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Quasi-Experimental Design, Constituency, and Advancing Women's Interests: Reexamining the Influence of Gender on Substantive Representation

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Abstract

Research investigating whether female legislators provide more effective substantive representation on women's issues than their male colleagues faces a significant methodological hurdle. Models used to estimate the effect of gender on representation inevitably omit constituency variables that affect the character of legislators' decisions and are also correlated with gender, potentially biasing the estimates of the effect of gender. Employing a quasi-experimental research design as an alternative strategy, the authors remove this hurdle and estimate the influence of gender on representation free from this potential bias. The authors find that gender does affect representation and observe critical mass effects.

Keywords

Congress, descriptive representation, substantive representation, women and politics, agenda setting

Do female members of Congress provide more effective representation of women's interests than their male colleagues? Prior research generally indicates "yes" with debate as to whether having a substantial number of women in the legislative body, or a "critical mass," encourages congresswomen to devote their scarce time and legislative resources to enhancing the quality of representation that they provide women (e.g., Bratton 2005; Grey 2006; Reingold 1992, 2000; Saint-Germain 1989; Thomas 1994). Scholars observe this benefit of descriptive female representation across a host of issues (Barrett 1995; Carey, Niemi, and Powell 1998) but especially with respect to women's and feminist issues (e.g., Burrell 1994; Carroll 2001; Dolan 1997; Swers 1998, 2002). Female legislators also voice distinct policy priorities from men (e.g., Thomas 1994; Thomas and Welch 2001), view women outside their district as constituents (Carroll 2002), and expand the agenda to include women's perspectives (e.g., Dodson 2005, 2006; Kathlene 1994; Reingold 1996; Rosenthal 2000; Wolbrecht 2002, 193). In summary, "women legislators tend, more often than men, to make priorities of issues important to women and to introduce and successfully usher those priorities through the legislative process" (Dodson 2001, 226).

This conclusion comes from studies that employ an array of methodological strategies. Nonetheless, these strategies regularly fail to differentiate between the effect of a legislator's gender on her or his representational behavior and the demands of one's constituency. The inferential dilemma is profound: "The omission of constituency preferences from models of gender differences in legislators' policy attitudes" may produce premature conclusions where "the impact of gender may have been overestimated" (Poggione 2004, 306). The relatively few studies that work to account for constituency demands in assessing women's legislative behavior thus offer a substantial leap forward for understanding women's substantive representation (e.g., Bratton and Haynie 1999; Poggione 2004; Swers 1998, 2000; Thomas and Welch 1991; Tolleson-Rinehart 2001). Inferential dilemmas persist, though, as these inquiries rely almost solely

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on standard sampling techniques and multivariate models. Valid data that accurately capture the full political landscape and constituency characteristics for all 435 districts simply do not exist. This means inevitable problems of measurement error and an inability to determine how omitted variable bias is at play (Clarke 2005, 2009).

This article capitalizes on experimental logic to offer a different empirical solution to the problem of controlling for constituency. We utilize a quasi-experimental research design composed of longitudinal data. The sample includes all pairs of House members serving during the 1970s, 1980s, and 1990s during which a district transitioned from having a woman serve it immediately before and/or immediately after a man (omitting transitions in which districts are changed due to redistricting). After identifying these pairs, we recorded the number of bills sponsored by the members on "social welfare" and "feminist" issues during the relevant congressional sessions employing Swers's (2002) criteria for coding these issues. Our research design, then, allows us to isolate variance in representatives' behavior explained by gender from the variance in behavior explained by representatives' constituents, providing an estimate of the effect of gender on behavior that is less clouded by concerns about correlation with omitted constituency variables than virtually any previous research.

In addition, we take advantage of our ability to control for the influence of lawmakers' constituencies on their behavior by extending our analysis to assess whether congresswomen provide enhanced representation to women when a higher percentage of their colleagues are women (see Kanter 1977). Although there is a considerable body of research on whether such a "critical mass" effect exists, this research does not examine this question with the effects of gender and constituency on representatives' behavior disconnected. Since our analysis does so, our findings provide novel insight on this "critical mass" question.

Quasi-experimental Design and the Empirical Foundations of Gender and Critical Mass Effects

The Difficulty of Isolating the Effects of Gender and Constituency on Representation

The difficulty for understanding how gender influences how well legislators represent women does not stem from a lack of outcome variables—almost all levels of legislative activity have been studied.[1] Rather, as Poggione (2004, 306) argues, the omission of constituency characteristics from these models leaves open the possibility that "systematic differences in men and women's constituencies explain the relationship between gender and legislators' preferences, rather than gender itself." Palmer and Simon (2008, 202) add more recent credence to this concern, finding that female candidates cluster in particular districts and that these districts share attributes that make them particularly favorable toward electing women. Therefore, it is plausible that women elected from "favorable" districts legislate differently because of the district's unique policy demands rather than their gender.

This possibility demonstrates the importance of separating the effect of gender and district characteristics and suggests why those studies that began to take constituency into effect advanced the literature. Thomas and Welch (1991), for example, controlled for the degree of urbanization in a district. Swers (1998, 2002) controlled for district effects by including in her models the percentage of district vote going to Clinton and Perot in 1992, an urbanicity measure, a southern dummy variable, the percentage of African Americans in the district, and the district's median household income. Poggione (2004) took district into account by controlling for average household income, the percentage unemployed in the district, the percentage of district residents who earned a college degree, and urbanicity. Some who push the representational question to simultaneously examine gender and racial effects have included the percentage of a district that is African American and district income (Bratton, Haynie, and Reingold 2006) as well as the size of the largest city in the district, majority black district, and various interaction variables (Bratton and Haynie 1999).

The question of whether women provide more effective substantive representation because of their gender, differences in constituencies, or both is not, however, settled. It is not credible to conclude from the independent variables listed that the degree to which members' constituencies are favorable toward the representational behaviors that serve as the dependent variables in these studies is measured exhaustively.[2] For example, does the level of support that President Clinton received (Swers 2002) indicate how predisposed constituents across House districts are to addressing social welfare policy challenges and/or endorsing feminist policies (the number of such bills sponsored by members is a dependent variable in some of the studies described previously)? President Clinton won states, and congressional districts within states, across the South, and received similar support in states, and congressional districts, in the Northwest and Northeast. But individuals who voted for Clinton in the South did not, on average, react similarly to Clinton voters in the northeast when the new president endorsed a "Don't ask, don't tell" policy regarding the ability of homosexuals to serve in the military.

