Gender Quotas and Far-Right Backlash

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Abstract

How do gender quotas affect support for far-right parties? We argue that the implementation of gender quotas may trigger a threat response in individuals who perceive efforts to include a historically disaffected group as an attack on their privileged societal position. The literature typically defines the far-right in terms of three major criteria: populism, nativism, and authoritarianism. Yet there is also an undercurrent of anti-feminism that is prominent among far-right supporters and in party platforms. Far-right parties tend to place "traditional" values in the spotlight, which emphasize the supportive and domestic roles of women. At times, these leanings hyperbolize a feminist oppression of men, which leaves men feeling deprived of their previous social standing. Gender quotas trigger backlash based on these perceived threats. Using a staggered difference in differences design and an instrumental variables design, we study how institutional change affects vote-share gains among far-right parties in Italy. This research suggests that there may be unintentional downstream consequences of gender quota implementation by provoking demand for far-right parties.

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1   Introduction

Universal suffrage is a cornerstone of liberal democracy. Still, universal suffrage is insufficient when there are barriers to entry, cultural or otherwise, that exclude or strongly discourage the participation of large swaths of the population. Historically, women have been excluded from the political process by limiting women's suffrage or not including women in the pipelines to power (including service in local government, nomination to party lists, etc.). Gender quotas, which increase the proportion of women represented in government and/or the pipelines to power, are the main policy tool used to correct this historic injustice. The evidence shows a) gender quotas work in increasing women's representation (Tripp and Kang, 2008; Weeks, 2022; O'Brien and Rickne, 2016), and that b) gender quotas have normatively positive knock-on effects, including increasing the quality of legislators (Weeks and Baldez, 2015), the perceived qualification of women as leaders (O'Brien and Rickne, 2016), and increasing political knowledge, especially among younger women (Venturini, forthcoming `https://strathprints.strath.ac.uk/90754/`). However, one cannot correct an injustice without backlash from those whom the injustice has historically privileged. In this paper, we investigate a normatively troubling knock-on effect, namely, a far-right backlash to the implementation of gender quotas.

Do gender quotas affect support for far-right parties? Mudde (2019) correctly points out that much of the literature on the intersection of far-right politics and gender focuses on the role of sexism and sexist beliefs in far-right parties as a factor *pushing* women away from the far-right instead of sexism as a factor *pulling* male voters towards the far-right. We argue that the implementation of gender quotas may further incentivize disaffected individuals to vote for far-right parties.[1] The literature typically defines the far-right in terms of three major criteria: populism, nativism, and authoritarianism (Mudde, 2007; Mudde and Kaltwasser, 2012). Yet there is also an undercurrent of anti-feminism that is prominent among far-

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[1] We hypothesize that these disaffected individuals will be disproportionately male, but we lack the data to empirically test this.

right supporters and in far-right party platforms (Klammer and Goetz, 2017). Far-right parties tend to place traditional values in the spotlight, emphasizing women's supportive and domestic roles. At times, these leanings hyperbolize feminist oppression of men, which leaves men feeling deprived of their previous social standing (Klammer and Goetz, 2017). Parties that capitalize on discontent surrounding gender quotas may attract support from citizens that a) reject modern feminism generally, and b) oppose institutional interventions in gender equality that threaten the role of men in the public sphere. For these reasons, we hypothesize that gender quotas may have unintended consequences by driving support to far-right parties.

In this paper, we examine municipal-level gender quotas for city councils and their effect on vote share for mayoral candidates supported by the far-right. This is a useful measure for several reasons. First, gender quotas have no direct *mechanical* influence on the mayoral office, and thus any systemic changes to vote share for mayoral candidates must come from a change in expressed preferences.[2] Second, this measure is plausibly continuous between 0-1, which gives us a variety of appropriate modeling choices. We then use a variety of matching algorithms, a difference in differences design, and an instrumental variables approach and find that implementing a municipal gender quota for city counselors is associated with a $\sim$ 1.8% - 7.8% increase in vote share for mayoral candidates supported by far-right parties. Our preferred estimation strategy, a staggered difference in differences design with an entropy-balanced sample (model 3), gives an estimate of a $\sim$ 5.31% increase in vote share for mayoral candidates supported by far-right parties.

2   Theory

What are gender quotas? According to the Global Database of Quotas for Women (*Quotas | International IDEA*, 2024), there are three types of gender quotas. 1. Reserved seats, 2. legal candidate quotas, and 3. political party quotas. Reserved seats codify the number or

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[2] See Spaziani (2022)

percentage of seats reserved for women. Legal candidate quotas require a certain number of women to be included on party lists. The Italian gender quota system we analyze in this paper uses these first and second types of quotas. Some countries, such as Argentina and Belgium, include so-called "double quotas", or mandates about the placement of women on these party lists to discourage parties from simply placing women at the bottom of the list. Political party quotas are voluntary quotas that political parties place upon *themselves*. Because these quotas are *voluntary*, and party-level, we largely ignore the impact of this type of quota in this paper. It is important to note that no Radical Right Party in Europe has introduced this third type of quota (Weeks, 2022). To date, eleven EU member states plus Norway have implemented gender quotas at the national level[3]. Quotas also exist at varying levels of government.

Far-right parties and voters often espouse sexist ideas that either seek to limit the role of women outside of the home or are overtly hostile to women. Sociologist Andreas Kemper argues that at its core, the far right espouses *familialism*, or the idea that traditional hetero-sexual family structure is the building block of the nation, and thus, the individual's rights to self-determination must be subjugated when they come into conflict with the 'reproduction' of the nation (Kemper, 2014). The traditional heterosexual majority ethnic group family is viewed as a microscopic version of the nation, and thus policies that support the traditional heterosexual majority ethnic group family are encouraged, while policies that support any deviation from that family model, such as migrant family reunification, are discouraged.

It is important to note that the importance of the traditional heterosexual family has been challenged in some iterations of far-right parties. Still, this nominal support for the LGBTQ community has largely been used as "a nationalist cudgel to attack Muslim immigrants and immigration (and, by association, all immigrants from the Global South regardless of faith) as threats to the nation state's defining liberal values" (Magni and Reynolds, 2023, 54). The

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[3] Italy had national-level gender quotas from 1993 to 1995, after which the quotas were deemed unconstitutional. Quotas for parliamentary seats were reinstated after constitutional reforms in 2003.

radical right parties[4] we investigate in this paper do not support LGBTQ rights, as is the case in much of Southern and Eastern Europe, due to the influence of the Catholic Church on matters of gender roles and sexuality (Indelicato and Magalhães Lopes, 2024).

Drawing from Ambivalent Sexism Theory (Glick and Fiske, 1997), Mudde (2019) suggests that far-right parties generally espouse *ambivalent sexism*, which is a mixture of two types of sexism, a so-called *benevolent sexism*, and *hostile sexism*. Benevolent sexism justifies the patriarchy in terms of *subjectively* positive attitudes towards women. Women are viewed as "morally good and physically weak" (Mudde, 2019). Benevolent sexism emphasizes the 'value' of the traditional roles of both men and women. Manly men must do the physical labor, protect 'their' women, and be aggressive. Womanly women must be the 'womb of the nation,' have children, and be motherly. The nature of sexual reproduction requires interactions across the sexes. This type of sexism suggests that women do not have power over their own lives beyond that which men give to them, or, what Secord (1983) terms *dyadic power*. Benevolent sexism recognizes that women do have some dyadic power, but suggests that women's roles are subsidiary and complementary to the roles of men.

