Is There Gender Bias Among Voters ? Evidence from the Chilean Congressional Elections

Francisco J. Pino
SBS-EM, ECARES, Université Libre de Bruxelles

December 2014

ECARES working paper 2014-53

ECARES
ULB - CP 114/04
50, F.D. Roosevelt Ave., B-1050 Brussels BELGIUM
www.ecares.org

Is There Gender Bias Among Voters? Evidence from the Chilean Congressional Elections

Francisco J. Pino*

December 2, 2014

Abstract

I exploit the unique institution of gender-segregated voting booths in Chile, allowing the use of actual voting data, instead of self-reported surveys, to test for gender bias among voters. Overall I find evidence of a small but significant negative gender bias: women overall are less likely than men to vote for female candidates. The effect is mainly driven by center-right voters. Selection, candidates' quality and districts' characteristics do not explain away the results. This evidence does not question whether female leaders have an effect on economic outcomes, but rather the mechanism through which this effect takes place.

JEL codes: D72, J16, O54.

Keywords: gender, voting, electoral system, discrimination, political economy.

---

*ECARES, Université Libre de Bruxelles, Avenue Roosevelt 42, 1050 Bruxelles Belgium, fpino@ulb.ac.be. I am indebted to my dissertation committee members Dilip Mookherjee, Daniele Paserman, Laurent Bouton and Robert Margo for their continuous support and advice. I also thank Manuel Bagues, Micael Castanheira, Paola Conconi, Stefano DellaVigna, Jordi Jaumandreu, Stephan Litschig, Ana Nuevo-Chiquero, Claudia Olivetti, Maurizio Zanardi and participants of seminars at Boston University, ECARES, Universitat de Barcelona, Universitat Pompeu Fabra, IAE-CSIC, CEA-University of Chile and FUNDP Namur for their helpful comments. Maria Bernarda Dall'Aglio provided excellent research assistance. All remaining errors are mine.

1 Introduction

"He's the girl in the race. Clinton came out tough; she voted for the war. Obama came out as the person bringing people together and offering messages of hope and reconciliation."

—Marie Wilson, head of the White House Project, a nonprofit that helps women move into positions of leadership. Quoted in *Why Didn't More Women Vote for Hillary?*, Amy Sullivan, TIME Magazine, June 8, 2008.

It is a well known fact that women are under-represented in leadership positions, particularly in the political arena.[1] It has been suggested that an increase in female representation could shift policies to better reflect women's preferences (Chattopadhyay and Duflo, 2004). However, the empirical evidence for this is mixed, particularly in contexts where there are no seats reserved for minorities.[2] Understanding the mechanism that links women's representation to policy outcomes becomes thus crucial.

There are two main theoretical approaches that have been used to analyze the effect on economic outcomes of having a female leader. On the one hand, the median voter theory (Downs, 1957), which assumes that candidates can commit to a specific policy platform, predicts policy convergence, with both parties choosing their policy platform to coincide with the median voter's bliss point. Thus, in Downs' model the gender of the politician is irrelevant for policy decisions. On the other hand, the citizen-candidate model (Osborne and Slivinski, 1996; Besley and Coate, 1997) assumes that candidates have a preferred

[1] According to data collected by the Inter-Parliamentary Union, as of October 2014 the average percentage of women in a parliament or congress was 22.2% (considering only the lower or single house).

[2] Chattopadhyay and Duflo (2004) look at the effect of having a local female leader elected in a reserved seat in the states of West Bengal and Rajastan, India. They find that women leaders choose policies more closely related to women's needs. In contrast, Bardhan, Mookherjee and Torrado (2010) find that reserved seats for women deteriorated the targeting of public goods to disadvantaged groups in West Bengal. Analyzing Indian state governments, Clots-Figueras (2011, 2012) finds a positive effect of female legislators on health and education investment, but the effect is particularly strong when women are elected in seats reserved for lower castes and disadvantaged tribes, for the case of health outcomes, and in urban areas for education. Regarding evidence from settings without reservation policies, Rehavi (2007) finds positive effects of female representation on health expenditures in the context of U.S. state legislatures, while Gagliarducci and Paserman (2012) and Ferreira and Gyourko (2014) do not find these effects in the context of Italian and U.S. municipalities, respectively.

policy platform and cannot commit to implementing a policy different from their preferred one once in office. When the following two additional assumptions are made, the citizen-candidate model predicts policy divergence, with the male and female candidates choosing policies equidistant from the median voter's bliss point: (1) women and men should have sufficiently different preferences over policies, and (2) the candidate's gender needs to be a good signal of a candidate's policy preferences, i.e. each candidate should receive more votes from her/his own group than from the opposite group. There is evidence corroborating the first assumption, namely that women and men have different preferences.[3] However, evidence of the second assumption is more elusive since, due to the introduction of the secret ballot, data on individual voting are impossible to obtain.

In this paper I focus on this second assumption, critically examining the traditional view that voters generally prefer candidates of their same gender. Using data from a unique setup—the Chilean congressional elections in 1989–2009, where men and women vote in separate voting booths—I am able to overcome the issues that have plagued previous studies, which relied on the use of surveys or exit polls to analyze the support for female candidates (Paolino, 1995; Dolan, 1998 and 2008).[4] Survey data can be misleading if there are differential response biases by gender.[5] Gender-segregated voting allows me to use actual voting data and avoid the shortcomings associated with surveys and exit polls.

I find a negative but small gender bias among voters, i.e. women voters are slightly less

[3] Chattopadhyay and Duflo (2004) show that women care about drinking water and roads more than men do. Other studies have shown that women have different preferences from men over child welfare policies (Thomas, 1990 and 1994; Duflo, 2003; Miller, 2008), as well as over environmental protection and defense spending (Funk and Gathmann, 2010).

[4] In his 1955's seminal work, Duverger argues that the reason why he finds no difference between the votes of men and women is because of a "tendency for husband and wife to vote in the same way". Dolan (2008) provides a comprehensive summary of the political science literature analyzing this phenomenon. She affirms that a variety of results shows that "the relationship between women voters and female candidates is often conditioned by forces beyond a shared sex identity".

[5] Surveys can present additional problems. Epstein (2006) argues that there is a systematic upward bias in turnout in surveys such as the National Election Studies (NES). There is also evidence that the gender of the interviewer can affect the responses differentially depending on the respondent's gender (Kane and Macaulay, 1993; Huddy, Billig, Bracciodieta, Hoeffler, Moynihan and Pugliani, 1997; Flores-Macias and Lawson, 2008; Benstead, 2014). Responses can be also affected by the interviewer's religiosity (Blaydes and Gillum, 2013) or the language of the interview (Lee and Pérez, 2014).

likely to vote for female candidates than men voters are.[6] This effect decomposes into a positive gender bias among center-left voters and a negative gender bias among center-right voters. These results are not explained away by the inclusion of socio-demographic controls at the municipality level or by controls for candidates' political experience. Additional evidence restricting the sample to close within-coalition elections rules out the possibility that these results are driven by candidates' unobservable characteristics.

To further examine these results and confront them with the existing literature, I analyze roll-call voting data from the Chilean Chamber of Deputies. Implementing a Regression Discontinuity (RD) framework, I find that female legislators in the center-right coalition vote in opposition to the majority of the coalition more often than their male counterparts do. More striking, this effect comes mainly from female legislators voting differently on so-called women's issues (family, health, labor, education and justice). Even though I am not able to identify pro-female legislation due to the lack of voting scores for the Chamber of Deputies, the evidence suggests that center-right female legislators deviate from their coalition by adopting a stronger pro-women stance on women's issues. Finally, using data from the *Latinobarómetro* Survey, I show that center-right women are not more right-wing than their male counterparts, particularly regarding their view of the role of women.

These results challenge the standard views of political competition, since they do not conform to the predictions of either *Downsian* models or citizen-candidate models. I discuss alternative models which could give rise to these results. Gender identity, defined as the existence of social norms about gender roles which could induce a distaste of women voters for female candidates (Akerlof and Kranton, 2000), does not find strong support in my data. Other policy-based explanations, such as female candidates "acting tough" to appeal to men voters but in turn losing the support of women voters as in the opening quotation above (Herrnson, Lay and Stokes, 2003) have less support in the data, as well as more sophisticated versions of the Downsian model ("neo-Downsians") where individuals may

[6]This definition of gender bias is different from the political gender gap, which compares the relative support of men and women for the left. The political science literature has used the terms "gender gap" and "gender affinity" to describe a preference of voters for same-sex candidates.

care not only about the policy platform's position, but also about its quality (Carrillo and Castanheira, 2008) or the quality of the candidates themselves (Kartik and McAfee, 2007). Taken together, these results provide evidence against a policy-based interpretation of the negative gender bias among center-right women voters.

By showing the existence of a small but statistically significant distaste of women voters for center-right female candidates in congressional elections, I add a new dimension to the discussion of gender quotas and reserved seats: women might not always prefer a female leader over a male one. This result is consistent with Beaman, Chattopadhyay, Duflo, Pande and Topalova (2009), who find that being exposed to a female leader improves male voters' evaluations of these leaders, but do not find an improvement in female voters' evaluations of them, even though female leaders invest more in public goods preferred by women. The results are also compatible with previous research showing that women do not always perform better when evaluated by a committee with a larger share of women (Bagues and Esteve-Volart, 2010; Bagues, Sylos-Labini and Zinovyeva, 2014).

The results also give an alternative interpretation to previous research on U.S. elections, which show that women feel more positively than men towards female Democratic candidates, but are indifferent towards female Republican candidates (Dolan, 2008). Dolan's conjecture is that women may experience "cross pressures", with gender considerations bringing them closer to female candidates but the candidate's party (Republican) pulling them away. I suggest that these cross pressures could have the opposite sign, with women supporters of the Republican party being pulled away from female candidates because of gender considerations. Finally, I provide a cautionary tale for the use of surveys when analyzing gender differences in voting behavior.

The remainder of the paper is organized as follows. Section 2 provides the context for female enfranchisement and gender-segregated voting and describes the Chilean electoral system. Section 3 describes the sources of my data, while section 4 presents the econometric framework. Section 5 presents the main results of the paper, while section 6 analyzes possible alternative mechanisms using roll-call voting data and the *Latinobarómetro* Survey.

Finally, section 7 states the conclusion.

2 The Chilean congressional electoral system

This section describes the several unique features of the Chilean Electoral System, which make this dataset unusually valuable. The first is that women and men vote in separate voting booths, which makes it possible to analyze voting data by gender. Secondly the structure of the two-member congressional districts and the special rules that determine the winners in each district mean that most political competition occurs within coalitions rather than across coalitions, which allows me to analyze gender bias for each coalition independently.

2.1 Women and the vote

Although the women's suffrage movement started in Chile as early as the 1870s, women were finally allowed to vote for the first time in the 1935 municipal elections, though not in either the presidential or congressional elections. As a result of this differentiation, two separate registries were created: The General Male Register, for men older than 21; and the Municipal Register, for women older than 21.[7] Both groups had to vote in different ballot booths and their votes were counted separately (Carrera and Ulloa, 2006).

