Political Dynasties, Term Limits and Female Political Representation: Evidence from the Philippines*

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Abstract

We investigate the effect of term limits on female political representation. Using data from Philippine municipalities where strict term limits have been in place since 1987, we show that term limits led to a large increase in the number of women running and winning in mayoral elections. However, we show that this increase is entirely driven by female relatives of the term-limited incumbents. We further show that the differential gender impact of this policy is driven by political dynasties' adaptive strategies to stay in power.

Keywords: female representation, dynasties, term-limits, elections.

Declarations of interest: none

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1 Introduction

The second half of the 20th century has seen women break gender roles and stereotypes in all spheres of life. Despite such dramatic shifts, they still lag in leadership positions, particularly in politics, in both developed and developing countries. In 2019, only 24.3 percent of national legislators around the world were women, and such low representation was even more severe for elected executive positions, with only 11 female Heads of State.[1]

Scholars and policy-makers are interested in understanding women's pathways to elected offices. Recent studies have brought attention to structural characteristics of the political environment such as electoral systems (Rosen, 2013), political competition (Escobar-Lemmon and Taylor-Robinson, 2005; Lawless and Pearson, 2008; Folke and Rickne, 2016) and the politicization of ethnicity (Arriola and Johnson, 2014) as well as cultural norms such as matrilineality (Robinson and Gottlieb, 2019). In parallel, a growing body of research has focused on how policies such as gender quotas (Krook, 2009; Pande and Ford, 2012; O'Brien and Rickne, 2016) or political reservations (Chattopadhyay and Duflo, 2004; Bhavnani, 2009; Cassan and Vandewalle, 2017) shape female political representation both substantively and descriptively.

We investigate the effect of a widely used policy on female political representation: term limits. Although the intended goal of term limits is not explicitly to improve women's representation, a plausible side effect could be a rise in female elected officials. After all, open-seat races – races where the incumbent is not running – are known to attract outsiders and lesser-known candidates to the political scene (Cain, Hanley and Kousser, 2006). The search for a context to elucidate these issues takes us to the rise of female politicians in mayoral positions in the Philippines.

The Philippines ranks among the top countries in the world in terms of female political representation. Women currently hold 29 percent of the seats in the Senate and the House of Representatives. However, the Philippines didn't always have such high share of female politicians. In the aftermath of the fall of Ferdinand Marcos' autocratic regime in the mid 1980s, only 9 percent of women were elected to the Senate and the House of Representatives. Other elected offices followed similar trends. Of relevance to the present study, in 1988 only 9 percent of the municipalities had a female

[1] https://www.unwomen.org/en/what-we-do/leadership-and-political-participation/facts-and-figures accessed on September 5, 2019.

mayor, climbing to 21 percent in 2010.

We provide evidence that this increase in female political representation is linked to the term limits introduced in the 1987 constitution. Our analysis relies on a difference-in-differences empirical strategy where we compare forced-open-seat races to other races, before and after term limits bind. Hereafter, we use the term forced-open-seat races to refer to open-seat races following binding term limits, to differentiate from open-seat races in which the incumbent was eligible to run but decided not to.

First, we show that forced-open-seat races are significantly more likely to have (i) a woman running for office, and (ii) a woman being elected to the mayoral office. Our estimates suggest that term limits can account for about two thirds of the increase in the share of female mayors in the Philippines. We next show that dynastic candidates, that is, relatives of previous mayors in the same municipality, are more likely to run in forced-open-seat races and that the increase in female representation following binding term limits is entirely driven by dynastic women. While non-dynastic women are slightly more likely to run in forced-open-seat races, they are not more likely to win. Thus, in our context, term limits increase the share of a very specific group of women in power: *dynastic women*. This is a striking and unexpected result given the low share of municipalities that had a female representative to start with.

What explains this dramatic increase in the number of female dynastic politicians? Recent studies have shown that family ties are an important vehicle for women's access to politics, in both developed and developing countries (Dal Bo, Dal Bo and Snyder, 2009; Jalalzai, 2013; Chandra, 2016; Folke, Rickne and Smith, 2016; Smith and Martin, 2017; van Coppenolle, 2017; Jalalzai and Rincker, 2018). If political dynasties are more likely to field female candidates, then the increase in the share of dynastic candidates in forced-open-seat races may explain the rise of female politicians. However, we show that term limits also affect the gender composition of these dynastic successions. We find that the share of dynastic candidates that are female increases from 15% prior to binding term limits to 45% in forced-open-seat races. A simple simulation exercise suggests that these two effects can help account for the rise of female politicians, though neither by itself can fully explain the observed change.

We provide evidence of two mechanisms that help explain why term-limited incumbents are more likely to nominate female relatives. First, term-limited incumbents who wish to return to office after waiting out one term, may be more likely to select female relatives who, given existing social norms, might be more willing to step aside after one term to restore their male relative's political career. Our data shows that female relatives are three times more likely than male relatives to hold office for one term and then retire to allow the termed-out incumbent to run again. Following Coronel et al. (2004), we refer to these politicians as "benchwarmers". Second, we show that given their age, term-limited incumbents may be constrained in the number of eligible male relatives. In particular, if term-limited incumbents are, on average, younger than those who retire voluntarily, they may not have any sons of eligible age when the term-limit binds and may thus have to rely on their wives to keep the mayoral position in the family.[2]

Our results are relevant for multiple reasons. Most recent research on female descriptive representation focuses on institutions such as quotas that mandate candidacy or incumbency and are often relevant for legislative bodies. Here, we bring attention to term limits, an electoral institution widespread across the world, and most common for executive offices, for which women are particularly underrepresented. Unlike quotas, term limits are not introduced with the deliberate goal of increasing female representation. Yet, as posited in the American setting, with state-level legislative offices, term limits may lead to an increase in female political representation, by providing winnable open-seat opportunities, where the incumbent is not running.[3] We find that in our context forced-open-seat races increase the probability of women accessing elected offices, but mainly for a small group of women: dynastic women.[4] This unintended consequence of term limits is related to a recent study by Cassan and Vandewalle (2017), where female reservations

[2]Our work is related to George, Siddharth (2020) who shows that in India, politicians with a son are twice as likely to establish a dynasty. We do not have data on the family structure of each politician to examine this issue. However, our argument is that family structure may impact the *gender* of the incumbent's successor, and not only the likelihood of starting a dynasty in the first place.

[3]Term limits are known to attract outsiders and lesser-known candidates to the political scene (Cain, Hanley and Kousser, 2006). The relation between term limits and female political representation is specifically explored in Thompson and Moncrief (1993); Reed and Schansberg (1995); Caress (1999); Carroll and Jenkins (2001*a*), and Carroll and Jenkins (2001*b*) .

[4]We thus contribute to existing work that explores the characteristics of women who run for public office. See for example, Escobar-Lemmon and Taylor-Robinson (2009) and Schwindt-Bayer (2011).

in India increased the representation of low-caste groups. In their case, the increase in low-caste representation is related to traditional gender norms prevalent among high caste people in Indian villages. Thus, while our empirical analysis is limited to local elections in the Philippines, the patterns we document are likely relevant for other democracies.

Our findings are also broadly related to the nascent literature emphasizing the importance of culture when exploring the effects of policies, see Ashraf et al. (2016) and Corno, Hildebrandt and Voena (2017). Ashraf et al. (2016) show that female education is more responsive to school construction programs for ethnic groups practicing bride price. Our paper highlights a similar point within a political context. Although we do not exploit cultural variation within the Philippines, our results provide strong evidence that the impact of term limits on political outcomes is tied to the dynastic nature of the Philippine political institutions and the high share of male term-limited incumbents. Absent a strong dynastic environment, the impact of term limits will most likely manifest itself through other channels such as increased competition.

Finally, we build on existing work by Querubin (2012) and show one particular way through which political elites can persist across time and adapt to specific institutional reforms, such as the introduction of term limits. By having relatives run – and win – in forced-open-seat races, political dynasties can retain control over key political offices and reproduce their power across time. Specifically, Querubin (2012) focuses on elections for higher levels of government (governors and members of congress) and documents that term limits are relatively ineffective in promoting the alternation in office of different families, since incumbents adapt to term limits by having relatives run to replace them. In this paper, we argue that these adaptive strategies of incumbents and their dynasties in response to term limits, have important implications for the gender composition of elected officials in the Philippines.

2 Term limits and the Rise of Female Politicians

The Philippines saw a dramatic increase in the share of female mayors in the last 30 years. Using official candidate-level results of all mayoral elections between 1988 and 2010, provided by the Commission of Elections (COMELEC), Figure 1 reports the share of municipalities with an elected female mayor in each of the eight elections between 1988-2010.[5] Starting at nine percent in 1988, the share of municipalities with a female mayor jumped to 16% in 1998 and kept on increasing to 21% by 2010. The discontinuous jump in 1998 is key. It hints at the importance of term limits in the rise of female politicians as this is the first election year with forced-open-seat races following binding term-limits.[6]

Term limits for all elected offices were introduced in the 1987 constitution, following the fall of Ferdinand Marcos. Under term limits, politicians can only be elected to the same office three times consecutively (not counting elections before 1987). Upon reaching a binding term limit, the politician can either immediately run for a different office, or wait out an election cycle and run for the same office. While the first senatorial and congressional elections were organized in 1987, those for provincial (governors, vice-governors and provincial legislators) and *municipal* (mayors, vice-mayors and councilors) officials took place in 1988. Yet, in accordance with the provisions of the 1987 constitution, the next elections for all congressional, provincial and municipal offices were organized in 1992. Thereafter, politicians at all levels of government were elected simultaneously every three years.[7] Thus, the 1998 election was the first election with potentially term-limited incumbents. In what follows, we formally test the hypothesis that term limits led to an increase in the number of women elected as mayors.

[5]The dataset includes the full names and votes received for all candidates. The only exception is 1988; we only have data on winning candidates for that election. We used the first name of the candidates in our sample to code their gender.

[6]Descriptive statistics for our main electoral variables are reported in Appendix Table A.1.

[7]Since elections for offices at all levels of government take place simultaneously, it is not uncommon for three-term incumbents to run for a different office immediately after reaching their term limit. See Querubin (2012) for an illustration of these phenomenon for legislators and provincial governors.

