¿La educación preescolar inclina la balanza a favor de las madres en el mercado laboral? Evidencia para Brasil.

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Does Preschool Tips the Balance in favour of Mothers in the Labour Market? Evidence for Brazil.[†]

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Abstract

This paper examines the impact of preschool enrolment on maternal labour market outcomes in Brazil, specifically focusing on the effect of preschool attendance of the youngest child and how this effect varies with the presence of other family members in the home. Using a fuzzy regression discontinuity design that exploits changes in preschool-entry age regulations, I find that enrolling the youngest child in preschool increases the probability of mothers' labour force participation by 40% and employment by 55% while also increasing family income by 70%. These positive effects are not observed for mothers who enrol children other than their youngest. Notably, the employment effects are more pronounced for mothers without other female relatives in the household, highlighting the role of informal childcare in alleviating maternal childcare responsibilities. The results suggest a positive effect on these other women's labour market decisions, though the effects are not statistically significant, possibly due to the reduced sample size. Additionally, the results indicate that fathers' labour market decisions remain unaffected by this policy.

JEL Classification: J13, J16, J22

Keywords: Gender inequality, preschool, regression discontinuity.

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1 Introduction

Women's increased participation in the labour market has contributed positively to narrowing gender gaps. However, their domestic and childcare responsibilities have largely remained unchanged. As recent evidence shows, motherhood has gained increasing importance over the last two decades, accounting for roughly 40% of the remaining gender earnings gap in Latin America (Marchionni & Pedrazzi, 2023). In this context, public policies that reduce this burden, such as the expansion of access to childcare at an early age, could not only encourage more mothers to enter the labour market but also improve their outcomes once employed. Indeed, Garcia et al. (2023) finds that the expansion of public childcare services in São Paulo, Brazil, resulted in a significant and lasting reduction in the motherhood effect: each additional seat per child led to a 20% increase in mothers' formal employment after the birth of their first child.

A substantial body of literature has investigated how maternal labour supply responds to increased childcare provision facilitated by preschool education. In developed countries, studies on the impact of preschool attendance on maternal labour market outcomes reveal overall positive effects, but primarily for mothers whose youngest child is affected by the policy. In the United States, Gelbach (2002) finds a 6% to 24% increase in maternal employment when their 5-year-old children are enrolled in school. Later on, Fitzpatrick (2010) finds a significant effect only for single mothers without additional young children. In Canada, enhanced childcare access in Quebec leads to a 7.7 percentage point increase in maternal employment (Baker et al., 2008; Lefebvre & Merrigan, 2008). Conversely, in Norway, Havnes & Mogstad (2011) observes a nearly null effect on maternal employment due to significant crowding out of informal childcare. In Europe, studies in France (Goux & Maurin, 2010), Germany (Bauernschuster & Schlotter, 2015), Hungary (Lovász & Szabó-Morvai, 2019), and Italy (Carta & Rizzica, 2018) consistently show increases in maternal employment, particularly among single mothers or those enrolling the youngest child at home, with employment gains ranging from 6 to 11.7 percentage points.

Less is known about the impact on developing countries, particularly in Latin America and the Caribbean (LAC). To the best of my knowledge, Berlinski & Galiani (2007) were the first to provide evidence from a Latin American country, showing that a large expansion of public preschools in Argentina boosts preschool attendance by approximately 7.5 percentage points. According to their results, this expansion led to an increase in maternal employment ranging from 7 to 14 percentage points, although the effect on the intensive

margin (i.e., hours worked) is not statistically significant. Later on, Berlinski et al. (2011), in a more closely related setting to this study, find that mothers with access to preschool are more likely to be employed and tend to work more hours. The availability of preschool care reduces the need for mothers to stay at home, thus enabling them to take on full-time or more stable jobs. Other evidence for Brazil assesses the effect of childcare on the labour market outcomes of mothers. For instance, Attanasio et al. (2022) shows that access to free daycare in Rio de Janeiro did not affect the mothers but notably increased the labour supply and income of grandparents (primarily grandmothers) living in the same household as the child attending daycare. Furthermore, Ryu (2019) finds that preschool enrolment significantly increased the time spent working rather than performing household chores among mothers living without additional younger children and other relatives.

This paper examines the causal impact children's preschool enrolment has on maternal labour market outcomes in Brazil by exploiting variation in preschool attendance induced by a school-entry age regulation and estimating a fuzzy regression discontinuity (FRD) model. In the past few years, Brazil has undergone several changes to provide more equal educational opportunities. Originally, compulsory education in Brazil started at age 7 and lasted 8 years. In 2006, a major reform lowered the minimum age for school entry from 7 to 6 years old and increased the duration of mandatory education from 8 to 9 years. In 2009, the compulsory school entry age was lowered even further to 4 years old, and the duration was extended to a total of 14 years. This last expansion, which includes preschool ages, provides a significant opportunity to evaluate the impact on mothers' labour force participation across the entire country and how this changes with the presence of other women in the household.

This paper aims at extending the literature in two key ways. First, I provide new causal evidence on the effects of preschool enrolment on mothers' labour market outcomes in a developing country, applying recent advances in regression discontinuity methods (Calonico et al., 2014). While Attanasio et al. (2022) focused on a lottery program for daycare access in Rio de Janeiro, my study broadens the scope to include a wider range of states in Brazil, providing robust evidence on the local causal impact of childcare policy in this context. Furthermore, my study builds on the work of Ryu (2019) by offering a more comprehensive analysis of preschool policies across different household structures and investigating how preschool attendance affects other co-residing relatives. Given the prevalence of informal childcare systems in Brazil, I explore whether formal childcare options, such as preschool, are particularly beneficial in households without other female relatives that can help alle-

viate the mother's childcare responsibilities. To address this, I compare households where the mother is the sole caregiver with households where an additional female aged 15 or older (e.g. a sister, aunt, or grandmother) resides. Finally, this study also examines the impact of preschool enrolment on fathers, as well as on other co-residing female relatives, to offer a comprehensive overview of how access to childcare influences the entire household.

I find that labour force participation increases by 40% and employment by 55% for mothers enrolling their youngest child in preschool and that this has an important impact on family income, which is 70% larger. In contrast, no effect is found when the child attending preschool is not the youngest. This is unsurprising, as having another younger child to care for may limit mothers' labour supply even after their eligible child is enrolled in preschool.

Moreover, for mothers who enrol their youngest child in preschool and do not live with other female relatives in the household, the effects on employment increase by 44%, and their weekly hours worked rise by an additional 20 hours. In contrast, there are no significant impacts on labour market outcomes for mothers who live with another female relative, indicating that the positive effects are primarily associated with the absence of support from other household members. While the results indicate qualitatively similar effects for these co-residing women, they are not statistically significant, likely due to a substantial loss of precision, which suggests the need for further research. This raises the question about the effect on fathers who do not appear to be affected, as no impact on their labour outcomes is found.

The rest of this paper is organized as follows. Section 2 details the data, the empirical approach, and validity checks. Section 3 presents and discusses the results and robustness checks related to mothers, while Section 4 focuses on the findings for other household members. Finally, Section 5 concludes with some final remarks.

2 Data and empirical strategy

2.1 Methodology

Estimations using conventional OLS to measure the relationship between mothers' labour market outcomes and attendance at preschool for their children would be biased and inconsistent, as mothers often simultaneously decide on their children's school enrolment and their participation in the labour force. For instance, more career-motivated women are more likely to choose to enrol their children in schools and participate in the work-

force. To sort out this problem, I exploit the cutoff date for preschool enrolment enforced by the Brazilian government as a source of exogenous variation in school attendance. In Brazil, children must turn 4 years old before the cutoff date of the school year to enrol in preschool. Those born one day after this cutoff must wait an additional year to be eligible for enrolment in that grade.

I use a fuzzy regression-discontinuity design (FRD) to assess the causal relationship between children's preschool enrolment and maternal labour market outcomes. Ideally, I would be interested in estimating the following equation:

$$Y_{ist} = \beta_0 + \beta_1 E_{\text{ist}} + f\left(X_{ist} - c_{st}\right) + v_{\text{ist}} \quad -h_n \leq \left(X_{ist} - c_{st}\right) \leq h_n, \tag{1}$$

where Y_{ist} represents the labour market outcomes of a woman i in state s in year t, E_{ist} is an indicator variable of the preschool enrolment of the 4-year-old child and $f\left(X_{ist} - c_{st}\right)$ is a polynomial function of the running variable that can vary on either side of the cutoff date. The running variable indicates the distance between the child's birthdate and the school cutoff date for state s in year t, c_{st}. $|h_n|$ controls the width of the neighbourhood around the cutoff that is used to fit the local polynomial approximation, where n indicates the bandwidth selection method used.

