# The effects of an elementary school reform on maternal labor outcomes

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### Abstract

This paper investigates the impact of an educational reform in Brazil that lowered the mandatory school entry age from seven to six years. Using a Local Average Treatment Effect (LATE) and a Differences-in-differences estimation, we estimate the effects of the reform on school enrollment and maternal labor outcomes. This reform allowed us to address the typical non-randomization of school enrollment. Our results show that the policy increased elementary school enrollment, primarily by shifting children from preschool to elementary education. This shift produced modest effects on maternal outcomes: labor force participation rose by 1.7 percentage points, and time spent on housework declined by 0.59 hours per week. Heterogeneity analysis reveals stronger effects for Black and rural mothers, whose likelihood of labor force participation increased by 4.78 and 6.93 percentage points, respectively. These findings underscore the complexity of maternal labor supply decisions and highlight the importance of considering intersectional factors in policy design.

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# 1 Introduction

Female labor force participation has globally increased over the past decades. Despite these advancements, relevant gender inequalities persist in the labor market (Goldin, 2014). Evidence shows that having children permanently affects women's career trajectories, whereas men's careers remain unaffected or face minor impacts (Kleven; Landais; Søgaard, 2019). Since women are often the primary caregivers for their children, maternal labor force participation and childcare responsibilities are closely intertwined, highlighting the critical role of care provision in shaping women's decisions and availability to work outside the home.

In Brazil, women 's employment rate falls 37% relative to men after having children (Kleven; Landais; Leite-Mariante, 2024). A disparity that persists even after the children have grown up. Beyond loss of employment, motherhood also drives women towards more flexible but precarious employment arrangements, such as part-time jobs, self-employment, and informal work (Berniell et al., 2023). Although this consistent empirical pattern underscores the disproportionate career impacts of motherhood, the mechanisms behind these outcomes and the potential of targeted policies to mitigate these effects remain less understood (Canavire Bacarreza; Yanez, 2024).

In this context, higher school enrollment for younger children has the potential to reshape mothers' work lives by influencing their ability to participate in the labor market and reducing housework burdens. This study examines how an educational reform that extended the duration of the elementary school in Brazil, effectively lowering the minimum school entry age from 7 to 6 years, affected mothers' labor market participation and time spent on housework. Six-year-old children could previously be enrolled in non-mandatory preschool, and due to the reform, should mandatorily be enrolled in elementary school. By alleviating childcare responsibilities through expanded access to schooling, these types of reforms may enable mothers to engage more fully in economic activities by entering the workforce or increasing their working hours.

Using data from the Brazilian Household Survey (PNAD, in the Portuguese Acronym), we employ two empirical strategies: a Local Average Treatment Effect (LATE) and a Differences-in-differences (DiD) estimation, leveraging this reform to understand how it impacted elementary (and preschool) school enrollment. The dataset provides information on mothers' labor market outcomes and children's school enrollment. We selected households with mothers aged 17 to 49 and only one child, between 5 and 7 years old to investigate whether lowering the age of school enrollment influenced maternal labor market outcomes and time use.

Our findings indicate that the reform led to an increase in elementary school enrollment, primarily driven by a shift from preschool to elementary education. The expansion of access to elementary school had modest but meaningful effects on mothers' labor market outcomes. Specifically, it increased the likelihood of maternal labor force participation by 1.7 percentage points. Additionally, it led to a reduction of 0.59 hours per week in mothers' time devoted to housework. Heterogeneity analysis reveals that these effects were not uniform: Black mothers and those residing in rural areas experienced the largest gains. The probability of participation in the labor force increased by 4.78 percentage points among Black mothers and by 6.93 percentage points among rural mothers. These larger effects likely stem from Black and rural mothers facing greater pre-existing barriers to labor market participation and a greater increase in school enrollment. These findings highlight the complexity of female labor force participation and emphasize how mothers' demographic and socioeconomic characteristics can affect policy interventions.

This study contributes to a growing body of literature that examines the relationship between early childhood education policies and maternal labor market outcomes. Previous literature found mixed results, with varying effects depending on institutional and socioeconomic contexts. While several studies report small or no impacts of preschool or daycare access on women's labor supply, often due to high pre-existing maternal employment rates, substitution with informal care, or limited labor supply elasticity (Baker; Gruber; Milligan, 2008; Finseraas; Hardoy; Schøne, 2017; Carta; Rizzica, 2018), other studies point to positive effects, particularly among subgroups such as single mothers or those without younger children (Goux; Maurin, 2010; Cascio; Haider; Nielsen, 2015; Contreras et al., 2016).

In the Latin American context, most studies suggest that access to formal childcare tends to improve maternal labor outcomes, although the magnitude of effects varies. Evidence from Argentina, Chile, Ecuador, and Nicaragua generally supports a positive association between early childhood education and mothers' employment (Berlinski; Galiani, 2007; Rosero; Oosterbeek, 2011; Lopez Boo; Hojman, 2019). However, some studies detect meaningful heterogeneity: effects are concentrated among lone mothers or mothers of only children in Argentina and Chile (Berlinski; Galiani; Ewan, 2011; Contreras et al., 2016), while Ivone Perazzo (2019) found no significant impact at all in Uruguay. More recent analyses expand the focus to unpaid labor, showing that the use of childcare services is linked to a reduction in household work and an increase in paid labor hours among women (Amarante; Rossel; Scalese, 2023).

The Brazilian literature remains relatively recent but aligns with broader regional findings. A study by Ryu (2020) on the 2009 reform that made preschool mandatory

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for four-year-olds finds that mothers of eligible children increased formal employment and reduced time spent on domestic chores, but only when the child was the youngest in the household and no other relatives were cohabiting. Similarly, for the same 2009 reform, Reyna and Costa Silva (2021) document unintended effects on older siblings, who increased their participation in household chores and, in some cases, reduced school attendance, especially among girls.

Attanasio et al. (2022) use a specific lottery assignment in 2007 to assess the long-run effects of public daycare in the city of Rio de Janeiro and find increases in household income driven not by mothers, but by other household members such as grandparents and adolescent siblings. They argue that this is related to the already reasonably high employment rates of mothers and fathers at 70% and 90%. This could be the issue for our case, given that over 70% [1] of mothers in our sample are already employed before the reform.

We expand the literature by also looking at mothers of multiple children, but whose youngest child is affected by the policy. Additionally, we more deeply explored heterogeneity among mothers, addressing key contextual factors that may shape the results in the Brazilian setting. Our study also contributes to the literature on nationwide education policies by examining an earlier national reform that targeted a different age group, children aged 5 to 7, compared to those analyzed in previous studies by (Ryu, 2020) and (Reyna; Costa Silva, 2021).

The remainder of this paper is structured as follows. Section 2 describes education in Brazil and the overall context of the reform. Section 3 explains the theory and mechanisms behind maternal labor and child school enrollment. Section 4 provides information on the data used in the paper. Section 5 shows the methodology, including an explanation of the empirical strategy, the construction of key variables, and the identification strategy. Section 6 presents the main results and discussions. Finally, we conclude with a summary of the results, policy implications, and suggestions for future research.

# 2   Education in Brazil

Child education is a right in Brazil and is provided in collaboration between the Union, States, Federal District, and Municipalities, and through private initiative. Municipalities are responsible for the provision of early childhood education (nursery and preschool) and elementary education, the latter being shared with the states.

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[1]This number is higher than countries such as Korea, Mexico, Italy, and Turkey (OECD, 2023)

Parents can opt for formal care/education in the public or private sector. Public centers are free and universal, depending on age requirements and upon availability, for all families, regardless of socioeconomic level. Private centers charge a market price and work independently of public centers while following the laws and guidelines on education in Brazil. The percentage of students in public schools in 2005 was high: 89.2% of elementary school students and 76.3% of preschool students (INEP, 2007).

During the period of the reform, the minimum length of the school day in Brazil was 4 hours for half-day programs and 7 hours for full-day programs. However, some schools offered longer instructional hours. According to data from the School Census, the share of elementary school enrollments in the Southeast region [2] attending schools with more than 5 hours of daily instruction nearly tripled between 2004 and 2006, rising from 6.3% to 18.5% (INEP, 2007). This trend contrasts sharply with other regions, which started from much lower baselines and experienced either stagnation or declines over this period [3]. Our dataset does not provide information on the length of the school day for the child, nor whether the program is of full or half day, which prevents us from controlling for time spent at school.

A law enacted in 2005 (Lei 11.114/05) changed the Law of Guidelines and Bases of National Education (LDB in Portuguese), making compulsory the start of elementary education at six years old starting in 2006. In 2006, another law (Lei 11.274/06) increased the duration of elementary education from 8 to 9 years, while lowering the official age of elementary school entry from 7 to 6 years old. The law required the new school length to be adopted by 2010, but made the enrollment of 6-year-olds mandatory since 2006. According to the Ministry of Education, the change of the entrance age and the expansion of compulsory education aimed to ensure all children spend more time in school, have more opportunities to learn, and have a quality education, objectives unrelated to female labor force participation.

Before the reform, preschools in Brazil served children between the ages of 4 and 6, while elementary education formally began at age 7. Public preschools, administered by either municipal or state governments, were tuition-free and had to admit all children unless demand exceeded available capacity. Regulatory provisions stipulated that parents were to enroll their children in public preschools located near their residences. Private preschools operated on a usual for-profit basis and retained full

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[2]The Southeast region comprises the states of Rio de Janeiro, Minas Gerais, Espírito Santo, and São Paulo.

[3]Percentage of enrollments in elementary schools with over five hours of instruction per day in 2006 and growth rate (in parentheses): North – 1.1% (-52%); Northeast – 1.0% (+11%); South – 1.3% (+44%); Center-West – 3.8% (-13%) (INEP, 2007).

discretion over tuition fees (Bastos; Straume, 2014). Preschool enrollment was not mandatory: enrolling the child in preschool was the parents' decision, and not all children were enrolled.

