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JEL Classification: D72

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Electoral systems and female representation in politics: evidence from a regression discontinuity

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Abstract

This work looks at the impact of electoral rules on female participation in local legislative bodies using a natural experiment involving a series of changes to electoral law in Poland. Using an exogenous population threshold dividing municipalities into ones with proportional and ones with majoritarian elections, we estimate the effect of each electoral system on female representation. Contrary to the literature on the national elections, we find that more females are elected to local councils under a majoritarian system. We link this observation to countering party bias in list placements and lower costs of electoral participation in the majoritarian system.

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1. Introduction

The question of female representation in politics has recently gained more attention, being important for at least two reasons: First, women are underrepresented on the political stage across the globe and scholarship is looking for policies that would promote a more gender-balanced political system. Second, a growing body of literature shows that female politicians make different

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decisions than their male counterparts. Understanding the factors behind differences in female representation across countries can help understand the differences in policies and their impact.

This work looks at the effects of electoral systems on promoting female participation in legislative bodies using a series of changes to electoral law in Poland.[3] Electoral rules are rarely introduced in view of how they affect female political candidates, contrary to other outcomes of electoral systems such as, for instance, fiscal performance and corruption (Persson and Tabellini, 2004; Persson et al., 2007). The effects of electoral systems on female representation are mostly a side effect and arguably unintended, minimizing the problem of reverse causality. To further mitigate the problem of omitted variable bias, we exploit a natural experiment occurring at the level of local elections in Poland, specifically elections to municipal councils. Several aspects of the law make it perfect for a well-identified causal analysis of how electoral rules affect female representation in politics. The law in place between 2002 and 2010 stipulated a population threshold (20,000) that divided the municipalities into ones with majoritarian and ones with proportional elections. We use this threshold to conduct a regression discontinuity (RD) analysis of an effect of electoral rules on female representation. Furthermore, the threshold has since been removed in 2014 allowing us to run placebo regressions.

There is a general consensus in the literature, that using proportional representation as compared to majoritarian elections is associated with a higher number of elected females (Norris, 1985; Matland and Studlar, 1996; Castles, 1981; Rule, 1981; Kenworthy and Malami, 1999). Several features of proportional representation are believed to be responsible for this phenomenon. Firstly, majoritarian elections are typically associated with choosing the strongest candidate to run for office in a particular district, whereas in the proportional elections the lists of candidates should reflect a wide spectrum of voters. In this, gender might be a more important factor in a candidate-centered majoritarian election (as males are typically perceived as stronger candidates), particularly if the voters are biased against females. Nevertheless, literature exists that states that

[3]Earlier research studying female representation in politics in Poland was mainly looking at the effects of gender quota in the context of elections to the national parliament (see, e.g., Górecki and Kukołowicz (2014); Jankowski and Marcinkiewicz (2019); Gendźwiłł and Żółtak (2020)).

higher competition tends to equalize male and female chances for both nomination and election if purely meritocratic features are all that matter (Folke et al., 2016). Secondly, district magnitude differs significantly between electoral systems: proportional systems have consistently higher district magnitudes, so parties can pull from deeper in their lists, which scholars have argued increases the chances of women being elected (Norris, 2006; King, 2002). Thirdly, electoral systems affect the competitiveness of particular seats (Profeta and Woodhouse, 2018). Since females are typically more competition-averse than males (Niederle and Vesterlund, 2007), this might translate into differences in female participation. Namely, females are less likely to compete for seats in majoritarian elections, which are typically more competitive or built around a competition between individuals rather than one between party lists. Finally, incumbency advantage, which is higher in majoritarian systems, is believed to be responsible for the strong persistence of individuals of the same gender (i.e. male) occupying elected positions (Lippmann, 2017; Schwindt-Bayer, 2005). Moreover, a recent study Gonzalez-Eiras and Sanz (2021) identified that more female candidates, councilors, and mayors are present under closed-list proportional representation and showed that this effect is driven by the supply of females and party bias.

All of the above arguments suggest that proportional elections have an advantage over majoritarian regimes in promoting female nomination and election. Yet, these arguments are mainly established for the national elections in older democracies, overlooking the fact that at the local level and in young democracies the nomination and election dynamics may differ. We propose another channel and explanation for why single-seat districts could be at times more favorable towards female candidates. Typically, the costs of entering electoral races are much lower in local elections than in national ones. In the Polish context, this is particularly true in the majoritarian elections, where practically anyone can take part in an electoral race as the backing of a particular political party or electoral committee is not mandatory and the ballot access requirements are very low (see Brancati (2008) for the argument about why low ballot access requirements favor independent candidates). In fact, majoritarian regimes, where support of political parties is not required, are known to create more opportunities for independent candidates to be elected. This situation may be particularly attractive for female candidates who, as stipulated before, have

a tendency to be more competition-averse. By individually putting themselves forward as candidates, they can avoid intra-party/intra-committee competition, first, for the nomination and, second, for favorable placement on the list. Generally females might prefer electoral systems where they can bypass party leadership, which is often perceived by women as providing female recruits with less strategic and financial support than the male recruits (Butler and Preece, 2016). The whole electoral system in Poland is also quite young, which means that stringent party elites are not yet developed, which gives more room for independent local committees, instead of established political parties. Independent local committees, which tend to rely on more decentralized procedures for candidate selection, contrary to national party leadership, might be less biased against female participation. Furthermore, the local independent committees may be more focused on the competencies of candidates, contrary to political parties, which recruit the candidates who can then potentially compete in elections at the higher level, thus the competitiveness for them may be of crucial importance. All this leads to a situation where, paradoxically, at the local level, majoritarian elections may be more favorable towards female representation. Which argument prevails – whether the majoritarian or proportional representation election type most increases female representation – is an empirical question, and we aim to test it by using the quasi-experimental setting at the local level in Poland.

While our work is mostly contributing to the literature examining the causal effects electoral systems have on female representation, it indirectly also speaks to other streams of literature inasmuch as female participation affects policy outcomes and economic performance. Policy outcomes have been found to be affected by the gender of the elected official by e.g., Chattopadhyay and Duflo (2004); Adams and Funk (2012); Bhalotra and Clots-Figueras (2014); Brollo and Troiano (2016); Hicks et al. (2016); Clayton and Zetterberg (2018). Female decision-makers have been found to prioritize policies related to e.g., public health (Bhalotra and Clots-Figueras, 2014) and child health (Schwindt-Bayer, 2006; Miller, 2008) and deprioritize military spending (Clayton and Zetterberg, 2018). On the other hand, education investments have been found to be more supported by males by Chattopadhyay and Duflo (2004) yet by females by Clots-Figueras (2012). On the local level, policy shifts have been found by e.g., Svaleryd (2009); Slegten and Heyndels (2019);

Bratton and Ray (2002). On the other hand, Ferreira and Gyourko (2014) find no policy effects of female leaders when it comes to the size of local government, the composition of municipal spending and employment, or crime rates. Geys and Sørensen (2019); Bagues and Campa (2021); Rigon and Tanzi (2012) do not find any evidence for policy shifts at the local level but Hessami and Baskaran (2019) show that at least the discussions in the local council are affected by the gender composition with females having a greater preference for child-care investments.

Our main results point to several policy-relevant conclusions. First, majoritarian elections seem to help females become elected, compared to proportional elections. The mechanism in place touches upon the distinction between national parties and independent local committees. Nationwide party lists, more prevalent in proportional elections, typically place females on lower places - less than 30 percent of first-list placements in proportional elections are occupied by females. On the other hand, in single-seat districts in the majoritarian system, females are, by definition, placed on first places and since voters in the analyzed elections do not seem to prefer any gender, it is easier for women to get elected from single-seat districts. And since political parties have fewer females elected than independent committees, and the latter are more likely to become elected in majoritarian elections, this results in an overall higher number of females elected in the majoritarian elections. Second, lower costs of participation lead to more female candidates, who do not have as much backing from a political party. These two results suggest that female participation can be promoted by single-seat districts, as long as these counteract the penetration of local political markets by national parties, and, on the contrary, promote local electoral committees. Our results are in line with the recent findings of Gonzalez-Eiras and Sanz (2021), which identify party bias and supply effects as driving the results between closed-list and open-list proportional representation. Also Le Barbanchon and Sauvagnat (2019) found, for a majoritarian electoral system in France, that intrinsic party bias is an important driver of lower female representation.

This paper is structured as follows: the next section gives an overview of the institutional set-up in Poland. Section 3 describes the details of the data and the empirical approach. Section 4 presents the main results; channels of transmission for the main observations are discussed in

Section 5. Robustness of the main results is tested in Section 6. Section 7 concludes the paper. Additional tables, figures and methodological information are presented in the Appendices.

