# Do Female Leaders Promote Gender-Sensitive Policies? **

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**Abstract**

We study whether female-headed local governments in Spain are more likely to engage in gender sensitive policies such as long-term care support, pre-schooling, or work and family life balancing services. Using a fuzzy regression discontinuity design estimated on the set of mixed electoral races, we find no evidence of female mayors being more likely to implement these policies at the local level. We do find evidence of differences between parties in the probability of implementing these policies, suggesting that the gender of the politician is less important than their partisan or ideological position when it comes to these policy levers.

*Keywords:* Female politicians, gender policies, long-term care.
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## 1. Introduction

Women are under-represented in top political positions in the vast majority of countries, both in national politics and at the local level.[1] Yet recent decades have seen an increase in the representation of women in government leadership positions. This increase in female representation is important for (at least) two reasons. First, it can have symbolic consequences, by changing the attitudes of voters and the general public towards other women inside and outside politics (Beaman et al., 2012; Iyer et al., 2012; Beaman et al., 2009; De Paola, Scoppa and Lombardo, 2010). Second, if the identity of politicians matters for policy decisions - as in Besley and Coate (1997), Osborne and Slivinski (1996) - female leaders may implement different policies from their male counterparts.

Women are more likely to experience sexual and domestic violence (Aizer, 2010), labour discrimination (Goldin and Rouse, 2000) and pay and employment gaps (Weichselbaumer and Winter-Ebmer, 2005; Olivetti and Petrongolo, 2008). This happens in a context in which the burden of caring activities within the family is disproportionately borne by women in most households. Government policy can alleviate or reinforce these gender disparities. For example, public pre-schooling or long-term care services may substitute for care-giving time at home. If the identity of politicians matters for policy, it might well be the case that women in office promote interventions to deal with these issues, be it because of individual preference for these type of policies, more precise information on the problems in question, or concerns about the welfare of other women.[2]

Our paper attempts to test this hypothesis by studying how mayoral gender influences spending in three gender-sensitive policies in Spanish municipalities: pre-schooling, long-term care (LTC), and work and family-life balancing services. Our empirical strategy focuses on mixed-gender electoral races to implement a close-election regression discontinuity (RD) design which provides credible exogenous variation in the gender of the municipal mayor. We find that the gender of the mayor does not have a significant effect on municipal engagement with gender-sensitive policies. To analyze whether spending in these policy levers is orthogonal to the political process altogether during our period of study we use a similar strategy to estimate the effect of political parties. We find significant differences by party, with centre-right governments being less likely to adopt one of these policies. Hence, according to our findings, party lines matter for these policies but mayoral gender does not. We confirm the robustness of all our results using a panel specification requiring a different set of identifying assumptions.

The choice of the three policies of interest is motivated by the specific impact these may have on the livelihood of women. All three are mentioned in the European Parliament's Gender Equality reports as key policies directed at enhancing the position of women in society. They can be seen as providing a substitute to household activities which are disproportionately carried out by women. Regarding long-term care, Spanish micro-data from *Instituto Nacional de Estadística* show that it is women who disproportionately perform these activities. In 2008, before our period of study, 76%

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[1]Women represent 12% of legislative seats worldwide (Kanthak and Woon, 2015). To cite three specific examples, only 12%, 19% and 40% of legislators are women in India, the United States and Spain. Under-representation is also present at the local level, with only 18% and 17% of female mayors in the United States and Spain, respectively.

[2]Interestingly, previous research has shown that, at the household level, women attend better the necessities of other female members of the family (Duflo, 2003; Thomas, 1990). Furthermore, there is substantial evidence that men and women hold different preferences (see for example Alesina and La Ferrara, 2005; Miller, 2008; Lott and Kenny, 1999; Funk and Gathmann, 2015), also among professional politicians (Lovenduski and Norris, 2003).

of professional and non-professional caregivers of long term patients were women. In 2004, 84% of non-professional caregivers were women. As Crespo and Mira (2014) show, this has an impact on the labour market decisions of mature women in Southern Europe. On average, 20% of women between the ages 50 and 60 whose parents lack autonomy take up daily care, with half of them dropping out of the labour force as a result. This disproportionate role of women is also observed in childcare, with early-years care being carried out mainly by mothers. The 2009-2010 wave of the Spanish time use survey indicates that women engage more often in childcare activities than men, and that they devote more time per day to those activities.[3] Moreover, maternity also impacts labour market decisions of women. According to Landwerlin, Balsas and Saura (2012), only 55% of women in Spain return to full-time employment after maternity, compared with nearly 100% of men. Thirty-five percent return to part-time work or take part-time leave, 7% give up paid work altogether and 3% lose their job. As in the case of LTC, this may also lead to statistical discrimination in hiring and promotion decisions.

The Spanish context is well-suited to study the engagement of local leaders with gender-sensitive policies for two reasons. First, there is substantial discretion in the implementation of these interventions at the local level. In 2006, the national government passed a long-term care law that granted discretion to Spanish municipalities in the provision of supplemental services for long-term patients. Moreover, the 2006 Education Law (*Ley Orgánica 2/2006*), which does not guarantee universal access to preschooling, explicitly recognizes the role for municipalities in the task of providing this service to an increasing number of families. As for other education-related policies that may contribute to work and family life balancing services – such as school busing services for children and subsidized school meals –, they are entirely carried out by regional and local governments. Hence, local governments in Spain are relevant actors on the deployment of gender-sensitive policies. The second reason why Spain is well-suited for our purposes relates to the rules of administrative information disclosure in force between 2009 and 2014. The budget structure law passed in 2008 (*ORDEN EHA/3565/2008*) required municipalities to separately report spending in LTC, preschool and work and family life balancing services. This allows us to directly identify the municipalities which are executing these policies from administrative records and, in doing so, to dig beyond the broad spending categories that have been studied in much of the related literature.

Our main results show no evidence of female politicians being more likely to implement gender-sensitive policies. Point-estimates are generally very close to zero – especially when using our intensive margin measures of policy engagement – and are always statistically insignificant. Panel estimates corroborate these findings, yielding tightly estimated zeroes. Hence, our results indicate that there are no differences between male- and female-led local governments in the engagement with gender-sensitive policies. When turning to the effect of parties, we do find statistically significant differences. Right-wing mayors are roughly 10 percentage points less likely to implement these policies. Comparisons of both sets of estimates are not straightforward as they are based on different samples. Still, our results suggest that party platforms are more important

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[3]According to this survey, 22.2% of women report engaging in childcare activities while only 16.7% of men do so. Average time devoted to childcare conditional on it being positive is 2h 22m for females while it is 1h 46m for males. Data from the 2015 American Time Use Survey shows similar qualitative patterns for the United States, with women devoting 125% more time to caring activities for household members than males.

than mayoral gender in determining gender-sensitive policies.

These results have important implications. In recent years, there have been rising concerns about the under-representation of women in position of power. In some cases, this has prompted the implementation of gender quotas, the effectiveness of which has been studied in a growing literature on the matter (see for example Esteve-Volart and Bagues, 2012; Fréchette, Maniquet and Morelli, 2008; Casas-Arce and Saiz, 2015; Bagues and Campa, Forthcoming; Gonzalez-Eiras and Sanz, 2018). According to the results of our study, having more female politicians will not result in increased attention to gender-related issues. That is, while increasing female representation continues to be a policy priority which can be defended on different grounds, our results indicate it can have limited effects on policy decisions. Conversely, the fact that right-wing parties devote less resources to these issues is consistent with female voters favouring left-wing parties in elections (Edlund and Pande, 2002; Iversen and Rosenbluth, 2006).

Previous studies on the effect of female leaders on policy choices are abundant and cover both a wide variety of policies and countries. Chattopadhyay and Duflo (2004) use randomization of women-reserved mayoralties in Indian villages to find that female leaders are more responsive to issues raised by female constituents. Also in the context of India, Clots-Figueras (2011) shows that female politicians that occupy seats reserved for low castes and scheduled tribes favour female-friendly policies. In particular, she studies reforms of the Hindu succession act that give women the same succession rights as men. A common characteristic shared by both of these papers is that they study policies which have a specific gender component. Yet both the policies under study and the institutional design of the political arena are somewhat specific to the Indian context. This limits the scope to extrapolate the conclusions of these studies to other jurisdictions.[4] Our results differ from those in these papers both in the context under consideration and in that we find no effect of gender on policy decisions. One important consequence of our close-election RD strategy is that our estimates come from female leaders elected in competitive elections where they run against a male candidate. This stands in contrast with the context discussed in Chattopadhyay and Duflo (2004), where villages are randomly assigned to have a female leader and only women can run for election.

Much of the remaining literature studying the effect of politicians' gender on policy has *not* focused on specific policies that systematically favour (or are systematically prioritized by) women. Ferreira and Gyourko (2014) find no effect of gender of the local mayor on broadly defined policies such as size of local government or the broad composition of municipal spending. Likewise, Bagues and Campa (2018) exploit the introduction of gender quotas in party lists and find no effect of the gender composition of Spanish municipal councils on broadly defined social policies. Conversely, Brollo and Troiano (2016) use a close election RD for Brazilian municipalities and find significant effects on corruption and measures of patronage indicating that female mayors are less likely to engage in these practices. Clots-Figueras (2012) focuses on outcomes and finds electing a female politician has a positive impact on educational outcomes in Indian urban areas. While these studies are of interest in their own right, they do not consider whether female politicians system-

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[4]Focusing on an arguably more general set of policies, Svaleryd (2009) uses a municipal fixed-effects panel specification to show that the fraction of female politicians in Swedish municipal councils correlates with relative spending on different gender-sensitive policies.

atically choose gender-sensitive policies when in power. The policies or outcomes they analyse are not clearly linked to gender-specific problems. Our paper attempts to tackle this question in a context where we can both identify key gender-sensitive policies that are widespread in many economies and implement a state-of-the-art identification strategy to obtain credible estimates of the relationship between policy decisions and gender.

