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JEL Cassification: H4, J1, J2, O1

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1. Introduction

Women earn less income and are less likely to participate in the labor market, especially in poorer countries (World Bank, 2012; International Labor Organization, 2019). The international community has jointly called for improvements to gender inequality in economic activities, which are highlighted in various targets and indicators with the Sustainable Development Goals (SDG).[1] There are several ways to increase women's involvement in such activities, such as micro-credit programs, self-help groups, or programs that are specially designed to help improve their access to infrastructure and technology (see, e.g., Duflo (2012) and Quisumbing, Meinzen-Dick, and Malapit (2021)).[2] In this paper, we study a simple but important factor—childcare—that can release women from domestic work and encourage them to participate in the labor market. Specifically, we examine the impacts on women's labor market outcomes of pre-school (age 1-5) childcare in Vietnam.[3]

Analyzing the most recent household surveys for Vietnam in the past decade, we find that childcare has strong and positive effects on women's probability of working in a formal wage-earning job. Sending children to childcare centers helps mothers move from self-employed farm work to wage-earning job. These results are robust to a battery of sensitivity analyses and falsification tests. Childcare also helps increase women's total annual wages and household income per capita and reduce household poverty (although these effects are marginally statistically significant). Moreover, a medium-term effect exists after two years for younger children, increasing the probability that women have a wage-earning job by 38 percentage points and reducing self-employed farm work. We also find heterogeneous effects of childcare; these effects on the probability of having a wage job are larger for more educated women and women in ethnic majority groups. These effects are (somewhat) larger for younger children or areas with higher income levels.

We offer the first study that rigorously examines the impacts of childcare for Vietnam, which represents an interesting case to study for several reasons. Despite being a poorer country, Vietnam's annual economic growth rate of around 6% during the past two decades ranks among the fastest in the world. Yet, gender inequality remains a challenge for

[1] See SDG number 5 at https://sustainabledevelopment.un.org/sdg5.
[2] See also Card, Kluve, and Weber (2018) for a recent meta-analysis of other labor market programs.
[3] We refer to childcare and preschool interchangeably in this paper.

the country. The proportion of men working in a wage job is 42%, but the corresponding figure for women is lower at 30%. Furthermore, Vietnam started joining the ranks of low-middle income countries about 10 years ago, and concerns have recently been raised that the country is prone to being caught in the "middle-income trap" (World Bank and MPI, 2016). Since reducing gender inequality is not only an end by itself but can also offer a promising way to achieve better long-run economic growth, our findings are particularly relevant to gender and labor policies.[4]

A key challenge in measuring the effects of childcare is the endogeneity issue. Women who send children to childcare services differ from those who do not (e.g., they may be richer or simply have less time for childcare). We address this problem by using the quasi-experimental regression discontinuity design (RDD) method that exploits plausibly exogenous thresholds in the birth months of children to identify the causal impacts of childcare. In Vietnam, the school year for public preschool and primary school starts in September. Children's enrollment in preschool is based on their age simply according to the calendar year (i.e., their running/ current age), regardless of whether they have reached their birth day in that year (i.e., their full/ completed age). We can thus compare labor outcomes for women whose children were born in adjacent months in two contiguous, but different, years. In particular, a child who was born in January in any given year is more likely to start preschool one year later than a child who was born in December of the preceding year, despite an age difference of only one month.

Beyond a specific country study, we offer new analysis to the literature by investigating both the *quantity* and *quality* aspects of women's labor outcomes. The latter is increasingly pertinent to a developing country context, given that on average two out of three workers in emerging economies are in the informal sector and significant gaps exist in the returns to experience between the informal sector and the formal sector (World Bank, 2019). Besides women's labor force participation (LFP) and number of work hours, we look at a wide range of quality indicators, including employment outcomes and household-level outcomes both in the short term and the medium term. Specifically, the former outcomes

[4] See, e.g., Lagerlöf (2003) and Diebolt and Perrin (2013) for studies on the role of gender empowerment and economic growth. Duflo (2012), Bandiera and Natraj (2013), and Silva and Klasen (2021) provide recent reviews of this literature. The recent Covid-19 pandemic also exacerbates gender inequality issues (e.g, Alon *et al.* (2021) and Dang and Nguyen (2021)).

include self-employment, wage work, farm and non-farm work, skilled employment, and formal work, and the latter outcomes include household income, poverty, household size, migration, and co-residence with grandparents. To our knowledge, hardly any other studies examine such diverse and comprehensive outcomes, particularly regarding the quality of female employment, in a poorer country context.[5]

We analyze a rich and nationally representative survey spanning the early to mid-2010s, with each survey round covering more than 46,000 households. Rather than focusing on a specific age (subpopulation) group or a smaller sample as most previous studies do, our sample for analysis consists of all preschool children age 1-5. This larger survey sample for Vietnam has rarely, if at all, been analyzed in an academic study before.

While a large literature exists on the impacts of childcare subsidies for richer countries, far fewer studies rigorously investigate the developing country context. The effects of childcare on parental employment can vary significantly between the former and the latter countries because of their systematic differences in childcare and labor market institutions. Furthermore, the empirical findings on the effect of childcare on parental employment appear inconclusive even for richer countries. Indeed, while most recent studies find a significant, positive effect of childcare use on women' labor supply (e.g., Bauernschuster and Schlotter, 2015; Martínez and Perticará, 2017; Berthelon *et al.*, 2018), a number of other studies do not (e.g., Cascio, 2009; Havnes and Mogstad, 2011). Reviews by Blau and Currie (2006) and Akgunduz and Plantega (2018) show a large variation in the elasticity of maternal employment to childcare costs across different studies, resulting from differences in samples of women and children, estimation methods, and country contexts.

We briefly review some recent studies on poorer countries that are most relevant to our studies. Berlinski, Galiani, and McEwan (2011) find a positive impact for childcare use on women's LFP and work hours in Argentina when the youngest child in the household attends preschool (but no effects when a child who is not the youngest in the household attends preschool). Calderon (2014) shows that access to childcare in Mexico enables women to obtain more stable jobs and increases their labor incomes. Yet, while Clark *et al.* (2019)

[5] Clark *et al.* (2019) offer an exception. Besides women's employment and income, they also investigate the impacts of childcare on other outcomes such as women's autonomy and fertility intentions.

find that subsidized early childcare in Kenya increases the likelihood of employment for married women and helps single mothers move to jobs with more regular hours, they find mixed evidence on women's earnings. These findings perhaps concur with Halim, Johnson, and Perova (2021)'s results that public preschool expansion in Indonesia has positive impacts on women's LFP mostly through unpaid family work and but has no impacts on earnings or hours of work.[6]

This paper consists of six sections. We describe the data in the next section, before presenting in Section 3 descriptive analysis of childcare and maternal employment in Vietnam. We discuss the estimation method and provide the empirical results in Sections 4 and 5 and finally conclude in Section 6.

2. Data

The data that we analyze include four rounds of the Vietnam Household Living Standards Survey (VHLSSs) between 2010 and 2016. The VHLSSs have been conducted biennially by the General Statistics Office of Vietnam (GSO) with technical support from the World Bank since 2002. We restrict analysis to the 2010 survey round onwards, since the more recent (survey) rounds contain more information on employment consisting of monthly wages (i.e., regular salaried work) and formal employment.

The sampling frame for the VHLSSs is based on the 2009 Population and Housing Census. Each VHLSS covers around 46,000 households and is divided into two sample types: (1) The sample for the income survey includes around 36,000 households and collects information to assess monetary living standards at the national, regional, and provincial (city) levels; (2) The sample for the income-expenditure survey includes around 9,000 households

[6] Notably, two common limitations exist with most of the few existing studies on developing countries. First, these studies mostly focus on women's LFP decision and work hours rather than the quality of work (such as work sectors) or other household outcomes (such as total household income or poverty status). Second, these studies usually restrict analysis to a population subgroup (e.g., children age 3-4 in Berlinski *et al.* (2011) or age 3-6 in Halim *et al.* (2021)) or collect data in a specific location (e.g., an informal settlement in Nairobi, Kenya as in Clark *et al.* (2019)) rather than the whole population. Besides the cited studies, other studies include Maurer-Fazio *et al.* (2011) who focus on urban China and Attanasio *et al.* (2017) who mostly focus on slum households in the city of Rio de Janeiro, Brazil. In contrast, we study a number of employment outcomes for women and their households, using several nationally representative survey rounds with data on all pre-school children age 1-5 in Vietnam.

and collects richer information on both monetary and non- monetary living standards at the national and regional levels. While the (smaller) income-expenditure sample has been widely used, the (larger) income sample has not been analyzed, mostly due to restricted data access.

We use the full sample of the VHLSSs to obtain the maximum data points on children birth months. The total numbers of sampled households and household members in the VHLSSs are as follows: 46,995 households with 185,696 household members in 2010, 46,996 households with 182,042 household members in 2012, 46,335 households with 178,267 household members in 2014, and 46,380 households with 175,340 household members in 2016.

The VHLSSs collect detailed data on individuals, households, and communes. Household-level data include information on household assets, income and expenditures, and participation in government programs. Individual-level data include information on demographics, education, employment, and migration. However, the VHLSSs collect data on the birth year and the birth month, but not the full birth day, for each individual.

3. Childcare systems and descriptive analysis

3.1. Childcare in Vietnam

We study childcare (either in childcare centers or preschools) for children age 1-5. We do not consider children younger than one year old, since almost no such children attend childcare in our data. We also exclude children age 6 (or older), since this is the age when most children start attending the first grade of primary school.[7] The admission age for school entrance in Vietnam is based on a child's current age rather than completed age. The school year in the country starts from September each year. Thus, for example, if a child was born in 2000 (regardless of the birth month), this child can attend a preschool from September 2003 and attend a primary school from September 2006.[8] Following this practice, hereafter, when we discuss a child's age, we refer to the current age rather than the completed age.

[7] According to the 2005 Law on Education, the starting age for entering primary school is 6. Only children with special difficulties can enter primary school later than 6 years old.

[8] Children attend childcare centers and preschools from Monday to Friday. The school day at these institutions and at primary schools often starts at 7.30 a.m. and ends at 4.30 p.m. But this time schedule is not fixed. Some

Children under the age of three can attend early childcare centers, but access to early childhood care centers remains limited in Vietnam, with only 26% of villages in rural Vietnam providing such centers.[9] Preschools for children age 3 to 5 are more available, with 49% of villages having at least a preschool. The education system in Vietnam is mostly public, with 90% of children age 3-5 attending public preschools and the rest going to private preschools. More children below the age of 3, however, are enrolled in private childcare centers.[10] The proportions of children age below 3 attending private and public childcare centers were 27% and 73% in 2016, respectively. For simplicity, hereafter we refer to childcare centers for small children below 3 and preschools for children age 3-5 as "childcare (centers)".

The VHLSSs do not collect data on the availability of childcare centers for urban areas. However, using individual-level data, we can estimate the proportions of urban and rural children attending childcare centers and preschools. Figure 1 presents the percentage of children attending childcare by age. Less than 1% of children below the age of 1 attended childcare in 2016. The number of children age 1 attending childcare is also small, at around 3%.[11] Childcare increases significantly by age; specifically, 48% of children age 3, 69% of children age 4, and 80% of children age 5 attended preschools in 2016. Figure 1 also shows an increase in the enrollment rate of children over time. For all age groups, the percentage of children attending childcare was significantly higher in 2016 than 2010.

3.2. Maternal employment

We examine the employment outcomes of women with at least one child age below 6 over the period 2010-2016 in Table 1. These outcomes are measured by different variables, including women's current working status, whether they have a wage-earning job, a skilled

childcare centers (and preschools) admit children on Saturday and allow them to be picked up later than 4.30 p.m. Some preschools also admit younger children (National Assembly of Vietnam, 2005).

[9] There are 63 provinces and provincial-level cities in Vietnam, which are split into districts, and each district is split further into communes. Communes are the smallest administrative units in Vietnam. In 2016, there were 713 districts and 11,164 communes in Vietnam; each commune contains around 3-15 villages. Communes are called wards in urban areas. In 2016, 44% of urban children age below 6 attended childcare centers and preschools, while the corresponding figure for rural children was lower at 35%.

[10] These mostly appear to be formal private childcare centers, since the VHLSSs may not capture another emerging form of private childcare that is informal and unregistered.

[11] Women are given up to 6 months' maternity leave in Vietnam (National Assembly of Vietnam, 2014).

occupation, a formal job, or their main occupation during the past 12 months is in the farm or non-farm sector.[12] These women's average age hovers around 32 and ranges from 17 to 58.