To address the endogeneity in the preschool attendance of children, I model the probability of mothers enrolling their children in preschool as follows:

$$E_{ist} = \gamma + \delta T_{ist} + f\left(X_{ist} - c_{st}\right) + \varepsilon_{ist} \quad -h_n \leq \left(X_{ist} - c_{st}\right) \leq h_n, \tag{2}$$

where E_{ist} is a binary variable indicating preschool enrolment of a child i in state s in year t, which takes the value 1 if the child is enrolled in preschool and 0 otherwise. T_{ist} is an indicator variable, which is equal to 1 for students who turn 4 years old in year t after the cutoff date corresponding to state s and are thus ineligible to start preschool and 0 for those who were born before the cutoff date and are eligible to start preschool. The primary coefficient of interest in equation (2) is δ, which captures the discontinuity in the probability of school enrolment at the cutoff. I anticipate that $\delta < 0$ since children born after the cutoff date are not eligible for preschool enrolment. In practice, the point estimate is likely between -1 and 0 since younger children may already be attending daycare or nurseries, and some older children may ignore compulsory attendance rules.

Once I estimate equation (2), I can consistently estimate equation (1) by replacing the enrolment rate with the estimated $\hat{E_{ist}}$. I am interested in β_1 from equation (1) that rep-

resents the Local Average Treatment Effect (LATE) of a child's enrolment in preschool on the maternal labour outcome. The underlying assumption for the validity of this strategy is that maternal outcome variables would be continuous if there was no discontinuity of school enrolment around the cutoff date. In the next sections, I conduct several tests to assess the validity of this assumption.

The most recent methods in regression discontinuity use local polynomials to approximate the regression function near the cutoff (Cattaneo & Titiunik, 2022). Therefore, I estimate both stages using a shared bandwidth chosen to minimize the mean squared error. Within this bandwidth, it is common practice to adopt a weighting scheme to ensure that the observations closer to the cutoff receive a higher weight than those further away. In line with standard practice in this literature, I use a triangular kernel function in all main specifications. Additionally, I follow Calonico et al. (2020) for robust inference.

2.2 Data

The analysis is based on Brazil's main national cross-sectional household surveys from 2001 to 2015. The *Pesquisa Nacional por Amostra de Domicílios* (PNAD) is managed by the Brazilian Institute of Geography and Statistics (IBGE) and was conducted annually until 2015, except for 2010 due to the census. The data from 2001-2015 allow me to know whether or not children go to preschool. However, starting in 2016, this question is only posed to children aged 5 and older; therefore, my analysis focuses on the period from 2001 to 2015. To ensure consistent definitions and adjustments over time, I follow the Socioeconomic Database for Latin America and the Caribbean (SEDLAC) protocol, a collaborative effort between CEDLAS at the Universidad Nacional de La Plata and the World Bank (CEDLAS and The World Bank, 2024).

From the PNAD, I am able to obtain social, labour, and demographic variables at the individual level. Specifically, I focus on five key labour market outcomes, namely: (i) participation in the labour market (coded as 1 if an individual is employed or looking for a job, and 0 otherwise), (ii) employment (coded as 1 if an individual is employed, and 0 otherwise), (iii) weekly working hours, (iv) log of hourly wages, and (v) log of per capita family income, where outcomes (iv) and (v) are expressed in real prices of 2011. In addition, the data provides every household member's exact date of birth as well as the mother's identifier for each child, which allows me to identify both the preschool-eligible child and her mother.

For my main results, I use a sample of women between the ages of 18 and 49 who have

at least one child around the age eligibility rule in the survey year. Therefore, the final sample includes mothers of children who were either 3 or 4 years old as of the cutoff date. Children who turned 4 years old by the cutoff date were eligible to start preschool in that survey year, while those who were only 3 years old had to wait another year before they could enrol in preschool. This setup enables a comparison between the mothers of children who could start preschool as soon as they turn 4 and the mothers of those who experienced a one-year delay in their entry due to being only 3 years old at the cutoff.

In Brazil, like in many other countries, a cutoff date is established for those who can enrol in a given academic year. The Law of Guidelines and Bases of National Education (*Lei de Diretrizes e Bases da Educação*) grants teaching and administrative autonomy to the different states and municipalities in Brazil. Before the 2009 reform, this implied significant heterogeneity in the admission criteria for preschool and primary education, as well as in how strictly these rules were enforced. After the reform and to address this variation between states, the Basic Education Chamber of the National Education Council attempted to standardize the cutoff date across states by establishing that students must be 4 years old by March 31 to enter preschool education.[1,2] Despite this, some states appealed the decision and adopted different cutoff dates.[3] Further efforts were made to enforce uniformity, such as the enactment of Law No. 12.796 in 2013, which required all states and municipalities to adopt the new system by 2016. Ultimately, in 2018, the Basic Education Chamber of the National Education Council mandated March 31 as the uniform cutoff date for all states, effectively ending the variation across the country.

To gather information about the cutoff date used in every state and over time, I obtained official resolutions from the Basic Education Chamber of the National Education Council and resolutions from Subnational Educational Ministries. Thereafter, cutoff dates were rescaled to unify enrolment rules and to generate a continuous variable indicating dates of birth relative to the cutoff date established by each state and period. During the period under analysis, 75% of states followed the March 31 cutoff date. For a detailed list of the thresholds applied in each state, see Table A.1 in the Appendix.

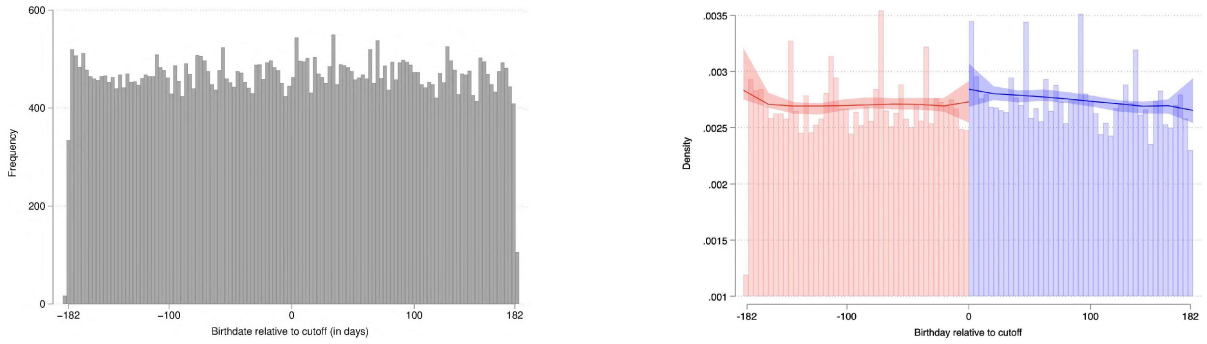
[1]See CNE/CEB N° 5/2009, N° 20/2009, N° 6/2010, 12/2010, N°17/2012.
[2]Other cutoff dates that have worked as threshold points are June 30 or the beginning of the school year.
[3]Rio de Janeiro's state has maintained the June 30 cutoff until 2018, for example. See Table A.1 for a detailed review of enrolment rules by state.

2.3 Validity of the empirical strategy

One potential concern about the validity of the fuzzy regression discontinuity design I adopt is the potential manipulation of the running variable, for example, if mothers can manipulate the birthdate of their children. To evaluate this, I first explore the distribution of birthdates using a histogram. Figure 1a shows the frequency of birthdates in the sample. Although there are some jumps, there is no evidence of any particular discontinuity in the distribution of birthdates around the cutoff dates.

In addition, I formally test for the presence of manipulation around the cutoff dates using a density test constructed using local polynomial density estimators proposed by Calonico et al. (2017). Figure 1b shows the density of birthdates around the cutoff for the distribution on either side of the cutoff. I fail to reject the null hypothesis of no discontinuity—the t-statistic is 0.73, and the corresponding p-value is 0.46—providing suggestive evidence that results are unlikely driven by sorting or systematic manipulation of birthdate around the cutoff point.

Figure 1. Histogram and estimated density of birthdates



(a) Histogram

(b) Estimated density

Notes: In (a) the bars represent the frequency of the birthdate relative to the cutoff dates in my sample, grouped in 3-day intervals. Figure (b) shows the manipulation tests based on density discontinuity proposed in Cattaneo et al. (2018).

As a second validation check, I assess the balance of several mothers' observable characteristics to understand whether those just above the cutoff are comparable to those just below it and discard the possibility that differences in these predetermined variables drive any observed effects. Table 1 presents the results from estimating discontinuities around the cutoff on predetermined characteristics of the mother's samples. Except for the years of education, all predetermined covariables are smooth around the threshold date, suggest-

ing that treated and control mothers do not differ systematically in these characteristics. In Section 3.3, I show that my results do not appear to be influenced by differences in education but rather by the discontinuity in enrolment generated by the eligibility criteria.

Table 1. Balance analysis of predetermined characteristics of mothers

Variables	MSE optimal bandwidth	RD estimator	Robust Inference		Observations Left,Right
			p-value	Confidence Interval	
Age	61.83	0.019	0.92	[-0.45 ; 0.50]	[9476, 9918]
Years of education	46.42	-0.496	0.00	[-0.89 ; -0.20]	[7071, 7538]
Child is male	61.02	-0.023	0.14	[-0.06 ; 0.01]	[9476, 9918]
No. children	72.53	0.067	0.17	[-0.03 ; 0.17]	[9423, 9853]
Married	50.68	0.005	0.91	[-0.04 ; 0.04]	[7673, 8181]
Urban	54.54	-0.025	0.09	[-0.06 ; 0.00]	[8332, 8842]
White	51.15	0.002	0.79	[-0.03 ; 0.05]	[7872, 8400]
Youngest child	60.92	0.011	0.41	[-0.02 ; 0.05]	[9333, 9778]
Other female (+15) at home	72.20	0.004	0.67	[-0.02 ; 0.04]	[11190, 11750]

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. Bandwidth is the MSE optimal based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff.

