Female mayors and child health*

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Abstract

Millions of children continue to die in low- and middle-income countries. A large share of these deaths is considered preventable, for instance through better access to basic healthcare services. Growing evidence suggests that greater female political representation could improve child health by ensuring public good provision. We analyze close elections between male and female mayoral candidates in Brazil to assess the role of women as political leaders on child health. Sharp regression discontinuity design estimates indicate that marginally electing a woman as mayor has no statistically significant effect on child death in the whole sample. However, we identify a decrease in mortality in municipalities with a higher share of legislative power held by women, and where most of the population is illiterate. As results contrast with existing findings, we conclude that increasing women's political representation might not necessarily represent a viable avenue for child health promotion in middle-income countries.

JEL classifications: D72, I15, I18, J16
Keywords: Child mortality, gender, women's political representation Brazil, South America

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1 Introduction

Women's political leadership is increasingly acknowledged as playing a key role in shaping social and economic development, and enhancing women's political participation (Hessami and da Fonseca, 2020). Greater political involvement of women has been shown to improve public good provision in education (Clots-Figueras, 2012), health (Bhalotra and Clots-Figueras, 2014), and infrastructure (Chattopadhyay and Duflo, 2004), as well as institutional quality and corruption (Brollo and Troiano, 2016; Afridi et al., 2017), particularly in developing countries.[1] In addition, higher female political representation was shown to induce woman empowerment – greater participation of women in politics (Brown et al., 2022), education (Clots-Figueras, 2012), or household decision-making (Bargain et al., 2019), and lower gender-based violence (Delaporte and Pino, 2022) – policy goals that are considered means to unlock greater development potentials in low- and middle-income countries (Duflo, 2012).

Besides public good provision and institutional quality, empirical works on the role of women's political leadership have focused on its effects on health – specifically, child and maternal health. Miller (2008), for instance, evidenced that the introduction of women's suffrage at the state-level in the United States (US) led to substantial reductions in child mortality, as a result of increases in state's public health spending. Bhalotra and Clots-Figueras (2014) estimated decreased neonatal mortality as a result of minimum quota laws in India, that raised female representation shares in state legislatures. Exploiting a similar setting, Kalsi (2017), that higher shares of female state legislators reduced child mortality rates for girls, while increasing death of first-born boys. Bhalotra et al. (2022) found that, across the globe, raising the share of women in parliament gender quotas reduced maternal mortality, but not infant mortality, likely because of greater skilled birth

[1] This relation is often explained by the fact that women, and their elected representatives, display greater preferences for redistributive policies compared to men. Indeed, women were found to be more altruistic (Eckel and Grossman, 2009), inequality-averse (Andreoni and Vesterlund, 2001), or more involved in family and children matters, such as health or education (Thomas, 1991), than men. Provided that politicians mirror the citizens they represent, it is expected that they would behave accordingly to citizens' interests, which would, in turn, influence policy choices. And, given the fact that individual preferences are known to influence policy (Besley and Coate, 1997), the gender of the elected representative should influence political decisions. A greater share of women holding executive or legislative power is thus expected to raise attention to women's preferential policy matters.

attendance and prenatal care utilization.

While millions of children continue to die in low and middle-income countries, children's health is commonly used as indicator of social and economic development,[2] its importance being confirmed by its inclusion as a sustainable development goal.[3] A large share of these deaths is, nonetheless, considered preventable by implementing low-cost interventions such as better access to basic healthcare services (Cutler and Lleras-Muney, 2006). The interest in the plausibly positive impacts of women's policy-making on child health is explained by the fact that child health and mortality are strongly related to the quality and availability of healthcare services, access to nutrition, and environmental risks – access to safe water and sanitation (Black et al., 2017) – and that female leaders could trigger action by improving the provision of these public goods, the primary policy challenge lying in addressing early life health and survival of children in poor families.[4]

In this article, we analyze Brazilian mayoral elections between 2000 and 2019 using a sharp regression discontinuity (RD) design. Municipality data on child health outcomes are matched to electoral data and mayoral characteristics, focusing on those municipalities with mixed-gender races. By comparing municipalities where a woman 'barely' won an election to municipalities where a woman 'barely' lost, i.e. by a narrow margin, this design allows us to isolate the causal effects of a woman winning office from the spurious correlation between a mayor's gender and municipality child death indicators – a correlation that may arise, for example, if the gender of a competitively elected leader is associated with voters' preferences in the first place.

Brazil, our country of interest, offers an ideal case to study the role of female local leadership. It is a highly decentralized federal system (Samuels, 2003), where municipalities have gradually been gaining autonomy since the 1970s, and are now viewed as the most autonomous sub-national units in Latin America (Nickson, 1995). Municipalities enjoy major policy control, e.g. land

[2] For instance, life expectancy at birth, as included in the human development index, is strongly and negatively associated with child mortality.