On the other hand, hostile sexism justifies the patriarchy through derogatory characterizations of women. Hostile sexism characterizes women as "incompetent, overly emotional, and attempting to manipulate men to gain power" Hammond et al. (2018). Unlike benevolent sexism, which largely does not view women as a threat, hostile sexism views women as a threat to the patriarchy, although *largely* not a physical threat to men. This line of thinking has largely been popularized in online communities (incels, etc.) and has been the driving ideology behind mass casualty femicides in communities like Isla Vista in 2014 and Tallahassee in 2018. Mudde (2019) argues that hostile sexism has become more normalized in far-right spaces and parties, not just online, but offline. Thierry Baudet, leader of the Dutch *Forum voor Democratie* (FvD) has appeared on multiple podcasts with Andrew Tate, and supported him after his arrests for sex trafficking.

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[4] Namely, the *Lega* or *Lega Nord* and the *Fratelli d'Italia*.

Both hostile and benevolent sexism may cause one to disagree with measures designed to increase women's participation in politics. As Glick and Fiske (1997) point out, "both presume traditional gender roles and both serve to justify and maintain patriarchal social structures" (p. 121).

The analysis of gender *quotas* in relation to far-right support is an area yet unexplored in the literature. The growth and success of far-right parties since the 1990s coincides with a similar proliferation of countries implementing gender quotas. Before 1985, only four countries had instituted reforms that included quotas, rocketing up to 55 countries by 2005 (Tripp and Kang, 2008). These reforms occurred during a period of societal change that saw needs and values shift from security, well-being, and physical integrity to quality-of-life improvements (Inglehart, 1977). While on one hand opening the door to the political sphere to more women, these changes also saw increased demand for nationalistic and populist parties.

This project also brings together two parallel discussions in the literature on far-right support. First, there is a gendered component of far-right support. Men are more likely than women to support far-right parties (Givens, 2004). This is not to say that there are no women in far-right movements. Harteveld et al. (2015) find that men consider the issue positions of far-right parties to be more salient than women, meaning that men vote for the far-right more often because they place higher value on the programmatic core of these parties. Second, support for the far-right often stems from a backlash to multiculturalism and immigration; perceived cultural and social threats from outgroups are strong predictors of far-right support (Betz, 1993; Kitschelt, 1996).

To be clear, we are not arguing that all far-right voters are so opposed to women in power that they would not vote for a party that fields female candidates. The election of Giorgia Meloni as prime minister of Italy in October 2022 is one example of the fact that the party and a significant portion of voters *tolerate* female leadership. A native-born woman can still be the torch-bearer of far-right policies so long as she espouses the core ideological concerns

of the far-right, including the protection of traditional gender roles which also translate to the opposition of gender quotas (Indelicato and Magalhães Lopes, 2024).

Far-right parties are also strategic in their selection of female candidates. Weeks et al. (2023) argue that these parties appeal to women voters when they experience a gap in gender support and are struggling electorally by including more women on the ballot. As a result, more women have been included on far-right ballots in recent years.[5] However, this does not mean that all far-right supporters are on board with increased descriptive representation.[6] Despite these strategic incentives to include more women on the ballot, the far-right still consistently includes fewer women on average compared with other parties. For example, at the time of writing, the party with the lowest percentage of female MPs in the German parliament is the AfD, even though the AfD's current co-leader, Alice Weidel, is a woman. Further, there is a difference between the strategic incentives for a party to include more women on the ballot and the incentives for voters to select those candidates. In other words, just because the party leadership is acting strategically or progressively, that does not necessarily mean that all members of the voter base hold those same values.

We argue that institutional change that challenges the role of men in the public sphere triggers the threat response described here among a subset of voters. Voters are likely *not* reacting to there being more women in the legislature but are reacting to there being *restrictions* placed on the number of men permitted to run for local office.

The link between far-right vote share and opposition to gender quotas lies in perceived threat. In two nationally representative surveys in the US and Britain, Gest, Reny and Mayer (2018) find that supporters of the far-right are driven by a perceived discrepancy between their current social status and their views on past social standings. They coin the term "nostalgic deprivation" to refer to the cognitive evaluation that far-right supporters make of their place in the world. Crucially, perceived social and cultural threats are principal

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[5] See Guinaudeau and Jankowski (2024), and the response from Weeks et al. (2024) for a discussion on the impact of a statistical outlier, power, and modeling choices in Weeks et al. (2023).

[6] It may also be the case that as women's representation increases on far-right party ballots, the same individuals who react negatively to gender quotas may withdraw their support.

drivers of far-right success. Previous work on how perceived threat drives support for far-right parties has focused largely on immigration and multiculturalism, and how far-right movements have capitalized on xenophobia to attract support (Betz, 1993). Higher levels of migrant inflows are also correlated with far-right support (Knigge, 1998; Edo et al., 2019), which leads to perceived encroachment on economic opportunities or changes to the cultural fabric of society.

In Kitschelt's 1996 study on the far-right in contemporary Europe, he highlights that institutional change and control by a perceived corrupt elite that controls the lives of the average citizen as a catalyst for the rise in popularity of far-right parties. Further, there is evidence that far-right supporters oppose gender quotas specifically when introduced. In Finland, the Gender Equality Act of 1995 introduced mandatory gender quotas for publicly elected positions (Holli, 2022). Public opinion surveys found that 54% of men oppose the quotas, and this opposition is strongest amongst supporters of the far-right Finns Party Ylä-Anttila and Luhtakallio (2015).

Expectations about privilege and status that accompany citizenship or ethnicity (see Gest, Reny and Mayer, 2018) may be similar to expectations about the role of gender in the public sphere. Threats to these assumptions about maleness may trigger voting behavior similar to the kind we see when cultural threats are perceived. The causal mechanism behind the expected increase in far-right vote share is the belief among a subset of the male population that the status quo is shifting, thus limiting their influence in public life. This nostalgic deprivation that describes the comparison an individual makes between their objective position in life and where they think they should be (either at an individual or group level), prompts anti-immigrant voting behavior among white voters (Gest, Reny and Mayer, 2018). We anticipate a similar cognitive process occurring among male voters. When observing institutional change that increases women's participation in politics, we predict that some male voters will perceive this as a form of gender-based discrimination which will then trigger a threat response in the form of voting for parties that emphasize traditional

family values.

This leads to our main hypothesis:

H1: Implementing a gender quota will increase the vote share for mayoral candidates supported by the far-right

3   Context: The Italian Institutional Environment

Italy was one of the first EU member states to introduce gender quotas, with the passing of Law 81/1993, which mandated that in municipalities with fewer than 15,000 residents, neither gender could represent more than three-quarters of the candidates on each party list, and in municipalities with more than 15,000 residents, neither gender could represent more than two-thirds of the candidates on each party list.[7] This quota system was short-lived, as the Italian Constitutional Court ruled in 1995 that this particular part of Law 81/1993 was unconstitutional (Spaziani, 2022).

In the early 2000s, the Italian Parliament added amendments to the constitution that required that all levels of government actively promote gender equality.[8] This opened the doors for the two gender quota laws that we use to gain empirical leverage, namely Law 215/2012 and Law 56/2014, which are both still in place.

Law 215/2012 mandates that neither gender can represent more than two-thirds of councilor candidates on party lists in municipalities with more than 5000 residents, and both men and women must be represented on executive councils. Sanctions against municipalities take effect only after a successful legal challenge to a party list that violates the law. (Baraldi et al., 2023; Spaziani, 2022).

