Estimating the gender penalty in House of Representative elections using a regression discontinuity design

Anastasopoulos, Lefteris

April 2016

Electoral Studies xxx (2016) 1–8

Contents lists available at ScienceDirect

Electoral Studies

journal homepage: www.elsevier.com/locate/electstud

Estimating the gender penalty in House of Representative elections using a regression discontinuity design

Lefteris Anastasopoulos

UC Berkeley School of Information, 102 S. Hall Road, Berkeley, CA 94704, USA

ARTICLE INFO

Article history:
Received 12 September 2015
Received in revised form
9 November 2015
Accepted 21 April 2016
Available online xxx

ABSTRACT

While the number of female candidates running for office in U.S. House of Representative elections has increased considerably since the 1980s, women continue to account for about only 20% of House members. Whether this gap in female representation can be explained by a gender penalty female candidates face as the result of discrimination on the part of voters or campaign donors remains uncertain. In this paper, I estimate the gender penalty in U.S. House of Representative general elections using a regression discontinuity design (RDD). Using this RDD, I am able to assess whether chance nomination of female candidates to run in the general election affected the amount of campaign funds raised, general election vote share and probability of victory in House elections between 1982 and 2012. I find no evidence of a gender penalty using these measures. These results suggest that the deficit of female representation in the House is more likely the result of barriers to entering politics as opposed to overt gender discrimination by voters and campaign donors.

1. Introduction

After a more than three-fold increase in the number of women in the U.S. House of Representatives since the 100th Congress (1987–89), women continue to account for under 20% of House members. Moreover, there are signs that growth in the number of female representatives is slowing. As the number of women nominated by the major parties since the 1980's surged (Dolan, 2008), the number of females holding office increased from 30 in 1992, referred to as the "Year of the Woman" (Carpini and Fuchs, 1993; Dolan, 1998), to 62 in 2002. Yet the following ten year period saw an increase of only 14 females.

This paper brings a regression discontinuity design to bear on the question of whether the limited growth of female election to the House of Representatives can be explained by a gender penalty faced by female candidates. While over two decades of observational research has attempted to estimate the effect of candidate gender on general election outcomes (Dolan, 2008; Milyo and Schosberg, 2000; Darcy et al., 1985; Kahn, 1992; McDermott, 1997; Herrnson et al., 2003) and campaign contributions (Burrell, 1985; Uhlaner and Schlozman, 1986; Dabelko and Herrnson, 1997; Crespin and Deitz, 2010), this paper is the first to provide causal estimates of the effects of chance selection of female candidates to general election ballots (see Fig. 1).

Many studies have cast doubt on the role of that gender bias plays in limiting the number of female representatives. Using a series of experiments, Brooks (2013) calls the gender penalty into question and finds little evidence that gender stereotypes harm female candidates. In most observational studies exploring the effects of candidate gender on election outcomes and campaign contributions, female candidates raise as much money and win general elections at the same rate as male candidates (Burrell, 1985; Seltzer et al., 1997; Uhlaner and Schlozman, 1986). As Sanbonmatsu (2002) describes, scholars have thus emphasized factors leading to candidate selection and nomination rather than voter dispositions toward female candidates. One limiting factor is the relatively large share of female incumbents (Burrell, 1994). Another is the relatively small number of women in professions that often serve as a springboard to running for local office (Thomas, 1994). Sanbonmatsu (2002) suggests that party recruitment practices and relevant beliefs about women's likelihood of electoral success held by party and interest group leaders helps explain patterns of female participation in electoral politics.

Recent innovative research revives the debate over whether voters' gender bias limits female office-holding (Fulton, 2012). Theoretically, as Fulton points out, it is hard to square null findings from vote-share models with survey evidence that many voters

E-mail address: jasonanastas@ischool.berkeley.edu.

http://dx.doi.org/10.1016/j.electstud.2016.04.008

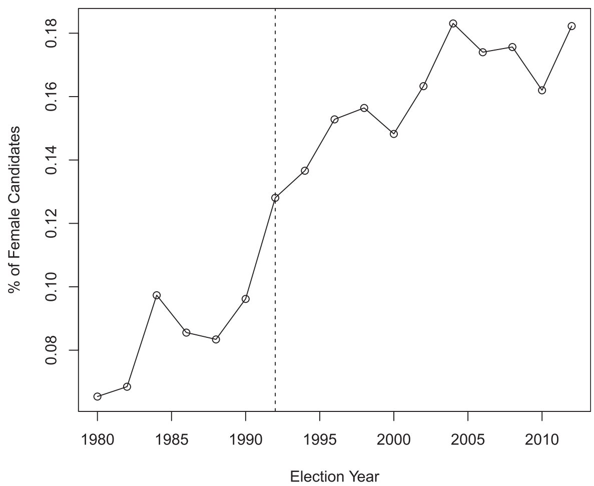

Fig. 1. Proportion of female candidates running in House of Representative general elections, 1980–2012.

hold uncongenial stereotypes about female candidates (Kahn, 1996; Sanbonmatsu and Dolan 2009). Empirically, Fulton takes the critical step of introducing a measure of candidate quality into models of vote share, thus addressing a potential source of omitted variable bias that emerges if the pool of female candidates is on average more electable than the pool of male candidates. Given that women may face a more trying route to the nomination (Lawless and Pearson, 2008; Anzia and Berry, 2011), it is sensible to expect such bias and indeed Fulton shows that once candidate quality is controlled for using her measure, a three percentage point gender penalty emerges. Without this control, no female disadvantage is statistically apparent.