[3] Target 3.21 aims at reducing under-five mortality rate (U5MR) by 2030.

[4] In poor developing countries specifically, early childhood death accounts for 30 percent of all deaths, compared to 1 percent in richer societies (Cutler and Lleras-Muney, 2006). In addition, there are sharp health inequalities within poor societies, most child deaths happening in poor families.

parceling, local public service provision, including public transportation, education, or health services (IBGE, 2002). Moreover, the number of competitive elections frequently held, and the size of the country, offer enough statistical power to identify effects in close, mixed-gender races. Discontinuities in close elections between female and male mayoral candidates in Brazil have been exploited to study the role of women as political leaders. For instance, Brollo and Troiano (2016) find that female mayors are less likely to engage in corruption compared to male mayors, to hire fewer temporary public employees, and to attract less campaign contributions when running for reelection, concluding that male incumbents are more likely to engage in strategic behavior in order to improve their electoral performance. Electing a female mayor is also shown to increase non-preterm births and prenatal visits. Bruce et al. (2022) reveal that, during the Covid-19 pandemic crisis, female leadership reduces Severe Acute Respiratory Infection (SARI) deaths and hospitalizations, and increases enforcement of non-pharmaceutical interventions, concluding that female mayors outperformed male mayors when dealing with a global policy issue. Delaporte and Pino (2022) evidence that electing female mayors reduces episodes of gender violence.

Still, gender-based violence or SARI deaths resulting from a global health crisis are not clear measures of child health dynamics. And, while preterm births – children born before week 37 of gestation – might be a relevant indicator of population health, they do not inform on child health *per se*, rather on mothers' physiological conditions, exposure to risk factors, and access to healthcare.[5] However, a strong, negative correlation has recently been evidenced between the election of female mayors and U5MR in Brazil between 2000 and 2015, likely explained by expanded access to primary health care and conditional cash transfer programs (Hessel et al., 2020). We build on existing works, and contribute to the literature by estimating the causal effect of women's local leadership on child health, before the Covid-19 pandemic.

We find that marginally electing a woman as mayor has no statistically significant effect on

5 Preterm births tend to follow spontaneous preterm labor by the mother, maternal or fetal infections, or premature rupture of the membranes – the first risk factors for preterm labor being maternal infections or inflammation, birth history and genetic factors (Goldenberg et al., 2008).

the number of under-five child deaths per 1,000 live births in the whole sample of Brazilian municipalities. Baseline results mask the presence of heterogeneity. Breaking up estimates, we identify a decrease in child mortality in municipalities with a greater share of legislative power held by women, where electing a female mayor causes a decrease of 9.51 in under-five child deaths per 1,000 live births, which represents 41.47 percent of child death mean among municipalities where a man was elected as mayor. We also evidence a decrease of 6.83 in child deaths in municipalities where most of the population is illiterate, equivalent to 24.99 percent of the child death mean in control areas; as well as a moderately significant decrease of 5.79 deaths in municipalities with a greater population of eligible voters – 28.91 percent of the child death mean where a man holds office.

Our results, robust to numerous checks, contrast with existing findings, bringing nuance to growing research on the role of women as policymakers on child health. They suggest that increasing women's political representation might not systematically represent a cost-effective avenue for child health promotion in middle-income countries, except for some subsets of the population. While there is a strong and robust correlation between women's political representation, e.g. in national parliaments, and child mortality, causal evidence on women holding executive power remains scarce and inconclusive. Most of the empirical evidence on the role of female leadership relying on quasi-experimental methods is geographically limited to India, and tends to exploit the introduction of quota laws, leading to sudden and substantial increases in the proportion of female legislators (at the state level). Albeit of essential value, existing findings might have limited external validity, as the assumptions that quotas strengthen women's identity, and influence their policy preferences might not hold elsewhere (Broockman, 2014). Neither do they clearly inform on the possible role of women holding executive power. Competitively running for (local) executive power might alter their incentives, becoming orthogonal to their own political preferences. We complement this literature by providing credible evidence on the impact of electing women to local executive power on child health.

Moreover, the absence of a significant effect of women's political representation in the full sample of Brazilian municipalities is consistent with evidence from the US (Ferreira and Gyourko, 2014), Italy (Gagliarducci and Paserman, 2011), or India (Ban and Rao, 2008); and, even, a worsening of health indicators (Anukriti et al., 2022). The lack of clear effects in Brazil might be explained by the combination of (i) women outperforming men because they are supported by a stronger presence of women holding legislative power (Thomas, 1991); (ii) female mayors accessing power through electoral races, rendering them obliged to address general demands of the Brazilian population;[6] and (iii) the presence of regional inequalities in child health (Alves and Belluzzo, 2004), and greater participation in elections by otherwise disenfranchised groups in Brazil, leading to discrepancies in healthcare spending and utilization across the country (Fujiwara, 2015). Our analysis thus emphasizes the need for more research to uncover the mechanisms underlying the policy preferences of female politicians, and corresponding effects on child health.