The working rate of women in 2016 was 93%. For comparison, Table A.1 in the Appendix also reports men's employment rate, which is 6 percentage points higher with almost all men (99%) working. This gender gap in the employment rate, however, was stable during the period. The 2014 and 2016 VHLSSs include questions concerning the reasons for not working, which are very different for men and women. Figure A.1 in the Appendix presents the distribution of women and men by the reasons for not working. In 2016, 90% of women did not work outside the home because they were occupied with housework. These activities include childcare, caring for older people, or caring for one's own home. Childcare is an important reason why women stay home; indeed, Figure 2 shows a large gap in labor market outcomes between women with children attending childcare and those with children not attending childcare. Compared with women, the proportion of men not working outside the home for the same reason was much lower at 23% in 2016.[13]

Among working people, men work more hours (209 hours) in a month than women (190 hours) in 2016. The gender gap for wage-earning jobs is larger, although this gap decreased over time. In 2016, around half of all men (51%) had a wage-earning job, while the corresponding figure was much lower at 38% for women. This means that more than half of all women (55%) were self-employed. We further disaggregate self-employment into farm and non-farm employment; in 2016, 18% of women worked in the non-farm sector and 37% of women worked in the farm sector.

We also examine the quality of employment, which we categorize into skilled employment and formal employment. While men were more likely to have a skilled job than

[12] Skilled occupation is defined based on 2-digits occupation codes of workers provided in the VHLSS. People with skilled occupations include managers, professionals, experts, office staffs, mechanic and other skilled workers. Formal jobs are jobs with social insurance benefits.

[13] For men, the main reasons for not working were retirement, sickness, and disability. The unemployment rate (the unemployed comprise those who were unable to find a job during the past year) was very low, less than 1% in 2016. Furthermore, although Vietnam has accomplished almost universal primary school enrollment, more than half (53%) of children age 1-5 do not attend childcare. This gender mismatch can worsen without the appropriate policies, since there was already a reverse gender gap at the upper secondary school level in Vietnam starting from 2005 (Dang and Glewwe, 2018). See also Heath and Jayachandran (2017) for a general discussion on rising female school enrolment and LFP in developing countries.

women (i.e., 54% for women in Table 1 versus 62% for men in Table A.1), women had a slightly higher rate of formal employment than men (i.e., 24% for women in Table 1 versus 21% for men in Table A.1).[14]

The VHLSSs collect data on respondents' wages over the past 30 days and their total wages over the past 12 months. Since a number of workers have more than one job, we compute total wages from the main and secondary jobs for our analysis.[15] The real hourly wage of women increased from 18,100 to 23,500 VND during 2010-2016, as did their monthly and yearly wages. However, the gender gap in wages also increased over time. In 2010, the average annual wage for men was 9% higher than that for women, but in 2016, this gender wage gap widened to 18%.[16]

3.3. Childcare and maternal employment

Figure 2 compares by children's age several employment variables for women whose children attend childcare versus those whose children do not. The difference in the working rate between the two groups is small. However, there is a clear gap in terms of wage-earning jobs, which is larger for those with smaller children than those with older children. Similarly, women whose children attend childcare have a higher proportion of formal jobs, more working hours, and higher wages than women with children who do not attend childcare. The difference in these employment variables also tends to be larger for younger children than older children.

Figure 2 shows a correlational—rather than a causal—relationship between women's employment and childcare because there can be unobserved factors that affect both childcare and maternal employment. The next sections discuss our estimation method and empirical findings on the causal effects of childcare on maternal employment.

[14] The social contribution or payroll tax in Vietnam is equal to 34% of the monthly salary, of which 23.5% is paid by employers and 10.5% by workers. Workers making a social insurance contribution are eligible for health insurance, employment subsidies, and pensions (on retirement).

[15] We also conduct analysis using wages from the main jobs alone, and estimation results are very similar because wages from secondary jobs equal only around 4% of those from the main jobs.

[16] Further breakdown of women's employment by urban/ rural areas (Table A.2) suggests that while the percentage of women's LFP is higher in rural areas, the percentage of urban women working in non-farm work, skilled work, or formal work is higher. Urban women also work more hours and earn more wages per hour than rural women.

4. Estimation method

To measure the impacts of childcare on maternal employment, we use the quasi-experimental regression discontinuity design (RDD) method.[17] In the absence of randomized control trial (RCT) evidence, the RDD method arguably offers the most rigorous experimental evidence on the causal impacts of childcare. This empirical strategy was employed by Gelbach (2002) and Fitzpatrick (2010) in previous studies for the U.S. and Berlinski *et al.* (2011) for Argentina. In RDDs, a group is selected for treatment if the value of at least one observed variable crosses a cut-off threshold. In particular, there exists a conditioning variable *Z,* such that the treatment variable *D* is equal to 1 if and only if *Z* is larger than a specific value *c*. We can identify the treatment effect by comparing outcomes *Y* for individuals just above and below the threshold *c*: $\tau_{RD} = Y^{+} - Y^{-}$.

The proportion of children attending childcare increases by age, and children's current age is used to determine their enrollment eligibility as discussed earlier. We thus use birth months as the conditioning variable that determines childcare. Specifically, we compare the employment of women whose children were born in December and January in two contiguous years. Since the school year starts in September, a child born in January of a given year likely starts attending childcare one year later than a child born in December of the previous year, even though the two children differ in age by only one month.[18] In the RDD terminology, the bandwidth is one month.

Figure 3 shows the proportion of children age 1-5 attending childcare by birth month in two consecutive years for two groups of children: the older group were born from July to December, while the younger group were born from January to June of the following year. Panel A of Figure 3 presents a graph using data from the pooled sample of children. Other panels of the figure present graphs for different age groups. For example, for children age 1-2, panel B of the figure shows the percentage of childcare of children age 2, born in July to

[17] Recent detailed discussion of RDD methods are provided in Lee and Lemieux (2010) and Cattaneo, Idrobo, and Titiunik (2019).

[18] The VHLSSs contain data on age (year and month), but not an individual's full date of birth. Thus, we cannot use the date of birth as the conditioning variable. But we also offer robustness checks in Section 5 where we vary the month bandwidths around the cut-off threshold.

December, and of children age 1 and born from January to June. Older children unsurprisingly have a higher rate of childcare. However, there is an obvious, large gap in the incidence of childcare between children born in December and those born in January.[19]

Since birth month does not strictly determine childcare attendance, we apply fuzzy regression discontinuity to measure the effect of childcare on maternal employment. Fuzzy regression discontinuity identifies the local effect of enrollment at the threshold, as follows:

$$\tau_{FRD} = \frac{Y^+ - Y^-}{D^+ - D^-}, \quad (1)$$

where Y^+ and Y^- are the employment outcomes of women of children born in December and those born in January in two consecutive years, respectively. D^+ and D^- represent the probability of being enrolled in childcare.

To estimate the effect of childcare on maternal employment in equation (1), we use an instrumental variable (IV) regression, which consists of two stages. In the first stage, we estimate the effect of being born in December on the probability of attending childcare:

$$D_{i,j} = \alpha + \varphi Dec_{i,j} + \gamma' X_{i,j} + \epsilon_{i,j} \quad (2)$$

where $D_{i,j}$ is a dummy variable that indicates a mother i with a child j who currently attends a childcare center. $Dec_{i,j}$ is a dummy variable which equals 1 if the child was born in December and 0 if the child was born in January of the following year. $X_{i,j}$ and $\epsilon_{i,j}$ are vectors of observed and unobserved characteristics of women, respectively. We use a small set of exogenous control variables, including age, gender, ethnic minorities, women's number of years of schooling, and year dummy variables, which serve as basic demographic explanatory variables of employment and wages. These control variables should be exogenous and unaffected by the treatment variable of interest (Angrist and Pischke, 2009; Heckman *et al.*, 1999), which is childcare in this case. We aim to estimate the total effects of

[19] This gap no longer exists for primary-school age or older children; Figure A.2 in the appendix shows an illustration for children age 7 to 9. Gelman and Imbens (2019) suggest that local linear and local quadratic regressions should be used instead of higher-order polynomials. In our case, the linear regression line and quadratic regression line produce very similar graphs, especially the difference in outcomes for December and January. For interpretation, we use graphs from the quadratic regressions, which are generated with the Stata command "*rdplot*".

childcare on outcomes rather than the partial effects with other variables held constant (Duflo *et al.*, 2007).

In the pooled sample of the VHLSSs, there are 3,869 children age 1-5 born in December and January. There are six twins, and we drop these twins from the sample, resulting in the final estimation sample of 3,863 observations. We focus on a 1-month bandwidth since parents can choose seasons or perhaps even a window of a few months around the birth for their children. While a wider bandwidth can improve efficiency by allowing for a larger number of observations, it likely results in bias. For example, children born in October to December may differ from those born in January to March in different aspects such as health and non-cognitive skills, which can affect mothers' employment through other channels rather than childcare alone. Thus we use the results from a 1-month bandwidth for interpretation, but we also discuss robustness checks using wider bandwidths.

In the second stage, we regress women's employment outcomes on children's childcare as follows

$$Y_{i,j} = \delta + \theta D_{i,j} + \pi' X_{i,j} + u_{i,j} \qquad (3)$$

where $Y_{i,j}$ is the employment variable of interest. Equation (3) is estimated together with Equation (2) in an IV model, where the instrumental variable for childcare is a dummy variable indicating whether the child is born in December.

Our dependent variables include both continuous and dummy variables. For continuous variables, such as (log of) wages and (log of) the number of working hours, we use the 2SLS method. Where the dependent variable is binary, 2SLS regressions can be applied for the linear probability model with a dummy endogenous variable (e.g., Angrist, 2001). However, a major limitation of the 2SLS method is that its predicted outcomes can be unrealistically smaller than -1 or larger than 1. This problem likely arises when the value of dependent variables is close to 0 or 1 (e.g., Long, 1997). To address this issue, we use the bivariate probit model (see, e.g., Wooldridge, 2010), which jointly estimates Equations (2) and (3) with maximum likelihood methods. We also apply 2SLS regressions for binary dependent variables as robustness checks but we use the bivariate probit model as our preferred model for interpretation.

For robustness checks, another way to estimate the probit model with endogenous variables is employing a control function method (Rivers and Vuong, 1988), where we first estimate the residuals from a regression of childcare on the instrument and other explanatory variables. We subsequently estimate a probit model of maternal employment on the childcare, controlling for the predicted residuals from the first stage regression and other explanatory variables. Furthermore, we perform two additional robustness checks. Firstly, we extend the bandwidth to 2 and 3 months and use the same IV model to estimate the effects of childcare on women's employment. For example, a 2-month bandwidth means that we compare women with children born in November and December with women whose children were born in January and February of the following year.

Secondly, we further extend the bandwidth to 6 and 11 months, and apply a local linear regression developed by Hahn *et al.* (2001) as follows:

$$Y_{i,j} = \beta_0 + \beta_1 D_{i,j} + \beta_2 (Z_{i,j} - c) I_{\{Z_{i,j} - c\}} + \beta_3 (Z_{i,j} - c) + \xi' X_{i,j} + v_{i,j} \qquad (4)$$

where Z is the month of birth, sorted for a time window from February of a year to November of the following year, and *c* is the threshold at January in that window. $I_{\{Z_{i,j}-c\}}$ is a dummy variable indicating whether the child was born from January onward. We estimate Model (4) using IV regressions, using $I_{\{Z_{i,j}-c\}}$ as an IV for the treatment variable $D_{i,j}$. The local average treatment effects (LATE) of childcare is estimated by β_1. For this local linear regression, the standard errors are clustered by enumeration areas and months of birth.

5. Empirical results

5.1. Testing the IV

The RDD method relies on the assumptions that the threshold of the conditioning variable is exogenous (or random), and this conditioning variable is relevant. Thus, the key identification strategy in our study is the exogeneity of being born in December versus January for children age 1-5, which means that the variable "born in December" can affect maternal employment only through the channel of childcare (given the control variables). To

test the exogeneity of being born in December, we first compare the proportion of children born in different months. Figure A.3 in Appendix shows that the proportion of children born in December is slightly lower than the proportion of children born in January. However, the difference, as we can see from this figure, is not statistically significant at the 5% level. We can examine the assumption more formally by selecting a small neighborhood around the cutoff, for example being born in November to February of the following year, and conducting a simple Bernoulli test within the neighborhood that the probability of being born in November and December equal to 1/2 (Cattaneo *et al.*, 2019). The p-value of the two-side test is 0.6173, which suggests that there is no evidence of "sorting" around the threshold and the threshold can be considered as random in the neighborhood.