Clearly, there is a lot more to a district than measures of vote share, income, urbanization, and demographics. Although the scholars cited previously recognize this problem, and consciously use the best available data to confront it, there remains inevitable slippage between variables used to control for the effects of constituency on representational behavior, relegating some of these effects to the error term. Since these effects are likely to be correlated with representatives' gender, there is a need for additional analyses that more thoroughly control for the effects of constituency. But these measures are notoriously elusive. How does one obtain such measures for the representational behavior analyzed by scholars and ensure that these measures are valid for all 435 congressional districts for the various dependent variables that serve as indicators of the quality of substantive representation for women?

The quasi-experimental design employed in this article sidesteps these problems of measurement error and omitted variable bias by including only male and female "matched pairs" who represented the exact same district in immediate sequential ordering (Clarke 2005, 348-49). The experimental logic looks for differences before and after the treatment (gender change in representation) on the outcome variable (prioritization of women's issues and feminist issues) while keeping all district factors potentially related to the outcome variable the same or constant (same district). In doing so, we avoid potential bias in the gender coefficients presented in the following by fully accounting for representatives' constituencies.[3] Surely candidates may try to galvanize somewhat different electoral alignments within a district (Fenno 1978; Swers 2002), but by ensuring that each matched pair represented the same district, and in light of the fact that districts that elect women tend to demand more women's interests (Palmer and Simon 2008, 177-213), both men and women representing the same district in sequential ordering can reasonably be expected to have to respond to their district's unique constituent demand for women's interests.[4]

The Benefits of a Critical Mass of Women in the Legislature?

The question of whether high percentages of women in legislatures spur female representatives to provide effective representation of women's interests faces its own hurdles. There is evidence for critical mass effects when 15 percent to 20 percent of legislative bodies are composed of women (Dolan and Ford 1998, 77; Saint-Germain 1989; Thomas 1991, 1994). Others, however, only find these effects for some legislative outcomes (Berkman and O'Connor 1993), report that critical mass effects dissipate as more women, with more diverse viewpoints, enter the legislative body (Reingold 1992, 2000; Vega and Firestone 1995), or find that when very small numbers of female state representatives ("tokens") are present that these individual women successfully "take up" women's interests and do so more than when higher percentages of women are in the chamber (Bratton 2005; Crowley 2004). Indeed, the notion that a "single proportion holds the key to all representational needs of women" has increasingly fallen out of favor and been replaced by calls for more nuanced accounts (Childs and Krook 2006; Gray 2006, 492).

Such accounts could include properly controlling for contextual factors like state and/or national political culture and party control (e.g., Dodson 2005, 2006; Reingold 2008, 142; Rosenthal 1998; Thomas 1994), specifying what dependent variable(s) best captures representation, and answering the interpretive question of how, exactly, critical mass effects occur (Cammisa and Reingold 2004, 197-98). Our previous discussion of how constituency has been underestimated, however, suggests that at least in the case of the U.S. Congress this debate takes place on a considerable amount of quicksand. Without the full inclusion of constituency in any of the previous models, it is impossible to truly specify the importance, nonimportance, or conditional importance of women reaching a critical mass in Congress. Since our quasi-experimental design provides more firm empirical ground on the constituency–gender correlation issue, we can uniquely assess whether a critical mass of female legislators is necessary for congresswomen to provide more effective substantive representation than their male colleagues.

Substantive Representation in Congress: Sample Size and Generalizability

Another hurdle for understanding the effect of gender on legislative behavior occurs, ironically, as scholars craft creative data collection and sampling procedures. To better account for the influence of constituency on behavior, and to address the problem of too few women in Congress to allow for meaningful statistical analysis, researchers examine individual states (e.g., Kathlene 1994). Other researchers draw data from strategic groupings of states—usually to have variance in the percentage of women serving in the legislative bodies (e.g., Bratton 2005; Bratton and Haynie 1999; Bratton, Haynie, and Reingold 2006; Thomas 1991). On the national level, researchers look at particular legislative sessions (Swers 1998) or pool data from a few chronologically neighboring sessions (Swers 2002).

Trying to better account for constituency, or increase the "*n*," are both sound reasons for developing these data sources but generalizing from them remains problematic for understanding women's substantive representation in Congress. State studies suffer from the difficulty of generalizing about national legislative behavior when only examining unique local contexts. Focusing on the "Year of the Woman" or sessions surrounding the "Republican Revolution" may offer particular analytic advantages (higher numbers, party shift) but also may not be representative snapshots of congressional behavior. Our sampling strategy, however, allows for longitudinal analysis not easily swayed by unrepresentative years or legislative session(s) as it includes all female members serving in Congress from 1973 to 2002 who had a man immediately proceed or follow them in office and all these male representatives. This also alleviates the small sample size problem that leads many researchers to state level data, as we have 103 women and 103 men in our sample.

The advantages of this research design and subsequent analysis are fourfold. First, because our female–male pairs (and vice versa) of members represented the exact same district, we are able to assess the effect of gender on representational behavior "cleanly" in the sense that coefficient estimates of the effect of gender on the number of "feminist" and "social welfare" bills will not be biased by correlation with constituency factors unaccounted for in the models and therefore residing in the error terms. In this way, our analysis is similar to that of Gerrity, Osborn, and Mendez (2007). However, in examining pairs of representatives from the 1970s, 1980s, and 1990s redistricting rounds, our estimates of the effect of gender on representation are based on a larger sample size than these authors who examine twenty-one male-female pairs and note that their findings are limited by a small sample size. Second, our research design allows for examining whether the relative presence of other women in Congress influences women's legislative behavior while adequately controlling for the influence of representatives' constituencies on their behavior. Third, since we examine data on members over three decades, we avoid the "low *n*" problem of not having enough women in Congress to conduct meaningful statistical analysis at the federal level. Finally, our longitudinal data also guard against the possibility of generalizing about gender and representation from a potentially unrepresentative legislative session(s).

Data and Method

The dependent variables in our models, the number of bills members sponsor promoting feminist issue positions and the number of bills that address social welfare issues, follow, among others, Swers (2002), Bratton (2005), and Gerrity, Osborn, and Mendez (2007). Sponsorship behavior

> provides important insights about which members are working to place women's interests on the national agenda. In contrast to other legislative activities like offering amendments in which restrictive rules governing debate can prevent members from offering women's issue proposals, representatives have complete control over the number and content of the bills the sponsor. (Swers 2002, 32)

Indeed, sponsorship behavior provides "a central way that proposals for public policy enter the congressional agenda" (Wolbrecht 2002, 178).

