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Childcare and maternal employment: Evidence from Vietnam

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ABSTRACT

Little literature currently exists on the effects of childcare use on maternal labor market outcomes in a developing country context, and the few recent studies offer mixed results. We attempt to fill these gaps by analyzing several latest rounds of the Vietnam Household Living Standards Survey spanning the early to mid-2010s. Addressing endogeneity issues with an instrumental variable/ regression discontinuity design estimator based on children's birth months, we find sizable and positive effects of childcare on women's own labor market outcomes and their household income and poverty status. The effects of childcare differ by women's characteristics and are stronger for women in the ethnic majority group or women with daughters. These effects are also somewhat larger for areas with higher income levels. Furthermore, we also find that some positive effects last after two years.

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 formal pre-school (age 1–5) childcare in Vietnam.[3]

We make several new contributions to the literature in this paper. First, 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. Gender inequality was practically eliminated at the primary school level with almost universal enrolment; in fact, it has even been reversed at the secondary school levels in the past decade with girls' net enrolment rates leading those of boys by as much as 10 percentage points at the upper secondary level (Dang and Glewwe, 2018). Girls also tend to perform better, or equally as well as, boys on a number of different national and international assessments in various subjects ranging from math and science subjects to English and

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[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] In this paper, we refer to formal childcare as the care of children by a day-care center, kindergarten, or preschool or other formal institutions. Unfortunately, we do not have data on informal childcare (which is more broadly defined as unpaid childcare that children receive from relatives, grandparents, babysitters and other unregistered child care), so do not investigate this type of childcare. We refer to childcare and preschool interchangeably in this paper.

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reading (OECD, 2015; Singh and Krutikova, 2017; Aurino et al., 2019; Azubuike and Little, 2019; Ho et al., 2020).[4]

Yet, gender inequality remains a challenge for the country. The proportion of men working in a wage job is 42%, but the corresponding figure for women is lower at 30%.[5] Gender gaps in both the formal work sector and pay have long been recognized as an issue for the country (Gallup, 2004; Liu, 2004; Feeny et al., 2021), and it is particularly more severe among low-paid and high-paid wage earners (Vo et al., 2021). 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 & 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.[6]

We also add to the scarce literature on childcare in a developing country. 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., Gelbach, 2002; Bauernschuster and Schlotter, 2015; Martínez and Perticará, 2017; Berthelon et al., 2020), a number of other studies do not (e.g., Cascio, 2009; Fitzpatrick, 2010; 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.

Second, 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). Indeed, informal employment is one of the main indicators of International Labor Organization (ILO)'s decent work indicators (ILO, 2008). Yet, while women have made considerable advances regarding LFP in recent years, women are generally not perceived to have made sufficient progress into decent and high-quality work. In most low-income and middle-income countries, women still predominantly work in the agricultural sector and the informal sector, which offer low-productivity and low-paying jobs (World Bank, 2012; UN Women, 2015; ILO, 2019). Women also continue to suffer from significant pay gaps and continue to be under-represented in management and executive positions (ILO, 2016). Furthermore, increasing the share of household income controlled by women could result in various changes to household expenditures (such as more expenditures on education, health, nutrition, or housing) that benefit the whole households, particularly children (World Bank, 2012).

As such, besides analyzing employment quantity indicators on women's labor force participation (LFP) and number of work hours, we look at a wide range of welfare indicators regarding employment quality and household outcomes, both in the short term and the medium term. Specifically, the former outcomes can be roughly grouped under three categories: i) work sectors (including outcomes such as self-employment, wage work, farm and non-farm work, and formal work), ii) level of employment skills (including skilled employment), and iii) wages (including hourly, monthly, and annual wages). The latter outcomes include household income and poverty. 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.[7]

Finally, a key challenge in measuring the effects of (formal) 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 applying the instrumental variable (IV) method to rich and nationally representative Vietnamese survey data spanning the early to mid-2010s, with each survey round covering more than 46,000 households. This IV method is similar to the quasi-experimental regression discontinuity design (RDD) method, where we exploit plausibly exogenous thresholds in the birth months of children to identify the causal impacts of childcare. Rather than focusing on a specific age (subpopulation) group or a smaller sample as most previous studies do, we take full advantage of the survey's large sample sizes and study all preschool children age 1–5. This larger survey sample allows us to employ the RDD model and further examine the robustness of results for different population groups. In fact, this larger survey sample for Vietnam has rarely, if at all, been analyzed in an academic study before.

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 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. Moreover, a medium-term effect exists after two years for younger children, further increasing the probability that women have a wage-earning job and reducing self-employed farm work. Childcare also helps increase women's total annual wages and household income per capita and reduce household poverty (although these effects are mostly marginally statistically significant). 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 briefly review some recent studies on poorer countries that are most relevant to our studies. Berlinski, Galiani, and McEwan (2011) find positive impacts for childcare 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) 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

[4] More generally, despite Vietnam's modest position as a lower-middle-income country, it has recorded better education performance than what may be suggested from its income level, particularly for women. Dang et al. (2021) offer evidence that a key driving factor behind this success was the mass education program implemented by the government during the first Indochina War almost a century ago.

[5] These estimates are obtained from our analysis of the Vietnam Household Living Standards Surveys unless it is noted otherwise,

[6] 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, Dang and Nguyen (2021) and Alon et al. (2022)).

[7] 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.

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.

Yet, 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.

This paper consists of six sections. We provide the institutional background on child care in the next section before describing the data and descriptive statistics of childcare and a first look at its relationship with maternal employment in Vietnam in Section 3. We discuss the estimation method and provide the empirical results in Sections 4 and 5 and finally conclude in Section 6.

2. Background on childcare in Vietnam

Given Vietnam's age-old Confucian culture of lifelong learning, there is strong demand for education of all levels in Vietnam, including preschool. Sending a child to preschool is generally regarded not only as just supporting women to work, but also as having a more child-focused goal related to education and lifelong learning (Boyd and Dang, 2017). Yet, childcare is generally considered a woman's inherent duty in Vietnam; in particular, women undertake almost all childcare tasks, ranging from feeding and bathing to nursing a sick child (ISDS, 2015). We study formal 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.[8]

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 current/ calendar age), regardless of whether they have reached their birth day in that year (i.e., their completed/ actual age). 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.[9] Consequently, we can 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. Fig. A.1 in Appendix A further illustrates that while the percentages of children attending childcare steadily increase by children's age, they remain rather stable for each age cohort (i.e., the sizes of the circles, which are proportional to the enrolment rates, are rather similar for each age cohort but become larger for older age cohorts).

According to Vietnam's 2005 Law on Education (revised in 2019), children can attend childcare starting from being 3 months old. But attendance at kindergarten and preschool is not compulsory. When deciding to send children to a public preschool, parents typically take into account the calendar age instead of the completed age.[10] As a result, children born in December are still more likely to attend preschool than those born in January of the following year. However, the proportion of children below age two attending childcare is very small in practice (Fig. 1).

Fig. 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. Fig. 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. One possible reason for the significant increase in the childcare attendance rate between 2010 and 2016 is the implementation of the program on universalization of preschool education.[12] On the supply side, this program aims at constructing 86 preschools in poor communes, adding 11,600 preschool classrooms for 5-year-olds, and training more preschool teachers. On the demand side, the program creates incentives for parents by offering free lunch for around 400,000 children every year from poor households or living in remote and mountainous areas. This program aims to increase the rate of preschool attendance for children aged 5 to 95% in 2015. Furthermore, the Vietnamese government recently committed to increasing preschool attendance for children under 3 in preschool to 30% by 2020 and 35% by 2025 (Government of Vietnam, 2018).

Like other school levels, preschool enrolment is typically based on a residence registration (school catchment area) basis. That is, households living in a village (or an urban ward) can send their children to the preschool serving the village (ward). Households may also apply to send their children to a preschool in a different catchment area, but these applications are subject to the preschool's decision. 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.[13] Preschools for children age 3 to 5 are more available, with 49% of villages having at least a preschool. Access to childcare in industrial zones are even more limited. For example, preschool facilities in industrial parks and export-processing zones in Ho Chi Minh City only meet 2% of demand (UNICEF, 2017).

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 (see more detail in Table A.1 in the Appendix).

[8] 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.

