Consequential Inclusion: Gender Quotas and Social Class Diversity in Legislative Politics

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Abstract

Can policies aimed at increasing gender representation also shape the social class makeup of legislatures? This paper examines that question through a natural experiment in Italy, where some municipalities temporarily adopted mandatory gender quotas on candidate lists in the 1990s. Leveraging archival data, I show that these quotas led to more women being elected, especially from lower-middle and working-class backgrounds, without reducing the presence of working-class men. Quota municipalities also saw higher voter turnout and legislative seat share gains for locally rooted parties, suggesting that class diversification was driven by shifts in voter behavior. These findings show that gender quotas can generate broader representational change, reshaping the social profile of legislative bodies via demand-side change. The paper contributes to our understanding of the downstream effects of inclusion for democratic institutions.

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1 Introduction

In the past decades, gender quotas have become one of the most prevalent institutional tools for addressing inequalities in political representation. While these have helped increase the number of women in legislatures across the world[1], much less is known about which women benefit—and whether quotas alter the broader social makeup of political elites.

The question is increasingly salient in a political climate marked by growing mistrust of institutions and the rise of 'anti-rights' rhetoric (UN Women, 2025). Across Europe and the Americas, socio-economic elites have appealed to working-class voters by presenting themselves as champions of the 'ordinary citizen' and critics of identity-based reforms (Weeks and Allen, 2023). Gender quotas, in particular, have been framed as symbols of elite feminism: measures that advance well-educated, professional women while doing little to address the concerns of the working class.[2] These portrayals echo broader concerns raised by the literature, which has shown that quotas often benefit candidates with higher levels of education (Aldrich and Daniel, 2024; Baltrunaite et al., 2014), social status (Karekurve-Ramachandra and Lee, 2020), and more extensive personal networks (Folke et al., 2021).

This paper directly addresses that debate by examining whether gender quotas can enhance the descriptive representation of socially marginalized groups without producing zero-sum outcomes. Leveraging a natural experiment in 1990s Italy—where a subset of municipalities temporarily adopted mandatory gender quotas—I show that these reforms increased the number of lower-middle and working-class women in office, without reducing the representation of working-class men.

These findings support a broader theoretical argument: that institutional reforms aimed at gender inclusion can generate positive spillovers for class representation. While much of the existing literature emphasizes supply-side mechanisms—such as shifts in party nomination strategies—this study focuses on demand-side dynamics, highlighting how quotas can alter patterns of political selection through changes in voter behavior. Prior research shows that higher rates of female candidacy elevates the salience of gender issues on party platforms (Weeks, 2022) and increases various forms of womens' political participation (De Paola et al., 2014; Goyal, 2024). These dynamics can work to

[1] See e.g. Clayton (2021), Hughes et al. (2017), Hughes et al. (2019), Paxton et al. (2010), and Tripp and Kang (2008).

[2] This critique has also featured in feminist debates over parité laws in France (Bird, 2001; Lépinard, 2007) and all-women shortlists in the UK (Steer, 2023), and resonates with broader theoretical arguments that frame liberal pro-female inclusion efforts as masking deeper class inequalities (e.g. Fraser, 2020).

shift the identity of the median voter—by encouraging turnout among underrepresented groups, or changing electoral preferences. Where such shifts thus benefit parties with more socially inclusive candidate pools, quotas will indirectly improve class diversity in elected office.

To test this, I draw on historical data covering the universe of Italian municipal legislators and evaluate how quota exposure affected the political trajectories of both women and men from different class backgrounds. The empirical strategy exploits the fact that the quota was ruled unconstitutional in 1995, producing a natural comparison between municipalities that did and did not hold elections while the law was in effect (De Paola et al., 2010). Given parallel pre-treatment trends in gender-class representation, I use a difference-in-differences approach to estimate causal quota effects.

The results contribute to three strands of literature. First, they show that quotas increased representation primarily for lower-middle and working-class women, without displacing working-class men. This challenges common assumptions about zero-sum trade-offs in group-based inclusion and extends work on quota spillover effects for political minorities (see e.g. Barnes and Holman, 2020). While previous accounts emphasize norm diffusion—wherein growing female presence in office expands ideas about political competence (see also Alexander, 2012; Beauregard, 2017)—the findings here highlight a different mechanism: direct intersectional inclusion, whereby marginalized women themselves enter office.[3]

Second, refining the results of De Paola et al. (2010) and Baltrunaite et al. (2014), I uncover that the quota's effects on female office-holding were detectable even after quotas were dropped. More specifically, positive quota effects—especially for lower-middle class women—persisted up to two election cycles after quota abolition. Long-run effects such as these have been documented in other work (e.g. Beaman et al., 2009) and occur mainly because parties work to facilitate the re-election of 'new' female policymakers, even after quotas are scrapped (Bhavnani, 2009; Shair-Rosenfield, 2012).

Third, the results contribute to a growing literature on how descriptive representation reshapes electoral dynamics. Consistent with demand-side hypotheses of quota effects, mechanism tests indicate that improvements in gender-class representation were associated with an increase in voter turnout—an important finding in a period characterized by widespread electoral disengagement (Martini and Quaranta, 2015; Daniele et al., 2023). Additional evidence suggests that these shifts were

[3] Carnes (2015) similarly observes the importance of this intersectionality in his study of Californian local elections.

largely propelled by gains for centrist Civic List parties, which were locally organized and socially inclusive. This pattern aligns with the preferences of Italian female voters at the time, who were more likely than men to support centrist or conservative parties (Corbetta and Cavazza, 2008; Emmenegger and Manow, 2014) and exhibited low loyalty to mainstream partisan alternatives (Bull, 1997). Together, these dynamics suggest that gender quotas can mitigate political apathy by diversifying the candidate pool and expanding representational pathways for socially marginalized groups.

2 Women and workers in legislative politics

A large body of work has highlighted the substantive benefits of improving the descriptive representation of women in political legislatures. These studies have found that higher rates of female politicians serve to reduce gender stereotypes among voters (Beaman et al., 2009) and party elites (Bhalotra et al., 2018; Bhavnani, 2009), as well as moderate the prevalence of gender-based violence (Bochenkova et al., 2023; Kuipers, 2020). Female policymakers are also better at promoting political self-efficacy among women voters (Atkeson and Carrillo, 2007; Stauffer, 2021); at increasing female elections turnout (De Paola et al., 2014; Goyal, 2024), and at encouraging other forms of political participation among women (Barnes and Burchard, 2013). On the level of policy, studies find women legislators to be more prone towards supporting bills that are strongly favored by their female constituents (Bratton, 2005; Schwindt-Bayer, 2006). And, ultimately, a range of research shows that policies favored by female voters are more likely to be adopted when women are well-represented in elected government.[4]

In tandem with work on female representation, a separate yet related literature, has examined the substantive effects of working class representation in politics. While fewer in number, these studies uncover similar results to studies on gender. For one, they have found blue collar voters to hold more positive views of electoral institutions, when they are represented by legislators from their own social class (Barnes and Saxton, 2019; O'Grady, 2019; Poertner, 2023). This translates, in turn, to bolstered political participation, in the form of stronger political interest and higher election turnout (Poertner, 2023). Likewise, class studies find that blue-collar legislators are more likely to advocate and vote for policy programs that redistribute to low-income voters, who suffer from stronger employment insecu-

[4]See e.g. Bratton and Ray (2002); Chattopadhyay and Duflo (2004); Taylor-Robinson and Heath (2003), as well as Clayton (2021) for an extensive review on this link.

rity (Carnes, 2012; Carnes and Lupu, 2015; O'Grady, 2019). Finally, results from additional studies also suggest that descriptive working-class representation in formal politics positively affects redistributive social spending—the benefits of which disproportionately accrue to lower income households (e.g. Alexiadou, 2022; Hemingway, 2020).

If the numerical representation of both gender and social class matters for policy responsiveness, this raises questions about how policies to improve gender equity in legislatures, in turn, affect class representation. By definition, women will benefit from these efforts. But only where policies bolster the representation of women from lower class strata, and where they refrain from harming the political opportunity of working class men, will they function to improve descriptive class representation. The issue is arguably of considerable consequence for democratic efficacy, as both women and workers are relative majority groups within most electorates.

2.1 Secondary effects of female inclusion on class representation

How does the inclusion of women affect the social class composition of democratically elected legislatures? Resolving this question in an empirically robust manner is difficult, as polities represented by higher rates of female policymakers are also—for a multitude of reasons—more likely to produce legislatures in which workers are numerically well-represented. Here statistical inference is deceiving, because female and working class representation may well be correlated, without the former factor influencing the latter. An ideal experimental set-up for this question would be to randomly assign the share of a legislature's seats that are given to women: a factor we can't—nor want to—manipulate as social scientists. But the adoption of gender quotas allows us to approximate this experimental ideal, as their adoption induces an exogenous positive shock to the number of women that run for office.

A number of studies have leveraged gender quota enactment to examine how quotas affect the socio-economic composition of political legislatures. Almost none look at the outcome I focus on in this paper: social class, conceptualized as a person's occupation. However, their findings on how quotas affect the numerical representation of caste, ethnic and educational groups have informed this study greatly, as these variables are typically correlated with occupational class. Table 1 provides an overview of causally identified studies: the cases they test, the secondary dimension of descriptive representation that they examine, and the direction of effects that they uncover, if any. With the

exception of Aldrich and Daniel (2024), all studies examine outcomes in local level politics. And in each case, quotas are set on party lists used in PR systems, which are prepared in advance of elections.

Table 1: Secondary effects of gender quotas on descriptive representation

Study (by secondary group analyzed)	Case	Effect
Higher-level occupational class		
Lassébie (2020)	France, 2010-2016	Null
Low caste groups		
Karekurve-Ramachandra and Lee (2020)	Delhi, 2007-2017	Negative
Educational attainment		
Aldrich and Daniel (2024)	EU Parliament, 1994-2019	Positive
Baltrunaite et al. (2014)	Italy, 1983-1996	Positive
Bagues and Campa (2021)	Spain, 2012-2018	Null
Immigration-background		
Folke et al. (2015)	Sweden, 1982-2010	Null
Pre-political career income		
Besley et al. (2017)	Sweden, 1982-2014	Positive
Dynasty legislators		
Folke et al. (2021)	Sweden, 1982-2010	Positive

All-in-all, aforementioned studies have yielded a mix of findings, and theoretically they focus on what is considered *supply-side constraints* to political selection. Supply-side accounts emphasize that gender quotas can change how parties advocate for individual candidates—either as quotas affect biases in intra-party selection processes, which give certain social groups an advantage in candidacy contests, or because quotas alter the propensity of disparate socio-economic groups to run for office.

Several of these studies posit that gender quotas can have a regressive effect on the overarching diversity of legislatures. Here one set of theory has highlighted that party-gatekeeping dynamics may bias political selection committees towards selecting female candidates that come from the higher social classes. This occurs as party selection bodies place disproportional value on the secondary traits of female candidates who enter politics post-quota adoption. In other words, the biases of party selectorates force female candidates to display stronger competence signals than their male counterparts. Studies in this vein have found that gender quotas bolster office-holding among women who retain familial ties to extant or former politicians (Folke et al., 2021). Other work on Sweden, affirms that these women also display higher pre-political career incomes (Besley et al., 2017).[5] Finally, two studies—on

[5]The authors of this study analyze pre-political income adjusted for age, gender, locality, education, and employment sector. Inferences about quotas harming the underprivileged must thus assume these individuals receive lower wages after accounting for these factors.

1990s Italy (Baltrunaite et al., 2014) and European Parliament (Aldrich and Daniel, 2024)—suggest that female legislators who enter office post-quota adoption are in possession of a higher level of formal education than the men, which they replace.

Quotas can also skew the socio-economic pool of female candidates *absent* party gatekeeping. In India, Karekurve-Ramachandra and Lee (2020) find that gender quotas harm social caste diversity, as women from under-privileged casts are bound by socially conservative household norms that bar them from entering into a career in politics. In reserving a higher proportion of party lists to women, the 'new' female candidate pool thus consists disproportionately of women from higher castes. This increases the overall affluence displayed by political candidates, which in turn affects the socio-economic makeup of the legislature.

Alternatively, other supply-side accounts find that gender quotas have no effects on secondary dimensions of descriptive representation. In Sweden, Folke et al. (2015) show that the introduction of a Social Democratic zipper quota had no impact on the numerical representation of individuals with immigration backgrounds in municipal councils. Another study, on France, suggests that gender quota adoption did not affect the occupational diversity of legislators aside from increasing the number of pensioners that entered office (Lassébie, 2020). And finally, one account on Spain fails to uncover that quotas impacted the educational backgrounds of local policymakers (Bagues and Campa, 2021). Overall, null-effects imply that the introduction of quotas, as well as an ensuing increase of women in politics, do not always impact on the overarching recruitment and/or selection norms of internal party bodies.

2.2 Linking Gender Quotas to Social Class Representation

To understand how gender quotas can alter the class composition of elected legislatures, we must distinguish between two sets of mechanisms: *supply-side* and *demand-side* dynamics. As discussed, much of the existing literature focuses on supply-side effects—particularly how quotas influence the candidate selection process within parties. In these accounts, the social class profile of legislators shifts because quotas change who enters the candidate pool or who is selected by party gatekeepers. By contrast, this paper advances a broader theoretical argument: that quotas can also reshape the demand-side of political representation by influencing voter behavior—specifically, electoral participation and

party support—in ways that indirectly alter class representation. Table 2 summarizes the distinction between this perspectives.

Table 2: Theoretical perspectives on secondary effects of gender quotas on descriptive representation

Theoretical Perspective	Voter Response	Effects on Secondary Traits in Office
Supply-side dominant (selection)	No effects on electoral behavior	*Socially regressive*; reflects shifts in supply of potential legislators only.
Demand-side relevant (selection + election)	Changes in turnout and party-specific vote shares	*Socially regressive or progressive*; reflects both changes in candidate supply and voter behavior following quota adoption.

In supply-side dominant theories, gender quotas are assumed to leave voter behavior unchanged. Quotas alter descriptive representation primarily through internal party selection processes, by encouraging selectorates to nominate more elite women to meet quota targets and by shifting intra-party competition in ways that disadvantage lower-status men. These frameworks typically predict *regressive* social effects, whereby gender quotas promote more educated or affluent candidates without improving broader social diversity.

By contrast, a demand-side framework incorporates both selection and election stages. Gender quotas increase the number of female candidates on party lists, which in turn reshapes electoral campaigns. Greater female candidacy elevates the salience of gender-related issues in party competition (Weeks, 2022) and increases the involvement of women in grassroots campaigning, as female candidates are more likely to mobilize gendered networks (Goyal, 2024). These dynamics can enhance the political efficacy of female voters (Atkeson and Carrillo, 2007; Stauffer, 2021), leading to higher turnout among previously disengaged women (De Paola et al., 2014) and shifts in their partisan preferences. Similar shifts may also occur among men, either as acts of solidarity with women or as backlash against equity reforms.

