Did Rosie the Riveter Send Ms. Smith to Washington?*

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Abstract

Scholars have long argued that women's labor force participation increases their descriptive representation. Labor participation plausibly affects both demand- (e.g., voter norms) and supply-side (e.g., women in political feeder professions) determinants of descriptive representation. While findings are mixed, on balance, women's labor force participation and descriptive political representation correlate, at least in cross-national panel studies. But establishing a causal connection between labor participation and descriptive representation is difficult. The two variables are likely endogenous and relationships between them are almost certainly confounded. I take advantage of a plausibly exogenous instrument for labor participation—county-level variation in U.S. World War II casualties that differentially increased women's work-force participation across localities—to test whether women's labor participation *causes* descriptive representation.

WORKING PAPER

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"*Women administer the home. They set the rules, enforce them, mete out justice for violations. Thus, like Congress, they legislate; like the Executive, they administer; like the courts, they interpret the rules. It is an ideal experience for politics.*"

— Margaret Chase Smith, US Senator

1 Working Women and Descriptive Representation

Margaret Chase Smith was not the first woman elected to a full term in the U.S. Senate but, unlike Hattie Wyatt Caraway, she initially entered the Senate without a widow's appointment.[1] Ironically, she uttered the above words during her 1948 Senate primary campaign, after serving four terms in the House, and on the heels of nearly twenty years working jobs ranging from schoolteacher to newspaper circulation manager (Dismore, 2020). She should not have needed to argue that household administration was good preparation for political office. While Smith was, herself, a trailblazer for white-collar working women, she won her primary, and that year's general election, in the aftermath of a war that increased women's work-force participation by more than 50 per cent. Over the next few decades both working women and female politicians became increasingly common. It seems likely that World War II accelerated the pace of women's entry into the labor force,[2] but did it also establish a beachhead for women in politics? Did the rapid transformation of the gender distribution of the labor force during the war help women to obtain office earlier, and at greater rates, than would otherwise have been possible?

This argument that the war accelerated women's entry into politics fits with prevailing wisdom about the relationship between labor demographics and women's engagement in politics. Recent research finds that women's labor force participation (WLF-participation) correlates with women's tendency to run for—and win—office in democracies (Iversen and Rosenbluth, 2008; Stockemer and Byrne, 2012; Meserve, Pemstein and Bernhard, 2020). The-

[1] Gladys Pyle was the first woman elected to the Senate without first being appointed, although she served only two months, having won a special election.

[2] The debate over this contention in the labor economics literature is voluminous, focuses mostly on the direct effect of wartime labor on subsequent work, and contentious, but recent research tends to provide support—if circumscribed—for this claim (see e.g., Goldin and Olivetti, 2013; Rose, 2018).

oretically, WLF-participation potentially affects both the demand- and supply-side processes determining women's descriptive political representation (WDP-representation). On the demand side, because gender stereotypes affect candidate evaluations (Dolan, 2010), experience with working women plausibly alters attitudes about appropriate roles for women, increasing voter support for female politicians. Women in the labor force also approximate a precondition for a supply of female political candidates, notwithstanding Senator Smith's contention regarding the isomorphism between household and political administration. Politicians are members of the labor force, after all, and professional histories—such as success in business or blue collar bona fides—often feature heavily in campaign pitches. Women may also be inspired to pursue careers, including political ones, by other working women. Evolving norms about WLF-participation are also likely to increase supply by improving familial support for female candidates' aspirations.

But does WLF-participation *cause* women to obtain public office? While it is theoretically plausible that WLF-participation causes WDP-representation, there are other explanations for the observed correlation. First, the causal arrow might point in the opposite direction (Lv and Yang, 2018; West and Duell, 2020; Priyanka, 2020). Alternative causes might also confound the relationship. Most obviously, evolving gender norms that alter women's propensity to enter the work-force (Fernández, Fogli and Olivetti, 2004; Fernández, 2013) could simultaneously change both women's willingness to compete in elections and their tendency to win. As a result, deeper cultural change could cause WLF-participation and WDP-representation to trend together, with neither variable acting as a root cause of the other. Finally, while more recent and comprehensive cross-national studies tend to find positive relationships between WLF-participation and WDP-representation, the overall empirical record is mixed (Stockemer and Byrne, 2012). Despite substantial study, we still do not know if WLF-participation independently causes WDP-representation.

Since WLF-participation and WDP-representation are likely endogenous, and confounding variables almost certainly abound, I use exogenously generated variation in WLF-participation to test if women entering the labor force causes improvements in WDP-representation. Notably, county-level variation in male World War II casualties affected women's propensity to seek, and obtain, employment. High-casualty counties experienced persistently higher women's WLF-participation in the decades following the war (Lee, 2021). I therefore use World War II casualties as an instrument for WLF-participation to test whether high WLF-participation encourages better WDP-representation. Future versions of this draft paper will expand on this analysis by examining demand-side mechanisms, testing the contentions that WLF-participation increases voters'—especially working women voters'—support for women's involvement in politics.

2 What We Know So Far

A deep literature explores the determinants of women's descriptive representation (for a review see Wängnerud, 2009). This work is wide-ranging and examines numerous correlates of WDP-representation, such as electoral rules, economic organization, and religion. Within comparative politics, work has focused especially on WDP-representation's relationships with electoral rules—notably district magnitude (see, e.g. Thames and Williams, 2010) and gender quotas (see, e.g. Krook, 2009)—and socioeconomic factors; most notably WLF-participation. Electoral institutions, especially district magnitude, are robustly correlated with women's representation. While it is difficult to conclusively establish a causal link, given institutional stickiness, the question of whether or not electoral institutions affect WDP-representation generates little controversy.[3] Whether WLF-participation affects WDP-representation is more contested, although the balance of evidence shows that the two are related (Stockemer and Byrne, 2012). But the work on this topic is observational, relying primarily on cross-national panel data (for somewhat recent examples, see Wolbrecht and Campbell, 2007;

[3]The mechanisms underlying this relationship are more contested, and may vary across contexts (Meserve, Pemstein and Bernhard, 2020).

Iversen and Rosenbluth, 2008; Stockemer and Byrne, 2012),[4] and I know of no example of a study that uses a natural experiment or randomized intervention to test WLF-participation's influence on WDP-representation. Indeed, very little—if any—within-case work attempts to estimate a causal relationship between WLF-participation and WDP-representation[5] And, as I note above, there are numerous plausible arguments that this relationship reflects omitted variable bias, or reverse causality, or both.

Nonetheless, theory provides an intuitive and compelling case that WLF-participation should cause WDP-representation. The demand side of the story, as articulated by authors including Norris (1987) and Matland (1998) posits that entering the labor force exposes women to politics, causing them to agitate for political reforms and to seek descriptive representation, while scholars like Dolan (2010) argue that, by normalizing women's work, WLF-participation increases voters' acceptance of women politicians. These mechanisms work together to increase demand for women politicians and increase their viability as electoral candidates. In turn, participating in the labor force produces a supply of women candidates by providing them with the skills, and networks, necessary to effectively compete for political office (Iversen and Rosenbluth, 2008), and, as in the story underpinning the demand side, growing interest in political activism (Matland, 1998).

