Let the Voters Choose Women*

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Abstract

We study the effectiveness of a novel measure to reduce gender gaps in political empowerment: double preference voting conditioned on gender, coupled with gender quotas on candidate lists. This policy was introduced in 2012 in Italian municipal elections. Using a regression discontinuity design, we find that the share of female councilors rises by 18 percentage points. The result is mainly driven by an increase in preference votes cast for female candidates, suggesting a salient role of double preference voting. We also detect changes in voters' behavior in casting preferences in higher level elections, suggesting the presence of spill-over effects of the double preference voting policy.
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1 Introduction

Gender gaps dominate the political arena. According to the Global Gender Gap Index (World Economic Forum, 2018), the world has closed only 22% of the gender gap in politics. In Europe, women represent 30% of politicians in legislative bodies and 29.5% in government cabinets (EIGE, 2018). In Italy, women represent approximately 35.7% of members of Parliament.

How to promote female political empowerment? This paper examines a new policy in Italy, which in municipal council elections introduces double preference voting conditioned on gender, whereby voters can express two preferences, instead of one, if they vote for candidates of different gender. In addition, the policy foresees gender quotas on candidate lists to municipal councils to guarantee a substantial presence of female candidates. The law targets all Italian municipalities with more than 5,000 residents, and we use it to implement a regression discontinuity design around this threshold. We first estimate that the policy introduced by Law 215/2012 leads to a 18 percentage point increase in the share of elected female politicians. We show that the result is robust to a number of specification changes and does not depend on pre-existing differences between municipalities below and above the 5,000 resident cut-off. To investigate the mechanisms behind the working of the policy, we hand-collect new data on candidate lists and preference votes, and find that the latter play an important role in promoting female political empowerment. We also analyze voters' behavior in casting preference votes in higher level elections to study the potential presence of spill-over effects of the policy. We find some evidence of more preference votes cast for women in regional elections. This suggests that even soft policy measures, such as double preference voting, may have effects beyond their direct target.

Female under-representation in politics may result from various obstacles in a multi-step ladder process of political recruitment (Norris and Lovenduski, 1995). First, women may not be willing to or may not be interested in competing for political seats, for instance due to time constraints associated with child care duties (e.g., Schlozman et al., 1994). Alternatively, lack of self-confidence or external encouragement (Fox and Lawless, 2004) or lower returns on the political market for women (Júlio and Tavares, 2017) may motivate their absence from politics. Second, parties, in their role of gatekeepers, may not put women forward as candidates (e.g., Kunovich and Paxton, 2005). Third, voters may be biased against female candidates and not cast votes for them (e.g., Schwindt-Bayer et al., 2010; Black and Erickson, 2004).

The promotion of female participation in politics is justified on the grounds of equity considerations (Stevens, 2007), since women represent 50% of the overall voting population. Moreover, female politicians appear less corrupt and show higher cooperation and team working skills (Epstein et al., 2005; Brollo and Troiano, 2016). Female participation in politics may also create role models for other women, who may decide to pursue a political career (Gilardi, 2015). In addition, a gender-balanced political body may impact public policy and the allocation of resources across different programs, as documented in, e.g., Chattopadhyay and Duflo (2004), Duflo and Topalova (2004), Beaman et al. (2010), Funk and Gathmann (2015).[1]

In this work we study the introduction of double preference voting, coupled with gender quotas, as a new tool to increase female presence in political institutions. The novelty of this policy measure is that it concentrates on voters' preferences, in addition to the more common gender quota requirement on candidate lists. Preference votes allow voters to select one candidate (or more) on a list in proportional representation systems and they were introduced in a number of countries[2] in past decades. Preference votes are argued to create a direct link between voters and candidates and raise accountability, due to a "threat" that politicians in top list positions are surpassed by candidates below them. In addition, parties may use preference votes cast for candidates in open list systems to test the popularity of politicians and then promote them to more powerful positions (Folke et al., 2016). However, preference votes appear to be highly ineffective, as voters continue to cast their preferences for candidates at the top of the list (Farrell, 2001; Gallagher and Mitchell, 2005). There is evidence of general voters' predisposition to vote for male over female candidates or viceversa, which is often context-specific (Sanbonmatsu, 2002; Black and Erickson, 2003; Schwindt-Bayer et al., 2010).[3] Conditioning double preference voting on gender thus may be a promising way to raise effectiveness of preference votes in promoting female political representation. Up to date, there is no causal evidence on the effectiveness of policies targeting voters' preferences in achieving stronger female political empowerment.[4]

[1] Yet, this result varies considerably across different contexts. For examples, Clots-Figueras (2011), Gagliarducci and Paserman (2012), Ferreira and Gyourko (2014) and Bagues and Campa (2018) find that gender has no (or limited) effect on policies.

[2] Austria, Belgium, the Czech Republic, Denmark, Estonia, Lithuania, Norway, the Netherlands, Slovakia, and Sweden. Since 2013, in French subnational elections voters can elect two members of the opposite sex on a "binôme"or tandem ballot, whose names are arranged in alphabetical order. This new system of nomination of both female and male candidates ("binôme") guarantees the achievement of parity in departmental councils.

[3] The fact that women do not necessarily vote for other women is in line with evidence from other contexts outside politics. For example, in academics, Bagues et al. (2017) find that the presence of women in selection committees does not lead to more female professors being promoted.

[4] In terms of descriptive analysis, Kunovich (2012) shows that in the Polish open-list system, preference votes cast by the electorate shift females higher up in the post-election ranking, compared with the original one proposed by the party, and that these shifts result in a higher number of elected women. Shair-Rosenfield and Hinojosa (2014) show evidence from Chile which is consistent with a negative gender (female) bias among parties, but not among voters.

Gender quotas are the most common policy for tackling gender imbalances and are in place in a few countries, either at the national or the subnational level (Krook, 2009). They are often accompanied by additional measures to further support female political representation, such as zipping, i.e. a man and a woman alternate in the list of candidates, placement mandates (Schmidt, 2009; Schwindt-Bayer, 2009) or list-proportional representation systems (Tripp and Kang, 2008). However, their effectiveness is under scrutiny (see Dahlerup and Freidenvall, 2008 for a discussion). There is evidence from Italy showing that gender quotas on candidate lists increase the share of female municipal councilors (De Paola et al., 2010) and voters' turnout (De Paola et al., 2014); they also promote the election of younger politicians (Baltrunaite et al., 2015) and improve the quality of municipal councilors (Baltrunaite et al., 2014). The positive effect on quality is documented also for Sweden by Besley et al. (2017), who show that quotas do not stand at odds with meritocracy, as they raise male politicians' competence precisely where effects on female representation are the largest. However, Bagues and Esteve-Volart (2012) study the case of the Spanish senate and find that women remain "pawns" in the political game. Similarly, Bagues and Campa (2018) and Casas-Arce and Saiz (2015) show that women's access to political institutions can be challenged by the strategic positioning of female candidates on male-dominated party lists.

Our paper contributes to the existing literature and to the policy debate on how to promote the presence of women in politics. Against the background of mixed evidence on the effectiveness of gender quotas, our results show that paying attention to voters, and not only to parties, may have immediate and sizable effects on female political empowerment. In local councils elected in municipalities with double preference voting conditioned on gender, the women to men ratio rises to 40/60, as compared to a ratio below 30/70 in municipalities not subject to the policy.

The effectiveness of double preference voting can be explained by the presence of a limited voters' bias against female candidates. With single preference voting, voters are more likely to cast their (single) preference vote in favor of a male candidate. Thanks to the expanded set of voters' choices due to double preference voting, also female candidates have a chance of getting a preference vote, provided that voters are not fully biased against women. In fact, the higher number of women elected suggests that some female candidates are ranked close enough to the voters' favorite male candidates. In addition, the effectiveness of the new system may also be consistent with the presence of voters who, irrespectively of their gender preferences, derive extra utility from having more, rather than less, choice.

Our findings suggest that a simple change in the rules of the voting game may affect voters' behavior in the direction of more gender balanced political representation. They are also consistent with the idea that the underrepresentation of women in politics is not purely an artifact of intrinsic gender biases of voters, but it is at least in part institution-driven, and thus modifiable.

The paper is organized as follows: Section 2 presents the institutional setting and the details of Law 215/2012, Section 3 studies the impact of the policy on female politicians and Section 4 explores the mechanisms behind the effects of the reform and the potential spill-overs of the policy in higher level elections. Section 5 concludes.

2 The institutional framework and data

2.1 Law 215/2012

The sub-national levels of government in Italy include regions, provinces and municipalities. There are 20 regions,[5] 97 provinces and approximately 8,100 municipalities. Electoral rules are set independently at each level of government and Law 215/2012 applies at municipal level.

Italian municipalities vary in terms of geographic, demographic and economic indicators. The municipal administration manages the registry of births and deaths, the registry of deeds, and decides over the level and allocation of local expenditure to different goals, such as administration, education and social services. Expenditure is financed via own taxes and tariffs and via transfers from the central government. Municipalities are headed by a mayor, who is assisted by a legislative body –the municipal council (*Consiglio Comunale*)–, and an executive body –the executive committee (*Giunta Comunale*). Municipal elections take place every five years and municipal governments cannot affect their schedule.

The electoral rules at municipal level change at the 15,000 resident threshold. In order to keep the electoral institutions constant, and considering that the law we are interested in applies at the cut-off of 5,000 residents, we focus on municipalities with less than 15,000 residents. In these municipalities, a mayor is elected according to a single-ballot system.[6] The mayoral candidate who gets the relative majority is appointed. Under this scheme, each candidate for the mayor position can be backed

[5]They are, in alphabetical order: Abruzzo, Basilicata, Calabria, Campania, Emilia-Romagna, Friuli-Venezia Giulia, Lazio, Liguria, Lombardia, Marche, Molise, Piemonte, Puglia, Sardegna, Sicilia, Toscana, Trentino-Alto Adige, Umbria, Valle d'Aosta, Veneto. 5 of these regions (Sicilia, Sardegna, Trentino-Alto Adige, Valle d'Aosta and Friuli-Venezia Giulia) have special autonomy (*Regioni a Statuto Speciale*).

[6]In municipalities above the 15,000 resident threshold the mayor is elected according to the run-off system.

by one list only, with a substantial victory bonus: the list supporting the winner gets 2/3 of the seats in the municipal council, while the rest of the seats is assigned to the remaining lists according to a proportionality criterion. Candidate lists to municipal councils are formed by the local organization of a given party or by independently organized groups of citizens. The list consists of at most as many candidates as the number of seats in the council and at least as many candidates as 3/4 of the number of seats. The number of seats in municipal councils varies between 6 and 16, depending on the size of the resident population. The electoral system prescribes semi-open lists, whereby voters vote for a party and can also cast a preference vote for an individual candidate from their preferred list, by writing down a candidate name on the ballot. After the election, for each party, candidates are re-ranked according to the number of preference votes they receive. The number of seats each party wins determines the number of candidates who get elected according to this ranking.

Italian Law 215 was passed in 2012 with the aim of increasing women's presence on municipal councils. The measures prescribed by the law apply to municipalities with more than 5,000 residents. The law introduces double preference voting conditioned on gender, which gives the voters the following options: they vote for a list by crossing the related symbol, and may choose, among candidates of that list, one candidate of any gender for whom to express one preference vote, or two candidates of different gender for whom to express one preference vote each. Voters may also express no preference vote for any specific candidate. More specifically, for each party, the ballot displays two empty lines, rather than one, to write down up to two names of candidates of different gender.[7] To ensure the presence of candidates of both sexes, the law also establishes that neither gender can represent more than 2/3 of the total number of candidates on party lists for municipal councils. In practice, parties have to reserve at least 1/3 of the total number of positions for female candidates. In municipalities with resident population between 5,000 and 15,000, non-compliance is punished by removing the names of male candidates exceeding 2/3 of the total. The law was in force for the first time in the municipal elections in 2013.

[7]When a voter expresses a preference for a candidate, the candidate gains one preference vote. When a voter expresses preferences for two candidates, both candidates get one preference vote if they are of different gender; if the two candidates are of the same gender, only the candidate whose name is written in the first line gains one preference vote, and the other gets zero preference votes. When a voter does not express any preference, no preference votes are assigned to any candidate. Note that the vote expressed for a party holds independently of the expression of preference vote.

2.2 Data

We collect data on elected politicians, candidate lists and preference votes cast. More precisely, we gather publicly available data on elected politicians in the 4,599 Italian municipalities with less than 15,000 residents, which voted in 2013, 2014 and 2015,[8] and the corresponding previous elections. Since municipal elections take place every five years, the previous elections span the period 2008-2010.[9] For these municipalities we use information on the total number and identity of elected councilors, the number of female elected councilors, and the political orientation of the majority party. In addition, we collect information on the number of registered and effective voters, overall and by gender, as well as the number of invalid votes, for the elections taking place in 2013-2015, and the corresponding previous election.[10]

Table 1, Panel A shows that 3,628 municipalities are below the threshold of 5,000 residents and are not subject to the provisions of Law 215/2012; 971 municipalities are above this threshold and must therefore comply with the law. We refer to the former group of municipalities as control, and to the latter as treated. In terms of geographical distribution, both treated and control municipalities are spread all over the country. Table 1, Panel B shows the share of elected female councilors in treated and control municipalities, and provides descriptive evidence suggesting that the reform leads to a higher presence of female councilors in municipalities subject to it: in these municipalities, municipal councils are more gender balanced, with women representing between 39% (in 2013) and 42% (in 2015) of the total number of councilors, against corresponding values of 22% and 27% in municipalities which were not subject to the law.