Research exploring female candidates' ability to raise campaign funds have arrived at a number of different conclusions. Uhlaner and Schlozman (1986) find that women tend to raise less money than men on average. Burrell (1985), Dabelko and Herrnson (1997) and Crespin and Deitz (2010) find that while women tend to be better at raising campaign contributions than men, they also have to work harder to secure these contributions. Contradictory findings between more recent and prior research on this topic demonstrates the hazards that omitted variable bias can pose. Fulton's (2012) control for quality, which is based on informants' ratings of 176 men and 24 women incumbents, is subject to potential biases of its own which could emerge if the foundations of assessments of males and females differ. For example, if assessments of male and female candidate quality are differentially responsive to informants' perceptions of how potential voters or donors were responding to these candidates up to the time of the interview, the estimated effect on gender could be biased in either direction. Furthermore, there is no guarantee that other omitted variables related to differences between districts in which male and female candidates happened to run are driving these effects.

2. Estimating the gender penalty in House of Representative elections

A cursory examination of differences between Congressional districts in which female candidates competed for House seats reveals that voters in these districts tended to be younger, more educated, wealthier and more Democratic. Due to a variety of demographic and political factors which contribute to the ability of women to run in House general elections, identifying the effect of candidate gender on election outcomes is a difficult task.

A correctly specified regression discontinuity design (RDD), however, can potentially eliminate these sources of bias. Regression discontinuity designs can provide valid causal estimates with comparatively weak assumptions (Hahn et al., 2001; Lee and Lemieux, 2010; Thistlethwaite and Campbell, 1960). In more recent research, Lee (2008) argues that producing valid causal estimates using RDDs requires only the assumption that agents are unable to manipulate their value of the forcing variable around the cutpoint.

Lee (2008) pioneered the use of vote share as a forcing variable for the purpose of estimating incumbency advantage in House of Representative elections. His work was followed by an explosion of similar political science research. Eggers and Hainmueller (2009) used candidate vote share from members of the Labour and Tory party to estimate the effect of holding office on monetary gain in the UK. In the American context, Gerber and Hopkins (2011) used mayoral vote share to estimate the effect of mayoral partisanship on policy outcomes and more recently Broockman (2014) used party vote share in male-female state legislator elections to estimate the effect of candidate gender on female political participation.[1] While Caughey and Sekhon (2011) cast doubt on studies relying on RDDs which use close general House elections, recent empirical evidence demonstrates that the same pathologies do not necessarily apply to close House primaries (Eggers et al., 2014; Hall, 2015).

2.1. Model and setup

The majority vote share requirement in 2-candidate, partisan House primaries presents an inherent RDD. If a candidate is challenged in their party's primary, their ability to run in the general election becomes a deterministic function of their primary vote share. For two-candidate male/female primaries, the focus of this study, a female candidate will represent her party in the general election if she receives greater than 50% of the vote share.

$$B^f_{ipt} = 1\left[\Pi^f_{ipt} - \Pi^m_{ipt} > 0\right] \tag{1}$$

In Equation (1), Π^f_{ipt} represents primary vote share for the female candidate in two-candidate male/female primaries in district i for party p=[*Democratic*, *Republican*] in year t and Π^m_{ipt} represents primary vote share for the male candidate in the same primary contest. A female represents her party in her district in the general election in year t when $\Pi^f_{ipt} - \Pi^m_{ipt} > 0$. Alternatively, when $\Pi^f_{ipt} - \Pi^m_{ipt} < 0$ in these primaries, the male candidate runs in the general election.

Fig. 2 provides a graphical overview of the RDD used for this analysis. On the x-axis, the running variable is House primary vote share for the female candidate (female candidate vote share minus male candidate vote share in House primaries) and general election outcomes used to measure voting behavior are on the y-axis. From the perspective of the potential outcomes framework (Rubin, 2005) the “treated” group are districts in which female candidates won their party's primary and ran in the general election while the “control” group are districts in which a male candidate ran in the general election.

Formally, I estimate:

[1] For summaries of papers employing regression discontinuity designs using vote share see Caughey and Sekhon (2011, 390–391).

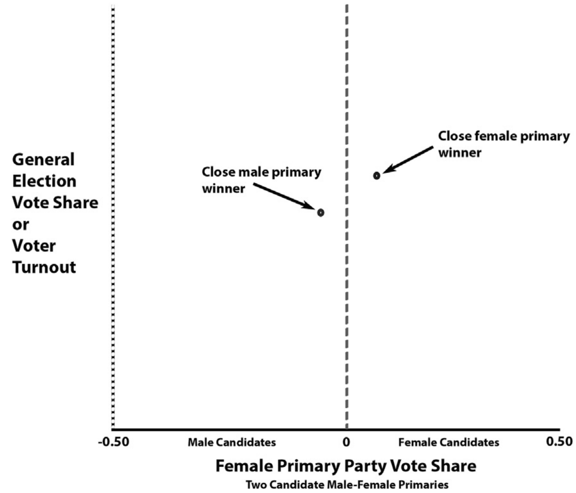

Fig. 2. Regression discontinuity design setup.

$$\tau_{SRD} = \lim_{\Pi^f - \Pi^m \downarrow 0} E\left[V(1)_i \middle| \Pi_i^f - \Pi_i^m = 0\right] - \lim_{\Pi^f - \Pi^m \uparrow 0} E\left[V(0)_i \middle| \Pi_i^f - \Pi_i^m = 0\right] \quad (2)$$

where τ_{SRD} represents the local average treatment effect of chance female nomination on a number of outcomes discussed below.