## 2. Institutional Setting & Data

### 2.1. Institutional Setting

Spain had 8,116 municipalities in 2011. Municipalities are the lowest level of territorial administration of the Spanish state and have autonomy in managing their interests as recognized in the Spanish constitution (Article 140). Municipal financing is based on municipal taxes (the largest of which are a property tax and a business tax) and transfers from the national and regional governments. The functions of the municipal government are partly dependent on size but encompass waste disposal, water and sewage services, lighting, transport network upkeep, public parks, and, crucially, the provision of some local public services.[5] Each year, municipal governments have to approve a budget for the next calendar year recording all planned expenses. These budgets are our source of data on the policies of interest. In this study we focus on the provision of three services that are considered to promote female empowerment by substituting for caring activities in the household: long-term care, pre-scholing and work and family life balancing services.

On December 2006 the Spanish parliament passed the law *Ley 39/2006*, popularly known as *Ley de Dependencia*. This law established the role of the public administration in providing long-term care to the elderly and other dependent individuals. However, in practice the new policies laid out in the *Ley de Dependencia* were poorly funded and had a limited scope. These limitations were reflected in the 2007 CIS national survey, in which 61.6% of women and 53.9% of men considered the national and regional coverage of long-term care needs to be poor/limited.[6] In this context, local authorities, whose complementary role was recognized in the 12th article of the law, had plenty of room to intervene at their discretion. Some examples of the interventions carried out by municipalities included policies that targeted the caregiver in the family, such as psychological counselling services or subsidies to acquire home medical equipment. They also included policies trying to professionalize the service, by providing publicly provided home care services or subsidies to hire private caregivers.

The case of pre-schooling is similar. The 2006 Education law (*Ley 2/2006*) established two voluntary stages of education for children from 0 to 3 and from 3 to 6 years of age. However, while access to the second stage at no cost was granted at the national level, access to the first stage was not. To cope with this limitation, article 15 of the law states the goal of promoting access at no cost also at the first stage. To that end, it gives a key role to local governments, which at their discretion could offer public daycare centres or subsidies to acquire these services privately.

Finally, work and family life balancing services include services such as busing, lunch at school, early drop and breakfast, extra-curricular activities in the afternoon and boarding schools. None of

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[5]See details in law number 7/1985 (2 of April 1985). *Ley reguladora de las bases del régimen local.*
[6]The difference in male and female answers is significant at all conventional levels.

these services are granted at the national level and therefore local governments play a crucial role in their provision. These expenditures are recorded in the budget as complementary educational spending.

Municipalities are governed by a municipal council (*pleno municipal*) and a mayor (*alcalde*). Municipal council members are directly elected by residents in municipal elections held every four years. The electoral system varies with population. Municipalities with populations over 250 use a single-district, closed list, proportional electoral system. Municipalities with populations under 250 inhabitants have an open list system with voters able to express multiple preferences for different candidates. These municipalities will not be used in our analysis precisely because of this difference and its implications for our empirical strategy. In municipalities using the closed-list system, municipal council seats (from a minimum of 7 to a maximum of 57 in Madrid) are assigned following the D'Hondt rule featuring a 5% vote share entry threshold. The municipal mayor is elected by the council under a majority rule. The closed list proportional system prevalent in most Spanish municipalities promotes party discipline of council members. Crucially, the mayor is often the leader of the local party branch, with provides substantial power in defining these lists. As a result, party identification is strong and mayors are influential. In a cross-country comparative analysis of local government leaders, Mouritzen and Svara (2002) classify Spanish mayors as strong, where a strong mayor is "an elected official who is the primary political leader of the governing board and possesses considerable executive authority". Below, the *ruling party* refers to the party of the mayor.

Female participation in national and local Spanish politics experienced a sustained increase in the period between the democratic transition in the late 70s and the present. In 1979, only 5% of all members of parliament, 1.2% of all mayors and 4.5% of council members were women (see Fernández, Fernández and de Ulzurrun 2003, Giol 1992). By 2015, these numbers had increased to 39%, 22% and 35%, respectively. While the increase in participation is consistent throughout the period – and has continued subsequently – women continue to be under-represented in Spanish politics.

### 2.2. Data and Descriptives

We construct a municipal panel for the period 2010-2014 combining data from four different sources. Electoral results for Spanish municipalities in the 2007 and 2011 municipal elections are obtained from the Spanish *Ministerio del Interior* which collects and disseminates electoral data. This data includes both municipal level results for all running parties and the list of candidates parties presented in every municipality. It includes the ordering of these lists and, importantly, the gender of each candidate. Data on mayors was facilitated as well by the *Ministerio de Hacienda y Administraciones Públicas* upon request.

We obtain data on municipal characteristics from the 2001 Census of Population. These include average household size, fraction of population with tertiary (college) education and fraction of female homemakers for 2001.[7] Data from *Estadística del padrón Continuo* include yearly information on population and population by age categories for all the sample period.

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[7]Data is also available from the 2011 census. Given that this falls within our sample period, the characteristics themselves could be outcomes of the treatment so we focused on demographics measured in 2001 as controls. Preliminary analyses suggested that this decision does not have an effect on our main results.

Data on yearly municipal budgets is obtained from the database on local authorities budgets published by the *Ministerio de Hacienda y Administraciones Públicas*. These include information on revenues and spending disaggregated by spending category. The fine level of disaggregation during the period 2009-2014 is crucial to identify the policies we analyze in this paper.[8] Also from the *Ministerio de Hacienda y Administraciones Públicas* we take the outstanding debt by municipality in 2009.

Merging data from these sources, we construct a panel of municipalities for the period 2010-2014 indicating the vote shares obtained by all parties, gender of the municipal mayor, information on municipal spending by program (including LTC, preschool and life-balancing services), outstanding debt in 2009 and municipal characteristics.

Changes in the law regulating the level of disaggregation in municipal budgets allow us to identify data on specific gender-sensitive programs. The availability of this data is restricted to the period 2009-2014. Hence, we will restrict to these years and to municipal elections taking place in 2007 and 2011. Some, typically small, municipalities fail to report their budget in time every year, so we loose 1,141 further year-municipality observations for that reason (roughly 4.2% of the final sample). In addition, as mentioned above, we restrict our sample to municipalities with populations above 250 inhabitants in electoral years.

Municipal descriptives for our sample are presented in tables 1 and 2. In Table 1 we present mean and standard deviation for both our outcome variables (the three selected policies) and municipal characteristics. We show these statistics for all municipalities, for municipalities ruled by a female mayor, by a male mayor, and for municipalities where a mixed race took place (male and female candidates running for election) in columns 1 through 4, respectively. We measure our outcome variables both in terms of the extensive (whether spending in a policy takes place) and intensive (what is the fraction of total spending allocated to that policy) margins. Additional characteristics include population, fraction of municipalities ruled by the centre-left *Partido Socialista* (PSOE) and the center-right *Partido Popular* (PP), as well as selected demographic variables. The average population of municipalities in our sample is 8.65 thousand inhabitants. We can observe that municipalities governed by a woman are usually larger but otherwise not very different from municipalities ruled by men in terms of observables. The average fraction of homemakers in Spanish municipalities is 33%. In terms of our outcome variables, municipalities ruled by women tend to devote slightly more resources to long term care than the average (0.66%). But both the dummy indicating whether a policy is funded and the intensity measures for other policies captured by budget shares have similar means across groups.

In Table 2 we present similar descriptives for municipalities split by ruling party. Female mayors are more common in towns ruled by PSOE than in towns ruled by PP although the difference is small (18% vs. 17%). There is substantial difference in the three gender policies by party with PSOE being more likely to implement these policies and to allocate a larger fraction of spending to them. While this difference is suggestive, other variables in Table 2 indicate there are other differences between PP and PSOE municipalities which highlight that these mean comparisons

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[8]The promulgation of the law (*ORDEN EHA/3565/2008*) in 2008 enforced municipalities to report their budget data with a high level of detail. This law was in force until 2014, when it was replaced by *Orden HAP/419/2014* that returned to a situation similar to that in place pre 2009, with budget data reported at broader categories.

TABLE 1

TABLE 1

MUNICIPAL DESCRIPTIVES BY GENDER

| | All | Female Mayor | Male Mayor | Mixed Races |
|---|---|---|---|---|
| **Gender-sensitive Policies** | | | | |
| Long-Term Care share | 0.61 | 0.66 | 0.60 | 0.63 |
| | (1.9) | (2.0) | (1.9) | (1.9) |
| Long-Term Care dummy | 0.22 | 0.22 | 0.22 | 0.22 |
| | (0.4) | (0.4) | (0.4) | (0.4) |
| Preschool share | 1.16 | 1.15 | 1.17 | 1.20 |
| | (2.2) | (2.2) | (2.2) | (2.3) |
| Preschool dummy | 0.36 | 0.36 | 0.36 | 0.37 |
| | (0.5) | (0.5) | (0.5) | (0.5) |
| Work-life balancing serv. share | 0.32 | 0.32 | 0.33 | 0.32 |
| | (0.9) | (0.9) | (0.9) | (0.9) |
| Work-life balancing serv. dummy | 0.25 | 0.24 | 0.25 | 0.25 |
| | (0.4) | (0.4) | (0.4) | (0.4) |
| **Demographics** | | | | |
| Population 2007 (000s) | 8.65 | 11.39 | 8.09 | 9.52 |
| | (58.4) | (91.8) | (48.8) | (69.9) |
| Pop % above 80 | 7.29 | 7.18 | 7.31 | 7.13 |
| | (3.8) | (3.8) | (3.8) | (3.8) |
| Pop % under 4 | 3.81 | 3.86 | 3.80 | 3.89 |
| | (1.8) | (1.9) | (1.8) | (1.9) |
| Fraction housewives 2001 Census | 0.33 | 0.34 | 0.33 | 0.34 |
| | (0.1) | (0.1) | (0.1) | (0.1) |
| **Parties** | | | | |
| PSOE mayor (%) | 34.87 | 37.82 | 34.26 | 35.01 |
| | (47.7) | (48.5) | (47.5) | (47.7) |
| PP mayor (%) | 39.88 | 40.69 | 39.71 | 41.05 |
| | (49.0) | (49.1) | (48.9) | (49.2) |
| Observations | 26255 | 4468 | 21787 | 8143 |
| Municipality*Elections | 10911 | 1826 | 9085 | 3347 |

*Notes:* Descriptives for all municipalities with populations over 250 for which budget data is available. Column 1 includes all municipalities, column 2 includes municipalities led by a female mayor, column 3 includes municipalities led by a male mayor and column 4 presents descriptives for municipalities with mixed races (male and female candidates run for election). A total of 62 municipalities had no budget data for the first legislature (2007-2010) and 104 of these municipalities had no budget data in the second legislature (2010-2014). Standard deviations of the selected variables presented in parentheses.