After the reform, preschools were to enroll only students between 4 and 5 years old. Elementary schools opened a new grade for 6-year-old students, designated as the 1st grade. Daycare continued enrolling students from 0 to 3 years old. Different from preschool, elementary school enrollment was mandatory. Preschool enrollment only became mandatory in Brazil in 2013, with an even lower entrance age of 4 years.

Although the new law required 6-year-old children to begin elementary school, each state had the freedom to regulate implementation. States and schools adopted their own rules on the entrance age and birth date cut-off. The state of São Paulo, for example, implemented the change in 2006, stating that children who had already turned 6 by the end of the present year would be allowed to enroll in elementary school the following year. The lack of a national birth date cut-off made the adoption of a regression discontinuity design, used by Ryu (2020) and Reyna and Costa Silva (2021), not feasible for this paper.

Later, the Brazilian government approved a constitutional amendment (EC 59/09) in 2009 that went further with the expansion of schooling in the country. The amendment expanded mandatory education by extending the compulsory enrollment to students between the ages of 4 and 17. Universalization had to be implemented by states and municipalities by 2016. Figure 1 shows the timeline of the law that made school enrollment mandatory for 6-year-olds (Lei 11.114/05) and both school reforms (Lei 11.274/06 and EC 59/09) and deadlines.

Figure 1: School Reforms Timeline
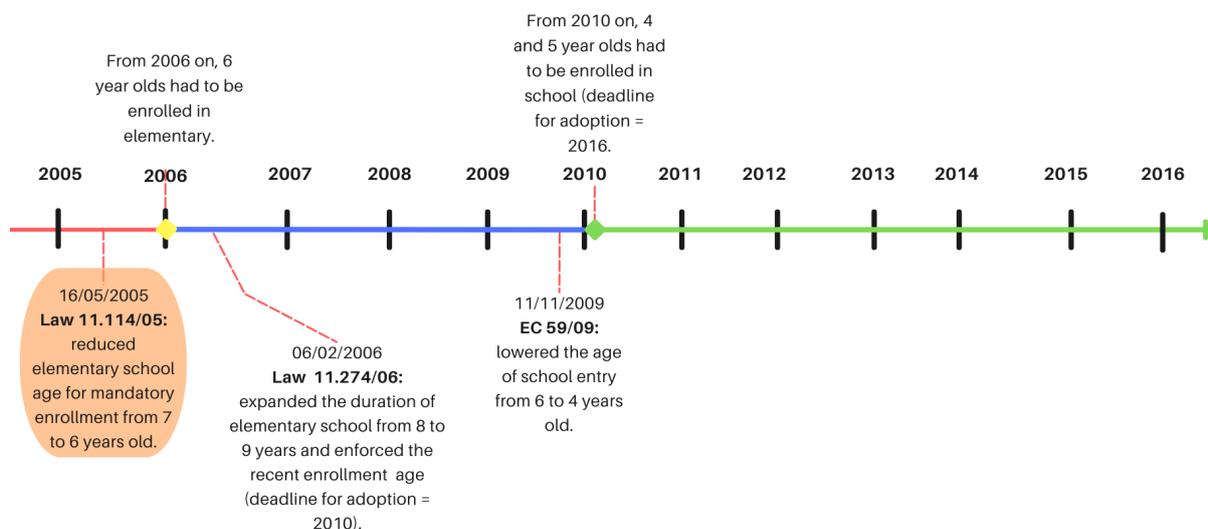


Notes: Timeline of approval of the main laws that changed Brazilian education from 2001 to 2011.

The effects of this secondary reform have been studied by Ryu (2020) and Reyna and Costa Silva (2021). The former found that preschool attendance increased the average weekly working hours of the mother without additional younger children and other relatives. The latter also found positive effects, but observed that the impact is larger for mothers from families formed by couples than for lone mothers.

Six-year-old children were the target of the educational reform we addressed. They were not previously subjected to school legally, but could be enrolled in preschool. 7-year-olds were obligated by law to study and should be enrolled in elementary school. 5-year-old children should not be affected by the change in theory. Figures 2 to 4 show the share of children at each type of school by age. Figure 2 shows us that children who started the year as five-year-olds were affected by the policy, changing from preschool to elementary school. The proportion of children who started the year as six-year-olds in elementary school grew around 20 percentage points until 2010, while the percentage at preschool lowered in an even bigger magnitude (See Figure 3). Seven-year-olds were unaffected: they were the previous entrance age and already had a near-total elementary school enrollment (See Figure 4).

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Figure 2: School enrollment of children who started the year as 5-year-olds, 2001-2011
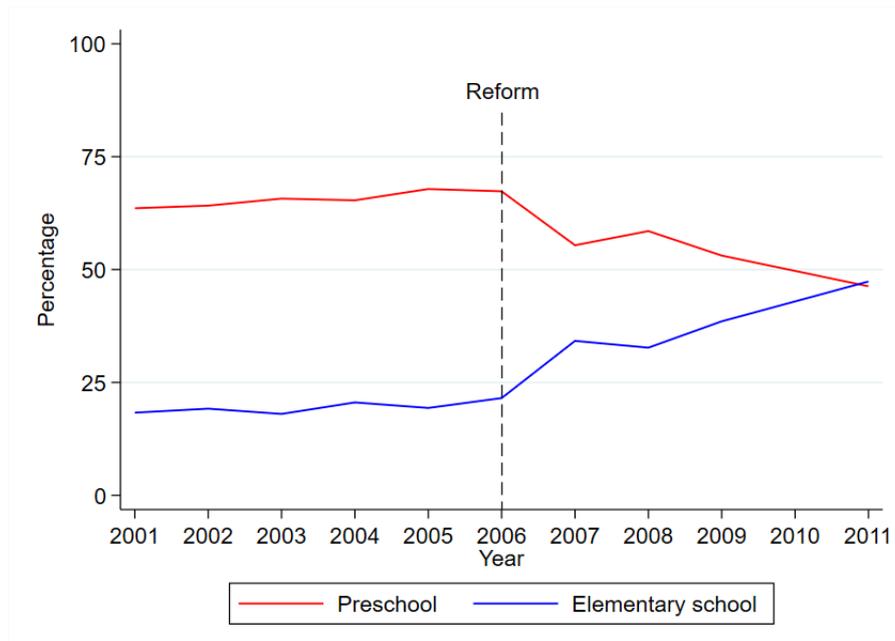


Notes: Percentage of children who started the year as 5-year-olds by school level from 2001 to 2011. Data from PNAD.

Figure 3: School enrollment of children who started the year as 6-year-olds, 2001-2011
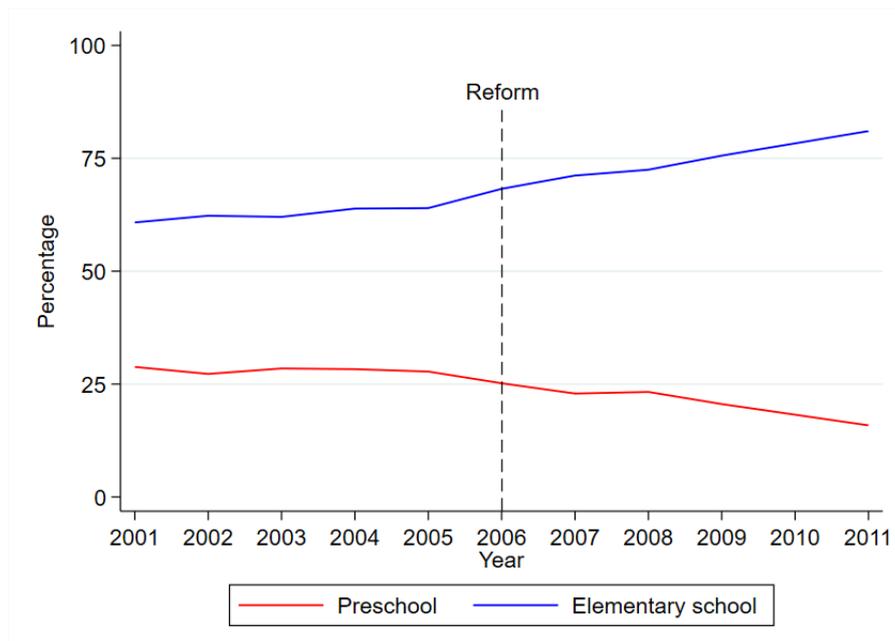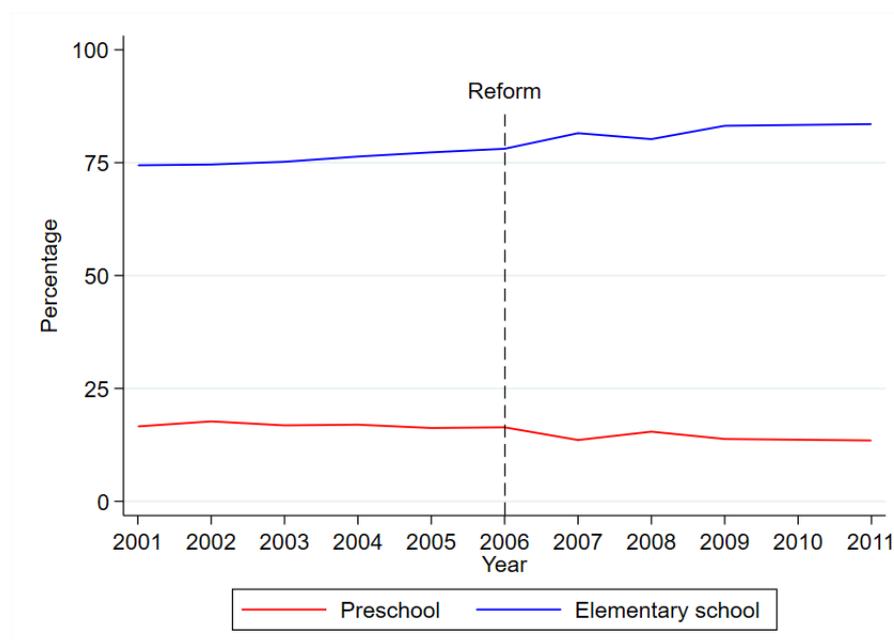


Notes: Percentage of children who started the year as 6-year-olds by school level from 2001 to 2011. Data from PNAD.

