Do Female Leaders Promote Gender-Sensitive Policies? **

Felipe Carozzi*

London School of Economics

Andrés Gago**

Universidad Torcuato Di Tella

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. We interpret these results through a model of political selection in which strong parties can impose their agendas on candidates despite primitive differences in policy preferences across genders.

Keywords: Female politicians, gender policies, long-term care.

JEL codes: J14, J16

This version: August 2021

*Address: Dept. of Geography and the Environment. London School of Economics. Houghton Street. London WC2A 2AE. Email: F.Carozzi@lse.ac.uk.

**Address: Business School. Av. Pres. Figueroa Alcorta 7350. CABA C1425CKB. Email: agago@utdt.edu

**We would like to thank María Inés Berniell, Guillermo Caruana, Mónica Martínez-Bravo and seminar participants at CEMFI, SAEe 2016 and the Uppsala Political Economy Workshop 2017 for useful comments and suggestions. Funding from European Research Council under the European Union's Horizon 2020 Programme (ERC Starting Grant agreement 638893 – CompSChoice) and from Spain's Ministerio de Economía y Competitividad (María de Maeztu Programme for Units of Excellence in R&D, MDM-2016-0684) is gratefully acknowledged. A previous version of this paper was distributed under the title "*Female Mayors and Gender Policies in a Developed Country*".

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.[2] At the same time, they also endure a disproportionate burden of caring activities in most households. These caring activities can be partially provided by governments, alleviating the amount of unpaid work carried out by women. For example, public pre-schooling services alleviate the burdens of early childhood care and adult day centres substitute for elderly care activities at home. Can the increased presence of women in government leadership positions promote these types of policies?

In citizen candidate models in the tradition of Osborne and Slivinski (1996) and Besley and Coate (1997), politicians' individual traits – such as gender – can influence policy. Lack of commitment to party platforms means elected politicians implement the policies they prefer.[3] In contrast, many classical models of policy formation focus exclusively on party competition, paying no attention to individual candidates, their traits or their preferences (see Downs 1957, Alesina 1988). It follows from this variety that the influence of politicians' gender on policy selection is ultimately an empirical matter, yet an important one to understand how representation affects policy in democratic governments.

In this paper, we investigate how the gender of government leaders affects policies by looking at three gender-sensitive policies in Spanish municipalities: pre-schooling, long-term care (LTC), and work and family-life balancing services. Our empirical analysis begins by studying how mayoral gender affects engagement with these policies. For this purpose, we focus 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 then use a similar strategy to estimate the effect of political parties on policies. Finally, we use a simple model of political selection to rationalize the

[2]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.

[3]See also Levy 2004. A large group of recent studies has convincingly shown that the identity of the elected politician can matter for which policies are adopted, even after accounting for the party the politician in question belongs to. See for example Besley, Montalvo and Reynal-Querol (2011), Hodler and Raschky (2014), Alesina, Cassidy and Troiano (2019), and the references therein.

different results obtained from both of these analyses.

The choice of the three policies of interest, all of them mentioned in the European Parliament's Gender Equality reports as enhancing the position of women in society, is motivated by the specific impact they may have on the livelihood of women. They provide a substitute to household activities which are disproportionately carried out by women. Regarding LTC, Spanish micro-data from *Instituto Nacional de Estadística* show that it is women who disproportionately perform these activities. As Crespo and Mira (2014) show, this has an impact on the labor market decisions of mature women in Southern Europe. Similarly, a disproportionate role of women is also observed in childcare, with early-years care being carried out mainly by mothers.[4] This also influences labor market decisions of women in Spain, who according to Landwerlin, Balsas and Saura (2012), only return to full-time employment after maternity in 55% cases – compared to nearly 100% for men. As in the case of LTC, this may lead to statistical discrimination in hiring and promotion decisions, which can also affect women who are not heavily involved in caring activities.

The substantial discretion that local governments have in the implementation of these interventions means the Spanish context is especially well-suited to study the engagement of female leaders with gender-sensitive policies. 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. Similarly, 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 balance – such as public school busing services and subsidized school meals –, it is the regional and local governments who are entirely in charge of their provision. Crucially, the rules of administrative information disclosure in force between 2009 and 2014, required municipalities to separately report spending in LTC, preschool and

[4]The 2009-2010 wave of the Spanish time use survey indicates that 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. Regarding long-term care, Spanish micro-data from *Instituto Nacional de Estadística* show that it is women who disproportionately perform these activities: In 2004, before our period of study, 84% of non-professional caregivers were women. 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

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. Complementary estimates show no systematic impact of other characteristics of the politician such as education attainment or occupation on gender-sensitive policies either. Conversely, 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. For two out of three policies, results are confirmed using panel regressions with municipal fixed effects. Hence, in our context, party platforms appear to be more important than gender in the deployment of gender-sensitive policies. In section 6, we build a candidate selection model to show this result is consistent with the presence of strong parties at the local level in Spain, arguably a consequence of the closed-list system that characterizes Spanish elections.

These results have important implications. In recent years, there have been rising concerns about the under-representation of women in positions 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). A separate set of papers document role model effects, whereby female representation prompts more female participation or improved results in elections (see for example De Paola, Scoppa and Lombardo, 2010; Beaman et al., 2009; Baskaran and Hessami, 2018). According to the results of our study, access of female politicians to leadership positions in this context does not by itself result in increased attention to gender-related issues.[5] Instead, our finding that right-wing parties devote less resources to these policies looks consistent with female voters favoring left-wing parties in elections (Edlund and Pande, 2002; Iversen and Rosenbluth, 2006) and with Democrat US representatives receiving higher rating from the League of Women Voters (Lee, Moretti and Butler, 2004).

Previous studies on the effect of female politicians on policy choices are abundant and cover both a wide variety of policies and countries.[6] 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 parlamentaries that occupy seats reserved for low castes and scheduled tribes favor 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. Our results differ from those in these papers both in the context under consideration, in that we find no effect of gender on policy decisions and in that we also look at the role of parties in shaping these decisions. Moreover, 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.

Closer to our institutional setting, Bagues and Campa (2021) exploit the introduction of gender quotas in party lists to explore how they affect the gender composition of Spanish municipal councils, the gender of the mayor, and broadly defined social policies. They do not find any effect on the gender of the mayor, but they do find an effect on the composition of the *pleno municipal*, the legislative body in Spanish municipalities. Moreover, af-

[5]This is not to say that some policies may simultaneously affect the representation of women in office and change the policy landscape. For example, quotas could still affect the deployment of gender policies through changes in entry of candidates and the conditions of political competition.

[6]For a recent literature review on this topic see Hessami and da Fonseca (2020).

ter classifying budget spending categories as *male* or *female* using survey information on policy priorities, they find no effect of quotas on policy. In contrast, Lippmann (2019) finds that after the introduction of quotas in France, female legislators present more female related initiatives than men. Relative to these studies, we study changes in the gender of the head of the executive which are not directly or indirectly related to the introduction of gender quotas.[7] Moreover, differences in the sample period relative to Bagues and Campa (2021) allow us to focus more precisely on policies that are generally viewed as gender-sensitive.[8] We find that the gender of the leader of the executive has no effect on the deployment of gender sensitive policies and rationalize those findings and the findings for party differences using our political selection model.

Much of the remaining literature studying the effect of politicians' gender on policy has *not* focused on specific policies that systematically favor (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. 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 economic outcomes and finds electing a female politician has a positive impact on education attainment in Indian urban areas. While these studies are of interest in their own right, they do not consider whether female politicians systematically choose gender-sensitive policies when in power. The policies or outcomes they analyse are not clearly linked to gender-specific problems. Our paper tackles this question in a context where we can both identify key gender-sensitive policies that are widespread in many economies and implement an identification strategy yielding credible estimates of the relationship between policy decisions and leaders' gender.

[7]Quotas in Spain – and in most countries – may shape the composition of electoral lists for the council but impose no restrictions on the identity of the first candidate in that list. It is typically this candidate who is the one that is appointed to mayoral role by the local council. Evidence from Bagues and Campa (2021) shows these quotas have no effect on the prevalence of female mayors.

[8]Their sample period begins in 2003, before the budget structure law of 2008. As a result, their policy categories are more broad. For example, all local spending in infrastructure or agriculture is classified as *male*, while all spending in pensions or health services is classified as *female*.

2. Institutional Setting & Data

2.1. Institutional Setting

Our empirical analysis is based on data from Spanish municipalities. There were a total 8,116 municipalities in 2011, each ruled by a separate local government. 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 local 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.[9]

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. We focus on municipalities with populations over 250, which use a single-district, closed list, proportional electoral system.[10] In those cases, 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. If the most voted party obtains a majority of seats, it can appoint the mayor directly. If this is not the case, there is a coalition building period in which candidates need to secure the support of the council to be elected. If no candidate can secure this support, the candidate from the most voted party is appointed as mayor. 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

[9]See details in law number 7/1985 (2 of April 1985). *Ley reguladora de las bases del régimen local.*

[10]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.

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. Gender Sensitive Policies

This study leverages on the budget structure law passed in 2008, which required municipalities to separately report spending on LTC, preschool and work and family-life balancing services. We consider these to be good examples of gender sensitive policies in the context of Spain and other OECD countries for a variety of reasons. First, the policies are widely regarded as having a disproportionate effect on women and often discussed as vehicles for increased gender equality. For example, the policies are featured in the Handbook of Gender Budgeting edited by the Council of Europe (Quinn, 2009) and are classified as key policies in the Gender Equality Reports of the European Parliament.[11] Secondly, in meetings we held with diverse agents of the civil society before our study, they coincided in the relevance of said policies for pursuing municipal gender budgeting. Third, municipal governments have substantial discretion in the implementation of these policies. Finally, these policies have differential effects for male and female citizens and are more demanded by the latter. We discuss the last two points in the following.

Long-Term Care

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 for the elderly and other dependent individuals. In practice, the new policies laid out in the *Ley de Dependencia* were poorly funded and only partially implemented. In this context, municipalities, whose complementary role was recognized in the 12th article of the law, had ample room to intervene at their discretion.

[11] According to the OECD, gender budgeting is a way for governments to promote gender equality through the budget process. By 2018, 17 OECD countries (including Spain) had introduced gender budgeting practices.

Some examples of programs that lay under the umbrella of long-term care expenses are: *Taking care of the caregiver* and *Weekend rest* – targeted at the family members that provide care –, public provision of home care services, funding to hire these services privately, and subsidies to install home medical equipment. These policies have a differential effect on women in three dimensions. First, according to the survey conducted by the Spanish Center of Sociological Research, in 2006, before our period of study, daughters were 4.5 times more likely than sons to take care of their parents in old age (CIS, 2006). Thus, policies benefiting non-professional caregivers and policies that alleviate their burden of work, have a differential impact on women. Second, the same survey indicates that before our period of study elderly women needed help to perform their daily activities 2.3 times more often than elderly men. In consequence, in January 2015, after our period of study, the ratio of elderly women to elderly men benefiting directly from these policies was 1.92 to 1. Finally, from the perspective of job creation, interventions targeted at professionalizing care-giving represent an increase in employment opportunities for female workers. In 2015, 93% of professional caregivers in Spain were women.[12]

Consequently, when we look at who demands these policies, we also observe an asymmetry by gender. In 2004, before the congress passed the *Ley de Dependencia*, 78.5% of women versus 73.6% of men regarded it as a priority (CIS, 2004), while in 2007, right after the passage of the law, 61.6% of women versus 53.9% of men regarded the national coverage as poor/limited, leaving room for municipalities to complement it (CIS, 2007) (both differences significant at the 1% level).

Pre-Schooling and Work and Family-Life Balancing Services

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 address this limitation, article 15 of the law stated the goal of promoting access at no cost also at the first stage. To that end, it gave a key role to local governments, which at their discretion could offer public daycare centers or subsidies to acquire these services

[12]Data on the ratio of female to male beneficiaries from the Spanish System for the Attention to Dependent People (SAAD). Data on the gender of professional caregivers from Spanish Statistical Institute (INE) – see category 57 in the Spanish Labour Force Survey (*Encuesta de Población Activa).*

privately. At the same time, work and family-life balancing services, which include busing, lunch at school, early drop and breakfast and extra-curricular activities, were entirely provided by regional and local authorities.[13]

Time use surveys indicate that women benefit disproportionately from the provision of pre-schooling and work-life balancing services. According to CIS (2014), 37.5% of Spanish families with children between 0 and 3 do not use pre-schooling services because of affordability issues. Considering that the person that takes care of 0 to 3 year old children is the mother 75.9% of the time (CIS, 2014), it is easy to see how extending the coverage of free pre-schooling can benefit women in particular. Likewise, the policies that lay under the umbrella of work and family-life balancing services provide an alternative to families that cannot take care of their children early in the morning, during lunchtime, in the afternoon and/or on holidays. According to CIS (2014), 76% of mothers vis-a-vis 15.9% of fathers take care of their children when they are not at school. Hence, providing free or subsidized care services for those children also alleviates substantially the burden of non-remunerated work on women.

