates in the alternative DiD model. However, the most parsimonious specifications from the two DiD models yield very similar results, 0.031 and 0.029, which means that the extensive set of fixed effect controls successfully absorb most of the biases in both models. Lastly, the triple-differences estimate is 0.027, which is slightly smaller than the results from both of the DiD models, but it remains positive and significant. This model further accounts for different changes of formal employment across genders and age groups, so the estimate is slightly more conservative than those of both DiD models.

The fact that the estimates are considerably similar to each other and remain statistically significant when using different control groups, models, or controlling for different levels of fixed effects, is reassuring evidence that any biases from local labor market conditions or factors unrelated to the maternity leave extension are relatively small. That is, under the assumption that the other provisions of the 2012 Amendments and the structural transformation are likely to affect both genders, the alternative DiD model using men age 25–44 would yield unbiased estimates. Given that both DiD models yield very similar results, the main DiD estimates are also likely to be unbiased. Similarly, the triple-differences model is likely to be unbiased because it accounts for both gender and age group-specific shocks. Since the estimates from both of the DiD models and the triple-differences model are very close to each other, the biases are, again, likely to be insignificant.

Panels B, C, and F suggest that the maternity leave extension had no effect on informal waged employment, not working, or non-agricultural household work. The results are insignificant across the board, and the estimates are close to zero. Different DiD and triple-differences models yield the same conclusion. These results suggest that workers are not moving from these labor market options to formal employment.

In contrast, we find substantial evidence that workers move out of agriculture household work in response to the maternity leave extension. In the baseline estimate of the main DiD model, the treatment effect is −0.032 but is not significant. However, when we control for province–age group and province–year fixed effects, the estimate rises to −0.043 and becomes significant. Further controlling for local labor market conditions also increases the estimated treatment effect to −0.046. The alternative DiD model yields estimates between −0.025 in the base specification to −0.029 in the least parsimonious specification. These results again suggest that local labor market shocks may bias the DiD estimates downward since the treatment effect becomes larger as we control for fixed effects at a more granular geographic level. Lastly, the triple-differences estimate is −0.032, which is in between the estimates of the two DiD models. These robust estimates suggest that the maternity leave extension decreases employment in the agricultural household sector. The magnitude of this negative effect is comparable to the magnitude of the positive effect on formal employment, implying that workers mainly move from agricultural household work to formal employment.

Unlike the results for employment, the evidence for the effect on log monthly income is mixed across different models. Without controlling for local labor market shocks, the estimate from the main DiD model is 0.108. When adjusting for shocks at the province level, the model yields a smaller and insignificant estimate, which is 0.072. When we adjust for shocks at the district level, the estimate becomes negative and still insignificant. These results indicate that local labor market shocks bias the main DiD model's estimate upwards. In contrast, the alternative DiD model yields similar estimates, which range between 0.084 to 0.090 and are all statistically significant. The triple-differences model yields an estimate of 0.019, which is also insignificant.

The considerable inconsistency between the three models warrants a further investigation. Fig. 2 shows monthly wage increases throughout the study period for all genders and age groups, although women of childbearing age appear to experience a slightly larger increase relative to other groups. This difference, however, can be driven by the local labor market conditions; as more firms enter the market over time, demands for labor also increase and so do wages. The triple-differences model addresses this problem by accounting for changes for both gender and age group over time, so the estimate is smaller and insignificant. Taken together, these results provide some evidence that the extension increased monthly wages, but the evidence is not robust across different estimation methods.

The event study estimates in Fig. 3 provide a more detailed picture of the dynamic effects of the maternity leave extension. For each outcome, we show three sets of results which correspond to the two DiD models and the triple-differences model. For the formal employment outcome, we observe that the 2010-term is statistically insignificant for all three models, allowing us to conclude that the main finding is unlikely to be driven by differential pre-treatment trends. The estimates are close to zero for the treatment effect in 2014, but become positive, larger, and significant for 2016 and 2018. In other words, the extended maternity leave increases the formal employment among women of childbearing age over time.

The evidence for the treatment effects on agricultural household work is less clear. The main DiD model using older women as the control group and the triple-differences model do not suffer from the pre-trends problem, while the alternative DiD model using men aged 25–44 as the control group appears to have negative pretrends. All three models indicate no effect in 2014. The effect become negative and significant in 2016 in the alternative DiD model. In 2018, the effects are negative and significant across all three models. These event-study estimates do confirm that the maternity leave extension decreases agricultural household work, although the effects do not become apparent until 2018.

In short, the main DiD and triple-differences results as well as the event-study estimates indicate that women of childbearing age respond to the maternity leave extension of the 2012 Amendment by switching out of agricultural household work and into formal employment. Accounting for local labor market shocks at the district level appears to account for most biases from factors unrelated to the maternity leave extension. We rule out the possibility that our findings reflect the effects of the local labor market or the other components of the 2012 Amendment, e.g., changes in the strike conditions or minimum wage definition, which are likely to be gender-neutral. By comparing women of childbearing age with older women as well as with men of the same age group, we find that our results are driven specifically by women of childbearing age instead of by younger workers relative to older workers.

Fig. 4 shows our estimates for treatment effects by age groups. We present two sets of results for each outcome from estimating a DiD model and a triple-differences model that allows treatment effects to vary by age group. For the formal employment outcome, the treatment effects from the first model are smallest among the 40–44 age group and largest among the 35–39 age group, and the estimates are all statistically significant. Although the estimate for the 40–44 age group is relatively small compared to other age groups, one may still be concerned that the effect size might be unreasonably large given that this group has a very low, albeit non-negative, birth rates (see Fig. 1). In other words, some unrelated factors might be at play that are driving these estimates.