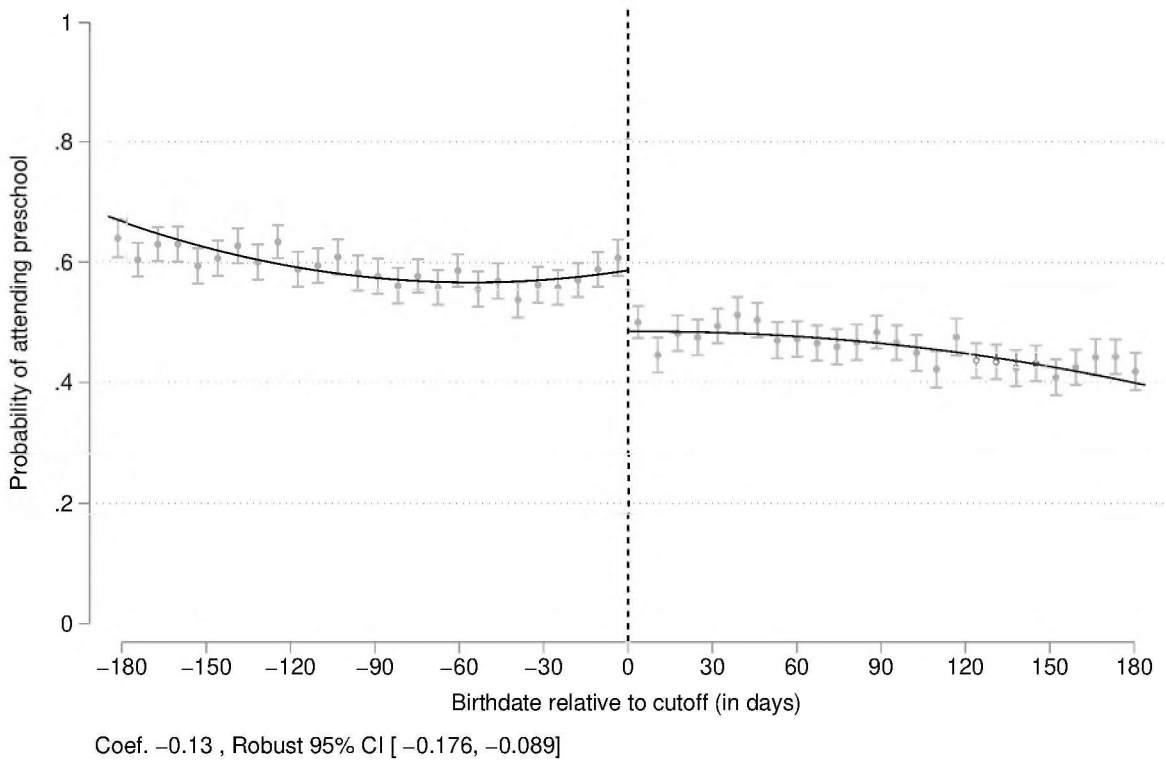
3 Main results

3.1 Effect of the eligibility rule on preschool enrolment

To examine the relevance of the eligibility rule, Figure 2 illustrates the discontinuity in the probability of preschool enrolment for eligible children, as presented in equation (2). Specifically, a child who turned 4 years old before the cutoff date is 13 percentage points more likely to be enrolled in preschool than a child who did not meet the minimum age requirement. Given that the average preschool enrolment rate for non-eligible children between 2001 and 2015 was approximately 48%, these estimates suggest an increase in enrolment rates of about 27%. Figure A.4 in the Appendix shows similar results when splitting the sample of mothers according to whether the child turning 4 years old around the cutoff is the mother's youngest child or not. The significant discontinuity in both cases suggests that compliance with the enrolment rule does not change with respect to the birth order of the children. In other words, what matters for increasing enrolment is adherence to the age rule, rather than the parity order of the children.

Additionally, similar to the weak instruments problem in the IV literature, if the eligibility criteria has a non-zero but very small effect on the probability of being enrolled in preschool, the estimates become unreliable and statistical inferences turn invalid when examining the Local Average Treatment Effect (LATE). I address this concern using tests of weak instruments, and all F-statistics exceed the standard values (see Table A.2).

Figure 2. Effect of preschool eligibility on the probability of enrolment



Notes: Figure shows second-order polynomial approximation using a triangular kernel and 95% confidence intervals. The dots in the scatterplots represent the average value of school enrolment rates in 7-day birthdate bins. This figure is based on the total sample of mothers aged 18 to 49.

3.2 Preschool effects on mothers

In this subsection, I present and discuss the estimated effects of preschool on mothers' labour market outcomes. First, Table 2 shows the reduced form results to analyse the effect the eligibility rule has on mothers' labour outcomes. Panel A refers to the sample of mothers whose youngest child is between 3 and 4 years old relative to the cutoff, while

Panel B refers to mothers whose child in this age range is not the youngest.

Columns (1) to (6) of Table 2 present the effects of the eligibility rule on maternal outcomes. Panel A shows that for mothers whose youngest child is just above the cutoff, the eligibility rule significantly influences their labour market decisions. Specifically, mothers of children below the age requirement are 6.6 percentage points less likely to participate in the labour force and 6.8 percentage points less likely to be employed. Examining weekly hours worked, column (3) indicates that ineligible mothers work 2.5 fewer hours than those eligible. However, when focusing on hours worked among already employed mothers in column (4), no significant effect on hours is observed, suggesting that the entire effect on hours in column (3) stems from increased employment in column (2). Additionally, no statistically significant difference is observed in wages earned among employed mothers.

Finally, households with eligible children report a per capita family income nearly 10% higher than those with ineligible children. This result raises important questions about household dynamics: Is this increase primarily due to maternal employment, or do other household members also enter the labour market because the children are now in preschool (e.g., grandmothers)? Conversely, are there household members who reduce their work hours because the mothers are now working (e.g., fathers)? Moreover, the positive effect on family income prompts further inquiries regarding the potential reduction of poverty associated with this policy. I plan to explore the household dynamics in the following sections.

In contrast to the effect found on the youngest child's mothers, Panel B confirms that, despite the significant discontinuity in preschool enrolment rates, the eligibility rule has no significant effect on labour market outcomes for mothers whose child is not the youngest.

Table 2. Effects of preschool eligibility on mothers' labour market outcomes

	Intention to Treat					
	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage (cond)	Family income
	(1)	(2)	(3)	(4)	(5)	(6)
PANEL A: Youngest child						
RD Estimate	-0.066***	-0.068***	-2.496***	-0.737	-0.040	-0.099**
	(0.024)	(0.024)	(0.938)	(0.728)	(0.043)	(0.039)
Observations	[3523;3799]	[3825;4171]	[4764;5161]	[3989;4178]	[3689;3846]	[5768;6181]
Mean	0.67	0.58	20.91	35.82	1.37	5.78
Robust 95% CI	[-.129 ; -.024]	[-.132 ; -.024]	[-4.952 ; -.697]	[-2.453 ; .975]	[-.146 ; .055]	[-.196 ; -.013]
Robust p-value	0.00	0.00	0.01	0.40	0.37	0.03
Bandwidth (h)	33.69	36.74	45.65	64.71	69.06	55.67
PANEL B: Non-youngest child						
RD Estimate	-0.011	-0.000	-0.176	-0.477	-0.123	-0.040
	(0.030)	(0.028)	(1.092)	(1.333)	(0.112)	(0.056)
Observations	[2685;2780]	[2916;2987]	[2864;2933]	[1412;1484]	[756;776]	[3160;3228]
Mean	0.51	0.42	13.41	31.75	1.12	5.10
Robust 95% CI	[-.085 ; .053]	[-.073 ; .059]	[-3.015 ; 1.961]	[-3.934 ; 2.162]	[-.408 ; .114]	[-.184 ; .076]
Robust p-value	0.65	0.84	0.68	0.57	0.27	0.42
Bandwidth (h)	55.61	59.43	58.17	67.29	48.02	64.30

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff. Conventional standard error of local-polynomial RD estimator in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

I now turn to the effect on the subgroup of mothers who are induced to enrol their children after the eligibility rule to learn about the effect of children attending preschool on mothers' labour decisions. Table 3 rescales the intent-to-treat effect by dividing it by the estimate obtained in the first stage, shown in column (1). Under the monotonicity assumption, this is known as the Local Average Treatment Effect (LATE). Once again, statistically significant impacts are present only for the group of mothers enrolling their youngest child. Column (2) shows that mothers who are influenced by the cutoff date to enrol their children in preschool are 27 percentage points more likely to be employed or looking for a job. In addition, columns (3) and (4) show that mothers are 33 percentage points more likely to be working and work 15 more hours per week. However, I do not find any increase in hours for mothers who are already employed (column 5), suggesting that all the effect on hours worked seems to be driven by mothers who enter the labour market rather than mothers who were already working.