2 Data

Our analysis focuses on municipal administrations in Brazil. Brazilian municipalities are minor federal units with an autonomous local government, ruled by a mayor, directly elected by citizens to a four-year mandate, and a legislative body, directly elected by voters as well. Mayors of municipalities above 200,000 voters are elected by a majority run-off rule, while mayors of municipalities below 200,000 voters are elected by a plurality rule. We focus on five municipal administration mandates in municipalities below 200,000 voters: 2001-2004, 2005-2008, 2009-2012, 2013-2016, and 2017-2020.

Electoral data and information on mayoral characteristics, including gender, age, education, or marital status, come from *Tribunal Superior Eleitoral* (Superior Electoral Court), the highest judicial body of the Brazilian electoral justice. Brazil having 5,567 municipalities during our period of analysis, we select those municipalities with mixed gender mayoral races in municipal

[6] In other words, politicians select interventions according to the median voter's preferences, disregarding their own, rendering politicians' individual characteristics irrelevant for policy decisions.

elections held in 2000, 2004, 2008, 2012, and 2016. We further restrict our sample to 'close' races with two candidates, where one candidate is a woman, and the other is a man. Further restricting the sample to elections where there are two candidates of mixed gender leads to a sample of 2,466 races, occurring in 1,639 unique municipalities.

Our outcome of interest is under-five child mortality, i.e. the number of 0-4-year-old deaths per 1,000 live births. It is obtained from the Brazilian Ministry of Health.[7] It is measured at the municipality-year level, specifically at the $t + h$ year, where t stands for a year during which mayoral elections are hold, and h, varies from 1 to 3. Years of electoral campaigns – 2000, 2004, 2008, 2012 and 2016 – are excluded from the estimation sample.[8] Data at the municipality-election year level are merged with under-five child mortality at the municipality-mandate level, as well as a cross-section of predetermined (2000) municipality variables, and elected mayors' features. As shown in Table 1, our final dataset has 5,514 municipality-mandate year observations.

To study how a woman holding mayor's office affects child health in mixed-gender races, we use a sharp regression discontinuity design, where all units (municipality-year observations) have a score (the margin of victory obtained by a female candidate in a given election), and those whose score exceeds a known cutoff receive the treatment (a female candidate winning a mayoral election), while those below the cutoff, do not. Under appropriate assumptions, a comparison of units with and without the treatment close to the cutoff might be used to study the causal effect of the treatment on some outcome of interest. Assignment to treatment is formalized as:

$$Female_{it} = 1[MV_{it} \geq 0] \tag{1}$$

where MV_{it} is the female mayor candidate's margin of victory in municipality i during term t,

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[7] Data are available at `https://datasus.saude.gov.br/`.

[8] This is because it has been shown that the vote share of candidates aligned with the president is lower in municipalities where more beneficiaries received penalties shortly before the elections for not complying to Bolsa Familia program conditions, an intervention known to influence health indicators, as well as weaker enforcement of Bolsa Familia conditions before elections in municipalities where mayors from the presidential coalition run for reelection (Brollo et al., 2019).

and $1[.]$ the indicator function. MV_{it} is specified as the vote share of the female candidate $minus$ the vote share of the male candidate. The cutoff that determines a woman's electoral victory is normalized to zero: she wins the election when her vote margin is positive, and loses otherwise. MV_{it} thus takes positive values if the mixed-gender electoral race resulted in a female mayor, and negative if it resulted in a male mayor. At the zero threshold, $MV_{it} = 0$, the gender of the mayor $Female_{it}$ sharply changes from zero to one.

Table 1 summarizes outcome and control variables according to the gender of the mayor. The first panel of the table shows predetermined (2000) municipal characteristics; the second panel, mayoral features; and the last panel, the outcome variable. While we observe statistically significant differences between male and female candidates – the former being more likely married, and less educated than the latter – we do not find any dissimilarities in predetermined municipality covariates, nor in child deaths.

3 Estimation strategy

We estimate a sharp regression discontinuity design. We rely on local polynomial methods to fit two separate regression functions above and below the cutoff. The estimated regression discontinuity effect is calculated as the difference between the two estimated intercepts.

