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Balancing family and career: The effect of public childcare on fertility in Japan*

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Abstract

In this study, we examine the impact of public childcare provision on fertility in Japan between 1990 and 2020. Using micro data from Vital Statistics and the Population Census, we reveal the heterogeneous effects of increasing childcare slots on birth rate by mothers' age and birth parity. Our analysis shows that a 100% increase in the capacity of childcare centers increases the birth rate by 0.14 points. First births account for 53% of this effect, second births for 41%, and third births for the remaining 6%. The effect of having a first child was particularly large among women in their late 20s to early 30s, while the effect of having a second child was notable among women in their late 20s and late 30s. The effect of having a third child was limited to women in their early 30s, suggesting that it was particularly large for those who had their first and second children at a relatively younger age. Additionally, our analysis shows that the policy effect is larger for university graduates than for non-graduates, suggesting that public childcare alleviates challenges in balancing work and family responsibilities among career-oriented women.

Keywords: Childcare, fertility, work-life balance, birth timing, birth parity
JEL classification: J13, J18

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1 Introduction

Many developed countries have faced population aging and have struggled to boost fertility rates over the past few decades. Recently, the United States, which had previously maintained a total fertility rate (TFR) above two, also experienced a sharp decline, with the TFR dropping to 1.78 in 2023. While various factors such as increased female employment opportunities and more gender-neutral social norms contribute to the low fertility rate, a key issue is that parents, particularly mothers, face significant challenges in balancing career and family responsibilities. Research shows that childbirth has a substantial and lasting negative impact on maternal labor market outcomes, a phenomenon known as the *child penalty* or the *motherhood dip*, which is widely observed across developed and developing countries (Fernández-Kranz et al., 2013; Kleven et al., 2019a,b, 2024). Therefore, an important policy goal is to support parents by balancing work and family responsibilities.

Typical policies designed to support preschool childcare include parental leave, the right to work part-time, and center-based daycare. However, the duration of parental leave often does not fully cover the preschool years, and some studies have suggested that excessively long parental leave can have negative consequences for both maternal labor market outcomes and child development (Rossin-Slater, 2018; Canaan, 2022). Similarly, Fernández-Kranz and Rodríguez-Planas (2021) demonstrate that a Spanish policy allowing mothers with children aged less than six to work part-time triggered negative employer responses, leading to a reduction in permanent contracts for women. These findings suggest that policies offering mothers labor supply flexibility during the preschool years are not always effective.

In this context, center-based daycare becomes crucial, as it provides reliable childcare while parents are at work. The literature suggests that high-quality formal care is effective for both maternal employment and child development (Datta Gupta and Simonsen, 2010; Havnes and Mogstad, 2011b; Cornelissen et al., 2018; Yamaguchi et al., 2018a,b), although children may be worse off when the quality of center-based care is not sufficiently high (Baker et al., 2008; Havnes and Mogstad, 2015; Cornelissen et al., 2018). Although some studies find no significant impact of center-based daycare on maternal employment, especially when informal childcare is replaced by formal childcare (Fitzpatrick, 2010; Havnes and Mogstad, 2011a), it still helps parents balance work and family responsibilities by providing more stable childcare arrangements.

This study aims to reveal the effects of public-center-based daycare on fertility rates. To achieve this, we exploit the staggered expansion of public childcare slots in Japan. As a leading country experiencing population aging, Japan has seen a significant decline in the TFR over the past half-century. In 1970, the TFR was 2.13, slightly above the

replacement level, but dropped to 1.54 in 1990 and further declined to 1.29 in 2004. Although the TFR saw a slight recovery afterward and was stable at approximately 1.4, it reached its historically lowest level of 1.20 in 2023. The TFR in Japan is lower than that in most OECD countries, but similar to Italy and Spain, whose TFR is 1.25 and 1.19 in 2021, respectively, according to OECD family database.

In response to the substantial decline in TFR, the Japanese government implemented a series of policies to expand public childcare availability. In fact, the number of slots increased by approximately 70% from 1990 to 2020 on average. However, because public childcare centers are highly regulated by the teacher-child ratio and facility area per child, as well as the price of childcare services, some municipalities cannot offer sufficient slots despite the high demand for daycare services owing to local budget constraints and even facing waiting lists of children. We utilized this staggered expansion of public childcare slots using within-municipality variation in the slots to identify the causal effect of public childcare provision on the fertility rate.

Our analysis draws on micro data of birth record registration and population census from 1990 to 2020. Our analysis shows that the expansion of public childcare capacity has a moderate impact on birth rate. In particular, a 70% increase in slots, which was the average growth rate between 1990 and 2020, resulted in an increase in the yearly birth rate of 0.28 points or 6.5% compared with the average birth rate. We examined the heterogeneous effects of maternal age, and our empirical evidence shows that the policy effect is most pronounced among women in their late 20s. Overall, the pattern of heterogeneity closely mirrored the distribution of birth rates by maternal age, suggesting that the effects of public childcare provision were observed across a wide range of maternal age groups.

To further understand how public childcare provision influences childbirth behaviors, we examined the heterogeneous policy effects of parity and maternal age. Our findings indicate that the effect on first births is broadly observed across their 20s and 30s. By contrast, the effect on second births followed an M-shaped pattern, peaking in women's late 20s and late 30s. Although the magnitude of the policy effect on second births was slightly smaller than that on first births, it remained comparable. The effect on third births was smaller than the effects on first and second births and was limited to women aged 28–31 years. Therefore, combining these findings, those who would plan to have a third child due to public childcare policy seem likely to give second birth in their late 20s, whereas those who give second birth in their late 30s are unlikely to give third birth. This policy effect is somewhat surprising given that the fertility rate of third births peaks at age 32 years, suggesting that childcare policy encourages women to have higher-order births at relatively young ages.

Using these estimation results, we also inferred the policy effect on the hazard rate of childbirth. Our results indicate that the effect on the hazard rate is pronounced for women in their 30s for first and second births. In particular, the expansion of public childcare has increased the number of late-in-life pregnancies in these birth parities. In contrast, however, we did not find an increase in the hazard rate of third births in women in their late 30s; instead, the hazard rate increased in women in their late 20s. This analysis of the hazard rates highlights the substantial heterogeneity in the policy effects of birth parity and maternal age.

Because we estimated the policy effects by birth parity and maternal age, we used these estimates to infer the policy effect on the TFR and decomposed it by birth parity. We found that a 70% increase in public childcare expansion increased the TFR by 0.097. Considering that the TFR declined by 0.2 points between 1990 and 2020, this effect was sizable. The decomposition results demonstrate that first births accounted for 53% of the total effect, second births for 41%, and third births for 6%. Therefore, while d'Albis et al. (2017) emphasize the effect of childcare policy on second births, our findings underscore that the policy effect on first births is also sizable; therefore, the policy reduces childlessness.

Additionally, our heterogeneity analysis of maternal education revealed that childcare policies were more effective for university graduates than for non-graduates. The difference was most pronounced for first births. However, we also observed significant heterogeneity in birth timing for higher-order births between these educational groups. The expansion of childcare slots led to an increase in higher-order births for non-graduates in their 20s, while the policy effect was more substantial for university graduates in their late 30s. This finding aligns with the theoretical predictions of Caucutt et al. (2002) and Doepke et al. (2023), which suggest that career-oriented women are incentivized to delay childbirth to avoid early career interruptions.

Finally, we examine whether public childcare provision has adverse effects on birth outcomes, as studies have found that employment opportunities have non-negligible effects on birth outcomes and maternal health (Dehejia and Lleras-Muney, 2004; Currie et al., 2015; van den Berg et al., 2020; De Cao et al., 2022; Akesaka and Kikuchi, 2024). As the Japanese public childcare system requires maternal employment to access daycare services, working during pregnancy may have adverse effects on birth outcomes; alternatively, stable employment may reduce maternal malbehavior and improve birth outcomes. In addition, working women may delay childbirth because of career concerns, and late-in-life pregnancies are associated with a higher risk of adverse birth outcomes. Thus, we analyzed the policy's impact on birth weight and gestational duration. However, we did not find any adverse policy effects on birth outcomes. We found that, if any, public

childcare provision slightly reduces premature pregnancy, which is likely due to increased birth rates by highly educated women, those who face a lower risk of birth outcomes.

The contributions of this study are twofold. First, we provide new evidence for analyzing the effect of public childcare on fertility. A majority of the literature relies on cross-country data for empirical analysis (e.g. Castles 2003; Del Boca et al. 2008; Borck 2014; Baizan et al. 2016; d'Albis et al. 2017; Olivetti and Petrongolo 2017). However, as noted by Olivetti and Petrongolo (2017), the causality of a cross-country analysis needs careful interpretation because country-level policy variation is likely to be confounded, for example, by family policies other than public childcare provisions or changes in social norms. Therefore, evidence from a single country with a clean research design is valuable. Notable exceptions include Rindfuss et al. (2010), Bauernschuster et al. (2015), Fukai (2017), Wood and Neels (2019) and Schuss and Azaouagh (2023). Rindfuss et al. (2010) and Wood and Neels (2019) conduct discrete-time hazard analysis with Norwegian and Belgian data, respectively, and found positive policy impact on birth hazard rates. Bauernschuster et al. (2015) and Schuss and Azaouagh (2023) analyze the same policy reform in West Germany, and found the positive policy effects. Fukai (2017) analyzes Japanese data, and found positive effect, particularly in region with strong female labor demand.

In particular, our study contributes to the literature by providing a rich set of heterogeneous policy effects, namely, birth parity, maternal age, and maternal education. While previous studies have not analyzed heterogeneity in these three dimensions simultaneously, such a heterogeneity analysis is essential to understand the mechanisms by which public childcare provision affects the birth rate. As public childcare services enable mothers of preschool-aged children to continue working, the policy effects may vary across different age groups. For instance, Caucutt et al. (2002) and Doepke et al. (2023) show that women have an incentive to delay childbirth because of career considerations because having a child early in their careers hinders human capital accumulation. In such cases, women who have had a child and quit their jobs may choose to delay childbirth in order to pursue their careers when public childcare is available. Moreover, public childcare could encourage women who would otherwise not have a child to reconsider, with its effects likely to be more pronounced among women in their 30s than among those in their 20s, again because of career considerations. This argument also suggests that analyzing heterogeneity by education is crucial for understanding how public childcare provision alleviates challenges in balancing work and family responsibilities, as the career cost of children is higher for university graduates than for non-graduates. Additionally, the decline in birth rates in many developed countries is driven by lower birth rates among highly educated women (Raute, 2019), a phenomenon referred to as the "baby gap," The share of childlessness

is particularly high among highly educated women (Doepke et al., 2023). Given these differences in fertility decisions across socioeconomic groups, the welfare implications of childcare policies vary depending on the extent to which the policy is effective across birth parities and socioeconomic statuses. Our study reveals that the policy is most effective for first births among university graduates and, as such, is effective in reducing the baby gap.