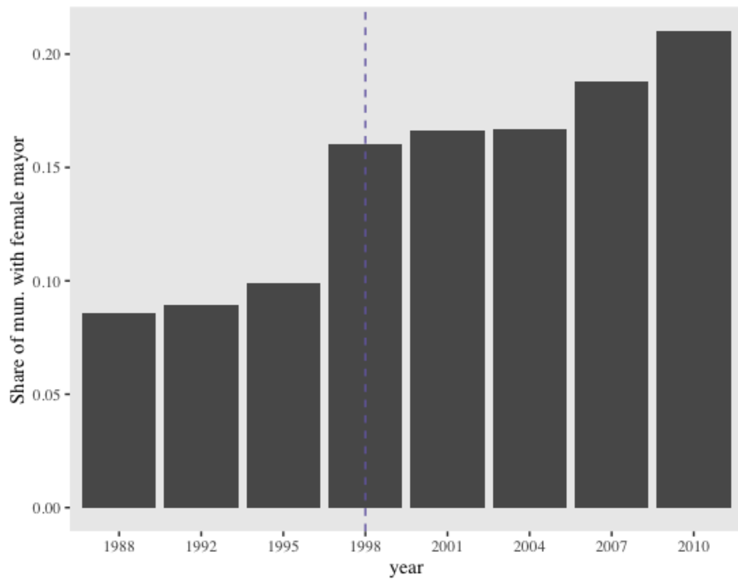

Figure 1: Share of municipalities with a female mayor

2.1 Empirical Strategy and Main Results

To test the impact of term limits on the rise of female politicians, we use a difference-in-differences estimation strategy. We compare municipalities with a term-limited incumbent to municipalities without term-limited incumbents, before and after the forced-open-seat race. Our basic estimating equation is:

$$y_{mt} = \alpha + \beta Term\_Limit_{mt} + \mu_m + \rho_t + \epsilon_{mt} \tag{1}$$

where y_{mt} is one of two indicators: a dummy for whether a woman ran in a mayoral election in municipality m and election year t and a dummy for whether a woman won a mayoral election in municipality m and election year t. $Term\_Limit_{mt}$ is a dummy equal to 1 if at time t municipality m had a forced-open-seat race and zero otherwise. Our main coefficient of interest is β. We cluster standard errors at the municipal level.[8] We include election-year dummies to account flexibly for common shocks or trends at the national level and a full set of municipality fixed effects to account for all time-invariant municipal characteristics. We also estimate a more demanding specification

[8] We also clustered at the province level and the results remain unchanged.

by including interactions between election-year dummies and province fixed effects.

Table 1 presents our main results. The dependent variable is a dummy for whether a woman ran (Columns 1-2) or won the race (Columns 3-4). Municipalities are 17 percentage points more likely to have a woman running for the mayoral office in a forced-open-seat race (Column 1). Women are not only more likely to run, they are also 10.5 percentage points more likely to *win* in forced-open-seat races (Column 3). These results are robust to using the interactions between election-year dummies and province fixed effects instead of country-wide election-year fixed-effects (Columns 2 and 4).[9] Contrasting our point estimates to the jump in the share of female politicians in 1998 (the first year with forced-open-seat races), binding term limits could account for more than two thirds of the 6.09 percentage points increase in the share of female mayors between 1995 (9.9%) and 1998 (16.0%). Given that 42% of the municipalities had a forced-open-seat in that year, our estimates would imply a 4.3 percentage points increase in the share of municipalities with female mayors.

To sum up, the estimates in Table 1 show that women are more likely to *run* and *win* in forced-open-seat races. This results from (both) an increase in the likelihood of women running, and to a minor extent in the likelihood of women winning conditional on running. Prior to 1998 (before term limits bind), conditional on at least one woman running, the likelihood that a woman was elected mayor was 0.45. In forced-open-seat races, that likelihood increases slightly to 0.5.

2.2 Identifying Assumptions

Before proceeding with the mechanisms, we discuss a number of important points related to the validity of, and the source of variation behind the difference-in-difference estimates in Table 1. Specifically, we first provide evidence supporting the parallel-trend assumption. We investigate the sources of variation behind the timing of term limits, and other potential confounders. We finally engage with the recent literature exploring heterogeneous treatment effects, and its implications

[9] In all of our regression analyses, we drop the 1988 election from the sample since for this year we only have information on winning candidates and thus we cannot establish whether a woman ran for office (unless she won the election). Furthermore, sample sizes are a bit smaller in Columns 3 and 4 due to missing information on winning candidates for some races. The results are unchanged when we restrict all regressions to the subset of races where the information about the candidate winning is provided.

Table 1: Term Limits and the Rise of Female Politicians

	Female Candidate		Female Mayor	
	(1)	(2)	(3)	(4)
Term Limit	0.172***	0.167***	0.106***	0.105***
	(0.014)	(0.014)	(0.011)	(0.012)
Municipal FE	Yes	Yes	Yes	Yes
Election FE	Yes	No	Yes	No
Province × Election FE	No	Yes	No	Yes
Observations	10,434	10,434	10,377	10,377
R-squared	0.308	0.353	0.342	0.376
Mean Dep. Var.	0.329	0.329	0.154	0.154

Notes: Results from municipality*elections regressions and include municipal and election year fixed effects. The dependent variable is a dummy for whether a woman ran for mayor (Columns 1-2), or a dummy for whether a woman was elected mayor (Columns 3-4). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

for the validity of the difference-in-difference estimator. This subsection serves two purposes. On one hand, it provides evidence supporting crucial identifying assumptions. On the other hand, it helps us gain a better understanding of our setting, and potential mechanisms driving our results.

Testing for the Parallel-Trend Assumption: We first provide evidence of the key identifying assumption in difference-in-difference designs: namely, that municipalities with and without forced-open-seat races were on similar trends before binding term limits. To do so, we estimate an extended version of equation (1) where we include leads and lags of the term-limit variable. More concretely, we estimate the following equation:

$$y_{mt} = \alpha + \sum_{S=-3}^{2} \beta^S Term\_Limit_{mt}^S + \mu_m + \rho_t + \epsilon_{mt}, \tag{2}$$

where β^S are the estimates on the full set of leads and lags (up to three election years). For example, $Term\_Limit_{mt}^1$ is the first lead of the term-limit dummy. It is equal to one for all term-limited municipalities one election before they became term-limited and zero otherwise.

Figure 2 provides strong evidence in support of the parallel trends assumption. The coefficients on the one and two period leads are small and statistically insignificant, which suggests that treated and control groups follow a common trend prior to term limits. This reinforces our confidence that the results discussed above capture the causal effect of binding term limits on female political participation.

In addition to documenting parallel trends, Figure 2 highlights an interesting feature. The effect of term limits becomes much weaker one and two periods *after* the forced-open-seat race. The coefficients on the one and two-period lags, while positive, are substantially muted compared to the contemporary term-limit dummy. We return to this finding below in Section 4.

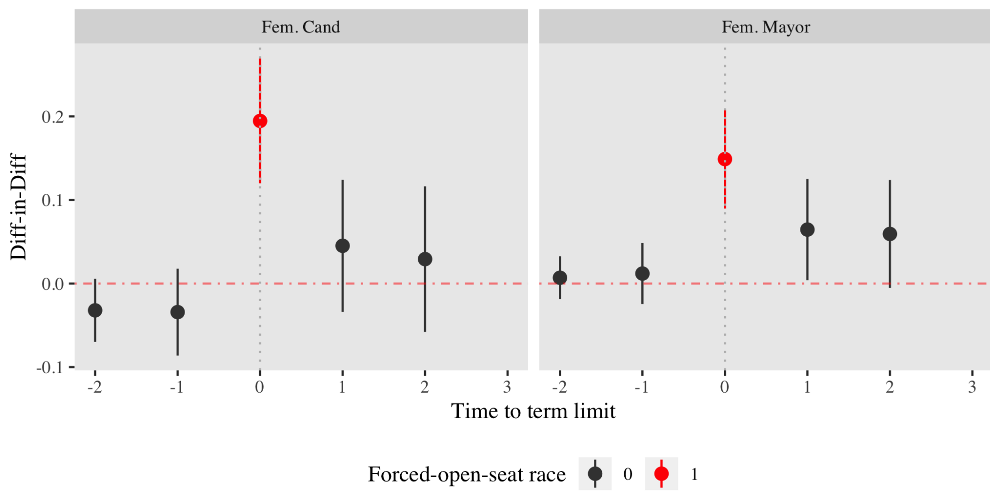

Figure 2: Estimates and 95% Confidence Intervals for Leads and Lags of Term Limit Variable

Exploring Potential Confounders: An important concern in any difference-in-difference strategy is that the treatment variable may be confounding another factor that ultimately drives the estimates. First, a substantial share (89.9%) of municipalities become term-limited at some point within our sample period. As such, the main difference between treated and control groups is the timing of forced-open-seat races. Second, from the previous discussion we know that the effect of term limits tends to short lived. These two points strongly suggest that any confounder would need to correlate with the timing of the forced-open-seat race, yet its effect should not persist after one election.

To explore the potential selection in the timing of forced-open-seat races, we construct five dummy variables for whether a municipality had a forced-open-seat race in 1998, 2001, 2004, 2007 or 2010. We then run a regression for each of the dummy variables on a set of socio-economic variables at the municipality level. We do so both on the full sample and on the sample of municipalities that became term-limited at some point. The socio-economic characteristics are measured using the 1990 census. Appendix Tables A.3 and A.4 report the results and show very little differences between municipalities term-limited in different election years. We also show, in Appendix Table A.5, that our main results are robust to controlling for these characteristics interacted with election-year fixed effects.[10]

Exploring Heterogeneous Treatment Effects as a Source of Bias: We also engage with the recent difference-in-difference literature that highlights heterogeneity in the treatment effect as a potential source of bias (Goodman-Bacon, 2018; Gibbons, Serrato and Urbancic, 2018).

Goodman-Bacon (2018) decomposes the two-way fixed effects regression difference-in-difference coefficients, which rely on variation in treatment timing, in a weighted average of all the simple two-by-two difference-in-difference estimators, i.e., subsamples with two groups (treated/control) and two periods (pre/post). This decomposition highlights that such estimators will estimate the average treatment effect, only when the treatment effects of all two-by-two difference-in-difference estimators are *homogeneous*. In the presence of heterogeneity, the OLS estimates are biased and one would need to re-weigh the data appropriately.[11]

To gauge for the presence of heterogeneity across election years, and whether a specific election year might be driving the results, we ran our main specification excluding one election year at a time. We show in Figures 3 and 4 that our estimates, both for the probability of a woman running and winning an election, are robust and stable to such exclusion, suggesting very little heterogeneity across time. In particular, one might have worried that our estimates relied heavily

[10]The regressions include a comprehensive set of municipal-level socio-economic characteristics measured in 1990: population, gender ratio, urbanisation, migration, asset ownership (land, house, ration, tv, phone, car and fridge) and education levels broken down by gender. Descriptive statistics for all 1990 census variables used in the analysis are reported in Appendix Table A.2.