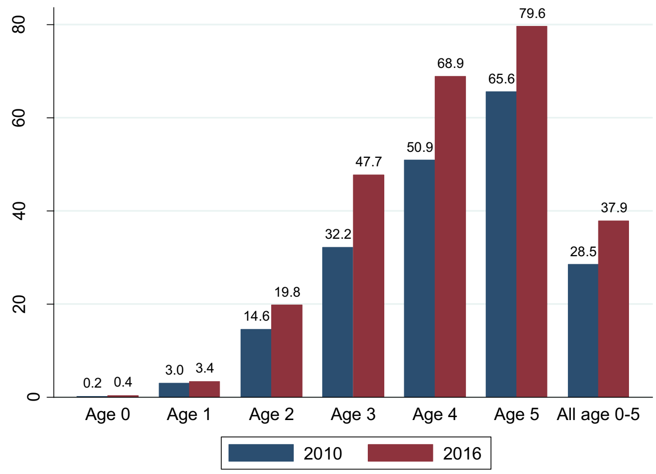
[9] When we discuss a child's age in this paper, we refer to the calendar age rather than the completed age. Children attend childcare centers and preschools from Monday to Friday. The school day at these institutions and at primary schools often starts at 7.30 a.m. and ends at 4.30p.m. But this time schedule is not fixed. Some childcare centers (and preschools) admit children on Saturday and allow them to be picked up later than 4.30p.m. Some preschools also admit younger children (National Assembly of Vietnam, 2005).

[10] For private preschools, parents can enroll their children any time during the time.

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

[12] Decision No. 239/QD-TTG of the Prime Minister "Approving the Project on universalizing preschool education for five-year-old children for the period 2010–2015", which is Available at https://vanban.chinhphu.vn/default.aspx?pageid=27160&docid=93227.

[13] 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.



Note: This figure presents the level of childcare by children age 0-5.

Fig. 1. Percentage of children attending childcare.

More children below the age of 3, however, are enrolled in private childcare centers.[14] The proportions of children age below 3 attending private and public childcare centers were 27% and 73% in 2016, respectively. The cost of private childcare is likely significantly higher than that of public childcare. Table A.4 in the Appendix shows that in 2016 the average annual expenditure on private childcare was 9,356 thousand VND per child (around 426 USD), which is almost five times the corresponding figure for public childcare at 1,998 thousand VND (around 91 USD).[15] 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. In 2016, 44% of urban children age below 6 attended childcare centers and kindergartens, while the corresponding figure for rural children was lower at 35%. Ethnic minority groups still lag behind in terms of childcare. Out of all children attending pre-school, the percentage of ethnic minority children was roughly 2% in the school year 2012–2013 (MOET, 2015), which is far below the share of ethnic minority groups in the country population of about 15%.[16]

3. Data and descriptive analysis

3.1. Data description

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. The surveys were conducted in several months during the years from March to December. 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 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, incomes 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.

The information on childcare is obtained from household heads' or parents' responses to questions on whether the child is currently going to preschool (or whether the child has been going to preschool in the past year if the child is no longer enrolled in school). While the VHLSSs do not collect information on hours of attending childcare, they provide information on whether a public childcare center exists in the commune. There is also information on household expenditure on childcare for each child in the household.

3.2. Maternal employment

We analyze indicators regarding women's employment that are roughly grouped under two aspects: quantity and quality of work. The quantity indicators include outcomes such as women's LFP, including working at least one hour in economic activities during the past 12 months and working fulltime, and the number of work hours. The working rate is thus defined as the proportion of people who have been working at least one hour in economic activities during the past 12 months. For fulltime work, we follow General Statistical Office, 2020 and define fulltime work as working at least 35 hours per week.

We roughly group the quality employment indicators under three categories: i) work sectors (including outcomes such as self-employment, wage work, farm and non-farm work, and formal work), ii) level of employment skills (including skilled employment), and iii) wages (including hourly, monthly, and annual wages). There are several advantages to analyzing these employment categories. First, these different aspects of quality employment offer more comprehensive and granular analysis of gender inequality that have not been done before. Second, they are directly relevant to the areas where there is still much room to improve gender equality as pointed out by studies discussed earlier. Finally, since reducing gender inequality through employment is a complex, multi-faceted undertaking that depends on country specific socio-economic contexts (Kabeer, 2021), it is helpful to examine different aspects of employment. This last advantage is especially consistent with economic studies showing that women's

[14] 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.

[15] The exchange rate is 21,935 VND for 1 USD in 2016 according to World Bank's World Development Indicators database.

[16] See Boyd and Dang (2017) and Vu (2021) for recent discussion on other aspects of early childhood education (such as curricula, staffing, legal regulations) and its history in Vietnam.

Table 1
Employment outcomes of women and men.

Variables	Employment of women				Employment of men
	VHLSS 2010	VHLSS 2012	VHLSS 2014	VHLSS 2016	VHLSS 2016
% working	91.8	93.2	92.5	93.5	98.9
	(0.4)	(0.5)	(0.5)	(0.5)	(0.3)
% working fulltime	65.0	67.6	65.9	67.4	81.7
	(0.5)	(0.9)	(1.0)	(1.0)	(0.9)
% in a wage-earning job	30.9	33.4	35.5	37.6	52.2
	(0.7)	(1.0)	(1.0)	(1.1)	(1.1)
% self-employed in a nonfarm job	16.5	14.6	13.7	18.0	16.0
	(0.6)	(0.7)	(0.7)	(0.8)	(0.8)
% self-employed in a farm job	44.4	45.1	43.4	37.9	30.6
	(0.8)	(1.0)	(1.0)	(1.1)	(1.0)
% in a skilled job	45.0	47.2	49.3	53.5	61.9
	(0.8)	(1.0)	(1.0)	(1.1)	(1.1)
% in a formal job	15.1	18.6	21.5	23.7	20.5
	(0.7)	(0.8)	(1.0)	(0.9)	(0.9)
Number of working hours per month	180.0	187.2	188.7	188.0	206.9
	(1.3)	(1.5)	(1.6)	(1.6)	(1.6)
Hourly wage (thousand VND)	18.1	19.5	20.4	24.2	27.0
	(0.8)	(0.7)	(0.5)	(0.0)	(1.3)
Monthly wage (thousand VND)	3252.4	3554.0	3845.2	4404.3	5360.3
	(89.1)	(99.1)	(79.9)	(0.0)	(128.9)
Yearly wage (thousand VND)	39013.0	41878.6	46334.3	52749.0	62592.6
	(1434.5)	(1331.3)	(1131.8)	(0.0)	(1690.6)
Number of observations	4,525	3,806	3,405	3,059	3,011

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. Working fulltime is defined as working at least 35 hours per week during the last month. 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. All estimates are weighted with the survey weights. Standard errors of the mean are in parentheses. Wages are measured in 2016 prices.
Source: Authors' estimation from VHLSSs 2012, 2014, and 2016.

increased control over household income could result in various beneficial changes to the households, particularly children (World Bank, 2012). Furthermore, we also consider more general quality of life outcomes such as household income and poverty. For heterogeneity analysis, we consider other related outcomes such as household size, migration, and co-residence with grandparents.

We examine the employment outcomes of women with at least one child aged below 6 over the period 2010–2016 in Table 1. These variables are defined in detail in Table A.2 in the Appendix. These women's average age hovers around 32 and ranges from 17 to 58. The working rate of women in 2016 was 94%. The proportion of women with fulltime work is lower at 67% for the same year. For comparison, Table 1 also reports men's employment rate in 2016, which is 5 percentage points higher with almost all men (99%) working.[17] The proportion of men working fulltime is also remarkably higher at 82%. This gender gap in the employment rate, however, was stable during the period under study. The reasons for not working are very different for men and women. Fig. A.2 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.

The proportion of men not working outside the home for the same reason was much lower at 23% in 2016.[18]

Among working people, men work more hours (207 hours) in a month than women (188 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 (52%) 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 (56%) 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 38% of women worked in the farm sector.

We also categorize employment into skilled employment and formal employment. Skilled employment is defined based on workers' 2-digits occupation codes provided in the VHLSS (which uses the International Standard Classification of Occupations (ISCO-08) (ILO, 2012)). People with skilled employment include managers, professionals, experts, office staffs, mechanic and other skilled workers. Formal jobs are jobs with social insurance benefits, and workers with a formal job are eligible for health insurance, unemployment allowances, and pensions.[19] While men were more likely to have a skilled job than women (i.e., 54% for women versus 62% for men), women had a slightly higher rate of formal employment than men (i.e., 24% for women versus 21% for men).