Figure 4: School enrollment of children who started the year as 7-year-olds, 2001-2011



Notes: Percentage of children who started the year as 7-year-olds by school level from 2001 to 2011. Data from PNAD.

# 3   How would the reform affect mothers?

This section examines the mechanisms through which access to early childhood education and care can affect maternal labor outcomes. While the exact consequences of changes in the relative price of childcare on maternal participation depend on an underlying model of labor supply, improved access to free public early childhood education constitutes an implicit childcare subsidy for parents (Carta; Rizzica, 2018). Theoretically, the *ex ante* impact of lowering the price of childcare on the number of hours worked is generally ambiguous due to the opposite forces of positive substitution and negative income effects (Carta; Rizzica, 2018), employment, however, increases as a consequence of the higher indirect utility in the case of work.

However, the effects of lower childcare costs depend not only on economic incentives but also on existing access and availability. In developing countries like Brazil, where households often include multiple family members (UN, 2022), the presence of informal caregivers, such as grandparents or older siblings, introduces important nuances. In such contexts, the introduction of public childcare may simply displace existing informal arrangements. This crowding out of family-provided care could lead to a lack of labor market responses among mothers. Indeed, in Latin America, informal care provided by household members remains the primary alternative to formal,

center-based care (Lopez Boo; Hojman, 2019). In Brazil, relatives play a particularly important role: for example, (Reyna; Costa Silva, 2021) found that an earlier education reform increased preschool attendance and mothers' working hours, but also led to greater household responsibilities for older siblings, reducing their school attendance in low-income families.

Family composition also matters. The presence or absence of a second parent in the household shapes maternal labor decisions. Lone mothers face two opposing forces in our context: they are more vulnerable by being the sole provider for the child, but as the sole provider, they often cannot afford to be jobless or work few hours. The net effect of these forces is inconclusive and will depend on specific circumstances faced by lone mothers.

In addition to direct labor market responses, it is important to consider how mothers reallocate their time across other non-market activities. A reduction in childcare costs may free up time previously spent on direct childcare, but this time could be redirected toward housework, especially in the short term. As entering or re-entering the labor market takes time and effort, a constant stream of tasks that go beyond childcare may temporarily absorb part of the time freed up by reduced childcare responsibilities.

Beyond household dynamics, structural constraints in the education system also play a role. For parents seeking education for their six-year-old children, the choice is typically between public and private education. Although overall enrollment rates were relatively high even before the reform, a significant number of children were still not attending school. This may reflect several factors: some parents may not have considered early education essential, others may have been unable to secure a spot in the public system, or were financially constrained from choosing private options.

While our data do not include detailed information on children's school enrollment before the reform, the 2004 PNAD survey provides relevant insights. Among five-year-olds not enrolled in school that year, 22.15% cited parental or guardian preference as the main reason. An additional 14.63% reported the absence of nearby schools, and 11.82% indicated a lack of available places. These figures underscore both demand and supply-side barriers to enrollment that may have shaped parental decisions before the reform.

## 4 Data

We use data from the Brazilian Household Survey (PNAD, Portuguese Acronym), collected by the Brazilian Bureau of Statistics (IBGE). The survey assessed annual sociodemographic data and household information for more than 100,000 Brazilian

households and 300,000 individuals. PNAD was the Brazilian national household sample survey until 2015, and its sample is representative at the state and metropolitan region level.

We use data covering the period from 2001 to 2011. The PNAD was not collected during Census years, which was the case in 2010. We tried to use Census data for the year 2010, but the survey differs on variables important to this analysis, forcing us to ignore this year in our sample. Although PNAD looks at the time spent on housework, the Census does not collect this information. Moreover, the Census only looks at employment on a specific week of the survey month (July), whereas PNAD has data on employment during the year of analysis.

Following Ryu (2020), we restrict our sample based on mothers' age and the number of children. We selected households with mothers aged 17 to 49 and only one child, between 5 and 7 years old. This age range of the mothers corresponds to their usual working age, and the focus on only having one child stems from the possible interference of other children's needs affecting mothers' employment (Ryu, 2020). Since siblings can change household dynamics and be differently affected by an educational reform (Reyna; Costa Silva, 2021), we focus primarily on mothers of only children to ease and strengthen identification and causal inference.

The data provide detailed information on mothers' labor market outcomes, which we used as the primary dependent variables: the average number of hours spent working on the market and on housework per week [4], whether the mother was employed or actively seeking work, and whether she had a formal or informal labor contract. A formal labor contract is defined as having a formal labor card (*carteira de trabalho*) registered with an employer [5]

# 5 Empirical Framework

In this study, our objective was to test the impact of changes in mandatory elementary education rules on mothers' labor market participation. Specifically, we analyze how the policy reform, which lowered compulsory enrollment at elementary school, influenced maternal labor supply. We also explore how the reform impacted enrollment in preschools and net school enrollment.

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[4]Even though the variable's name is "House work", it also includes time spent on care for children and the elderly. In the PNAD survey, the question about the housework was direct, without explaining the activities to be considered as domestic tasks. Housework was defined in the interviewer's guidelines as the performance of non-economic tasks in the home: cleaning, cooking or preparing food, washing or dishwashing, ironing, guiding or directing domestic workers in the execution of household tasks, and caring for children, minor residents, and the elderly. Thus, this variable also includes hours of childcare.

Our strategy addresses two main identification challenges: (1) non-random selection of parents (or caregivers) into early education and (2) the potential for confounding factors, such as childcare costs or availability of childcare, and reverse causality, as maternal labor force participation could influence school enrollment decisions.

To address the potential endogeneity of school enrollment, we propose a two-stage approach where we estimate our model by Two-Stage Least Squares (2SLS). Under standard IV assumptions, this yields a Local Average Treatment Effect (LATE). First, we analyze the impact of the reform on school enrollment ($Enroll_{ist}$) for eligible students ($EligStudents_{ist}$) after the reform ($Post_t$). This is what we call our first-stage regression. We run the same regression for elementary, preschool, and overall school enrollment to understand the transition from the reform.

In this first stage of regression, we exploit the policy change that mandated education at a specific age: we compared the students' school enrollment for mothers of 5 or 6-year-olds (treated group) and 7-year-olds (control group) before and after the reform. The "pre-reform" period covers the years 2001 to 2005 in the PNAD data, while the "post-reform" period spans from 2006 to 2011, excluding 2010, as no survey was conducted that year.

We argue that the policy reform was exogenous for the three student cohorts and induced enrollment decisions for eligible students. This approach mitigates selection bias by ensuring that enrollment is driven by the reform rather than parental choice.

$$Enroll_{ist} = \alpha_1 + \delta(EligStudents_{ist} * Post_t) + \beta X_{1ist} + \lambda_t + \mu_s + \epsilon_{ist} \tag{1}$$

Where $EligStudents$ is a dummy variable indicating if the child starts the school year being six years; $X_{1ist}$ is a vector of socio-demographic covariates that affect school enrollment and represent the same cohort of mothers[6]; and $\epsilon_{ist}$ is an error term assumed to be independent and identically distributed. The model includes year fixed effects, captured by $\lambda_t$, to control for possible year-specificities and state fixed effects, captured by $\mu_s$, to control for time-invariant differences across states. We cluster standard errors at the state level.

For the second stage, we run the estimated enrollment driven by the policy on several mother labor market outcomes:

$$Y_{ist} = \alpha_2 + \tau E\hat{nroll} + \eta X_{2ist} + \psi_t + \mu_s + \varepsilon_{ist} \tag{2}$$

Where $Y_{ist}$ stands for one of the following the outcome variable for mother $i$ in year

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[6]The covariates include: age, race, years of mother's education, age of the child, number of people that reside in the house, if there is a relative at home and if they live in an urban or rural area.

$t$ and state $s$: average number of hours spent working and on chores, if the mother worked or was looking for a job, and if they have a formal or informal labor contract. In the cases of dummy outcome variables, such as labor force participation, it is assigned one if she is in the labor force and zero otherwise. $\hat{Enroll}$ is the estimated enrollment driven by the policy. $X_{2ist}$ is a vector of covariates that affect maternal labor market outcomes, and $\varepsilon_{ist}$ is an error term assumed to be independent and identically distributed. The model also includes year and state fixed effects, captured by $\psi_t$ and $\mu_s$. The parameter of interest, $\tau$, represents the mean effect of a child's elementary enrollment on the maternal labor market outcome.

We also estimate a reduced-form equation using OLS (Ordinary Least Squares), simulating a Difference-in-Differences (DiD) approach to study the direct impact of the policy in a quasi-experimental setting. This method allows us to investigate the induced effect of the reform on maternal labor outcomes.

$$Y_{ist} = \alpha_3 + \tau^*(EligStudents_{ist} * Post_t) + \eta X_{2ist} + \theta_t + \mu_s + \varepsilon_{ist} \tag{3}$$

This specification is similar to the others. $X_{2ist}$ is a vector of variables that affect maternal labor market outcomes. The covariates include: age, race, years of education, number of people who reside in the house, if there is a relative at home, and if they live in an urban or rural area. The model also includes year and state fixed effects, captured by $\theta_t$ to control for possible year-specific trends and unobserved factors. The parameter of interest, $\tau$, represents the mean effect of a child's elementary enrollment on the maternal labor market outcome. We use the reduced-form because these regressions provide intention-to-treat effects, trying to avoid possible inference problems due to treatment compliance behavior varying in complex ways across individuals. We also cluster standard errors at the state level.