Additionally, there appears to be no difference in hourly wages between mothers who enrol their children and those who do not. This result implies that the effect of children's preschool enrolment has only an effect on the extensive margin of labour participation

and employment.[4] Households where mothers enrol their children also experience a 70% increase in per capita family income, likely due to a household member entering the labour force, which has positive implications for poverty reduction. The conclusions also hold when including controls for year and state (see Table A.3).

The results suggest that the possibility of enrolling children in preschool does not lead to women securing jobs with longer working hours or better pay; rather, it facilitates their entry into the labour market. These effects likely stem from the increased free time, which allows women greater opportunities to engage in employment without significantly affecting the nature or quality of the jobs they obtain.

Table 3. Effects of preschool enrolment on mothers' labour market outcomes

	First Stage	Local Average Treatment Effect					
	Enrolment	Participation	Employment	Hours worked (cond)	Hours worked (cond)	Hourly wage (cond)	Family income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: Youngest child							
RD Estimate	-0.140***	0.271**	0.334**	15.410**	7.648	0.393	0.707**
	(0.023)	(0.131)	(0.140)	(6.023)	(7.519)	(0.447)	(0.277)
Observations	[4142;4500]	[6365;6733]	[6049;6434]	[6049;6434]	[3571;3719]	[3331;3457]	[5667;6067]
Mean	0.49	0.67	0.58	20.91	35.82	1.37	5.78
Robust 95% CI	[-.202 ; -.097]	[-.004 ; .592]	[.042 ; .694]	[1.95 ; 29.874]	[-9.813 ; 24.052]	[-.624 ; 1.471]	[.115 ; 1.321]
Robust p-value	0.00	0.05	0.03	0.03	0.41	0.43	0.02
Bandwidth (h)	39.30	60.22	57.21	57.41	57.86	62.33	54.01
PANEL B: Non-youngest child							
RD Estimate	-0.090***	0.113	0.002	2.925	2.811	0.749	0.460
	(0.027)	(0.306)	(0.313)	(11.406)	(10.521)	(0.756)	(0.617)
Observations	[3172;3237]	[3172;3236]	[2864;2933]	[3420;3481]	[1257;1316]	[1184;1203]	[2961;3031]
Mean	0.38	0.67	0.58	20.91	35.82	1.37	5.78
Robust 95% CI	[-.157 ; -.03]	[-.537 ; .91]	[-.678 ; .803]	[-18.76 ; 33.709]	[-20.861 ; 28.685]	[-.845 ; 2.717]	[-.905 ; 2.006]
Robust p-value	0.00	0.61	0.87	0.58	0.76	0.30	0.46
Bandwidth (h)	64.60	64.42	58.66	69.16	59.33	74.83	60.63

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

3.3 Is education driving the results?

As years of education seem to present some discontinuity around the cutoff, I want to rule out that changes in mothers' educational attainment are driving the effects. Since women with more years of education face better labour market conditions, this could be behind the observed impacts. While one could consider controlling for education, this strategy is not recommended as does not allow to identify the treatment effect (Cattaneo et al., 2023).

[4] I find no effect on the likelihood of having an informal job, conditional on being employed. Results are available upon request.

Therefore, to assess that schooling is not driving the effects, I estimate equations (1) and (2) for a subsample of mothers for whom all characteristics are balanced around the cutoff.

After further inspecting the data, it appears that the imbalance in years of education occurs only in the years 2004, 2005, and 2007 (the results of the balance analysis by year are presented in Figure A.2 in the Appendix). Therefore, I exclude these years and re-estimate equations (1) and (2). Table 4 presents the results for compliant mothers. The effects on the probability of being employed and weekly hours worked are positive and significant at conventional levels, similar to those found for the entire sample. Additionally, the effects observed for labour force participation and family income are also in the same direction, although they are not statistically significant at usual levels. The fact that the effects found on this subsample are similar to those in the entire sample suggests that the labour market impact on mothers is not driven by differences in education.[5]

Table 4. Effects of preschool enrolment on mothers' labour market outcomes for the subsample with balanced characteristics

	First Stage	Local Average Treatment Effect					
	Attendance	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage (cond)	Family income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: Youngest child							
RD Estimate	-0.129***	0.213	0.295*	14.274*	5.866	-0.121	0.548
	(0.024)	(0.155)	(0.167)	(7.300)	(8.592)	(0.538)	(0.333)
Observations	[4010;4313]	[5714;6017]	[5158;5384]	[4918;5129]	[3214;3375]	[2671;2757]	[4961;5179]
Mean	0.49	0.67	0.58	20.91	35.82	1.37	5.78
Robust 95% CI	[-.189 ; -.079]	[-.115 ; .601]	[-.037 ; .755]	[-1.564 ; 32.944]	[-14.024 ; 24.858]	[-1.347 ; 1.135]	[-.195 ; 1.312]
Robust p-value	0.00	0.15	0.08	0.07	0.58	0.87	0.15
Bandwidth (h)	48.01	68.19	61.61	58.44	65.20	62.66	59.16
PANEL B: Non-youngest child							
RD Estimate	-0.090***	0.368	0.263	6.799	1.633	0.073	0.209
	(0.031)	(0.365)	(0.378)	(13.031)	(12.460)	(0.965)	(0.682)
Observations	[2475;2574]	[2217;2299]	[2010;2098]	[2321;2404]	[1177;1232]	[936;952]	[2434;2528]
Mean	0.38	0.51	0.42	13.41	31.75	1.12	5.10
Robust 95% CI	[-.166 ; -.022]	[-.406 ; 1.315]	[-.521 ; 1.247]	[-18.3 ; 41.635]	[-25.949 ; 32.54]	[-1.992 ; 2.558]	[-1.357 ; 1.875]
Robust p-value	0.01	0.30	0.42	0.45	0.83	0.81	0.75
Bandwidth (h)	66.87	59.38	54.29	62.22	73.83	75.13	65.38

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

[5]I conducted a second exercise focusing on the subperiod from 2011 to 2015, with the hypothesis that compulsory preschool would balance the sample. While the results are qualitatively similar, I do not find statistically significant effects, likely due to the reduction in sample size.

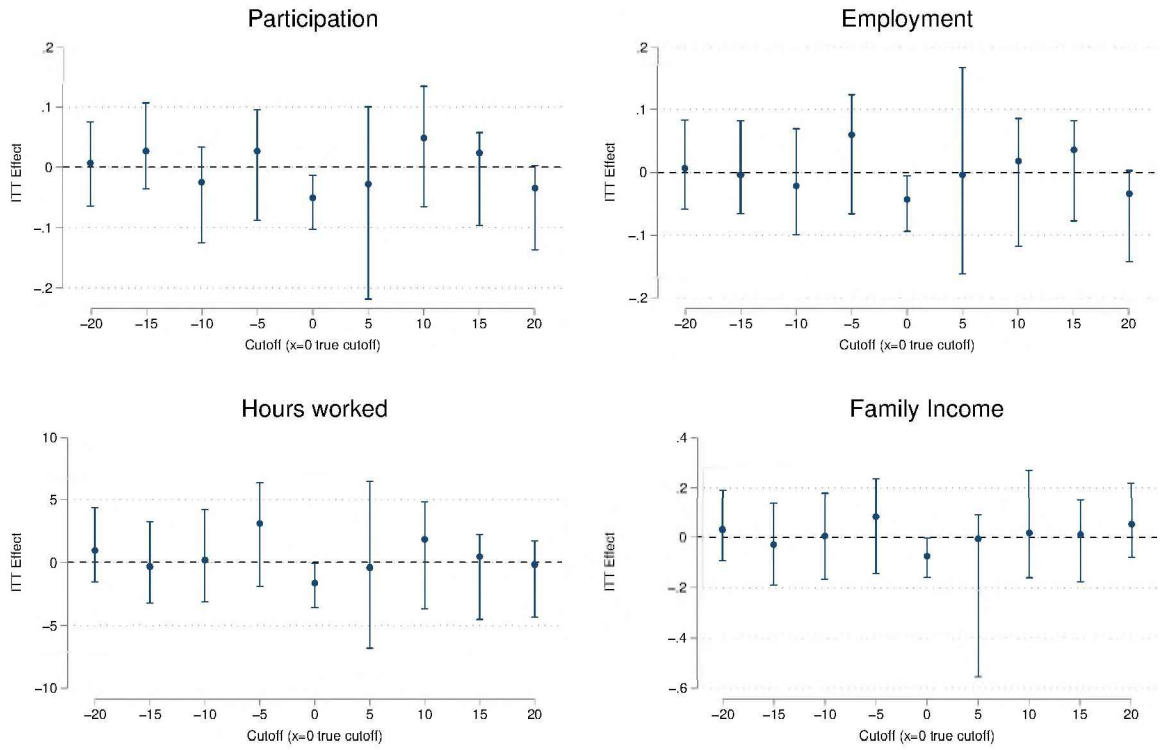
3.4 Robustness

This subsection reports a series of robustness checks to increase the confidence in the results obtained using the fuzzy regression discontinuity approach presented so far.