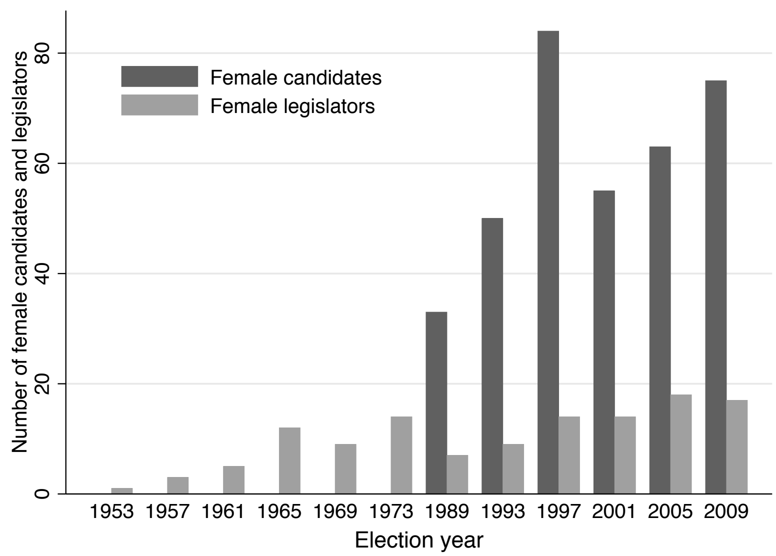
Eighteen years later in 1949, when women were allowed to vote in presidential, congressional, as well as municipal elections, the separate registers for men and women continued to be maintained. Lewis (2004, p. 720) argues that segregated polling was kept "in order to allow women more freedom to vote according to their preferences".

By the 1969 election women constituted almost half of the electorate. Figure 1 shows that the number of congresswomen increased gradually, from 1 out of 147 legislators in 1953, to 14 out of 150 in 1973. When the new binominal system was put in place in 1989, the authorities decided to still keep the gender-segregated registers, as well as the

[7] The Municipal Register also included (men and women) foreigners.

gender-segregated polling stations. Since 1989 both the number of candidates and of elected congresswomen have shown an upward trend, though the number of elected women now seems to be stalled at a sixth of the seats (20 out of 120).

Figure 1: Number of female candidates and elected members, lower house.



Notes: Figure shows the number of female candidates and elected members for the lower house. There were no congressional elections between 1973 and 1989. The total number of seats was 147 between 1953 and 1965, 150 between 1969 and 1973, and 120 between 1989 and 2009. Data on the number of female candidates is available only after 1989.

2.2 The binominal system

The Chilean National Congress consists of two chambers: The *Senado* (Senate or Upper Chamber) and the *Cámara de Diputados* (Chamber of Deputies or Lower Chamber). The former has 36 members that represent 18 two-member Senate districts while the latter has 120 members representing 60 two-member congressional districts. Candidates running for these offices are presented by coalitions, which are nationwide conglomerates of parties running on a common policy platform. A coalition cannot present more than 2 candidates per district.[8] Each list is open, so voters can cast their vote directly for their preferred

[8]Independent candidates can either join a coalition (in which case their party name is displayed as "Independent") or they can run as completely independent. Figure A-1 in online Appendix A shows a

candidate (they can only vote for 1 candidate).

The two winners in each district are determined by the D'Hondt method, which stimulates most of the political competition to occur *within coalitions* rather than *across coalitions*. In this method, the first seat always goes to the coalition with the largest share of votes (and within this coalition, to the candidate receiving the most votes). Typically, the second seat goes to the coalition with the next highest number of votes, and within this coalition the seat goes to the candidate with more votes. The only exception is in cases where the coalition with the most votes receives more than twice as many votes as any other coalition; in this situation, the coalition with the most votes receives both seats.[9] Since this case is unusual and difficult to obtain (roughly 12% of elections end up in a "doubling"), candidates in the two largest coalitions are forced to compete against their coalition "partner" instead of competing against candidates of other coalitions.[10] This feature allows me to analyze gender bias for each coalition independently, abstracting from ideology considerations.[11] Therefore in some specifications the sample will be restricted to one of the two main coalitions: Center-left (*Concertación*) and Center-right (*Alianza*). These are the two largest coalitions and its party members have remained unchanged in most of the elections since 1989.[12]

An additional feature of the Chilean system is that during the period of analysis voting was mandatory for registered voters, which makes selective turnout less of a concern.[13]

Previous research using data from the Chilean elections has focused mainly on the

ballot with an example of the former, in which candidate No. 21 is running as an independent within coalition B.

[9] Table A-1 in online Appendix A shows a set of examples to illustrate all possible election outcomes.

[10] Siavelis (2002) argues that party leaders seek to avoid intralist confrontation by pairing candidates who will not engage in it. However, accounts of disputes between candidates of the same coalition, including fights among supporters and the destruction of the other candidate's posters, abound in the media.

[11] Alemán and Saiegh (2007) analyze the voting behavior of Chilean legislators in years 1997–2000, and find that the median policy position of legislators in a party is indistinguishable from the median of the coalition to which the party belongs.

[12] These two coalitions always present candidates in all districts, as opposed to smaller coalitions.

[13] Cerda and Vergara (2009) analyze voters' turnout in Chile using both aggregate and individual data, and conclude that the observed decline in turnout is mainly due to low participation of the youth. This in turn is due to under-registration of this group (registration is voluntary, though voting is mandatory once one has registered), and not due to a lower level of participation once registered to vote. A new change to the electoral law enacted in 2012 established voluntary voting and automatic registration.

innate bias of the electoral system towards the second-largest coalition. To the best of my knowledge, there are only two papers that take advantage of the segregated voting system in Chile. Lewis (2004) uses aggregate data from the Chilean presidential elections in 1952-1999 to analyze the political gender gap (i.e. the relative support of men and women for the left); it documented women's bias towards conservatism. Carrera and Ulloa (2006) use data from the Chilean municipal elections in 1992-2004 to show that this bias decreased in more recent elections.

3 Data

This unique dataset comes from various sources. In this section I describe the data gathering process, and present summary statistics of the variables used in the analysis.

3.1 Votes by gender

Voting data for the Chamber of Deputies at the ballot booth level was obtained from the *Tribunal Calificador de Elecciones* (Election Qualifying Court). Since controls are available at the municipality level (see below), I aggregate voting data coming from ballot booths of the same gender, within a municipality.[14] The data contain the total number of votes per candidate in each of the 6 congressional elections held every 4 years since 1989. They also include an identifier for whether the ballot booth is a male or female one.

3.2 Candidates' characteristics

The determination of candidate gender was done manually. Fortunately, names in Spanish are easy to classify across genders. Ambiguous cases were looked for in the website of the *Servicio Electoral* (Electoral Service). Panel A of Table 1 shows the summary statistics for the Chamber of Deputies' elections. On average a district has 6.78 candidates, slightly increasing from 6.98 in 1989 to 7.15 in 2009. The average number of 1 female candidate

[14] A district can contain between 1 and 15 municipalities, with an average of 6.

per district masks significant variation across years, with 0.55 female candidates in 1989 to 1.22 in 2009. Regarding the two largest coalitions (Center-left and Center-right) they have on average one female candidate in every five districts, increasing from less than one in every 7 districts to more than one in every 4 districts. The average voteshare for women candidates is 12% on average, but it increases to about 50% when they run in one of the two largest coalitions, showing that female candidates are competitive when running for either the Center-left or Center-right.[15]

Table 1: Summary statistics: districts and candidates.

Election year	1989	1993	1997	2001	2005	2009	Average
Panel A. Districts							
No. candidates	6.98	6.40	7.37	6.35	6.43	7.15	6.78
No. female candidates	0.55	0.83	1.40	0.92	1.05	1.22	0.99
From Center-left	0.12	0.18	0.25	0.23	0.33	0.28	0.23
From Center-right	0.15	0.13	0.20	0.23	0.25	0.27	0.21
Voteshare female candidates	0.12	0.12	0.08	0.14	0.16	0.14	0.12
Within Center-left	0.49	0.54	0.49	0.58	0.53	0.48	0.52
Within Center-right	0.44	0.47	0.37	0.53	0.49	0.53	0.48
Panel B. Candidates							
Age (males)	45.99	45.99	45.63	47.98	48.62	48.08	47.01
Age (females)	46.42	44.18	42.01	45.24	46.56	46.70	44.97
Incumbent (males)	0.00	0.25	0.22	0.25	0.25	0.21	0.19
Incumbent (females)	0.00	0.08	0.06	0.20	0.19	0.21	0.13

Notes: Panel A: Sample size is 360 (6 election years and 60 districts). All numbers are district averages. Panel B: Sample size is 2,441. Age is in years, and incumbent= 1 if the candidate was elected in the same district he/she is currently running.

Few information is available for both elected and non-elected candidates. The candidates' age was obtained from the *Servicio Electoral*. In addition, I construct the dummy variable *incumbent*, which takes the value of 1 if the candidate was currently in office and running for a subsequent term in the same district. Panel B of Table 1 shows statistics for these controls. On average male candidates are 2 years older than female candidates. 19% of male candidates are incumbent, compared to 13% of female candidates.

[15]The only exception is year 1997, in which female candidates of the Center-right received on average 37% of the vote in the coalition.

3.3 Demographic Data

I construct socio-demographic controls at the municipality level, which is a finer level of detail than electoral district level. Average age, education, share of urban and indigenous population, income, labor force participation (LFP), LFP gap (defined as the difference between male LFP and female LFP), share of women in the municipality and share of married population are constructed using the *Encuesta CASEN* (Survey of National Socio-economic Characterization), a nationally representative survey (see the online appendix B for details and summary statistics).

4 Empirical framework

4.1 Specification

The goal is to analyze whether there is gender bias among voters, and in particular, whether women vote more often for female candidates than do men. Let SV_{ibmt} be the share of votes to candidate i in b-type ballots ($b \in \{female, male\}$) in municipality m and in election t (I drop the sub-index t hereafter for simplicity). This vote-share is computed at the gender-municipality level as follows:

$$SV_{ibm} = \frac{V_{ibm}}{\sum_i V_{ibm}} \tag{1}$$

where V_{ibm} is the number of votes that candidate i gets. Now define ΔSV_{im} as the difference in the vote-share between female voters and male voters:

$$\Delta SV_{im} = SV_{i,b=female,m} - SV_{i,b=male,m} \tag{2}$$

I therefore consider the following specification:

$$\Delta SV^F_{im} = \beta + X_m\,\gamma + Z_i\,\lambda + \eta_t + \mu_d + \epsilon_{im} \tag{3}$$

where X_{mt} are municipality controls and Z_{it} are candidate controls (age and incumbency). η_t and μ_d are election year and district dummies, respectively. The supra-index F in the dependent variable indicates that the model is estimated for the subsample of female candidates. In addition, all controls are demeaned, and therefore the estimate for the constant term $\hat{\beta}$ gives the average gender bias between female and male voters, which does not change once controls are included. This is useful since it allows me to directly compare the average gender bias across the different samples analyzed in the next section.[16]

Equation (3) is estimated including candidates from all coalitions. I redefine the dependent variable to estimate the model for each coalition separately. Specifically, I restrict the denominator in equation (1) to candidates in the same coalition and re-compute ΔSV^F. I present the results for the center-left (Concertación) and center-right (Alianza) coalitions, since as mentioned before, these are the two largest coalitions which have remained relatively constant over time, as opposed to smaller left and right coalitions which sometimes run together but split afterwards.

4.2 Identification

The main identification strategy considers all congressional elections. Given that the focus of this paper is on the constant term $\hat{\beta}$ of equation (3), identification is straightforward since the constant term does not suffer from omitted variable bias. Nonetheless, women do not run in all districts, and even if they do, the sorting of female candidates into districts is likely to be non-random. In particular, this sorting can be due to unobservable characteristics of the candidate that in turn correlate with her relative support among men and women voters. If this is the case, $\hat{\beta}$ will not be informative of the average relative support of female candidates. In addition, party leaders might be less inclined to choose competitive female candidates if they have a preference for male candidates and the likelihood of getting only

[16] The specification shown in equation (3) gives identical results to regressing the share of votes computed as in equation (1), on a dummy variable taking the value of 1 when the votes come from a female ballot booth (FEM): $SV_{ibm} = \tilde{\beta} + \tilde{\delta}\, FEM_{bm} + FEM_{bm} \times (X_m\, \tilde{\gamma} + Z_i\, \tilde{\lambda} + \eta_t + \mu_d) + \epsilon_{ibm}$. The estimate for $\tilde{\beta}$ is identical to the one for β.

one seat is high.

