The Impact of Public Election Funding on Women Candidates: Comparative Evidence From State Elections*

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Harvard University President Larry Summers may have found himself in hot water for even suggesting that there are innate differences between men and women, but an entire literature on women in politics has grown around the question of whether women bring unique perspectives, goals, and talents to legislatures and executive offices (even if no one is quite sure just what, precisely, those perspectives, goals and talents are, or whether they are a direct function of gender or can be attributed to other factors, such as political experience or party affiliation). Although one thread of this literature holds that portraying women as a cohesive group is too limited a conception of the role that gender plays in the political process, there is little doubt that women, as a group, show distinctive behavioral characteristics once in office.[1]

Extensive evidence has documented that women legislators, on average, show distinctive views on key issues. In their review essay, Cammisa and Reingold note that

> Female legislators have been found to be more likely than their male colleagues to represent women's interests in two ways. First, state legislators' policy preferences and roll call votes tend to reflect (and, in some cases, magnify) gender gaps in public opinion. Female legislators are more likely than male legislators to take liberal positions on a wide array of issues, such as gun control, social welfare, civil rights, environmental protection, and public health and safety [citations omitted]. Second, even though public opinion on many women's rights issues (e.g., abortion and the Equal Rights Amendment) is not split along gender lines, women in state legislatures are more likely than men to lend their support to the feminist side of such issues [citations omitted]. Together, these studies cover every state and, in almost every instance, the sex differences they report withstand controls for party and district-level factors.[2]

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[1] For a useful review of the subfield, see Anne Marie Cammisa and Beth Reingold, "Women in State Legislatures and State Legislative Research: Beyond Sameness and Difference," *State Politics and Policy Quarterly*, 4:181-210 (No. 2, Summer 2004), and Sue Thomas, *How Women Legislate* (New York: Oxford University Press, 1994). Cammisa and Reingold cite, among other results, findings that women enter politics through distinct career tracks, emphasize different issues once the are in office, show more concern with constituent relations.

[2] Cammisa and Reingold, "Women in State Legislatures and State Legislative Research," 191-2.

Women also tend to adopt distinctive campaign strategies, particularly in the content and style of television advertising. They also face gender-specific expectations about how they should behave during the campaign.[3]

Whether or not these differences extend to the ability to raise campaign funds is something of an open question. On the one hand, women do not appear to face gender-specific barriers. The most detailed investigation of gender and state legislative campaign funding concluded that "there are no significant differences" in the ability of male and female candidates – whether they are incumbents, challengers, or running in open seats – "in their ability to attract campaign resources."[4] In fact, the authors of that study found that in many instances women raised *more* money than their male counterparts, especially in less professionalized "citizen legislatures."[5] These findings are consistent with other data showing that, on the whole, women raise the same amount as men in similar circumstances (open, incumbent, challenger), although this parity is offset by the fact that men who face women raise *more* than men who are running against other men. There is, then, a possibility that women as candidates motivate their opponents.[6]

On the other hand, women, despite comprising over 50 percent of the electorate, still make up only a minority of state legislators, and tend to hold far less than a proportional share

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[3] Virginia Sapiro and Katherine Cramer-Walsh, "Doing Gender in Congressional Campaign Advertising," Paper delivered at the Annual Meeting of the International Society for Political Psychology, Berlin, July 2002.
[4] Joel A. Thompson, Gary F. Moncrief, and Keth E. Hamm, "Gender, Candidate Attributes, and Campaign Contributions," in Joel A. Thompson and Gary F. Moncrief, eds., *Campaign Finance in State Legislative Elections* (Washington, DC: CQ Press, 1998), 130.
[5] Thompson, Moncrief, and Hamm, 135.
[6] Brian L. Werner, "Financing the Campaigns of Women Candidates and Their Opponents: Evidence From Three States, 1982-1990," *Women & Politics* 18:81-98 (No. 1, 1997).

of local political offices as well.[7] This means that women benefit less from the clear advantage of incumbency – since by definition, winning a higher proportion of seats requires knocking off male incumbents[8] – and that women are less likely to have the sort of political experience that makes them credible candidates.[9] As Duerst-Lahti put it in 1998, "Political scientists know that the incumbency advantage is a primary explanation for the paucity of women in office. . . Women have a tougher time winning elections not because they are women, but because they are not incumbents."[10] At the same time, there may well be some gender-specific affects to the initial decision to run for office. A survey study of an "eligibility pool" of prominent lawyers, businesspeople, educators – that is, people whose backgrounds suggested that they would be make credible political candidates – found that women were less likely to take seriously the idea of running for office.[11]

The consensus position is that women are less likely to come forward and run for office in the first place. Once they do make this decision, they raise as much money as their male counterparts. How campaign money figures into the initial decision is hard to capture, since by definition we do not "observe" a decision not to run for office and cannot know how many people in the pool of potential candidates – women or men – choose to forego a campaign. We

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[7] According to the Center for American Women in politics, women currently hold 15 percent of seats in Congress, 25 percent of statewide elected offices, 22 percent of state legislative seats, and roughly 14 percent of mayoral offices in cities with populations greater than 30,000. Center for American Women in Politics, *Women in Elective Office, 2005* (http://www.cawp.rutgers.edu/Facts/Officeholders/elective.pdf, accessed February 28, 2005).

[8] Or alternatively, to wait until an incumbent retires and then run for the open seat.

[9] Experience in elected office is a crucial element of candidate quality; such experience gives candidates campaign experience, an existing political base, and name recognition. For the classic account of these benefits, see Gary C. Jacobson and Samuel Kernell, *Strategy and Choice in Congressional Elections* (New Haven: Yale University Press, 1981).

[10] Georgia Duerst Lahti, "The Bottleneck: Women Becoming Candidates," in *Women and Elective Office: Past, Present, and Future*, Sue Thomas and Clyde Wilcox, eds. (New York: Oxford University Press, 1998), 15.

[11] Richard L. Fox, "Gender, Political Ambition, and the Decision Not to Run for Office," undated manuscript Center for Women and American Politics, http://www.cawp.rutgers.edu/Research/Reports/Fox2003.pdf

do not really know, therefore, whether women and similarly situated men use a different metric in coming forward.

But public funding programs may give us a toehold on the question. Does public election funding play increase women's ability to run for office? One justification for public election funding is that it can increase the diversity of the candidate pool, by eliminating a hurdle that bars those without easy access to campaign funds from running for office.[12] In theory, public funding provides a way out of the competitive catch-22 that challengers find themselves in: they cannot run competitive campaigns unless they can raise at least a threshold amount of campaign money, but they cannot raise campaign money unless they are perceived to be competitive. Since there are fewer potential female candidates with political experience than potential male candidates, women start off with a disadvantage that public funding could lessen.

This prospect suggests that additional work can shed light on the relationship between barriers to entry and gender. Public funding, in which candidates receive state subsidies for their campaign in return for agreeing to spending limits, is widely credited with increasing the diversity of the candidate pools. Access to public money gives candidates an alternative to private fundraising and allows those not connected to existing political networks or parties to organize a run more easily. This effect is not limited to women but applies to all challengers, who receive a larger percentage of their campaign funds from public accounts than do incumbents.