[11]Goodman-Bacon (2018) also highlights that, in the presence of dynamic effects, where the treatment effects change over time within municipalities, using already treated units as controls biases the estimates of the treatment effect. In such situations, a better approach would be an event study estimation, in line with Figure 2. As we noticed, in our context there is very little persistence and our treatment effect estimate is driven by the election year with a forced-open-seat race.

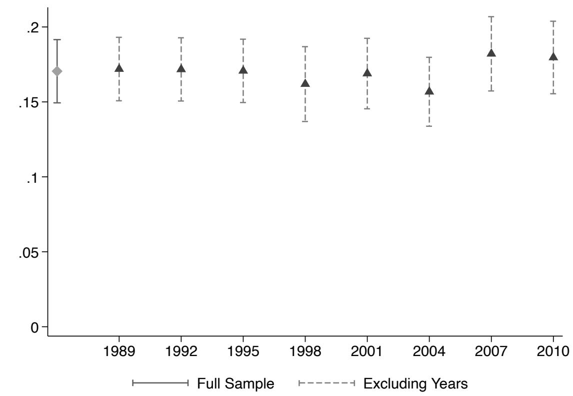

Figure 3: Estimates and 95% Confidence Intervals for Women Running - Excluding One Election Year at a Time

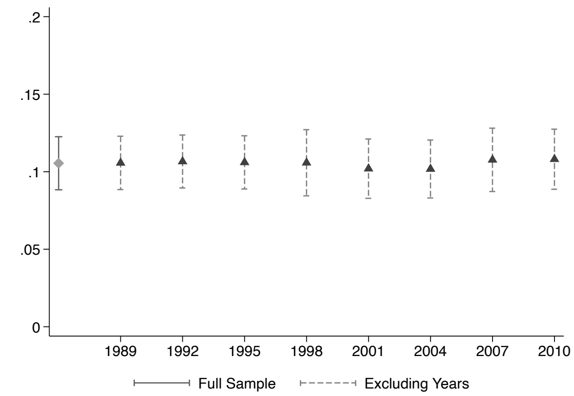

Figure 4: Estimates and 95% Confidence Intervals for Women Winning - Excluding One Election Year at a Time

on the 1998 election (the first election year with forced-open-seat races), potentially confounding other macroeconomic shocks in that year such as the Asian financial crisis. Finally, as mentioned above, in Goodman-Bacon (2018), the difference-in-difference estimates are a weighted average of estimates obtained when comparing early adopters and late adopters before late adopters are treated, early adopters and late adopters when the early adopters are treated and those treated at some point vs. the group never treated. Using the STATA command bacondecomp, we compute those estimates, and find that they range from .170 to .191 for the female candidate dummy and from .104 to .120 for the female elected dummy. These results show very little heterogeneity by election year. Overall, these two pieces of evidence suggest that our main findings are robust and are not due to heterogeneity by election year.

Heterogeneity through time is not the only source of heterogeneity. Gibbons, Serrato and Urbancic (2018), taking a more general approach, show that in the presence of heterogeneous treatment effects, adding fixed effects could lead to a biased difference-in-difference estimator. The underlying intuition is that this bias originates since the difference-in-difference estimator averages the different treatment effects with the wrong weights. Note that, in contrast to Goodman-Bacon (2018), which provides the source of heterogeneity — timing of the treatment — Gibbons, Serrato and Urbancic (2018) do not take a stance on the source of such heterogeneity. Rather than mine for potential sources of heterogeneity, in Table 2 we investigate the relevance of this source of bias by comparing the estimates of various specifications differing in the combination of fixed effects included. The first column has no fixed effects, the second column has only year fixed effects, the third column has only municipality fixed effects, and the last column is our main specification with both year and municipality fixed effects. The coefficients remain robust to these specifications. Although we can't fully rule out heterogeneity in the treatment effect, under Gibbons, Serrato and Urbancic (2018) setting, the estimated coefficients in Table 2 should be unstable as we move from one specification to another, given the different estimators will be using different weights. The stability of our estimates across these specifications gives us confidence about the main findings in Table 1.

Table 2: Term Limits and the Rise of Female Politicians — Robustness

Panel A	Female Candidate			
	(1)	(2)	(3)	(4)
Term Limit	0.179***	0.162***	0.188***	0.172***
	(0.012)	(0.013)	(0.013)	(0.014)
Municipal FE	No	No	Yes	Yes
Election FE	No	Yes	No	Yes
Observations	10,434	10,434	10,434	10,434
R-squared	0.022	0.027	0.303	0.308
Panel B	**Female Mayor**			
Term Limit	0.120***	0.105***	0.121***	0.106***
	(0.010)	(0.010)	(0.011)	(0.011)
Municipal FE	No	No	Yes	Yes
Election FE	No	Yes	No	Yes
Observations	10,377	10,377	10,377	10,377
R-squared	0.017	0.024	0.335	0.342

Notes: Results from municipality*elections regressions. The dependent variable is a dummy for whether a woman ran for mayor (Panel A), or a dummy for whether a woman was elected mayor (Panel B). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

2.3 Close Elections and Term Limits

Finally, we show the robustness of our findings by using an alternative empirical strategy: close elections. We focus on the set of municipalities in which an incumbent was the winner or runner-up in his/her race for a third consecutive term and the race was decided by a narrow margin.

Intuitively, these are municipalities in which the underlying level of popularity of the incumbent is similar. However, in some of them the incumbent wins (plausibly for idiosyncratic reasons that determine the outcome of close races) and thus the municipality ends up with a forced-open-seat race, while in others the incumbent loses and the municipality ends up with a new, non term-limited incumbent. Thus, these municipalities provide us with a natural experiment to study the impact of term limits on the electoral success of women.[12]

An important limitation of this methodology in our context, is that the sample of races where an incumbent faces a close election in its attempt for a third consecutive term is small. Intuitively, these races tend to be rather noncompetitive since strong challengers may prefer to wait one term for the seat to become open, rather than spend resources in challenging an incumbent. In fact, the average margin of victory in races featuring an incumbent running for a third consecutive term is 35 percentage points, and only 12% of these races have a winning margin of 5 percentage points or lower. In these exceptionally competitive races however, incumbents win close to half of them (52%) which is consistent with the outcome of these races being driven by idiosyncratic factors.

In Table 3 we report the coefficient on the dummy for whether the incumbent won its third consecutive race (which resulted in a forced-open-seat), on dummy variables for whether a woman ran or won in the following election. We focus on races with a winning margin smaller than 5% (columns 1-2) or 10% (columns 3-4). Reassuringly, the point estimates are similar to those reported for the entire sample in Table 1.

3 The Dynastic Nature of the Rise of Female Mayors

What drives the increase in female politicians during forced-open-seat races? Are the patterns consistent with the original goal of the policy which was to promote alternation in office and curb the power of incumbent political elites, namely political dynasties? Political dynasties play an important role in politics at both the national and local levels (McCoy, 2009) and have persisted across many decades. It is common for relatives to take turns holding the same office, and for the same family to control multiple elected offices at the same time (Querubin, 2012). To understand how term-limits interact with this prevalent role of political dynasties, we next explore

[12]See Eggers et al. (2015) for a discussion on the validity of research designs based on close elections.

Table 3: Close Elections: Term Limits and the Rise of Female Politicians

	Female Candidate	Female Mayor	Female Candidate	Female Mayor
	(1)	(2)	(3)	(4)
Incumbent Won	0.212***	0.075*	0.234***	0.116***
	(0.056)	(0.042)	(0.040)	(0.031)
Observations	271	270	549	547
R-squared	0.088	0.030	0.065	0.027
Mean Dep. Var.	0.299	0.146	0.299	0.146
Bandwidth	5%	5%	10%	10%

Notes: Results from municipality*elections regressions and include election year fixed effects. The dependent variable is a dummy for whether a woman ran for mayor (Columns 1 and 3), and a dummy for whether a woman was elected mayor (Columns 2 and 4). Sample is restricted to election years in which in the previous election, the incumbent ran for a third consecutive term and was either the winner or runner-up of the election. In columns 1 and 2 we restrict the analysis to cases in which the previous incumbent 's reelection attempt was decided by a margin lower than 5 percentage points. In columns 3 and 4 we restrict the analysis to cases in which the previous incumbent 's reelection attempt was decided by a margin lower than 10 percentage points. The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

the composition of candidates across different races between dynastic and non-dynastic. Contrary to the original intent of the policy, we provide evidence that forced-open-seat races lead to an increase in the number and share of dynastic politicians, particularly of incumbent relatives and show that the increase in female representation following binding term limits is driven by the entry of female relatives of the incumbent into the mayoral races.

A key step to analyzing changes in the composition of mayoral races is to identify dynastic candidates and the relatives of the incumbents in our sample. Individuals in the Philippines carry two family names: the mother's maiden name and father's family name (men and single women) or the father's family name and husband's family name (married women). We follow Querubin (2016) and Fafchamps and Labonne (2017) and classify a candidate as *dynastic* if s/he shares at least one family name with any current or previous mayor and as an *incumbent relative* if s/he shares at least one family name with the current mayor.

This approach assumes that a shared family name indicates an actual family tie, which is valid given the historical way in which family names were allocated in the Philippines. In 1849, Governor Narciso Claveria y Zaldua assigned a different family name to every household in each municipality. As a consequence very common family names are not as prevalent in the Philippines as in other countries and thus, sharing a family name is very strongly correlated with an actual family tie. For example, using biographical data, Querubin (2016) documents that the rate of false positives (i.e. candidates who share a family name but are not related to each other) amongst candidates for provincial offices (a larger sub-national level) is around 5%. This rate is likely lower at the municipal level since unique family names were originally allocated at this administrative level.[13] Fafchamps and Labonne (2017) and Cruz, Labonne and Querubin (2017) validate and discuss this method for tracing relatives in more detail.

Our empirical strategy is identical to regression (1), but we use a different set of outcome variables related to whether candidates or elected mayors are dynastic.[14] In Table 4 we focus on the composition of candidates and mayors, between dynastic and non-dynastic, independently of their gender. In Column 1, we use as a dependent variable the number of candidates in the mayoral race and find, as expected, that forced-open-seat races attract a larger number of candidates. However, once we turn to the composition of the pool of candidates between dynastic and non-dynastic candidates in Column 2, we find that the share of candidates that are dynastic (relative to all candidates) increases by 6 percentage points in forced-open-seat races. This is driven by the higher likelihood of dynastic candidates (Column 3) and in particular incumbent relatives (Column 4), running in forced-open-seat races. The point estimates show that incumbent relatives are almost 50 percentage points more likely to run in forced-open-seat races relative to other races. Finally, the dependent variables in Columns 5 and 6 are dummies for whether dynastic candidates or incumbent relatives *win the election*, respectively. The point estimate in Column 6 shows that incumbent relatives are 33 percentage points more likely to get elected following binding term limits. Contrasting the estimated coefficient to the share of mayors that are relatives of the incumbent in non-open seat races (4.7%), incumbent relatives are 7 times (33/4.7) more likely to

[13]For example, in principle there should be one Aquino family per municipality, even if the Aquino family name may exist in several municipalities. Two individuals with the Aquino family name living in the same municipality are thus much more likely to share an actual family tie than two individuals with the Aquino family name from different municipalities in the same province.