In practice, we estimate a series of local linear regressions of child death at $t + h$ on a female mayoral candidate's margin of victory at t. Weights are computed through a kernel function to the distance between each observation's score and the cutoff. To be implemented, kernel-based estimators need a bandwidth, with observations outside the bandwidth generally receiving zero weight. Following common practice, we choose an optimal bandwidth that minimizes the mean squared error (MSE). However, since minimizing MSE generates bandwidths too large for conventional confidence intervals to be valid, we use Calonico et al.'s (2014) robust confidence intervals:[9]

[9] Calonico et al.'s (2014) robust confidence intervals account for the asymptotic bias often ignored by conventional inference, and correct standard errors, ensuring valid inferences, even for large bandwidths. For implementation,

$$Y_{i,t+h} = \alpha + \beta Female_{it} + f(MV_{it}) + \epsilon_{i,t+h} \tag{2}$$

where $Y_{i,t+h}$ stands for the outcome of interest, child mortality; α is a constant; $Female_{it}$ is the treatment variable; MV_{it} is the forcing variable, i.e. the female mayoral candidate's margin of victory; $f(MV_{it})$ is specified as an nth-order polynomial on MV_{it} that captures continuous changes in the dependent variable with respect to the margin of victory. Specification (2) allows for a flexible relation between the outcome and the margin of victory. In a subset of specifications, we add predetermined municipality and mayor covariates to increase estimate precision.

The parameter of interest, β, is the causal effect of a female mayor under the assumption that observables and unobservables are not correlated with marginal victory; it quantifies the difference in average child death outcomes of municipalities with a female candidate who barely won or barely lost election. If municipalities where a woman barely wins the election at t (the treatment group) are not suddenly different from municipalities where the a woman barely loses the election at t (the 'control group'), this design enables us to study the (local) average effect at the cutoff of a woman winning office at t on its mandate period child health at $t + h$.[10,11]

4 Results

4.1 Baseline results

Table 2 documents estimates of a sharp regression discontinuity design. Columns (1) and (2) present results without (predetermined) covariates; columns (3) and (4) include predetermined

we use Stata -rdrobust- software.

[10] Note that the RD effects we estimate are local by construction, as they represent the effects of a female candidate gaining access to office in 'extremely' close elections.

[11] In other words, the exogenous variation in a mayor's gender comes from variation in that, around the zero cutoff, winning just above the cutoff generates a sharp regression discontinuity design. Municipalities where a woman was elected mayor just above the 0 victory cutoff, and those where a man was marginally elected, are identical on observable and unobservable characteristics, except for the gender of the elected mayor.

municipality and mayor covariates listed in Table 1. Instead of choosing an arbitrary bandwidth value, we implement the most conservative algorithm, as suggested by Calonico et al. (2014), that also provides biased corrected estimates and robust confidence intervals. Standard errors are clustered at the municipality level in columns (1) and (3). The heteroskedasticity-robust plug-in residuals variance estimator (with $hc3$ weights) is used to compute standard errors in columns (2) and (4).

Albeit coefficient estimates being negative in sign, in line with existing evidence, Table 2 indicates that, when a female candidate barely wins the t election, i.e. by a small margin, there is no statistically significant change in under-five child mortality during her mandate, compared to when she barely loses – women holding local executive power does not decrease significantly under-five child mortality. These results hold whether changing how standard errors are computed, or including predetermined covariates. In addition, point estimates are similar in sign and statistical significance regardless of the exact econometric specification used; and including covariates slightly decrease their magnitude in absolute value, validating our regression discontinuity design.

The absence of an impact is confirmed by Figure 1, where we plot the binned outcome means against the female mayor candidate's margin of victory (score), where the dots are optimally-chosen binned means and the solid line is a second-order polynomial fit. We do not observe any jumps at the cutoff: the average under-five child mortality per 1,000 live births does not appear to decrease significantly in those municipalities where a female candidate obtained a barely positive margin of victory than in those municipalities where the female candidate barely lost. There does not appear to be a break in the trend within the data at the threshold. In other words, there is no visible discontinuity at the cutoff: Figure 1 corroborates the limited evidence that the gender of the winning mayor translates into better child health highlighted by Table 2.

Importantly, the lack of an effect of a politician's gender has been highlighted elsewhere in the literature: in the US (Ferreira and Gyourko, 2014), Italy (Gagliarducci and Paserman, 2011), or India (Ban and Rao, 2008; Clots-Figueras, 2012).[12] Worsened health outcomes have even been

[12] Gagliarducci and Paserman (2011) and Ferreira and Gyourko (2014) found that female politicians had no effect on expenditures in Italy and the US, respectively. Ban and Rao (2008) showed that female parliament leaders did

evidenced by Anukriti et al. (2022) who report an increase in intimate partner violence when women are elected in Indian state legislatures, likely due to conflicts around contraceptive use. At this stage, results suggest that, at best, female mayors do not perform 'worse' than men.

4.2 Validity of the research design

The above results indicate that municipalities where women won the mayoral election at t by a small margin do not present significantly different child death rates during their mandates. Surprisingly contrasting with existing works, we should ensure that there is no threat to their validity. The regression discontinuity design estimates would be invalid if, for instance, (fe)male candidates could precisely manipulate close elections to their advantage. In this case, observations on either side of the zero cutoff could not be comparable.