3. Data

Data containing information about House primary elections between 1982 and 2012 were collected from two sources: (1) the *CQ Elections and Voting Collection* for primary election returns between 1994 and 2012 and; (2) primary election returns provided by David Brady of Stanford University for primaries between 1982 and 1992. Both data sets contain information about candidate name, state, district, primary year, primary vote share and primary vote totals. For competitive two-candidate elections, candidate gender was determined and verified independently by two coders using candidate first name. When candidate gender was ambiguous, Google searches were conducted to learn more about the candidate. If a search revealed no further relevant information, they were counted as missing and that primary was not included in the analysis.

Outcome variables were collected from a variety of sources. Party vote share and votes cast were obtained from the *CQ Elections and Voting Collection* and information about campaign contributions and candidate ideology were obtained from the Database on Ideology, Money in Politics and Elections (Bonica, 2013).[2] Two-candidate partisan primaries included in the analyses below are those in which *only* a female candidate ran against a male candidate. For primaries between 1982 and 2012, there were $N = 429$ such primaries.

Table 1 contains characteristics of male and female winners of close primaries. Female winners of close primaries in the sample are slightly more Republican on average, are more likely to win in the general election and have more contributions from PACs. None of these differences, however, are statistically significant.

[2] http://data.stanford.edu/dime/.

Table 1
Characteristics of male and female close primary winners (±5%).

	Male winners	Female winners	N	T-test p-value
% Democrats	53.33	42.11	49	0.45
% Republicans	46.67	57.89	49	0.45
Vote share (Gen. Elect)	39.10	40.64	49	0.70
P(*Victory*) (Gen. Elect)	0.10	0.21	49	0.29
Avg. contrib. ($)	572,255	551,129	49	0.96
Avg. PAC Contrib. ($)	72,545	120,499	49	0.28
Avg. indv. contrib. ($)	202,850	376,021	49	0.37
Avg. CF-Score	0.008	−0.126	49	0.71

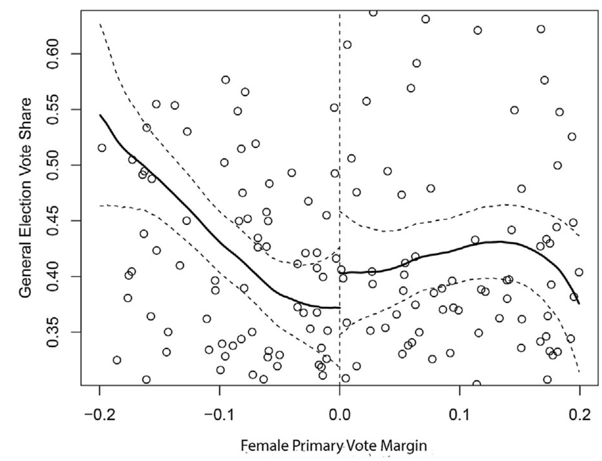

Fig. 3. General election vote share for male v. female winners of close primaries, N = 125. Dotted lines are 95% confidence bands.

4. Do female candidates face voter bias?

I first assess whether female candidates that win nomination to the general election over male candidates by chance face voter bias:

$$V_{ipt} = \alpha + \beta_1 1\left[\Pi_{ipt}^f - \Pi_{ipt}^m > 0\right] + f\left(\Pi_{ipt}^f - \Pi_{ipt}^m\right) + \eta_{ipt} \quad (3)$$

In Equation (3) V_{it} represents two dependent variables: general election party vote share and general election victory. $1\,[\Pi_{ipt}^f - \Pi_{ipt}^m > 0]$ is a female candidate dummy variable equal to one if a female candidate won a male-female House primary and zero if the male candidate won. β_1 is an estimate of the local average treatment effect of chance female nomination on general election vote share and general election victory. $f(\Pi_{ipt}^f - \Pi_{ipt}^m)$ is a function of the forcing variable, female primary vote margin, which takes the form of a non-parametric kernel or pth order polynomial. Triangular kernels are typically used as the default method of estimation in software programs because they are boundary optimal (Cheng et al., 1997; McCrary, 2008). I estimate β_1 for both dependent variables using a triangular kernel as seen in Fig. 3[3] and a second-order polynomial as shown in Fig. 4[4] as seen in Fig. 4 and find no evidence that chance nomination of female candidates

[3] Kernel regression estimates are performed using the **rd** and **rdrobust** packages in Stata.

[4] Second-order polynomials are often used to measure treatment effects in regression discontinuity designs when the data are binned and are also useful for illustration purposes (Imbens and Lemieux, 2008). Research by Gelman and Imbens (2014) finds that higher-order polynomials (above 2) should not be used for estimating treatment effects in regression discontinuity designs because results tend to be very sensitive to choice of the polynomial order.