Second, this study addressed the temporal effects of birth timing. Goldstein et al. (2009) and Bongaarts and Sobotka (2012) argue that part of the declining trend in the TFR is an artifact of delayed childbirth. For instance, when women choose to have children at age 40 years rather than at age 30 years, the TFR temporarily declines and rises again a decade later. However, the complete fertility rate, which the TFR aims to approximate, remains unaffected. This suggests that if childcare policies influence birth timing, fertility rate estimates may be biased unless complete fertility is considered. In fact, Bauernschuster et al. (2015) and Fukai (2017) do not address this issue. In contrast, the extended duration of our sample period allowed us to address this issue by focusing on women whose entire birth histories were fully observed.

2 Institutional background

2.1 Formal childcare system in Japan

In this subsection, we describe the formal Japanese preschool childcare system, which primarily consists of two types of institutions: kindergartens and accredited daycare centers. Kindergartens focus on early childhood education and typically offer half-day care for children aged 3-–5 years, regardless of the parents' employment status. In contrast, accredited daycare centers are designed to assist parents who face challenges in caring for their children, such as dual-income families or those with sick family members. Therefore, these daycare centers offer full-time care for children aged 0-–5 years, primarily catering to families in which the mother works full-time. Importantly, daycare centers are available for 0–2-year-old children while kindergartens are not available for them. Given that the duration of Japanese parental leave is up to one year, it is crucial for parents that full-time care services are available immediately after parental leave.[1] In this study, we focused on the latter type of daycare center, as there is significant policy variation in the availability of their slots, driven by evident social requirements to support women's career advancement after childbirth, as will be discussed in the next subsection.

Accredited daycare centers in Japan are highly regulated by the government. For

[1]When parents cannot use daycare services due to capacity constraint of daycare centers, they are allowed to extend parental leave duration up to one and half years.

example, the maximum child-teacher ratio was three for 0-year-old children, six for 1–2-year-old children, 20 for 3-year-old children, and 30 for 4–5-year-old children.[2] In addition, regulations exist regarding the size of childcare facilities. The minimum area for infant room per child is 1.65m^2, room for crawling 3.30m^2, nursery room 1.98m^2, and play field 3.30m^2. Furthermore, the minimum capacity of childcare centers was set at 20 until 2015; therefore, childcare providers were not allowed to open small-scale centers. In terms of the educational attainment of caregivers, typical nursery teachers graduate from junior college or university specialized for nursery teachers; otherwise, they are required to pass the exam for a nursery teacher license after graduating (junior) college. Due to these regulations, childcare centers have faced a shortage of caregivers and a lack of sufficient areas to establish new daycare centers, resulting in a shortage of childcare slots, particularly for 0–2-year-old children.

Excess demand for daycare services is partly driven by price regulations coupled with generous government subsidies. Because the local government sets the price for daycare services, centers are not allowed to raise childcare fees even under high demand. Moreover, although the price is adjusted according to family income, it remains relatively low. In fact, in 2018, the average monthly cost of daycare for 2-year-old and 5-year-old child in the 23 wards of Tokyo was 25,000 JPY and 15,000 JPY, respectively. These rates are significantly lower than the average monthly full-time earnings of women in Tokyo, which was 397,000 JPY according to the Basic Survey on Wage Structure 2018. From 2019, public childcare for children aged 3–5 years has become free, regardless of their family income. In addition, low-income households are exempt from fees for childcare services for children aged 0–2 years.

2.2 Expansion of subsidized childcare centers

In the face of severe population aging, the Japanese government has struggled to increase both the fertility rate and female labor force participation. Among the various family policies, including parental leave and the right to work part-time for mothers, a key objective has been to expand the availability of accredited daycare centers since 1990s. This is particularly important, as the number of slots for children aged 0-–2 years has been limited, with some municipalities facing waiting lists for these age groups.

Following the rising demand for childcare, since 1995, substantial efforts have been made to expand public childcare facilities. That year marked the launch of a five-year plan ("Angel Plan") to develop childcare centers. In 2000, the *Child Welfare Act* further

[2]The government tighten this restriction to improve childcare quality in 2024. In particular, maximum ratio for 4–5-year-old children was reduced to 25. In 2025, maximum ratio for 1-year-old children will be reduced to 5.

increased the supply of public childcare services by relaxing entry restrictions for private care providers. In 2006, childcare services incorporating kindergarten elements were introduced in response to diverse demands. However, despite a series of policy plans to increase childcare coverage for children aged 0–2 years, the expansion of childcare capacity has remained relatively stagnant owing to strict government regulations. Between 1990 and 2010, the number of childcare slots per child increased from 0.21 0.28.

A significant shift occurred with a landmark legal amendment in 2015 that introduced a more flexible small-scale public childcare system aimed at accommodating children aged 0-–2 years. In particular, the minimum capacity was changed from 20 to 6, and these small-scale childcare centers could be opened without the need for playfields. As a result, this reform led to a substantial expansion of childcare coverage. Since these centers are compact, they are typically located in residential or office areas, or even within offices in some cases, allowing parents to easily drop off and pick up their children before and after work. By 2020, the number of slots per child increased to 0.53. This demonstrates that government regulations had a significant impact on childcare providers' ability to expand their capacity, particularly in establishing new childcare centers, until the 2015 reform.

3 Data

Our analysis mainly draws on Japanese birth record data, Vital Statistics for the period of 1990–2020. As these data are based on administrative birth records, they cover all births in Japan with accurate information, such as mother's age, birth parity, birth weight, and gestational weeks. Birth record data were most suitable for our analysis for four reasons. First, it covered all births during our analysis period, and the large sample size allowed us to conduct a rich heterogeneity analysis by birth parity and maternal age. This heterogeneity analysis is crucial for understanding the underlying mechanisms through which public childcare provision affects fertility rates. Second, the extended sample period allowed us to analyze completed fertility. Third, birth record data included birth outcomes, such as gestation duration and birth weight. Considering that mothers' employment opportunities can influence birth outcomes (De Cao et al., 2022; Akesaka and Kikuchi, 2024) and that access to public childcare services in Japan is restricted to working mothers, it is worth investigating the policy's effect on birth outcomes. Fourth, since it is part of the administrative data, it provides precise information about birth date, birth parity, parents' age, and birth outcomes.

However, because the Vital Statistics do not include maternal education, we supplemented these data with micro data from the Japanese population census to analyze the heterogeneous effects of public childcare provision by educational attainment and house-

hold type. The census takes place every five years and covers the entire population, with the advantage that it is a household survey that collects a rich set of demographic characteristics. Importantly, the census includes birth year and month as well as municipalities where a household resides, and using the 1990–2020 census, we can count the number of births in each municipality and census year by household characteristics. As education information is collected only in decennial censuses, we used census data from 1990, 2000, 2010, and 2020 to analyze the heterogeneous policy effects of maternal education.

We defined birth rate as the number of births divided by the number of women aged between 15 and 49 years in each municipality, where the number of births was obtained from the Vital Statistics and the number of women was obtained from the census. Because the census was collected every five years, we used the birth rate every five years between 1990 and 2020. When we investigate the policy impact on a specific age X of women, the birth rate is defined as the number of births given by mothers aged X divided by the number of women aged X in each municipality: Similarly, the birth rate of k-th birth and maternal age X are defined as the number of k-th births of mothers aged X divided by the number of women aged X in each municipality.

To measure the availability of public childcare, we used the Survey of Social Welfare Institutions 1990–2020 collected by the Ministry of Health, Labour and Welfare. The survey compiled the number of public childcare slots in each municipality in Japan. To measure childcare slots, we used the number of accredited childcare center slots that provide full-time daycare.[3] While childcare slots are set by child age, we measure the capacity of childcare centers by the total number of slots owing to data availability. In addition, we did not count the number of kindergartens, as they provided half-day care. We merged these data with birth records at the municipal level.

[Figure 1 about here.]

Figure 1 shows the growth in the number of public childcare slots per child under 6 (capacity rate) between 1990 and 2020. As the y-axis of the figure shows the rank of the capacity rate in 1990, the red points indicate the distribution of the capacity rate in 1990. The capacity rate in 1990 was limited, with a median slightly above 0.2, or, putting it differently, only one childcare slot for each of the five children. Furthermore, the capacity rate in most cities is less than 0.5. The blue points in the figure represent the capacity rate in 2020, and the blue bars represent the growth in the capacity rate between 1990 and 2020. Two facts are noteworthy. First, we find a substantial increase in the capacity rate

[3]From 1990 to 2010, we used the slot of public childcare centers (*ninka-hoikusho*). For 2015 and 2020, in addition to these public childcare centers, the slot included certified childcare facilities, such as *nintei kodomo-en*, small-scale childcare services (*shokibo-hoiku-jigyo*), and workplace childcare centers (*jigyoshonai-hoiku-jigyo*), all approved by the Ministry of Health, Labour and Welfare.

of almost all cities, which is consistent with a series of Japanese policy efforts to increase public childcare availability. As seen in Table 1, the mean capacity rate increased from 21% to 53% during this period. In particular, the expansion of coverage between 2010 and 2020 was most notable due to government de-regulation for small-scale childcare centers. Second, we observe a huge variation in the size of growth across cities, and the growth rate is not necessarily strongly related to the initial capacity rate. This variation reflects strict government regulations, with which some cities have struggled to expand childcare capacity despite the high demand for full-time childcare services, as discussed earlier.

[Table 1 about here.]

Table 1 shows the descriptive statistics of the birth outcomes, along with covariates that are used as control variables in the regression analysis. The birth rate remained relatively stable at approximately 4% throughout the sample period. Rows 2–5 show annual birth rates by maternal age. The birth rate is highest among mothers aged 25–34 years, and approximately 10% give birth each year. It is also notable that late-in-life pregnancy has increased substantially during the last decades, and the birth rate between the ages of 35 and 49 years has more than tripled, whereas the birth rate of young women aged 15–24 years has decreased. The remaining rows in Panel A show birth weights and gestation weeks with evidence of negative time trends, indicating that these birth outcomes worsened during the sample period. In fact, compared to 1990 and 2020, the share of low-birthweight infants (less than 2500g) increased from 6.3% to 9.4%, and the share of births with gestation periods shorter than 37 weeks increased from 4.6% to 5.7%. These negative trends in birth outcomes may be linked to an increase in maternal age at childbirth because late-in-life pregnancy is associated with a higher risk of adverse birth outcomes. Given that the Japanese public childcare system requires maternal employment, it is possible that women who choose to have a child due to improved childcare availability delay childbirth due to career concerns, and may experience adverse effects on birth outcomes as a result of late-in-life pregnancy. Furthermore, employment during gestation can negatively impact maternal health and birth outcomes (Akesaka and Kikuchi, 2024).