We conduct two tests to assess the validity of this design. First, covariate tests aim at showing null regression discontinuity effects on key predetermined (municipality and mayor) characteristics, included as control variables to increase precision. In theory, these variables should be prior to or unaffected by the gender of the elected mayor at t. Second, density tests intend to confirm that the approximate number of observations just above the zero cutoff does not significantly diverge from the number of observations just below this cutoff (McCrary, 2008). Below, we present evidence supporting the validity of our design as per these two tests.

We first examine whether, near the cutoff, treated units are similar to control units in terms of observable characteristics. We use the set of predetermined variables used for covariate adjustment, as listed in Table 1. We run regression discontinuity specifications, and generate corresponding plots. Table A1 indicates that, on average, point estimates are small, and all 95% robust confidence intervals contain zero, with p-values ranging from 0.108 to 0.975. Table A1 confirms that there is no empirical evidence that these predetermined covariates are discontinuous at the cutoff.[13]

not perform better than men in South India, and that women occupying reserved seats did not converge to the preferences of Indian women. While Clots-Figueras (2012) showed that raising female political representation in state assemblies increased a student's probability to finish school in urban areas, no effect was evidenced in rural areas, nor in the whole sample.

[13] Note that the number of observations used in the analysis varies for each covariate; this occurs because the MSE-

Figure A1 presents corresponding plots. Consistent with the formal statistical results, the graphical analysis within the optimal bandwidth shows that the right and left intercepts in the local linear fits are very close to each other. This is evidence of the regression discontinuity design, in spite of the absence of a large, non-zero effect for the outcome of interest.

In addition to a graphical illustration of the density of the running variable, in Figure A2, we then explore this assumption more formally using a statistical test – a density test. This strategy tests whether the number of treated observations in the chosen neighborhood is compatible with what would have been observed if units had been assigned to the treatment group, i.e. being above the cutoff, with a 50% probability. We only keep observations with $MV_{i,t} \in$ [-0.02, 0.02], and find that, in this neighborhood, there are 233 control observations and 258 treated observations. Using this information and setting a probability of success equal to 1/2, we perform a binomial test. The p-value is 0.2787. This means that under the continuity-based approach, we fail to reject the null hypothesis of no difference in the density of treated and control observations at the cutoff – the numbers of treated and control observations are consistent with what would be expected if municipalities were assigned to a female win or loss by the flip of an unbiased coin. Figure A2 provides a graphical representation of the continuity in density test approach, exhibiting the actual density estimate with shaded 95% confidence intervals. It indicates that the density estimates for treated (blue) and control groups (red) at the cutoff are very near each other, and the confidence intervals (shaded areas) overlap. This plot is consistent with the results from the formal test; it confirms no manipulation around the cut-off.

4.3 Heterogeneity effect analysis

We next explore plausible differences in the relative performance of male and female mayors in subsamples according to election and mayors' features, and 2000 municipality characteristics. Discontinuities in close elections between mayoral candidates in Brazil have been used to highlight that becoming the incumbent party results in large subsequent electoral losses

optimal bandwidth is different for every covariate analyzed.

(Klašnja and Titiunik, 2017), and that experienced and educated mayors tend to devote a smaller fraction of the budget to current and personnel expenditures (Rocha et al., 2018). In pre-election years, municipalities where mayors were affiliated with the president's political party were found to receive about one-third greater discretionary transfers for infrastructure (Brollo and Nannicini, 2012); and politicians able to run for re-election and part of the presidential coalition, to manipulate strategically the enforcement of conditional welfare programs to influence electoral outcomes (Brollo et al., 2019). In other words, if aligned mayors have more public resources to work with, being favored by Brazil's government, and if child health indicators are influenced by female mayors making better (or worse) use of these resources, there might be stronger differences in female over male led municipalities in the sample of municipalities aligned with the President's party.

Table 3 indicates that a female mayor's victory has, overall, an effect indistinguishable from zero on child death depending on mandate years or mayor predetermined characteristics, such as past political experience, or alignment with the President's political party. Still, column (4) suggests an 11.57 child death decrease by 1,000 live births following the 2012 election, between 2013 and 2015, corresponding to 51.85 percent of child mortality in municipalities where a man marginally won election. Interestingly, this improvement in child health follows the presidential election year of 2010, a year of intense presidential election campaign, where expenditures might have been reallocated to the detriment of policies affecting child health, as suggested by the positive coefficient estimate in column (3), indicating a possible reverse to the pre-presidential election child mortality mean. Column (14) also indicates that, among mayors without tertiary education, female mayor appears to decrease child deaths by 7.78 per 1,000 live births. Being statistically significant at 10 percent, this estimate is nonetheless informative. Its confidence interval indicates that we are able to reject a relatively large positive effect, greater than a 0.55 child death increase, of non-tertiary educated female mayors with a 95 percent probability. Last, column (16) confirms that incumbency affects child health: non-incumbent female mayor happens to decrease child mortality by 7.52 deaths, corresponding to 32.68 percent of child mortality in municipalities where a man was elected for the first time.