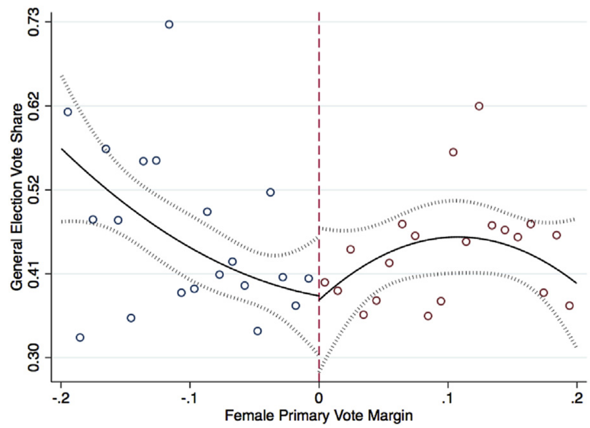

Fig. 4. General election vote share for male v. female winners of close primaries with 1% bins. Dotted lines are 95% confidence bands.

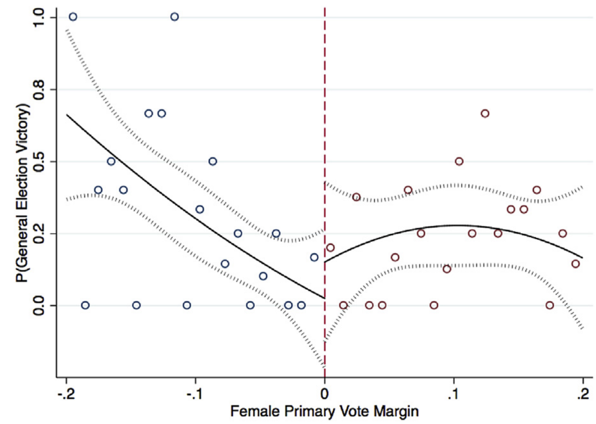

Fig. 5. General election probability of victory for male v. female winners of close primaries, 1% bins. Dotted lines are 95% confidence bands.

results in decreases in party vote share or decreased probability of electoral success.

While there are several methods of estimating local average treatment effects using RDDs, visual evidence of a discontinuity in plots of the forcing variable versus outcomes provide the strongest evidence of an effect (Imbens and Lemieux, 2008). Figs. 3 and 4 are plots of female primary vote margin vs. general election vote share. The first plot displays predictions from non-parametric local regressions using a triangular kernel and the second plot uses average vote share with 1% bins and a quadratic fit. Both provide no evidence that female candidates nominated by chance do worse in general elections. These findings are confirmed by estimates of β_1 presented in Table 2 using the Imbens and Kalyanaraman (2011) optimal bandwidth which minimizes mean-squared error.

Using probability of general election victory as an outcome produces similar results as can be seen in Fig. 5 above and also included in Table 2.

5. Are female candidates financially disadvantaged?

Evidence from the previous section suggests that female candidates nominated to run in the general election by chance do not suffer from voter bias. However, female candidates may still be at a disadvantage during the election process because of individuals or groups that hold discriminatory beliefs about females or their likelihood of electoral success (Crespin and Deitz, 2010; Sanbonmatsu, 2002). If this is true, we would expect female candidates who won the nomination by chance to have fewer resources available to them in the form of campaign contributions from individuals and political action committees. If we find that these female candidates were at a financial disadvantage during their campaign yet still managed to perform on par with men in the general election, this would provide support for Fulton's (2012) claims that female candidates that win nomination by chance are of higher quality than male candidates but suffer from discrimination by contributors.

To assess whether female candidates are at a financial disadvantage during the course of their campaigns, I matched candidates from Bonica's (2013) Database on Ideology, Money and Politics (DIME) to candidates used in the analysis and re-estimated β_1 from Equation (3) using contributions from individuals and contributions from political action committees as outcomes.

As seen in Fig. 6 and 7 and in Table 3 I find no evidence to suggest that female candidates that won nomination by chance received fewer contributions from individuals or from PACs.

6. RDD assumptions

Regression discontinuity designs are generally considered the "gold-standard" among observational studies because they offer the ability to estimate causal treatment effects with few assumptions. Problems with RDDs arise when agents with values of the forcing variable near the cutpoint are able to manipulate selection into treatment (Lee, 2008). Caughey and Sekhon (2011) demonstrated that, in close House elections, incumbents appear to be able

Table 2
Estimates of female candidacy on general election vote share and general election victory at 3 bandwidths and robust nonparametric estimates using Calonico et al., (2014) estimation methods.

Variables	(1) Vote share	(2) Victory	(3) Vote share	(4) Victory
Female Candidate (LATE)	0.001	−0.196	*	*
	(0.053)	(0.255)	*	*
Female Candidate (LATE, Half-BW)	−0.061	−0.398		
	(0.070)	(0.360)	*	*
Female Candidate (LATE, Double-BW)	−0.001	0.096	*	*
	(0.052)	(0.198)	*	*
Female Candidate (LATE, Bias Corrected, Robust)	*	*	−0.271	−0.411
	*	*	(0.229)	(0.557)
IK Bandwidth	0.059	0.036	0.048	0.036
Observations	49	49	45	35

Standard errors in parentheses.
***p < 0.01, **p < 0.05, *p < 0.1.