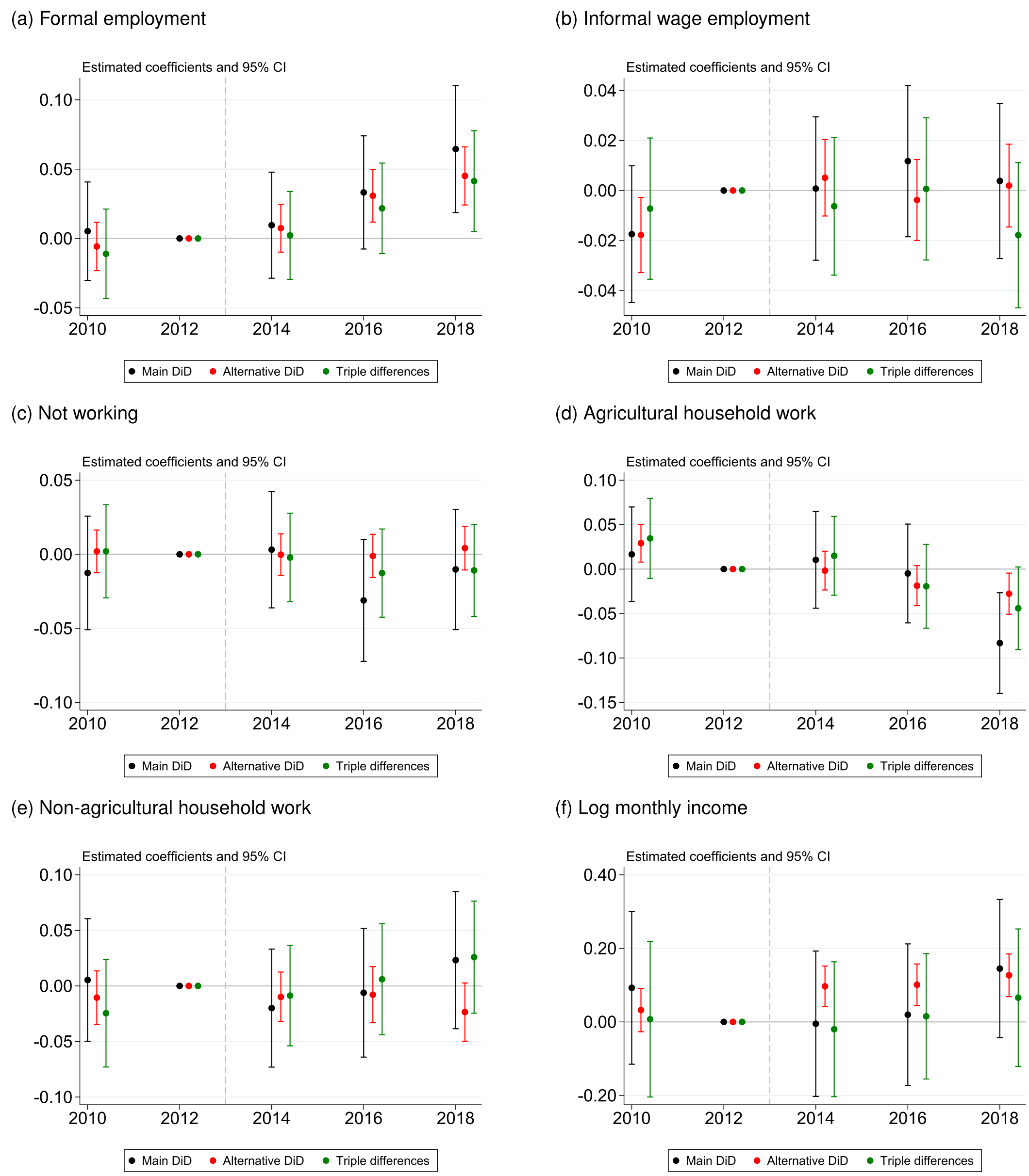

Fig. 3. Event-study estimates for effects on women's labor market outcomes. (Note: standard errors are clustered at the commune-year level. Source: authors' estimating event study specifications for the main DiD, the alternative DiD, and the triple-differences models. In the main DiD model, the control group is women aged 45–54; the model controls for district–year and district–age-group fixed effects. In the alternative DiD model, the control group is men aged 25–44; the model controls for district–year, district–female, and district–age group fixed effects. In the triple-differences model, we control for province–year–age group, district–group–female, and district–female–year fixed effects. All models control for the individual covariates described in text.)

The triple-differences model appears to address some of this concern by adjusting downward these estimates. The treatment effect for the formal employment outcome estimated from this model are smaller and the results are significant only for the 25–29 and 35–39 age groups.

The treatment effect estimates for agricultural household work are all negative, but are significant only in the main DiD model. The triple-differences estimates have similar magnitude, but wider confidence intervals. In other words, the triple-difference model yields less precise estimates, likely because of the extensive set of fixed effects as control variables. Interestingly, the main DiD model also estimates a negative and significant effect on non-agricultural household work among the 30–34 and 35–39 age group. However, these results are likely biased by other factors, which are accounted for by the triple-differences model.

5.2. Results from the second research design

As discussed in Section 3.1, the main difference-in-differences design relies on the assumption that the formal employment outcome of women of childbearing age and older women would have followed the same trends if the maternity leave requirement had not been extended. This parallel trends assumption would be violated if other factors also affect the labor market outcomes of younger women relative to older women (or workers in general).

In Eq. 5, we consider a different DiD design where we compare directly women between age 25–44 in low birth-rate groups with those in high birth-rate groups before and after the law was implemented. This allows us to identify the effect using the variation in birth rates conditional on age group–year fixed effects, which absorb the differential effects of any unrelated factors on different

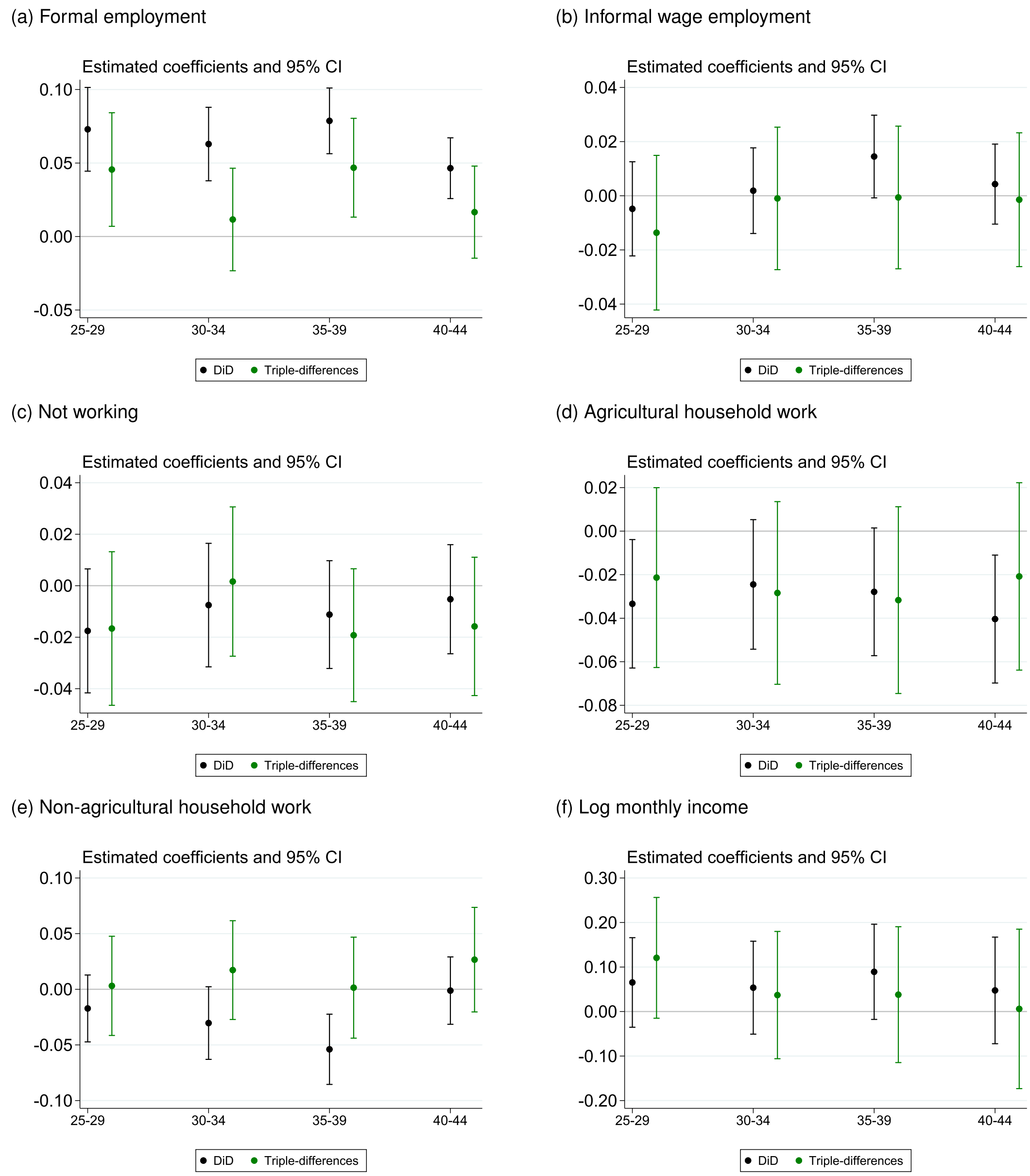

Fig. 4. DiD estimates for effects by age. (Source: authors' estimating the DiD model in Eq. 3 and the triple-differences model, in which treatment effects are allowed to vary by age group. In the main DiD model, the control group is women aged 45–54; the model controls for district–year and district–age-group fixed effects. In the triple-differences model, we control for province–year–age group, district–group–female, and district–female–year fixed effects. All models control for the individual covariates described in text.)

age groups over time. The plots in Fig. 5 summarizes the estimates for the effects on the four groups with higher birth rates relative to the group with the lowest birth rates.