To identify the pairs of female–male members serving in identical districts in consecutive congressional sessions, we identified every woman who served in the U.S. House during the 1970s (1973-1982), 1980s (1983-1992), and 1990s (1993-2002) rounds of redistricting. We then identified the members serving immediately prior to and after these female representatives. When the member preceding/succeeding them was a man, we classified the two members as a pair for inclusion in our sample.[5] We omitted instances in which members of different genders served successively during the same session on the grounds that only members serving the same amount of time are comparable. This process yielded 103 pairs (206 observations). These pairs included 88 of the 122, or 72.13 percent, of women who served in the House during the 1973-2002 period. In addition, 57.28 percent of the members were Democrats while 42.72 percent were Republicans, figures that roughly approximate the Democratic advantage in the House for most of this period. Given that our sample includes such a high percentage of women who served during the period, we believe our findings are generalizable to the differences in behavior between male and female representatives.[6] One limitation of the analysis is that many women elected in 1992 ("The Year of the Woman") are excluded from our sample because they were elected in the first year of the 1990s redistricting cycle (ten of the twenty-four women elected in 1992 are included in our analysis because when they departed Congress, they were replaced in their districts by men). This means that we cannot compare their behavior to men who they replaced, since the districts that they represented were not identical to the districts of these male representatives. This omission is noteworthy because these women may have been particularly likely to champion women's interests in their bill sponsorship behavior. Since these women may have been more likely to prioritize women's interests than their male predecessors, their omission works against finding that there is a difference between male and female representatives' bill sponsorship behavior.

To measure the number of feminist and social welfare bills members sponsored, we examined the legislative record made available at *Thomas*, the Library of Congress's legislative information page. We identified each bill for which members were the primary sponsors during the congressional sessions for which they were the first/second member of their pair. A coder then employed the Swers (2002, 36-38) criteria for identifying bills involving feminist and social welfare legislation to determine whether the bills involved feminist/social welfare issues. Feminist bills promote "role equity and/or role change for women" whereas social welfare bills "include both liberal and conservative proposals concerning issues with which women have historically been concerned in their role as caregiver" (Swers 2002, 37). In employing these criteria, feminist legislation sponsored by members in the sample included bills such as H.R. 503 in the 104th Congress seeking to require telecommunications providers to purchase products by businesses owned by women in addition to many others seeking to engineer role equity/change through federal law. Legislation considered to fall into the "social welfare" rubric included bills seeking to enhance health care for rural areas (H.R. 3909) in the 103rd Congress and expand the Head Start program to include child care services (H.R. 3) in the 101st Congress in addition to a host of proposals related to fostering an improved standard of health and living for U.S. citizens.

To produce counts of the number of feminist and social welfare bills members sponsored during the session in which they are included in our analysis, we summed the number of such bills sponsored by members. To check the reliability of the individual coding decisions used to produce these counts, a second coder also classified all bills sponsored by members serving during the sessions spanning the 1990s round of redistricting. Intercoder agreement kappa scores were significant at $p < .001$, demonstrating that the data used to compile these count variables are reliable.[7]

We present estimates from negative binomial maximum likelihood models, which are appropriate for count data that is overdispersed.[8] In addition to controlling for constituency experimentally, we also control statistically for other factors that may affect the number of bills that members sponsor. We employ a dummy variable to measure partisanship (1 = *Democrat*; 0 = *Republican*) and first dimension DW Nominate coordinates to measure ideology, or more specifically, members' liberalism on economic issues, which range from –1 to 1 with higher values indicating higher levels of conservatism (McCarty, Poole, and Rosenthal 1997).[9] We do not control for both partisanship and ideology in the same model due to collinearity between them.[10] Since the opportunity costs of sponsoring legislation are lower when one serves on a committee with jurisdiction over the policy area with which the legislation is involved (Hall 1996), we control for members' committee assignments, creating dummy variables equal to 1 when members serve on committees with jurisdiction over issues related feminist and social welfare policy; 0 otherwise.[11] Additionally, since members with greater experience are more informed on many policy issues than members with less experience, more senior members may be more likely to sponsor bills; therefore, we include a variable that is a count of the number of terms served by members.[12] Since members departing Congress lack the electoral incentive to engage in the type of representational behavior that bill sponsorship entails, we control for whether a member has announced the intention to retire from the House at the end of the session using a dummy variable (1 if the member made such an announcement; 0 otherwise).[13] Recognizing that additional factors, such as a member's personal interest in making public policy as opposed to devoting more attention to other goals (Fenno 1973; Hall 1996), may affect the decision to participate in the production of policy, we include a variable that is the total number of bills sponsored by members in the congressional session.[14] This variable controls statistically for member's "personal taste" for sponsoring legislation in general, ensuring that if we observe a positive and significant relationship between gender and sponsorship, it is real and not due to the gender coefficient displaying an effect that is really due to a member's propensity to "load the deck" by sponsoring legislation in all domains—regardless of their interest in women's issues.

Finally, we include a variable that is the interaction between members' gender and the percentage of women serving in the congressional session during which the member served. If gender affects bill sponsorship conditional on the level of women serving in the House, the coefficient for this variable should be positive and significant. Such a finding would constitute important evidence for the critical mass perspective when explaining women's substantive representation in the U.S. Congress since our analysis is the first to assess this perspective having isolated the effects of constituency on members' behavior from the analysis.[15]

In discussing our findings, we focus primarily on the additive effect of gender, and the effect of gender conditional on the volume of women serving in the chamber, on sponsorship of feminist and social welfare bills. This focus is due to the nature of the sample, which facilitates inferences about the effect of gender on representation. Because the observations are sampled to accomplish this goal, we do not examine the universe of members and their representational decisions. As such, this analysis is less capable of sorting effects of other variables in the

Table 1. Negative Binomial Estimates of the Number of Feminist Bills Sponsored by Representatives

Variable	(1)	(2)	(3)	(4)
Gender	1.319****	0.272	1.267****	0.208
	(0.340)	(0.714)	(0.330)	(0.718)
Democrat	–0.102	–0.104		
	(0.317)	(0.331)		
1st dimension nominate coordinate			–0.133	–0.140
			(0.427)	(0.421)
Number of terms served	0.144****	0.154****	0.136****	0.146****
	(0.044)	(0.044)	(0.046)	(0.045)
Member on a women's issues committee	0.378	0.340	0.320	0.278
	(0.298)	(0.293)	(0.304)	(0.299)
Retiring member	–0.164	–0.172	–0.114	–0.120
	(0.428)	(0.420)	(0.440)	(0.432)
Percentage of women serving in the House	2.955	–6.890	3.408	–6.570
	(4.270)	(7.431)	(4.328)	(7.521)
Gender × Percentage of Women Serving in the House		13.784*		13.943*
		(8.423)		(8.514)
Total number of bills sponsored	0.035****	0.035****	0.036****	0.035****
	(0.007)	(0.007)	(0.008)	(0.007)
Constant	–3.420	–2.719	–3.464	–2.746
	(0.546)	(0.658)	(0.533)	(0.653)
N	206	206	206	206
Log-Likelihood	–142.16	–140.80	–142.16	–140.80
Likelihood Ratio Chi-Squared	67.39****	70.11****	67.39****	70.11****
LR Test vs. Poisson	7.58***	5.72***	8.50***	6.41**

Note. The standard errors are in parentheses.
*p < .10, ***p < .01, ****p < .001 (one-tailed tests).

model on the dependent variables even though it is essential to include them as independent variables to control for their effects on sponsorship among the members in our sample.