One concern about the instrument "born in December" is parents' manipulation of birth dates. A well-known hypothesis suggests that being the oldest in the class is an advantage for children (Bedard and Dhuey, 2006). To the extent that parents are aware of these advantages, they may delay registering the birth of a December child by a few weeks to January so that the child can be slightly older than the other kids in his/her class. However, manipulation of birth date is not simple. According to Vietnam's Law on Civil Status 2014, to register for a birth certificate for a child, parents have to submit a document certifying the birth of the child issued by a hospital, commune health center, or midwifery house.[20]

The number of children born in October is larger than those for other months. Without in-depth studies on this issue, it is difficult to provide an accurate explanation for the higher rate of births in October. But in Vietnam, traditional New Year festivals often take place in late January and early February. People have a long holiday during these festivals and may possibly have sexual relations during this time, which can result in fertility rate increases

[20] Furthermore, the month of birth is stated on birth certificates and is difficult to manipulate. Moreover, if a number of people manipulate their children's birth month on their birth certificate in order to send their children to childcare earlier, the proportion of children with a reported birth month in December will be higher than the proportion of children with a reported birth month in January. Thus, there is no evidence of manipulation of birth months in our data. If the conditioning variable is continuous, we can use the manipulation test developed by McCrary (2008) to test the exogeneity of the conditioning variable. In our study, since the conditioning variable is binary (children born in December versus January), we simply compare the proportions of children born in these two months.

nine months later. The higher proportion of children born in October warns against using the 3-month bandwidth for the RDD model.

To further test the exogeneity of the IV, we run OLS regression of this variable on the exogenous demographic characteristics of women (age, gender, ethnicity and the number of years of schooling). Table A.3 reports the regression results, where the dependent variables indicate women whose children were born in the specified first half of the bandwidth in the sample of women whose children were born in the whole bandwidth. For example, the dependent variable in column 1 indicates "women whose children were born in December in the sample of women who gave births in December/January," column 2 indicates "women whose children were born in November and December in the sample of women who gave births from November to February," and so on. The explanatory variables are variables of women rather than variables of children.

Table A.3 shows that in the sample with a 1-month bandwidth (i.e., children born in December and January of consecutive years), being born in December is not correlated with maternal characteristics. All the explanatory variables are of small magnitude and not statistically significant at the conventional levels. However, in the samples with 2-month or 3-month bandwidths, the ethnicity and years of schooling variables are statistically significant, although these variables have very small magnitudes. For example, a 1-year increase in the number of years of schooling is associated with an increase of 0.005 in the probability of having children born in November and December (compared with children born in January and February). For the 3-month bandwidth sample, the corresponding figure is estimated at 0.003. Again, this finding advises against using 2-month or 3-month bandwidths in our RDD estimates.[21]

The second condition of the IV is a strong correlation between the IV (being born in December) and childcare. In Table 2, for easier interpretation we report the marginal effects of a probit regression of childcare on the IV and other maternal control variables (the full regression results are shown in Appendix A, Table A.4). Compared with a January birth,

[21] Do and Phung (2010) observe that children born in auspicious years, according to the Vietnamese horoscope, have two extra months of schooling and are more likely to have been planned. We pool children born in December and January across different years for analysis, thus our estimates are not affected by specific birth years. We also widen this bandwidth to include more months in our robustness checks in Section 5.2, which lend further support to our estimation results.

being born in December of the previous year increases the probability of attending childcare by 0.092, and the estimate is strongly statistically significant at the 1% level.

Since younger children need more care and attention than older children, we estimate the effect of childcare for children of different ages for children age from 1 (and born in December) to 3 (born in January), and children age from 3 (born in December) to 5 (born in January). In these two separate samples of children, the instrument is significant at the 1% level. We also perform Cragg-Donald and Kleibergen-Paap weak identification tests on the instruments. All the test statistics are much higher than the rule-of-thumb F value of 10 (Staiger and Stock, 1997), indicating that the instruments are strong (Table A.4).

5.2. Main estimation results

Panel A of Table 3 reports the estimated impacts of childcare on maternal employment outcomes, using the IV regressions. There are 10 outcome variables, and we estimate the effects of childcare on these outcomes for three different samples of children of different ages. Each cell in Table 3 shows only the marginal effects of the estimated effects of childcare. Tables A.5 to A.8 in Appendix A present the full regression results.

We find that childcare has a statistically insignificant effect on women's LFP (Table 3, Panel A, row 1 for the first outcome). A possible reason is that the female work rate in Vietnam is very high at 94% in 2016 (Table 1), and a large proportion of these workers are self-employed. If children cannot attend childcare, women can care for them and at the same time be self-employed. Consequently, the effects of childcare on women's working status are not statistically significant. This result is consistent with the result that childcare has no significant effects on women's number of monthly working hours (row 7). But interestingly, our imprecise point estimate suggests that childcare raises women's number of monthly work hours by roughly 30 hours (=0.16*188, based on Table 1 and Table 3), which is close to Berlinski *et al.*'s (2011) estimate for Argentine women.[22]

The effects of childcare on engaging in skilled work are small and statistically insignificant (row 5). Obtaining work skills takes a long time; as such, children's attendance

[22] Berlinski *et al.*'s (2011) estimate is 7.8 additional hours of work per week, which roughly translates into slightly more than 30 hours of work per month.

at childcare does not likely improve women's work skills. Moreover, skilled workers may be self-employed and provide childcare for their children at the same time. In our sample, 50% of skilled workers were self-employed.

However, we find strong effects of childcare on women's wage-earning employment (row 2). Childcare increases the probability of having a wage-earning job by 0.41 (or 41 percentage points). Childcare also has negative effects on self-employed farm work (row 4), and somewhat negative effects on self-employed non-farm work (row 3) but the latter effects are mostly statistically insignificant. These results suggest that women may switch from self-employed farm work to wage-earning work, which can provide higher incomes and more job stability.[23] There is also a significant positive effect of childcare on women's formal jobs (row 6). Having a child in childcare increases the probability of women having a formal job by 0.26. Thus, these results suggest that sending children to childcare helps women find higher-quality and more stable employment.

For wage earners, the effects of childcare on hourly and monthly wages are positive but not statistically significant (rows 8 and 9). While our imprecise point estimate for a monthly wage increase of 53% is larger than the corresponding estimate of 20% for Mexican women in Calderon (2014), it appears consistent with an increase of nearly 50% for Kenyan women obtained by Clark *et al.* (2019). We also find a marginally statistically significant effect on annual wages (row 10). The increase in total wages may be due to an increase in the number of women in formal jobs as well as an increase in productivity.

In Panel A of Table 3, we examine the contemporaneous relationship between childcare and maternal employment (i.e., in the same year). But another policy-relevant question is whether childcare has an ongoing effect on maternal employment. To examine the medium-term effect, we regress the current employment outcomes of women on childcare recorded in the previous survey round, using panel data from two consecutive survey rounds.[24] This approach aims to measure the 2-year lagged effects of childcare on parental employment outcomes, using the same model specification as with Table 3, Panel A. The

[23] Dang (2012) and Cunningham and Pimhidzai (2018) offer further discussion on the differences between self-employment and wage work, particularly between ethnic majority and minority groups in Vietnam.
[24] The VHLSSs have a rotating panel design where 50% of households sampled in one survey round are re-interviewed in the next one. For example, roughly 50% of households in the 2010 VHLSS are resampled in the 2012 VHLSS. The attrition rate in two consecutive surveys is around 8%.

estimation results, shown in Panel B of Table 3, do not suggest a significant effect from childcare on the working status or wages of women after two years. However, children's attendance in childcare has strong lagged effects on the probability of women taking a wage-earning job: it increases this probability by 0.38 after two years. The effects on self-employed farm work are negative, indicating a movement from farm work to wage-earning employment.

Still, part of the impacts shown in Table 3, Panel B may be attributed to the direct contemporaneous impacts of children being in childcare in the current survey round. To address this concern, we restrict the estimation sample to mothers of the children that are currently sent to childcare. Our assumption is that, if these contemporaneous impacts of childcare are similar for children that are currently sent to childcare, the estimated impacts could pick out the 2-year lagged effect. Estimations results, shown in Appendix A, Table A.9, remain very similar.[25]

5.3. Robustness and further falsification analysis

In this section, we report a number of robustness checks and falsification tests. Firstly, we examine the sensitivity of the estimation results to different sets of control variables. We estimate a bivariate probit model without any control variables and a bivariate probit model with additional control variable and province fixed-effects. Table A.10 in Appendix shows that results are very similar to the models which are used for interpretation (Table 3).

Secondly, we explore different estimation models. In Table 3, we show the estimates from the bivariate probit model for binary dependent outcomes. This model is suitable and efficient for models with a binary dependent variable and a binary endogenous variable, but it relies on specific parametric assumptions of the distribution of errors. Thus, we re-estimate the effects of childcare using 2SLS and control function models for robustness checks. For the 2SLS model, both the dependent and endogenous variables are estimated using linear

[25] To keep a reasonable estimation sample size, we only run these regressions for the pooled sample. An alternative to restricting the estimation sample to those that are currently sent to childcare would be to consider a sample of children that were sent to childcare two years ago, but are not currently in childcare. In other words, we could compare mothers with children attending children two years ago and those with children not attending two years ago, and both the two groups of children are not attending child care or school in the current year. However, since there is (almost) universal primary school enrolment in Vietnam, these data would not exist.

probability models. For control function models, we implement two different types.[26] We also estimate the local linear regression in Equation (4) using the bandwidth of 6 and 11 months. Table A.11 in Appendix A report the estimation results, which are qualitatively similar to those shown in Table 3. Childcare has a positive effect on the probability of having a wage-earning job and a formal job. The sign and magnitude of the effects are rather similar for different models.

Third, we further examine the sensitivity of the estimates to the bandwidth selection. We use a 2-month bandwidth (i.e., comparing children born in November and December with those born in January and February in the following year) and a 3-month bandwidth (i.e., comparing children born in October to December with those born in January to March in the following year). Table A.12 in the Appendix shows that the estimates of the effects of childcare on women's employment are similar using different bandwidths. In fact, the estimated effects on wages are more significant in models using 2- and 3-month bandwidths than in models using a 1-month bandwidth, perhaps because there are more observations in models using 2- and 3-month bandwidths.

Fourth, Table A.13 reports the reduced-form regressions of maternal employment on the instrument (i.e., children born in December), using the sample of children born in December and those born in January of the following years. It shows that women who have children in December are more likely to have a wage-paying job and are less likely to engage in self-employed farm work than women with children born in January. Table A.13 reports the coefficient of the instrument from 30 regressions (3 samples multiplied by 10 outcomes). Childcare is significant at the 1% level in 6 regressions, accounting for 20% of the total number of regressions.

Fifth, an issue with our instrument is that on average, children born in December are still 1 month older than those born in January. One may argue that a 1-month difference in age is small but may still affect maternal employment. To test this argument, we run

[26] In the first type, following Rivers and Vuong (1988), we first regress childcare on the instrument and other explanatory variables using OLS, and estimate the residuals from this regression. Next, we run a probit model of maternal employment on the childcare variable, the predicted residuals, and other explanatory variables. In the second type, we regress a probit model of childcare on the instrument and other explanatory variables, and estimate the generalized residuals (Wooldridge, 2015). We then run a probit model of maternal employment on the childcare variable, the generalized residuals, and other explanatory variables. We estimate the standard errors using bootstraps with 200 replications.

regressions of women's employment on a dummy variable indicating women who have children 1 month older than others. In these regressions, children are of the same age. For example, we use a sample of women with children born in January and those with children born in February. We repeat the analysis for each pair of months up to the sample of women with children born in November and those with children born in December. The instrument in this case is "children born one month earlier." We conduct regressions for all 10 outcomes and estimate the percentage of regressions in which childcare is significant at the 1% significance level. We repeat this analysis for different gaps in children's birth month, including gaps of 2 months, 3 months, and 1 to 3 months.

Figure A.4 in Appendix shows the distribution of the p-values of the variable "born earlier" in these regressions. For the birth month gap from 1-3 months, only 3.3% of regressions show a significant effect of "being born earlier" on parental outcomes. For a gap of 1 month, 1.8% of regressions show a significant effect of "being born 1 month earlier" on parental outcomes. No regressions show that the effect of being born earlier is significant at the 1% level. Thus, there is no evidence of the effect of being born earlier (for children of the same age) on parental outcomes.

Finally, we conduct a balance test by comparing variables (for both the demographics and outcomes) between mothers who have children aged 1 and born in December and those who have children aged 0 and born in January in the following year. Almost all of these children are not attending childcare. Thus, the differences in the variables between the two groups of mothers are not affected by childcare. The results, presented in Table A.14 in Appendix A, show that the differences in all the variables between the two mother groups are not statistically significant at the 1% level or higher. These results confirm the exogeneity of the IV and similarity between mothers with children born in December and those with children born in January of the following year.

5.4. Spill-over and heterogeneous effects

Table 4 presents the estimates of the impacts of childcare on several household outcomes. Taking advantage of childcare has a positive effect on per capita income. Childcare increases the quality of women's employment and their wages and as a result, increases household

income. This income effect can in turn help reduce poverty.[27] The probability of being poor is reduced by 0.22 when children are placed in childcare. But these effects are only marginally statistically significant at the 10% level.

It is common in Vietnam for grandparents who live with their children to take care of the grandchildren. In the 2016 VHLSS, in approximately 21% of households, grandparents live with their children. We test whether formal childcare in a center can substitute for informal childcare from grandparents by running a regression of a dummy variable "living with grandparents" on childcare. The effect of childcare is negligible and is not statistically significant. There are no significant effects on parental migration and household size. Thus, childcare affects maternal employment but not household demographic composition and migration outcomes.