First, one useful robustness exercise is to examine the treatment effects at artificial or placebo cutoff values. This test replaces the true cutoff value with another value at which the treatment status does not change, such as 5 days, 10 days, or 15 days before or after the actual cutoff. The expectation is that no significant treatment effect will occur at placebo cutoff values.

The formal implementation of this test is depicted in Figure 3, which presents the estimates for each outcome derived from various placebo cutoffs. The figure clearly shows that the estimate corresponding to the true cutoff date differs from those obtained using false cutoff dates. Except for the estimate at the true cutoff, all other regression discontinuity estimates are not statistically different from zero. Consequently, the absence of discontinuous jumps at the artificial cutoffs evaluated reinforces the validity of my main results.

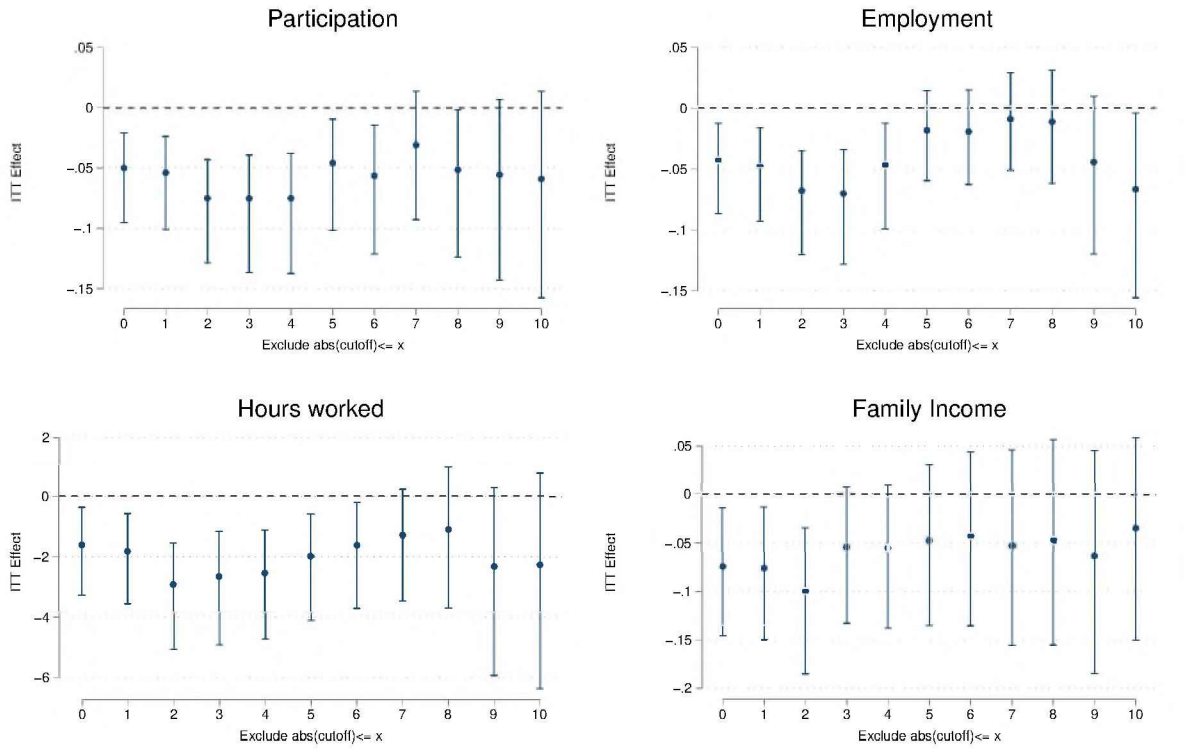
Figure 3. Estimates from artificial cutoff dates



Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. For artificial cutoffs above the real cutoff, I only use treated observations, and for artificial cutoffs below the real cutoff, I only use control observations. Bandwidth is the MSE optimal based on Calonico et al. (2014).

Another robustness check consists of evaluating the sensitivity of results to the exclusion of units located very close to the cutoff. This strategy, commonly referred to as the "Donut Hole" method, involves estimating the unknown regression function while excluding observations within a specific narrow range around the threshold. This approach is also valuable for evaluating how sensitive the results are to the inherent extrapolation required in local polynomial estimation. Figure 4 reports the estimates resulting from gradually excluding observations around the cutoff (the x-axis indicates the number of days excluded). Even after excluding mothers of children with birthdates within one week on either side of the cutoff, the estimated effects continue to be significant at the 10% level, reinforcing the robustness of the results.

Figure 4. Sensitivity to the exclusion of observations around cutoff date



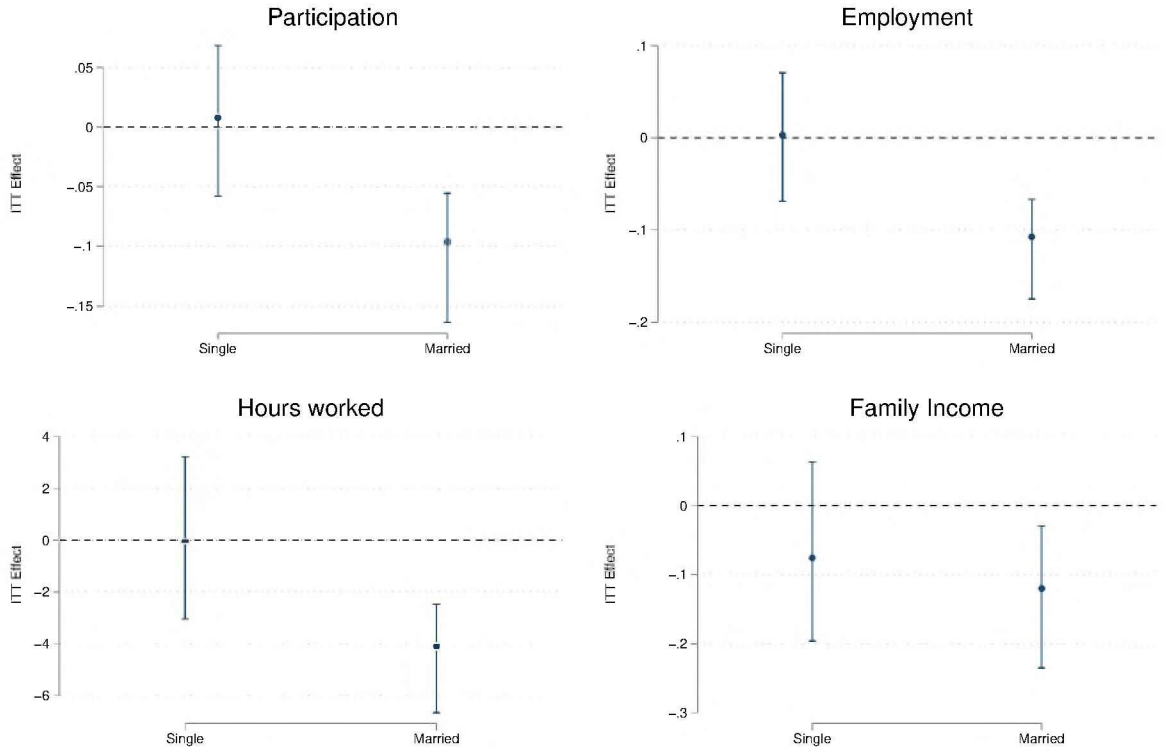
Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. Bandwidth is the MSE optimal based on Calonico et al. (2014).

3.5 Single versus married mothers

Building on previous evidence from developed countries that highlights the role of marital status in mothers' labour market decisions (Carta & Rizzica, 2018; Fitzpatrick, 2010), I examine how my main findings differ between single and married mothers using the pooled sample of mothers with the youngest child affected by the policy. I explore how the impact of the minimum age requirement varies by marital status by estimating the reduced form for single and married women separately. For instance, married mothers may have more financial and caregiving support from their partners, which can alleviate the burden of child care. In this case, although children's preschool attendance may provide some relief, this may not be as crucial for married mothers as for single mothers, who are often solely responsible for both child care and financial support. Conversely, single mothers may face

a greater need to participate in the labour market regardless of whether their children are in preschool, as they cannot afford to be out of the labour market. This is precisely what Figure 5 suggests. For married mothers, labour market participation and employment are significantly affected by the minimum age requirement, with ineligible mothers less likely to participate. In contrast, for single mothers, the estimates are both virtually null and not statistically significant. This pattern suggests that the observed effects primarily stem from married mothers, as single mothers are probably compelled to participate in the labour market regardless of their children's preschool eligibility.[6]

Figure 5. Effects of preschool eligibility on mothers' labour market outcomes by marital status



Notes: Figure presents the point estimates and 90% confidence intervals from regressing the outcome on birthdate relative to the cutoff date. Each estimate corresponds to a different regression. The estimates are based on a local linear polynomial with a triangular kernel and robust inference. In each estimation, the bandwidth is established as the optimal MSE according to Calonico et al. (2014). Married refers to women in formal and consensual unions.