[14]We show the results for our preferred specification with municipality and election year fixed effects. The point estimates are very similar when we include province-specific election-year fixed effects.

become mayors in forced-open-seat races. In sum, our findings show that while forced-open-seat races attract more candidates, they do not necessarily lead to an increase of "new blood" in the system.

Table 4: Term Limits and the Rise of *Dynastic* Politicians

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A			All Candidates:			
	No. Cand.	Share Dyn.	Dyn. Ran	Inc. Rel. Ran	Dyn. Won	Inc. Rel. Won
Term Limit	0.615***	0.061***	0.200***	0.467***	0.139***	0.336***
	(0.032)	(0.006)	(0.012)	(0.013)	(0.011)	(0.012)
Observations	10,437	10,437	10,437	10,437	10,437	10,437
R-squared	0.494	0.502	0.491	0.375	0.483	0.339
Mean Dep. Var.	2.890	0.202	0.419	0.215	0.267	0.116

Notes: Results from municipality*elections regressions and include municipality and election-year fixed effects. The dependent variable is the number of candidates (Column 1), the share of candidates who are dynastic relative to the pool of candidates (Column 2), a dummy equal to one if a dynastic candidate ran (Column 3), a dummy equal to one if an incumbent relative ran (Column 4), a dummy for whether the elected mayor is dynastic (Column 5) and a dummy for whether the elected mayor is related to the previous incumbent (Column 6). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Turning to the sample of female candidates in Table 5, in Columns 1-3 we use as dependent variables dummies for whether a dynastic, incumbent relative or non-dynastic woman ran in the race, respectively. The estimates show that the increased probability of women running for office documented in Table 1 is mostly driven by dynastic candidates. Dynastic women are 17 percentage points more likely to run in forced-open-seat races (Column 1). The effect is even larger for female relatives of the incumbent who are 24 percentage points more likely to run following binding term limits (Column 2). The effect for non-dynastic women (Column 3), while positive and statistically significant, is substantially smaller than that for women related to previous or current incumbents.

Investigating the composition of winning candidates (female mayors), in Columns 4-6 we use dummies for whether dynastic, incumbent relatives or non-dynastic women won the election, respectively. The estimates show that the increase in the number of female politicians following binding term limits is driven entirely by the entry of dynastic female politicians, in particular, by women related to the term-limited incumbents (Column 5) who are 16 percentage points more likely to win in forced-open-seat races. In fact, close to 75 percent of the women elected in the first wave of forced-open-seat races in 1998 were related to the term-limited incumbent and out of the 342 dynastic women winning an election in open-seat races, only 15 are dynastic but unrelated to the term-limited incumbent. On the other hand, the coefficient in Column 6 shows that non-dynastic women are not more likely to win forced-open-seat races.

In sum, the descriptive patterns and regressions presented in Figure 1 and Tables 1 - 5 provide strong evidence that forced-open-seat races constitute critical junctures in which women are disproportionately more likely to access political office. However, not all women benefit from term-limits: relatives of the term-limited incumbent are substantially more likely to win in forced-open-seat races.[15]

Table 5: Term Limits and Female *Dynastic* Politicians

	(1)	(2)	(3)	(4)	(5)	(6)
	Female Candidates:					
	Dyn. Ran	Inc. Rel. Ran	Non-Dyn Ran	Dyn. Won	Inc. Rel. Won	Non-Dyn Won
Term Limit	0.171***	0.244***	0.044***	0.105***	0.162***	0.000
	(0.011)	(0.011)	(0.012)	(0.009)	(0.009)	(0.007)
Observations	10,436	10,436	10,437	10,434	10,434	10,377
R-squared	0.324	0.259	0.307	0.335	0.230	0.363
Mean Dep. Var.	0.127	0.0705	0.227	0.0788	0.0420	0.0751

Notes: The dependent variable is a dummy for whether a dynastic woman ran (Column 1), a dummy for whether a woman related to the incumbent ran (Column 2), a dummy for whether a non-dynastic woman ran (Column 3), a dummy for whether a dynastic woman won (Column 4), a dummy for whether a woman related to the incumbent won (Column 5), and a dummy for whether a non-dynastic woman won (Column 6). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

4 The Gendered Nature of Dynastic Politics and Term Limits

So far we've shown that term limits led to an increase in female political representation in mayoral positions and that most of these women "succeeded" their term-limited relatives in office. This suggests that a key factor behind the rise of female mayors in the Philippines is related to political successions and the response of term-limited incumbents when faced with the constraint imposed

[15]The fact that voters often vote for and elect female relatives of term-limited incumbents is consistent with the argument of Dolan (2014) that traditional political forces (such as membership in the incumbent dynasty) are often a more important determinant of voter behavior than a candidate's gender.

by term limits on their political careers. In this section, we explore two plausible explanations for the role of term limits in the increase in the probability of female relatives of the incumbent running for office. We focus on the likelihood of running for office, rather than on the likelihood of winning, since incumbents only have full control on whether a relative runs for office (not on whether they actually win). Moreover, recall from Section 2.1 that the increase in female representation in forced-open-seat races is mostly driven by the higher likelihood of women running for office, rather than of women winning conditional on running.

Specifically, the share of female relatives running for office is the product of i) the likelihood of incumbent relatives running for office and ii) the gender composition of these incumbent relatives. Let *Prob*(*FemRel*) be the probability that a female relative of the incumbent runs for mayor. We can use the following identity to illustrate the point:

$$Prob(FemRel) = Prob(Rel) \times Prob(Fem|Rel),$$

where *Prob*(*Rel*) is the probability that a relative of the incumbent runs for mayor and *Prob*(*Fem*|*Rel*) is the probability that, conditional on a relative of the incumbent running, the relative is female. In what follows, we will explore quantitatively the significance of each of these components in explaining the change in the share of female relatives running in forced-open-seat races.

4.1 Term Limits and the Likelihood of Dynastic Successions

Based on the above expression, a first explanation behind the rise of female relatives running for office is that, as shown in Table 4, term limits are associated with an increase in the likelihood of incumbent relatives running for office (an increase in the likelihood of dynastic successions). As documented by Querubin (2012), this reflects the adaptive strategies of term-limited incumbents who want to continue their political careers and preserve their political power. In a context like the Philippines, where incumbents enjoy a very large electoral advantage (Querubin, 2016), it is essential to prevent opponents from gaining access to office in forced-open-seat races. Given the weakness of political parties in the Philippines (Montinola, 1999) and the importance of families in politics, term-limited incumbents often attempt to maintain political control by having a relative run to replace them. Moreover, as mentioned above, previous studies have documented the importance of political dynasties as a vehicle for women's access to office both for executive positions and for legislatures in other developed and developing countries (Dal Bo, Dal Bo and Snyder, 2009; Jalalzai, 2013; Chandra, 2016; Folke, Rickne and Smith, 2016; Smith and Martin, 2017; van Coppenolle, 2017; Jalalzai and Rincker, 2018). In our setting, during 1988-1995, prior to any binding term-limits, the female share amongst candidates related to the incumbent ($Prob(Fem|Rel)$) was 0.15. Can the increase in the likelihood of dynastic candidates running for office (documented in Table 4) explain the increased probability of female relatives running and the ensuing rise in female political representation? Specifically, what would be the predicted change in the likelihood of female relatives running for mayor ($Prob(FemRel)$) if we held the female share of dynastic successions constant at 0.15, but account for the increased likelihood of incumbent relatives running for office?

Recall from Table 4, Column 4, that forced-open-seat races are associated with an increase of 46 percentage points (0.46) in the probability of an incumbent relative running for office. Thus, $\Delta Prob(FemRel) = 0.46 \times 0.15 = 0.069$. Contrasting this number with the estimated difference-in-difference coefficient in Column 2 of Table 5 we can see that the change in the likelihood of dynastic successions can account for 29% of the 24 percentage increase in the probability of female relatives of the incumbent running for office (0.069/0.24=0.29). This suggests that the increase in the number of "dynastic successions" following term-limits, holding everything else constant, plays an important role, but leaves 71% of the estimated change in the likelihood of female relatives running for office unexplained.

4.2 Term Limits and the Gender Composition of Dynastic Successions

This brings us to our second explanation: the change in the gender composition of dynastic successions. Recall that the share of female relatives running for office (conditional on a relative running for office) is 0.15 during the baseline period (1988-1995), prior to any binding term limits. This share increases to 0.45 in forced-open-seat races. This observed difference roughly coincides with the difference-in-difference estimate reported in Column 2 of Table 5. In other words, the share of female relatives differs substantially in dynastic successions following binding term limits compared to other dynastic successions (that occur due to different causes such as death or retirement, for example). Why do term limits change the gender composition of dynastic successions? We investigate two plausible explanations rooted in the design of the term limits policy in the

Philippines.

4.2.1 Age Profile of Term-Limited Incumbents and Availability of Eligible Male Relatives

A first explanation is related to the fact that term-limits force incumbents out of office after three terms. As a consequence, strong incumbents are forced to exit their positions at a younger age, on average, compared to strong incumbents in standard (non-forced) succession decisions that do not follow binding term-limits. We argue that the younger age of the exiting incumbent can generate large differences in the gender composition of political successions in line with what we observe in our setting. This is because the term-limited incumbent is less likely to have children (in particular sons) of eligible age (which in the Philippines is 21) compared to other successions, and hence he may resort to his wife to succeed him in office. This argument is better exemplified by the story of one prominent political family in the Philippines. Jejomar Binay was elected mayor of Makati in 1988. He was re-elected in both 1992 and 1995 and thus was term-limited in 1998 at age 55. At that time, Jejomar did not have any sons of eligible age. His kids were Nancy (25, female), Abby (23, female), Jun (20, male), Marita (19, female) and Joanna (10, female). Elenita, his wife, ran and was elected for the position but did not run in 2001. She was replaced by her husband who was again re-elected in 2004 and 2007. In 2010, when he became term-limited again, his son Jun (now of eligible age) ran and was elected.