When new voters, or voters who defect from their original party preference, disproportionately back parties with more socially inclusive slates, such behavioral changes indirectly benefit candidates from socio-economically disadvantaged backgrounds. Crucially, this implication holds regardless of whether the new or defecting voters are women or men. In this view, quotas do not merely affect candidate supply; they also reshape the electorate, at times creating incentives for parties to support

more socially diverse candidates.[6] Such demand-side pathways are especially likely to operate under conditions of party system volatility, where weak partisan attachments heighten the electoral consequences of shifts in turnout and party preference. This framework informs three core expectations that I evaluate in the empirical analysis:

H1: Gender quotas increase the descriptive representation of socio-economically marginalized women in legislatures.

H2: Gender quotas increase voter turnout.

H3: Gender quotas shift electoral outcomes in favor of parties with more socially inclusive candidate pools.

3 Research Design

To evaluate my research questions, I analyze how the country-wide enactment of gender quotas affected the class composition of Italian municipal councils between 1985 and 2008. This within-country design offers three key advantages. First, the national-level imposition of the quota minimizes concerns about endogeneity, as municipalities could not self-select into treatment. Second, the large number of municipalities in Italy enables the use of large-N statistical methods ($N = 7904$).[7] Third, I analyze the effects of a temporary quota, which allows me to estimate both short- and long-run changes in intersectional representation—including periods after the quota was repealed. Before presenting the empirical strategy, I outline the institutional context of the quota reform and the electoral rules that shape political competition in Italian local government.

[6] One might speculate that these dynamics also disproportionately raise the candidacy propensity of working-class women. There are, however, no published studies that substantiate this claim. On the contrary, several studies suggest the opposite: that gender quotas tend to bolster candidacy among women with higher socio-economic status (Besley et al., 2017; Karekurve-Ramachandra and Lee, 2020).

[7] This figure holds in 2024, but has remained stable since the early 1980s.

3.1 Institutional and Policy Context

The 'Second Republic' is a broad term used to describe Italy's contemporary electoral system, which emerged from a wave of institutional reform in the early 1990s. Most changes to electoral rules were enacted as part of a 1993 reform package, passed in the wake of a far-reaching political corruption scandal known as *Mani Pulite*, or 'Clean Hands' (Donovan, 1995). The scandal centered on revelations that politicians from Italy's dominant postwar parties—Democrazia Cristiana (DC) and Partito Socialista Italiano (PSI)—had accepted widespread bribes. Beyond triggering legal reforms, the crisis effectively dismantled the party system that had governed Italy since World War II, as corruption-afflicted parties from both the left and right collapsed in its wake (Katz, 1996).

Among the reforms introduced was a temporary gender quota for municipal elections—the first national policy aimed explicitly at improving women's inclusion in elected office. While a feminist movement had grown in Italy since the 1970s, it was largely 'anti-institutionalist' in orientation (Campus, 2016; Willson, 2010) and had made only modest efforts to expand women's presence within formal party politics prior to this period (Guadagnini, 2005).

The quota mandated that no more than two-thirds of a party's candidates could be of the same gender. It remained in effect until late 1995, when Italy's Constitutional Court ruled it unconstitutional. As a result, only municipalities that held elections between April 1993 and September 1995 were subject to the quota. A subset of 317 municipalities did not hold elections during this window, due to earlier government breakdowns that had pushed them off-cycle. Since election timing was fixed and the Court's decision was unpredictable, self-selection into quota exposure is unlikely.

This variation in electoral timing forms the basis of a difference-in-differences (DiD) design, in which I compare changes in representational outcomes between municipalities that held elections under the quota and those that did not, before and after the reform. Figure 1 illustrates the geographical distribution of the two groups.[8]

[8] Municipalities that underwent mergers or were governed by delegated commissioners are excluded, as are those in the German-speaking region of Trentino-Alto Adige, which was subject to ethno-linguistic quotas that superseded the national law.

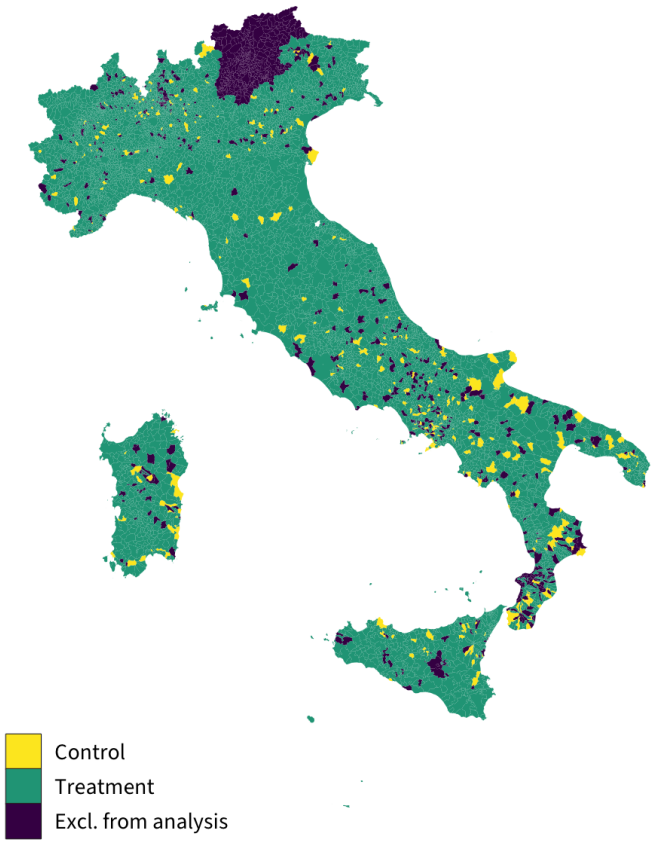


Figure 1: Sampled municipalities, by treatment assignment
Notes: Map based on municipality borders of 2001 (N = 7353).

3.2 Electoral System and Local Governance

Italy has approximately 8,000 municipalities, which together control around 10% of total public spending. Municipal governments are responsible for a broad range of policy areas, including land zoning, water and waste management, local transportation, cultural programming, social housing, and early childhood education. Local governance is divided between an elected mayor, who appoints and leads the executive committee, and a municipal council, which holds legislative authority.

Mayoral and council elections are institutionally linked. In municipalities with 15,000 inhabitants or fewer—92% of those in my sample—two-thirds of council seats are automatically awarded to the party list of the winning mayoral candidate. The remaining third are distributed proportionally across other party lists using the D'Hondt formula. In larger municipalities, the mayoral list receives 60% of council seats, and voters may split their vote between different parties for mayor and council. In practice, however, incentives to split votes remain limited, as doing so weakens the governing capacity of the preferred mayoral candidate.

3.3 Data

This study draws on annual data compiled by the Italian Ministry of Internal Affairs between 1986 and 2008.[9] These data include information on the gender, age, birthplace, party affiliation, educational attainment, and occupational background of all Italian municipal councilors. I focus on the 1985–2008 period, as gender quotas—both mandatory and voluntary—were reintroduced after 2008, raising concerns about comparability in later years.[10] Between 1993 and 2008, municipal elections were held every four years, which ensures that all municipalities held at least two elections following the quota reform.

Information on candidate gender is complete; occupational data has a slightly higher non-response rate of 6%. The final dataset includes 551,520 unique council posts across 35,178 municipal elections. Council size varies from 1 to 80 seats, depending on local population. Only 33 of these elections produced councils with fewer than five members (<0.1% of cases).

3.3.1 Descriptive Gender-Class Representation

To capture the social class of legislators, I code a series of dummy variables based on their occupational titles, as recorded by the Ministry. I begin by assigning each occupation to one of four social classes, using a schema informed by Wright's (1997) conceptual framework and operationalized by Oesch (2006). Occupation-based class measurement offers several advantages. First, occupational identity tends to be stable over time, unlike more transient indicators such as income. Second, occupations reveal important information about individuals' workplace conditions, authority structures, and exposure to labor market risks—all of which are strongly linked to class-based patterns in political behavior and attitudes (Evans, 2000; Manza and Brooks, 2008; Rennwald, 2020; Carnes and Lupu, 2015).

Oesch's schema distinguishes five classes based on four key dimensions of socio-economic vulnerability: (1) relation to the means of production, (2) level of authority, (3) skill scarcity, and (4) number of employees supervised (2006, pp. 14–15). These include the higher-grade service class, the lower-grade service class, small business owners, skilled workers, and unskilled workers. Due to data

[9] Some records extend back to 1985 or 1983, as many councils elected in 1986 took office earlier. Systematic data from before 1985 is not available.

[10] In 2009, the *Partito Democratico* became the first major party in Italy to adopt voluntary gender quotas. Mandatory quotas were reintroduced in 2012, requiring party lists to include at least 40% women (IDEA, 2021).

limitations, I make two adjustments to this framework. First, the available data do not allow me to distinguish between business owners and employed workers. Second, I collapse skilled and unskilled workers into a single category, labeled 'working class', as both subgroups are heavily underrepresented in local office.[11]

Table 3 provides examples of how raw occupational entries are coded into this four-category scheme, alongside their corresponding Oesch classification. While the categories are hierarchically ordered, they differ somewhat from traditional 'upper', 'middle', and 'working' class divisions. In most political sociology, the upper class refers to rentiers whose income derives largely from assets; this group is rare in local elections. I therefore refer to the higher-grade service class as the *upper-middle class*, the lower-grade service class as the *lower-middle class*, and workers simply as the *working class*. I adopt these terms throughout the analysis for clarity and consistency.

Table 3: Social class coding based on occupational belonging

Occupations ***(examples from raw data)***	**Oesch (2006) coding**	**Four cat. coding**
Company directors, Senior civil servants, Lawyers, Engineers, Financial analysts, University professors, Journalists	Higher grade service class	Upper middle class
Hospitality managers, Police inspectors, Technicians, Sales agents, Designers, Librarians and curators	Lower grade service class	Lower middle class
Bank-tellers, Travel consultants, Legal secretaries, Customer service reps, Fire fighters, Craftsmen	Skilled workers	Working class
Drivers, Machine operators, Assemblers, Farm workers, Waiters, Hairdressers, Cleaners	Unskilled workers	Working class
Homemakers, students, retirees, unemployed	—	Inactive in labor market

Having classified legislators by social class, I combine these categories with a binary gender indicator to create eight intersectional class-gender groups. I focus my analysis on individuals with clearly identifiable class positions; those who were inactive in the labor market—such as students, retirees, homemakers, or the unemployed—are included when constructing denominators but not analyzed as a separate group.[12]

[11] Appendix A.1 provides additional elaboration.

[12] Results for labor market inactive subgroups are reported in appendix C.2.

My main outcome variables capture the share of elected councilors from each class-gender group, measured at the municipality-election level. For each election year, I compute the proportion of seats held by, for example, working-class women, lower-middle-class men, and so on. Figure 2 plots the annual average descriptive representation of these groups over time. The figure reveals a persistent under-representation of both women and working-class individuals in municipal politics throughout the 1990s and early 2000s. In 2008—the final year of the sample—women across all classes held just 15% of council seats, despite comprising roughly 40% of the labor force. Similarly, workers reached a peak of 17% representation in 1992, far below their estimated 40% share of the active workforce.[13]

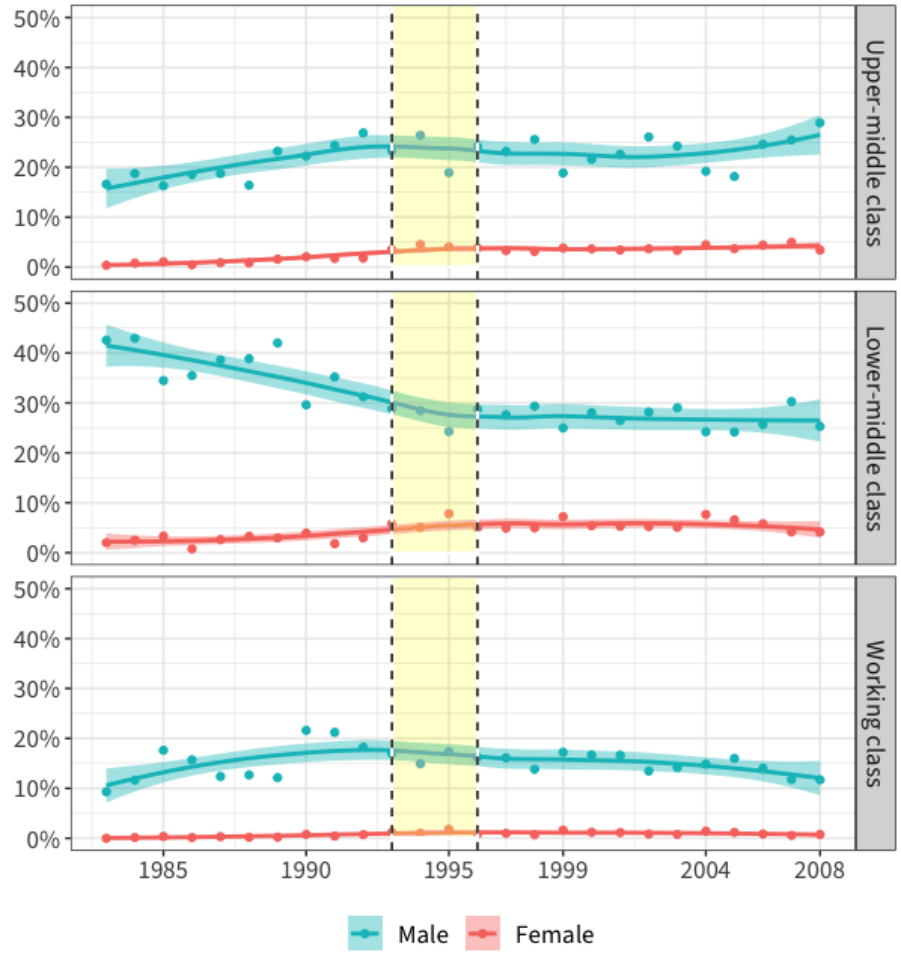


Figure 2: Annual average class-gender shares of Italian municipal councils
Notes: As the number of elections varies noticeably per year, trend lines indicate smoothed conditional averages, attained by means of loess estimation. Bands indicate 95% confidence intervals. Calendar years denoted on x-axis are regular cycle election years (electoral terms were temporarily shortened from five to four years between 1993 to 1998). Period in between dashed lines indicates time during which gender quotas were in force.

[13] Author's own calculation based on ILO data (2022).

Voter Turnout

To investigate the electoral consequences of quota exposure, I examine whether gender quotas affected voter turnout in local elections. Turnout is measured as the fraction of registered voters who cast a ballot in a given municipal election. To compile this outcome, I rely on archival records from the Italian Ministry of Interior Affairs. The Ministry has systematically collected municipal turnout data since 1993, covering 91% of quota-round elections in my sample (N = 6629). For earlier years, I digitize turnout figures from official print volumes, allowing me to recover data for over 99% of the elections immediately preceding quota adoption (N = 7298).