2. The institutional background

Municipalities (Polish: *gmina*) are the principal units of administrative division in Poland, and constitute the lowest tier of government.[4] There are currently 2,477 municipalities, varying in size between 1,400 and 1.7 million inhabitants. The municipalities in Poland enjoy relatively high fiscal autonomy on both what comes in and how it is spent. They are responsible for primary education, health care, local road infrastructure, provision of utilities and spatial planning, among other concerns. On the income side, municipalities are free to raise their own funds, which stem from municipal assets, from local taxes and fees (of which property taxes are the most important), and from a revenue sharing mechanism.

This relatively high fiscal autonomy is matched by a high degree of democratic accountability. The legislative and controlling body of each *gmina* is the elected municipal council (*rada gminy*) or, in a town, the town council (*rada miasta*). Executive power is, since 2002, held by the directly elected mayor of a municipality. The position of the mayor is rather strong in the Polish context. First, this is the body which puts forward all local resolutions, including budgetary resolutions, which are then voted on during the plenary sessions by the members of municipal councils. Second, mayors typically have big support within the municipal councils as they typically represent the largest local committees (parties).

We exclude a certain group of municipalities from the analysis, which have special rights of districts (*gminy na prawach powiatu*, typically larger cities). Besides all municipal competencies, they are vested with competencies of districts, and thus with the competencies of an intermediary level of government.

[4] In Poland there are two intermediary levels of governments called districts (powiaty) and voivodeships (wojewodztwa), respectively.

2.1. Elections in the years 2002 to 2010

In small municipalities (below 20,000) between 2002 and 2010, the councilors were elected in small districts via plurality rule. In most municipalities that meant elections from single-seat districts. The electoral law prescribed, however, that a maximum of five members should be elected from one district, hence, in larger districts representatives were elected through block voting, which meant that a voter could cast as many votes as there were seats to be filled. As a result, the most popular party might have been able to win every seat, creating even larger disproportionality between votes and seat distribution than in the case of single-seat districts (Lijphart, 2012).

In municipalities larger than 20,000 inhabitants, proportional elections were in place. Since 2002 the d'Hondt method has been implemented to translate votes into seats; before 2002, the Sainte-Laguë method had been applied. Moreover, in the municipalities with the proportional systems, more members were elected from each district than in the municipalities with majoritarian regimes, i.e. five to eight, as stipulated by the law.

It is important to stress that starting in 1998 candidates for the municipal councils needed to be supported by an electoral committee. The committee can be established by political parties, public associations and organizations. It can also be created by a group of five voters, of whom one is nominated as a proxy of the committee. Furthermore, in municipalities with proportional elections the registration of candidates' list in a district was conditional on gathering, at minimum, 150 voter signatures in a district. Another rule stated that the candidates' list in proportional elections had to contain, at minimum, the same number of candidates as the number of mandates assigned to a given district. It turns out that the registration of candidates' list was much easier in majoritarian elections. In smaller municipalities only 25 signatures from the district population were required to register a list, and the list was eligible for elections if it contained just one name. These differences in the ballot access requirements are not uncommon and are in line with the argumentation that the majoritarian elections promote independent candidates who should not face too high ballot access requirements. These requirements are naturally more stringent in the proportional elections where competition occurs between committees or parties and not individual

candidates. Given that in municipalities with proportional elections the costs of entering political markets were much higher (Szczepanowska, 2010) and the fact that proportional systems introduce the idea of competing elites, ideologies or sectoral interest rather than geographical interests (Shugart and Carey, 1992; Gendzwill and Zoltak, 2014), it was much easier for national political parties to penetrate local political markets under proportional elections[5]. Kantorowicz (2017) and Kantorowicz and Köppl-Turyna (2019) show that there is indeed a substantial difference in the share of council members affiliated with political parties between majoritarian and proportional systems. The final difference between the two systems is the size of the council. As will be shown in the paper, while elections in 2014 used the same electoral systems above and below the threshold, and two different council sizes, there has been no difference in the fraction of females elected or any other analyzed outcomes. This suggests that the council size has no effect on the chances of females (further robustness checks of the effect of the council are performed in Section 6. The summary of the institutional features of the municipal electoral systems used in the years 2002 to 2010 is presented in Table 1.

Table 1: Institutional details of electoral systems in the years 2002 to 2010

	Majoritarian	**Proportional**
Population size	<20,000	>=20,000
Electoral rule	Plurality	Proportional (d'Hondt)
District magnitude	1 to 5	5 to 8
No of signatures to register a party list in a district	Min. 25	Min. 150
No. of candidates on the list	Min. 1	Min. 5
Confounding factors		
Council size	15	21
Campaign spending limitation	750PLN	1000PLN

2.2. Elections in 2014

According to the electoral law enacted in 2011[6], the local elections conducted in November 2014 would use an entirely new election procedure. In particular, the 20,000-inhabitants threshold

[5]It is important to underscore that before the 1998 elections national political parties were banned from putting forth their candidates in local elections. They could merely support particular candidates or local committees (Kotarba, 2016).

[6]*Ustawa z dnia 5 stycznia 2011 r. Kodeks wyborczy, DzU 2011, nr 21, poz. 112.*

was removed and all members of the councils in all municipalities had been elected via plurality rule in single-seat districts. The sizes of the council, however, remained the same as in the previous years.

3. Data and the empirical model

3.1. Data

The data used comes from the National Electoral Commission of the Republic of Poland. It contains information about all candidates to local councils, including their gender, age, citizenship, position on the electoral list, the fraction of obtained votes, and whether the person has been elected to the council or not. We match this information with data about the institutional set-ups of each municipality, i.e., the electoral system in place, the size of the council, number of mandates from each electoral district in each municipality and other. Data is available for the years 2002[7] to 2014, and includes a total of 786,880 candidates, among which 228,155 were females. Elections take place every four years (2002, 2006, 2010, 2014) and are held in all municipalities simultaneously. In municipalities above 20,000 inhabitants, proportional representation were in place, while in those below the threshold single-seat districts and block voting were used. Thus, within one municipality below 20,000 inhabitants some districts are single-seat, while in others between two and five candidates are elected using block voting, which results in mixed cases. In 2014 all municipalities used single-seat districts.

In the 2002 municipal election, a total of 299,827 candidates competed for 46,805 seats in local councils, 76,496 of which were female. A fraction of 24.4% of candidates were female below the 20,000 threshold, and 26.5% above the threshold. In 2006, 198,136 candidates competed for 39,944 seats and 57,992 candidates were females. The fractions below and above the threshold are at 28.1% and 30.7%, respectively. In the 2010 municipal election, a total of 159,863 candidates, among which were 51,035 women, competed for 37,818 seats in local councils. The fractions of female candidates above and below the 20,000 inhabitants threshold are almost the same at

[7] We would like to thank Adam Gendzwiłłfor providing us with the data for the years 2002 and 2006.

31.6% and 32.4%, respectively. In 2014, we observe 131,799 candidates, among which 43,527 were women, who competed for 36,109 councils seats. Also in this case, the fractions of females at both sides of the population threshold of 20,000 inhabitants are almost the same at 33.8% and 29.9%, respectively.

3.2. Empirical approach and identification

The question of how electoral rules (and other institutional and political factors) affect female participation in politics can be addressed in two ways:

1. What is the probability of a female candidate being elected?
2. What is the probability of encountering a woman among elected candidates?

The relationship between the two approaches can be easily summarized by the Bayes' formula:

$$P(female|elected) = \frac{P(elected|female) \times P(female)}{P(elected)}, \tag{1}$$

where $P(female)$ is the pool of females among all candidates and $P(elected)$ is the general, gender-independent probability of getting elected in a particular electoral system. As it is *a priori* unclear which approach gives a better answer to the question of female representation in politics, we shall look at both probabilities and try to establish the institutional reasons behind the differences. Using our data, we can separately estimate each of the probabilities in the formula. We estimate the conventional (sharp) regression-discontinuity (RD) estimator using the the non-parametric approach. If $\mu_t(x) = \mathbb{E}[Y|X = x]$ is the expectation of Y given x, the conventional RD estimator is given by:

$$\tau = \lim_{x \to 0^+} \mu(x) - \lim_{x \to 0^-} \mu(x) = \mu_+ - \mu_-. \tag{2}$$

The main dependent variables are the fraction of females in the council, and the fraction of elected officials among female candidates in a municipality. In these cases, the dependent variable is continuous, but obviously taking values between 0 and 1. Standard linear or polynomial

estimations do not guarantee in this case that the fitted values would remain within the $[0, 1]$ interval. A standard solution to this problem is to apply the log-logit transformation (Papke and Wooldridge, 1996) of the form $\log \frac{y}{1-y}$, where y is the fractional response. In this case, however, we cannot recover the $E(y|x)$, but we can still interpret the sign and the significance of the discontinuity, without interpreting the actual size of the effect. We consider this approach as a robustness check to the standard polynomial model. In the main text, we focus on the actual policy-relevant outcome, which is the fraction of females in the councils in each analyzed election. This outcome, however, depends, on the individual probabilities of each female actually being elected. Therefore, we also look at the individual-level probabilities of becoming elected as a female, and of encountering a female among elected officials. This is done using an RD estimator for categorical outcomes, which is described in more detail in Appendix B.