Findings

Because the dependent variables are counts of the number of feminist and social welfare bills sponsored by members, and because diagnostic tests revealed that variance of these counts exceeded their means, we employ negative binomial estimators to model the relationships between the dependent and independent variables. Table 1 provides estimates of the effect of gender, as well as the other independent variables, on the number of feminist bills that members sponsored. In model 1, we estimate the number of feminist bills sponsored using the additive specification for gender—without the interaction with the percentage of women serving in the House. We observe a positive and significant relationship between gender and the number of feminist bills sponsored, providing support for the hypothesis that gender affects substantive representation of women's interests. This finding is noteworthy in that we observe this relationship and can credibly argue that we have controlled for all of the district characteristics that also affect such behavior through the quasi-experiment as well as the other key independent variables. We do not observe a significant relationship between partisanship and sponsorship of feminist bills, however. As expected, there are positive and significant coefficients for the variables measuring the number of terms served by members and the total number of bills sponsored by members during the session. The other independent variables are not significantly related to the volume of feminist bills sponsored by members.[16]

In model 2 of Table 1, we estimate the effect of gender on feminist bill sponsorship conditional on the percentage of women serving in the House and expect the coefficient for the interaction between gender and the percentage of women serving will be positively and significantly related to the volume of bill sponsorship. In fact, we observe this relationship, as evidenced by the positive and significant interaction coefficient in model 2 (p = .051). Additionally, we calculated conditional coefficients and standard errors for gender across the range of the percentage of women serving in the House from 1973 to 2002 (Friedrich 1982; Brambor, Clark, and Golder 2006). These findings reveal that female members were significantly more likely to sponsor feminist legislation when at least 5 percent of their House colleagues were women. However, the gender variable alone is no longer significant. This suggests that gender certainly matters for sponsoring feminist bills—that women are more apt to do so than men—but this effect is interlinked with how many other women serve in the chamber. Reaching a critical mass of women in the chamber, therefore, appears necessary to spur or make it feasible for individual congresswomen to enhance substantive representation in this policy area. The other findings in this model are identical to those observed in model 1.

Models 3 and 4 are similar to models 1 and 2, respectively, except that the ideology of members is substituted for their partisanship. Briefly, and consistent with models 1 and 2, we observe that gender affects the volume of bill sponsorship in the positive and significant gender coefficient in model 3 and the positive and significant interaction coefficient for gender and the percentage of women in the House in model 4 (p = .051). Here again then, women are more apt than men to sponsor feminist legislation but this relationship is conditioned by having a sizeable presence of women in the chamber.

Some readers may wonder whether the findings supporting the critical mass hypothesis may be due to changes in the congressional agenda over time. In particular, was it that the agenda was composed of more social welfare and feminist items in the 1990s when the House was composed of a higher percentage of women? If this was the case, the critical mass findings could be spurious. Rather than a high percentage of female colleagues spurring female representatives' sponsorship behavior, it could simply be that there were more of such bills sponsored during this period. This concern is not warranted based on the data in our sample. In fact, there were more social welfare and feminist bills sponsored by members in our sample who served during the earlier period covered by our data. This means that although more bills were sponsored in the 1970s when female representation in Congress was low, female members in the sample themselves were more likely to sponsor feminist (models 2 and 4 in Table 1) and social welfare (models 2 and 4 in Table 2) legislation in the 1990s when a higher percentage of their colleagues were women. Therefore, far from the critical mass finding being driven the congressional agenda, we observe this finding in spite of what the agenda looked like in terms of the volume of feminist and social welfare sponsored over time.

But what of the substantive effect of the findings? Figures 1 and 2 illustrate this by charting the substantive effect of gender on the sponsorship of feminist bills. Figure 1 presents the number of bills that model 1 predicts members to offer based on their gender, holding the other variables at their modal or mean values.[17] Although the coefficient for partisanship is not significantly related to feminist bill sponsorship, we present the predictions for male and female Democrats and Republicans to emphasize the primacy of gender and the peripheral role of partisanship in influencing bill sponsorship on feminist issues to readers, something that is highlighted by the lines for Democrats and Republicans running nearly together. The figure shows that model 1 predicts that Democratic and Republican men will offer .08 and .07 bills, respectively.[18] However, the model predicts more than a 300 percent increase in the number of bills female Democrats and female Republicans are expected to offer compared to their male colleagues—.26 and .25, respectively. Indeed, as Figure 1 tellingly shows, the predictions for Democratic and Republican men and Democratic and Republican women are virtually identical, emphasizing the important role of gender and the negligible role of partisanship in determining the sponsorship of feminist legislation.

The conditional models also told us that these gender effects are tightly woven with the percentage of women serving in the House. Once gender was interacted with this factor, the resulting variable was statistically significant. Gender alone was not. To illustrate the interactive effect for gender across levels of female membership in the U.S. House on the number of feminist bills sponsored, we chart the number of such bills that model 2 predicts male and female members will offer in Figure 2.[19] The figure reveals that men and women are predicted to offer nearly an identical number of feminist bills at low levels of female membership in the House. However, as female membership increases, the number of bills that women are predicted to sponsor rises substantially, increasing approximately threefold at the maximum value of female membership. The predictions for their male colleagues, on the other hand, drop slightly as the House's female population increases.