Next, we use a simple quantitative simulation to show that, even in a world with a strong male preference in dynastic succession decisions, this channel can generate changes in the share of female candidates similar in magnitude to the ones observed in the data, under plausible assumptions. To that end, suppose that incumbents prefer a male succession: the exiting incumbent chooses an eligible male successor as long as there is one available, otherwise he chooses amongst the potential set of eligible female successors (consisting of the wife and daughters). Consider first a standard succession in which a male politician voluntarily retires or dies. In this case, such politician is likely to be of older age and succeeded by one of his children. Fertility data for wealthy Filipino couples in the 70s suggests an average of 3 children per family, which given a gender ratio of 50%, would imply a share of male relatives in this succession of around 87.5%.[16] Interestingly, this unconstrained male biased scenario is close to what we observe in non-term-limited successions

[16]The estimates for the number of children per family comes from Table 5 in Hirschman and Guest (1990). In 1975-79,

where the share of women related to the incumbent is around 15%.

As mentioned above, dynastic successions following term limits differ from standard successions as they likely affect politicians at a younger age. As a consequence, their children might not be old enough to run for office and thus, their wives may become important (or the only) contenders in the succession. As politicians get older, they become more likely to have older children who can run to replace them. To illustrate this point, assume the same male-bias applies to dynastic successions following binding term limits. In other words, term-limited incumbents select an eligible male relative when available, otherwise they select their wives or eligible daughters. In order to account for the fact that sons may not yet be of eligible age, we introduce parameters p_1, p_2 and p_3, as the probability that, respectively, the first, second and third child is eligible to run. Thus, the share of potential male successors becomes: $p_1 * 1/2 + p2 * 1/4 + p_3 * 1/8$.[17] Using statistics on the age distribution of fathers from Official Vital Statistics of the Philippines at the end of the 70s, we can estimate p_1, p_2 and p_3 for different age values of the term-limited politician, given a representative politician with three kids and an age difference of two years between the kids.[18] We can then simulate the probability of having a male relative of eligible age for incumbents aged 45 to 64 years old. Figure 5 shows the simulated (predicted) share of male relatives running in our sample. As the age of the term-limited incumbent increases, the share of male relatives running converges towards the share in other non-term-limited successions. Unfortunately, we do not have data on the age of incumbents in our sample. However, assuming that politicians are uniformly distributed between 45 - 64 years old at the moment of first becoming term-limited, the probability of having an eligible male relative would be, on average, 54% which coincides with the actual (observed) fraction of male relatives running for office in term-limited successions.[19]

the Total Fertility Rate of women with at least some College education (as a proxy for wealth) was 3.08. 87.5% comes from the probability of having at least one son, conditional on having three children: i.e., $1 - (1/2)^3$.

[17]The formula comes from the probability of having a first son weighted by the probability of the first son being of eligible age ($p_1 * 1/2$), the probability of having a middle son weighted by the probability that the middle son is of eligible age ($p_2 * 1/4$), and the probability of having a third son weighted by the probability that the third son is of age ($p_3 * 1/8$).

[18]See `https://psa.gov.ph/sites/default/files/1978%20Vital%20Statistics%20Report.pdf` accessed on September 5, 2019.

[19]These numbers would be in line with mayors being first elected when they are 35 – 45 years old, which is consistent

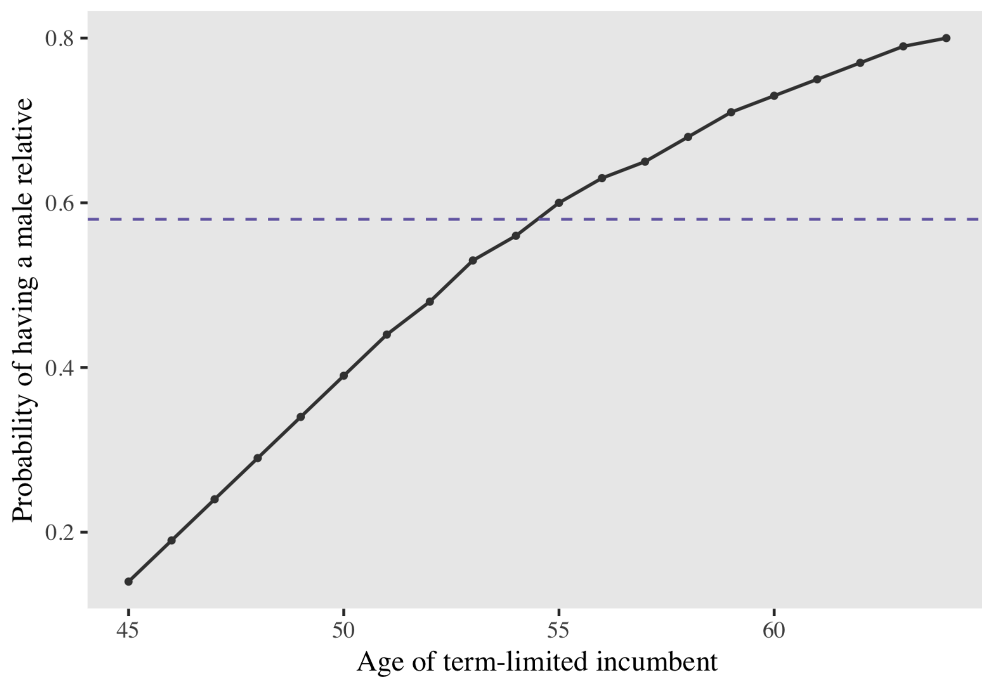

Figure 5: Simulation probability of having a male relative running

It is worth emphasizing the importance of wives for this channel, which is a by-product of the high share of male incumbents in the Philippines. For example, close to 92% of term-limited incumbents in 1998 were men. This high share of male politicians elected in unconstrained races is necessary for term-limits to generate a discrete jump in the gender composition of dynastic successions. Had the majority of term-limited incumbents in 1998 been female, then term-limits would have increased the likelihood of having *male* mayors, since term-limited female incumbents would have had eligible male relatives (their husbands) available for the succession.

4.2.2 Women as Benchwarmers

For our second explanation behind the change in the gender composition of dynastic successions, recall from Section 2 that term limits in the Philippines impose a short-term constraint on the incumbent, forcing the politician out of office for only one term. This contrasts with other standard dynastic successions (following retirement or death) where the incumbent's exit is more permanent. Thus, since term limits generate only a short-run (i.e. one term) constraint on the incumbent, those interested in running again after one term may be more likely to appoint a female relative in office to

with statistics for other countries.

serve as *benchwarmers*. This is a Filipino expression for incumbent relatives who serve for only one term after the forced-open-seat race, and retire immediately after to allow their relative to return to office (Coronel et al., 2004). Elenita, the wife of Jejomar Binay, is an example of a benchwarmer. We argue that female relatives might be more attractive than male relatives as benchwarmers due a number of reasons. First, consider a term-limited incumbent who wants to remain influential during his term out of office and to run again immediately after. Given gender norms in the Philippines, women may be more willing to comply with the relatives' directives while in office and to retire immediately after to allow his return. Similarly, in a context where the careers of men are perceived as more important than those of women, short-term distortions on the life (and career) of the female relatives might be perceived as less costly. For example, according to the 2009 Labor Force Survey, female labor force participation is 19 percentage points lower for women than for men (gap amongst those with college education is similar). Moreover, 60% of Filipino respondents in the 6^{th} wave of the World Values Survey (2010-2014) agree that whenever jobs are scarce, men should have more right to a job than women. Finally, 56.4% of respondents agree with the statement that "men make better political leaders than women do" (fraction who agree amongst male respondents is 63.7%). This is also consistent with findings by Fox and Lawless (2014) and Lawless and Fox (2005) that women are significantly less likely than men to demonstrate ambition to run for elective office. If incumbents are interested in returning to office after one term, they may prefer to select female relatives who will be less inclined to remain in office and continue a political career of their own.

In line with this explanation, we find that female mayors who replaced their term-limited relative are more likely to be one-termers – that is, to exit after serving only one term – than other mayors. In particular 57.7% of women who replaced their term-limited relative serve for only one term. This is higher than the share of one-termers amongst non-dynastic women (44.5%) or men (38%) elected in forced-open-seat races. It is also higher than the share of one-termers amongst women elected in non-forced-open-seat races (37%).[20] Female mayors who replaced their term-limited relative are also more likely to be *benchwarmers*. This refers specifically to cases in

[20]These differences in the share of one-termers between female relatives of the incumbent elected in forced-open-seat races and other mayors are statistically significant at the 1% level.

which an incumbent relative replaces a term-limited incumbent, serves for only one term *and is then immediately replaced by the previously term-limited incumbent.* Out of 114 benchwarmers in our dataset, 80 (70%) are female. Moreover, out of the 259 women who replaced a relative in a forced-open seat, 31% are benchwarmers, while the corresponding fraction for men is 10%. As a consequence, male relatives of the incumbent serve on average a larger number of terms (2.23) than female relatives (1.81). This may also explain the small coefficient on the one and two period lags of the term-limit dummy reported in Figure 2. The high prevalence of benchwarmers amongst dynastic women is in line with the effect of term-limits on female representation being short-lived (a result to which we return below when we test for the existence of electoral “role model" effects on other local offices).

The benchwarmer effect and the lower average number of terms served by female politicians elected in forced-open-seat races, point at two important forces behind the steady increase in the share of municipalities with a female mayor in the Philippines in the last 30 years. While there is an increase in the number of female mayors accessing office during forced-open-seat races, there is a decrease in the average number of terms these women stay in power (as well as an increase in the number of one-termers), compared to other races. In other words, while following binding term limits a larger number of municipalities experience a female mayor, these female mayors are on average less likely to be re-elected. These two forces, the number of forced-open-seat races and the share of women exiting after one term, generates a delicate balance of inflows and outflows that translate into the observed increase in the share of municipalities with a female mayor. For example, the number of female mayors in 1995 was 147 and it increased to 237 in 1998. This net increase in 90 female mayors consists of 170 new female mayors (of which 133 were elected in forced-open-seat races) and 80 outgoing female mayors who either did not run or were not re-elected. A similar pattern can be observed in 2001: out of the 237 female mayors, 94 decided not to run, while 40 were term-limited (for an outflow of 134 female mayors). Importantly, 59 out of the 94 female mayors who decided to not run another term, were relatives of the term-limited incumbent in 1998 and served for only one term. Despite the higher share of female mayor outflows in 2001 (40% or 94 out of 237) than in 1998 (23% or 30/133) there was still a net increase in the total number of female mayors to 246 due to the inflow of 143 new female mayors (83 of them elected in forced-open-seat races). The key observation is that absent the benchwarmer effect (or higher likelihood of one-termers amongst female relatives of term-limited incumbents), the net increase in the number of municipalities with a female mayor would have been much larger.