To address these concerns, as an alternative identification strategy I re-estimate equation (3), but now restricting the sample of female candidates to those who face close elections within a coalition, that is that win or lose by a small margin against her coalition partner.[17] In these elections, candidates who win or lose should exhibit similar characteristics. This argument has been used extensively when implementing Regression Discontinuity (RD) designs, and in particular when comparing mixed-gender races (e.g. Gagliarducci and Paserman, 2012; Ferreira and Gyourko, 2014; Brollo and Troiano, 2014).[18] An additional argument for analyzing the outcome of close elections is that the citizen-candidate model predicts that elections with two candidates should be highly contested, since candidates' positions are symmetric around the median (Osborne and Slivinski, 1996; Besley and Coate, 1997). It is possible that in elections where one of the candidates wins in a "landslide" gender does not convey much information about the candidates' preferences.

To provide further support for this argument, Table C-1 in online appendix C shows that male and female candidates' characteristics (age and incumbency) do not significantly differ when focusing on candidates running in close elections, particularly on elections with a margin of victory below 10 percent.[19] Age and incumbency are the only covariates available for all candidates and all elections. An imbalance on incumbency, however, can be thought as a proxy for imbalances on other characteristics.[20] Given that this procedure significantly reduces the sample size, its outcome should be regarded as a robustness exercise.

[17]As mentioned previously, the binominal system forces competition to occur within coalitions rather than across coalitions. Another type of close election occurs when the largest coalition has nearly double the votes of the second largest coalition (Pino, 2011). Competition among candidates of the same coalition in this case, however, need not be close.

[18]Both Gagliarducci and Paserman (2012) and Brollo and Troiano (2014) note that most of the candidates' observable characteristics converge in close elections, even those that appear unbalanced for the average sample. See Eggers, Fowler, Hainmueller, Hall and Snyder Jr (2014) for a discussion on the validity of this argument.

[19]Notice that standard validity tests employed in the RD literature (Imbens and Lemieux, 2008) are not useful in this case, since rather than comparing outcomes of male and female candidates at either side of the threshold (and therefore only of those elected), the focus here is on comparing outcomes of winning and losing male candidates to their female analogs. The standard RD test are discussed later when I analyze the behavior of legislators.

[20]Eggers *et al.* (2014) show that incumbency can account for most of the imbalance in U.S. House elections found by Caughey and Sekhon (2011).

5 Gender bias for female candidates

5.1 All elections

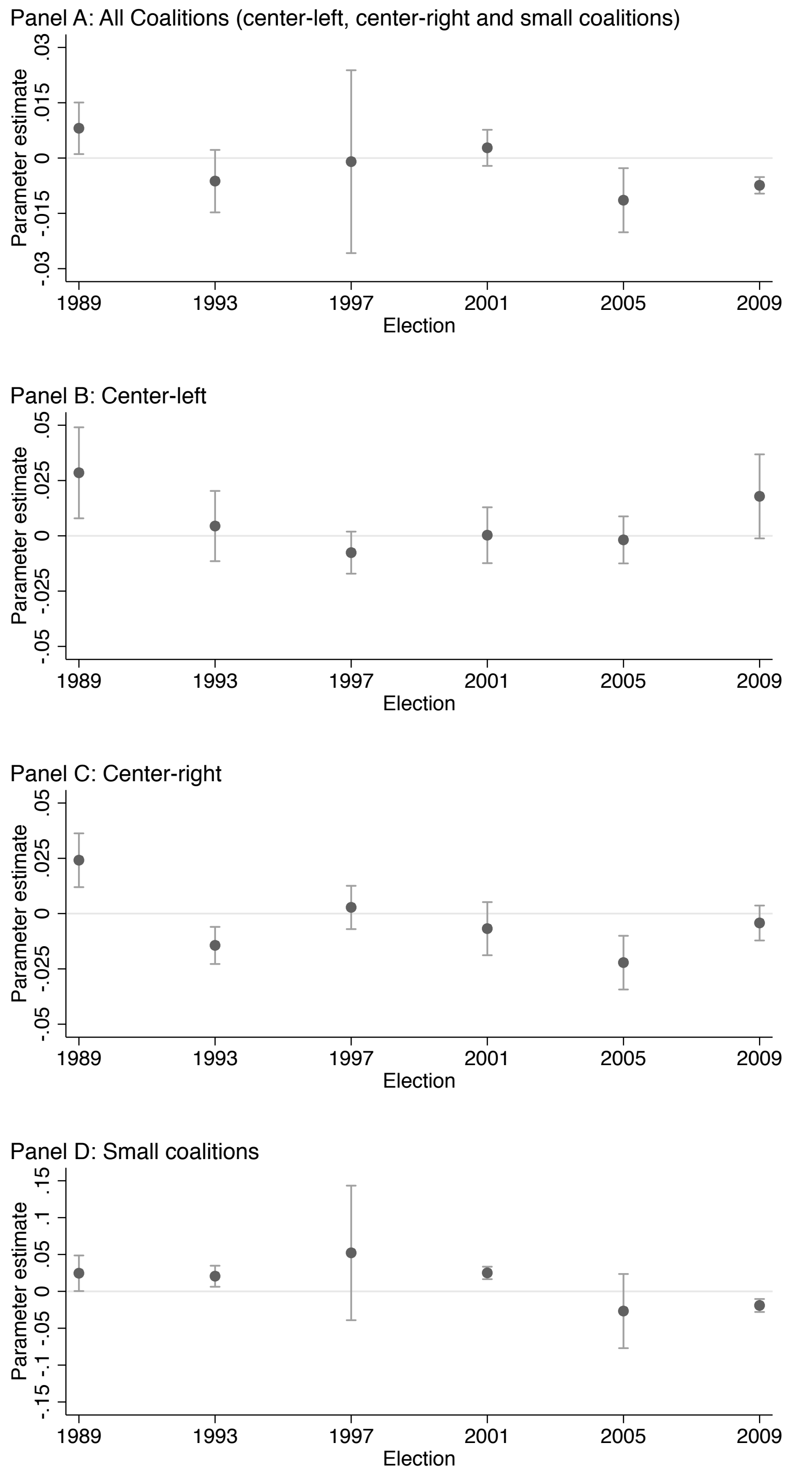
I start with a graphical representation of the results. Figure 2 shows the average gender bias for female candidates for each election in 1989-2009. The figure plots estimates of the coefficients η_t when none of the other regressors in equation (3) are included. A larger support for female candidates from women voters than from men voters corresponds to a positive coefficient.

Panel A shows the estimates when all coalitions are included. Of the six elections considered, only the first one in 1989 displays a positive and statistically significant gender bias. In the next three elections (1993, 1997 and 2001) the bias is not significantly different from zero. Finally, the estimated gender bias for the 2005 and 2009 elections is negative and statistically significant at conventional levels.

As explained before, the binominal system induces most of the electoral competition to occur within coalitions (as opposed to across coalitions). Recall that coalitions can put up at most two candidates per district, and they usually do. Therefore I now estimate equation (3) but redefining the dependent variable so that the share of votes for female candidates ΔSV^F is computed at the gender-coalition-municipality level. Panels B and C show the average gender bias for the main two coalitions: center-left (*Concertación*) and center-right (*Alianza*). These are the two largest coalitions and its party members have remained unchanged in most of the elections since 1989. Both coalitions display a positive gender bias in the 1989 election. Nonetheless, while the bias is not significantly different from zero for the center-left in the following elections, it is negative for the center-right in the 1993 and 2005 elections. Except for the 1989 election —which was the first election after the end of the dictatorship, and the first where the binominal system was in place— there is no evidence of a positive gender bias towards center-right female candidates. For completeness I also show the average gender bias for small coalitions in Panel D.[21] The four

[21]Notice that here ΔSV^F is computed taking into account all female candidates from small coalitions. Therefore, panels B, C and D are not an exact decomposition of Panel A.

Figure 2: Gender bias for female candidates, 1989-2009.



Notes: The figures show the estimated coefficients of regressing ΔSV^F on election dummies. Panel A includes all coalitions, while panels B and C include only the center-left and center-right coalitions, respectively. Error bars are $\pm$ 95% confidence intervals.

first elections feature a positive gender bias (though not statistically significant in 1997), to latter become negative and significant in 2009.

To further analyze the gender bias for female candidates, Table 2 presents regression results for all elections since 1989. The first two columns present the coefficients from estimating equation (3) including all coalitions in the sample. I add district dummies in the second column to exploit changes across municipalities while keeping district-level characteristics —such as the number of candidates on the ballot— constant.[22] The dependent variable is ΔSV^F, i.e. the difference between the vote-share that a female candidate obtains from female and male voters. The average gender bias, measured by the constant term and presented at the top of the table, is -0.3 percent, and it is significant at the 1 percent level. This means that on average female candidates get 0.3 percent less votes in female voting booths than in male voting booths. That is, they get 0.3 percent less votes coming from female voters, compared to male voters.

Columns 3 and 4 show that the gender bias for center-left female candidates is positive and statistically significant when district dummies are included (column 4). On average female candidates on the center-left get 0.5 percent more votes from women than from men voters. On the other hand, columns 5 and 6 show a negative and statistically significant gender bias for center-right female candidates. Female candidates on the center-right get 0.6 percent less votes from women than from men voters.

The table also shows how candidates and municipality characteristics affect gender bias —a positive effect on the gender bias is given by a positive coefficient. Few of the controls have a significant effect on gender bias, particularly when district dummies are included. For instance, women voters are more supportive than men of center-left female candidates in municipalities with higher income, lower average labor force participation (LFP), as well as lower LFP gap (column 3), but these effects become non-significant when exploiting the within-district variation (column 4). From the results in column 4, women voters are less

[22]Since districts are formed by neighboring municipalities, the variance of the controls drops significantly. There is also evidence that districts were gerrymandered to increase the representation of the right (Rahat and Sznajder, 1998).

Table 2: Determinants of gender bias for female candidates

Dep. Variable:	ΔSV^F (Difference in vote-shares to female candidates)					
Sample:	All		Center-left		Center-right	
	(1)	(2)	(3)	(4)	(5)	(6)
Average gender bias	-0.003***	-0.003***	0.005	0.005***	-0.006*	-0.006***
	(0.001)	(0.000)	(0.005)	(0.000)	(0.003)	(0.000)
Candidate age	0.000	0.001	0.006	0.013	0.000	-0.003
	(0.001)	(0.001)	(0.004)	(0.010)	(0.003)	(0.005)
Candidate incumbent	-0.004	-0.003	-0.014	-0.006	0.012*	0.026*
	(0.004)	(0.004)	(0.009)	(0.012)	(0.006)	(0.015)
Age	0.020	-0.007	-0.029	-0.044	0.053	-0.063
	(0.027)	(0.026)	(0.093)	(0.086)	(0.067)	(0.055)
Age^2	-0.000	0.000	-0.000	0.000	-0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Education	-0.008	0.004	-0.093	-0.059*	-0.068	0.008
	(0.018)	(0.016)	(0.076)	(0.031)	(0.053)	(0.051)
Urban	0.237	-0.269	1.608	0.216	-0.227	-2.465*
	(0.436)	(0.302)	(2.161)	(1.453)	(1.358)	(1.449)
Income	0.002	0.000	0.020*	0.006	0.002	-0.004
	(0.003)	(0.002)	(0.011)	(0.006)	(0.011)	(0.013)
$Income^2$	-0.000	0.000	-0.001*	-0.000	0.000	0.001
	(0.000)	(0.000)	(0.001)	(0.000)	(0.001)	(0.001)
LFP	-1.858	-0.928	-13.608**	-8.267	-0.087	-7.115
	(1.163)	(1.491)	(6.474)	(7.615)	(6.615)	(6.812)
LFP gap	-0.824	0.552	-10.859**	-3.070	-3.254	-4.123
	(0.806)	(0.779)	(5.019)	(2.969)	(3.442)	(3.162)
Sex ratio	3.191	3.644	-8.077	-0.499	7.803	4.150
	(3.030)	(2.687)	(12.331)	(10.851)	(11.525)	(13.489)
Married	-1.206	-1.437	-5.896	-4.149	-8.319	-7.394
	(1.620)	(1.363)	(5.411)	(4.683)	(6.067)	(5.471)
Log(Population)	-0.000	0.000	-0.003	-0.002	0.003	0.002
	(0.001)	(0.001)	(0.002)	(0.002)	(0.003)	(0.004)
District dummies	no	yes	no	yes	no	yes
Observations	1699	1699	389	389	306	306
R-squared	0.157	0.260	0.076	0.240	0.212	0.472

Notes: The table reports municipality-level regressions for elections where at least one candidate in the district (columns 1–2) or in the coalition (columns 3–6) is female. Robust standard errors are reported in parentheses, adjusted for clustering at the district level. The dependent variable ΔSV^F is defined as the difference between women and men vote-shares to female candidates. All regressions include election year dummies. All controls shown except for Candidate age and candidate incumbent, are municipality averages. All controls are centered around their sample mean. ***, ** and * indicate statistical significance at 1%, 5% and 10%, respectively.

likely than men to vote for female candidates in municipalities with higher education (a 1 standard deviation increase in age increases gender bias by 1 percent). On the other hand, female candidates running for the center-right receive more votes from women than men when they are incumbent (the negative bias turns to positive when an incumbent female candidate is running), and from women that live in municipalities with less urban population (an decrease of 1 standard deviation in the share of urban population increases the gender bias in 0.8 percent). None of the other controls explains the observed gender biases with statistical significance at conventional levels.