Furthermore, as in the case of LTC, the jobs associated with these policies are disproportionately occupied by women – 97.6% of preschool teachers and 93% of caregivers are female.[14] Hence, it seems reasonable to think that an increase in the amount provided of these services would result in new employment opportunities for women.

Taken together, this evidence indicates that the policies under consideration – long term care, preschooling and work and family life balancing services – are especially important for Spanish women. Moreover, considering the division of competences between the different levels of the Spanish administrations, they represent the main area for municipalities to deploy gender sensitive policies. As a result, we expect female politicians may display a preference for these types of policies.

The focus on our policies of interest is, in our view, a significant improvement relative to the literature. The issue with the policies studied in previous papers is that whether or not they can be associated to the preferences of a specific group of voters – e.g. males

[13]These services, targeted at all children in school age, are recorded in the municipal budget as complementary educational spending

[14]Data from Spanish Statistical Institute (INE) for 2015. Caregivers include those professionals in the kindergarten that are not teachers but look after children.

or females – is not always clear and often context-dependent. For example, Bagues and Campa (2021) classify spending in infrastructure as *male* spending. However, in other contexts, it is actually women who are concerned with such issues as road improvement (Chattopadhyay and Duflo, 2004). Similarly, they classify agriculture as *male*, which might be appropiate for some Spanish regions, but not for others which are specialized in crops that are almost entirely harvested by female workers (e.g. the strawberry fields in Huelva, Spain). Likewise, given the low literacy rate among Indian women, one could interpret the positive effect of female policitians on educational outcomes in India (Clots-Figueras, 2012) as driven by female politicians caring about gender issues. However, according to Chattopadhyay and Duflo (2004) it is actually Indian males who favor spending in education. Relative to these studies, we use policies which are widely considered as effective tools for gender budgeting. This allows us to test more accurately whether female politicians actually favor gender issues, as well as to study the partisan determinants of gender policies.

2.3. Data and Descriptives

To conduct our analysis, we build a municipal panel for the period 2010-2014 combining data from four different sources. Electoral results for Spanish municipalities in the 2007 and 2011 local elections are obtained from the Spanish *Ministerio del Interior*. This information includes both municipal level results for all running parties and the list of candidates parties presented in every municipality. It reflects the ordering of these lists and, importantly, the gender of each candidate. Data on demographics of the candidates 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.[15] Data from *Estadística del padrón Continuo* include yearly information on population and population by age categories for all the sample period.

[15] 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.

Data on yearly municipal budgets is obtained from the database on local authority 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.[16] 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 including the vote shares obtained by all parties, demographics of the politicians, information on municipal spending by program (including LTC, preschool and life-balancing services), outstanding debt in 2009 and other municipal characteristics. Municipalities approve budgets for a given year on December of the previous year, so we allocate budgets for t to the mayor in $t-1$.[17] 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 the mean and standard deviation for both our outcome variables (the three selected policies) and municipal characteristics. These statistics are reported for all municipalities in column 1, for municipalities ruled by a female mayor in column 2, by a male mayor in column 3, and for municipalities where a mixed race took place (male and female candidates running for election) in column 4. We measure our outcome variables both in terms of the extensive margin (whether spending on a policy takes place) and intensive margin (what is the fraction of total spending allocated to that policy). Municipalities ruled by female and male mayors do not differ substantially in terms of engagement with the policies of interest. The only statistically significant difference can be found in the share of spending devoted to long-term care policies where we see female-led local gov-

[16] 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 in broader categories.

[17] This means, for example, that budgets corresponding to 2011 are associated to the government in power at the end of 2010.

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 homemakers 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. Observations, indicated in the table foot, correspond to municipality-year pairs. Municipality*Elections correspond to municipality-term pairs (we have at most two terms, the one beginning in 2007 and the one beginning in 2011).

ernments having slightly larger spending (associated p-value= 0.048). In the second panel of Table 1, we report demographic characteristics for our municipalities including population, age structure and the fraction of homemakers in the 2001 census. The average population of municipalities in our sample is 8.65 thousand inhabitants. We can observe that municipalities governed by women are usually larger (associated p-value= 0.064) but otherwise not very different from municipalities ruled by men in terms of these observables. Finally, we also report the fraction of municipalities ruled by the centre-left *Partido Socialista* (PSOE) and the center-right *Partido Popular* (PP) in each column.

TABLE 2
MUNICIPAL DESCRIPTIVES BY PARTY

	All	PP Mayor	PSOE Mayor
Municipal Characteristics			
Female Mayor	0.17	0.17	0.18
	(0.4)	(0.4)	(0.4)
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)
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 census	0.33	0.36	0.35
	(0.1)	(0.1)	(0.1)
Observations	26257	10470	9155
Municipality*Elections	10911	4125	3844

Notes: Column 1 includes all municipalities in our sample, 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. Observations, indicated in the table foot, correspond to municipality-year pairs. Municipality*Elections correspond to municipality-term pairs (we have at most two terms, the one beginning in 2007 and the one beginning in 2011).

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. **??**This is especially the case for preschool and work-life balancing services, where differences by party are significant at all conventional levels in both the intensive and extensive margins. Other variables in Table 2 indicate there are additional differences between PP and PSOE municipalities. Therefore, mean comparisons in policy engagement 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 Appendix Table B.1. 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.

3. Female Mayors and Gender Policies

Empirical Strategy

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 many cases, these confounding factors may be unobservable, so matching and regression-based estimates may be biased.

To correctly estimate the effect of interest, we need an 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 vast 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 results 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.[18] 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 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 $FemaleVoteMargin_{it} = VS^F_{it} - VS^M_{it}$ where VS^F_{it} is the vote share of the female candidate, and VS^M_{it} 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. We will also report results for the reduced-form effect of a female candidate winning the election.

[18]The person appointed as mayor is, with very few exceptions, the first-in-the-list candidate of one of the parties in the council.

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 labor force. Year dummies are represented by η_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 as dependent variables. The parameter of interest in each specification is δ, 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.

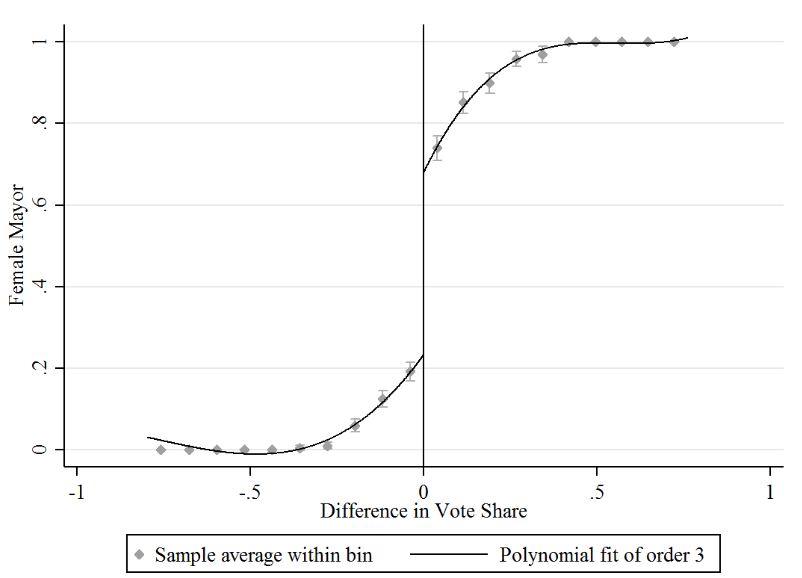
To ensure that our estimates for δ are obtained from variation at the threshold, we will restrict the sample to a bandwidth of the running variable around 0. A large methodological literature on the implementation of RD designs proposes different methods to choose i) the bandwidths around the threshold 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.

Validity of Research Design

Before discussing the results, 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 just under 50 percentage points – when $FemaleVoteMargin_{it}$ goes above 0. This is confirmed in Appendix Table B.2, 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% and correspond to the optimal bandwidths selected for the second stage regressions reported below. Using these bandwidths, 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. 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

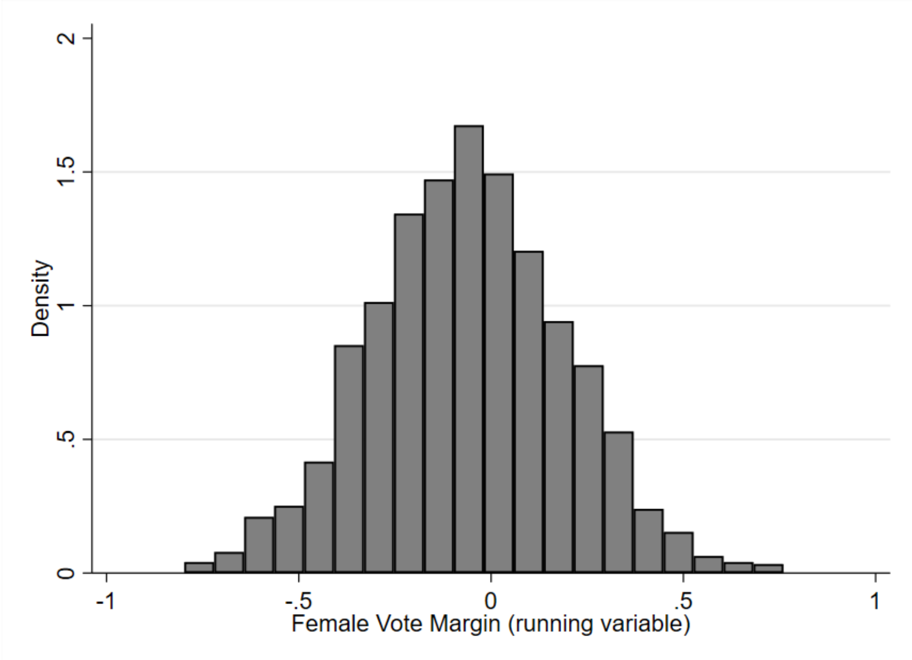


Note: Graph indicates the discontinuity in the probability of having a female mayor when a woman barely wins a mixed race. Vertical axis represents fraction of female mayors. Horizontal axis represents female winning vote share margin, negative if female candidate lost election. Points indicate averages 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, a key assumption required for the validity of RD designs. This is not particularly surprising as it is unlikely that political parties can precisely determine vote shares.

FIGURE 2

RUNNING VARIABLE HISTOGRAM - FEMALE MAYOR

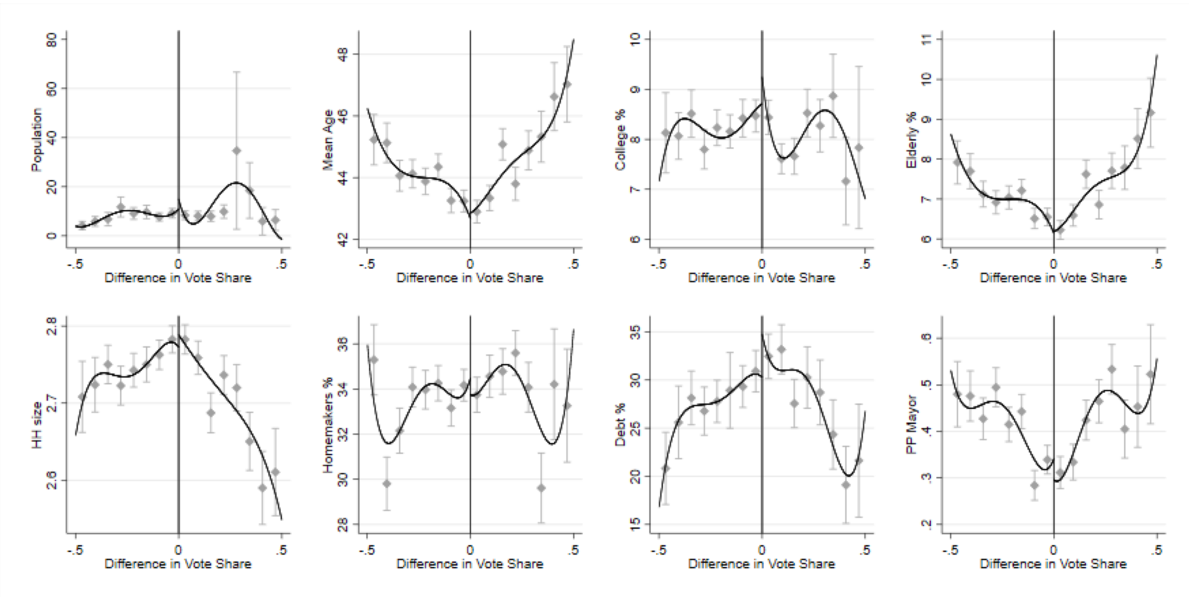


Note: Distribution of female vote shares in mixed races in the 2007 and 2011 elections (pooled).