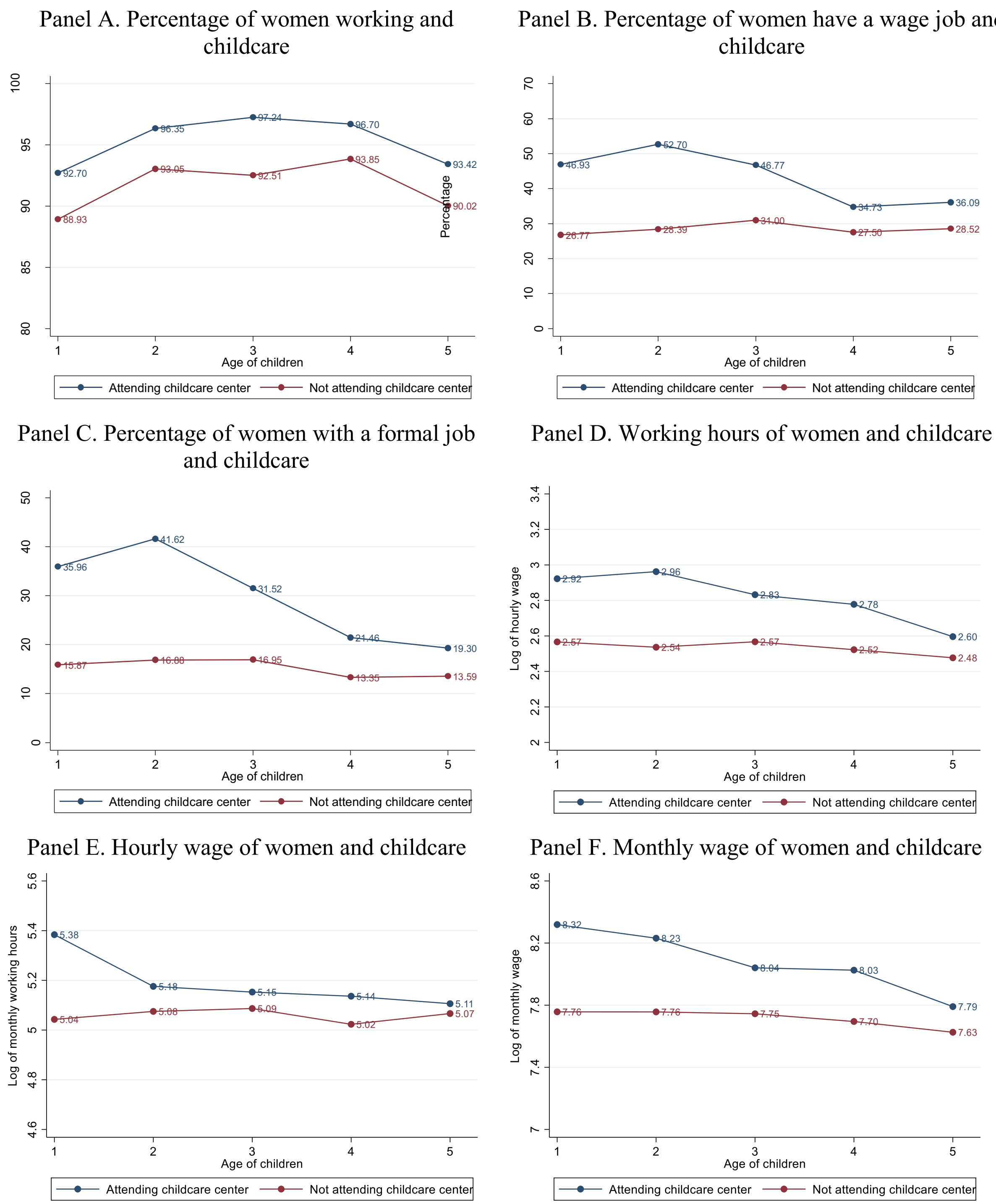
The VHLSSs collect data on respondents' wages over the past 30 days and their total wages over the past 12 months. Since a number of workers have more than one job, we compute total wages from the main and secondary jobs for our analysis.[20] The real hourly wage of women increased from 18,100 to 24,200 VND during 2010–2016, as did their monthly and yearly wages. However, the gender gap in wages also increased over time. In 2010, the aver-

[17] The estimates of men's employment outcomes in other years are reported in Table A.3 in the Appendix.

[18] 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. See also Heath and Jayachandran (2017) for a general discussion on rising female school enrolment and LFP in developing countries.

[19] 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).

[20] 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.



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

Fig. 2. Maternal employment by children's attendance at childcare centers.

age annual wage for men was 9% higher than that for women, but in 2016, this gender wage gap widened to 19%.

3.3. First look at childcare and maternal employment

Fig. 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 regarding 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.

Fig. 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. Indeed, Fig. A.3 in the Appendix further illustrates that mothers who send their children to childcare have more schooling than those who do not. Urban and ethnic majority (Kinh) mothers are more likely to have their children attend children than rural and ethnic minority mothers. The next sections discuss our estimation method and empirical findings on the causal effects of childcare on maternal employment.

4. Estimation method

As discussed earlier, the proportion of children attending childcare increases by age, and children's calendar age is used to determine their enrollment eligibility. Since the VHLSSs contain data on age (year and month), but not an individual's full date of birth, we 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.

Fig. 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 Fig. 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 December, and of children age 1 and born from January to June. Older children unsurprisingly have more childcare attendance. However, there is an obvious, large gap in the incidence of childcare between children born in December and those born in January.[21] As discussed earlier (Section 2), parents use the calendar age instead of the completed age when deciding to send children to preschool. This explains why children born in December still have a higher rate of childcare than those born in January of the following year.

The difference in childcare attendance between months of birth suggests that this variable can be used as the instrumental variable for childcare. To estimate childcare effects on maternal employment, we use an instrumental variable (IV) regression, which consists of two stages. In the first stage, we estimate the effects of being born in December on the probability of attending childcare:

$$D_{i,j} = \alpha + \varphi Dec_{i,j} + \gamma' X_{ij} + \varepsilon_{i,j} \tag{1}$$

where $D_{i,j}$ is a dummy variable that indicates a mother i with a child j who currently attends childcare. $Dec_{i,j}$ 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 $\varepsilon_{i,j}$ are vectors of observed and unobserved characteristics of women, respectively. We use a small set of exogenous control variables, including a mother's age, gender, ethnic minorities, number of years of schooling, and year dummy variables. We also control for age and gender of the children. For children, we measure their age in months using the information on month and year of birth and the month of interview. We also control for month of interview. It should be noted that the 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, rather than the partial effects, of childcare on outcomes 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 estimation 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.

$$Y_{i,j} = \delta + \theta D_{i,j} + \pi' X_{ij} + u_{i,j} \tag{2}$$

where $Y_{i,j}$ is the employment variable of interest. Eq. (2) is estimated together with Eq. (1) 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 (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 Eqs. (1) and (2) with maximum likelihood methods as our preferred model for interpretation. But we also apply 2SLS regressions for binary dependent variables as robustness checks.

We cluster the standard error at the level of primary sampling units, i.e., the village level in this case. The VHLSSs contain rotated panel data. In our final data set, only 20% of women were surveyed at least in two adjacent surveys. For robustness check, we also tried to cluster the standard errors at the individual levels. However, the number of individuals in the panel is small compared with the whole sample. Thus we use the results with standard errors clustered at the village level 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 childcare, controlling for the predicted residuals from the first stage regression and other explanatory variables. Furthermore, we perform two additional robustness checks. First, 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.

[21] This gap no longer exists for primary-school age or older children; Figure A.4 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*".

