Organized crime and women in politics: Evidence from a quasi-experiment in southern Italy

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Evidence from a Quasi-Experiment in Southern Italy

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Abstract

This paper provides new and unexplored evidence of a negative link between an increasing female participation in politics and the infiltration of organized crime in government. We perform an empirical analysis of about 1,700 Southern Italian municipalities between 1985 and 2013 exploiting two Italian laws: law no. 164/1991, which allows measure of mafia infiltration in the Italian municipalities, and law no. 81/1993, which creates an exogenous source of variation in the share of women on the council that allows for correction of endogeneity bias. Increasing the female proportion on the city council of 10 percentage points reduces the probability of dissolution for mafia infiltration of about 1.8 p.p.; the result is confirmed when considering a female mayor. This negative effect remains across several robustness checks. This research adds a further reason in favour of the reduction of the gender gap in politics. In fact, policies aimed at legitimizing democracy, such as gender quotas in electoral law, also have the effect of strengthening institutions in the fight against organized crime, which is always a key government agenda.

JEL Classification: D72, D78, J16, K42

Keywords: organized crime, gender gap, quasi-experiments, panel probit model.

1. Introduction

Arguments in favour of an increasing female participation in politics address the legitimation of democracy and the adoption of practices that improve the quality of institutions (Epstein et al., 2005). Although the number has increased over the past 20 years, women are still underrepresented in several contexts, especially in government.[1] Academic research on the gender gap in politics has investigated the effect of policy measures aimed at encouraging female presence in politics. The rationale of those policies relies on the notion that stimulating a more equal representation of women in government leads to a gain in performance and quality of government, because men and women tend to perform differently in similar contexts. Nevertheless, no conclusive evidence describes whether women's behaviour is better than men's one in policy setting.

Relying on a quasi-experiment, we contribute to this debate by providing evidence of a further argument in favour of a greater representation of women in political institutions, related to organized-crime infiltration in politics. To the best of our knowledge, this issue has not yet been investigated in the gender literature. Political economy literature referred to sociological and psychological characteristics of women to justify that they tend to be less corrupt than men and to promote more honest government (Swamy et al., 2001; Dollar et al., 2001). The several studies that follow underlined that women, as the "fairer sex", have a 'higher moral nature and propensity to bring their finer moral sensibilities on public life, and particularly on the conduct of politics' (Goetz, 2007); moreover, they are less likely to sacrifice the common good for personal (material) gain. This role of women in government may be particularly relevant because one of the most significant difficulties faced by governors is the design of institutions that discourage their agents from acting opportunistically, at the expense of the public interest. Thus, increasing direct female participation in government could help mitigate the lack of responsiveness of governments and enhance the quality of institutions and organizations (Staudt, 1998).

One of the consequences of stronger institutions can be deterrence toward all the form of crime and, in particular, toward organized crime that, in different ways, tries to infiltrate politics. Indeed, although criminal organizations are usually involved in a wide range of illegal activities in the economy (e.g., supplying illicit goods and services, extorting individuals or firms, and in some cases, offering private security) they also try to influence politics. Organized crime, through violence (in rare cases), threats, and, above all, corruption, tries to infiltrate political bodies; the infiltration

[1]At the beginning of 2019, approximately one fourth of the members of lower or single houses of parliament were women. At the end of 2018, women's membership in parliaments rose to 24%, representing an increase of 13% with respect to the past two decades. In 2018, the number of women presiding over houses of national parliaments was almost 20%; similarly, only 18% of appointed ministers are women, and they are usually assigned portfolios related to environment, natural resources, and energy, followed by social sectors. This means women are still largely excluded from the executive branches of government (European Parliamentary Research Service).

consists of constraining and influencing the choices of civil servants to oversee management of city life in public procurements, public works, urban plants, housing, etc. The result is a huge economic gain for criminal organizations and a loss for society. Thus, on one hand, policies aimed at reducing the underrepresentation of women in politics, beyond equality reasons, lead to institutional improvements. On the other hand, stronger institutions discourage the proliferation of crime (Buonanno et al., 2015). Literature has not yet investigated the effect of lowering the gender gap in politics on organized crime involvement in government; our paper fills this gap by addressing whether a greater share of women in political bodies may prevent the influence exercised by criminal organizations on politicians. Because women have higher standards of ethical behaviour and are more concerned with the common good, they may be naturally less tolerant of all the forms of violation or alteration of morality, like corruption (as well documented) and, in particular, crime (as we assert).

We perform an empirical analysis on the effect of an increase in the fraction of women in political offices on the probability of mafia infiltration. We focus on southern Italian municipalities where criminal organization has been traditionally concentrated. We exploit law no. 164/1991 that prescribes the dissolution of the local government due to mafia infiltration, to capture the presence of organized crime in political institutions. This law states that to dissolve a municipality, the "Commissione Parlamentare Antimafia" must ascertain, above all, the direct or indirect links of local administrators with organized crime or, alternatively, the conditioning that the mafia imposes on administrators. We estimate linear probability model (LPM) and probit models, where the binary dependent variable takes a value of 1 if the municipality has been dissolved and 0 otherwise (according to different operationalizations) and the regressor of interest is the share of women on political city councils. To resolve the potential endogeneity problem in the relationship of interest, we exploit law no. 81/1993, introducing gender quotas in local electoral law, as a quasi-experiment. The reform constitutes an exogenous source of variation in the share of women in a municipal body. Thus, we instrument the share of women with a dummy variable taking the value of 1 for municipalities affected by the gender quota law and 0 otherwise; this instrument has predictive power for women's presence in municipal political bodies but does not directly affect mafia infiltration.

The panel data analysis of about 1,700 Italian southern municipalities from 1985 to 2013 finds that an increase in the share of woman in a municipal body decreases its probability of dissolution due to mafia infiltration. The average marginal effect (AME) shows that an increase in the female participation in a council of 10 percentage points, decreases the probability of dissolution for mafia infiltration by about 1.8 p.p. These findings are robust to different operationalizations of the binary dependent variable. Moreover, we address the idea that women as leaders may enhance institutions by promoting more effective measures to deter crime infiltration. Accordingly, we test if a female mayor affects the probability of interest. Analogous to previous results, the probability of municipal dissolution decreases by about 5 p.p. if the mayor is female rather than male. Findings are robust to the sample restriction to more homogeneous units and to different thresholds of populations, according to appropriate criteria. As in several recent contributions on related topics, we focus on the local (micro) level to exploit the similarity of institutional settings to strengthen identification issues. Results provide additional reasoning against the underrepresentation of women in politics. Therefore, policies targeted to the empowerment of women in government may have the further effect of deterring mafia infiltration in public organizations. A key related issue concerns the efficacy of policies such as gender quota laws. De Paola et al. (2010) and Baltrunaite et al. (2014) demonstrated, in the same scenario of Italian municipalities, that gender quota increased the female share on city councils and improved the quality of elected politicians. We tried to evaluate the effect of the gender quota law. Indeed, according to the estimates of the increase in the share of women in municipal bodies due to the gender quota law made by De Paola et al (2010) (almost 2 p.p.), our results predict a decrease in the probability of dissolution of about 6 p.p.; this decrease in probability of dissolution reaches almost 26 p.p. if we consider the estimated growth in the number of female mayor (3.1 p.p.). The rest of the paper is organized as follows. Section 2 discusses the related literature. Section 3 describes the Italian institutional framework, variables, and data. Sections 4 and 5 present the empirical strategy and results, respectively. Section 6 discusses the issues related to results and performs some robustness checks. Finally, Section 7 concludes.

2. Literature Review

The underrepresentation of women is especially evident in political institutions. The causes lie in a number of reasons: the higher entry cost for participating in political life, the violation of the traditional role of women, the choice made by political parties not to nominate women because voters may dislike female participation in politics, etc. In addition to ethical and equality reasons as rationales to reduce the gender gap, the debate on whether gender matters in policymaking is still open and evidence is mixed. On one hand, reducing the underrepresentation of women in government, achieving ethical or social justice goals, contributes to legitimate democracy (Stevens, 2007); moreover, recent works have shown that it may also impact performance and governmental quality. Specifically, women are geared toward certain kinds of public spending (Rigon and Tanzi, 2011); they are more likely to implement policies such as childcare, water provision, health, and environmental sustainability (Clots-Figueras, 2011; Funk and Gathmann, 2015; Rehavi, 2007); they increase electoral participation (De Paola et al., 2014) and the quality of elected politicians (Baltrunaite et al., 2014); being more liberal than men, female legislators are more likely to support women's issues (Swers, 1998; Washington, 2008). In contrast, when women run local administrations, those systems tend to be less stable (Gagliarducci and Paserman, 2012); actions enhancing a more equal participation of sexes could, instead, promote less qualified individuals at the expense of performance and efficiency (Holzer and Neumark, 2000). A mayor's gender is uncorrelated with local government size and with the composition of municipal spending (Ferreira and Gyourko, 2014).