Please cite this article in press as: Anastasopoulos, L., Estimating the gender penalty in House of Representative elections using a regression discontinuity design, Electoral Studies (2016), http://dx.doi.org/10.1016/j.electstud.2016.04.008

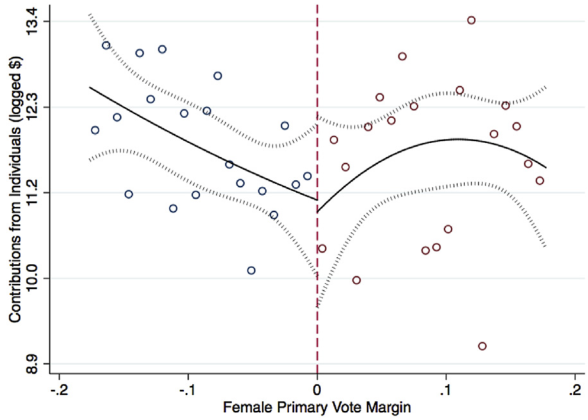

Fig. 6. Log total individual contributions for male and female winners of close primaries, 1% bins. Dotted lines are 95% confidence bands.

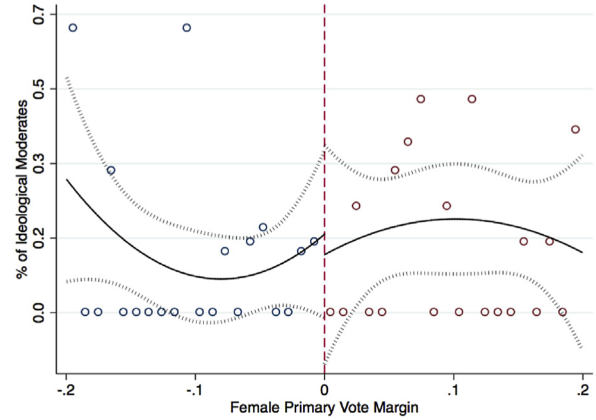

Fig. 8. Proportion of moderate male and female winners of close primaries (1% bins). Dotted lines are 95% confidence bands.

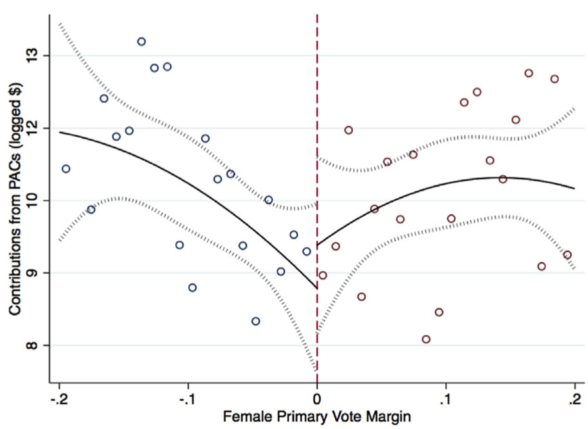

Fig. 7. Log Total PAC contributions for male and female winners of close primaries, 1% bins. Dotted lines are 95% confidence bands.

to manipulate vote share in order to win reelection. As discussed above, close House primaries are very different from close general elections and thus should not automatically be dismissed as viable sources of data for use in regression discontinuity designs.

That being said, unique conceptual issues are present in this regression discontinuity design. These issues apply to similar designs in which individuals on either side of the cut-point are assumed to differ by only one trait, whether that trait be race, gender, ethnicity or ideological extremity. While it is intuitively conceivable that a female candidate can win a House primary contest over a male candidate by chance, evidence of discrimination against female candidates in the past makes the proposition that female and male candidates of the same party are similar except for their gender more difficult to accept.

6.1. *Candidate quality*

The most frequently cited distinction between similar male and female candidates is candidate quality, a factor which Fulton (2012) argues is rooted in gender discrimination. If stereotypes and biases inherent in the political process make it more difficult for women to run against, and beat, males in primaries, then female candidates that barely win against male candidates will presumably be of higher quality, on average, than their male counterparts. This explanation, however, does not square with the findings above. If females that won House primaries by chance against males were of higher quality, we would expect that this quality differential would be reflected in the amount and types of donations that these female candidates received. Overall, I find no evidence that female candidates received differing amounts of campaign donations from any source.

6.2. *Candidate ideology*

Another concern regarding differences between candidates is perceived candidate ideology. Biases and stereotypes among voters toward female candidates of either party might lead them to believe that they are ideologically distinct from male candidates in

Table 3
Estimates of female candidacy on logged PAC and individual contributions at 3 bandwidths and robust nonparametric estimates using Calonico et al. (2014) estimation methods.