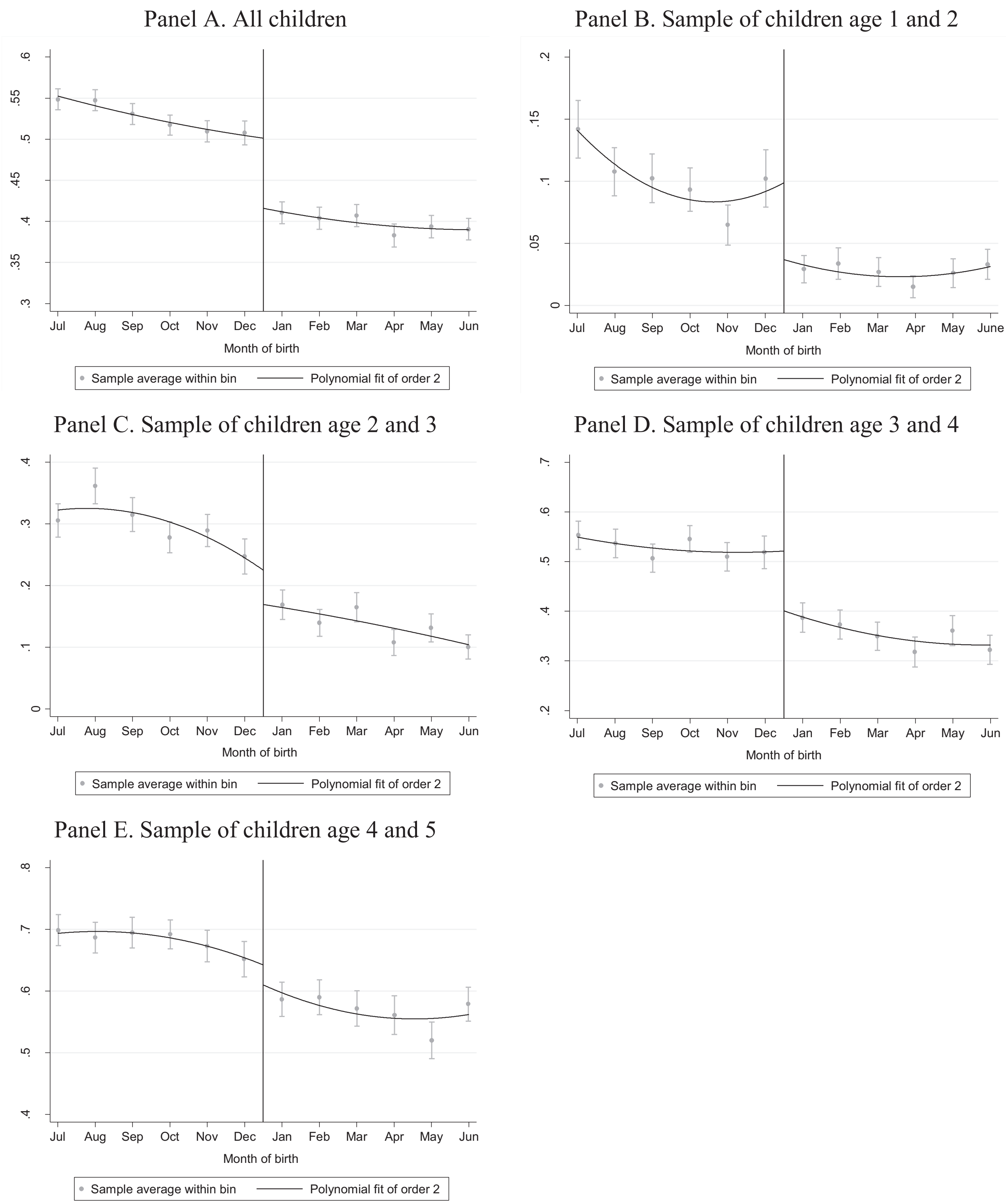


Fig. 3. The proportion of enrolled children age 1–5 and month of birth.

Second, our IV model can be regarded as a regression discontinuity design (RDD) model, with month of birth used as the conditioning variable.[22] While we use the bandwidth of one month to estimate Eq. (2), we also extend the bandwidth to 6 and 11 months and apply more standard RDD for robustness analysis. Since birth month does not strictly determine childcare attendance, we apply the fuzzy RDD method to measure childcare effects on maternal employment. Specifically, we use a local linear regression developed by Hahn et al. (2001) as follows:

$$Y_{i,j} = \beta_0 + \beta_1 D_{i,j} + \beta_2 (Z_{i,j} - c) I_{\{Z_{i,j}-c\}} + \beta_3 (Z_{i,j} - c) + \xi Xij + v_{i,j} \quad (3)$$

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

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

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 Eq. (3) 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 the survey 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). We provide various analysis to validate this identification strategy. To test the exogeneity of being born in December, we first compare the proportion of children born in different months. Panel A of Fig. A.5 in the Appendix shows that the proportion of children born in December is slightly lower than the proportion of children born in January. However, the difference is not statistically significant at the 5% level. Panel B of Fig. A.5 presents the manipulation plot test, which shows that the difference between the proportion of children born in December and the proportion of children born in January is not statistically significant at the conventional levels.

We further 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 equals 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 dates is not simple. According to Vietnam's Law on Civil Status 2014, to register for a child's birth certificate, parents have to submit a document certifying the birth of the child issued by a hospital, commune health center, or midwifery house. As such, the month of birth is stated on birth certificates and is difficult to manipulate. Another related concern is that caesarean section procedures have become more popular, accounting for around 34% of births in the country (General Statistics Office and UNICEF, 2021), so parents may choose their child's birthdates using these procedures. Yet, these procedures typically allow choosing birthdates by a few days rather than a month.[23]

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. 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 the higher fertility rate nine months later. The higher proportion of children born in October warns against using larger (e.g., 3-month) bandwidths 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.4 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.5 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, an additional year 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.[24]

We can examine whether women with children born in December and those with children born in January are different before having the children. In Table A.6 in the Appendix, we analyze panel data on women who gave birth to children in December and January in the next two years. We then regress the outcomes of the current year on "having a child born in December in the next two years" and other control variables. The variable "having a December child" has small magnitude and is not statistically significant in all the regressions, indicating the validity of the instrument "being born in December." We also use the same approach to examine whether there is selection bias and reverse causality issues with a naïve OLS estimator for the effects of childcare. We regress women's employment on a variable indicating 'sending their children to childcare in the next two years'. This childcare variable is statistically significant in the regressions of fulltime working, wage job, working hours and wages, indicating the existence of these issues. This provides further supporting evidence for our use of IV and RDD models instead of the OLS model.

Finally, we test the exogeneity of the instrument by graphing several exogenous variables across children's birth month in two consecutive years. Fig. A.6 shows that there are no significant differences in these exogenous variables between mothers with chil-

[23] Moreover, if a number of people manipulate their children's birth month on their birth certificate or employment C-section procedures in order to send their children to childcare earlier, the proportion of children with a reported birth month in December would 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.

[24] 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 5.3, which lend further support to our estimation results.

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

Explanatory variables	Probit model (marginal effect)			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)	0.087***	0.056***	0.087***	0.066***	0.059***	0.080***
	(0.018)	(0.016)	(0.025)	(0.015)	(0.018)	(0.022)
Age	0.038***	0.033***	0.041**	0.029***	0.027***	0.034**
	(0.013)	(0.013)	(0.018)	(0.010)	(0.010)	(0.015)
Age squared	−0.595***	−0.554***	−0.608**	−0.452***	−0.457***	−0.505**
	(0.202)	(0.197)	(0.260)	(0.143)	(0.154)	(0.214)
Ethnic minority	0.018	−0.030*	0.058*	0.020	−0.017	0.052*
	(0.024)	(0.018)	(0.033)	(0.019)	(0.020)	(0.031)
Number of years of schooling	0.021***	0.011***	0.023***	0.017***	0.013***	0.021***
	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.003)
Gender of child (boy = 1, girl =)	−0.009	−0.006	−0.006	−0.007	−0.007	−0.005
	(0.018)	(0.015)	(0.024)	(0.014)	(0.017)	(0.022)
Age of child (in month)	0.036***	0.014	0.004	0.009***	−0.021*	0.008
	(0.005)	(0.012)	(0.032)	(0.003)	(0.011)	(0.028)
Age of child (in month) squared (multiplied by 100)	−0.020***	−0.005	0.012	0.010**	0.064***	0.007
	(0.006)	(0.023)	(0.033)	(0.004)	(0.023)	(0.028)
Interview month	0.013***	0.006**	0.019***	0.009***	0.007**	0.017***
	(0.003)	(0.003)	(0.005)	(0.003)	(0.003)	(0.004)
Dummy year 2010	*Reference*					
Dummy year 2012	0.028	−0.015	0.059*	0.027	−0.004	0.056*
	(0.025)	(0.020)	(0.034)	(0.020)	(0.023)	(0.031)
Dummy year 2014	0.113***	0.046*	0.136***	0.088***	0.048**	0.128***
	(0.027)	(0.024)	(0.034)	(0.020)	(0.025)	(0.032)
Dummy year 2016	0.135***	0.049**	0.165***	0.111***	0.056**	0.154***
	(0.027)	(0.025)	(0.033)	(0.021)	(0.026)	(0.032)
Constant				−0.913***	−0.382*	−1.038
				(0.170)	(0.217)	(0.738)
Weak identification test						
Cragg-Donald Wald F statistic				24.7	13.6	15.3
Kleibergen-Paap rk Wald F statistic				19.9	11.0	12.6
Observations	3,863	1,718	2,145	3,863	1,718	2,145
Pseudo R-squared	0.239	0.178	0.092	0.269	0.131	0.120

This table reports the probit and OLS regressions of childcare on the instrumental variable and control variables of women. The observations in these regressions are women of children age 1–5. For the probit regression, this table reports marginal effects.
Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the village 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).
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.

dren born in December and those with children born in January of the following year.

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. Compared with a January birth, being born in December of the previous year increases the probability of attending childcare by 0.087, 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 2).