While there has been some debate about whether the type of WLF-participation matters—for instance, Kenworthy and Malami (1999) contend that only high-status professions open doors to WDP-representation, particularly on the supply side—these arguments point to

[4]This cross-national work is, of course, complemented by a host of single-country studies such as, for the U.S. case, Lawless and Fox (2005).

[5]West and Duell (2020) use an experimental design to show something very close to the converse relationship (they manipulate perceived WDP-representation, not actual WDP-representation) in a recent working paper and find that U.S. women are more likely to apply to jobs when they believe that more women serve in Congress. Also, in related work, Breuer and Asiedu (2017) use baseline survey data to argue that gender-focused employment interventions can improve women's community involvement in Togo, in anticipation of a randomized trial of such interventions (I could not find a paper presenting the results of this trial). Again focusing on the converse causal story, Priyanka (2020) uses an election regression discontinuity design to show that exposure to women in politics likely caused increases in women's labor force participation in India. Saying work doesn't exist is always a dangerous—and admittedly ill-advised—proposition and the author especially welcomes feedback that includes citations to papers using strong designs to test the causal links between WLF-participation and WDP-representation.

three broad hypotheses. First, the demand mechanism implies:

Hypothesis 1. *Voters support women's participation in politics more as women's labor participation increases.*

More specifically, given that working women are likely to be exposed to politics and develop gender-specific political demands:

Hypothesis 2. *Working women are especially supportive of women's political participation.*

Next, turning to candidate supply:

Hypothesis 3. *Women run for political office more often when labor participation is higher.*

Finally, the core policy-relevant implication of both theory and observational work is that:

Hypothesis 4. *Women's descriptive political representation increases when women's labor force participation increases.*

This draft focuses on hypotheses 3 and 4. Future iterations of this paper will test the remaining hypotheses to see how well the data support the supply-side mechanisms that theoretically drive the relationship between WLF-participation and WDP-representation.

3 Research Design

I exploit a natural experiment that exogenously perturbed WLF-participation across the U.S. Specifically, geographic variation in World War II draftee casualties (see figure 1) semi-randomly altered the extent of women's labor force participation during the war. Crucially, while many women exited the labor force when hostilities ended (Goldin, 1991; Rose, 2018), both because men came home to work and because industrial production decreased, casualties produced labor disruptions that outlived the war. Women and other historically under-represented groups, such as African Americans, filled these labor gaps (Ferrara,

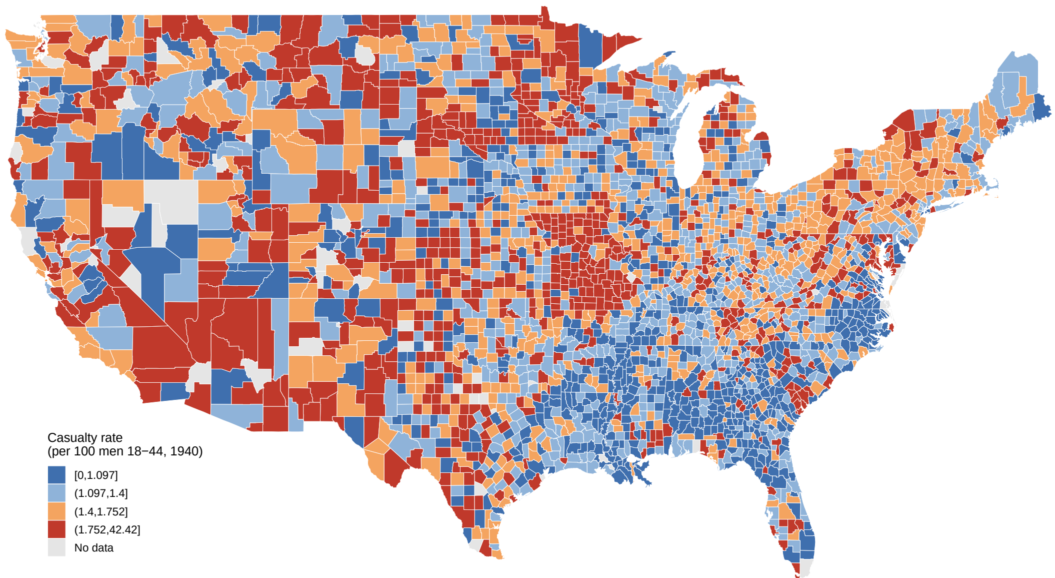

Figure 1. WWII draftee casualty rates, by county. Counties are grouped into quartiles by casualty rate (deaths per 100 men aged 18–44 in 1940). Source: ICPSR 38927; 1940 Census.

2020).[6] In turn, casualty-induced variation in WLF-participation generated sustained geographic differences in the proportion of working women across localities. Lee (2021) uses a differences-in-differences design to show that a one percentage-point increase in a county's WWII draftee casualty rate increased WLF-participation in that county by 2.5 percentage points in 1950, followed by a slight drop-off in 1960, and then a slow increase to a peak of 10 percentage points in 2000. The initial effect was driven by single women, while the longer-term trend represented the entry of married women into the work-force. Lee (2021) argues that this pattern reflects an acceptance of working women through learned experience. In short, exogenous shocks gave people in certain counties more experience with working women, resulting in bigger changes in gender norms about work in those counties than in others. Over time, these war-driven differences in norms compounded, leading to substantial cross-county variation in WLF-participation.

This is a fascinating finding, in and of itself, but it also provides leverage over the question of whether or not WLF-participation causes WDP-representation. WWII draftee casualty rates predict, and likely cause, post-WWII WLF-participation. They were also generated by a process that plausibly affects WDP-representation only through draftee casualty rates' effect on WLF-participation.[7] In other words, WWII draftee casualty rates are a plausible instrument for WLF-participation. Given the size of the effect that Lee (2021) finds, they are also likely to provide a strong instrument. I should therefore be able to use the county-level draftee casualty rates depicted in figure 1 to conduct a more convincing test of the causal supposition in hypothesis 4 than has been possible in previous work.

[6]They also became more important to local economies in other ways. For instance, Cook et al. (2021) collect data from the Negro Motorist Green Books to show that, during Jim Crow, hotels accepted more Black clients in areas where WWII casualties among white draftees were greater, compensating for slack in demand for hospitality services from white men.

[7]Casualty rates do exhibit some spatial auto-correlation and are sensitive to the size of counties' African American populations Ferrara (2020), raising the possibility that geo-correlated county characteristics that might affect WDP-representation (e.g., population density) might confound casualty rates. I discuss these issues, and solutions, in the section on control variables, below.

Data

Independent Variable: WLF-Participation

I use U.S. decennial census data to measure the independent variable, WLF-participation. Summarizing the consensus in the literature, Stockemer and Byrne (2012) argue that one should measure women's labor force participation relative to that of men, rather than in absolute terms, when investigating the relationship between WLF-participation and WDP-representation. A relative measure captures women's contribution to the labor force, while washing out broader determinants of employment, such as local economic development. It therefore provides a context-independent picture of how integral women are to the labor force in any given place or time. I calculate WLF-participation as the number of women in the labor force, divided by the total size of the labor force, for each county, in each decade.