In this paper we are interested in a particular effect of the underrepresentation of women in politics that concerns the proliferation of crime. Politicians are one of the targets of organized crime in several countries. In recent years, the study of organized crime has become an important research topic among social scientists in discerning its origins linked to the distribution of natural resources and to the weakness of institutions (Dimico et al., 2012; Konrad and Skaperdas, 2012; Buonanno et al., 2015), its destructive effect on physical and human capital (Pinotti, 2015), and its effect on politicians' quality (Daniele and Geys, 2015). Dal Bó and Di Tella (2003) and Dal Bó et al. (2006, 2007) showed that once elections have taken place and the winner takes office, criminal organizations influence policymaking by "inducing a given policy maker to change his action from that preferred by society to that preferred by the criminal groups" (Dal Bó and Di Tella, 2003). Kugler et al., (2005) showed that criminal organizations with sufficient economic and military power may affect policies by bribing or intimidating politicians in office. The empirical evidence has confirmed that the influence of organized crime on politics (Acemoglu et al., 2013) increases immediately after elections (Daniele and Dipoppa, 2017). Moreover, criminal organizations may distort the political selection process by inducing the election of less able politicians (Daniele, 2017).

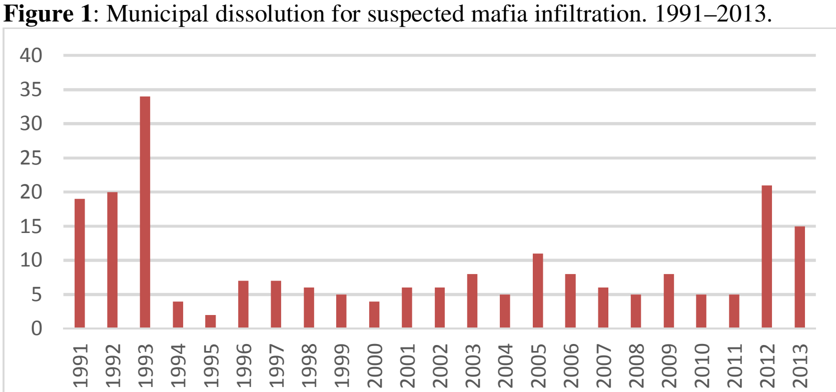
The relationship between the presence of women in political organizations and crime has been analysed referring to corruption (as a form of criminality) in two leading papers by Dollar et al. (2001) and Swamy et al. (2001). In cross country analysis, they find a negative correlation between women's presence in parliament and corruption. Very recently, a paper by Jha and Sarangi (2018) first addressed the question of causality, providing robust evidence that women's presence in parliaments have a causal negative impact on corruption. We contribute to this field of literature, averring that more women in politics, promoting loyal and honest behaviours in governments, could constrain criminal organizations from infiltrating. This argument provides further motivation in favour of the adoption of policy measures aimed at increasing the presence of women in politics.

3. Institutional Framework, Variables and Data

To examine the impact of women's participation in politics on the scale of mafia infiltration in government, we focus on Italian municipalities in Calabria, Campania, Puglia, and Sicilia, the southern Italian regions where organized crime (in the form of, respectively, 'ndrangheta, camorra, sacra corona unita, and mafia) has traditionally been concentrated. Restricting the sample to the south of Italy allows us to work with sufficiently similar municipalities, in terms of unobserved characteristics (e.g. political culture and social capital) that can affect estimations. We construct a data set including yearly observations for about 1,700 Italian municipalities over the period 1985–2013.

To capture organized crime infiltration in Italian municipalities, we consider local government dissolutions due to mafia infiltration (see, among others, Acconcia et al., 2014; Daniele and Geys, 2015). In 1991, the Italian national government, in a period of intense mafia-related killings, imposed an emergency measure. Law no. 164/1991 states that the national government can decree the dissolution of municipal administration "*when evidence emerges regarding direct or indirect links between members of the local government and criminal organisations [. . .] jeopardising the free will of the electoral body and the sound functioning of the municipal administration*".[2] In particular, the dissolution of an administration is proposed by a parliamentary commission in the Ministry of Interior ("Commissione Parlamentare Antimafia") to the President of the Republic, who issues the official decree of municipal dissolution after approval of the government cabinet. Upon the removal of a city council, three external commissioners are assigned to govern the administration during the next 12–18 months, possibly extended to 24. Then, a new local government is elected.[3] Figure 1 shows the number of municipal dissolutions for (suspected) mafia infiltration per year in the timeframe 1991–2013. Figure 2 illustrates the number of municipal dissolutions from 1991 to 2013, divided by the four regions.[4]

Figure 1: Municipal dissolution for suspected mafia infiltration. 1991–2013.



Notes: on the vertical axis there is the number municipal dissolution. Our elaboration.

Between 1991 and 2013, 216 local governments were dismissed because of (suspected) mafia infiltration.[5] Their greatest number is in Campania; Sicilia and Calabria show almost the same number of dissolutions and Puglia counts just 7 dissolutions. As suggested by Pinotti (2015), the relatively

[2] http://www.gazzettaufficiale.it/eli/id/1991/07/25/091A3362/sg

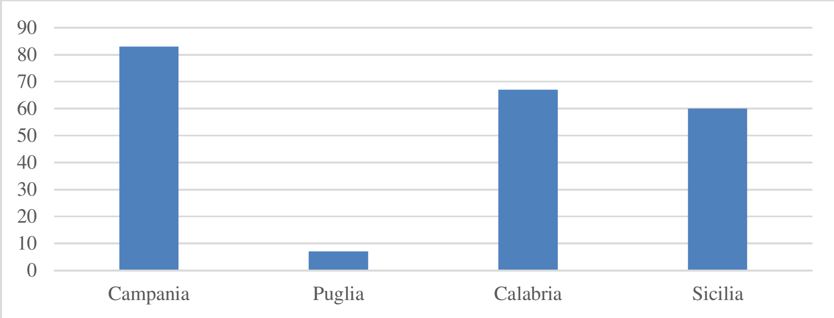
[3] Data referring to the municipal dissolution for mafia infiltration according to law no. 164/1991 are publicly available on the Italian Parliament's "Commissione Parlamentare Antimafia" website. http://www.camera.it/_bicamerali/leg15/commbicantimafia/documentazionetematica/23/schedabase.asp.

[4] Note that, during the considered time span, only seven municipalities were dissolved in the rest of the Italy outside the examined regions.

[5] Among these, 40 municipalities were dissolved more than once.

low number of municipal dissolutions in Puglia is justified by the notion that, although in Sicilia, Calabria and Campania criminal organizations date back more than 150 years, in Puglia criminal activities began around 1970.

Figure 2: Municipal dissolution for suspected mafia infiltration by region. 1991–2013



Notes: on the vertical axis there is the number municipal dissolution. Our elaboration.

The variable representing municipal dissolution (the dependent variable in the empirical analysis) is a dummy taking value of 1 if the municipality has dissolved and 0 otherwise. To assess the robustness of our results, we experiment with a number of different implementations of the dependent variable. The regressor of interest is the share of women in the municipal political body. The political structure of a municipality comprises three bodies: the municipal council, the mayor, and the municipal executive. The municipal council issues municipal laws and is elected (along with the mayor[6]) by citizens every five years[7] and its size is statutory according to municipalities' population size. The mayor is a member of the municipal council responsible for governance of the administration and to maintain public order, civil defence, electoral and registry offices, and other duties delegated by a higher order political body; moreover, the mayor issues decrees and ordinances. The municipal executive council is selected by the mayor and cooperates in municipal management; the municipal executive has a residual role in carrying out all the tasks not directly attributed by law to the municipal council or mayor. The municipal executive is smaller than the municipal council and its size is decided by the mayor with a statutory maximum number.