## 6 Results

This section presents the main findings of the empirical analysis. We begin by analyzing our sample by treatment, then describing the baseline estimates, followed by a discussion of the robustness checks and alternative specifications. Lastly, we explore potential heterogeneous effects.

### 6.1 Sample description

Table 1 presents summary statistics for the mothers in our sample, comparing the treatment group (mothers of children who begin the year aged 5 or 6) to the control

group (mothers of children who begin the year aged 7) for our whole period of analysis from 2001 to 2011 (See Table 21 on the Appendix for the table before the reform).

The sample consists primarily of White, urban women with an average age of around 32. Mothers in the control group are, on average, older than those in the treatment group. This difference is expected, as the control group includes mothers of older children. Racial composition is similar across groups: 52% of mothers in the control group and 51% in the treatment group identify as White. Location is also similar between treatment and control, both being more located in urban areas.

There is a notable difference in family structure: 15% of mothers in the treatment group are lone mothers, compared to only 4% in the control group. To investigate whether this gap could be explained by age, assuming older mothers in the control group may have had more time to remarry, we compared the ages of lone mothers across groups. However, we found no significant age difference. Lone mothers in both groups are more similar in age to each other than to the broader sample of mothers.

Household characteristics such as the presence of relatives and the overall number of people in the household are similar. On average, households contain about three people, likely reflecting the nuclear family structure—mother, partner, and child.

Labor market outcomes are also broadly comparable. Labor force participation rates are almost identical between groups, with 73% in the control group and 72% in the treatment group, while average work hours are slightly lower in the treatment group, 24.45, compared to the control group, 25.35. Formal employment rates are similar, and job search rates are marginally higher in the treatment group. Time spent on housework is also slightly higher among treated mothers, but the difference is not statistically significant.

Expected differences arise in children's school enrollment, which aligns with the reform's age cutoffs. Only 0.7% of control group mothers have a child in preschool, compared to 45% in the treatment group. Conversely, 96% of control group mothers have a child in elementary school, versus 49% in the treatment group. These patterns are consistent with the eligibility criteria: children in the control group were already subject to compulsory elementary education, while those in the treatment group were younger and only enrolled due to the 2006 reform. Overall school attendance is high in both groups, 97% in the control group and 94% in the treatment group.

Table 1: Comparison of Treatment and Control Groups (2001-2011)

|  | Control | Treatment | Difference |
|---|---|---|---|
| Age | 32.93 | 31.13 | 1.8*** |
| White | 0.52 | 0.51 | 0.01* |
| Urban | 0.89 | 0.88 | 0.01*** |
| Lone Mother | 0.04 | 0.15 | -0.10*** |
| Relatives Present | 0.08 | 0.09 | -0.00 |
| People on the Household | 2.99 | 3.02 | -0.03*** |
| Labor Force Participation | 0.73 | 0.72 | 0.009* |
| Work Hours | 25.35 | 24.45 | 0.90*** |
| Formal Work | 0.31 | 0.30 | 0.00 |
| Job Search | 0.06 | 0.07 | -0.009*** |
| Housework | 26.82 | 27.02 | -0.20 |
| Child in Preschool | 0.007 | 0.45 | -0.44*** |
| Child in Elementary | 0.96 | 0.49 | 0.47*** |
| Child in School | 0.97 | 0.94 | 0.02*** |
| Observations | 8,382 | 19,356 | |

Notes: The table compares the characteristics of mothers in the Control group, i.e., with children who start the year as seven-year-olds, and the Treatment group, i.e., with children who start the year as six or five-year-olds. Column "Control" presents the average control mothers' characteristics, column "Treatment" reports the average characteristics for treated mothers, and column "Difference" reports the differences between both groups and the t-test statistical significance of this result. The variables analyzed are Dummy for whether the mother is White; Dummy for whether the mother lives in an urban area; Dummy for whether the mother is a lone parent; Dummy for whether other relatives cohabit in the household; Number of people on the household; Dummy for whether the mother participates on the labor force; Number of hours worked on the labor market; Dummy for whether the mother had a formal work attachment; Dummy for whether the mother was looking for a job; Percentage of mothers whose children were in preschool, elementary school or any type of school.
P-values: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

## 6.2 Impact of the reform on school enrollment

Our first-stage results confirm that the policy reform had a significant impact on elementary school enrollment among eligible children (see Column 1 of Table 2). Specifically, the interaction term is positive and statistically significant (F-statistic of the excluded instrument is 63.94), indicating that the reform effectively increased the likelihood of elementary school enrollment for six-year-olds after implementation: children

of eligible age were 16.2 pp. more likely to be enrolled in elementary school.

Likewise, the results in Column 2 of Table 2 confirm the expected negative impact of the policy reform on preschool enrollment among eligible children. In this case, the interaction term is negative and statistically significant (F-statistic of the excluded instrument is 32.75), indicating that the reform effectively decreased the likelihood of elementary school enrollment for six-year-olds post-implementation: eligible children were 14.4 pp. less likely to be enrolled in preschool.

Column 3 of Table 2 shows that the reform led to a 1.8 percentage point increase in overall school enrollment among eligible children. This net effect reflects a shift in enrollment from preschool to elementary school, as evidenced by the corresponding increase in elementary enrollment and decrease in preschool attendance. While the reform primarily reallocated children across school levels, it also resulted in a modest but significant rise in overall enrollment.

Table 2: Impact of the reform on school enrollment (2001–2011)

| | (1) Enrolled – elementary | (2) Enrolled – preschool | (3) Enrolled – all types |
|---|---|---|---|
| 1st-stage | 0.162*** | -0.144*** | 0.018*** |
| | (8.00) | (-7.36) | (3.66) |
| Observations | 27,738 | 27,738 | 27,738 |
| $R^2$ | 0.24 | 0.23 | 0.04 |
| F-stat | 63.94 | 54.22 | 13.37 |
| p-value | 0.00 | 0.00 | 0.00 |

Notes: The table presents 1st stage coefficients from separate 2SLS regressions. The dependent variables are the dummies for whether the mother had a child in enrolled in preschool (Column 1), elementary school (Column 2), or any type of school (Column 3). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

These results validate that the reform induced increased participation of children in elementary schools in Brazil but reduced participation in preschools. The net impact on mothers' outcomes is, therefore, not obvious, which we analyze in the next subsection using second-stage as well as reduced form regressions.

## 6.3 Impact of school enrollment on mothers' labor market outcomes

Building on the first-stage findings, the second-stage analysis examines the relationship between school enrollment and maternal labor market outcomes. Table 3

shows the results for our second-stage regression, where we run the estimated preschool, elementary, and general school enrollment driven by the policy on all mother labor market outcomes. All estimates were calculated using year and state fixed effects, along with control variables including mothers' years of education, age, and race, and household information such as location and presence of relatives. We only observe a significant impact of being eligible for elementary enrollment on the number of weekly hours of housework.

The results of the first and second stages for elementary enrollment show an impact on the average time spent on housework. Mothers whose children become eligible for elementary school spend 3.63 hours less on housework per week. Having a child who stays at home instead of at school may not only impede the mother from working but can also mean more time spent on housework since the child requires more food and uses the house's resources more intensively.

Even though the first stage shows the reform had an impact on preschool and overall school enrollment, the second stage results presented no statistically significant impact on any variable. We continue our analysis focusing on elementary enrollment.

Table 3: Impact of enrollment on mothers' labor market outcomes (2001-2011)

|  | (1) LFP | (2) Work hours | (3) Wage/hour | (4) Formal work | (5) Job search | (6) House work |
|---|---|---|---|---|---|---|
| Elementary | 0.108 | 1.318 | -0.0517 | -0.00920 | 0.0410 | -3.639** |
|  | (1.64) | (0.29) | (-0.25) | (-0.10) | (1.28) | (-2.08) |
| Observations | 27,738 | 27,738 | 27,738 | 27,738 | 27,738 | 27,738 |
| $R^2$ | 0.04 | 0.06 | 0.04 | 0.16 | 0.01 | 0.06 |
| 1st stage F-stat | 63.94 | 63.94 | 63.94 | 63.94 | 63.94 | 63.94 |
| Preschool | -0.122 | -1.487 | 0.0583 | 0.0104 | -0.0462 | 4.106 |
|  | (-1.54) | (-0.39) | (0.17) | (0.13) | (-1.02) | (1.43) |
| Observations | 27738 | 27738 | 27738 | 27738 | 27738 | 27738 |
| $R^2$ | 0.04 | 0.06 | 0.04 | 0.16 | 0.01 | 0.06 |
| 1st stage F-stat | 54.22 | 54.22 | 54.22 | 54.22 | 54.22 | 54.22 |
| Any school | 0.952 | 11.58 | -0.454 | -0.0809 | 0.360 | -31.99 |
|  | (1.38) | (0.29) | (-0.25) | (-0.10) | (1.38) | (-1.50) |
| Observations | 27738 | 27738 | 27738 | 27738 | 27738 | 27738 |
| R2 | . | 0.06 | 0.04 | 0.16 | . | . |
| 1st stage F-stat | 13.37 | 13.37 | 13.37 | 13.37 | 13.37 | 13.37 |

Notes: The table presents 2nd stage coefficients from separate 2SLS regressions. The dependent variable are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is estimated enrollment driven by the policy (calculated on the 1st stage, see Table 2). All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Using a Differences-in-Differences (DiD) approach (See equation 3), we examine the reduced form relationship between policy reform and maternal labor outcomes, which indicates intent-to-treat effects (See Table 4). All the estimates were calculated with year and state fixed effects, and all the control variables.