We also gather information on a large number of observable municipal characteristics, which we use to test the validity of the regression discontinuity design in Section 3.1. From the 2011 Italian Census we collect information on gender and age composition and density of the resident population, shares of males and females with upper secondary education or higher, and share of employed males and females. We also use geographical indicators provided by the National Institute for Geophysics and Volcanology (*Istituto Nazionale di Geofisica e Vulcanologia*), such as the municipality geographical location in different macro areas of the country, surface area in square

[8]Regions with special autonomy, with the exception of Sardinia, do not apply Law 215/2012. Therefore, we exclude municipalities in these regions (i.e., Sicily, Valle d'Aosta, Friuli-Venezia Giulia and Trentino-Alto Adige) from our sample.

[9] We note that municipal councils may terminate their mandate earlier due to factors such as the unexpected death of the mayor or the resignation of the majority of the councilors and therefore there are some municipalities that vote in intervals shorter than five years.

[10]The data are provided by the Ministry of Interior.

kilometers, gradient calculated as the difference between maximum and minimum altitude over the surface, degree of seismicity on a scale 0-4 and mountain area on a scale 1-3. We also use the information provided by provincial Chambers of Commerce (*Infocamere*) to compute the number of limited liability firms in a given municipality, as well as indicators based on tax records and compiled by the Ministry of Economy and Finance on average municipal income and share of taxpayers.

In order to better understand how the policy works, we collect data on candidate lists. These data are difficult to obtain, as they are gathered only by local electoral offices and they are not published by the Ministry of Interior or made available on the Internet. To the best of our knowledge, this is the first time these data are systematically collected. We restrict our attention to municipalities which voted in 2013. We contact all electoral offices of these municipalities in order to request candidate lists presented by every party with the original (party-composed) candidate ordering and the number of preference votes each candidate on the lists obtained, for the 2013 election and for the previous one.[11] Table 2 summarizes the sample coverage in terms of number of municipalities and party lists in the 2013, and in the previous election, happening for most of these municipalities in 2008.

[Tables 1 and 2 here]

To study the existence of broader effects of the policy, we complement the dataset with information on female candidates' performance in higher level regional elections. In particular, we collect preference votes cast in municipalities in our sample for candidates in regional elections. Regional elections are ruled by regional electoral laws, which vary across regions.[12] Some regional electoral laws prescribe double preference voting. We consider all regional elections taking place after the introduction of Law 215/2012, excluding regions which adopt double preference voting. The resulting sample consists of municipalities voting in regional elections in Basilicata (2013), Calabria

[11]If there was no response, we searched for candidate lists published in local newspapers, or directly contacted members of the municipal council or local politicians. On several occasions, the lists could only be obtained by watching parties' electoral campaign video material. We have verified that there are no statistically significant differences in the observable characteristics between municipalities for which we were able to obtain candidate lists for the election with the policy and the previous one, and those for which we were not.

[12]The element in common is that members of the regional assemblies are elected according to a proportional system combined with a majority premium. 4/5 of the members of the assembly are elected from lists formed at the province level, with a proportionality rule, allowing for preference votes. All the municipalities in a given province face the same candidate list by party. The remaining seats are assigned according to the majoritarian system, with regional closed candidate lists.

and Piemonte (2014), Liguria, Marche, Puglia and Veneto (2015).[13] The sample consists of 1,930 municipalities, of which 1,582 are in the control and 348 in the treated group.[14]

3 The impact of the policy on female politicians

In this section we investigate the effects of double preference voting conditioned on gender and gender quotas on the election of women to municipal councils.

3.1 Empirical strategy

We adopt a sharp regression discontinuity design in order to estimate the effect of Law 215/2012 on female presence in local politics. We exploit the fact that the measures included in the law, gender quotas and double preference voting conditioned on gender, only apply to municipalities with more than 5,000 residents. This results in a discontinuous variation in the institutional framework for municipalities of different size along a smoothly increasing forcing variable, namely, municipal population size. Our main regression equation is:

$$\begin{aligned} y_i = \alpha + \gamma_{01}\widetilde{x}_i + \gamma_{02}\widetilde{x}_i^{\,2} + \cdots + \gamma_{0p}\widetilde{x}_i^{\,p} + \psi Treatment_i + \\ \gamma_{11}\widetilde{x}_i * Treatment_i + \gamma_{12}\widetilde{x}_i^{\,2} * Treatment_i + \cdots + \\ \gamma_{1p}\widetilde{x}_i^{\,p} * Treatment_i + \varepsilon_i \end{aligned} \tag{1}$$

where y_i is the outcome variable of interest, e.g., the share of elected female councilors in municipality i; $\widetilde{x}_i$ is the resident population size in municipality i, centered at the 5,000 resident threshold; p is the order of the control polynomial function, with $p = 1, 2, 3, 4$; and $Treatment_i$ is an indicator for municipalities with more than 5,000 residents ("treated municipalities"). The coefficients on the polynomial terms γ are also indexed by 0 and 1 because we allow for different polynomial coefficients on the two sides of the cut-off. The main coefficient of interest is ψ, which estimates the local average treatment effect of the reform.

We rely on three sets of results:

[13] After the introduction of Law 215/2012, regional elections were also held in Abruzzo and Molise. However, it is not possible to include them in our analysis because the data on preference votes are not reported at municipal level for these two regions.

[14] To assess the absence of pre-existing differences, we also collect data on preference votes cast in municipalities in our sample for female candidates in lists in the previous regional election. For these elections, we have data for Basilicata, Liguria, Piemonte and Veneto, which all voted in 2010, because data on preference votes are not collected at municipal level in the other regions.

1. We graphically investigate the existence of the discontinuity around the 5,000 resident cut-off. For this purpose, we plot local sample means of the dependent variable in small equidistant non-overlapping bins over the support of the resident population size $\widetilde{x}_i$, together with the quadratic polynomial fit for municipalities below and above the threshold, and the 95 per cent confidence interval.

2. We estimate Equation (1) using polynomials of different orders, ranging from 1 to 4, for the entire sample of municipalities (*parametric approach*).

3. We implement local linear regressions using the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) (*non-parametric approach*).

While these different specifications serve the purpose of transparently showing the robustness of the results, we will focus on the estimates from local linear regressions when commenting on the magnitudes of the effects.

For the validity of the regression discontinuity, we first verify that there are no discontinuities at the 5,000 resident threshold in the distribution of demographic (male and female shares, children and elderly share, population density), geographical (a dummy indicator North for the geographical location, surface in squared Kilometers, gradient, degree of seismicity on a scale 0-4, mountain area on a scale 1-3), educational (share of males and femalse with upper secondary education or higher) and economic (shares of employed females and males, average income, share of taxpayers, and the number of firms) characteristics for municipalities voting in 2013-2015. The results of the graphical analysis in Figure 1 and those of the local linear regressions in Table 3 show that municipal characteristics vary continuously with municipal population size.

We then test the potential presence of sorting, i.e. the tendency of municipalities to strategically manipulate their population to fall on the preferred side of the cut-off. We implement a McCrary test (McCrary, 2008) and find no evidence of manipulation of the population size in the sample of Italian municipalities which voted in the period 2013-2015, as shown in Figure 2.

[Figures 1, 2 and Table 3 here]

3.2 Results

We examine the share of elected female councilors (i.e. the number of elected female councilors over the total number of councilors) around the 5,000 resident threshold.

Figure 3 shows a discontinuous jump in the share of elected female councilors in the municipalities above the cut-off, which were subject to the policy.[15]

[Figure 3 here]

We next estimate the magnitude of the change in the share of female councilors using the control polynomial (parametric) approach. Specifically, we use observations both close to and far from the cut-off point and estimate equation (1) with polynomials of orders 1 to 4 in the four Columns of Table 4, Panel A. Polynomials are allowed to differ on the two sides of the cut-off. The results show that the estimated coefficient on the indicator *Treatment* is positive and remains statistically significant in all Columns.

[Table 4 here]

To test the existence of a discontinuity in the share of elected female councilors non-parametrically, we implement local linear regressions using a triangular kernel density estimator. In Table 4, Panel B, conventional estimates with conventional standard errors are presented in row 1. The results are consistent with the coefficients presented in Panel A. Moreover, the point estimate increases as we concentrate on observations closer to the 5,000 resident threshold. We also show biased-corrected estimates with conventional standard errors, and biased-corrected estimates with robust standard errors in rows 2 and 3 in Table 4, Panel B. The point estimate of the coefficient on the variable *Treatment* is 0.183 in these last specifications and implies that municipalities that voted under the provisions of Law 215/2012 elected municipal councils with 18 percentage points more women. This corresponds to two more women in municipal councils, which is a rather sizable effect. The increase in female elected politicians is confirmed when we conduct the analysis separately in the subsample of municipalities in the North, Centre and South of Italy, which are characterized by a marked divide in female empowerment, as shown in Table A.1.

3.3 Robustness checks

In Table 5 we present robustness checks of non-parametric estimation. In particular, we investigate sensitivity of the estimated parameters to the choice of the bandwidths, as well as to the use of alternative placebo cut-offs in the municipality size. As before, the dependent variable is the share of female councilors over the total number of councilors in elections in 2013-2015 and we report conventional RD estimates with

[15] The discontinuity in the share of female councilors is robust and evident in analogous figures with polynomial fits of orders 1, 3 and 4.

a conventional variance estimator, bias- corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator. In Panel A we consider cut-offs alternative to the 5,000 resident cut-off which determines whether the policy applies. Namely, we consider cut-offs of 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents. The results show that the only significant change in the share of elected female councilors is at the correct 5,000 resident cut-off, while there are no significant changes at alternative placebo cut-offs. In Panel B we consider other bandwidths alternative to the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017). In particular, we consider bandwidths of 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000 residents. Our estimate of the treatment effect is not sensitive to the use of these alternative bandwidths.

[Table 5 here]

As a placebo exercise, we assess whether there are pre-existing differences in the share of female politicians that could confound our estimates of the policy effect. We thus examine the potential discontinuity in the share of female councilors in the previous election. Table 6 and Figure 4 show that the share of female elected politicians does not exhibit any discontinuity at the cut-off in the previous election.[16]

[Table 6 and Figure 4 here]

We also deal with the threats to the interpretation of regression discontinuity design results, coming from "confounding policies" (Eggers et al., 2017). The only relevant confounding policy concerning local electoral outcomes is the legislation which imposes a variation in the salary of the mayor at the same cut-off of 5,000 residents.[17] However, we point out that our analysis focuses on municipal councilors, and not mayors, and compensation of municipal councilors is not regulated by the Italian law. Furthermore, the change in the mayor's salary at the 5,000 resident cut-off precedes the introduction

[16]In the Appendix, we also show that this zero result is not sensitive to the use of alternative cut-offs and bandwidths, as shown in Table A.2.

[17]We note that in 2013 the rules of the Internal Stability Pact, regulating local public finances, vary at the 5,000 resident cut-off, while they are the same in 2014 and 2015. We find that the effects on the share of female councilors are present in 2013, 2014 and 2015, considered separately. This evidence points against the presence of confounding effects stemming from differences in the rules of the Internal Stability Pact. We also note that the size of the municipal council and, hence, the length of the candidate list, change at the cut-off in 2013, whereas they are the same in 2014 and 2015. We take this into account by defining our dependent variables in terms of shares, instead of absolute values. In addition, we point out that, also in this case, the results hold for each year analyzed in isolation.

of Law 215/2012 and it was present long before 2013-2015. As argued above, there are no discontinuities in the share of female councilors (or of female candidates, as will be shown in Section 4.1.) in the previous election, confirming that the observed effects are not driven by differences in the mayor's salary. Finally, we also show that the result on elected female politicians are robust to adopting a difference-in-discontinuities design. Following the specification adopted by Grembi et al. (2016), we estimate the following linear model:

$$y_{it} = \delta_0 + \delta_1 \widetilde{x}_i + Treatment_i(\gamma_0 + \gamma_1 \widetilde{x}_i) + After_t[\alpha_0 + \alpha_1 \widetilde{x}_i + Treatment_i(\beta_0 + \beta \widetilde{x}_i)] + \epsilon_{it} \tag{2}$$

where y_i is the outcome variable of interest, namely the share of elected female councilors in municipality i, $\widetilde{x}_i$ is the resident population size in municipality i, centered on the 5,000 resident threshold, $Treatment_i$ is an indicator for municipalities with more than 5,000 residents ("treated municipalities") and $After_t$ is an indicator equal to 1 for the election with the policy (i.e., in the 2013-2015 period) and 0 for the previous election. We further augment the regression specification with municipality fixed effects to account for time-constant observable and unobservable municipal-level characteristics. The main coefficient of interest is β_0, which estimates the local average treatment effect of the reform. The difference-in-discontinuities analysis hinges on the additional identifying assumption of the presence of parallel trends in the dependent variable prior to the reform. To test it, we complement our data with the information on the share of elected female councilors in municipal elections since 1997. Figure 5 plots the share of female councilors in treated and control municipalities in elections up to 2013 (which corresponds to election 0 on the horizontal axis). It shows that the share of elected female councilors develops in a parallel manner for the two groups of municipalities over the long period before the introduction of the double preference voting policy, validating the use of difference-in-discontinuities design.