Relative to the lowest birth-rate group, the treatment effects on formal employment for the other four higher birth-rate groups are positive and statistically significant. The two highest birth-rate groups also have substantially higher treatment effects than the rest, indicating that the 2012 Amendment has a larger effect among women with higher expected birth rates. This is consistent with our main findings that this is driven by the maternity leave extension.

Surprisingly, we find that the treatment effects on informal wage employment are negative and larger among the groups with higher birth rates. The treatment effects are significant for the two highest birth-rate groups. The estimates for the agricultural household work are negative but no longer significant across all birth-rate groups. These are different from the main findings, where we find that the maternity leave extension decreases agricultural household work and does not affect informal wage employment. In contrast, the estimates for not working, non-agricultural household work, and log monthly income are relatively small and insignificant, which is consistent with the results from the main DiD design.

To assess whether this specification absorbs any confounders unrelated to birth rates, we re-estimate the same model for a sample of only men between age 25 and 54 as a placebo test and present the results in Fig. A2. The estimates here represent the treatment effects on men in district–age groups with different expected birth rates. Since these birth rates should not have any effect on men (conditional on the fixed effects), we would expect all estimates to be statistically insignificant if this DiD design is valid. Indeed, we find that the treatment effects on men are mostly

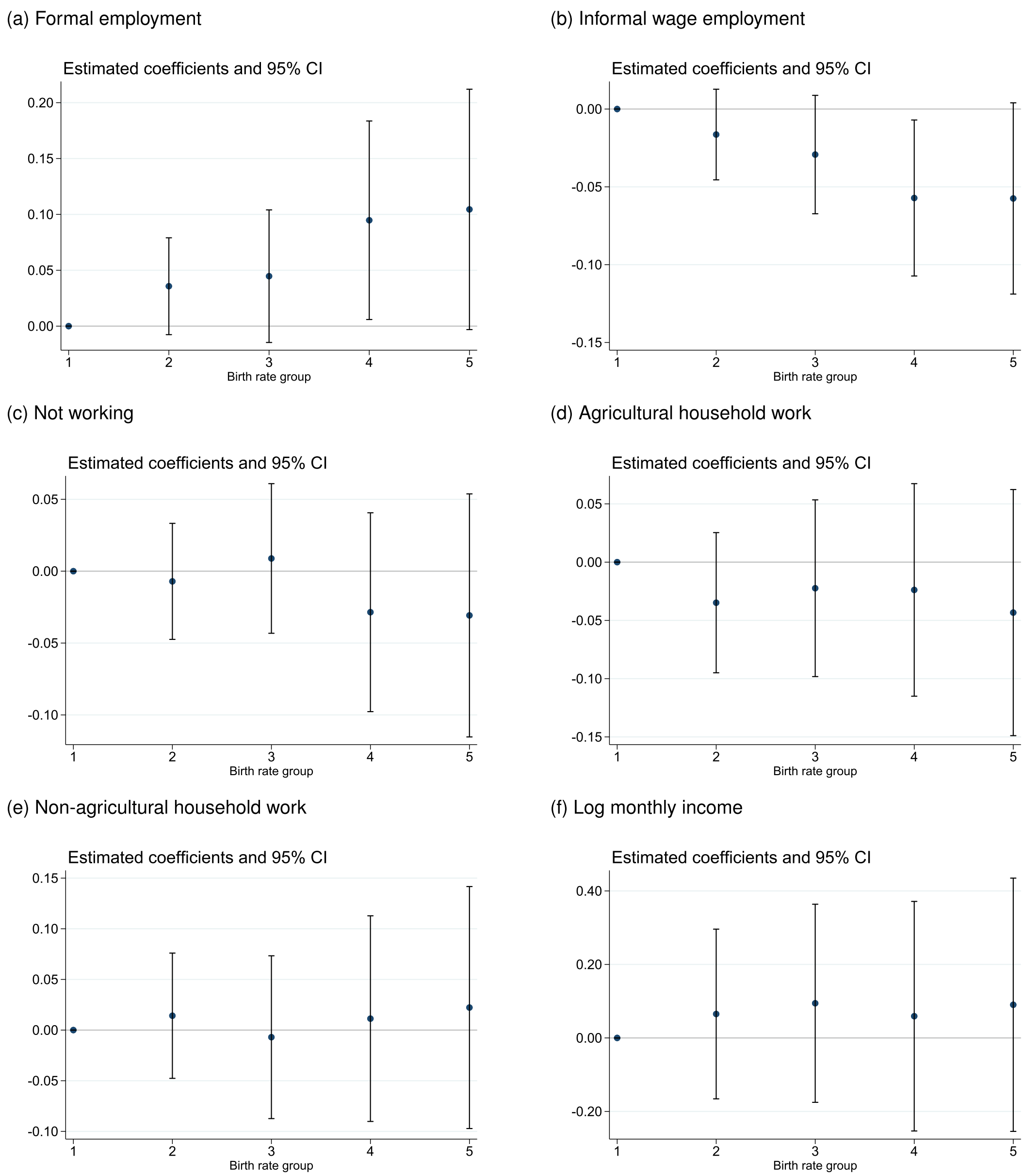

Fig. 5. DiD estimates for effects by district–age-group birth-rate bin. (Note: the plots report the results from estimating the DiD model in Eq. 5, in which treatment effects are allowed to vary by birth-rate groups. Birth rates vary by district and age group, and are binned into five equal groups. The lowest birth-rate group is the reference group. All models control for province–age group fixed effects, year–age group fixed effects, and province–year fixed effects. The sample includes women aged 25–44 in the VHLSS 2010–18 sample. Standard errors are clustered at the commune–year level and sampling weights are applied in the regressions.)

insignificant and relatively close to zero. The result for the informal wage employment in the male sample are similar to those from the female sample, but they are not precisely measured. Overall, the results from this placebo test support the conclusion that the increase in formal employment among women in the higher birth-rate group is indeed driven by the maternity leave extension and not the other provisions of the 2012 Amendment.

In summary, the results for the formal employment outcome are robust against different specifications and research designs, which allow us to confirm that the maternity leave extension provides an incentive for women of childbearing age to move into the formal sector. The evidence for informal wage employment and agricultural household work is less conclusive; the research design using the variation in age and gender suggests that women move out of the agricultural sector, while the research design using the variation in birth rates suggest that women move out of informal wage employment.