Legislature Seat Shares by Party Ideology

In addition to outcomes pertaining to descriptive representation and voter turnout, I construct outcome variables that capture the ideological composition of municipal legislatures. Because official vote shares are not consistently available for the pre-1993 period, I use the seat shares of ideological party blocs—directly observable from council composition—as a proxy for local electoral preferences.[14]

Rather than coding individual parties, I classify party lists into broader ideological blocs, reflecting the nature of Italian municipal elections. These elections are often contested by coalitions of parties that share a mayoral candidate, and many party lists are locally specific. In total, the dataset includes 773 unique party list or coalition names. Where possible, I assign these lists to ideological families using the ParlGov classification scheme (Döring et al., 2022).[15] However, this approach does not cover the large number of Civic Lists—locally organized party formations often unaffiliated with national politics. To classify Civic Lists, I apply a dictionary-based method: if a list name contains ideological keywords (e.g., 'Workers', 'Radical', 'Monarchist'), I assign it to the corresponding bloc. If a Civic List is part of a coalition with a national party, I adopt the ideological position of that party. In cross-bloc coalitions, I code the list as 'Left-Right Coalition'.

This procedure yields eight ideological blocs: (a) Left, (b) Conservative, (c) Liberal, (d) Green, (e) Left-Right Coalition, (f) Special Issue, (g) Right-Populist, and (h) Civic Lists. Additional details and keyword dictionaries used in the coding process are provided in appendix A.5.

[14] While the electoral formula used to allocate seats was also reformed in 1993 for some municipalities, both treated and control units were subject to these changes within each election round.

[15] ParlGov is preferred here due to its broader coverage of parties historically active in national-level politics. Other sources, such as the Chapel Hill Expert Survey or the Manifesto Project, cover a narrower range of parties.

3.4 Empirical Strategy

To estimate the effects of quota exposure, I use a difference-in-differences (DiD) design adapted from De Paola et al. (2010) and Baltrunaite et al. (2014). I update this approach to account for differences in legislature size and to evaluate how treatment effects vary across election rounds.

The baseline specification is:

$$R_{ijt} = \beta_1 QuotaMunicipality_j + \beta_2 ElectionRound_t + \beta_3 (QuotaMunicipality_j \times ElectionRound_t) + \beta_4 X_{jt} + \epsilon_{jt} \quad (1)$$

Here, R_{ijt} denotes the proportion of council seats held by class-gender group i in municipality j during election round t. $QuotaMunicipality_j$ is a binary indicator for whether municipality j held an election while the quota was in force. Municipalities that held elections between April 1993 and September 1995 are coded as treated; those whose next election occurred after the quota's repeal serve as controls. $ElectionRound_t$ is a categorical variable denoting the election round, centered on the 1993–1997 window (i.e. the quota election round itself).

I use election round—as opposed to election year—as the time dimension, since treatment is assigned based on election timing. This setup avoids conflating treatment with broader reforms introduced in 1993, which affected municipalities in both groups. The interaction term captures the group-time average treatment effect (β_3), comparing outcomes across treated and control municipalities, before and after quota implementation.

The vector X_{jt} includes municipal-level controls: council size, population, population squared, unemployment rate, and tertiary education rate. These are indexed to the closest available census year. Standard errors are clustered at the municipality level.

For the mechanism tests, I restrict the sample to two election rounds: the quota round and the immediately preceding one. In this two-period setup, I replace the election round fixed effects with a binary $PostQuota_t$ indicator. Outcomes include (a) voter turnout and (b) the share of council seats held by each ideological bloc.

3.5 Testing the Parallel Trends Assumption

The credibility of the DiD strategy hinges on the assumption that treated and control municipalities would have followed similar representational trends in the absence of quota exposure. To assess this, I plot pre-treatment trends in the class-gender composition of municipal councils, comparing group-specific means across treated and control units.[16]

Figure 3 shows adjusted trends for women and men, respectively, with municipality-level covariates held constant. Across nearly all subgroups, pre-quota trajectories are closely aligned. The only partial exception is for upper-middle-class men, who are slightly less represented in municipalities that later receive treatment. However, even in this case, trends unfold in parallel prior to the 1993 reform. These visual diagnostics support the identifying assumption of the DiD design: in the absence of the quota, treated and control municipalities were likely to evolve similarly with respect to class-gender representation.[17]

[16] Systematic data on municipal council composition are only available from 1985 onward, which precludes tests over a longer pre-treatment period.

[17] Both pre-treatment trends and main results are substantively similar in estimations without statistical controls (not reported), but I opt for their inclusion as they improve pre-treatment outcomes in terms of levels.

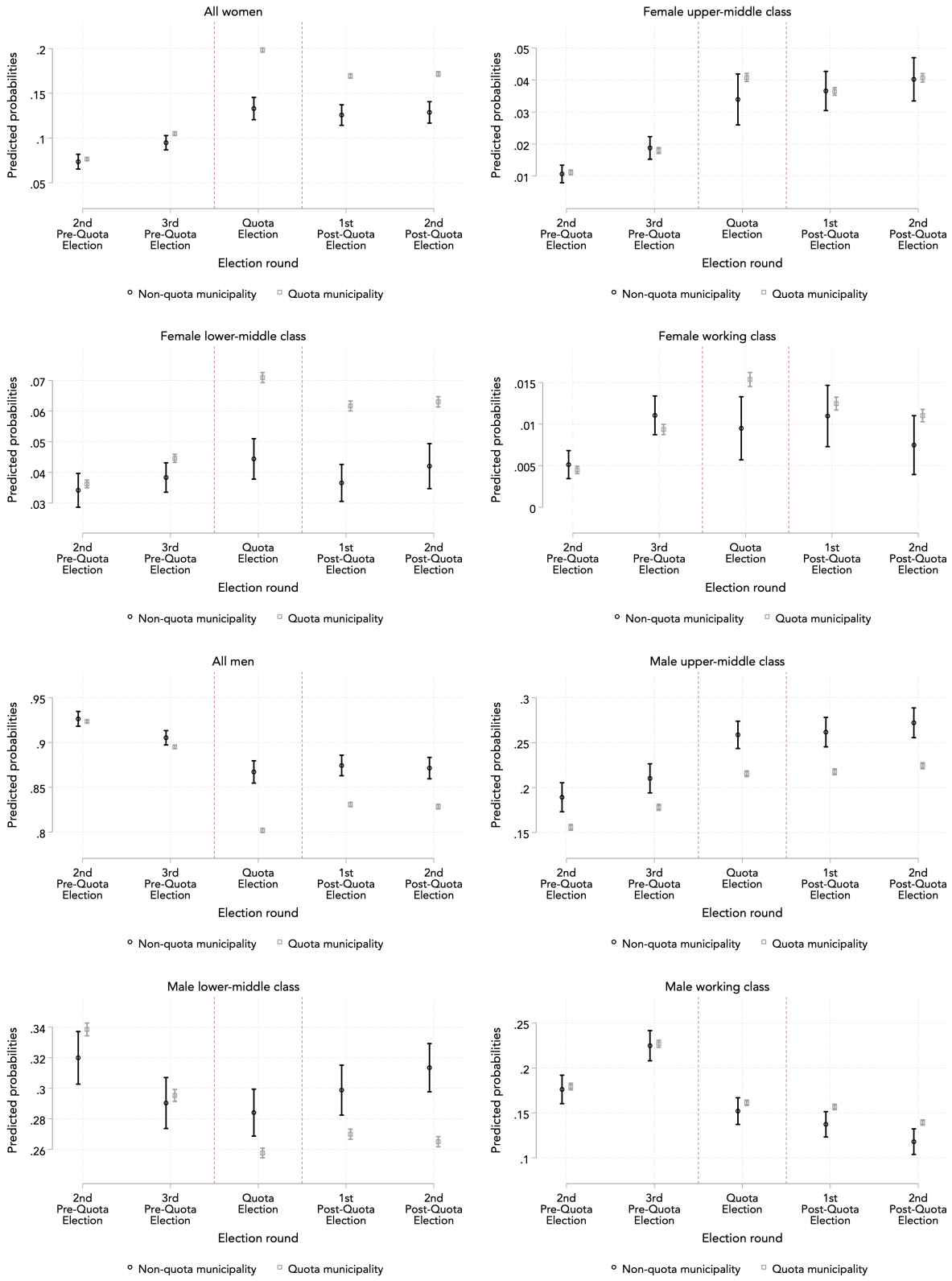


Figure 3: Over-time trends in subgroup representation in municipal councils

Notes: Group means adjusted for municipality-level controls. Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots. Standard errors are clustered at the municipality level.

Quota Effects on Intersectional Gender-Class Representation

How did gender quotas affect the intersectional composition of municipal councils? Figure 4 presents DiD estimates of the quota's impact on women's representation by social class. Prior to the reform, women—regardless of class—were equally likely to win council seats in treated and control municipalities. There were no statistically significant differences for upper- or lower-middle-class women, and only a small disparity for working-class women, who were 0.22 percentage points less likely to be elected in municipalities that would later implement quotas ($p < 0.05$). However, pre-treatment trends were parallel across all subgroups.

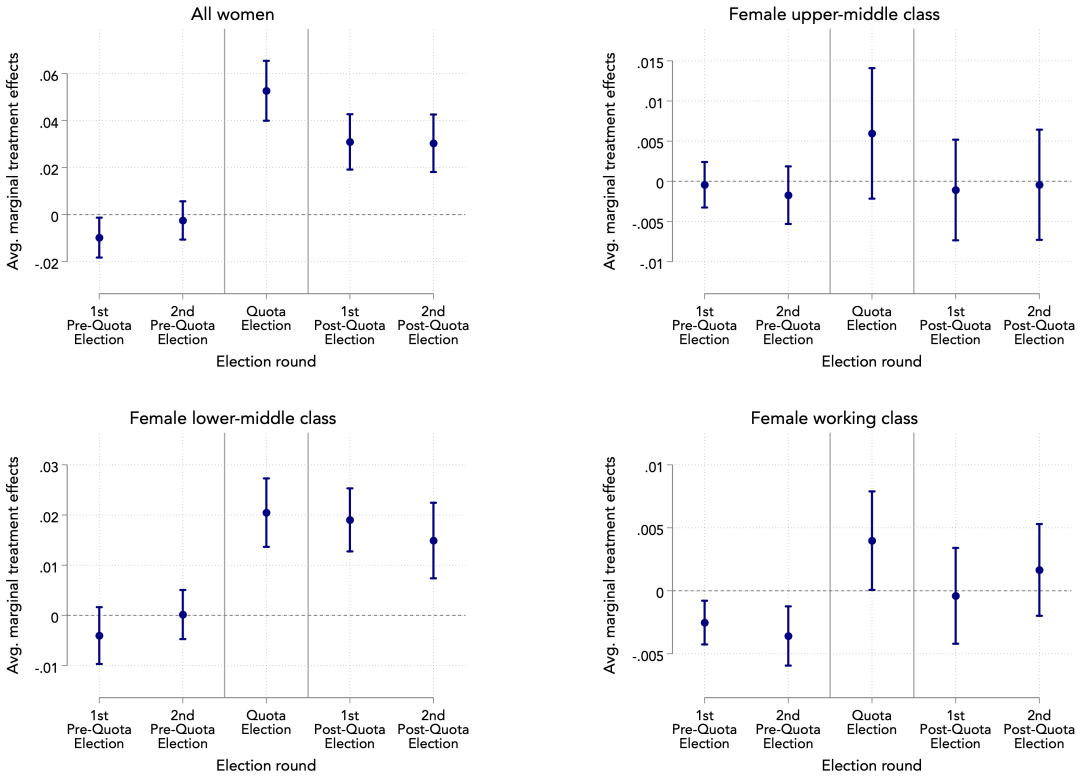


Figure 4: Treatment effects on female representation in municipal councils, by social class
Note: Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots, to better display effect sizes and confidence intervals. All model specifications include municipal-level controls. Standard errors are clustered at the municipality level. Regression tables in appendix C.1.

Quota adoption led to significant gains for lower-middle-class women, whose probability of election increased by 2.0 percentage points ($p < 0.01$). This effect remained detectable in subsequent elections. Working-class women also saw a smaller but statistically significant increase of 0.76 percentage points during the quota round ($p < 0.05$). While modest in absolute terms, these gains represent sharp relative increases given low baseline representation: in 1990, lower-middle-class women made

up just 3.9% of councilors, and working-class women only 0.8%. The quota thus produced increases of 53% and 100%, respectively.

For men, Figure 5 shows that working-class representation remained stable following quota implementation. The only group to experience a measurable decline was lower-middle-class men, whose seat share dropped by 3 percentage points ($p < 0.05$). This suggests that female entrants from the same social class may have displaced these men, while working-class male representation remained unaffected.

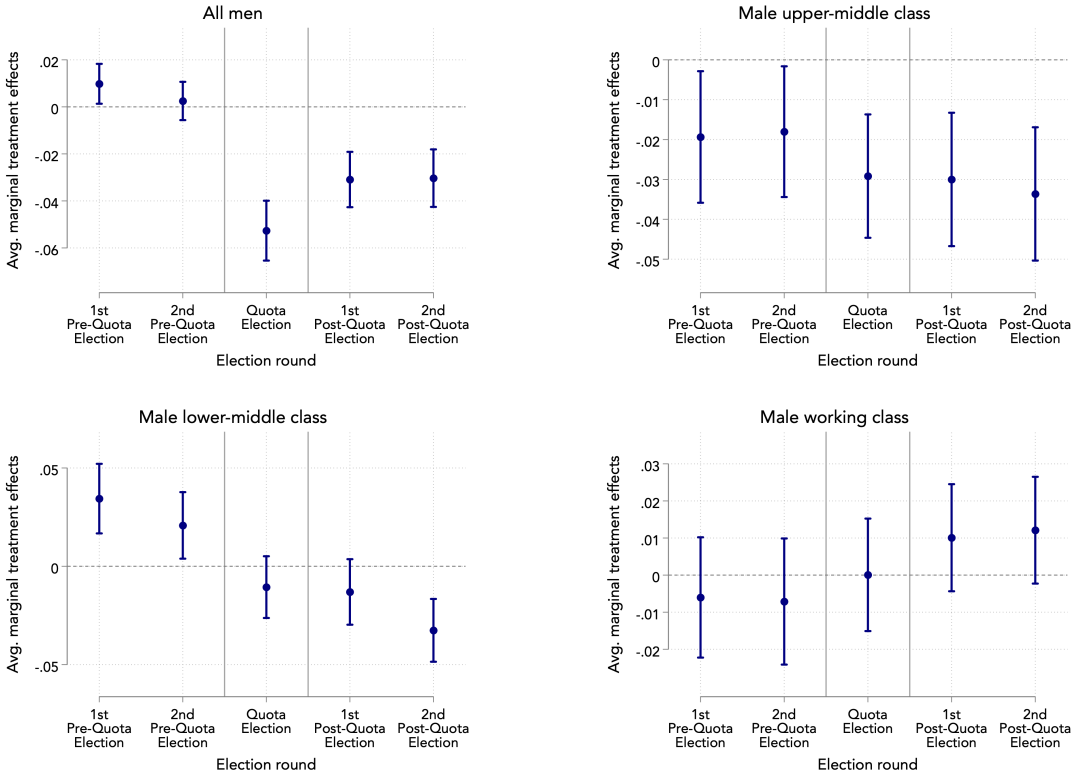


Figure 5: Treatment effects on male representation in municipal councils, by social class
Note: Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots, to better display effect sizes and confidence intervals. All model specifications include municipal-level controls. Standard errors are clustered at the municipality level. Regression tables in appendix C.1.