Dependent Variable: WDP-Representation

For the post-1967 period I use the universe of state legislative election outcomes assembled by Bernhard and de Benedictis-Kessner (2021), mapped to counties via geographic crosswalks from de Benedictis-Kessner and Warshaw (2020). For the pre-1967 period I use Galofré-Vilà's digitization of Cox (1996), which records the name, state, county, and terms served for 6,466 women elected to state legislatures in 48 continental states between 1895 and 1997. The post-1967 data cover 141,116 lower- and upper-house elections from 1967 to 2016 while the Cox data add pre-1967 coverage back to the early twentieth century, allowing me to obtain a pre-war baseline for a differences-in-differences analysis.

Like Bernhard and de Benedictis-Kessner (2021), I focus on state and (potentially, in future drafts) local races because "...these races are the right place to assess when women [enter or] drop out of the "pipeline" to power, and can give us a sense for what the gender composition of the national candidate pool may look like in years to come." These data provide a picture of where women first become politically active as candidates and major office-holders. Because the barriers to entry are lower, these races also provide a more sensitive measure of women's ability to win elections than national-level races. For one thing, the single member district plurality elections used for Congressional elections in the U.S. make it difficult for women to break into politics at the national level—even when demand for women candidates is as strong as it is for men—because incumbency advantage is so overwhelming in the U.S. that it makes it hard for anyone to unseat the (mostly) men holding those seats (Schwindt-Bayer, 2009; Meserve, Pemstein and Bernhard, 2020). While most state and local elections use similar institutions, the lower value of these offices, and resulting higher turnover, provides a more sensitive measure of women's ability to win office, simply because incumbents are less likely to squat on these seats. State legislative districts are also smaller than Congressional districts, providing a more fine-grained mapping between counties and political offices than would be possible if I focused on national offices.

My unit of observation is the county-decade. Legislative districts cross counties in myriad ways, somewhat complicating the coding of the dependent variable. In this draft I code the DV two ways. First, I code a binary variable that equals one if any district that overlapped the county elected a woman in the given decade, and zero otherwise. This is a crude measure—for example, consider a county where a woman was elected in one election in a district covering 2 per cent of the county's population, and all other elections were won by men—but is easy to understand and is sensitive to women's entry into office. The second version of the DV attempts to measure the proportion of elections in which the typical county was represented by a woman, across each decade. Specifically, I code women winners as 1 and men as 0, multiply these values by the proportion of the county covered by the given district, and then average across all races contested in districts overlapping the county during the given decade. In a county covered by exactly one district, this variable would measure the proportion of elections won by women candidates in each decade. Under more complicated configurations, it weights this proportion by the percentage of the county population represented by each district.

Instrument: WWII Draftee Casualty Rates

I draw my measure of WWII draftee casualty rates from (Ferrara, 2020, 2024), who combined WWII Enlistment Records—which provide information on draftee name, serial number, age, education, race, marital status, residence, date and place of enlistment, and pre-war occupation—with the WWII Honor List of the Dead and Missing, which include name, state and county of residence, cause of death, and serial number—to create a county-level measure of draftee casualties during the war. These data cover Army, Army Air Force, Navy, Marines, and Coast Guard and include around 88 per cent of total US WWII deaths.

I matched these data with U.S. census data to compute county-level casualty rates, specifically I divided each county's casualty count by its population of men between 18 and 44, as recorded by the 1940 census. I also generate a logged version of this variable, given its highly skewed distribution.

Control Variables

While I cannot control for all variables that potentially confound the relationship between WLF-participation and WDP-repreesentation, Palmer and Simon (2012) identify urban environments with educated, wealthy, and diverse voters as those most likely to elect women to Congress. Therefore I include logged county population and the proportion of white-identifying adults in the county in all models to capture urban areas and (inverse) diversity. I will include measures of education and income in future analyses.[8] I also include county-level democratic vote share in state legislative elections, as calculated by de Benedictis-Kessner and Warshaw (2020), to account for the common finding that left parties elect more women than those on the right.[9]

While county draftee assignment to casualty-causing battles was largely random, Lee

[8] I was focused on census variables with good pre-WW2 coverage when I first assembled my data and therefore do not have these variables in my current dataset. Converting population to population density would also make sense.

[9] I currently only have this variable post-1967, so this control only appears in one analysis that focuses only on the post-1967 data.

Variable	Mean	Median	Std. Dev.	Min	Max
Any Woman Rep.	0.30	0.00	0.39	0.00	1.00
Wt. Pct. Woman Rep.	0.12	0.00	0.22	0.00	1.00
WLF Participation	0.47	0.47	0.10	0.00	1.00
Log Population	9.87	9.83	1.07	3.74	15.22
Prop. White	0.88	0.98	0.18	0.14	1.00
Casualty Rate	1.62	1.40	1.42	0.00	42.42
Log Casualty Rate	0.32	0.34	0.57	-3.47	3.75

Table I. Summary statistics.

(2021) finds that casualty rates were higher in counties with more whites, fewer working-age women, a more urban setting, fewer educated men, and fewer men in agriculture. These results were likely driven primarily by two factors. First, the military systematically discriminated against Black soldiers and tended to put them in non-combat roles; this meant that Black soldiers were less likely than their white counterparts to become casualties of war (Ferrara, 2020). Second, because draftees were assigned to units at the state level, counties located close to one another tended to have similar casualty rates; while which unit was sent to the bloodiest battles was idiosyncratic, geographically proximate counties tended to experience similar fates. Looking only at Southern states, (Ferrara, 2020) argues that county casualty rates for white soldiers are "as-if random" once one includes state fixed effects, effectively controlling for the geographic autocorrelation induced by state-level mustering.[10] Thus I include state—or county—fixed effects in all analyses.

Table I provides summary statistics for the variables that I used in the analysis.

Estimation

This section describes my estimation strategy. I estimate every model described here twice, once for each coding of the dependent variable. Every model is linear. This is not necessarily the appropriate functional form for the DVs; future iterations of this paper will include

[10]He tests this proposition by regressing casualty rates on a host of county characteristics and shows that only the African American share of the population, and the share of volunteers, systematically predict casualty rates after including state fixed effects. In the full sample, county size also matters, but county fixed effects also adjust for this issue during analysis.

robustness checks to ensure that this decision is not driving results.

Baseline Regressions

To set a baseline, I start by estimating a sequence of decade-specific regression models. So, for each t, I fit

$$y_{ct} = \alpha_{t,s(c)} + \beta_t \cdot \text{WLF-participation}_{ct} + \mathbf{x}_{ct}\boldsymbol{\gamma}_t + \boldsymbol{\epsilon}_t, \tag{1}$$

where y_{ct} is the WDP-representation value for county c in decade t, $\alpha_{t,s(c)}$ is a decade-specific state intercept, β_t is the decade-specific coefficient for the marginal relationship between WLF-participation and WDP-representation, $\mathbf{x}_t$ is a set of control variables for county c in decade t, and $\boldsymbol{\gamma}_t$ is a decade-specific vector of control coefficients. These models serve as a baseline for the instrumental variables regressions. Fitting multiple cross-sectional models also provides for a simple way to allow for the possibility that the relationship between WLF-participation and WDP-representation changes over time, at the cost of some efficiency in estimation.

I also fit a two-way fixed effects regression,

$$y_{ct} = \alpha_c + \lambda_t + \beta \cdot \text{WLF-participation}_{ct} + \mathbf{x}_{ct}\boldsymbol{\gamma} + \epsilon_t, \tag{2}$$

to estimate the average marginal relationship between WLF-participation and WDP-representation across the observation period.[11] This would arguably be the industry-standard model for these data, if one were attempting to estimate a causal relationship between WLF-participation and WDP-representation in the absence of a good instrument.