To investigate the heterogeneous effects of childcare, we include a number of interaction terms between childcare and explanatory variables. If the interacted variable is discrete, we convert it into a set of dummy variables and include the interactions between childcare and the set of dummy variables in one regression. These interaction variables are also endogenous, so we use the interaction terms between the instrument for childcare (children born in December) and the interacted variables as instruments for the interaction variables. For simplicity, we estimate models with interactions using the control function method. In this method, the endogenous part of childcare is controlled for by the residuals from the first-stage regression.

Since our estimation model is probit, the interaction effects and their significance level are not the same as the marginal effects of the interaction terms in the linear model and can vary across different observations (Ai and Norton, 2003; Norton *et al.*, 2004). Consequently, we follow the cited methods to estimate the effects of interactions between childcare and explanatory variables.[28] Since the interaction effects vary across observations, we report the average size and the average z-statistics of the interaction effects and use these

[27] A household is defined as poor if its per capita income is below a poverty line. In 2016, this poverty line was set by the government of Vietnam to equal VND 700,000 and 900,000/person/month for rural and urban households, respectively. The poverty rate in the full sample of the 2016 VHLSS is 7.7%. We further explore whether the poverty-reducing impacts of childcare are stronger for certain disadvantaged groups such as ethnic minorities or unemployed individuals by interacting these variables with childcare. Estimation results (not shown), however, are not statistically significant.

[28] The interaction effects are computed using command 'inteff' in Stata (see Norton *et al.* 2004).

averages for interpretation. Table 5 presents the averages of the interaction effects and z-statistics. For simplicity, we investigate the heterogeneous effects of childcare only on the labor market participation of women (i.e., the dependent variable is women with a wage-paying job).

We first include the interactions between childcare and women's demographic variables. The effect of childcare does not differ for age (Model 1 in Table 5). However, more educated women are more likely to have a wage-earning job than less educated women (Model 2), which is consistent with earlier findings by Schlosser (2005). The effect of making use of childcare centers is lower for ethnic minority women than for Kinh women (Model 3), perhaps because ethnic minorities are less likely to have similar job opportunities.

We also examine whether childcare effects differ for the gender of children. Model 4 in Table 5 shows that the effect on maternal employment of boys attending childcare is slightly lower than that of girls. In our data set, the rate of childcare is 48% for boys and 49% for girls. The difference is small but still statistically significant. Vietnam is a country with a preference for boys, especially in rural areas (e.g., Guilmoto, 2012; Nguyen and Tran, 2017), which may result in women having to spend more time taking care of boys than girls. As a result, the effect on maternal employment of boys attending childcare is smaller than that of girls attending childcare.

Children's order of birth is negatively correlated with maternal employment, since a higher birth order implies a larger number of children and having more children is associated with a lower probability of labor market participation. However, the interaction between childcare and the birth order of the child is negative but not statistically significant (Model 5). Children may receive care from grandparents. Several studies show that informal childcare provided by grandparents can increase women's labor supply in China, a neighboring country (Maurer-Fazio *et al.*, 2011). Thus, we include interactions between childcare use and living with grandparents, but this interaction effect is not statistically significant (Model 6), which is consistent with the estimation results with Table 4 discussed earlier. There can be differences in quality between public childcare and private childcare. We thus interact childcare and a variable indicating that the commune has a public childcare center, but this interaction effect is not statistically significant (Model 7).

We find a smaller effect for childcare in communes which are far from town (Model 8). One possible reason is that employment opportunities and wages are higher in areas that are closer to towns, which can help increase the effects of childcare on women's employment in these areas. Indeed, the effects of childcare on maternal employment may depend on the opportunity costs of staying at home (i.e., not participating in the labor market) to take care of children. In Model 9, we test the interaction of childcare with the districts' average income. The interaction marginally significant but is positive, indicating that the effect of childcare use is greater in areas with higher income levels.[29]

6. Conclusion

In this paper, we offer the first study that rigorously investigates the prevalence of childcare, and the effect of pre-school childcare on maternal employment in Vietnam. We find that the percentage of children attending childcare is less than 1% and 3% respectively, for children younger than age 1 and at age 1, although this figure improves for older children. We find childcare to have a small, insignificant effect on women's LFP, which may be due to the high rate of self-employment in the country. However, we find that childcare has a strong effect on women's quality of employment. Specifically, the use of childcare increases the probability of women having a wage-earning job by 41 percentage points and increases the probability of their having a formal job by 26 percentage points. We also find that childcare has heterogeneous effects and differs for women of different characteristics. In particular, these effects are greater for ethnic majority and more educated women, younger children, and for areas with higher wages or with greater opportunity costs for not participating in the labor market.

These findings point to the importance of accessible childcare services in both enhancing women's labor market outcomes and reducing the gender gaps. This has important policy implications, especially given that in Vietnam, women are currently given at most 6

[29] Mean wages and per capita income at the district levels are obtained from Lanjouw *et al.* (2017). We also test the interaction of the childcare and commune-level variables such as availability of preschools in the commune and whether the village is accessible by car during the previous 12 months. The VHLSSs contain commune-level data for rural areas but not for urban areas. Thus, we use the rural sample to estimate the models, including interactions with commune-level variables. Both the interactions are negative, but not statistically significant (not shown).

months' maternity leave and the existing supply of public childcare may be inadequate. In particular, providing childcare in areas with higher wages can be particularly beneficial for women's access to a wage job. The opportunity costs for not participating in the labor market will be larger for women as the economy develops, which is likely to amplify the beneficial impacts of childcare.

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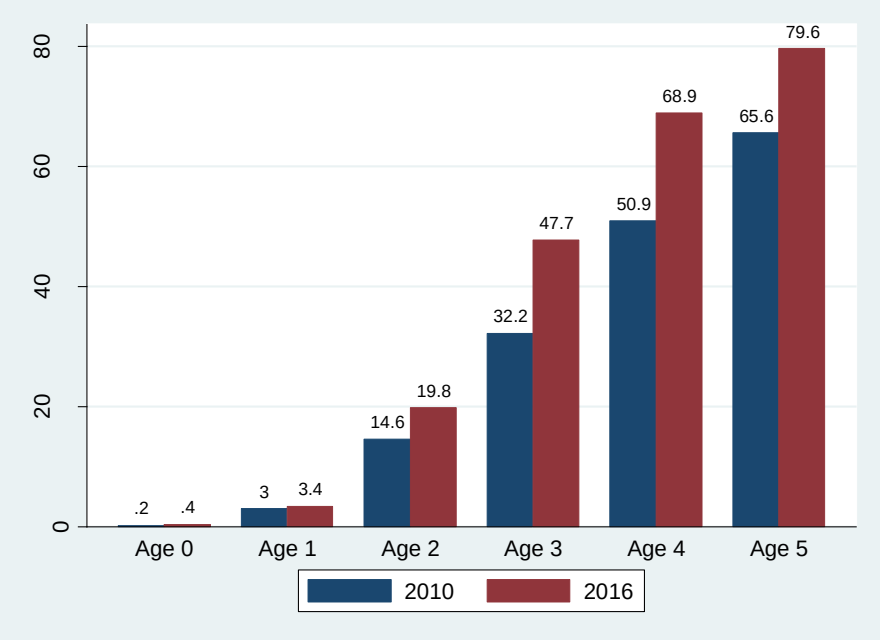
Figure 1. Percentage of children attending childcare

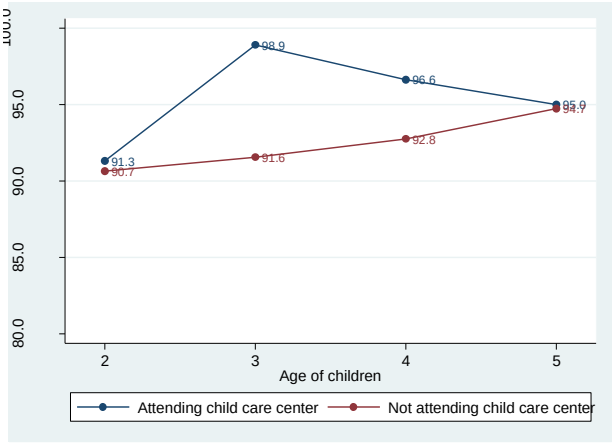


Note: This figure presents the level of childcare by children age 1-5.

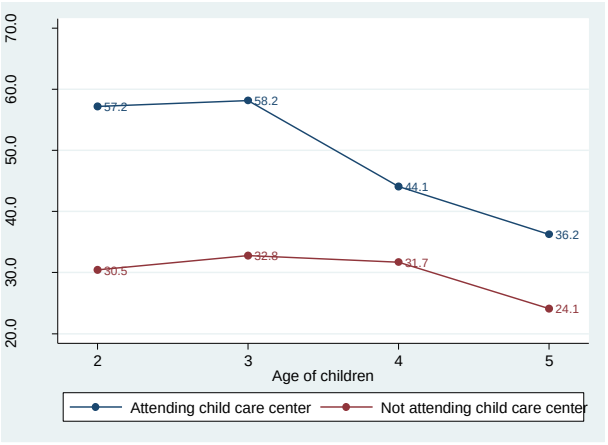
Source: Authors' estimation from VHLSS 2010 and 2016.

Figure 2. Maternal employment and children's attendance at childcare centers

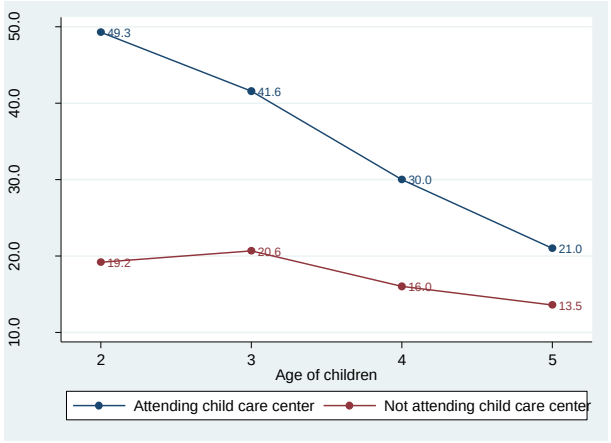
Panel A. Percentage of women working and childcare



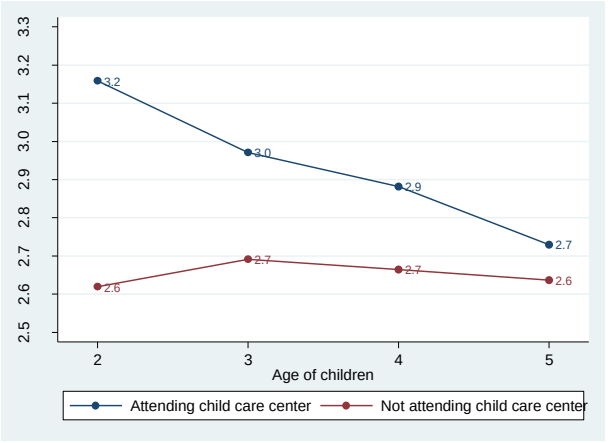
Panel b. Percentage of women have a wage job and childcare



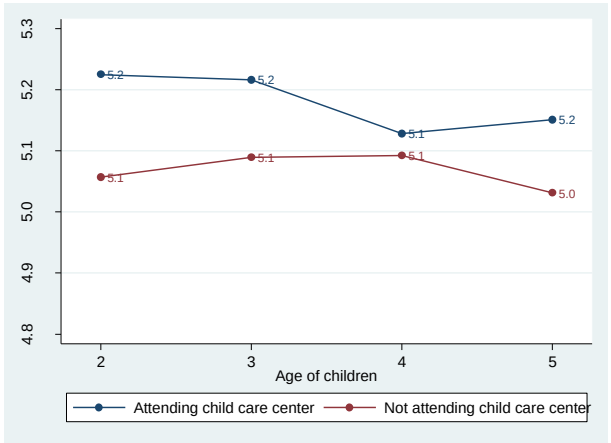
Panel C. Percentage of women with a formal job and childcare



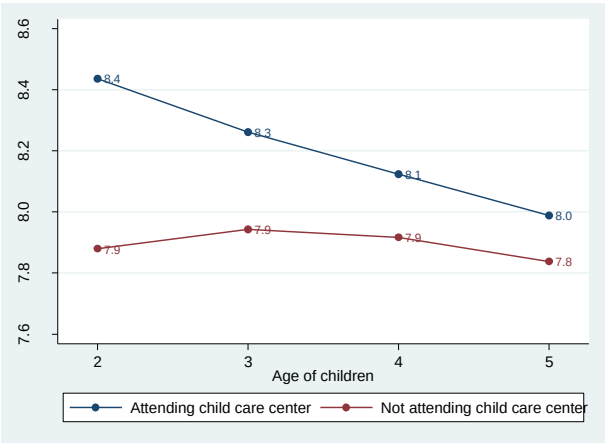
Panel D. Working hours of women and childcare



Panel E. Hourly wage of women and childcare



Panel F. Monthly wage of women and childcare

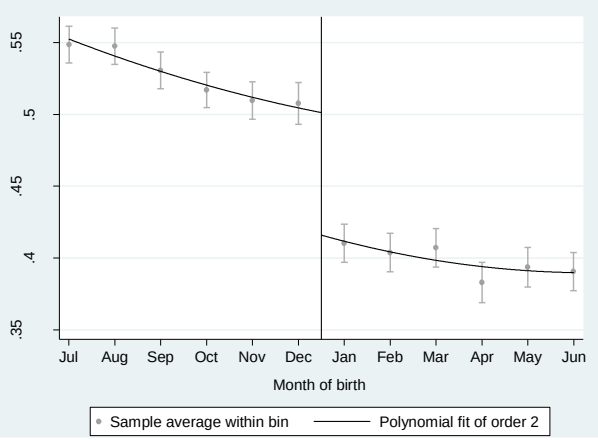


Note: The vertical axis indicates the employment variables of the parent, and the horizontal axis gives the children's age.