Law 56/2014 addresses the gender balance of executive committees, requiring that municipalities with more than 3000 residents have at least 40% of either gender, thus ensuring that women make up at the very least, a large minority of the executive committees. However,

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[7] Law 81/1993 also instituted direct election of mayors, changed the size of municipal councils, etc. See (Spaziani, 2022, 8)

[8] See Spaziani (2022) footnote 16 for more details on the exact measures passed in 2001 and 2003.

there are no real consequences when this 40% threshold is not achieved. If the threshold is not achieved, the mayor must write an official letter outlining their efforts to achieve the threshold (Spaziani, 2022).

Law 267/2000 sets up two other institutional factors that change at the 3000 resident threshold: the size of the executive committees *and* mayoral pay (Eggers et al., 2018).[9] This likely will lead to *sorting* and *compound treatment* across this threshold, which we address by following the advice in Eggers et al. (2018).

While these gender quota laws have been established, Italy has several special statute regions, namely Valle d'Aosta, Friuli Venezia Giulia, Sardinia, Sicily, and Trentino-Alto Adige, where Law 215/2012 and Law 56/2014 do not apply (Baraldi et al., 2023; Spaziani, 2022; Ferri and Grieco, 2022).[10]. These special statute regions were created as part of the Constitution of 1948 to protect their highly concentrated ethnic minorities, such as the significant German and Ladin-speaking minorities in Trentino-Alto Adige (Fessha, Kössler and Palermo, 2022, chp 8). These institutional thresholds create a particularly fruitful environment for empirically investigating the impact of gender quotas on far-right support. Many social scientists are using these population thresholds to study a variety of outcomes, including women's representation, and mafia-related crime (Eggers et al., 2018; Baraldi et al., 2023; Spaziani, 2022; Ferri and Grieco, 2022). Figure 3 shows the distribution of municipalities subject to gender quotas under Law 56/2014.

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[9] Mayoral Pay Increases and the Size of the Executive Council increases from 12 to 16.

[10] Spaziani (2022) does not include Sardinia as one of the regions exempt from Law 215/2012 (Spaziani, 2022, See footnote 17)
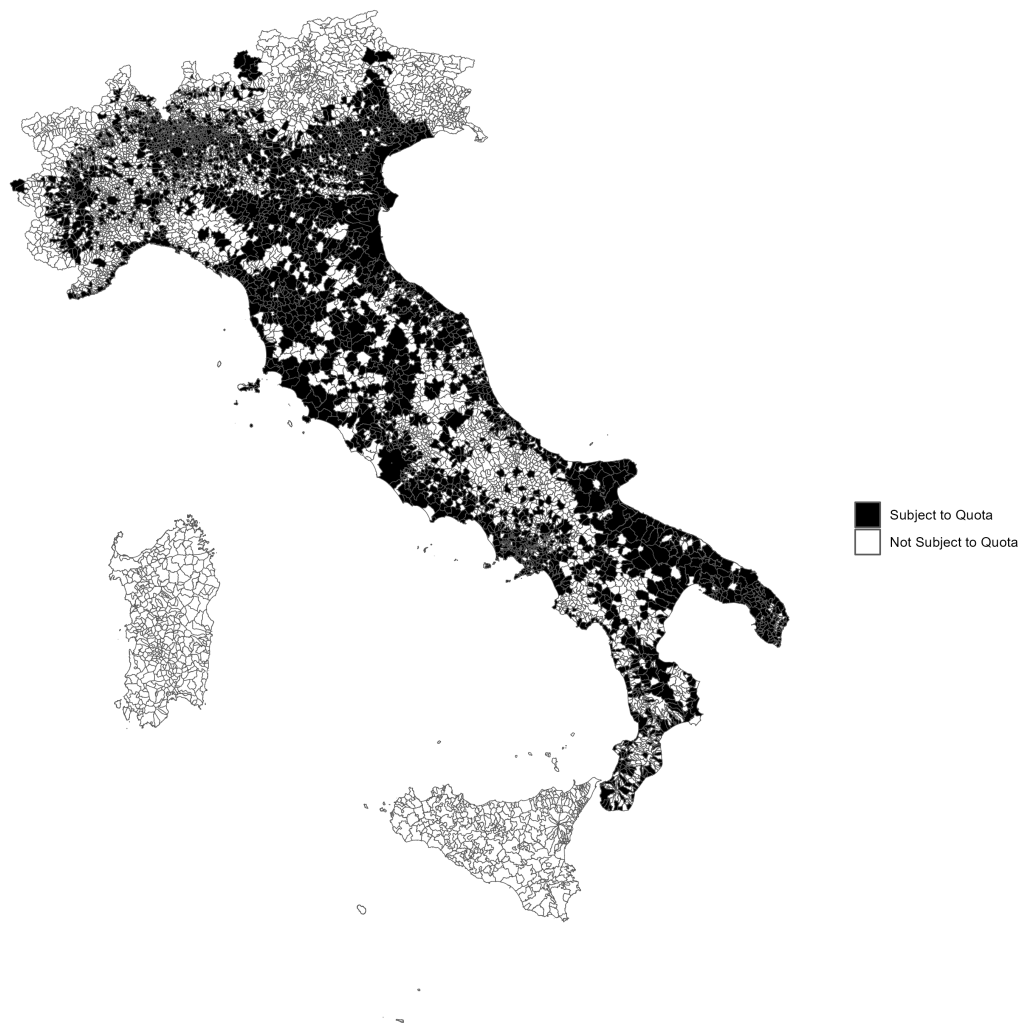
Distribution of Municipal Gender Quotas Note that any municipality subject to a gender quota under Law 56/2014 is coded as subject to a gender quota, even if they were not subject to a gender quota for the entire period covered by the data (i.e. treatment reversals are not shown).

Gender equality was and is a pressing issue. Google Trends data show that, directly prior to the passage of Law 56/2014 but after any elections in 2013, the relative importance of Gender Equality spiked. Law 56/2014 was passed on April 7th, 2014, and implemented the day after. The first municipal elections after the passage and implementation of Law 56/2014 happened on May 25th, 2014. Figure 3 shows the relative importance of searches

regarding Gender Equality and Law 56/2014 in Italy in 2014.
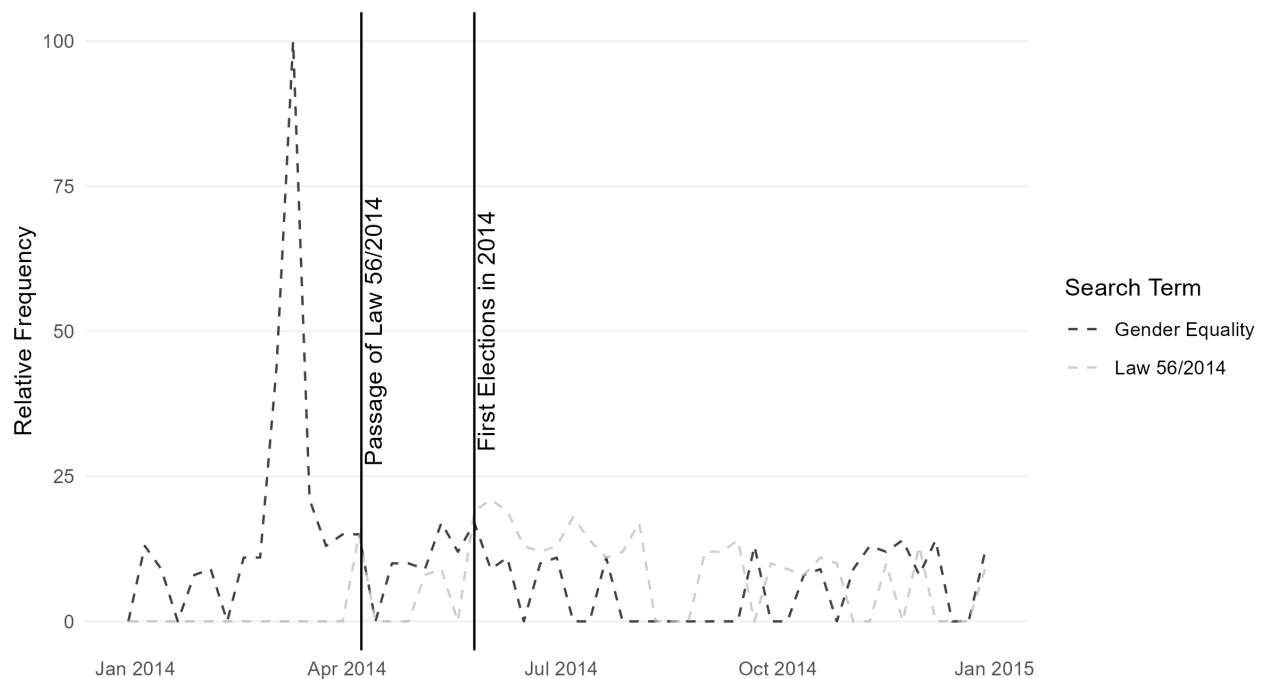
Google Trends for "Gender Equality" and "54/2014"