In summary, the multivariate models reveal that gender exerts a substantively large influence over the volume of feminist bills that members sponsor. We observe this effect in controlling for constituency characteristics with certainty and eliminate any endogeneity problems. Nuance about the relationship between gender

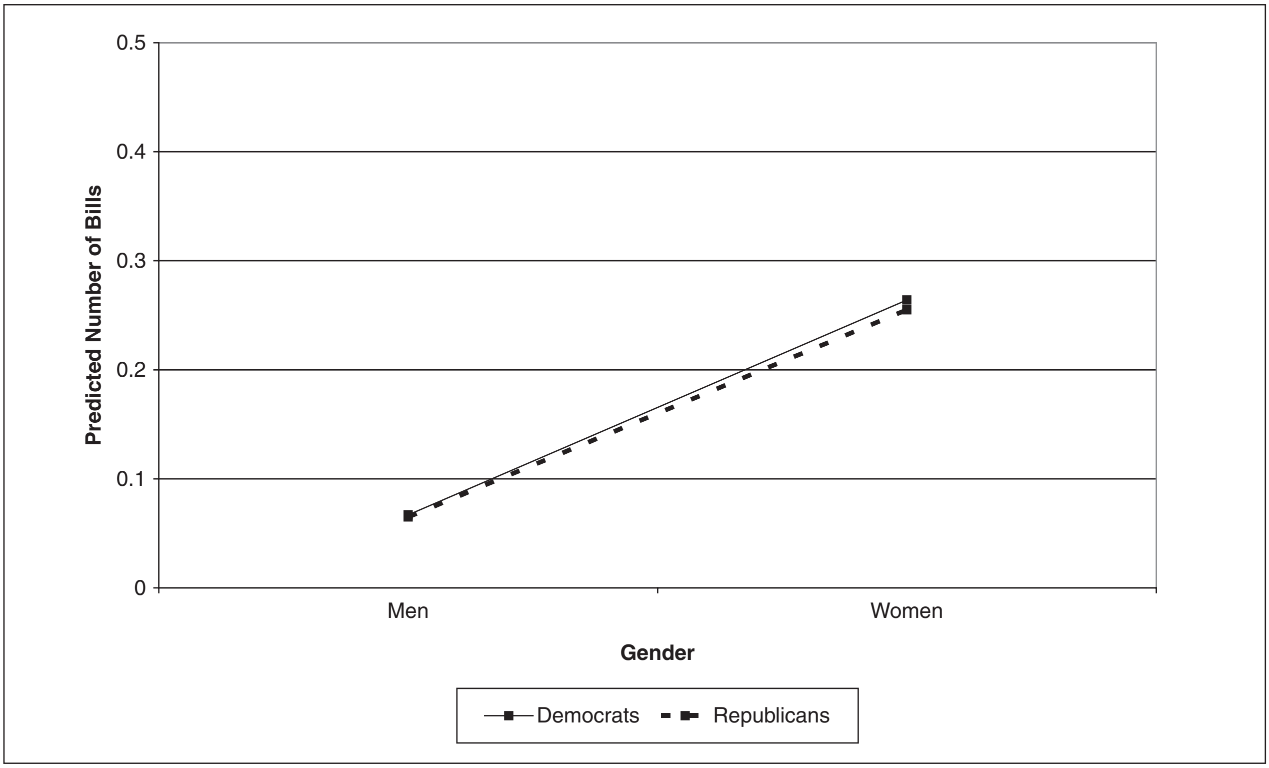

Figure 1. The predicted number of feminist bills sponsored

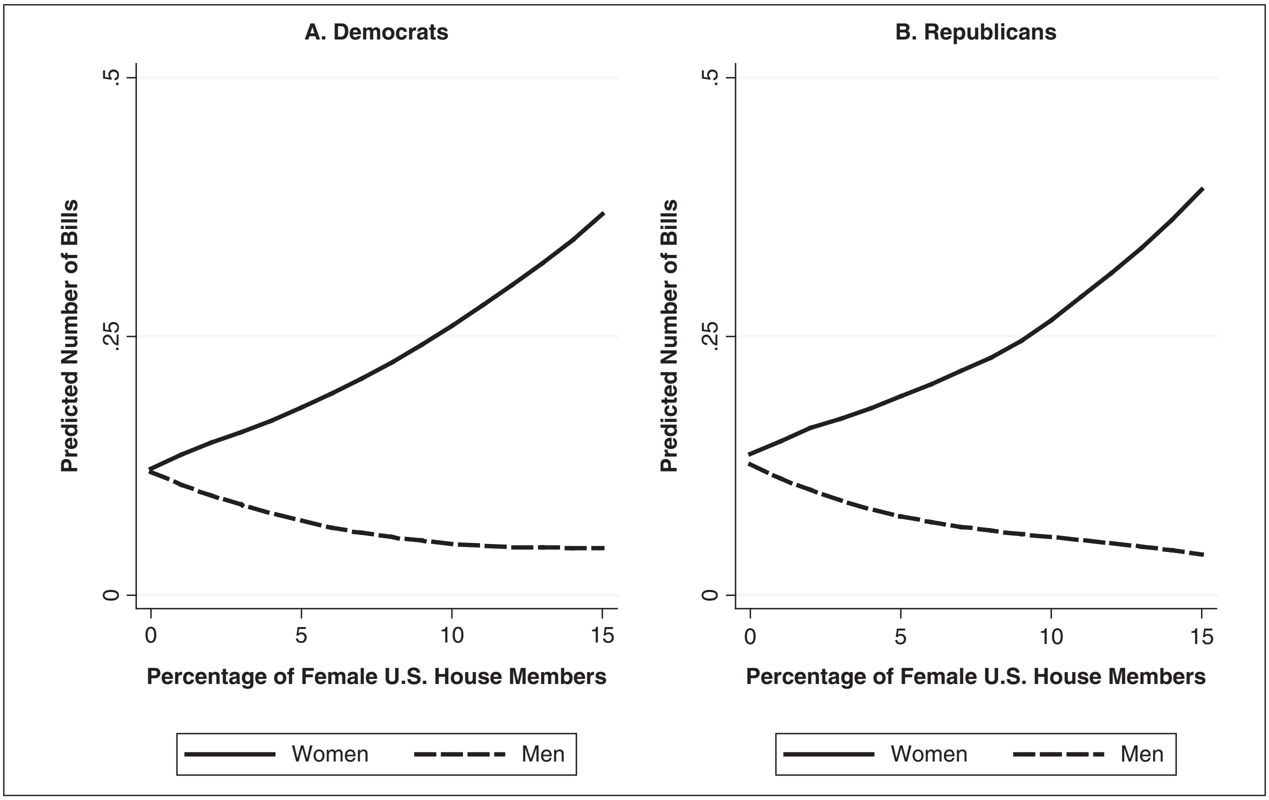

Figure 2. The predicted number of feminist bills sponsored
Note. Conditional coefficients for gender are significant starting at 5 percent of female House members.