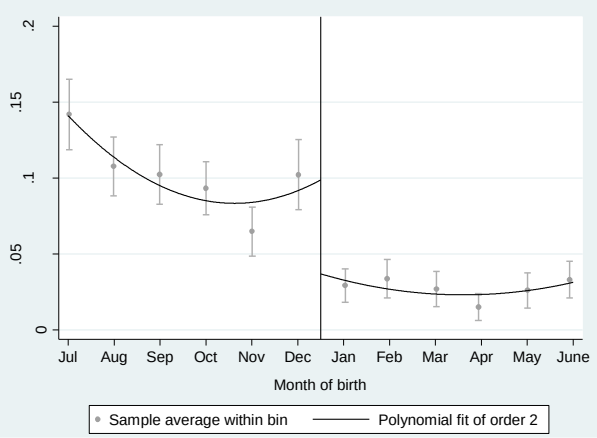
Sources: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Figure 3. The proportion of enrolled school-age children and month of birth

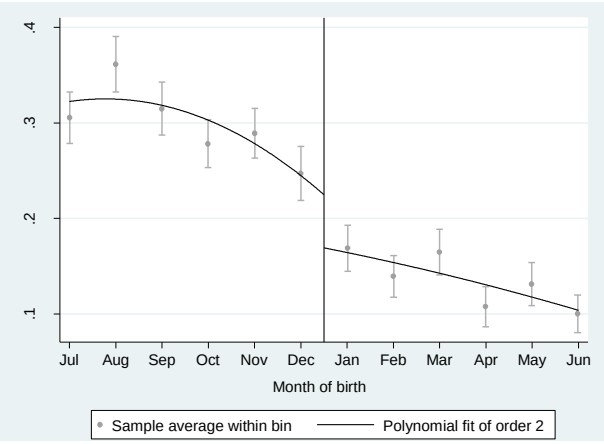
Panel A. All children



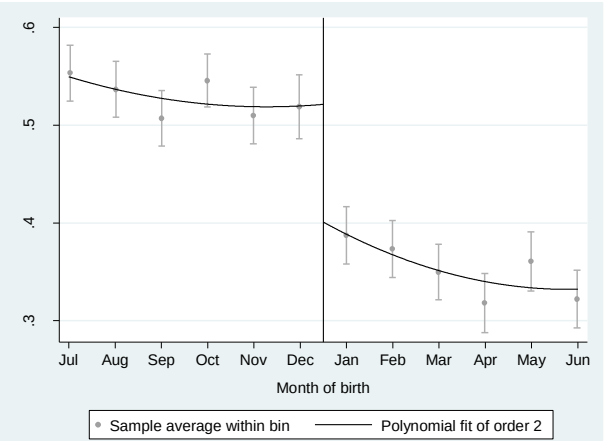
Panel B. Sample of children age 1 and 2



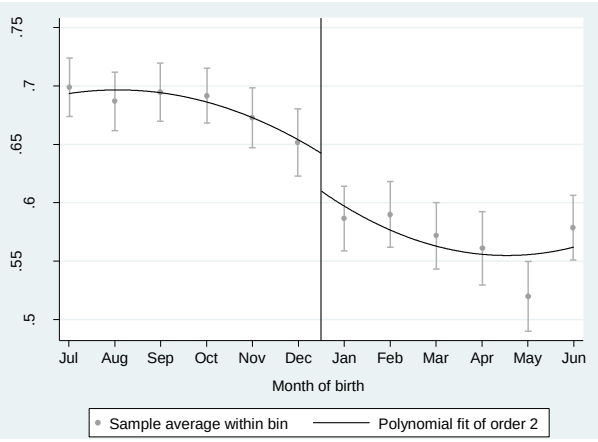
Panel C. Sample of children age 2 and 3



Panel D. Sample of children age 3 and 4



Panel E. Sample of children age 4 and 5



Note: The vertical axis gives the proportion of children attending childcare, and the horizontal axis presents the birth months of children with contiguous birth months. Children born in December are 1 month older than those born in January.

Sources: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table 1. Employment outcomes of women

Variables	VHLSS 2010	VHLSS 2012	VHLSS 2014	VHLSS 2016
% working	91.8	93.2	92.5	93.5
	(0.4)	(0.5)	(0.5)	(0.5)
% in a wage-earning job	30.9	33.4	35.5	37.6
	(0.7)	(1.0)	(1.0)	(1.1)
% self-employed in a nonfarm job	16.5	14.6	13.7	18.0
	(0.6)	(0.7)	(0.7)	(0.8)
% self-employed in a farm job	44.4	45.1	43.4	37.9
	(0.8)	(1.0)	(1.0)	(1.1)
% in a skilled job	45.0	47.2	49.3	53.5
	(0.8)	(1.0)	(1.0)	(1.1)
% in a formal job	15.1	18.6	21.5	23.7
	(0.7)	(0.8)	(1.0)	(0.9)
Number of working hours per month	180.0	187.2	188.7	188.0
	(1.3)	(1.5)	(1.6)	(1.6)
Hourly wage (thousand VND)	18.1	19.5	20.4	24.2
	(0.8)	(0.7)	(0.5)	(0.0)
Monthly wage (thousand VND)	3252.4	3554.0	3845.2	4404.3
	(89.1)	(99.1)	(79.9)	(0.0)
Yearly wage (thousand VND)	39013.0	41878.6	46334.3	52749.0
	(1434.5)	(1331.3)	(1131.8)	(0.0)

Note: This table reports the employment variables of women with children age 1 to 5. Variables of wage-paying jobs, skilled jobs, and formal jobs are defined using the main occupation over the past 12 months. Employment consists of wage-paying employment, self-employed non-farm work, and self-employed farm work. Wages are defined as the total wages, including main and secondary jobs. Standard errors of the mean are in parentheses. Wages are measured in 2016 prices.
Source: Authors' estimation from VHLSS 2012, 2014, and 2016.

Table 2. First-stage probit regression of childcare on the instrumental variable (marginal effects)

Explanatory variables	Dependent variable is childcare		
	Pooled sample	Children age 1-3	Children age 3-5
Instrument (child born in December)	0.092***	0.080***	0.097***
	(0.017)	(0.018)	(0.024)
Age	0.046***	0.033**	0.048***
	(0.013)	(0.014)	(0.017)
Age squared	-0.639***	-0.548**	-0.697***
	(0.189)	(0.213)	(0.249)
Ethnic minority	0.021	-0.029	0.049
	(0.022)	(0.021)	(0.032)
Number of years of schooling	0.016***	0.012***	0.022***
	(0.002)	(0.002)	(0.003)
Dummy year 2010	*Reference*		
Dummy year 2012	0.025	-0.033	0.013
	(0.021)	(0.021)	(0.032)
Dummy year 2014	0.039*	0.015	0.089***
	(0.022)	(0.024)	(0.033)
Dummy year 2016	0.078***	0.025	0.088***
	(0.023)	(0.024)	(0.032)
Observations	3,863	1,718	2,145
Pseudo R2	0.029	0.072	0.038

This table reports the marginal effects from the logit regression of childcare on the instrumental variable and control variables of women. The observations in these regressions are women of children age 1-6. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level. The marginal effects are computed using the "margin" command in Stata. For childcare (or other dummy variables), the marginal effect is the estimated discrete change in the probability of the dependent variable due to the change in the childcare (or other dummy variables). For a continuous explanatory variable, it is the estimated partial derivative of the dependent variable with respect to the explanatory variable (evaluated at the mean value of all the other explanatory variables).
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table 3. The effect of childcare on women's employment outcomes

Dependent variables	Panel A. Short-term effects			Panel B. Medium-term effects		
	All children	Children age 1-3	Children age 3-5	All children	Children age 1-3	Children age 3-5
Bivariate probit model (marginal effects)						
Working	-0.110	-0.170	-0.128	-0.016	0.037	0.146
	(0.126)	(0.144)	(0.090)	(0.110)	(0.060)	(0.124)
In wage-paying job	0.411***	0.490***	0.408***	0.377***	0.477***	0.333***
	(0.010)	(0.033)	(0.021)	(0.024)	(0.038)	(0.087)
In self-employed	-0.103	-0.240**	0.070	0.043	-0.004	0.089
nonfarm work	(0.105)	(0.092)	(0.149)	(0.108)	(0.150)	(0.145)
In self-employed farm	-0.454***	-0.563***	-0.440***	-0.419***	-0.384***	-0.297***
work	(0.011)	(0.053)	(0.008)	(0.032)	(0.078)	(0.103)
In skilled work	0.108	-0.146	0.043	-0.055	0.187	-0.239
	(0.835)	(1.260)	(0.238)	(0.384)	(0.143)	(0.157)
In a formal job	0.257***	0.172	0.264***	0.149	0.382	0.017
	(0.035)	(0.229)	(0.077)	(0.206)	(0.349)	(0.296)
2SLS						
Log of number of	0.155	0.378	-0.009	0.293	0.489	0.206
monthly working hours	(0.209)	(0.358)	(0.255)	(0.312)	(0.470)	(0.463)
Log of hourly wage	0.572	0.948	0.141	-0.275	-0.104	-0.421
	(0.460)	(0.649)	(0.568)	(0.478)	(0.511)	(0.842)
Log of wage for the last	0.525	0.951	0.113	-0.078	0.071	-0.286
month	(0.410)	(0.586)	(0.521)	(0.523)	(0.580)	(0.895)
Log of total wage for the	0.903*	1.165	0.645	-0.068	0.397	-0.527
past 12 months	(0.524)	(0.743)	(0.666)	(0.678)	(0.733)	(1.183)

This table reports estimation results from the bivariate probit and 2SLS regressions of maternal employment outcomes on childcare and the control variables shown in Table 2. The marginal effect is the estimated discrete change in the probability of the dependent variable due to the change in the childcare. The observations in these regressions are women of children age 1-5. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table 4. 2SLS regression of household-level outcomes on childcare

	Log of income per capita	Household is poor	Living with grandparents	Women are migrating	Household size
Childcare	0.428*	-0.222*	0.009	0.029	0.047
	(0.237)	(0.124)	(0.053)	(0.050)	(0.363)
Ethnic minority	-0.970***	0.547***	0.021***	-0.017***	0.527***
	(0.030)	(0.018)	(0.008)	(0.005)	(0.058)
Dummy year 2010	*Reference*				
Dummy year 2012	0.328***	-0.011	0.039***	-0.008	0.112**
	(0.034)	(0.019)	(0.006)	(0.006)	(0.050)
Dummy year 2014	0.530***	-0.070***	0.034***	-0.007	0.094*
	(0.039)	(0.021)	(0.007)	(0.007)	(0.057)
Dummy year 2016	0.678***	-0.106***	0.041***	0.005	0.127**
	(0.041)	(0.021)	(0.009)	(0.009)	(0.061)
Constant	9.316***	0.323***	-0.008	0.014	4.193***
	(0.101)	(0.053)	(0.022)	(0.021)	(0.153)
Observations	3,863	3,863	3,863	3,863	3,863

Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table 5. The effect of interactions between child schooling and demographic variables on the probability o job (probit models)

Interaction variables	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
Childcare * age	-0.003						
	(-0.330)						
Childcare * schooling years		0.010**					
		(2.222)					
Childcare * ethnic minority			-0.071*				
			(-1.744)				
Childcare * boy				0.004*			
				(1.794)			
Childcare * birth order					-0.038		
					(-1.439)		
Childcare * grandparents living in household						-0.063	
						(-1.028)	
Childcare * public childcare center							-0.104
							(-1.415)
Childcare * distance to nearest town							
Childcare * log of district per capita income							
Control variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3,863	3,863	3,863	3,863	3,863	3,863	3,863
Pseudo R2	0.103	0.104	0.103	0.103	0.106	0.106	0.105

Note: This table reports the interaction effect between childcare and explanatory variables in from probit regressions of the probability of women v jobs on childcare and the interactions between childcare and other explanatory variables. We first run a probit model of childcare on the instrumen variables, and estimate the generalized residuals (Wooldridge, 2015). We subsequently run a probit model of maternal employment on the childcare v residuals, the interactions, and other explanatory variables.