This spike in searches for Gender Equality suggests that there *may* have been some anticipation of the law. Still, there was no election before the passage of Law 56/2014 that this anticipation would have affected.

4   Empirical Analysis

To investigate the effects of institutional change on voting patterns, we focus on municipal elections in Italy. Our empirical analysis draws largely on publicly available local election data from the Italian Ministry of the Interior's Department of Business — Internal and Territorial[11] as well as municipal-level demographic data from the National Institute of Statistics Population and Households census[12]. Our data covers the time period from 1993 to 2021.

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[11] https://dait.interno.gov.it/elezioni/open-data
[12] https://www.istat.it/en/

We measure our dependent variable in two ways, including 1. vote share for mayoral candidates supported by far-right parties, and 2. the proportion of councilors supported by far-right parties. Our preferred measure, vote share for mayoral candidates supported by a far-right party, has several advantages. First, it is more *precise* than the proportion of councilors supported by far-right party lists. Second, it is unaffected by one of the major institutional changes at the population threshold where gender quotas are implemented, namely, the size of municipal councils (Eggers et al., 2018). This means that the likelihood of *compound treatment* decreases. Mayoral elections are also not subject to gender quotas, and the implementation of council-level quotas has not increased women's representation (or electoral support) in mayoral elections (Spaziani, 2022), meaning that any change in vote share for far-right candidates is not driven by a change in women's representation.[13] The share of votes received by far-right mayoral candidates is measured as a continuous proportion between 0 and 1. We identify far-right mayoral candidates as those who run on lists supported by either Lega (Nord) or the Brothers of Italy parties. Our results are robust to our alternative measure, whose results are reported in Appendix B.

Our primary independent variable is a binary indicator of quota implementation, equalling 1 if a municipality has reached the minimum population threshold for executive committee quotas (3,000) after the Law 56/2014 reforms (which went into effect in April 2014), and 0 otherwise. Some regions have special provisions which allow additional autonomy with regards to Law 56/2014 (Baraldi et al., 2023), meaning they are not subject to the same institutional change as other regions and creating a group of never-takers in addition to small towns never reaching the population threshold.

4.1   Matching

We begin by truncating the sample to only include municipalities with fewer than 15,000 residents. At that threshold, additional electoral rule changes take effect that complicate

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[13] Women are less likely to be candidates for far-right parties, thus changes in the gender makeup of mayoral candidates support would mechanically change the support for far-right mayors.

the analysis. First, voters can express an additional preference vote on a candidate list beyond the first. Second, elected mayors can select executive councilors from the electorate (Spaziani, 2022), thus changing the proportion of city councilors. Third, mayoral elections shift from a single-ballot system to a run-off/double-ballot system (Benedetto, Destefanis and Guadalupi, 2023). We are left with 2,823 treated units (towns above the 3,000 threshold after April, 2014) and 35,133 control units. Since the implementation of a gender quota is not randomly assigned, we need to address systematic differences that may exist between treatment (municipalities that implement quotas) and control (those municipalities that do not implement quotas). To ensure that we can responsibly compare treated and control units, we pre-process the data to match based on several covariates. We match units based on pre-2013 (i.e. pre-treatment) values of population, far-right mayoral vote-share, and the number of male voters in the electorate[14]. Matching on pre-treatment indicators avoids post-treatment bias in the final results.

For robustness, we present two matching strategies. First, we use 5-to-1 nearest neighbor matching due to the sample size differential between treatment and control. This propensity score weighting scheme allows us to imagine probabilistic assignment to treatment and control based on our set of pre-treatment covariates. 14,115 control units are matched to treated units under this scheme. Second, we employ entropy balancing (Hainmueller, 2012). Instead of discarding unmatched control observations like in the propensity score method in 5-to-1-nearest neighbor matching, entropy balancing allows us to retain all information from each control unit. The reweighting scheme in entropy balancing maintains proximity to the base weights, thus ensuring that control observations that most closely match the treated units are given the most weight.

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[14] For summary statistics, see Appendix C.

4.2   Estimation Strategy

After April 2014, the implementation of gender quotas is based on population size, meaning in any given year a municipality may cross the requisite 3,000 resident threshold. Due to the staggered nature of treatment assignment, we opt for a staggered difference-in-difference design (see Callaway and Sant'Anna, 2021). This approach allows us to apply the difference-in-difference framework to a context where treatment assignment occurs in multiple time periods. As communities grow and sometimes move from control to treatment status after the blanket quota adoption in 2014, researchers must consider additional time periods beyond the canonical pre- and post- treatment framework. This approach also assumes the absence of treatment reversal. In our context, this would insist that no municipalities that had implemented a gender quota after clearing the 3,000 threshold fell below that mark and dismantled the quota in a subsequent election. This is a stout assumption. However; we identify only three towns that qualify as reversal cases in the truncated sample. We therefore omit these three municipalities and their elections from the analysis ($N = 16$).

The staggered DiD design differs from the canonical DiD framework in several crucial ways. First, the main reported causal parameter is considered a "group-time average treatment effect" (Callaway and Sant'Anna, 2021), where groups are identified by their first-treated time period. This treatment effect is identified for members of each of these groups (g) at time t. We report the aggregation of these effects, which are the averaged effects of participating in treatment among all units who ever participated in treatment.

Second, a typical concern with DiD designs is that of anticipation (see Abbring and van den Berg, 2003). In our case anticipation of treatment effects is very likely for two reasons. First, the blanket national Law 56/2014 was applied to all municipalities aside from the special statute regions at the same time. Well-informed voters would have likely known about the passage of the law before its implementation. Second, municipalities approaching the 3,000 population threshold that would trigger quota requirements are likely aware of the impending institutional change. Municipalities near that threshold in 2022 could have

anticipated treatment as far back as 2014 and adjusted behavior accordingly. However, the staggered DiD design allows researchers to minimally account for anticipation effects in the estimation by introducing a conditional parallel trends assumption and estimating the ATT nonparametrically accounting for a maximum value of possible treatment anticipation.