Additional tests confirm that parallel trends also hold for councilors who were inactive in the labor market, such as pensioners (appendix C.2). Among these, female representation increased after quota adoption, while male representation remained unchanged. The only remaining uncertainty concerns whether lower- or upper-middle-class men were the principal losers of the reform; but even if these groups were more affected than the estimates suggest, the broader picture is consistent with the main hypothesis: the quota had an equalizing effect on descriptive class representation.

To probe the robustness of these results, I rerun the analyses using a conventional two-way fixed-effects (TWFE) design with municipality fixed effects. This approach is problematic in the first instance, given the long time frame analyzed and the temporary nature of the quota: effects naturally weaken after quota abolition, violating the TWFE assumption of time-invariant treatment effects.[18] To assess the robustness of the *direct* quota effect, however, I restrict TWFE estimations to quota-round elections in the post-quota period, which yields average treatment effects for the treated (ATET). Results, presented in figure 6 and contrasted with group-time average effects from two-period models (controlling for election round variation), remain highly consistent with the baseline estimates.[19]

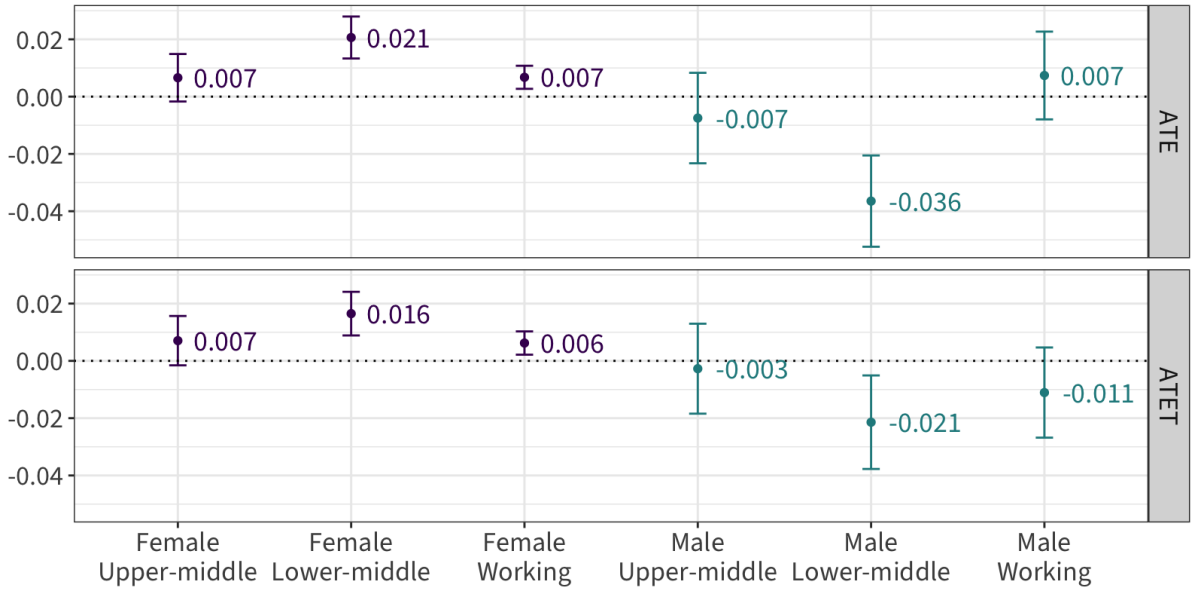


Figure 6: Group-time ATEs vs. ATETs
Note: ATE estimates adjusted for covariates and election round effects; ATETs additionally account for within-municipality fixed effects. Bars indicate 95% confidence intervals. N = 20,584. Regression tables in C.3.

4 Demand-Side Mechanism Tests

The main analysis shows that gender quotas most strongly benefited women from lower social classes—a result that runs counter to much of the existing literature, which often emphasizes the regressive social effects of quota adoption. Particularly uncommon is the finding that the policy expanded political op-

[18] See e.g. Callaway and Sant'Anna (2021) and De Chaisemartin and d'Haultfoeuille (2020).
[19] Regression tables and joint significance tests for pre-treatment trends are presented in appendix C.3.

portunities for working-class women, albeit only during the quota election round itself. I hypothesize that the overall pattern reflects changes in voter behavior—specifically, a demand-side shift triggered by quota-induced changes in candidate pools. To probe this mechanism, I therefore examine two observable implications: quota effects on voter turnout (H2) and on party support (H3). This analysis evaluates whether quota adoption was associated with systematic demand-side changes in electoral participation and party competition, which correspond to the observed gains for lower-middle- and working-class women.

Voter Turnout. To test whether quota exposure influenced overall voter participation, I replicate the analysis of De Paola, Scoppa and De Benedetto (2014) on the same reform. Like the authors, I estimate the baseline model, along with specifications that include year fixed effects and a control for off-cycle municipalities (i.e., those that experienced early council collapses prior to quota installation).[20] Results are presented in Figure 7.

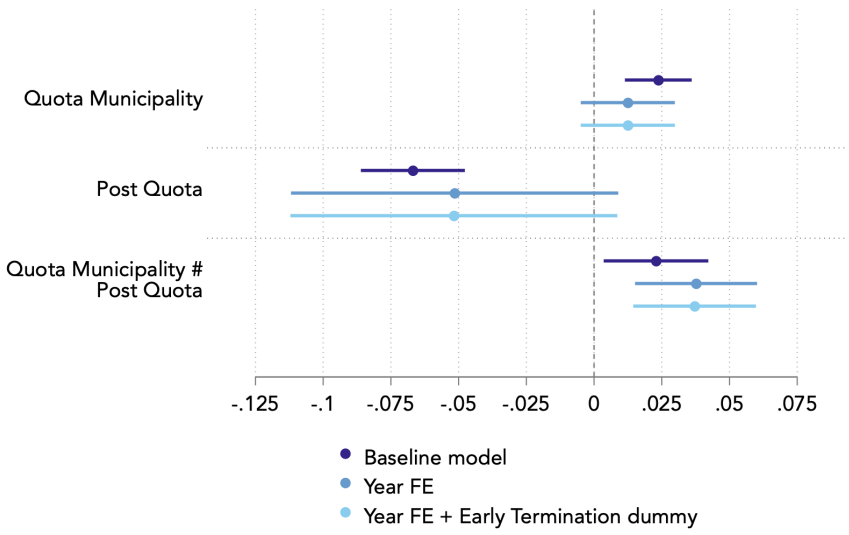


Figure 7: Average quota effects on turnout in local elections
Notes: All coefficients indicate changes between pre-quota and quota election rounds, adjusting for municipality-varying controls. Standard errors are clustered at the municipality level. Bands display 95% confidence intervals. Regression table in C.10.

The *Quota Municipality* coefficient shows that pre-treatment turnout levels did not differ significantly across treatment groups once year-specific variation is accounted for. As expected, the *Post Quota* coefficient captures a general decline in turnout during the quota round (5.2–6.7 percentage

[20] As none of the sampled local elections coincided with national elections, observed turnout shifts cannot be attributed to national-level coattail effects.

points, $p < 0.1$), consistent with the broader post-Mani Pulite climate of disengagement. Crucially, the interaction term *Quota Municipality* × *Post Quota* indicates that quotas mitigated this decline: treated municipalities experienced a statistically significant increase in turnout of 2.3–3.8 percentage points. These effects are nontrivial in a context where many mayoral elections were decided by narrow margins.[21]

Although gender-disaggregated turnout data are unavailable for the purposes of this study, the observed pattern is consistent with theoretical expectations: by diversifying candidate pools, quotas reshaped the demand side of politics and mobilized voters who might otherwise have abstained. Supporting this interpretation, evidence from De Paola et al. (2014) indicates that such mobilization is not confined to one gender; rather, quota adoption increased voting propensities among both men and women, albeit more strongly among women. Whether this response reflects the behavior of newly mobilized women, men reacting in solidarity, or even backlash dynamics, the key implication is that quota exposure counteracted apathy and broadened political participation during a period of institutional crisis.

Party Composition. Having established effects on voter participation, I turn to whether quotas also altered the distribution of party support. Figure 8 presents quota effects on the seat shares of ideological party blocs. Right-wing parties such as Lega Nord gained modestly (2 percentage points, $p < 0.05$), but the most pronounced effects were for Civic Lists—locally rooted, centrist parties with weak national ties. Civic Lists gained 11 percentage points ($p < 0.01$), primarily at the expense of mainstream Left and Conservative parties (both $p < 0.01$).

This realignment is consistent with known trends in female voter preferences at the time. The collapse of the Christian Democrats (DC) and the Italian Socialist Party (PSI) in the wake of the Tangentopoli scandal led to widespread defection from mainstream parties across the ideological spectrum. However, women were especially likely to shift their support to smaller, locally focused parties in the local elections that directly followed (Bull, 1997). Two factors shaped this pattern. First, older generations of Italian women—often religious and with low labor market attachment—exhibited a traditional gender gap, tending to vote more conservatively than men.[22] Second, women's repre-

[21] In the quota round, a quarter of municipalities with single-round elections chose a mayor by fewer than five percentage points.

[22] See e.g. Corbetta and Cavazza (2008), Emmenegger and Manow (2014), Giger (2009), Inglehart and Norris (2003) and Morgan (2013).

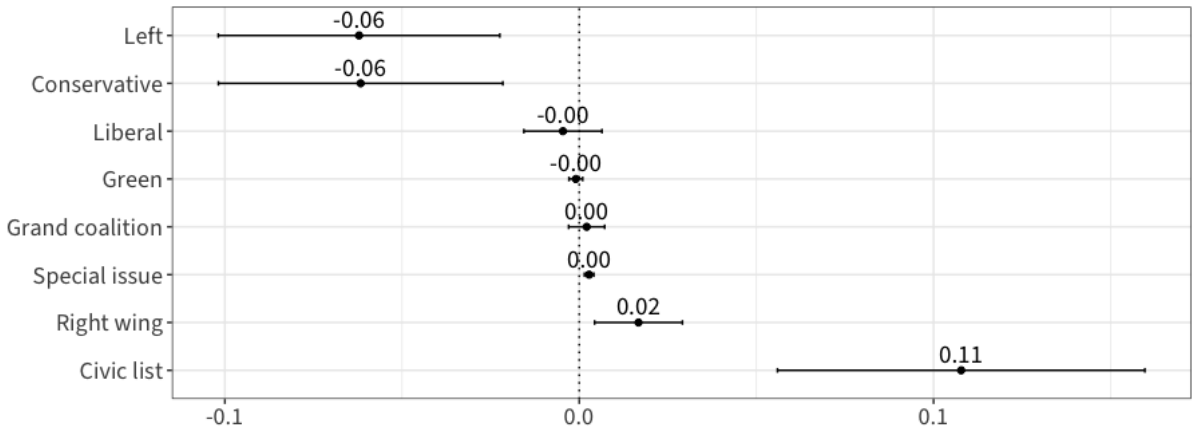


Figure 8: Quota effects on legislature seat shares, by party ideological bloc
Notes: Coefficients indicate changes between pre-quota and quota election rounds, adjusting for municipality-varying controls. Standard errors are clustered at the municipality level. Bands indicate 95% confidence intervals. Regression tables in C.4.

sentation—whether numerical or substantive—was largely absent from party platforms across the spectrum prior to the 1990s, contributing to low partisan loyalty among female voters (Bull, 1997).

To clarify the relationship between civic-list electoral success and the election of women from lower social classes, Figure 9 reports the share of female councillors, by social class and election round, who entered office via Civic Lists. In elections directly prior to quota adoption, most women entered office through mainstream parties, although working-class women were already more likely than women from other social classes to enter office via Civic Lists. A similar class gradient is observed among male legislators (see appendix B.2). Following quota adoption, this class gradient persists; however, Civic List entry becomes more prevalent across all female social classes, accounting for a majority of women's successful candidacies regardless of class background.

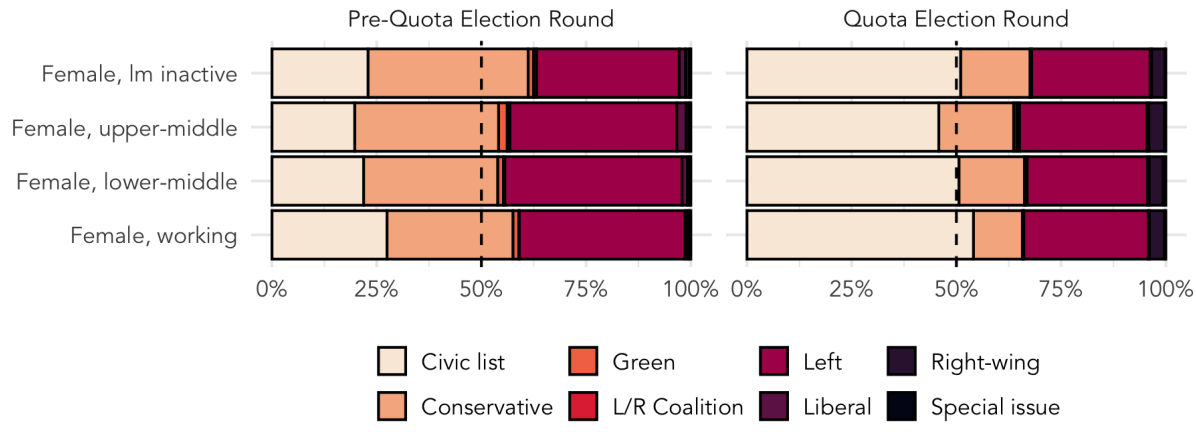


Figure 9: Party affiliation of female legislators, by social class and election round.

In light of the previously documented increase in Civic List seat shares under quota exposure, this pattern suggests that quotas broadened access through an existing entry channel rather than inducing class- or party-specific shifts in candidate selection. Evidence on the social-class composition of party delegations further indicates that lower-middle- and working-class women, as well as working-class men, constitute a larger share of representatives within Civic Lists than within mainstream parties, pointing to greater social diversity within these lists (see appendix B.2).

Taken together, these findings support the argument that gender quotas can activate demand-side dynamics in political selection. By altering the structure of electoral competition, quotas appear to have mitigated political disengagement in the aftermath of institutional crisis, while simultaneously favoring party vehicles with lower barriers to entry. The gains observed for lower-middle and working class women—particularly through Civic Lists—are therefore arguably best understood as arising from changes in voter participation and party competition, rather than from class-specific targeting in candidate selection.