We assess the impact of public funding in the five states with public funding programs for state legislative elections – Arizona, Hawaii, Maine, Minnesota, and Wisconsin. Although

[12] Jamin Raskin and John Bonifaz call this a "wealth primary," arguing that a candidate's wealth has become an additional (and unconstitutional) qualification for public office. "[The] tyranny of private money corrupts the democratic relationship of one person/one vote by making it exceedingly difficult for poor or middle-class persons to run for office." Jamin Raskin and John Bonifaz, "Equal Protection and the Wealth Primary," *Yale Law and Policy Review* 11:273-332 (1993), 277.

we cannot make any claim that these states are representative of any national trends (indeed, quite the opposite is true, since these states are distinguished by merely the existence of legislative public funding, which by definition puts them in an unrepresentative class), we can claim that they present an opportunity to assess the impact of public funding in a variety of contexts. These legislatures range from the non-professional citizen type (Maine), to the middle range (Arizona and Hawaii), to the highly professionalized (Minnesota and Wisconsin). Two of the programs have been relatively stable over several decades (Minnesota and Wisconsin), two others went from no programs at all to "clean elections" programs with full public funding (Arizona and Maine), and Hawaii moved from a providing trivial grants ($50) to one in which candidates can receive small, but noticeable, grants amounting to a few thousand dollars. The states span a broad geographic and demographic range, from the Northeast to the Midwest, Southwest, and Pacific.

If the conventional wisdom about public funding is correct, we would expect that public funding should increase the number of women in the candidate pool, and also (though this effect might not be as strong) increase the success of women candidates. The reasoning is straightforward: if women have a tougher time raising money and establishing credible candidacies (which are, in most cases, part of the same process), then public funding should eliminate one key barrier and increase the probability of women emerging as candidates, as well as give women a better chance of winning.[13]

[13] It is possible that the size of the candidate pool and victory probabilities work against each other. If public funding lowers barriers to entry, then lower quality candidates – those with less experience, or poorer political skills – might decide to run. Presumably, these less skilled candidates would do worse at the polls.

Alternatively, if women have already reached campaign finance parity, we would expect to see less of an effect. With parity, women would face the same barriers as men, with gender itself having little to do with decisions to run or accept public funding, or outcomes.

Aggregate Evidence

Based on initial accounts of the public funding programs, particularly in Maine and Arizona, we expected to see a surge in both the number of women who get elected, and also in the number who run. But that has not happened.[14] We show the direct impact on the number of women running and winning in figures 1 and 2. In figure 1a and 1b, we graph the makeup of the candidate pool, by showing the percent of women running on the general election ballot in each of the five states, in both the lower chamber and the Senate.[15] The key points in the graph are 2000 for Maine and Arizona (the first cycle in which public funding was available) and 1998 for Hawaii. There are few strong trends in these graphs. In figure 1a (lower chamber), the percentage of candidates who are women remain stable in Minnesota and Maine, drop substantially in Arizona from 2000 to 2002, and show stochastic variation in Wisconsin. The trend for Hawaii is in the expected direction, as the percentage of women candidates has risen from about 20 percent in 1998 to nearly 30 percent in 2002.

Similar patterns emerge from figure 1b, which shows the results for the state Senates. Again, the only state with a consistent trend in the expected direction is Hawaii. Even here, though, the effect appears to be small. Despite an increase in the percentage of women candidates from 1998 to 2002 (from about 20 percent to over 30 percent), the 2002 percentage is lower than the 1994 percentage.

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[14] The Clean Elections Institute, an advocacy group that supports public funding, claims that the Arizona program did benefit women and minorities. See *The Road to Victory: Clean Elections Shapes 2002 Arizona Elections* (December 2002), accessed at http://www.azclean.org/documents/ 2002RoadtoVictory-Final.pdf

[15] The numbers change slightly if we include primary elections, but the overall trends are the same.

Finally, in figure 2a and 2b, we show the percentage of legislative seats held by women in both chambers.

From a macro perspective, there is little evidence that public funding has increased representation of women in the state legislature or the number of women who run for office. Post-1998 results for Maine and Arizona – that is, after full public funding was adopted – show no consistent pattern, and in some cases even run in the opposite direction. The results from Hawaii, while intriguing, are probably not an effect of public funding, since only a handful of candidates accept it, and of these, only a few are women.[16]

Individual Level Analysis

We extend the analysis of overall effects to examine individual level decision making by candidates. We are interested in an empirical question that may be stated easily, but which has a number of interesting ramifications: are women more likely to accept public election funding than men? We are especially interested in the question of whether gender has an independent effect on this decision, controlling for other factors that may play a role, such as party identification, political experience, and incumbency.

We collected the data for this project from various state-level campaign finance and election commissions over the past year. We gathered the data for Maine and Arizona online, and we visited Hawaii, Minnesota, and Wisconsin in person to access the relevant filings. Although our database extends back to the early 1990s, we limit our analysis here to the 2000 and 2002 election cycles since Maine and Arizona did not have public funding programs prior to these elections. Further, we restricted our analysis to major-party general election candidates in contested races. We instituted these restrictions for several reasons. First, several of the

[16] In 2002, for example, only 14 candidates (out of 203) for the state legislature accepted public funding, and of these, only 3 were women.

states do not have public funding for primary elections. Second, we eliminated third-party candidates since they consisted of a small fraction of the candidates and even smaller proportion of the winners—third party candidates won eight out of the 1,325 races under analysis or 0.6 percent. Finally, the electoral dynamics of contested versus uncontested races are drastically different, particularly in regard to campaign finance.

Figures 3 and 4 provide a revealing portrait of individual level participation in public funding programs. These pie charts break down participation by state, election cycle, and gender. As they demonstrate, there is significant cross-state and cross-gender variation in the decision to take public funding. We consider the cross-state differences an epiphenomenon of the differences in program structure and the actual amount of funding provided to candidates. Appendix A provides a table that details the differences between the five states' programs during the 2000-2002 election cycles. Just a quick glance at the pie charts indicates that participation is much higher in the three states that provide full public funding (Arizona and Maine) or near-full public funding (Minnesota). Although some trends are more visible in one chamber, we believe these charts reinforce our decision to analyze public funding at the level of the individual candidate.

To analyze the impact of different factors that we expected to affect the decision to take public funding, we estimated separate logistic regression models for the lower- and upper-chambers, pooling candidates across states and cycles. Since several of the candidates are repeat observations in 2000 and 2002 (either incumbents or repeat candidates who are challengers or running in an open seat), we computed our regressions using robust standard errors that clustered on a candidate unique identifier indicator. This precaution addresses

concerns of non-independence across observations and provides for accurate estimates of the regression coefficient's standard errors.

Our dependent variable is a binary indicator coded (1) for candidates who accepted public funding for their general election campaign and (0) for those who did not accept public funding. In 2000, 55.2 percent of major-party general election candidates accepted public funding; in 2002, the level of participation increased to 65.8 percent. Over the two-cycle period under study 60.7 percent of major-party general election candidates in these five states accepted public funding.

Our independent variable of interest is a dummy variable for gender. We coded female candidates (1) and male candidates (0). Since our expectation is that female candidates are *ceteris paribus* more likely than male candidates to accept public funding, we expect the coefficient for this variable to be positively signed across chambers.