To sum up, in this section we provide evidence that the age-profile of term-limited incumbents and the use of benchwarmers by term-limited incumbents interested in returning to office after one term can help us explain the change in the gender composition of dynastic successions in forced-open-seat races. Returning to our original question, what fraction of the change in the fraction of women running for office can be explained by the increase in the fraction of women in dynastic successions? We conduct a similar exercise to the one reported in Section 4.1, but use the baseline (1988-1995) fraction of incumbent relatives running ($Prob(Rel) = 0.08$) and predict the change in the likelihood of women running for office, using the change in the fraction of incumbent relatives that are female: $\Delta Prob(FemRel) = 0.08 \times 0.3 = 0.024$. Thus, the change in the female ratio of dynastic successions by itself can account for about 10% (0.024/0.24) of the change in the likelihood of women running for office. This suggests that the interaction of both forces – an increase in the likelihood of dynastic successions and in the female share of these successions– provides a compelling account of how term limits led to increased female political representation in the Philippines, as each force by itself is not enough to explain the jump in the fraction of women running for office.

5 Female Dynastic Mayors: Policy Impacts and Role Model Effects?

An important question is whether, in our context, the election of dynastic women translates into different policy choices or encourages other women to join the labor force, enroll in school or run for public office. On one hand, having a woman in office might lead to the enactment of pro-women policies, in line with the literature on female political representation (Chattopadhyay and Duflo, 2004). On the other hand, the sole fact of having a woman in office could lead to demonstration or role model effects via greater aspirations or political participation by other women (Phillips, 1995, 1998; Bhavnani, 2009; Beaman et al., 2009, 2012; Broockman, 2014; Bhalotra, Clots-Figueras and Iyer, forthcoming). Yet, in contrast to our setting, in these studies, women accessed their political positions via channels with a clear gender mandate such as reservation seats or gender quotas.

This is an important distinction. In our setting, women's political positions are tied to their relatives, as a strategy for the dynasty to hold on to power. This could generate a different set of incentives and ultimately lead to little differences in policy outcomes or role model effects between male and female mayors, related to the term-limited incumbent. To start with, female mayors related to the term-limited incumbent might have policy preferences aligned with those of their family members. Similarly, the fact that many of these women are benchwarmers may limit their ability to shift policies towards their own preferences (or those of other women). This could further feed back to their (lack of) ability to become role models for other women in society. We further explore these questions empirically.

We focus on the sample of municipalities in which an incumbent was term limited and replaced by a relative.[21] This allows us to partial out the effect of term limits (election of new incumbents) and the effect of having a dynastic incumbent (male or female), as all municipalities are term-limited and have a relative replacing the incumbent. We perform a difference-in-difference analysis in which we compare policy and other economic outcomes during the first term of the dynastic mayor to those of the last term of the term-limited incumbent in municipalities in which female vs. male relatives were elected. We restrict our analysis to cases where the term-limited incumbent was *male*. This makes the interpretation easier as cases in which a female relative is elected correspond to transitions from a male to a female mayor. However, the results do not depend on this restriction.[22]

More concretely, we estimate the following municipal fixed-effect regression model:

$$Y_{mt} = \alpha + \beta Post\_Term\_Limit_{mt} + \gamma Post\_Term\_Limit_{mt} x Female_{mt} + \mu_m + \rho_t + u_{mt}, \quad (3)$$

[21]Notice that we do not compare male and female mayors in general, but rather male and female mayors *related to an incumbent*. Although an overall comparison of policies by male and female mayors is interesting per se, the relevant predictions in our context are related to the difference between male and female mayors related to an incumbent (not across male or female politicians in general). Furthermore, the overall difference between male and female mayors presents a number of identification challenges that are less pronounced when one compares male and female relatives instead. Specifically, a simple comparison between municipalities run by female and male mayors would likely confound many other municipal characteristics. Richer or more educated municipalities may be more likely to elect women. Furthermore, most female mayors are dynastic and elected following binding term limits. It would be difficult to disentangle the effect of gender, from that of having a new incumbent, and having a dynastic mayor. Note that this could also be a concern in an research design focused on close elections.

[22]A disproportionate majority of term-limited incumbents are male (around 90 percent).

where Y_{mt} is the outcome of interest in municipality m, *Post_Term_Limit* is a dummy that equals 1 during the incumbent relative's first term (i.e. in the term immediately after the forced-open-seat race), and $Female_{mt}$ is a dummy that equals 1 if a female relative of the incumbent was elected.[23] Our coefficient of interest is γ.

We begin by looking at economic policies, more concretely public finance variables such as the amount and sources of revenue and the composition of expenditures. These variables are directly under the control of the municipal mayor. Detailed budget data are only available for the years 2000-2009, and thus we can only study the election of term-limited incumbents' relatives starting in 2001.[24] The results, reported in Table 6, provide no evidence that female mayors raise more or less revenues (Column 1), rely less on transfers from the central government (Column 2) or raise more local taxes (Column 4) than their male counterparts. Nor do female mayors run smaller or larger budget deficits (Column 3). Most importantly, Column 5 shows that female mayors are not more likely to spend on education and health, policy issues often associated with female preferences.[25]

Next, we study variables that may reflect the potential role model effect of dynastic women on other women in the municipality. In Table 7 we test whether the election of a female incumbent relative has an impact on the share of working-age women: who join the labor force (Column 1), become employed (Column 2), work in the public sector (Column 3) and overseas foreign workers (OFW, Column 4) or who have a permanent contract (Column 5). The coefficients are small relative to the sample mean and are only significant at the 10% level in one case.

In Table 8 we show that the election of female dynastic mayors does not influence overall (Column 1), elementary (Column 2) or secondary (Column 3) female school enrollment.

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[23]Since we restrict our analysis to male term-limited incumbents, $Female_{mt}$ can only equal 1 after the binding term limit.

[24]For incumbent relatives who entered office in 2001, the "pre" years of the term-limited incumbent's last term are 2000 and 2001, while the relative's first term is 2002-2004.

[25]We use survey data to compare the policy preferences of men and women collected for 3,400 households in 12 municipalities in the provinces of Ilocos Norte and Ilocos Sur (Cruz, Keefer and Labonne, 2017). Respondents were asked to indicate the share of the municipal budget that they would like to spend across 10 sectors. The results, reported in Appendix Table A.6, show that women report a stronger preference than men for expenditures on health, education, emergencies and business loans, and a lower preference for roads, community facilities and agriculture assistance.

Table 6: Policy Impacts: Public Finance

	(1) Log Revenues	(2) IRA/Revenues	(3) Exp/Rev	(4) Loc. Rev/Rev	(5) % Health & Ecuc
Post Term Limit x Female	0.01	0.66	-1.15	0.28	0.00
	(0.02)	(0.86)	(1.70)	(0.44)	(0.31)
Observations	1,360	1,360	1,360	1,360	1,359
R-squared	0.59	0.04	0.06	0.03	0.03
Mean Dep. Var.	17.7	81.9	90	14.9	10.6

Notes: Results from municipality*year regressions and include municipal and year fixed effects. The dependent variables are the (log) municipal budget (Column 1), IRA as a share of the municipal budget (Column 2), total expenditures as a share of the municipal budget (Column 3), revenues collected locally as a share of the municipal budget (Column 4), and education and health expenditures as a share of total expenditures (Column 5). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Table 7: Role Model Effects: Female Employment

	(1) Fem. Lab. Force	(2) Fem Emp.	(3) Fem. Pub. Sec.	(4) Fem. OFW	(5) Fem. Perm.
Post Term Limit x Female	1.41	1.24	0.07	-0.42*	0.60
	(1.15)	(1.17)	(0.36)	(0.22)	(1.97)
Observations	823	823	823	823	823
R-squared	0.01	0.01	0.02	0.04	0.04
Mean Dep. Var.	49.9	46.2	4.62	2.03	76.3

Notes: Results from municipality*year regressions and include municipal and year fixed effects. The dependent variables are the share of the female working-age population that is in the labor force (Column 1), employed (Column 2), employed in the public sector (Column 3), working as a overseas foreign workers (OFW) (Column 4) and employed on a permanent contract (Column 5). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Table 8: Role Model Effects: Female School Enrollment

	(1)	(2)	(3)
	Enrollment of Girls Aged:		
	5-16	5-11	12-16
Post Term Limit x Female	-0.37	-0.48	0.27
	(1.04)	(1.15)	(1.40)
Observations	823	822	823
R-squared	0.95	0.93	0.91
Mean Dep. Var.	59.4	59.3	59.5

Notes: Results from municipality*year regressions and include municipal and year fixed effects. The dependent variables is share of girls age 5-16 who are enrolled in school (Column 1), share of girls age 5-11 who are enrolled in school (Column 2), share of girls age 12-16 who are enrolled in school (Column 3). The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Finally, in Table 9 we test whether female dynastic mayors encourage other women to run for public offices. We estimate equation (3) but focus on whether women are more likely to run or win elections for Vice-Mayor (Columns 1-4) or the Municipal Council (Columns 5-8) in the election *following* the forced-open seat race when the incumbent relatives got elected.[26] All point estimates are very small and not statistically significant, which suggest electoral-role-model effects were limited.

Although these results suggest no policy or role-model effect, one potential concern with these results is that the gender of the relative of the term-limited incumbent may confound other socio-economic and political characteristics correlated with the electoral success of women. To investigate this, we regress $Female_{mt}$ on a broad set of municipality characteristics from the 1990 population census. In Appendix Figure A.1 we report the coefficients for all of these variables and show that they are small, and they are not statistically significant. We also show that female relatives of the incumbent face a similar number of challengers and win by similar margins as their male counterparts (see Appendix Table A.7). Finally, the electoral success of the term-limited incumbent, measured by the win margin in the last race, or the average win margin in the three races, is similar for incumbents replaced by male and female relatives. These patterns suggest that the variation

[26]Thus, in this case $Post\_Term\_Limit_{mt} = 0$ during the forced-open-seat race in which incumbent relatives got elected and $Post\_Term\_Limit_{mt} = 1$ in the following election.