[6]Results exploring differential effect by educational level or geographic area are found in Figure A.3, although estimations are not precise for all of the groups.

4 The role of and effects on other household members

4.1 Effect of having additional women in the household

As mentioned before, in settings with high levels of informal childcare, the preschool enrolment of the youngest child may free up time not only for mothers of young children but also for other relatives who might have been previously taking care of children in the household, such as grandmothers (Evans et al., 2017; Pinto, 2023). If this is the case, then one might expect the effect on the labour outcomes of mothers to be larger in households where there are no other female relatives who can provide childcare. Because of that, I explore the potential heterogeneity in the effect depending on the presence or absence of other 15-year-old or older female relatives in the household.

Results presented in Table 5 show strong and significant employment effects only for mothers of youngest children with no other women at home.[7] Turning to the local effect on mothers who are induced to enrol their youngest child, the point estimates are larger than the findings of the previous section. More precisely, results from Table 6 indicate that mothers without female relatives in the household, who now have additional free time as their children attend preschool, are 37 percentage points more likely to participate in the labour force. The probability of being employed increases by 53 percentage points, which is also reflected in approximately 23 hours of increase in weekly hours worked on average. Consistent with previous findings, there seems to be no effect on the intensive margin (i.e. hours worked for employed mothers and hourly wage).

On the contrary, the results reveal no significant impact on labour market outcomes for the subsample of mothers living with another female relative, despite the 15 percentage point discontinuity in children's preschool enrolment (Panel B in Tables 5 and 6). These results suggest that the findings from the previous section are primarily driven by the impact on mothers who lack support from other women in the household. The absence of significant effects for mothers raises important questions about the dynamics within these households, particularly regarding whether changes are occurring for other women in the household.

[7] Even in the absence of other female relative residing in the household, mothers may receive support from non-cohabiting relatives, such as a grandmother who does not live in the same home. However, these findings suggest that the critical factor is the presence of someone within the household who can provide assistance.

Table 5. Effects of preschool eligibility on mothers' labour market outcomes by presence of other women (15+) in the household

	Intention to Treat					
	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage	Family income
	(1)	(2)	(3)	(4)	(5)	(6)
PANEL A: No other female (+15) at home						
RD Estimate	-0.081***	-0.100***	-3.961***	-0.796	-0.059	-0.106**
	(0.027)	(0.027)	(1.105)	(0.823)	(0.058)	(0.046)
Observations	[2805;3035]	[3063;3262]	[3378;3595]	[3261;3339]	[2167;2208]	[4291;4534]
Mean	0.66	0.58	20.23	35.06	1.42	5.77
Robust 95% CI	[-.153 ; -.032]	[-.171 ; -.05]	[-6.866 ; -1.911]	[-2.879 ; .944]	[-.199 ; .074]	[-.222 ; -.004]
Robust p-value	0.00	0.00	0.00	0.32	0.37	0.04
Bandwidth (h)	35.10	39.00	42.57	70.29	54.25	54.69
PANEL B: Other female (+15) at home						
RD Estimate	0.020	0.042	1.186	-0.273	0.026	-0.084
	(0.032)	(0.040)	(1.670)	(1.530)	(0.094)	(0.065)
Observations	[1867;2058]	[1439;1597]	[1605;1764]	[772;883]	[629;735]	[1820;1997]
Mean	0.71	0.60	22.89	37.92	1.26	5.78
Robust 95% CI	[-.05 ; .102]	[-.053 ; .134]	[-2.393 ; 5.404]	[-3.341 ; 3.637]	[-.198 ; .241]	[-.24 ; .067]
Robust p-value	0.51	0.40	0.45	0.93	0.85	0.27
Bandwidth (h)	72.02	56.34	62.60	51.46	47.45	70.96

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff. Conventional standard error of local-polynomial RD estimator in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6. Effects of preschool enrolment on mothers' labour market outcomes by presence of other women in the household

	First Stage	Local Average Treatment Effect					
	Attendance	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage	Family income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: No other female (+15) at home							
RD Estimate	-0.143***	0.377**	0.532***	22.694***	9.128	0.595	0.800**
	(0.026)	(0.157)	(0.171)	(7.190)	(9.249)	(0.565)	(0.331)
Observations	[3298;3517]	[5405;5675]	[5747;6046]	[5651;5960]	[3354;3450]	[2363;2394]	[4935;5151]
Mean	0.49	0.66	0.58	20.23	35.06	1.42	5.77
Robust 95% CI	[-.211 ; -.093]	[.056 ; .793]	[.176 ; .981]	[7.885 ; 41.765]	[-11.243 ; 31.925]	[-.695 ; 1.97]	[.104 ; 1.598]
Robust p-value	0.00	0.02	0.00	0.00	0.35	0.35	0.03
Bandwidth (h)	41.77	68.59	72.16	71.48	72.30	59.32	62.81
PANEL B: Other female (+15) at home							
RD Estimate	-0.150***	0.003	-0.256	-10.143	-5.929	-0.246	0.393
	(0.043)	(0.255)	(0.298)	(13.719)	(16.257)	(0.890)	(0.525)
Observations	[1173;1331]	[1273;1429]	[1221;1377]	[1173;1331]	[620;721]	[583;691]	[1170;1323]
Mean	0.51	0.71	0.60	22.89	37.92	1.26	5.78
Robust 95% CI	[-.263 ; -.073]	[-.531 ; .634]	[-.885 ; .457]	[-40.502 ; 20.506]	[-47.742 ; 25.59]	[-2.257 ; 1.887]	[-.822 ; 1.496]
Robust p-value	0.00	0.86	0.53	0.52	0.55	0.86	0.57
Bandwidth (h)	46.80	50.88	48.90	46.64	41.10	44.73	46.60

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

4.2 Effects on the outcomes of other women in the household

In this subsection, I address the question raised in the previous subsection and present the estimated effects on labour market outcomes for cohabiting female household members aged 15 and above (Table 7). The results suggest a positive effect on labour force participation and employment for these women, albeit statistically insignificant. The lack of statistical significance at conventional levels may be due to a substantial reduction in sample size, as only 25% of mothers in the sample live with another female relative, limiting the precision of the estimates. Nevertheless, the point estimates remain comparable in magnitude to those observed for mothers. Further research is needed to better understand the effects on these co-residing women.

Table 7. Effects of preschool enrolment on mothers' labour market outcomes by presence of other women in the household

	First Stage	Local Average Treatment Effect					
	Attendance	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage	Family income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: Youngest child							
RD Estimate	-0.129***	0.443	0.489	15.256	-13.478	2.249	0.108
	(0.036)	(0.298)	(0.298)	(11.646)	(31.925)	(2.905)	(0.572)
Observations	[1715;1965]	[1827;2080]	[1827;2080]	[1827;2080]	[802;843]	[775;836]	[1587;1843]
Mean	0.49	0.49	0.43	15.46	36.14	1.12	5.75
Robust 95% CI	[-.223 ; -.065]	[-.303 ; 1.062]	[-.274 ; 1.093]	[-14.35 ; 38.915]	[-84.504 ; 59.017]	[-4.798 ; 8.977]	[-1.271 ; 1.287]
Robust p-value	0.00	0.28	0.24	0.37	0.73	0.55	0.99
Bandwidth (h)	50.34	53.31	53.60	53.47	52.69	59.72	47.47
PANEL B: Non-youngest child							
RD Estimate	-0.137***	-0.193	-0.126	-0.407	9.122	0.466	0.254
	(0.049)	(0.290)	(0.281)	(11.536)	(13.666)	(1.296)	(0.462)
Observations	[982;1005]	[1517;1565]	[1545;1585]	[1571;1613]	[570;607]	[340;355]	[1285;1316]
Mean	0.38	0.53	0.44	16.00	36.46	0.82	5.13
Robust 95% CI	[-.259 ; -.029]	[-.852 ; .523]	[-.825 ; .508]	[-27.467 ; 27.293]	[-20.606 ; 44.772]	[-2.634 ; 3.455]	[-.869 ; 1.307]
Robust p-value	0.01	0.64	0.64	1.00	0.47	0.79	0.69
Bandwidth (h)	47.84	72.56	73.38	74.31	67.26	49.21	62.18

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

4.3 What about fathers?

Lastly, I present the results concerning labour market outcomes for fathers. For this analysis, I restrict the sample to those households where the mother of the child is either the household head or the household head's spouse, and thus, the household head or spouse is expected to be the child's father.[8]

Overall, Table 8 shows that the estimated effects for fathers are not statistically significant. In contrast to the substantial changes observed in mothers, who experienced nearly a 40% increase in participation relative to the mean outcome (see Table 3), the impacts on fathers' labour force participation, employment, hours worked, or hourly wage are not significant. This is expected, given the already high employment rate among fathers.

These results suggest that caregiving responsibilities are predominantly seen as a duty of mothers, highlighting the enduring nature of traditional gender roles within the household. Moreover, the lack of any impact on fathers' labour market outcomes, combined with the significant and positive effect on family income, highlights the importance of this policy not just for mothers but for the entire household. By enabling mothers to enter the labour market, preschool contributes substantially to household income, enhancing economic stability and potentially lowering poverty levels.