5.3. Formal employment in the public and private sectors

In this section, we apply the same approaches to understand more about sectors (i.e., public or private), industries, and occupations that women move into as a result of the maternity leave extension. First, we note from Fig. 6(a) that public employment has been declining among women of both age groups during the entire study period, but the decrease for women of childbearing age has slowed down since 2014. Similarly, public employment also decreases among older men during 2010–18, while that of younger men only decreases in 2018. The formal employment in the private sector in Fig. 6(b) are more similar to the patterns of

Fig. 6. Public and private formal employment for women aged 25–54 by year. (Source: authors' calculation based on the VHLSS 2010–18.)

overall formal employment that we observed in Fig. 2(a). There is an upward trend among all age and gender groups, but the increases among younger men and women appear to be larger than those among older men and women. Within the 25–44 age group, women also experience a larger increase than men, especially after 2014. These trends suggest that formal employment in the private sector of women of childbearing age increases following the maternity leave extension, relative to the other demographic groups. Despite these changes in employment among women aged 25–44, there is little evidence that their monthly wages are affected in both the public sector and the private formal sector. Log monthly wages follow an upward trend for all genders and age groups, although those among the older age group are more volatile for both genders.

We now turn to the regression estimates for the effects on these outcomes in Table 3. The main DiD model using older women as the control group yields positive and significant estimates for the effect on public sector employment. The baseline estimate is 0.029, and controlling for province–year and province–age group fixed effect yields a similar result. Controlling for district–year and district–age group, on the other hand, yields slightly lower estimate of 0.024.

Surprisingly, the estimates from the alternative DiD model using men of the same age as the control group are all close to zero and statistically insignificant. This is consistent across different levels of fixed effects as control variables. The triple-differences estimate is −0.007 and also insignificant. In contrast, the estimates for the effect on formal employment in the private sector are small and insignificant in the main DiD model, but are positive and significant in the alternative DiD model and also the triple-differences model. The baseline estimate of the alternative model is 0.036, while controlling for district–year and district–age group fixed effects reduces the estimate to 0.031. The triple-differences model also indicates that the treatment effect on private formal employment is positive and significant. The magnitude of the treatment effect on formal employment in the private sector is also comparable to the effect on overall formal employment. Unlike employment outcomes, the estimates for both wage outcomes are all small and statistically insignificant.[13]

The event-study estimations provide a similar story as the main models (see Figs. 7(a) and (b)). The main DiD model indicates a positive effect on public employment, while the alternative DiD and the triple-differences models suggest that the effects are close to zero. The opposite is true with private formal employment; the main DiD estimates are positive but insignificant; the alternative DiD and triple-differences' estimates are positive and significant.

The inconsistency between the main DiD model and the two other models raises a concern about whether the maternity leave extension increases formal employment in the public or the private sector. We find it unlikely that the maternity leave extension increased public employment. First, when estimating the treatment effects by age group, the estimates for the public employment outcome follow a puzzling pattern: the estimates are larger for older age groups, but the estimate for the 25–29 age group, the youngest individuals in the sample, is negative and significant (see Fig. 7(c)). We would expect younger individuals (with higher birth rates) to experience larger effect than older individuals (with lower birth rates) if these estimates represent the treatment effects of the maternity leave extension. More importantly, these estimates are also not robust to the triple-differences model. When we switch to the second DiD design that uses birth rate groups as the treatment variable (see Eq. 5), the estimated treatment effects are also small and statistically insignificant across all birth

[13] The sample size for log monthly wages decrease substantially when we consider public and private sectors separately.

Table 3
Difference-in-differences estimates for effects on formal employment in public and private sectors.

	DiD model			DiD model			DDD
	Women age 45–54 as control			Men age 25–44 as control			model
Specification	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Public sector							
	0.029***	0.030***	0.024***	−0.004	−0.005	−0.003	−0.007
	(0.008)	(0.008)	(0.009)	(0.004)	(0.004)	(0.004)	(0.009)
N	34701	34701	34422	45294	45294	45215	67739
Panel B: Private formal sector							
	0.009	0.009	0.006	0.036***	0.037***	0.031***	0.032***
	(0.011)	(0.011)	(0.012)	(0.006)	(0.005)	(0.006)	(0.009)
N	34701	34701	34422	45294	45294	45215	67739
Panel C: Log monthly income (public)							
	0.061	−0.102	0.596*	−0.020	−0.018	−0.052	0.154
	(0.110)	(0.148)	(0.346)	(0.047)	(0.056)	(0.123)	(0.446)
N	1388	1265	349	2093	2075	805	621
Panel D: Log monthly income (private formal)							
	−0.047	−0.062	−0.172	0.036	0.030	0.061	−0.071
	(0.089)	(0.098)	(0.184)	(0.029)	(0.033)	(0.052)	(0.281)
N	2419	2302	1432	3776	3721	2724	2130
Additional controls							
Age group FE	✓			✓			
Year FE	✓			✓			
Female FE				✓			
Province × Year FE		✓			✓		
Province × Age group FE		✓			✓		
Province × Female					✓		
District × Year FE			✓			✓	
District × Age group FE			✓			✓	
District × Female						✓	
District × Female × Year FE							✓
District × Age group FE × Female							✓
Province × Age group FE × Year FE							✓

This table reports the difference-in-differences estimate for the effects of the maternal leave extension on labour market outcomes. In columns 1–3, we use women of age 45–54 as the control group, and the results are reported for the interaction term Post-2013 × Age 25–44. The sample for these estimates are non-college females aged 25–54. In column 4–6, we use men of age 25–44 as the control group, and the results are reported for the interaction term Post-2013 × Female. The sample for these estimates includes all non-college individuals aged 24–44. Standard errors are clustered at the commune-year level and reported in parentheses and p-value is reported in brackets; sampling weights are applied. All models control for year FE, age group FE, urban, ethnicity, household size, number of children age under 10 in household, educational attainment, marital status, and year-specific cohort linear trends. Data is drawn from the VHLSS 2010–2018.

rate groups. In other words, the effect on public employment is also not robust to our second DiD design.

The results for the private formal employment are similar to our main conclusion about the overall formal employment. The estimated treatment effects by age, as shown in Fig. 7(d), show that the effects are largest among women between age 25 and 29, those that have the highest birth rate among all age groups (see Fig. 1). This is consistent with what we would expect if women switch their career in response to the incentive of the maternity leave extension. When estimating the model using birth rate groups as the treatment variable, the estimates for public employment are small and statistically insignificant. In contrast, the treatment effects on private formal employment are positive, significant, and larger among higher birth-rate groups. Our conclusion is that it is more likely that the maternity leave extension increases formal employment in the private sector.

Instead of splitting by public and private sectors, we can also split the formal employment by industry and occupation. That is, we consider which industry and occupation women of childbearing age are switching to as a result of the maternity leave extension.[14] We present the triple-differences estimates for the treatment effect on 20 industries and 10 occupations along with the means of the outcome variables in Fig. 8. First, we observe a relatively large, positive and significant effect on formal employment in the manufacturing sector (which also includes the textile industry), where the estimate is about 0.04. The large magnitude of the effect on this sector is perhaps not surprising, given the fact that it is also the largest sector in terms of formal employment. We also observe a positive and significant effect on the entertainment sector (including arts, sports, and other entertainment industries), although the estimate is smaller than 0.01. Surprisingly, we also find negative and significant effect on the mining and politics sectors, although the estimates are also 0.01 or lower. For other industries, the effects are close to zero and statistically insignificant. These results are consistent with our earlier conclusion that the effect is mainly on the formal employment in the private sector, as these industries are mainly located in the private sector.[15]

The results on occupation outcomes yield a similar conclusion. The treatment effects are positive and largest among typical occupations in the manufacturing sector, including operators of plants and machines and textile workers. The formal employment of women is relatively large among these occupations, and they likely require lower skills from their workers (relative to technicians and professionals). Therefore, there are likely more opportunities to switch from the informal sector. There is also a positive and significant effect on the clerical occupation, which typically include office jobs such as secretaries and customer service jobs. Consis-

[14] We thank an anonymous reviewer for suggesting that we look at the effects by industry and occupation.