Village variables (public childcare center, preschool, distance to the nearest town, accessible by car) are available only for the rural sample. Thus, the in regressions using the interaction between childcare and these village variables is lower than other regressions.

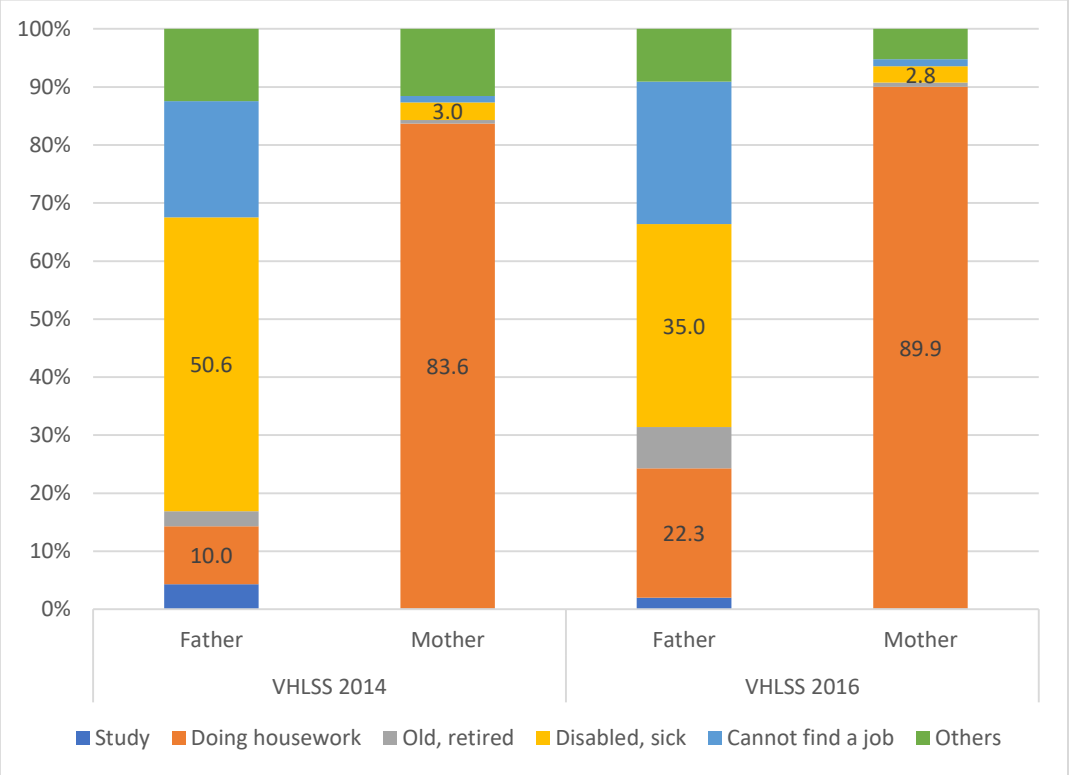
The interaction effects are computed using command 'inteff' in Stata. This table presents the average of the interaction effects across the observation statistic is reported in parentheses.

*** the absolute value of Z-statistic > 2.57, ** >1.96, * > 1.65.

Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Appendix A: Additional Tables and Figures

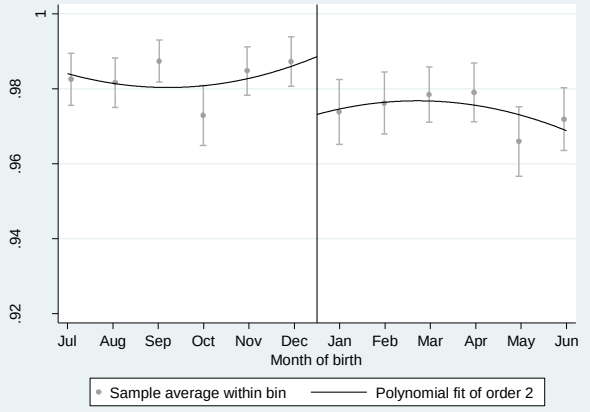
Figure A.1. The main reasons for not working



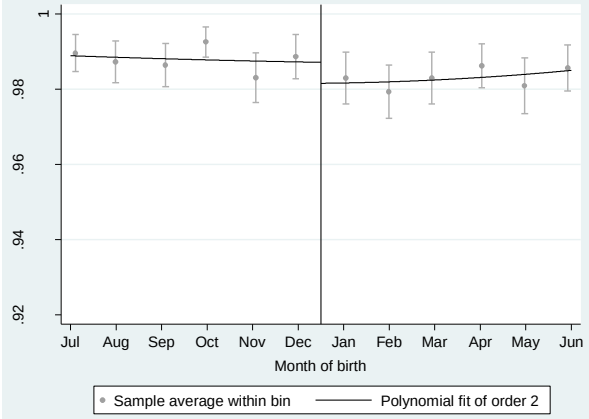
Source: Authors' estimation from VHLSSs 2014 and 2016.

Figure A.2. The proportion of enrollment for children aged 7-9 and month of birth

Panel A. Children aged 7 and 8

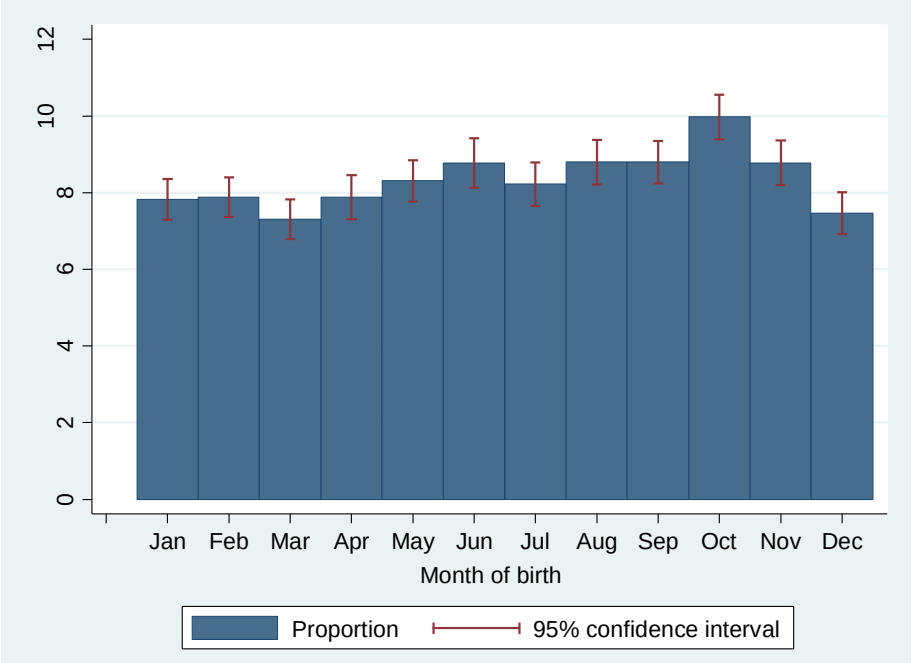


Panel B. Children aged 8 and 9



Sources: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Figure A.3. Distribution of children by month of birth

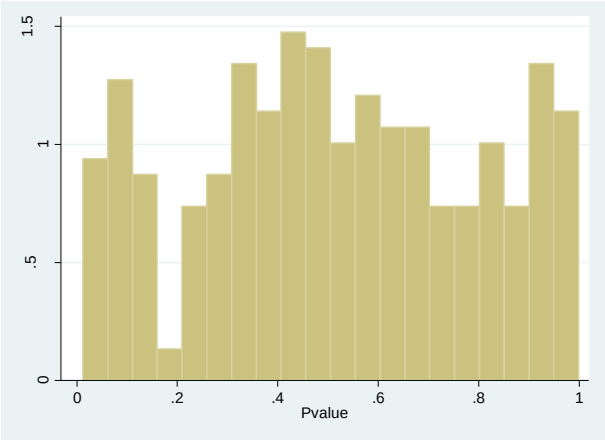


Note: The figure presents the proportion and the 95% confidence interval of the proportion of children age 1-6 by month of birth.

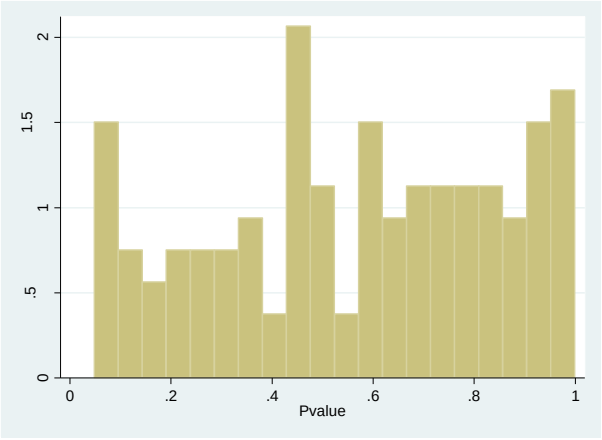
Sources: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Figure A.4. P-value in the placebo analysis

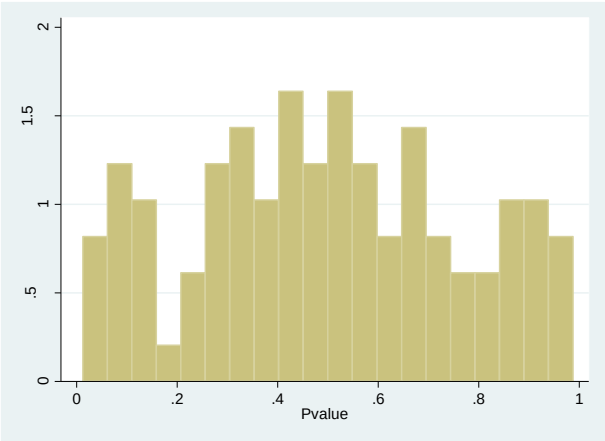
Panel A. 1-3 months difference: 3.3% with P-value<=0.05



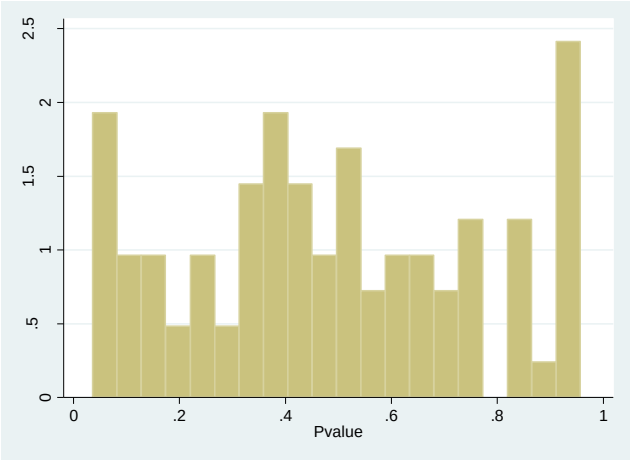
Panel B. 1-month difference: 1.8% with P-value<=0.05



Panel C. 2-month difference: 3.0% with P-value<=0.05



Panel D. 3-month difference: 5.5% with P-value<=0.05



Note: This figure shows the distribution of the p-value of the variable "born earlier" in the reduce-form regressions of maternal employment on "born earlier" (1, 2, or 3 months earlier).
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.1. Employment outcomes of men

Variables	VHLSS 2010	VHLSS 2012	VHLSS 2014	VHLSS 2016
% working	98.9	99.1	99.2	98.9
	(0.2)	(0.2)	(0.2)	(0.3)
% in a wage-earning job	43.3	48.8	50.8	52.2
	(0.8)	(1.0)	(1.1)	(1.1)
% self-employed in a nonfarm job	13.7	14.3	13.1	16.0
	(0.5)	(0.7)	(0.7)	(0.8)
% self-employed in a farm job	41.9	36.0	35.3	30.6
	(0.8)	(1.0)	(1.0)	(1.0)
% in a skilled job	58.8	59.8	61.2	61.9
	(0.8)	(1.0)	(1.0)	(1.1)
% in a formal job	15.8	18.0	18.3	20.5
	(0.6)	(0.8)	(1.1)	(0.9)
Number of working hours per month	199.8	209.0	207.2	206.9
	(1.2)	(1.5)	(1.5)	(1.6)
Hourly wage (thousand VND)	17.5	21.8	23.6	27.0
	(0.6)	(0.7)	(0.6)	(1.3)
Monthly wage (thousand VND)	3500.4	4374.1	4684.8	5360.3
	(97.7)	(102.9)	(97.5)	(128.9)
Yearly wage (thousand VND)	42233.3	50885.2	53548.0	62592.6
	(1468.9)	(1506.1)	(1321.4)	(1690.6)