[8]Within my sample, 75% of households have the father as the head or spouse's head.

Table 8. Effects of preschool enrolment on father's labour market outcomes

	First Stage	Local Average Treatment Effect					
	Attendance	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage	Family income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: Youngest child							
RD Estimate	-0.167***	0.112	0.104	-0.337	-5.517	0.366	0.875***
	(0.027)	(0.058)	(0.074)	(5.237)	(4.301)	(0.277)	(0.282)
Observations	[3005;3261]	[3327;3578]	[3620;3912]	[3408;3667]	[3193;3409]	[4628;4839]	[4316;4621]
Mean	0.48	0.96	0.93	42.96	46.12	1.54	5.81
Robust 95% CI	[-.238 ; -.116]	[-.02 ; .245]	[-.064 ; .267]	[-12.618 ; 10.762]	[-15.52 ; 3.677]	[-.218 ; 1.04]	[.284 ; 1.512]
Robust p-value	0.00	0.10	0.23	0.88	0.23	0.20	0.00
Bandwidth (h)	38.78	42.36	46.72	43.28	43.80	64.52	55.51
PANEL B: Non-youngest child							
RD Estimate	-0.106***	-0.211	-0.207	-4.327	6.836	-0.012	0.596
	(0.033)	(0.124)	(0.164)	(10.300)	(9.762)	(0.746)	(0.665)
Observations	[2067;2150]	[2301;2373]	[2211;2298]	[2581;2650]	[2024;2120]	[1818;1904]	[1809;1900]
Mean	0.37	0.96	0.93	42.96	46.12	1.54	5.81
Robust 95% CI	[-.191 ; -.039]	[-.483 ; .087]	[-.5760 ; .178]	[-28.096 ; 19.407]	[-17.438 ; 28.045]	[-1.648 ; 1.874]	[-.848 ; 2.278]
Robust p-value	0.00	0.17	0.30	0.72	0.65	0.90	0.37
Bandwidth (h)	57.82	63.10	61.53	70.78	60.87	56.01	50.45

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

5 Concluding remarks

In this paper, I contribute to the growing literature on the effectiveness of formal childcare as a public policy to tip the balance in favour of mothers in the labour market. Exploiting exogenous variation in preschool attendance induced by the school-entry age regulation, I estimate a fuzzy regression discontinuity model to assess the effects of children's preschool attendance on their mother's labour market outcomes and how this changes with the presence of other female family members in a developing country like Brazil with high levels of informal childcare arrangements.

The findings reveal significant increases in labour market participation following the enrolment of the youngest child in preschool. Mothers who enrol their youngest child in preschool are 27 percentage points more likely to participate in the labour market, which represents a 40% increase. Additionally, there is a proportional rise in employment and hours worked; however, no effects were observed on the intensive margin (e.g., working hours and hourly wages for employed mothers). Furthermore, the impact on the extensive margin has substantial implications for family income, which increases by 70%. These results suggest that the ability to enrol children in preschool does not necessarily lead

women to secure jobs with longer working hours or better pay; rather, it facilitates their entry into the labour market. This effect likely arises from the time constraint release that allows mothers to work without significantly altering the nature or quality of the jobs they obtain. Conversely, no effect was found for mothers with younger children at home, as having an additional child to care for limits mothers' labour supply, even after their eligible child is enrolled in preschool.

The results are primarily driven by married mothers whose youngest child is affected by the policy and mothers who do not live with other female relatives. In these households, preschool leads to a 56% increase in mothers' labour force participation, raising weekly hours worked to as much as 23 hours and significantly boosting family income. These positive effects also underscore the greater vulnerability of single mothers or mothers with multiple children, highlighting a need for additional policies to support their labour market engagement and financial stability.

Alternatively, the absence of significant effects between mothers with no other female relatives in the household raises important questions about the dynamics within these households, particularly regarding whether changes occur for other women. However, my ability to accurately identify these effects is limited, as the regression discontinuity design focuses on local estimates and becomes less reliable with smaller sample sizes. Future research should delve deeper into understanding the underlying mechanisms behind these caregiving arrangements, contributing to the rigidity of gender gaps.

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A Appendix

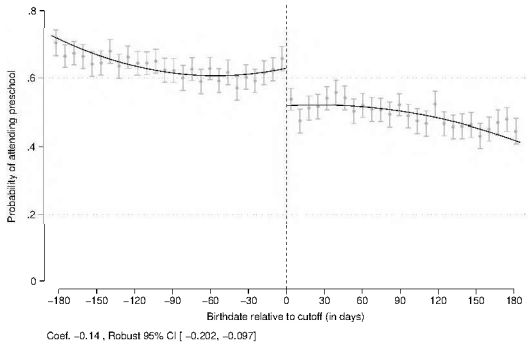
Table A.1. Changes in cutoff dates for Brazilian states (2001-2009 and 2011-2015)

State	**2001-2009**	**2011-2015**
Acre	December 31	March 31
Alagoas	June 30	March 31
Amapá	March 31	March 31
Amazonas	December 31	March 31
Bahia	March 31	March 31
Ceará		March 31
Espírito Santo	December 31	March 31
Federal District	March 31	March 31
Goiás	March 31	March 31
Maranhão	March 31	March 31
Mato Grosso		March 31
Mato Grosso do Sul		March 31 (until 2013)
Minas Gerais	June 30	June 30
Pará	December 31	March 31
Paraíba	December 31	
Paraná		December 31
Pernambuco	December 31	March 31
Piauí		March 31
Rio de Janeiro	June 30	December 31
Rio Grande do Norte	June 30	March 31
Rio Grande do Sul		March 31
Rondônia	March 31	March 31 (until 2013)
Roraima	June 30	June 30
Santa Catarina	December 31	March 31
São Paulo	December 31	June 30

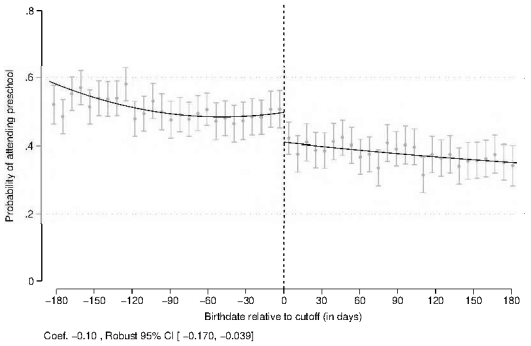
Source: Compilation of resolutions from the National Council of Education - Chamber of Basic Education of Brazil and various state ministries of education.

Notes: Sergipe and Tocantins were excluded due to lack of information about the eligibility rule.

Figure A.1. Effect of preschool eligibility on the probability of enrolment



(a) Youngest child



(b) Non-youngest child

Notes: Second-order polynomial approximation using a triangular kernel with a 95% confidence interval. The dots in the scatterplots represent the average value of school enrolment rates in 7-day birthdate bins. Figure A.1a is based on the subsample of mothers with the youngest child eligible to enrol and Figure A.1b with the subsample of mothers with non-youngest children eligible for preschool.

Table A.2. Weak identification test results

F-statistic	**Value**
Montiel-Pflueger Robust	56.50
Cragg-Donald Wald	62.19
Kleibergen-Paap rk Wald	61.75

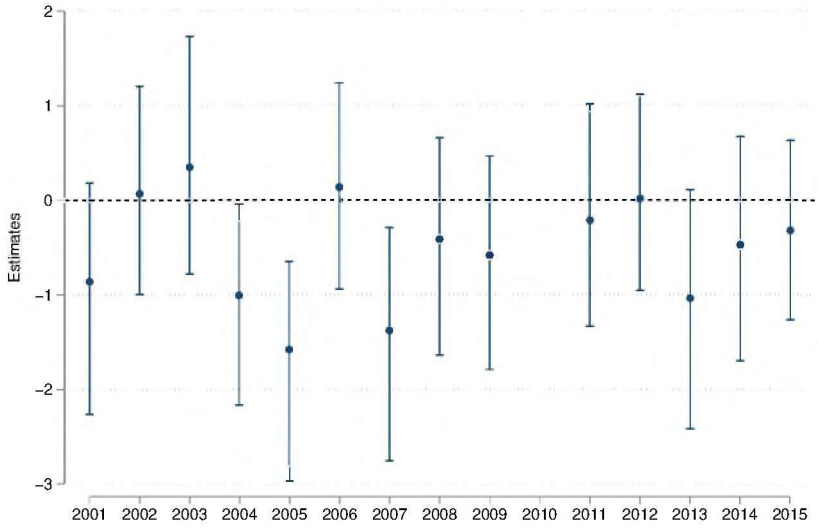
Notes: The estimates are for mothers between 18 and 49 years of age.