One issue worth mentioning is the fact that results of elections are conditional on whether candidates decide to candidate. In such a case, Lee (2009) and Anagol and Fujiwara (2016) proposed a method of estimating a bounds on treatment effect. At the same time, if females and males have the same likelihood to compete again in the following election after not receiving a mandate in the previous one, the effects identified in this paper are still valid. Following Anagol and Fujiwara (2016) this relies on the assumption that those candidates who decided not to compete would have had the same probability of winning - which is unobservable. The evidence on this assumption is mixed: e.g., Bernhard and de Benedictis-Kessner (2021) show that that women candidates are no more dissuaded from seeking office again after losing than men are, while Wasserman (2018) shows that females become more discouraged than males after losing elections. Still, there is no systematic evidence on how these probabilities would be different between different electoral systems.

3.3. Validity of the RD design

Several conditions have to be met for the sharp RD design to be valid. First, the running variable must deterministically predict the treatment, which in our case is fulfilled, as the law prescribes the population threshold of 20,000 inhabitants that groups the municipalities into one

of the distinctive electoral systems. Second, no manipulation of treatment should be possible - which is analyzed in Section 3.3.1. Third, treatment must not be correlated with any outcome-determining factor - this assumption is considered in Section 3.3.2. Finally, confounding factors should not be present at the analyzed population threshold. As indicated in Table 1, two policies change at the 20,000 threshold: the size of the council and the limit to campaign financing. Regarding campaign expenditure limits, there is evidence that these are not strictly enforced, and thus are not binding (Szyszko, 2014). It is a well-known practice in Poland for parties to engage in so-called "pre-campaigning" in order to circumvent expenditure limits. This means that politicians begin agitation before the beginning of the official campaign, i.e., when the expenses of campaigning go unreported (Szyszko, 2014; Kantorowicz and Köppl-Turyna, 2019). Moreover, to confirm the validity of the result, we can exploit the fact that the campaign limit changes again at the threshold of 40,000 inhabitants from 1000 to 1200 PLN (and this is the only change at this threshold). We address this issue in the robustness section and find no evidence that our outcome variable changes at this threshold, which is strong evidence of validity. When it comes to the council size, we test the validity in two ways. First, as described in 2, in 2014 the changes to the electoral system were removed, but the change in the council size remained in place. Observing no differences in female participation in 2014 provides evidence that council size does not play a role in determining the results. Secondly, in the robustness section, we also look at the changes to female representation at a different threshold. We exploit the fact that in municipalities above 50,000 inhabitants, the council size increases to 23 persons, and above 100,000 to 25. We find no evidence that these institutional factors play a role in our results.

3.3.1. Sorting

For the regression-discontinuity assumptions to be valid, we need to establish whether sorting around the cutoff does not appear. We first look at the McCrary (2008) test. Following that, however, there is evidence that the test is sensitive to the choice of the bin sizes and assignment of borderline integer observations (as in the case of population) to the first bin on the left of the threshold (Eggers et al., 2017). To account for that, Table 2 presents the results of density testing

with different bin sizes, to control for sensitivity. In all tests and figures in the following sections, the number of inhabitants is expressed in thousands. For exemplification, several specifications for the year 2010 are presented graphically in Figure 1. The figures for 2014 look essentially the same.

Table 2: McCrary density tests - p-values reported

Bin size	2002	2006	2010	2014
100	0.2611	0.8141	0.3096	0.2269
200	0.2684	0.8257	0.3040	0.2329
300	0.2958	0.7488	0.3850	0.2121
400	0.3101	0.9322	0.3103	0.2097
500	0.1514	0.9006	0.2214	0.2266
1000	0.3330	0.9875	0.2784	0.1962

Bandwidths in thousand.

Alternatively, one can test density continuity with local polynomial density estimators (see, Cattaneo et al., 2016, 2019), which avoids pre-binning of the data.[8] Results are summarized in Table 3 and similarly point to no sorting present at the population threshold.

Table 3: Polynomial density tests - p-values reported

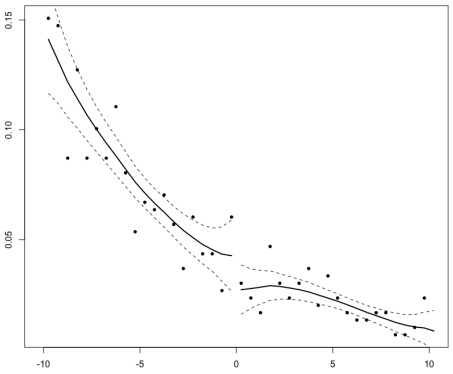
	2002		2006		2010		2014	
Order	Tri	Uni	Tri	Uni	Tri	Uni	Tri	Uni
1	0.4916	0.5506	0.5994	0.4715	0.3256	0.3776	0.1151	0.074
2	0.4139	0.7017	0.8387	0.6941	0.4722	0.5664	0.4289	0.2486
3	0.9851	0.9466	0.4353	0.2924	0.3754	0.7862	0.1945	0.2452
4	0.905	0.7955	0.8487	0.8515	0.6309	0.7129	0.3017	0.2191

3.3.2. Continuity of the municipal characteristics

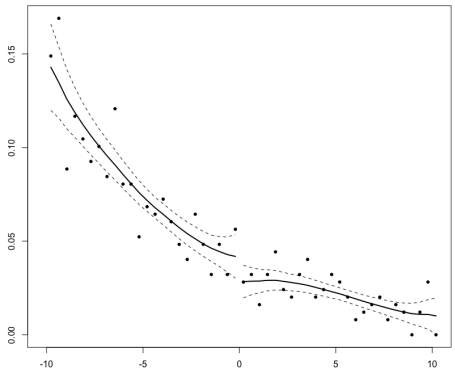
Furthermore, we need to make sure that variables potentially affecting the female representation - determined prior to the realization of the assignment variable - have the same distribution just above and just below the threshold of interest, so that local randomization is secured. In the Appendix (Figure A.9), we show a variety of continuity checks of several important variables. Since the variables are very stable over time, we only report the values for the year 2010, as other years look virtually the same. Population variables could affect the probability of voting for females independent of the electoral system: young people could vote more "progressively" (young defined as population between 18 and 30 years of age), as opposed to an older population (defined

[8]Implemented in R with *rddensity* package.

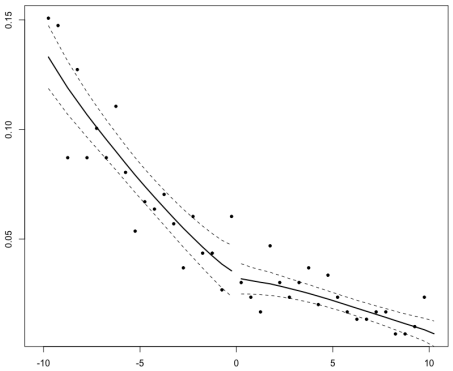
Figure 1: McCrary Test of manipulation of reported number of municipality inhabitants



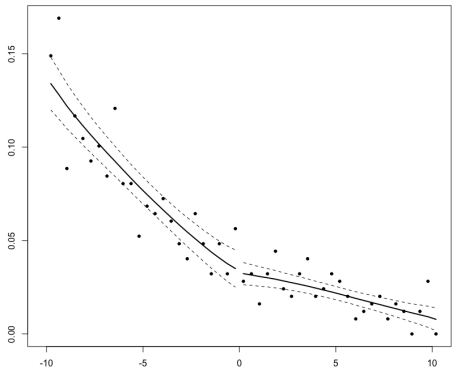
(a) BW=5, bin=500



(b) BW=5, bin=416



(c) BW=10, bin=500



(d) BW=10, bin=416

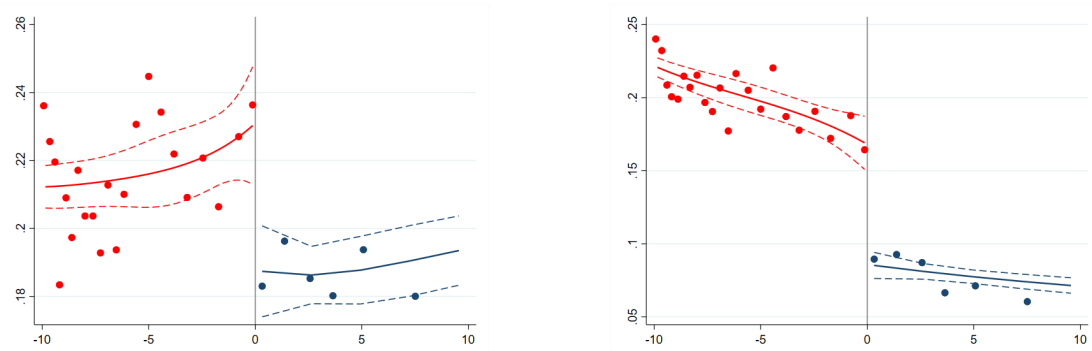
as above 65). However, we find no discontinuity in the distribution of age at the population threshold of interest. More dense areas, that is, cities, could be more positive towards females, but there is no visible discontinuity in the population density. Finally, percentage of females in the overall population could matter. Also in this case, we do not observe any discontinuity.