5.2. Main estimation results

Panel A of Table 3 reports the estimated impacts of childcare on maternal employment outcomes, using the IV bivariate probit regressions. There are 7 outcome variables in this table, 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.8 to A.11 in Appendix A present the full regression results. For comparison, we also show the univariate (non-instrumental) probit regression results.[25]

As discussed earlier (Section 4), the probit method suffers from endogeneity and provides biased estimates. This explains why for several outcomes, the estimation results switch from being statistical significant under the probit regressions to being statistical insignificant under the bivariate probit regressions (for working, working fulltime, working in a skilled job and a formal job). In other cases, the probit estimates are smaller in absolute magnitude than the bivariate probit estimates (for having a wage-paying job or a self-employed farm job), suggesting that the former are biased downward toward zero. Consequently, we focus on the bivariate probit estimates when interpreting the estimation results.

We find that childcare has statistically insignificant effects 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

[25] The Probit and OLS regression results for subgroups of children are reported in Table A.7 in the Appendix.

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

Dependent variables	Panel A. Short-term effects					Panel B. Medium-term effects				
	Probit model	*Bivariate probit model*			*Number of obs. (all children)*	*Probit model*	*Bivariate probit model*			*Number of obs. (all children*
	All children	All children	Children age 1–3	Children age 3–5		All children	All children	Children age 1–3	Children age 3–5	
Working	0.022**	−0.070	−0.041	−0.113	3,863	0.016*	−0.119	−0.302	0.108	1,375
	(0.011)	(0.069)	(0.135)	(0.089)		(0.009)	(0.246)	(0.299)	(0.302)	
Working fulltime	0.087***	0.173	0.312	0.316	3,863	0.095***	0.294	0.451	0.191	1,375
	(0.019)	(0.145)	(0.203)	(0.203)		(0.031)	(0.238)	(0.295)	(0.323)	
In wage-paying job	0.069***	0.254**	0.408***	0.427***	3,863	0.002	0.467***	0.728**	0.363*	1,375
	(0.021)	(0.123)	(0.079)	(0.019)		(0.033)	(0.164)	(0.289)	(0.203)	
In self-employed nonfarm work	0.022	−0.034	−0.185	−0.028	3,863	0.043*	0.187	0.361	0.021	1,375
	(0.015)	(0.090)	(0.113)	(0.125)		(0.025)	(0.247)	(0.264)	(0.315)	
In self-employed farm work	−0.079***	−0.384***	−0.415**	−0.441***	3,863	−0.041	−0.455***	−0.772***	−0.120	1,375
	(0.022)	(0.065)	(0.156)	(0.046)		(0.039)	(0.139)	(0.169)	(0.344)	
In skilled work	0.048**	0.074	−0.174	0.109	3,863	0.022	0.050	0.172	−0.084	1,375
	(0.019)	(0.205)	(0.126)	(0.140)		(0.031)	(0.258)	(0.392)	(0.304)	
In a formal job	0.045***	0.010	−0.131	0.284***	3,863	0.012	−0.093	0.170	−0.228	1,375
	(0.015)	(0.181)	(0.177)	(0.051)		(0.021)	(0.247)	(0.316)	(0.264)	

This table reports marginal effects of childcare from the probit and bivariate probit 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 village level. For the medium-term effects (Panel B), we run regression of women's outcome in the current survey round (t = 2) on childcare attendance in the preceding survey round (t = 1).
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

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. Furthermore, the effects of childcare on mothers working fulltime are positive but not statistically significant (row 2). These results are consistent with the result that childcare has no significant effects on women's number of monthly working hours discussed later (Table 4).

Indeed, these two factors (high LFP and prevalence of part-time work) are considered to be responsible for insignificant to low impacts of childcare on maternal employment in several richer countries (Lundin et al., 2008; Bettendorf et al., 2015; Akgunduz and Plantega, 2018). The argument is that further reductions in child care prices will have diminishing effects due to high maternal employment. The effects can be smaller particularly if other structural factors are supportive as in the case of Vietnam, where there is more working hour flexibility due to the high self-employment rate in the agricultural sector. The availability of part-time work can lower the demand for formal, full-time child care, while at the same time allowing women to provide informal child care. These two effects of part-time work are expected to increase both the demand for and the supply of informal child care which can be substituted for formal child care, which potentially results in smaller effects on maternal employment.

The effects of childcare on engaging in skilled work are small and statistically insignificant (row 6). Obtaining work skills takes a long time; as such, children's attendance 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 3 of Table 3). Childcare increases the probability of having a wage-earning job by 0.25 (or 25 percentage points). Childcare also has negative effects on self-employed farm work (row 5), and somewhat negative effects on self-employed non-farm work (row 4) but the latter effects are 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.[26] There are also significant positive effects of childcare of children aged 3–5 on women having formal jobs (row 7). Having a child aged 3–5 in childcare increases the probability of women having a formal job by 0.28. These results are qualitatively consistent with case studies in various developing countries including Guatemala, South Africa, and Uganda, which point to lack of childcare as a key reason for women being unable to take formal-economy jobs (Cassirer and Addati, 2007; Alfers, 2016; Devercelli and Frances Beaton-Day, 2020). These studies also suggest that in the absence of childcare, women are more willing to accept more poorly paid, insecure, and precarious types of work in the informal sector to increase flexibility, such that they can take care of their children.

In Table 4, we look at the effects of childcare on working hours and wage. For wage earners, the effects of childcare on hourly and monthly wages are positive but not statistically significant (rows 2 and 3). While our imprecise point estimate for a monthly wage increase of 65% 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 marginally statistically significant effects on annual wages (row 4). The increase in total wages may be due to women having more participation in formal jobs as well as increased productivity.

Panels A of Tables 3 and 4 examine the contemporaneous relationship between childcare and maternal employment (i.e., in the same year). But another policy-relevant question is whether childcare has ongoing effects on maternal employment. To examine the medium-term effects, we regress women's outcome in the current survey round (t = 2) on childcare attendance in the preceding survey round (t = 1), using panel data from two consecutive survey rounds.[27] This approach aims to measure the 2-year lagged effects

[26] 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.

[27] 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, by design 50% of households in the 2010 VHLSS are resampled in the 2012 VHLSS. The attrition rate in two consecutive surveys is around 8%.

Table 4
The effects of childcare on women's working hours and wage.

Dependent variables	Panel A. Short-term effects					Panel B. Medium-term effects				
	OLS	*2SLS*			*Number of obs. (all children)*	*OLS*	*2SLS*			*Number of obs. (all children)*
	All children	All children	Children age 1–3	Children age 3–5		All children	All children	Children age 1–3	Children age 3–5	
Log of number of monthly working hours	0.063*** (0.022)	0.185 (0.263)	0.502 (0.477)	0.013 (0.302)	3,638	0.053 (0.039)	0.384 (0.410)	0.592 (0.595)	0.245 (0.486)	1,325
Log of hourly wage	0.058 (0.046)	0.688 (0.548)	1.229 (0.941)	0.163 (0.696)	1,345	−0.028 (0.063)	−0.271 (0.537)	0.043 (0.621)	−0.470 (1.037)	537
Log of wage for the last month	0.064 (0.042)	0.652 (0.496)	1.352 (0.884)	0.168 (0.646)	1,379	0.032 (0.068)	−0.010 (0.590)	0.213 (0.727)	−0.056 (1.035)	537
Log of total wage for the past 12 months	0.163*** (0.052)	1.122* (0.646)	1.646 (1.106)	0.900 (0.857)	1,381	0.134 (0.084)	−0.009 (0.761)	0.785 (0.952)	−0.345 (1.387)	537

This table reports estimate of childcare from OLS and 2SLS regressions of maternal employment outcomes on childcare and the control variables shown in Table 2. 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 village level.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

of childcare on parental employment outcomes, using the same model specification as with Panel A. The estimation results, shown in Panels B of Table 3 and Table 4, do not suggest significant childcare effects on the working status or wages of women after two years. However, childcare has strong lagged effects on the probability of women taking a wage-earning job: it increases this probability by 0.47 after two years. The effects on self-employed farm work are negative, indicating a movement from farm work to wage-earning employment. These medium-term effects are around 6 times higher than the short-term effects.

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.12, remain very similar.[28]

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.13 in Appendix shows that results remain very similar (Tables 3 and 4).

Secondly, as mentioned in the method section, we tried to cluster the standard error at the individual level instead of the village level. The results are reported in Table A.14 in the Appendix. The magnitude of the standard errors is similar.

Thirdly, we examine whether the estimates are sensitive to women's age. In the previous section, we use the whole sample of women with children. The age of women ranges from 17 to 58. In Table A.15 in the Appendix, we limit the sample to women age 20 to 50 (during their main reproductive years). It shows that the estimated childcare effects on women's employment are almost the same as those in Tables 3 and 4.