5 Robustness and Alternative Explanations

In a final step, I evaluate the credibility of my primary findings by conducting further robustness checks and probing for potential confounding explanations.

Robustness to alternative model specifications. To ensure that findings are not sensitive to modeling assumptions, I re-estimate the main models under alternative specifications. First, I include NUTS-2 (regional) or, alternatively, NUTS-3 (provincial) fixed effects to account for unobserved geographical variation. In a separate set of estimations, I also vary the census-based timing of control variables, since the quota round occurred midway between two decennial censuses. None of these adjustments substantively alter the main results.[23]

Alternative outcome: education. To verify that the quota's effects on social class representation are not specific to my occupation-based coding, I re-estimate models using educational attainment as an alternative proxy for socio-economic status. The results—distinguishing between university and non-university educated candidates, including labor market inactive groups—mirror those from the

[23] Full regression output is available upon request.

main analysis (see appendix D.2).

Controlling for other electoral reforms. A key identifying assumption is that changes in class-gender representation stem from quota exposure, rather than other electoral reforms implemented during the same period. The 1993 reform package also introduced two changes: (1) a shift from proportional to hybrid majoritarian electoral rules in municipalities over 5,000 residents, and (2) the introduction of two-stage mayoral elections in municipalities over 15,000. Although treatment assignment is based on election timing—exogenous to population size—I include dummy variables indicating exposure to these rule changes. The inclusion of these controls does not meaningfully affect the estimated treatment effects (see appendix D.1).

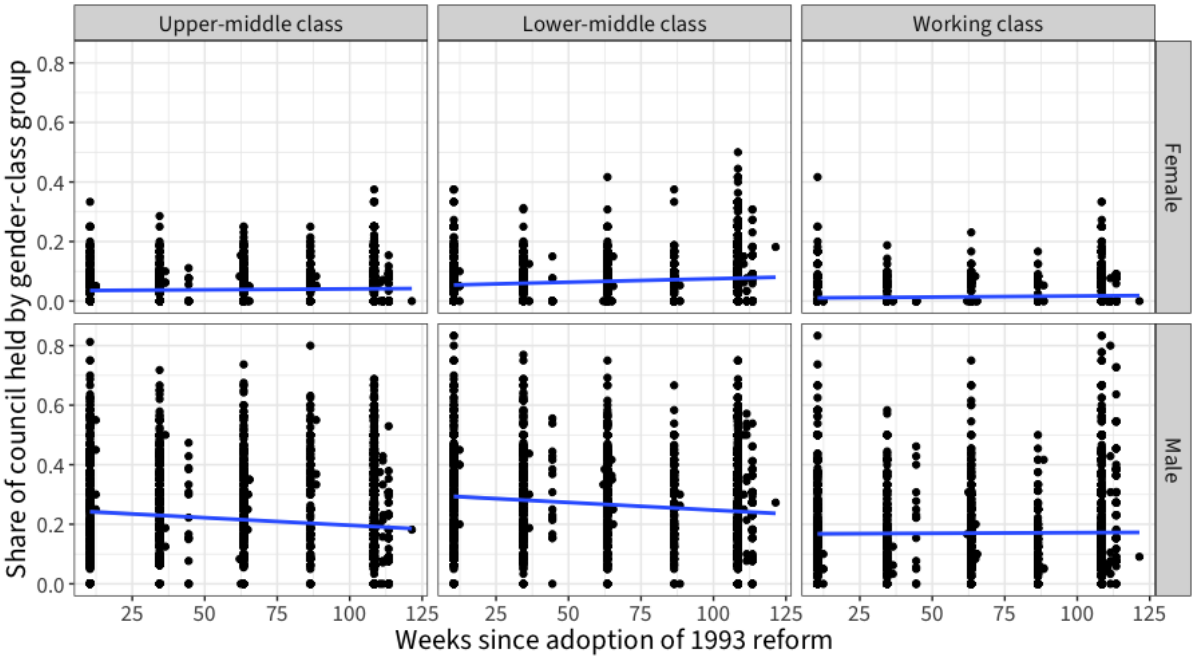
Excluding elections around the 1994 national shock. Another concern is that quota effects may be confounded by the broader political upheaval surrounding the March 1994 parliamentary election, which marked the collapse of Italy's postwar party system and the rise of Forza Italia. While all municipalities were exposed to these national dynamics, some held local elections in close proximity to the general election, raising the possibility that short-term anti-elite sentiment influenced election results. To assess this, I exclude all municipalities with local elections occurring within six months before or after the 1994 national vote. The estimated quota effects remain unchanged, indicating that results are not driven by contemporaneous political realignment (see appendix D.7).

Anti-elite sorting. Finally, I test whether the estimated effects reflect sorting dynamics linked to the Tangentopoli scandal, which disproportionately undermined Italy's traditional political elite. If anti-elite sentiment was strongest immediately after the scandal erupted (i.e. immediately after quota adoption), treated municipalities might simply be those more inclined to vote out middle-class men. To probe this, I restrict the sample to councils elected under the quota regime and regress the seat share of each gender–class group on the number of weeks since the 1993 reform.

Contrary to the sorting hypothesis, women of all classes gain the further elections occur from the crisis, while working-class men's representation remains stable or slightly rises (see Figure 10). Upper- and lower-middle-class men, by contrast, fare worse in later elections. Taken together, these patterns imply that any sorting bias would dampen—rather than inflate—estimated quota effects, suggesting that my results should be interpreted as lower-bound estimates. Repeating the analysis using both voter turnout and party-level seat share outcomes yields consistent patterns. Across all outcomes

examined in mechanism tests, time trends further support the interpretation that the estimated effects are lower bounds rather than artifacts of immediate anti-elite backlash (see appendix D.4).

Figure 10: Over-time change in probability of attaining municipal council by class-gender group, April 1993- Sept. 1995



Note: Results are robust to the inclusion of municipality-level control variables. Full regression results in appendix D.4.

Discussion

This paper shows that gender quotas can expand political opportunity for lower-middle and working-class women—without displacing working-class men. Instead, the group most likely to lose ground under quota-induced inclusion is lower-middle-class men. Mechanism tests suggest these outcomes are driven by demand-side shifts: quotas affect the electoral success of parties, leading to changes in voter turnout and legislative party composition.

As with all single-country studies, these findings raise questions of generalizability. Comparative work is needed to identify the scope conditions under which quotas reshape the social profile of elected officials. A key feature of the Italian case is its institutional context: the collapse of mainstream parties in the wake of the *Mani Pulite* scandal produced a period of exceptional party system volatility. While this volatility affected the entire sample—limiting concerns about treatment bundling—it likely amplified the downstream effects of quotas by weakening partisan attachments. This may help

explain why the results diverge from prior studies conducted in more stable, party-centered systems. Importantly, this dynamic is not unique to Italy. Similar conditions have been documented in Latin America—for example, in Peru and Argentina during periods of party system collapse and voter realignment (Lupu and Stokes, 2010; Roberts and Wibbels, 1999)—and in parts of Asia, such as the Philippines and Indonesia (Croissant and Völkel, 2012). In such contexts, quotas may be more likely to generate downstream effects on voter turnout and electoral choice.

Second, further research should clarify when and how quotas mobilize voters. Prior studies suggest quota effects on participation may decline over time (e.g. Gilardi, 2015), especially where female voters are already highly engaged. In such cases, null findings may reflect ceiling effects rather than genuine absence of demand-side mechanisms.

Third, while this study focuses on local elections, the findings carry broader implications for national politics. High-income individuals are more likely to cluster in large urban areas, making the descriptive representation of social class comparatively more feasible at the local level (Folke and Rickne, 2024). However, in many political systems—including Italy—local office serves as a key recruitment channel for higher-level positions (Profeta and Woodhouse, 2021). At the same time, women's descriptive representation at the local level often lags behind that in national politics (Smith et al., 2025; Holman, 2017), reflecting factors such as conservative recruitment norms and the smaller size of local legislatures (Crowder-Meyer, 2013). In this sense, gender quotas that diversify local councils may also shape the long-term composition of higher-level legislatures—and even cabinets[24]—provided that parties draw from these more inclusive pools.

Finally, while the theoretical framework emphasizes demand-side mechanisms, this study focuses on aggregate outcomes such as turnout and party composition. I do not directly measure changes in party strategy, platform content, or voter psychology. As such, the results should be interpreted as consistent with, but not conclusive of, the proposed causal pathways. Future work using individual-level or experimental data could help illuminate the micro-foundations behind the observed effects.

[24] See e.g. Barnes et al. (2024).

In closing, the findings point to a broader equity rationale for gender quotas. Not all women enter politics on equal footing. Working-class women, in particular, lack the institutional footholds—such as trade unions or elite professional networks—that often facilitate candidacy for men (Crowder-Meyer, 2013). Formal gender quotas may help level that playing field, enhancing access to office for groups who represent a large share of the electorate but remain severely underrepresented in elected institutions.

Competing interests: The author(s) declare none.

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Online appendix

A Data and coding

A.1 Policymaker occupational class

While my coding is similar to Oesch's class schema, it has been modified in three ways. First, I collapse both skilled and unskilled workers into a singular group, simply dubbed 'workers'. Second, I am unable to code the social class of persons with low labor market attachment, as there is no data available on other individual-level variables, such income, marital status and pre-pension occupation. As such, homemakers, pensioners and students constitute a residual category of legislators.

Third, my coding only accounts for three of the four axes of socio-economic vulnerability, as theorized by Oesch. Ministry data on policymaker backgrounds allow me to account for fine-grained sectoral employment (i.e an individual's relation to means of production), individual-level managerial competency (authority) as well as skill-specificity. It does not, however, allow me to directly differentiate between business-owners, the self-employed and employees. That means that e.g. an individual coded to belong to a managerial class may either be employed in a management position by a larger enterprise, or simply act as a manager in a family-owned firm. Equally, individuals in some high-skill service professions—such as architects or accountants—are automatically coded as belonging to the upper grade service class, as opposed to being owners of small firms.

It is likely that firm ownership generates mixed economic vulnerabilities, as firm owners face economic insecurities that stem both from household and business finances. It remains an open question, however, whether their social class should be treated as entirely distinct from those of employees. Some work has shown that the economic policy preferences of business owners overlap significantly with those of the employed (e.g. Ares and Häusermann, 2023). Concerning social policy, for instance, small business owners generally have preferences that are more progressive than the lower grade service class, but less so than skilled and unskilled workers. Regardless of what position one takes in this debate, it is important to stress that my class coding is noticeably more differentiated than that which is commonplace in the literature on descriptive representation and social class.[25]

[25] See e.g Carnes and Lupu (2016); Carnes (2012); Carnes and Lupu (2015); Lassébie (2020).

A.2 Municipality units

Based on information provided by ISTAT, I exclude municipalities that a) engaged in municipal mergers or splits during the time period analyzed; b) were forced into government oversight via the appointment of municipal commissioners; and c) had two elections during the quota election round, as a result of pre-term government collapses (0.9 percent of municipalities in sample).

A.3 Census data

Control variables in the study were coded using census data from the censuses of 1991, 2001 and 2011. Table A.1 presents an overview of coding decisions related to the operationalization of these variables. Results are robust to the use of alternative temporal cut-off points for census merging.

Table A.1: Census data used for municipality-level control variables

Variable operationalization	**Census wave**	**Used for period:**
Population size		
Absolute number of inhabitants	Census 2011	2005-2008
Absolute number of inhabitants	Census 2001	1995-2004
Absolute number of inhabitants	Census 1991	1985-1994
Absolute number of inhabitants	Census 1981	1983-1984
Educational attainment		
Fraction of population aged 9 and above with tertiary degree	Census 2011	2005-2008
Fraction of population aged 6 and above with tertiary degree	Census 2001	1995-2004
Fraction of population aged 6 and above with tertiary degree	Census 1991	1983-1994
Unemployment		
Fraction of population in unemployment, aged 15 and above	Census 2011	2005-2008
Fraction of population in unemployment, aged 15 and above	Census 2001	1995-2004
Fraction of population in unemployment, aged 15 and above	Census 1991	1983-1994

Census data from 2011 and 2001 was extracted from online repositories of Italian statistical authorities. Data from the 1991 is collated from provincial census dossiers.[26]

[26] URL: https://ebiblio.istat.it/SebinaOpac/resource/IST0071571. (Last accessed 26.04.2022).

A.4 Election rounds

The number of unique municipal councils included in sample are displayed in table A.2, split by election year and election round. The number of elections is lower in the first election round as the Ministry of Interior only started collating data in 1987. It is also somewhat declines somewhat over time, as municipalities that engage in mergers/splits in the post-treatment period are dropped only once the territorial reform has been enacted.

Table A.2: Municipal councils, by election year and election round

	Election Round					
Election year	1st pre-quota round	2nd pre-quota round	**Quota round**	1st post-quota round	2nd post-quota round	*Total*
1983	23	0	0	0	0	23
1984	104	1	0	0	0	105
1985	5615	16	0	0	0	5631
1986	43	0	0	0	0	43
1987	63	34	0	0	0	97
1988	84	1046	0	0	0	1130
1989	9	215	0	0	0	224
1990	79	5713	0	0	0	5792
1991	0	93	0	0	0	93
1992	0	180	0	0	0	180
1993	0	1	1474	0	0	1408
1994	0	0	618	0	0	618
1995	0	0	5008	42	0	5050
1996	0	0	81	153	0	234
1997	0	0	154	1387	6	1547
1998	0	0	1	679	20	700
1999	0	0	10	4410	36	4456
2000	0	0	1	352	189	542
2001	0	0	0	121	1075	1196
2002	0	0	0	75	765	840
2003	0	0	0	8	361	369
2004	0	0	0	54	4197	4251
2005	0	0	0	1	336	337
2006	0	0	0	4	120	124
2007	0	0	0	3	83	86
2008	0	0	0	0	23	23
Total	6020	7299	7280	7289	7211	35099

A.5 Coding party list ideology

Party list ideology is, in the first hand, coded based on the ideological classifications of ParlGov (Version — 29 April 2022), which has classified the ideological leaning of parties both currently and historically active in national level politics (Döring et al., 2022). Table A.3 presents ParlGov party ideological family classifications, as well as the simplified coding used in the analyses of this paper. As the analysis focuses only elections that take place latest 1997, the overview contains only names of parties founded before 1998 and delimits parties that were dissolved before the founding of the First Republic (1946).