4. Results

4.1. Years 2002 to 2010

Tables 4 and 5 show the results at the municipality level: the percentage of women in the local council (Table 4) and the percentage of elected among females in the municipality (Table 5). In both cases, we observe a discontinuity at the population threshold. Above the threshold, where proportional elections are used, there are about three to six percentage points fewer females in the councils and this is significant at the 5% level. Similarly, above the threshold, there are less elected among female candidates. The size of the effect equals about eight percentage points and is significant at the 1% level. Nevertheless, as will be explained later on, this second effect is driven by the overall lower probability of election in the proportional system. The first effect, though, cannot be explained by this feature, and will be analyzed in further detail. Figure 2 shows the discontinuity in the years 2002 to 2010 and confirms the existence of a significant jump at the analyzed threshold, and the size of the gap.

Figure 2: Elections in years 2002 to 2010



(a) Percentage of females in the council

(b) Percentage of elected among females

Table 4: Percentage of females in the local council, Years 2002 to 2010, $P(female|elected)$

	(1)	(2)	(3)
RD Estimate	-0.035**	-0.032	-0.064**
	[0.014]	[0.019]	[0.029]
Robust 95% CI	[-.071 ; -.003]	[-.075 ; .013]	[-.135 ; -.013]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	1128	1488	992
Eff. Observations R	408	426	392
Conventional p-value	0.015	0.096	0.024
Robust p-value	0.035	0.168	0.017
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	8.473	9.773	7.869
BW Bias (b)	12.048	11.937	10.289

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table 5: Percentage of elected among females in municipality, Years 2002 to 2010, $P(elected|female)$

	(1)	(2)	(3)
RD Estimate	-0.078***	-0.074***	-0.081***
	[0.013]	[0.018]	[0.023]
Robust 95% CI	[-.108 ; -.046]	[-.113 ; -.032]	[-.135 ; -.036]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	1024	1268	1246
Eff. Observations R	398	412	411
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.001
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	8.061	9.034	8.930
BW Bias (b)	11.413	11.263	10.977

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

4.2. Year 2014

Tables 6 and 7 do not display any discontinuity at the 20,000 population threshold in 2014. There are no differences in the fraction of females in the local council, nor are there any differences in the probabilities of being elected among female candidates. This is an expected result as the 20,000 population threshold was cancelled in 2014 and plurality rule was introduced uniformly in all municipalities. Figure 3 illustrates the 2014 discontinuity. We can generally observe that no discontinuity can be found in any of the variables at the 20,000 threshold. For a percentage of females elected from a municipality reported in Table 7 there is a significant jump for the third polynomial, but since no other formulations result in significant effects this can be considered an outlier.

Table 6: Percentage of females in the local council, Year 2014 $P(female|elected)$

	(1)	(2)	(3)
RD Estimate	0.038	0.044	0.039
	[0.043]	[0.053]	[0.070]
Robust 95% CI	[-.056 ; .146]	[-.073 ; .167]	[-.121 ; .188]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	89	137	143
Eff. Observations R	53	79	84
Conventional p-value	0.375	0.406	0.583
Robust p-value	0.380	0.443	0.672
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	2.825	4.051	4.240
BW Bias (b)	4.199	5.277	4.958

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

5. Channels of transmission

5.1. Individual-level probabilities

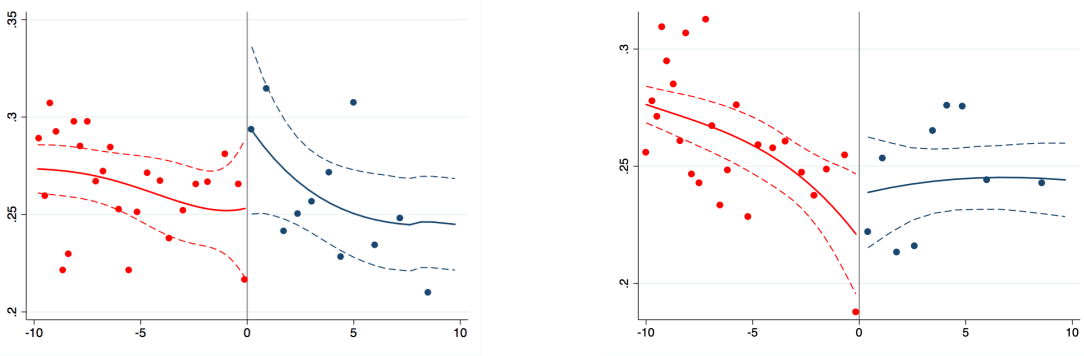
The analysis so far has concentrated on the aggregate outcome of the election, represented by the female representation in the municipal council. The aggregation at the municipality level

Table 7: Percentage of elected among females in municipality, Year 2014 $P(elected|female)$

	(1)	(2)	(3)
RD Estimate	0.023	0.034	0.115***
	[0.025]	[0.031]	[0.042]
Robust 95% CI	[-.039 ; .075]	[-.03 ; .105]	[.037 ; .213]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	223	347	215
Eff. Observations R	114	134	113
Conventional p-value	0.352	0.279	0.006
Robust p-value	0.532	0.272	0.006
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	5.920	7.798	5.780
BW Bias (b)	10.116	10.837	8.080

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Figure 3: Election in 2014

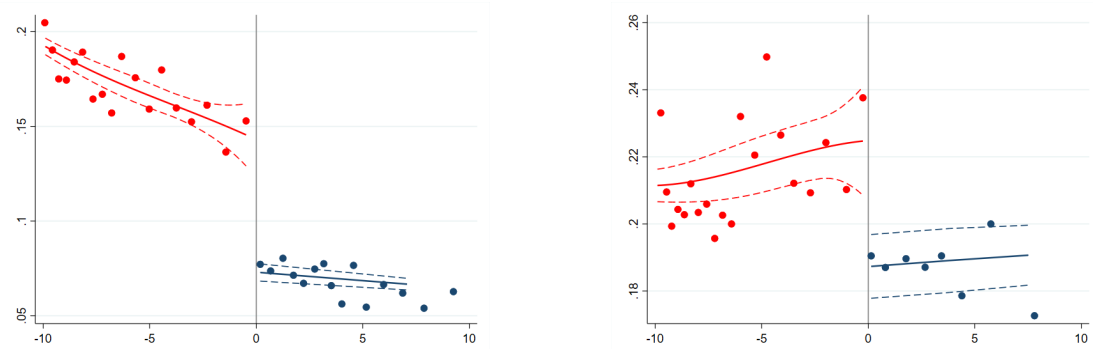


(a) Percentage of females in the council

(b) Percentage of elected among females

results in averaging the results over electoral districts of different sizes. A question remains, however, whether the observed differences are reflected in individual-level probabilities, by which we mean a probability of being elected as each individual female, i.e., $P(elected|female)$, and the opposite, i.e., probability that an elected candidate is a female $P(female|elected)$. Appendix B analyzes this question in more depth, and explains the details of the estimation. Here we present a summary of the main results in a graphical illustration. The following pattern emerges: in years 2002 to 2010, the majoritarian system was associated with higher individual probabilities of election for females, or of encountering a female among elected candidates. Corresponding to the municipality-level result, the probability of finding a female among elected candidates drops by three to seven percentage points above the threshold. This means that averaging over the districts within each municipality does not change the conclusions.

Figure 4: Individual-level probabilities in the years 2002 to 2010



(a) Individual probability that a female candidate gets elected.(b) Individual probability that an elected candidate is a female.

5.2. Probability of election

The unconditional probability $P(elected)$ defines the overall probability of being elected from a totality of candidates in a municipality, independent of gender. As we can see from Table 8, there is a visible jump in election probability at the 20,000 threshold, in which seats below the threshold are generally less contested. This is the channel which finds that females are more often candidates for less-contested seats, which is consistent with the previous literature. These results suggest that the probability that the female candidate gets elected is mostly dependent on

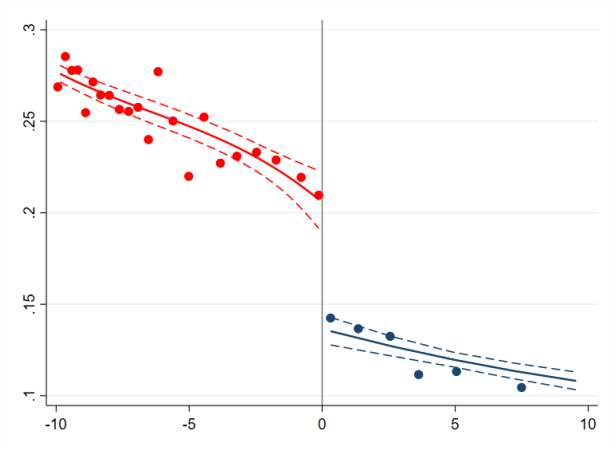
the overall probability of election, which is much lower in the proportional system. The jump in the probability of election does not explain, however, the overall lower fraction of females in the council.