A related assumption required for the validity of this estimation method is that no municipal characteristics (other than the mayor's gender) change discontinuously at the threshold. To explore this, we use our set of controls X_{it} and estimate whether these vary discontinuously when 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 Appendix Table B.3. In all cases, we observe that the coefficients measuring discontinuities in the covariates are small and not significantly different from zero.[19]

[19]Estimates reported in Table B.3 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 3
Covariate Balancing – Female Winners



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.

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. Yet 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 representation. Hence, we do not control for them in our regressions. We report estimates for the change at the threshold of these 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.4. The reduced-form effect of having a female winner only has a statistically significant effect on the mayor's age.

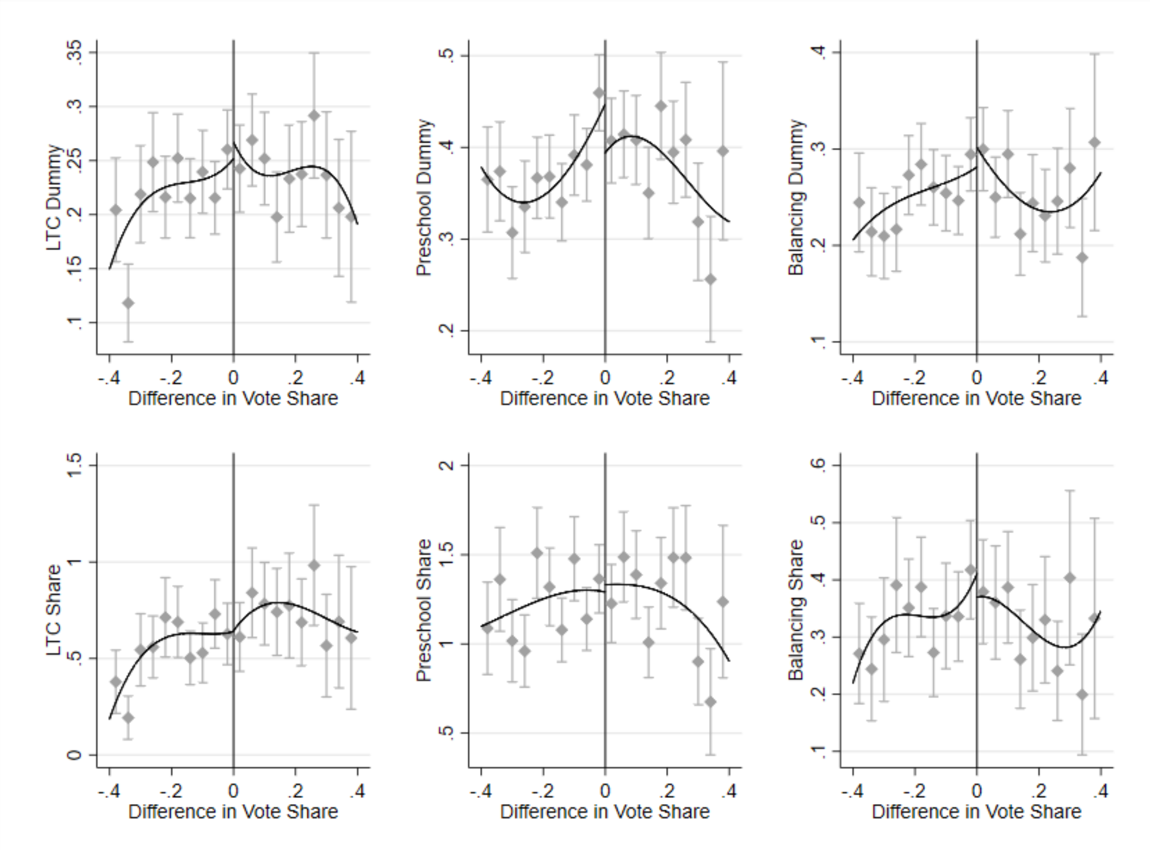
Results: Female Mayors and Gender Policies

We turn now to the estimates of the effects of mayoral gender on our set of gender-related policies obtained using our RD specification. We start by providing reduced-form graphs illustrating 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.

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.5 indicate that none of these differences are statistically significant.

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.

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 3 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 3
2nd Stage - Female Mayor

VARIABLES	(1) D. LTC	(2) D. Preschool	(3) D. Bal. Serv.	(4) LTC (%)	(5) Preschool (%)	(6) Bal. Serv. (%)
Female Mayor	0.0189	-0.0964	0.0453	0.0576	0.0988	-0.0154
	(0.0795)	(0.0913)	(0.0871)	(0.351)	(0.389)	(0.203)
Observations	3947	3634	3482	4422	4264	2782
Clusters	1684	1551	1487	1918	1855	1442
p-value	0.812	0.291	0.603	0.870	0.800	0.940
Bandwidth	0.174	0.155	0.148	0.204	0.195	0.144

Note: Dummy outcomes in columns 1 through 3 take value one if spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services respectively is above zero. Percentages in columns 4 through 6 calculated as the fraction spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services as a percentage of total spending. 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 0.0576 coefficient indicates that having a female mayor increases the fraction of the budget devoted to long-term care spending by roughly half a percentage point. This is less than one-thirtieth 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 percentage points. 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 in column 1 would indicate that having a female mayor increases the probability of engaging in LTC policies by 1.89 percentage points. This is less than 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. Across specifications, we can exclude effects of roughly 1/3 of the standard deviation of the outcome with 90% confidence. We address potential concerns about the precision of these estimates with our panel exercise in section 5.

We conclude from these results that female leaders do not promote gender-sensitive policies at the local level. These findings are in line with previous studies looking at the composition of spending in developed countries. For example, Ferreira and Gyourko (2014) use data on US mayors to report no effect of politicians' gender on broadly defined policy areas which are not necessarily gender specific. However, these results stand in contrast with those obtained for developing countries, which indicate large and significant effects of politician's gender on gender sensitive policies in India (see Clots-Figueras 2012, Chattopadhyay and Duflo 2004). This difference between the findings across contexts is something that has already been highlighted in Hessami and da Fonseca (2020), and may relate to important differences in the policies under consideration, the social norms surrounding gender and the institutional design of political competition. In section 6, we build a theoretical model to rationalize these heterogeneous empirical findings based on differences in the strength of parties across contexts. With strong parties, it is parties and not politicians' characteristics which shape policy choice. We provide evidence consistent with this hypothesis in the next section.

4. Political Parties and Gender Policies

Empirical Strategy

In this section we estimate the effect of political parties on gender-sensitive policies. Specifically, we study whether municipalities ruled by centre-right *PP* mayors differ in their engagement with these policies. If parties operate as policy platforms and candidate commitment to these platforms is strong, then the party in power may dictate the policies implemented by the local government. Under these assumptions we would also have that individual candidate preferences may not matter at all for policy.

To test whether parties differ in their application of gender sensitive policies we conduct a close election *fuzzy* RD similar to the one described in section 3. 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 Popular* 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. Estimates are obtained using local linear regressions with optimal bandwidth as in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016).

Validity of the Research Design

This empirical strategy is suitable to study the effect of partisan affiliation on gender-policies. The instrument is strong, with Appendix Figure A.1 showing a sizeable jump at the threshold in the first stage, and Table B.6 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.2 and suggests no manipulation at the threshold. In addition, Appendix Figure A.3 shows no substantial discontinuity for covariate values at the threshold, a result confirmed by the insignificant coefficients reported in Appendix Table B.7. Taken together, these results indicate our RD strategy is suitable to estimate the impact of parties on gender-sensitive policies.

Results: Political Party and Gender Policies

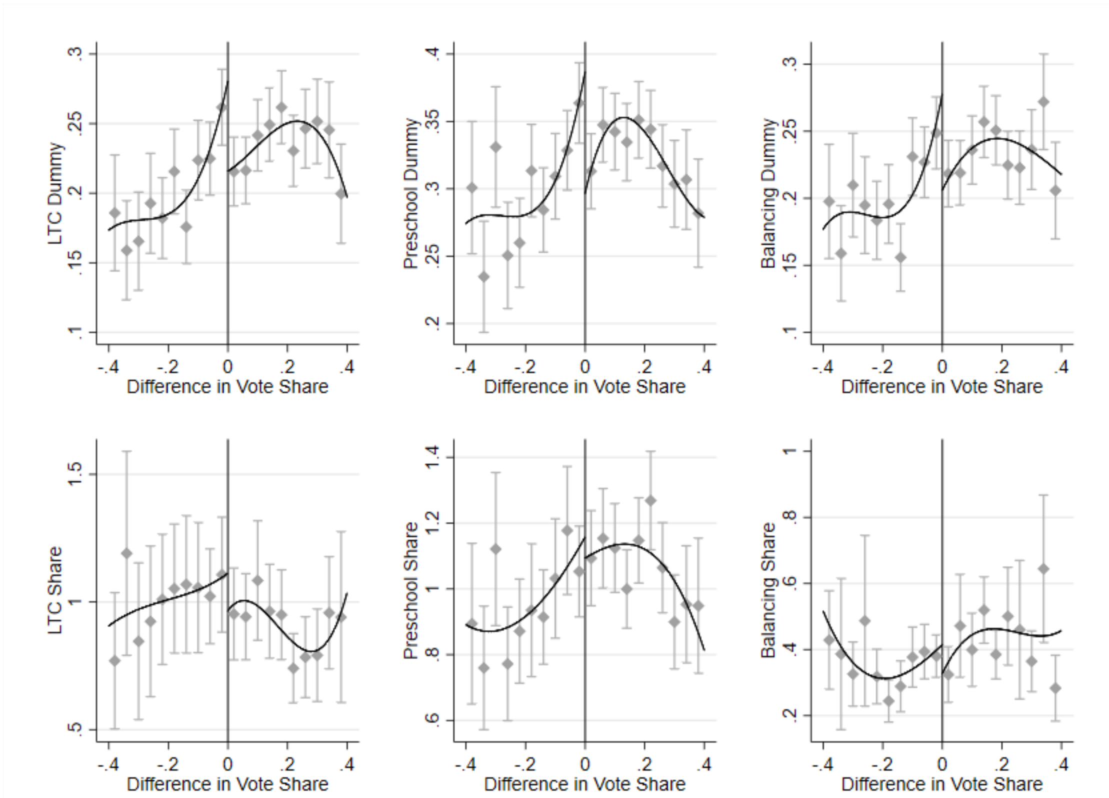
Figure 5 illustrates the reduced-form effect of *PP* winning an election on extensive and intensive margin measures of engagement with gender policies. The first row shows an apparent discontinuity at the threshold in the extensive margin outcomes, indicating local governments led by a centre-right mayor are less likely to promote these policies. Second stage estimates of the effect of having a PP mayor on policies are presented in Table 4. We find statistically significant effects for our extensive margin outcomes – columns 1 through 3 – showing that mayors from the centre-right *PP* are less likely to engage in gender-friendly policies. 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 the point estimates obtained for female mayors in columns 1 through 3 of Table 3. 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.

While a growing literature 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.[20] Arguably, these differences in policy can have important implications for voters. The weaker commitment of right-wing parties to gender policies has been

[20]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.

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 and Pande 2002). In the case of Spain, it might explain the emerging gender gap in favor of PSOE (center-left party) in the general elections of 2011 (CIS, 2011). Likewise, a stronger commitment of left-wing parties in deploying gender sensitive policies could explain the positive effect of a Democratic state house reducing the gender wage and unemployment gap (Kuk and Hajnal, 2020) in the US.

One potential concern regarding the estimates reported in Table 4 is that gender could vary discontinuously at the threshold with the identity of the ruling party. This is plau-

Table 4

2nd Stage - PP Mayor

VARIABLES	(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

Note: Dummy outcomes in columns 1 through 3 take value one if spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services respectively is above zero. Percentages in columns 4 through 6 calculated as the fraction spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services as a percentage of total spending. 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.

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

sible, given that there is a slight difference in the fraction of municipalities with female mayors by party (see Table 2). A similar issue could arise in our analysis for the effect of female mayors. Table 5 reports results of 2SLS estimates of the effects in both cases. Column 1 estimates obtained by replacing the PP Mayor dummy as the outcome in the RD specification for the effect of gender. Column 2 estimates obtained by replacing the PP Mayor dummy as the outcome in the RD specification for the effect of party. The resulting coefficients are statistically insignificant at conventional levels. This indicates that the negative effects of centre-right mayors on engagement with these policies are not confounded by changes in the gender of the mayor and vice-versa.[21]

We conclude that parties are more important than mayoral gender in determining engagement with our policies of interest. This is consistent with parties operating as policy platforms, leaving only a limited role for the individual characteristics of politicians in determining policy. The model proposed in section 6 helps us understand how this result can arise with an institutional design that promotes strong party discipline.

[21]Note that the main results for the effects of party and gender are unaffected by including the other variable as a control (not shown).