[Figure 2 about here.]

Figure 2 descriptively examines the relationship between the birth rate and the number of childcare slots. To partial out the time-invariant unobserved heterogeneity of each city, we plotted these variables after subtracting the city-specific mean for each 5 years.[4]

[4]We plotted separately by each year to avoid spurious negative correlation driven by time trends. Due to decreasing trend in fertility rate and increasing trend in childcare slots, if we pool all years to plot these two variables, we observe negative relation. As this is because of national yearly trend, this effect is removed by year fixed effects in regression analysis.

We found a positive association between birth rate and childcare capacity, suggesting a positive effect of the policy on fertility. In the next section, we confirm that this relationship is not a correlation, but a causal one via regression analysis.

4 Estimation method

To analyze the effect of public childcare provision on fertility, we estimated the following equation:

$$y_{ajt} = \beta \ln Slot_{jt} + \gamma \ln Women_{jt} + x'_{jt}\delta + \eta_{aj} + \pi_t + u_{jt}, \tag{1}$$

where a, j and t indicate maternal age, municipality, and year, x_{jt} is the prefecture-specific year trend with a second-order polynomial; and y_{ajt} is the age-specific birth rate. $Slot_{jt}$ is the number of slots in public childcare centers and $Women_{jt}$ is the number of women aged between 15 and 49 years.[5] Because we control for the number of women who can potentially give birth, the coefficient of interest β represents the effect of increasing the number of public childcare slots while holding the number of potential mothers constant. Because our model also includes age-municipality fixed effects η_{aj} and year–fixed effects π_t, the identification relies on within-municipality variations in the number of slots.

To further facilitate the interpretation of the coefficient, we note that our model is equivalent to the model with the potential capacity rate defined as the number of slots divided by the number of women. The model can be rewritten as

$$y_{ajt} = \beta \ln \left(\frac{Slot_{jt}}{Women_{jt}} \right) + \tilde{\gamma} \ln Women_{jt} + x'_{jt}\delta + \eta_{aj} + \pi_t + u_{jt},$$

where $\tilde{\gamma} = \gamma + \beta$. In this sense, β is interpreted as the effect of an increase in childcare slots per woman on the fertility rate, holding the number of women constant. Including the number of women as a control variable is indispensable, because it affects fertility and potential capacity rates.

Our key identification assumption is that within-municipality variations in childcare slots are exogenous to shocks that affect fertility rates. A major concern is the reverse causality, in which local governments adjust their childcare capacity in response to the fertility rate. This reverse causality would be more problematic in rural areas, where childcare slots decrease in response to a decline in birth rates. While it is possible that a local government increased childcare capacity to counteract the declining fertility rate, the policy effect would be underestimated. Thus, our estimation result would serve as the lower bound of the true policy effect. By contrast, in urban areas, there is a substantial

[5]To account for different population sizes across municipalities, we weight the observations by the number of women in each municipalities in the regression.

excess demand for childcare services due to strict government regulations. Some cities have successfully increased their capacity, while others have struggled despite high demand. This provides exogenous policy variation in the expansion of childcare slots to identify the causal effect of public childcare provision on the fertility rate. Therefore, our main sample consisted of cities and excluded small villages and towns.[6] However, we also use a sample that includes all municipalities to confirm the robustness of our empirical evidence.

To facilitate the interpretation of our estimation results, we introduce a method for translating our estimates into hazard rates. Letting $T^{(k)}$ be a discrete random variable for the timing of k-th birth, we denote the proportion of women who gives k-th birth at t as $p_t^{(k)} = \Pr\left(T^{(k)} = t\right)$. Then, since the proportion of women who have not given k-th birth yet at $t-1$ is $1-\sum_{s=1}^{t-1} p_s^{(k)}$, the hazard rate is written as $\Pr(T^{(k)} = t | T^{(k)} > t-1) = \frac{p_t^{(k)}}{1-\sum_{s=1}^{t-1} p_s^{(k)}}$.[7] Because $p_t^{(k)}$ corresponds to the dependent variables in equation (1), they are used to construct the baseline hazard rate. Furthermore, we have $\Delta p_t^{(k)} = \beta_a^{(k)} \Delta \ln Slot$, where $\beta_a^{(k)}$ is the age-and birth parity-specific policy effect estimated using equation (1). Hence, defining $\ddot{p}_t^{(k)} = p_t^{(k)} + \beta_t^{(k)} \Delta \ln Slot$, the hazard rate after the childcare expansion is written as $\frac{\ddot{p}_t^{(k)}}{1-\sum_{s=1}^{t-1} \ddot{p}_s^{(k)}}$. To evaluate the policy effect on hazard rates, we considered a policy scenario with a 70% increase in childcare slots, which is approximately the average growth in childcare slots between 1990 and 2020.

5 Effects of public childcare provision on fertility

5.1 Baseline results

[Table 2 about here.]

Table 2 presents the estimation results of equation (1). Column 1 shows our preferred specifications using only cities as the analysis sample. We find a significantly positive effect of public childcare slots on the birth rate. Between 1990 and 2020, the number of slots increased by approximately 73%, which is associated with 0.28 points increase, or a 6.5% increase in the birth rate. Therefore, this estimation result implies that the effect of public childcare on birth rate is moderate. Column 2 shows the estimation results using marriage rate as a control variable. Although the marriage rate could be a bad control, since marriage is a mediator for a couple to have a child, we conducted a robustness check by controlling for the marriage rate. As expected, the magnitude of the

[6] A municipality is required to have at least 50,000 population to become a city.

[7] To see this, let $N_t^{(k)}$ be the number of women who have not given k-th birth at t, where N_0 is population size. Then, the number of k-th birth at t is $n_t^{(k)} = p_t^{(k)} N_0$, and we have $N_t^{(k)} = N_{t-1}^{(k)} - n_t^{(k)}$. Thus, it follows that $N_t^{(k)} = N_0 \left(1 - \sum_{s=1}^{t} p_t^{(k)}\right)$.

estimated coefficient decreases; however, we still find a significant positive effect. Because this regression indicates that the effect of childcare slots is positively significant, holding the marriage rate constant, it suggests that family formation or marriage is not the main mechanism through which childcare provision affects the fertility rate, and public childcare provision encourages married couples to have additional children. As discussed later, we found a policy effect in both the extensive margin (first births) and the intensive margin (second and third births).

We also test the robustness of our findings using data from all municipalities (Column 3). In this specification, the estimated effect was larger than our baseline estimate. This is likely due to the reverse causality. In rural areas, the demand for public childcare has been decreasing because of severe population aging and the overall population decline. As a result, municipalities in these areas may reduce childcare slots—for example, by closing daycare centers in response to lower demand—which could introduce an upward bias in our estimates. By contrast, our baseline specification focuses on urban areas, where we exploit the staggered expansion of childcare capacity. These municipalities are more likely to face higher demand for childcare services but struggle to expand their capacity due to the regulation of childcare quality and local budget constraints. Consequently, the timing of childcare expansion in urban areas does not necessarily align with that of increased demand. Therefore, we preferred sample restrictions that excluded towns and villages.

As the types of childcare centers expanded differently before and after 2015, we divided our sample period into two segments: 1990–2010 and 2010–2020. As explained in Section 2, public childcare centers were highly regulated by the government until 2015, which led to a sluggish expansion of childcare slots during this period. The capacity rate, defined as the number of slots per child, increased by 31% between 1990 and 2010. However, in 2015, the government lifted some regulations on public childcare services, notably allowing small-scale childcare centers to care for children aged 0–2 years. As a result, the capacity rate increased by 38% from 2010 to 2015, surpassing 50% in 2020, and nearly doubling the capacity compared with 2010. As the expansion of childcare capacity after 2015 is exclusively for 0–2 year-old children, given the excess demand for daycare for infants, we expect that childcare expansion after 2015 will have a larger impact than the expansion before 2015.

Columns 4 and 5 present the results estimated separately for the two sample periods. Although we found positively significant estimates in both periods, the differences in policy effects were sizable. During 1990 and 2010, a 100% increase in childcare slots increased the fertility rate by 0.23 percentage points, or 5.4% compared with the mean fertility rate, whereas it increased the fertility rate by 0.5 percentage points or 11.9% during 2010–2020. Put it differently, the magnitude of the policy effect in the latter period is more than twice

as large as the effect in the former period. This finding highlights that providing public childcare for infants is critical for increasing the fertility rate.

[Figure 3 about here.]

To explore the heterogeneous effects of the policy by maternal age, we estimate equation (1) for each mother's age (Figure 3a). Our results show that the policy's effect is most pronounced among women in their late 20s. Overall, the pattern of heterogeneity appears to be similar to the distribution of birth rates by maternal age (Figure 3b), suggesting that the effect of public childcare provision is observed across a wide range of maternal ages. However, the effect on late-in-life births is substantial compared with the baseline birth rate shown in Figure 3a. This is likely because women working full time may hesitate to have children early in their careers because of concerns about career progression (Caucutt et al., 2002; Doepke et al., 2023). As the primary objective of Japanese public childcare services is to support women in maintaining full-time employment after childbirth, they may be incentivized to delay having children and pursue their careers instead of quitting their jobs at the time of childbirth. Considering that career interruptions during pregnancy and parental leave can hinder human capital accumulation, public childcare provision may lead them to choose to have children later after establishing their career path.

[Figure 4 about here.]

Figure 4a shows the estimation results for first births by maternal age. We found that the effect of public childcare provision on first birth rate was positive and statistically significant across a wide range of maternal ages. While the birth rate of first births peaks at age 27 years (Figure 4b), the magnitude of the policy effect is similar across women in their 20s and early 30s. The estimated coefficients indicate that a 70% increase in childcare slots, which is the average increase between 1990 and 2020, is associated with 0.28 point ($= 100 \times 0.004 \times 0.70$) increase in age- and birth parity-specific birth rates. Figure 5a converts the birth probability and policy effects into a hazard rate, assuming a 70% capacity increase, to draw the blue line in the figure. We found a disproportionately large impact on birth hazards in women in their 30s compared with those in their 20s, reconfirming our finding that public childcare provision increases late-in-life pregnancy.