Table 9: Role Model Effects: Election of Women to Other Offices

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Vice-Mayor				Council			
	# Cand.	Fem. Run	% Fem Cand.	Fem Won	# Cand.	Fem. Run	% Fem Cand.	Fem Won
Post Term Limit x Female (t-1)	0.11	-0.02	-0.02	0.01	0.75	0.03	-0.00	-0.06
	(0.12)	(0.06)	(0.02)	(0.03)	(0.76)	(0.02)	(0.01)	(0.04)
Observations	872	872	872	872	872	872	872	872
R-squared	0.17	0.02	0.00	0.01	0.27	0.01	0.01	0.01
Mean Dep. Var.	2.64	0.28	0.12	0.097	23.6	0.97	0.16	0.77

Notes: Results from municipality*year regressions and include municipal and election-year fixed effects. The dependent variables are the number of candidates for vice-mayor (Column 1), a dummy for whether a woman ran for vice-mayor (Column 2), the share of vice-mayoral candidates that are women (Column 3) a dummy for whether a woman was elected vice-mayor (Column 4), the number of candidates for councilor (Column 5), a dummy for whether a woman ran for councilor (Column 6), the share of council candidates that are women (Column 7) and a dummy for whether a woman was elected councilor (Column 8). The standard errors (in parentheses) account for potential correlation within municipality. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

in the gender of the incumbent's relative may be driven by idiosyncratic variables such as the incumbent's family structure, and does not confound other features of the municipality or the incumbent's political career, in line with our argument about the age structure of the term-limited incumbent.

6 Conclusion

In this paper we show that term limits - a widely used institution across several democracies - can lead to an increase in female representation and explain the dramatic rise of female mayors in the Philippines. However, we show that in the Philippines this effect was mostly restricted to female relatives of the incumbent. This finding has potential implications for the extent to which an increase in female descriptive representation translates into substantive representation or the empowerment of other women in dynastic contexts. Previous research has shown that women who were elected to office via reservation seats or gender quotas have enacted pro-women policies and have had demonstration or role model effects on the political participation of other women (Chattopadhyay and Duflo, 2004; Bhavnani, 2009). However, our findings suggest that the channel through which women access office mediates their impact on other policy or electoral outcomes.

What explains the limited impact of female dynastic mayors on electoral and economic outcomes in our context? Indian women who accessed office via reservations were also often relatives of previous incumbents: 17% were the spouse of a former Panchayat councilor or Pradhan. They also reported governing with the help of their spouse and were more likely to claim that they will not run again.[27] Yet, these female Panchayat councilors or Pradhans implemented different policies, and had lasting role model effects. Why? One possibility is that in contrast to our setting, in these studies, women accessed their political positions via a gender mandate such as reservations or gender quotas.

This is an important distinction. In our setting, women's political positions are tied to their relatives, as a strategy for the dynasty to hold on to power. This could generate a different set of incentives, ultimately leading to little differences in policy or role model outcomes between male

[27] There is also emerging evidence of similar patterns in Indonesia where parties comply with gender quotas by using relatives of male politicians Choi (2019); Hillman (2018).

and female mayors related to the term-limited incumbent. To start with, female mayors related to the term-limited incumbent might have policy preferences aligned with those of their family members. As argued by Jalalzai (2013), women might derive their political identities through their close male relatives and may be expected to further the political goals of the male relatives they succeeded. Similarly, if women are elected through the dynastic channel, they may not have a mandate that enables (or encourages) them to prioritize women's needs and preferences, in contrast to contexts where women are elected via gender quotas (Franceschet and Piscopo, 2008) or reservations (Chattopadhyay and Duflo, 2004).

Finally, the new pool of elected dynastic women may not be representative of the broader pool of women. This could impact whether their policy choices will benefit the majority of women. For instance, Clots-Figueras (2011) find that in India, female legislators of lower castes, but not those of higher castes, are more likely to enact female-friendly policies.

While the evidence presented in this paper is specific to the Philippines, we believe that many of the insights are relevant for other democracies where term limits are common and where parties are weak and/or political dynasties play an important role. Moreover, some of the specific channels we have explored (such as the differential age profile induced by term limits) likely apply to many other contexts. Our findings, however, highlight the importance of understanding how the cultural and social context mediates the effect of widely used political institutions such as term limits. In dynastic contexts, term limits may, perhaps inadvertently, increase female representation but restrict this higher access to office to dynastic women. And this may have important consequences for the role that these newly elected female politicians can play. As stated by Jalalzai (2013), "in spite of the rising numbers of women executives, we should question women's ultimate progress in achieving powerful positions" (p. 114). Where women's increased access to elected office is mainly driven by family connections, the consequences for substantive representation may be very different.

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Political Dynasties, Term Limits and Female Political Representation: Evidence from the Philippines

By Julien Labonne, Sahar Parsa and Pablo Querubin

Online Appendix

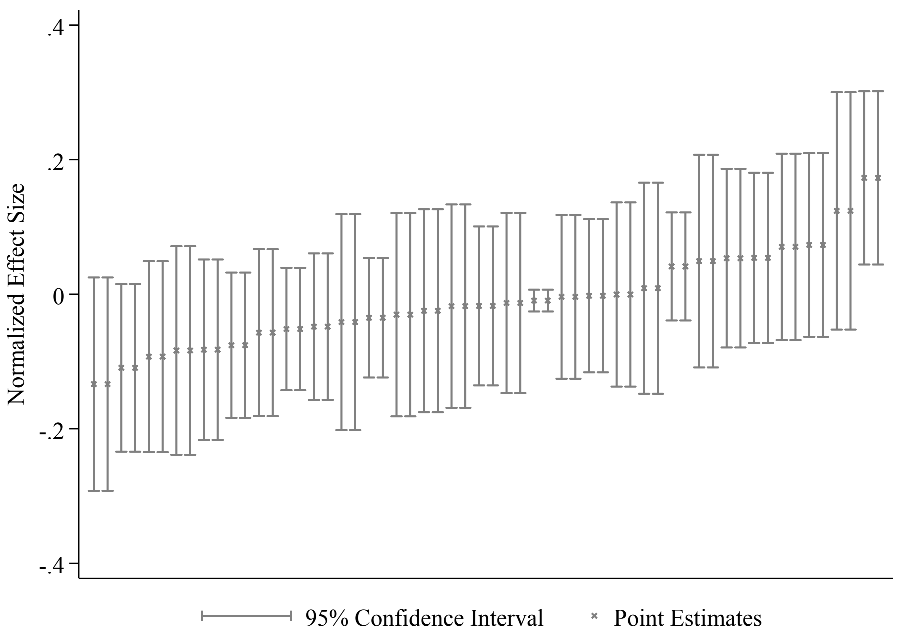

Figure A.1: Balance: Coefficient on Female Relative Dummy for a Broad Set of Municipal Characteristics (All variables listed on Table A.1)

Table A.1: Descriptive Statistics (Political Variables)

Year	Obs.	Term Limit		Female		Dyn. Run		Dyn. Win		Dyn. Fem. Run		Dyn. Fem Win		Inc. Rel. Win		Inc. Rel Fem. Run		Inc. Rel Fem. Win	
		Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev	Mean	Std. Dev
1992	1,491	0.00	0.00	0.09	0.29	0.13	0.33	0.04	0.20	0.02	0.14	0.01	0.10	0.04	0.20	0.02	0.14	0.01	0.10
1995	1,491	0.00	0.00	0.10	0.30	0.17	0.38	0.08	0.27	0.03	0.18	0.01	0.12	0.04	0.20	0.02	0.13	0.01	0.09
1998	1,491	0.42	0.49	0.16	0.37	0.41	0.49	0.24	0.43	0.14	0.35	0.08	0.28	0.17	0.38	0.13	0.33	0.07	0.26
2001	1,491	0.20	0.40	0.17	0.37	0.47	0.50	0.31	0.46	0.15	0.36	0.09	0.29	0.14	0.35	0.08	0.27	0.05	0.22
2004	1,491	0.13	0.34	0.17	0.37	0.52	0.50	0.34	0.47	0.15	0.36	0.10	0.30	0.10	0.29	0.06	0.23	0.03	0.18
2007	1,491	0.29	0.46	0.19	0.39	0.60	0.49	0.41	0.49	0.19	0.39	0.11	0.32	0.16	0.37	0.10	0.30	0.06	0.24
2010	1,491	0.25	0.43	0.21	0.41	0.64	0.48	0.45	0.50	0.20	0.40	0.13	0.34	0.16	0.36	0.09	0.29	0.06	0.23

Table A.2: Descriptive Statistics (1990 Census)

	N	Mean	Std. Dev.	Min	Max
	(1)	(2)	(3)	(4)	(5)
Share with relatives overseas	1,426	0.010	0.010	0.000	0.060
Share who migrated recently	1,426	0.870	0.270	0.000	1.000
Muslim	1,426	0.050	0.190	0.000	1.000
Share urban	1,426	0.300	0.250	0.000	1.000
Gender ratio	1,426	0.510	0.010	0.410	0.550
Population	1,426	37,931	77,877	632	1,660,000
Share own house	1,426	0.890	0.080	0.440	1.000
Share own housing lot	1,426	0.570	0.190	0.000	1.000
Share house with iron roof	1,426	0.410	0.250	0.000	1.000
Share own radio	1,426	0.610	0.110	0.140	0.860
Share own TV	1,426	0.170	0.190	0.000	0.830
Share own Fridge	1,426	0.110	0.100	0.000	0.630
Share own phone	1,426	0.010	0.020	0.000	0.280
Share own car	1,426	0.050	0.040	0.000	0.270
Share own residential land	1,426	0.170	0.110	0.000	0.840
Share own agricultural land	1,426	0.350	0.190	0.000	0.970
Share own other land	1,426	0.100	0.120	0.000	1.750
Share some primary schooling (Male)	1,426	0.530	0.120	0.100	0.830
Share primary graduate (Male)	1,426	0.250	0.070	0.060	0.510
Share some secondary schooling (Male)	1,426	0.090	0.040	0.010	0.380
Share Secondary graduate (Male)	1,426	0.020	0.020	0.000	0.130
Share some college (Male)	1,426	0.050	0.040	0.000	0.270
Share some primary schooling (Female)	1,426	0.520	0.110	0.070	0.820
Share primary graduate (Female)	1,426	0.220	0.060	0.040	0.480
Share some secondary schooling (Female)	1,426	0.070	0.030	0.010	0.310
Share Secondary graduate (Female)	1,426	0.020	0.010	0.000	0.090
Share some college (Female)	1,426	0.080	0.040	0.000	0.300

Notes: Descriptive Statistics from the 1990 census.