[11] Note that most of the control variables are time-invariant, and fall out. Only democratic vote share, in the specification using the second DV specification, is time-varying.

Lee Replication

Before using WWII draftee casualty rates to instrument for WLF-participation, I verify that the instrument functions as Lee (2021) describes: casualty rates should predict female labor force participation within counties over time, but not in simple cross-section. I replicate her event-study first stage using individual-level IPUMS microdata, estimating the model

$$\mathrm{LFP}_{ict} = \alpha_c + \lambda_t + \sum_{t \neq 1940} \beta_t \cdot \tilde{C}_c + \mathbf{x}'_{it}\boldsymbol{\gamma} + \epsilon_{ict}, \tag{3}$$

where $\mathrm{LFP}_{ict} \in \{0, 1\}$ is a labor-force participation indicator for woman i in county c in census decade t; α_c and λ_t are county and decade fixed effects; and $\tilde{C}_c$ is the standardized log draftee casualty rate for county c, which is time-invariant. Interacting $\tilde{C}_c$ with decade indicators (with 1940 as the reference category) captures how the within-county association between casualty rates and individual WLF-participation changes, relative to the pre-war baseline, over time. Individual controls $\mathbf{x}_{it}$ include age group and race. Standard errors are clustered at the county level and observations are weighted by census person weights. This specification follows Lee's equation (21) and her Table A6 / Figure 11A.

Event Study Differences-in-Differences

Having established that casualty rates predict within-county growth in WLF-participation (spoiler), I use this variation to identify the causal effect of WLF-participation on WDP-representation. I proceed in two steps: a first-stage event study linking casualty rates to county—as opposed to individual, as in Lee (2021)—WLF-participation, and a reduced-form event study linking casualty rates directly to WDP-representation.

First Stage

The first stage establishes the instrument at the county-decade level. For county c in census decade t, I estimate

$$\text{WLFP}_{ct} = \alpha_c + \lambda_t + \sum_{t \neq 1940} \delta_t \cdot \log(\text{casualty}_c) + \epsilon_{ct}, \tag{4}$$

where WLFP_{ct} is the female share of the labor force in county c in decade t, α_c and λ_t are county and year fixed effects, and $\log(\text{casualty}_c)$ is the log draftee casualty rate, which is fixed at its 1940 value and does not vary across decades. The county fixed effects absorb all time-invariant county characteristics, including casualty_c. The coefficients δ_t quantify the within-county trajectory of women's LFP relative to the 1940 reference year, with standard errors clustered by county.

Reduced Form

The reduced form estimates the effect of casualty rates directly on WDP-representation, bypassing the endogenous WLF-participation variable:

$$y_{ct} = \alpha_c + \lambda_t + \sum_{t \neq t_0} \beta_t \cdot \log(\text{casualty}_c) + \mathbf{x}'_{ct}\boldsymbol{\gamma} + \epsilon_{ct}, \tag{5}$$

where y_{ct} is one of the two WDP-representation outcomes for county c in decade t; α_c and λ_t are county and year fixed effects; and $\mathbf{x}_{ct}$ is a vector of time-varying controls (Democratic vote share in state legislative elections, available post-1967). The reference decade t_0 is 1940 for the binary any-woman-representative outcome, which is available from 1940 onward, and 1970 for the weighted percentage outcome, which is available only from 1970 onward. The coefficients β_t represent the average difference in WDP-representation between counties with higher versus lower casualty rates, relative to the reference decade, holding county and year fixed effects constant. If WWII-induced WLF-participation gains caused increases in WDP-representation, we would expect $\beta_t > 0$ for post-war decades. The ratio $\hat{\beta}_t/\hat{\delta}_t$ provides a decade-specific implied IV estimate of the effect of WLF-participation on WDP-representation.

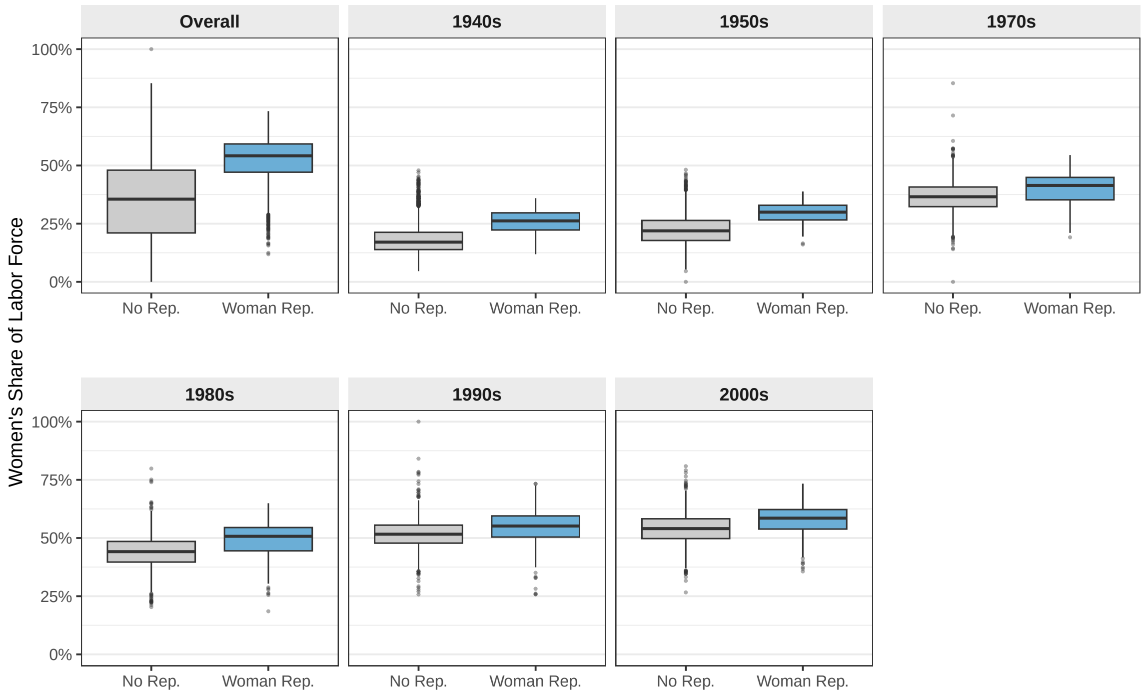

Figure 2. WLF-participation by WDP-representation, overall and by decade. Counties are classified by whether any woman was elected to the state legislature in districts overlapping the county during the given decade.

4 Results

Figure 2 serves as a sanity-check for this initial data analysis. The pooled "Overall" panel shows that, as expected, WLF-participation correlates with WDP-representation at the county level across the full 1940–2000 period: counties represented by at least one woman in the state legislature exhibit higher average WLF-participation. The decade-by-decade panels show that this correlation is visible in each cross-section but tends to be quite subtle, particularly in the later decades.[12] While far from conclusive, this pattern is consistent with the possibility that WLF-participation and WDP-representation trend together over time, rather than causing one another within decades.