There are two potential explanations for this finding: tempo effect and composition effect. First, public childcare provision can influence the timing of birth for mothers with children, even without access to public childcare. For example, a woman may quit her job after childbirth if childcare is not available, and have a child at a relatively young

age, which may be more biologically favorable. In contrast, when public childcare is available, she may continue her job after childbirth, delaying childbirths to pursue her career. Second, public childcare provision can encourage women who would otherwise remain childless to have a child by making it easier to balance work and family responsibilities. Considering that these women are more likely to be career-oriented and have higher than average levels of education, it is expected that their childbirth timing will tend to be later than that of other women.

To discuss the relative magnitude of these two effects, we note that the difference between the tempo and composition effects is that the former does not increase completed fertility, whereas the latter does. Therefore, if the temporal effect dominated the composition effect, we would observe negative policy effects for some age groups to offset the temporal increase in late-in-life pregnancies. However, as we observed positive effects for almost all ages, it seems that the composition effect dominates the temporal effect. In other words, public childcare reduces childless women, and those affected by the policy tend to have children later in life. While d'Albis et al. (2017) emphasized the role of public childcare in increasing higher-order births, our findings underscore that it is effective in increasing first births as well. We further address the issue of tempo effect in Section 5.4.

Figure 4c presents the estimation results for second births by maternal age. In contrast to the first-birth results, we found an M-shaped pattern that peaked in the late 20s and late 30s. The magnitudes of the estimates are comparable to those of the estimates for first births, although they tend to be slightly smaller. Similar to the effect on first births, the policy effect on the hazard rate of second births was more notable for women in their 30s than for women in their 20s (Figure 5b).

Figure 4e presents the estimation results for third births by maternal age. The effect on third births is much smaller than that on first or second births, and the effect is limited to those aged 28–31 years, which is earlier than the peak fertility rate of third births (Figure 4f). The hazard rate in Figure 5c indicates that the policy has no effect on late-in-life pregnancies for third births. This finding is somewhat interesting, given that we found a positive impact on second births for those in late 30s as well as those in their late 20s. Combining these two findings suggests that public childcare provision does not encourage those who gave a second birth in their 30s to give a third birth, but is only effective for those who gave birth at a younger age in terms of third births. This might explain why the effect on second births is M-shaped, as those who do not plan to have a third child would delay their second birth; thus, the policy impact on their second birth is most evident in their late 30s. Meanwhile, those who plan to have three or more children would not delay their second birth to have a biologically favorable birth spacing; thus, the policy impact on their second birth is most evident in their late 20s.

[Figure 5 about here.]

5.2 Policy effects on total fertility rate

Because we estimated the policy impact by maternal age and birth parity, we can use these estimates to infer the policy impact on total fertility rate (TFR) and decompose it into the contribution of first, second, and third births. In particular, let $n_a^{(k)}$ be the number of women who gave k-th birth at age a, and N and N_a be total numbers of women and women aged a. Then, the number of k-th child is $n^{(k)} = \sum_{a \in A} n_a^{(k)}$, and total number of children is $n = n^{(1)} + n^{(2)} + n^{(3)}$, so $TFR = \sum_{a \in A} \frac{n_a}{N_a} = \sum_{a \in A} \sum_{k=1}^{3} \frac{n_a^{(k)}}{N_a}$. Because $\frac{n_a^{(k)}}{N_a}$ is the dependent variable of our estimation equation (1), when we analyze the policy effect of birth parity and maternal age, the policy effect on completed fertility is calculated by

$$\frac{\partial TFR}{\partial \ln Slot} = \sum_{k=1}^{3} \sum_{a \in A} \frac{\partial}{\partial \ln Slot} \frac{n_a^{(k)}}{N_a} = \sum_{k=1}^{3} \sum_{a \in A} \beta_a^{(k)}, \tag{2}$$

where $\beta_a^{(k)}$ is the parameter estimated in equation (1). In addition to the effect on the TFR, equation (2) allows us to calculate the contribution of each birth parity to the total effect, denoted as $\frac{\sum_{a \in A} \beta_a^{(k)}}{\sum_{k=1}^{3} \sum_{a \in A} \beta_a^{(k)}}$.

[Table 3 about here.]

Table 3 shows that a 70% increase in the number of childcare slots increased the TFR by 0.097 ($= 0.7 \times 0.139$). From 1990 to 2020, the TFR decreased by 0.2 points from 1.54 to 1.34, and thus, our estimate suggests that the decline of the TFR would have been an additional 0.1 points without the expansion of childcare.[8] Therefore, although childcare expansion is not a panacea for increasing fertility rates, our findings demonstrate that its contribution is non-negligible.

Furthermore, our decomposition analysis revealed that the increase in first births accounted for 53% of the total effect, while second births accounted for 41% and third births accounted for 5%. Thus, the impact of childcare provision is largely driven by the first and second births. The sizable impact on second births is probably due to access to public childcare services, reduced childcare costs for women, and reconciling disagreements within a couple regarding family planning. Doepke and Kindermann (2019) report that

[8] Unayama (2023) analyzes the effect of public childcare provision on the marriage rate and infers the policy's impact on the TFR to be 0.1 between 2005 and 2020, based on a back-of-the-envelope calculation using the estimated policy effect on *marriage rate*, and using the average fertility rate of married couples. Given that the number of childcare slots increased by 61% during this period, our more rigorous analysis, which directly estimated the policy effect on fertility, suggests that the TFR increased by 0.085 (i.e., 0.139×0.61). Our findings support this argument.

a wife and husband tend to have different preferences over children. Typically, husbands prefer to have more children than wives, and this disagreement becomes more pronounced for second child than first child. Furthermore, the large impact on second births is consistent with the findings of d'Albis et al. (2017), who theoretically and empirically analyzed the positive effect of public childcare services on second births using cross-country data.

Meanwhile, our finding contrasts with Bauernschuster et al. (2015), who find that the policy effect is particularly remarkable for second and third births in West Germany, rather than first births, whereas we found the policy impact most remarkable for first births. This may be attributed to differences in childcare services between Japan and Western Germany. As Japanese public childcare centers offer full-time care, as opposed to half-day care in West Germany, women who would otherwise be childless to pursue their careers are also encouraged to have a child, suggesting that full-time care is an essential feature of public childcare that incentivize working women to have a child.

5.3 Heterogeneous effects by maternal education

In this subsection, we analyze the heterogeneity of maternal education using census data. Heterogeneity by education is of interest, because highly educated women may face higher career costs associated with having a child, which creates challenges in balancing work and family responsibilities. In fact, Raute (2019) show that the decline in birth rates in many developed countries can be explained by lower birth rates among highly educated women. In line with her findings, Doepke et al. (2023) show that childlessness is particularly prevalent among highly educated women. Public childcare policies are expected to assist career-oriented women by providing affordable and reliable childcare; thus, they may be more effective for highly educated women.

[Table 4 about here.]

Before conducting the heterogeneity analysis, we replicate our baseline results reported in Table 2 to validate the census data. Column 1 of Table 4 uses quinquennial census data. As education information is collected only in the decennial census, we rerun the regression using only the decennial census data (Column 2). Although the estimates from the census are slightly smaller than those from Vital Statistics, the estimated coefficients are both positive and statistically significant. The magnitude of the policy effect was smaller in the census than in the Vital Statistics, likely because of survey non-response in the census by couples with children. In fact, the average yearly birth rate in the census is approximately 3.6 percent, whereas the corresponding figure in Vital Statistics is 4.3 percent. Nevertheless, keeping this limitation in mind, we confirmed that the census is a reliable data source for analyzing birth rates.

To conduct the heterogeneity analysis, we divided our sample according to maternal education. Columns 3 and 4 show the estimation results for non-graduates and university graduates, respectively. In this analysis, we restricted our sample to those aged 23 years or older to facilitate comparison between education groups. While both estimates were positive and statistically significant, we found that childcare policy was more effective for university graduates than for non-graduates. In fact, the policy effect on university graduates was twice as large as that on non-graduates.

[Figure 6 about here.]

Figure 6 provides a more nuanced analysis of heterogeneity based on maternal education, age, and birth parity. We found that the effect of first births was most pronounced at the age when the birth rate peaked: 23–27 years for non-graduates and 28–32 years for university graduates. However, we also found a relatively large effect on late-in-life first births in both educational groups. When we add the estimates by age group, the effect on first births is larger for university graduates than for non-graduates, indicating that the childcare policy is effective in reducing childlessness, particularly among university graduates.

Regarding higher-order births, the results show a clear contrast between the education groups. For non-university graduates, the effect of a second birth was positive at a relatively young age. In fact, while the second birth rate peaks at age 30 years among non-graduates, the effect is comparable for ages 23–27 and 28–32 years. However, no significant effect was found for second births in the late 30s. For university graduates, however, we found a sizable positive effect on second births in the late 30s, whereas the policy had no impact on those aged 23–27 years.

These heterogeneous policy effects align with the M-shaped pattern of second births shown in Figure 4. The first peak in the policy effect is likely driven by non-graduates, whose lower career interruption costs seem to motivate childbearing at a younger age, with biological considerations outweighing career concerns. In contrast, university graduates tend to delay childbirth, accepting a trade-off between delayed fertility and career advancement. This observation aligns with theoretical predictions regarding the tradeoff between career and biological clocks (Caucutt et al., 2002; Doepke et al., 2023). Consequently, our findings suggest that although childcare policies may reduce childbearing costs and boost birth rates, they are not sufficient to fully overcome this trade-off.

5.4 Robustness check

Tempo effect

One concern in our analysis is that some of our estimates of policy effects could be influenced by the tempo effect, —changes in the timing of births. As we focused on age-specific birth rates rather than completed fertility rates, any shifts in birth timing within cohorts whose completed fertility had not yet been observed could bias our estimates. For example, childcare policies may encourage women to delay childbirth to pursue their careers, which could lead to an underestimation of policy effects if delayed births are not considered. Ideally, complete fertility should be analyzed to avoid this issue. However, this approach presents challenges, as it is difficult to assign childcare availability to individuals, especially considering that childcare access changes over time, and couples may move between municipalities, which we cannot fully track.