Table 7 shows the results of the difference-in-discontinuities estimations. The coefficient of interest is positive, large and statistically significant, confirming that the effect of the reform on women's empowerment holds true even when controlling for the discontinuity in the mayor's salary.

[Figure 5 and Table 7 here]

4 Mechanisms

In this section, we investigate the role of parties and the way they select candidates, as well as the role of voters and their preferences in determining the increase in the presence of female municipal councilors. Moreover, we analyze whether the policy induces broader effects related to voting behavior and political selection.

4.1 The working of the policy

Our purpose is to shed light on how the expanded set of voters' choices interacts with party selection of candidates in fostering female presence in local politics. To this end, we examine the gender composition of candidate lists, which are formed by parties, and preference votes cast for female candidates by the electorate. More specifically, we consider the share of female candidates on the electoral lists composed by political parties, the party-determined ranking of women on the candidate list, the preference votes cast for female candidates and the preference-vote-determined ranking of female candidates. We restrict our attention to the 2013 election, for which we use hand-collected data on these outcomes.

We run party-level regressions as in (1), where the subscript i is replaced by is and all variables are defined for party list s in municipality i.[18] In the interest of space, for this set of outcomes we report results of non-parametric estimations and graphical evidence.[19]

Since for municipalities with more than 5,000 residents the law requires that at least 1/3 of the candidates on each list are female, we start by investigating the behavior of the share of female candidates on party list s in municipality i around the 5,000 threshold. Non-parametric estimates for the share of female candidates are shown in Table 8, Column 1: there is no significant discontinuity at the threshold, indicating that parties do not set the gender composition of the lists differently across the cut-off.[20] This evidence is confirmed by the graphical analysis in Figure 6, Panel A. The pattern is also similar when we look at the election prior to the introduction of the policy. Table 9, Column 1 and Figure 7, Panel A, show that the share of female candidates does not exhibit any discontinuity at the cut-off in the previous election.[21]

[Table 8 and Figure 6 here]

[Table 9 and Figure 7 here]

[18] Civic lists can also run for seats. They are also considered under the wording "party lists".

[19] The results of parametric estimations are in line with the evidence presented in the paper and are available upon request.

[20] Table A.3 shows that this zero result is not sensitive to the use of alternative cut-offs and bandwidths. We replicate this robustness check also for the other dependent variables in Table 8, Columns 2-4, in Tables A.4, A.5 and A.6 respectively, which all confirm the robustness of the results.

[21] Table A.7, Column 1 shows the difference-in-discontinuity estimation, which also confirm the result. We replicate this estimation also for the other dependent variables in Table 8, Columns 2-4, in Table A.7, Columns 2-4, respectively. All of the results appear robust.

The absence of a significant increase in the share of female candidates may appear in contrast with other contributions showing the effectiveness of gender quotas in promoting female presence in politics (for Italy, see De Paola et al., 2010; and Baltrunaite et al., 2014). One potential explanation of the lack of effectiveness of the quota may be the fact that the latter is non-binding, i.e., it imposes a requirement which is equal to or smaller than the existing share of female candidates. Yet, the data do not support this hypothesis, since the share of women on candidate lists stood below the 33% quota requirement imposed by the law in the election before the reform for most municipalities in our sample (Figure 7). Interestingly, our evidence reveals that over a five year period of time, there was an overall increase in the share of women candidates in all municipalities, not only those subject to the reform. This may reflect a general positive trend in female political participation. These contextual differences may help to reconcile the seemingly contrasting evidence on the effectiveness of gender quotas.

Although the share of female candidates does not change at the threshold, the likelihood of being elected may depend on the ranking of candidates, as politicians at the top of the list tend to obtain more preference votes and are therefore more likely to be elected (Farrell, 2001). Some studies (Bagues and Esteve-Volart, 2012; Casas-Arce and Saiz, 2015) show that, when constrained by gender quotas, parties manipulate the ranking of the candidates, placing women at the bottom, so that there is little change in the chances of being elected for male candidates, who usually form the existing party élite. On the contrary, Shair-Rosenfield (2012) shows that parties in India often place women on their lists higher than required by the law. Therefore, we investigate whether parties below and above the 5,000 resident threshold rank male and female candidates differently. If this is the case, the discontinuity we observe in the number of elected females at the cut-off may partially result from party decisions regarding the ranking of candidates.[22] We rely on Borda ranking which attributes a decreasing number of points to each candidate on the list, i.e. in a list with five candidates, the first one gets five points, the second one – four points, etc., and the last one – one point. We define a Borda score of female candidates as the sum of Borda points of female candidates over the total number of Borda points of all candidates on a given list. This measure exploits the information on the full ranking of candidates to detect systematic differences in candidates' placement, across lists of different length. The results of the regression analysis in Table 8, Column 2 show that there is no change at the threshold.[23] Overall, parties do not appear to be strategic in deciding the ranking

[22]We point out that 42% of the lists in our sample are ranked alphabetically and, therefore, are not very likely to exhibit a strategic placement of candidates by parties.

[23]We also consider an alternative measure of candidate placement based on the presence of at least one female candidate on the top two positions of the list. Once more, we do not find a discontinuity

of female candidates under the new constraints imposed by the policy. Interestingly, this is the case also in the previous election, as shown in Table 9, Column 2, and Figure 7, Panel B. All the above evidence suggests that it is unlikely that the gender composition of candidate lists or differences in ranking can solely explain the large increase in the share of female councilors.

We then turn to analyzing preference votes to examine the role of double preference voting conditioned on gender in promoting female politicians. The regression results in Table 8, Column 3 show that the share of preference votes cast for female candidates on lists presented in municipalities subject to the policy increases by 15 percentage points. Figure 6 confirms that there is a visible positive discontinuity at the cut-off. These results provide evidence that the policy was effective in attracting more preference votes for women. This is further confirmed by the analysis of the preference votes in the previous election: Table 9, Column 3 and Figure 7, Panel C show no discontinuities at the 5,000 resident cut-off.

We further investigate how preference votes cast for female candidates affect women's presence on municipal councils. In the Italian semi-open lists system, the original party ranking of candidates is re-ordered according to preference votes cast by the electorate. This post-election ranking determines which candidates are elected and reflects the influence of the voters' decisions on the ultimate electoral outcome. To capture this influence, we calculate the Borda score of female candidates using the post-election ranking of all female candidates (elected and non-elected) and use it as a dependent variable in the analysis. Table 8, Column 4 and Figure 6, Panel D, show that there is a positive discontinuity in this measure at the cut-off. Similar to other outcomes, this is not an artifact of pre-existing differences in outcomes across the two groups of municipalities: Table 9, Column 4 and Figure 7, Panel D show no changes at the threshold in the election prior to the introduction of the policy. Overall, as parties do not compile candidate lists - either in terms of gender composition or ranking - differently across the threshold, the results in this section strongly support that preference votes elicited by the reform have an important role in promoting female presence on municipal councils.

4.2 Other voting outcomes

To strengthen our analysis of the mechanisms behind the effectiveness of the reform, in this section we investigate further outcomes which may influence the increase in the share of female councilors, such as voters' turnout, use of preference votes and the quality of councilors. Results are presented in Table 10 and Figure 8.

at the cut-off. The results are available upon request.

We first ask whether the reform changes voters' turnout –overall and by gender– in municipal elections, which we measure as the share of actual voters (i.e., those who turn out to vote in a given municipality) over eligible voters. Table 10, Columns 1 and 2 and Figure 8, Panels A and B show that there is no discontinuous change in overall voters' turnout and voters' turnout by gender.[24]

[Table 10 and Figure 8 here]

Next, we examine the use of preference votes measured as the ratio between the total number of preference votes cast for candidates of a given list and the number of actual voters for that list.[25] Results shown in Table 10, Column 3 and in Figure 8 Panel C indicate that preference votes are used more actively in treated municipalities. In particular, the figure shows that in municipalities below the threshold, roughly 7 out of 10 voters choose to express a preference. Under the assumption that voters' turnout in expressing preference votes does not change due to the reform, the full adoption of the double preference voting policy would imply 14 preference votes every 10 voters, whereas no adoption of the policy would imply no change in the number of preference votes per voter, i.e. 7 preference votes every 10 voters. In municipalities above the threshold, we observe roughly 9 preference votes every 10 voters. This suggests that preference votes are used more, though their potential is not fully exploited. Note that we find that there is no discontinuity at the cut-off in the number of votes cast for male candidates,[26] thus double preference voting does not subtract preference votes from them.

The increase in preference votes cast for women may come from a change in the selection of politicians, which increases the quality of candidates running for office. We cannot test this effect directly, because data on the personal characteristics of candidates are not collected. Hence, we study the quality of the elected councilors, as measured by the average years of education (Galasso and Nannicini, 2011; Baltrunaite et al., 2014).[27] The following possibilities can arise. If the quality of both male and female candidates increases, the higher number of preferences for female candidates at the threshold cannot be explained by changes in quality. If only the quality of female candidates increases, we should expect that better-quality women obtain more

[24] In addition, we find no evidence that voters are "confused" by this policy: the number of invalid ballots is not significantly different at the cut-off. Results are available upon request.

[25] We rely on this measure because electoral data do not register whether a voter has expressed 0, 1, or 2 preferences. We also point out that the number of actual voters - used as the denominator of this ratio - is continuous across the cut-off.

[26] The results are available upon request.

[27] Note that not only the researcher but also the voters are not systematically provided with information on the level of education or job held by candidates to the municipal council.

preference votes, independently of the double preference voting mechanism, and are hence elected. However, we do not find any significant discontinuity at the cut-off in the quality of elected female councilors, as shown in Table 10, Column 4 and Figure 8, Panel D.[28] Finally, if only the quality of male candidates increases, we should expect an increase in the number of votes cast for male candidates, which we do not observe, as argued above. Therefore, changes in the selection of politicians do not appear to be consistent with the observed patterns in the data.[29]

To further investigate effects related to the quality of elected councilors, in Table 11 we perform a difference-in-discontinuities estimation, using as outcomes the difference between male and female average years of education and the average years of education of men and women, separately. There is no evidence of significant changes, except for a marginally significant increase in educational attainment for men, which is however not confirmed in the local estimation.

To ensure that our results in Table 10 – overall and female voters' turnout, use of preference votes and quality of politicians – do not depend on pre-determined conditions, we consider the four outcomes in the election prior to 2013 and show that no discontinuity arises both in the non-parametric estimation in Table 12 and in the graphical analysis in Figure 9.

We also look at the political orientation of the majority party elected in municipalities below and above the 5,000 resident threshold. Interestingly, in most municipalities that held elections in the period 2013-2015 (4,195 out of 4,599) civic lists obtained the majority of seats and the shares of municipalities with a civic list, left-wing, center-left and right-wing majority are smooth around the 5,000 resident threshold.[30]

Having established the key role of preference votes in promoting female political empowerment, we are now interested in understanding whether the positive effects of the reform on female political empowerment last beyond the first election in which voters were given the opportunity to express more than one preference vote. To this end, we consider elections where voters voted under the double preference voting system for

[28]The robustness of the results in Table 10 to the use of alternative cut-offs and bandwidths is assessed in Tables A.8 - A.11. The overall evidence indicates that the results are not systematically sensitive to the choice of the bandwidth. Similarly, there are no spurious relationships between the municipal population and our outcome variables, with an exception of Table A.10 revealing a drop, rather than an increase, in the use of preference votes at some cut-offs. Regardless, the effect documented at the actual 5,000 resident cut-off is of the opposite sign, substantially larger and much neater, as shown in Figure 8, Panel C.

[29]Rather than changes in the selection of politicians, the increase in preference votes can be linked to a change in the behavior of candidates who, in the presence of the policy, increase their effort in political campaigning. If this were the case, we would expect an increase of turnout and/or turnout by gender, which instead is not confirmed by the data (see Table 10, Columns 1 and 2).

[30]All results are available upon request.

a second time, when the implications of the policy are expected to be learnt by parties and voters. We collect data on the share of female councilors for municipalities which voted in 2013 and a second time afterwards, which for most municipalities happens in 2018. Figure A.1 shows that the policy remains effective in promoting female political empowerment and its effects do not die out over time.