How large are these biases? Compared to the average margin of victory in center-right and center-left races where a female candidate competes (31 and 44 percent, respectively), gender biases are small. There are, however, two caveats. First, margins of victory and gender biases are not directly comparable. The former measures the difference in the share of votes of two candidates, taking into account both female and male voters. The gender bias, by contrast, compares the share of votes of female and male voters for the female candidate. Therefore, large margins of victory can display a very small gender bias, and the opposite is also possible. Second, gender biases become larger (and therefore comparable in size to the average margin of victory) when I restrict the sample to close elections, as I show in the next section.

5.2 Evidence from close elections

As explained earlier, unobservable characteristics of the candidates might bias the results if they are correlated with the likelihood of a female candidate running in a district and with her relative support among men and women voters. This issue can be addressed by analyzing the gender bias in elections where the female candidate won or lost by a small margin. To this end, I re-estimate equation (3) restricting the sample to female candidates in elections where the margin of victory was smaller than a certain threshold (5, 10 or 15 percent of the total votes in the coalition). The exercise is performed for center-left and center-right coalitions only, since the other smaller coalitions do not always put up two candidates, and therefore it is not always possible to compute the margin of victory.

Table 3 shows the results. In columns 1 and 4 the sample is restricted to elections whose margin is smaller than 5 percent of the votes for the coalition, while in columns 2 and 5 (resp. 3 and 6) the sample considers a 10 percent (resp. 15 percent) margin of victory. The main result shown in Table 2 is confirmed: Elections of center-left female candidates display a positive and statistically significant gender bias, while the opposite is true for center-right female candidates. For female candidates on the center-left, the positive gender bias ranges between 0.5 and 1.9 percent, depending on the specification. Female candidates on the center-right, on the other hand, exhibit a negative gender bias that ranges between 0.4 and 1.2 percent.

These results show that gender biases are not reduced when analyzing close elections in which competitive female candidates run. Even more so, the evidence suggests that gender biases increase in close elections. Another interesting result is that candidates' characteristics have significant effects on the relative support of men and women voters, despite being balanced across candidates on average. For instance, incumbent female candidates on the center left receive significantly less support from female voters than male voters (columns 1 and 2). The same is true for center-right female incumbents though only in elections where the margin of victory is smaller than 15 percent. Also, older center-right candidates receive significantly less support from women voters than from men voters in very close races (columns 4 and 5); however this result reverses when races with a margin of victory of up to 15 percent are included (column 6).[23]

The magnitude of the effects observed in Table 3 are comparable to the sample average margin of victory, displayed at the bottom of the table. The positive gender bias towards female center-left candidates ranges between 18 and 38 percent of the average margin of victory, while the negative gender bias towards female center-right candidates ranges between 8 and 41 percent of the average margin of victory, depending on the specification.

[23]Given that all covariates are centered around their mean value, the large and significant effects of incumbency and age could partially account for the gender bias being larger in close elections than in all elections. For example, for the sample in column 2 of Table 3 the share of incumbent candidates is 22 percent, compared to 33 percent for the sample in column 3 of Table 2.

Table 3: Determinants of gender bias for female candidates, close elections

Dep. Variable:	ΔSV^F (Difference in vote-shares to female candidates)					
Sample:	Center-left			Center-right		
Margin of victory:	$\leq 5\%$	$\leq 10\%$	$\leq 15\%$	$\leq 5\%$	$\leq 10\%$	$\leq 15\%$
	(1)	(2)	(3)	(4)	(5)	(6)
Average gender bias	0.005***	0.018***	0.019***	-0.012***	-0.004***	-0.012***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Candidate age	-0.001	0.011	0.045	-0.025**	-0.023***	0.004***
	(0.002)	(0.006)	(0.049)	(0.007)	(0.006)	(0.000)
Candidate incumbent	-0.042***	-0.078***	0.055		-0.003	-0.041**
	(0.004)	(0.012)	(0.054)		(0.008)	(0.014)
Age	0.372	0.477	-0.039	-0.034	0.265	-0.235
	(0.300)	(0.338)	(0.110)	(0.282)	(0.361)	(0.173)
Age^2	-0.001	-0.001	0.000	0.000	-0.000	0.000
	(0.000)	(0.001)	(0.000)	(0.000)	(0.001)	(0.000)
Education	-0.095	-0.040	-0.078	0.120	0.067	0.143*
	(0.054)	(0.069)	(0.057)	(0.194)	(0.044)	(0.076)
Urban	4.581	2.841	1.962	-4.413**	-10.601*	-2.859
	(2.511)	(2.139)	(1.935)	(1.346)	(4.983)	(3.312)
Income	-0.001	-0.005	0.008	-0.003	-0.031	-0.024
	(0.034)	(0.011)	(0.010)	(0.022)	(0.028)	(0.014)
Income^2	0.000	0.000	-0.000	-0.000	0.002	0.001
	(0.007)	(0.000)	(0.000)	(0.001)	(0.002)	(0.001)
LFP	18.764	9.926	5.892	-20.587*	-51.773*	-17.813
	(9.806)	(9.288)	(10.358)	(8.259)	(23.673)	(11.947)
LFP gap	10.775	4.009	0.748	-4.881	-16.073*	1.279
	(7.108)	(5.119)	(6.409)	(2.934)	(8.157)	(4.538)
Sex ratio	-17.802	-8.006	-4.051	5.075	-29.666**	-21.905
	(16.928)	(10.629)	(11.937)	(29.938)	(9.726)	(14.342)
Married	-12.723	-13.164	-18.138*	-18.148*	-21.519	-18.410*
	(9.590)	(10.191)	(10.436)	(7.919)	(12.840)	(9.715)
Log(Population)	-0.009	-0.008*	-0.006	-0.005	0.006	-0.008
	(0.007)	(0.005)	(0.004)	(0.016)	(0.003)	(0.008)
District dummies	yes	yes	yes	yes	yes	yes
Observations	51	88	151	22	42	79
R-squared	0.569	0.651	0.519	0.953	0.917	0.786
Average margin of victory	0.028	0.047	0.078	0.029	0.048	0.091

Notes: The table reports municipality-level regressions for elections where at least one candidate in the coalition is female, and where the margin of victory of the winning candidate is either 5, 10, or 15 percent. Robust standard errors are reported in parentheses, adjusted for clustering at the district level. The dependent variable ΔSV^F is defined as the difference between women and men vote-shares to female candidates. All regressions include the log of population and election year dummies as controls. All controls shown except for Candidate age and candidate incumbent, are municipality averages. All controls are centered around their sample mean. ***, ** and * indicate statistical significance at 1%, 5% and 10%, respectively.

6 Possible confounding factors

The results in the previous section show that women do not always prefer the female candidate over the male candidate, neither in elections at large, nor when elections are close and thus candidates' characteristics other than gender are comparable. In this section I discuss possible explanations, as well as other confounding factors that could be driving these results.

6.1 Gender identity

One possibility is that the gender of the candidate not only signals the candidate's preferred policy platform, but also has an effect on voters' behavior which may depend on voters' identity. Akerlof and Kranton (2000) postulate that social categories, such as "man" and "woman", are associated with physical attributes and prescribed behaviors. From these prescribed behaviors individuals form their identity as a man or as a woman. Acting in a way that differs from these behaviors generates discomfort (i.e. a negative payoff) in oneself and others. In this context, women supporters of a more conservative party, such as the ones in the center-right coalition, could have conservative views of a woman's identity. They might feel their identity threatened when a female candidate runs for office (which they see as outside appropriate behaviors for a woman), and may therefore refrain from voting for her to validate or preserve the social norm. Gender identity can therefore explain why women vote less often for female center-right candidates than men do.[24]

The regressions shown in the previous section include the Labor Force Participation (LFP) gap as a regressor. This variable should be a good predictor of the prevalence of traditional gender roles.[25] The evidence from Table 2 shows that female candidates from

[24]Men might also feel their identity of "breadwinners" threatened when a female candidate runs for office. Therefore an additional condition is that the gender identity effect must be larger for women than men.

[25]The LFP gap seems to be a good measure of what both men **and** women think their roles should be. Evidence to support this argument comes from Booth and Van Ours (2009), who analyze the relationship between full- and part-time work and family wellbeing. They find that women part-time workers are more satisfied with working hours than full-time women, but that their satisfaction increases if their partners work full-time. On the other hand, male's satisfaction is unaffected by their partners work decision but it

both center-left and center-right coalitions receive less votes from women voters than from men voters in municipalities with a larger LFP gap. However, the estimate is only significant in column 3, for the center-left and when district dummies are not included. Indeed, the point estimate is reduced to a third when district dummies are included (column 4), and it is no longer significant at conventional levels. The results for close elections (Table 3) are more auspicious, with the coefficient for LFP gap being negative and significant at 10 percent for elections in which the center-right coalition won by a margin of less than 10 percent (column 5). The effect is sizeable: One standard deviation increase in the LFP gap can reduce the votes to female candidates coming from women voters by 2.7 percent, compared to the votes received from men. The effect is still negative but less precisely estimated when the sample is restricted to elections won with a 5 percent margin (column 4), but it disappears when the margin of victory is increased to 15 percent (column 6). Overall there is some, but not robust, evidence of gender identity having an effect on the support for female candidates, particularly in close elections.

It is possible that the LFP gap does not capture the complexity of gender identity. It might also be the case that men voters feel that their identity as breadwinners is also threatened when a female candidate runs for office and therefore, even though it imperfectly measures traditional gender roles, the LFP gap decreases both men and women voters' likelihood of voting for a female candidate.

6.2 Legislators' Behavior

Do the findings in the previous section reflect the fact that female legislators do not differentiate themselves from their male counterparts by, for instance, not supporting different bills, or by supporting bills that are not preferred by the female electorate? An alternative to the traditional Downsian framework is a model in which (center-right) female candidates "act tough" in order to appeal to men voters but who in turn alienate women voters

increases if they themselves work full-time. This can be regarded as evidence consistent with the gender identity hypothesis, and generates a prediction for the LFP gap (once one has controlled for the average level of LFP).

(Herrnson *et al.*, 2003). To investigate this possibility I collected roll-call voting data from the Chamber of Deputies to test whether female legislators vote differently from their male counterparts. I describe the data and present summary statistics in online appendix B.