TABLE 2

MUNICIPAL DESCRIPTIVES BY PARTY

| | All | PP Mayor | PSOE Mayor |
|---|---|---|---|
| **Female Mayor** | 0.17 | 0.17 | 0.18 |
| | (0.4) | (0.4) | (0.4) |
| **Gender-sensitive Policies** | | | |
| Long-Term Care share | 0.61 | 0.64 | 0.67 |
| | (1.9) | (1.9) | (2.0) |
| Long-Term Care dummy | 0.22 | 0.22 | 0.23 |
| | (0.4) | (0.4) | (0.4) |
| Preschool share | 1.16 | 0.90 | 1.11 |
| | (2.2) | (1.9) | (2.2) |
| Preschool dummy | 0.36 | 0.30 | 0.35 |
| | (0.5) | (0.5) | (0.5) |
| Work-life balancing serv. share | 0.32 | 0.26 | 0.30 |
| | (0.9) | (0.8) | (0.9) |
| Work-life balancing serv. dummy | 0.25 | 0.21 | 0.23 |
| | (0.4) | (0.4) | (0.4) |
| **Demographics** | | | |
| Population 2007 (000s) | 8.65 | 10.68 | 7.83 |
| | (58.4) | (79.8) | (37.6) |
| Pop % above 80 | 7.29 | 7.79 | 7.32 |
| | (3.8) | (4.1) | (3.8) |
| Pop % under 4 | 3.81 | 3.52 | 3.72 |
| | (1.8) | (1.9) | (1.8) |
| Fraction homemakers 2001 Census | 0.33 | 0.36 | 0.35 |
| | (0.1) | (0.1) | (0.1) |
| Observations | 26255 | 10470 | 9155 |
| Municipality*Elections | 10911 | 4125 | 3844 |

*Notes:* All municipalities for which budget data is available. Column 1 includes all municipalities, column 2 includes municipalities led by a PP mayor, column 3 includes municipalities led by a PSOE mayor. Standard deviations of the selected variables presented in parentheses.

can hardly be given a causal interpretation. We will return to this in the next sections.

Finally, a summary of the characteristics of the municipalities that have positive spending on each policy category can be found in Table 3. The most frequently implemented policy is preschool education, which is also the policy with the highest average share of spending among municipalities that have it in place (3.36%). On the other hand, Work-Life Balancing services is the one with the smallest average spending share among those municipalities which offer it (1.36%). The percentage of observations with a positive spending share on LTC, Preschool, and Work-Life Balancing Services is 22%, 35.6% and 24.7% respectively. The three of them are observed in municipalities across the full range of populations.[9] The number of observations in this table refers to the number of year-municipality pairs with positive spending in the corresponding category.

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[9]The minimum is in all cases below 250 people. This may seem puzzling given that we restrict the sample to municipalities with more than 250 inhabitants. Notice that we implement this sample restriction in electoral years (when the different electoral rules above and below the 250 threshold apply). If a municipality loses population after the elections, we keep it in the sample for the intervening electoral period. Excluding these municipalities from our sample has no effect on qualitative results and a negligible effect on point estimates.

TABLE 3

MUNICIPAL DESCRIPTIVES BY POLICY ENGAGEMENT

|  | LTC | Preschool | Balancing Services |
|---|---|---|---|
| Avg Spending | 2.92% | 3.36% | 1.36% |
| Std Spending | 3.22% | 2.64% | 1.47% |
| Avg Population | 25,098 | 18,039 | 22,669 |
| Max Population | 3,273,049 | 3,273,049 | 3,273,049 |
| Min Population | 220 | 217 | 232 |
| Observations | 5,423 | 8,964 | 6,166 |
| Percentage of Observations | 22% | 35,6% | 24,7% |
| Municipality*Elections | 2,737 | 4,487 | 3,160 |

*Notes:* Column 1 includes all municipalities that spend in LTC, column 2 includes municipalities that spend in Preschooling, column 3 includes all municipalities that spend in Work and Family Life Balancing Services.

## 3. Empirical Analysis - Female Mayors and Gender Policies

### 3.1. Empirical Strategy – Close Election Regression Discontinuity

We now turn to study whether local governments led by female mayors are more likely to implement gender-sensitive policies – long-term care, pre-schooling or work and family life balancing services. We measure the degree of engagement of a municipality with a policy using either an indicator taking value 1 if there is any reported spending for that category in the municipal budget, or the share of the budget assigned to the policy in question. Descriptive statistics reported in Table 1 indicate that, on average, municipalities led by male or female mayors have very similar engagement with these policies. Yet several confounding factors may simultaneously affect both the probability of having a female mayor and the amount of spending in gender-sensitive programs. For example, municipalities populated by younger citizens may be more likely to vote for a party headed by a female leader and, simultaneously, demand more spending in pre-schooling. Conversely, municipalities where the hold of traditional or religious values is stronger may be less likely to elect a female mayor but simultaneously have higher demand for long-term care spending linked to disability or other health problems. In some cases, these confounding factors may be observable. But this will not be the case in general.

To correctly estimate the effect of interest, we need an alternative empirical strategy that is not subject to these sources of bias. For this purpose, we implement a close election regression discontinuity design (as in Lee (2001) and a large subsequent literature) focusing on *mixed races* – municipal elections in which a female candidate runs against a male candidate. Several studies have used this strategy to look at gender or party differences in policies (see for example Ferreira and Gyourko (2009), Beland (2015), Brollo and Troiano (2016) or Solé-Ollé and Viladecans-Marsal (2013)) and other political outcomes (Gagliarducci and Paserman, 2012). With this method, we exploit the stochastic nature of electoral outcomes to obtain quasi-random variation in the gender of the elected mayor. As a result, the identification assumptions invoked here are arguably weaker than those required in alternative strategies based on conditioning on observables or municipality fixed-effects.

Given the specificities of the Spanish electoral system, we define as *mixed* electoral races those in which first-in-the-list candidate of the most voted and second most voted parties are of different

gender. The person appointed as mayor is, with very few exceptions, the first-in-the-list candidate of one of the parties in the council. Henceforth, we refer to these individuals as the candidates of each party. Note that being the most voted candidate does not necessarily imply becoming mayor (see section 2.1). It is the council that ultimately elects this position depending on coalition and majority formation strategies.

We will use the female victory margin as the running variable in our regression-discontinuity design. Focusing on mixed races only, we define this victory margin as $FemaleVoteMargin_{it} = VS_{it}^F - VS_{it}^M$ where $VS_{it}^F$ is the vote share of the female candidate, and $VS_{it}^M$ the vote share of the male candidate. Given that the probability of having a female mayor is not zero below the threshold nor one above it, we estimate the effect of mayoral gender on policies using a *Fuzzy* RD design. That being said, we will also report results for the reduced-form effect of a female candidate winning the election.

The first stage in our IV estimation is:

$$Female_{it} = \pi_0 + \pi_1 \mathbf{1}(FemaleVoteMargin_{it} > 0) + f(FemaleVoteMargin_{it}) + \gamma_1' X_{it} + \eta_t + u_{it} \quad (1)$$

Where variable $Female_{it}$ is a dummy taking value 1 if municipality $i$ is ruled by a female mayor in year $t$, $\mathbf{1}(FemaleVoteMargin_{it} > 0)$ is a dummy taking value 1 if the party headed by a female candidate was the most voted party in the 2007 or 2011 municipal election and $f(FemaleVoteMargin_{it})$ are polynomials in the female vote margin fitted at both sides of the threshold. Estimates are reported using first-order polynomials estimated on either side of the threshold value as suggested in Imbens and Lemieux (2008). Vector $X_{it}$ corresponds to our set of controls including population, the fraction of citizens with a college degree, the fraction of population above 80 years of age, the share of debt as a proportion of current spending in the municipal budget and the fraction of women of working age not in the labour force. Year dummies are represented by $\eta_t$. These controls are included to reduce residual variance and improve estimate precision. The second stage is given by:

$$Policy_{it} = \alpha + f(FemaleVoteMargin_{it}) + \delta Female_{it} + \gamma_2' X_{it} + \eta_t + \epsilon_{it}$$

We provide results using positive spending dummies and spending shares ($PolicyShare_{it}$) as dependent variables. The parameter of interest in each specification is $\delta$, capturing the effect of having a female mayor on either spending shares (intensive and extensive margins) or a program spending dummy (extensive margin only). Given that most municipalities only change mayors in election years, we cluster standard errors at the municipality-term level.

To ensure that our estimates for $\delta$ are obtained from variation at the threshold, we will restrict the sample to a bandwidth of the running variable around 0. A large methodological literature on the implementation of RD designs suggests different methods to choose i) the bandwidths around the threshold which determines the discontinuity in treatment probability, ii) length of the polynomials in function $f(.)$ and iii) varying weights given to observations around the cut-off. In our paper we follow Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016) both in selecting the optimal bandwidth for our regressions and for conducting inference based on this optimal bandwidth selection.

## 3.2. Validity of Research Design

Before discussing the results for the effects of gender on the outcomes of interest, a few notes are due regarding the validity of the regression discontinuity design in this context. Figure 1 illustrates our first stage: the probability of having a female mayor jumps substantially – by roughly 50 percentage points – when *FemaleVoteMargin$_{it}$* goes above 0. This is confirmed in Table 4, which reports first stage regression coefficients by estimating equation 1 for different sub-samples corresponding to observations at different bandwidths around the threshold. These bandwidths vary between 14% and 20.4% and correspond to the optimal bandwidths selected for the second stage regressions reported below. We observe that the probability of having a female mayor jumps by between 45 and 47 percentage points at the threshold, a magnitude comparable to the jump displayed in Figure 1. The instrument is strong, with the associated F-statistics being in all cases above 100. Different bandwidths lead to very similar estimates. This is also the case across specifications without controls or for higher order polynomials in the running variable (not shown).