There are natural questions of endogeneity when examining observational data, particularly when studies focus on institutional reform. Namely, quotas can be accompanied by other institutional changes that could drive support for the far-right. For example, the Italian constitutional reform in 1993 that implemented the first round of gender quotas was triggered by a series of corruption scandals (Weeks and Baldez, 2015) that could also fuel populist voting patterns. Further, we have only demonstrated that quota implementation has a positive effect on far-right vote share among municipalities that have implemented a gender quota *on paper*, but not whether councils are compliant with these reforms.

To address these concerns, we identify an instrumental variable in the form of *compliance* with gender quota implementation. We create a binary measure of compliance that equals 1 if a municipality has implemented a gender quota *and has achieved the legal 40% minimum proportion of elected women councilors*, and 0 otherwise. This identifies which municipalities have *de jure* equality, and which ones achieve *de facto* equality. This is a matter of treatment exposure. If a town meets the 3,000 resident threshold but for some reason (whether internal or external) does not implement the gender quota, then residents may not even be aware that the quota should be in place. We argue that gender quotas should only influence support for the far-right when a municipality has achieved that minimum legal compliance.

The exclusion restriction implies that the instrument should not be correlated with the outcome measure, and should only impact the dependent variable through the explanatory variable. Thus compliance with quotas should only affect voting for the far-right when quotas are actually in place. Indeed, we find that compliance with quotas correlates with quota implementation ($\phi = 0.58$) and only weakly correlates with far-right mayoral vote

share in a town in given year ($\phi = 0.13$)[15]. Further, compliance with gender quotas should not be affected by the proportion of votes received by far-right mayors during the same year they are elected; mayors have no control over the composition of the council itself. Municipal noncompliance with the quotas may in fact be unintentional, with election results determining the final composition. As a result, we consider compliance in terms of treatment exposure; voters will only react to policies fully in practice.

We use this instrument in a two-stage least squares (2SLS) analysis. We estimate the first stage via logistic regression, with the treatment indicator of quota implementation regressed on the compliance instrument as well as the pre-treatment covariates and municipal and year fixed-effects. In the second stage we regress our outcome variable on the fitted values from the first stage in addition to the controls and fixed effects listed above.

5   Results

5.1   Difference-in-Difference

We report our results from the staggered DiD models in Table 1. All reported models cluster standard errors at the municipal (i.e. treatment assignment) level. Estimates report the Average Treatment Effect for the Treated (ATT). Model 1 reports the results for the unmatched sample, while models 2 and 3 provide the ATT for the 5-1 nearest neighbor matched (restricted) sample and the entropy balanced (weighted) sample, respectively. While the unmatched sample does not achieve statistical significance, both of the matched samples show a positive and statistically significant effect of gender quota implementation on the vote-share for far-right mayoral candidates. This substantive increase ranges between 5.3% and 7.3%. The restricted sample reports the smallest treatment effect at 5.3%, while the weighted sample demonstrates a larger 7.3% ATT.

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[15] See Appendix B.1 for details.

Tab. 1: Effect of Gender Quotas on Far-Right Vote Share

	Dependent variable:		
	Far-Right Vote Share		
	Unmatched	5-1 Matched	Entropy Balanced
	(1)	(2)	(3)
ATT	0.0506	0.073*	0.0531*
	(0.0299)	(0.0308)	(0.0249)
	[-0.0081, 0.1092]	[0.0128, 0.1333]	[0.0043, 0.1019]
Observations	37,940	13,839	37,940

* 95% CI's do not cross 0.

Clustered standard errors with 1,000 bootstrapped iterations in parentheses. Dependent variable: far-right vote share, measured 0 to 1, continuous. Independent variable: Implementation of gender quota, equalling 1 if a municipality has achieved the 3,000 resident threshold after the 2014 reforms, 0 otherwise. Matching variables include: pre-treatment vote share of far-right mayors, pre-treatment proportion of men in the electorate, and pre-treatment census population.

Figure 1 plots the aggregated ATT for each time period after Law 56/2014 went into effect. The first panel plots the effect for the restricted sample and the second panel plots the effects for the weighted sample. The largest ATT's appear in the one to two years immediately following quota implementation. This suggests that the reaction toward quota implementation may in fact be a result of the national blanket institutional reform rather than realized changes in local conditions.
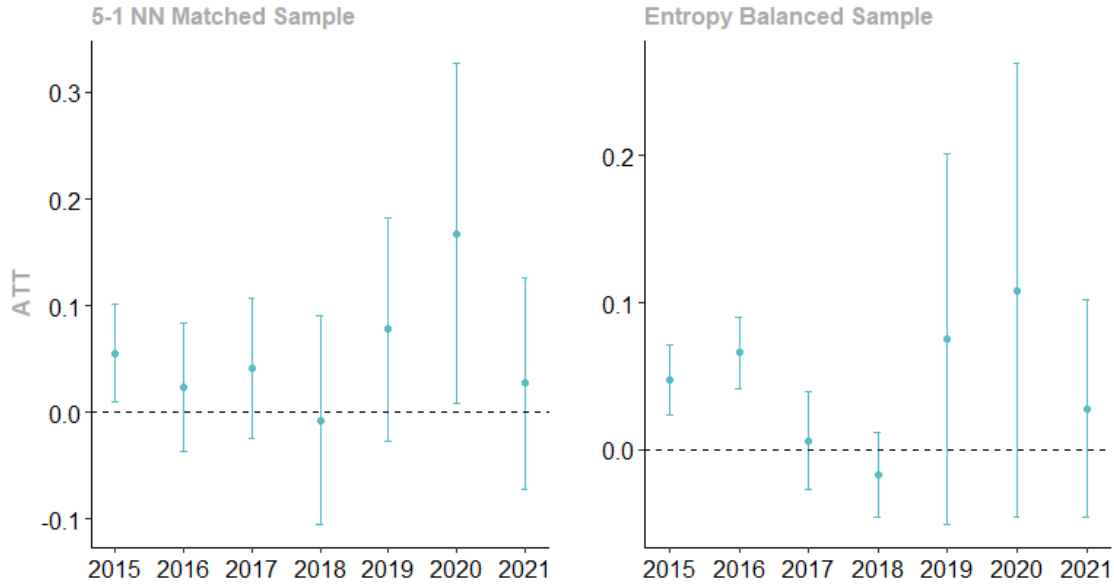
Fig. 1: ATT Over Time

5.2 Endogeneity

The difference-in-difference approach assumes equal exposure to the intervention across all individuals within the treatment groups. This is a strong assumption; in the case of gender quotas it assumes that individuals are politically knowledgeable enough to recognize a quota has been implemented, even without the municipality complying with the mandate. Among the 2,819 treated units for which we have compliance data, 2,129 are in compliance with Law 56/2014. This means the proportion of women on executive councils meets or exceeds 40%. The other 690 treated units have towns above the requisite population threshold but have not met the 40% minimum. An absence of enforcement mechanisms means that individuals living in noncompliant communities may not even be aware that a quota is *de jure* but not *de facto*.

To address this issue, we instrument on compliance with gender quotas. The compliance indicator equals 1 if a town is above the minimum population threshold after Law 56/2014 went into effect *and* if it has achieved the 40% proportion of women on the executive council,

0 otherwise. We then estimate a two-staged least squares (2SLS) model for the unmatched sample, the nearest-neighbor matched sample, and the entropy balanced sample. Table 2 reports the results from these models[16]. The 2SLS estimates report a Local Average Treatment Effect (LATE)[17], or the effect of implementing a gender quota on the proportion of the vote received by far right-supported mayors among the towns where the impact of quotas on women's representation is most clearly visible.