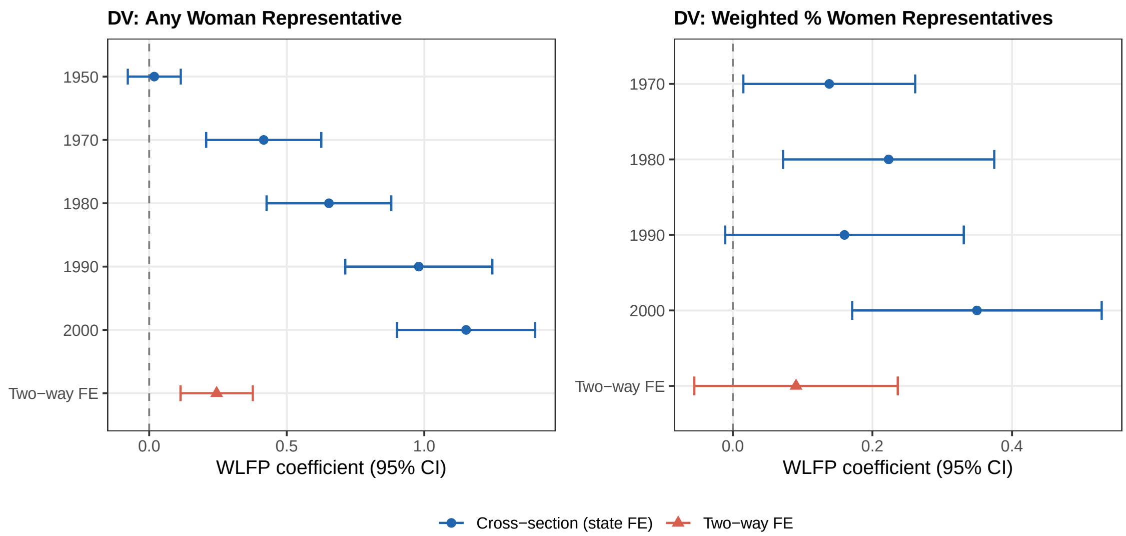

Figure 3. WLF-participation coefficients (95% CIs) from decade-specific cross-sectional OLS (state fixed effects, county population and Black population share controls) and a two-way fixed effects panel model (county and decade fixed effects). Left panel: binary any-woman-representative DV (1950–2000). Right panel: weighted percentage DV (1970–2000 only). Standard errors clustered by county.

Baseline Regressions

Figure 3 presents the WLF-participation coefficients—as a reminder, this is coded as the number of women in the labor force divided by the total number of people in the labor force—for the baseline cross-sectional regressions and two-way fixed effects models. These include models for both versions of the dependent variable. Note that the decade-by-decade models include state fixed effects, in addition to controls. The appendix provides full regression tables.

The results of the regressions reinforce the preliminary visualizations in the box-plots. First, excepting 1950, the coefficients in each of the pooled models are statistically significant. Overall, as in figure 2, we find that WLF-participation and WDP-representation are correlated, and especially when using the any-woman-representative DV, this effect grows over time. This relationship even survives two-way fixed effects with the binary variable. Thus, at

[12]1960 is missing because I ran into an issue pulling census data for 1960 that was comparable to the rest of the census data. I may fix this issue in future iterations of the paper.

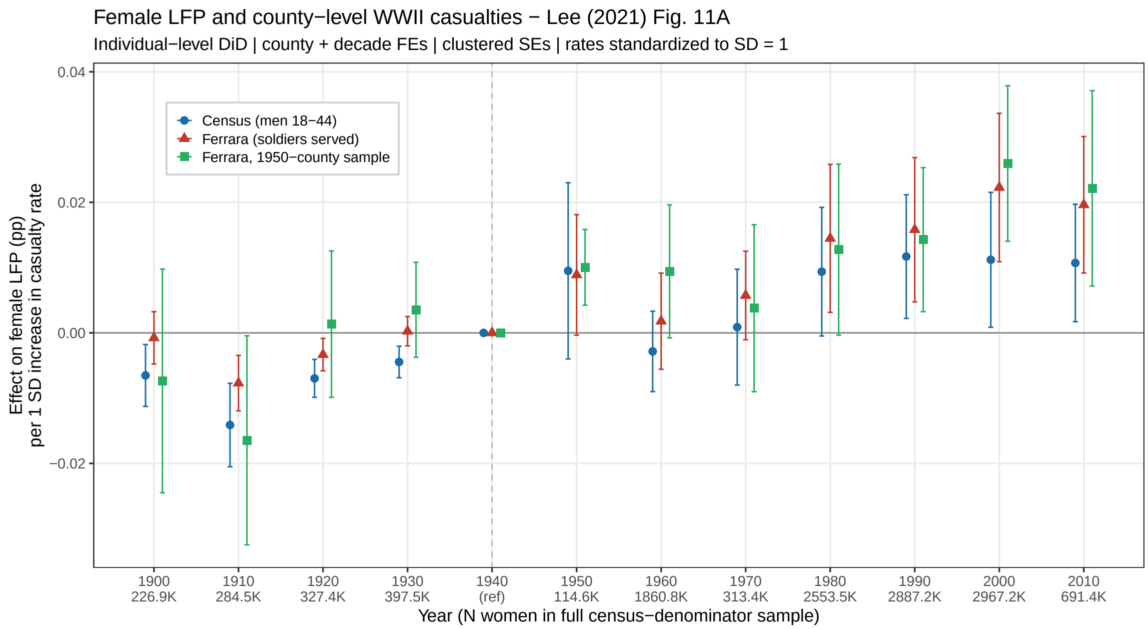

Figure 4. Replication of Lee (2021) Figure 11A. Estimated decade-specific coefficients ($\hat{\beta}_t$) from equation 3: the within-county association between the standardized log draftee casualty rate and individual female labor force participation, relative to 1940. Three specifications are shown: the census denominator (casualties per 100 men aged 18–44), the Ferrara denominator (casualties per 100 soldiers served), and the Ferrara denominator restricted to counties identifiable in the 1950 census. Full coefficients for the census-denominator specification are reported in Appendix Table V. Error bars are 95% confidence intervals based on standard errors clustered by county.

first glance, the argument that WLF-participation causes WDP-representation seems plausible. In particular, the two-way fixed effects model is the traditional test of this hypothesis, and finds a positive and statistically significant, relationship, at least with one DV. That said, two-way fixed effects cannot rule out that the correlation is driven by county-specific trends in a third variable that causes both WLF-participation and WDP-representation (e.g., evolving gender norms), nor can it rule out reverse-causality.