The results for the DiD on Table 4 show a small effect on labor force participation. Treated mothers have a 1.76 percentage point increase in the probability of being employed. Eligibility for the reform had a small impact on hours spent on chores at home: mothers of eligible children spent 0.59 fewer hours on housework per week. While significant at 10%, the effect only represents a reduction of approximately 40 minutes per week.

The remaining outcomes show no impact. We found no evidence of the reform on the strength of mothers' attachment to the workforce. The reduction of hours spent on housework did not translate into longer working hours or transition into positions that require longer shifts. In addition, mothers whose children became eligible for elementary school admission did not show a statistically significant change in their wages per hour.

The lack of translation into a more permanent attachment [7] to the labor market can also be seen in the absence of a statistically significant change in the probability of having formal work. The absence of an effect on formal employment may be related to the service sector, a major employer of women in the country, and a sector characterized by a high prevalence of informal work.

We do not observe the impact of compulsory education on the probability of looking for a job. Women have a lower probability of engaging in job search in general (Del Boca, 2015), especially if they are married, which is the case for the majority of our sample[8].

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[7]We consider having more working hours and having a formal contract as indicators of a more permanent attachment to the workforce.

[8]Our data does not include information on the status of the relationship between the head of the house and their partner, but we infer that couples who live together are similar to married couples. 80% of our sample consists of women who cohabit with their partners.

Table 4: Impact of elementary enrollment on mothers' labor market outcomes (2001-2011)

| | (1) LFP | (2) Work hours | (3) Wage/hour | (4) Formal work | (5) Job search | (6) House work |
|---|---|---|---|---|---|---|
| Eligible mothers | 0.0176* | 0.214 | -0.00838 | -0.00149 | 0.00664 | -0.590* |
| | (1.74) | (0.30) | (-0.24) | (-0.10) | (1.34) | (-1.95) |
| Observations | 27,738 | 27,738 | 27,738 | 27,738 | 27,738 | 27,738 |
| $R^2$ | 0.05 | 0.06 | 0.04 | 0.16 | 0.01 | 0.06 |
| Mean of dep. var. (pre-treatment) | 0.721 | 24.143 | 0.469 | 0.258 | 0.070 | 28.466 |

Notes: The table presents coefficients from separate OLS regressions, our reduced-form approach. The dependent variable are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

## 6.4 Robustness

Our main specification consisted of mothers of only one child. We relax this restriction to see how it affects results. The first two columns of Table 5 are the results for mothers who had multiple children and at least one child between 5 and 7 years old. These mothers could have 1, 2, 3, or any number of children. For mothers with any number of children, school enrollment does not significantly affect labor force participation, nor lead to a reduction in time spent on housework.

In theory, a mother's youngest child is the one who requires the most care, and we test this hypothesis. Columns 3 and 4 of Table 5 present the results for mothers of multiple children whose youngest child is either 5, 6, or 7 years old. For mothers whose youngest is in our target age, the effects on labor force participation remain insignificant. Ryu (2020) also found no significant effects of a reform of preschool eligibility rules on any maternal labor market outcomes if the child was not the youngest member of the household or there were other relatives present.

On the other hand, there is still a statistically significant effect on housework time, although the reduction is small. Mothers whose youngest child is either 5, 6, or 7 decrease their housework time by 0.31 hours, similar to the main specification. This could indicate that the affected child, being the youngest, is more impactful than the number of children for hours of housework.

Table 5: Impact of elementary enrollment on mothers' labor market outcomes on alternative samples (2001-2011)

| | Multiple children | | Youngest child | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| | LFP | House work | LFP | House work |
| Eligible mothers | -0.00336 | -0.248 | -0.00340 | -0.313* |
| | (-0.77) | (-1.28) | (-0.59) | (-1.92) |
| Observations | 133,403 | 133,403 | 96,328 | 96,328 |
| $R^2$ | 0.05 | 0.06 | 0.04 | 0.05 |
| Mean of dep. var. (pre-treatment) | 0.634 | 32.95 | 0.633 | 33.07 |

Notes: The table presents coefficients from separate OLS regressions. The first two columns show results for regression on a sample of mothers of multiple children, while the last two columns show results for mothers whose youngest children were 5, 6, or 7 years old. The dependent variables were a dummy for whether the mother participated in the labor force (Columns 1 and 3) and the number of hours spent on housework per week (Columns 2 and 4). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

We define treatment as having a child who starts the year at 5 or 6 years old. To test the robustness of this specification, we ran regressions assigning treatment to children who start the year at 4 years old, which should be unaffected by the school reform. Table 6 shows results for the first-stage regression of the reform on enrollment and finds positive effects only for preschool. This is expected since preschool went from offering spots to 4-6 year old children to just 4 and 5 year olds, which could have increased spot availability for children who start the year as 4-year-olds.

Table 7 presents second-stage results for all outcomes on elementary and preschool. We see no effects of the reform on any of the outcomes. The reduced form results are present on Table 8 and show no effects besides a small effect on formal employment. Eligible mothers were 0.9 percentage points more likely to be in formal employment. These results strengthen the argument of our treatment assignment. Table 18 on the Appendix shows another treatment comparison, where we set a reduced date of birth window to assign treatment. It also provides results similar to our main specification.

Table 6: Impact of the reform on school enrollment (2001-2011) 4-year-olds

|  | (1) Enrolled - elementary | (2) Enrolled - preschool |
|---|---|---|
| 6-years after the reform | 0.00636 | 0.0572*** |
|  | (1.25) | (4.00) |
| Observations | 19389 | 19389 |
| R2 | 0.90 | 0.40 |
| F-stat | 1.56 | 16.01 |
| p-value | 0.22 | 0.00 |

Notes: The table presents 1st stage coefficients from separate 2SLS regressions for a different sample with 4-year-olds as treated. The dependent variables are the dummies for whether the mother had a child in enrolled in preschool (Column 1) and elementary school (Column 2). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Impact of elementary enrollment on mothers' labor market outcomes (2001-2011) 4-year-olds

|  | (1) LFP | (2) Work hours | (3) Wage per hour | (4) Formal work | (5) Job search | (6) House work |
|---|---|---|---|---|---|---|
| Enrollment - elementary | 0.133 | -127.9 | -5.635 | -2.739 | 1.530 | 55.09 |
|  | (0.06) | (-0.71) | (-0.56) | (-0.87) | (0.90) | (0.71) |
| Observations | 19389 | 19389 | 19389 | 19389 | 19389 | 19388 |
| R2 | 0.05 | . | . | . | . | . |
| Enrollment - preschool | 0.0148 | -14.23 | -0.627 | -0.305 | 0.170 | 6.116 |
|  | (0.07) | (-1.30) | (-0.63) | (-1.36) | (1.31) | (0.76) |
| Observations | 19389 | 19389 | 19389 | 19389 | 19389 | 19388 |
| R2 | 0.06 | . | 0.04 | 0.08 | . | 0.04 |

Notes: The table presents 2nd stage coefficients from separate 2SLS regressions for a different sample with 4-year-olds as treated. The dependent variables are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is estimated enrollment driven by the policy (calculated on the 1st stage, see Table 2). All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Impact of elementary enrollment on mothers' labor market outcomes (2001-2011) 4 years

|  | (1)<br>LFP | (2)<br>Work hours | (3)<br>Wage/hour | (4)<br>Formal work | (5)<br>Job search | (6)<br>House work |
|---|---|---|---|---|---|---|
| Eligible mothers | 0.000848 | -0.813 | -0.0358 | -0.0174 | 0.00973* | 0.350 |
|  | (0.06) | (-0.95) | (-0.61) | (-1.16) | (1.89) | (0.78) |
| Observations | 19389 | 19389 | 19389 | 19389 | 19389 | 19388 |
| R2 | 0.05 | 0.07 | 0.05 | 0.15 | 0.01 | 0.07 |

Notes: The table presents coefficients from separate OLS regressions, our reduced-form approach, for a different sample with 4-year-olds as treated. The dependent variable are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

## 6.5   Heterogeneities

To better understand the results, we use the reduced-form to analyze the effects of the reform by different mothers' characteristics. We decompose our results by characteristics of our sample of mothers to understand how different mothers were impacted by the reform. Since the impacts were only observed for the rate of labor force participation and average time spent on housework, we continue our inquiries by looking at these two variables.

Table 9 shows the effects of the reform based on the race of the mothers. Columns 1 and 2 show the results on LFP and housework for White mothers, respectively, and columns 3 and 4 for Black mothers. While PNAD registers race in other categories such as Yellow and Brown (*pardo* in Portuguese), we grouped Yellow to White and Brown to Black due to the low number of observations and similarities with the biggest categories.

Black mothers are the only ones affected by the policy: eligible mothers are 4.78 percentage points more likely to participate in the labor force than ineligible ones. Time spent on housework seems to be negatively impacted, but shows no statistical significance for both White and Black mothers. The 1st stage regression results on Table **??** show that net school enrollment for Black mothers was similar in percentage points to all mothers, suggesting that the difference in impact on LFP is not due to bigger school enrollment.