I look at the likelihood that a legislator voted differently from the majority of her coalition.[26] To identify the bills where it is relevant to cast a vote in opposition to the coalition I restrict the sample to include only *party unity votes*, in which the majority of the center-left coalition voted differently from the majority of the center-right coalition.[27]

Comparing outcomes of legislators of opposite gender by estimating OLS regressions is likely to produce biased results, since the gender of legislators is not randomly assigned (see Gagliarducci and Paserman, 2012 for a thorough discussion). To account for this I implement a Regression Discontinuity (RD) design, which exploits the discontinuity in the treatment assignment to causally identify the effect.[28] In this setup, the forcing variable is the margin of victory of a female candidate relative to a male candidate, in a within-coalition election. When the margin of victory is above the cutoff $c = 0$, the gender of the elected legislator is female. The opposite is true for elections with a margin of victory below c. This leads to the following specification:

$$DIF_{iv} = \gamma_1 + \gamma_2\, FEMALE_i + f(MV_i) + \epsilon_{iv}, \text{ with } MV_i \in (-h, h). \tag{4}$$

The dependent variable DIF is a dummy for whether the legislator voted differently from the majority of her/his coalition. The dummy FEMALE (the treatment in the RD terminology) takes the value of 1 for female legislators. MV_i, the margin of victory, is the forcing variable, and $f(\cdot)$ is a control function, in this case a polynomial in MV_i on each side

[26] A similar strategy was used in Rehavi (2007) to analyze roll-call data from U.S. State Assemblies. Previous papers analyzing the voting behavior of U.S. congressmen relied on voting scores for "pro-female" legislation, or analyzed a subset of bills with a clear position on a particular issue (e.g. Lee, Moretti and Butler, 2004; Washington, 2008). This information is not available for bills discussed in the Chilean Congress.

[27] This is a well established measure of discrepancy between parties. The term was introduced by Congressional Quarterly (CQ), a company that produces reports of roll-call voting statistics.

[28] See Imbens and Lemieux (2008) for details regarding identification in RD designs. Gagliarducci and Paserman (2012), Ferreira and Gyourko (2014) and Brollo and Troiano (2014) have implemented RD designs in the context of mixed-gender elections. In this formulation I follow Meyersson (2014).

of the cutoff. Finally, h is the bandwidth. I follow Meyersson (2014) in presenting results of local linear regressions with an optimal bandwidth $\hat{h}$ and a linear control function as the main specification. The optimal bandwidth is estimated using Imbens and Kalyanaraman (2012)'s procedure. I also present results of increasing the bandwidth to $2\hat{h}$ and of including a quadratic control function.[29]

In online appendix C I show standard validity tests employed in the RD literature (Imbens and Lemieux, 2008). In particular, the McCrary (2008) density test does not show evidence of manipulation of the margin of victory. Regarding the balance of covariates, three of the four graphs do not reveal any significant jump at the cutoff, the only exception being incumbency in Center-left close elections, where female candidates are less likely to be incumbents than male candidates. However, only in 1 of 3 regression specifications I find this relationship to be statistically significant at 10 percent.

Table 4 shows the results for elections in the Center-left coalition in Panel A and for the center-right coalition in Panel B. Columns 1 and 2 present the results of OLS estimates, without (column 1) and with additional controls (column 2).[30] The next four columns present the results of the RD regressions. Starting with Panel A, both the results of the OLS and RD estimates indicate a small and imprecise effect for female legislators. While in column 1 there is some evidence of a negative effect that is significant at 10 percent, in column 4 there is some evidence of a positive and significant effect. This positive effect, however, is not confirmed on any of the other RD specifications. This suggests that female legislators from the Center-left coalition vote similar to their male peers.

The results from Panel B suggest something different for Center-right female legislators. The OLS results in the first two columns show a positive but not statistically significant effect. This effect is corroborated in the RD regressions, particularly when additional controls are included (columns 4 and 6). These results indicate that female legislators from the Center-right coalition are more likely to vote different from the majority of their coalition

[29] I do not present results of using a smaller bandwidth since it significantly reduces the sample. I also abstain from using higher-order polynomials as control functions (Gelman and Imbens, 2014).

[30] These additional controls are incumbency, candidate age, a dummy for whether the bill requires a special quorum, year and issue dummies.

Table 4: Probability of different vote

Dep. Variable:	Vote different from the majority of the coalition					
Control function:	None		Linear			Quadratic
Bandwidth:	Global		$\hat{h}$		$2\hat{h}$	$\hat{h}$
Covariates:	No	Yes	No	Yes	Yes	Yes
	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Center-left					
FEMALE	-0.031*	-0.026	-0.001	0.090**	0.048	-0.034
	(0.017)	(0.020)	(0.024)	(0.033)	(0.047)	(0.138)
Bandwidth	1.000	1.000	0.156	0.156	0.313	0.156
R^2	0.003	0.033	0.005	0.050	0.042	0.052
Observations	16802	16802	6524	6524	10421	6524
	Panel B: Center-right					
FEMALE	0.034	0.051	0.104	0.131**	0.071	0.187***
	(0.037)	(0.039)	(0.117)	(0.048)	(0.072)	(0.056)
Bandwidth	1.000	1.000	0.285	0.285	0.571	0.285
R^2	0.004	0.051	0.023	0.127	0.066	0.130
Observations	13949	13949	5296	5296	9466	5296
	Panel C: Test for coefficient equality between center-left and center-right					
p-value	0.105	0.071	0.371	0.473	0.792	0.129
	Panel D: Center-left, women's issues					
FEMALE	-0.036	-0.028	0.048	0.190***	0.125*	-0.423
	(0.026)	(0.029)	(0.045)	(0.044)	(0.069)	(0.267)
Bandwidth	1.000	1.000	0.130	0.130	0.260	0.130
R^2	0.004	0.072	0.007	0.075	0.078	0.079
Observations	7481	7481	2845	2845	4441	2845
	Panel E: Center-right, women's issues					
FEMALE	0.051	0.077	0.198	0.291***	0.152	0.296***
	(0.050)	(0.053)	(0.191)	(0.080)	(0.105)	(0.088)
Bandwidth	1.000	1.000	0.220	0.220	0.440	0.220
R^2	0.008	0.086	0.062	0.279	0.150	0.276
Observations	6224	6224	2138	2138	3487	2138
	Panel F: Test for coefficient equality between center-left and center-right					
p-value	0.113	0.080	0.433	0.255	0.842	0.009

Notes: The table reports estimates of equation (4). Panel A shows regression results for the Center-left coalition, while Panel B shows the corresponding results for the Center-right coalition. Panel C presents p-values from Seemingly Unrelated Regression (SUR) tests of equality of the coefficient FEMALE in panels A and B. Similarly, panels D and E show results for Center-left and Center-right coalitions for the subset of bills in which a women's issue (education, justice, health, labor or family) is the primary issue. Panel F presents p-values from SUR tests of equality of the coefficient FEMALE in panels D and E. Robust standard errors in parentheses, adjusted for clustering at the legislator level. Covariates include seniority, candidate age, quorum and year dummies. ***, ** and * indicate statistical significance at 1%, 5% and 10%, respectively.

than their male peers. Depending on the specification, they can be between 10 and 19 percent more likely to vote different from the majority of the coalition than men. However, the p-value from testing the difference in the treatment effect for Center-left and Center-right coalitions, shown in Panel C, fails to reject the null hypothesis of no difference in all cases but in column 2.

In the following panels of Table 4 I repeat the exercise, but now restricting the sample to those bills where a "women's issue" was the primary issue. I consider as women's issues education, government, justice, health, labor and family.[31] Panels D and E, showing results for the Center-left and Center-right coalitions, respectively, confirm the findings in panels A and B. Regarding panel D, there is a larger and significant treatment effect in columns 4 and 5. Column 6, however, still displays a negative but insignificant effect. As of Panel E, there is a large positive and statistically significant effect in columns 4 and 6. These results suggest that Center-right female legislators can be up to 30 percent more likely to vote different from their peers when voting on women's issues.

By and large, these results show that female center-right legislators do differentiate themselves from male legislators, particularly when voting on bills related to women's issues. Even though candidates could deviate from their campaign platform once elected, it is unlikely that the "female running as male" hypothesis can explain the negative gender bias in the Center-right coalition.

6.3 Are right-wing women more right-wing than men?

In the previous section I have shown that female legislators from the Center-right coalition do differentiate themselves from male legislators by voting differently more often than their peers, particularly on bills whose primary issue is a women's issue. An alternative explanation for the negative gender bias, also consistent with the previous results, is that

[31] There is a long literature discussing the definition of women's issues. I adopt the definition of Volden, Wiseman and Wittmer (2013), who determine women's issues as those where a larger proportion of women introduce bills in the U.S. House than do men. These issues are civil rights and liberties; education; health; labor, employment and migration; and law, crime and family.

Center-right female voters might choose to vote for the male candidate instead of the female candidate because they are more conservative than Center-right men. As mentioned earlier, Lewis (2004) and Carrera and Ulloa (2006) have documented that women are more likely to vote for conservative parties than men. However, conditioning on being a right-wing supporter, are they more rightist than men?

I analyze this possibility by looking at questions regarding attitudes towards democracy, free market and gender roles in the *Latinobarómetro* Survey. Table 5 shows regressions where the dependent variable is a dummy taking the value of 1 for a right-wing answer in different questions from the *Latinobarómetro* Survey. The details regarding sample restrictions and the choice of questions can be found in Online Appendix D.

Panel A presents results of regressions for a sample of right-wing voters, defined as the set of individuals that place themselves between 6 and 9 in a left-right scale ranging from 0 to 10. In the first column the dependent variable is a dummy that takes the value of 1 if the respondent favors a market economy; in columns 2 and 3 it takes the value of 1 if the respondent expresses less favorable attitudes towards democracy, while in columns 4 and 5 the dependent variable equals 1 if the respondent reveals a more conservative attitude towards the role of women. This set of questions aims at capturing the economically-liberal/morally-conservative nature of the post-dictatorship right in Chile.[32] The main regressor is a dummy for the respondent's gender. In all regressions but in column 2 right-wing women significantly differ from right-wing men in their attitudes towards the economy, democracy and gender roles. Particularly interesting is the result of column 5, which shows that women are 11 percent less likely to agree with the statement "it is better that women concentrate at home and men at work". In panel B I reproduce these results, but now restricting the sample to individuals that would vote for a party of the Center-right coalition in the next elections. The number of observations in each column is smaller since for some years it is not possible to identify the party. Despite exhibiting larger standard errors, the point estimates go in the same direction of those in panel A (except for column

[32] Pollack (1999) describes the new right in Chile as a mixture of neo-liberals and neo-conservatives that united despite their differences, with conservatism arising as a response to the primacy of liberalism.

Table 5: Conservatism among Center-right voters

Dep. Variable:	Marketecon (1)	Demprob (2)	Gobauto (3)	Womenopp (4)	Womenrole (5)	Menleaders (6)
Panel A: Score between 6 and 9 in the left-right scale						
Woman	-0.056**	0.015	-0.037**	-0.123***	-0.109***	-0.302***
	(0.026)	(0.022)	(0.018)	(0.021)	(0.033)	(0.069)
R^2	0.315	0.246	0.173	0.152	0.254	0.402
Observations	1205	1790	2897	967	759	346
Panel B: Vote for Center-right in next elections						
Woman	0.015	0.026	-0.024	-0.090	-0.101**	-0.246***
	(0.040)	(0.025)	(0.028)	(0.055)	(0.045)	(0.059)
R^2	0.312	0.290	0.239	0.188	0.353	0.479
Observations	842	1391	1688	548	536	312

Notes: The table presents results of regressions where the dependent variable is a dummy created from questions of the Latinobarómetro Survey for Chile in 1995–2011. Marketecon takes the value of 1 if the respondent agrees that a market economy is the only system through which the country can become developed. Demprob takes the value of 1 if the respondent does not agree that democracy is the best system of government. Gobauto takes the value of 1 if the respondent agrees that under some circumstances an authoritarian government can be preferred to a democratic one. Womenopp takes the value of 1 if the respondent agrees that women have the same opportunity than men to earn the same salary. Womenrole takes the value of 1 if the respondent agrees that it is better that women concentrate at home and men at work. Finally, Menleaders takes the value of 1 if the respondent agrees that men are better political leaders than women. The variable Woman takes the value of 1 if the respondent is female. In panel A the sample is restricted to those individuals who place themselves between 6 and 9 in a left-right scale from 0 to 10. Panel B restricts the sample to individuals who would vote for the Center-right in the next elections. All regressions include the respondent's age and age squared, marital status and education, as well as city and year dummies. Robust standard errors in parentheses, adjusted for clustering at the city level. ***, ** and * indicate statistical significance at 1%, 5% and 10%, respectively.