Table A.3: Ideological family of party lists — Coding scheme

Collapsed coding	Party names (Abbrev)
Right-wing	Lega d'Azione Meridionale, Lega Nord, Fratelli d'Italia, Fiamma Tricolore, Movimento Sociale Italiano, Partito Nazionale Monarchico
Conservative	Alleanza Democratica, Alleanza Nazionale, Centro Cristiano Democratico, Cristiani Democratici Uniti, Cristiano Sociali, Centro-destra, Democrazia Cristiana, Democrazia Europea, Forza Italia — Il Popolo della Liberta, I Democratici, Patto Segni, Popolari per Prodi, Movimento per la Democrazia – La Rete, Unione Democratici per l'Europa
Liberal	Democrazia e Liberta — La Margherita, Italia dei Valori, Polo Liberal-Democratico, Partito Liberale Italiano, Partito Repubblicano Italiano, Unione Democratica, Unione Democratica di Centro
Green/Ecologist	Federazione delle Liste Verdi, Federazione dei Verdi, Verdi Arcobaleno
Left	Lista Dini — Rinnovamento Italiano, Democrazia Proletaria, Democratici di Sinistra, Lista per Trieste, Partito dei Comunisti Italiani, Partito di Unita Proletaria, Partito Socialisti Democratici Italiani, Radicali, Partito Comunista Italiano, Partito della Rifondazione Comunista, Partito Sardo d'Azione, Partito Socialista Democratico Italiano, Partito Socialista Italiano, Partito Socialista di Unita Proletaria
Special issue	Federalismo — Europa dei Popoli, Liga Lombarda, Liga Veneta, Movimento Indipendentista Siciliano, Partito dei Contadini d'Italia, Partito Nazionale Pensionati, Partito Pensionati, Suedtiroler Volkspartei, Union Valdotaine

Notes: Parties coded as Left are categorized as either 'Communist/Socialist' or 'Social Democratic' by ParlGov. Parties coded as Conservative are categorized as either 'Conservative' or 'Christian Democratic'.

Based on the above categorization, I use a list of terms to capture party ideology (see table A.4). If a party list contains none of the below-mentioned strings, it gets coded as a Civic List. In cases where party coalitions cross the right-left spectrum, I classify them as a Left-Right coalition.

Table A.4: String terms used for coding of party list ideology

Ideology	**Terms**
Left parties (Communist)	[COMUNIS*], [MAOIST], [MARX], [P.C.I], [PCI], [PDUP], [PDU], [PROGRESSISTI], [PROLETAR], [RIF.COM.], [ROBIN HOOD].
Left parties (Soc Dem)	[.SIN.*], [ALL. POP*], [ALLEANZA REPUBBLICA*], [CEN-SIN*], [CENTRO SINISTRA], [DEM.PROL*], [DEM.SIN*], [DEMOCRATICI SIN*], [DEMOCRATICI*], [DEMOCRAZIA-LIB*], [DINI*], [I DEMO*], [IND.SIN*], [L'UNIONE*], [LABURIST*], [LAVATOR*], [LAVORATOR*], [LISTA PER TRIESTE*], [MANIFEST*], [OPERAI*], [P.DEM*], [PANELLA*], [PART. DEMOCRAT*], [PART.DEM*], [PARTITO DEMOCRATICO*], [PARTITO SARDO D'AZIONE*], [PARTITO SOCIALISTA*], [PATTO DEM*], [PDL*], [PDS*], [PROGRESS*], [PROGRESSIST*], [PSI*], [RIFORMIST*], [RINNOVAMENTO IT*], [SDI*], [SINISTRA*], [SI], [SOC.*], [SOC.DEM*], [SOCIAL*], [SOLIDAR*], [ULIVO*], [UN.POP.*].
Green parties	[.VER], [ALL. VERD], [ALL.VERD], [ECOLOG], [VERDI].
Liberal	[CDR], [CENTRO DEMOCRAT], [CENTRO UNITO], [CONCENTRAZIONE DEMO], [EUROPE], [LA MARG], [LIB.DEM.], [LIBERAL], [LIBERT], [LISTA RAD.], [MARGHERITA], [P.RAD], [PANNELL], [PLI], [POLO B], [POLO P], [PRI], [RIFORMATORI SARD], [RINN.IT], [RINNOVAMENTO], [UDC], [UNIONE DEMOCRAT], [VALORI].
Conservative	[ADP], [ALL. NAZ.], [ALL.NAZ], [ALL.ZE DEMO], [ALLEANZA DEMOCRAT], [ALLEANZA NAZ], [AZZURR], [CASALINGHE], [CATTOL], [CATTOLICA], [CATTOLICI], [CCD], [CDL], [CDU], [CENTRO DEM], [CONTADINI], [CRIST.], [CRISTIAN], [DC], [DELLE LIBERTA], [DEM.CRIST], [-DES], [DESTR], [DESTRA], [DN-CD], [FI-], [FORZA], [FRONTE], [IL POPOLO DELLA LIBE], [LA RETE], [MODERATI], [MOVIMENTO PER L'AUTO], [P.D.C], [P.POPOLARE], [PARTITO POPOLARE], [POLO], [POPOLAR], [PPI], [PRODI], [SEGNI], [U.D.EUR], [UDEUR], [UDR], [UDS].
Right-wing	[AN-], [CONCENTRAZIONE], [DEMOCRAZIA NAZIO], [FASCISMO], [FIAMMA], [L.NORD], [L.VEN], [LEGA], [LG.], [LIGA VEN], [MON.NAZ.], [MONARCHI], [MSI-DN], [MUSSOLINI], [NAZ.INQ.], [NAZIONALIST], [P.NAZ.], [TRICOLOR].
Special Issue	[.PENS], [AUTONOM], [DEMOCRATICI POPO], [FEDERALISMO], [FRIUL.], [PATT], [PENSIONAT], [SVP], [TIROL], [UV], [VENETO], [VESUVIO].

B Descriptive statistics

B.1 Covariate balance

Table B.1: Covariate balance across treatment and control groups

	Treatment $N = 7358$	Control $N = 317$	All $N = 7036$	Min	Max
NUTS-2 level regions					
North-West	0.40	0.23	0.38	0	1
	(0.006)	(0.02)	(0.02)		
North-East	0.15	0.10	0.19	0	1
	(0.006)	(0.01)	(0.02)		
Center	0.13	0.07	0.13	0	1
	(0.004)	(0.01)	(0.003)		
South (incl. islands)	0.30	0.56	0.30	0	1
	(0.005)	(0.03)	(0.01)		
Municipality size (1991 Census)					
+200K inhabitants	0.0026	0.0032	0.002	0	1
	(0.0006)	(0.002)	(0.0006)		
50-200K inhabitants	0.013	0.028	0.013	0	1
	(0.001)	(0.02)	(0.001)		
15-50K inhabitants	0.11	0.14	0.011	0	1
	(0.004)	(0.02)	(0.003)		
-15 K inhabitants	0.88	0.83	0.88	0	1
	(0.004)	(0.02)	(0.004)		
Unemployment (1991 Census)					
Above 15 rate	0.16	0.19	0.15	0	0.89
	(0.001)	(0.007)	(0.001)		
Educational attainment (1991 Census)					
Tertiary schooling rate	0.02	0.03	0.02	0	0.46
	(0.0002)	(0.001)	(0.0002)		

B.2 Party affiliations of local legislators

Figure B.1: Party affiliation of male legislators, by social class and election round

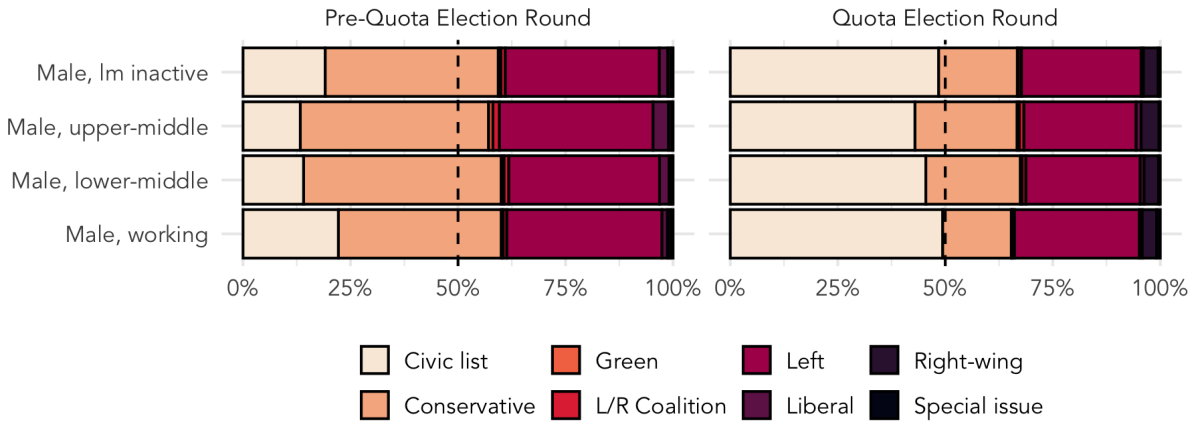
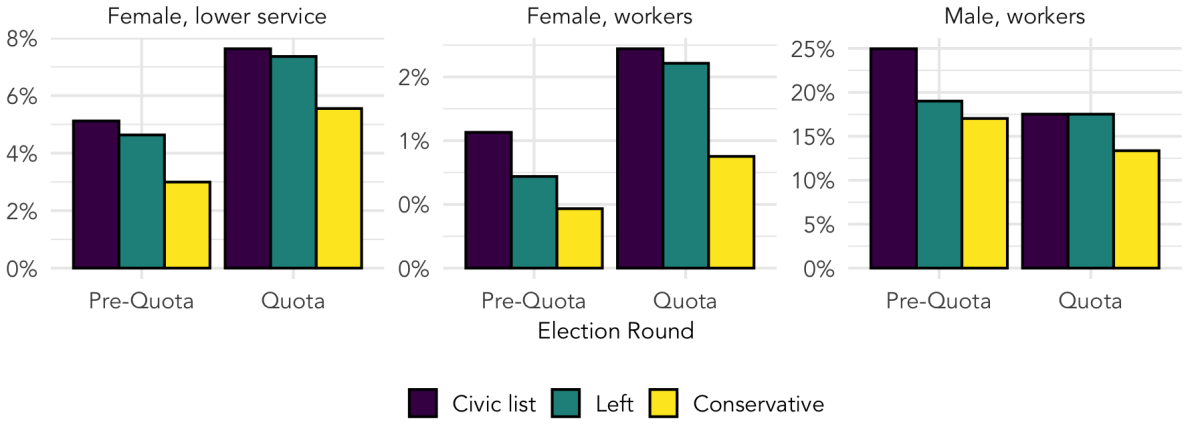


Figure B.2: Mean share of lower social class legislators, by party ideology and election round



C Regression tables

C.1 Quota effects on intersectional representation

Table C.1: Quota effects on female representation in municipal councils

	(1) All women	(2) Female upper-middle	(3) Female lower-middle	(4) Female working
Quota municipality	-0.0107*	-0.00105	-0.00447	-0.00272**
	(0.00449)	(0.00153)	(0.00299)	(0.000912)
2nd Pre-Quota	0.0211***	0.00805***	0.00333	0.00610***
	(0.00452)	(0.00203)	(0.00304)	(0.00117)
Quota	0.0610***	0.0205***	0.0119**	0.00630**
	(0.00747)	(0.00439)	(0.00441)	(0.00212)
1st Post-Quota	0.0544***	0.0236***	0.00467	0.00781***
	(0.00671)	(0.00319)	(0.00400)	(0.00199)
2nd Post-Quota	0.0566***	0.0251***	0.00961*	0.00529**
	(0.00740)	(0.00382)	(0.00486)	(0.00191)
Quota municipality	0.00715	-0.00131	0.00498	-0.00117
× 2nd Pre-Quota	(0.00462)	(0.00207)	(0.00311)	(0.00119)
Quota municipality	0.0617***	0.00769	0.0238***	0.00594**
× Quota	(0.00754)	(0.00443)	(0.00450)	(0.00215)
Quota municipality	0.0398***	-0.000471	0.0225***	0.00210
× 1st Post-Quota	(0.00678)	(0.00323)	(0.00409)	(0.00200)
Quota municipality	0.0398***	0.00217	0.0190***	0.00321
× 2nd Post-Quota	(0.00740)	(0.00382)	(0.00488)	(0.00190)
Constant	0.150***	0.00524**	0.0699***	0.0197***
	(0.00519)	(0.00202)	(0.00337)	(0.00117)
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table C.2.

Table C.2: Marginal quota effects on female representation in municipal councils, by election round

	(1) All women	(2) Female upper-middle	(3) Female lower-middle	(4) Female working
1st Pre-Quota	-0.0107*	-0.00105	-0.00447	-0.00272**
	[-0.0195,-0.00190]	[-0.00405,0.00196]	[-0.0103,0.00138]	[-0.00451,-0.000935]
2nd Pre-Quota	-0.00355	-0.00235	0.000511	-0.00389**
	[-0.0120,0.00491]	[-0.00609,0.00138]	[-0.00453,0.00555]	[-0.00633,-0.00145]
Quota	0.0510***	0.00664	0.0194***	0.00322
	[0.0381,0.0639]	[-0.00157,0.0149]	[0.0125,0.0262]	[-0.000743,0.00718]
1st Post-Quota	0.0291***	-0.00152	0.0180***	-0.000619
	[0.0172,0.0411]	[-0.00781,0.00477]	[0.0117,0.0244]	[-0.00444,0.00320]
2nd Post-Quota	0.0291***	0.00113	0.0145***	0.000489
	[0.0166,0.0416]	[-0.00574,0.00800]	[0.00693,0.0221]	[-0.00321,0.00419]
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. 95% confidence intervals in brackets. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table C.3: Quota effects on male representation in municipal councils

	(1) All men	(2) Male upper-middle	(3) Male lower-middle	(4) Male working
Quota municipality	0.0107*	-0.0214*	0.0349***	-0.00349
	(0.00449)	(0.00875)	(0.00942)	(0.00859)
2nd Pre-Quota	-0.0211***	0.0189*	-0.0258**	0.0446***
	(0.00452)	(0.00779)	(0.00921)	(0.00865)
Quota	-0.0610***	0.0663***	-0.0356***	-0.0133
	(0.00747)	(0.00968)	(0.0101)	(0.00980)
1st Post-Quota	-0.0544***	0.0699***	-0.0208	-0.0287**
	(0.00671)	(0.0102)	(0.0113)	(0.0100)
2nd Post-Quota	-0.0566***	0.0801***	-0.00286	-0.0431***
	(0.00740)	(0.0108)	(0.0120)	(0.0108)
Quota municipality	-0.00715	0.00358	-0.0190*	0.00405
× 2nd Pre-Quota	(0.00462)	(0.00792)	(0.00942)	(0.00880)
Quota municipality	-0.0617***	-0.00706	-0.0461***	0.00167
× Quota	(0.00754)	(0.00977)	(0.0102)	(0.00982)
Quota municipality	-0.0398***	-0.00871	-0.0465***	0.0141
× 1st Post-Quota	(0.00678)	(0.0103)	(0.0114)	(0.0100)
Quota municipality	-0.0398***	-0.0121	-0.0691***	0.0110
× 2nd Post-Quota	(0.00740)	(0.0107)	(0.0119)	(0.0105)
Constant	0.850***	-0.00189	0.233***	0.326***
	(0.00519)	(0.0101)	(0.0103)	(0.00954)
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table C.4.