To control for candidate experience, first we employ a series of dummy variables related to incumbency status. We include in our model, the dummies for incumbents and open-seats, omitting the dummy for challengers. Our expectation is that incumbents and candidates in open-seats will both be less likely than challengers to take public funding. Incumbents will have less of a need since they are likely to have established and experienced campaign organizations and are also likely to have more access to private donors, in contrast to challengers who are almost certain to be at a disadvantage financially to the incumbent they are trying to unseat. Since public financing programs provide money in exchange for a candidate agreeing to a spending limit and since open-seats are often the most toughly fought and expensive elections, we expect fewer open-seat candidates than challengers to opt-in to public financing programs.

Second, we also control for candidate experience by including a dummy variable to capture whether (1) or not (0) the candidate had previously run for the office at stake. For incumbents, this variable will naturally be coded 1. The real leverage for such an indicator comes from what it tells us about challengers. To take full advantage of this information, we calculated an interaction term by multiplying this variable and the challenger dummy, to see if previous competitive experience has a particular effect on challengers. Our expectation for the previous run variable and the interaction term between previous run and challenger is that their coefficients will be negative, as most candidates running for the second time will be more established politically and thus have more access to private sources of campaign cash.[17]

Finally, in our Senate models, we included a dummy variable for members of the lower chamber who are seeking elevation to the Senate. We expect that these candidates will behave much like incumbents and will thus shun public funding in favor of private financing.

Democratic candidates are widely perceived to be both more supportive of the institution of public financing and more likely to take public funding when it is available. This perception stems from the broader array of differences between the two major parties on issues of campaign finance. Thus, we expect a positive effect on the likelihood of taking public funding for Democratic candidates.

As the pie charts above illustrate, we clearly need to account for cross-state differences. To do so, we employ of series of dummies for candidates in Arizona, Hawaii, Maine, and Minnesota. We omitted Wisconsin and made it our base category. We expect these dummies to soke-up the differences between the different state programs and their effects on the likelihood of a candidate in that state taking public funding. Since Wisconsin's fairly

[17] Alternatively, this indicator may pick-up so-called "vanity candidates" who run for the same office repeatedly, despite having no real chance to win the election—and such candidates may be likely to "abuse" public funding programs.

inconsequential public funding program is our baseline, we expect the coefficients for Arizona, Maine, and Minnesota to all be positive. Only in Hawaii, which gives an amount, in terms of overall spending, that is as comparably meager as Wisconsin's, do we expect the effect on the likelihood of taking public funding to be negative or statistically indistinguishable.

All of our variables, their summary statistics, and their expected effects are listed in tables 1 and 2.

Additionally, we estimated a second set of logistic regression models that included information regarding each candidate's party's previous performance in the district. To do so, we had to eliminate several cases from our data set. In Arizona, Hawaii, Minnesota, and Wisconsin, state legislative redistricting occurred between 2000 and 2002. Thus, information on a party's previous performance in open seat districts in 2002 in these four states is unavailable. (In non-open seats, we calculated previous two-party performance from the incumbent candidate's previous share of the two-party vote. Although this measure remains imperfect, we believe it is preferable to eliminating all but the Maine candidates from the 2002 portion of the data set.) In our upper chamber models, our *N* actually remained unchanged at 536; in our lower chamber models, our *N* decreased dramatically from 1,440 to 496. The actual variable is the absolute value of 50 minus the candidate's party's share of the two-party vote percentage in the previous election. We expect the likelihood of a candidate taking public funds to increase as this amount increases—that is as the party's distance from victory grows in either direction. We believe that candidates in tight races do not want to limit their spending (a requirement for accepting public funding), and thus, in those races where the difference between the two parties is minimal to rely on private financing.

Results

Our first model for the lower chambers provides evidence for an independent effect for gender on the decision to take public funding. As table 3 shows, after controlling for candidate experience, partisanship, and program characteristics we found significant evidence ($\alpha = 0.01$) that woman are much more likely than men to participate in public funding programs. This finding is consistent with previous research that suggests that female candidates follow different tracks than male candidates in the electoral process, and extends this "exceptionalism" argument deeper into the campaign finance literature. Further, it provides support for the conventional wisdom that women are more likely to accept public funding (at least in races for the lower chamber) than men.

We illustrate the independent effect of gender by plotting the predicted probabilities of taking public funding for men and women separately versus their party's prior performance in the district in figure 5. The independent effect of gender in this particular illustration is to increase the probability of a candidate taking public funds by roughly 15 percent.

The overall fit of the first house model is strong as well. The Wald χ^2 and pseudo-r^2 calculations reject the null-hypothesis that the coefficients are jointly equally to zero and show that the model explains approximately 44 percent of the variation in the decision to accept public funding, respectively. Of our control variables that account for candidate quality, only the dummy for incumbents is significant, and it has a negative sign as expected, indicating that incumbents are less likely than challengers to accept public funding. In terms of partisanship, Democrats are significantly more likely to participate in public funding programs than Republicans. No systematic difference seems to exist between candidates running in Maine or Arizona and Wisconsin. This finding is surprising given how generous the Maine and Arizona

programs are compared to Wisconsin's. However, the lack of an effect may be due to a learning curve. The two elections in our dataset are the first two in which Arizona and Maine provided public funding; perhaps candidates need more time to adapt to such a change in the electoral environment before their behavior better accords to rational expectations. For instance, candidates in Minnesota and Hawaii, which have both had public funding programs since the 1970s, exhibit behavior that appears to be much more systematic and rational.

When we included the previous difference between the two parties in the district (and the accompanying interaction terms), we ran into collinearity problems. As a result, the dummies for Maine and Minnesota were dropped from the model. Nonetheless, the overall fit of the model still appears to be good, but not as good as model one's fit. Unfortunately, we cannot compare the two models using a likelihood-ratio test because of the difference in the number of cases between them. We had to drop almost 1,000 cases to estimate model two.

Nonetheless, the independent effect of gender remains. In fact, the coefficient on female approximately doubled in size, even if its level of significance declined slightly. We believe that our findings regarding the independent effect of gender on the decision to take public funding is strengthened by its continuing presence after we introduced further controls to take into account district-level party strength. The other effects we saw in model one still exist: Democrats remain more likely than Republicans and candidates in Hawaii are less likely than candidates in Wisconsin to accept public funding. None of the three variables dealing with the previous electoral difference between the parties in the district even approaches conventional levels of statistical significance.

The overall fit of our two upper chamber models (one without prior party performance, etc. and one without) is as strong as the fit of our lower chamber models. In both

specifications, we can reject the null-hypothesis that the coefficients are jointly equal to zero and both models have pseudo-r^2s of approximately 0.45.

However, the independent effect of gender disappears in our state senate models (tables 5 and 6); gender does not have a statistically significant effect in either of our two specifications. We believe this difference between the chambers reflects aspects of candidate quality that we are unable to capture despite the multiple indicators we employed. Our belief is that candidates for the state senate are likely to be more prominent individuals, regardless of gender, who are more active in partisan politics. This type of experience would likely lead to greater access to private financing channels. Additionally, candidates for the state senate may also face more competitive and more costly races, two factors that are likely to further lower the probability that they will accept public funding.