Table 5

RD Balancing Checks: Party and Gender

	(1) Mayor PP		(2) Female Mayor
Female Mayor	-0.0421	PP Mayor	-0.00478
	(0.0992)		(0.0540)
Observations	3884	Observations	7228
Clusters	1657	Clusters	2964
p-value	0.671	p-value	0.929
Bandwidth	0.170	Bandwidth	0.145

Note: Estimate in column 1 obtained by replacing the PP Mayor dummy as the outcome in the RD specification for the effect of gender. Conversely, estimate in column 2 obtained by replacing the Female Mayor dummy as the outcome in the RD specification for the effect of party. Bandwidths coincide with average bandwidths reported in Tables 3 and 4. Clustered standard errors in parentheses.

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

5. Robustness Checks

In this section we test the robustness of our main results. We change details of the method used to obtain our RD estimates, we employ an alternative empirical strategy with a different set of identifying assumptions, and we run the main specification with a different definition of gender policies. In all cases, qualitative results remain the same as in the sections above.

Robustness of the RD

Several methodological decisions need to be made when estimating the RD estimates discussed in Sections 3 and 4. We have followed the estimation method in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016), controlling for a first order polynomial fitted at each side of the threshold, using a uniform weighting kernel and including a set of controls to reduce variance and increase the precision of our estimates. We can show the qualitative results in Tables 3 and 4 are robust to the inclusion of higher order polynomials, weighting observations with a triangular kernel and the exclusion of controls. As an illustration, Tables B.9 and B.10 in the appendix shows estimated effects obtained without the additional set of controls (panel A) and fitting a global third order polynomial with a bandwidth that includes all elections won/lost by a margin of +/- 50% vote share (panel B).[22] In Table B.9 all results remain as in the main specification. The coefficients for female mayor remain non-significant and of a similar order of magnitude than those reported in Table 3. Similarly, Table B.10 shows that results for party are also robust to these changes.

It is important to note that our baseline RD estimates for the effects of female mayors and PP mayors are obtained using different samples. There are two reasons for this. First, the definition of *mixed* race is not the same in both exercises – e.g. some electoral races in which PP is one of the two most voted parties may not feature any female candidates. Second, our parameters of interest are estimated after restricting the sample to observations close to the threshold given by a data-driven bandwidth selection method. We can restrict these samples to be identical by selecting only electoral races which are mixed for both PP candidacy and gender and fixing a common bandwidth. We provide estimates using these sample restrictions in Appendix Table B.11. The sample restriction reduces estimate precision substantially, but we continue to find significant negative effects of PP Mayors in 2 out of 3 policies. We find no significant effects of female mayors.

Panel Estimates

We can use longitudinal variation in our dataset to test whether the main conclusions of our analysis are confirmed using a different set of identifying assumptions. Specifically, we can exploit time variation in mayoral gender and party to estimate a panel specification with municipality fixed effects:

$$Policy_{it} = \beta_1 PP_{it} + \beta_2 Female_{it} + \gamma pop_{it} + \eta_t + \alpha_i + \epsilon_{it}$$

where PP_{it}, $Female_{it}$ and η_t are defined as above, pop_{it} is the population of municipality i in year t, and α_i represents a municipality fixed-effect.[23] 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.

[22] Notice that, with the optimal bandwidth selected following Calonico et al. (2016), we use between one third and half of available observations to produce the RD estimates of the impact of mayoral gender on the three policies of interest, and roughly 38% of available observations when studying the impact of party affiliation. These numbers rise to 94% and 93%, respectively, when we use a vote margin of +/- 50%.

[23] Given that all other controls included in vector X_{it} defined above are fixed over time, we only include population as a control in our panel specifications.

By including α_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 or party of the mayor and with spending in gender-sensitive policies (long-term care, preschool services or life-balancing services). For example, more progressive voters may be more likely to vote for a party headed by a female leader and, simultaneously, demand more spending in gender-sensitive policies. Causal interpretation of the resulting estimates requires assuming time-varying unobservable determinants of policies are not correlated with mayoral gender or party. Note that this assumption is arguably stronger than the one invoked when interpreting our RD estimates.

Results from our fixed effects specification for the six outcomes of interest are reported in Table 6. The coefficient for $Female_{it}$ is statistically indistinguishable from zero in all specifications. For our discrete outcomes, these coefficients are in fact negative and small. For example, the point estimate in column 2 would indicate that having a female mayor decreases the probability of engaging in pre-schooling policies by under 0.8 percentage points (the baseline probability is 45 times larger). In the case of our continuous outcomes, our coefficients of interest are also smaller than those reported in Table 3. In all specifications, we can reject effects of over 4% of the standard deviation of the outcome with 90% confidence. In the case of the coefficients for the centre-right mayor dummy, we find negative effects on engagement with gender sensitive policies in the case of LTC and Work Life Balancing Services. We also find a negative and significant effect on the fraction of spending corresponding to LTC. We interpret these results to be broadly consistent with the effect for PP Mayors reported in our RD analysis.[24] Taken together, the estimates reported in Table 6 are in line with the politician's party being more important than their gender in determining engagement with these policies (p-values for the test with null hypothesis $H_0 : \beta_1 = \beta_2$ in each specification provided in the table foot).

The additional precision of these estimates is likely to result from the larger sample used in our longitudinal analysis. Moreover, they do not rely on the selected sample of

[24]It is important to note that the samples used to produce our panel and RD estimates are quite different. In Appendix Table B.8 we estimate our panel specification using the sample of municipalities which experience a mixed race (mixed by gender and mixed by party) in either the 2007 or the 2011 election. Qualitative results are similar when using the full sample of municipalities.

TABLE 6

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)
Female Mayor	-0.000	-0.008	-0.010	0.038	-0.054	0.021
	(0.007)	(0.007)	(0.008)	(0.046)	(0.038)	(0.019)
$H_0\beta_{PP} = \beta_{Fem}$	0.07	0.04	0.68	0.08	0.07	0.12
Observations	26257	26257	26257	25995	25995	25995

Notes: Results of estimating our panel specifications including municipality fixed effects, year dummies and a population control. Columns 1 through 3 correspond to dummy outcomes. Columns 4 to 6 correspond to the share of all spending corresponding to each policy. Standard errors in parentheses clustered at the municipal-election level. We report p-values for test of equality between the coefficients for the female and PP dummies.

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

competitive races that are used in the close election RD reported above. [25]

Other Policies

As explained in section 2.2, our choice of gender-sensitive policies rests on three facts: They are framed as gender sensitive in the policy debate, the individuals affected by them are mostly female, and they are prioritized by female voters. Yet, it is possible that mayoral gender affects a different set of policies. To address this point, we classify spending categories in municipal budgets into *male* and *female* following the alternative criteria proposed in Bagues and Campa (2021).[26] We then calculate female spending as a fraction of total classifiable spending and use this as the outcome variable in a modified version of our RD design.

Results for this analysis are reported in Table 7, where column 1 presents the effect of gender and column 2 the effect of center-right mayors. Coefficients indicate that having a female or a PP mayor results in a reduction of the fraction of spending classified as

[25]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 in sections 3 and 4 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.

[26]The classification is based on observed gender differences in the priority given to different issues as recorded in survey by the Spanish Center for Sociological Research. These are mapped to the program classification in the municipal budgets. See Bagues and Campa (2021) for details.

female. The point estimate in column 1 is roughly twice the size of the point estimate in column 2, although both are imprecisely estimated and fall in each other's 95% confidence intervals. We arrive to one key conclusion from this analysis: female mayors do not appear to promote spending in these broadly defined *female* programs either. If anything, they reduce spending in these policies, as does the election of a centre-right mayor.

TABLE 7

BROAD POLICY CATEGORIES: PARTY AND GENDER

	(1) Female spending		(2) Female spending
Female Mayor	-10.78**	PP Mayor	-5.740*
	(4.338)		(3.129)
Observations	4364	Observations	7242
Clusters	1871	Clusters	2994
p-value	0.013	p-value	0.067
Bandwidth	0.198	Bandwidth	0.146

Note: RD estimates for of the effect of gender categories in broadly defined "female" spending. The outcome variable is the fraction of female spending as a part of all spending classified by gender (female + male) Column 1 corresponds to the effect of having a female mayor on female spending and column 2 corresponds to the effect of having a PP mayor on female spending. Clustered standard errors in parentheses.

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

One cautionary note is due regarding aggregated results in Table 7. The categories involved in the classification of policies present in Bagues and Campa (2021) are quite broad. For example, all local spending in infrastructure or agriculture is classified as *male*, while all spending in pensions or health services is classified as *female*.[27] Hence, to identify which of these spending categories drive the negative effect for female mayors, we separate female and male aggregates into their components in Appendix table B.13.

Panels A and B correspond to the *female* and *male* categories, respectively. We observe that only the coefficient for infrastructure spending is large and statistically significant at conventional levels, indicating that female mayors spend 4.5 percentage points more resources in this *male* category. We find this result hard to interpret, as it is difficult to attribute a clear link between infrastructure spending and gender. For instance, Bagues and Campa (2021) classify infrastructure spending as male as a result of a difference in

[27] The sample period used in Bagues and Campa (2021) extends from 2003 to 2015 so harmonized budget categories for this period are necessarily broader than in our sample.

the fraction of males and females citing infrastructure as a political issue they are concerned with in regular sociological surveys. Yet according to outcomes of this survey, infrastructure is a second order concern for both males and females, and the difference appears to be driven by inhabitants of relatively large municipalities. Moreover, if we look at a different country, we observe that Chattopadhyay and Duflo (2004) find some infrastructure categories are indeed not a male but a female category in India. Hence, we find hard to interpret this single significant coefficient as decisive evidence of differences in policy driven by differences in preferences between female and male mayors.

6. Theoretical Framework

Results in the previous sections indicate that the party in power is more important than the gender of the mayor in determining whether a municipal government engages in gender-sensitive policies. This stands in apparent contradiction with both the evidence that Spanish women disproportionately favor these policies – see section 2.2 – and with findings for developing countries in previous research, which shows substantial differences by gender in related policies. To rationalize this, we explore under which conditions a different distribution of policy preferences between male and female agents translates into different policies implemented by male and female mayors. We do so by specifying a candidate selection model with endogenous policy choices and heterogeneous preferences by gender. We use the model to show that with strong parties – parties that can impose big punishments if elected officials defect from party lines –, the preferences of the agents that self-select into running for a party are not representative of the preferences of the agents' gender. Conversely, when parties have little capacity to punish defection, heterogeneity in preferences by gender can result in substantial differences in implemented policy.

Agents and Parties

In our model, some citizens – henceforth, agents – opt whether to be a candidate with one of two parties. Agents differ in their preferred policy $\hat{\theta}_i$, which is private information to them, and in competence or charisma. If agents postulate to be party candidates, are selected to run by the party, and become mayors, they select a policy along a one-dimensional policy space.

Agents derive a positive utility for being in office and negative utility for implementing policies that are distant from their individual preferences. When in office, parties – which act as principals – can punish agents who implement policies that are distant from the party line.

$$V_i(\theta) = \omega - \alpha(\theta - \hat{\theta}_i)^2 - \gamma(\theta - \theta_p)^2 \tag{2}$$

The parameter ω captures rents from office, $\hat{\theta}_i$ is the agent's preferred policy, α denotes how painful is for the agent to depart from their preferred policy, θ_p is party p preferred policy or bliss point – the 'party line' – and γ denotes how harshly can parties punish mayors that depart from party lines. The agent decides whether to run as a candidate for party A, for party B, or not running at all. Without loss of generality, we assume that $\theta_A > \theta_B$.

We assume there is a continuum of male and female agents with identical distribution of ability. We also assume that the distribution of preferences and ability is orthogonal for both genders and that both preferences and ability are continuously distributed. Primitive differences in policy preferences by gender are built into the model by assuming that the distribution of agent preferences for females (f) first order stochastic dominates the distribution of preferences for males (m).

$$\hat{\theta}_i \sim f_f(\hat{\theta}) \;\; FOSD \;\; \hat{\theta}_i \sim f_m(\hat{\theta}) \rightarrow E(\hat{\theta}_i|f) > E(\hat{\theta}_i)|m)$$

Following much of the literature on probabilistic voting models, we assume there is a unit mass of atomistic voters who vote in the elections according to the competence of the candidate, to their individual preference for one of the parties, and to an aggregate preference party shock. As agents' preferred policy $\hat{\theta}_i$ is private information, parties respond to this voting behavior by selecting agents only in terms of competence (see the Theoretical Appendix for further details).

Timing

The sequence of actions is as follows: i) Eligible agents choose whether or not they opt in to the pool of potential candidates for each party, ii) parties draw a male and a female candidate at random from the pool of potential candidates, iii) parties simultaneously select a candidate who runs for election from these two options, iv) elections are held where voters select candidates according to their competence, v) elected candidates implement their optimal policy θ_i^*, vi) parties punish candidates if they defect from party lines.