Lee Replication

Figure 4 displays a replication of her dynamic difference-in-differences analysis of the relationship between casualty rates and WLF-participation, using individual-level census data. The analysis computes casualty rates using a variety of denominators, which have subtle

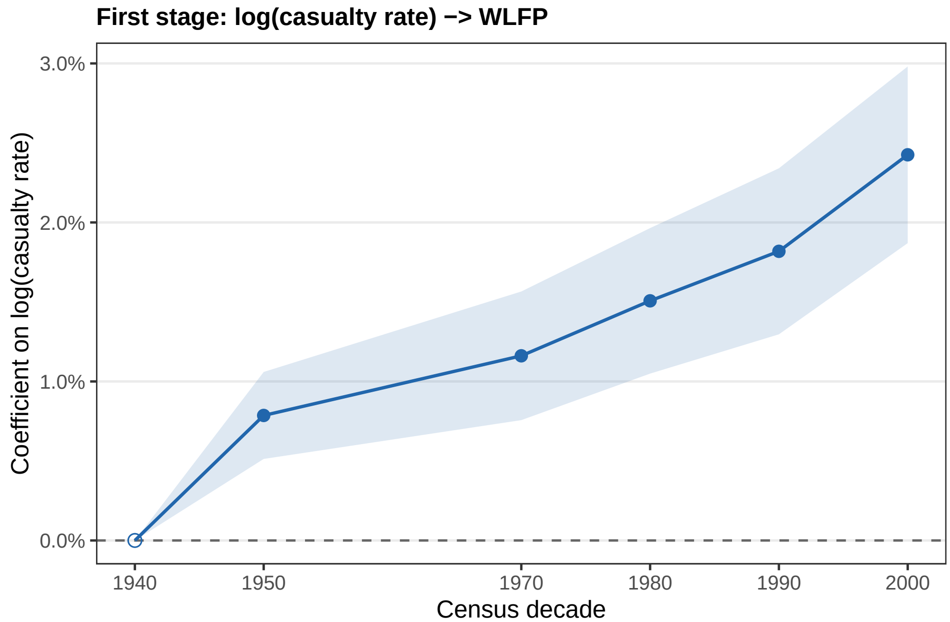

Figure 5. First-stage event-study DiD: estimated decade-specific coefficients ($\hat{\delta}_t$) from equation 4, giving the within-county association between log draftee casualty rate and county female labor force participation relative to 1940. County and year fixed effects absorbed. Shaded band is a 95% confidence interval; standard errors clustered by county. Open circle marks the 1940 reference year (coefficient normalized to zero). Full regression output in Appendix Table VI.

effects on the results. While Figure 4 does not exactly match Lee's figure, it recovers the same overall relationship, especially in later decades, and verifies that women in counties with higher WWII casualty rates tended to enter the labor force at higher rates than their counter-parts, and this effect compounded over time.

Event Study Differences-in-Differences

Figure 5 displays the first-stage results from the event study differences-in-differences analysis. As in Lee (2021), there is a strong positive relationship between WWII casualty rates and WLF-participation, even when we measure this relationship broadly, at the county level. This helps verify that WWII casualties are a good instrument for WLF-participation.

Figures 6 and 7 display the reduced-form estimates of the effect of casualty rate di-

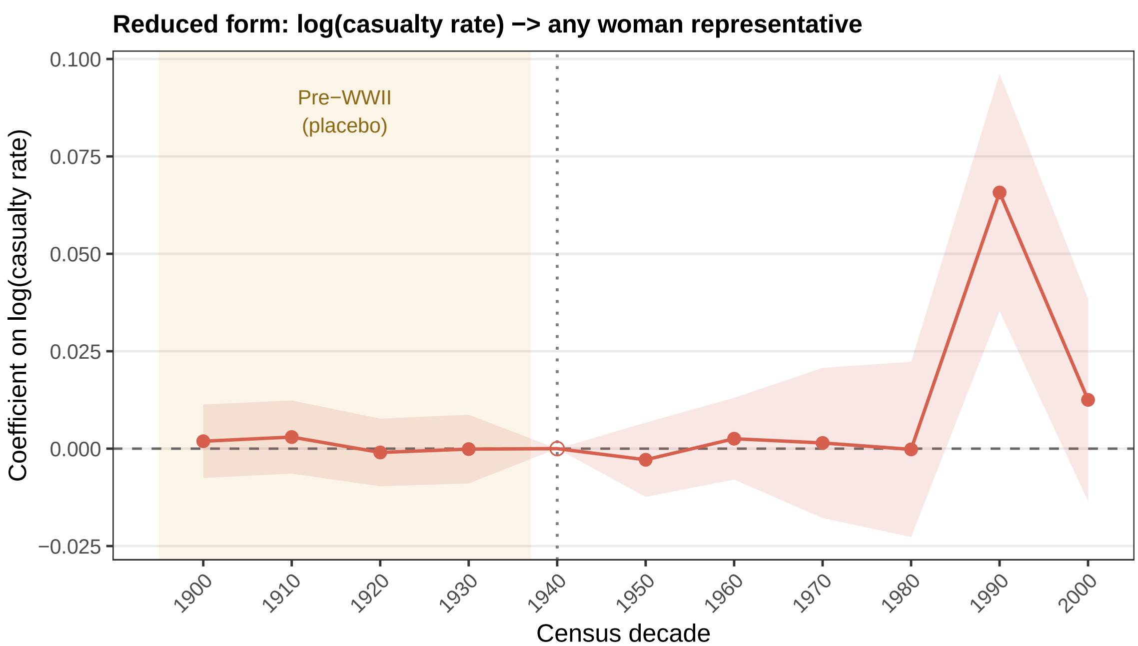

Figure 6. Reduced-form event-study DiD: estimated decade-specific coefficients ($\hat{\beta}_t$) from equation 5 with the binary any-woman-representative outcome, relative to 1940. County and year fixed effects absorbed. Shaded band is a 95% confidence interval; standard errors clustered by county. Open circle marks the 1940 reference year. Shaded region (pre-WWII) extends the event-study back to 1900 as a parallel trends test; coefficients for all pre-war decades are near zero and statistically insignificant. Full regression output in Appendix Table VII.

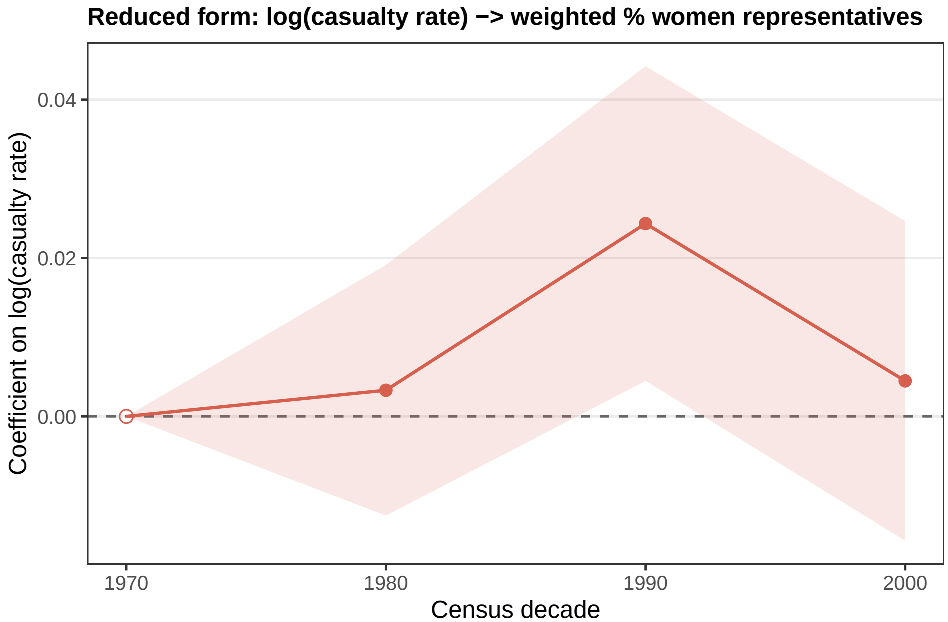

Figure 7. Reduced-form event-study DiD: estimated decade-specific coefficients ($\hat{\beta}_t$) from equation 5 with the weighted percentage of women representatives outcome, relative to 1970. County and year fixed effects and Democratic vote share (time-varying) absorbed. Shaded band is a 95% confidence interval; standard errors clustered by county. Open circle marks the 1970 reference year. Full regression output in Appendix Table VII.