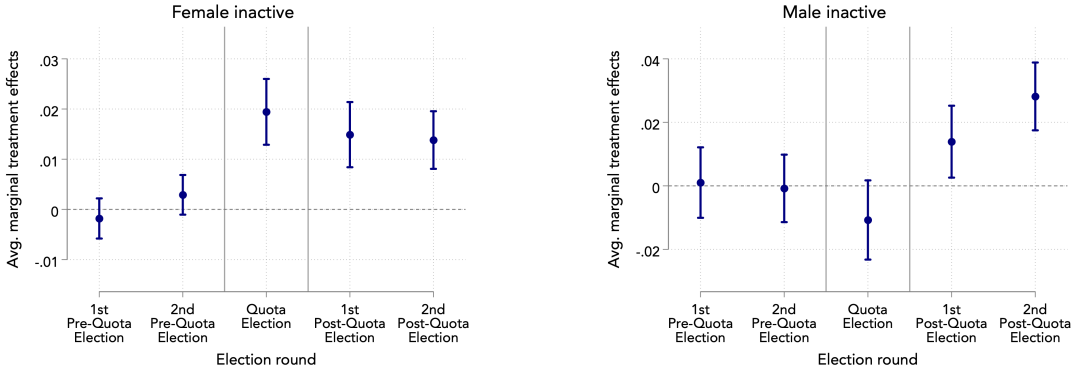
Table C.4: Marginal quota effects on male representation in municipal councils, by election round

	(1) All men	(2) Male upper-middle	(3) Male lower-middle	(4) Male working
1st Pre-Quota	0.0107*	-0.0214*	0.0349***	-0.00349
	[0.00190,0.0195]	[-0.0385,-0.00421]	[0.0164,0.0533]	[-0.0203,0.0134]
2nd Pre-Quota	0.00355	-0.0178*	0.0159	0.000558
	[-0.00491,0.0120]	[-0.0350,-0.000626]	[-0.00143,0.0332]	[-0.0164,0.0175]
Quota	-0.0510***	-0.0284***	-0.0112	-0.00183
	[-0.0639,-0.0381]	[-0.0440,-0.0129]	[-0.0272,0.00468]	[-0.0170,0.0134]
1st Post-Quota	-0.0291***	-0.0301***	-0.0116	0.0106
	[-0.0411,-0.0172]	[-0.0468,-0.0133]	[-0.0284,0.00513]	[-0.00373,0.0249]
2nd Post-Quota	-0.0291***	-0.0335***	-0.0343***	0.00754
	[-0.0416,-0.0166]	[-0.0505,-0.0165]	[-0.0505,-0.0180]	[-0.00710,0.0222]
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. 95% confidence intervals in brackets. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

C.2 Quota effects on labor market inactive sub-groups

Figure C.1: Quota effects on labor market inactive sub-groups



Note: Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots, to better display effect sizes and confidence intervals. All model specifications include municipal level controls. Standard errors are clustered on a municipality level. Regression coefficients presented in table C.6; calculated marginal effects presented in table C.5.

Table C.5: Marginal quota effects on labor market inactives in municipal councils, by election round

	(1) Female inactives	(2) Male inactives
1st Pre-Quota	-0.00181 [-0.00581,0.00219]	0.00103 [-0.0101,0.0121]
2nd Pre-Quota	0.00291 [-0.00105,0.00686]	-0.000799 [-0.0114,0.00983]
Quota	0.0194*** [0.0129,0.0260]	-0.0108 [-0.0232,0.00172]
1st Post-Quota	0.0149*** [0.00840,0.0214]	0.0139* [0.00258,0.0252]
2nd Post-Quota	0.0138*** [0.00808,0.0196]	0.0282*** [0.0175,0.0388]
N	35071	35071
Controls	✓	✓

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table C.6: Quota effects on labor market inactives in municipal councils

	(1) Female inactives	(2) Male inactives
Quota municipality	-0.00181	0.00103
	(0.00204)	(0.00566)
2nd Pre-Quota	0.00483*	0.00197
	(0.00214)	(0.00504)
Quota	0.0211***	-0.00534
	(0.00358)	(0.00696)
1st Post-Quota	0.0151***	-0.0162*
	(0.00359)	(0.00720)
2nd Post-Quota	0.00759*	-0.0396***
	(0.00358)	(0.00710)
Quota municipality	0.00471*	-0.00183
× 2nd Pre-Quota	(0.00221)	(0.00518)
Quota municipality	0.0212***	-0.0118
× Quota	(0.00369)	(0.00707)
Quota municipality	0.0167***	0.0129
× 1st Post-Quota	(0.00368)	(0.00730)
Quota municipality	0.0156***	0.0271***
× 2nd Post-Quota	(0.00364)	(0.00713)
Constant	0.0454***	0.192***
	(0.00244)	(0.00654)
N	35071	35071
Controls	✓	✓

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table C.5.

C.3 Quota effects on intersectional representation, TWFE models

Estimations below run on elections taking place in pre-quota election rounds, as well as the quota round itself. All estimates adjusted for covariates, time effects (election round) and panel effects (i.e. municipality fixed effects).

Table C.7: Average effects on the treated (ATETs) - Gender-class representation

	(1) Female upper-middle class	(2) Female lower-middle class	(3) Female working class	(4) Male upper-middle class	(5) Male lower-middle class	(6) Male working class
Quota in force	0.00705	0.0165***	0.00622**	-0.00273	-0.0214*	-0.0111
	(0.00439)	(0.00388)	(0.00207)	(0.00801)	(0.00833)	(0.00803)
Observations	20584	20584	20584	20584	20584	20584
Election round FE	✓	✓	✓	✓	✓	✓
Municipality FE	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table C.8: F-tests for joint significance of pre-treatment coefficients — TWFE models

	Female upper-middle	Female lower-middle	Female working	Male upper-middle	Male lower-middle	Male working
$F(1, 7296)$	0.02	1.77	0.70	1.85	3.95	0.20
p-value	0.900	0.184	0.403	0.174	0.047	0.658

Notes: As all tests are run on the same sample, degrees of freedom are identical across estimations. Lower p-values indicate that pre-treatment trends diverge between municipalities assigned to treatment and control.

C.4 Quota effects on legislature ideological composition

Table C.9: Quota effects on legislature seat shares, by party ideological group

	(1) Left	(2) Conservative	(3) Liberal	(4) Green	(5) G.Coalition	(6) Special issue	(7) Right wing	(8) Civic List
Quota municipality	0.0485***	0.0376*	-0.00201	0.00217*	-0.00204	0.00203***	-0.0250***	-0.0612***
	(0.0125)	(0.0151)	(0.00315)	(0.000881)	(0.00211)	(0.000392)	(0.00511)	(0.0153)
Quota in force	0.0254	-0.158***	0.000796	0.00150	0.00107	-0.00489***	0.0214***	0.113***
	(0.0204)	(0.0205)	(0.00566)	(0.000981)	(0.00266)	(0.000800)	(0.00642)	(0.0266)
Quota municipality	-0.0621**	-0.0617**	-0.00460	-0.000963	0.00210	0.00279***	0.0167**	0.108***
× Quota in force	(0.0203)	(0.0205)	(0.00562)	(0.000988)	(0.00259)	(0.000676)	(0.00632)	(0.0264)
Constant	0.0859***	0.336***	-0.00507	-0.0105***	-0.00919**	0.0116***	0.0147*	0.576***
	(0.0163)	(0.0183)	(0.00368)	(0.00136)	(0.00291)	(0.00207)	(0.00653)	(0.0192)
N	14216	14216	14216	14216	14216	14216	14216	14216
Controls	✓	✓	✓	✓	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

C.5 Quota effects on voter turnout

Table C.10: Quota effects on turnout in municipal elections

	Dependent variable: Turnout		
	(1)	(2)	(3)
Quota Municipality	0.0237***	0.0125	0.0125
	(0.00629)	(0.00886)	(0.00886)
Post Quota	-0.0669***	-0.0515*	-0.0518*
	(0.00979)	(0.0308)	(0.0308)
Quota Municipality × Post Quota	0.0229**	0.0377***	0.0371***
	(0.00985)	(0.0115)	(0.0115)
N	13861	13861	13861
Controls	✓	✓	✓
Year FE		✓	✓
Early govt. termination control			✓

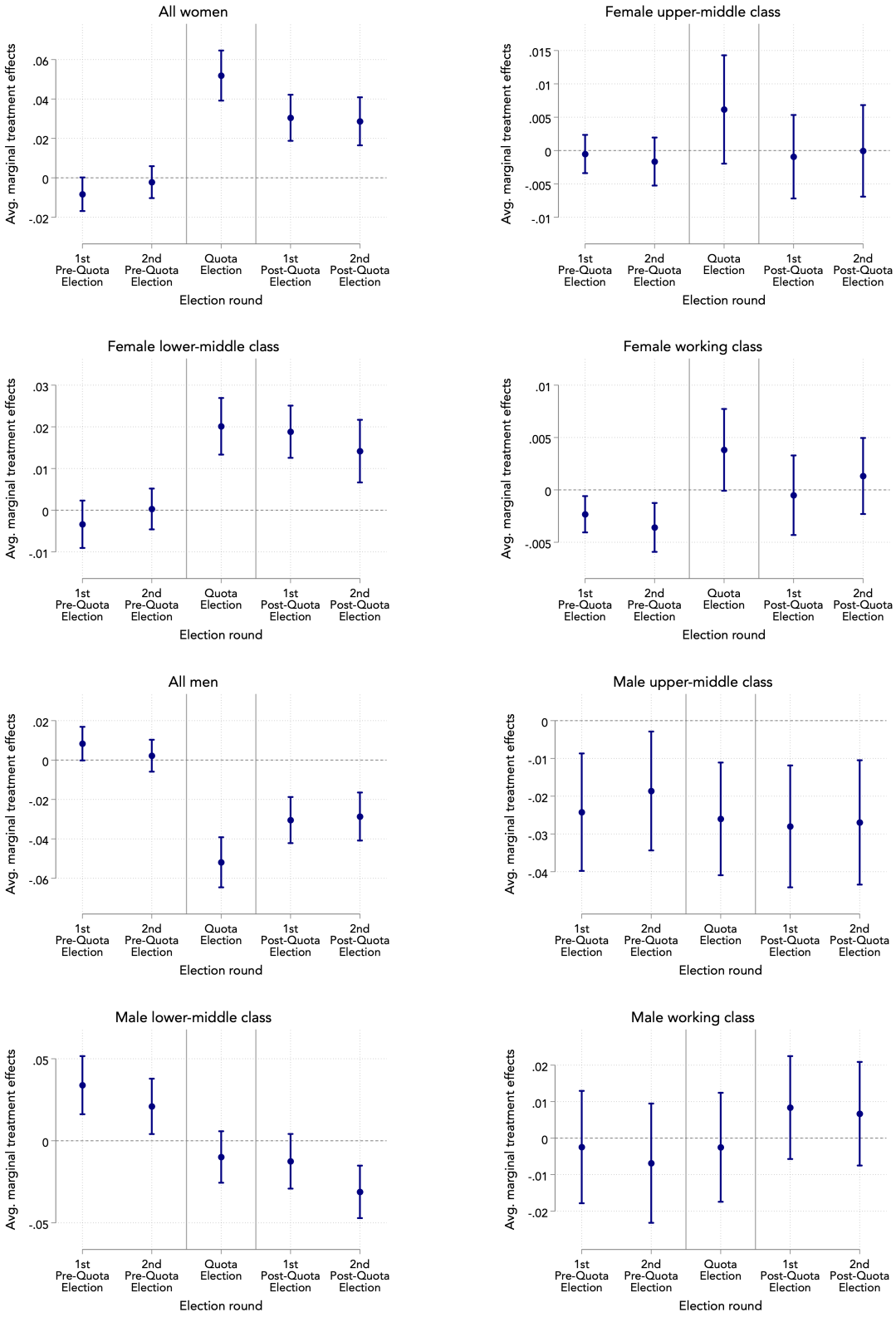
Notes: Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

D Robustness checks

D.1 Accounting for electoral system change

The 1993 gender quota was introduced concomitant with a major electoral reform for local elections. First, municipalities with +5000 inhabitants shifted from a proportional to quasi-majoritarian electoral system. Municipalities below the 5000 inhabitant threshold were unaffected by the change, as they maintained quasi-majoritarian electoral rules pre-reform. Second, municipalities with +15'000 inhabitants introduced dual-stage (or run-off) elections for the mayorship. As the population size of municipalities is exogenous to election-timing, municipalities that were affected by aforementioned reforms are prevalent in both treatment and control groups. To ensure that treatment affects are not fully mediated by these other reforms, I control for whether municipalities were subject to any of the two reform changes, by including dummy variables to my baseline statistical model. Estimates from this exercise highlight that my findings are robust to the inclusion of these controls (see figure D.1).

Figure D.1: Quota effects on intersectional representation, when controlling for electoral reform changes



Note: Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots, to better display effect sizes and confidence intervals. All model specifications include municipal level controls. Standard errors are clustered on a municipality level. For regression coefficients, see tables D.1 (female) and D.2 (male). For calculated marginal effects see tables D.3 (female) and D.4 (male).

Table D.1: Quota effects on female representation in municipal councils (adj. for electoral reforms)

	(1) All women	(2) Female upper-middle	(3) Female lower-middle	(4) Female working
Quota municipality	-0.00930*	-0.00115	-0.00386	-0.00253**
	(0.00451)	(0.00155)	(0.00300)	(0.000909)
2nd Pre-Quota	0.0186***	0.00822***	0.00225	0.00576***
	(0.00451)	(0.00202)	(0.00303)	(0.00116)
Quota	0.0663***	0.0196***	0.0143**	0.00726***
	(0.00745)	(0.00437)	(0.00441)	(0.00212)
1st Post-Quota	0.0596***	0.0226***	0.00698	0.00876***
	(0.00669)	(0.00319)	(0.00400)	(0.00199)
2nd Post-Quota	0.0632***	0.0238***	0.0126**	0.00655***
	(0.00739)	(0.00383)	(0.00485)	(0.00192)
Quota municipality	0.00602	-0.00114	0.00449	-0.00136
× 2nd Pre-Quota	(0.00460)	(0.00207)	(0.00310)	(0.00119)
Quota municipality	0.0594***	0.00810	0.0228***	0.00553**
× Quota	(0.00752)	(0.00442)	(0.00448)	(0.00214)
Quota municipality	0.0380***	-0.000184	0.0217***	0.00179
× 1st Post-Quota	(0.00674)	(0.00323)	(0.00407)	(0.00200)
Quota municipality	0.0366***	0.00285	0.0175***	0.00259
× 2nd Post-Quota	(0.00736)	(0.00382)	(0.00485)	(0.00190)
Reformer:	-0.0208***	0.00740***	-0.00958***	-0.00511***
Quasi-maj. formula	(0.00182)	(0.000894)	(0.00112)	(0.000402)
Reformer:	-0.0125***	-0.00875***	-0.00445*	0.00190***
Dual-stage elections	(0.00301)	(0.00143)	(0.00179)	(0.000575)
Constant	0.129***	0.00661**	0.0608***	0.0169***
	(0.00547)	(0.00218)	(0.00357)	(0.00121)
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table D.3.