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Table 9: Impact of elementary enrollment on mothers' labor market outcomes by race (2001-2011)

|  | White | | Black | |
|---|---|---|---|---|
|  | (1) | (2) | (3) | (4) |
|  | LFP | House work | LFP | House work |
| Eligible mothers | -0.00311 | -0.354 | 0.0478** | -0.847 |
|  | (-0.24) | (-1.13) | (2.53) | (-1.34) |
| Observations | 14,335 | 14,335 | 13,403 | 13,403 |
| $R^2$ | 0.05 | 0.08 | 0.05 | 0.05 |
| Mean of dep. var. (pre-treatment) | 0.728 | 28.170 | 0.712 | 28.815 |

Notes: The table presents coefficients from separate OLS regressions. The first two columns show results for regression on a restricted sample of White mothers, while the last two columns show results for Black mothers. The dependent variables were a dummy for whether the mother participated in the labor force (Columns 1 and 3) and the number of hours spent on housework per week (Columns 2 and 4). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
$^*$ $p < 0.10$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$

While there is no municipal information on the data, we can analyze if the household is at a rural or urban area. Table 10 shows the reduced form results distinguished by mothers' location of residence. Columns 1 and 2 report the effects on LFP and housework time for urban mothers, while the corresponding columns for rural mothers follow.

For mothers who live in urban areas, there is no significant change in labor force participation or effect on housework time. In contrast, rural mothers experience a significant increase in the likelihood of labor force participation of 6.93 percentage points. The labor force participation of mothers living in rural areas, 75%, was already bigger than that of mothers living in urban areas, 71% (See the last row on Table 10). Preschool enrollment is higher amongst urban mothers: 32% compared to 27% of rural mothers. Looking at the entire sample, 57% of 5-year-old children of rural mothers are enrolled in preschool, compared to 72% of urban children.

Table 20 in the Appendix presents first-stage regression results for rural mothers: rural eligible children were 7.4 p.p. more likely to be in school due to the reform. This value is more than 4 times bigger than the estimate for all mothers (See Column "Any school" on Table 2). This suggests that the bigger impact on rural mothers derives from the higher school enrollment, which could have freed up rural mothers and allowed for a greater effect on labor force participation.

Table 10: Impact of elementary enrollment on mothers' labor market outcomes by location (2001-2011)

|  | Urban | | Rural | |
|---|---|---|---|---|
|  | (1) | (2) | (3) | (4) |
|  | LFP | House work | LFP | House work |
| Eligible mothers | 0.0119 | -0.454 | 0.0693** | -1.763 |
|  | (1.10) | (-1.41) | (2.07) | (-1.57) |
| Observations | 24,649 | 24,649 | 3,089 | 3,089 |
| R² | 0.05 | 0.06 | 0.09 | 0.07 |
| Mean of dep. var. (pre-treatment) | 0.7166 | 28.1306 | 0.7563 | 31.4621 |

Notes: The table presents coefficients from separate OLS regressions. The first two columns show results for regression on a restricted sample of urban mothers, while the last two columns show results for rural mothers. The dependent variables were a dummy for whether the mother participated on the labor force (Columns 1 and 3) and the number of hours spent on housework per week (Columns 2 and 4). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

We also examine whether the effects of the reform vary by mothers' educational attainment. We define high-skilled mothers as those who have at least a high school diploma or higher, and low-skilled mothers as the remaining ones. Table 11 shows the results of the reduced form regression. Columns 1 and 2 report the effects on LFP and housework time for high-skilled mothers, while columns 3 and 4 present the results for low-skilled mothers. The results suggest that education level did not significantly influence the transmission of the reform's effects: for both skill groups, the estimated impacts on labor and housework outcomes are not statistically significant at conventional levels.

Table 11: Impact of elementary enrollment on mothers' labor market outcomes by skill levels (2001-2011)

|  | High | | Low | |
|---|---|---|---|---|
|  | (1) | (2) | (3) | (4) |
|  | LFP | House work | LFP | House work |
| Eligible mothers | 0.0317 | -0.868 | 0.00820 | -0.426 |
|  | (1.63) | (-1.22) | (0.52) | (-0.88) |
| Observations | 11,805 | 11,805 | 15,933 | 15,933 |
| $R^2$ | 0.04 | 0.05 | 0.05 | 0.08 |
| Mean of dep. var. (pre-treatment) | 0.7705 | 26.5972 | 0.6864 | 29.7484 |

Notes: The table presents coefficients from separate OLS regressions. The first two columns show results for regression on a restricted sample of more highly educated mothers, while the last two columns show results for less educated mothers. The dependent variables were a dummy for whether the mother participated in the labor force (Columns 1 and 3) and the number of hours spent on housework per week (Columns 2 and 4). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Brazil's five regions — North, Northeast, Southeast, South, and Center-West — display significant socioeconomic disparities that shape children's education and maternal employment. During our period of analysis, while there was a regional convergence of income (Monteiro Neto, 2014), the more developed Southeast and South still had better infrastructure, higher incomes, and broader access to education and childcare services (Loyola, 2017). In contrast, the North and Northeast faced greater poverty and limited public service provision. The Center-West experienced growth but remained uneven in access (Loyola, 2017). These regional differences could affect how the reform impacted families across the country. There are municipal differences that would be important to the analysis, but our dataset does not provide more granular information on location

Table 12 presents the impact of elementary school enrollment on mothers' labor market outcomes across Brazil's regions. Panel A reports the effects on LFP, while Panel B examines changes in time spent on housework. Only mothers living in the South and Midwest regions showed effects of the reform on LFP. However, the effects were in opposite directions. In the South, mothers experienced a significant increase of 3.75 percentage points in labor force participation. In the Midwest, the likelihood of a mother being employed declined by 4.29 percentage points. This suggests that, in the Midwest region, school enrollment may not be sufficient to enable greater employment for mothers, possibly due to limited job opportunities or other factors affecting

mothers' work decisions. This could also be related to the change in enrollment between preschool and elementary school in the region. If kids spent more time per day at preschool, the change into elementary could end up reducing mothers' available time and have negative impacts on employment. Moreover, region-specific policies could be affecting mothers' labor force participation and housework time.

In the South, where labor markets are more developed and formal employment opportunities are more accessible, mothers may be more likely to translate time savings from children's school enrollment into labor market participation. The region also has an impact on housework: mothers from the South reduce their time spent on housework by 1.09 hours. Effects for the other regions are not significant.

Table 12: Impact of elementary enrollment on mother's labor market outcomes by region (2001-2011)

|  | (1) North | (2) Northeast | (3) Southeast | (4) South | (5) Midwest |
|---|---|---|---|---|---|
| **Panel A: LFP** | | | | | |
| Eligible mothers | -0.0166 | 0.0354 | -0.0108 | 0.0375* | -0.0429** |
| | (2.36) | (1.59) | (-0.42) | (0.14) | (-4.20) |
| Observations | 2872 | 7892 | 5383 | 8616 | 2975 |
| $R^2$ | 0.06 | 0.04 | 0.03 | 0.04 | 0.04 |
| Mean of dep. var. (pre-treat.) | 0.7292 | 0.6874 | 0.7844 | 0.6962 | 0.7495 |
| | | | | | |
| **Panel B: House work** | | | | | |
| Eligible mothers | -0.0156 | -0.961 | 0.319 | -1.095*** | 1.209 |
| | (-0.01) | (-1.21) | (0.74) | (-7.36) | (1.21) |
| Observations | 2,872 | 7,892 | 5,383 | 8,616 | 2,975 |
| $R^2$ | 0.06 | 0.04 | 0.05 | 0.05 | 0.04 |
| Mean of dep. var. (pre-treat.) | 24.5228 | 30.1034 | 27.3788 | 29.7527 | 25.6702 |

Notes: The table presents coefficients from separate OLS regressions. The columns show separate results for mothers of each of Brazil's regions: North, Northeast, Southeast, South, and Midwest. Panel A shows the dependent variable of a dummy for whether the mother participated in the labor force, and Panel B shows the dependent variable of the number of hours spent on housework per week. The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

The results differ when we divide the sample by certain groups. This heterogeneity analysis suggests that overall effects on all women seem to be driven by specific groups of women, an important finding for policy design.

## 6.6 Impact on lone mothers' labor market outcomes

Since lone mothers represent a unique group of mothers, we ran the regression separately for lone mothers to try and isolate the effects on this subgroup. Table 13 presents the estimated effects of elementary school enrollment on all original labor market outcomes for lone mothers. The findings indicate that, unlike the general population of mothers, lone mothers experience a decline in labor market engagement due to the reform. Lone mothers' eligible children experience a decrease of 6.71 p.p. in their likelihood of being employed. Similarly, they reduce their weekly work hours by approximately 5.33 hours. Unlike the general results for the main specification and other subgroups, there is no effect on time spent on housework. We do not observe a shift from working hours to housework hours.

The results show that rather than increasing their labor supply when their child starts school, lone mothers may be adjusting their working conditions, possibly shifting toward jobs with fewer hours or leaving the labor market entirely. As supported by Baker, Gruber, and Milligan (2008), these women could have been affected and supported by welfare and child benefit policies during the period.

It is worth noting that lone mothers have the highest labor force participation in our entire sample, with 87% employed compared to 72% in the main sample (see Table 4). They work longer hours before the reform, averaging 31.61 hours per week, whereas mothers in the main sample work 24.14 hours on average. Furthermore, lone mothers dedicate less time to housework, spending 22.28 hours per week compared to 28.46 hours among all mothers in the sample. Reyna and Costa Silva (2021) also found a negative effect of eligibility on LFP or weekly work hours of lone mothers, but the effect was not statistically significant. The authors attribute the lack of effect to lone mothers being the main source of income and already needing to work before the reform.