1), and in column 5 the estimate is still significant at 5 percent.

These results suggest that it is unlikely that differences between Center-right women and men voters in their attitudes towards the role of women can explain the distaste of women voters for women candidates.

7 Conclusion

Recent literature has stressed the importance of leaders, in particular female leaders, to economic outcomes. The gender of the leader can be thought of as a signal of the leader's preferences. Therefore, if women share similar policy preferences, as a group they should show a larger support for female leaders than for male leaders. This is difficult to observe because of the secrecy of the ballot. In order to analyze women's support for female candidates, this paper makes use of a unique dataset from Chilean congressional elections where women and men vote separately.

I find that on average women vote slightly less often for female candidates than men do. This negative gender bias breaks down into a positive gender bias among center-left voters and a negative gender bias among center-right voters, both statistically significant. The results do not seem to be explained away by focusing on competitive elections; on the contrary, the biases increase in magnitude. Moreover, the analysis of the legislators' voting records provides evidence against a policy-based explanation of this gender bias.

If within-coalition elections were modeled using the citizen-candidate model, its predictions would not find support in my data. The data does not support Downsian models either, since center-right female legislators move away from their coalition's preferred policy and vote differently, particularly on social issues. These results underline the complexity of the process of aggregation of individual preferences.

The evidence presented in this paper does not question whether female leaders have an effect on economic outcomes, but instead challenges the mechanism through which this effect takes place. Thus, women leaders might matter not because they implement policies which are closer to female preferences, but because they perform better in other dimensions such as being less susceptible to corruption (e.g. Brollo and Troiano, 2014), or being better at negotiating for others (e.g. Bowles, Babcock and McGinn, 2005).

The paper contributes to the literature in two other aspects. First, it underlines the importance of actual voting data instead of surveys or exit polls, especially when differential response biases could be present. And second, it adds to the growing literature which finds that women do not perform better when evaluated by their gender-peers.

References

Akerlof, G. and Kranton, R. (2000). Economics and identity. *Quarterly Journal of Economics*, **115** (3), 715–753.

Alemán, E. and Saiegh, S. (2007). Legislative Preferences, Political Parties, and Coalition Unity in Chile. *Comparative Politics*, **39** (3), 253–272.

Bagues, M. F. and Esteve-Volart, B. (2010). Can Gender Parity Break the Glass Ceiling? Evidence from a Repeated Randomized Experiment. *Review of Economic Studies*, **77**, 1301–1328.

—, Sylos-Labini, M. and Zinovyeva, N. (2014). Do gender quotas pass the test? Evidence from academic evaluations in Italy. *Unpublished Manuscript*.

Bardhan, P., Mookherjee, D. and Torrado, M. P. (2010). Impact of Political Reservations in West Bengal Local Governments on Anti-Poverty Targeting. *Journal of Globalization and Development*, **1** (1).

Beaman, L., Chattopadhyay, R., Duflo, E., Pande, R. and Topalova, P. (2009). Powerful women: Does exposure reduce bias? *Quarterly Journal of Economics*, **124** (4), 1497–1540.

Benstead, L. J. (2014). Effects of Interviewer–Respondent Gender Interaction on Attitudes toward Women and Politics: Findings from Morocco. *International Journal of Public Opinion Research*, **26** (3), 369–383.

Besley, T. and Coate, S. (1997). An economic model of representative democracy. *Quarterly Journal of Economics*, **112** (1), 85–114.

Blaydes, L. and Gillum, R. (2013). Religiosity-of-Interviewer Effects: Assessing the Impact of Veiled Enumerators on Survey Response in Egypt. *Politics and Religion*, **6** (3), 459–482.

Booth, A. L. and Van Ours, J. C. (2009). Hours of work and gender identity: Does part-time work make the family happier? *Economica*, **76** (301), 176–196.

Bowles, H., Babcock, L. and McGinn, K. (2005). Constraints and triggers: Situational mechanics of gender in negotiation. *Journal of Personality and Social Psychology*, **89** (6), 951.

Brollo, F. and Troiano, U. (2014). What Happens When a Woman Wins an Election? Evidence from Close Races in Brazil. *Unpublished Manuscript*.

Camara de Diputados de Chile (2011). *Detalle de Votaciones*. http://www.camara.cl/trabajamos/sala_votaciones.aspx (accessed September 14, 2011).

Carrera, C. and Ulloa, J. (2006). Evolución del voto femenino y las candidaturas de mujeres en los procesos electorales locales en Chile: 1992-1996- 2000 y 2004. *Avances de Investigacion, Fundacion Carolina*, pp. 1–95.

Carrillo, J. D. and Castanheira, M. (2008). Information and strategic political polarisation. *Economic Journal*, **118** (530), 845–874.

Caughey, D. and Sekhon, J. S. (2011). Elections and the Regression Discontinuity Design: Lessons from Close U.S. House Races, 1942-2008. *Political Analysis*, **19** (4), 385–408.

Cerda, R. and Vergara, R. (2009). Voter Turnout: Evidence from Chile. *Mimeo*.

Chattopadhyay, R. and Duflo, E. (2004). Women as Policy Makers: Evidence from a Randomized Policy Experiment in India. *Econometrica*, **72** (5), 1409–1443.

Clots-Figueras, I. (2011). Women in politics. Evidence from the Indian States. *Journal of Public Economics*, **95** (7-8), 664–690.

— (2012). Are Female Leaders Good for Education? Evidence from India. *American Economic Journal: Applied Economics*, **4** (1), 212–244.

Dolan, K. (1998). Voting for women in the "year of the woman". *American Journal of Political Science*, **42** (1), 272–293.

— (2008). Is There a "Gender Affinity Effect" in American Politics? Information, Affect, and Candidate Sex in U.S House Elections. *Political Research Quarterly*, **61** (1), 79–89.

Downs, A. (1957). *An Economic Theory of Democracy*. New York: Harper and Row.

Duflo, E. (2003). Grandmothers and Granddaughters: Old-Age Pensions and Intrahousehold Allocation in South Africa. *The World Bank Economic Review*, **17** (1), 1–25.

Duverger, M. (1955). *The Political Role of Women*. Paris, France: UNESCO.

Eggers, A. C., Fowler, A., Hainmueller, J., Hall, A. B. and Snyder Jr, J. M. (2014). On the validity of the regression discontinuity design for estimating electoral effects: New evidence from over 40,000 close races. *American Journal of Political Science*.

Epstein, W. M. (2006). Response bias in opinion polls and american social welfare. *The Social Science Journal*, **43** (1), 99–110.

Ferreira, F. and Gyourko, J. (2014). Does gender matter for political leadership? The case of US mayors. *Journal of Public Economics*, **112**, 24–39.

Flores-Macias, F. and Lawson, C. (2008). Effects of interviewer gender on survey responses: Findings from a household survey in Mexico. *International Journal of Public Opinion Research*, **20** (1), 100–110.

Funk, P. and Gathmann, C. (2010). Gender Gaps in Policy Making: Evidence from Direct Democracy in Switzerland. *Unpublished Manuscript*.

Gagliarducci, S. and Paserman, M. D. (2012). Gender interactions within hierarchies: Evidence from the political arena. *Review of Economic Studies*, **79** (3), 1021–1052.

Gelman, A. and Imbens, G. W. (2014). Why high-order polynomials should not be used in regression discontinuity designs. *NBER Working Paper*, (20405).

Herrnson, P., Lay, J. C. and Stokes, A. (2003). Women running "as women": Candidate gender, campaign issues, and voter-targeting strategies. *Journal of Politics*, **65** (01), 244–255.

Huddy, L., Billig, J., Bracciodieta, J., Hoeffler, L., Moynihan, P. J. and Pugliani, P. (1997). The effect of interviewer gender on the survey response. *Political Behavior*, **19** (3), 197–220.

Imbens, G. W. and Kalyanaraman, K. (2012). Optimal bandwidth choice for the regression discontinuity estimator. *Review of Economic Studies*, **79** (3), 933–959.

— and Lemieux, T. (2008). Regression discontinuity designs: A guide to practice. *Journal of Econometrics*, **142** (2), 615–635.

Kane, E. W. and Macaulay, L. J. (1993). Interviewer gender and gender attitudes. *Public opinion quarterly*, **57** (1), 1–28.

Kartik, N. and McAfee, R. P. (2007). Signaling character in electoral competition. *American Economic Review*, **97** (3), 852–870.

Lee, D. S., Moretti, E. and Butler, M. J. (2004). Do Voters Affect Or Elect Policies? Evidence From The U. S. House. *Quarterly Journal of Economics*, **119** (3), 807–859.

Lee, T. and Pérez, E. O. (2014). The persistent connection between language-of-interview and latino political opinion. *Political Behavior*, **36** (2), 401–425.

Lewis, P. (2004). The 'Gender Gap' in Chile. *Journal of Latin American Studies*, **36** (04), 719–742.

McCrary, J. (2008). Manipulation of the running variable in the regression discontinuity design: A density test. *Journal of Econometrics*, **142**, 698–714.

Meyersson, E. (2014). Islamic rule and the empowerment of the poor and pious. *Econometrica*, **82** (1), 229–269.

Miller, G. (2008). Women's Suffrage, Political Responsiveness, and Child Survival in American History. *Quarterly Journal of Economics*, **123** (3), 1287–1327.

Ministerio de Planificación, Gobierno de Chile (2000, 2006, 2009). *Encuesta CASEN*. http://www.mideplan.gob.cl/casen/ (accessed August 4, 2010).

Osborne, M. and Slivinski, A. (1996). A model of political competition with citizen-candidates. *Quarterly Journal of Economics*, **111** (1), 65–96.

Paolino, P. (1995). Group-salient issues and group representation: Support for women candidates in the 1992 senate elections. *American Journal of Political Science*, **39** (2), 294–313.

Pino, F. (2011). $2 + 1 < 3$: Strategic voting in the Chilean Congressional Elections. *Mimeo.*

Pollack, M. (1999). *The New Right in Chile, 1973-97*. New York: St. Martin's Press.

Rahat, G. and Sznajder, M. (1998). Electoral engineering in Chile: The electoral system and limited democracy. *Electoral Studies*, **17** (4), 429–442.

Rehavi, M. M. (2007). Sex and politics: Do female legislators affect state spending? *Mimeo.*

Siavelis, P. (2002). The hidden logic of candidate selection for chilean parliamentary elections. *Comparative Politics*, **34** (4), 419–438.

Subsecretaría de Desarrollo Regional y Administrativo, Gobierno de Chile (2001-2009). *Sistema Nacional de Información Municipal*. http://www.sinim.gov.cl (accessed January 31, 2011).

Thomas, D. (1990). Intra-household resource allocation: An inferential approach. *Journal of Human Resources*, **25** (4), 635–664.

— (1994). Like father, like son; like mother, like daughter: Parental resources and child height. *Journal of Human Resources*, **29** (4), 950–988.