### FIGURE 1
### FIRST STAGE - FEMALE MAYOR
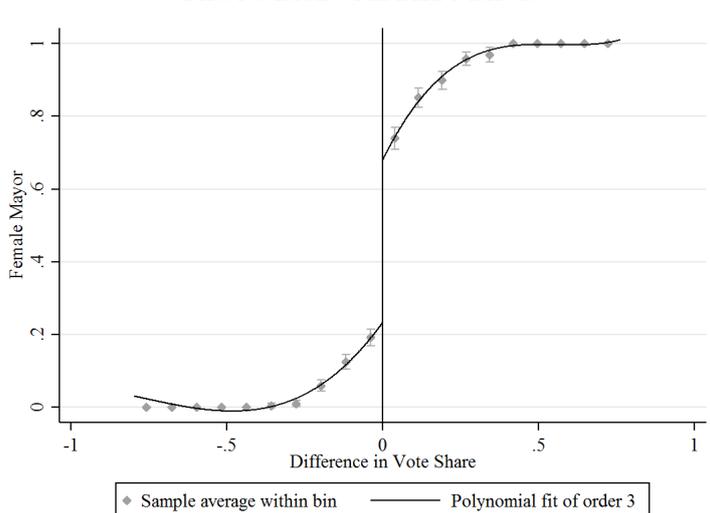


*Note*: Graph indicates the discontinuity in the probability of having a female mayor when a woman barely wins a mixed race. Vertical axis represents fraction of female mayors. Horizontal axis represents female winning vote share margin, negative if female candidate lost election. Points indicate averages within bins of the female victory margin. Line represents a third-degree polynomial fit on either side of the threshold value. Sample restricted to mixed races.

Figure 2 shows the histogram of the running variable *FemaleVoteMargin$_{it}$*. The distribution of *FemaleVoteMargin$_{it}$* has no obvious discontinuity at the 0 threshold. A McCrary test estimating a discontinuity in mass around the threshold yields a p-value of 0.27. We interpret this as indicating there is no perfect manipulation of the running variable, which is one of the key assumptions required for the validity of RD designs. This result is not particularly surprising as it is unlikely that political parties can affect vote shares in a deterministic way (in the vast majority of cases, pre-election polls are not available at the municipal level).

An additional assumption required for the validity of this estimation method is that no municipal characteristics (other than the mayor's gender) jump discontinuously at the threshold. To explore this, we use our set of controls $X_{it}$ and estimate whether these vary discontinuously when

TABLE 4

FIRST STAGE - FEMALE MAYOR

| | (1) Female Mayor | (2) Female Mayor | (3) Female Mayor | (4) Female Mayor | (5) Female Mayor | (6) Female Mayor |
|---|---|---|---|---|---|---|
| Female Winner | 0.461*** | 0.458*** | 0.457*** | 0.474*** | 0.474*** | 0.452*** |
| | (0.0394) | (0.0414) | (0.0424) | (0.0364) | (0.0374) | (0.0433) |
| Observations | 3952 | 3628 | 3479 | 4503 | 4352 | 3315 |
| Clusters | 1686 | 1548 | 1485 | 1918 | 1855 | 1415 |
| F-stat | 137 | 122 | 116 | 168 | 160 | 108 |
| Bandwidth | 0.174 | 0.155 | 0.148 | 0.204 | 0.195 | 0.141 |

*Note*: Sample restricted to mixed races. The dependent variable is a Female Mayor dummy in all columns and the coefficient displayed corresponds to a dummy taking value 1 if a party headed by a female candidate won the municipal election. In all columns we report results obtained restricting the sample to bandwidths around the female vote margin threshold selected using the methods described in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016). Columns 1 through 3 do not include controls other than the running variable on either side of the threshold. Columns 4 to 6 include controls as described in the text. Standard errors clustered at the level of town-electoral period.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

the female candidate wins the election by a small margin. The results of the exercise are illustrated in the graphs provided in Figure 3. We observe that in none of these variables there is a substantial discontinuity at the threshold. To test this formally, we also provide results using local linear regressions in Table B.2. In all cases, we observe that the coefficients measuring discontinuities in the covariates are small and not significantly different from zero.[10]

Finally, interpretation of the effects of interest will depend on whether other characteristics of the mayors themselves vary with gender. Female politicians are a selected sample of the population of local level officials in Spain. For example, given the traditionally low involvement of women in Spanish politics, female politicians are typically younger than their male counterparts. Other differences in, for example, prior occupation or educational levels may also exist. An increase in female participation in politics will bring with it a relative over-representation of the characteristic features of this population. If we are seeking to understand the implications of increased female participation in politics, it is important to allow these variables to vary at the threshold too, as they are part of the bundle of characteristics that changes with this increased participation. Hence, we do not control for these variables in our regressions. We do however explore whether these characteristics jump at the threshold. We report estimates for the change at the threshold in mayors' characteristics – age and indicators for blue collar workers, white collar workers, college graduates, mayors with no previous reported studies, previously unemployed, and housekeepers – in Appendix Table B.3. We observe that coefficients are in most cases not statistically significant, although some are imprecisely estimated. The reduced-form effect of having a female winner only has a statistically significant effect on the mayor's age.

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[10]Estimates reported in Table B.2 are based on optimal bandwidths calculated separately for each outcome variable. Alternative estimates using the second stage bandwidths lead to very similar results and are available upon request.

## FIGURE 2
### RUNNING VARIABLE HISTOGRAM - FEMALE MAYOR
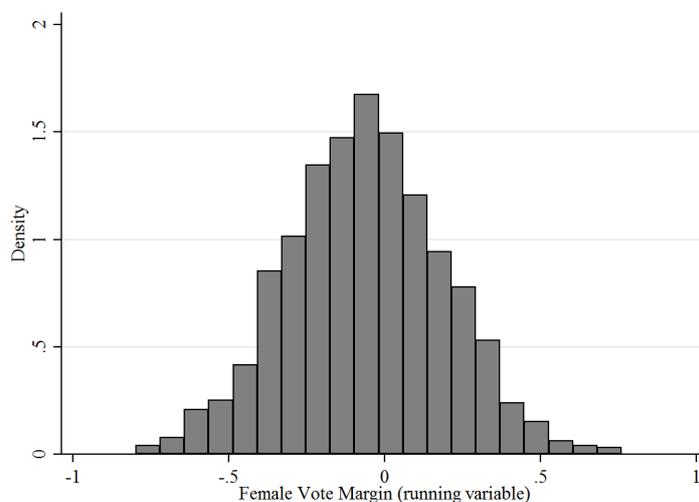


*Note*: Distribution of female vote shares in mixed races in the 2007 and 2011 elections (pooled). Smooth density estimated with an Epanechnikov kernel.

## FIGURE 3
### FEMALE MAYOR - COVARIATES
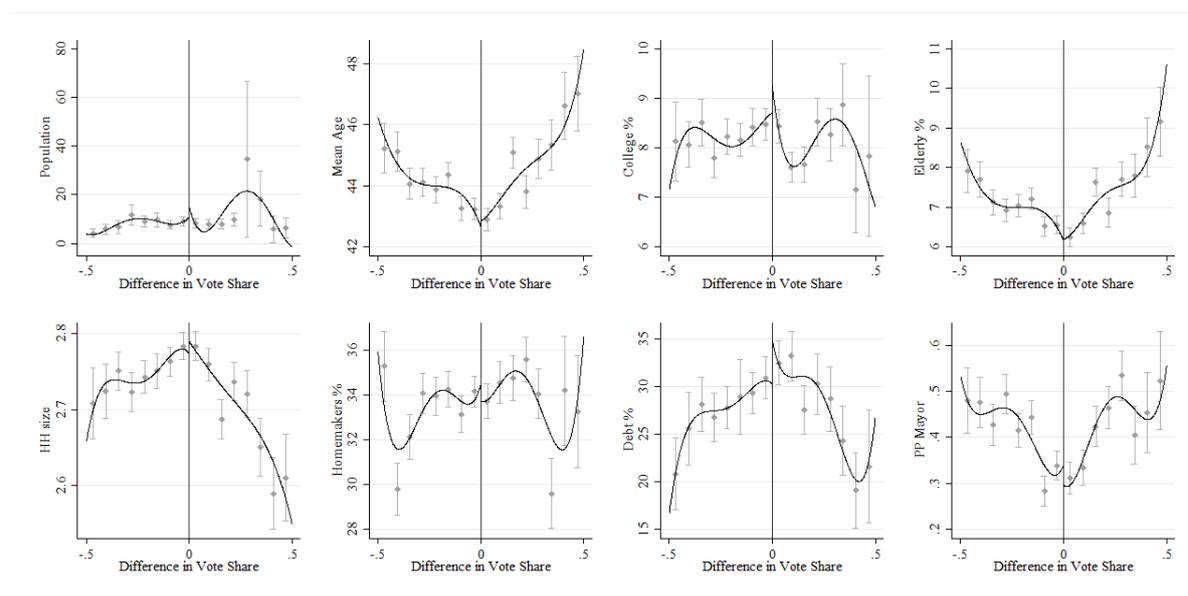


*Note*: Horizontal axis represents the vote share difference between the most voted female and male candidates. From left to right and top to bottom the outcome variables are population, mean age, fraction with college education, percentage of population above 80 years of age, average household size (2001 census), percentage of female homemakers (2001 census), percentage of outstanding debt of municipality before sample period, and probability of having a PP (centre-right) mayor. Solid lines represent third degree polynomials in the running variable estimated separately for positive and negative polynomials. Gray dots correspond to averages for bins of the running variable. Vertical lines correspond to 95% confidence intervals around these averages.

*3.3. Results*

We turn now to reporting the main results of this paper, pertaining to the effects of mayoral gender on our set of gender-related policies as estimated using our RD specification. We start by providing reduced-form graphs measuring the effect of a female candidate victory in the elections on extensive and intensive margin variables measuring the municipality's engagement with these policies. These are reported in Figure 4. The first row corresponds to dummy outcomes indicating whether a municipality recorded any spending in the specified category or not. In the second row, the vertical axes measure the share of total municipal spending assigned to each policy. Solid lines represent third degree polynomials in the running variable estimated separately on both sides of the threshold. Gray dots correspond to average values of the outcome calculated for bins of the running variable. Segments around these dots represent 95% confidence bands.