The effects of the candidate experience variables we included are different in the upper chamber models. First, neither the dummy for challenger nor the dummy for incumbent are significant, though they are correctly signed across both models. Second, our indicator that measures whether or not the candidate served in the lower chamber is significant and signed as expected. Senate candidates that previously served in the lower chamber are significantly less likely than those that did not serve to accept public funding. This finding squares with the general thrust of the public funding literature: higher quality candidates are less likely to accept (or rely) on public funds than lesser quality candidates. Although, we would have expected this effect to show up in the incumbent/challenger/open series of dummies as well, we are not surprised that previous service in the lower chamber of the legislature has an independent and negative effect on the probability of taking public funding.

The rest of the effects in the upper chamber models are largely consistent with those found in the lower chamber models. Partisanship remains a strong driver in the choice to take or reject public funding, with senate Democratic candidates being far more likely than Republican candidates to accept public funding. Figure 6 demonstrates the effect of partisanship on the probability of a candidate taking public funds. The effects of the differences in the states' programs are also the same: candidates in Maine and Minnesota are more likely to take public funding than candidates in Wisconsin, while candidates in Hawaii are significantly less likely. Senate candidates in Arizona are no more likely than those in Wisconsin to take public funding. These findings hold up in both specifications.

Like the second lower chamber model, in the second upper chamber model, we included the variables related to the prior two-party vote breakdown in the district. Unlike, the lower chamber model however, in the upper chamber model, one of the variables (the previous difference and incumbent interaction) is approaching statistical significance ($z = -1.93$). Interestingly, this suggests that as the margin between the two parties grows an incumbent is *less* likely to accept public funding. We suspect two causal mechanisms may be behind this finding. First, incumbents may perceive that being a so-called "clean" candidate may help them in an election if they represent a marginal district. Alternatively, incumbents may reject public funding more readily if their district is safe (i.e. the previous two-party difference is greater), and they have little need for campaign funds whether they be public or private.

Conclusion

Our analysis suggests that female candidates are substantially more likely than male candidates to accept public funding in races for the lower chamber of state assemblies. This finding holds even after controlling for a variety of partisan and candidate experience variables;

thus, we believe that gender does have an independent effect on the probability of a candidate participating in a public funding program. Further, we believe that this finding supports those who claim that public funding serves to broaden the pool of candidates willing to run for office. Although we cannot make that inferential claim with the evidence presented here, it is not hard to read our findings as evidence for such an argument.

Additionally, this work supports in whole or part various existing claims in the campaign finance literature (or in conventional wisdom). First, the effect of partisanship is clear and distinct across all of our models: Democrats are more likely than Republicans to take public funding. Second, candidate experience plays a major role in the decision: incumbents are less likely than candidates in open seats to take public funding and candidates for the state senate who previously served in the lower chamber are also less likely. Finally, the nature of the public funding program is also crucial in determining participation. A clear trend implies that candidates who run in states with more generous (or full) public financing programs are much more likely to participate than those running in states with relatively trivial funding programs.

Our most intriguing finding remains the sign on the interaction term between incumbency and previous two-party vote margin in the state senate model. This variable behaves in the exact opposite manner of what we expected and such a finding calls for a further exploration as to why this is.

We believe that this paper contributes much needed empirical evidence to support claims that are often made regarding participation in public funding. Most importantly, it confirms the presence of a participation gender gap, at least in elections to the lower house.

Investigating this inter-chamber difference as well as the effect of participation on candidate success are two avenues of future inquiry that stem from these results.