Table 8: Probability of election, years 2002 to 2010 (municipal level): $P(elected)$

	(1)	(2)	(3)
RD Estimate	-0.070***	-0.068***	-0.067***
	[0.012]	[0.016]	[0.018]
Robust 95% CI	[-.099 ; -.039]	[-.103 ; -.032]	[-.104 ; -.027]
Kernel Type	Triangular	Triangular	Triangular
BW Type			
Eff. Observations L	675	1113	1994
Eff. Observations R	349	405	459
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.001
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	6.303	8.419	11.058
BW Bias (b)	8.947	10.423	13.386

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Figure 5: Probability of election $P(elected)$, year 2010



5.3. Female representation on electoral lists

As required by Equation 1, the conditional probability of encountering a female among candidates elected to local councils also depends on the overall fraction of females $P(female)$ on

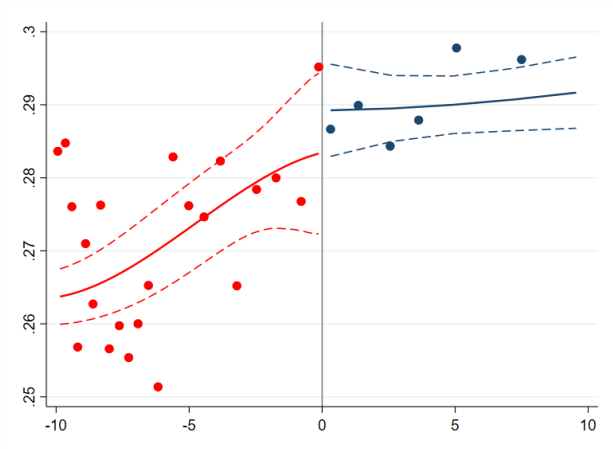
electoral rules. We have seen in Section 3 that on average there is a comparable number of females in both systems; now we turn to looking more specifically at the population threshold. Table 9 shows that no discontinuity is present at the 20,000 threshold regarding the fraction of female candidates. This implies that the overall lower female representation in the council is not driven by the lower representation of females on electoral lists between the two systems. Figure 6 shows the percentage of females on electoral lists in the years 2002 to 2010.

Table 9: Percentage of females on electoral lists (municipal level): $P(female)$, years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	0.004	0.003	0.005
	[0.010]	[0.011]	[0.014]
Robust 95% CI	[-.017 ; .029]	[-.022 ; .03]	[-.022 ; .037]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	620	1474	2151
Eff. Observations R	340	426	469
Conventional p-value	0.655	0.810	0.690
Robust p-value	0.619	0.765	0.611
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	6.037	9.709	11.428
BW Bias (b)	8.998	12.282	13.950

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Figure 6: Percentage of females on electoral lists $P(female)$, years 2002 to 2010



The fraction of females on the electoral lists does not differ between the two electoral systems in the years 2002 to 2010. This means that, a priori, there is a comparable number of females who stand for election in both systems.[9] This raises the question of what drives the difference in representation of females in the council, when there is a comparable fraction of females in both systems.

5.4. List placements

Since there is a comparable number of females on electoral lists in both systems, an obvious candidate for a driver of the results is list placement. To answer the question of how list placement changes, we look at the *relative placement* of females on lists, to account for the fact that there are differences in district magnitude - and list lengths - between the electoral systems. To do so, first, we normalize the placements. As the lengths of electoral lists are different, we normalize the position on the list with respect to the list length: *relative position* is defined as the actual position divided by the length of the electoral list, in which a lower number corresponds to a higher position on the electoral list. Then to avoid the problem of this method not working for single-member districts, we define the difference between a candidate's relative position on the list and mean relative position. This outcome takes value 0 for candidates on a single-member list, reflecting the fact that they are both the first and the last and are thus best compared to the mean candidate on other lists (also taking the value of 0). The results are presented in Table 10. There is strong evidence that list placement is a significant predictor of female electoral success. At the 20,000 threshold, the relative position on the list of a female changes significantly. So although females are equally represented on lists in both systems, they tend to be placed on significantly lower places in the proportional system.

Yet, most of the elected candidates are those placed on the top of the list. In fact, in the proportional system 46 percent of councilors are those placed on the top of the list. 85 percent of

[9] The fact that there is a comparable number of females at both sides of the threshold does not preclude the possibility that the types of females (e.g. their education levels) are different. Unfortunately, we do not have information about further characteristics in the dataset.

Table 10: Relative position on a list (individual level), years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	0.019***	0.026***	0.029**
	[0.005]	[0.008]	[0.009]
Robust 95% CI	[.007 ; .032]	[.011 ; .044]	[.012 ; .051]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	16734	16280	22674
Eff. Observations R	20952	20883	23480
Conventional p-value	0.000	0.001	0.001
Robust p-value	0.002	0.001	0.001
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.054	6.945	8.623
BW Bias (b)	10.572	9.899	11.247

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

all places are filled with candidates from the top five list positions. So in order to better compare the two systems, we include list placements of only the top five candidates in both systems, and compare the relative placement. Results are presented in Table 11. Taking the first five positions on electoral lists, we find that females are placed lower in this case as well, which negatively affects their probability of election.

Finally, we can compare what happens at the very top of the list (or on the actual list in single-seat districts), which gives the best possible comparison group between the two systems. First, Table 12 shows that in the proportional system there is a significantly higher probability of becoming elected if placed on the top of the list than there is in the majoritarian system. This can be explained by a lower number of electoral lists in the proportional system that stand for election. Second, there are significantly fewer females on the top of the list in the proportional system, as shown in Table 13. Finally, this results in a significantly lower percentage of females elected from the top of the list in the proportional system,[10] as shown in Table 14.

[10]No such differences are present from the positions two to five; the corresponding results can be obtained upon request.

Table 11: Relative position on a list - top five positions (individual level), years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	0.017***	0.024**	0.027**
	[0.005]	[0.007]	[0.009]
Robust 95% CI	[.006 ; .029]	[.01 ; .042]	[.011 ; .047]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	16349	15030	21747
Eff. Observations R	10967	10705	12126
Conventional p-value	0.001	0.001	0.001
Robust p-value	0.004	0.002	0.002
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	6.971	6.603	8.404
BW Bias (b)	10.491	9.579	10.834

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table 12: Probability of election from the top position on the list (individual level): years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	0.218***	0.218***	0.203***
	[0.014]	[0.016]	[0.022]
Robust 95% CI	[.186 ; .251]	[.185 ; .256]	[.153 ; .244]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	29697	61535	48141
Eff. Observations R	6624	8456	7946
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	5.583	9.036	7.788
BW Bias (b)	8.419	12.340	10.373

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table 13: Females on top position on the list (individual level): years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.089***	-0.077***	-0.099***
	[0.015]	[0.015]	[0.019]
Robust 95% CI	[-.13 ; -.062]	[-.113 ; -.052]	[-.143 ; -.063]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	17554	64104	58665
Eff. Observations R	4720	8511	8427
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	3.650	9.247	8.761
BW Bias (b)	6.785	14.208	11.735

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table 14: Elected females from the top position on the list (individual level): years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.068***	-0.076**	-0.143***
	[0.019]	[0.023]	[0.033]
Robust 95% CI	[-.117 ; -.026]	[-.131 ; -.027]	[-.223 ; -.086]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	6822	13002	8903
Eff. Observations R	2907	3584	3210
Conventional p-value	0.000	0.001	0.000
Robust p-value	0.002	0.003	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	5.871	8.600	6.936
BW Bias (b)	8.790	10.829	9.949

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

5.5. Partisanship

In Kantorowicz and Köppl-Turyna (2019), it is shown that the 20,000 population threshold is responsible for a jump in the partisanship of local councils. In particular, it can be established that the proportional system is associated with a much higher percentage of local councilors and mayors who are members of nationwide parties, as opposed to independent voter committees which prevail in the majoritarian system. We demonstrate this result, for the sake of clarity of interpretation of our results, also in Table 15, which presents the percentage of independent candidates in the council dropping by about 15 to 18 percentage points above the 20,000 population threshold.