[15] According to the 2010–2018 VHLSS, almost 90% of manufacturing workers work for the private sector.

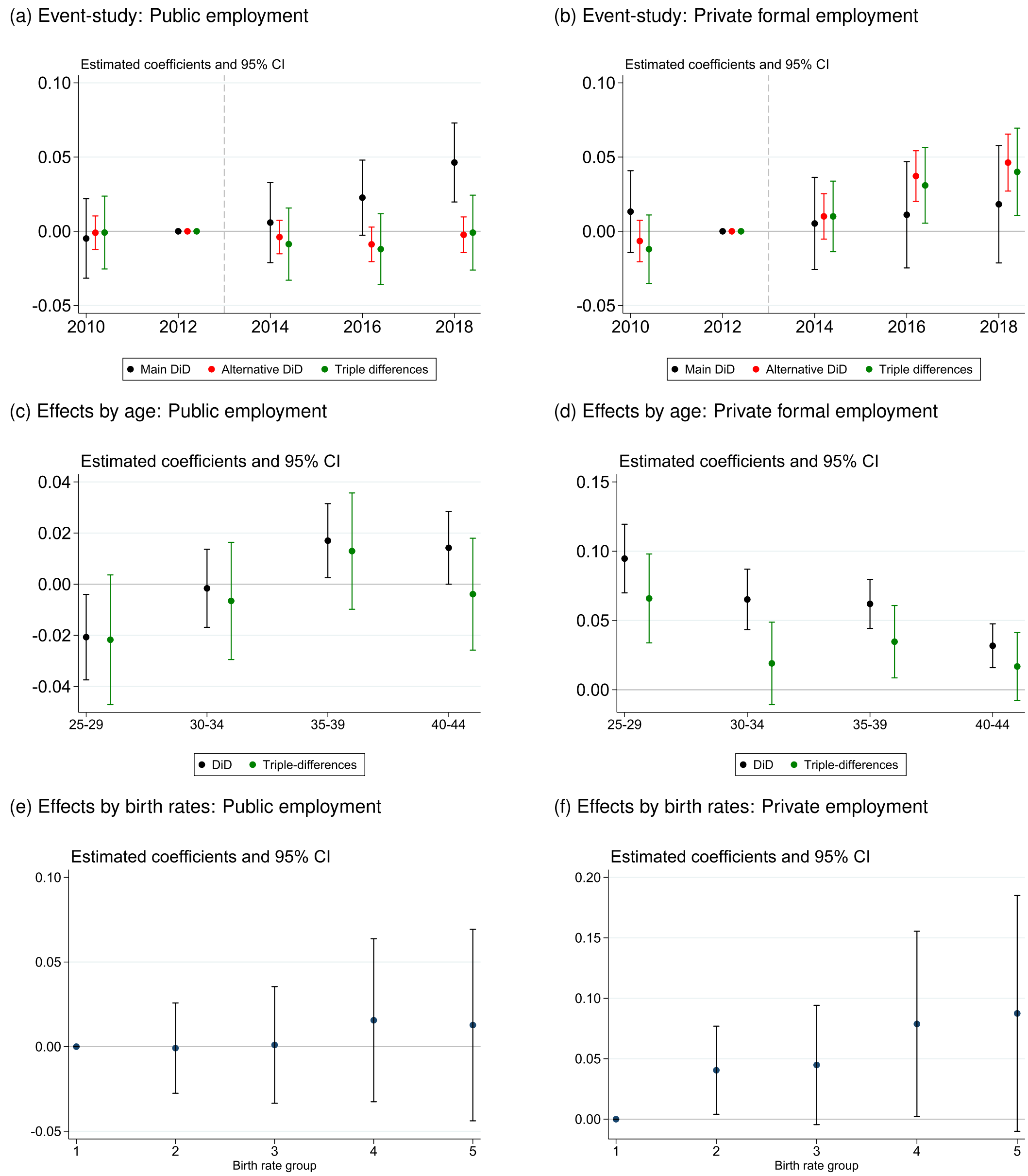

Fig. 7. Event-study estimates for effects on public and private formal employment. (Note: standard errors are clustered at the commune–year level. Source: Figure (a) and (b) report the event-study estimations for the main DiD, the alternative DiD, and the triple-differences model. In the main DiD model, the control group is women aged 45–54; the model controls for district–year and district–age-group fixed effects. In the alternative DiD model, the control group is men aged 25–44; the model controls for district–year, district–female, and district–age group fixed effects. In the triple-differences model, we control for province–year–age group, district–group–female, and district–female–year fixed effects. Figure (c) and (d) report the estimations for the main DiD and the triple-differences models, in which the treatment effects vary by age group. Figure (e) and (f) report the DiD estimates where the treatment effects vary by birth-rate groups. All models control for the individual covariates described in text.)

tent with what we find for the industry outcomes, these are the occupations that are mainly found in the private sector.[16]

In short, these results suggest that the maternity leave extension of the 2012 Amendment provides an incentive for women of childbearing age to move into formal employment in the private sector. The switch primarily happens in manufacturing where the formal employment is more common than in other industries. The effects are also concentrated among middle-skilled occupations such as machine operators, plant workers, craft workers, as well as clerks. This is likely because the labor demands for these occupations have been rising recently as a result of more private firms entering the manufacturing industries (McCaig Pavcnik, 2017), and the skills for these occupations are relatively easy to acquire relative to those in the high-skilled occupations.

6. Conclusion

The 2012 Amendment to the Labor 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,

[16] According to the 2010–2018 VHLSS, about 86% of the plant and machine operators/workers and 82% of the craft and related workers are hired by the private sector. The clerical occupation is more balanced, as only 53% of the occupation are employed by the private sector.

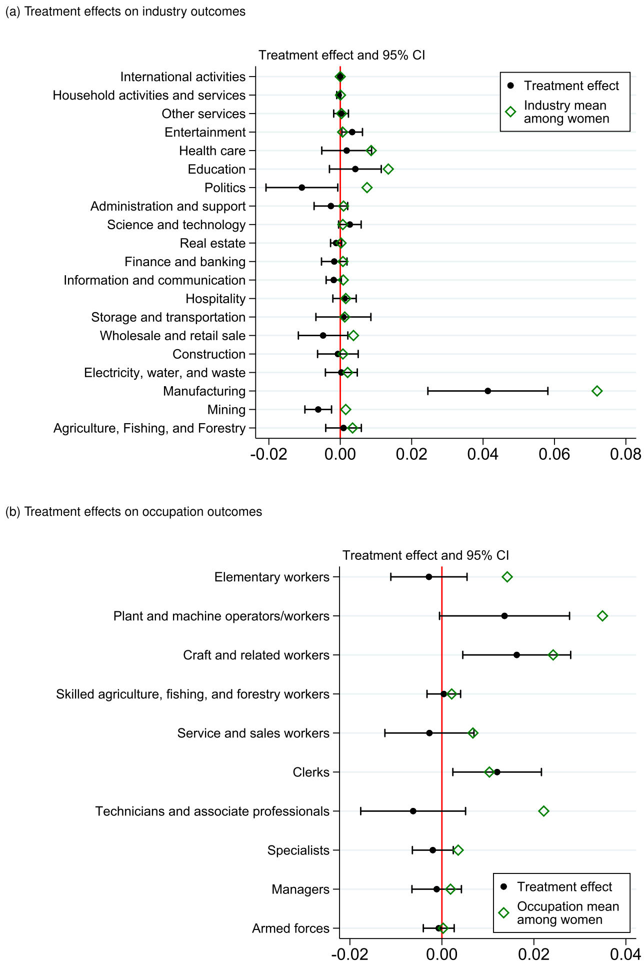

Fig. 8. Triple-differences estimates on industry and occupation outcomes. (Note: the plots report the results from estimating the triple-differences model in Eq. 4. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, marital status, province–age group–year fixed effects, district–age group–female fixed effects, and district–female–year fixed effects. The sample is all non-college men and women age 25–54 in the 2010–18 VHLSS sample. Standard errors are clustered at the commune–year level and sampling weights are applied in the regressions.)

creating an incentive for these women to switch from informal work, such as farm or non-farm household work, to formal employment.