Fourthly, we explore different estimation models. In Tables 3 and 4, 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 probability models. For control function models, we implement two different types.[29] We also estimate the local linear regression in Eq. (3) using the bandwidth of 6 and 11 months (Table A.16 in the Appendix). The last two columns of Table A.16 show the 'donut' RDD, in which we drop children born in December and January to examine whether the effects are sensitive to the sample around the threshold. The estimation results are qualitatively similar to those shown in Table 3. Childcare has positive effects 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. In addition, the RDD shows positive effects on working fulltime and formal jobs.

Fifthly, we further examine the sensitivity of the estimates to 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.17 in the Appendix shows that the estimated childcare effects 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. When using the 3-month bandwidth, we include a set of dummy variables to the regressions, indicating "children born in adjacent months", e.g. a dummy variable that equals 1 if child was born in either Dec 2010 or Jan 2011.

[28] 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.

[29] 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.

Table 5
2SLS regression of household-level outcomes on childcare.

Dependent variables	Sample of households		
	All households with children aged 1–5	Households with children aged 1–3	Households with children aged 3–5
Log of income per capita	0.702**	0.992**	0.528
	(0.296)	(0.446)	(0.384)
Household is poor	−0.281*	−0.476*	−0.158
	(0.151)	(0.249)	(0.190)
Living with grandparents	0.026	0.024	0.027
	(0.061)	(0.100)	(0.073)
Women are migrating	0.011	−0.053	0.062
	(0.060)	(0.087)	(0.082)
Household size	−0.439	−0.621	−0.349
	(0.441)	(0.716)	(0.564)
Number of observations	3,841	1,721	2,120

This table reports estimation results of the effect of childcare on household-level outcomes from 2SLS regressions of the household-level outcomes on childcare and the control variables shown in Table 2. The observations in these regressions are households with children are age 1–5 and born in December and January of the following years. Heteroskedasticity-robust standard errors in parentheses. Standard errors are corrected for sampling weights and cluster correlation at the village level.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

It shows that inclusion of these control variables does not change the results.

Sixthly, Table A.18 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. In addition, we plot the outcome variables across months of birth of children in Fig. A.7, which clearly shows different probabilities of having a wage-earning job and a self-employed farm job between women with December-born children and women with January-born children.

Seventhly, a potential 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 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.

Fig. A.8 in the Appendix shows the distribution of the p-values of the variable "born earlier" in these regressions. For the birth month gap from 1 to 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 effects of being born earlier is significant at the 1% level, indicating no supportive evidence of "being born earlier" effects (for children of the same age) on maternal outcomes.

Finally, we conduct a balance test by comparing variables (for both the demographics and outcomes) between mothers who have children age 1 and born in December and those who have children age 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.19 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 5 presents the estimated impacts of childcare on several household outcomes. Taking advantage of childcare increases household income (row 1), which can in turn help reduce poverty (row 2).[30] The probability of being poor is reduced by 0.28 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. Grandparents live with their children in approximately 21% of households in 2016. Informal childcare provided by grandparents was shown to increase women's labor supply in China, a neighboring country (Maurer-Fazio et al., 2011). We test whether formal childcare in a center can substitute for informal childcare from grandparents by regressing a dummy variable "living with grandparents" on childcare. The estimated childcare effects are negligible and statistically insignificant (row 3). There are no significant effects on maternal migration and household size (rows 4 and 5). 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 (chil-

[30] 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.

Table 6
The effects of interactions between child schooling and demographic variables on the probability of mothers having a wage job (probit models).

Interaction variables	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10
Childcare	0.262	−0.075	0.128	0.153	0.157	0.572**	0.261*	0.096	−0.675***	−0.456*
	(0.181)	(0.154)	(0.149)	(0.151)	(0.155)	(0.228)	(0.153)	(0.165)	(0.174)	(0.257)
Childcare * age	−0.004									
	(0.003)									
Childcare * schooling years		0.015***								
		(0.006)								
Childcare * ethnic minority			−0.109**							
			(0.045)							
Childcare * boy				−0.063*						
				(0.033)						
Childcare * birth order					−0.028					
					(0.022)					
Childcare * grandparents living in household						−0.106				
						(0.083)				
Childcare * public childcare center							−0.110			
							(0.180)			
Childcare * distance to nearest town								−0.005***		
								(0.002)		
Childcare * log of district per capita wage									0.130***	
									(0.047)	
Childcare * log of district per capita income										0.063*
										(0.037)
Control variables										
Observations	3,863	3,863	3,863	3,863	3,863	1,107	3,863	2,853	3,821	3,863
Pseudo R2	0.138	0.140	0.139	0.139	0.145	0.128	0.141	0.0908	0.137	0.162

Note: This table reports the interaction effect between childcare and explanatory variables in from probit regressions of the probability of women working in wage-paying jobs on childcare and the interactions between childcare and other explanatory variables. We first run a probit model of childcare on the instrument and other explanatory variables, and estimate the generalized residuals (Wooldridge, 2015). We subsequently run a probit model of maternal employment on the childcare variable, the generalized 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 number of observations 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 observations. The average of the Z-statistic is reported in parentheses.
*** the absolute value of Z-statistic > 2.57, ** >1.96, * > 1.65.
Source: Authors' estimation from VHLSSs 2010, 2012, 2014 and 2016.

dren 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, childcare endogeneity 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 interaction effects between childcare and explanatory variables.[31] Since the interaction effects vary across observations, we report in Table 6 the average size and the average z-statistics of the interaction effects and use these averages for interpretation. 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. Childcare effects do not differ for age (Model 1 in Table 6). However, we find heterogeneous effects across the education levels. The effects of childcare on wage job tend to be larger for women withmore school years (Model 2). Possibly, the opportunity cost of staying at home to take care of children is larger for highly-educated women. The availability of childcare likely encourages highly-educated women to participate in the labor market and find a wage job. The effects are, however, 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 boys and girls. Model 4 in Table 6 shows that the effects on maternal employment of boys attending childcare are slightly lower than those of girls. 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 effects on maternal employment of boys attending childcare are smaller than that of girls attending childcare.

Children's order of birth tends to negatively correlate 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. Thus, we include interactions between childcare and living with grandparents, but this interaction effect is not statistically significant (Model 6), which is consistent with the estimation results with Table 5 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 smaller effects 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 childcare effects on women's employment in these areas. Indeed, such effects may depend on the opportunity costs of staying at home (i.e., not participating in

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

the labor market) to take care of children. In Model 9, we test the interaction of childcare with the districts' average wage. The opportunity costs of not participating in the labor market are higher in districts with a higher average wage. The interaction is marginally significant but positive, indicating that childcare effects are greater in areas with higher wage levels.[32] Consistently, we also find stronger effects of childcare in areas with higher income levels (Model 10).

6. Discussion and conclusion

In this paper, we offer the first study that rigorously investigates the effects of pre-school childcare on maternal employment in Vietnam. We find that the percentage of children attending childcare is 3% for children age 1, although this figure improves for older children. We find childcare to have small, insignificant effects on women's LFP, which may be due to the high self-employment rate in the country.

However, we find that childcare has strong effects on women's labor market outcomes. Specifically, childcare increases the probability of women having a wage-earning job by 25 percentage points and reduces their probability working in a self-employed farm job by 38 percentage points. The medium-term effects (after two years) are even stronger. Childcare also helps increase women's total annual wages and household income per capita and reduce household poverty (although these effects are mostly marginally statistically significant). We also find that childcare has heterogeneous effects and differs for women of different characteristics. In particular, these effects tend to be greater for Kinh majority, more educated women, and those with daughters, and for areas with higher wages or with greater opportunity costs for not participating in the labor market.

We acknowledge that the VHLSS data have limitations (which only offer birth years and birth months but not birth days) that prevent us from richer analysis. As such, it can be useful to improve the survey future design to collect more detailed data on birth dates. The VHLSSs do not contain long-run panel data, thus we cannot estimate the long-term effects of childcare. Estimating such effects is out of the scope of this study, but holds much potential for future studies.

Our findings on the positive effects of childcare are generally qualitatively consistent with those in studies cited earlier for developing countries (Section 1). They are also supported by a recent global review of studies on both richer and poorer countries (Devercelli and Beaton-Day, 2020). 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, the existing supply of childcare appears inadequate. In particular, providing childcare in areas with higher wages can be particularly beneficial for women's access to a wage job. This agrees with evidence from other poorer countries where the lack of childcare can form a barrier that prevents women from taking jobs in the formal sector. 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.