We now turn to possible heterogeneity driven by predetermined municipality characteristics, in Table 4. First, a greater share of legislative seats (*vereadores*) won by women, or a larger share of the female population, might favor female mayors, leading to women outperforming men because they are supported by a stronger presence of women holding legislative power (Thomas, 1991). In column (3), we find evidence of favoritism along the lines of gender – in areas with the greatest share of women holding legislative power, that women barely winning elections results in a 9.51 child death decrease, which represents 41.47 percent of child deaths in municipalities led by male mayors. Second, column (5) indicates that, in municipalities with the greatest eligible voter population, barely electing women as mayors decreases under-five child mortality by 5.79 per 1,000 live births, equivalent to a 28.91 percent decrease compared to child mortality in municipalities led by men. Albeit statistically significant at 10 percent, we are able to reject with 95 percent probability an effect bigger than a 0.52 death increase. Third, we see a 6.83 child death decrease in areas with the lowest rates of literacy as a result of female candidates marginally winning elections – 24.99 percent of control municipality mortality.

In addition to women outperforming men because they are supported by a stronger presence of women holding legislative power (Thomas, 1991), the lack of clear effect of a mayor's gender on child mortality in Brazil might be explained by the fact that female mayors are elected through electoral races, rendering them obligated to address general demands of the Brazilian population. This is consistent with Downs' (1957) median voter theorem (MVT) (Carvalho, 2017),[14] suggesting that, under certain conditions, politicians select interventions as per the median voter's preferences, disregarding their own preferences. This might explain decreased child mortality in areas with a lower share of literacy. Further support for this hypothesis comes from the existence of regional inequalities in child health (Alves and Belluzzo, 2004), and Fujiwara

[14] This theorem predicts that the identity and preferences of the candidate might not determine their policy choices. It follows that gender (among any other features) of the political leader should not influence corresponding policy outcomes as they address the median voter's interests. They are concerned with winning elections. The interventions that raise their likelihood of victory are those in line with the median voter's preferences – the options the majority of voters favor.

(2015) who shows that, in Brazil, greater participation in elections by otherwise disenfranchised groups, especially lower educated individuals, led to significant increases in healthcare spending and utilization, e.g. prenatal visits, as well as reductions in the number of children born with low birth weight.

5 Conclusion

We explored how women accessing to local executive power affects under-five child mortality in Brazil. Sharp regression discontinuity estimates showed that a female candidate (barely) becoming mayor resulted in decreases in child deaths in predominantly illiterate municipalities, and areas with a greater share of women holding legislative power, in the next three years following election. Our results, robust to numerous checks, contrast with existing findings. We proposed an explanation for these findings, by which female mayors are favored by a stronger presence of women with legislative power, and, simultaneously, rendered obligated to address general demands of the Brazilian population, disregarding, at times, their own preferences. Given the presence of heterogeneity, we conclude that raising women's political representation might not necessarily represent a cost-effective avenue for child health promotion in middle-income countries. Our analysis emphasizes the need for more research to uncover the mechanisms underlying the policy preferences of female politicians, and corresponding effects on child health.

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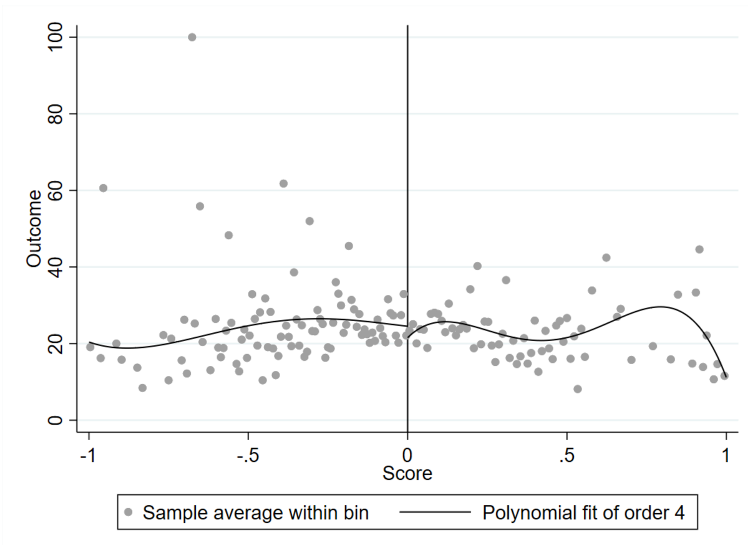
Figure 1: Regression discontinuity plot of 0-4-year-old deaths per 1,000 live births



Notes: Regression discontinuity plot of female mayor candidate margin of victory and under-five mortality per 1,000 live births. The intervention cutoff is set to 0; it is represented by the vertical black line.