To compute the share of women in a municipality, we divide the number of women in these three bodies by the sum of all members, considering relevant for our analysis the full political body of municipalities. Indeed, law no. 164/1991 states that a municipality can be dissolved because of emerging evidence regarding direct or indirect links between local administration and criminal organizations. According to Decree 267/2000 (Art. 77), local administrators are the mayor, municipal council members ("consiglieri"), and aldermans ("assessori" – municipal executive members).

[6] Before 1993 the Mayor was elected by the members of municipal council.

[7] The term was shortened to four years between 1993 and 2000 (L 25/3/1993, no.81, art.2; DLgs 18/8/2000, no.267, art.51).

The Italian Ministry of Interior provides data on local administrators.[8] In particular, information on every local election from 1985 to 2017 is publicly available on the website. For every municipality, the website provides information about year of elections, size of the municipal council and executive, name, age, gender, role in the administration, political party, job, and education level of each member of the administration.

In the time span of this analysis, a relevant institutional change in municipal elections occurred: the introduction of the gender quota with law no. 81/1993. The reform prescribed, first, *inter alia*, the election of the mayor by universal suffrage; then, it established (art.5, subsequently modified by L 15/10/1993, no. 415, art. 2) that, for municipalities with more than 15,000 inhabitants, no more than 2/3 of the candidates on an electoral list could be of the same sex, whereas for municipalities with less than 15,000 inhabitants, the threshold was fixed at 3/4. The main reason for the introduction of the gender quota law was to overcome unequal political participation. Hence, the provision of the law was aimed at promoting gender equality in politics and, to let the voter choose freely, the quotas related only to participation in the electoral lists and did not ensure women would be elected (Parliamentary Debate – "Discussione Parlamentare" – n° 33724).[9] In September 1995, the Constitutional Court (Sentence no. 422) declared the gender quota unconstitutional because prejudicial to art. 3 and art. 51 of the Italian Constitution that enshrine the fundamental principle of equal access to elective offices; hence, no preferential treatment could be performed on the basis of sex.[10] Therefore, the law was only enforced during municipal elections between March 25, 1993 and September 12, 1995.[11] In the four regions of interest, 85% of municipalities voted during the enforcement of the gender quotas law.

Figure 3 shows the mean, across municipalities, of the share of women in the political body for municipalities affected/unaffected by the gender quota law, from 1985 to 2013. We identify the first as the treatment group and the others as the control group. Until 1992, the two lines overlap; then, in 1993, the share of woman in the treatment group sharply increases. In line with the findings of De Paola et al., (2010), even if in force only for two years, the effects of the gender-quota legislation persisted after the abrogation of the law and municipalities unaffected by the gender quota law experienced a continuously growth in the share of women.

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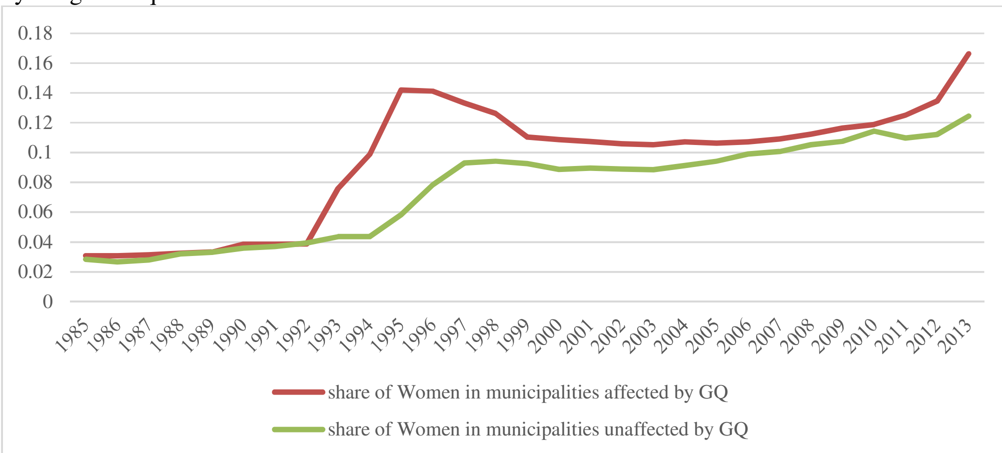
[8] https://dait.interno.gov.it/elezioni/open-data?f%5B0%5D=node%253Afield_argomento%3A180.
[9] https://storia.camera.it/lavori/sedute/4-dicembre-1992-s11-33724.
[10] All other provisions of law no. 81/1993 were unaffected by the Judgement of Constitutional Court.
[11] A potential mixing between the two groups of municipalities may emerge. Some municipalities preparing electoral lists before the period when the gender quota was in place, voted with the gender quota; some municipalities that had prepared electoral lists during the period when the gender quota was in force voted after its abolition. However, considering that electoral campaigns officially last for 30 days, we have no such mixing in our sample because no municipalities voted during the 30 days after March 25, 1993 and in the 30 days after September 12, 1995.

Figure 3: Mean, across municipalities, of the share of women in the political body for municipalities affected/unaffected by the gender quota law. 1985–2013



Notes: on the vertical axis there is the mean (over municipalities) of the share of women in political body from 1985 to 2013. Red line: municipalities affected by the gender quota law; green line: municipalities unaffected by the gender quota law. Our elaboration.

Table 1.1 shows the results of mean difference tests in changes in the share of women between the treatment group and the control group of municipalities. *Difference* is a measure of variation in the share of female presence on city councils, driven by the gender quota law, that is the mean difference test by comparing the share of women's variations across the two groups.

Table 1.1: Share of women. Mean difference test

	All municipalities (1)	*Dissolved municipalities* (2)	*Undissolved municipalities* (3)
Difference	0.022^{***}	0.006^{***}	0.024^{***}
	(0.000)	(0.002)	(0.002)

Notes: The table reports one-sided mean difference test results for the share of women in municipality changes between the treatment and control groups; column 1 shows the test over all municipalities; columns (2) and (3) restrict the sample to dissolved and undissolved municipalities for mafia infiltration, respectively. The treatment group consists of municipalities affected by the gender quota reform; the control group consists of the rest of the sample. Data are annual from 1993 to 2013 at the municipal level. Standard errors are reported in brackets. * shows significance at the 1% significance level.

The mean difference in the share of women in local councils between treatment and control group is 2.2 p.p.. Interestingly, this difference is stronger in undissolved municipalities than in dissolved ones. Regardless of the gender quota reform, table 1.2, column 1, shows the mean difference test relative to changes in the share of women between undissolved and dissolved municipalities. The *t*-test confirms that the mean of the share of women in municipalities that have never been dissolved is statistically greater than that in dissolved municipalities (at the 1% significance level).

Table 1.2: Share of women. Mean difference test

	All Municipalities (1)	*Dissolved municipalities* (2)
Difference	0.017^{***}	0.022^{***}
	(0.000)	(0.002)

Notes: column 1 reports the one-sided mean difference test results for the share of women in municipality changes between undissolved and dissolved municipalities. Column 2 shows the mean difference test results for the share of women in municipalities in the years outside the dissolution and in the year of dissolution. Data are annual from 1991 (when the municipalities' dissolution for mafia infiltration starts) to 2013 at the municipal level. Standard errors are reported in brackets. *** shows significance at the 1%.

Column 2, instead, considers only dissolved municipalities and shows the mean difference test of the change in the share of women in the years outside the dissolution and in the year of dissolution. The test reveals that the share of women at a time outside the dissolution is 2.2 p.p. higher than the mean of the share of women in the year of dissolution (at the 1% significance level). Therefore, all these tests support intuition about a relationship between women's presence in political administrations and mafia infiltration. We also tested that the mean of the share of women in municipalities dissolved more than once is not significantly different from the share of women in municipalities dissolved only once.

For empirical analysis we use a set of variables controlling for municipality characteristics (Daniele and Geys, 2015). They are:[12]

1) population size (in natural log) to control for the size of municipalities (*ln_pop*);
2) municipal unemployment rate in 2001, in natural log (*ln_unemployment*);
3) ratio of young to old inhabitants in 2001, calculated as the number of >65 for every 100 inhabitants under age 15, in natural log (*ln_young/old*).