FIGURE 4

FEMALE MAYOR & GENDER POLICIES - REDUCED-FORM GRAPHS
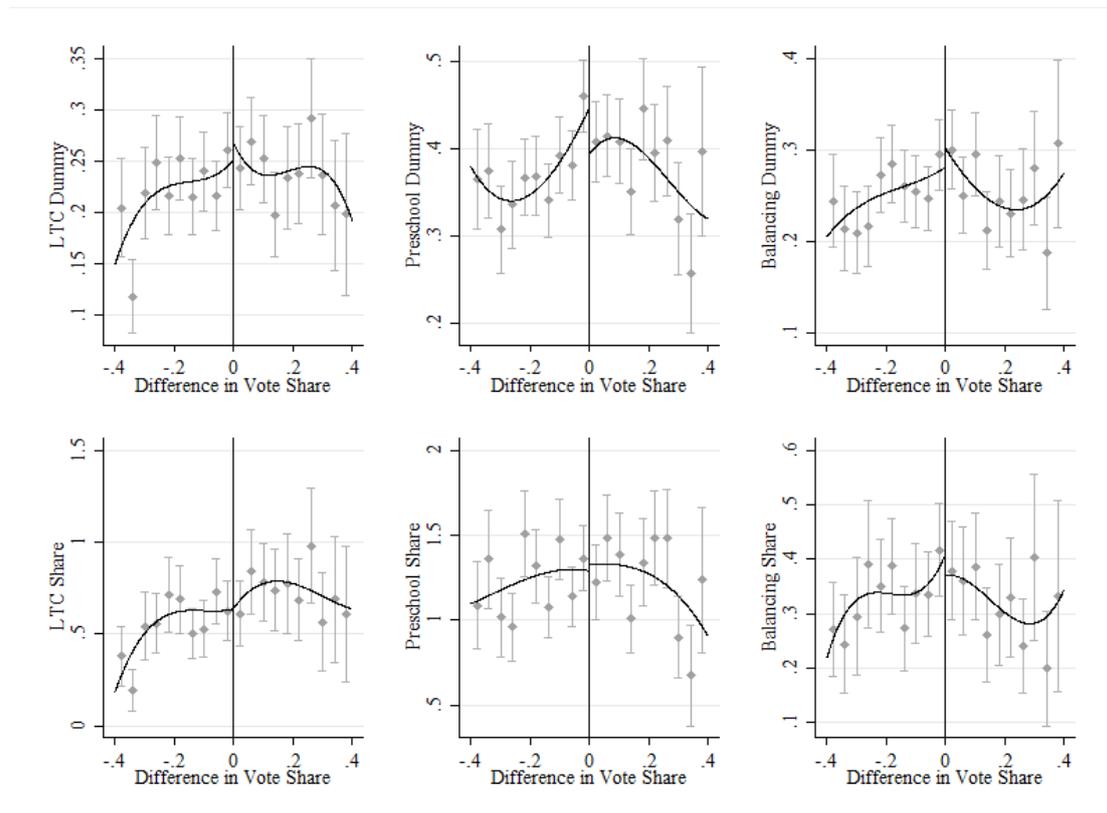


*Notes*: Horizontal axis represents the vote share difference between the most voted female and male candidates. Vertical axes correspond to a long term care spending dummy (top-left), a preschool spending dummy (top-center), a balancing services spending dummy (top-right), the share of LTC spending (bottom-left), the share of Preschool spending (bottom-center) and the share of balancing service spending. Spending shares calculated relative to total municipal spending in that year. Solid lines represent third degree polynomials in the running variable estimated separately for positive and negative polynomials. Gray dots correspond to averages for bins of the running variable. Vertical lines correspond to 95% confidence intervals around these averages.

Visual inspection of Figure 4 reveals that discontinuities at the threshold are generally small, and often imperceptible. A slight discontinuity is observed in the case of the preschool dummy

outcome, perhaps indicating a *negative* effect of female mayors on this policy. Yet reduced-form coefficients reported in Appendix Table B.4 indicate that none of these differences are statistically significant.

To find the effects of *Female$_{it}$* on our set of gender-policies we need to re-scale these reduced-form coefficients by the change in the probability of having a female mayor at the threshold. Table 5 reports these estimates using our discrete and continuous outcomes. Results for LTC policies are presented in columns 1 and 4. Results for preschool education are presented in columns 2 and 5. Finally, results for work and family life balancing services are presented in columns 3 and 6. Estimates in columns 1 through 3 are obtained using a discrete outcome, and those in columns 4 through 6 correspond to continuous outcomes.

TABLE 5

2ND STAGE - FEMALE MAYOR

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Dependent Variable: | D. LTC | D. Preschool | D. Bal. Serv. | LTC (%) | Preschool (%) | Bal. Serv. (%) |
| | | | | | | |
| Female Mayor | 0.0226 | -0.0874 | 0.0376 | 0.0831 | 0.0988 | -0.0253 |
| | (0.0799) | (0.0928) | (0.0882) | (0.362) | (0.389) | (0.208) |
| | | | | | | |
| Observations | 3935 | 3583 | 3413 | 4274 | 4264 | 2727 |
| Clusters | 1679 | 1530 | 1457 | 1854 | 1855 | 1415 |
| p-value | 0.777 | 0.346 | 0.670 | 0.818 | 0.800 | 0.903 |
| Bandwidth | 0.173 | 0.152 | 0.145 | 0.195 | 0.195 | 0.141 |

*Note*: Dummies take value one if spending in Long-term Care, Pre-schooling and Work and Family Life Balancing Services respectively is above zero. Shares equal spending share over total budget in Long-term Care, Pre-schooling and Work and Family Life Balancing Services. In columns we report local linear regressions with uniform kernel and polynomials of order 1 fitted at the two sides of the discontinuity. Standard errors clustered at the level of town-electoral period.

$^{*}$ p < 0.1, $^{**}$ p < 0.05, $^{***}$ p < 0.01.

We observe none of these coefficients is significantly different from 0. When using our continuous outcomes (in columns 4 to 6) we find tightly estimated zeroes. In the case of LTC, the 0.0831 coefficient indicates that having a female mayor increases the fraction of the budget devoted to long-term care spending by less than a tenth of a percentage point. This is less than one-twentieth of a standard deviation in the LTC share (see Table 1). The point estimate for pre-school spending is roughly 0.1, indicating that having a female mayor increases the share of the budget devoted to these policies by 0.1 percent. This continues to be less than 1/10 of a standard deviation in the outcome. We arrive at a similar conclusion when looking at the effect on the spending share for work and family-life balancing services which is negative and small. Coefficients in columns 1 to 3 correspond to our binary outcomes. We also find none of the coefficients are statistically significant. The point estimate for LTC would indicate that having a female mayor increases the probability of engaging in LTC policies by 2.3 percentage points. This is roughly one tenth of the baseline probability of conducting LTC spending (see Table 1). In the case of balancing services, the coefficient is of similar magnitude, standing at 1/8 of the baseline probability. The only point estimate that is somewhat larger in magnitude is that for the effect of having a female mayor on pre-schooling, but it has a negative sign. Given the limited variation in our discrete outcomes, these estimates are somewhat imprecise. We address potential concerns about estimate precision

of these estimates in the next section.

We conclude from these results – and the robustness results below – that female leaders do not promote gender-sensitive policies at the local level. These findings contrast with those in the literature studying developing countries, which reports large and significant effects of politician's gender on gender sensitive policies in India (see Clots-Figueras (2012), Chattopadhyay and Duflo (2004)). It is important to highlight that the policies at hand, the social norms around gender and the institutional design of the political competition are very different from those considered here. Our results are consistent with those reported in Ferreira and Gyourko (2014) for US mayors, which find no effect of politicians' gender on broadly defined policy areas which are not necessarily gender specific. For completeness, we have also tested whether the gender of the politician affects broad spending categories such as basic services, social spending, transport or health spending among others. We continue to find little evidence of systematic and interpretable differences in broad policy categories by gender with only a few marginally significant coefficients, probably resulting from multiple testing rather than substantial and robust policy heterogeneity.

As shown below, we do find differences in gender policies by mayoral party. This indicates that other characteristics of leaders and/or governments may matter for the implementation of these policies. Before turning to those results, we discuss the robustness of our estimates for mayoral gender in the following.

### 3.4. Panel Estimates & Robustness Checks

Our baseline specification follows the approach proposed in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016) with a first order polynomial fitted at each side of the threshold, a uniform kernel and a set of controls to reduce variance and increase the precision of our estimates. Nonetheless, results are robust to the inclusion of higher order polynomials, triangular kernels and the exclusion of controls. As an illustration, Table B.6 in the appendix shows the results for the main regression without controls and with a second order polynomial. All coefficients are non-significant and of a similar order of magnitude than those reported in table 5.

We can also test whether the main conclusions of our analysis are confirmed using a different empirical strategy. For that purpose, we exploit the longitudinal variation in a panel specification with municipality fixed effects, which relies on a different set of assumptions for identification.[11] We estimate:

$$Policy_{it} = \beta Female_{it} + \gamma pop_{it} + \eta_t + \alpha_i + \epsilon_{it}$$

where $Female_{it}$ and $\eta_t$ are defined as above, $pop_{it}$ is the population of municipality $i$ in year $t$, and $\alpha_i$ represents the municipality fixed-effect. Following the analysis in the previous sections, $Policy_{it}$ can represent binary or continuous outcomes for the three policies of interest, measuring the extensive and intensive margins of policy implementation. By including $\alpha_i$ in our model we want to account for possible fixed unobserved factors at the municipal level that could be simultaneously correlated with the gender of the mayor and with spending in gender-sensitive policies (long-term care, preschool services or life-balancing services). For example, more progressive vot-

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[11]The identifying assumptions of the RDD are discussed in section 3.2. The validity of the panel approach requires that the unobserved characteristics of municipalities that correlate both with the dependent and the independent variables remain constant throughout the estimation period.

ers may be more likely to vote for a party headed by a female leader and, simultaneously, demand more spending in gender-sensitive long-term care policies. As long as these confounding factors are fixed over time, we can exploit the longitudinal variation in the gender of mayors to estimate the effect of gender on these policies. Given that most of the controls included in vector $X_{it}$ defined above are fixed over time, we only include population as a control in our panel specifications.

Results from our fixed effects specification for the six outcomes of interest are reported in Appendix Table B.5. Again, all coefficients are statistically indistinguishable from zero. For our discrete outcomes, the coefficients are in fact negative and very close to zero. For example, the point estimate in column 1 would indicate that having a female mayor decreases the probability of engaging in long-term care policies by a puny 0.1 percentage points (the baseline probability is over 220 times larger). In the case of our continuous outcomes, our coefficients of interest are also smaller than those reported in Table 5.