Note: This table reports the employment variables of men with children age 1 to 5. Variables of wage-paying jobs, skilled jobs, and formal jobs are defined using the main occupation over the past 12 months. Employment consists of wage-paying employment, self-employed non-farm work, and self-employed farm work. Wages are defined as the total wages including main- and secondary jobs. Standard errors of the mean are in parentheses. Wages are measured in 2016 prices.
Source: Authors' estimation from VHLSSs 2012, 2014, and 2016

Table A.2. Employment outcomes of women by residence area

Variables	Rural			Urban		
	VHLSS 2012	VHLSS 2014	VHLSS 2016	VHLSS 2012	VHLSS 2014	VHLSS 2016
% working	95.9	94.6	95.2	87.8	90.0	86.9
	(0.4)	(0.5)	(0.6)	(0.0)	(0.0)	(0.0)
% in a wage-earning job	27.1	30.9	31.3	50.0	50.0	54.0
	(1.0)	(1.1)	(1.2)	(0.0)	(0.0)	(0.0)
% self-employed in a nonfarm job	12.2	10.9	16.1	23.7	23.5	22.2
	(0.7)	(0.7)	(1.0)	(0.0)	(0.0)	(0.0)
% self-employed in a farm job	56.5	52.8	47.8	14.1	16.6	10.7
	(1.1)	(1.2)	(1.3)	(0.0)	(0.0)	(0.0)
% in a skilled job	40.7	42.0	45.2	72.1	71.6	70.9
	(1.1)	(1.2)	(1.3)	(0.0)	(0.0)	(0.0)
% in a formal job	12.4	14.9	17.4	36.4	37.0	40.8
	(0.7)	(0.9)	(1.0)	(0.0)	(0.0)	(0.0)
Number of working hours per month	181.1	185.4	184.4	202.3	204.5	203.5
	(1.8)	(1.8)	(1.9)	(0.0)	(0.0)	(0.0)
Hourly wage (thousand VND)	15.1	16.9	18.8	26.6	27.1	31.1
	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)
Monthly wage (thousand VND)	2850.0	3216.9	3607.1	5004.2	5044.3	5561.0
	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)
Yearly wage (thousand VND)	30899.6	36132.2	40304.2	66569.1	67196.2	73370.4
	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)	(0.0)

Note: This table reports the employment variables of women with children age 1 to 5. Variables of wage-paying jobs, skilled jobs, and formal jobs are defined using the main occupation over the past 12 months. Employment consists of wage-paying employment, self-employed non-farm work, and self-employed farm work. Wages are defined as the total wages, including main and secondary jobs. Standard errors of the mean are in parentheses. Wages are measured in 2016 prices.
Source: Authors' estimation from VHLSS 2012, 2014, and 2016.

Table A.3. OLS regression of the instrument on demographic variables of women

Explanatory variables	Dependent variables		
	Children born in December (one-month bandwidth)	Children born in November and December (two-months bandwidth)	Children born in October to December (three-months bandwidth)
Age	0.000	-0.010	-0.009
	(0.012)	(0.008)	(0.006)
Age squared	-0.012	0.122	0.121
	(0.178)	(0.122)	(0.091)
Ethnic minority	-0.037	-0.033**	-0.023*
	(0.024)	(0.016)	(0.012)
Number of years of schooling	0.003	0.005***	0.003**
	(0.002)	(0.002)	(0.001)
Dummy year 2010	*Reference*		
Dummy year 2012	-0.036	-0.015	-0.000
	(0.024)	(0.016)	(0.013)
Dummy year 2014	-0.065***	-0.018	0.000
	(0.025)	(0.017)	(0.013)
Dummy year 2016	-0.020	0.015	0.016
	(0.025)	(0.017)	(0.013)
Constant	0.488**	0.650***	0.663***
	(0.197)	(0.134)	(0.102)
Observations	3,863	8,159	12,730
R-squared	0.004	0.004	0.002

Heteroskedasticity-robust standard errors in parentheses. The standard errors are corrected for sampling weights and cluster correlation at the commune level.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.4. First-stage of school attendance on the instrument

Explanatory variables	Probit model			OLS		
	Pooled sample	Children age 1- 3	Children age 3-5	Pooled sample	Children age 1- 3	Children age 3- 5
Instrument (child born in December)	0.090***	0.081***	0.094***	0.248***	0.390***	0.246***
	(0.017)	(0.018)	(0.023)	(0.046)	(0.085)	(0.061)
Age	0.041***	0.026***	0.044***	0.122***	0.164**	0.120***
	(0.011)	(0.010)	(0.015)	(0.034)	(0.069)	(0.043)
Age squared	-0.569***	-0.435***	-0.637***	-1.715***	-2.725**	-1.757***
	(0.158)	(0.146)	(0.218)	(0.507)	(1.072)	(0.628)
Ethnic minority	0.021	-0.015	0.046	0.056	-0.150	0.125
	(0.021)	(0.020)	(0.031)	(0.060)	(0.118)	(0.081)
Number of years of schooling	0.015***	0.013***	0.021***	0.042***	0.061***	0.054***
	(0.002)	(0.002)	(0.003)	(0.006)	(0.011)	(0.008)
Dummy year 2010	*Reference*					
Dummy year 2012	0.023	-0.029	0.013	0.067	-0.173	0.033
	(0.020)	(0.022)	(0.031)	(0.056)	(0.116)	(0.080)
Dummy year 2014	0.037*	0.014	0.086***	0.104*	0.074	0.226***
	(0.021)	(0.025)	(0.032)	(0.059)	(0.113)	(0.084)
Dummy year 2016	0.074***	0.028	0.086***	0.206***	0.121	0.225***
	(0.022)	(0.025)	(0.032)	(0.059)	(0.112)	(0.084)
Constant	-0.548***	-0.395**	-0.451*	-3.034***	-4.228***	-2.578***
	(0.173)	(0.168)	(0.251)	(0.554)	(1.119)	(0.711)
Weak identification test						
Cragg-Donald Wald F statistic				34.9	35.6	24.5
Kleibergen-Paap rk Wald F statistic				28.3	29.4	20.0
Observations	3,863	1,718	2,145	3,863	1,718	2,145
Pseudo R-squared	0.036	0.055	0.051	0.029	0.072	0.038

Dependent variable is the school enrolment. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level. Cragg-Donald Wald F statistic and Kleibergen-Paap rk Wald F statistic are test statistics of weak instruments. As a rule of thumb, if a F statistic is under 10, the instruments might be weak (Staiger and Stock 1997).

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.5. Probit and OLS regressions of maternal employment outcomes on childcare (all sample)

Explanatory variables	Probit (marginal effects)							
	Working	Have a wage job	Have a self-employed nonfarm work	Have a self-employed farm work	Have a skilled work	Have a formal job	Log of total monthly working hours	Log o hourly wage
Childcare	0.028***	0.056***	0.025*	-0.064***	0.088***	0.031**	0.054***	0.060
	(0.010)	(0.019)	(0.013)	(0.019)	(0.021)	(0.012)	(0.021)	(0.041
Age	0.001	0.025*	0.024**	-0.038***	0.077***	0.040***	0.073***	0.048
	(0.008)	(0.013)	(0.009)	(0.014)	(0.016)	(0.011)	(0.016)	(0.037
Age squared	-0.001	-0.390*	-0.285**	0.515**	-1.081***	-0.640***	-1.051***	-0.544
	(0.112)	(0.201)	(0.138)	(0.208)	(0.242)	(0.166)	(0.236)	(0.558
Ethnic minority	0.074***	-0.016	-0.145***	0.334***	-0.311***	-0.061***	-0.017	-0.270*
	(0.009)	(0.025)	(0.011)	(0.023)	(0.023)	(0.015)	(0.028)	(0.066
Number of years of schooling	0.002	0.028***	-0.001	-0.033***	0.051***	0.037***	0.009***	0.068**
	(0.001)	(0.002)	(0.001)	(0.002)	(0.003)	(0.002)	(0.002)	(0.005
Dummy year 2010	*Reference*							
Dummy year 2012	0.067***	0.065***	-0.004	0.046*	0.054**	0.051**	0.037	0.182**
	(0.009)	(0.025)	(0.016)	(0.027)	(0.027)	(0.020)	(0.029)	(0.069
Dummy year 2014	0.045***	0.059**	-0.034**	0.055**	0.020	0.048**	0.031	0.340**
	(0.010)	(0.026)	(0.016)	(0.027)	(0.028)	(0.020)	(0.029)	(0.068
Dummy year 2016	0.041***	0.088***	-0.005	-0.020	0.042	0.073***	0.022	0.457**
	(0.011)	(0.026)	(0.017)	(0.027)	(0.028)	(0.021)	(0.027)	(0.069
Constant							3.739***	0.761
							(0.272)	(0.612
Observations	3,863	3,863	3,863	3,863	3,863	3,863	3,638	1,345
R-squared	0.0546	0.0592	0.0556	0.151	0.207	0.260	0.022	0.275

For probit models, the marginal effects are computed using the "margin" command in Stata. For childcare (or other dummy variables), the m discrete change in the probability of the dependent variable due to the change in the childcare (or other dummy variables). For a continuous estimated partial derivative of the dependent variable with respect to the explanatory variable (evaluated at the mean value of all the Heteroskedasticity-robust standard errors in parentheses. The standard errors are corrected for sampling weights and cluster correlation at the c *** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.6. Bivariate probit and 2SLS regression of maternal employment outcomes on childcare (all sa

Explanatory variables	Bivariate probit							
	Working	Have a wage job	Have a self-employed nonfarm work	Have a self-employed farm work	Have a skilled work	Have a formal job	Log of total monthly working hours	Log o hourly wage
Childcare	-0.637	1.472***	-0.439	-1.540***	0.350	1.212***	0.155	0.572
	(0.610)	(0.054)	(0.426)	(0.052)	(2.743)	(0.156)	(0.209)	(0.460
Age	0.035	-0.010	0.130***	-0.011	0.189	0.134***	0.069***	0.013
	(0.052)	(0.033)	(0.044)	(0.031)	(0.137)	(0.051)	(0.018)	(0.051
Age squared	-0.393	0.050	-1.593**	0.131	-2.638	-2.216***	-0.995***	-0.011
	(0.770)	(0.506)	(0.645)	(0.461)	(1.929)	(0.771)	(0.262)	(0.780
Ethnic minority	0.689***	-0.323***	-0.848***	0.645***	-0.849***	-0.333***	-0.018	-0.250**
	(0.115)	(0.065)	(0.112)	(0.063)	(0.075)	(0.094)	(0.028)	(0.072
Number of years of schooling	0.024**	0.049***	0.006	-0.034***	0.127**	0.151***	0.008*	0.061**
	(0.011)	(0.007)	(0.010)	(0.007)	(0.054)	(0.014)	(0.004)	(0.008
Dummy year 2010	*Reference*							
Dummy year 2012	0.561***	0.102*	-0.005	0.114**	0.134	0.198**	0.036	0.194**
	(0.094)	(0.061)	(0.074)	(0.055)	(0.096)	(0.084)	(0.029)	(0.073
Dummy year 2014	0.373***	0.085	-0.136*	0.145**	0.045	0.177**	0.028	0.308**
	(0.093)	(0.062)	(0.080)	(0.058)	(0.119)	(0.085)	(0.030)	(0.075
Dummy year 2016	0.367***	0.096	0.022	0.077	0.095	0.224**	0.015	0.404**
	(0.092)	(0.062)	(0.080)	(0.057)	(0.221)	(0.087)	(0.031)	(0.084
Constant	0.206	-1.075**	-3.213***	0.703	-4.426**	-4.785***	3.784***	1.175
	(0.850)	(0.548)	(0.697)	(0.512)	(1.907)	(0.864)	(0.288)	(0.765
Observations	3,863	3,863	3,863	3,863	3,863	3,863	3,638	1,345

The sample used for this regression consists of women with children being born in December and January of two consecutive years. Children l older than those born in January of the following year. The instrument for the childcare is children being born in December. Heteroskedas parentheses. The standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.7. Bivariate probit and 2SLS regression of maternal employment outcomes on childcare (samp