Table 2 below summarizes the descriptive statistics of all the variables.

Table 2: Variables statistics

Variable	Mean	Std. Dev.	Min	Max	Observations
Women	0.102	0.091	0	0.58	N = 41620 (n=1695; T=25)
Ln_pop	8.351	1.176	5.11	13.9	N = 42219 (n=1692; T=25)
Ln_unemployment	3.058	0.374	0.39	3.93	N = 37582 (n=1642; T=22)
Ln_young/old	4.748	0.447	3.09	6.35	N = 37582 (n=1642; T=22)

Notes: the statistics are calculated starting from 1989 because of the definition of the the binary dependent variable used for the baseline analisis (see below).

4. Empirical strategy

We analyse the impact of the female share in political bodies in southern Italian municipalities on the probability of dissolution for mafia infiltration. As shown in figure 1, dissolution did not take place at the same time for each municipality; the panel structure of our dataset allows us to take this into account and to consider the possibly unobserved time-specific event.

The baseline specification follows (with subscript *i* referring to municipalities and *t* to time)

$$\Pr\left(y_{i,t} = 1 \middle| W_{i,t}, X_{i,t}\right) = \beta_0 + \beta_1 W_{i,t} + \beta_2 X_{i,t} + \delta_i + \delta_t + T\delta_r + \varepsilon_{i,t} \quad (1)$$

where $y_{i,t}$ is a binary variable taking value 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution, due to (presumed) mafia infiltration, and 0 otherwise (that is, it takes value 1 from 1 to maximum 5 times according to the fact that the dissolution occurred respectively at the beginning or at the end of the legislative period of the dissolved administration, and 0 otherwise). This is because the law no. 164/1991 prescribes dissolving the municipalities if the local administrator *in force* had direct and indirect links with criminal organizations. $W_{i,t}$ is the share of women in the municipal political body (thereafter *Women*) of municipality *i* at time *t*, $X_{i,t}$ is the set of controls for characteristics of municipality *i* at time *t* (listed above), δ_i form a vector of provincial and municipal dummies, δ_t are year fixed effects; $T\delta_r$ is a region-

[12] The source of those variables is ISTAT.

specific time trend (to control for any potential differential temporal developments across regions) and $\varepsilon_{i,t}$ is the idiosyncratic error term. We hypothesize that the effect of female presence on city councils is time-invariant.

It is well known that a precise measure of crime is impossible to obtain. Municipal dissolution provides an imperfect measure of mafia infiltration in municipalities because it refers to the detection of the phenomenon. In fact, 1) not all local administrations infiltrated by organized crime are detected and dissolved; 2) a local administration can be erroneously dissolved for suspected mafia infiltrations and the decree of dissolution is not cancelled. Those aspects, luckily, do not affect the identification assumption. Indeed, in the first case, these municipalities will be erroneously part of the undissolved group determining a bias of the estimated effect of women's presence in municipalities' political bodies towards zero; in the second case, municipalities erroneously belonging to compulsory administration group (the dissolved ones) also contribute to bias estimates towards zero.[13] Hence, if our findings should confirm that women affect mafia infiltration, they will only represent a lower bound of the investigated effect; in fact, excluding any possible mismatch listed above, results using a "real" measure of organized crime infiltration should be stronger (see Daniele and Geys, 2015).

We run a first set of municipality-level regressions using linear probability and probit models. Estimating Eq. (1) through LPM and probit can produce biased results because of possible endogeneity. Endogeneity issues can appear if some unobservable characteristics at the municipal level correlate with both the gender of the appointed members of the political body and the probability of dissolution. In addition, results may be biased by the reverse causality issue, because higher/lower probability of municipal dissolution could affect voters' predilection regarding the gender of politicians.

Potential endogeneity is treated by exploiting the introduction, in 1993, of the gender quota reform in municipal elections, as a quasi-experiment. The introduction of the gender quota law, as described in section 3, creates an exogenous source of variation in the gender composition of the municipal political body (between municipalities and over time) that potentially correlates with the gender of the elected politicians, but does not correlate with our outcomes of interest.[14] Thus, we measure the share of women in the municipal political body with the implementation of the gender quota law (Braga and Scervini, 2017). Because the effectiveness of the instrument was well documented in a previous study (De Paola et al., 2010), we have at least two reasons to believe in the validity of the instrument. First, neither in the text of the law nor in parliamentary debates, the introduction of gender

[13] Moreover, no problem arises for dissolved municipalities whose decree of municipal dissolution is cancelled because insufficient proofs link criminal groups and municipal administrations: these municipalities have been removed from the compulsory administration group and included in the undissolved group.

[14] The same exogenous event has already been exploited by other scholars (De Paola et al., 2010; Rigon and Tanzi, 2011; Gagliarducci and Paserman, 2012; De Paola et al., 2014; Baltrunaite et al., 2014; Braga and Scervini, 2017).

quotas was considered a possible tool to affect organized crime infiltration in municipalities. Secondly, local elections take place every 5 years[15] and municipalities cannot change their scheduled time. In specific circumstances, the legislature ends before the natural terms and anticipated elections take place; from then on, the elections (always after 5 years) will be mismatched compared to the regular year of voting in other municipalities. Therefore, it is hardly possible that a municipality adapted the election schedule to vote according (or not) to the gender quotas law. For these reasons, we can safely maintain that the instrument is uncorrelated with the dependent variable. Moreover, findings by De Paola et al. (2010) and the path depicted in figure 3 confirm that the gender quota law affects the share of women over the entire period of analysis.

We address the endogeneity of the share of women in municipal councils with the control function approach, consistent in non-linear models (Rivers and Vuong 1988; Wooldridge 2002).[16] The approach consists of a two-stage procedure: in the first stage, potential endogenous variables are regressed on all the assumed exogenous explanatory variables and the instrument; in the second stage, the predicted residuals are used as an additional regressor in the structural equation with the potential endogenous variables.

In the empirical model, the instrumental variable is a dummy taking the value of 1 for all municipalities during the years affected by the gender quota law and 0 otherwise (thereafter *GQ*).[17] We estimate the following two-stage model:

$$W_{i,t} = \alpha_0 + \alpha_1 GQ_{i,t} + \alpha_2 X_{i,t} + \delta_i + \delta_t + T\delta_r + u_{i,t} \quad (2)$$

$$\Pr(y_{i,t} = 1) = \beta_0 + \beta_1 W_{i,t} + \beta_2 X_{i,t} + \beta_3 Res + \delta_i + \delta_t + T\delta_r + \varepsilon_{i,t} \quad (3)$$

In the first stage, the endogenous variable, $W_{i,t}$ is regressed on the instrumental variable, $GQ_{i,t}$ and the set of regressors $X_{i,t}$; in the second stage, the dummy variable $y_{i,t}$ is regressed on the $W_{i,t}$, the same set of regressors $X_{i,t}$, and the predicted residuals (*Res*).

5. Results

First, we show estimation results of LPM and probit models. LPM is easier to use and interpret because the coefficient of the regressor of interest directly expresses variation in the probability that $y = 1$ for a given variation in the regressor. However, the LPM is not very useful for predictive purposes because the probability can take a value below 0 or above 1. Therefore, the probit model correct this non-linearity of the binary variables. Given the definition of the binary dependent variable,

[15] In the time span 1993–1999, the electoral period was shortened to four years.

[16] This approach allows us to consider endogeneity in panel data probit estimations.

[17] For municipalities that never voted for their council under gender quota law, the *GQ* takes the value of 0 for the entire period; for municipalities in which the council were elected under gender quotas, the *GQ* takes a value of 1 in that election and 0 for previous and following elections.

the estimation period starts in 1989 (because before 1989 the value of the binary dependent variable is 0 for all municipalities) and ends in 2013. Results appear in table 3.

The first four columns of table 3 show the results of LPM, whereas the last four show probit estimations. For both estimation techniques, columns (1) and (5) report estimates of Eq. (1) without considering control variables, region specific time trend, and year fixed effect (FE); columns (2) and (6) include year FE; columns (3) and (7) add region specific time trend and, finally, columns (4) and (8) use the full specification including control variables to minimize the possibility of omitted variable bias.