Volden, C., Wiseman, A. E. and Wittmer, D. E. (2013). Women's issues and their fates in congress. *Unpublished Manuscript.*

Washington, E. (2008). Female socialization: How daughters affect their legislator fathers' voting on women's issues. *American Economic Review*, **98** (1), 311–332.

ONLINE APPENDIX (NOT FOR PUBLICATION)

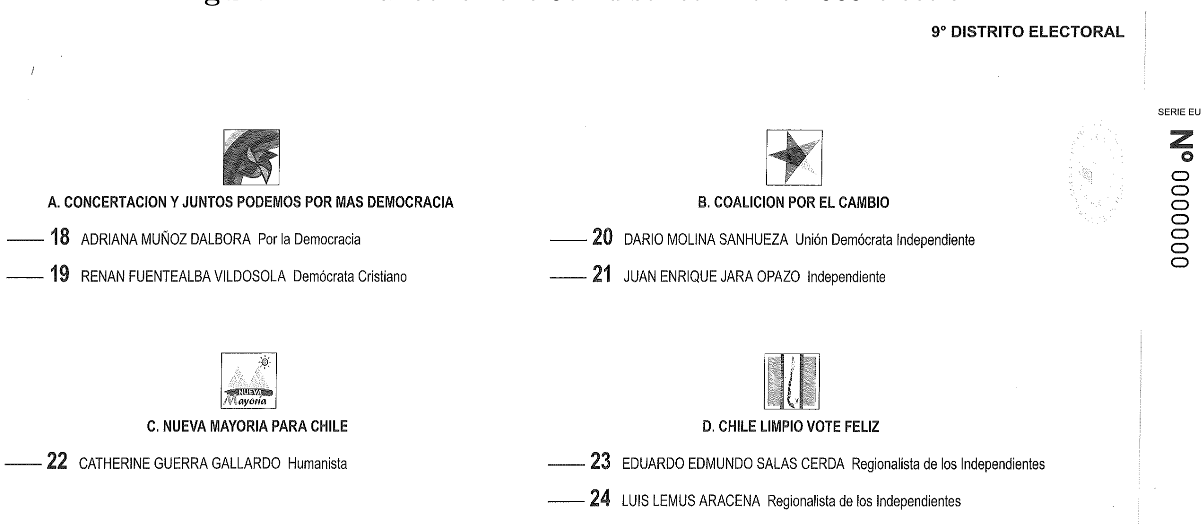
Is There Gender Bias Among Voters?

Evidence from the Chilean Congressional Elections

Francisco Pino

A. Chilean Elections: the binominal system

Figure A-1: Ballot for the 9th district in the 2009 election.


9° DISTRITO ELECTORAL

SERIE EU

Nº 0000000

A. CONCERTACION Y JUNTOS PODEMOS POR MAS DEMOCRACIA

—— 18 ADRIANA MUÑOZ DALBORA Por la Democracia

—— 19 RENAN FUENTEALBA VILDOSOLA Demócrata Cristiano

B. COALICION POR EL CAMBIO

—— 20 DARIO MOLINA SANHUEZA Unión Demócrata Independiente

—— 21 JUAN ENRIQUE JARA OPAZO Independiente

C. NUEVA MAYORIA PARA CHILE

—— 22 CATHERINE GUERRA GALLARDO Humanista

D. CHILE LIMPIO VOTE FELIZ

—— 23 EDUARDO EDMUNDO SALAS CERDA Regionalista de los Independientes

—— 24 LUIS LEMUS ARACENA Regionalista de los Independientes

Source: Chilean Electoral Service.

Table A-1: Four examples of election outcomes.

	Case 1	Case 2	Case 3	Case 4
Coalition A	40%	50%	60%	70%
Candidate A1	**30%**	**30%**	**35%**	**60%**
Candidate A2	10%	20%	**25%**	**10%**
Coalition B	40%	30%	30%	20%
Candidate B1	**22%**	**18%**	18%	18%
Candidate B2	18%	12%	12%	2%
Coalition C	20%	20%	10%	10%
Candidate C1	11%	11%	6%	6%
Candidate C2	9%	9%	4%	4%

Notes: Vote shares of elected candidates are in boldface. In cases 1 and 2 one legislator in each of the largest coalitions is elected. In case 2 coalition A "fails to double", since even though its two candidates obtain the first and second larger shares of votes, the coalition obtains less than double the votes of coalition B. In cases 3 and 4 coalition A "doubles" coalition B and gets candidates A1 and A2 elected. Case 4 is a peculiar one, since candidate B1 has a larger share of votes than candidate A2. Even though the four cases are possible, during the last 6 elections cases 1 and 2 have been the most frequent outcome with 87% of the total cases.

B. Details on data colletion

Demographic controls

The *Encuesta CASEN* is not performed every year, so I use the survey conducted the year closest to the election.[1] The survey is designed to evaluate the impact and focus of social policies, and therefore it aims at covering the entire Chilean territory. Nonetheless, even in the 2009 round there were 12 municipalities (3.5 percent of the total) that were not covered by the survey. Table B-1 shows the availability of controls by election year. I imputed the average of neighboring municipalities when they were in the same *Provincia* (Province), the next administrative unit. This procedure works for 11 municipalities in year 2009, leaving only 1 municipality without demographic controls.

Table B-1: Encuesta CASEN: availability of controls.

	Year						
Election year	1989	1993	1997	2001	2005	2009	Total
Controls at municipal level	151	244	243	304	335	334	1,611
Percent	45.07	72.84	71.05	88.89	96.82	96.53	78.74
Controls at province level	163	60	91	35	7	11	367
Percent	48.66	17.91	26.61	10.23	2.02	3.18	17.94
No controls	21	31	8	3	4	1	68
Percent	6.27	9.25	2.34	0.88	1.16	0.29	3.32
Total	335	335	342	342	346	346	2,046
Percent	100.00	100.00	100.00	100.00	100.00	100.00	100.00

Table B-2 present summary statistics for all demographic controls included in the regressions.

[1]In particular, I use rounds 1990, 1994, 1998, 2000, 2006 and 2009.

Table B-2: Demographics: summary statistics.

Election year	1989	1993	1997	2001	2005	2009
Age (years)	28.70	29.92	30.54	31.23	33.76	35.10
Education (years)	8.08	8.00	8.46	8.43	8.84	9.06
Urban (percent)	67.79	63.22	67.47	62.76	63.54	64.01
Income (10,000 pesos)	6.90	12.76	20.09	20.21	23.76	31.07
Labor Force Participation, LFP (percent)	54.13	56.69	56.19	55.98	59.16	56.22
LFP gap (percent)	62.54	61.30	54.95	54.76	49.58	48.29
Women (percent)	50.94	50.93	50.55	50.09	50.60	50.91
Married (percent)	62.83	62.89	61.75	61.53	59.12	58.72

Notes: All variables are averages constructed at the municipality level. LFP and LFP gap are computed for individuals of age between 18 and 60. Married is computed for individuals of age 18 and above.

Roll-call voting data

Data on recorded votes comes from the website of the Chamber of Deputies. I collected data for votes that took place between March 2002 and September 2011, corresponding to the 2002-2006, 2006-2010, and 2010-2014 legislative periods. The total number of votes is 6,163, but this number reduces to 4,969 when considering votes for bills that were assigned to a specific committee, which determines the broader issue of the bill (*materia*).[2] In each of these votes I record the vote (favor, against, or abstain) of each deputy that exercises her right to vote. I also have information on the name of the bill, date and time, quorum required, and the vote of each legislator present in the room (in favor, against, or abstention). Table B-3 shows summary statistics for the proportion of votes in each broader issue. Particularly interesting for my analysis are the votes on Family issues, which accounts for 6% of the votes in the first legislative period. This is due to the reform to the Civil Code in 2004, which among other changes legalized divorce, and the creation of the family courts.

Table B-4 presents the average of this variable for each of the congressional periods (2002–2006, 2006–2010 and 2010–2014), each of the coalitions and by gender of the legislator. Overall the variable ranges from 4% to 10%, but it shows important differences across

[2] The Cámara de Diputados has 25 committees, and therefore there are 25 *materias*. I aggregate some of these to finally obtain 12 broader issues.

coalitions and gender. The table also shows that in party unity votes around 10% of the legislators vote differently from their coalition, as opposed to 4% for non-party unity votes.

Table B-3: Votes in the Chamber of Deputies, summary statistics.

Congress	2002-2006	2006-2010	2010-2011[a]
Agriculture	8.92	8.89	12.14
Defense	5.11	1.94	1.97
Education	10.56	14.21	15.86
Finance	12.89	26.73	23.80
Government	10.41	8.94	9.33
Justice	14.53	12.27	12.80
Mining	4.81	3.78	4.59
Public works	5.90	5.96	5.91
Foreign relations	6.79	5.96	2.45
Health	8.38	4.02	4.04
Labor	5.65	5.56	4.39
Family	6.05	1.74	2.71
Total number of votes	2,017	2,013	951

Notes: Each number indicates the proportion of votes in each category. [a] Data up to 9/7/2011.

Table B-4: Proportion of votes differing from own coalition

Congress	2002–2006	2006–2010	2010–2014[a]
all legislators	0.063	0.066	0.077
center-left	0.043	0.065	0.121
men	0.044	0.066	0.123
women	0.037	0.059	0.109
center-right	0.086	0.067	0.028
men	0.085	0.067	0.028
women	0.101	0.069	0.029
no party unity vote	0.044	0.045	0.057
party unity vote	0.102	0.101	0.112
Observations	173,483	182,124	90,862

Notes: Each number corresponds to the proportion of votes where the legislator voted different from the majority of her own coalition. [a] Data up to 9/7/2011.

C. Close within-coalition elections: Sample averages, density and balance tests

This appendix presents sample averages for close within-coalition elections, as well as standard validity tests employed in the RD literature (Imbens and Lemieux, 2008).

Table C-1 shows sample averages and mean-comparison tests for age and incumbency of male and female candidates running in close within-coalition elections. Columns 1-3 include elections with a margin of victory below 5 percent, while in columns 4-6 and 7-9 the margin of victory increases to 10 and 15 percent, respectively. The table shows that candidates' age and incumbency are balanced across gender for margins of victory of 5 and 10 percent. When less competitive elections are included, such as elections with a margin of victory up to 15 percent (columns 7-9), the table shows that the share of incumbent male candidates is significantly higher than the share of female candidates in the Center-Left. Age, however, remains similar across gender, as well as age and incumbency of Center-Right candidates.