Figure 1a: Number of Women Candidates by State, 1990-2002 – Lower Chamber

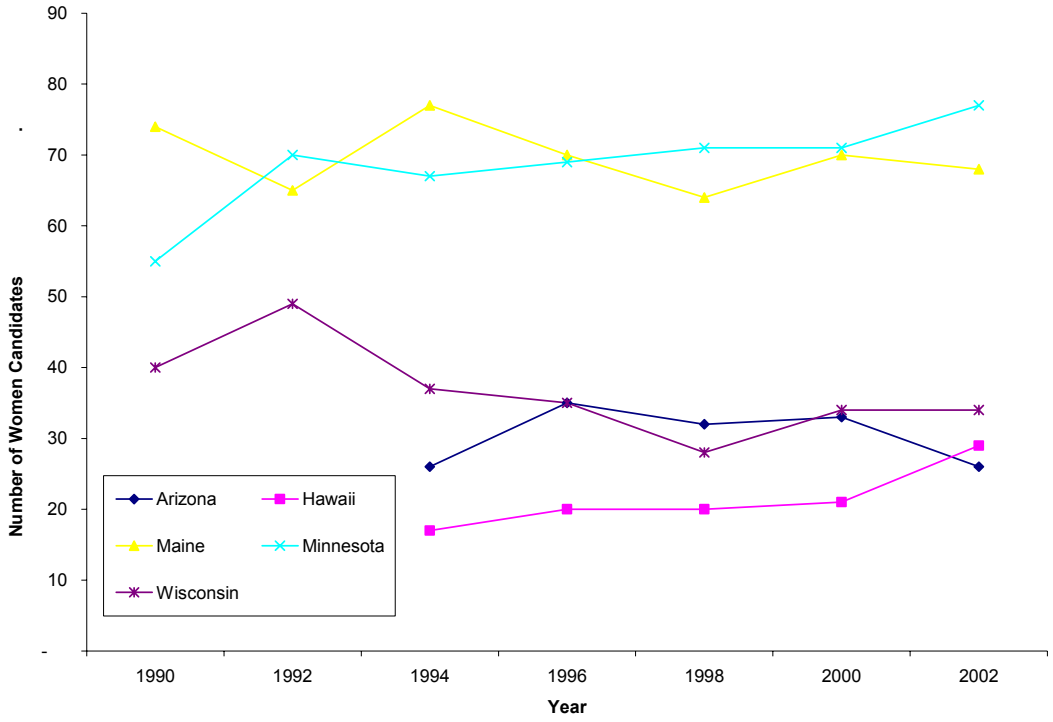


Figure 1b: Number of Women Candidates by State, 1990-2002 – Upper Chamber

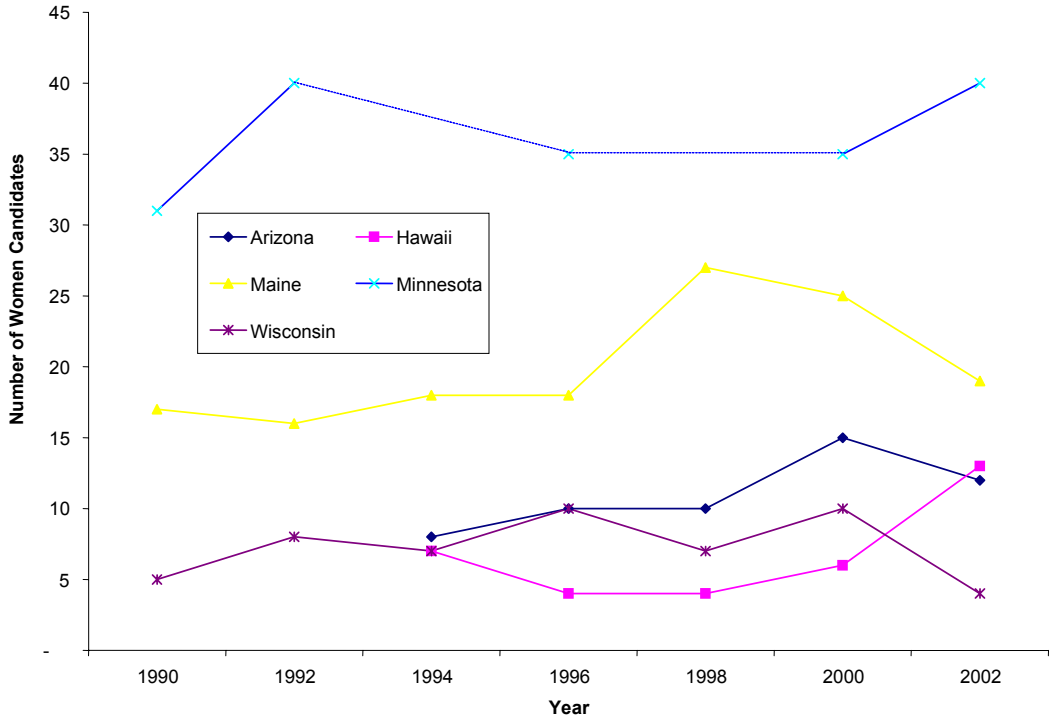


Figure 2a: Percent Women Members by State, 1991-2003 – Lower Chamber

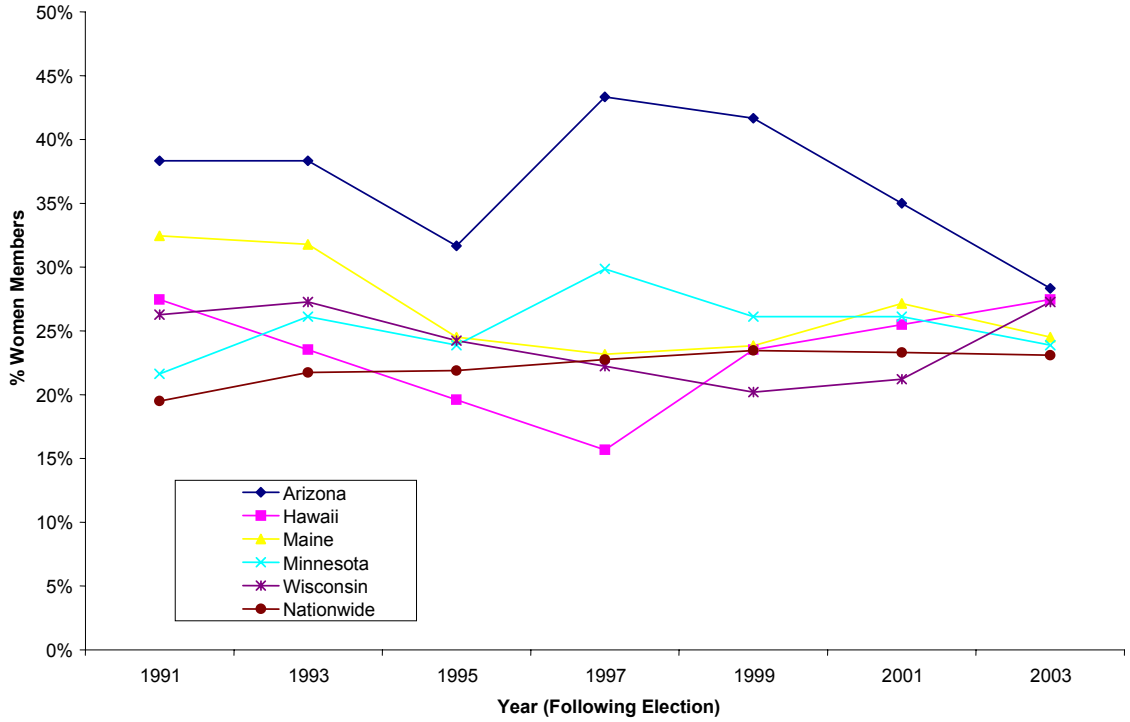


Figure 2b: Percent Women Members by State, 1991-2003 – Upper Chamber

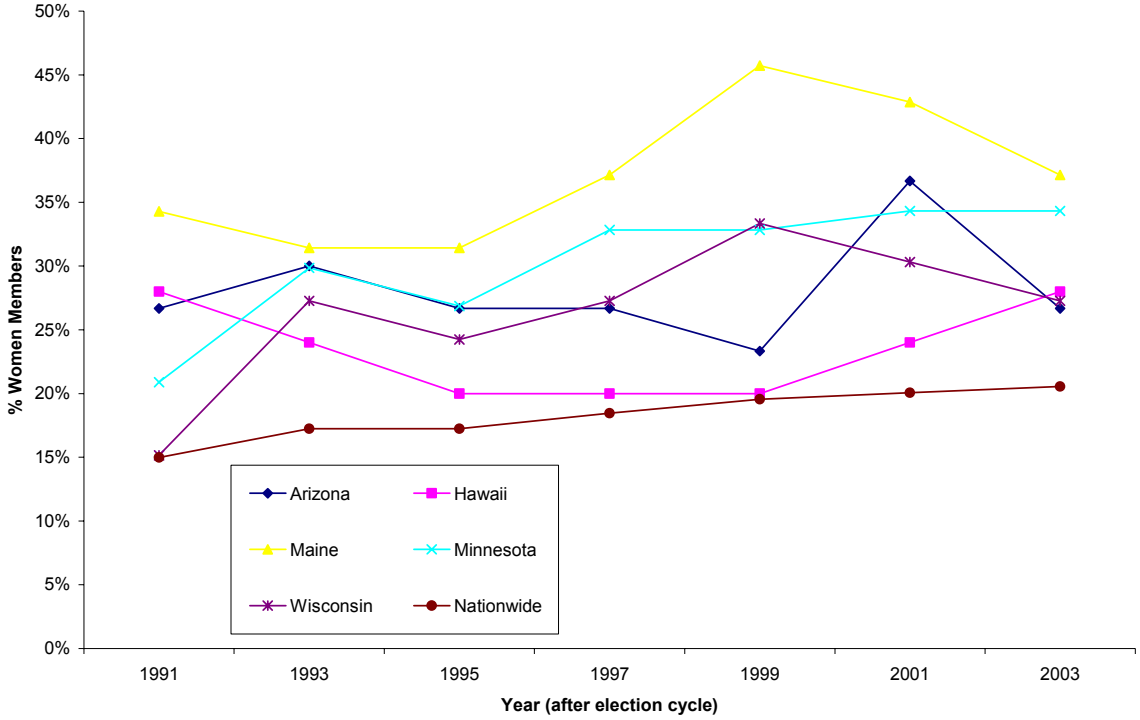


Figure 3: Proportion of Major-Party General Election Candidates Accepting Public Funds by Gender, Cycle, and State – Lower Chamber