Table 2. Negative Binomial Estimates of the Number of Social Welfare Bills Sponsored by Representatives

Variable	(1)	(2)	(3)	(4)
Gender	0.085	–0.491	0.081	–0.478
	(0.171)	(0.391)	(0.173)	(0.392)
Democrat	0.397**	0.381**		
	(0.174)	(0.174)		
1st dimension nominate coordinate			–0.490**	–0.457**
			(0.230)	(0.228)
Number of terms served	0.071***	0.074***	0.069***	0.072***
	(0.026)	(0.026)	(0.027)	(0.027)
Member on a women's issues committee	0.498***	0.487***	0.487***	0.481***
	(0.168)	(0.167)	(0.170)	(0.168)
Retiring member	–0.350*	–0.355*	–0.331	–0.338
	(0.271)	(0.269)	(0.273)	(0.271)
Percentage of women serving in the House	3.655*	–0.229	3.910*	–0.526
	(2.383)	(3.154)	(2.416)	(3.195)
Gender × Percentage of Women Serving in the House		6.973*		6.811*
		(4.267)		(4.285)
Total number of bills sponsored	0.028****	0.028****	0.028****	0.028****
	(0.005)	(0.005)	(0.005)	(0.005)
Constant	–1.142	–0.871	–0.959	–0.701
	(0.299)	(0.335)	(0.280)	0.317
N	206	206	206	206
Log-Likelihood	–315.32	–314.00	–315.63	–314.38
Likelihood Ratio Chi-Squared	82.56****	85.21****	81.93****	84.45****
LR Test vs. Poisson	29.66****	27.88****	29.95****	28.18****

Note. The standard errors are in parentheses.
*$p < .10$, $p < .05$, ***$p < .01$, ****$p < .001$ (one-tailed tests).

and feminist bill sponsorship emerges though with additional specifications. With constituency fully accounted for, we observe that female members increase the level of substantive representation afforded to feminist issues as the number of women in the House increases. Female legislators may consistently desire to sponsor more feminist legislation, but our results show that it is when this desire combines with higher numbers of women serving in the House that female legislators are most capable of delivering.

Turning to the analysis of the number of social welfare bills proposed by members, the coefficient for gender in model 1 of Table 2 does not reveal a positive and significant relationship between gender and the number of social welfare bills that members propose. One cannot reject the null hypothesis that there is no relationship between the gender of members and the level of substantive representation that they provide with respect to social welfare policy issues. This finding is mirrored in model 3 that employs ideology rather than partisanship to measure members' policy preferences.

The models present a somewhat different story when we turn to those that include the interaction between gender and the percentage of the House whose members are women (models 2 and 4). Regardless of whether it is party or ideology we control for, the coefficient for this variable is positively and significantly related to the number of social welfare bills sponsored by members. As such, women provide higher levels of substantive representation on social welfare issues than their male colleagues—but *only* when women populate the House at relatively high percentages. Additionally, we calculated conditional coefficients and standard errors for gender across the range of the percentage of women serving in the House from 1973 to 2002. These findings, illustrated in Figure 3, reveal that female members were significantly more likely to sponsor social welfare legislation when at least 13 percent of their House colleagues were

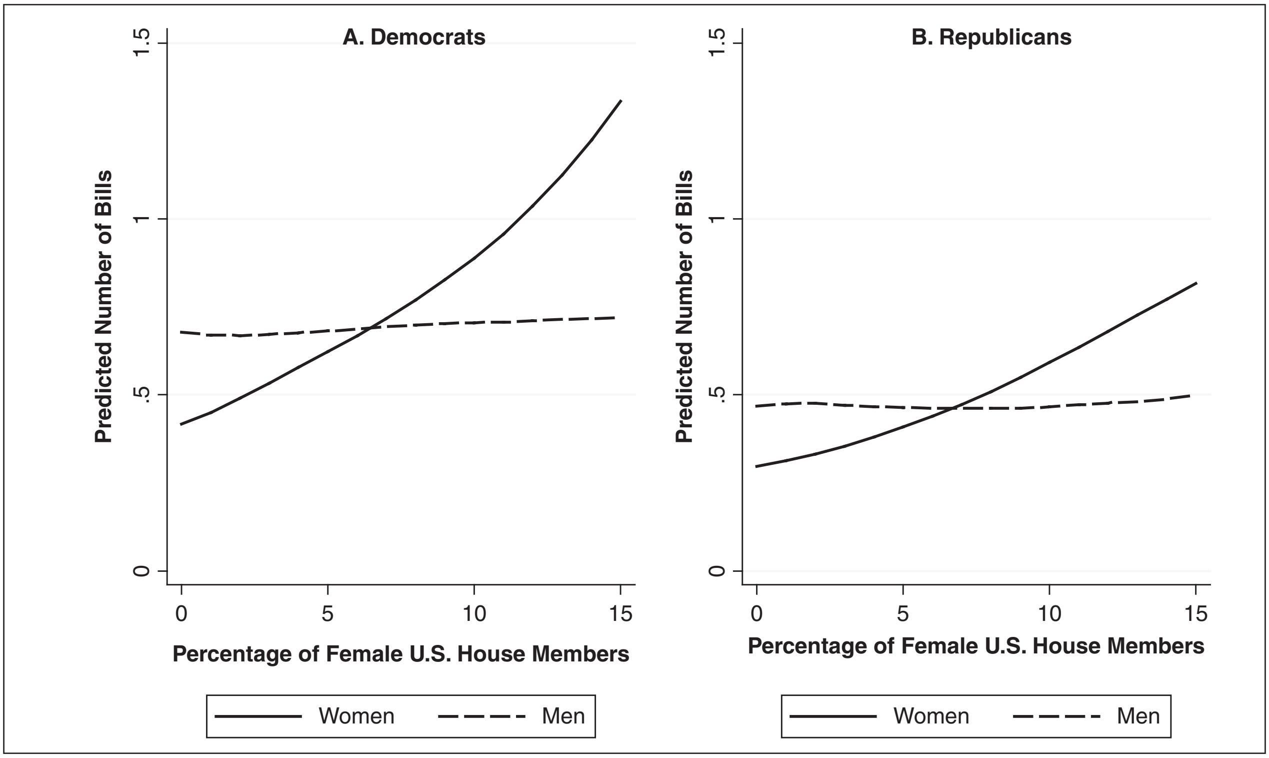

Figure 3. The predicted number of social welfare bills sponsored
Note. Conditional coefficients for gender are significant starting at 13 percent of female House members.

women. Figure 3 provides a sense of the magnitude of this relationship while including the partisanship variable as it is statistically significant in all the specifications. We plot in the number of social welfare policy bills that model 2 of Table 1 predicts Democratic and Republican congresswomen and men to sponsor across the observed range of the presence of women in the House. We observe that at low levels of women serving in the House, men in both parties are actually predicted to sponsor more social welfare bills than women. However, as a higher percentage of members serving in the House are women, the model's prediction for the number of bills sponsored by women increases while that of men remains stable. Female Republicans experience steady increases starting when the chamber becomes 10 percent female. The jump for female Democrats in social welfare sponsorship at this percentage is even more dramatic. At the highest levels of female presence in the House, the predicted number of bills sponsored by women is well above that of men of the same party—twice as high for both Democratic and Republican women compared to their male colleagues. Thus, although there is no additive effect of gender on the sponsorship of social welfare bills, gender affects such sponsorship conditionally on how many women are serving in the House. What is more, this effect is sizeable in magnitude, as Figure 3 indicates.