Table 15: Percentage of candidates in the council elected from independent committees (municipal level), years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.153***	-0.175***	-0.184***
	[0.033]	[0.043]	[0.048]
Robust 95% CI	[-.238 ; -.088]	[-.281 ; -.091]	[-.291 ; -.082]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	557	719	1212
Eff. Observations R	322	359	410
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	5.427	6.502	8.778
BW Bias (b)	8.772	8.741	10.547

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

We link this observation to lower costs of participation in elections below the threshold. The question arises of whether the observed patterns with regard to female participation can be linked to different partisanship. To answer this question we calculate the percentage of females in the electoral council, looking separately at candidates elected from independent lists versus nationwide parties. Tables 16 and 17 present the results.

According to the results, partisanship is an important driver of the election outcomes. In years 2002 to 2010, no difference can be observed in the percentage of councilors from indepen-

dent committees in the two systems. Above the threshold from 2002 to 2010, there are about 10 percentage points fewer females elected from nationwide parties. So, for the independent committees the overall lower representation of females above the threshold is essentially driven by the lower electoral success of these committees above the threshold. But, as can be observed also within the nationwide parties, there is a slightly lower probability of having a female councilor. This, as presented in Table 18, is additionally driven by a significantly lower list placement of females above the threshold in the first five places.[11] So while most of the partisanship effect is driven by the representation of independent committees, some additional negative effect comes from lower placements of females on lists within nationwide parties in the proportional representation system.

Table 16: Percentage of females in the council elected from independent committees (municipal level), years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.006	-0.000	-0.019
	[0.019]	[0.025]	[0.033]
Robust 95% CI	[-.043 ; .044]	[-.056 ; .057]	[-.1 ; .044]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	813	1188	1101
Eff. Observations R	369	409	403
Conventional p-value	0.767	0.997	0.563
Robust p-value	0.992	0.977	0.447
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.007	8.715	8.377
BW Bias (b)	10.827	10.754	10.195

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

5.6. District magnitude

As mentioned in Section 2, below 20,000 inhabitants, both block voting and single-seat districts are being used. We expect that these two systems might produce different effects. Typically, block

[11] This relationship is also significant for all places.

Table 17: Percentage of females in the council elected from nationwide parties (municipal level), years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.089**	-0.093**	-0.110*
	[0.043]	[0.045]	[0.054]
Robust 95% CI	[-.201 ; -.004]	[-.208 ; -.01]	[-.234 ; .009]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	410	1177	1798
Eff. Observations R	296	410	445
Conventional p-value	0.038	0.039	0.042
Robust p-value	0.042	0.031	0.071
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	5.532	10.277	12.154
BW Bias (b)	8.562	13.854	13.649

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table 18: Relative position on a list - top five positions (individual level) for nationwide parties, years 2002 to 2010

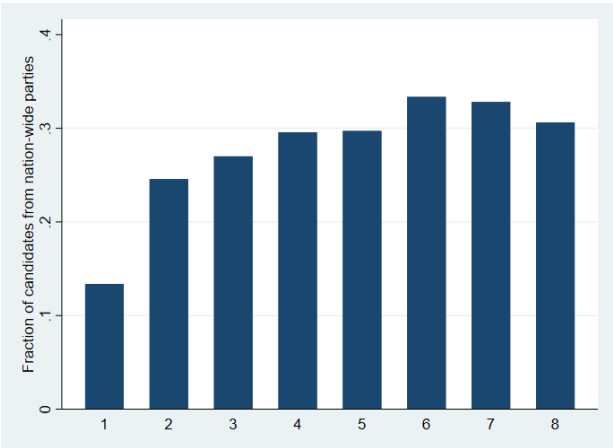
	(1)	(2)	(3)
RD Estimate	0.024**	0.030**	0.029*
	[0.010]	[0.014]	[0.015]
Robust 95% CI	[.003 ; .049]	[0 ; .063]	[-.004 ; .061]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	4120	4279	8724
Eff. Observations R	3486	3505	4325
Conventional p-value	0.013	0.035	0.055
Robust p-value	0.026	0.049	0.087
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.277	7.427	11.119
BW Bias (b)	10.856	9.728	13.220

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

voting would be more similar to proportional representation. We base this hypothesis on findings of, e.g. Carey and Shugart (1995), who find that if the electoral system fosters party reputation, the effect is stronger in districts of higher magnitude.

When we change the threshold of the minimum fraction of single-seat districts in a municipality to be included in the sample, two effects take place: on the one hand, with a more selective sample we expect stronger effects, which we associate with elections in single-seat districts. On the other hand, a too restrictive sample results in a very low number of observations, which makes a reliable estimate impossible. Figure 8 shows the size of the discontinuity dependent on what minimum fraction of single-seat districts is included in the sample of municipalities below 20,000 inhabitants. We can also observe that the fraction of nationwide parties is much higher in districts of higher magnitude, and increasing in magnitude, as presented in Figure 7. In fact, there is almost no difference between the block-voting districts and the (smaller) districts in proportional representation.

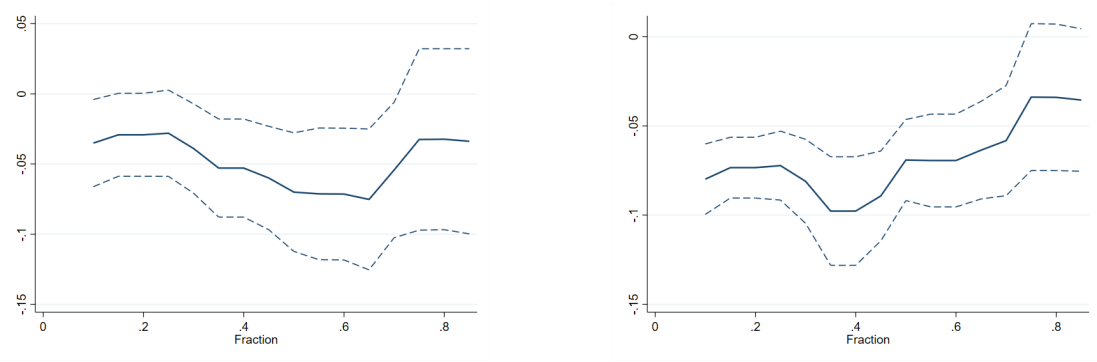
Figure 7: Fraction of candidates from nationwide parties dependent on district magnitude



6. Robustness analysis

As mentioned before, one of the assumptions of a valid RD design is a lack of confounding factors. At the analyzed threshold of 20,000, the causal interpretation of our results could be confounded by the change in the size of the council. Nevertheless, given that the size of the council

Figure 8: Discontinuity dependent on the fraction of single-seat districts included.



(a) Percentage of females in the council

(b) Percentage of elected among females

increases at the 20,000 threshold, we see no reason to believe that this would somehow *decrease* the chances of women to be elected, given that they are typically placed lower on electoral lists. This has been mentioned, e.g., by Matland and Brown (1992). This suggests that, if anything, our results would *underestimate* the effect of the electoral system. Empirically, however, no change in the fraction of elected females or any other analyzed outcomes can be observed at the 20,000 threshold in 2014, in which both above and below the 20,000 threshold the same electoral system was used, but the council sizes remained as in the previous years. Moreover, at the 50,000 inhabitants threshold, there is a further increase in the size of the council and we will look at the change in female representation at this threshold. Notice, however, that there are many fewer municipalities around the 50,000 mark. Tables C.23 and C.24 in the Appendix show the results. No visible discontinuity is present, but since there are only a few observations available, these results could be not reliable.

The size of the council has, however, an additional effect. It indirectly affects the impact of the percentage of females on electoral lists $P(female)$, as the same percentage of females is elected to councils of different sizes. As a robustness check, we therefore look at the *number of female candidates per seat*, in which we normalize the number of females by the size of the council. The results are shown in Table C.27 in the Appendix, and show that the different council sizes do play a role. While there are no differences in the fraction of females on the electoral lists, the seats above the threshold are more contested between the female candidates, as there are fewer females

per available seats.

Secondly, as mentioned in the previous sections, a confounding factor is present at the 20,000 threshold, which is the limit on campaign financing. We use the fact that a similar change is present at the threshold of 40,000 inhabitants to show that campaign financing is not associated with any change in the outcome variables. We present both the municipality-level as well as individual-level observations, as the number of observations at the 40,000 threshold is quite low, in the Appendix in Tables C.28 to C.31. For no specification there is any evidence of a change in the outcome at the 40,000 threshold, for which only the campaign finance limits change. We consider this strong evidence that the results are not driven by this factor.

The results of the log-logit transformation model for the municipal level in years 2002 to 2010 are presented in Tables C.25 and C.26 in the Appendix and confirm the main conclusions.

Figure C.10 shows the sensitivity of the regression discontinuity estimate for the percentage of females in the council in year 2010 to the choice of bandwidth. As can be observed, the results are very stable across bandwidths between 1,000 and 10,000 inhabitants.