The additional precision of these estimates is likely to result from the larger sample used in our longitudinal analysis. Another distinctive feature of these estimates is that they do not rely on the selected sample of competitive races that are used in the close election RD reported above. In competitive scenarios, governments or politicians may be less able to exercise discretion in the implementation of policies. This is to say that the local average treatment effects reported above may not be informative about effects of gender on policy far from the threshold. Yet even when relatively uncompetitive elections are included in the analysis, we continue to find no evidence of an effect of mayoral gender on our policies of interest.

## 4. Political Parties and Gender Policies

A possible explanation to the null results in previous sections is that municipal governments have no discretion to implement gender-sensitive policies. While local governments do have legal authority to implement these policies, it is possible that budgetary or political constraints preclude them from exercising this authority. To investigate this, we estimate whether *political parties* have an effect on the implementation of gender-sensitive policies. If parties operate as policy platforms, then the party in power may determine the policies implemented by the local government. Moreover, if parties can influence policies, then this means that municipal governments have some discretion in policy-making.

We test whether parties differ in their application of gender sensitive policies by conducting a close election *fuzzy* RD similar to the one described in section 3.1. Specifically, we test whether municipalities ruled by center-right PP where less likely to engage in gender-sensitive policies. We chose *Partido Popular* as our reference party because it was the party controlling the largest number of municipal governments in our sample and had been the most important centre-right party in Spain since the late 1980s. Our running variable for this analysis $PPVoteMargin_{it}$ is defined as the difference between the vote share of PP and the vote share of the most voted party (other than PP). This difference is positive when PP was the most voted party in a municipal election and negative if some other party beat PP in the polls. Using the aforementioned variable we estimate:

$$Policy_{it} = \sigma + f(PPVoteMargin_{it}) + \delta_{PP}PP_{it} + \theta_2'X_{it} + \omega_t + \epsilon_{it}$$

where $PP_{it}$ is a dummy variable taking value 1 if the municipality is ruled by the *Partido Pop-ular* and *Policiy$_{it}$* can represent a dummy or a share for the three spending categories of interest, capturing the extensive or the extensive and intensive margins respectively. As in the analysis for gender we use $1(PPVoteMargin_{it} > 0)$ as an instrument for $PP_{it}$. Also mimicking what we did for gender, we use first-order polynomials in the vote margin fitted at both sides of the threshold, a time fixed effect, and the same set of controls described in section 3.1. Estimates are obtained using local linear regressions with optimal bandwidth and inference as in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016).

This empirical strategy is suitable to study the effect of partisan affiliation on gender-policies. The instrument is strong, with Appendix Figure A.3 showing a sizeable jump at the threshold in the first stage, and Table B.1 reporting the first stage coefficients for the four optimal bandwidths used in the second stage. The first stage estimates are all highly significant (F-statistic above 300) and the estimated jump in probabilities is always above 53%. An histogram of the running variable is displayed in Appendix Figure A.1 and suggests no manipulation at the threshold. In addition, Appendix Figure A.2 shows no substantial discontinuity for covariate values at the threshold, a result confirmed by the insignificant coefficients reported in in Table B.7. Taken together, these results indicate our RD strategy is suitable to estimate the impact of parties on gender-sensitive policies.

Visual inspection of the first row in Figure 5 shows an apparent discontinuity at the threshold in the extensive margin of the three policies under study. Estimation results for the effect of having a PP mayor on the likelihood of implementing a gender policy are presented in Table 6. We find statistically significant effects indicating that mayors from the centre-right PP are less likely to engage in gender-friendly policies. Results are robust to using quadratic polynomials on both sides of the threshold and using different weighting kernels. The coefficients are significant both statistically and economically, with a negative effect in the probability of spending above 10 percentage points in any of the categories. This is a substantial number in relation to the proportion of municipalities spending in long-term care (22%), life-balancing services (25%) and preschooling (36%). Moreover, coefficients are larger in absolute value than those reported in columns 1 through 3 of table 4. We do not find a significant effect of conservative mayors on intensive margin measures of the amount spent on the policies of interest, suggesting that parties decide whether or not to engage in a policy but might face tighter constraints in the amount of spending committed once the extensive margin decision is made.

These results show that the commitment to gender-sensitive policies may vary with charac-teristics of the local government. Furthermore, they indicate that, in this context, centre-right parties are less likely to engage in gender-sensitive policies. While a growing literature in public economics has focused on estimating the effect of parties on policies, to our knowledge these are the first estimates indicating partisan differences in gender-sensitive policies.[12] Arguably, these differences in policy can have important implications for voters. The weaker commitment of right-wing parties to gender policies has been mentioned as one of the explanations for the increase in the fraction of women voting left over the last decades (see Iversen and Rosenbluth 2006, Edlund

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[12]Studies of partisan differences in policies include Pettersson-Lidbom (2008), Ferreira and Gyourko (2009), Freier and Odendahl (2015) and Solé-Ollé and Viladecans-Marsal (2013).

FIGURE 5

POLITICAL PARTY & GENDER POLICIES - REDUCED-FORM GRAPHS
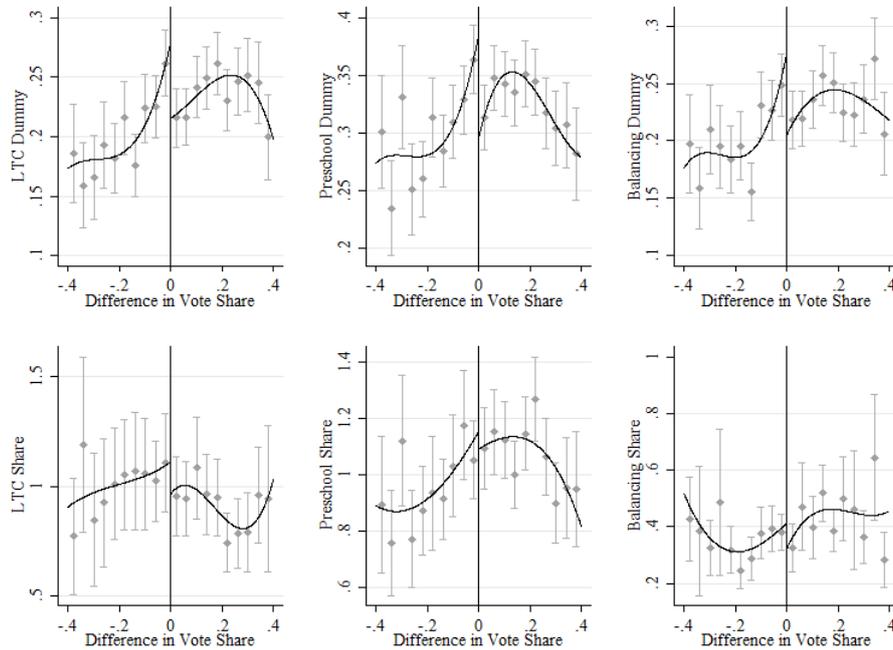
*Note*: Horizontal axis represents the vote share difference between PP and the most voted party other than PP. Vertical axes correspond to a long term care spending dummy (top-left), a preschool spending dummy (top-center), a balancing services spending dummy (top-right), the share of LTC spending (bottom-left), the share of Preschool spending (bottom-center) and the share of balancing service spending. Spending shares calculated relative to total municipal spending in that year. Solid lines represent third degree polynomials in the running variable estimated separately for positive and negative polynomials. Gray dots correspond to averages for bins of the running variable. Vertical lines correspond to 95% confidence intervals around these averages.

and Pande 2002). It may also explain the emerging gender gap in favour of PSOE (center-left party) in Spanish general elections of 2011 (CIS, 2011).

In order to evaluate the robustness of our results for partisan differences in policy, we provide alternative estimates based on a panel fixed effects specification. These are reported in Appendix Table B.8. We find negative and significant effects of our PP mayor dummy on extensive margin measures of engagement with long-term care and life-balancing services, echoing the RD results in Table 6. However, in contrast with the RD results, they suggest a positive and marginally significant effect on the probability of spending in pre-schooling. Insofar as the identifying assumptions involved in our RD estimates are weaker, we believe they provide a more trustworthy measure of this effect.

## 5. Conclusions

Despite enduring gaps in political representation between women and men, the fraction of female politicians leading in public office has increased in most countries. This change in representation came alongside an increase in the debate about gender policies; policies specifically

## Table 6
### 2nd Stage - PP Mayor

| Dependent Variable: | (1) D. LTC | (2) D. Preschool | (3) D. Bal. Serv. | (4) LTC (%) | (5) Preschool (%) | (6) Bal. Serv. (%) |
|---|---|---|---|---|---|---|
| PP Mayor | -0.111** | -0.107* | -0.104** | -0.261 | 0.0388 | -0.0983 |
| | (0.0487) | (0.0550) | (0.0485) | (0.387) | (0.269) | (0.168) |
| Observations | 7441 | 6693 | 7403 | 7461 | 7721 | 6665 |
| Clusters | 3053 | 2741 | 3038 | 3061 | 3160 | 2731 |
| p-value | 0.023 | 0.051 | 0.032 | 0.500 | 0.885 | 0.559 |
| Bandwidth | 0.149 | 0.134 | 0.149 | 0.150 | 0.155 | 0.133 |

Dummies take value one if spending in Long-term Care, Preschooling and Work and Family Life Balancing Services respectively is above zero. Shares equal spending share over total budget in Long-term Care, Preschooling and Work and Family Life Balancing Services. In columns we report local linear regressions with uniform kernel and polynomials of order 1 fitted at the two sides of the discontinuity. Standard errors clustered at the level of town-electoral period.

$^{*}$ p < 0.1, $^{**}$ p < 0.05, $^{***}$ p < 0.01.

directed to reduce gender inequalities in, for example, labour market participation. A natural emerging question in this context is whether increased female representation fosters the application or extension of these gender-sensitive policies.