Explanatory variables	Bivariate probit							
	Working	Have a wage job	Have a self-employed nonfarm work	Have a self-employed farm work	Have a skilled work	Have a formal job	Log of total monthly working hours	Log o hourly wage
Childcare	-0.910	1.706***	-0.990***	-1.723***	-0.459	0.831	0.378	0.948
	(0.711)	(0.134)	(0.321)	(0.162)	(3.953)	(1.140)	(0.358)	(0.649
Age	-0.118	-0.021	0.045	-0.016	0.154**	0.168*	0.068***	0.002
	(0.075)	(0.053)	(0.061)	(0.051)	(0.066)	(0.088)	(0.026)	(0.070
Age squared	1.942*	0.301	-0.454	0.101	-2.180*	-2.567*	-0.975**	0.273
	(1.172)	(0.806)	(0.939)	(0.774)	(1.160)	(1.349)	(0.405)	(1.111
Ethnic minority	0.567***	-0.290***	-0.868***	0.724***	-0.764***	-0.362**	0.004	-0.164
	(0.153)	(0.104)	(0.156)	(0.094)	(0.238)	(0.165)	(0.045)	(0.125
Number of years of schooling	0.026*	0.065***	0.007	-0.046***	0.135***	0.180***	0.006	0.052**
	(0.013)	(0.011)	(0.014)	(0.010)	(0.012)	(0.036)	(0.006)	(0.012
Dummy year 2010	*Reference*							
Dummy year 2012	0.462***	0.165*	-0.107	0.094	0.150	0.252*	0.062	0.165
	(0.131)	(0.096)	(0.109)	(0.092)	(0.220)	(0.132)	(0.043)	(0.133
Dummy year 2014	0.294**	0.172*	-0.062	0.040	0.001	0.280**	0.012	0.261*
	(0.132)	(0.098)	(0.101)	(0.094)	(0.115)	(0.125)	(0.044)	(0.109
Dummy year 2016	0.264**	0.219**	-0.079	-0.002	0.088	0.382***	0.016	0.357**
	(0.123)	(0.101)	(0.108)	(0.098)	(0.135)	(0.145)	(0.046)	(0.107
Constant	2.553**	-0.958	-1.666*	0.675	-3.773***	-5.612***	3.827***	1.430
	(1.193)	(0.863)	(0.980)	(0.827)	(0.942)	(1.521)	(0.419)	(1.084
Observations	1,718	1,718	1,718	1,718	1,718	1,718	1,589	593

The sample used for this regression consists of women with children being born in December and January of two consecutive years. Children b older than those born in January of the following year. The instrument for the childcare is children being born in December. Heteroskedas parentheses. The standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.8. Bivariate probit and 2SLS regression of maternal employment outcomes on childcare (samp

Explanatory variables	Bivariate probit							
	Working	Have a wage job	Have a self-employed nonfarm work	Have a self-employed farm work	Have a skilled work	Have a formal job	Log of total monthly working hours	Log o hourly wage
Childcare	-0.784**	1.382***	0.299	-1.567***	0.139	1.191***	-0.009	0.141
	(0.398)	(0.093)	(0.631)	(0.049)	(0.776)	(0.284)	(0.255)	(0.568
Age	0.132**	0.011	0.172**	-0.020	0.248***	0.127*	0.080***	0.050
	(0.066)	(0.044)	(0.071)	(0.042)	(0.069)	(0.075)	(0.025)	(0.060
Age squared	-1.847*	-0.194	-2.188**	0.193	-3.499***	-2.134*	-1.125***	-0.565
	(0.947)	(0.663)	(1.029)	(0.620)	(1.022)	(1.126)	(0.351)	(0.918
Ethnic minority	0.839***	-0.358***	-0.875***	0.650***	-0.898***	-0.321***	-0.023	-0.309*
	(0.157)	(0.082)	(0.129)	(0.078)	(0.105)	(0.116)	(0.037)	(0.084
Number of years of schooling	0.031**	0.036***	0.002	-0.024***	0.131***	0.140***	0.010	0.068**
	(0.013)	(0.010)	(0.016)	(0.008)	(0.020)	(0.027)	(0.006)	(0.012
Dummy year 2010	*Reference*							
Dummy year 2012	0.590***	0.125	0.011	0.071	0.101	0.219*	0.029	0.266**
	(0.124)	(0.081)	(0.103)	(0.075)	(0.087)	(0.113)	(0.038)	(0.088
Dummy year 2014	0.522***	-0.020	-0.247**	0.299***	0.115	0.058	0.048	0.417**
	(0.129)	(0.083)	(0.122)	(0.076)	(0.116)	(0.125)	(0.043)	(0.140
Dummy year 2016	0.453***	0.054	0.037	0.096	0.135	0.163	0.023	0.532**
	(0.142)	(0.082)	(0.118)	(0.077)	(0.113)	(0.127)	(0.039)	(0.123
Constant	-1.219	-1.557**	-4.218***	1.148	-5.360***	-4.717***	3.639***	0.606
	(1.139)	(0.748)	(1.121)	(0.711)	(1.021)	(1.307)	(0.390)	(0.881
Observations	2,145	2,145	2,145	2,145	2,145	2,145	2,049	752

The sample used for this regression consists of women with children being born in December and January of two consecutive years. Children b older than those born in January of the following year. The instrument for the childcare is children being born in December. Heteroskedas parentheses. The standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

Table A.9. The medium-term effect of childcare on women's employment outcomes

Dependent variables	All children
Bivariate probit model (marginal effects)	
Working	0.011
	(0.043)
In wage-paying job	0.367***
	(0.041)
In self-employed nonfarm work	0.215
	(0.177)
In self-employed farm work	-0.403***
	(0.052)
In skilled work	0.163
	(0.164)
In a formal job	0.219*
	(0.119)
2SLS	
Log of monthly working hours	0.080
	(0.387)
Log of hourly wage	-0.298
	(0.423)
Log of wage for the last month	-0.072
	(0.525)
Log of total wage for the past 12 months	0.213
	(0.623)
Number of observations	592

This table reports the marginal effects from the bivariate probit regression and 2SLS of dummy employment variables on childcare and the control variables of women. For the dependent variables of wages and working hours, the regressions are 2SLS. The observations in these regressions are women of children age 3-5. The estimation samples are restricted to children that were sent to childcare both currently and two years ago. For probit models, the marginal effect of childcare is the estimated discrete change in the probability of the dependent variable due to the change in the childcare. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.10. The effect of childcare on maternal employment outcomes using different sets of control variables

Dependent variables	Small model specification (no control variables)	Large model specification with province fixed-effects
Working	-0.145	-0.150
	(0.166)	(0.102)
In a wage-earning job	0.426***	0.398***
	(0.042)	(0.016)
In self-employed nonfarm work	-0.054	-0.176
	(0.125)	(0.092)
In self-employed farm work	-0.445***	-0.407***
	(0.037)	(0.021)
In skilled work	0.221	-0.289
	(0.157)	(0.216)
In a formal job	0.266***	0.224***
	(0.082)	(0.047)

This table reports estimation results from the bivariate probit of maternal employment outcomes on childcare with different sets of control variables. In the small model, there are no control variables. In addition to control variables used in previous tables, the large model also controls for birth order and gender of children, urban and province fixed-effects.
The marginal effect is the estimated discrete change in the probability of the dependent variable due to the change in the childcare. The observations in these regressions are women of children age 1-5. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.11. The effect of childcare on maternal employment outcomes using different models

Dependent variables	2SLS (linear probability model)	Control function with the first step of linear probability model (marginal effects)	Control function with both probit (marginal effects)	Regression discontinuity design using the bandwidth of 11 months	Regression discontinuity design using the bandwidth of 6 months
Working	-0.160	-0.149	-0.213	-0.026	-0.038
	(0.123)	(0.166)	(0.169)	(0.081)	(0.127)
In a wage-earning job	0.526***	0.511***	0.393***	0.316***	0.271**
	(0.199)	(0.087)	(0.129)	(0.076)	(0.131)
In self-employed	-0.104	-0.124	-0.099	-0.043	0.038
nonfarm work	(0.141)	(0.109)	(0.123)	(0.046)	(0.051)
In self-employed farm	-0.582***	-0.495***	-0.446***	-0.299***	-0.458***
work	(0.202)	(0.060)	(0.084)	(0.110)	(0.130)
In skilled work	0.029	0.079	0.002	0.096	0.040
	(0.177)	(0.154)	(0.158)	(0.092)	(0.115)
In a formal job	0.244*	0.262*	0.227	0.276***	0.201***
	(0.146)	(0.140)	(0.146)	(0.042)	(0.054)

For probit in the control function models, the marginal effect of childcare is the estimated discrete change in the probability of the dependent variable due to the change in the childcare.
Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level. For the control function estimators, standard errors are estimated by bootstrap with 200 replications.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.12. The effect of childcare on maternal employment outcomes using different models and bandwidths

Dependent variables	2-month bandwidth	3-month bandwidth
Bivariate probit model (marginal effects)		
Working	-0.031	-0.031
	(0.073)	(0.059)
In a wage-earning job	0.405***	0.398***
	(0.008)	(0.007)
In self-employed nonfarm work	-0.073	-0.061
	(0.064)	(0.050)
In self-employed farm work	-0.409***	-0.374***
	(0.019)	(0.024)
In skilled work	0.233**	0.155
	(0.130)	(0.138)
In a formal job	0.255***	0.265***
	(0.026)	(0.018)
2SLS		
Log of monthly working hours	0.242	0.207*
	(0.147)	(0.107)
Log of hourly wage	0.489*	0.490**
	(0.294)	(0.223)
Log of wage for the last month	0.603**	0.519**
	(0.298)	(0.221)
Log of total wage for the past 12 months	0.705*	0.773***
	(0.378)	(0.287)

This table reports the marginal effects from the bivariate probit regression of dummy employment variables on childcare over the previous 2 years. For the dependent variables of wages and working hours, the regressions are 2SLS. This table reports only coefficients of childcare. For probit models, the marginal effect of childcare is the estimated discrete change in the probability of the dependent variable due to the change in the childcare. The observations in these regressions are women of children age 1-5. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.13. Reduced-form regression of maternal employment outcomes on the instrument

Dependent variables	1-month bandwidth	2-month bandwidth	3-month bandwidth
Probit model (marginal effects)			
Working	-0.016	0.037	0.146
	(0.110)	(0.060)	(0.124)
In a wage-earning job	0.377***	0.477***	0.333***
	(0.024)	(0.038)	(0.087)
In self-employed nonfarm work	0.043	-0.004	0.089
	(0.108)	(0.150)	(0.145)
In self-employed farm work	-0.419***	-0.384***	-0.297***
	(0.032)	(0.078)	(0.103)
In skilled work	-0.055	0.187	-0.239
	(0.384)	(0.143)	(0.157)
In a formal job	0.149	0.382	0.017
	(0.206)	(0.349)	(0.296)
2SLS			
Log of monthly working hours	0.293	0.489	0.206
	(0.312)	(0.470)	(0.463)
Log of hourly wage	-0.275	-0.104	-0.421
	(0.478)	(0.511)	(0.842)
Log of wage for the last month	-0.078	0.071	-0.286
	(0.523)	(0.580)	(0.895)
Log of total wage for the past 12 months	-0.068	0.397	-0.527
	(0.678)	(0.733)	(1.183)

Note: This table reports estimation results from the regressions of maternal employment outcomes on the instrument variable and the control variables shown in Table 2. The 1-month bandwidth sample includes women with children born in December and in January of the following year, and the instrument is a dummy indicating a child born in December. The 2-month bandwidth sample includes women with children born in November-December and in January-February of the following year, and the instrument is a dummy indicating a child born in November-December. The 3-month bandwidth sample includes women with children born in October-December and in January-March of the following year, and the instrument is a dummy indicating a child born in October -December. The marginal effect of childcare is the estimated discrete change in the probability of the dependent variable due to the change in the childcare. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the commune level.

*** p<0.01, ** p<0.05, * p<0.1.

Source: Authors' estimation from VHLSS 2010, 2012, 2014 and 2016.

Table A.14. Difference in characteristics and outcomes between mothers with children born in December and those with children born in January

Variables	Mothers with child born in December	Mothers with child born in January	Difference
Age	30.4	30.1	0.3
	(0.3)	(0.4)	(0.5)
Ethnic minority	20.2	25.5	-5.2
	(1.9)	(2.4)	(3.3)
Working	86.1	86.2	-0.1
	(2.0)	(2.0)	(3.0)
In a wage-earning job	27.2	27.5	-0.3
	(2.6)	(2.6)	(3.8)
In self-employed nonfarm work	14.7	13.4	1.3
	(2.1)	(2.1)	(3.0)
In self-employed farm work	44.3	45.3	-1.1
	(2.8)	(2.9)	(4.1)
In skilled work	45.9	44.9	1.0
	(2.9)	(2.9)	(4.3)
In a formal job	18.0	18.2	-0.2
	(2.3)	(2.3)	(3.3)
Log of monthly working hours	145.3	136.3	9.0
	(5.1)	(5.1)	(7.4)
Log of hourly wage	7.5	5.9	1.7
	(1.4)	(0.6)	(1.5)
Log of wage for the last month	951.0	899.4	51.6
	(107.0)	(99.1)	(149.1)
Log of total wage for the past 12 months	11161.7	10057.3	1,104.4
	(1365.5)	(1202.9)	(1,852.2)
Number of observations	354	348	

Note: This table reports the mean of variables of mothers with children aged 1 and born in December and mothers with children aged 0 and born in January. These children are not attending childcare. Thus, the differences in the variables between the two groups of mothers are not affected by childcare.
Standard errors of means in parentheses. The standard errors are corrected for sampling weights and cluster correlation at the commune level.
*** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.