Finally, although the policy directly targets elections to the municipal council, one may expect gender salience originated by the policy to affect also other electoral outcomes within the local government. To test this possibility, we first focus on mayors, and check whether the policy results in higher chances of women being elected to the top executive position in municipalities subject to the law. We find no conclusive evidence in favor of the latter hypothesis; if anything – there is a slight decline in the number of female mayors in affected municipalities. Thus, in the short-term, the 2012 policy does not help women to gain easier access to executive positions in municipalities.

In summary, there is evidence that voters do make use of the expanded set of choices guaranteed by double preference voting and that the latter plays an important role in ensuring that more women are elected to municipal councils. Although with single preference voting voters are more likely to choose a male candidate as a single preference, double preference voting also gives female candidates a chance of getting preference votes, provided that voters are not fully biased against women (as in Shair-Rosenfield and Hinojosa, 2014). In fact, the higher number of women elected suggests that female candidates are ranked close enough to the voters' favorite men. Moreover, the effectiveness of the double preference voting policy may be explained by the presence of voters who, irrespectively of their gender preferences, derive extra utility from having more, rather than less, choice.[31] Both interpretations suggest that the underrepresentation of women in politics may be determined, at least in part, by voting rules constraining voters' choices.

4.3 Spill-over effects on regional elections

Given the salient role of double preference voting documented in earlier sections, we next analyze whether this affirmative action measure introduced by the 2012 policy affects outcomes beyond its direct scope of application. This influence may result, for instance, from voters' behavioral reactions due to learning, mimicking or habit

[31] This explanation builds on the theory of expressive voting in explaining voting behavior, according to which voters enjoy benefits from the act of voting itself. These benefits may stem from the possibility to express one's opinion, confirm one's identity and follow moral norms (Hamlin and Jennings, 2018).

formation. In general, this analysis may contribute to the debate on the use of hard versus soft measures to achieve gender balance in political decision making (Dahlerup, 2006).

To study the potential presence of spill-over effects of the policy, we investigate voters' behavior in casting preference votes in higher level elections. We consider all regional elections taking place after the introduction of Law 215/2012, excluding regions which adopt double preference voting in the regional electoral law, as explained in Section 2. We define our dependent variable as the average number of preference votes cast in municipality i for female candidates on a given list in regional elections. This allows us to study voters' behavior in casting preference votes, in isolation from any effect of candidate supply.[32]

Figure 10, Panel A shows a positive discontinuity at the threshold in the average number of preference votes cast for female candidates in municipalities voting in regional elections in the period subsequent to the local election with the double preference voting policy. Such discontinuity is not present in the previous regional election in 2010 (Figure 10, Panel B). To quantify the effect, we perform difference-in-discontinuities analysis and report the results in Table 13. The coefficient is positive, large and statistically significant in Column 1, in which the analysis uses the entire sample of the data. It amounts to roughly three preference votes more, on average, cast for female candidates. The effect is sizable with respect to the sample mean of roughly 4 preference votes for an average candidate on a party list cast in a given municipality. We note, however, that there is no immediate link between a large effect in a given treated municipality and the electoral outcome at the regional level, as the latter is determined by candidates' success in all municipalities within the region. In Table 13, Column 2, we restrict the analysis to a narrower window around the cut-off. The effect is positive, yet smaller, and loses its statistical significance in this specification. We note, however, that this may at least partially be driven by the fact that a very stringent regression specification is used in a substantially reduced sample. Overall, since the policy is very recent, the documented effects may be interpreted as a lower bound in the presence of habit formation regarding women presence in politics.

[Figure 10 and Table 13 here]

[32]In particular, this measure is unaffected by the presence or absence of gender quotas on regional candidate lists.

5 Conclusions

This paper shows that the policy which introduces double preference voting conditioned on gender and guarantees a minimum presence of both genders on candidate lists has a large and robust impact on women's political representation in Italian municipalities. Specifically, our causally identified estimates suggest an increase of 18 percentage points in the share of female councilors. We provide evidence that the effect, to a large extent, comes from preference votes in favor of female candidates expressed by electorate in municipalities subject to the policy. In other words, if voters are given the option of casting a preference vote for one candidate of each gender, they do select female candidates more often.

The design of policies to promote women in politics has so far mostly focused on selection made by parties, prescribing, mainly, gender quotas on candidate lists. However, gender quotas are not always effective, and when they are, the increase in female representation is often of limited size. Our results show that a policy which targets voters, such as double preference voting, leads to stronger effects on female representation and brings the municipal council composition closer to gender equality. In addition, they suggest that even soft policy measures, imposing no obligation on parties or voters, but rather acting through the expansion of the set of choices available to the latter, may spill-over beyond their direct target. This result is particularly encouraging for the evaluation of the effectiveness of affirmative action measures, such as or similar to the one analyzed in this paper.

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Tables and figures

Table 1: Descriptive statistics: municipalities and elected councilors

Panel A: Geographical coverage			
No. of municipalities voting in 2013:	Control	Treated	Total
North	132	65	197
South and islands	153	63	216
Center	34	21	55
Total	319	149	468
No. of municipalities voting in 2014:	Control	Treated	Total
North	2023	493	2,516
South and islands	473	99	572
Center	392	117	509
Total	2,888	709	3,597
No. of municipalities voting in 2015:	Control	Treated	Total
North	94	32	126
South and islands	295	74	369
Center	32	7	39
Total	421	113	534
Panel B: Share of female councilors			
Municipalities voting in 2013:	Control	Treated	Total
	0.22	0.39	0.28
	(0.19)	(0.11)	(0.19)
Municipalities voting in 2014:	Control	Treated	Total
	0.29	0.40	0.31
	(0.14)	(0.10)	(0.14)
Municipalities voting in 2015:	Control	Treated	Total
	0.27	0.42	0.30
	(0.14)	(0.09)	(0.14)

Notes. The table reports descriptive statistics on voting municipalities and share of female councilors. Panel A reports the number of municipalities with less than 15,000 residents that held elections in 2013, 2014 and 2015, distinguishing between treated municipalities (those with at least 5,000 residents) and control municipalities (those with less than 5,000 residents), overall and separately for each geographical area. Panel B reports the means of the share of elected female councilors (with standard errors in parentheses) in municipalities with less than 15,000 residents that held elections in 2013, 2014 and 2015, distinguishing between treated and control group.

Table 2: Descriptive statistics: candidate lists

Panel A: 2013 election			
No. of municipalities:	Control	Treated	Total
voted	319	149	468
with all lists available	276	134	378
with preference votes available	255	126	381
with pre-election ranking available	213	116	329
No. of party lists:	659	475	1,134
with pre-election ranking available	560	444	1,004
with non-alphabetical ranking	302	277	579
Panel B: Previous election			
No. of municipalities:	Control	Treated	Total
voted	319	149	468
with all lists available	178	93	271
with preference votes available	178	93	271
with pre-election ranking available	178	93	271
No. of party lists	437	300	737
with pre-election ranking available	437	300	737
with non-alphabetical ranking	311	230	541

Notes. The table reports sample numerosity for the 2013 municipal election (Panel A) and for the previous one (Panel B), distinguishing between treated and control municipalities. Elections take place every 5 years, thus, in most cases, previous election took place in 2008. See the main text for details. Panel A and B report, for 2013 and the previous election, respectively, the number of municipalities that voted (for which we have data on all elected councilors), the number of municipalities with all party lists available, those with post-election preference votes available, and those with pre-election party-composed ranking of candidates available. It also reports the total number of party lists, the number of party lists with pre-election party-composed ranking of candidates available and, among them, those with non-alphabetical ranking of candidates.

Table 3: Balance checks of covariates

Panel A: Demographic characteristics					
	% Female	% Male	% Children	% Elderly	Population density
Treatment	0.005	−0.004	0.002	−0.004	41.610
	(0.004)	(0.003)	(0.002)	(0.006)	(60.518)
Bias-corrected	0.006	−0.003	0.003	−0.006	53.125
	(0.004)	(0.003)	(0.002)	(0.006)	(60.518)
Treatment (bias-corrected, robust SE)	0.006	−0.003	0.003	−0.006	53.125
	(0.005)	(0.004)	(0.003)	(0.007)	(72.398)
Cut-off	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,249	2,142	1,744	1,256	1,565
Observations on the left	400	782	606	401	520
Observations on the right	242	388	323	243	294
Panel B: Geographical characteristics					
	North	Area	Gradient	Seismicity	Mountain area
Treatment	−0.077	−6.851	4.450	−0.017	−0.032
	(0.072)	(6.607)	(4.133)	(0.170)	(0.112)
Bias-corrected	−0.090	−6.995	5.554	0.020	−0.033
	(0.072)	(6.607)	(4.133)	(0.170)	(0.112)
Treatment (bias-corrected, robust SE)	−0.090	−6.995	5.554	0.020	−0.033
	(0.088)	(8.290)	(4.770)	(0.200)	(0.135)
Cut-off	5,000	5,000	5,000	5,000	5,000
Bandwidth	2,047	1,721	1,435	1,798	2,264
Observations on the left	732	598	473	595	844
Observations on the right	370	320	268	316	404
Panel C: Education characteristics					
	% Female HS+	% Male HS+			
Treatment	−0.003	−0.001			
	(0.003)	(0.003)			
Bias-corrected	−0.004	−0.002			
	(0.003)	(0.003)			
Treatment (bias-corrected, robust SE)	−0.004	−0.002			
	(0.003)	(0.003)			
Cut-off	5,000	5,000			
Bandwidth	1,364	1,574			
Observations on the left	441	520			
Observations on the right	257	294			
Panel D: Economic characteristics					
	% Female employed	% Male employed	Average income	% Taxpayers	# Firms
Treatment	−0.009*	−0.007	−177.530	−0.023	−2.114
	(0.005)	(0.005)	(450.627)	(0.014)	(8.143)
Bias-corrected	−0.010**	−0.008	−224.842	−0.025*	−2.384
	(0.005)	(0.005)	(450.627)	(0.014)	(8.143)
Treatment (bias-corrected, robust SE)	−0.010*	−0.008	−224.842	−0.025	−2.384
	(0.006)	(0.006)	(548.946)	(0.017)	(9.878)
Cut-off	5,000	5,000	5,000	5,000	5,000
Bandwidth	2,021	1,617	1,711	1,980	1,655
Observations on the left	720	549	593	710	554
Observations on the right	365	299	316	358	301

Notes. The table shows the results of non-parametric estimation for demographic, geographic, education and economic characteristics at municipal level, as dependent variables. Panel A reports female, male, children (0-9) and elderly (70+) shares of the population and the population density. Panel B reports a dummy North for geographical location, area in squared Km, gradient, degree of seismicity on a scale 0-4, mountain area on a scale 1-3. Panel C reports the share of female and male population with upper secondary education or higher. Panel D reports the share of employed females and males, average income (total municipality income over the number of taxpayers), share of taxpayers, and the number of limited liability companies. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015, within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017). *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. * p<0.1, ** p<0.05, *** p<0.01.

Table 4: Female presence on municipal councils

	Panel A: Parametric Approach			
Dependent variable:	Share of female councilors			
	(1)	(2)	(3)	(4)
Treatment	0.135***	0.130***	0.151***	0.162***
	(0.009)	(0.014)	(0.018)	(0.023)
Polynomial order	1	2	3	4
Observations	4,599	4,599	4,599	4,599
R-Squared	0.122	0.122	0.123	0.124

	Panel B: Non-parametric Approach
Dependent variable:	Share of female councilors
	(1)
Treatment	0.174***
	(0.021)
Bias-corrected	0.183***
	(0.021)
Treatment (bias-corrected, robust SE)	0.183***
	(0.024)
Bandwidth	1,132
Observations on the left	353
Observations on the right	219

Notes. The table shows the results of parametric and non-parametric estimations. The dependent variable is the share of female councilors over the total number of councilors. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. In Panel A, the sample includes all municipalities with less than 15,000 residents that held elections in the period 2013-2015. Columns 1-4 include polynomials of orders 1-4, respectively, in the resident population, centered on the 5,000 resident threshold. Polynomials are allowed to differ on the two sides of the cut-off. In Panel B, conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. * p<0.1, ** p<0.05, *** p<0.01.

Table 5: Female presence on municipal councils: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Share of female councilors						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.009	−0.009	−0.001	0.174***	−0.005	−0.016	−0.014
	(0.018)	(0.018)	(0.024)	(0.021)	(0.018)	(0.022)	(0.025)
Bias-corrected	0.009	−0.011	0.010	0.183***	−0.012	−0.013	−0.003
	(0.018)	(0.018)	(0.024)	(0.021)	(0.018)	(0.022)	(0.025)
Treatment (bias-corrected, robust SE)	0.009	−0.011	0.010	0.183***	−0.012	−0.013	−0.003
	(0.022)	(0.021)	(0.027)	(0.024)	(0.021)	(0.027)	(0.027)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	727	1,212	727	1,132	1,767	1,471	1,883
Observations on the left	709	801	299	353	436	251	276
Observations on the right	494	476	211	219	265	190	194
	Panel B: Alternative bandwidths						
Dependent variable:	Share of female councilors						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.173***	0.165***	0.150***	0.145***	0.140***	0.137***	0.136***
	(0.022)	(0.019)	(0.016)	(0.014)	(0.013)	(0.012)	(0.011)
Bias-corrected	0.163***	0.179***	0.181***	0.167***	0.160***	0.154***	0.148***
	(0.022)	(0.019)	(0.016)	(0.014)	(0.013)	(0.012)	(0.011)
Treatment (bias-corrected, robust SE)	0.163***	0.179***	0.181***	0.167***	0.160***	0.154***	0.148***
	(0.030)	(0.026)	(0.023)	(0.021)	(0.019)	(0.018)	(0.017)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	300	495	718	983	1,338	1,798	2,392
Observations on the right	203	278	360	437	494	555	609

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the share of female councilors over the total number of councilors. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within each bandwidths. * p<0.1, ** p<0.05, *** p<0.01.