rectly on WDP-representation, using both forms of the dependent variable.[13] In both cases, we find statistically insignificant relationships for all decades except 1990. Thus, on the balance, these results are inconsistent with the idea that WLF-participation causes WDP-representation. The positive and statistically significant coefficient in 1990 warrants more investigation, but perhaps reflects something about the 1992 "year of the woman" election. In any case, the overall result is largely inconsistent with hypothesis 4. This stands in contrast to the two-way fixed effects results. This analysis has some substantial advantages over the two-way fixed effects specification, because it isolates plausibly exogenous variation in WLF-participation and tests to see if it predicts WDP-participation. It therefore has the potential to rule out the third-variable and reverse causation concerns that persist in the two-way fixed effects analysis. At the same time, it sacrifices substantial power, and

[13]Figure 6 includes 1960 because this analysis does not rely on census data for 1960.

therefore may not be able to identify small positive effects.[14]

Figure 8 decomposes the WDP outcome into its supply and selection components using candidate-level data (1970–2000). The left panel examines whether WWII-induced WLF-participation expanded the supply of women entering politics. The right panel shows WWII's effect on female win rate conditional on candidacy, capturing demand-side or gatekeeping effects. Neither effect is statistically significant in any decade (Appendix Table VIII provides full regression results). The null effect for candidacy share contradicts the supply-side mechanism described in Hypothesis 3. Counties where WWII casualties induced larger WLF-participation gains did not produce more female candidates. The null for win rate contradicts a potential demand-side story in which more women ran but faced higher electoral barriers. The 1990 effect visible in Figures 6 and 7 disappears here, further implying that this effect is spurious.

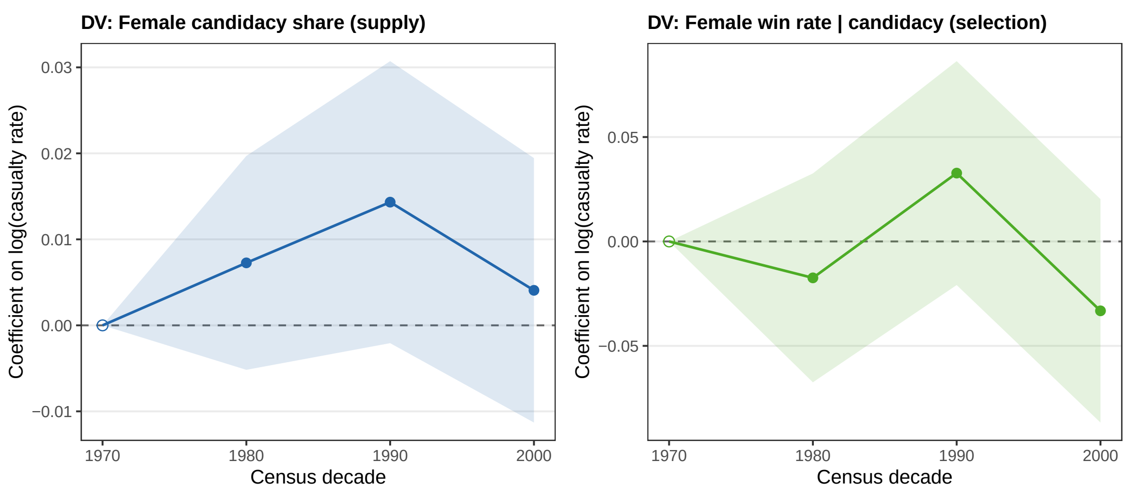

Figure 8. Supply and selection decomposition: reduced-form event-study DiD coefficients ($\hat{\beta}_t$) from equation 5 with female candidacy share (left) and female win rate conditional on candidacy (right) as outcomes, relative to 1970. County and year fixed effects absorbed; standard errors clustered by county. Data from BdBK lower-house candidate records, 1970–2000. Full regression output in Appendix Table VIII.

[14] I ran an analysis, which I will report in future iterations of the paper, that suggests that this analysis is under-powered to detect effects of the magnitude found in the baseline regressions.

2SLS IV Estimates

As a complement to the reduced-form event study, Table IX in the appendix reports panel 2SLS estimates instrumenting WLF-participation with the full set of log(casualty rate) $\times$ decade interactions, using county and year fixed effects. The first-stage F-statistics are strong (26.0 for `any_woman_rep`; 29.1 for `pct_woman_rep`), confirming the instrument is relevant in the within-county panel design. The 2SLS estimates are null in both cases: $b = 1.136$ (SE $= 0.706$, $p = 0.11$) for `any_woman_rep` and $b = 0.292$ (SE $= 0.789$, $p = 0.71$) for `pct_woman_rep`. These null IV estimates contradict the positive and significant OLS panel estimates (0.245 and 0.113, respectively), consistent with the interpretation that the OLS correlation reflects selection — counties where more women work tend to elect more women for confounding reasons — rather than a causal effect of labor force participation on representation.

The Sargan overidentification test rejects consistency of the full instrument set ($p < 0.001$ for `any_woman_rep`; $p = 0.01$ for `pct_woman_rep`), pointing to the 1990 decade-instrument as the source of inconsistency. Dropping the 1990 interaction and re-estimating flips signs (to -0.76 and -0.44) but these estimates remain statistically insignificant, implying that the 1990 anomaly in the reduced form estimates reflect confounding rather than a real effect. Appendix Section B provides full 2SLS results and robustness checks.

5 Discussion

The core question here is: does increasing the proportion of women in the labor force boost women's share of political offices? This question matters for policy because, while difficult to change, WLF-participation is potentially more malleable than other strong correlates of WDP-representation, such as electoral district magnitude. Prevailing wisdom supports this relationship, at least based on the balance of cross-national studies (Stockemer and Byrne, 2012). The data that I examine here have the potential to improve the quality of our answer to this question. They provide a granular picture of the relationship between WLF-participation and entry into lower-level political office as both practices have expanded across four decades in the US. I find little support for the argument that WLF-participation causes WDP-representation. Rather, my analysis is more consistent with the argument that both variables are jointly caused by other factors, such as evolving gender norms about women working. Nonetheless, I stress that this finding is tentative. While the design that I use here has the potential to isolate the causal relationship between WLF-participation and WDP-representation, it may not be powerful enough to identify small effects. This analysis also focuses only on one case, that of the US, which may be unusual. It also focuses broadly on WLF-participation and future work might narrow the focus to white collar jobs, which are arguably more relevant to the supply side of WDP-representation. Nonetheless, the results that I present here call into question the supply-side story underlying the cross-national literature. And, from an historical political economy perspective, they also imply that WWII labor shocks were largely unrelated to women's entry into public office in the US. It is unlikely that Rosie the Riveter sent Ms. Smith to Washington.

Future iterations of this paper will seek to improve on these findings in multiple ways. I plan to:

1. Work to fill in gender data missing from Bernhard and de Benedictis-Kessner (2021) and resolve issues with merging these data with the census,
2. expand the analysis to mayoral and California county races,
3. work to disaggregate WLF-participation by job type,
4. examine the demand side of the story with ANES data about attitudes towards women politicians.