To mitigate the potential tempo effect, we re-estimate equation (1) using only birth cohorts from 1970 to 1980, as we observe birth behavior between the ages of 20 and 40 years. Column 6 of Table 2 presents the estimation results based on the Vital Statistics. Despite the differences in the cohorts analyzed, the estimate is similar in magnitude to the baseline estimate in Column 1, which includes the birth cohorts from 1941 to 1995. If there are any differences, the estimate is slightly smaller, likely because the analysis sample excludes more recent cohorts that experienced a significant increase in public childcare availability after 2015. In fact, the estimation result is very similar to the estimate from the sample period between 1990–2010 (Column 4 of Table 2) This analysis supports the robustness of our finding that policy effects are positive, even when accounting for the tempo effect.

Policy endogeneity

Another important factor is policy endogeneity. Specifically, it is possible that the expansion of childcare slots was targeted at municipalities with a high excess demand for childcare, that is, those areas that experienced a significant increase in fertility rates. In this case, the expansion might respond to higher fertility rates, which leads to a positive correlation between policy and birth rates, creating a potential upward bias in our estimates. Alternatively, the increase in childcare slots could have been implemented as a countermeasure to declining birth rates, in which case public childcare is more accessible in municipalities with lower fertility, potentially introducing a downward bias in our estimates. To address this issue of reverse causality, we re-estimate equation (1) using the number of childcare slots from five years prior as an instrumental variable (IV). This approach helps mitigate the concern that fertility trends influence policy expansion. Column

1 of Table 5 presents the estimation results of the FE-IV regression using Vital Statistics. The estimation results are virtually identical to the baseline estimates (see Column 1 of Table 2). Therefore, this analysis confirmed that policy endogeneity was not the primary driver of our results.

[Table 5 about here.]

Selective migration

We also address the issue of selective migration, which could have biased our estimates. Selective migration can occur in at least two ways. First, couples planning to have a child may migrate to a city with better access to public childcare services. In this case, selective migration may have caused an upward bias in the estimate. Second, couples planning to have a child may move to an urban area with better employment opportunities because the Japanese public childcare services require maternal employment. Given that the capacity of public childcare centers tends to be lower in urban areas, this type of selective migration could result in a downward bias in our estimates.

To assess the extent of selective migration, we used census data to calculate the share of women who migrated from other cities. The census asked respondents where they lived five years ago, which allowed us to calculate the share of migrants in each city. Then, we estimate equation (1) using this migration share as the dependent variable. A positive coefficient for the number of slots suggests selective migration based on childcare availability and a negative coefficient suggests selective migration based on employment opportunities. The estimated coefficient is -0.0238 with a standard error of 0.0069, indicating that a 100 percent increase in the number of slots reduces the inflow of women by 2.4 percentage points. Therefore, selective migration is more likely to bias our estimate downward, suggesting that our results represent the lower bound of the policy effect.

To address potential biases from selective migration, we constructed an instrument for the number of slots based on cities where women lived five years ago. Formally, let $j_{(i,t)}$ denote the city where woman i lived in year t and I_{jt} denote the set of women living in city j in year t. The instrument is defined as $\widetilde{Slot}_{jt} = \frac{1}{|Ijt|}\sum_{i \in I_{jt}} Slot_{j_{(i,t-5)}t}$, which is the average number of slots weighted by the distribution of cities in which women in city j lived five years ago. Given the possibility of negative selective migration, couples planning to have children are likely to migrate from cities with a higher childcare capacity to cities with a lower capacity. This suggests that the unobserved factors affecting birth rates are positively correlated with $Slot_{j_{(i,t-5)}t}$, leading to an upward bias in the IV estimate. Therefore, the IV estimate represents the upper bound of the policy effect, whereas the baseline estimate represents the lower bound.

By estimating equation (1), using $\ln \widetilde{Slot}_{jt}$ as an instrument for $\ln Slot_{jt}$, we find that the estimated coefficient is 0.0076 (see Column 2 of Table 5). As expected, this estimate is larger than the baseline estimate of 0.0038. Therefore, by combining these two estimation results, the true policy effect is likely to be between 0.0038 and 0.0076. This analysis confirms that public childcare services have a positive impact on birth rates.

6 Effects of public childcare provision on maternal and child health

Because Japanese public childcare centers require mothers to use childcare services, women continue to work during pregnancy, which could potentially deteriorate maternal and child health at birth. For example, Akesaka and Kikuchi (2024) show that better employment opportunities for women lead to an increase in low-birth weight children, whereas other studies find that stable employment opportunities improve birth outcomes and maternal health (Currie et al., 2015; De Cao et al., 2022). In addition to these direct effects, public childcare provision may also have an indirect effect on birth outcomes through changes in the composition of mothers. Considering that childcare availability increases fertility rates, the characteristics of the marginal women most affected by the policy may differ from those of average mothers. For instance, public childcare provision may be particularly effective for high (low) income families, who are typically associated with a low (high) risk of birth complications (Currie, 2009). Because it is difficult to separate these two channels, our estimand is the net effect of both channels.

To address these potential concerns regarding birth outcomes, we analyzed how maternal and child health regarding childbirth is affected by the provision of public childcare. Birth weight and gestational age were used as measures of health outcomes. We defined low birth weight as birth weight less than 2,500g, 1500g or 1000g, and premature birth as birth at gestation weeks less than 37 or 34. Because birth weight and gestation duration are recorded in birth records, they are measured precisely. We estimate equation (1) using individual-level birth outcomes as a dependent variable; thus, the unit of observation of this analysis is individuals, as opposed to cities.

[Table 6 about here.]

[Table 7 about here.]

Table 6 presents the estimation results for birth weight. We observed no adverse or positive effects on birth weight. While the estimated coefficients indicate that the incidence of low birth weight is decreased by public childcare provision, the estimates are

not only statistically insignificant but also economically insignificant. In fact, the 95% confidence interval for the low birth weight (2500g) indicator is [-0.0031, 0.0017], such that even with a 100% increase in public childcare slots, the adverse effect on low birth weight is, if any, limited to 0.17 points, or 2% compared to the mean, at most. Therefore, our finding of no impact on birth weight is not due to imprecise estimates, and the magnitude of the policy impact is negligible, even considering the estimation errors. Even when we analyzed extremely low birth weights of less than 1500g or 1000g, we did not find any policy effects.

In contrast to birth weight, we found a positive policy effect on gestation weeks (Table 7). In particular, the expansion of public childcare has decreased the incidence of premature births. The size of the effect is non-negligible, and a 70% increase in childcare slots decreases the number of premature births with gestation weeks less than 37 and 34 by 3.4% and 6.4%, respectively, compared with the means. This effect was probably due to the change in the proportion of women who gave birth. As shown earlier, childcare policies are more effective in increasing birth rates among university graduates than among non-graduates. Consequently, as highly educated women generally face lower risks related to child health, this shift in demographic composition may lead to improved average birth outcomes. To summarize, our analysis shows no evidence of the adverse effects of public childcare provision on birth outcomes and, if any, it has a positive effect.

7 Conclusion

This study analyzes the effects of public childcare provision on fertility. Using micro data of birth records and the population census between 1990 and 2020, we found that a 70% increase in childcare slots, which is the average growth rate between 1990 and 2020, increases the yearly fertility rate by 0.28, or 6.5%. Furthermore, we investigated the heterogeneity by birth parity and maternal age. Our empirical evidence shows that the effect on first births was particularly large among women in their late 20s to early 30s, while the effect on second births had an M-shaped pattern and was notable among women in their late 20s and late 30s. The effect on third births was limited to those in early 30s. Furthermore, using these estimation results, we infer the effect of public childcare provision on total fertility rate, finding that a 70% increase in childcare slots is associated with an increase in TFR by 0.097, and this effect is mostly explained by increases in first and second births, as opposed to third births. Additionally, our heterogeneity analysis of maternal education demonstrated that the policy effect was particularly large among highly educated women. Finally, we also examined the policy impact on health outcomes regarding childbirth; we did not find any adverse impact on gestational duration or birth

weight.

References

Akesaka, Mika and Nobuyoshi Kikuchi, "The effects of gender-specific local labor demand on birth and later outcomes," *Labour Economics*, 2024, *90*, 102546.

Baizan, Pau, Bruno Arpino, and Carlos Eric Delclòs, "The effect of gender policies on fertility: The moderating role of education and normative context," *European Journal of Population*, 2016, *32*, 1–30.

Baker, Michael, Jonathan Gruber, and Kevin Milligan, "Universal Child Care, Maternal Labor Supply, and Family Well-Being," *Journal of Political Economy*, 2008, *116* (4), 709–745.

Bauernschuster, Stefan, Timo Hener, and Helmut Rainer, "Children of a (Policy) Revolution: The Introduction of Universal Child Care and Its Effect on Fertility," *Journal of the European Economic Association*, 11 2015, *14* (4), 975–1005.

Boca, Daniela Del, Silvia Pasqua, and Chiara Pronzato, "Motherhood and market work decisions in institutional context: a European perspective," *Oxford Economic Papers*, 12 2008, *61* (suppl_1), i147–i171.

Bongaarts, John and Tomáš Sobotka, "A Demographic Explanation for the Recent Rise in European Fertility," *Population and Development Review*, 2012, *38* (1), 83–120.

Borck, Rainald, "Adieu Rabenmutter—culture, fertility, female labour supply, the gender wage gap and childcare," *Journal of Population Economics*, 2014, *27*, 739–765.

Canaan, Serena, "Parental leave, household specialization and children's well-being," *Labour Economics*, 2022, *75*, 102127.

Castles, Francis G., "The World Turned Upside Down: Below Replacement Fertility, Changing Preferences and Family-Friendly Public Policy in 21 OECD Countries," *Journal of European Social Policy*, 2003, *13* (3), 209–227.

Caucutt, Elizabeth M., Nezih Guner, and John Knowles, "Why Do Women Wait? Matching, Wage Inequality, and the Incentives for Fertility Delay," *Review of Economic Dynamics*, 2002, *5* (4), 815–855.

Cornelissen, Thomas, Christian Dustmann, Anna Raute, and Uta Schönberg, "Who Benefits from Universal Child Care? Estimating Marginal Returns to Early Child Care Attendance," *Journal of Political Economy*, 2018, *126* (6), 2356–2409.

Currie, Janet, "Healthy, Wealthy, and Wise: Socioeconomic Status, Poor Health in Childhood, and Human Capital Development," *Journal of Economic Literature*, March 2009, *47* (1), 87–122.

_ , Valentina Duque, and Irwin Garfinkel, "The Great Recession and Mothers' Health," *The Economic Journal*, 2015, *125* (588), F311–F346.