We have sought to answer this question using budget information on municipal spending for Spanish local governments. The Spanish context has several advantages because recent policy changes have expanded the set of discretionary gender-related policies available for municipal governments and made disaggregated administrative data on spending available for research. In our analysis, we have implemented close-election regression discontinuity design and a panel specification, both leading to broadly consistent conclusions. Our findings indicate that there is no evidence that female mayors are more likely to engage in gender sensitive policies. The evidence is robust across specifications and policies, with estimated coefficients being insignificant and small. This result contrasts with some of the previous studies which had focused on context-specific gender policies in developing countries and found strong effects (Chattopadhyay and Duflo, 2004; Clots-Figueras, 2012). Relative to these papers, we look at a professional and competitive political setting in which male and female politicians compete for the position of mayor. Moreover, we focus on policies that are relevant to developed countries and generally accepted as gender sensitive: preschooling, long-term care and work and family life balancing services. We find no evidence of female mayors favouring these policies. On the other hand, we do find a significant difference in spending decisions in such policies related to political affiliation.

From an institutional design perspective, our results suggest that establishing a gender quota will not be enough to promote gender sensitive policies. Gender equality and women empowerment constitutes one of the UN goals in their 2030 Agenda, which stresses the public provision of care-giving services as one of its six specific targets. Our analysis focuses on widespread care-giving policies and shows that female political leadership at the local level might not suffice to ensure their provision. Instead, national (or European) laws establishing a level of minimum coverage could be more effective at granting universal access to these services.

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# Appendices

## A.  Appendix Figures

RUNNING VARIABLE HISTOGRAM - PP MAYOR
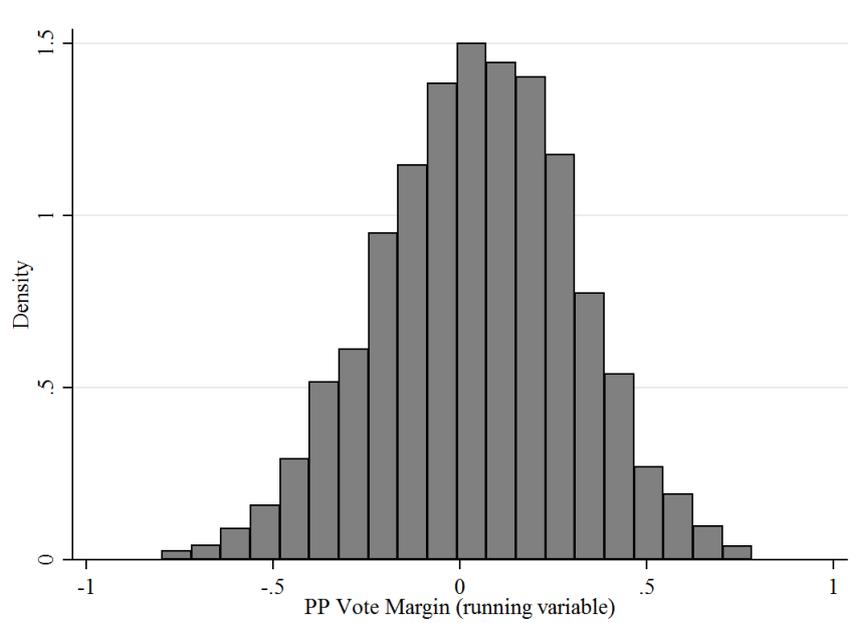


*Notes:* Distribution of PP candidate winning vote share margin, negative if PP lost election, in the 2007 and 2011 elections (pooled). Smooth density estimated with an Epanechnikov kernel.

## Figure A.2
### PP Mayor - Covariates
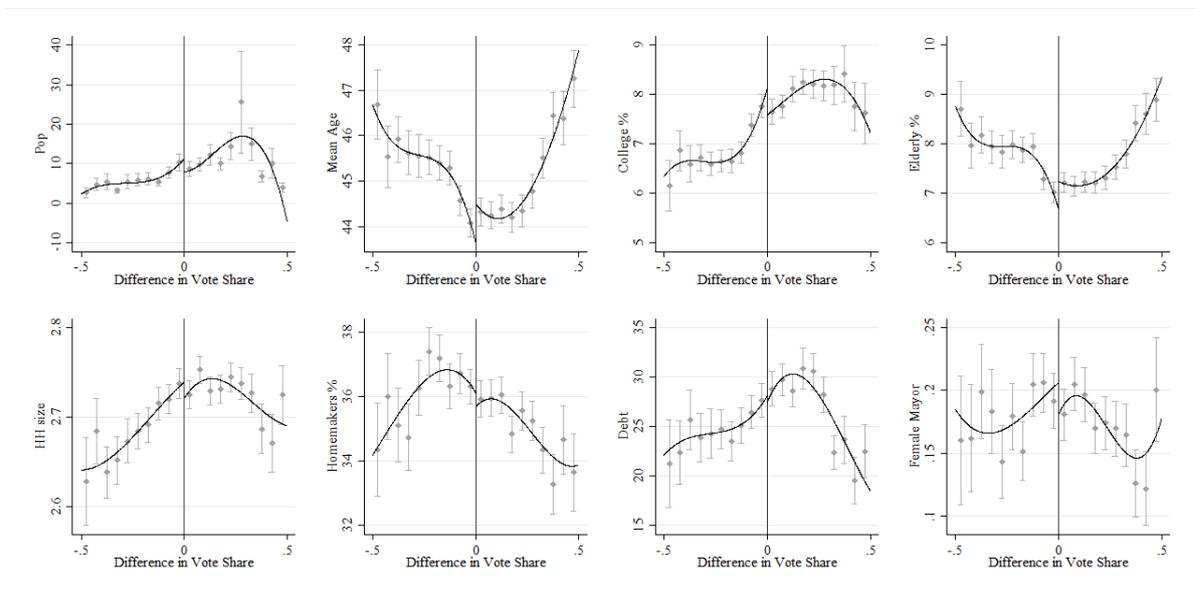


*Note*: Horizontal axis represents the vote share difference between PP and the most voted party other than PP. From left to right and top to bottom the outcome variables are population, mean population age, fraction with college education, percentage of population above 80 years of age, average household size (2001 census), percentage of female homemakers (2001 census), percentage of outstanding debt of municipality before sample period, and probability of having a female mayor. Solid lines represent third degree polynomials in the running variable estimated separately for positive and negative polynomials. Gray dots correspond to averages for bins of the running variable. Vertical lines correspond to 95% confidence intervals around these averages.
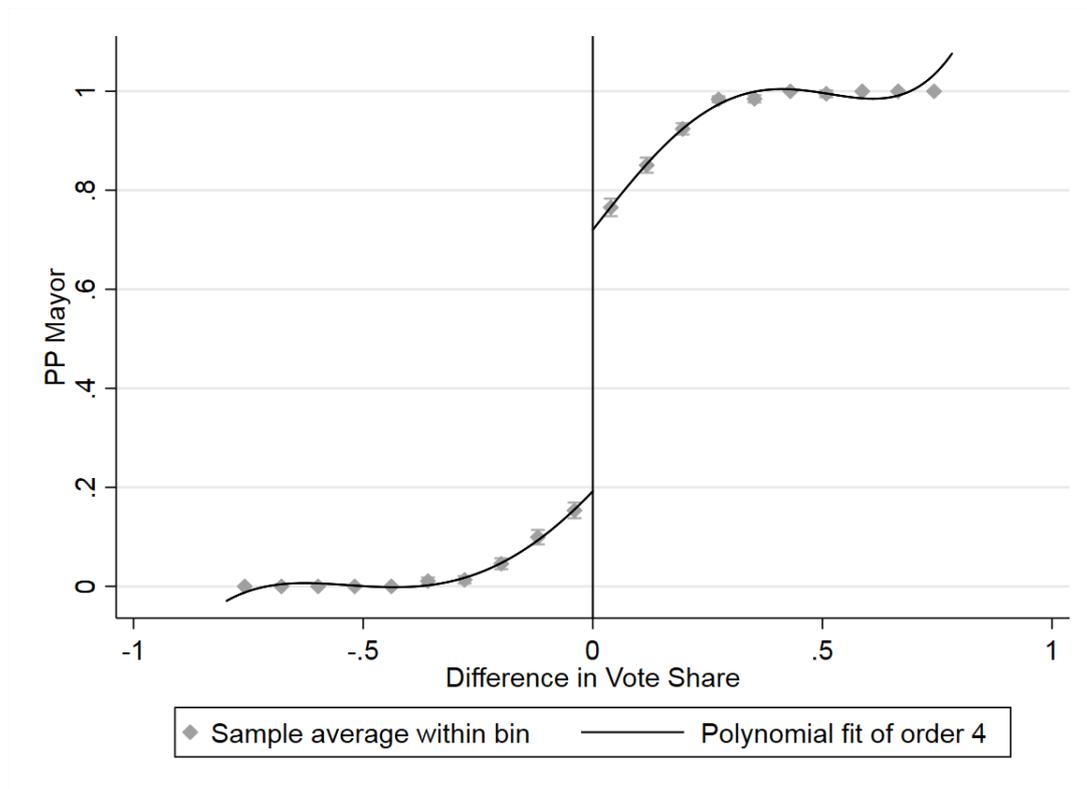
*Notes:* Graph indicates the discontinuity in the probability of having a mayor from Partido Popular (PP) when a PP candidate barely wins an election. Vertical axis represents fraction of PP mayors. Horizontal axis represents PP candidate winning vote share margin, negative if PP lost election. Points indicate averages within bins of the PP victory margin. Line represents a quartic fit on either side of the threshold value.

## B. Appendix Tables

### TABLE B.1
#### FIRST STAGE - PP MAYOR

|  | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
|  | PP Mayor | PP Mayor | PP Mayor | PP Mayor |
| PP Winner | 0.531*** | 0.533*** | 0.536*** | 0.545*** |
|  | (0.0265) | (0.0274) | (0.0285) | (0.0298) |
| Observations | 7914 | 7476 | 6995 | 6548 |
| Clusters | 3237 | 3067 | 2868 | 2684 |
| F-stat | 401.000 | 379.000 | 352.000 | 335.000 |
| Bandwidth | 0.160 | 0.150 | 0.140 | 0.130 |

Sample restricted to mixed races. The dependent variable is a PP Mayor dummy in all columns and the coefficient displayed corresponds to a dummy taking value 1 if a party headed by a PP candidate won the municipal election. In all columns we report results obtained restricting the sample to bandwidths around the female vote margin threshold selected using the methods described in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016). Columns 1 through 3 do not include controls other than the running variable on either side of the threshold. Columns 4 to 6 include controls as described in the text. Standard errors clustered at the level of town-electoral period.

$^{*}$ $p < 0.1$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$.