It may be useful to consider some back-of-the-envelope cost-benefit analysis. Our study finds that childcare reduces the probability of working in the agricultural sector and increases the probability of having a wage job for women. We assume that if a woman sends her child to childcare, she can move from working in the self-employed agricultural sector to a wage job. According to the 2016 VHLSS, the average income from the self-employed agricultural sector and a wage job were 3.1 and 4.4 million VND, respectively. It suggests that this woman can earn an additional income of 1.3 million VND per month or 15.6 million VND per year. Even if she send her child to private childcare and has to pay a higher childcare cost of 9.4 million VND in 2016 (Table A.4 in the Appendix), the net gain is considerable at 6.2 million VND. This would equal two months of her income in the self-employed agricultural sector. While our calculations are rather simplistic (and do not consider other issues such as whether there are enough wage jobs for every woman who would like to switch their employment), it helps illustrate that if the government provides childcare for a fee, even at the higher private childcare cost, there can be substantial gains for women.

Our simple cost-benefit estimation is consistent with some recent case studies for Vietnam, which suggests that the business benefits of employer-supported childcare can help attract the top female talent, reduce turnover and unplanned absenteeism, improve productivity, and comply with international practices (International Finance Corporation, 2020). In particular, this study documents that a footwear manufacturer (with close to 7,500 workers)'s success in halving average monthly turnover from 4.1 percent in 2011 to 2 percent in 2018 could create annual savings of up to $US 537,000. Another footwear manufacturing calculates that an average unplanned absenteeism rate of 0.6 percent across a workforce of 33,000 in its Dong Nai factories costs the company around $US 945,000 each year. These firms credit at least parts of these savings to their childcare programs. Scaling up these employer-provided childcare programs could offer a win–win situation for both the employers and women workers.

CRediT authorship contribution statement

Hai-Anh H. Dang: Conceptualization, Formal analysis, Funding acquisition, Methodology, Writing – original draft, Writing- review & editing. **Masako Hiraga:** Conceptualization, Funding acquisition. **Cuong Viet Nguyen:** Conceptualization, Data curation, Formal analysis, Methodology, Writing – original draft.

Declaration of Competing Interest

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

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[32] Mean wages 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).

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.worlddev.2022.106022.

References

Ai, C., & Norton, E. C. (2003). Interaction terms in logit and probit models. *Economics Letters, 80*(1), 123–129.

Akgunduz, Y. E., & Plantenga, J. (2018). Childcare prices and maternal employment: A meta-analysis. *Journal of Economic Surveys, 32*(1), 118–133.

Alfers, L. (2016). *Our children do not get the attention they deserve. WEIGO Child Care Initiative Research Report*. Cambridge, MA: WEIGO.

Alon, T., Coskun, S., Doepke, M., Koll, D., & Tertilt, M. (2022). From mancession to shecession: Women's employment in regular and pandemic recessions. *NBER Macroeconomics Annual, 36*(1), 83–151.

Angrist, J. D. (2001). Estimation of limited dependent variable models with dummy endogenous regressors: Simple strategies for empirical practice. *Journal of Business & Economic Statistics, 29*(1), 1–28.

Angrist, J. D., & Pischke, J. (2009). *Mostly Harmless Econometrics: An Empiricist's Companion*". New Jersey: Princeton University Press.

Aurino, E., Schott, W., Behrman, J.R. and Penny, M., 2019. Nutritional status from 1 to 15 years and adolescent learning for boys and girls in Ethiopia, India, Peru, and Vietnam. (2019). *Population Research and Policy Review*, *38*(6): 899-931.

Azubuike, O. B., & Little, A. W. (2019). Learning Motivations, Learning Outcomes and Gender in Vietnam. *Young Lives Working paper series No. 181*. Oxford.

Attanasio, O., Paes de Barro, R., Carneiro, P., Evans, D., Lima, L., Olinto, P., & Schady, N. (2017). *Impact of free availability of public childcare on labour supply and child development in Brazil*. New Delhi: International Initiative for Impact Evaluation (3ie).

Bandiera, O., & Natraj, A. (2013). Does Gender Inequality Hinder Development and Economic Growth? Evidence and Policy Implications. *World Bank Research Observer, 28*, 2–21.

Bauernschuster, S., & Schlotter, M. (2015). Public childcare and women' labor supply—Evidence from two quasi-experiments. *Journal of Public Economics, 123*, 1–16.

Bedard, K., & Dhuey, E. (2006). The persistence of early childhood maturity: International evidence of long-run age effects. *Quarterly Journal of Economics, 121*(4), 1437–1472.

Berlinski, S., Galiani, S., & Mc Ewan, P. J. (2011). Preschool and maternal labor market outcomes: Evidence from a regression discontinuity design. *Economic Development and Cultural Change, 59*(2), 313–344.

Berthelon, M., Kruger, D., Lauer, C., Tiberti, L., & Zamora, C. (2020). Longer School Schedules, Childcare and the Quality of Mothers' Employment: Evidence from School Reform in Chile. *Partnership for Economic Policy Working Paper*, (2020-07).

Bettendorf, L. J., Jongen, E. L., & Muller, P. (2015). Childcare subsidies and labour supply—Evidence from a large Dutch reform. *Labour Economics, 36*, 112–123.

Blau, D., & Currie, J. (2006). Preschool, daycare, and after-school care: Who is minding the kids? *Handbook of the Economics of Education*, ed. by E. A. Hanushek, and F. *Welch, 2*, 1163–1278.

Boyd, W., & Dang, T. P. (2017). Early childhood education in Vietnam: History and evaluation of its policies. In E. P. J. C. H. Li (Ed.), *Early childhood education policies in Asia Pacific: Advances in theory and practice* (pp. 263–283). Singapore: Springer.

Calderón, G. (2014). The effects of childcare provision in Mexico, *Working Papers No. 2014-07*, Banco de México, Ciudad de México.

Card, D., Kluve, J., & Weber, A. (2018). What Works? A Meta Analysis of Recent Active Labor Market Program Evaluations. *Journal of the European Economic Association, 16*(3), 894–931.

Cascio, E. U. (2009). Maternal labor supply and the introduction of kindergartens into American public schools. *Journal of Human Resources, 44*(1), 140–170.

Cassirer, N., & Addati, L. (2007). *Expanding women's employment opportunities: Informal economy workers and the need for childcare*. Geneva: ILO (International Labour Organization).

Cattaneo, M. D., Idrobo, N., & Titiunik, R. (2019). *A practical introduction to regression discontinuity designs: Foundations*. Cambridge University Press.

Clark, S., Kabiru, C. W., Laszlo, S., & Muthuri, S. (2019). The impact of childcare on poor urban women's economic empowerment in Africa. *Demography, 56*(4), 1247–1272.

Cunningham, W., & Pimhidzai, O. (2018). *Vietnam's Future Jobs: Leveraging Mega-Trends for Greater Prosperity*. Washington, DC: The World Bank.

Dang, H. A. H. (2012). Vietnam: A Widening Poverty Gap for Ethnic Minorities. In G. Hall & H. Patrinos (Eds.), *Indigenous Peoples, Poverty and Development*. New York: Cambridge University Press.

Dang, H. A. H., & Glewwe, P. W. (2018). Well begun, but aiming higher: A review of Vietnam's education trends in the past 20 years and emerging challenges. *Journal of Development Studies, 54*(7), 1171–1195.

Dang, H. A. H., & Nguyen, C. V. (2021). Gender inequality during the COVID-19 pandemic: Income, expenditure, savings, and job loss. *World Development, 140* 105296.

Dang, H. A. H., Hoang, T. X., & Nguyen, H. (2021). The Long-Run and Gender-Equalizing Impacts of School Access: Evidence from the First Indochina War. *Economic Development and Cultural Change, 70*(1), 453–484.

Devercelli, A. E., & Beaton-Day, F. (2020). *Better Jobs and Brighter Futures: Investing in Childcare to Build Human Capital*. Washington, DC: World Bank.

Diebolt, C., & Perrin, F. (2013). From stagnation to sustained growth: The role of female empowerment. *American Economic Review, 103*(3), 545–559.

Do, Q. T., & Phung, T. D. (2010). The importance of being wanted. *American Economic Journal: Applied Economics, 2*(4), 236–253.

Duflo, E. (2012). Women empowerment and economic development. *Journal of Economic Literature, 50*(4), 1051–1079.

Duflo, E., Glennerster, R., & Kremer, M. (2007). Using randomization in development economics research: A toolkit. *Handbook of development economics, 4*, 3895–3962.

Feeny, S., Mishra, A., Trinh, T. A., Ye, L., & Zhu, A. (2021). Early-Life exposure to rainfall shocks and gender gaps in employment: Findings from Vietnam. *Journal of Economic Behavior & Organization, 183*, 533–554.

Fitzpatrick, M. D. (2010). Preschoolers enrolled and mothers at work? The effects of universal prekindergarten. *Journal of Labor Economics, 28*(1), 51–85.