Equilibrium Policy

The model is solved by backward induction. The election winner chooses θ so as to maximize $V_i(\theta)$. Taking first order conditions in this concave objective function yields chosen policy θ^* as a function of $\hat{\theta}_i$ and θ_p. We define $\theta_i^* \equiv h(\hat{\theta}_i, \theta_p)$ (see solved expressions in the Appendix). The set of agents that opt in to the pool of candidates who run for party p can be characterized in terms of their candidate types $\hat{\theta}_i$ as:

$$\mathbf{C}_p = \{\hat{\theta}_i : V(h(\hat{\theta}_i, \theta_p)) > 0, V(h(\hat{\theta}_i, \theta_p)) > V(h(\hat{\theta}_i, \theta_{-p}))\}$$

The first statement in the set definition is equivalent to a participation constraint (whether to run for a party or not) and the second relates to the preferred party to run for. From the first inequality we can derive a lower and upper bound for $\hat{\theta}_i$ as a function of θ_p and γ so that agents are willing to run for election with party p if $\hat{\theta} \in [\hat{\theta}_{1p}(\gamma), \hat{\theta}_{2p}(\gamma)]$. As shown in the theoretical Appendix, this interval is symmetric around θ_p. It is also narrower for greater values of γ, as stronger parties are better able to constrain the actions of their elected officials. The second condition in the definition of $\mathbf{C}_p$ is satisfied if $|\hat{\theta}_i - \theta_p| < |\hat{\theta}_i - \theta_{-p}|$. That is, individuals choose the party whose bliss point is closest to their own.[28]

The assumption that competence is orthogonal to $\hat{\theta}_i$ implies that agents selected in the candidate selection stage are, on average, a random sample of the agents that opt into eligibility. Therefore, the average choice of θ in equilibrium conditional on gender (g) and party (p) can be computed as:

$$E(\theta^*|g, p) = \frac{\int_{\hat{\theta}_{1p}(\gamma)}^{\hat{\theta}_{2p}(\gamma)} h(x, \theta_p) f_g(x) dx}{F_g(\hat{\theta}_{2p}(\gamma)) - F_g(\hat{\theta}_{1p}(\gamma))} \tag{3}$$

[28]For low enough values of γ there is a mass of agents for whom this constraint is binding. See theoretical Appendix.

Differences in Policy by Party and Gender

We can use the expression of $E(\theta^*|g, p)$ above to study how average policies vary by party and by gender under different institutional conditions. To do this, let us first consider the limiting cases of $\gamma = 0$ and $\gamma \to \infty$. These polar cases serve to illustrate how party strength can be related to individual and partisan differences in policy.

If $\gamma = 0$, then parties cannot impose any costs of defections to their elected officials. As a result, the conditional distribution of candidate preferences for each gender coincides with primitive distribution $f_g(\hat{\theta})$, mayors always select their preferred policy and there is no difference in policy by parties.[29] Given the assumption of stochastic dominance above we have that $E(\theta^*|f, p) > E(\theta^*|m, p) \quad \forall p$. So, if parties are weak and have no ability to impose punishments on mayors who defect, then the differences in the agents' preferences by gender translate into differences in policy.

Conversely, if $\gamma \to \infty$, then mayors cannot defect from the party line and mayors select θ_p. Given that we have assumed that $\theta_A > \theta_B$, in equilibrium there will be differences in implemented policy between parties but no difference in the policies implemented by male and female mayors.

Moving beyond the polar cases, we can establish a set of conditions under which we can obtain sharp predictions of how changes in party strength (i.e. changes in γ) translate into differences by mayoral gender. For this purpose we will assume that functions $f_f(\hat{\theta})$ and $f_m(\hat{\theta})$ are unimodal with modes $\overline{\theta}_f$ and $\overline{\theta}_m$. Note that, given the assumption of stochastic dominance above, we will have that $\overline{\theta}_f > \overline{\theta}_m$. Moreover, we assume γ is bounded by below so that, $\gamma > \underline{\gamma}$ with $\underline{\gamma} \equiv max\left[\frac{\alpha\omega}{\alpha(\overline{\theta}_f - \theta_A)^2}, \frac{\alpha\omega}{\alpha(\theta_B - \overline{\theta}_m)^2}\right]$.

Proposition

Assume the distribution of preferences is unimodal for both genders, $\theta_p \in (\overline{\theta}_m, \overline{\theta}_f) \ \forall p$ and $\gamma < \underline{\gamma}$. Define the difference in expected implemented policy by gender as $\Delta_G(\gamma)$. We then have that $\frac{\partial \Delta_G(\gamma)}{\partial \gamma} < 0$.

Proof: See Theoretical Appendix.

The proposition states that contexts with stronger parties (i.e. higher γ) lead to a

[29]Formally, with $\theta = 0$, the integration limits in 3 coincide with the bounds on the support of $\hat{\theta}_i$.

smaller difference in equilibrium policy between mayors of different gender. We consider this to be consistent with the Spanish context, where the closed-list electoral system and other institutional elements, such as a long-lived two party system present at different levels of the administration, lead to a situation in which parties have substantial capacity punish mayors who defect from the party line. Note that this context differs substantially from the Indian context that has been the subject of much of the preceding literature.

One key insight in our model is that it is because of the selection of candidates into parties that primitive differences in policy preferences by gender do not translate into differences in policy. The implication is that, even if female citizens across the ideological spectrum demand more gender policies than male citizens, right-wing female mayors will deliver less gender-sensitive policies than other female politicians. To provide an informal test of these predictions, we conduct two exercises. First, we document differences by gender in answers to opinion polls from CIS (2004) and CIS (2007), which include questions about recent changes in gender-sensitive policies. Then, we replicate our findings for party effects in races where both contestants are women.

Table 8

Citizens Preferences

	Male	Female	p-value	Female-Left	Female-Right	p-value
Proportion declaring Ley de Dependencia was very necessary before the law was passed (2004) .	77.9 (1.23)	83.4 (1.07)	0.00	82.9 (1.57)	84 (1.45)	0.60
Proportion declaring Ley de Dependencia was insuficient after the law was passed (2007) .	59.30 (1.71)	67.8 (1.58)	0.00	66.4 (2.2)	69.4 (2.29)	0.35
Proportion declaring Ley de Igualdad was insuficient after the law was passed (2007) .	34.1 (1.5)	57.1 (1.57)	0.00	56.8 (2.16)	57.5 (2.3)	0.81

Note: This table shows the responses to three questions in CIS (2004) and CIS (2007): (i) How necessary was to pass a *Ley de Dependencia*, (ii) whether the coverage of the aforementioned law was enough, and (iii) whether the coverage of the *Ley de Igualdad* was enough. The panel on the left show the proportion of men and women who responded: (i) that it was very necessary and that (ii) and (iii) were insufficient – after removing *Don't Know/No Answer* –, along with the p-value of the t-test comparing the proportions. The panel on the right does the same comparison between women who identify their political views as left or right.

In Table 8 we report the proportion of men and women who deemed the elaboration of the *Ley the Dependencia* as "very necessary" and the proportion of men and women who found its coverage "insufficient" after the law was sanctioned. Both figures indicate that women demand significantly more public support for long-term care than men. On the other hand, when we compare the answers of women who self-identify as left-wing or right-wing, we find no significant difference. These patterns are also observed in questions referring to the more general *Ley de Igualdad* (Ley Orgánica 3/2007). The *Ley de Igualdad* was meant to "guarantee effectively the equality between men and women". We find a significant difference between male and female participants who consider this law "insufficient", but no significant difference between right- and left-wing women.[30]

By contrast, when we restrict our RD analysis of party effects to races in which both contestants are women, we find that right-wing female mayors are indeed less likely to deploy gender-sensitive policies than their left-wing counterparts (see Table 9). Taken together, these results are consistent with preferences over gender issues being different for male and female citizens, but political parties having the ability to impose their agenda to their female candidates, as proposed in the model.

Table 9

PP Mayor - Female Mayor & Female Oppositor

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	D. LTC	D. Preschool	D. Bal. Serv.	LTC (%)	Preschool (%)	Bal. Serv. (%)
PP Mayor	-0.618***	-0.601***	-0.229	-0.978	-1.236*	-0.753
	(0.204)	(0.162)	(0.162)	(1.128)	(0.687)	(0.590)
Observations	328	357	359	532	348	341
Clusters	134	145	146	218	142	139
p-value	0.002	0.000	0.158	0.386	0.072	0.202
Bandwidth	0.113	0.125	0.126	0.217	0.123	0.116

Note: Sample restricted to races between female candidates. Dummy outcomes in columns 1 through 3 take value one if spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services respectively is above zero. Percentages in columns 4 through 6 calculated as the fraction spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services as a percentage of total spending. 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.

Moreover, another implication of the model is that the presence of strong parties means other idiosyncratic characteristics would also be irrelevant for policy. In Table

[30] If we regress the demand for elaborating the *Ley de Dependencia* and the subjective sufficiency of *Ley de Igualdad* on political ideology measured in a scale from 1 to 10, we find that going right is correlated with less demand for gender policies, but only for male citizens.

B.14 in the Appendix, we report additional tests using a modified version of the RD specification with four other characteristics of the mayor.[31] The reported coefficients do not exhibit any clear pattern associating politicians' personal traits to gender-sensitive policies.[32] Furthermore, we find no relationship between traits and policies in the panel specifications reported in Appendix Table B.12.

7. Conclusions

Despite enduring gaps in political representation between women and men, the fraction of female politicians in leadership positions has increased in most countries over the past decades. This change in representation came alongside a debate about gender policies; policies specifically directed to reduce gender inequalities in, for example, labor market participation. A natural emerging question in this context is whether increased female representation in leadership positions fosters the application or extension of these gender-sensitive policies. We answer this question using budget information on municipal spending for Spanish local governments. We focus on Spain 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 implement a 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. Instead, we do find significant differences in the implementation of these policies by political affiliation, consistent with a setting with

[31] These characteristics are: sector of employment, white collar occupation, retired status, and college education. We construct mayoral characteristics as follows. *Health/Social/Edu* includes mayors that are medical doctors, nurses, social workers and teachers. *College education* includes mayors that have earned a college degree or higher. *White collar* and *blue collar* are standard. *Retired* includes all retired workers. As above, samples are restricted to mixed races, where mixed races are defined as municipal elections in which the most voted and second most voted candidates differ in the characteristic under consideration. Estimates of the effect of each personal trait on discrete and continuous measures of the policies under consideration are reported in Table B.14.

[32] Two coefficients are significant at conventional levels, but this is not surprising in a table with 24 coefficients in total.

strong commitment by candidates to party platforms, where the individual preferences of elected politicians do not influence policy. This result contrasts with some of the previous studies which focused on context-specific gender policies in developing countries and found strong effects for the gender of the politician (Chattopadhyay and Duflo, 2004; Clots-Figueras, 2012).

Our model illustrates a mechanism which can explain this divergence in results. In contexts in which parties are strong, female leaders that self-select into politics are not representative of their gender group. This is in line with the claim that women in politics are often pushed to adopt what has been labelled as normative masculine behavior in order to thrive (Despentes and Wynne, 2021; Fridkin and Kenney, 2014; Jones, 2016). In the model, this phenomenon would be particularly pronounced for one of the parties. Consistently with this prediction, previous empirical evidence has shown that right-wing female politicians tend to conform more often with normative masculinity when they run as presidential candidates in US primary elections (Jones, 2017), and, unlike men, are twice as likely to portray themselves as insiders in the US Senate than their left-wing female counterparts (Gulati, 2004). When we explore this in our data, we also find that, while female citizens are equally likely to demand gender policies irrespectively of their political orientation, right-wing female mayors are significantly less likely to adopt gender-sensitive policies than other female mayors.

An avenue for future research may be found in the study of the factors that promote the application of gender-sensitive policies under different institutional settings. Uncovering these factors could reveal a flexible agenda suitable to promote the interests of women in different contexts.

References

Alesina, Alberto. 1988. "Credibility and policy convergence in a two-party system with rational voters." *The American Economic Review*, 78(4): 796–805.

Alesina, Alberto, Traviss Cassidy, and Ugo Troiano. 2019. "Old and young politicians." *Economica*, 86(344): 689–727.

Bagues, Manuel, and Pamela Campa. 2021. "Can gender quotas in candidate lists empower women? Evidence from a regression discontinuity design." *Journal of Public Economics*, 194: 104315.

Bagues, Manuel, and Pamela Campa. Forthcoming. "Women and Power: Unpopular, Unwilling, or Held Back? Comment." *Journal of Political Economy.*

Baskaran, Thushyanthan, and Zohal Hessami. 2018. "Does the election of a female leader clear the way for more women in politics?" *American Economic Journal: Economic Policy*, 10(3): 95–121.