Tab. 2: Instrumental Variable Estimates

	Dependent variable:		
	Far-Right Vote Share		
	Unmatched	5-1 Matched	Entropy Balanced
	(4)	(5)	(6)
LATE	0.078***	0.018**	0.078***
	(0.006)	(0.007)	(0.008)
Observations	33,639	12,851	33,639
Wald	440, $p = 2.2e^-16$	8.331, $p = 0.0039$	710.6, $p = 2.2e^-16$
RSE	0.111	0.159	0.06
Adj. R^2	0.024	0.004	0.041

*p<0.05; **p<0.01; ***p<0.001

Clustered standard errors by municipality in parentheses. Municipal and year fixed effects included. Dependent variable: far-right vote share, measured 0 to 1, continuous. Independent variables: Implementation of gender quota, equalling 1 if a municipality has achieved the 3,000 resident threshold after the 2014 reforms, 0 otherwise; pre-treatment values of far-right vote share, population, and proportion of male voters in the electorate. Matching variables include: pre-treatment vote share of far-right mayors, pre-treatment proportion of men in the electorate, and pre-treatment census population.

In all samples, the effect is in the expected direction and significant at the 0.01 level. Comparing the estimates with the ATT's reported in Table 1, the effect sizes are 2% larger

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[16] First-stage estimates are reported in Appendix B.1.

[17] Model 6 of Table 5 maintains the entropy weights in the estimation, and so reports a *weighted* LATE.

in the unmatched and entropy balanced (full) samples, indicating that the boost in vote share that far-right mayoral candidates experience is more impactful among constituencies in compliance with their legal equality mandates. The 5-1 matched sample's effect size is smaller in the 2SLS estimate than in the staggered DiD, which may be a result of the truncated sample size. These results are robust to swapping the mayoral vote-share dependent variable with a proportion of far-right councilors variable, see Appendix B.1 for details.

In our setup, the only way for compliance with gender quotas to affect far-right vote share is through the introduction of a quota in the first place. Municipalities that both implement and adhere to the quota requirements after their town crosses the minimum population threshold experience an average increase in far-right vote share between 1.8% and 7.8%. In towns below the 15,000 population threshold, elections do not use the majoritarian run-off system, meaning candidates who win a plurality of the vote are capable of winning. In some cases, arguably negligible vote shares may be enough to clinch victory. For example, in 2012 Mayor Marco Facchinotti of Mortara, Lombardia, who was supported by the Lega party, won his election with a vote share of just 22.9% (1,716 of the 7,487 votes cast in the election). A boost of 7% can have major impacts on the outcomes of an election.

Wald tests for all samples are statistically significant, indicating that the instrument is correlated with the endogenous variable (quota implementation) and has an effect on the outcome (far-right vote share). In the truncated sample, the Wald estimate is low, possibly indicating a weak instrument, but the result still indicates that the instrument affects far-right vote through quota implementation. This demonstrates that municipalities that have direct exposure to quota implementation through compliance generate higher vote shares for the far-right.

Taken together these results suggest that there are downstream backlash effects to the implementation of gender quotas. This Far-right vote share improves in municipalities that have implemented a gender quota. Our exogenous variable, compliance, indicates that municipalities that enforce a quota experience generally larger shares of the vote moving to

candidates supported by far-right party lists.

6   Conclusion

Gender quotas are purposed to remove obstacles to women's representation in government. We argue in this paper that there are downstream backlash effects to efforts to improve women's representation among individuals who perceive gender equality as a threat to status quo gender norms and power dynamics. This backlash manifests in the form of voting for the parties that most vehemently position themselves as guardians of the traditional family structure. We demonstrate these backlash effects in the context of municipal elections in Italy, where candidates supported by far-right lists experience a vote share increase between 5% and 7% after quotas are implemented.

Potential far-right voters are reacting to institutional change that limit male agency. The far-right relies on claims of merit; that no institutional support should be required for minority or women politicians to secure office. When this viewpoint is challenged, and when institutions are crafted to improve access to authority among underrepresented groups, votes are channeled from those who feel most threatened by these changes to the parties whose platforms most closely reflect a reversion to the status quo.

Our results indicate that far-right mayoral candidates benefit from blanket institutional reforms that restrict access to positions of authority amongst the male demographic that is key to far-right success. This has important implications for future policy reforms. First, reformers are likely not considering how quota policy might disproportionately affect one party family over others. Second, reform efforts may be jeopardized by an increased vote share for parties that affirm traditional values that may limit women's access to positions of authority.

This project contributes to the literatures on women's representation and "demand-side" explanations of far-right voter behavior (see Golder, 2016). While modest, the results we present are consequential for our understanding of how far-right voters react to institutional

change. Our conceptual contribution combines existing theoretical work on ambivalent sexism and applies it to institutional reform.

It is important here to reaffirm the importance and effectiveness of gender quotas in achieving gender equality, in addition to the positive downstream outcomes that they also produce. While quotas may be provoking a specific subset of the population to vote in a particular direction, quotas are only an addition to a long list of societal grievances that potential far-right voters may latch on to.

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A   Balance

Table 3 provides the summary and balance statistics of the pre-treatment patching variables. There are clear differences between municipalities that do and not implement gender quotas in all three of our covariates — far-right vote share, number of male voters, and population. This justifies and necessitates our matching approach.

Tab. 3: Balance Statistics for Pre-Treatment Covariates

	Mean (Control)	SD (Control)	Mean (Treat)	SD (Treat)	p-Value
FR Vote Share	0.050	0.134	0.116	0.183	$1.855075e-73$
Male Voters	995.747	996.378	2,078.323	941.720	0
Population	2,975.754	3,065.884	6,319.236	2,976.154	0

B   Alternative Measures & Specifications

Our results in the main analysis are robust to alternative measures of the dependent variable. As an additional measure, we consider the proportion of far-right councilors elected in a given municipality. This is computed by first identifying which councilors run under far-right supported lists. Due to the fact that two-thirds of Italian districts nominate candidates under a closed-list proportional representation system, we are unable to identify which councilors are nominated by specific parties. Rather, the measure that we create reflects more specifically the proportion of councilors that have won a seat on a municipal council who run on lists that far-right parties have endorsed[18].

We opt for a proportional variable, ranging between 0 and 1, over a simple count measure of far-right supported councilors for two reasons. First, in addition to gender quota implementation at the 3,000 population threshold, municipalities also increase the size of city councilors (Eggers et al., 2018), thus potentially confounding observed increases in raw

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[18] Multiple parties may support the same list, making it impossible to separate those councilors who are actively supported by far-right parties from those whose parties who may cooperate enough with the far-right to jointly support a candidate list. However, we argue that a *list* which is supported by a far-right party winning seats on a municipal council is a decent proxy for far-right supported *candidates*.

counts of far-right councilors. Any change in the *proportion* of far-right supported coun-
cilors after quota implementation should be an artifact of the institutional change and not
an arbitrary increase in the total number of councilors.

These analyses use an alternative set of covariates for matching as well. We match
units within the same province as well as based on pre-2012 (i.e. pre-treatment) values of
population, the proportion of far-right councilors, and the proportion of women councilors.
We are left with 4,501 treated units (towns above the 3,000 threshold after April, 2014) and
43,490 control units.