Table 3. Panel estimation.

	LPM - FE				***Probit***			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Women	-0.07***	-0.03**	-0.03**	-0.03*	-3.59***	-2.20***	-2.66***	-2.66***
	(-4.78)	(-2.27)	(-2.16)	(-1.88)	(-4.30)	(-2.65)	(-3.17)	(-2.84)
Constant	0.02***	0.02***	0.02***	1.19	-3.25***	-3.40***	-3.46***	0.41
	(17.52)	(7.27)	(8.51)	(1.26)	(-10.09)	(-9.90)	(-9.62)	(-0.69)
Year FE	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Region time trend	No	No	Yes	Yes	No	No	Yes	Yes
Municipality controls	No	No	No	Yes	No	No	No	Yes
AME					**-0.087***	**-0.051***	**-0.059***	**-0.058***
N. obs.	41420	41420	41270	35515	37709	37709	37709	33384
Municipalities	1695	1695	1695	1642	1539	1539	1539	1488

Notes. The dependent variable is the dummy *y* taking the value of 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution due to mafia infiltration, and 0 otherwise. *t*-test values and standardised normal *z*-test values are in parentheses, respectively, for LMP and probit model; robust standard errors cluster at the municipal level. Probit estimations contain provincial FE; in the LPM they have been dropped because of multicollinearity. AME: average marginal effect. Significant coefficients are indicated by * (10% level), ** (5% level) and *** (1% level).

For LPM, we reject the null of the Hausman test; therefore we report only FE estimations. The available probit estimation for a binary dependent variable fits only random-effects models; we use the Chamberlain and Mundlak correction to fit the (pseudo) fixed-effects model.[18] Every estimated equation has robust standard errors clustered at the municipal level.

A first look at table 3 shows that the coefficient on *Women* is negative and highly significant everywhere, meaning that an increase in the share of women in the municipal political body decreases the probability of municipal dissolution for mafia infiltration. LPM estimates show a decrease in the coefficient of *Women* when we gradually strengthen the model from (1) to (4). More precisely, an increase in the share of women in a municipal body by 10 p.p. decreases the probability of dissolution by 0.7 p.p. in the most parsimonious specification and by 0.3 p.p. in the most complete one. Given that an improved specification should allow removing biases, this pattern is consistent with the

[18] We used the Stata command `xtprobit` for RE probit estimation. A random effects approach is valid if the regressors do not correlate with individual effects. If such a correlation exists, the estimated coefficients are biased. Therefore, we used Chamberlain's random effects probit model (Chamberlain, 1982; Chamberlain, 1984; Mundlak, 1978). The Chamberlain and Mundlak correction prescribes the inclusion in the estimated equation of the mean terms, over the full time period, of regressors that should capture the correlation between unobserved heterogeneity and the covariates that render the random effect model inconsistent. We also performed RE probit estimations that do not change the main results. They are available upon request.

presence of an upward bias in the most parsimonious specification. Consistently with the findings of the linear estimation, in the probit model we find a significant and negative association between the share of women on municipal councils and the probability of dissolution. To provide an easier interpretation of the coefficient of interest in the probit model, in table 3 we report the estimation of the AME. Because probit is a non-linear model, the investigated effect will differ from individual to individual; the AME computes the effect for each individual and then computes the average. In column (5), the average impact of a 10 p.p. increase of women on a council is a decrease in the probability of dissolution of 0.87 p.p. The relationship of interest is robust to the improvement of the model specification by including year FE, regional trend, and controls; the full specification in (8) reduces the AME: a rise of 10 p.p. in the share of women decreases the probability of dissolution by 0.58 p.p.

5.1 Endogeneity

As said above, we addressed the endogeneity problem using the control function approach. Table 4 shows the results of the estimations shown in Eq. (2) and (3) according to the different definitions of the dependent variable we use. In columns (1)–(3) the binary dependent variable is defined as in table 3: it takes the value of 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution due to (presumed) mafia infiltration, and 0 otherwise. Here, as in table 3, the presentation of estimations starts with the most parsimonious model; then, the specification is gradually strengthened.

The definition of the dependent binary variable given above creates an asymmetry across municipalities. Moreover, it restricts mafia infiltration only to the dissolved administration. However, despite of the time of detection by the "Commissione Parlamentare Antimafia", organized crime could have been infiltrated some time before. To take these features into account, we propose three further definitions of the dependent variable (see Daniele and Geys, 2015): 1) $y_{i,t}$ takes the value of 1 in the year of commissioners and in the 4 years preceding the dissolution of the government, and 0 otherwise; 2) $y_{i,t}$ takes the value of 1 in the year of commissioners and in the 9 years preceding the dissolution of the government, and 0 otherwise; 3) $y_{i,t}$ takes the value of 1 in the year of commissioners and in the entire period back to 1985, and 0 otherwise. The first definition completely removes the asymmetry across municipalities; the second one removes the asymmetry for municipalities put under commissioners starting from 1994; the third one maintains a sort of asymmetry and considers that the mafia infiltrated local government during the entire period, dating back to 1985. All three definitions still hypothesize a time-invariant effect for female participation in municipal administrations on mafia infiltration. According to these new operationalizations of the dependent variable, in columns (4)-(6) of table 4 we show the results of the full specification.

In the upper part of table 4 we show the results of the first (panel FE) stage estimation where the variable *Women* is regressed on the gender quota dummy. In all regressions, standard errors are clustered at the municipal level and robust to heteroscedasticity.

In line with previous research (De Paola et al., 2010), the coefficient of *GQ* is positive and highly significant everywhere, meaning that municipalities whose council was elected when the gender quota law was in force have a higher share of women in charge. Table 4 shows a set of tests for the validity of instruments for all specifications. The *F*-test of weak identification assures that the power of the instrument is extremely high. The Kleibergen–Paap test of under-identification rejects the null hypothesis that the equation is under-identified. Finally, the Anderson–Rubin test for the significance of endogenous regressors in the structural equation rejects the null hypothesis that the coefficients of the endogenous regressors in the structural equation are jointly equal to zero and that the over-identifying restrictions are also valid. Therefore, the tests presented do not show any sign of a weak instrument problem.

The second stage estimation is a panel data probit with the Chamberlain and Mundlak correction. In all regressions, standard errors are calculated using the delta method.[19] The first three columns of table 4 show strong evidence that the probability of dissolution of a municipality decreases when it has a larger share of women in its administration. When we progressively include controls, the magnitude of the estimated AME of *Women* decreases: an increase of 10 p.p. in the share of women reduces the probability of dissolution by 1.84 to 1.78 p.p. Results in (3) tells us that one standard deviation (0.091) increase in the share of women on the council lowers the probability of dissolution by 1.61 p.p. Finally, the variation in the mean of the share of women between dissolved and undissolved municipalities (respectively 0.091 and 0.109, as in table 1) decreases the probability of dissolution by 4.25p.p.[20]

Applying the Rivers–Vuong test, we check the endogeneity of the variable of interest: the coefficient of the residuals (*Res*) are significantly different from zero (except in (1)), therefore IV estimation is the most appropriate one.

Columns (4)–(6) differ in the definition of the dependent variable. Results show that when we hypothesize an organized crime infiltration in a municipality for a very long time (columns (5) and (6)), the average marginal effect for women's presence in municipalities on mafia infiltration decreases compared to the baseline case: it could be harder for women to break long-lasting ties between the administration and mafia.

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[19] We also perform estimations of specification (1), (2) and (3) where standard errors are calculated by the bootstrap method. The AME does not change. Results are not shown.

[20] For Women = 0.091, $\Pr(y = 1|W = 0.091) = \Phi(-8.22*0.091) = \Phi(-0.75) = 0.2266$. For Women = 0.109, $\Pr(y = 1|W = 0.109) = \Phi(-8.22*0.109) = \Phi(-0.90) = 0.1841$. Therefore, the variation in the probability is equal to 0.1841 - 0.2266 = -0.0425.