Table D.2: Quota effects on male representation in municipal councils (adj. for electoral reforms)

	(1) All men	(2) Male upper-middle	(3) Male lower-middle	(4) Male working
Quota municipality	0.00930*	-0.0261**	0.0344***	-0.0000332
	(0.00451)	(0.00828)	(0.00943)	(0.00815)
2nd Pre-Quota	-0.0186***	0.0274***	-0.0249**	0.0384***
	(0.00451)	(0.00770)	(0.00921)	(0.00837)
Quota	-0.0663***	0.0459***	-0.0390***	0.00262
	(0.00745)	(0.00949)	(0.0101)	(0.00961)
1st Post-Quota	-0.0596***	0.0499***	-0.0243*	-0.0131
	(0.00669)	(0.00993)	(0.0113)	(0.00982)
2nd Post-Quota	-0.0632***	0.0539***	-0.00769	-0.0223*
	(0.00739)	(0.0105)	(0.0120)	(0.0106)
Quota municipality	-0.00602	0.00776	-0.0183	0.000850
× 2nd Pre-Quota	(0.00460)	(0.00783)	(0.00942)	(0.00851)
Quota municipality	-0.0594***	0.00163	-0.0446***	-0.00512
× Quota	(0.00752)	(0.00955)	(0.0102)	(0.00961)
Quota municipality	-0.0380***	-0.00179	-0.0454***	0.00874
× 1st Post-Quota	(0.00674)	(0.01000)	(0.0114)	(0.00978)
Quota municipality	-0.0366***	0.000713	-0.0668***	0.000863
× 2nd Post-Quota	(0.00736)	(0.0104)	(0.0119)	(0.0103)
Reformer: Quasi-maj. formula	0.0208***	0.0939***	0.0238***	-0.0796***
	(0.00182)	(0.00313)	(0.00298)	(0.00224)
Reformer: Dual-stage elections	0.0125***	0.00495	-0.0226***	0.0153***
	(0.00301)	(0.00605)	(0.00512)	(0.00329)
Constant	0.871***	0.0697***	0.240***	0.274***
	(0.00547)	(0.00981)	(0.0108)	(0.00923)
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table D.4.

Table D.3: Marginal quota effects on female representation in municipal councils, by election round (w. controls for electoral reforms)

	(1) All women	(2) Female upper-middle	(3) Female lower-middle	(4) Female working
1st Pre-Quota	-0.00930*	-0.00115	-0.00386	-0.00253**
	[-0.0181,-0.000462]	[-0.00419,0.00190]	[-0.00974,0.00201]	[-0.00431,-0.000747]
2nd Pre-Quota	-0.00327	-0.00228	0.000623	-0.00389**
	[-0.0117,0.00516]	[-0.00603,0.00147]	[-0.00442,0.00567]	[-0.00631,-0.00147]
Quota	0.0501***	0.00696	0.0190***	0.00300
	[0.0372,0.0631]	[-0.00125,0.0152]	[0.0121,0.0258]	[-0.000951,0.00696]
1st Post-Quota	0.0287***	-0.00133	0.0178***	-0.000740
	[0.0168,0.0405]	[-0.00761,0.00495]	[0.0115,0.0242]	[-0.00455,0.00307]
2nd Post-Quota	0.0273***	0.00170	0.0137***	0.0000650
	[0.0148,0.0397]	[-0.00517,0.00857]	[0.00610,0.0212]	[-0.00362,0.00375]
N	34226	34226	34226	34226
Run-off reformer FE	✓	✓	✓	✓
Quasi-maj reformer FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. 95% confidence intervals in brackets. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table D.4: Marginal quota effects on male representation in municipal councils, by election round (w. controls for electoral reforms)

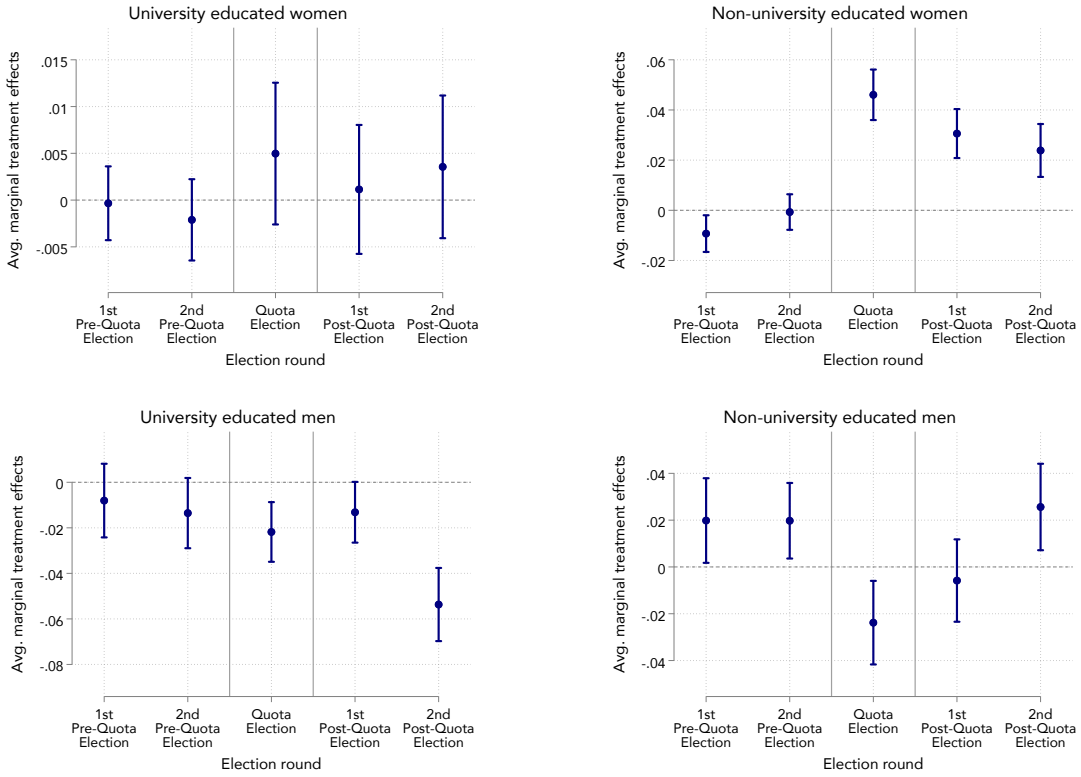
	(1) All men	(2) Male upper-middle	(3) Male lower-middle	(4) Male working
1st Pre-Quota	0.00930*	-0.0261**	0.0344***	-0.0000332
	[0.000462,0.0181]	[-0.0424,-0.00992]	[0.0159,0.0529]	[-0.0160,0.0159]
2nd Pre-Quota	0.00327	-0.0184*	0.0160	0.000816
	[-0.00516,0.0117]	[-0.0349,-0.00188]	[-0.00124,0.0333]	[-0.0155,0.0171]
Quota	-0.0501***	-0.0245**	-0.0102	-0.00516
	[-0.0631,-0.0372]	[-0.0395,-0.00954]	[-0.0261,0.00567]	[-0.0201,0.00982]
1st Post-Quota	-0.0287***	-0.0279***	-0.0110	0.00870
	[-0.0405,-0.0168]	[-0.0441,-0.0118]	[-0.0278,0.00571]	[-0.00526,0.0227]
2nd Post-Quota	-0.0273***	-0.0254**	-0.0324***	0.000830
	[-0.0397,-0.0148]	[-0.0422,-0.00869]	[-0.0487,-0.0161]	[-0.0136,0.0153]
N	34226	34226	34226	34226
Run-off reformer FE	✓	✓	✓	✓
Quasi-maj reformer FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. 95% confidence intervals in brackets. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

D.2 Quota effects on representation of gender, by educational class

To ensure robustness I re-run my analysis on a second outcome associated with social class: university-level educational attainment (see figure D.2). In spite of accounting for a full set of legislators, including the labor market inactive, finding from these tests display a pattern consistent with gender-class specific effects. Results may seem contrary to those of Baltrunaite et al. (2014), who use the same data and case to analyze gender quota effects on the educational attainment of Italian legislators. Importantly, these authors examine educational attainment as a continuous variable indicating schooling-years. Thus their results may diverge from mine as their effects stem from variation brought about by within-group schooling-year differences among legislators that are either university educated or not.

Figure D.2: Treatment effects on statistical representation in municipal councils by educational attainment



Notes: Bandwidths indicate 95% confidence intervals. Y-axis scales vary across plots, to better display effect sizes and confidence intervals. All model specifications include municipal level population controls. Standard errors are clustered on a municipality level. For regression coefficients, see table D.5. Estimated marginal effects presented in table D.6.

Table D.5: Quota effects on statistical representation in municipal councils by educational attainment

	(1) University ed. women	(2) Non-university ed. women	(3) University ed. men	(4) Non-university ed. men
Quota municipality	-0.000340	-0.00929*	-0.00803	0.0198*
	(0.00201)	(0.00373)	(0.00825)	(0.00923)
2nd Pre-Quota	0.00540*	0.0151***	-0.00582	-0.0172*
	(0.00244)	(0.00386)	(0.00689)	(0.00825)
Quota	0.0208***	0.0377***	0.0282**	-0.0958***
	(0.00410)	(0.00592)	(0.00858)	(0.0110)
1st Post-Quota	0.0192***	0.0326***	0.0253**	-0.0906***
	(0.00370)	(0.00567)	(0.00808)	(0.0108)
2nd Post-Quota	0.0264***	0.0280***	0.0805***	-0.152***
	(0.00413)	(0.00620)	(0.0102)	(0.0124)
Quota municipality	-0.00177	0.00859*	-0.00550	-0.0000889
× 2nd Pre-Quota	(0.00249)	(0.00396)	(0.00704)	(0.00841)
Quota municipality	0.00531	0.0553***	-0.0138	-0.0436***
× Quota	(0.00415)	(0.00599)	(0.00858)	(0.0111)
Quota municipality	0.00148	0.0399***	-0.00513	-0.0257*
× 1st Post-Quota	(0.00375)	(0.00574)	(0.00807)	(0.0108)
Quota municipality	0.00390	0.0331***	-0.0457***	0.00579
× 2nd Post-Quota	(0.00414)	(0.00616)	(0.00999)	(0.0122)
Constant	0.00619*	0.139***	-0.0566***	0.899***
	(0.00249)	(0.00431)	(0.00968)	(0.0105)
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table D.6.

Table D.6: Marginal effects of gender quota on statistical representation in municipal councils, by educational attainment

	(1) Uni ed. female	(2) Non-uni ed. female	(3) Uni ed. male	(4) Non-uni ed. male
1st Pre-Quota	-0.000340	-0.00929*	-0.00803	0.0198*
	[-0.00428,0.00360]	[-0.0166,-0.00197]	[-0.0242,0.00814]	[0.00175,0.0379]
2nd Pre-Quota	-0.00211	-0.000695	-0.0135	0.0197*
	[-0.00646,0.00223]	[-0.00778,0.00639]	[-0.0290,0.00189]	[0.00359,0.0359]
Quota	0.00497	0.0460***	-0.0218**	-0.0238**
	[-0.00260,0.0125]	[0.0360,0.0561]	[-0.0349,-0.00871]	[-0.0417,-0.00596]
1st Post-Quota	0.00114	0.0306***	-0.0132	-0.00582
	[-0.00575,0.00803]	[0.0208,0.0404]	[-0.0265,0.000169]	[-0.0234,0.0118]
2nd Post-Quota	0.00356	0.0239***	-0.0537***	0.0256**
	[-0.00407,0.0112]	[0.0133,0.0344]	[-0.0698,-0.0376]	[0.00715,0.0441]
N	34226	34226	34226	34226
Controls	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment. Calculated marginal effects are presented in table D.6.

D.3 Excl. elections around the 1994 national shock

Table D.7: Average effects on the treated - Excl. election around 1994 national shock

	(1) Female upper-middle class	(2) Female lower-middle class	(3) Female working class	(4) Male upper-middle class	(5) Male lower-middle class	(6) Male working class
Quota in force	0.00765	0.0169***	0.00615**	-0.00575	-0.0204*	-0.0139
	(0.00440)	(0.00389)	(0.00208)	(0.00804)	(0.00837)	(0.00800)
Observations	19791	19791	19791	19791	19791	19791
Election round FE	✓	✓	✓	✓	✓	✓
Municipality FE	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓

Notes: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

D.4 Distance to *Mani pulite*: Regression results

Table D.8: Distance to Mani Pulite — Numerical female representation in councils

	Dependent variable: Share of municipal council					
	Female Upper-middle class		Female Lower-middle class		Female Working class	
	(1)	(2)	(3)	(4)	(5)	(6)
Distance in weeks	0.0001***	0.00002	0.0002***	0.00002	0.0001***	0.00002
	(0.00002)	(0.00002)	(0.00002)	(0.00003)	(0.00001)	(0.00001)
Intercept	0.035***	0.018***	0.052***	0.114***	0.010***	0.034***
	(0.002)	(0.004)	(0.002)	(0.005)	(0.001)	(0.002)
Observations	7,036	7,008	7,036	7,008	7,036	7,008
Controls		✓		✓		✓
R^2	0.002	0.014	0.014	0.053	0.005	0.026
Adjusted R^2	0.001	0.013	0.014	0.052	0.005	0.025

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table D.9: Distance to Mani Pulite — Numerical male representation in council

	Dependent variable: Share of municipal council					
	Male Upper-middle class		Male Lower-middle class		Male Working class	
	(1)	(2)	(3)	(4)	(5)	(6)
Distance in weeks	−0.001***	−0.0003***	−0.001***	−0.0002***	0.00004	0.0002***
	(0.00004)	(0.0001)	(0.00004)	(0.0001)	(0.00004)	(0.0001)
Intercept	0.247***	0.014*	0.299***	0.194***	0.167***	0.316***
	(0.004)	(0.009)	(0.004)	(0.009)	(0.004)	(0.008)
Observations	7,036	7,008	7,036	7,008	7,036	7,008
Controls		✓		✓		✓
R^2	0.018	0.193	0.020	0.051	0.0001	0.119
Adjusted R^2	0.018	0.193	0.020	0.050	0.00000	0.119

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the municipality level. Controls, all indexed to municipality and time, include council size, population size, population size squared, educational attainment, and unemployment.

Table D.10: Distance to manipulite — Mechanisms

	Dependent variable:		
	Voter Turnout	Civic List Seat Share	Mainstream Party Seat Share
	(1)	(2)	(3)
Distance in weeks	0.0002***	0.002***	−0.002***
	(0.00003)	(0.0001)	(0.0001)
Intercept	0.816***	−0.017**	0.926***
	(0.003)	(0.008)	(0.009)
Observations	6,522	6,986	6,986
R^2	0.010	0.053	0.052
Adjusted R^2	0.010	0.053	0.052

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome *Mainstream Party Seat Share* is a composite measure of the seats held by Conservative and Left parties, shown to be negatively affected by gender quota adoption (see Table C.9).