We report two sets of analyses here. First, we reestimate the staggered DiD reported
in Table 4 using the proportion of far-right supported councilors instead of far-right vote
share. Second, we report results from traditional generalized difference-in-difference, or two-
way fixed-effects (2WFE) models to account for time-invariant and municipality specific
confounders. The municipal fixed-effects account for varying sizes of the municipal council,
the propensity to nominate (or screen out) far-right politicians, economic conditions, and
other unobservables. This generalized approach necessitates relaxing certain timing assump-
tions in the canonical difference-in-difference setup as well as the parallel trends assumption
(Callaway and Sant'Anna, 2021). We estimate DiD models first accounting for the bivariate
relationship between gender quota implementation and change in the proportion of far-right
councilors on the matched samples. We then include the pre-treatment covariates using the
specification below:

$$y_{it} = \alpha_i + \gamma_t + \mathbf{X}_{it}\beta + w_{it} + \epsilon_{it}$$

Where our outcome of far-right councilors y_{it} is a function of municipality fixed-effects α_i,
year fixed-effects γ_t, the presence of a quota in effect during the election, the pre-treatment

covariates of the proportion of far-right and women councilors and population size, and finally the series of weights w_{it} derived from the matching algorithms described earlier. Table 4 reports the results.

Model 1 reports the results from the bivariate generalized (TWFE) DiD. As predicted, the effect of institutionalizing a gender quota on the proportion of far-right councilors is positive and statistically significant at the 0.01 level. The model reports an increase of 5% in the proportion of far-right councilors. Model 2 reports the results from the 5-to-1 matched sample. Similarly, we find a positive change to the proportion of far-right councilors, significant at the 0.01 level. Model 2 predicts a positive change of 2.3%. Model 3 reports the results from the entropy balanced sample. The estimate remains similar, predicting an increase of 2.5%, while the standard error increases slightly resulting in a decrease in significance to the 0.1 level.

Tab. 4: Difference-in-Difference Estimates

	Dependent variable:		
	Far-Right Councilor Proportion		
	Bivariate DiD	5-1 Matched	Entropy Balanced
	(1)	(2)	(3)
Quota	0.050***	0.023***	0.025*
	(0.004)	(0.004)	(0.014)
Far-Right Proportion		0.298***	0.305***
		(0.011)	(0.064)
Women Proportion		0.015*	0.017**
		(0.008)	(0.008)
Population		0.000***	0.000***
		(0.000)	(0.000)
Constant	0.019***	−0.049***	−0.072***
	(0.001)	(0.010)	(0.023)
Observations	47,948	27,006	47,948
Adjusted R^2	0.0113	0.2284	0.2566
F-statistic	549.5	18.92	31.84

*p<0.1; **p<0.05; ***p<0.01

Clustered standard errors in parentheses. Individual municipal and election year fixed-effects output omitted. Dependent variable: proportion of far-right councilors in a given municipality in an election year, measured 0 to 1, continuous. Independent variable: Implementation of gender quota, equalling 1 if a municipality has achieved the 3,000 resident threshold after the 2012 reforms, 0 otherwise. Control variables include: pre-treatment proportion of far-right municipal councilors, pre-treatment proportion of women councilors, and pre-treatment census population.

Naturally, the pre-treatment proportion of far-right councilors has a positive effect on this proportion in the post-treatment period. This effect is large, reporting a change of roughly 30% per unit increase in proportion of far-right councilors. This may reflect some path dependency — as parties win seats on municipal councils, they may enjoy positive

incumbency effects that make it easier to field additional successful candidates.). In the Italian PR context specifically, individual legislators have experienced positive incumbency effects when seeking relisting on a party's ticket and party elites enjoy reelection benefits (Golden and Picci, 2015). This simply means that parties who successfully field candidates are likely to re-field winners. At the local level, our results suggest that these renomination benefits may also translate to campaign entry for additional partisans.

Towns that elect councilors from a certain party may also simply be more likely to elect additional co-partisan councilors in future elections. Indeed, recent work has shown that Italian mayors who present themselves as "outsiders" or "anti-establishment" are more likely to be selected with the split vote (Freschi and Mete, 2020). This may give parties who successfully utilize their populist appeal to secure a first councilor an advantage in fielding additional candidates later on.

Controlling for the proportion of women councilors in the pre-treatment period also reveals a positive change on the proportion of far-right councilors. These results fit the narrative that far-right voters may perceive changes in the proportion of women councilors to be some existential threat. Framing gendered political issues in terms of expectations regarding the roles of men and women both at home and in society (Donà, 2021, see) presents tension when these roles are challenged (e.g. when women join the political sphere). A 1% increase in the proportion of women on a municipal council leads to a 1.5% to 1.7% increase in the proportion of far-right councilors, adding an additional far-right half-councilor for every woman who joins.

Population size also a positive effect. Larger towns have higher proportions of far-right councilors. This of course may be due to the institutional rules that increase municipal council sizes at certain population thresholds, further necessitating our instrumental variable approach to account for the endogeneity inherent to institutional change based on population.

B.1 2SLS

Tab. 5: Instrumental Variable Estimates

	Dependent variable:		
	Far-Right Councilor Proportion		
	Unmatched	*5-1 Matched*	*Entropy Balanced*
	(4)	(5)	(6)
Quota (IV Fitted)	0.020***	0.024***	0.026**
	(0.004)	(0.007)	(0.008)
Far-Right Proportion	0.304***	0.297***	0.304***
	(0.003)	(0.004)	(0.003)
Women Proportion	0.005	0.015**	0.017***
	(0.003)	(0.006)	(0.005)
Population	0.00000***	0.00000***	0.00000***
	(0.00000)	(0.00000)	(0.00000)
Constant	−0.037	−0.049***	−0.072
	(0.062)	(0.007)	(0.079)
Observations	47,948	27,006	47,948
Adjusted R^2	0.217	0.228	0.257
Wald Test	25.64 on 537 and 47453 DF	18.87 on 446 and 26559 DF	31.74 on 537 and 47453 DF

*p<0.1; **p<0.05; ***p<0.01

Clustered standard errors in parentheses. Individual municipal and election year fixed-effects output omitted. Dependent variable: proportion of far-right councilors in a given municipality in an election year, measured 0 to 1, continuous. Independent variable: fitted values from first-stage logistic regression of binary treatment indicator on compliance instrument and pre-treatment covariates. Control variables include: pre-treatment proportion of far-right municipal councilors, pre-treatment proportion of women councilors, and pre-treatment census population.

Table 5 reports a similar 2% to 2.6% increase in the proportion of far-right councilors after

quotas are implemented. Consistent with expectations, the estimates for the matched samples increase in comparison to those in Table 1, but only by less than 1%. Importantly, precision improves in the entropy balanced sample when compared to the difference-in-difference estimates, raising the significance level from 0.1 to 0.05. Estimates for the pre-treatment covariates are comparable with those in Table 1. The proportion of far-right councilors maintains a large effect, between 29.7% and 30.4%, reflecting some path-dependence. The estimates for women's proportionality and the effects of population change remain unchanged in comparison with the difference-in-difference estimates.