Table 13: Impact of elementary enrollment on lone mothers' labor market outcomes (2001-2011)

|  | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
|  | LFP | Work hours | Wage/hour | Formal work | Job search | House work |
| Eligible mothers | -0.0671* | -5.338** | 0.250 | -0.00827 | 0.0250 | -0.288 |
|  | (-1.88) | (-2.51) | (1.62) | (-0.16) | (0.91) | (-0.17) |
| Observations | 3,341 | 3,341 | 3,341 | 3,341 | 3,341 | 3,341 |
| $R^2$ | 0.06 | 0.05 | 0.05 | 0.19 | 0.02 | 0.08 |
| Mean of dep. var. (pre-treatment) | 0.8707 | 32.0395 | 0.6245 | 0.3279 | 0.0558 | 22.2857 |

Notes: The table presents coefficients from separate OLS regressions, our reduced-form approach, for a restricted sample composed of lone mothers. The dependent variable are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

## 6.7 Impact on fathers' labor market outcomes

The main hypothesis of the paper is that mothers are differently affected than fathers by childcare policies since they are the primary carers and are constrained by the double workday (Reyna; Costa Silva, 2021). We run the reduced-form regressions for fathers to test this hypothesis.

Table 14 shows fathers' characteristics by treatment. Fathers are more than 10 years older than mothers. They are more attached to the labor market: fathers' LFP is much higher, they work longer hours, and are more formally employed. They spend fewer hours on housework than mothers.

Table 14: Comparison of Treatment and Control Groups for Fathers (2001-2011)

|  | Control | Treatment | Difference |
|---|---|---|---|
| Age | 35.25 | 33.80 | 1.44*** |
| White | 0.49 | 0.50 | -0.008 |
| Urban | 0.88 | 0.87 | 0.009* |
| Relatives Present | 0.07 | 0.07 | -0.003 |
| Num. People on the Household | 3.10 | 3.12 | -0.02*** |
| Labor Force Participation | 0.97 | 0.97 | -0.00 |
| Work Hours | 43.97 | 44.22 | -0.25 |
| Formal Work | 0.53 | 0.51 | 0.01** |
| Job Search | 0.01 | 0.00 | 0.001 |
| Housework | 5.51 | 5.62 | -0.10 |
| Child in Preschool | 0.008 | 0.45 | -0.44*** |
| Child in Elementary | 0.96 | 0.49 | 0.47*** |
| Child in School | 0.97 | 0.94 | 0.02*** |
| Sample size | 6,656 | 15,850 | |

Notes: The table compares the characteristics of fathers in the Control group, i.e., with children who start the year as seven-year-olds, and Treatment group, i.e., with children who start the year as six or five-year-olds. Column "Control" presents the average control fathers' characteristics, column "Treatment" reports the average characteristics for treated fathers, and column "Difference" reports the differences between both groups and the t-test statistical significance of this result. The variables analyzed are Dummy for whether the father is White; Dummy for whether the father lives in an urban area; Dummy for whether the father is a lone parent; Dummy for whether other relatives cohabit in the household; Number of people on the household; Dummy for whether the father participates on the labor force; Number of hours worked on the labor market; Dummy for whether the father had a formal work attachment; Dummy for whether the father was looking for a job; Percentage of mothers whose children were in preschool, elementary school or any type of school.
P-values: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 15 shows the results for our reduced-form regression for all labor market outcomes for fathers. The effects of the reform were insignificant in all variables besides the probability of having formal work. Eligible fathers were 3.38 p.p. less likely to be in formal work. Having a child in school could reduce fathers' demand for formal work dedication due to the reduction of childcare costs. Moreover, the slight increase in maternal labor market participation could allow men to change to less formal employment. Before the reform, men already represented a bigger share of formal work than women: 44.9% compared to 25.8%.

Table 15: Impact of elementary enrollment on fathers' labor market outcomes (2001-2011)

| | (1) LFP | (2) Work hours | (3) Wage/hour | (4) Formal work | (5) Job search | (6) House work |
|---|---|---|---|---|---|---|
| Eligible fathers | 0.00312 | -0.263 | 0.0186 | -0.0338*** | -0.00182 | 0.170 |
| | (0.74) | (-0.65) | (0.40) | (-3.45) | (-0.79) | (0.83) |
| Observations | 22,506 | 22,506 | 22,506 | 22,506 | 22,506 | 22,506 |
| $R^2$ | 0.01 | 0.01 | 0.04 | 0.19 | 0.01 | 0.02 |
| Mean of dep. var. (pre-treatment) | 0.979 | 44.71 | 0.865 | 0.449 | 0.010 | 5.68 |

Notes: The table presents coefficients from separate OLS regressions, our reduced-form approach, for the sample of fathers. The dependent variable are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

# 7 Final Remarks

This study investigates the impact of the education reform that lowered the entry age in Brazil on mothers' labor market outcomes. We used PNAD data and applied both a Local Average Treatment Effect (LATE) and a differences-in-differences methodology. We find that lowering the mandatory school age in Brazil had impacts on eligible children's enrollment, increasing the probability of their elementary enrollment and reducing preschool enrollment. The net overall school enrollment increase was 1.8 percentage points. Differences-in-differences results showed that school enrollment had small impacts on mothers' labor market participation and housework provision. We find that it increased the likelihood of labor force participation by 1.7 percentage points and reduced time spent on housework by 0.59 hours per week. Our heterogeneity analysis reveals that the impact of the reform was not evenly distributed for all types of mothers: Black and rural mothers experienced the most pronounced gains in labor force participation, 4.78 and 6.93 percentage points, respectively.

These highlight the importance of considering intersectional factors such as race, geographic location, and educational attainment in policy design. The small magnitude of overall effects also shows that while child caring is a relevant aspect of mothers' labor market decisions, it is intertwined with other variables that may hinder labor market participation.

Establishing causality posed a significant challenge in this study. Therefore, future research should explore how other changes within the educational system may have influenced children's school enrollment, especially the 2009 reform. Another important next step is to analyze heterogeneous effects by survey year, since implementation of the reform may have been more loosely enforced at the beginning.

Despite those limitations, our study's results highlight the potential and constraints of policies aimed at expanding access to formal education as tools for reducing the gender gap in both paid and unpaid labor. These findings contribute to the growing body of literature on the interaction between childcare availability and maternal labor market outcomes, offering insights from a developing country context.

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# A   Appendix

Table 16: Impact of elementary enrollment on mothers' labor market outcomes - public school (2001-2011)

|  | (1) LFP | (2) House work |
|---|---|---|
| Eligible mothers | 0.0000447 | 0.615 |
|  | (0.00) | (0.90) |
| Observations | 13129 | 13129 |
| $R^2$ | 0.05 | 0.06 |
| Mean | 0.7199017 | 29.65953 |

Notes: The table presents coefficients from separate OLS regressions on a restricted sample of mothers whose children were enrolled in public school. The dependent variables were a dummy for whether the mother participated in the labor force (Column 1) and the number of hours spent on housework per week (Column 2). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 17: Impact of elementary enrollment on mothers' labor market outcomes - relative present (2001-2011)

|  | (1) LFP | (2) House work |
|---|---|---|
| Eligible mothers | 0.0245 | -1.539 |
|  | (0.76) | (-0.77) |
| Observations | 2517 | 2517 |
| $R^2$ | 0.04 | 0.06 |
| Mean | 0.7538803 | 24.63193 |

Notes: The table presents coefficients from separate OLS regressions on a restricted sample of mothers who have relatives cohabiting the household. The dependent variables were a dummy for whether the mother participated in the labor force (Column 1) and the number of hours spent on housework per week (Column 2). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
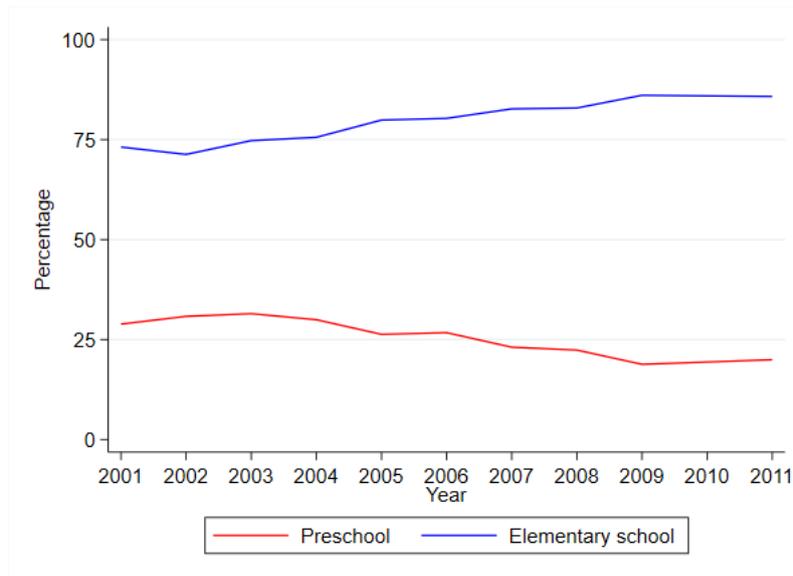* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