Beaman, Lori, Raghabendra Chattopadhyay, Esther Duflo, Rohini Pande, and Petia Topalova. 2009. "Powerful women: does exposure reduce bias?" *The Quarterly journal of economics*, 124(4): 1497–1540.

Beland, Louis-Philippe. 2015. "Political Parties and Labor-Market Outcomes: Evidence from US States." *American Economic Journal: Applied Economics*, 7(4): 198–220.

Besley, Timothy, and Stephen Coate. 1997. "An economic model of representative democracy." *The Quarterly Journal of Economics*, 85–114.

Besley, Timothy, Jose G Montalvo, and Marta Reynal-Querol. 2011. "Do educated leaders matter?" *The Economic Journal*, 121(554): F205–227.

Brollo, Fernanda, and Ugo Troiano. 2016. "What happens when a woman wins an election? Evidence from close races in Brazil." *Journal of Development Economics*, 122: 28–45.

Calonico, Sebastian, Matias D Cattaneo, and Rocio Titiunik. 2014. "Robust Nonparametric Confidence Intervals for Regression-Discontinuity Designs." *Econometrica*, 82(6): 2295–2326.

Calonico, Sebastian, Matias D Cattaneo, Max H Farrell, and Rocıo Titiunik. 2016. "Regression Discontinuity Designs Using Covariates." working paper, University of Michigan.

Casas-Arce, Pablo, and Albert Saiz. 2015. "Women and power: unpopular, unwilling, or held back?" *Journal of political Economy*, 123(3): 641–669.

Chattopadhyay, Raghabendra, and Esther Duflo. 2004. "Women as policy makers: Evidence from a randomized policy experiment in India." *Econometrica*, 72(5): 1409–1443.

CIS. 2004. "2581 - Opinion sobre la actuacion de la Administracion con las personas en situacion de dependencia." *Centro de Investigaciones Sociologicas.*

CIS. 2006. "2647 - Condiciones de Vida de las Personas Mayores." *Centro de Investigaciones Sociologicas.*

CIS. 2007. "2745 - Discriminaciones y su percepcion." *Centro de Investigaciones Sociologicas.*

CIS. 2011. "2920 - Postelectoral elecciones generales. Panel (segunda fase)." *Centro de Investigaciones Sociologicas.*

CIS. 2014. "3009 - Cuidados a Dependientes." *Centro de Investigaciones Sociologicas.*

Clots-Figueras, Irma. 2011. "Women in politics: Evidence from the Indian States." *Journal of Public Economics*, 95(7): 664–690.

Clots-Figueras, Irma. 2012. "Are Female Leaders Good for Education? Evidence from India." *American Economic Journal: Applied Economics*, 4(1): 212–44.

Crespo, Laura, and Pedro Mira. 2014. "Caregiving to elderly parents and employment status of European mature women." *Review of Economics and Statistics*, 96(4): 693–709.

De Paola, Maria, Vincenzo Scoppa, and Rosetta Lombardo. 2010. "Can gender quotas break down negative stereotypes? Evidence from changes in electoral rules." *Journal of Public Economics*, 94(5-6): 344–353.

Despentes, V., and F. Wynne. 2021. *King Kong Theory.* Farrar, Straus and Giroux.

Downs, Anthony. 1957. "An economic theory of political action in a democracy." *Journal of political economy*, 65(2): 135–150.

Edlund, Lena, and Rohini Pande. 2002. "Why have women become left-wing? The political gender gap and the decline in marriage." *The Quarterly Journal of Economics*, 117(3): 917–961.

Esteve-Volart, Berta, and Manuel Bagues. 2012. "Are women pawns in the political game? Evidence from elections to the Spanish Senate." *Journal of Public Economics*, 96(3-4): 387–399.

Fernández, Luis Ramiro, Celia Valiente Fernández, and Laura Morales Diez de Ulzurrun. 2003. "Mujeres en el Parlamento: un análisis de las desigualdades de género en el Congreso de los Diputados." *Revista de estudios políticos*, , (121): 179–208.

Ferreira, Fernando, and Joseph Gyourko. 2009. "Do Political Parties Matter? Evidence from US Cities*." *The Quarterly journal of economics*, 124(1): 399–422.

Ferreira, Fernando, and Joseph Gyourko. 2014. "Does gender matter for political leadership? The case of US mayors." *Journal of Public Economics*, 112: 24–39.

Fréchette, Guillaume R, Francois Maniquet, and Massimo Morelli. 2008. "Incumbents' interests and gender quotas." *American Journal of Political Science*, 52(4): 891–909.

Freier, Ronny, and Christian Odendahl. 2015. "Do parties matter? Estimating the effect of political power in multi-party systems." *European Economic Review*, 80: 310–328.

Fridkin, Kim L., and Patrick J. Kenney. 2014. *The Changing Face of Representation: The Gender of U.S. Senators and Constituent Communications.* University of Michigan Press.

Gagliarducci, Stefano, and M Daniele Paserman. 2012. "Gender interactions within hierarchies: evidence from the political arena." *The Review of Economic Studies*, 79(3): 1021–1052.

Giol, Jordi Capo. 1992. "La élite política local en España." *Revista de estudios políticos*, , (76): 127–144.

Gonzalez-Eiras, Martin, and Carlos Sanz. 2018. "Women's representation in politics: Voter bias, party bias, and electoral systems." *Banco de España Working Paper No. 1834.*

Gulati, Girish J. 2004. "Members of Congress and presentation of self on the World Wide Web." *Harvard International Journal of Press/Politics*, 9(1): 22–40.

Hessami, Zohal, and Mariana Lopes da Fonseca. 2020. "Female political representation and substantive effects on policies: A literature review." *European Journal of Political Economy*, 101896.

Hodler, Roland, and Paul A Raschky. 2014. "Regional favoritism." *The Quarterly Journal of Economics*, 129(2): 995–1033.

Imbens, Guido W, and Thomas Lemieux. 2008. "Regression discontinuity designs: A guide to practice." *Journal of econometrics*, 142(2): 615–635.

Iversen, Torben, and Frances Rosenbluth. 2006. "The Political Economy of Gender: Explaining Cross-National Variation in the Gender Division of Labor and the Gender Voting Gap." *American Journal of Political Science*, 50(1): 1–19.

Jones, Jennifer J. 2016. "Talk "like a man": The linguistic styles of Hillary Clinton, 1992–2013." *Perspectives on Politics*, 14(3): 625–642.

Jones, Jennifer Jean. 2017. "Talk" Like a Man": Feminine Style in the Pursuit of Political Power." PhD diss. UC Irvine.

Kanthak, Kristin, and Jonathan Woon. 2015. "Women don't run? Election aversion and candidate entry." *American Journal of Political Science*, 59(3): 595–612.

Kuk, John, and Zoltan Hajnal. 2020. "Democratic Party Control Reduces Gender Inequality." *Legislative Studies Quarterly*.

Landwerlin, Gerardo Meil, Pedro Romero Balsas, and Dafne Muntanyola Saura. 2012. "El uso de los permisos parentales en España The Social Use of Parental Leaves in Spain." *Working Paper.*

Lee, David S. 2001. "The Electoral Advantage to Incumbency and Voters' Valuation of Politicians' Experience: A Regression Discontinuity Analysis of Elections to the US." National Bureau of Economic Research.

Lee, David S, Enrico Moretti, and Matthew J Butler. 2004. "Do voters affect or elect policies? Evidence from the US House." *The Quarterly Journal of Economics*, 119(3): 807–859.

Levy, Gilat. 2004. "A model of political parties." *Journal of Economic theory*, 115(2): 250–277.

Lippmann, Quentin. 2019. "Gender and Lawmaking in Times of Quotas." *Available at SSRN 3330617.*

Mouritzen, Poul Erik, and James H Svara. 2002. *Leadership at the apex: politicians and administrators in Western local governments.* University of Pittsburgh Pre.

Osborne, Martin J, and Al Slivinski. 1996. "A model of political competition with citizen-candidates." *The Quarterly Journal of Economics*, 65–96.

Pettersson-Lidbom, Per. 2008. "Do parties matter for economic outcomes? A regression-discontinuity approach." *Journal of the European Economic Association*, 6(5): 1037–1056.

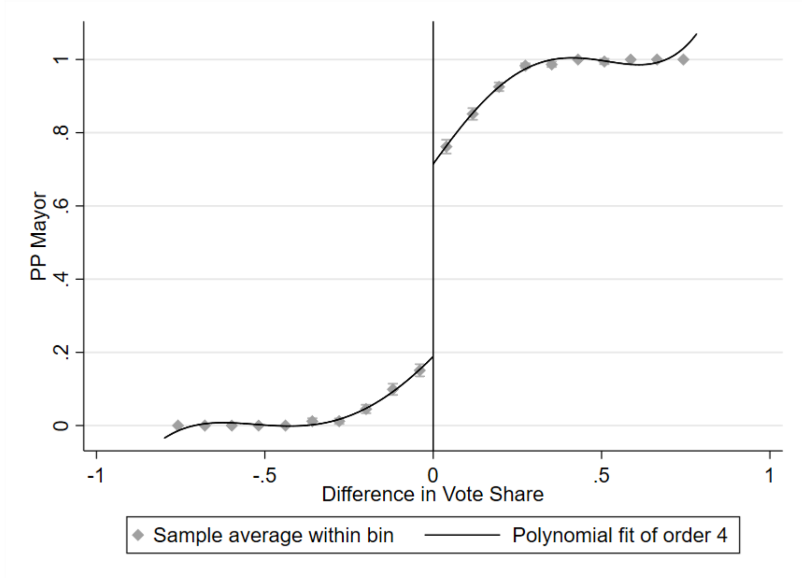
Quinn, Sheila. 2009. "Gender budgeting: practical implementation. Handbook." *Council of Europe.*

Solé-Ollé, Albert, and Elisabet Viladecans-Marsal. 2013. "Do political parties matter for local land use policies?" *Journal of Urban Economics*, 78: 42–56.

Appendices

A. Appendix Figures

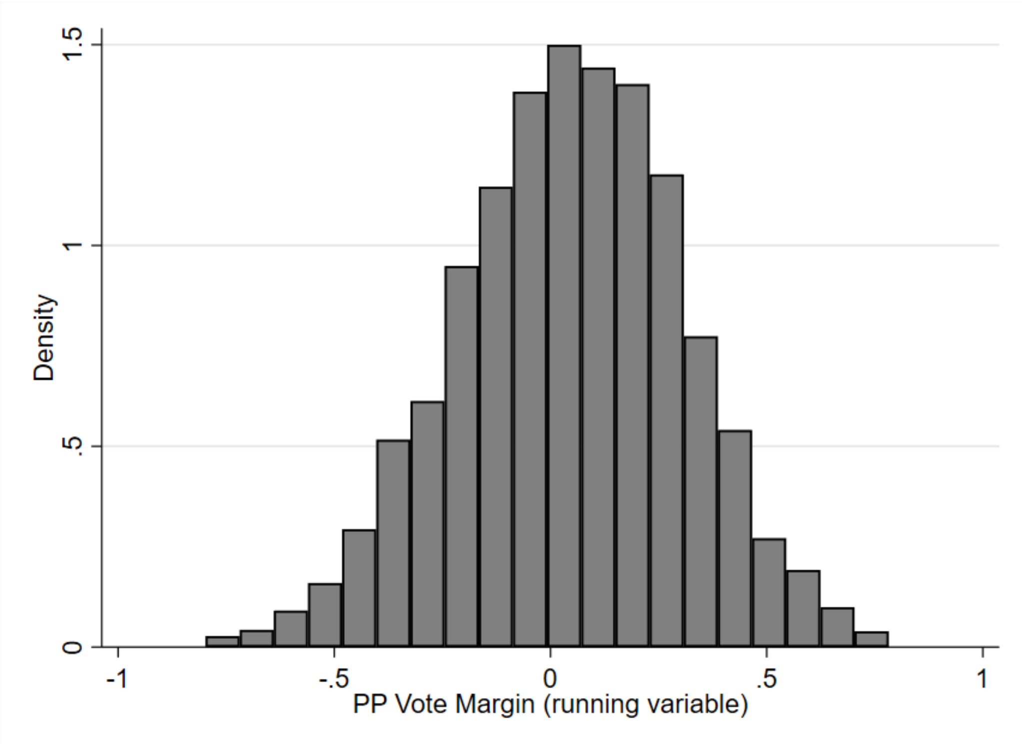
FIGURE A.1
FIRST STAGE - PP MAYOR



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.