We assess the impacts of the maternity leave extension on women's labor market outcomes using two different DiD designs. First, we compare 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 3.3 percentage points in the probability of formal employment. Our results are robust to switching to using men of the same age group as the control group as well as a triple-differences model. Second, we compare employment outcomes among women by their expected birth rates and arrive at similar conclusions. Groups with higher expected birth rates are more likely to move into formal employment in response to the new law. However, the two designs point to different sources of informal employment that women of childbearing age move out of. The first design suggests that women move out of the agricultural household work, while the second design suggests that women move out of informal wage employment. We also observe that the extension encourages women to switch into the private formal sector, specifically in the manufacturing industry and middle-skilled occupations.

Interestingly, our findings are opposite of those of Uribe et al. (2019), who use a similar research design to evaluate a similar law in Colombia and find that extending maternity leave has a negative effect on women's formal employment. The difference might be due to the differences in the labor market and regulations between Vietnam and Colombia. First, the private formal sector in Vietnam has undergone significant expansion in the last two decades, especially in the manufacturing industry. The 2000 Enterprise Law, for example, reduced the administrative burden to register new business, contributing to the growth of private formal firms (Malesky & Taussig, 2009). Changes in the trade policies such as the US-Vietnam Bilateral Trade Agreement (2001) and the Vietnam's accession to the WTO (2007) shifted the markets away from agriculture and towards manufacturing exports (McCaig Pavcnik, 2017). This considerable economic development likely provides more opportunities for workers to move out of the informal sector and into the manufacturing sector. Along with the increase in labor demands, female workers in Vietnam also appear to value sick and maternity leaves as a reason to be employed in a company (Tran & Jeppesen, 2016).

Second, in Vietnam, the salary received by female workers during a maternity leave is completely paid for by social insurance.[17] In the case of Colombia, the employer has to pay 73.06% of the workers' salary during leave, while the rest is covered by social security (Uribe et al., 2019). The critical difference between the two countries in how salary is paid during the maternity leave may help explain why our findings are different from those in Uribe et al. (2019). These differences highlight the importance of conducting more studies on the labor market effects of maternity leave mandates in developing countries. As countries differ in terms of labor laws and labor market conditions, the effects on female labor market outcomes may also vary.

To close, we note that our study comes with two major caveats. First, the VHLSS data allow us to study about formal employment only during 2010 or later because the survey did not ask for information about social insurance before that, which means that we can examine the pre-treatment trends for only two periods: 2010 and 2012. Although our event study results support the assumption of parallel trends between the treatment group and two different control groups during these two years, it is possible that the trends may not be parallel before 2010. As a result, the counterfactual trends (in absence of the maternity leave extension) may also be non-parallel.

However, we observe that the shares of formal employment of all age groups and genders follow relatively stable trends up until 2016 when the trend of women of childbearing age diverged substantially (see Fig. 2). More importantly, the trends of the control groups (men between age 25–44 and women between age 45–54) continue to be very stable over time. These patterns in the raw data provide reassuring evidence that the changes in the labor market outcomes respond to the 2012 Amendment rather than other factors. Our second DiD design further suggests that the results are driven by birth rates among the childbearing-aged women. Therefore, we can conclude that the intervention is related to fertility-related incentives, which can only be the maternity leave extension.

Second, the sample size for individuals with monthly wage information is likely too small to detect any effect of the leave extension on this outcome, especially for wages separately in the public and private formal sectors. Although on-leave employees

[17] Employers do have to contribute 18% of their payroll and employees have to contribute 8% of their salary towards social insurance.

are paid through their social insurance, firms are still likely to incur indirect costs for such leaves by leaving the position unfilled or hiring temporary workers. Hence, one may expect the maternity leave extension to have some effect on monthly wages as the required leave is increased from 4 to 6 months. It is possible that the 2-month increase is relatively small so the effects on monthly wages are also small.

CRediT authorship contribution statement

Khoa Vu: Conceptualization, Formal analysis, Methodology, Validation, Visualization, Writing – original draft, Writing – review & editing. **Paul Glewwe:** Conceptualization, Resources, Writing – review & editing, Validation.

Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

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Appendix A. Extra tables and figures

A.1. Overview of the demographic composition of the labor market in Vietnam

Compared to other low- to middle-income countries, Viet Nam has a relatively high female labor force participation rate and gender equality (Klasen et al., 2020). In this section, we provide an overview of the demographic compositions and changes of the labor market in Vietnam during 2010 to 2018 using the Vietnam Household Living Standard Survey (VHLSS).

The labor force participation rate among women aged 25–54 is slightly lower than that of men in the same age group. The female labor 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 Fig. A3. The labor 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.

Fig. A1. Reasons for not working, 2014–18. (Source: authors' calculations based on the VHLSS 2014–18 for individuals aged 22–65.)

The labor market composition is also remarkably different across gender and college education, as illustrated in Fig. A4. 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 labor 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.

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 Fig. A5. 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.[18] One potential explanation for this shift 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 (Fig. A6).

[18] 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.