Variables	(1)	(2)	(3)	(4)
	PACs	Individuals	PACs	Individuals
Female Candidate (LATE)	−1.146	−0.688	*	*
	(1.590)	(0.988)		
Female Candidate (LATE, Half-BW)	−0.614	−1.098	*	*
	(2.114)	(1.339)	*	*
Female Candidate (LATE, Double-BW)	−0.724	−0.771	*	*
	(1.195)	(0.979)	*	*
Female Candidate (LATE, Bias Corrected, Robust)	*	*	−0.281	5.098
	*	*	(2.790)	(7.565)
IK Bandwidth	0.029	0.055	0.027	0.049
Observations	45	46	26	43

Standard errors in parentheses.
***p < 0.01, **p < 0.05, *p < 0.1.

Table 4
Estimates of female candidacy on % moderate and ideology as measured by cf-scores at 3 bandwidths and robust nonparametric estimates using Calonico et al. (2014) estimation methods.

Variables	(1)	(2)	(3)	(4)
	% Moderate	Ideology	% Moderate	Ideology
Female Candidate (LATE)	−0.284	−0.611	*	*
	(0.240)	(1.028)	*	*
Female Candidate (LATE, Half-BW)	−0.391	−0.639	*	*
	(0.381)	(1.697)	*	*
Female Candidate (LATE, Double-BW)	−0.133	0.244	*	*
	(0.147)	(0.699)	*	*
Female Candidate (LATE, Bias Corrected, Robust)	*	*	−0.561	−0.316
	*	*	(0.550)	(2.120)
IK Bandwidth	0.035	0.025	0.034	0.024
Observations	49	49	33	24

Standard errors in parentheses.
***p < 0.01, **p < 0.05, *p < 0.1.

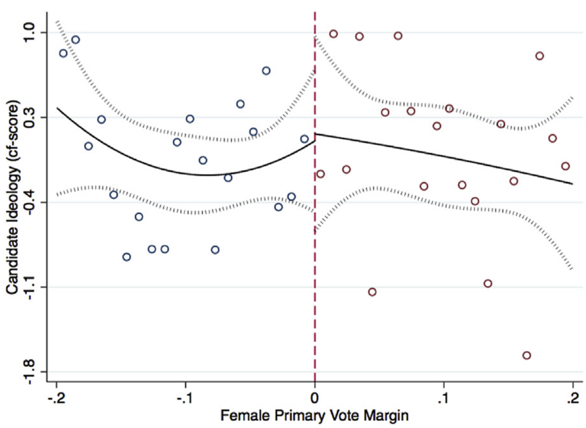

Fig. 9. Candidate ideology (cf-scores) for male and female winners of close primaries (1% bins). Dotted lines are 95% confidence bands.

similar districts. For example, we know from Hall (2014) that moderate candidates have an electoral advantage over extremist candidates. If female candidates that win nominations by chance are perceived as being more moderate that males, this ideological moderation should give female candidates a distinct electoral advantage over male candidates and might explain why we observe positive probabilities of female electoral success in the previous section.

To explore this possibility I identified "moderate" male and female candidates using Bonica's (2013) cf-scores and re-estimated Equation (3) using this measure. Moderates were defined as having a cf-score between the 25th and 75th percentile of the cf-score distribution. Fig. 8 and point estimates in Table 4 clearly show that female candidates are not more likely than male candidates to be moderate. In addition to this, I explored whether candidate ideology differed as measured by the cf-scores themselves and find no difference as shown in Fig. 9.

Finally to address more general concerns about close House primaries, I conducted density tests of the running variable as recommended by McCrary (2008) for female primary vote share and explored covariate balance on a number of general election and primary covariates (see Appendix). Neither suggests that the RDD assumption of non-manipulation in close elections was violated for this subset of primaries.

7. Discussion

Do female candidates face an electoral or financial gender penalty when running in House of Representative general elections? Results from the analyses above seem to suggest that the answer to this question is "no." Over the three decades of this study (1982–2012), there seems to be no evidence that female candidates nominated by chance faced disadvantages in terms of vote share or campaign funding. To the contrary, some results suggest that these female candidates had a slight electoral advantage.

A question that remains, however, is why the proportion of female candidates competing for House of Representative seats has held steady at around 20% despite substantial gains in female representation and the relative absence of a gender penalty. While there may be no single answer to this question, research exploring the career paths of female candidates and party recruitment practices may yield more insights into why gender disparities persist despite the incredible strides that female candidates have made over the past three decades.

Acknowledgements

I would like to thank Morris Levy, Matt Baum, Maya Sen, Ryan Sheely, Tarek Masoud, Sean Gailmard and Jasjeet Sekhon for their helpful comments and suggestions.

Appendix

Table 5
Estimates of Female Candidacy on All Outcome Variables at 3 Bandwidths Plus Robust Nonparametric Estimates Using Calonico, Cattaneo and Titiunik (2014) Estimation Methods (CCT)

Variables	(1)	(2)	(3)	(4)
	Vote share	Victory	PACs ($)	Individual ($)
Female Candidate (LATE)	0.001	−0.196	−1.146	−0.688
	(0.053)	(0.255)	(1.590)	(0.988)
Female Candidate (LATE, Half-BW)	−0.061	−0.398	−0.614	−1.098
	(0.070)	(0.360)	(2.114)	(1.339)
Female Candidate (LATE, Double-BW)	−0.001	0.096	−0.724	−0.771
	(0.052)	(0.198)	(1.195)	(0.979)
Observations	49	49	45	46
Female Candidate (LATE, Bias Corrected, Robust)	−0.271	−0.411	−0.281	5.098
	(0.229)	(0.557)	(2.790)	(7.565)
Observations	45	35	30	33

Standard errors in parentheses.
***p < 0.01, **p < 0.05, *p < 0.1.