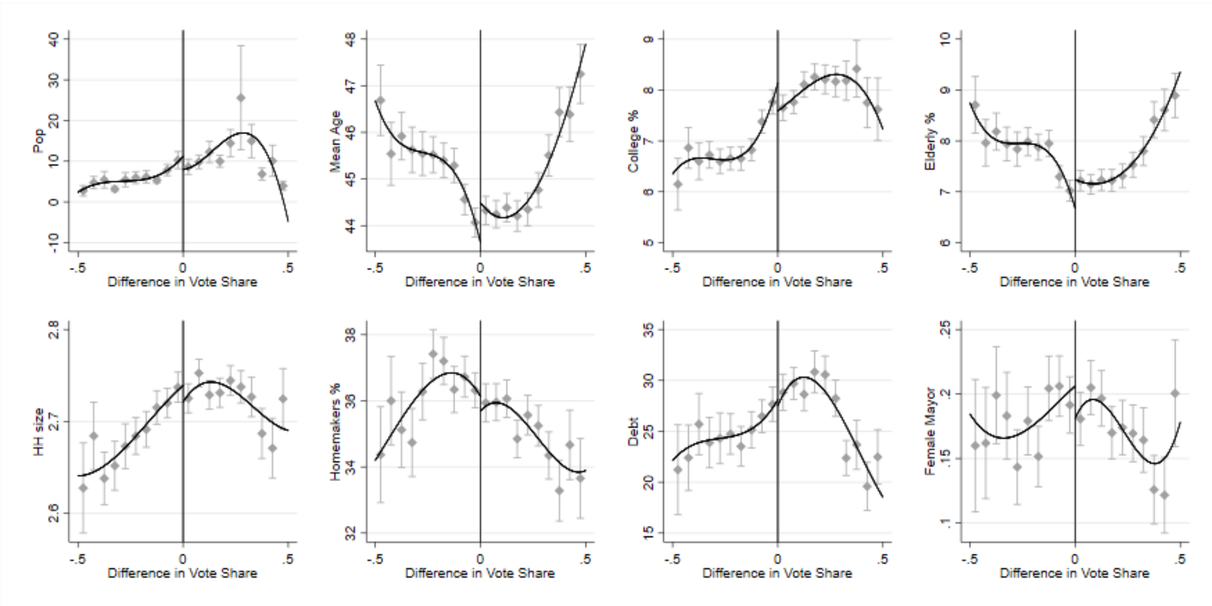
Figure A.2

Running Variable Histogram - PP Mayor



Note: 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.3
Covariate Balancing – PP Winners



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.

B. Appendix Tables

TABLE B.1

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
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.

TABLE B.2

FIRST STAGE - FEMALE MAYOR

VARIABLES	(1) Female Mayor	(2) Female Mayor	(3) Female Mayor	(4) Female Mayor
Female Winner	0.454***	0.461***	0.465***	0.478***
	(0.0435)	(0.0410)	(0.0389)	(0.0370)
Observations	3305	3717	4064	4433
Clusters	1411	1585	1735	1888
F-stat	109	126	143	166
Bandwidth	0.140	0.160	0.180	0.200

Note: Sample restricted to mixed races (by gender). 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. 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 Municipality

VARIABLES	(1) Population	(2) Mean Age	(3) College %	(4) Elderly %	(5) HH size	(6) Homemakers %	(7) Debt %
Female Mayor	-3.765	-1.048	-0.271	-0.875	0.0568	-0.649	6.506
	(7.597)	(1.018)	(0.986)	(0.674)	(0.0529)	(2.102)	(6.976)
Observations	1902	4381	4224	4202	4301	4339	4215
Clusters	805	1866	1802	1792	1834	1849	1798
p-value	0.620	0.303	0.783	0.194	0.283	0.758	0.351
Bandwidth	0.074	0.197	0.189	0.188	0.192	0.194	0.189

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.4

Female Mayor - Covariates Mayor

VARIABLES	(1) Age	(2) B Collar	(3) W Collar	(4) College	(5) No Studies	(6) Unemp.	(7) Houskp.
Female Mayor	-4.820**	0.0248	-0.118	0.107	-0.0430	-0.0113	0.0848
	(2.090)	(0.117)	(0.119)	(0.111)	(0.0464)	(0.0337)	(0.0535)
Observations	3542	2559	2909	3253	2592	2626	2373
Clusters	1792	1616	1808	1830	1481	1655	1512
p-value	0.021	0.832	0.319	0.336	0.353	0.738	0.113
Bandwidth	0.188	0.164	0.190	0.192	0.148	0.170	0.151

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.5

Reduced-Form – Female Winner and Gender Policies

VARIABLES	(1) D. LTC	(2) D. Preschool	(3) D. Bal. Serv.	(4) LTC (%)	(5) Preschool (%)	(6) Bal. Serv. (%)
Female Winner	-0.00568	-0.0450	-0.0162	-0.0346	0.0700	-0.0622
	(0.0412)	(0.0434)	(0.0441)	(0.172)	(0.204)	(0.0923)
Observations	3163	3443	2915	3616	3464	3165
Clusters	1351	1471	1249	1573	1511	1369
p-value	0.890	0.300	0.714	0.841	0.732	0.500
Bandwidth	0.134	0.147	0.123	0.158	0.151	0.135

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. 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.6

First Stage - PP Mayor

VARIABLES	(1) PP Mayor	(2) PP Mayor	(3) PP Mayor	(4) 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 (PP/No PP). 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. Different columns show different bandwidths.

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

Table B.7

PP Mayor - Covariates Municipality

VARIABLES	(1) Pop	(2) Mean Age	(3) College %	(4) Elderly %	(5) HH size	(6) Homemakers %	(7) Debt
PP Mayor	-5.525	0.648	-0.948	0.566	-0.00604	-1.295	1.809
	(4.337)	(0.801)	(0.610)	(0.534)	(0.0386)	(1.369)	(4.136)
Observations	7259	6538	6965	6548	7933	7885	7851
Clusters	2977	2680	2855	2684	3245	3226	3211
p-value	0.203	0.419	0.120	0.289	0.876	0.344	0.662
Bandwidth	0.146	0.129	0.139	0.130	0.161	0.159	0.158

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 and (7) Outstanding debt share of total budget in 2009. 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

Panel Estimates: Reduced Sample

	D.LTC	D.Preschool	D. Bal. Serv.	LTC	Preschool	Bal. Serv.
PP Mayor	-0.026***	0.005	-0.015*	-0.156**	0.011	0.002
	(0.009)	(0.009)	(0.009)	(0.063)	(0.043)	(0.019)
Female Mayor	-0.006	0.003	-0.006	0.006	0.018	0.022
	(0.009)	(0.009)	(0.009)	(0.061)	(0.038)	(0.020)
Observations	8929	8929	8929	8840	8840	8840

Notes: Results of estimating our panel specifications including municipality fixed effects, year dummies and a population control. Sample restricted to municipalities were a mixed race (party and gender) took place. Columns 1 through 3 correspond to dummy outcomes. Columns 4 to 6 correspond to the share of all spending corresponding to each policy. Standard errors in parentheses clustered at the municipal level. We report p-values for test of equality between the coefficients for the female and PP dummies.
* p < 0.1, ** p < 0.05, *** p < 0.01.

Table B.9

Robustness Checks – Female Mayors

	(1)	(2)	(3)	(4)	(5)	(6)
A) No Controls	D. LTC	D. Preschool	D. Bal. Serv.	LTC (%)	Preschool (%)	Bal. Serv. (%)
Female Mayor	0.0189	-0.0964	0.0453	0.0576	0.0988	-0.0154
	(0.0795)	(0.0913)	(0.0871)	(0.351)	(0.389)	(0.203)
Observations	3947	3634	3482	4422	4264	2782
p-value	0.812	0.291	0.603	0.870	0.800	0.940
Bandwidth	0.174	0.155	0.148	0.204	0.195	0.144
B) Degree 3 Pol.	D. LTC	D. Preschool	D. Bal. Serv.	LTC (%)	Preschool (%)	Bal. Serv. (%)
Female Mayor	0.0571	-0.124	0.0487	0.178	0.144	-0.0159
	(0.0973)	(0.108)	(0.101)	(0.428)	(0.514)	(0.217)
Observations	7664	7664	7664	7537	7535	7578
Clusters	3227	3227	3227	3227	3227	3227
p-value	0.558	0.248	0.630	0.677	0.780	0.942
Bandwidth	0.500	0.500	0.500	0.500	0.500	0.500

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.10

Robustness Checks – PP Mayors

	(1)	(2)	(3)	(4)	(5)	(6)
A) No Controls	D. LTC	D. Preschool	D. Bal. Serv.	LTC (%)	Preschool (%)	Bal. Serv. (%)
PP Mayor	-0.0753*	-0.0940	-0.129**	-0.208	-0.0194	-0.178
	(0.0422)	(0.0574)	(0.0540)	(0.388)	(0.290)	(0.173)
Observations	10978	7914	7044	7566	7497	7034
p-value	0.074	0.101	0.017	0.592	0.947	0.305
Bandwidth	0.230	0.160	0.142	0.152	0.151	0.141
B) Degree 3 Pol.	D. LTC	D. Preschool	D. Bal. Serv.	LTC (%)	Preschool (%)	Bal. Serv. (%)
PP Mayor	-0.115**	-0.125**	-0.102*	-0.0603	-0.110	-0.213
	(0.0572)	(0.0614)	(0.0563)	(0.456)	(0.309)	(0.187)
Observations	17163	17163	17163	17163	17163	17163
Clusters	6947	6947	6947	6947	6947	6947
p-value	0.044	0.042	0.070	0.895	0.721	0.253
Bandwidth	0.500	0.500	0.500	0.500	0.500	0.500

Note: Second-stage coefficients for the effect of having a PP 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.11

Same Samples: Party and Gender

A) Mayor Gender	(1) D. LTC	(2) D. Preschool	(3) D. Bal. Serv.
Female Mayor	0.110	-0.170	-0.0343
	(0.117)	(0.124)	(0.111)
Observations	2378	2378	2378
Clusters	983	983	983
p-value	0.343	0.171	0.757
Bandwidth	0.145	0.145	0.145

B) Party	(1) D. LTC	(2) D. Preschool	(3) D. Bal. Serv.
PP Mayor	-0.204**	-0.114	-0.167*
	(0.101)	(0.108)	(0.100)
Observations	2378	2378	2378
Clusters	983	983	983
p-value	0.044	0.293	0.095
Bandwidth	0.145	0.145	0.145

Note: RD estimates restricting the sample to mixed races in party and gender. Dummies take value one if spending in Long-term Care, Pre-schooling and Work and family-life Balancing Services respectively is above zero. 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.12

Other Identity Traits - Panel Estimates

	D. LTC	D. Pres.	D. Bal. Serv.	LTC (%)	Pres. (%)	Bal. Serv. (%)
Panel A						
H., SS. & E. Mayor	-0.002	0.002	0.017	0.024	0.044	0.041
	(0.011)	(0.010)	(0.012)	(0.074)	(0.057)	(0.032)
Observations	18958	18958	18958	18762	18774	18781
Panel B						
W. Collar Mayor	0.004	-0.003	-0.011	0.033	0.008	-0.011
	(0.008)	(0.007)	(0.008)	(0.049)	(0.039)	(0.020)
Observations	18958	18958	18958	18762	18774	18781
Panel C						
Retired Mayor	-0.005	0.002	0.016	0.046	0.022	0.002
	(0.017)	(0.014)	(0.016)	(0.139)	(0.062)	(0.036)
Observations	18958	18958	18958	18762	18774	18781
Panel D						
High Edu Mayor	0.007	-0.020**	0.003	0.006	-0.041	-0.013
	(0.008)	(0.009)	(0.009)	(0.063)	(0.044)	(0.021)
Observations	20613	20613	20613	20393	20405	20415

Notes: Results of estimating our panel specifications including municipality fixed effects, year dummies and a population control for other mayoral characteristics. Columns 1 through 3 correspond to dummy outcomes. Columns 4 to 6 correspond to the share of all spending corresponding to each policy. Standard errors clustered at the level of town-electoral period.

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

Table B.13

Other Policies - Disaggregated Categories

Female Spending	(1) Education share	(2) Health share	(3) Social share	(4) Pensions share
Female Mayor	-0.00767	0.00216	-0.00974	-0.00624
	(0.0109)	(0.00717)	(0.0157)	(0.00580)
Observations	3662	3443	4517	4049
Clusters	1562	1471	1924	1729
p-value	0.482	0.763	0.536	0.282
Bandwidth	0.157	0.147	0.205	0.179

Male Spending	(1) Environ. share	(2) Agricul. share	(3) Housing share	(4) Infrast. share
Female Mayor	0.00112	0.000256	-0.00921	0.0446**
	(0.00560)	(0.00567)	(0.0202)	(0.0207)
Observations	3889	4201	4359	3612
Clusters	1659	1792	1858	1542
p-value	0.841	0.964	0.648	0.031
Bandwidth	0.170	0.188	0.195	0.154

Notes: Second-stage coefficients for the effect of having a female mayor on different spending categories. All outcome variables are spending in each category as a share of all classifiable spending (male or female). Spending categories classified as *female* and *male* in Bagues and Campa (2021) reported in panels A and B, respectively. Standard errors clustered at the level of town-electoral period.