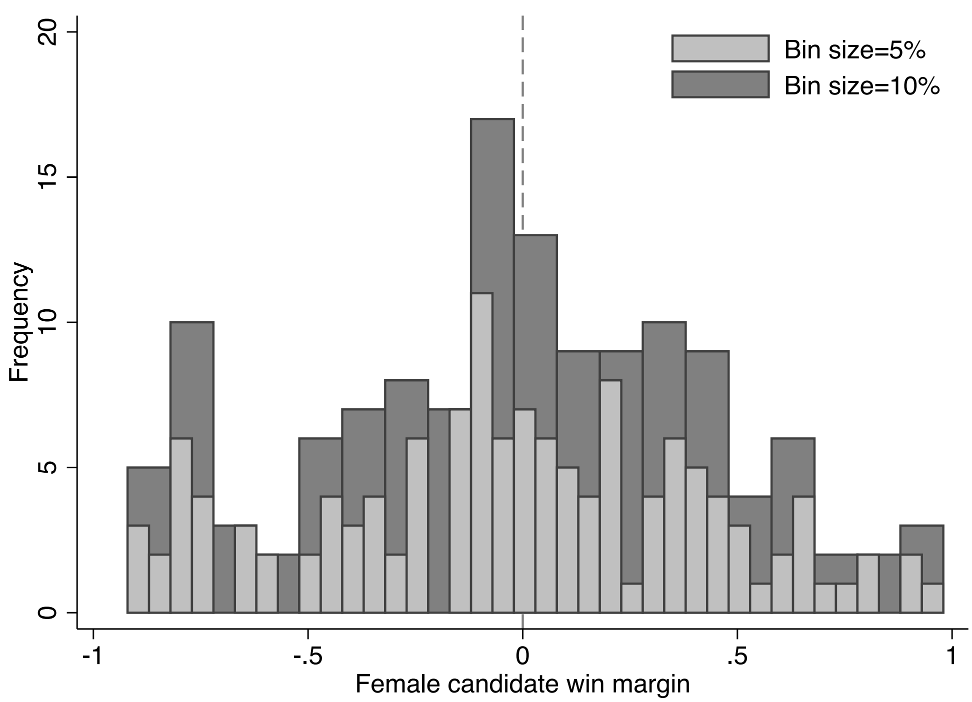
The rest of tables and figures show validity tests for the RD analysis presented in section 6.2 of the paper. Figure C-1 presents a histogram of the margin of victory, while Figure C-2 shows the McCrary (2008) density test of a jump at the discontinuity. The estimated jump at the discontinuity is not significant at conventional levels, providing evidence of no manipulation of the margin of victory. Figure C-3 and the corresponding regression results in Table C-2 examine the behavior of the available covariates (age and incumbency) in the vicinity of the threshold. Three of the four graphs in Figure C-3 do not reveal any significant jump at the cutoff, the only exception being incumbency in Center-left close elections, where female candidates are less likely to be incumbents than male candidates. However, only in 1 of 3 specifications in Table C-2 I find this relationship to be statistically significant at 10 percent.

Table C-1: Sample averages and mean-comparison tests

	Margin:	5%			10%			15%		
		Men	Women	Diff. p-value	Men	Women	Diff. p-value	Men	Women	Diff. p-value
Coalition	Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Center-left	Age	43.750	49.250	0.384	46.294	48.294	0.600	47.846	48.423	0.855
	Incumbent	0.500	0.375	0.642	0.471	0.235	0.160	0.500	0.192	0.019
Center-right	Age	43.200	44.600	0.844	44.909	43.727	0.783	44.933	43.000	0.620
	Incumbent	0.400	0.200	0.545	0.182	0.364	0.362	0.333	0.333	1.000

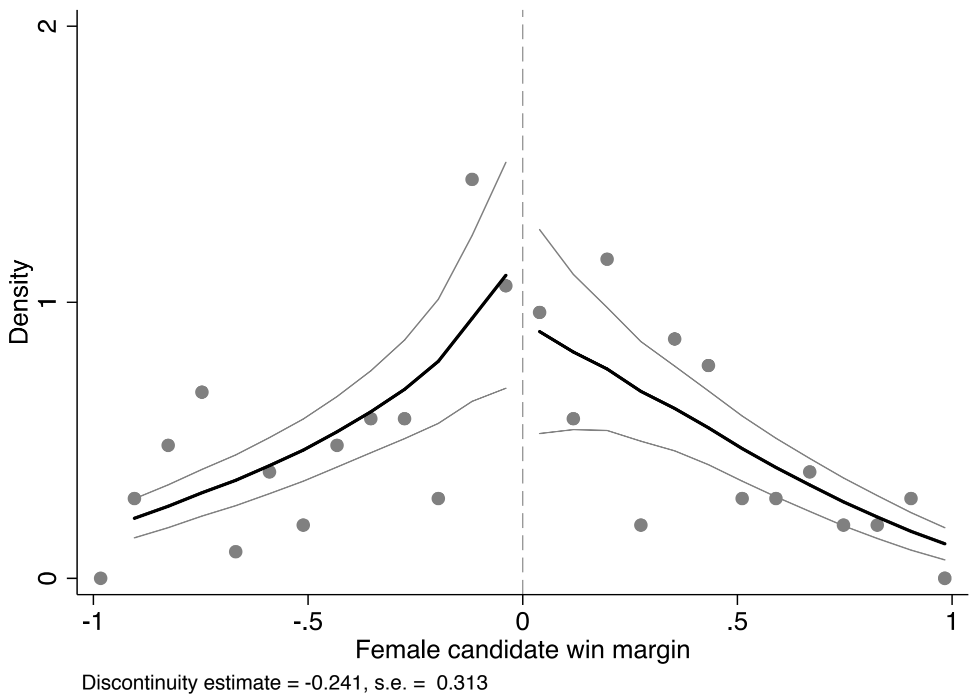
Notes: The table shows sample averages for male and female candidates running in mixed-gender elections, as well as the p-value of a mean-comparison test, for various margins of victory. Age is the age of the candidate, in years. Incumbent is a dummy that takes the value of 1 if the candidate was in office and running for a subsequent term in the same district.

Figure C-1: Histogram of margin of victory.



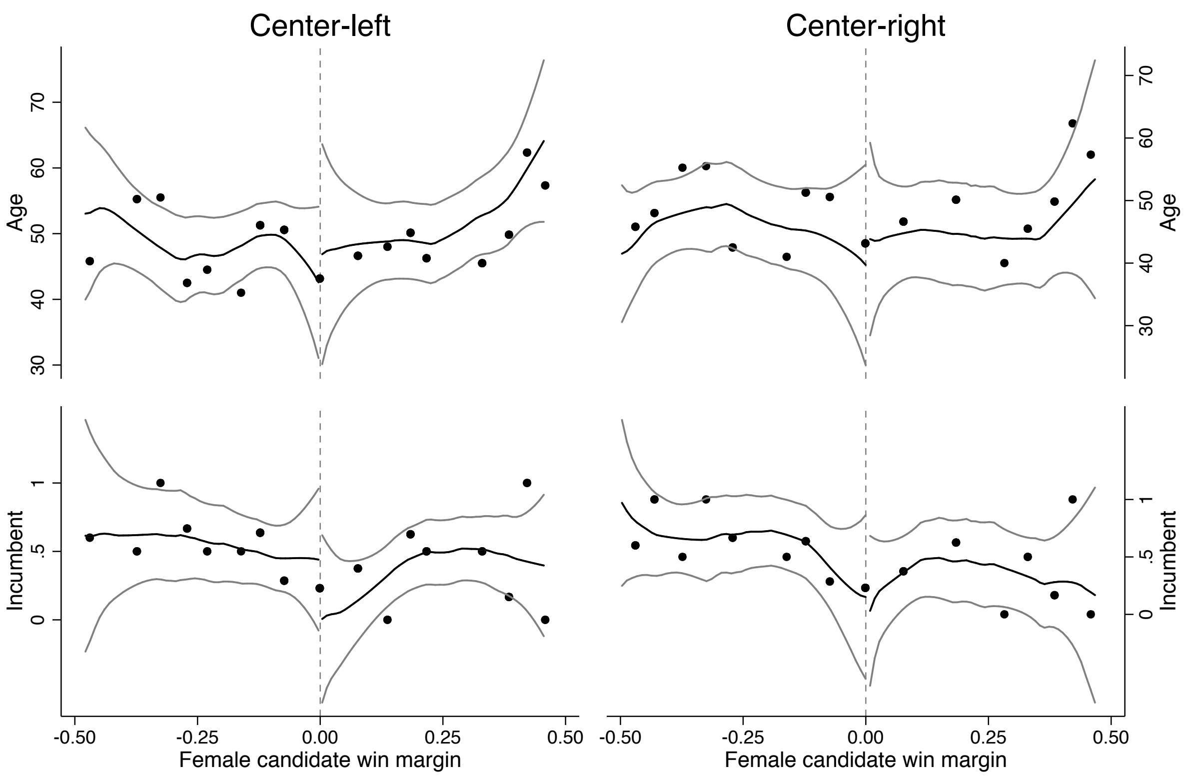
Notes: The figure shows a histogram of the margin of victory. The sample includes center-left and center-right mixed-gender elections.

Figure C-2: Density test.



Notes: The figure shows the McCrary (2008) test of a discontinuity at the female candidate win margin. The sample includes center-left and center-right mixed-gender elections.

Figure C-3: Balanced covariate tests.



Notes: The figure shows tests for a jump in age and incumbency at the zero margin of victory, for the center-left and center right separately. The sample includes center-left and center-right mixed-gender elections. Each dot plots the variable average in bins of 5 percent of the margin of victory. The black line presents smoothed values of local linear regressions estimated separately on each side of the zero margin. Gray lines present 95 percent confidence intervals.

Table C-2: Balanced covariate tests: regression results

Coalition:	Center-left		Center-right	
Outcome:	Age	Incumbent	Age	Incumbent
	(1)	(2)	(3)	(4)
A. Local linear regression with optimal bandwidth $\hat{h}$				
Female candidate elected	-1.245	-0.224	0.570	-0.231
	(6.598)	(0.206)	(6.496)	(0.203)
Bandwidth	0.284	0.547	0.531	1.000^a
Observations	40	60	38	64
B. Local linear regression with optimal bandwidth $2\hat{h}$				
Female candidate elected	-3.399	-0.327*	-3.361	-0.231
	(5.358)	(0.171)	(5.707)	(0.203)
Bandwidth	0.568	1.000^a	1.000^a	1.000^a
Observations	60	68	64	64
C. Local linear regression with optimal bandwidth $\hat{h}$ and quadratic polynomial in margin of victory				
Female candidate elected	4.433	-0.492	4.888	-0.195
	(10.975)	(0.345)	(8.331)	(0.330)
Bandwidth	0.284	0.547	0.531	1.000^a
Observations	40	60	38	64

Notes: The table shows local linear regression results of having a female candidate elected. Robust standard errors in parentheses, are adjusted for clustering at the district level. Panel A restricts the sample to the optimal bandwidth $\hat{h}$, computed using the Imbens and Kalyanaraman (2012) algorithm. Panel B restricts the sample to $2\hat{h}$, while panel C includes a quadratic polynomial in the margin of victory as control function. ***, ** and * indicate statistical significance at 1%, 5% and 10%, respectively.
[a] Optimal bandwidth greater than 1.

D. Latinobarómetro Survey

The Latinobarómetro Survey is a public opinion survey covering 18 Latin American countries. The sample size for Chile is 1,200 individuals and spans over years 1995–2011, except for year 1999. However, the same questions are not asked during all years. The following table describes the dependent variables included in the regressions of Table 5, with the details of the questions used to create them and the years in which these were available.

Table D-1: Latinobarómetro Survey, description of variables

Variable	Question	Coding	Years available
Marketecon	"The market economy is the only system through which Chile can become developed"	1 if respondent strongly agrees or agrees	1998, 2000, 2002, 2007, 2009, 2010, 2011
Demprob	"Democracy may have problems but it is the best system of government"	1 if respondent strongly disagrees or disagrees	2002–2011
Gobauto	"Under some circumstances, an authoritarian government can be preferable to a democratic one"	1 if respondent chooses this statement instead of "Democracy is preferable to any other kind of government"	1995–2011
Womenopp	"Do you think that women are given as much opportunity as men to earn the same salary?"	1 if respondent answers yes"	1995, 1996, 1997, 2000, 2005, 2006
Womenrole	"It is better that women concentrate at home and men at work"	1 if respondent strongly agrees or agrees	1997, 2000, 2004, 2009
Menleaders	"Men are better political leaders than women"	1 if respondent strongly agrees or agrees	2004, 2009

The key regressor is the variable Woman, that takes the value of 1 if the respondent is female. Additional controls are the respondent's age and age squared, a dummy equal to 1 if the respondent is married or lives with his/her partner, a dummy equal to 1 if the respondent went to college, and city and year dummies.

To determine the set of Center-right voters I follow two strategies. First I restrict the sample to include individuals who place themselves between 6 and 9 in a left-right scale from 0 to 10. Alternatively, I select respondents who would vote for a Center-right party if elections were held the next Sunday. This last procedure reduces the sample of years due to discrepancies in the codes for parties in years 1996, 1997, 1998 and 2001.