Datta Gupta, Nabanita and Marianne Simonsen, "Non-cognitive child outcomes and universal high quality child care," *Journal of Public Economics*, 2010, *94* (1), 30–43.

De Cao, Elisabetta, Barry McCormick, and Catia Nicodemo, "Does unemployment worsen babies' health? A tale of siblings, maternal behaviour, and selection," *Journal of Health Economics*, 2022, *83*, 102601.

Dehejia, Rajeev and Adriana Lleras-Muney, "Booms, Busts, and Babies' Health*," *The Quarterly Journal of Economics*, 08 2004, *119* (3), 1091–1130.

Doepke, Matthias and Fabian Kindermann, "Bargaining over Babies: Theory, Evidence, and Policy Implications," *American Economic Review*, September 2019, *109* (9), 3264–3306.

_ , Anne Hannusch, Fabian Kindermann, and Michèle Tertilt, "The economics of fertility: a new era," in Shelly Lundberg and Alessandra Voena, eds., *Handbook of the Economics of the Family*, Vol. 1, North-Holland, 2023, pp. 151–254.

d'Albis, Hippolyte, Paula E. Gobbi, and Angela Greulich, "Having a second child and access to childcare: Evidence from European countries," *Journal of Demographic Economics*, 2017, *83* (2), 177–210.

Fernández-Kranz, Daniel, Aitor Lacuesta, and Núria Rodríguez-Planas, "The motherhood earnings dip: Evidence from administrative records," *Journal of Human Resources*, 2013, *48* (1), 169–197.

_ and Núria Rodríguez-Planas, "Too family friendly? The consequences of parent part-time working rights," *Journal of Public Economics*, 2021, *197*, 104407.

Fitzpatrick, Maria Donovan, "Preschoolers Enrolled and Mothers at Work? The Effects of Universal Prekindergarten," *Journal of Labor Economics*, 2010, *28* (1), 51–85.

Fukai, Taiyo, "Childcare availability and fertility: Evidence from municipalities in Japan," *Journal of the Japanese and International Economies*, 2017, *43*, 1–18.

Goldstein, Joshua R., Tomáš Sobotka, and Aiva Jasilioniene, "The End of "Lowest-Low" Fertility?," *Population and Development Review*, 2009, *35* (4), 663–699.

Havnes, Tarjei and Magne Mogstad, "Money for nothing? Universal child care and maternal employment," *Journal of Public Economics*, 2011, *95* (11), 1455–1465. Special Issue: International Seminar for Public Economics on Normative Tax Theory.

_ and _ , "No Child Left Behind: Subsidized Child Care and Children's Long-Run Outcomes," *American Economic Journal: Economic Policy*, May 2011, *3* (2), 97–129.

_ and _ , "Is universal child care leveling the playing field?," *Journal of Public Economics*, 2015, *127*, 100–114. The Nordic Model.

Kleven, Henrik, Camille Landais, and Gabriel Leite-Mariante, "The Child Penalty Atlas," *The Review of Economic Studies*, 10 2024, p. rdae104.

_ , _ , and Jakob Egholt Søgaard, "Children and Gender Inequality: Evidence from Denmark," *American Economic Journal: Applied Economics*, 2019, *11* (4), 181–209.

_ , _ , **Johanna Posch, Andreas Steinhauer, and Josef Zweimuüller**, "Child Penalties across Countries: Evidence and Explanations," *AEA Papers and Proceedings*, 2019, *109*, 122–26.

Olivetti, Claudia and Barbara Petrongolo, "The economic consequences of family policies: Lessons from a century of legislation in high-income countries," *Journal of Economic Perspectives*, 2017, *31* (1), 205–230.

Raute, Anna, "Can financial incentives reduce the baby gap? Evidence from a reform in maternity leave benefits," *Journal of Public Economics*, 2019, *169*, 203–222.

Rindfuss, Ronald R., David K. Guilkey, S. Philip Morgan, and ØYstein Kravdal, "Child-Care Availability and Fertility in Norway," *Population and Development Review*, 2010, *36* (4), 725–748.

Rossin-Slater, Maya, "Maternity and family leave policy," in Laura M. Argys Susan L. Averett and Saul D. Hoffman, eds., *Oxford Handbook on the Economics of Women*, Oxford University Press, 2018.

Schuss, Eric and Mohammed Azaouagh, "The expansion of early childcare and transitions to first and second birth in Germany," *Bulletin of Economic Research*, 2023, *75* (2), 476–507.

Unayama, Takashi, "Building for Babies: Impacts of childcare expansion on fertility rates and marriage (in Japanse)," *RIETI Discussion Paper*, 2023.

van den Berg, Gerard J., Alexander Paul, and Steffen Reinhold, "Economic conditions and the health of newborns: Evidence from comprehensive register data," *Labour Economics*, 2020, *63*, 101795.

Wood, Jonas and Karel Neels, "Local childcare availability and dual-earner fertility: Variation in childcare coverage and birth hazards over place and time," *European Journal of Population*, 2019, *35* (5), 913–937.

Yamaguchi, Shintaro, Yukiko Asai, and Ryo Kambayashi, "Effects of subsidized childcare on mothers' labor supply under a rationing mechanism," *Labour Economics*, 2018, *55*, 1–17.

_ , _ , **and** _ , "How does early childcare enrollment affect children, parents, and their interactions?," *Labour Economics*, 2018, *55*, 56–71.

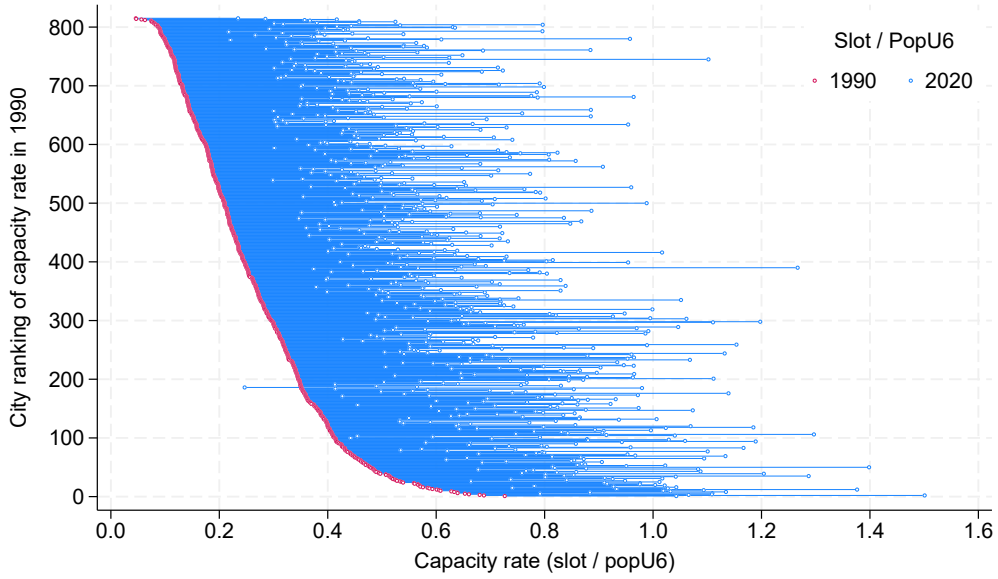


Figure 1: Changes in capacity rate (slot / popU6) from 1990 to 2020

Note: This figure shows the capacity rates of each city in 1990 and 2020. The capacity rate is defined by the number of childcare slots divided by the number of children aged 0–5 years.

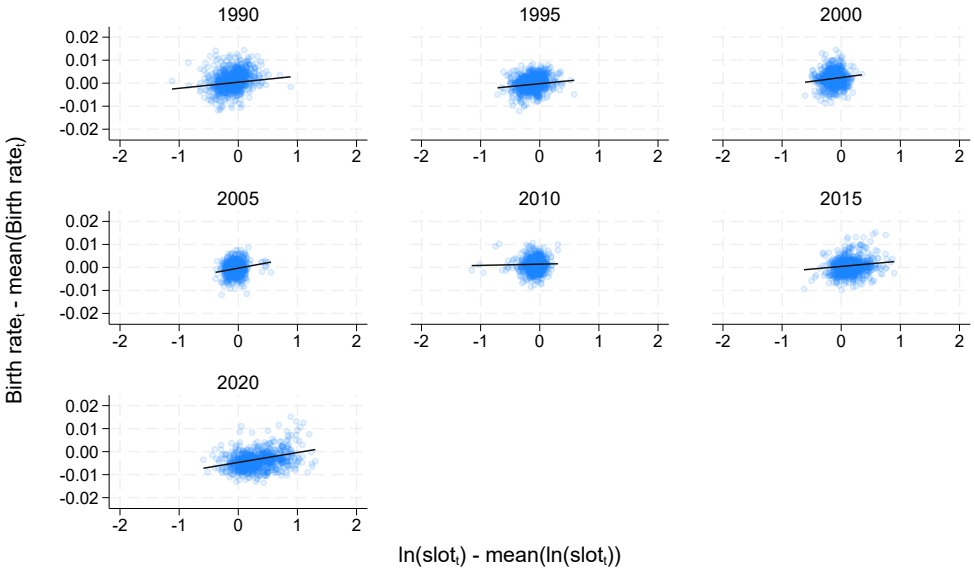
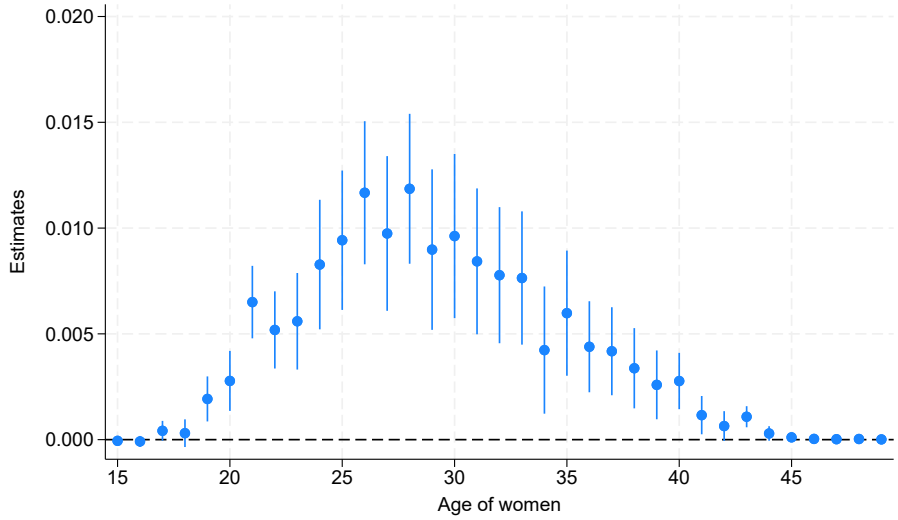
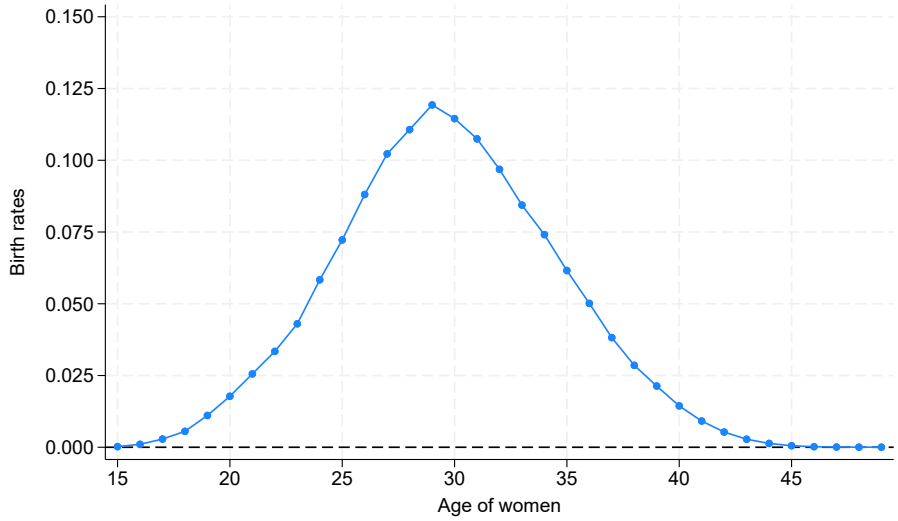