C   Balance

Figure 2 plots the distribution of propensity scores from the 5-to-1 nearest neighbor matching scheme.
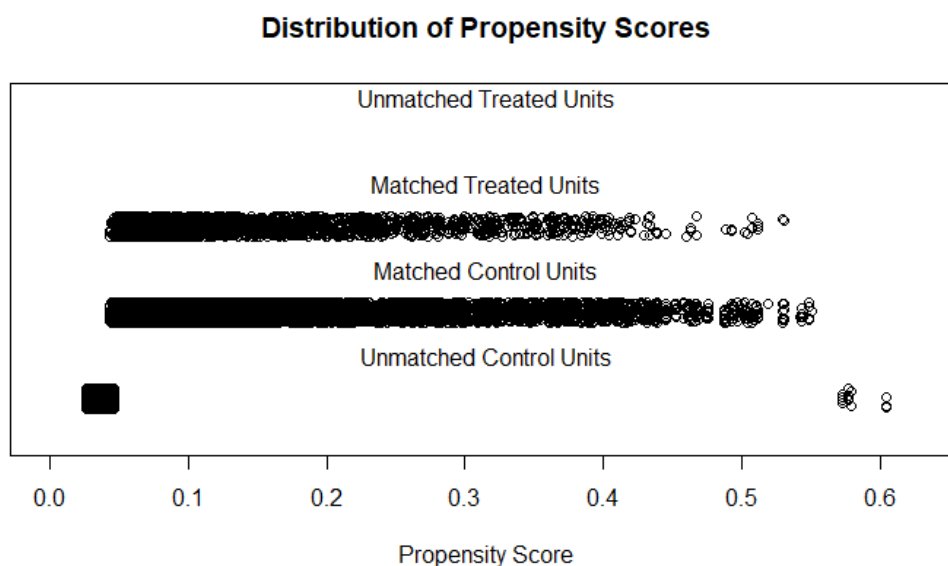

Fig. 2: 5-to-1 Nearest Neighbor Propensity Scores

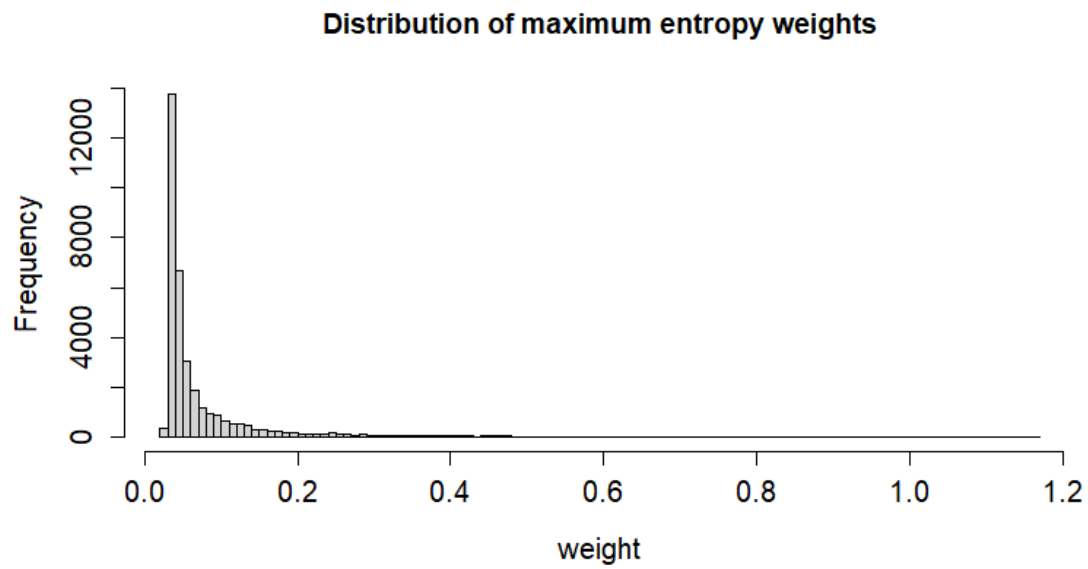Figure 3 plots the distribution of entropy-balanced weights.

Fig. 3: Entropy Balancing Weights

D   2SLS

Figure 4 Shows the correlation between the implementation of gender quotas and compliance with quotas.
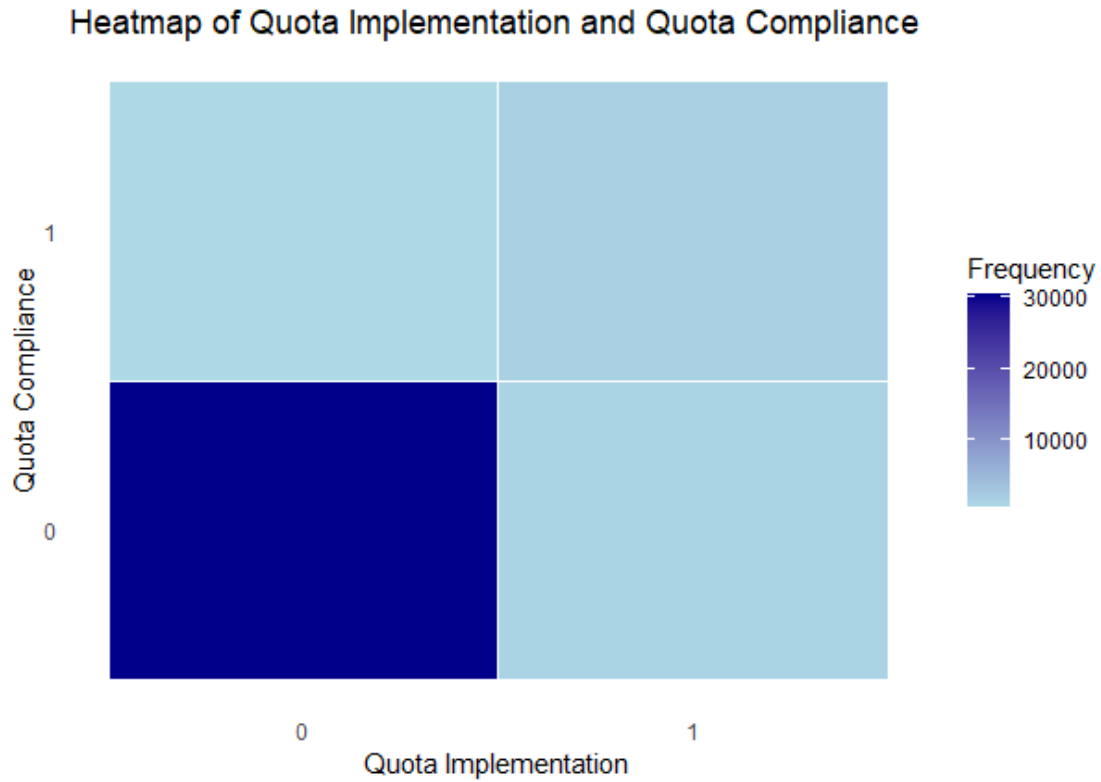
Fig. 4: Correlation Heatmap of Quota Implementation and Quota Compliance

For our alternative dependent variable, the proportion of far-right councilors elected in a given municipality, we report first-stage estimates predicting treatment assignment as a function of compliance with gender quotas and pre-treatment covariates are reported in Table 6. Compliance with gender quotas does not predict implementation with the quotas themselves.

Tab. 6: First-Stage Estimates

	Dependent variable:		
	Quota Implementation		
	(1)	(2)	(3)
Compliance	37.581	34.419	27.776
	(1,598.622)	(2,905.131)	(1,800.460)
Far-Right Proportion	1.724***	1.512***	1.630***
	(0.171)	(0.505)	(0.442)
Women Proportion	−0.496**	2.312***	−1.711***
	(0.224)	(0.763)	(0.506)
Population	0.001***	0.008***	0.002***
	(0.00002)	(0.0005)	(0.0001)
Constant	−50.754	−107.864	−51.324
	(31,098.700)	(175,933.400)	(33,269.210)
Observations	47,948	28,200	47,948
Log Likelihood	−2,978.515	−314.540	−317.454
Akaike Inf. Crit.	7,033.031	1,523.081	1,710.909

*p<0.1; **p<0.05; ***p<0.01