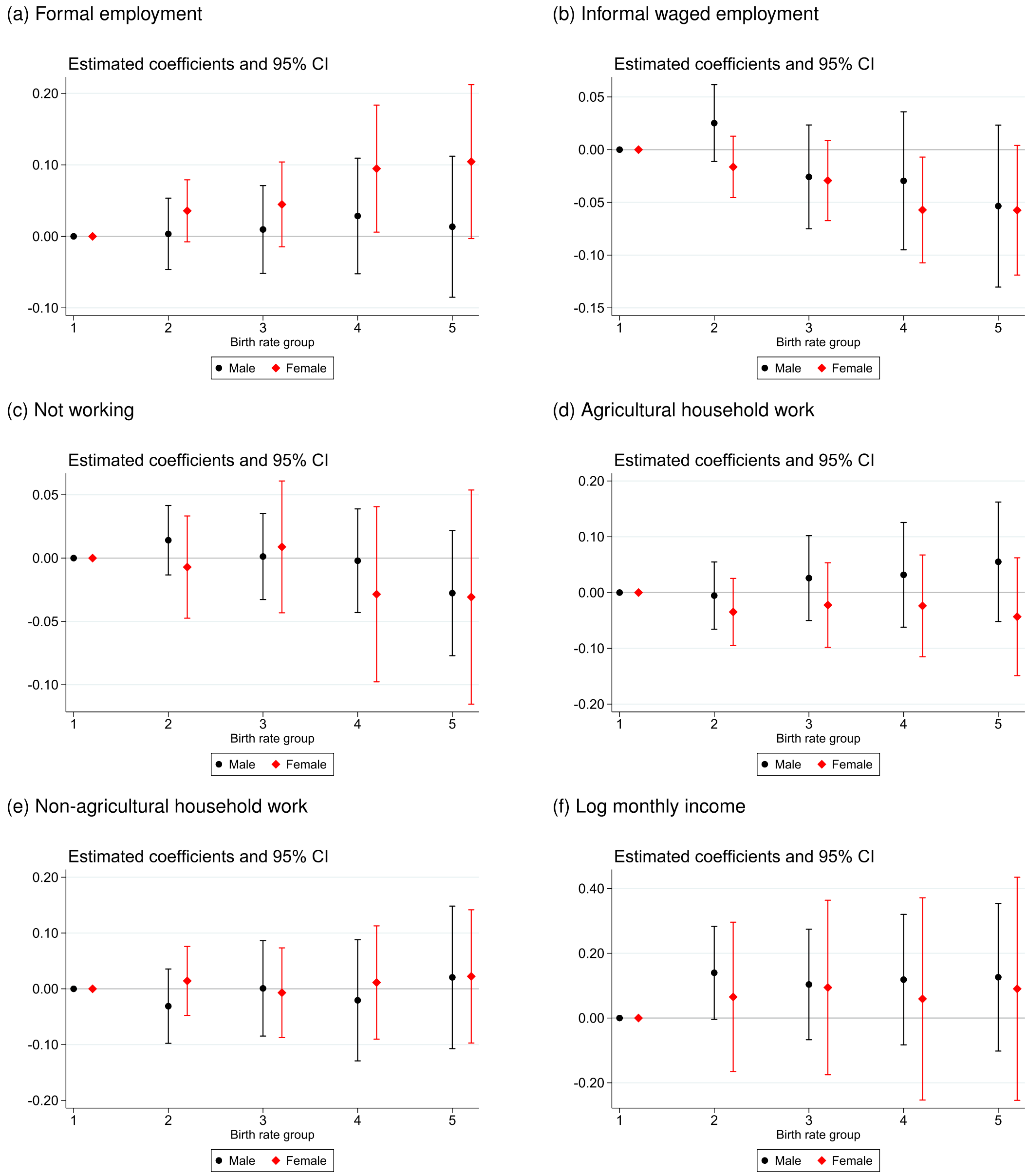

Fig. A2. DiD estimates for effects by district–age-group birth-rate bin on men and women. (Note: the plots report the results from estimating the DiD model in Eq. 5, in which treatment effects are allowed to vary by birth-rate groups. Birth rates vary by district and age, and are binned into five equal groups. All models control for urban, ethnicity, household size, number of children under age 10 in the household, educational attainment, and marital status. The model is estimated separately for men and women aged 25–45 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.)

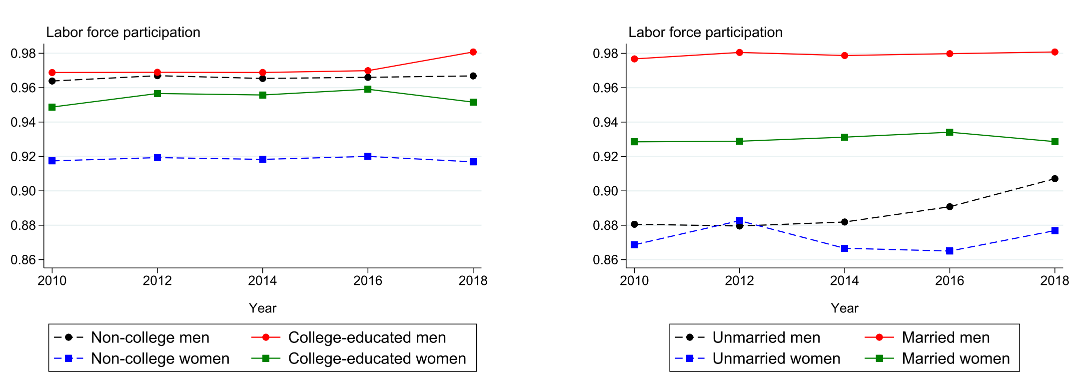

Fig. A3. Labor force participation of men and women aged 25–54 in 2018 by college education and marital status. (Note: the sample includes all men and women aged 25–54. Source: authors' calculations based on the Viet Nam Household Living Standard Survey (VHLSS) 2010–18 (see text for description).)

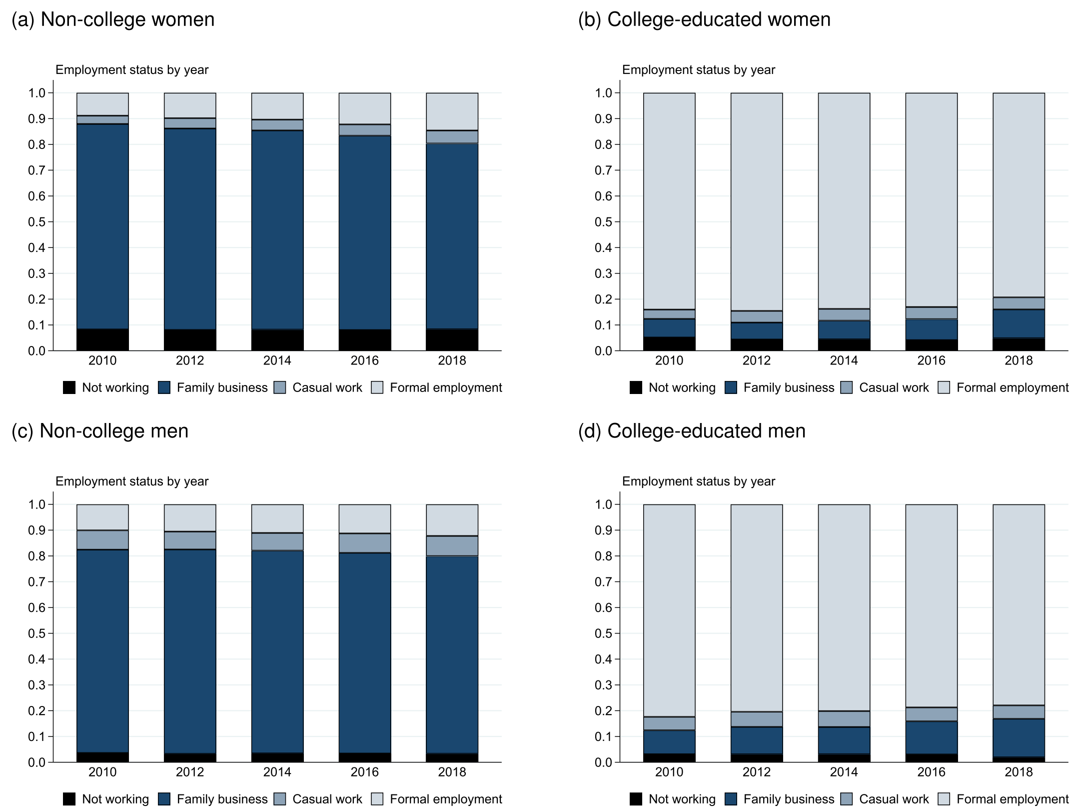

Fig. A4. Labor market composition among men and women aged 25–54 in 2018 by college education. (Note: the sample includes all men and women aged 25–54. Source: authors' calculations based on the VHLSS 2010–18.)