Municipality controls are the log of municipal population size, the log of municipal unemployment rate, and the log of the young-to-old inhabitants' ratio. Population size does not play a role in the probability of dissolution. The unemployment rate has a positive and significant sign meaning that an increase in unemployment increases the probability of mafia infiltration; the young-to-old inhabitants' ratio has a significant and reverse effect. Moreover, given that Sicilia is one of the five Italian regions with special statutes,[21] we control for this aspect with a dummy taking the value of 1 for municipalities belonging to Sicilia, and 0 otherwise. We run the full specification as in column 3 and the coefficient of interest does not change in sign, significance, and magnitude (the AME is equal to -0.178); the dummy for Sicilia is not significantly different from zero.[22]

Table 4. IV estimations.

Dep.Var. Women	***Panel FE - 1° stage regression***					
	(1)	(2)	(3)	(4)	(5)	(6)
GQ	0.04***	0.07***	0.07***	0.07***	0.07***	0.07***
	(19.5)	(30.4)	(31.1)	(32.05)	(32.05)	(32.1)
Provincial FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	No	Yes	Yes	Yes	Yes	Yes
Municipality controls	No	No	Yes	Yes	Yes	Yes
N. obs.	38641	38641	34524	38845	38845	38845
F-test	379.9***	926.1***	969.3***	1027.3***	1027.3***	1027.3***
Chi^2 Kleibergen-Paap	300.5***	926.7***	970.1***	1028.1***	1028.1***	1028.1***
Chi^2 Anderson-Rubin Wald test	5.65**	18.22***	15.51***	9.22***	12.83***	4.82**
Dep.Var. y	***Probit - 2° stage estimation***					
Women	-8.59***	-8.23***	-8.22***	-7.91***	-7.65***	-4.30***
	(-2.34)	(-3.49)	(-3.41)	(-3.80)	(-4.36)	(-3.17)
Constant	-3.07***	-2.71***	3.34	3.57	4.45*	0.68
	(-5.21)	(-7.06)	(1.50)	(1.61)	(1.02)	(0.61)
Provincial FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	No	Yes	Yes	Yes	Yes	Yes
Municipality controls	No	No	Yes	Yes	Yes	Yes
AME	**-0.184****	**-0.181*****	**-0.178*****	**-0.185*****	**-0.166*****	**-0.059*****
N. obs.	35081	35081	31292	31292	31292	30825
Res	5.19	6.25***	6.06***	5.80***	6.22***	3.03**

Notes. The dependent variable of the 1° stage estimation: *Women*. Columns (1)–(3): the dependent variable of the 2° stage is the dummy taking the value of 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution, due to mafia infiltration, and 0 otherwise. Column (4): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the 4 years preceding the dissolution of the government, and 0 otherwise. Column (5): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the 9 years preceding the dissolution of the government, and 0 otherwise. Column (6): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the entire period back to 1985, and 0 otherwise. *T*-test values and standardised normal *z*-test values are in parentheses, respectively, for the first and second stage. Standard errors are calculated with the delta method, are clustered at the municipal level, and are robust to heteroscedasticity. AME: average marginal effect. Significant coefficients are indicated by * (10% level), ** (5% level) and *** (1% level).

One can argue that women in positions of power in their organizations might design and implement more stringent laws and policies against organized crime than men and, thus, could be perceived as strengthening the organization. This argument is particularly suitable for municipalities with female

[21] The Italian regions with special statutes enjoy particular forms and conditions of financial, administrative, and legislative autonomy.
[22] Coefficients of control variables and the dummy for Sicilian are not shown in table 4.

mayors. In our sample, 3.36% of mayors are women; this percentage is equal to 3.4%, for undissolved municipalities and 1.7%, for dissolved municipalities. We estimated the effect of a female mayor on the probability of dissolution for mafia infiltration. In this case, our regressor of interest is a dummy variable taking the value of 1 if the mayor of a municipality is a woman, and 0 if it is a man (thereafter *Mayor*). Table 5 shows the results; columns (1) to (6) replicate the definitions of the dependent variable as the corresponding columns in table 4.

These estimations corroborate previous results. A female mayor negatively affects the probability of dissolution. Results using the baseline definition of the dependent variable in the first three columns show an almost constant AME equal to -0.056. Thus, on average, municipalities led by female mayors are 5.6 p.p. less likely to be dissolved than are municipalities with male mayors. When considering other definitions of the dependent variable, in the first two cases the AME is higher and in the last one, as before, it is lower.

Table 5. IV estimations.

Dep.Var. Mayor	***Panel FE - 1° stage regression***					
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
GQ	0.008*	0.01***	0.01***	0.01***	0.01***	0.01***
	(1.62)	(3.43)	(3.35)	(3.35)	(3.35)	(3.35)
Provincial FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	No	Yes	Yes	Yes	Yes	Yes
Municipality controls	No	No	Yes	Yes	Yes	Yes
N. obs.	38585	38585	34464	34464	34464	34464
F-test	2.63*	11.76***	11.24***	11.24***	11.24***	11.24***
Chi^2 Kleibergen-Paap	2.63*	11.69***	11.18***	11.18***	11.18***	11.18***
Chi^2 Anderson-Rubin Wald test	6.27***	19.32***	16.30***	18.66***	19.37***	14.63***
Dep.Var. y	***Probit - 2° stage estimation***					
Mayor	-41.58**	-32.15***	-33.81***	-33.61***	-33.79***	-18.02***
	(-2.28)	(-3.53)	(-3.44)	(-3.85)	(-4.41)	(-3.13)
Constant	-2.03***	-2.67***	3.07*	4.05**	5.04**	1.03
	(-2.77)	(-7.71)	(1.93)	(2.00)	(2.26)	(1.11)
Provincial FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	No	Yes	Yes	Yes	Yes	Yes
Municipality controls	No	No	Yes	Yes	Yes	Yes
AME	**-0.056****	**-0.056*****	**-0.055*****	**-0.059*****	**-0.069*****	**-0.041*****
N. obs.	35037	35037	31244	31244	31244	30780
Res	41.28**	31.94***	33.67***	33.52***	33.25***	17.94***

Notes. The dependent variable of the 1° stage estimation: *Mayor*. Columns (1)–(3): the dependent variable of the 2° stage is the dummy taking the value of 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution due to mafia infiltration, and 0 otherwise. Column (4): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the 4 years preceding the dissolution of the government, and 0 otherwise. Column (5): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the 9 years preceding the dissolution of the government, and 0 otherwise. Column (6): the dependent variable of the 2° stage is the dummy taking the value of 1 in the year of commissioners and in the entire period back to 1985, and 0 otherwise. *T*-test values and standardised normal *z*-test values are in parentheses, respectively, for the first and second stage. Standard errors are calculated with the delta method; are clustered at the municipal level and robust to heteroscedasticity. AME: average marginal effect. Significant coefficients are indicated by * (10% level), ** (5% level) and *** (1% level).

In the first-stage estimations, the instrument is positive and significantly different from zero and first-stage tests confirm the power of the instrument. The municipalities' controls are the same as before.

5.2 Effect of gender quota reform

In light of our findings, the investigation of policies promoting female participation in politics should be of great interest. In particular, the scenario in the analysis advocates the investigation of the effect of gender quota reform that interested Italian municipalities during the period under consideration. The difficulty of measuring organized crime infiltration is well known; the proxy we used makes unsuitable a direct evaluation of the impact that the introduction of the gender quota at municipal elections in 1993 has had on the penetration of organized crime in city councils. Indeed, the sub-sample of municipalities that were dissolved due to mafia infiltration but did not vote under the gender quota law is not sizable. This small number of observations does not allow implementation of a credible diff-in-diff methodology. However, results of our analysis allow this very interesting investigation by exploiting some findings from the De Paola et al. (2010) study.

As shown in figure 3, starting with gender quota institutional reform (in 1993), an increasing trend in the share of women in political bodies of all municipalities has emerged, partially due to the introduction of the gender quota law (De Paola et al., 2010). First, we can evaluate, according to the estimated coefficients in column (3) table 4, the effect of this increasing trend in the share of women on the probability of interest. Indeed, the growth in the share of women in municipal bodies between 1992 and 2013 (in mean over all municipalities) decreases the probability of dissolution for mafia infiltration by 28.3 p.p.[23] However, this sizable result includes not only the effect due to gender quota reform but also other confounding factors related to temporal trends. To disentangle the impact of the reform, we rely on the De Paola et al. (2010) paper. They show that the fraction of women elected in municipalities directly affected by the gender quota law increased by around 2 p.p. more than in municipalities not affected by the reform;[24] thereby, increasing the share of women in 1992 by 2 p.p., we calculate that the introduction of the gender quota reduced the probability of dissolution of 5.9 p.p.[25]

Secondly, we can do the same exercise according to the findings of column (3) table 5. We evaluate the effect of the increase in the elected female mayors due to the gender quota reform. De Paola et al. (2010) estimated that the gender quota reform grew the probability of electing a female mayor by 3.1p.p. As before, starting from the share of female mayors in our sample at 1992 (0.0126) and increasing it by 3.1 p.p., we calculate that the effect on the probability of dissolution is of -26.64 p.p.