Gallup, J. L. (2004). The wage labor market and inequality in Vietnam. In P. Glewwe, N. Agrawal, D. Dollar, & W. Bank (Eds.), *Economic growth, poverty, and household welfare in Vietnam* (pp. 53–94). Washington DC: World Bank. Regional and Sectoral Studies.

Gelbach, J. B. (2002). Public schooling for young children and maternal labor supply. *American Economic Review, 92*(1), 307–322.

Gelman, A., & Imbens, G. (2019). Why high-order polynomials should not be used in regression discontinuity designs. *Journal of Business & Economic Statistics, 37*(3), 447–456.

General Statistical Office (2020). *Labor and Employment Report 2020*. Vietnam: General Statistical Office of Vietnam.

General Statistics Office and UNICEF (2021). *Survey measuring Viet Nam Sustainable Development Goal indicators on Children and Women 2020–2021, Survey Findings Report*. Ha Noi, Viet Nam: General Statistics Office.

Guilmoto, C. Z. (2012). Son preference, sex selection, and kinship in Vietnam. *Population and Development Review, 38*(1), 31–54.

Government of Vietnam. (2018). Plan on Development of Pre-School Education for 2018 – 2020. https://thuvienphapluat.vn/van-ban/giaoduc/Quyet-dinh-1677-QD-TTg-2018-De-an-Phat-trien-Giao-duc-mamnon-2018-2025-401806.aspx.

Hahn, J., Todd, P., & Van der Klaauw, W. (2001). Identification and estimation of treatment effects with a regression-discontinuity design. *Econometrica, 69*(1), 201–209.

Halim, D., Johnson, H. C., & Perova, E. (2021). Preschool Availability and Women's Employment: Evidence from Indonesia. *Economic Development and Cultural Change*. https://doi.org/10.1086/714439.

Havnes, T., & Mogstad, M. (2011). Money for nothing? Universal childcare and maternal employment. *Journal of Public Economics, 95*(11–12), 1455–1465.

Heath, R., & Jayachandran, S. (2017). The causes and consequences of increased female education and labor force participation in developing countries. In *Oxford Handbook of Women and the Economy* (pp. 345–367). Oxford University Press.

Heckman, J., LaLonde, R., & J; and Smith Jeffrey A. (1999). The economics and econometrics of active labor market programs. In A. Ashenfelter & D. Card (Eds.). *Handbook of Labor Economics 1999* (volume 3). Elsevier Science.

Ho, M. T., La, V. P., Nguyen, M. H., Pham, T. H., Vuong, T. T., Vuong, H. M., ... Vuong, Q. H. (2020). An analytical view on STEM education and outcomes: Examples of the social gap and gender disparity in Vietnam. *Children and Youth Services Review, 119* 105650.

International Finance Corporation (2020). *Tackling Childcare: The Business Case for Employer-Supported Childcare in Vietnam*. Washington, DC: World Bank.

International Labour Organization. (ILO). (2008). *Measurement of decent work based on guidance received at the Tripartite Meeting of Experts on the Measurement of Decent Work (September 2008)*. Available at https://www.ilo.org/wcmsp5/groups/public/—dgreports/—integration/documents/meetingdocument/wcms_115402.pdf.

International Labour Organization. (ILO). (2012). *International Standard Classification of Occupations 2008 (ISCO-08): Structure, group definitions and correspondence tables*. Geneva: International Labour Office.

International Labour Organization. (ILO). (2016). *Women at work: trends 2016*. Geneva: International Labour Office.

International Labour Organization. (ILO). (2019). *Quantum Leap for Gender Equality: For a Better Future of Work for All*. Geneva: International Labour Office.

Institute for Social Development Studies. (ISDS). (2015). *Social Determinants of Gender Inequality in Vietnam*. Retrieved from https://vietnam.embassy.gov.au/files/hnoi/ISDS_Report_Binh%20dang%20gioi_EN_PDF-2.pdf.

Kabeer, N. (2021). Gender Equality, Inclusive Growth, and Labour Markets. In *Women's Economic Empowerment: Insights from Africa and South Asia* (pp. 13–48). Routledge.

Lagerlöf, N.-P. (2003). Gender equality and long-run growth. *Journal of Economic Growth, 8*(4), 403–426.

Lee, D. S., & Lemieuxa, T. (2010). Regression discontinuity designs in economics. *Journal of Economic Literature, 48*(2), 281–355.

Lanjouw, P., Marra, M., & Nguyen, C. (2017). Vietnam's evolving poverty index map: Patterns and implications for policy. *Social Indicators Research, 133*(1), 93–118.

Liu, A. Y. (2004). Gender wage gap in Vietnam: 1993 to 1998. *Journal of Comparative Economics, 32*(3), 586–596.

Long, J. S. (1997). *Regression Models for Categorical and Limited Dependent Variables*. Thousand Oaks, California: SAGE Publications Inc.

Lundin, D., Mork, E., & Ockert, B. (2008). How far can reduced childcare prices push female labour supply? *Labour Economics, 15*, 647–659.

Martínez, C., & Perticará, M. (2017). Childcare effects on maternal employment: Evidence from Chile. *Journal of Development Economics, 126*, 127–137.
Maurer-Fazio, M., Connelly, R., Chen, L., & Tang, L. (2011). Childcare, eldercare, and labor force participation of married women in urban China, 1982–2000. *Journal of Human Resources, 46*(2), 261–294.
McCrary, J. (2008). Manipulation of the running variable in the regression discontinuity design: A density test. *Journal of Econometrics, 142*(2), 698–714.
Ministry of Education and Training (MoET) (2015). National education for all 2015 review: Viet Nam Retrieved from unesdoc.unesco.org/images/0023/002327/232770e.pdf.
National Assembly of Vietnam. (2005). *Law on Education*. National Assembly of Vietnam, Vietnam.
National Assembly of Vietnam (2014). *Law on Social Insurance*. National Assembly of Vietnam, Vietnam.
Nguyen, C., & Tran, A. (2017). The effect of having children on women's marital status: Evidence from Vietnam. *Journal of Development Studies, 53*(12), 2102–2117.
Norton, E. C., Wang, H., & Ai, C. (2004). Computing interaction effects and standard errors in logit and probit models. *Stata Journal, 4*(2), 154–167.
Oecd (2015). *The ABC of Gender Equality in Education: Aptitude, Behaviour, Confidence*. Paris: OECD Publishing.
Quisumbing, A., Meinzen-Dick, R., & Malapit, H. (2021). Women's empowerment and gender equality in South Asian agriculture: Measuring progress using the project-level Women's Empowerment in Agriculture Index (pro-WEAI) in Bangladesh and India. *World Development, 105396*.
Rivers, D., & Vuong, Q. H. (1988). Limited information estimators and exogeneity tests for simultaneous probit models. *Journal of Econometrics, 39*, 347–366.
Silva, M. S., & Klasen, S. (2021). Gender inequality as a barrier to economic growth: A review of the theoretical literature. *Review of Economics of the Household, 19*, 581–614.
Singh, A., & Krutikova, S. (2017). Starting together, growing apart: Gender gaps in learning from preschool to adulthood in four developing countries. *Young Lives Working Paper Series No. 174*. Oxford.
Staiger, D., & Stock, J. (1997). Instrumental variables regression with weak instruments. *Econometrica, 65*(3), 557–586.
UNICEF. (2017). The Apparel and Footwear Sector and Children in Vietnam. https://www.unicef.org/vietnam/reports/apparel-and-footwearsector-and-children-viet-nam.
U.N. Women (2015). *Progress of the world's women 2015–2016: Transforming economies, realizing rights*. UN Women.
Vo, D., Van, L. T. H., Tran, D. B., Vu, T. N., & Ho, C. M. (2021). The determinants of gender income inequality in Vietnam: A longitudinal data analysis. *Emerging Markets Finance and Trade, 57*(1), 198–222.
Vu, T. T. (2021). Early childhood education in Vietnam, history, and development. *International Journal of Child Care and Education Policy, 15*(1), 1–18.
Wooldridge, J. M. (2010). *Econometric Analysis of Cross Section and Panel Data*. Cambridge, Mass.: MIT Press.
Wooldridge, J. M. (2015). Control function methods in applied econometrics. *Journal of Human Resources, 50*(2), 420–445.
World Bank (2012). *World Development Report 2012: Gender Equality and Development*. Washington, DC: World Bank.
World Bank (2019). *World Development Report 2019: The Changing Nature of Work*. Washington, DC: World Bank.
World Bank & Ministry of Planning and Investment of Vietnam (2016). *Vietnam 2035: Toward prosperity, creativity, equity, and democracy*. Hanoi: World Bank Vietnam.