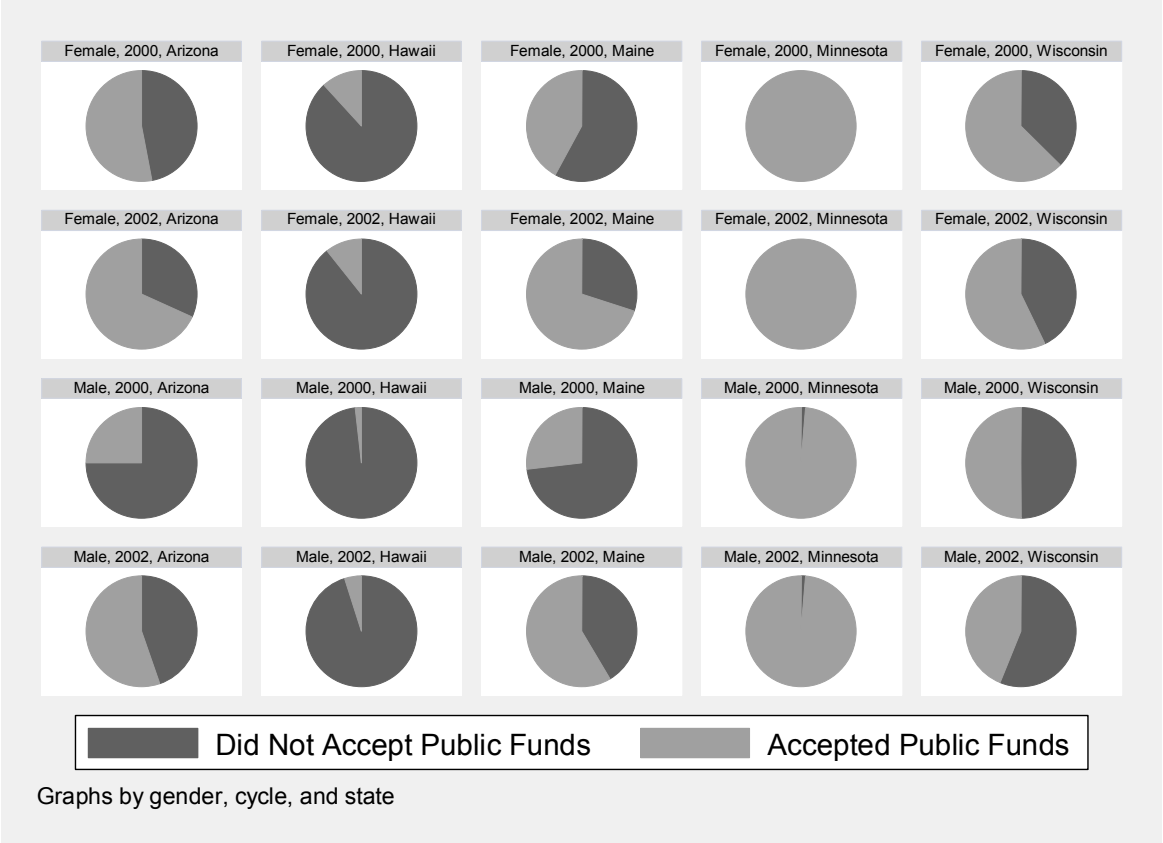


Figure 4: Proportion of Major-Party General Election Candidates Accepting Public Funds by Gender, Cycle, and State – Upper Chamber

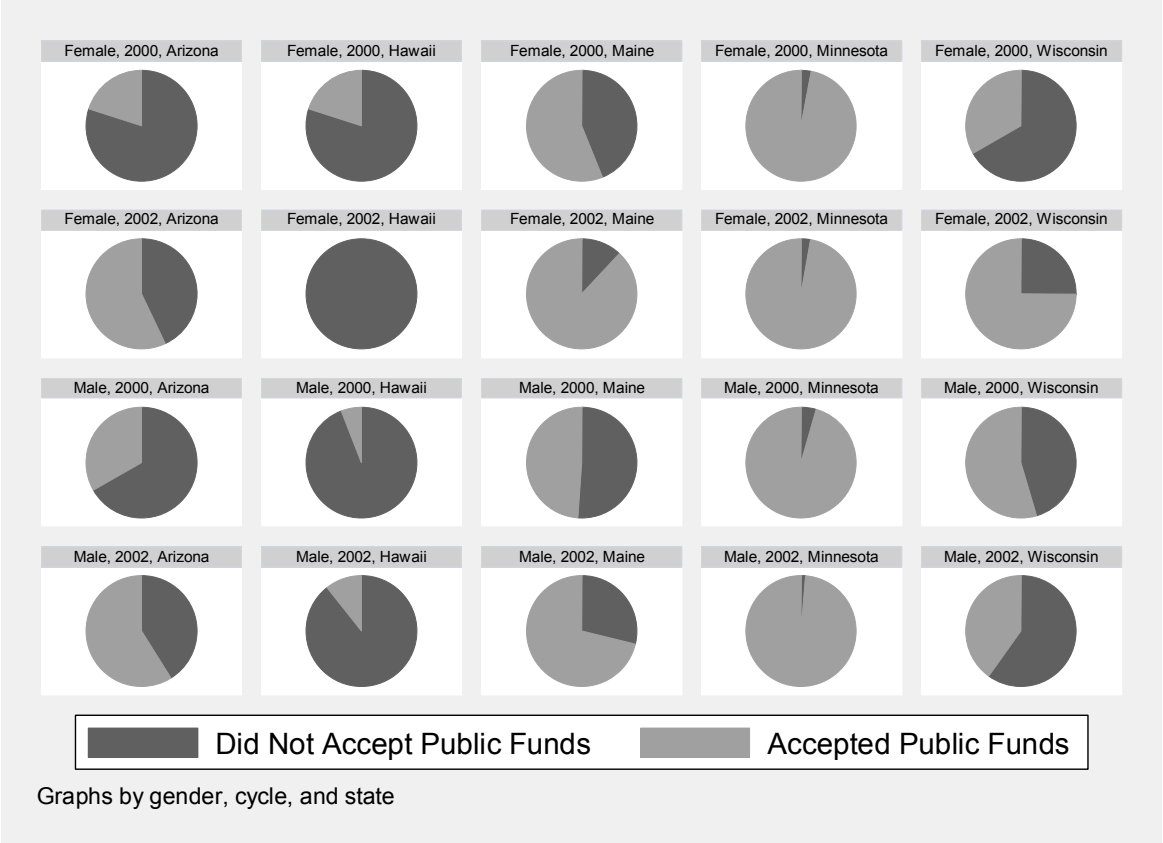


TABLE 1. Hypothesized Effects on Public Funding Acceptance and Summary Statistics Lower Chamber

	Expected Effect	Mean	Std. Dev.	Min	Max
Dependent Variable					
Public Funding		0.61	0.49	0	1
Independent Variables					
Female	+	0.27	0.44	0	1
Incumbent	-	0.36	0.48	0	1
Open Seat	-	0.27	0.44	0	1
Challenger	Omitted	0.37	0.48	0	1
Previous Run	-	0.30	0.46	0	1
Previous Run * Challenger	-	0.05	0.22	0	1
Democrat	+	0.50	0.50	0	1
Arizona	+	0.09	0.28	0	1
Hawaii	-	0.11	0.32	0	1
Maine	+	0.32	0.47	0	1
Minnesota	+	0.34	0.47	0	1
Wisconsin	Omitted	0.15	0.35	0	1
Previous Difference	+	23.67	19.21	0.10	50
Previous Difference * Incumbent	+	6.80	13.54	0	50
Previous Difference * Challenger	-	7.06	13.65	0	50
N = 1440					

TABLE 2. Hypothesized Effects on Public Funding Acceptance and Summary Statistics Upper Chamber