Figure 2: Correlation between birth rate and childcare slot

Note: This figure shows correlation between birth rate and childcare slot, where each variable is demeaned using city-specific mean to partial out city fixed effects.



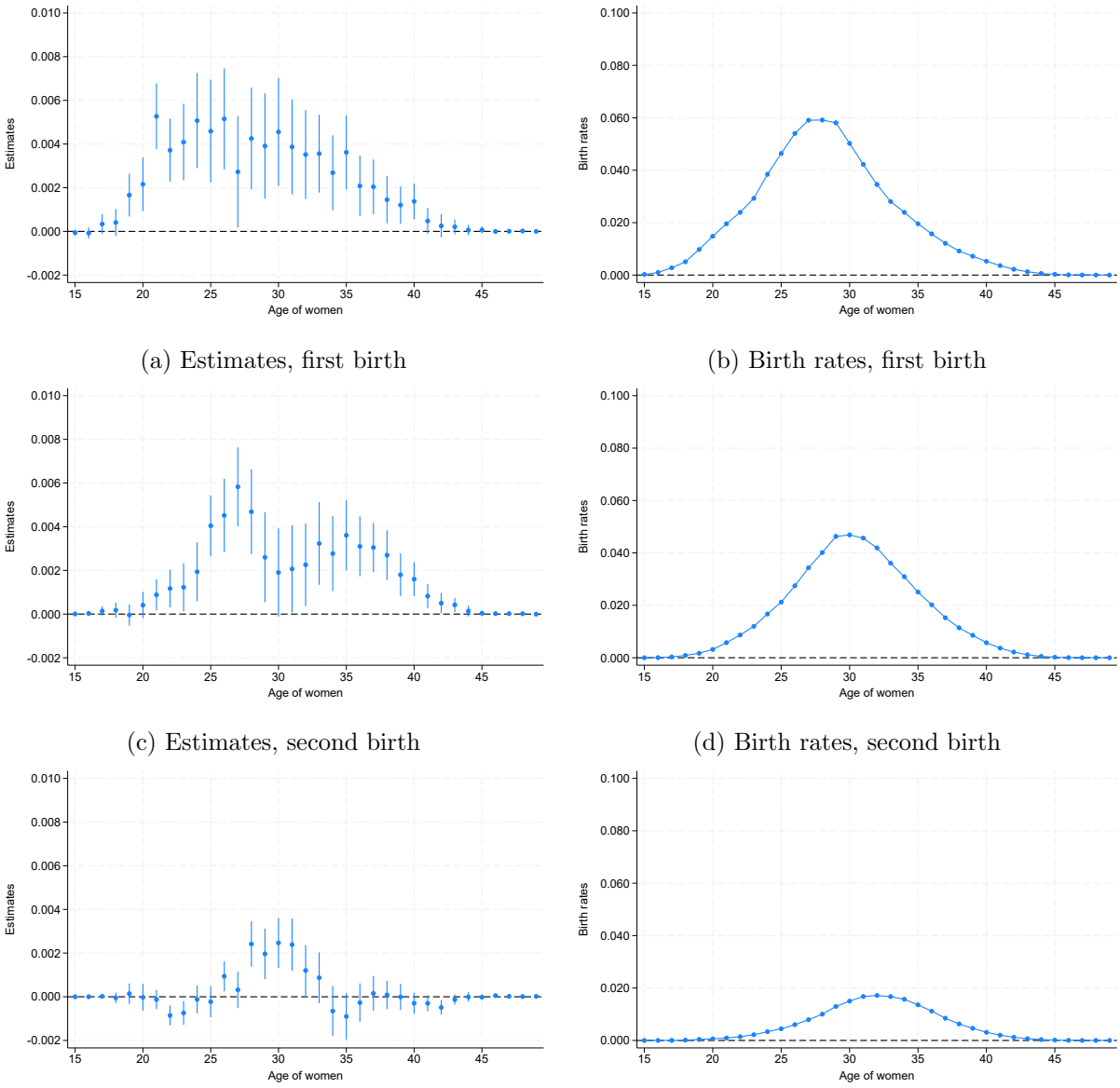
(a) Effects of childcare availability by age of mothers



(b) Birth rate by age of mothers

Figure 3: Effects of childcare availability on birth rate by age of mothers

Note: This figure shows estimation result of equation (1). Using birth rate at each age as the dependent variable, equation (1) is separately estimated. Points in Panel (a) show the estimates of β in equation (1), and blue bars show 95% confidence intervals.



(a) Estimates, first birth

(b) Birth rates, first birth

(c) Estimates, second birth

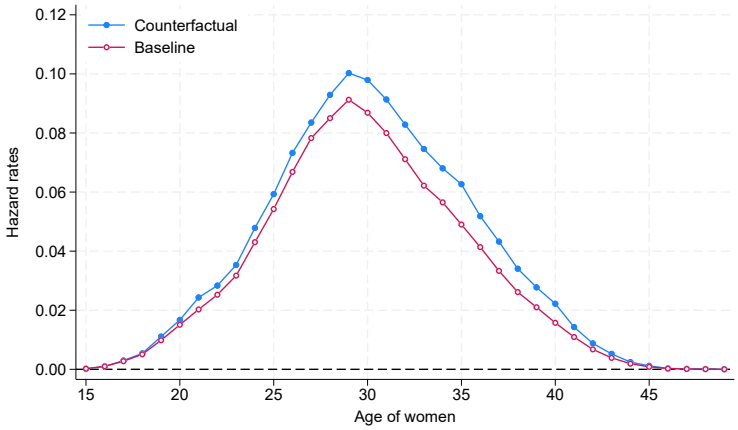
(d) Birth rates, second birth

(e) Estimates, third birth

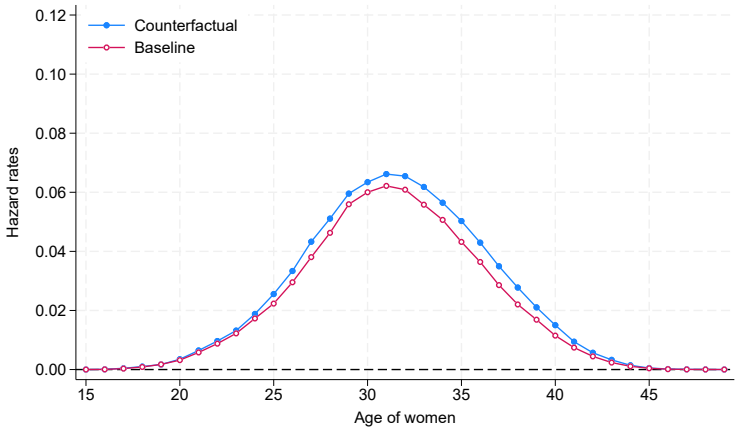
(f) Birth rates, third birth

Figure 4: Effects of childcare availability on birth rate by birth parity

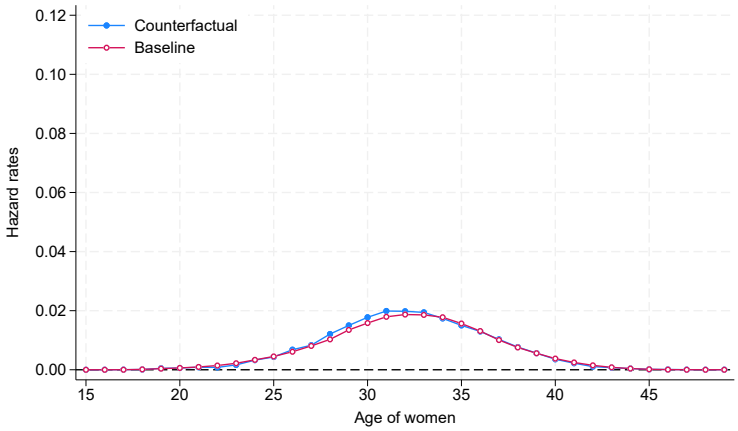
Note: This figure shows estimation result of equation (1). Using birth rate at each age as the dependent variable, equation (1) is separately estimated. Points in Panels (a), (c), and (e) show the estimates of β in equation (1), and blue bars show 95% confidence intervals.