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A Full Regression Tables

Table II. OLS by Decade: Any Woman Representative (DV). Standard errors clustered by county in parentheses. State fixed effects included in all columns.

	1950	1970	1980	1990	2000
WLFP	0.018	0.416^{***}	0.653^{***}	0.980^{***}	1.152^{***}
	(0.049)	(0.107)	(0.116)	(0.136)	(0.128)
Log pop. (1940)	0.054^{***}	0.077^{***}	0.120^{***}	0.102^{***}	0.135^{***}
	(0.006)	(0.008)	(0.008)	(0.008)	(0.007)
Prop. Black (1940)	0.005	-0.190^{***}	-0.181^{**}	-0.068	-0.002
	(0.018)	(0.040)	(0.056)	(0.076)	(0.059)
State FE	Yes	Yes	Yes	Yes	Yes
N	3,050	2,502	2,523	2,175	2,671
R^2	0.406	0.195	0.267	0.310	0.310

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table III. OLS by Decade: Weighted Proportion of Women Representatives (DV). Standard errors clustered by county in parentheses. State fixed effects included in all columns. Dependent variable not available for 1940 or 1950.

	1970	1980	1990	2000
WLFP	0.138^{*}	0.223^{**}	0.160	0.350^{***}
	(0.063)	(0.077)	(0.087)	(0.091)
Log pop. (1940)	-0.001	0.012^{*}	0.008	0.024^{***}
	(0.004)	(0.005)	(0.006)	(0.005)
Prop. Black (1940)	-0.101^{***}	-0.100^{*}	0.002	0.001
	(0.028)	(0.041)	(0.046)	(0.049)
Dem. vote share	-0.000	0.000	-0.001^{*}	0.001^{*}
	(0.000)	(0.000)	(0.000)	(0.000)
State FE	Yes	Yes	Yes	Yes
N	2,488	2,515	2,174	2,669
R^2	0.035	0.070	0.110	0.094

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table IV. Two-Way Fixed Effects (County and Year): Both Outcomes. Standard errors clustered by county in parentheses. Any woman representative covers 1940–2000; weighted proportion covers 1970–2000 only. Democratic vote share included as a time-varying control for the weighted proportion outcome (available 1970–2000 only).

	Any Woman Rep.	Wt. Pct. Women
WLFP	0.245***	0.091
	(0.067)	(0.074)
Dem. vote share		0.000
		(0.000)
County FE	Yes	Yes
Year FE	Yes	Yes
N	15,971	9,846
R^2	0.562	0.496

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table V. Lee Replication: Individual-Level DiD Coefficients. Decade-specific coefficients ($\hat{\beta}_t$) from equation 3, estimating the within-county association between the standardized log draftee casualty rate (census denominator: casualties per 100 men aged 18–44 in 1940) and individual female labor force participation, relative to 1940. Model includes county and decade fixed effects, age group and race controls, and census person weights. Standard errors clustered by county in parentheses. $N = 13{,}121{,}817$ individual observations; 3,049 counties. Additional specifications using the Ferrara denominator are shown in Figure 4.

Decade	$\hat{\beta}_t$
Pre-war (placebo)	
1900	-0.007^{**}
	(0.002)
1910	-0.014^{***}
	(0.003)
1920	-0.007^{***}
	(0.002)
1930	-0.005^{***}
	(0.001)
Post-war	
1950	0.010
	(0.007)
1960	-0.003
	(0.003)
1970	0.001
	(0.005)
1980	0.009
	(0.005)
1990	0.012^{*}
	(0.005)
2000	0.011^{*}
	(0.005)
2010	0.011^{*}
	(0.005)
Reference decade	1940
County FE	Yes
Decade FE	Yes
Age, race controls	Yes

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table VI. Event-Study DiD: First Stage. Decade-specific coefficients ($\hat{\delta}_t$) from equation 4: the within-county association between log draftee casualty rate and county female labor force participation, relative to 1940. County and year fixed effects included. Standard errors clustered by county in parentheses. N = 18,277 county-decades; 3,050 counties.

Decade	$\hat{\delta}_t$
1950	0.0079***
	(0.0014)
1970	0.0116***
	(0.0021)
1980	0.0151***
	(0.0023)
1990	0.0182***
	(0.0027)
2000	0.0243***
	(0.0028)
Reference decade	1940
County FE	Yes
Year FE	Yes

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table VII. Event-Study DiD: Reduced Form. Decade-specific coefficients ($\hat{\beta}_t$) from equation 5 for both WDP outcomes. Any woman representative uses 1940 as the reference decade and extends back to 1900 to display the parallel trends test (N = 31,221; 2,815 counties). Weighted percentage uses 1970 as the reference decade and includes Democratic vote share as a time-varying control (N = 9,846; 2,811 counties). County and year fixed effects included in both columns. Standard errors clustered by county in parentheses.

	Any Woman Rep. (ref. = 1940)	Wt. Pct. Women (ref. = 1970)
Pre-WWII decades (parallel trends test)		
1900	0.0019	
	(0.0048)	
1910	0.0030	
	(0.0048)	
1920	−0.0010	
	(0.0044)	
1930	−0.0001	
	(0.0045)	
Post-WWII decades		
1950	−0.0029	
	(0.0049)	
1960	0.0025	
	(0.0054)	
1970	0.0015	
	(0.0098)	
1980	−0.0002	0.0033
	(0.0115)	(0.0081)
1990	0.0657***	0.0243*
	(0.0156)	(0.0101)
2000	0.0125	0.0045
	(0.0133)	(0.0103)
Dem. vote share		0.0002
		(0.0002)
County FE	Yes	Yes
Year FE	Yes	Yes

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table VIII. Supply and Selection Decomposition: Reduced-Form Event-Study DiD

Dependent variables are female candidacy share (women candidates / total candidates) and female win rate conditional on candidacy, aggregated from BdBK lower-house candidate records. County and year fixed effects; reference year 1970. Standard errors clustered by county in parentheses.

	Cand. share (ref. = 1970)	Win rate (ref. = 1970)
1980	0.0073	−0.0174
	(0.0063)	(0.0255)
1990	0.0143	0.0327
	(0.0084)	(0.0274)
2000	0.0041	−0.0332
	(0.0078)	(0.0273)
County FE	Yes	Yes
Year FE	Yes	Yes
Observations	10,745	7,282

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

B 2SLS Results

Table IX. Panel 2SLS and OLS: WLFP → WDP (County and Year FEs)

	OLS	2SLS	First-stage F
Any woman rep. (1940–2000)	0.245***	1.136	26.0
	(0.067)	(0.706)	
Wt. pct. women (1970–2000)	0.113	0.292	29.1
	(0.074)	(0.789)	
County FE	Yes	Yes	
Year FE	Yes	Yes	

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors clustered by county.

Table X. Panel 2SLS Robustness: Drop 1990 Instrument

	Full instr.	Drop 1990	Sargan (full)	Sargan (drop 1990)
Any woman rep.	1.136 (0.706)	−0.761 (0.754)	$p < 0.001$	$p = 0.012$
Wt. pct. women	0.292 (0.789)	−0.442 (0.813)	$p = 0.010$	$p = 0.343$

Standard errors clustered by county in parentheses.