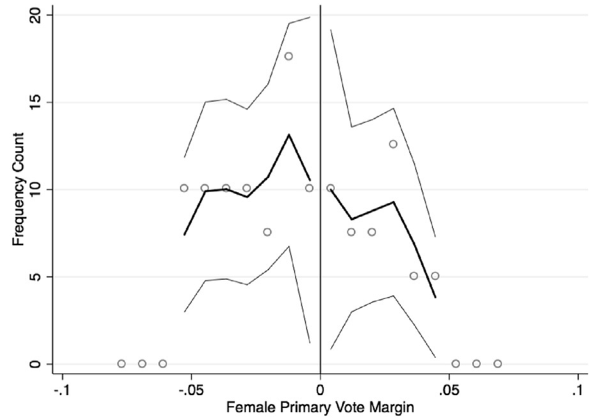

Fig. 10. McCrary (2008) Density test showing no jump in density of the assignment variable (female primary vote margin).

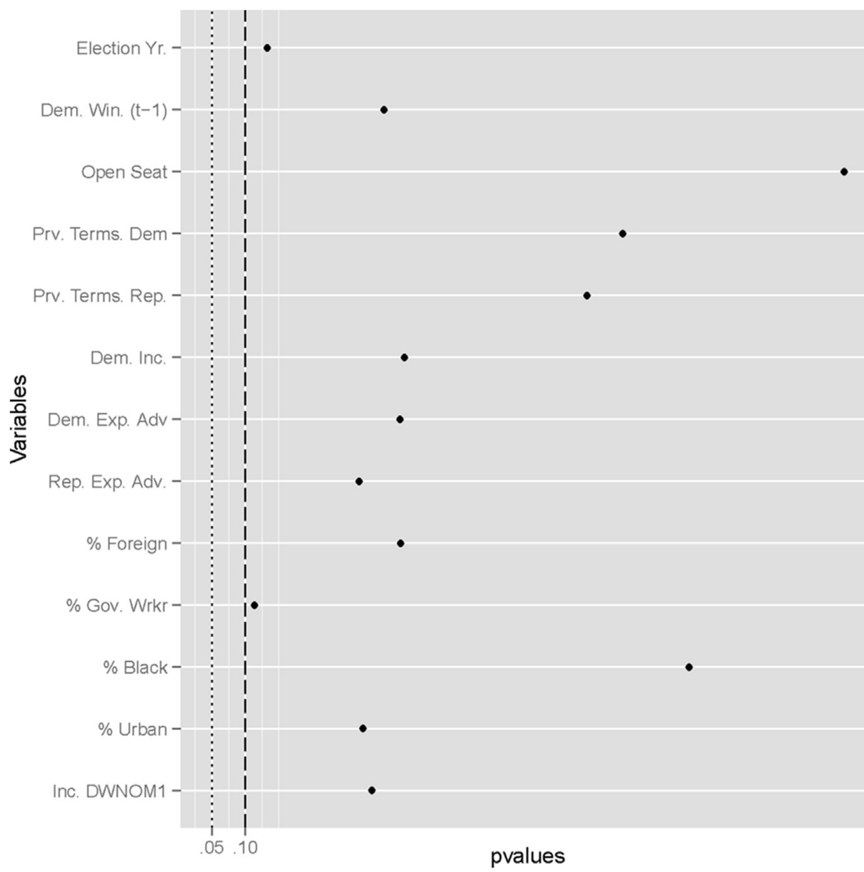

Fig. 11. Balance on covariates for close primaries (2% vote margin)*.
*Several covariates used in this analysis were from data made available by Caughey and Sekhon (2011).

References

Anzia, Sarah F., Berry, Christopher R., 2011. The Jackie (and Jill) Robinson effect: why do congresswomen outperform congressmen? Am. J. Political Sci. 55 (3), 478–493.

Bonica, Adam, 2013. Ideology and interests in the political marketplace. Am. J. Political Sci. 57 (2), 294–311.

Broockman, David E., 2014. Do female politicians empower women to vote or run for office? A regression discontinuity approach. Elect. Stud. 34, 190–204.

Brooks, Deborah Jordan, 2013. He Runs, She Runs: Why Gender Stereotypes Do Not harm Women Candidates. Princeton University Press.

Burrell, Barbara, 1994. A Woman's Place Is in the House. University of Michigan Press, Ann Arbor.

Burrell, Barbara C., 1985. Women's and Men's campaigns for the US house of representatives, 1972-1982 A finance gap? Am. Polit. Res. 13 (3), 251–272.

Calonico, Sebastian, Cattaneo, Matias D., Titiunik, Rocio, 2014. Robust nonparametric confidence intervals for regression-discontinuity designs. Econometrica 82 (6), 2295–2326.

Carpini, Michael X Delli, Fuchs, Ester R., 1993. The year of the woman? Candidates, voters, and the 1992 elections. Political Sci. Q. 29–36.

Caughey, Devin, Sekhon, Jasjeet S., 2011. Elections and the regression discontinuity design: lessons from close us house races, 1942–2008. Polit. Anal. 19 (4), 385–408.