Table 6: Female presence on municipal councils before the reform

	Panel A: Parametric approach			
Dependent variable:	Share of female councilors			
	(1)	(2)	(3)	(4)
Treatment	0.008	−0.010	0.006	0.003
	(0.008)	(0.012)	(0.017)	(0.021)
Polynomial order	1	2	3	4
Observations	4,599	4,599	4,599	4,599
R-Squared	0.012	0.013	0.014	0.014

	Panel B: Non-parametric approach
Dependent variable:	Share of female councilors
	(1)
Treatment	−0.009
	(0.016)
Bias-corrected	−0.011
	(0.016)
Treatment (bias-corrected, robust SE)	−0.011
	(0.019)
Bandwidth	2,037
Observations on the left	730
Observations on the right	368

Notes. The table shows the results of parametric and non-parametric estimations. The dependent variable is the share of female councilors over the total number of councilors in the election prior to 2013-2015. Elections take place every 5 years, thus, in most cases, previous election is in 2008-2010, see the main text for details. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. In Panel A, the sample includes all municipalities with less than 15,000 residents that held elections in the period 2013-2015. Columns 1-4 include polynomials of orders 1-4, respectively, in the resident population, centered on the 5,000 resident threshold. Polynomials are allowed to differ on the two sides of the cut-off. In Panel B, conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. * p<0.1, ** p<0.05, *** p<0.01.

Table 7: Female presence on municipal councils: diff-in-disc

Dependent variable:	Share of female councilors	
	(1)	(2)
Treatment × After	0.127***	0.186***
	(0.010)	(0.028)
Local		X
Observations	9,198	890
R-Squared	0.327	0.504

Notes. The table shows the results of difference-in-discontinuities estimation. The dependent variable is the share of female councilors over the total number of councilors. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. *After* is an indicator variable for elections in 2013-2015. Only the coefficient of interest *Treatment*After* is reported. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 and, correspondingly, in 2008-2010. In Column 1 the sample includes all municipalities; in column 2 the sample includes municipalities within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. Standard errors clustered at municipal level in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 8: Working of the policy

	Non-parametric approach			
Dependent variable:	Female candidates	Borda score	Preference votes	Post-election Borda score
	(1)	(2)	(3)	(4)
Treatment	0.002	−0.069	0.158**	0.121**
	(0.054)	(0.087)	(0.062)	(0.052)
Bias-corrected	−0.011	−0.082	0.151**	0.124**
	(0.054)	(0.087)	(0.062)	(0.052)
Treatment (bias-corrected, robust SE)	−0.011	−0.082	0.151**	0.124*
	(0.065)	(0.102)	(0.074)	(0.064)
Bandwidth	1,278	1,294	1,199	1,456
Observations on the left	82	65	74	78
Observations on the right	104	89	95	97

Notes. The table shows the results of non-parametric estimation. The dependent variable is the share of female candidates over the total number of candidates on list s in municipality i in column 1; the Borda score of female candidates on list s in municipality i in column 2; the share of preference votes cast for female candidates on list s in municipality i in column 3; the post-election Borda score of female candidates on list s in municipality i in column 4. See the main text for details on the definition of the variables. *Treatment* is an indicator variables for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. * p<0.1, ** p<0.05, *** p<0.01.

Table 9: Working of the policy before the reform

	Non-parametric Approach			
Dependent variable:	Female candidates	Borda score	Preference votes	Post-election Borda score
	(1)	(2)	(3)	(4)
Treatment	−0.066	−0.012	−0.009	−0.029
	(0.058)	(0.046)	(0.049)	(0.048)
Bias-corrected	−0.085	−0.021	−0.027	−0.047
	(0.058)	(0.046)	(0.049)	(0.048)
Treatment (bias-corrected, robust SE)	−0.085	−0.021	−0.027	−0.047
	(0.071)	(0.057)	(0.059)	(0.059)
Bandwidth	1,059	2,157	1,917	1,536
Observations on the left	38	128	109	77
Observations on the right	73	121	119	96

Notes. The table shows the results of non-parametric estimation. The dependent variable is the share of female candidates over the total number of candidates on list s in municipality i in column 1; the Borda score of female candidates on list s in municipality i in column 2; the share of preference votes cast for female candidates on list s in municipality i in column 3; the post-election Borda score of female candidates on list s in municipality i in column 4. See the main text for details on the definition of the variables. All outcome variables refer to the election prior to 2013, which is in most cases in 2008. *Treatment* is an indicator variables for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. The sample includes all lists presented in the previous election in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. * p<0.1, ** p<0.05, *** p<0.01.

Table 10: Other voting outcomes

	Non-parametric Approach			
Dependent variable:	Turnout	Female turnout	Use of preferences	Female education
	(1)	(2)	(3)	(4)
Treatment	−0.015	−0.011	0.858***	−0.130
	(0.011)	(0.010)	(0.157)	(0.287)
Bias-corrected	−0.018*	−0.014	0.854***	−0.151
	(0.011)	(0.010)	(0.157)	(0.287)
Treatment (bias-corrected, robust SE)	−0.018	−0.014	0.854***	−0.151
	(0.013)	(0.012)	(0.199)	(0.345)
Bandwidth	1,742	1,901	852	2,025
Observations on the left	605	679	62	663
Observations on the right	322	352	74	350

Notes. The table reports results of non-parametric estimation. The dependent variable is turnout, measured as the share of actual voters over eligible voters in municipality i in column 1; female turnout, measured as the share of actual female voters over eligible female voters in municipality i in column 2; the number of preference votes over the total number of actual voters for list s in municipality i in column 3; the average number of years of education of elected female councilors in municipality i in column 4. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. In column 1,2 and 4 the sample includes municipalities with less than 15,000 residents that held elections in 2013-2015, within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. In column 3, the sample includes municipalities with less than 15,000 residents that held elections in 2013 for which preference votes were available, within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. * p<0.1, ** p<0.05, *** p<0.01.

Table 11: Quality of elected councilors: diff-in-disc

Dependent variable:	Years of education					
	Male-Female	Male-Female	Male	Male	Female	Female
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment × After	0.338	−0.430	0.212*	−0.058	−0.141	0.273
	(0.215)	(0.422)	(0.118)	(0.220)	(0.183)	(0.345)
Local		X		X		X
Observations	8,456	1,700	9,141	2,032	8,474	2,044
R-Squared	0.004	0.008	0.023	0.037	0.006	0.013

Notes. The table shows the results of difference-in-discontinuities estimation. The dependent variable is the difference between the average years of education of male and female councilors in columns 1-2; the average years of education of male councilors in columns 3-4, the average years of education of female councilors in columns 5-6. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. *After* is an indicator variable for elections in 2013-2015. Only the coefficient of interest *Treatment*After* is reported. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 and, correspondingly, in 2008-2010. The results are computed for the entire sample in column 1, 3 and 5, and for the sample of municipalities within the optimal bandwidth selected by one common MSE-optimal bandwidth selector around the cut-off of 5,000 residents in column 2, 4 and 6. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table 12: Other voting outcomes before the reform

	Non-parametric Approach			
Dependent variable:	Turnout	Female turnout	Use of preferences	Female education
	(1)	(2)	(3)	(4)
Treatment	−0.014	−0.001	0.245	0.076
	(0.010)	(0.012)	(0.154)	(0.375)
Bias-corrected	−0.018*	−0.002	0.219	0.155
	(0.010)	(0.012)	(0.154)	(0.375)
Treatment (bias-corrected, robust SE)	−0.018	−0.002	0.219	0.155
	(0.012)	(0.015)	(0.184)	(0.448)
Bandwidth	1,702	2,032	997	1,586
Observations on the left	591	726	35	497
Observations on the right	313	366	73	278

Notes. The table reports results of non-parametric estimation. The dependent variable is turnout, measured as the share of actual voters over eligible voters in municipality i in column 1; female turnout, measured as the share of actual female voters over female eligible voters in municipality i in column 2; the number of preference votes over the total number of actual voters for list s in municipality i in column 3; the average number of years of education of elected female councilors in municipality i in column 4. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Outcome variables refer to the election prior to 2013-2015, which is in 2008-2010. In column 1,2 and 4 the sample includes municipalities with less than 15,000 residents that held elections in 2013-2015, within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. In column 3, the sample includes municipalities with less than 15,000 residents that held elections in 2013 for which preference votes for the previous elections were available, within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. * p<0.1, ** p<0.05, *** p<0.01.

Table 13: Preference votes in regional elections: diff-in-disc

Dependent variable:	Average preference votes	
	(1)	(2)
Treatment × After	2.838***	1.719
	(0.788)	(1.086)
Local		X
Observations	47,474	15,937
R-Squared	0.076	0.072

Notes. The table shows the results of difference-in-discontinuities estimation. The dependent variable is the average number of preference votes cast in municipality i for female candidates on list s in regional elections. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. *After* is an indicator variable for regional elections after 2013. Only the coefficient of interest *Treatment*After* is reported. The sample includes treated and control municipalities in regions which held elections after 2013 and, correspondingly, in 2010, which do not apply double preference voting at regional level. The results are computed for the entire sample in column 1, and for the sample of municipalities within the optimal bandwidth around the cut-off of 5,000 residents in column 2. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

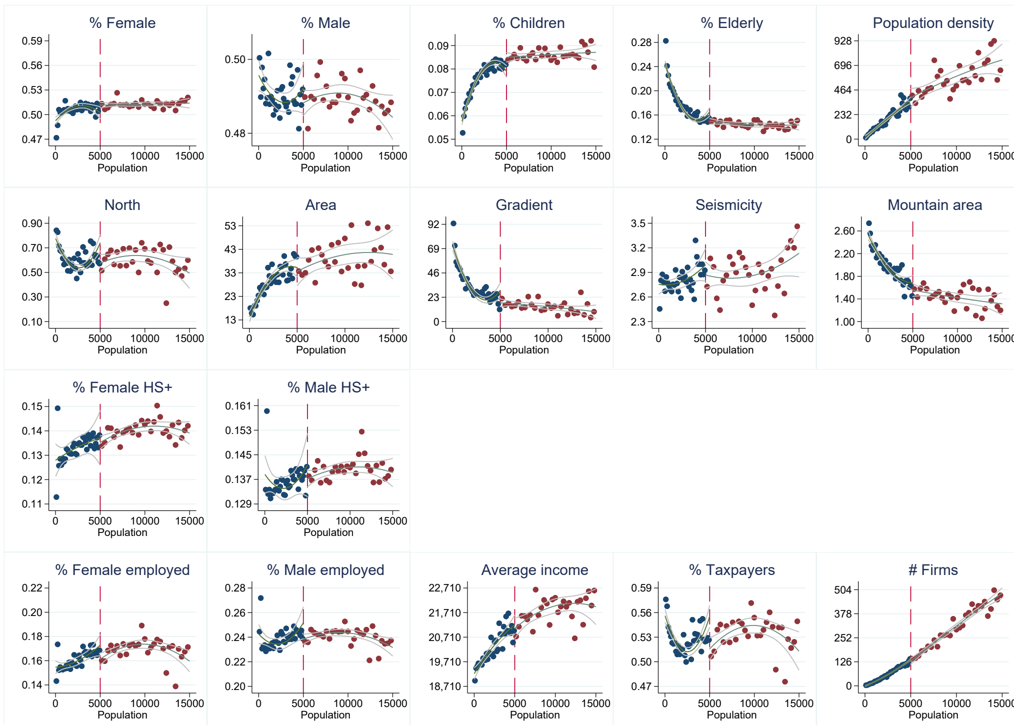

Figure 1: Balance check of covariates

Notes. The figure plots the binned averages of demographic, geographic, education and economic characteristics at municipal level against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. Row 1 reports female, male, children (0-9) and elderly (+70) shares of the population and the population density. Row 2 reports a dummy North for geographical location, area in squared Km, gradient, degree of seismicity on a scale 0-4, mountain area on a scale 1-3. Row 3 reports the share of female and male population with upper secondary education or higher. Row 4 reports the share of employed females and males, average income (total municipality income over the number of taxpayers), share of taxpayers, and the number of limited liability companies. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015.