7. Conclusions

Our results have important policy implications. Contrary to the literature for national level elections, we find that single-seat districts might be a better choice for promoting female representation. Firstly, in majoritarian elections, women have lower costs of entering the electoral race. Secondly, majoritarian elections give women the opportunity to free themselves from party nomination procedures and intra-party competition. The latter conjecture is clearly visible in our data: far fewer females are elected from nationwide parties compared to independent committees, and nationwide parties are more prevalent in proportional elections. Finally, in the proportional representation system there are ex ante as many females on electoral lists as in the majoritarian system, but they have a lower probability of becoming elected as they are placed lower on the electoral lists. This raises a question of whether, also in other set-ups, it could still be more favorable to promote females through single-seat districts - as opposed to the often found positive

effect of proportional representation - whenever these counteract the negative effect of penetration of local politics by nationwide parties.

While the established literature comes to a conclusion that proportional representation is more favorable to females, it turns out that this result does not hold in every context. In this study we analyzed a case, which is different to most of the literature in at least three aspects: we look at local as opposed to national-level elections; we look at a new democracy, in which a social status of females is generally quite high, but nevertheless certain institutions are not (yet) well established; and finally we look at a less partisan context, i.e., the context where the nationwide parties are not prominent or, in other words, where most of the candidates represent local and independent committees. We find a result, which is in opposition to most of the literature, and we do so using a clear identification strategy, which assures a very high level of internal validity. The question remains, and it should be placed on the agenda of future research, of why is this the case that in our context majoritarian elections promote more females.

We contribute to the discussion of how to design institutions, which would help get more females elected into electoral bodies. At the very least, our results point to the fact, that some established results from the literature do not hold always and everywhere. In the next step and further studies, we should aim at explaining the factors responsible for the observed results.

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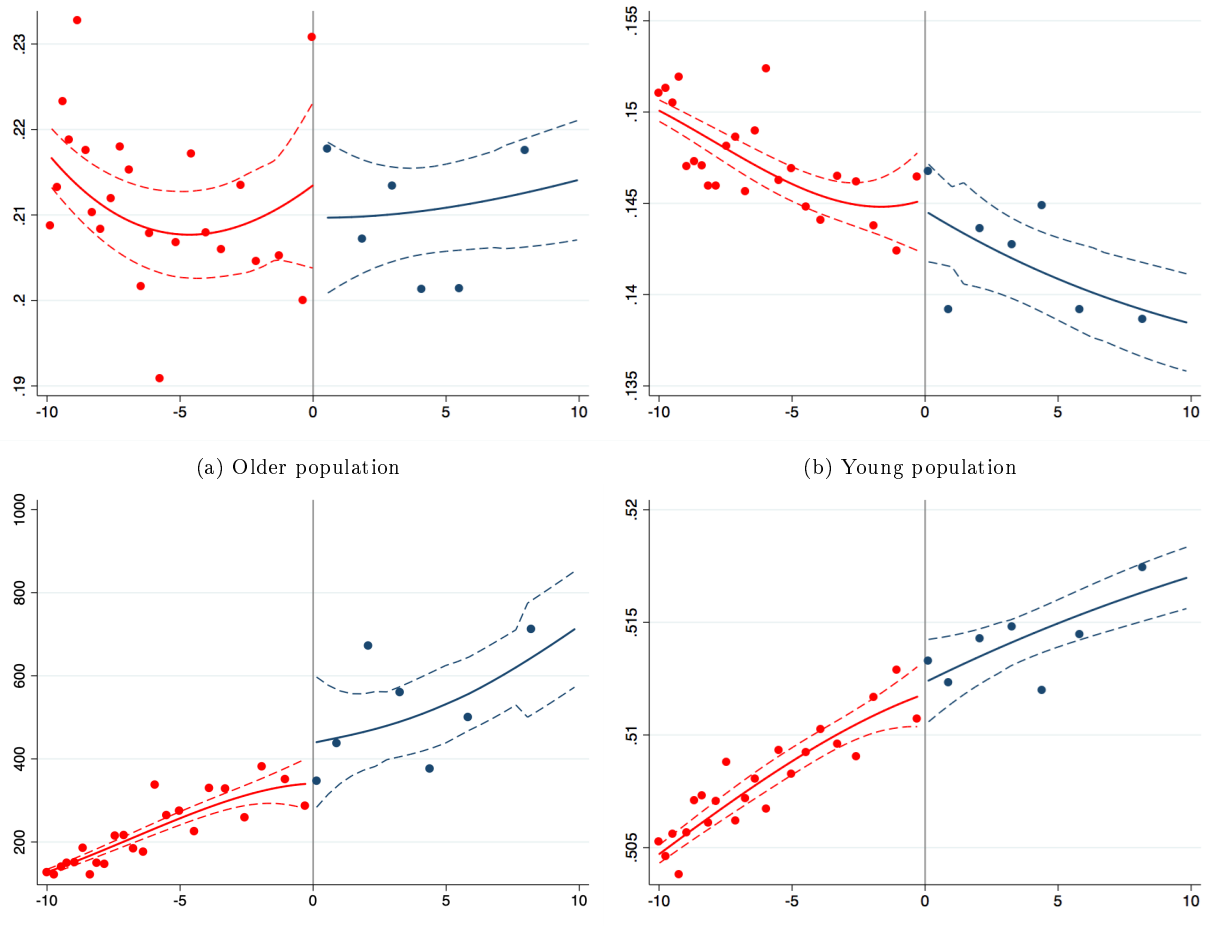
Appendix A. Continuity of covariates

Figure A.9: Continuity of the covariates



(a) Older population

(b) Young population

(c) Population density

(d) Percentage of females

Appendix B. Estimations at the individual level

Appendix B.1. Years 2002 to 2010

At the individual level, we need to estimate the binary choice of whether a female candidate is elected or not, or respectively, whether a elected candidate is a female. In this case, the dependent variable is binary, and we can apply an RD binary outcome model. For a categorical outcome model, which *a fortiori* can be applied to the binary outcome model, we have for an individual i and the outcome $\tilde{Y}_i$, which can take a value that belongs to $(J+1)$ mutually exclusive categories. In a (sharp) RD design, the binary treatment T_i is driven by a continuous variable $X_i \in R$ and a cutoff c, that is $T_i = \mathbf{1}(X_i \geq c)$. Let $\tilde{Y}_i(1)$ and $\tilde{Y}_i(0)$ be the potential outcomes for the treated $(T_i = 1)$ and untreated $(T_i = 0)$ groups, respectively. For j, the conditional outcome probabilities for the two groups are

$$\begin{aligned} \mathbb{P}(\tilde{Y}_i(1) = j | X_i = x) &= \mu_{+,j}(x) \\ \mathbb{P}(\tilde{Y}_i(0) = j | X_i = x) &= \mu_{-,j}(x), \end{aligned} \tag{B.1}$$

where the continuous functions $\mu(\cdot)$ are unknown. The average treatment effect τ can be defined as

$$\begin{aligned} \tau_j =& \mathbb{P}(\tilde{Y}_i(1) = j | X_i = c) - \mathbb{P}(\tilde{Y}_i(0) = j | X_i = c) \\ & \mu_{+,j}(c) - \mu_{-,j}(c) \\ & \lim_{x \to c+} \mathbb{P}(\tilde{Y}_i = j | X_i = x) - \lim_{x \to c-} \mathbb{P}(\tilde{Y}_i = j | X_i = x). \end{aligned} \tag{B.2}$$

Estimation of $\hat{\tau}_j$ can be performed using the standard non-parametric approach of Calonico et al. (2014), although as mentioned by Xu (2017) the bandwidth selecting procedure by Imbens and Kalyanaraman (2012) which is developed for the local linear estimator becomes suboptimal (although still with the optimal rate) for the local nonlinear estimator of the probability function.

For the binary outcome models, we thus additionally use the procedures proposed by Xu (2017), the results of which can be obtained upon request. In general, in the CCT specifications we apply the triangular kernel, as it leads to optimal variance and bias properties. While the choice of kernel typically should not affect the results too much, in those cases, in which estimation with linear kernel point to different conclusions, we shall report both.

According to Table B.19, the probability of a female candidate being elected is about six percentage points higher in the majoritarian system, or eight percentage points if we consider the second-order or a third-order polynomial. All results are significant at the 1% level. Similarly, the probability of encountering a female among elected candidates is also higher by about three to seven percentage points in the majoritarian system, significant at the 5% level.

Table B.19: Probability that a female candidate gets elected: $P(elected|female)$, linear (polynomial) probability model, Years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.066***	-0.067***	-0.085***
	[0.007]	[0.009]	[0.015]
Robust 95% CI	[-.084 ; -.049]	[-.088 ; -.047]	[-.122 ; -.056]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	18617	31589	15165
Eff. Observations R	21879	25897	20396
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.554	10.336	6.625
BW Bias (b)	11.181	13.195	8.640

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Appendix B.2. Year 2014

Tables B.21 and B.22 present the results of the regression discontinuity estimations for the year 2014 at the individual level. In both cases, we do not observe any discontinuity at the 20,000 threshold. Only in specification (4) in Table B.21 is the result significant. However, given all other results this seems to be a rogue result.