In terms of how other independent variables in the social welfare models compare with models for feminist bill sponsorship, we see striking similarity but for partisanship and ideology. The number of terms members served and the total number of bills members sponsored are again significantly related to both types of bill sponsorship, as is membership on a committee with jurisdiction over policy issues relevant to women's interests. Models 1 and 2 in Table 2, however, support the explanation that members' partisanship influences their behavior on social welfare issues in that the coefficient for the Democrat variable is positively and significantly related to social welfare bill sponsorship.[20] Including ideology, rather than their partisanship, tells the same story. In Table 2, models 3 and 4 demonstrate that higher levels of conservatism lead members to sponsor fewer social welfare bills in that the coefficient for members' first dimension DW Nominate coordinates is negatively and significantly related to the dependent variable. In contrast, neither the Democratic Party dummy nor the DW Nominate variables were significant for feminist bill sponsorship (see Table 1). This is consistent with Swers's (2002) observation that partisanship is a weaker predictor of representation on feminist issues than social welfare issues.[21] It also bodes well for the relevance of our results beyond gender as even though we sampled to fully account for the effect of constituency and gender we still witness triangulation with Swers's seminal results.[22]

Discussion

These findings are vital for answering the normatively important question of whether female legislators provide more effective substantive representation on women's issues than male congressional legislators. Much prior research has found that female legislators do so. Our results both support and refine this position at the congressional level.

Our methodological approach tackles the considerable challenge of fully accounting for constituency influences in examining women's substantive representation. Our research design ensured that the influence of constituency, which is theoretically likely to be correlated with gender, is removed from the error term of the models that we use to estimate the effect of gender on representation. Additionally, our quasi-experimental design negates any problems of inference due to endogeneity and longitudinal sampling removes both the need to go to state data and ensures we are not generalizing from potentially unique temporal snapshots.

These inferential advantages and full methodological accounting for constituency pay substantive dividends. Our findings show that female legislators advance women's interests more frequently than the male congressional colleagues but our unique data source also reveals that the influence of gender on representation is not always straightforward. When it came to feminist legislation, the gender variable reached levels of statistical significance. That we find such an effect, and that it is substantively large, should reduce concerns that the relationship between gender and representation is spurious due to the correlation between unaccounted for constituency factors and gender and the endogeneity of gender to constituency factors that are included in models of representation. However, it was also the case that for both feminist and social welfare legislation, once the models interact gender and the percentage of women serving in the legislature, only the interaction term was significant.

Across both issue domains then, women provided a higher volume of representation on women's interests; however, they do so when they are surrounded with a relatively high proportion of other women in Congress. Indeed, gender alone was never significant for the social welfare models—regardless of whether or not the interaction term was included. Substantively, these findings highlight how vital it is to take into account the role of women's relative presence in the chamber when it comes to understanding the probability of congresswomen advancing women's interest legislation.

Reports of the demise of critical mass theory may be exaggerated then. Some elements may be undertheorized and undercontextualized; and, interpretive questions about how exactly critical mass effects are realized remain (Beckwith and Cowell-Myers 2007; Childs and Krook 2006; Grey 2006; Reingold 2008). Nonetheless, in the U.S. context, our analysis suggests that as the percentage of women in Congress increases, female representatives are more likely to place women's interests on the agenda. This contrasts with analyses that cast doubt on critical mass effects when finding that individual female (state) representatives who serve with very few other women have more of an incentive/institutional ability to take up women's interests than those who serve with more women (Bratton 2005; see Crowley 2004). At the federal level, and with constituency fully accounted for, this is not the case for sponsoring women's interest legislation.

Of course, our data are drawn from a sample of districts that elected women to the U.S. House. Because these districts differ from "typical" districts (Palmer and Simon 2008), it is possible that the findings presented earlier apply only to members from such districts. With respect to the finding that female legislators provide enhanced substantive representation to women's interests, however, our findings buffet those from cross-sectional studies (e.g., Swers 2002). Whereas the cross-sectional studies are subject to concerns about omitted variable bias, our findings are not. And whereas our findings may be limited because we focus on members from districts that elect women, the cross-sectional studies are not. Together, then, multiple studies employing different methodological approaches reinforce the same conclusion: that female members of Congress enhance the substantive representation that women receive in government.

Authors' Note

The authors contributed equally to this research and are listed alphabetically.