	Expected Effect	Mean	Std. Dev.	Min	Max
Dependent Variable					
Public Funding		0.70	0.46	0	1
Independent Variables					
Female	+	0.30	0.46	0	1
Incumbent	-	0.37	0.48	0	1
Open Seat	-	0.22	0.42	0	1
Challenger	Omitted	0.40	0.49	0	1
Lower Chamber Member	-	0.07	0.26	0	1
Previous Run	-	0.26	0.44	0	1
Previous Run * Challenger	-	0.04	0.20	0	1
Democrat	+	0.50	0.50	0	1
Arizona	+	0.11	0.31	0	1
Hawaii	-	0.10	0.30	0	1
Maine	+	0.24	0.43	0	1
Minnesota	+	0.47	0.50	0	1
Wisconsin	Omitted	0.08	0.27	0	1
Previous Difference	+	20.04	17.12	0	50
Previous Difference * Incumbent	+	5.78	11.09	0	50
Previous Difference * Challenger	-	6.27	11.59	0	50
N = 536					

TABLE 3. Logistic Regression Model One
Lower Chamber

	Coefficient	Robust Std. Error	*z*-score	
Female	0.47	0.18	2.65	**
Incumbent	-0.83	0.22	-3.84	***
Open Seat	0.04	0.20	0.18	
Previous Run	-0.24	0.20	-1.18	
Previous Run * Challenger	-0.52	0.42	-1.25	
Democrat	1.32	0.16	8.29	***
Arizona	-0.30	0.25	-1.23	
Hawaii	-3.12	0.48	-6.44	***
Maine	-0.27	0.19	-1.46	
Minnesota	5.08	0.54	9.34	***
Constant	-0.30	0.20	-1.49	
Dependent Variable: Public Funding	N = 1440		Wald chi^2 (10) = 196.69	
Log-Likelihood = -540.58	Pseudo r^2 = 0.44		Prob > chi^2 = 0.00	

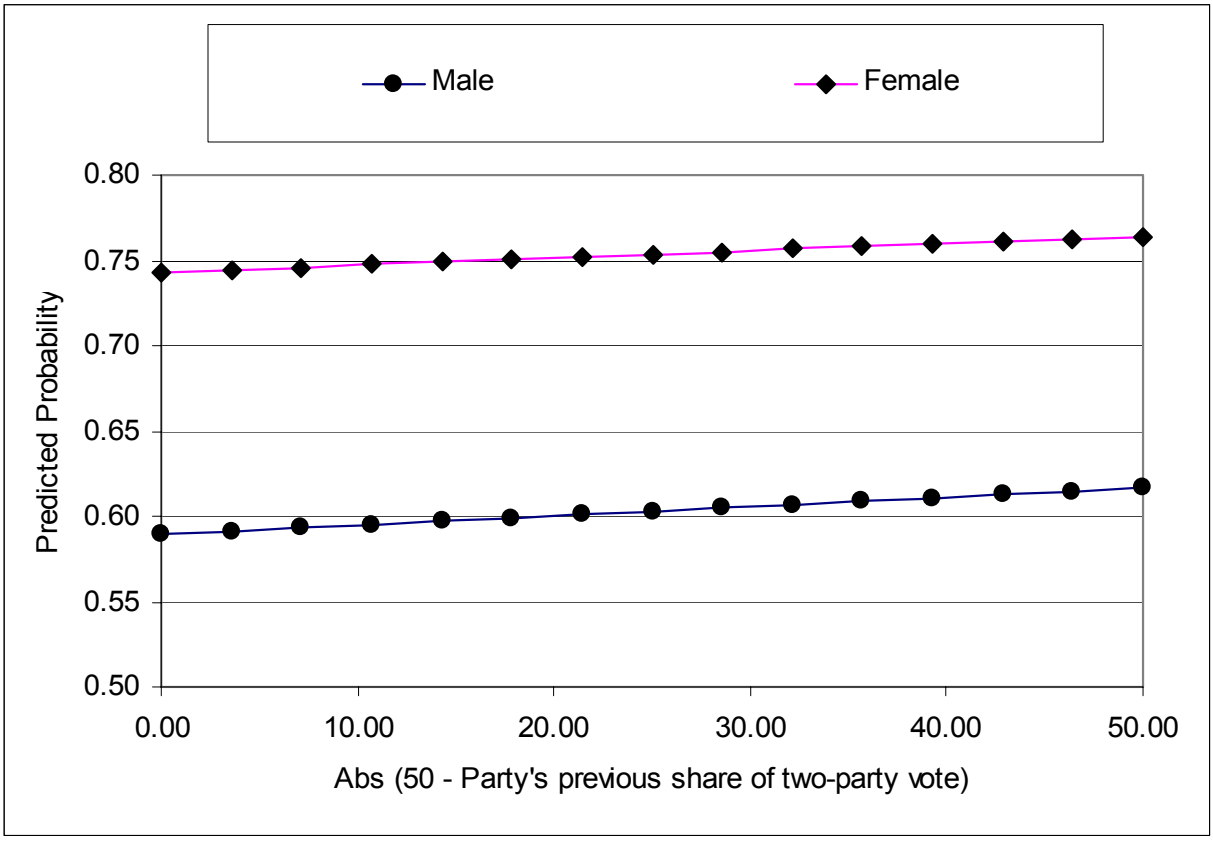
*** p < 0.001; ** p < 0.01; * p < 0.05 Robust Std. Errors clustered by candidate

TABLE 4. Logistic Regression Model Two
Lower Chamber

	Coefficient	Robust Std. Error	*z*-score	
Female	0.70	0.28	2.49	*
Incumbent	-1.59	0.50	-3.18	***
Open Seat	-1.12	0.60	-1.88	
Previous Run	0.46	0.35	1.32	
Previous Run * Challenger	-0.75	0.59	-1.27	
Democrat	1.92	0.28	6.86	***
Arizona	-0.01	0.29	-0.03	
Hawaii	-3.34	0.54	-6.16	***
Previous Difference	0.002	0.01	0.33	
Previous Difference * Incumbent	-0.002	0.02	-0.27	
Previous Difference * Challenger	0.001	0.02	0.09	
Constant	-0.44	0.35	-1.25	
Dependent Variable: Public Funding	N = 496		Wald chi^2 (10) = 79.28	
Log-Likelihood = -211.08	Pseudo r^2 = 0.34		Prob > chi^2 = 0.00	

*** p < 0.001; ** p < 0.01; * p < 0.05 Robust Std. Errors clustered by candidate
Maine and Minnesota dropped due to collinearity

Figure 5: Predicted Probability of Accepting Public Funds versus Previous Party Performance by Gender – Lower Chamber



Note: Assumes candidate is running in an open seat race for the Wisconsin Assembly and is a Democrat who has not previously run for the Assembly.