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

Table B.14

Other Identity Traits

VARIABLES	(1) D. LTC	(2) D. Pres.	(3) D. Bal. Serv.	(4) LTC (%)	(5) Pres. (%)	(6) Bal. Serv. (%)
Panel A						
Health/Social/Edu	0.0237	0.0369	0.0448	-0.0768	1.805**	0.172
	(0.120)	(0.123)	(0.121)	(0.607)	(0.778)	(0.239)
Obs.	1234	1029	1258	1063	814	1125
p-value	0.843	0.764	0.710	0.899	0.020	0.473
Band.	0.130	0.106	0.134	0.113	0.085	0.119
Panel B						
W. Collar	0.311***	0.0861	0.0461	0.774	0.650	0.341
	(0.117)	(0.112)	(0.116)	(0.621)	(0.590)	(0.229)
Obs.	1533	1913	1265	1667	1583	1072
p-value	0.008	0.442	0.691	0.212	0.271	0.137
Band.	0.169	0.213	0.129	0.188	0.176	0.111
Panel C						
Retired	-0.179	-0.0367	0.0105	-0.429	0.233	-0.310
	(0.136)	(0.130)	(0.108)	(0.759)	(0.784)	(0.529)
Obs.	942	1099	1006	909	637	877
p-value	0.188	0.778	0.922	0.571	0.767	0.558
Band.	0.164	0.194	0.177	0.162	0.124	0.158
Panel D						
Studies	0.0555	0.150	-0.0563	0.186	0.360	-0.215
	(0.0899)	(0.116)	(0.103)	(0.572)	(0.558)	(0.247)
Obs.	2368	2356	2228	2209	1954	1929
p-value	0.537	0.195	0.583	0.744	0.519	0.384
Band.	0.172	0.170	0.163	0.163	0.145	0.145

Note: Second-stage coefficients for the effect of having a mayor that works in Health, Social Sector or Education (Panel A), that is a white collar (Panel B), that is retired (Panel C) or that has voluntary studies (Panel D), 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 expressed in percentage terms. 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.

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

C. Theoretical Appendix

Voting

Voters are modeled following much of the literature on probabilistic voting models. There is a unit mass of atomistic voters. The individual utility of a voter i if electing party p is given by:

$$V_i(p) = c_p + \gamma_p^i + \nu_p \quad , \forall p = \{A, B\}$$

where c_p is the competence or charisma of the candidate run by party p, γ_p^i is an individual preference shock and ν_p is a aggregate preference shock in favor of party p. Voter i votes for the candidate run by party A if:

$$c_A - c_B > \gamma_B^i - \gamma_A^i + \nu_B - \nu_A$$

Re-labeling $\gamma_B^i - \gamma_A^i = \sigma^i$ and $\nu_B - \nu_A = \delta$ we obtain the standard expression from Persson and Tabelini. Individual i votes for party A if:

$$c_A - c_B > \sigma_i + \delta$$

In this expression, σ_i and δ correspond to shocks in favour of party B. As is customary, we specify a mean 0 distrubution for individual shock σ and a uniform distribution for δ such that $\delta \sim U[\frac{-1}{2\psi}, \frac{1}{2\psi}]$. The assumption on σ means that, conditional on δ, c_A and c_B, the probability of an election victory by A is given by $Pr(A \text{ win}|\delta, c_A, c_B) = Pr(c_A - c_b > \delta)$. Taking expectation over the distribution of δ, we are left with the probability of an A victory conditional on c_A and c_B is given by:

$$Pr(A \text{ win}) = \frac{1}{2} + \psi(c_A - c_B)$$

Parties

Parties are characterized by a preferred points in the policy space – a bliss point θ_p – and can punish mayors if they depart from it. Without loss of generality, we assume that $\theta_A > \theta_B$. As agents preferred policy $\hat{\theta}_i$ is private information, parties only select them based on their chance to win the election. Hence, they choose candidates so as to solve

$$Max \quad \frac{1}{2} + \psi(c_p^g - E(c_{-p})), \quad c_p^g = \{c_p^m, c_p^f\}$$

This mean ability is the only criteria to choose between male and female candidates, and party p chooses candidate g iff:

$$c_p^g \geq c_p^{-g}$$

Backward Induction

We start in the policy implementation stage. The election winner chooses θ_i^* so as to maximize $V(\theta_i)$ (see equation 2 in the text). Taking first order conditions in this concave objective function yields optimal policy as a function of $\hat{\theta}_i$ and θ_p:

$$\theta_i^* = h(\hat{\theta}_i, \theta_p) = \frac{\alpha\hat{\theta}_i + \gamma\theta_p}{\alpha + \gamma}$$

Notice that θ_i^* is a weighted average of the individual's and party's bliss points, and interior to the interval between both.

The set of agents that opt in to the pool of candidates who run for party p can be characterized in terms of their candidate types $\hat{\theta}$ as:

$$\mathbf{C}_p = \{\hat{\theta}_i : V(h(\hat{\theta}_i, \theta_p)) > 0, V(h(\hat{\theta}_i, \theta_p)) > V(h(\hat{\theta}_i, \theta_{-p}))\}$$

The first statement in the set definition is equivalent to a participation constraint (whether to run for a party or not) and the second relates to the preferred party to run in. The participation condition can be obtained by replacing θ_i^* in $V_i(\theta)$, and results in:

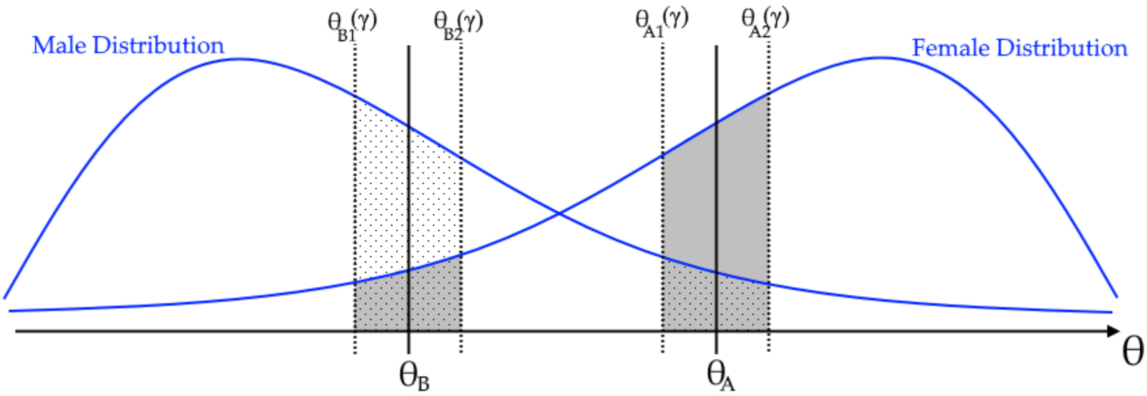
$$\frac{\alpha\gamma}{\alpha + \gamma}(\hat{\theta}_i - \theta_p)^2 - \omega < 0$$

Where $\dfrac{\alpha\gamma}{\alpha + \gamma}$ is positive and increasing in γ. Note that the left hand side of this inequality is a quadratic equation in vertex form, with the vertex of the parabola located where $\hat{\theta}_i = \theta_p$.

This gives us a lower and upper bound for $\hat{\theta}_i$ as a function of p and γ:

Figure C.1

Participation Regions in Agents' Type Space



Note: Regions of agents willing to participate. Gray areas correspond to the mass of female agents willing to postulate for a candidacy. Dotted areas correspond to the mass of male agents willing to postulate for a candidacy. Vertical dotted lines correspond to the limits of the participation regions for each party.

$$\hat{\theta_{1p}}(\gamma) = -(\frac{\omega(\alpha+\gamma)}{\alpha\gamma})^{\frac{1}{2}} + \theta_p$$

$$\hat{\theta_{2p}}(\gamma) = (\frac{\omega(\alpha+\gamma)}{\alpha\gamma})^{\frac{1}{2}} + \theta_p$$

The second condition in the definition of $\mathbf{C}_p$ is satisfied if $|\hat{\theta}_i - \theta_p| < |\hat{\theta}_i - \theta_{-p}|$. That is, individuals choose the party whose bliss point is closest to their own.

Figure C.1 illustrates the regions in θ space where agents are willing to postulate themselves for each party. Note that the limits of the participation regions are symmetric around θ_p, and are themselves a function of γ, with regions becoming narrower as γ increases.

Note that in this diagram we do not consider the case in which γ is sufficiently low so that the second condition above is binding for some agents (i.e. so low that some agents would obtain positive payoff from running with either party and have to choose between them). In that case the participation regions would meet at the mid point of θ_A and θ_B. The proof of proposition 1 below can be easily accommodated to include this case.

Equilibrium Policy

The average choice of θ in equilibrium can be computed by averaging over the optimal policy choices of the chosen candidates – the average $h(\theta, \theta_p)$ within the participation regions – for each party and multiplying these by the probability that each party is elected (1/2 in equilibrium).

$$E(\theta^*|g,p) = \frac{\int_{\hat{\theta}_{1p}(\gamma)}^{\hat{\theta}_{2p}(\gamma)} h(x,\theta_p) f_g(x) dx}{F_g(\hat{\theta}_{2p}(\gamma)) - F_g(\hat{\theta}_{1p}(\gamma))}$$

$$E(\theta^*|g) = \frac{1}{2}E(\theta^*|g,A) + \frac{1}{2}E(\theta^*|g,B)$$

Proposition 1

We assume that functions $f_f(\hat{\theta})$ and $f_m(\hat{\theta})$ are unimodal with modes $\overline{\theta}_f$ and $\overline{\theta}_m$. Note that, given the assumption of stochastic dominance above, we will have that $\overline{\theta}_f > \overline{\theta}_m$. Assume that $\theta_p \in (\overline{\theta}_m, \overline{\theta}_f)\ \ \forall p$ and $\gamma < \underline{\gamma}$. Finally, assume γ is bounded by below so that, $\gamma > \underline{\gamma}$ with $\underline{\gamma} \equiv max\left[\frac{\alpha\omega}{\alpha(\overline{\theta}_f - \theta_A)^2}, \frac{\alpha\omega}{\alpha(\theta_B - \overline{\theta}_m)^2}\right]$. If we define the difference in expected implemented policy by gender as $\Delta_G(\gamma)$, we will have that $\frac{\partial \Delta_G(\gamma)}{\partial \gamma} < 0$.

Proof:

When an agent is elected as mayor, they implement their optimal policy:

$$\theta_i^* = h(\hat{\theta}_i, \theta_p) = \frac{\alpha\hat{\theta}_i + \gamma\theta_p}{\alpha + \gamma}$$

Therefore, the average policy conditional on gender as a function of gamma can be expressed as:

$$E(\theta_i^*(\gamma)|g) = \frac{\alpha E(\hat{\theta}_i(\gamma)|g) + \gamma\theta_p}{\alpha + \gamma}$$

Taking derivatives:

$$\frac{\partial E(\theta_i^*(\gamma)|g)}{\partial \gamma} = \left(\alpha\frac{\partial E(\hat{\theta}_i(\gamma)|g)}{\partial \gamma} + \theta_p\right)\frac{1}{\alpha + \gamma} - (\alpha E(\hat{\theta}_i(\gamma)|g) + \gamma\theta_p)\frac{1}{(\alpha + \gamma)^2}$$

Taking differences between men and women:

$$\frac{\partial E(\theta_i^*(\gamma)|f)}{\partial\gamma} - \frac{\partial E(\theta_i^*(\gamma)|m)}{\partial\gamma} = \underbrace{\left(\frac{\partial E(\hat{\theta}_i(\gamma)|f)}{\partial\gamma} - \frac{\partial E(\hat{\theta}_i(\gamma)|m)}{\partial\gamma}\right)}_{I} \frac{\alpha}{\alpha+\gamma} - \underbrace{(E(\hat{\theta}_i(\gamma)|f) - E(\hat{\theta}_i(\gamma)|m))}_{II} \frac{\alpha}{(\alpha+\gamma)^2}$$

Note that $\frac{\partial \Delta_G(\gamma)}{\partial\gamma} = \frac{\partial E(\theta_i^*(\gamma)|f)}{\partial\gamma} - \frac{\partial E(\theta_i^*(\gamma)|m)}{\partial\gamma}$ so the proof of proposition 1 can be obtained by showing that the terms I and II in the equation above are negative and positive, respectively.

Term I is negative because the expected value of $\hat{\theta}$ is decreasing in γ for females and increasing in γ for males. This follows from the fact that the pdf is increasing in θ for females (and decreasing for males) in the relevant interval (see Figure C.1). Likewise, the assumption of first order stochastic dominance in the main text directly yields $E(\hat{\theta}_i(\gamma)|f, p) > E(\hat{\theta}_i(\gamma)|m, p)\ \forall p$, which shows II is positive.

■