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Notes

1. These measures include attitudinal self-reports of legislative priorities and subsequent bill introduction and passage rates (Thomas 1991; Thomas and Welch 2001; Bratton 2005); reports and interviews from samples of male and female state office holders (Carroll 2001; Dodson 2001), mayors (Tolleson-Rinehart 2001), and female members of Congress (Carroll 2002; Dodson 2006); roll call voting scores (e.g., Leader 1977; Burrell 1994; McCarty, Poole, and Rosenthal 1997; Dolan 1997; Swers 1998; Tatalicvich and Scheir 1993); floor debate (Walsh 2002); hearing behavior (Kathlene 1994); committee transcripts (Dodson 2006); and National Women's Political Caucus vote ratings (Burrell 1994).
2. Indeed, many of the (twelve) district variables Palmer and Simon (2008, 177-213) recently found significant for how friendly particular districts are to women (party and race taken into account) are not included in the studies. So, for example, they find that ten of the district attributes are significantly different in core Democratic districts that have elected a woman as opposed to those who have only elected men. As the listing of constituency controls in the previous paragraph demonstrates, this means that Swers (2002, 44-48) does not include seven of these twelve constituency variables. Poggione (2004, 308) is examining state-level data, and her dependent variables are legislator's policy preferences, but does not include seven of Palmer and Simon's twelve constituency characteristics. Thomas and Welch (1991, 449) do not include eleven of the twelve, Bratton and Haynie (1999, 665-67) do not include eight of the twelve, and Bratton, Haynie, and Reingold (2006, 78-82) do not include ten of the twelve.
3. Clarke (2005, 2009) suggests that because "textbook" determinations of whether or not omitted variable bias is at play are frequently not feasible in everyday research, scholars often cannot know whether additional control variables increase or decrease bias on the coefficient of interest. That we find our solution in experimental design reflects Clarke's (2005, 348-49) solution: "If the logic of control variables is flawed, experimental control must be achieved in another way. These include basing specification on theory, finding 'natural' experiments, and 'controlling' for unmeasured effects through careful sample stratification."
4. Even if it is still the case that female legislators incorporate women's interests into their reelection constituencies to a greater extent than male legislators within districts where all members need to be responsive to women's interests, our research design actually makes it more difficult to pick up this gender effect. This is because both female and male representatives will have the incentive to cater to reelection constituencies that prioritize women's issues. From a hypothesis testing standpoint, this makes it less likely that we will make a Type I error and reject the null hypothesis of no gender effect if it is in fact correct.
5. We employed the biographical directory of members of Congress (McKibbin 1997) available on the inter-university consortium for political and social research (ICPSR study number 7803) for members serving during and prior to the 104th Congress (1995-1996). Subsequently, we relied on descriptions in *CQ Weekly Report* to identify the gender of incoming members. Pairs of members were excluded if one of the pair served in California congressional delegations for the 93rd Congress (1973-1974) or Georgia's district for the 103rd Congress (1993-1994). Lawsuits delayed the creation of new districts in these states during these periods.
6. Another factor relevant to the generalizability of the findings presented in the following is that women are more likely to be elected from districts that are more liberal. Differences of means tests—available upon request—on all members serving in the House from 1973 to 2002 reveal that members from districts electing woman were more liberal than members from districts that did not elect women. Given the general leaning of these districts, and their demand for attention to women's issues (Palmer and Simon 2008), both men and women elected from them have an incentive to prioritize women's interests. Therefore, this makes it harder for us to observe gender differences in this sample, making it less likely that we will make a Type I error and reject the null hypothesis that gender is unrelated to substantive representation if it is in fact true.
7. Many bills introduced by members in the 1970s were "duplicate" or "identical" bills. When a member sponsored two, or three, or *n* identical bills, the bills were only counted as one bill. The intercoder reliability results are available in the "supplemental materials" for this article on *Political Research Quarterly*'s Web site: http://prq.sagepub.com/supplemental/.
8. Difference of means tests conducted by members' gender and then by members' gender and partisanship provide for baseline relationships and are available on *Political Research Quarterly*'s Web site: http://prq.sagepub.com/supplemental/.
9. Data on partisanship and ideology was obtained from the DW-Nominate data that Keith Poole makes available freely at www.voteview.com. We thank him for making these data available (downloaded June 26, 2006).
10. Controlling for party identification and ideology in the same model does not change any of the findings reported in Tables 1 and 2. The findings from analyses with both containing both models are available from the authors upon request.

11. Members were coded as serving on such committees if they served on the Ways and Means Committee, the Education and Labor/Education and Workforce Committee, the Judiciary Committee, and/or the Energy and Commerce/Interstate and Foreign Commerce/Commerce Committee. These are the committees that Swers (2002) identified most often as having jurisdiction over feminist and social welfare legislation.
12. This variable is obtained from the biographical directory (ICPSR 7803) for the 104th (1995-1996) Congress and congresses prior to it; subsequent data were obtained by counting the number of sessions served in the DW-Nominate data provided by Poole.
13. This variable was obtained using ICPSR 7803 and by examining *CQ Weekly Report*'s reporting on announced retirements after 1996.
14. We obtained this information from *Thomas* when collecting data on the dependent variables.
15. In analyses available upon request, we also control for the decade during which members served, employing dummy variables to indicate that members served after the 1970s redistricting cycle (from 1973-1982) and after the 1990s cycle (1993-2002) with the 1980s period serving as the reference category. No changes to findings reported in Tables 1 and 2 occur because of this specification. Additionally, we controlled for whether members were first elected in 1992 and 1994 on the grounds that the dynamics of these elections might have led members to have reelection constituencies (Fenno 1978) favorable to women's interests in the former case (1992) and less favorable to women's interests in the latter case (1994). These alternative specifications did not affect the findings reported in Tables 1 and 2 and are available from the authors upon request.
16. One issue that may occur to readers is that Republican women in the 1990s were, in general, more conservative than such members serving in the 1970s and 1980s. Therefore, we analyzed pairs containing Republican women separately, controlling for their ideology. These analyses show that the findings related to gender in Tables 1 and 2 are the same for this subset of members as they are for all members. These findings are available upon request.
17. The predictions are based on a member serving in his/her first term, who does not serve on a committee with jurisdiction over women's issues, who is not retiring, who sponsors a number of bills equal to the mean of all members and who serves in the House when it is composed of the percentage of women equal to that variable's mean.
18. Readers may be given pause because of the low number of predicted bills by the model. These predictions are due to the specification of values for the other independent variables, such as members are assumed to be serving in their freshman session, that depress the number of predicted bills; additionally, most members do not offer feminist bills, a fact reflected in the variable's median value of 0.
19. Since, as displayed in Figure 1, there is no effect of party on feminist bill sponsorship we do not plot Democrats and Republicans separately.
20. Figure 3 illustrated this showing how female and male Democrats are predicted to sponsor a higher volume of social welfare policy bills than Republicans of the same gender across the entire range of women's presence in the House.
21. Interestingly, in the additive models (1 and 3) of Table 1, the percentage of women in the House is (weakly at $p < .10$) positively and significantly related to the number of social welfare bills sponsored. This finding suggests that as the presence of women increase in the House, all members become more likely to address social welfare issues in the bills that they propose to their colleagues.
22. Readers may wonder whether findings differ across parties. We reestimated the models presented in Tables 1 and 2, analyzing pairs of Democratic women and the men replacing/replaced by them and Republican women and the men replacing/replaced by them separately. The result—available upon request—shows that Democratic women are more likely to sponsor feminist legislation than the men replacing them/replaced by them. This finding holds for Republican women/men. Additionally, the finding that female members are more likely to sponsor social welfare legislation at higher percentages of women serving in the House also holds across parties. The only difference is that the interaction term for gender and the percentage of women serving in the House is significant for Republican women pairs in the analysis of feminist bills but not for Democratic women. This finding suggests that Democratic women sponsor more feminist bills than the men replacing them/replaced by them regardless of the presence of women in the House, while Republican women are more likely to offer such legislation only when they have relatively high percentages of female colleagues. Regardless, the main finding that gender affects substantive representation holds across both parties.

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