TABLE 5. Logistic Regression Model One
Upper Chamber

	Coefficient	Robust Std. Error	*z*-score	
Female	-0.13	0.30	-0.43	
Incumbent	-0.51	0.37	-1.39	
Open Seat	-0.50	0.38	-1.31	
Lower Chamber Member	-1.24	0.50	-2.49	*
Previous Run	-0.33	0.34	-0.97	
Previous Run * Challenger	-0.17	0.80	-0.21	
Democrat	1.43	0.30	4.82	***
Arizona	-0.01	0.41	-0.02	
Hawaii	-2.53	0.60	-4.20	***
Maine	1.01	0.36	2.83	**
Minnesota	4.04	0.50	8.15	***
Constant	-0.47	0.33	-1.44	
Dependent Variable: Public Funding	N = 536	Wald chi^2 (10) = 111.35		
Log-Likelihood = -179.24	Pseudo r^2 = 0.45	Prob > chi^2 = 0.00		

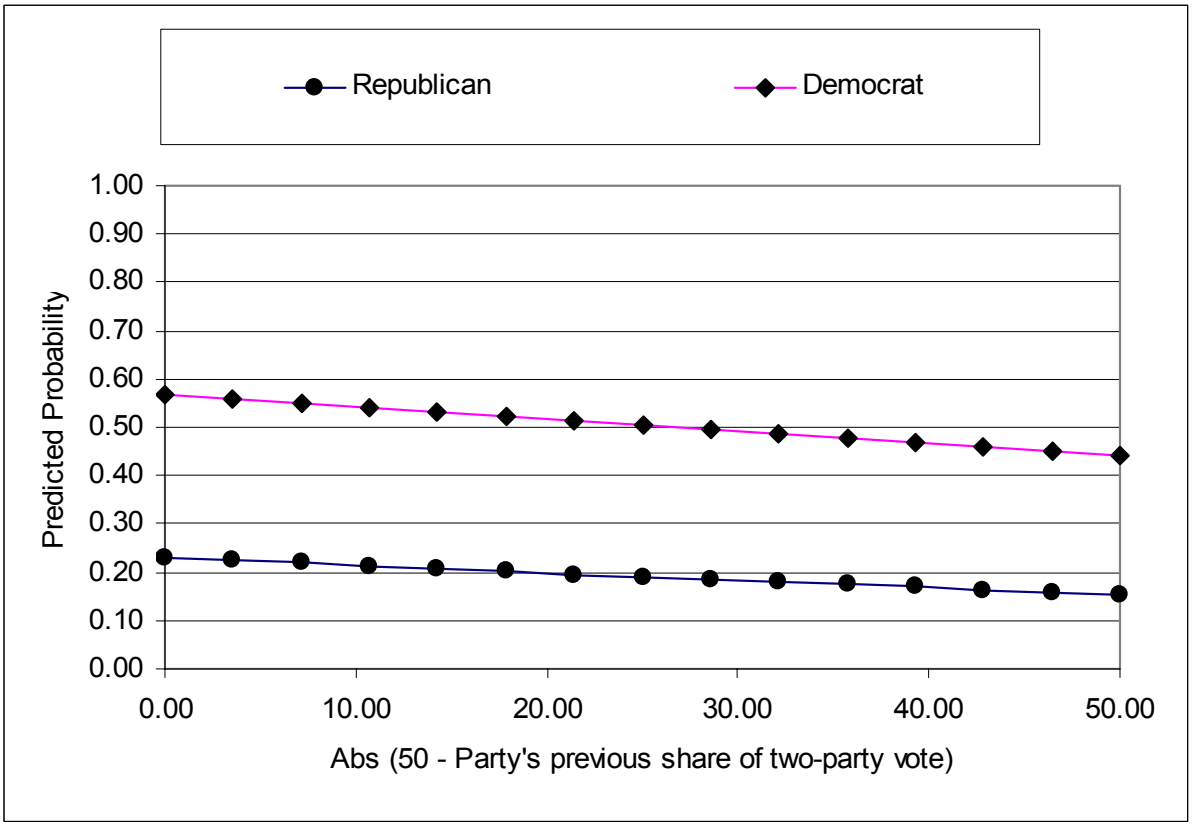
*** p < 0.001; ** p < 0.01; * p < 0.05 Robust Std. Errors clustered by candidate

TABLE 6. Logistic Regression Model Two
Upper Chamber

	Coefficient	Robust Std. Error	*z*-score	
Female	-0.12	0.30	-0.42	
Incumbent	0.19	0.53	0.36	
Open Seat	-0.32	0.63	-0.51	
Lower Chamber Member	-1.26	0.49	-2.55	*
Previous Run	-0.38	0.35	-1.10	
Previous Run * Challenger	-0.14	0.80	-0.17	
Democrat	1.47	0.31	4.80	***
Arizona	0.16	0.45	0.36	
Hawaii	-2.40	0.61	-3.92	***
Maine	0.96	0.37	2.59	**
Minnesota	4.05	0.51	7.92	***
Previous Difference	-0.01	0.01	-0.56	
Previous Difference * Incumbent	-0.04	0.02	-1.93	
Previous Difference * Challenger	0.01	0.02	0.28	
Constant	-0.46	0.40	-1.15	
Dependent Variable: Public Funding	N = 536	Wald chi^2 (10) = 113.73		
Log-Likelihood = -183.08	Pseudo r^2 = 0.44	Prob > chi^2 = 0.00		

*** p < 0.001; ** p < 0.01; * p < 0.05 Robust Std. Errors clustered by candidate

Figure 6: Predicted Probability of Accepting Public Funds versus Previous Party Performance by Party – Upper Chamber



Note: Assumes candidate is a female Democrat running in an open seat race for the Maine Senate and is a former member of the House who has not previously run for the Senate.

Appendix A – Characteristics of Public Funding Programs

	Arizona	Maine	Wisconsin	Minnesota	Hawaii
Effective	2000	2000	1978	1978	1996
Qualification	Threshold of 210 $5 contributions. These funds are deposited in public fund program and are not kept by the candidate	Threshold number of $5 qualifying contributions, 50 for House, 150 for Senate. These funds are deposited n the public fund program and are not kept by the candidate	• Win primary with at least 6% of total vote for office • Raise threshold amount in $100 contributions ($1,725 for Assembly, $3,450 for Senate)	Threshold amount raised in $50 contributions: $1,500 for House, $3,000 for Senate	Threshold amount raised ($1,500 for House, $2,500 for Senate)
Maximum Grant	• Up to spending limit • Matching funds (up to an additional $56,600) provided to participating candidates running against privately financed opponents, and to offset independent expenditures against • Independent candidates receive 70% of spending limit	• Up to spending limit • Matching funds (up to an additional 200% of original grant) provided to participating candidates running against privately financed opponents and to offset independent expenditures against	• $15,525 for Senate (2002) • $7,763 for Assembly (2002)	• amount of direct grants determined by dividing total funds by number of candidates, but may not exceed 50% of spending limit • Small contribution refund program reimburses individuals up to $50 for contribution to participating candidate	• Amount of grant limited to 15% of spending limit
Spending Limit (contested election)	• $28,300 for primary/general in both House and Senate elections • $11,320 for single party dominant districts	• $5,406 for primary/ general in House • $23,728 for primary/ general in Senate	• $17,250 for Assembly • $34,500 for Senate • Limits unchanged since 1986	• $34,100 for House (2004) • $64,866 for Senate (2002) • Separate spending limits for election and non-election years	Spending limit fixed as $1.40 x number registered voters in district 2004 range: House: approx. $14,000-$19,000 Senate: approx $23,000-$45,000
Special Conditions	• Unopposed candidates not eligible for public funds beyond qualifying contributions • Nonparticipating candidates faced additional reporting requirements	• Nonparticipating candidates face additional reporting requirements • Participating candidates permitted to raise small amounts	Spending limits apply only if all candidates accept public funds	• Spending limits increase by 10% for first time candidates and by 20% for candidates running in competitive primary • Spending limits waived when nonparticipating opponent exceeds threshold expenditures	