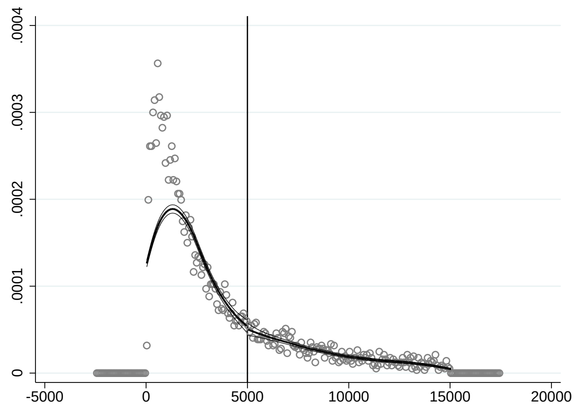

Figure 2: McCrary test

Notes. The figure plots the density of the municipal population. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015.

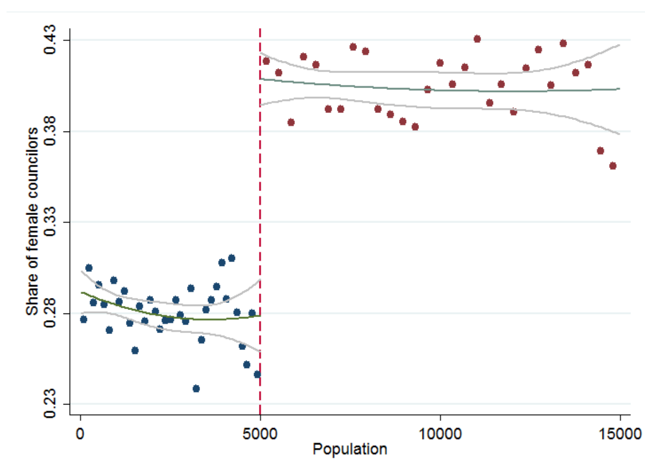

Figure 3: Female presence on municipal councils

Notes. The figure plots the binned averages of the share of female councilors against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015.

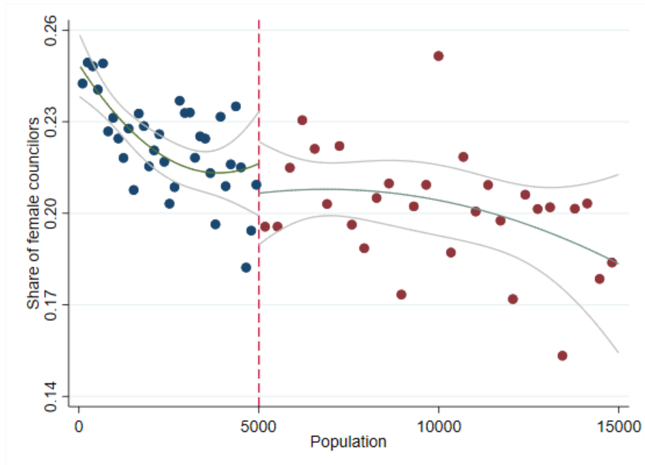

Figure 4: Female presence on municipal councils before the reform

Notes. The figure plots the binned averages of the share of female councilors in the election before the reform (2008-2010) against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015.

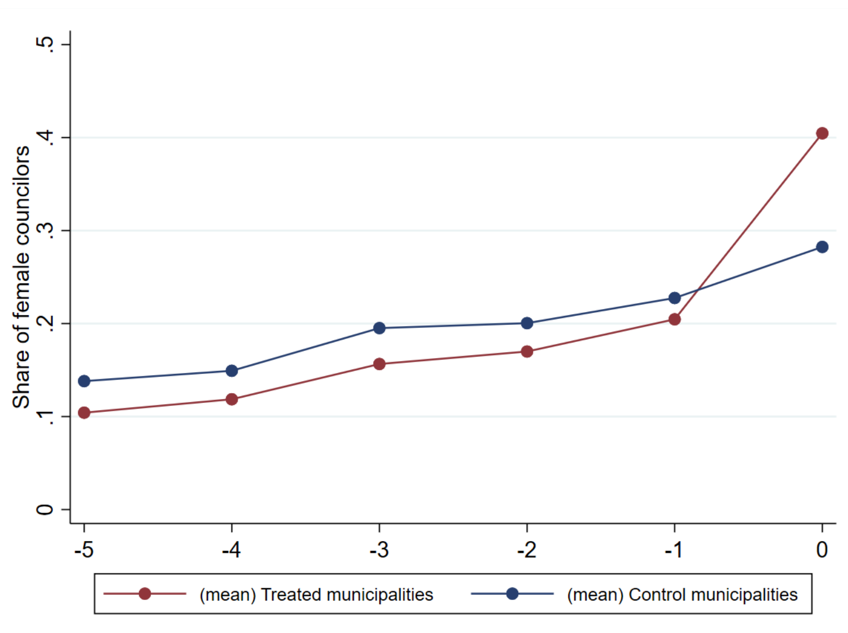

Figure 5: Female presence on municipal councils: parallel trends

Notes. The figure plots the share of female councilors in municipalities with more than 5,000 residents (treated) and municipalities with less than 5,000 residents (control). On the horizontal axis, which goes from -5 to 0, we report the outcomes of the elections prior to the 2013 election (coded as 0). The sample includes all the municipalities that held elections in 2013-2015. Only elections from 1997 onward are considered.

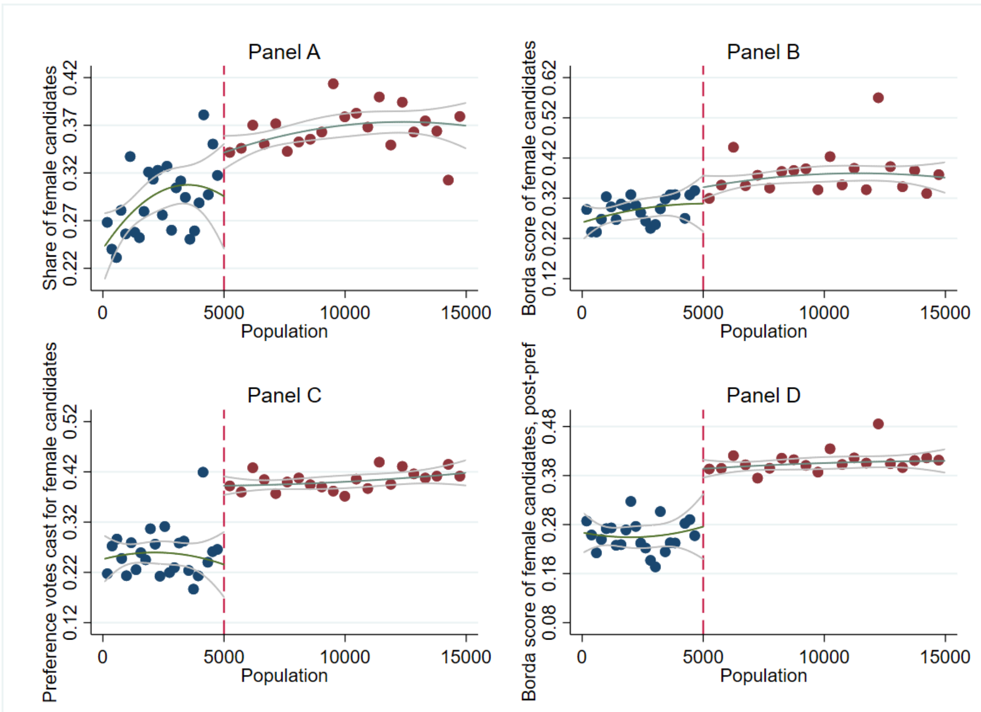

Figure 6: Working of the policy

Notes. The figure plots the binned averages of four outcomes against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. Panel A reports the share of female candidates over the total number of candidates on list s in municipality i; Panel B reports the Borda score of female candidates on list s in municipality i; Panel C reports the share of preference votes cast for female candidates on list s in municipality i; Panel D reports the post-election Borda score of female candidates on list s in municipality i. See the main text for details on the definition of the variables. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013.

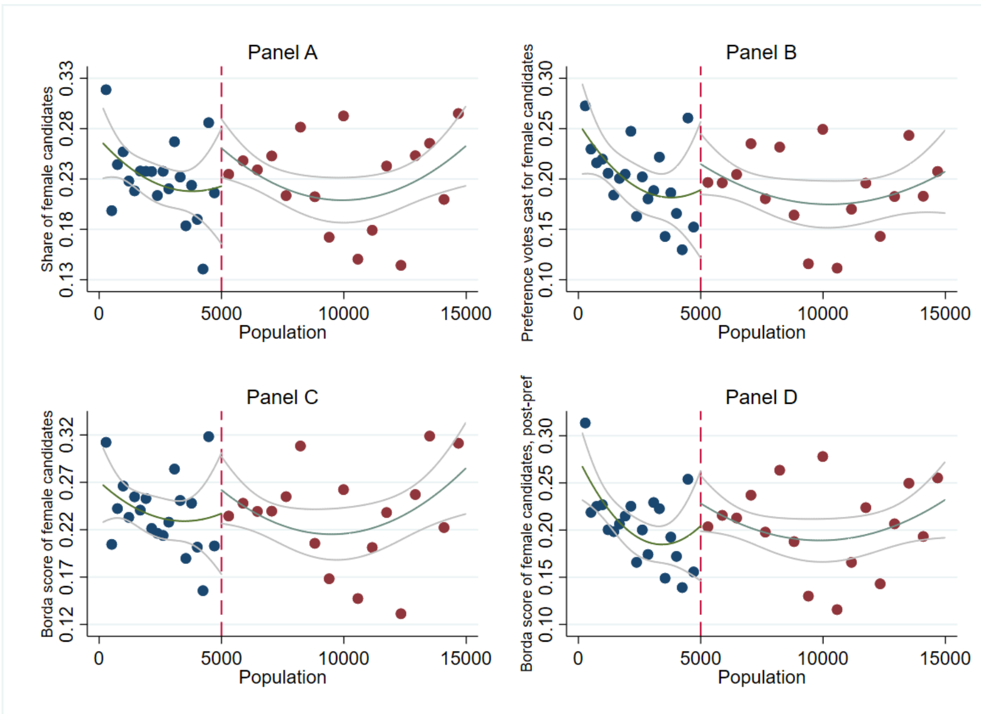

Figure 7: Working of the policy before the reform

Notes. The figure plots the binned averages of four outcomes against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. Outcomes are measured in the election before the reform, which in most cases is 2008. Panel A reports the share of female candidates over the total number of candidates on list s in municipality i; Panel B reports the Borda score of female candidates on list s in municipality i; Panel C reports the share of preference votes cast for female candidates on list s in municipality i; Panel D reports the post-election Borda score of female candidates on list s in municipality i. See the main text for details on the definition of the variables. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013.

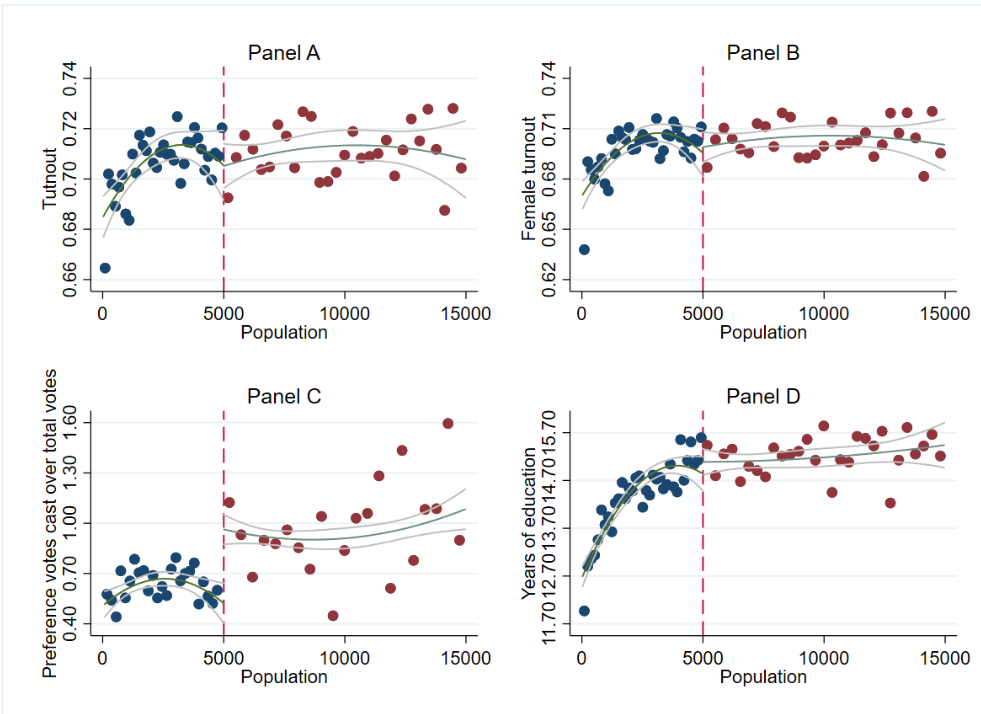

Figure 8: Other voting outcomes

Notes. The figure plots the binned averages of four outcomes against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. Panel A reports turnout, measured as the share of actual voters over eligible voters in municipality i; Panel B reports female turnout, measured as the share of actual female voters over eligible female voters in municipality i; Panel C reports the number of preference votes over the total number of actual voters for list s in municipality i; Panel D reports the average number of years of education of elected female councilors in municipality i. The sample includes all municipalities that held election in 2013-2015 in Panel A, B and D, and includes all municipalities that held election in 2013 for which preference votes were available in 2013 in Panel C.