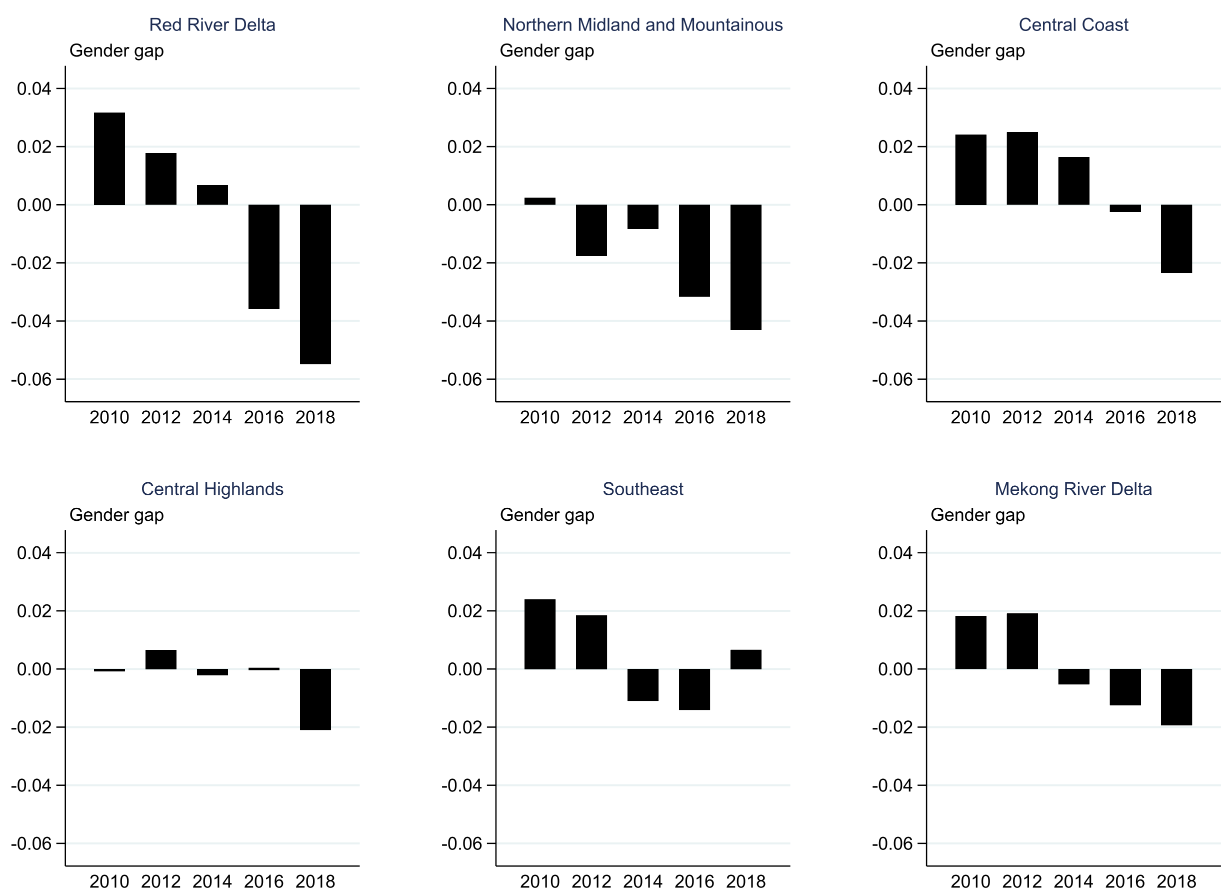

Fig. A5. Gender gap in formal employment by region and year. (Note: the sample includes all men and women aged 25–54. Source: authors' calculations based on the VHLSS 2010–18.)

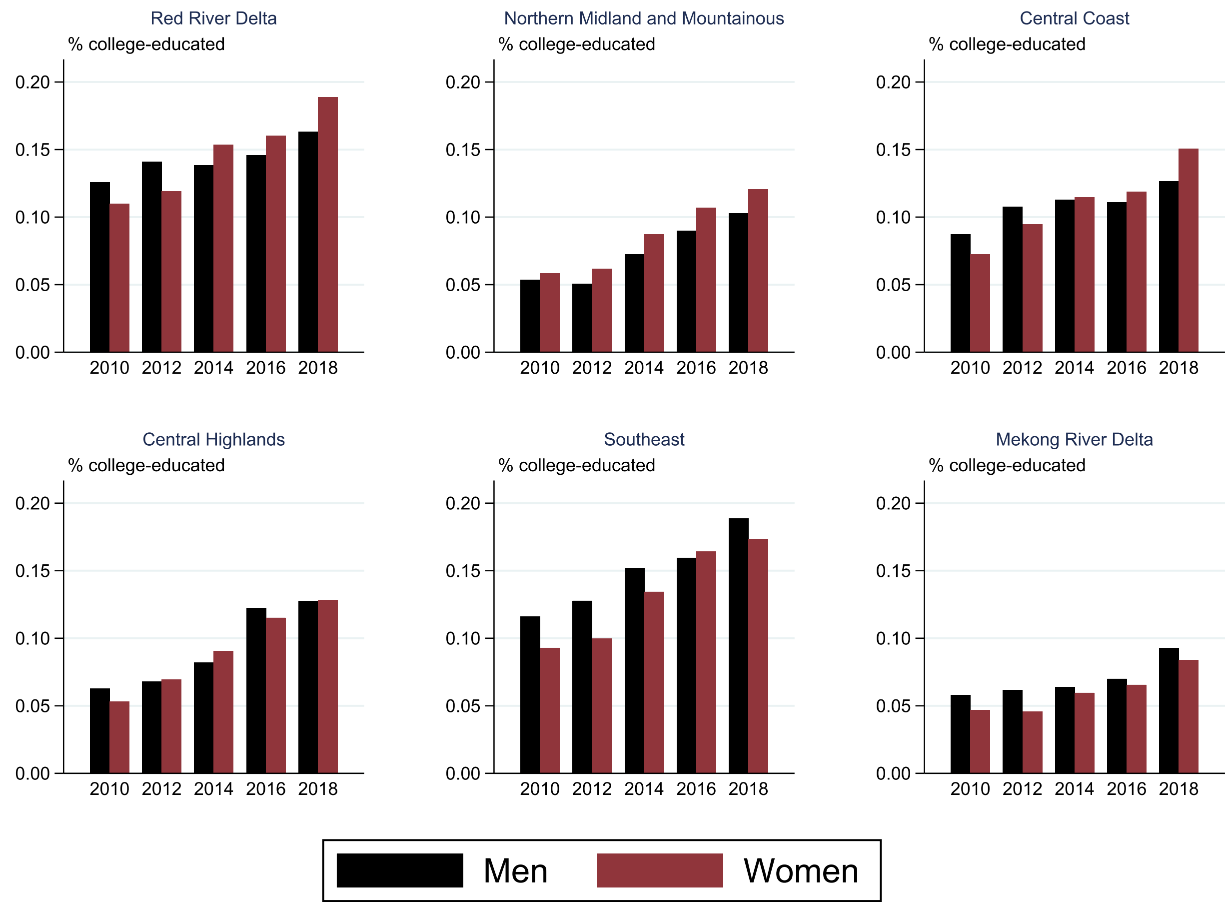

Fig. A6. Percentage of men and women who are college-educated by region and year. (Note: the sample includes all men and women aged 25–54. Source: authors' calculations based on the VHLSS 2010–18.)

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