Table B.20: Probability of encountering a female among elected: $P(female|elected)$, linear (polynomial) probability model, Years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.034**	-0.032	-0.071***
	[0.012]	[0.017]	[0.024]
Robust 95% CI	[-.064 ; -.006]	[-.068 ; .007]	[-.13 ;-.028]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	14969	19198	12974
Eff. Observations R	8232	8652	7875
Conventional p-value	0.006	0.054	0.004
Robust p-value	0.017	0.108	0.002
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.891	9.108	7.245
BW Bias (b)	11.525	11.591	9.847

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table B.21: Probability that a female candidate gets elected: $P(elected|female)$, linear (polynomial) probability model, Year 2014

	(1)	(2)	(3)
RD Estimate	0.069	0.065	0.089
	[0.042]	[0.048]	[0.062]
Robust 95% CI	[-.029 , .169]	[-.046 , .166]	[-.031 , .233]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	989	1371	1320
Eff. Observations R	488	927	927
Conventional p-value	0.101	0.179	0.152
Robust p-value	0.165	0.266	0.134
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	1.297	2.020	1.930
BW Bias (b)	1.797	2.674	2.561

Note: Regression discontinuity results at the threshold of 50,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns 1 to 3 correspond to polynomials 1 to 3; column 4 corresponds to polynomial 1 and bandwidth selection of Xu (2017).

Table B.22: Probability of encountering a female among elected: $P(female|elected)$, linear (polynomial) probability model, Year 2014

	(1)	(2)	(3)
RD Estimate	0.024	0.026	0.053
	[0.026]	[0.033]	[0.045]
Robust 95% CI	[-.040 ; .085]	[-.047 ; .104]	[-.037 ; .156]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	3985	6523	5217
Eff. Observations R	2562	2940	2835
Conventional p-value	0.355	0.428	0.237
Robust p-value	0.484	0.456	0.225
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	6.710	9.049	8.018
BW Bias (b)	10.085	11.224	10.129

Note: Regression discontinuity results at the threshold of 50,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns 1 to 3 correspond to polynomials 1 to 3; column 4 corresponds to polynomial 1 and bandwidth selection of Xu (2017).

Appendix C. Additional Tables and Figures

Table C.23: Percentage of females in the council at the placebo 50,000 threshold

	(1)	(2)	(3)
RD Estimate	0.003	0.026	0.048
	[0.090]	[0.121]	[0.153]
Robust 95% CI	[-.202 ; .212]	[-.228 ; .325]	[-.283 ; .394]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	27	42	51
Eff. Observations R	17	21	25
Conventional p-value	0.974	0.829	0.754
Robust p-value	0.961	0.730	0.748
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	4.145	5.265	6.171
BW Bias (b)	6.796	6.831	7.372

Note: Regression discontinuity results at the threshold of 50,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.24: Elected among females in a municipality at the placebo 50,000 threshold

	(1)	(2)	(3)
RD Estimate	0.002	0.019	0.039
	[0.041]	[0.042]	[0.067]
Robust 95% CI	[-.073 ; .102]	[-.06 ; .116]	[-.106 ; .185]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	22	31	26
Eff. Observations R	14	18	17
Conventional p-value	0.960	0.658	0.554
Robust p-value	0.748	0.531	0.592
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	3.003	4.362	3.983
BW Bias (b)	4.792	6.503	6.063

Note: Regression discontinuity results at the threshold of 50,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.25: Percentage of females in the council, years 2002 to 2010, log-logit transformation

	(1)	(2)	(3)
RD Estimate	-0.269**	-0.289**	-0.383**
	[0.099]	[0.125]	[0.167]
Robust 95% CI	[-.517 ; -.055]	[-.583 ; -.03]	[-.781 ; -.064]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	868	1427	1185
Eff. Observations R	373	419	402
Conventional p-value	0.007	0.021	0.022
Robust p-value	0.015	0.030	0.021
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.461	9.792	8.909
BW Bias (b)	11.059	12.734	11.306

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.26: Elected among females, years 2002 to 2010, log-logit transformation

	(1)	(2)	(3)
RD Estimate	-0.764***	-0.773***	-0.844***
	[0.106]	[0.134]	[0.186]
Robust 95% CI	[-1.021 ; -.513]	[-1.097 ; -.496]	[-1.267 ; -.467]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	1053	1985	1424
Eff. Observations R	394	452	419
Conventional p-value	0.000	0.000	0.000
Robust p-value	0.000	0.000	0.000
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	8.362	11.269	9.800
BW Bias (b)	11.927	14.428	11.837

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.27: Females per seat - years 2002 to 2010

	(1)	(2)	(3)
RD Estimate	-0.035**	-0.032	-0.064**
	[0.014]	[0.019]	[0.029]
Robust 95% CI	[-.071 ; -.003]	[-.075 ; .013]	[-.135 ; -.013]
Kernel Type	Triangular	Triangular	Triangular
BW Type			
Eff. Observations L	1128	1488	992
Eff. Observations R	408	426	392
Conventional p-value	0.015	0.096	0.024
Robust p-value	0.035	0.168	0.017
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	8.473	9.773	7.869
BW Bias (b)	12.048	11.937	10.289

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.28: Percentage of females in the council at the 40,000 threshold

	(1)	(2)	(3)
RD Estimate	0.020	0.027	0.025
	[0.059]	[0.069]	[0.076]
Robust 95% CI	[-.118 ; .161]	[-.123 ; .176]	[-.139 ; .18]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	40	88	189
Eff. Observations R	24	32	41
Conventional p-value	0.738	0.698	0.747
Robust p-value	0.765	0.729	0.799
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	7.609	13.051	18.741
BW Bias (b)	12.953	20.675	28.731

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.29: Percentage of elected among females at the 40,000 threshold

	(1)	(2)	(3)
RD Estimate	0.048	0.030	0.043
	[0.037]	[0.047]	[0.052]
Robust 95% CI	[-.035 ; .134]	[-.07 ; .128]	[-.068 ; .151]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	54	72	146
Eff. Observations R	26	30	36
Conventional p-value	0.201	0.521	0.411
Robust p-value	0.249	0.567	0.457
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	9.389	11.573	16.918
BW Bias (b)	17.959	21.252	26.558

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.30: Probability that a female candidate gets elected: $P(elected|female)$ at the 40,000 threshold, linear (polynomial) probability model

	(1)	(2)	(3)
RD Estimate	0.025	0.023	0.017
	[0.055]	[0.073]	[0.084]
Robust 95% CI	[-.103 ; .165]	[-.141 ; .181]	[-.161 ; .195]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	1070	1553	2967
Eff. Observations R	525	661	776
Conventional p-value	0.647	0.751	0.836
Robust p-value	0.649	0.808	0.852
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	9.058	12.285	16.594
BW Bias (b)	14.295	17.659	23.350

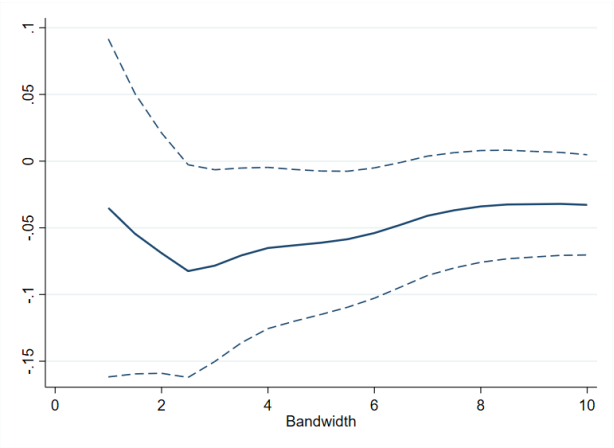
Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Table C.31: Probability of encountering a female among elected: $P(female|elected)$ at the 40,000 threshold, linear (polynomial) probability model

	(1)	(2)	(3)
RD Estimate	0.036	0.051	0.050
	[0.032]	[0.052]	[0.087]
Robust 95% CI	[-.04 ; .125]	[-.069 ; .168]	[-.138 ; .231]
Kernel Type	Triangular	Triangular	Triangular
BW Type	CCT	CCT	CCT
Eff. Observations L	2643	3875	2878
Eff. Observations R	1046	1151	1046
Conventional p-value	0.263	0.326	0.564
Robust p-value	0.316	0.413	0.622
Order Loc. Poly. (p)	1	2	3
Order Bias (q)	2	3	4
BW Loc. Poly. (h)	12.911	15.836	13.134
BW Bias (b)	19.027	22.459	18.052

Note: Regression discontinuity results at the threshold of 20,000 inhabitants. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in brackets. CCT corresponds to the bandwidth selection procedure of Calonico et al. (2014). Columns correspond to polynomials 1 to 3.

Figure C.10: Sensitivity with respect to bandwidth: percentage of females in the council



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