Cheng, M.Y., Fan, J., Marron, J.S., 1997. On automatic boundary corrections. Ann. Stat 25 (4), 1691–1708.

Crespin, Michael H., Dietz, Janna L., 2010. If you can't join 'em, beat 'em: the gender

gap in individual donations to congressional candidates. Political Research Quarterly 63, 581–593.
Dabelko, Kirsten La Cour, Herrnson, Paul S., 1997. Women's and Men's campaigns for the US house of representatives. Political Res. Q. 50 (1), 121–135.
Darcy, Robert, Welch, Susan, Clark, Janet, 1985. Women candidates in single and multi-member districts: American state legislative races. Soc. Sci. Q. 945–953.
Dolan, Kathleen, 1998. Voting for women in the "year of the woman". Am. J. Political Sci. 272–293.
Dolan, Kathleen, 2008. Is there a gender affinity effect in American politics? Information, affect, and candidate sex in US House elections. Political Res. Q. 61 (1), 79–89.
Eggers, A.C., Fowler, A., Hainmueller, J., Hall, A.B., Snyder, J.M., 2014. On the validity of the regression discontinuity design for estimating electoral effects: new evidence from over 40,000 close races. Am. J. Political Sci. 59 (1), 259–274.
Eggers, Andrew C., Hainmueller, Jens, 2009. MPs for sale? Returns to office in postwar British politics. Am. Political Sci. Rev. 103 (04), 513–533.
Fulton, Sarah A., 2012. Running backwards and in high heels the gendered quality gap and incumbent electoral success. Political Res. Q. 65 (2), 303–314.
Gelman, Andrew, Imbens, Guido, 2014. Why high-order polynomials should not be used in regression discontinuity designs. Technical report. National Bureau of Economic Research.
Gerber, Elisabeth R., Hopkins, Daniel J., 2011. When mayors matter: estimating the impact of mayoral partisanship on city policy. Am. J. Political Sci. 55 (2), 326–339.
Hahn, Jinyong, Todd, Petra, Klaauw, Wilbert Van der, 2001. Identification and estimation of treatment effects with a regression-discontinuity design. Econometrica 69 (1), 201–209.
Hall, Andrew B., 2015. What happens when extremists win primaries? Am. Political Sci. Rev. 109 (01), 18–42.
Hall, A.B., 2014. What happens when extremists win primaries? Am. Political Sci. Rev (Forthcoming).
Herrnson, Paul S., Celeste Lay, J., Stokes, Atiya Kai, 2003. Women running as women: candidate gender, campaign issues, and voter-targeting strategies. J. Polit. 65 (1), 244–255.
Imbens, G.W., Lemieux, T., 2008. Regression discontinuity designs: a guide to practice. J. Econom. 142 (2), 615–635.
Imbens, Guido, Kalyanaraman, Karthik, 2011. Optimal bandwidth choice for the regression discontinuity estimator. Rev. Econ. Stud. rdr043.
Kahn, Kim Fridkin, 1992. Does being male help? An investigation of the effects of candidate gender and campaign coverage on evaluations of US Senate candidates. J. Polit. 54 (02), 497–517.
Kahn, Kim Fridkin, 1996. The Political Consequences of Being a Woman.
Lawless, Jennifer L., Pearson, Kathryn, 2008. The primary reason for women's underrepresentation? Reevaluating the conventional wisdom. J. Polit. 70 (01), 67–82.
Lee, David S., 2008. Randomized experiments from non-random selection in US house elections. J. Econ. 142 (2), 675–697.
Lee, David S., Lemieux, Thomas, 2010. Regression discontinuity designs in economics. J. Econ. Lit. 48, 281–355.
McCrary, Justin, 2008. Manipulation of the running variable in the regression discontinuity design: a density test. J. Econ. 142 (2), 698–714.
McDermott, Monika L., 1997. Voting cues in low-information elections: candidate gender as a social information variable in contemporary United States elections. Am. J. Political Sci. 270–283.
Milyo, Jeffrey, Schosberg, Samantha, 2000. Gender bias and selection bias in House elections. Public Choice 105 (1–2), 41–59.
Rubin, D.B., 2005. Causal inference using potential outcomes: design, modeling, decisions. J. Am. Stat. Assoc 100, 322–331.
Sanbonmatsu, Kira, 2002. Gender stereotypes and vote choice. Am. J. Political Sci. 20–34.
Sanbonmatsu, Kira, Dolan, Kathleen, 2009. Do gender stereotypes transcend party? Political Res. Q. 62 (3), 485–494.
Seltzer, Richard, Newman, Jody, Leighton, Melissa Voorhees, 1997. Sex as a Political Variable: Women as Candidates and Voters in US Elections. Lynne Rienner Publishers.
Thistlethwaite, Donald L., Campbell, Donald T., 1960. Regression-discontinuity analysis: an alternative to the ex post facto experiment. J. Educ. Psychol. 51 (6), 309.
Thomas, Sue, 1994. How Women Legislate. Oxford University Press.
Uhlaner, Carole Jean, Schlozman, Kay Lehman, 1986. Candidate gender and congressional campaign receipts. J. Polit. 48 (01), 30–50.