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[23] The mean of the share of women in 1992 is 0.038; that in 2013 is 0.16. For Women = 0.038, $Pr(y = 1|W = 0.038) = \Phi(-8.22*0.038) = \Phi(-0.31) = 0.3783$. For Women = 0.16, $Pr(y = 1|W = 0.16) = \Phi(-8.22*0.16) = \Phi(-1.31) = 0.0951$. Therefore, the variation in the probability is equal to 0.0951 - 0.3783 = -0.283.

[24] De Paola et al. (2010) considered all Italian municipalities in their analysis but they showed no difference in the effect of the gender quota reform in the north and south of Italy.

[25] We increase the share of women in 1992 by 2 p.p. and make the same calculation as before: $Pr(y = 1|W = 0.038) - Pr(y = 1|W = 0.058)$.

Therefore, the effect of the gender quota reform, through its push toward a greater presence of women in municipal bodies, has strengthened local government against organized crime infiltration. The more sizable result relies on the increase in the number of female mayors with respect to the increase in the share of women on municipal councils; this result is probably due to the notion that a female mayor might exercise a stronger deterrence effect against organized crime infiltration by being the leader of the local administration. These findings corroborate the idea that introducing affirmative measures aimed at reducing women's underrepresentation in politics (as gender quotas) may promote further positive results from human/socio/economic perspectives linked to crime.

6 Discussion and Robustness checks

This analysis has provided evidence that female presence in local politics is more likely to link to fewer municipal dismissals for mafia infiltration. We now consider several possible explanations of this result. A natural argument links to the "nature" of women: they discourage organized crime infiltration because of their higher sense of morality, which translates into an improvement in the quality of institutions. Because women are more likely to exhibit "helping" behaviours (Eagly and Crowley, 1986), score more highly on "integrity tests" (Ones and Viswesvaran, 1998), and tend to vote according to social issues (Goertzel, 1983), a higher representation of women leads to the adoption of policies and practices that have a positive impact on the quality of institutions, organisations, and society as a whole; in this sense, women may be seen as an institutional strengthening and strengthened institutions make it difficult for the mafia to infiltrate. Another related argument concerns a positive externality exercised by women: they could positively influence their male colleagues by restraining, disciplining, and elevating the latter's behaviour.

A stronger argument for our results is that women in positions of power in their organizations, as mayors in local administrations, may design and implement more stringent laws and policies against organized crime than men. This explanation is in line with what the Italian system prescribes. That is, the system gives the mayor the greatest part of decision-making and executive power. Therefore, it is not difficult to believe in the effectiveness of measures discouraging criminal activities implemented by female mayors. However, our general result, showing the negative impact of increasing female participation in politics on organized crime infiltration, holds even when we restrict the sample only to municipalities led by a male mayor: this ensures that the influence of women in the political body on crime infiltration is not driven by the stronger effect of a female mayor. Therefore, recalling that councillors have important responsibilities over a wide range of city activities (e.g. urban life, roads, gardens, the environment, pollution, traffic conditions, local production, local services, cultural activities), women, as municipal councillors, could exercise a function of pressure and guidance toward certain "fair" policies, mostly if female councillors represent a local government majority. Moreover, our data show that the mean of the share of women in municipalities led by a female mayor is twice the mean of the share of women in municipalities led by a male mayor (respectively 0.175 and 0.092); taking the results of Baskaran and Hessami (2018), electing a female mayor can have spill-over effects for the likelihood of having women on municipal councils and, thus, can be seen as a further way to increase female participation in politics. One more reason for our results links to the differences in risk attitudes between genders; evidence indicates that women are less likely than men to engage in risky behaviours such as illicit drug use and criminal activities (Cooperstock and Parnell, 1982; Daly and Wilson, 1988; Gottfredson and Hirschi, 1990; Kandel and Logan, 1984; and Wilson and Herrnstein, 1985).

In addition, numerous recent studies concerning the policy implications of gender representation in local government find that women tend to allocate a greater budget share to education and childcare (Svaleryd, 2009), health care (Rehavi, 2007), and environmental sustainability (Funk and Gathmann, 2010). As the literature on corruption suggests (Mauro, 1998; Baraldi, 2008; Delavallade, 2006; Cordis, 2014), these kinds of public expenses cannot guarantee huge economic gains if managed under the influence of organized crime, rather than by public works, urban plans, etc. Therefore, where the share of women in local government increases, a higher pressure exists to devolve public resources toward these sectors (education, health, etc.), where it is less profitable for organised crime to infiltrate.

In the following section we provide further evidence for qualifying and supporting our results.

6.1 Impact on dissolutions unrelated to (suspected) mafia infiltration

An important issue concerns the interpretation of results linked to our measure of organized crime infiltration in city government, that is, dissolution due to mafia infiltration. City councils can also be dismissed for reasons unrelated to mafia infiltration. Reasons for dismissals include resignation by elected officials (resignation of the mayor/resignation of more than 50% of council members), failure to organize elections, special cases of ineligibility of the mayor, failure to pass the annual budget, political crisis in ruling coalitions. Such dissolutions are fairly common[26] and lots of municipalities witness at least one dissolution of their local government unrelated to mafia infiltration.[27] Independent of their effects on mafia infiltration, it might be possible that women on city councils have a negative impact on dissolution *per se*, as well as on dissolution due to mafia infiltration. If this is the case, our results could be led by confounding factors other than the "virtues" of women that, by strengthening institutions, discourage organized crime. Therefore, we checked this aspect by

[26] During the period 1990–2013, city councils dismissed for reasons unrelated to mafia infiltration were 47% of the entire sample. The most common reason was resignation by elected officials.
[27] We do not distinguish between different reasons for government dissolution; thus, we treat all non-mafia dissolutions as one group.

performing regressions where the dependent variable is a dummy taking the value of 1 in the year of local government's dissolution for reasons unrelated to mafia infiltration, and 0 otherwise. Results appear in table 6. In the first three columns the regressor of interest is the share of women on municipal councils whereas in the last three columns the regressor of interest is the dummy variable for the gender of the mayor. We present findings from several estimation strategies that control for regional trends and municipality controls. We show only the second stage; all the performed tests on the validity of the instrument (not shown) confirm no signal of weakness of the instrumental variable.

Table 6. Second stage IV estimations.

	Probit IV - 2° stage estimation					
Dep.Var. y	(1)	(2)	(3)	(4)	(5)	(6)
Women	9.77^{**}	5.61^{**}	5.11^{***}			
	(12.26)	(11.28)	(10.53)			
Mayor				56.19^{***}	22.53^{***}	21.50^{***}
				(12.52)	(11.27)	(10.58)
Constant	-3.89^{***}	-3.17^{***}	-6.15^{***}	-5.21^{***}	-3.44^{***}	-6.75^{***}
	(-11.89)	(-10.09)	(-10.46)	(-14.48)	(-10.98)	(-11.93)
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	No	Yes	Yes	No	Yes	Yes
Municipality controls	No	No	Yes	No	No	Yes
N. obs.	27449	27449	24728	27434	27434	24713
Res	-8.66^{***}	-4.25^{***}	-3.67^{***}	-56.15^{***}	-22.47^{***}	-21.43^{***}

Notes. The dependent variable of the 1° stage estimation in (1)-(3): *Women*. The dependent variable of the 1° stage estimation in (4)–(6): *Mayor*. In the first stage estimations we run the *F*-test, the Kleibergen–Paap test and the Anderson–Rubin test. The dependent variable is a dummy taking the value of 1 in the year of local government's dissolution for reasons unrelated to mafia infiltration, and 0 otherwise. Normal *z*-test values are in parentheses; robust standard errors are clustered at the municipal level. Significant coefficients are indicated by * (10% level), ** (5% level) and *** (1% level).