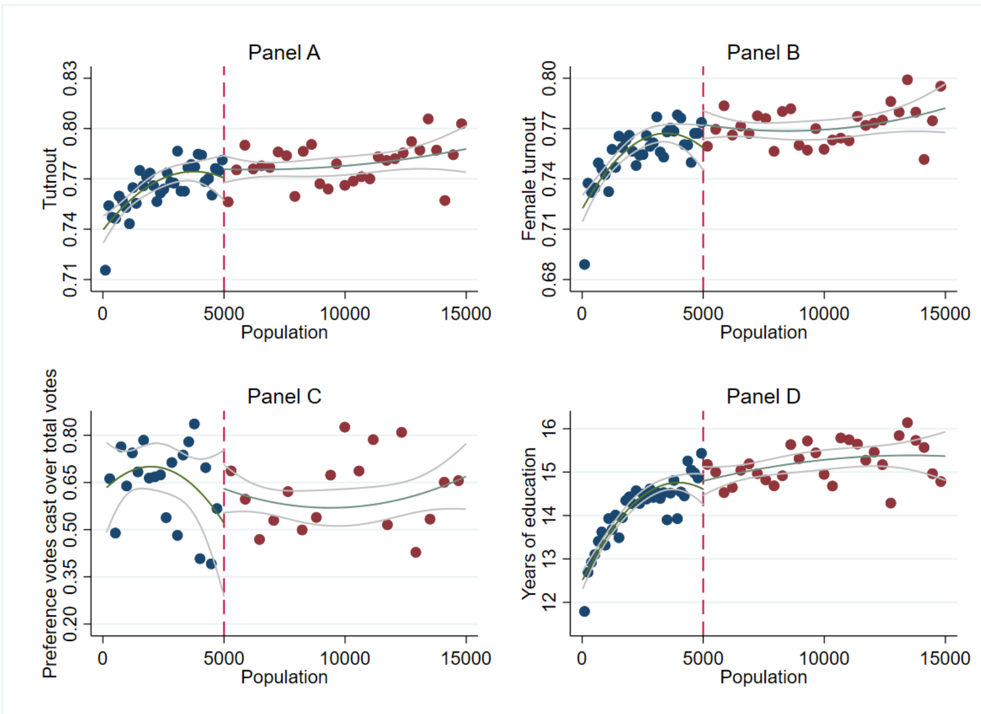

Figure 9: Other voting outcomes before the reform

Notes. The figure plots the binned averages of four outcomes against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence intervals. All outcomes are measured in the election before the reform, which is in 2008-2010. Panel A reports turnout, measured as the share of actual voters over eligible voters in municipality i; Panel B reports female turnout, measured as the share of actual female voters over eligible female voters in municipality i; Panel C reports the number of preference votes over the total number of actual voters for list s in municipality i; Panel D reports the average number of years of education of elected female councilors in municipality i. The sample includes all municipalities that held election in 2013-2015 in Panel A, B and D, and includes all municipalities that held election in 2013 for which preference votes were available in 2008 in Panel C.

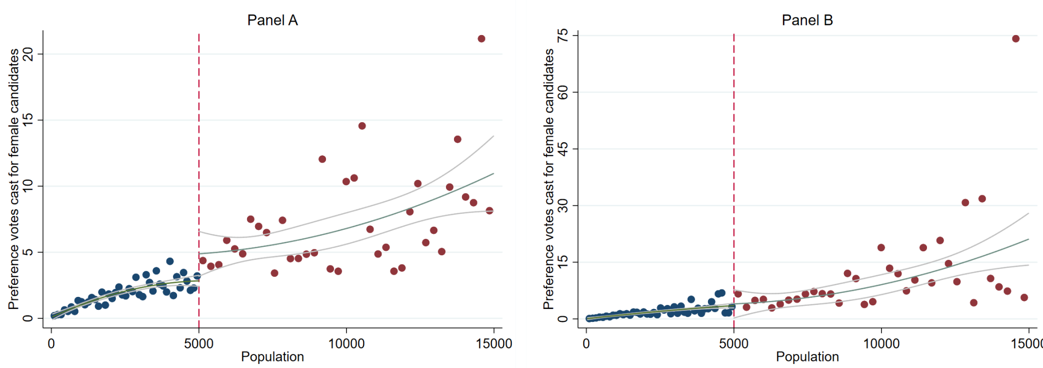

Figure 10: Preferences cast for female candidates in regional elections: before and after the reform

Notes. Panel A plots the binned averages of average preference votes cast in municipality i for female candidates on list s in regional elections in 2013-2015 against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cut-off and the 95% confidence. The sample includes lists presented in municipalities with less than 15,000 residents that held elections in the period 2013-2015 and that are in regions which do not apply double preference voting at regional level. Panel B shows the analogous plot for elections before the introduction of the reform (the relevant regional election is 2010). The sample includes lists presented municipalities with less than 15,000 residents that held elections in the period 2013-2015 and that are in regions which do not apply double preference voting at regional level.

Appendix

Table A.1: Female presence on municipal councils: geographical areas

	Non-parametric Approach		
Dependent variable:	Share of female councilors		
	(1)	(2)	(3)
Treatment	0.137***	0.159***	0.215***
	(0.028)	(0.060)	(0.028)
Bias-corrected	0.147***	0.160***	0.211***
	(0.028)	(0.060)	(0.028)
Treatment (bias-corrected, robust SE)	0.147***	0.160**	0.211***
	(0.032)	(0.077)	(0.034)
Area	North	Center	South
Bandwidth	986	2,061	1,886
Observations on the left	187	110	152
Observations on the right	118	46	94

Notes. The table shows the results of non-parametric estimation. The dependent variable is the share of female councilors over the total number of councilors. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Columns 1, 2 and 3 show the results for municipalities in the North, Center and South, respectively. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around the cut-off of 5,000 residents. *p<0.1, ** p<0.05, *** p<0.01.

Table A.2: Female presence on municipal councils before the reform: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Share of female councilors						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.013	0.015	0.006	−0.009	0.035	0.004	0.003
	(0.018)	(0.016)	(0.015)	(0.016)	(0.023)	(0.021)	(0.023)
Bias-corrected	0.017	0.017	0.011	−0.011	0.036	0.008	0.009
	(0.018)	(0.016)	(0.015)	(0.016)	(0.023)	(0.021)	(0.023)
Treatment (bias-corrected, robust SE)	0.017	0.017	0.011	−0.011	0.036	0.008	0.009
	(0.021)	(0.020)	(0.017)	(0.019)	(0.027)	(0.025)	(0.027)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	532	1,048	1,474	2,037	1,745	1,952	1,890
Observations on the left	491	671	667	730	429	351	277
Observations on the right	379	429	396	368	261	242	194

	Panel B: Alternative bandwidths						
Dependent variable:	Share of female councilors						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.001	−0.009	−0.009	−0.006	−0.006	−0.005	−0.005
	(0.023)	(0.019)	(0.016)	(0.014)	(0.013)	(0.012)	(0.011)
Bias-corrected	0.009	0.008	−0.007	−0.010	−0.009	−0.010	−0.009
	(0.023)	(0.019)	(0.016)	(0.014)	(0.013)	(0.012)	(0.011)
Treatment (bias-corrected, robust SE)	0.009	0.008	−0.007	−0.010	−0.009	−0.010	−0.009
	(0.034)	(0.028)	(0.024)	(0.021)	(0.020)	(0.018)	(0.017)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	300	495	718	983	1,338	1,798	2,392
Observations on the right	203	278	360	437	494	555	609

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the share of female councilors over the total number of councilors in the election prior to 2013-2015. Elections take place every 5 years, thus, in most cases, previous election is in 2008-2010. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015 within each bandwidth. Standard errors clustered at municipal level in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table A.3: Share of female candidates in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Share of female candidates						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.008	−0.008	0.055	0.002	0.025	0.024	0.014
	(0.052)	(0.050)	(0.035)	(0.054)	(0.025)	(0.019)	(0.019)
Bias-corrected	−0.021	0.004	0.066*	−0.011	0.033	0.033*	0.018
	(0.052)	(0.050)	(0.035)	(0.054)	(0.025)	(0.019)	(0.019)
Treatment (bias-corrected, robust SE)	−0.021	0.004	0.066	−0.011	0.033	0.033	0.018
	(0.066)	(0.059)	(0.043)	(0.065)	(0.029)	(0.022)	(0.023)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	724	1,160	1,639	1,278	1,477	1,411	2,166
Observations on the left	133	159	165	82	109	119	177
Observations on the right	112	83	123	104	124	91	76
	Panel B: Alternative bandwidths						
Dependent variable:	Share of female candidates						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.027	−0.012	−0.004	0.011	0.025	0.034	0.035*
	(0.070)	(0.043)	(0.034)	(0.029)	(0.025)	(0.023)	(0.021)
Bias-corrected	0.057	0.053	−0.019	−0.032	−0.026	−0.010	0.006
	(0.070)	(0.043)	(0.034)	(0.029)	(0.025)	(0.023)	(0.021)
Treatment (bias-corrected, robust SE)	0.057	0.053	−0.019	−0.032	−0.026	−0.010	0.006
	(0.176)	(0.095)	(0.070)	(0.053)	(0.045)	(0.039)	(0.035)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	62	105	145	207	277	389	462
Observations on the right	87	112	173	211	251	268	280

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variables is the share of female candidates over the total number of candidates on list s in municipality i. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within each bandwidth. Standard errors clustered at municipal level in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table A.4: Borda score of female candidates in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Borda score						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.014	−0.015	−0.120*	−0.069	0.074***	0.028	0.013
	(0.053)	(0.063)	(0.070)	(0.087)	(0.029)	(0.039)	(0.034)
Bias-corrected	−0.020	−0.033	−0.141**	−0.082	0.086***	0.046	0.008
	(0.053)	(0.063)	(0.070)	(0.087)	(0.029)	(0.039)	(0.034)
Treatment (bias-corrected, robust SE)	−0.020	−0.033	−0.141*	−0.082	0.086***	0.046	0.008
	(0.067)	(0.075)	(0.082)	(0.102)	(0.034)	(0.045)	(0.039)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	913	887	1,163	1,294	2,201	1,628	2,151
Observations on the left	135	105	75	65	132	127	161
Observations on the right	114	60	59	89	159	93	69
	Panel B: Alternative bandwidths						
Dependent variable:	Borda score						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.062	−0.048	−0.010	−0.003	0.007	0.021	0.028
	(0.108)	(0.069)	(0.048)	(0.040)	(0.035)	(0.031)	(0.029)
Bias-corrected	0.180*	−0.082	−0.086*	−0.045	−0.040	−0.041	−0.027
	(0.108)	(0.069)	(0.048)	(0.040)	(0.035)	(0.031)	(0.029)
Treatment (bias-corrected, robust SE)	0.180	−0.082	−0.086	−0.045	−0.040	−0.041	−0.027
	(0.332)	(0.153)	(0.113)	(0.082)	(0.066)	(0.055)	(0.049)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	54	80	116	173	233	315	379
Observations on the right	72	97	154	185	225	242	251

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the Borda score of female candidates on list s in municipality i. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.5: Share of preference votes for female candidates in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Preference votes						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.025	0.029	0.033	0.158**	0.050	−0.043*	0.033
	(0.045)	(0.054)	(0.044)	(0.062)	(0.033)	(0.022)	(0.027)
Bias-corrected	0.031	0.046	0.042	0.151**	0.061*	−0.044**	0.040
	(0.045)	(0.054)	(0.044)	(0.062)	(0.033)	(0.022)	(0.027)
Treatment (bias-corrected, robust SE)	0.031	0.046	0.042	0.151**	0.061	−0.044*	0.040
	(0.057)	(0.063)	(0.056)	(0.074)	(0.037)	(0.026)	(0.032)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	1,029	1,056	1,526	1,199	1,161	1,541	1,850
Observations on the left	184	144	145	74	87	134	156
Observations on the right	134	83	113	95	104	95	67
	Panel B: Alternative bandwidths						
Dependent variable:	Preference votes						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.176**	0.115***	0.119***	0.132***	0.140***	0.146***	0.146***
	(0.071)	(0.044)	(0.035)	(0.031)	(0.027)	(0.024)	(0.023)
Bias-corrected	0.105	0.218***	0.121***	0.101***	0.108***	0.118***	0.130***
	(0.071)	(0.044)	(0.035)	(0.031)	(0.027)	(0.024)	(0.023)
Treatment (bias-corrected, robust SE)	0.105	0.218**	0.121*	0.101*	0.108**	0.118***	0.130***
	(0.167)	(0.096)	(0.073)	(0.055)	(0.046)	(0.041)	(0.037)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	62	105	145	207	276	387	460
Observations on the right	87	112	173	211	251	268	280