Table A.3: Predicting Timing of Term Limits (Full sample)

	1998	2001	2004	2007	2010
	(1)	(2)	(3)	(4)	(5)
Share with relatives overseas	0.00	-0.00	0.00	-0.02	0.02
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Share who migrated recently	0.19	0.00	-0.08	-0.10	-0.10
	(0.15)	(0.12)	(0.10)	(0.09)	(0.07)
Share Muslim	-0.05	-0.02	0.00	0.02	-0.01
	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)
Share urban	-0.01	-0.00	-0.01	0.02	0.01
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Gender Ratio	-0.00	0.00	0.02	-0.01	0.02
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Population	0.00	0.01	-0.01	0.00	-0.01
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Share own house	0.04**	-0.02	0.02	-0.03***	0.00
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Share own housing lot	-0.00	-0.00	0.00	-0.01	-0.01
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share house with iron roof	-0.04	-0.02	-0.02	0.04*	0.00
	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
Share own radio	0.06**	0.02	-0.01	-0.02*	-0.02
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Share own TV	0.02	-0.06	0.06	-0.06	0.02
	(0.07)	(0.06)	(0.05)	(0.04)	(0.03)
Share own Fridge	-0.03	0.13**	-0.05	0.03	-0.03
	(0.07)	(0.05)	(0.05)	(0.04)	(0.03)
Share own phone	-0.01	-0.03	0.03*	-0.02	0.02
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own car	-0.00	0.00	-0.00	-0.01	-0.01
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own residential land	0.02	-0.05**	0.01	0.01	-0.01
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Share own agricultural land	-0.02	0.01	0.02	0.00	0.02
	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)
Share own other land	-0.03	0.04**	-0.01	0.00	0.00
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Share some primary schooling (Male)	-0.03	0.17*	-0.12	0.01	-0.01
	(0.11)	(0.09)	(0.08)	(0.07)	(0.05)
Share primary graduate (Male)	-0.06	0.11*	-0.03	-0.03	0.02
	(0.08)	(0.07)	(0.06)	(0.05)	(0.04)
Share some secondary schooling (Male)	0.02	0.13**	-0.06	-0.06	-0.03
	(0.07)	(0.05)	(0.05)	(0.04)	(0.03)
Share Secondary graduate (Male)	-0.01	-0.01	0.02	0.00	0.02
	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)
Share some college (Male)	-0.00	-0.00	-0.05	0.04	0.02
	(0.08)	(0.06)	(0.05)	(0.04)	(0.04)
Share some primary schooling (Female)	-0.04	-0.15**	0.10	0.04	0.03
	(0.09)	(0.07)	(0.06)	(0.05)	(0.04)
Share primary graduate (Female)	0.01	-0.10*	0.03	0.04	-0.01
	(0.07)	(0.05)	(0.05)	(0.04)	(0.03)
Share some secondary schooling (Female)	-0.03	-0.10**	0.05	0.05	0.05*
	(0.06)	(0.04)	(0.04)	(0.03)	(0.03)
Share Secondary graduate (Female)	0.01	-0.01	-0.03	-0.01	-0.01
	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)
Share some college (Female)	0.01	-0.03	0.06	-0.02	-0.02
	(0.07)	(0.05)	(0.05)	(0.04)	(0.03)
Observations	1,426	1,426	1,426	1,426	1,426
R-squared	0.02	0.02	0.02	0.03	0.02

Notes: Results from municipality regressions. The dependent variable is a dummy equal to one if the incumbent mayor was term-limited in 1998 (Column 1), 2001 (Column 2), 2004 (Column 3), 2007 (Column 4) or 2010 (Column 5). Standard errors (in parentheses). * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Table A.4: Predicting Timing of Term Limits (Term-limited sample)

	1998	2001	2004	2007	2010
	(1)	(2)	(3)	(4)	(5)
Share with relatives overseas	0.01	-0.00	-0.00	-0.03*	0.02
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share who migrated recently	0.28*	0.03	-0.09	-0.10	-0.11
	(0.16)	(0.13)	(0.11)	(0.09)	(0.08)
Share Muslim	-0.02	-0.01	0.01	0.03	-0.01
	(0.05)	(0.04)	(0.03)	(0.03)	(0.02)
Share urban	-0.01	-0.00	-0.01	0.02	0.01
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Gender ratio	-0.02	0.01	0.01	-0.02	0.02
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Population	0.00	0.01	-0.01	0.00	-0.01
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
Share own house	0.05**	-0.02	0.02	-0.04***	0.00
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Share own housing lot	0.00	0.00	0.01	-0.01	-0.01
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share house with iron roof	-0.03	-0.01	-0.02	0.04**	0.01
	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)
Share own radio	0.06**	0.02	-0.02	-0.03**	-0.02*
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own TV	0.03	-0.07	0.06	-0.05	0.02
	(0.08)	(0.06)	(0.05)	(0.05)	(0.04)
Share own Fridge	-0.07	0.14**	-0.06	0.02	-0.04
	(0.07)	(0.06)	(0.05)	(0.04)	(0.04)
Share own phone	-0.01	-0.03	0.04**	-0.02	0.02
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own car	0.01	0.00	0.00	-0.01	-0.01
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own residential land	0.03	-0.05**	0.01	0.02	-0.01
	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)
Share own agricultural land	-0.03	-0.00	0.01	0.00	0.02
	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
Share own other land	-0.04*	0.04**	-0.01	0.00	0.00
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Share some primary schooling (Male)	-0.05	0.23**	-0.16*	-0.01	-0.01
	(0.13)	(0.11)	(0.09)	(0.08)	(0.07)
Share primary graduate (Male)	-0.08	0.14*	-0.04	-0.05	0.03
	(0.09)	(0.08)	(0.07)	(0.06)	(0.05)
Share some secondary schooling (Male)	0.03	0.16**	-0.08	-0.08*	-0.03
	(0.08)	(0.06)	(0.05)	(0.05)	(0.04)
Share Secondary graduate (Male)	-0.03	-0.01	0.01	-0.00	0.03
	(0.04)	(0.04)	(0.03)	(0.03)	(0.02)
Share some college (Male)	-0.01	-0.00	-0.06	0.04	0.03
	(0.09)	(0.07)	(0.06)	(0.05)	(0.04)
Share some primary schooling (Female)	-0.02	-0.20**	0.12	0.06	0.04
	(0.11)	(0.09)	(0.07)	(0.06)	(0.05)
Share primary graduate (Female)	0.03	-0.12*	0.04	0.06	-0.01
	(0.07)	(0.06)	(0.05)	(0.04)	(0.04)
Share some secondary schooling (Female)	-0.05	-0.12**	0.06	0.06*	0.05*
	(0.06)	(0.05)	(0.04)	(0.04)	(0.03)
Share Secondary graduate (Female)	0.03	-0.00	-0.01	-0.00	-0.01
	(0.04)	(0.04)	(0.03)	(0.03)	(0.02)
Share some college (Female)	0.01	-0.04	0.07	-0.02	-0.03
	(0.07)	(0.06)	(0.05)	(0.04)	(0.04)
Observations	1,283	1,283	1,283	1,283	1,283
R-squared	0.02	0.03	0.02	0.04	0.02

Notes: Results from municipality regressions. The dependent variable is a dummy equal to one if the incumbent mayor was term-limited in 1998 (Column 1), 2001 (Column 2), 2004 (Column 3), 2007 (Column 4) or 2010 (Column 5). Standard errors (in parentheses). * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Table A.5: Term Limits and the Rise of Female Politicians - Interaction of Election Year Dummies with Baseline Municipal Characteristics

	Female Candidate	Female Mayor	Female Incumbent Relative
	(1)	(2)	(3)
Term Limit	0.17***	0.11***	0.17***
	(0.01)	(0.01)	(0.01)
Observations	11,358	11,303	11,358
R-squared	0.09	0.05	0.13

Notes: Results from municipality*elections regressions and include municipal and election year fixed effects. The dependent variable is a dummy for whether a woman ran for mayor (Column 1), a dummy for whether a woman was elected mayor (Column 2), or a dummy for whether a woman related to the incumbent won (Column 3). The regressions include a comprehensive set of municipal-level socio-economic characteristics measured in 1990: population, gender ratio, urbanisation, migration, asset ownership (land, house, ration, tv, phone, car and fridge) and education levels broken down by gender. Those variables are all interacted with a full set of election year dummies. The standard errors (in parentheses) account for potential correlation within municipalities. * denotes significance at the 10 percent, ** 5 percent and *** 1 percent levels.

Table A.6: Comparing Male and Female Preferences (Ilocos)

	Full Sample			Some College		
	Female	Male	OLS	Female	Male	OLS
	(1)	(2)	(3)	(4)	(5)	(6)
Health	19.34	18.31	1.00	19.83	17.80	2.02
	(13.49)	(12.42)	[0.02]	(13.04)	(10.98)	[0.01]
Education	18.15	17.29	0.91	19.12	19.62	-0.35
	(13.23)	(12.29)	[0.04]	(11.97)	(12.76)	[0.66]
Emergencies	9.28	7.01	2.29	8.55	6.25	2.32
	(10.73)	(8.39)	[0.00]	(9.25)	(7.36)	[0.00]
Water and Sanitation	8.16	8.17	0.01	8.22	8.14	0.12
	(7.91)	(8.81)	[0.96]	(7.62)	(8.82)	[0.82]
Roads	6.18	7.67	-1.49	6.92	8.03	-1.17
	(7.51)	(9.03)	[0.00]	(8.79)	(7.54)	[0.03]
Community Facilities	4.56	5.00	-0.40	4.93	5.65	-0.69
	(5.81)	(6.45)	[0.06]	(5.80)	(7.26)	[0.11]
Business Loans	6.17	5.28	0.91	5.78	4.64	1.16
	(10.39)	(9.02)	[0.01]	(9.34)	(6.85)	[0.03]
Agriculture Assistance	20.11	23.30	-3.28	17.71	21.33	-3.74
	(17.07)	(19.04)	[0.00]	(14.56)	(18.59)	[0.00]
Peace and Security	5.36	5.31	0.02	6.26	6.01	0.17
	(6.24)	(6.77)	[0.92]	(6.09)	(7.26)	[0.69]
Festivals	2.69	2.67	0.02	2.68	2.54	0.15
	(4.51)	(5.09)	[0.90]	(4.06)	(4.21)	[0.57]

Notes: n = 3,404 (Columns 1-3), n = 936 (Columns 4-6). Each variable is the share of the Local Development Fund that the respondent would like the municipal government to spend on that sector. The standard deviations are in parentheses (Columns 1-2 and 4-5). Each cell in Columns 3 and 6 is either the coefficient on the dummy variable indicating whether the respondent is a female from a different OLS regression with municipal fixed effects or the associated p-value in brackets.

Table A.7: Term-Limited Incumbents with Relatives Winning in the Subsequent forced-open-Seat Race and Relative's Gender

	Relative's Gender			
	Female	Male	p-value	N
Winning margin of incumbent's elected relative	0.27	0.27	0.71	731
Number of candidates in race	2.99	2.96	0.81	733
Incumbent's margin of victory in 3rd (last) race	0.46	0.46	0.72	733
Incumbent's margin of victory (average 3 races)	0.32	0.35	0.06	733