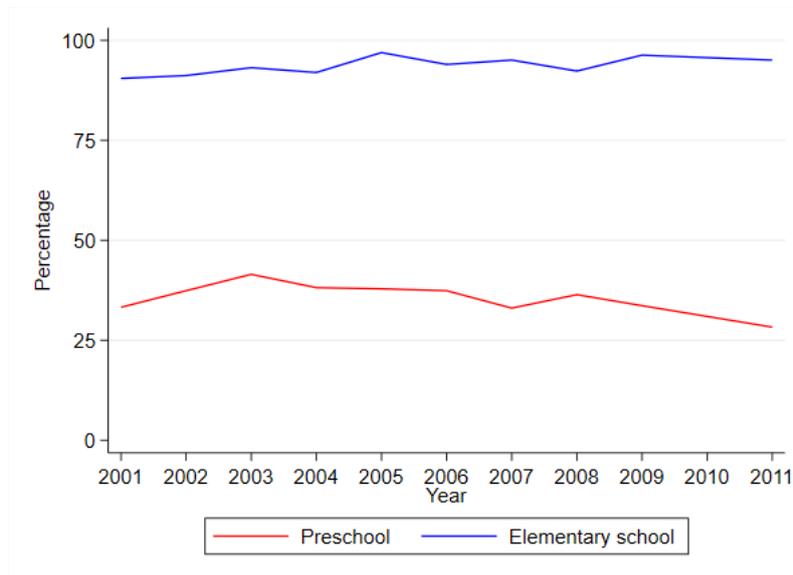Figure 5: School enrollment of children that started the year as 5-year-olds (lone mothers), 2001-2011
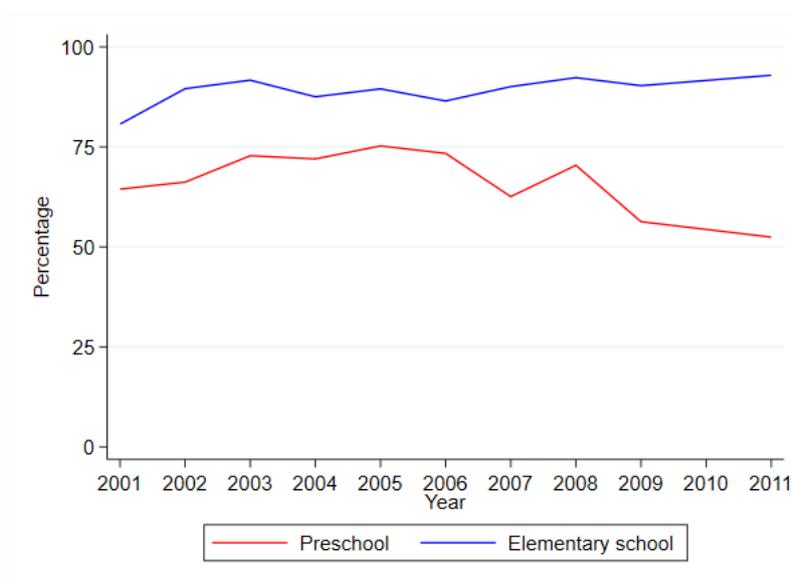


Notes: Percentage of children of lone mother who started the year as 5-year-olds by school level from 2001 to 2011. Data from PNAD.

Figure 6: School enrollment of children that started the year as 6-year-olds (lone mothers), 2001-2011



Notes: Percentage of children of lone mothers who started the year as 6-year-olds by school level from 2001 to 2011. Data from PNAD.

Figure 7: School enrollment of children that started the year as 7-year-olds (lone mothers), 2001-2011



Notes: Percentage of children of lone mothers who started the year as 7-year-olds by school level from 2001 to 2011. Data from PNAD.

Table 18: Impact of elementary enrollment on mothers' labor market outcomes by different windows (2001-2011)

|  | (1) LFP | (2) Work hours | (3) Wage/hour | (4) Formal work | (5) Job search | (6) House work |
|---|---|---|---|---|---|---|
| Panel A: main specification |  |  |  |  |  |  |
| Eligible mothers | 0.0176* | 0.214 | -0.00838 | -0.00149 | 0.00664 | -0.590* |
|  | (1.74) | (0.30) | (-0.24) | (-0.10) | (1.34) | (-1.95) |
| Observations | 27738 | 27738 | 27738 | 27738 | 27738 | 27738 |
| $R^2$ | 0.05 | 0.06 | 0.04 | 0.16 | 0.01 | 0.06 |
| Panel B: kids born from May 1st until April 30th |  |  |  |  |  |  |
| Eligible mothers | 0.00644 | -0.0900 | -0.0101 | -0.0168 | -0.000185 | -0.761** |
|  | (0.76) | (-0.15) | (-0.23) | (-1.13) | (-0.03) | (-2.11) |
| Observations | 27738 | 27738 | 27738 | 27738 | 27738 | 27738 |
| $R^2$ | 0.05 | 0.06 | 0.04 | 0.16 | 0.01 | 0.06 |

Notes: The table compares coefficients from separate OLS regressions, our reduced-form approach, for different samples of mothers. Panel A shows results for our main specification of mothers and panel B for mothers of children born from May st until April 30th. The dependent variables are a dummy for whether the mother participated on the labor force (Column 1), number of hours worked on the labor market per week (Column 2), wage per hour (Column 3), dummy whether the mother was formally employed (Column 4), dummy whether the mother was looking for a job (Column 5), number of hours spent on housework per week (Column 6). The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. All regressions include mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 19: Impact of the reform on school enrollment, all types for Black mothers (2001-2011)

|  | (1) Enrolled in school - any level |
|---|---|
| Child at school | 0.0173** |
|  | (2.65) |
| Observations | 13403 |
| R2 | 0.04 |
| F-stat | 7.04 |
| p-value | 0.01 |

Notes: The table presents 1st stage coefficients from a 2SLS regression on a restricted sample composed of Black mothers. The dependent variable is a dummy for whether the mother had a child in enrolled in any type of school. The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. The regression includes mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 20: Impact of the reform on school enrollment, all types for rural mothers (2001-2011)

|  | (1) Enrolled in school - any level |
|---|---|
| Child at school | 0.0742*** |
|  | (3.36) |
| Observations | 3089 |
| R2 | 0.08 |
| F-stat | 11.27 |
| p-value | 0.00 |

Notes: The table presents 1st stage coefficients from a 2SLS regression on a restricted sample composed of rural mothers. The dependent variable is a dummy for whether the mother had a child in enrolled in any type of school. The key explanatory variable is a dummy indicating whether the mother had a treated or control child due to the reform. The regression includes mothers' covariates: age, race, years of education, number of people in the household, presence of a relative in the home, and urban/rural location. All estimates include year and state fixed effects. Standard errors were clustered at the state level. $t$ statistics in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 21: Comparison of Treatment and Control Groups (2001-2006) Before the reform

|  | Control | Treatment | Difference |
|---|---|---|---|
| Age | 32.96 | 30.97 | 1.99*** |
| White | 0.56 | 0.54 | 0.02*** |
| Urban | 0.89 | 0.89 | 0.005 |
| Lone Mother | 0.05 | 0.15 | -0.10*** |
| Relatives Present | 0.08 | 0.09 | -0.00 |
| Num. People on the Household | 2.98 | 3.02 | -0.03*** |
| Labor Force Participation | 0.73 | 0.72 | 0.01* |
| Work Hours | 25.15 | 24.15 | 0.99** |
| Formal Work | 0.26 | 0.25 | 0.00 |
| Job Search | 0.06 | 0.07 | -0.005 |
| Housework | 27.90 | 28.19 | -0.29 |
| Child in Preschool | 0.009 | 0.52 | -0.51*** |
| Child in Elementary | 0.95 | 0.40 | 0.55*** |
| Child in School | 0.96 | 0.93 | 0.03*** |
| Sample size | 3,873 | 9,310 | |

Notes: The table compares the pre-reform characteristics of mothers in the Control group, i.e., with children who start the year as seven-year-olds, and Treatment group, i.e., with children who start the year as six or five-year-olds. Column "Control" presents the average control mothers' characteristics, column "Treatment" reports the average characteristics for treated mothers, and column "Difference" reports the differences between both groups and the t-test statistical significance of this result. The variables analyzed are Dummy for whether the mother is White; Dummy for whether the mother lives in an urban area; Dummy for whether the mother is a lone parent; Dummy for whether other relatives cohabit in the household; Number of people on the household; Dummy for whether the mother participates on the labor force; Number of hours worked on the labor market; Dummy for whether the mother had a formal work attachment; Dummy for whether the mother was looking for a job; Percentage of mothers whose children were in preschool, elementary school or any type of school.

P-values: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 22: Comparison of Treatment and Control Groups for Rural Mothers (2001-2011)

|  | Control | Treatment | Difference |
|---|---|---|---|
| Age | 32.15 | 30.12 | 2.02*** |
| White | 0.50 | 0.45 | 0.04* |
| Lone Mother | 0.02 | 0.05 | -0.04*** |
| Relatives Present | 0.05 | 0.07 | -0-.01 |
| Num. People on the Household | 3.04 | 3.06 | -0.01 |
| Labor Force Participation | 0.74 | 0.74 | -0.001* |
| Work Hours | 20.7 | 20 | 0.7 |
| Formal Work | 0.13 | 0.13 | -0.002 |
| Job Search | 0.03 | 0.04 | -0.005 |
| Housework | 29.9 | 30.36 | -0.46 |
| Child in Preschool | 0.009 | 0.37 | -0.36*** |
| Child in Elementary | 0.94 | 0.49 | 0.44*** |
| Child in School | 0.95 | 0.87 | 0.08*** |
| Sample size | 871 | 2,218 | |

Notes: The table compares the characteristics of rural mothers in the Control group, i.e., with children who start the year as seven-year-olds, and Treatment group, i.e., with children who start the year as six or five-year-olds. Column "Control" presents the average control rural mothers' characteristics, column "Treatment" reports the average characteristics for treated mothers, and column "Difference" reports the differences between both groups and the t-test statistical significance of this result. The variables analyzed are Dummy for whether the mother is White; Dummy for whether the mother lives in an urban area; Dummy for whether the mother is a lone parent; Dummy for whether other relatives cohabit in the household; Number of people on the household; Dummy for whether the mother participates on the labor force; Number of hours worked on the labor market; Dummy for whether the mother had a formal work attachment; Dummy for whether the mother was looking for a job; Percentage of mothers whose children were in preschool, elementary school or any type of school.

P-values: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$