### TABLE B.2
#### FEMALE MAYOR - COVARIATES MUNICIPALITY

|  | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Dependent Variable: | Population | Mean Age | College % | Elderly % | HH size | Homemakers % | Debt % | PP Mayor |
| Female Winner | -3.544 | -1.048 | -0.236 | -0.875 | 0.0507 | -0.955 | 6.506 | -0.0671 |
|  | (7.695) | (1.018) | (0.953) | (0.674) | (0.0548) | (2.090) | (6.976) | (0.103) |
| Observations | 1877 | 4381 | 4378 | 4202 | 4166 | 4416 | 4215 | 3817 |
| Clusters | 794 | 1866 | 1866 | 1792 | 1777 | 1882 | 1798 | 1628 |
| p-value | 0.645 | 0.303 | 0.804 | 0.194 | 0.354 | 0.648 | 0.351 | 0.514 |
| Bandwidth | 0.073 | 0.197 | 0.197 | 0.188 | 0.186 | 0.199 | 0.189 | 0.166 |

Dependent variables: (1) Population, (2) Mean age of inhabitants in the municipality, (3) Proportion of inhabitants with a college degree, (4) Fraction of inhabitants older than 80, (5) Household Size, (6) Fraction of female homemakers in 2001 census, (7) Outstanding debt share of municipal budget in 2009, and (8) PP (centre-right) mayor dummy. In columns we report local linear regressions with uniform kernel and polynomials of order 1 fitted at the two sides of the discontinuity. Standard errors clustered at the level of town-electoral period.

$^{*}$ $p < 0.1$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$.

TABLE B.3

FEMALE MAYOR - COVARIATES MAYOR

| Dependent Variable: | (1) Age | (2) B Collar | (3) W Collar | (4) College | (5) No Studies | (6) Unemp. | (7) Houskp. |
|---|---|---|---|---|---|---|---|
| Female Winner | -4.820** | 0.0572 | 0.0461 | 0.152 | -0.0367 | 0.00287 | 0.0603 |
| | (2.090) | (0.0857) | (0.113) | (0.112) | (0.0337) | (0.0205) | (0.0373) |
| | | | | | | | |
| Observations | 3542 | 3564 | 3246 | 3377 | 4115 | 4384 | 3523 |
| Clusters | 1792 | 1522 | 1385 | 1442 | 1756 | 1867 | 1504 |
| p-value | 0.021 | 0.504 | 0.684 | 0.173 | 0.276 | 0.889 | 0.106 |
| Bandwidth | 0.188 | 0.152 | 0.138 | 0.144 | 0.183 | 0.197 | 0.150 |

Dependent variables: (1) Age of the mayor, (2) Mayor is a blue collar worker, (3) Mayor is a white collar worker, (4) Mayor has a college degree, (5) Mayor with no degree, (6) Mayor is unemployed, (7) Mayor is a housekeeper. In columns we report local linear regressions with uniform kernel and polynomials of order 1 polynomials fitted at the two sides of the discontinuity. Standard errors clustered at the level of town-electoral period.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

TABLE B.4

REDUCED-FORM – FEMALE WINNER AND GENDER POLICIES

| Dependent Variable: | (1) D. LTC | (2) D. Preschool | (3) D. Bal. Serv. | (4) LTC (%) | (5) Preschool (%) | (6) Bal. Serv. (%) |
|---|---|---|---|---|---|---|
| Female Winner | -0.000834 | -0.0484 | -0.0133 | 0.0214 | 0.0700 | -0.0797 |
| | (0.0411) | (0.0456) | (0.0435) | (0.170) | (0.204) | (0.0947) |
| | | | | | | |
| Observations | 3197 | 3135 | 2988 | 3740 | 3464 | 3070 |
| Clusters | 1365 | 1341 | 1280 | 1626 | 1511 | 1331 |
| p-value | 0.984 | 0.288 | 0.759 | 0.900 | 0.732 | 0.400 |
| Bandwidth | 0.135 | 0.132 | 0.126 | 0.166 | 0.151 | 0.131 |

Reduced-form regression-discontinuity regression coefficients corresponding to the effect of the instrument (female winner) on our gender-policy outcomes. Columns 1 through 3 correspond to dummy outcomes indicating positive spending on long-term care, preschool services and work-life balancing services, respectively. Columns 4 to 6 correspond to our continues budget-share outcomes. In all specifications observations are restricted to the optimal bandwidth as described in Calonico et al. (2016). Order 1 polynomials fitted at both sides of the discontinuity included as controls in both regressions.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

TABLE B.5

FEMALE MAYOR - PANEL ESTIMATES

| | D. LTC | D. Preschool | D. Bal. Serv. | LTC (%) | Preschool (%) | Bal. Serv. (%) |
|---|---|---|---|---|---|---|
| Female Mayor | -0.001 | -0.008 | -0.010 | 0.037 | -0.053 | 0.021 |
| | (0.007) | (0.007) | (0.008) | (0.046) | (0.038) | (0.019) |
| Observations | 26255 | 26255 | 26255 | 25993 | 25993 | 25993 |
| Town Effects | Y | Y | Y | Y | Y | Y |
| Year Effects | Y | Y | Y | Y | Y | Y |

*Notes:* Estimates on the effect of mayoral gender on our three policies of interest obtained using our municipal panel specification with town fixed effects. Columns 1 through 3 use positive spending dummy as outcomes. Columns 4 to 6 use our spending share variables as outcomes. Standard errors clustered at the municipal level.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

## Table B.6
### Robustness Checks − Female Mayors

| A) No Controls | (1)<br>D. LTC | (2)<br>D. Preschool | (3)<br>D. Bal. Serv. | (4)<br>LTC (%) | (5)<br>Preschool (%) | (6)<br>Bal. Serv. (%) |
|---|---|---|---|---|---|---|
| Female Winner | 0.0428 | -0.0478 | 0.0500 | 0.0916 | 0.147 | -0.117 |
| | (0.0809) | (0.101) | (0.0906) | (0.372) | (0.416) | (0.201) |
| Observations | 4237 | 3696 | 3815 | 3952 | 4366 | 3190 |
| p-value | 0.597 | 0.635 | 0.581 | 0.806 | 0.725 | 0.562 |
| Bandwidth | 0.190 | 0.158 | 0.166 | 0.178 | 0.201 | 0.137 |
| B) Degree 2 Polynomial | D. LTC | D. Preschool | D. Bal. Serv. | LTC (%) | Preschool (%) | Bal. Serv. (%) |
| Female Winner | 0.00937 | -0.0498 | 0.00958 | -0.0372 | 0.226 | -0.167 |
| | (0.107) | (0.122) | (0.111) | (0.446) | (0.563) | (0.245) |
| Observations | 4528 | 4365 | 4452 | 4447 | 4474 | 4283 |
| p-value | 0.930 | 0.683 | 0.931 | 0.933 | 0.689 | 0.497 |
| Bandwidth | 0.206 | 0.196 | 0.201 | 0.206 | 0.207 | 0.194 |

Note: Second-stage coefficients for the effect of having a female mayor on our policy outcomes. Columns 1 through 3 correspond to dummy outcomes indicating positive spending on long-term care, preschool services and work-life balancing services, respectively. Columns 4 to 6 correspond to our continues budget-share outcomes. In panel A, we report estimates obtained excluding all of our demographic and budget controls. In panel B, we report estimates including our set of controls but using second-degree polynomials in the running variable estimated separately on both sides of the discontinuity.

$^*$ p < 0.1, $^{**}$ p < 0.05, $^{***}$ p < 0.01.

## Table B.7
### PP Mayor - Covariates Municipality

| Dependent Variable: | (1)<br>Pop | (2)<br>Mean Age | (3)<br>College % | (4)<br>Elderly % | (5)<br>HH size | (6)<br>Homemakers % | (7)<br>Debt | (8)<br>Female Mayor |
|---|---|---|---|---|---|---|---|---|
| PP Winner | -5.525 | 0.648 | -0.935 | 0.539 | -0.00586 | -1.295 | 1.809 | -0.00866 |
| | (4.337) | (0.801) | (0.610) | (0.540) | (0.0386) | (1.369) | (4.136) | (0.0546) |
| Observations | 7259 | 6538 | 6950 | 6377 | 7914 | 7885 | 7851 | 6948 |
| Clusters | 2977 | 2680 | 2849 | 2614 | 3237 | 3226 | 3211 | 2848 |
| p-value | 0.203 | 0.419 | 0.125 | 0.318 | 0.879 | 0.344 | 0.662 | 0.874 |
| Bandwidth | 0.146 | 0.129 | 0.139 | 0.127 | 0.160 | 0.159 | 0.158 | 0.139 |

Dependent variables: (1) Population, (2) Mean age of inhabitants in the municipality, (3) Proportion of inhabitants with a college degree, (4) Fraction of inhabitants older than 80, (5) Household Size, (6) Fraction of female homemakers in 2001 census, (7) Outstanding debt share of total budget in 2009, and (8) dummy taking value 1 if mayor is female. In columns we report local linear regressions with uniform kernel and polynomials of order 1 fitted at the two sides of the discontinuity. Standard errors clustered at the level of town-electoral period.

$^*$ p < 0.1, $^{**}$ p < 0.05, $^{***}$ p < 0.01.

## Table B.8

### PP Mayor - Panel Estimates

|              | D. LTC     | D. Preschool | D. Bal. Serv. | LTC (%)  | Preschool (%) | Bal. Serv. (%) |
|--------------|------------|--------------|---------------|----------|---------------|----------------|
| PP Mayor     | -0.018***  | 0.011*       | -0.014**      | -0.073*  | 0.039         | -0.014         |
|              | (0.006)    | (0.006)      | (0.006)       | (0.041)  | (0.031)       | (0.015)        |
| Observations | 26255      | 26255        | 26255         | 25993    | 25993         | 25993          |
| Town Effects | Y          | Y            | Y             | Y        | Y             | Y              |
| Year Effects | Y          | Y            | Y             | Y        | Y             | Y              |

*Notes:* Estimates on the effect of having a center-right (PP) mayor on our three policies of interest obtained using our municipal panel specification with town fixed effects. Columns 1 through 3 use positive spending dummy as outcomes. Columns 4 to 6 use our spending share variables as outcomes. Standard errors clustered at the municipal level.

$^{*}$ $p < 0.1$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$.