Table 1: Estimation sample descriptive statistics

Variables	Male mayor mean (1)	sd (2)	Female mayor mean (3)	sd (4)	Difference (3)-(1) (5)
Muncipality characteristics					
North	0.08	0.27	0.08	0.28	0.002
					(0.012)
South	0.13	0.33	0.12	0.32	-0.008
					(0.014)
2000 share of poverty	49.35	22.54	49.92	21.53	0.571
					(0.990)
2000 Gini index	55.12	6.66	55.47	6.91	0.346
					(0.305)
2000 share of employed	47.62	8.91	47.06	8.53	-0.560
					(0.379)
Elected mayor characteristics					
Age on the first day in office	47.43	10.18	46.71	9.20	-0.720
					(0.460)
Married	0.77	0.42	0.70	0.46	-0.074***
					(0.021)
University education	0.39	0.49	0.61	0.49	0.211***
					(0.023)
PMDB	0.18	0.39	0.21	0.41	0.030
					(0.019)
DEM	0.10	0.30	0.12	0.32	0.013
					(0.015)
Outcome					
Number of 0-4-year-old deaths per 1,000 live births	25.39	34.68	24.04	24.26	-1.350
					(1.072)
Observations	3,195		2,319		5,514

Notes: Standard errors clustered at the municipality level in parentheses. Significance levels: $^{*} < 10\%$ $^{**} < 5\%$ $^{***} < 1\%$

Table 2: Regression discontinuity estimates, 0-4-year-old deaths per 1,000 live births

Variables	(1)	(2)	(3)	(4)
Female	-4.0804	-3.4073	-2.7396	-2.5673
Confidence interval	[-11.8063, 2.9409]	[-9.0384, 1.6116]	[-10.3139, 4.1904]	[-8.1652, 2.3627]
Control group mean	25.3862	25.3862	25.3862	25.3862
Observations	5,514	5,514	5,514	5,514
Covariates	No	No	Yes	Yes

Notes: Regression discontinuity estimates. Robust 95% confidence intervals are in brackets. They are clustered at the municipality level in columns (1) and (3). The heteroskedasticity-robust plug-in residuals variance estimator (with $hc3$ weights) is used to compute standard errors in columns (2) and (4). *** p<0.01, ** p<0.05, * p<0.1

Table 3: Regression discontinuity estimates of 0-4-year-old deaths per 1,000 live births, by mandate and mayor characteristics

	2001-03	2005-07	2009-11	2013-15
	(1)	(2)	(3)	(4)
Female	6.9995	-5.6008	6.1978	-11.5660**
Confidence interval	[-5.5331, 18.0392]	[-16.1047, 7.6892]	[-2.9507, 16.8654]	[-25.5697, -0.9198]
Control group mean	35.8967	29.4077	21.7387	22.3070
Observations	860	915	1,078	1,347

	2017-19	Mandate year 1	Mandate year 2	Mandate year 3
	(5)	(6)	(7)	(8)
Female	-5.0159	-0.4655	-3.9494	-5.2206
Confidence interval	[-18.4587, 6.8706]	[-9.7940, 7.8334]	[-12.4229, 3.5157]	[-14.7268, 3.1939]
Control group mean	22.2154	25.9833	25.1813	25.0030
Observations	1,314	1,833	1,844	1,837

	Elected mayor has past political experience		Elected mayor is aligned with president's party	
	Yes	No	Yes	No
	(9)	(10)	(11)	(12)
Female	-1.0046	-2.6795	7.7251	-2.6554
Confidence interval	[-7.8865, 7.2590]	[-9.3535, 3.1594]	[-6.4433, 27.1048]	[-7.7258, 2.9390]
Control group mean	24.1178	25.7058	33.8527	24.2296
Observations	1,078	4,436	665	4,849

	University education		Reelection	
	Yes	No	Yes	No
	(13)	(14)	(15)	(16)
Female	2.4864	-7.7821*	4.0249	-7.5165***
Confidence interval	[-2.1893, 7.8731]	[-17.6942, 0.5520]	[-4.1989, 13.1978]	[-13.9714, -2.1221]
Control group mean	22.7690	27.0926	24.6496	23.0026
Observations	2,666	2,848	1,242	3,412

Notes: Regression discontinuity estimates. Specifications include control variables. Robust 95% confidence intervals are in brackets. The heteroskedasticity-robust plug-in residuals variance estimator (with $hc3$ weights) is used to compute standard errors. *** p<0.01, ** p<0.05, * p<0.1

Table 4: Regression discontinuity estimates of 0-4-year-old deaths per 1,000 live births, by municipality characteristics

	2000 share of female population		Share of seats won by female *vereadores*		Eligible voter population	
	Above median	Below median	Above median	Below median	Above median	Below median
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.8367	-3.7905	-9.5087**	0.2312	-5.7948*	0.2595
Confidence interval	[-4.6610, 3.7452]	[-13.4127, 4.3700]	[-20.2210, -1.0398]	[-5.5345, 5.4927]	[-12.3673, 0.5184]	[-7.7181, 7.1191]
Control group mean	20.3948	30.2145	22.9266	27.8141	20.0456	30.6570
Observations	2,756	2,758	2,743	2,768	2,757	2,757