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the share of preference votes cast for female candidates on list s in municipality i. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.6: Post-election Borda score of female candidates in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Post-election Borda score						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.061	0.057	−0.001	0.121**	0.005	−0.056**	0.028
	(0.056)	(0.057)	(0.054)	(0.052)	(0.026)	(0.022)	(0.024)
Bias-corrected	0.074	0.067	−0.011	0.124**	0.011	−0.062***	0.033
	(0.056)	(0.057)	(0.054)	(0.052)	(0.026)	(0.022)	(0.024)
Treatment (bias-corrected, robust SE)	0.074	0.067	−0.011	0.124*	0.011	−0.062**	0.033
	(0.069)	(0.068)	(0.068)	(0.064)	(0.030)	(0.026)	(0.028)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	838	1,021	1,447	1,456	1,405	1,513	2,224
Observations on the left	120	120	112	78	88	116	166
Observations on the right	107	62	89	97	110	88	69
	Panel B: Alternative bandwidths						
Dependent variable:	Post-election Borda score						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.147*	0.117**	0.114***	0.112***	0.114***	0.124***	0.126***
	(0.077)	(0.049)	(0.037)	(0.032)	(0.028)	(0.026)	(0.024)
Bias-corrected	0.106	0.168***	0.123***	0.122***	0.107***	0.095***	0.104***
	(0.077)	(0.049)	(0.037)	(0.032)	(0.028)	(0.026)	(0.024)
Treatment (bias-corrected, robust SE)	0.106	0.168	0.123	0.122**	0.107**	0.095**	0.104***
	(0.223)	(0.109)	(0.082)	(0.060)	(0.049)	(0.043)	(0.039)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	54	80	116	173	233	315	379
Observations on the right	72	97	154	185	225	242	251

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the post-election Borda score of female candidates on list s in municipality i *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.7: Working of the policy: diff-in-disc

		Panel A		
Dependent variable:	Female candidates	Borda score	Preference votes	Post-election Borda score
	(1)	(2)	(3)	(4)
Treatment × After	−0.008	0.026	0.103***	0.087***
	(0.026)	(0.031)	(0.030)	(0.029)
Local				
Observations	1,871	1,722	1,798	1,722
R-Squared	0.133	0.102	0.228	0.247
		Panel B		
Dependent variable:	Female candidates	Borda score	Preference votes	Post-election Borda score
	(1)	(2)	(3)	(4)
Treatment × After	−0.007	0.000	0.119**	0.088
	(0.050)	(0.061)	(0.058)	(0.055)
Local	X	X	X	X
Observations	550	502	549	502
R-Squared	0.171	0.128	0.334	0.346

Notes. The table shows the results of difference-in-discontinuities estimation. The dependent variable is the share of female candidates over the total number of candidates on list s in municipality i in column 1, the Borda score of female candidates on list s in municipality i in column 2, the share of preference votes cast for female candidates on list s in municipality i in column 3, the post-election Borda score of female candidates on list s in municipality i in column 4. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. *After* is an indicator variable for elections in 2013-2015. Only the coefficient of interest *Treatment*After* is reported. The sample includes all lists presented in municipalities with less than 15,000 residents that held elections in 2013 and, correspondingly, in 2008. Panel A reports results for the entire sample of municipalities; Panel B reports results for the sample of municipalities within the optimal bandwidth around the cut-off of 5,000 residents. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.8: Turnout in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Total turnout						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.013	0.011	−0.010	−0.015	−0.012	0.006	0.019
	(0.012)	(0.012)	(0.011)	(0.011)	(0.011)	(0.011)	(0.012)
Bias-corrected	−0.013	0.014	−0.011	−0.018*	−0.015	0.003	0.023*
	(0.012)	(0.012)	(0.011)	(0.011)	(0.011)	(0.011)	(0.012)
Treatment (bias-corrected, robust SE)	−0.013	0.014	−0.011	−0.018	−0.015	0.003	0.023*
	(0.014)	(0.014)	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	595	908	1,219	1,742	1,735	2,167	1,806
Observations on the left	566	548	529	605	426	414	264
Observations on the right	411	385	342	322	261	269	190
	Panel B: Alternative bandwidths						
Dependent variable:	Total turnout						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.026*	−0.018	−0.013	−0.011	−0.010	−0.009	−0.010
	(0.014)	(0.011)	(0.010)	(0.009)	(0.008)	(0.007)	(0.007)
Bias-corrected	−0.025*	−0.030***	−0.025**	−0.019**	−0.016**	−0.015**	−0.012*
	(0.014)	(0.011)	(0.010)	(0.009)	(0.008)	(0.007)	(0.007)
Treatment (bias-corrected, robust SE)	−0.025	−0.030*	−0.025*	−0.019	−0.016	−0.015	−0.012
	(0.021)	(0.017)	(0.015)	(0.013)	(0.012)	(0.011)	(0.010)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	300	495	718	983	1,338	1,798	2,392
Observations on the right	203	278	360	437	494	555	609

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is turnout, measured as the share of actual voters over eligible voters in municipality i in the period 2013-2015. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.9: Female turnout in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Female turnout						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.014	0.007	−0.010	−0.011	−0.013	0.007	0.018
	(0.012)	(0.011)	(0.011)	(0.010)	(0.011)	(0.011)	(0.012)
Bias-corrected	−0.014	0.009	−0.011	−0.014	−0.016	0.005	0.021*
	(0.012)	(0.011)	(0.011)	(0.010)	(0.011)	(0.011)	(0.012)
Treatment (bias-corrected, robust SE)	−0.014	0.009	−0.011	−0.014	−0.016	0.005	0.021
	(0.014)	(0.014)	(0.013)	(0.012)	(0.013)	(0.013)	(0.014)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	579	976	1,219	1,901	1,781	2,099	1,826
Observations on the left	545	600	529	679	438	392	268
Observations on the right	402	406	342	352	266	261	191
	Panel B: Alternative bandwidths						
Dependent variable:	Female turnout						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.023*	−0.015	−0.011	−0.009	−0.008	−0.008	−0.008
	(0.014)	(0.011)	(0.010)	(0.009)	(0.008)	(0.007)	(0.007)
Bias-corrected	−0.022	−0.027**	−0.022**	−0.017*	−0.014*	−0.012	−0.009
	(0.014)	(0.011)	(0.010)	(0.009)	(0.008)	(0.007)	(0.007)
Treatment (bias-corrected, robust SE)	−0.022	−0.027	−0.022	−0.017	−0.014	−0.012	−0.009
	(0.021)	(0.017)	(0.015)	(0.013)	(0.012)	(0.011)	(0.010)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	300	495	718	983	1,338	1,798	2,392
Observations on the right	203	278	360	437	494	555	609

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is female turnout, measured as the share of actual female voters over eligible female voters in municipality i in the period 2013-2015. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within each bandwidth. Standard errors clustered at municipal level in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table A.10: Use of preference votes in municipal elections: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Use of preferences						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.011	−0.101	−0.176	0.858***	−0.332***	−0.009	−0.277***
	(0.053)	(0.063)	(0.423)	(0.157)	(0.100)	(0.069)	(0.062)
Bias-corrected	0.042	−0.127**	−0.096	0.854***	−0.347***	−0.020	−0.288***
	(0.053)	(0.063)	(0.423)	(0.157)	(0.100)	(0.069)	(0.062)
Treatment (bias-corrected, robust SE)	0.042	−0.127*	−0.096	0.854***	−0.347***	−0.020	−0.288***
	(0.060)	(0.072)	(0.473)	(0.199)	(0.116)	(0.080)	(0.071)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	598	547	437	852	1,358	1,915	2,531
Observations on the left	122	69	38	62	96	168	212
Observations on the right	88	42	28	74	118	107	89
	Panel B: Alternative bandwidths						
Dependent variable:	Use of preferences						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.766***	0.732***	0.644***	0.573***	0.495***	0.454***	0.444***
	(0.144)	(0.090)	(0.069)	(0.061)	(0.055)	(0.050)	(0.047)
Bias-corrected	0.887***	0.793***	0.839***	0.855***	0.818***	0.716***	0.624***
	(0.144)	(0.090)	(0.069)	(0.061)	(0.055)	(0.050)	(0.047)
Treatment (bias-corrected, robust SE)	0.887**	0.793***	0.839***	0.855***	0.818***	0.716***	0.624***
	(0.359)	(0.195)	(0.146)	(0.109)	(0.091)	(0.079)	(0.073)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	62	105	145	207	277	389	462
Observations on the right	87	112	173	211	251	268	280

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the number of preference votes over the total number of actual voters for list s in municipality i. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in 2013 for which preference votes were available within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in 2013 for which preference votes were available within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.11: Years of education of female councilors: robustness checks

	Panel A: Alternative cut-offs						
Dependent variable:	Years of education						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	−0.003	0.143	0.488	−0.130	0.062	−0.026	0.049
	(0.340)	(0.348)	(0.345)	(0.287)	(0.373)	(0.366)	(0.409)
Bias-corrected	0.023	0.211	0.487	−0.151	0.180	−0.027	−0.081
	(0.340)	(0.348)	(0.345)	(0.287)	(0.373)	(0.366)	(0.409)
Treatment (bias-corrected, robust SE)	0.023	0.211	0.487	−0.151	0.180	−0.027	−0.081
	(0.408)	(0.417)	(0.422)	(0.345)	(0.436)	(0.442)	(0.479)
Cut-off	2,000	3,000	4,000	5,000	6,000	7,000	8,000
Bandwidth	608	1,071	1,589	2,025	1,354	1,875	2,022
Observations on the left	509	616	664	663	293	323	279
Observations on the right	381	392	396	350	205	215	192

	Panel B: Alternative bandwidths						
Dependent variable:	Years of education						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Treatment	0.020	−0.159	−0.132	−0.098	−0.064	−0.043	−0.074
	(0.417)	(0.336)	(0.288)	(0.260)	(0.238)	(0.220)	(0.204)
Bias-corrected	0.123	0.109	−0.097	−0.135	−0.143	−0.115	−0.021
	(0.417)	(0.336)	(0.288)	(0.260)	(0.238)	(0.220)	(0.204)
Treatment (bias-corrected, robust SE)	0.123	0.109	−0.097	−0.135	−0.143	−0.115	−0.021
	(0.617)	(0.500)	(0.430)	(0.380)	(0.344)	(0.318)	(0.297)
Cut-off	5,000	5,000	5,000	5,000	5,000	5,000	5,000
Bandwidth	1,000	1,500	2,000	2,500	3,000	3,500	4,000
Observations on the left	280	457	656	890	1,213	1,616	2,149
Observations on the right	194	267	346	416	469	527	576

Notes. The table shows the robustness checks of non-parametric estimation. The dependent variable is the average number of years of education of elected female councilors in municipality i. *Treatment* is an indicator variable for municipalities with more than 5,000 residents. Only the coefficient of interest *Treatment* is reported. Conventional RD estimates with a conventional variance estimator, bias-corrected RD estimates with a conventional variance estimator, and bias-corrected RD estimates with a robust variance estimator are reported. Panel A reports results for placebo cut-offs, namely 2,000, 3,000, 4,000, 6,000, 7,000 and 8,000 residents, in addition to the correct 5,000 one reported in Column 4. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within the optimal bandwidth selected by one common MSE-optimal bandwidth selector (Calonico et al., 2017) around each cut-off. Panel B reports results for alternative bandwidths, namely 1,000, 1,500, 2,000, 2,500, 3,000, 3,500 and 4,000. The sample includes municipalities with less than 15,000 residents that held elections in 2013-2015 within each bandwidth. Standard errors clustered at municipal level in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

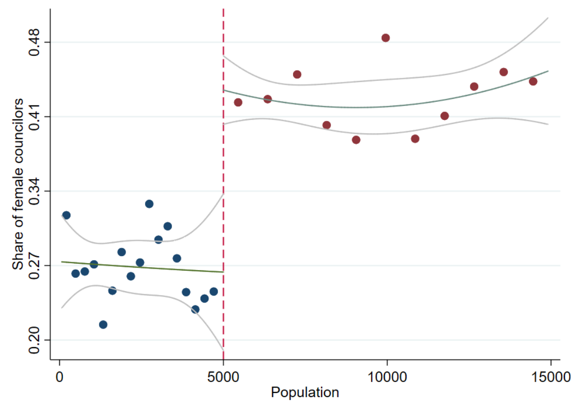

Figure A.1: Female councilors

Notes. The figure plots the binned averages of the share of female councilors elected in the second election with double preference voting against the municipal population, together with the quadratic polynomial fit on both sides of the 5,000 resident cutoff and the 95% confidence intervals. Elections take place every 5 years, thus, in most cases, the second election with double preference voting is in 2018. The sample includes municipalities with less than 15,000 residents that held elections in the period 2013-2015.