Different from previous results, the share of women on city councils as well as female mayors have positive and significant effects on local government dissolutions due to factors unrelated to mafia infiltration. This result is in line with that of Gagliarducci and Paserman (2012) and supports the interpretation of our findings.

6.2 Sample restriction

We undertook further robustness checks of our findings by restricting the sample to more homogeneous municipalities to control for other factors that might confound the impact of the share of women on the probability of municipalities' dissolution due to mafia infiltration. First, we restrict the undissolved group of municipalities to those one in the neighbourhoods of municipalities included in the dissolved group. To do that, we take the data by the National Institute of Statistics (ISTAT) on neighbouring municipalities [28] and, for each dissolved municipality, we take its neighbouring municipalities. We construct a new dataset made of 580 municipalities observed in the period 1985–2013. In columns (1) and (4) of table 7 we carry out the complete specifications of our model on this restricted sample where the regressors of interest are, respectively, *Women* and *Mayor*. The coefficient of the female share in municipal bodies is significant and the sign remains negative; the

[28] ISTAT, "Matrici di contiguità", https://www.istat.it/it/archivio/137333.

impact on the probability of dissolution for mafia infiltration is more than 3 times greater than in the full sample (shown by the AME). Instead, for female mayors, the AME increases, but less than double. Therefore, the magnitude of results was driven by comparing groups of municipalities that are more or less heterogeneous.

Table 7. Second stage IV estimations.

	Probit IV - 2° stage estimation					
	Neighbourhood	***Excluding Puglia***	***Population<15000***	***Neighbourhood***	***Excluding Puglia***	***Population<15000***
Dep.Var. y	(1)	(2)	(3)	(4)	(5)	(6)
Women	-9.15**	-7.99***	-5.32**			
	(-3.34)	(-3.29)	(-1.98)			
Mayor				-28.94***	-32.44***	-28.70**
				(-3.37)	(-3.35)	(-2.05)
Constant	3.34	3.84*	3.58	1.01	1.42	1.73
	(1.64)	(1.21)	(1.34)	(0.63)	(1.21)	(0.77)
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Region time trend	Yes	Yes	Yes	Yes	Yes	Yes
Municipality controls	Yes	Yes	Yes	Yes	Yes	Yes
AME	**-0.596***	**-0.187***	**-0.102**	**-0.083***	**-0.054***	**-0.052***
N. obs.	10658	28203	23575	10583	28162	23540
Res	6.94**	5.84**	2.93	28.82***	32.30***	27.94**

Notes. The dependent variable of the 1° stage estimation in (1)–(3): *Women*. The dependent variable of the 1° stage estimation in (4)–(6): *Mayor*. The dependent variable $y_{i,t}$ takes the value of 1 for municipalities put under commissioners every year from the appointment of the elected administration to the year of dissolution due to mafia infiltration, and 0 otherwise. In columns (1)-(3) the regressor of interest is *Women* whereas in columns (4)-(6) it is *Mayor*. Standardized normal *z*-test values are in parentheses; robust standard errors clustered at municipal level. AME: average marginal effect. Significant coefficients are indicated by * (10% level), ** (5% level) and *** (1% level).

In analysing a more homogeneous sample of municipalities, figure 2 suggests to exclude Puglia from the sample because of its very low number of municipalities in compulsory administration and the high number of undissolved municipalities. Columns (2) and (5) of table 7 show that the estimation results are in line with those of the baseline analysis in sign and magnitude of the AME.

A further robustness check focused on the threshold of population. According to law no. 164/1991 and decree no. 267/2000, to compute the share of women (*Women*), we considered all the local administrators: the mayors, municipal council members and members of municipal executive councils. However, they form the elective local political body only in municipalities under 15,000 inhabitants; in municipalities above 15,000 inhabitants, the elective body only comprises the mayor and municipal council members. Moreover, only the municipal council is the elective body affected by the gender quota law of 1993. Indeed, no gender restrictions exist on the mayor and on the municipal executive council that is appointed directly by the mayor. However, for municipalities under 15,000 inhabitants, the municipal executive body is appointed by the mayor among the members of the municipal council; thus, the gender quota law affects the entire political body. Therefore, we restrict the sample to municipalities whose population is below 15,000 inhabitants. In this way the sample perfectly fits the requirements of the gender quota and municipal dissolution laws we exploit. We expect no relevant changes in the results for two reasons: 1) municipalities below 15,000 inhabitants comprise almost 85% of the entire sample; 2) the size of the municipal executive is very small, with low variability in the share of women in that office. Columns (3) and (6) show estimation results. The significant negative impact of the presence of women in the political body and of a female mayor on mafia infiltration is confirmed. Although the AME of *Mayor* replicates the main analysis, the estimated AME of the share of women is lower. This may be due to the different gender thresholds stated by law no. 81/1993. Indeed, the law prescribes a different threshold for gender quotas according to the size of municipalities. For municipalities above 15,000 inhabitants the threshold for a gender presence in an electoral list is 2/3, whereas in municipalities below 15,000, this threshold is 3/4. Therefore, it is reasonable to believe that this different prescription led to different electoral outcomes in female presence in local political bodies; precisely, the threshold of 2/3 presumably led to a greater share of woman in the political body than the threshold of 3/4. Therefore, higher gender thresholds may imply greater advocacy for female presence in politics and, consequently, greater indirect effects in deterring organized crime infiltration.

6.3 Falsification test for the validity of the instrument

As a falsification test to check the validity of the gender quota law as an instrumental variable, we randomise the share of women appointed to the council in two ways: (i) we randomly assign observations on the share of women across the whole sample; (ii) we randomise the share of women appointed to the council in every municipality across years. Running the first stage of the full specification of the model, in both cases we find the gender quota is no longer significant; also, the *F*-test provides weak identification results (respectively, Prob > F = 0.8133 and Prob > F = 0.9820). This suggests we identified the right source of variability to explain changes in women's representation in local government.

7 Concluding remarks

In contributing to the still open debate about whether gender matters in policymaking, this paper offers a further motivation in favour of policies devoted to reduce the underrepresentation of women in politics at national and local levels, going beyond equality reasons. In recent years, a large number of countries introduced measures aimed at increasing the share of women in governments to legitimate democracy and stimulate voters to believe in female potential in politics. These measures, such as gender quotas, may be seen as institution-strengthening in light of the results of our research. Indeed, this work provides new and unexplored evidence of the effect of politician's gender on mafia infiltration. Given that the literature on organized crime ascribes its proliferation to the weakness of institutions, we showed that women, by improving institutions, can deter organized crime infiltration; thus, policies increasing female participation in political life also can be seen as enhancing public organization.

In the baseline analysis, we perform several estimations on about 1,700 southern Italian municipalities relating the probability of dissolution for mafia infiltration to the share of women in political body. We rely on two Italian laws; one, law no. 164/1991, allows us to measure mafia infiltration in Italian municipalities; second, law no. 81/1993 allows us to correct for endogeneity bias by creating an exogenous source of variation in the share of women in the municipal political body. We find that a 10 p.p. increase in the share of women decreases the probability of municipal dissolution by about 1.8 p.p. Results are confirmed when we consider the gender of the mayor: when the mayor is female, the probability of dissolution decreases of about 5 p.p. compared to a male mayor. This result holds over different estimation techniques and improvement in the specification of the model. It is also robust to the further definitions of the binary dependent variable and sample restrictions.

But, how do women reduce organized crime infiltration by being in politics? Here we propose possible explanations that advocate future research. The principle underlying gender studies is that women have different perspectives and preferences from those of men. Therefore, women in government implement new sets of views in decision-making and, consequently, affect policies, organizations, and institutions in the provision of good and services, pressures in committee, guidance in legislative activities, levels of corruption, and political participation (Dollar et al., 2001; Michelle Heath et al.; 2005; Lawless, 2004; Wängnerud, 2009). Although the number of female politicians is still small, a considerable increase in female representation at local and national elections is taking place worldwide. This evidence may reflect the awareness that participation in politics is more attainable for women and is desirable, also supported by our results. Thus, this analysis offers additional reasons to implement policies aimed at reducing women's underrepresentation in political life in that such policies may have a very large range of socio-economic benefits.

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