(a) first birth



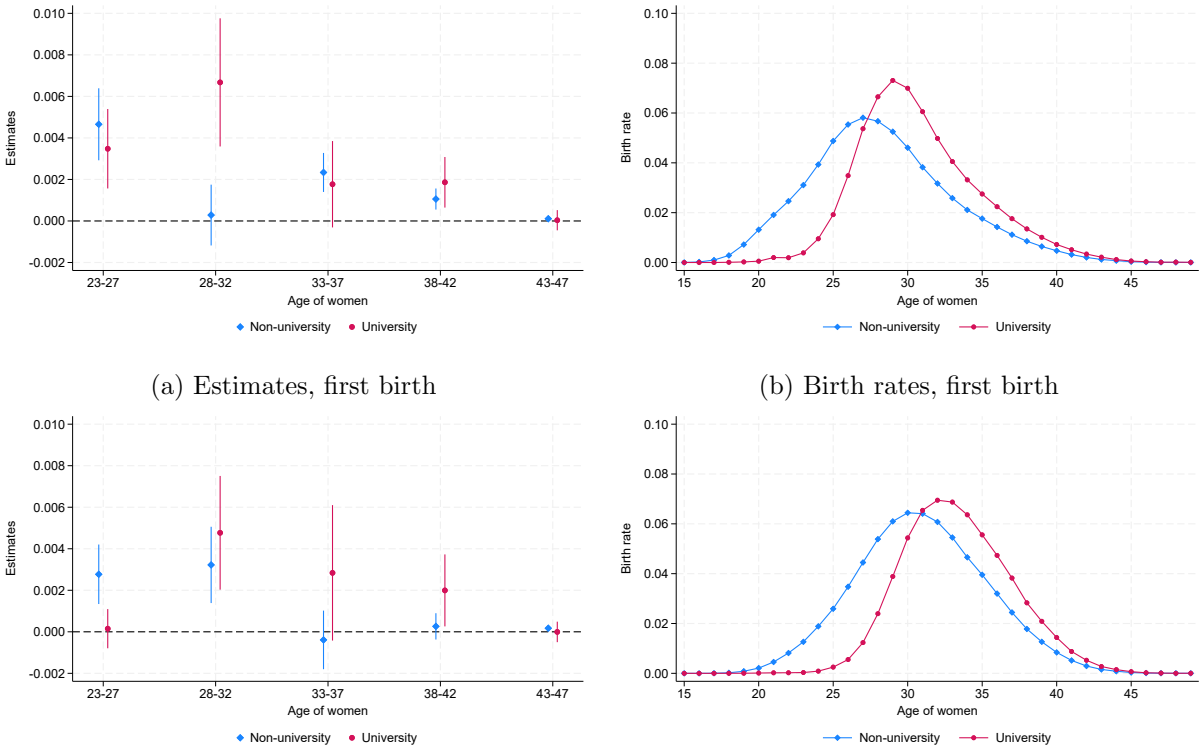
(b) second birth



(c) third birth

Figure 5: Effects of childcare availability on hazard rate of childbirth by birth parity

Note: This figure shows the hazard rate for each birth parity and at each age (red lines). The estimated policy effect on hazard rate is calculated by using estimates in Figure 4, and assuming policy scenario in which the childcare slots are increased by 70% (blue lines.



(a) Estimates, first birth

(b) Birth rates, first birth

(c) Estimates, second and more births

(d) Birth rates, second and more births

Figure 6: Effects of childcare availability on birth rate by education

Note: This figure shows estimation result of equation (1). Using birth rate at each age as the dependent variable, equation (1) is separately estimated by education level of women. Points in Panels (a) and (c) show the estimates of β in equation (1), and bars show 95% confidence intervals.

Table 1: Summary statistics

Year	1990	1995	2000	2005	2010	2015	2020
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Outcomes							
Birth rate (%)	3.94	3.87	4.13	3.88	4.09	4.05	3.70
	[0.51]	[0.49]	[0.54]	[0.48]	[0.48]	[0.47]	[0.47]
age 15–24 (%)	2.27	2.32	2.32	2.13	2.06	1.69	1.37
	[0.64]	[0.69]	[0.77]	[0.70]	[0.69]	[0.63]	[0.52]
age 25–34 (%)	11.68	10.41	9.51	8.40	9.23	9.76	9.21
	[1.36]	[1.45]	[1.41]	[1.22]	[1.24]	[1.19]	[1.25]
age 35–44 (%)	1.11	1.38	1.85	2.16	2.88	3.28	3.37
	[0.29]	[0.31]	[0.33]	[0.36]	[0.43]	[0.58]	[0.61]
age 45–49 (%)	0.01	0.01	0.01	0.02	0.02	0.03	0.04
	[0.01]	[0.01]	[0.01]	[0.02]	[0.02]	[0.03]	[0.03]
Birth weight (g)	3073.13	3073.05	3037.65	3013.03	3004.73	3003.97	3010.97
	[33.03]	[31.47]	[31.40]	[31.46]	[31.28]	[30.75]	[35.57]
B. weight < 2500 (%)	6.28	7.43	8.59	9.40	9.57	9.58	9.35
	[1.19]	[1.37]	[1.45]	[1.65]	[1.69]	[1.76]	[1.99]
Gestation week	39.02	38.94	38.87	38.83	38.75	38.72	38.68
	[0.15]	[0.14]	[0.14]	[0.13]	[0.15]	[0.14]	[0.16]
G. week < 37 (%)	4.55	4.95	5.46	5.69	5.79	5.64	5.65
	[1.15]	[1.17]	[1.27]	[1.38]	[1.39]	[1.37]	[1.73]
Panel B: Covariates							
Capacity rate (%)	21.46	23.05	23.59	26.93	28.16	38.88	52.91
	[9.72]	[10.39]	[10.56]	[11.41]	[11.63]	[13.12]	[14.62]
Potential capacity rate (%)	5.99	5.88	6.28	7.19	7.34	10.13	13.61
	[3.18]	[3.08]	[3,19]	[3.45]	[3.42]	[3.89]	[4.18]
Slot (1k)	2.09	2.04	2.06	2.24	2.21	2.90	3.61
	[2.86]	[2.80]	[2.87]	[3.18]	[3.24]	[4.68]	[6.00]
Pop (1k)	34.98	34.63	32.72	31.18	30.04	28.63	26.55
	[62.13]	[61.25]	[58.67]	[57.51]	[56.94]	[55.14]	[51.55]
Marriage rate (%)	59.04	56.15	53.50	51.48	50.77	50.46	50.02
	[4.77]	[4.87]	[4.93]	[4.61]	[4.12]	[3.32]	[3.08]
Age of mothers	28.36	28.51	28.84	29.55	30.35	30.98	31.23
	[0.52]	[0.54]	[0.59]	[0.69]	[0.75]	[0.75]	[0.68]
N. of cities	815	815	815	815	815	815	815

Table 2: Effects of childcare availability on birthrate using all age group of women

Region	Cities	Cities	All	Cities	Cities	Cities
Period	1990–2020	1990–2020	1990–2020	1990–2010	2010–2020	1990–2020
Cohort of women	1941–2005	1941–2005	1941–2005	1941–1995	1961–2005	1970–1980
	(1)	(2)	(3)	(4)	(5)	(6)
ln(slot)	0.0038	0.0024	0.0049	0.0023	0.0050	0.0027
	(0.0005)	(0.0005)	(0.0009)	(0.0005)	(0.0006)	(0.0007)
ln(pop women)	0.0024	0.0012	-0.0065	-0.0010	-0.0059	0.0031
	(0.0011)	(0.0013)	(0.0064)	(0.0013)	(0.0030)	(0.0033)
Mean of birth rate	0.0426	0.0426	0.0441	0.0427	0.0421	0.0474
Prefecture trend	X	X	X	X	X	X
Marriage rate		X				
N	199,640	199,640	412,804	142,590	85,575	57,854
N of municipalities	815	815	1,733	815	815	815

Note: This table shows estimation result of equation (1) using birth rate as the dependent variable. Standard errors clusters by municipality are reported in parentheses.

Table 3: Decomposition of the effects by birth parity

	Total	1st birth	2nd birth	3rd birth
	(1)	(2)	(3)	(4)
Effects on TFR	0.139	0.074	0.057	0.008
(%)		53.2	41.3	5.5

Note: This table shows estimation result of the effect of public childcare provision on the total fertility rate, and its decomposition into birth parity, using equation (2).

Table 4: Effects of childcare availability on birthrate using all age group of women using Census

Census timing	Quinquennial	Decennial	Decennial	Decennial
Period	1990–2020	1990–2020	1990–2020	1990–2020
Education	All	All	Low	High
	(1)	(2)	(3)	(4)
ln(slot)	0.0027	0.0028	0.0022	0.0039
	(0.0004)	(0.0005)	(0.0005)	(0.0007)
ln(pop women)	0.0049	0.0044	0.0042	0.0039
	(0.0010)	(0.0010)	(0.0013)	(0.0015)
Mean of birth rate	0.0359	0.0364	0.0478	0.0489
Prefecture trend	X	X	X	X
N	199,640	114,100	81,500	81,331
N of municipalities	815	815	815	815

Note: This table shows estimation result of equation (1) using birth rate as the dependent variable. Standard errors clusters by municipality are reported in parentheses.

Table 5: Robustness checks for policy endogeneity and selective migration

Robustness check	Lagged-IV	Migration adj.
Region	Cities	Cities
Period	1995–2020	1990–2020
	(1)	(2)
ln(slot)	0.0040	0.0076
	(0.0007)	(0.0008)
Outcome means	0.042	0.043
Prefecture trend	X	X
Method	IV	IV
Kleibergen-Paap F-stat.	377.8	28.4
N	171,080	114,100
N of municipalities	815	815

Note: This table shows estimation result of equation (1) using birth rate as the dependent variable. Standard errors clusters by municipality are reported in parentheses.

Table 6: Effect of childcare availability on birth weights

	ln(weight) (1)	weight<2500 (2)	weight<1500 (3)	weight<1000 (4)
ln(slot)	0.0015	-0.0007	-0.0004	-0.0002
	(0.0009)	(0.0012)	(0.0003)	(0.0002)
Outcome means	8.0039	0.0851	0.0068	0.0026
Prefecture trend	X	X	X	X
N	7,039,303	7,039,303	7,039,303	7,039,303
N of municipalities	815	815	815	815

Note: This table shows estimation result of equation (1) using birth weight as the dependent variable. The unit of observations in this analysis is individual-year. Standard errors clusters by municipality are reported in parentheses.

Table 7: Effect of childcare availability on gestation weeks

	ln(week) (1)	week<37 (2)	week<34 (3)
ln(slot)	0.0009	-0.0026	-0.0011
	(0.0003)	(0.0010)	(0.0004)
Outcome means	3.6587	0.0530	0.0120
Prefecture trend	X	X	X
N	7,039,303	7,039,303	7,039,303
N of municipalities	815	815	815

Note: This table shows estimation result of equation (1) using gestation weeks as the dependent variable. The unit of observations in this analysis is individual-year. Standard errors clusters by municipality are reported in parentheses.