Table A.3. Effects of preschool enrolment on mothers' labour market outcomes including fixed effects

	First Stage	Local Average Treatment Effect					
	Attendance	Participation	Employment	Hours worked	Hours worked (cond)	Hourly wage	Labour income
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PANEL A: Youngest child							
RD Estimate	-0.144***	0.325**	0.378***	17.017***	8.103	0.579	0.794***
	(0.022)	(0.131)	(0.141)	(6.037)	(6.761)	(0.393)	(0.249)
Observations	[3935;4278]	[5485;5901]	[5155;5590]	[5061;5475]	[3571;3719]	[3002;3153]	[5561;5964]
Mean	0.49	0.67	0.58	20.91	35.82	1.37	5.78
Robust 95% CI	[-.204 ; -.102]	[.057 ; .644]	[.093 ; .741]	[3.812 ; 31.105]	[-8.029 ; 22.696]	[-.309 ; 1.525]	[.247 ; 1.339]
Robust p-value	0.00	0.02	0.01	0.01	0.35	0.19	0.00
Bandwidth (h)	37.61	52.16	49.25	48.78	57.01	56.43	53.42
FE Year and State	YES	YES	YES	YES	YES	YES	YES
PANEL B: Non-youngest child							
RD Estimate	-0.100***	0.062	-0.050	0.881	1.389	0.339	0.293
	(0.027)	(0.274)	(0.279)	(10.335)	(10.571)	(0.722)	(0.501)
Observations	[2810;2880]	[3297;3380]	[2972;3038]	[3532;3593]	[1291;1363]	[1074;1109]	[3245;3333]
Mean	0.38	0.51	0.42	13.41	31.75	1.12	5.10
Robust 95% CI	[-.168 ; -.046]	[-.515 ; .77]	[-.648 ; .659]	[-18.809 ; 29.455]	[-20.675 ; 28.17]	[-1.243 ; 2.08]	[-.764 ; 1.573]
Robust p-value	0.00	0.70	0.99	0.67	0.76	0.62	0.50
Bandwidth (h)	57.51	67.19	60.19	71.30	61.12	68.55	66.95
FE Year and State	YES	YES	YES	YES	YES	YES	YES

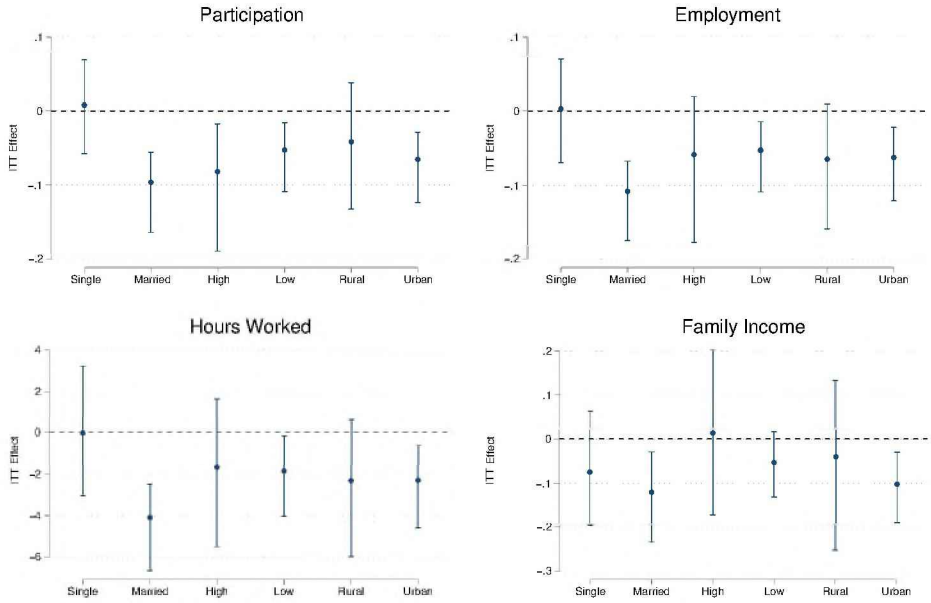
Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. h is the MSE optimal main bandwidth based on Calonico et al. (2014). Observations are sample sizes within the cutoff's main bandwidth to the right and left. Conventional standard error of local-polynomial RD estimator in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

Figure A.2. Effect of preschool eligibility on mothers' years of education



Notes: Figure presents the point estimates and 95% confidence intervals from regressing years of education on birthdates, relative to the cutoff date, by year. Each estimate corresponds to a different regression. The estimates are based on a local linear polynomial with a triangular kernel and robust inference. Each year, the bandwidth is established as the optimal MSE according to Calonico et al. (2014).

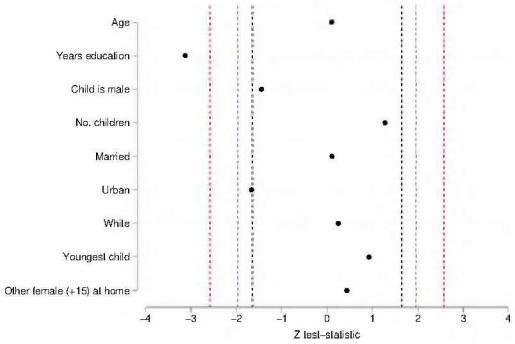
Figure A.3. Effects of preschool eligibility by marital status, educational level and geographic area



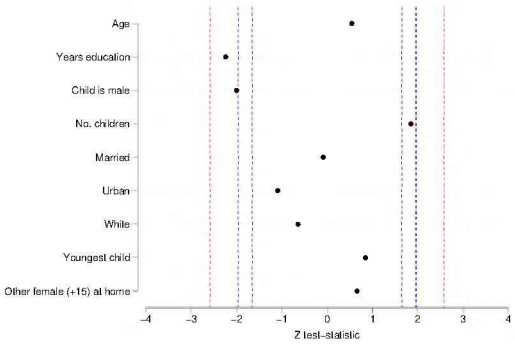
Notes: Figure presents the point estimates and 90% confidence intervals from regressing the outcome on birthdate relative to the cutoff date. Low-educated women are defined as having a high school degree or less, and high-educated women as having at least some college education. Each estimate corresponds to a different regression. The estimates are based on a local linear polynomial with a triangular kernel and robust inference. In each estimation, the bandwidth is established as the optimal MSE according to Calonico et al. (2014). Married refers to women in formal and consensual unions.

Figure A.4. Balance analysis of predetermined characteristics

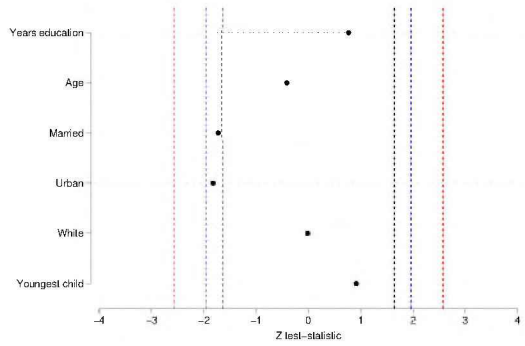
(a) Mothers



(b) Fathers



(c) Other (+15) female relatives



Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. Bandwidth is the MSE optimal based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff.

Table A.4. Balance analysis of predetermined characteristics of fathers

Variables	**MSE optimal bandwidth**	**RD estimator**	**Robust Inference**		**Observations Left,Right**
			p-value	**Confidence Interval**	
Age	53.78	0.146	0.62	[-0.57 ; 0.95]	[6083, 6482]
Married	74.30	-0.001	0.98	[-0.03 ; 0.03]	[8589, 8966]
Years of education	51.82	-0.453	0.02	[-0.90 ; -0.08]	[5826, 6228]
Urban	51.79	-0.020	0.30	[-0.06 ; 0.02]	[5850, 6258]
White	47.87	-0.015	0.51	[-0.06 ; 0.03]	[5404, 5773]
Child is male	67.57	-0.036	0.05	[-0.08 ; 0.00]	[7737, 8105]
No. children	68.03	0.108	0.05	[0.02 ; 0.21]	[7877, 8232]
Youngest child	56.22	0.010	0.46	[-0.03 ; 0.06]	[6423, 6823]
Other female (+15) at home	47.70	0.011	0.49	[-0.02 ; 0.04]	[5414, 5780]

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. Bandwidth is the MSE optimal based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff.

Table A.5. Balance analysis of predetermined characteristics of other (+15) female relatives

Variables	MSE optimal bandwidth	RD estimator	Robust Inference		Observations Left,Right
			p-value	Confidence Interval	
Age	67.86	-0.505	0.68	[-2.471 ; 1.484]	[3730, 4096]
Married	73.45	-0.039	0.09	[-0.082 ; -0.002]	[4059, 4427]
Years of education	66.41	0.158	0.45	[-0.231 ; 0.623]	[3661, 4024]
Urban	48.02	-0.039	0.07	[-0.085 ; -0.004]	[2647, 2908]
White	64.04	-0.003	0.99	[-0.048 ; 0.047]	[3591, 3928]
Youngest child	55.41	0.031	0.36	[-0.023 ; 0.082]	[3048, 3316]

Notes: Results from local linear polynomial estimation with a triangular kernel and robust inference. Bandwidth is the MSE optimal based on Calonico et al. (2014). Observations are sample sizes within the main bandwidth to the right and left of the cutoff.