	2000 rural population		2000 poverty		2000 literacy	
	Above median	Below median	Above median	Below median	Above median	Below median
	(7)	(8)	(9)	(10)	(11)	(12)
Female	-2.1942	-2.2095	-2.6291	-2.0802	-0.2264	-6.8275**
Confidence interval	[-11.6412, 5.3945]	[-7.5220, 2.3834]	[-10.6175, 5.0815]	[-7.9365, 5.0728]	[-7.3920, 6.4119]	[-13.2034, -0.9006]
Control group mean	29.0309	21.5835	29.0525	21.6620	23.4426	27.3261
Observations	2,744	2,755	2,755	2,759	2,756	2,758

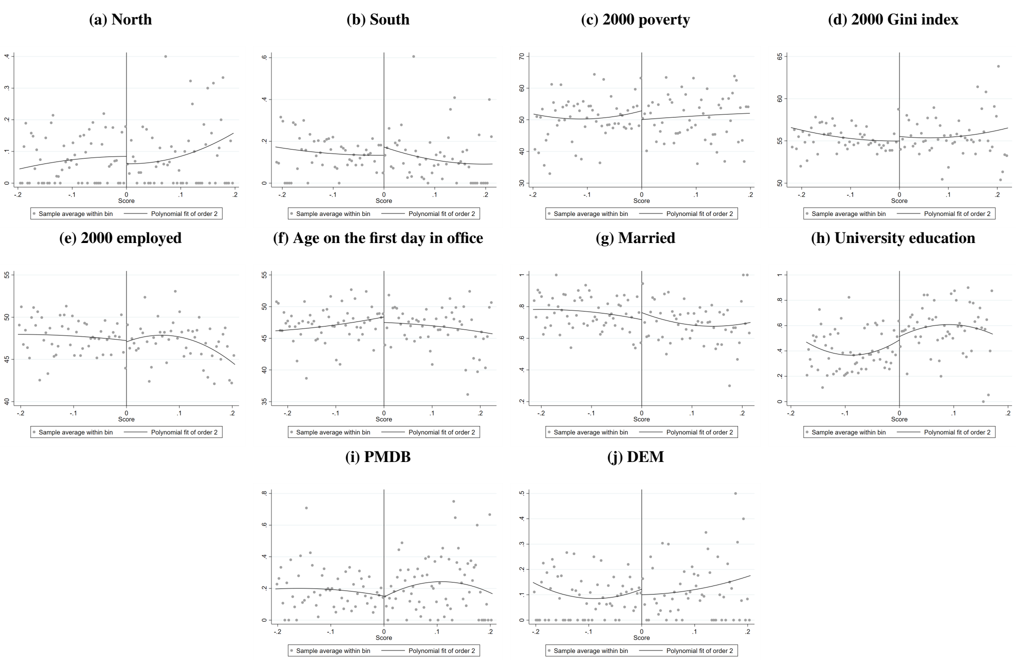
Notes: Regression discontinuity estimates. Specifications include control variables. Robust 95% confidence intervals are in brackets. The heteroskedasticity-robust plug-in residuals variance estimator (with $hc3$ weights) is used to compute standard errors. *** p<0.01, ** p<0.05, * p<0.1

Appendices

Table A1: Formal continuity-based analysis for covariates

Variables	MSE-optimal bandwith (1)	RD estimator (2)	Robust inference p-value (3)	Confidence interval (4)	Effective number of observations (5)
North	0.197	-0.02335	0.517	[-0.069568, 0.034982]	3,913
South	0.213	0.03943	0.258	[-0.029126, 0.108697]	4,057
2000 share of poverty	0.199	-2.7748	0.189	[-7.40553, 1.46263]	3,928
2000 Gini index	0.222	0.53695	0.462	[-0.826281, 1.82079]	4,146
2000 share of employed	0.205	-0.15488	0.785	[-2.08545, 1.57672]	3,972
Age on the first day in office	0.225	-0.87303	0.400	[-2.64307, 1.05408]	4,177
Married	0.217	0.04457	0.259	[-0.036441, 0.135296]	4,108
University education	0.172	0.02654	0.829	[-0.093358, 0.116544]	3,626
PMDB	0.204	-0.00601	.0759	[-0.082023, 0.059825]	3,966
DEM	0.205	-0.02197	0.421	[-0.087867, 0.036719]	3,972

Figure A1: Graphical illustration of local linear regression discontinuity effects for predetermined covariates



Notes: Regression discontinuity plots of female mayor candidate margin of victory and predetermined covariates. The intervention cutoff is set to 0; it is represented by the vertical black line.

Figure A2: Estimated density of the score