Closing Gender Gaps Through Workplace Diversity: The Intergenerational Effects of World War I*

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Abstract

This paper combines personnel records of the U.S. government with census data to study how exposure to female representation at work can persistently reduce intergenerational gender gaps in labor market outcomes. Exploiting city-by-department variation in the sudden expansion of female employment during World War I, we find that daughters of civil servants exposed to female co-workers are more likely to work later in life. This effect extends beyond public sector employment and clerical work, reducing the earnings gap by 12%. Consistent with a broader shift in attitudes toward working women, exposure to female co–workers also made male civil servants more likely to marry working women. We also show that cities exposed to larger increases in female federal workers saw persistently higher female labor force participation in both the public and the private sector. Increasing gender representation within the public sector can thus have broader labor market implications.

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1 Introduction

Labor market disparities by gender remain widespread despite significant progress (Goldin, 2006). A growing number of studies document non-egalitarian attitudes toward female work as one key factor reinforcing differences in labor force participation and occupational choice by gender (Fortin, 2005; Alesina et al., 2013; Bertrand, 2020; Giuliano, 2021).[1]

One approach to ameliorate such disparities in the labor market centers on enhancing gender representation and inclusion, particularly in environments where stereotypical gender roles are deeply ingrained. Scholars have found that the resulting exposure to women at workplace (Bertrand et al., 2019; Dahl et al., 2021; Greenberg et al., 2024) or in public office (Beaman et al., 2009) normalizes female work and leadership, with the potential to break down gender stereotypes. However, there is scant evidence on whether the effects of increased gender representation diffuse more broadly and persistently across generations and space. The presence of such "social multipliers" (Glaeser et al., 2003), however, is important for assessing the effectiveness of policies aimed at generating broader change through greater inclusion.

This paper studies how a sudden shock to the gender composition of a large public organization can persistently increase female labor force participation among those exposed to female white-collar work – daughters of male co–workers. We document the effects of exposing incumbent (primarily male) civil servants to female co–workers during World War I (WWI). This conflict induced a sudden increase in female workers within the nation's largest employer, the federal government. WWI, also known as the "Great War," entailed the nationwide mobilization of American resources to support the military of both the U.S. and its allies (Kennedy, 2004). The war thus increased demand for white-collar workers throughout the bureaucracy, including openings for stenographers, typewriters, and other clerical positions (Smith, 1928). These jobs were often filled by women (Gavin, 1997), exposing federal workers – many of whom were men – to female peers, often for the first time (Nienburg, 1920). This demographic change in the composition of the government workforce was persistent. Even after WWI ended, women played an increasingly prominent role in government work (Figure A.I), suggesting a potential critical juncture that set female work in the government on a path towards parity.

To study how the war-driven increase in female representation in the federal government created intergenerational effects on female labor market outcomes, we leverage micro-level data from the Official Registers of the United States ("Registers"). This source allows us to construct a rich dataset of the near-universe of federal employees for 1913–1921. The Registers provide detailed information on the city and departments where civil servants worked, allowing us to construct a granular measure of gender composition at the city-by-department ("office") level. Using this novel, granular measure capturing sharp changes in the office-level gender compo-

[1] A number of factors have been shown to contribute to past and contemporary gaps, including the unequal distribution of household work (Goldin, 2014; Kleven et al., 2019), differences in attitudes to risk and competition (Niederle and Vesterlund, 2007), and discrimination or harassment against women (Folke and Rickne, 2022; Kuhn and Shen, 2013).

sition, we study how labor force outcomes evolve for the daughters of incumbent civil servants in response to wartime increases in the presence of female co-workers, relative to sons. We link the digitized personnel records from the Registers to the U.S. full-count population census to identify children of civil servants. Linking across census rounds using a comprehensive set of state-of-the-art matching strategies, we then measure the children's labor force outcomes in adulthood in 1940. This approach allows us to observe gender gaps in later-life outcomes among children of civil servants who worked in offices with high vs. low exposure to female co–workers during WWI. Comparing cities with different degrees of exposure to female wartime entrants, we can also study any contemporaneous effects at the local labor market level.

The identification challenge in our setting is that the war-driven increase in female government workers may be correlated with geographical or workplace differences that directly affect gender-related outcomes. Our unique empirical setting allows us to overcome this limitation in several ways. The availability of *office-level* variation in exposure allows us to not only compare outcomes within the same city, but also within federal government departments. This is an advancement of the empirical literature, which relies on more aggregate geographical units of analyses (Giuliano, 2021). Furthermore, we can follow the "epidemiological approach" (see, e.g., Fernández (2011)) by studying movers who face the common local labor market conditions in their destination city but vary by their parental exposure to female co–workers depending on their origin city and parental department of work.

We organize our results into three parts. First, we document for the first time using high-frequency personnel records how WWI led to the rapid emergence of a new female workforce within the federal bureaucracy. Using biennial personnel records, we show that WWI increased the share of the female federal workforce by 13 p.p. – a doubling within two years (1917–1919). Linking personnel records to the full-count census, we show that these female entrants are younger, less likely to be married, more mobile, and more educated when compared to women entering before WWI. At the same time, they are less likely to have been part of the labor force prior to entering the civil service, suggesting that the war helped mobilize a previously untapped source of labor.

Second, we provide micro-level evidence for the intergenerational effects of greater gender representation in public office on women's labor force participation. Our main finding shows that exposure to female wartime entrants decreases the gender gap in labor force participation for the children of incumbent civil servants. Comparing labor force participation of children of exposed and non-exposed civil servants in adulthood, we find that a one standard deviation (SD) increase in parental exposure decreases the gender gap in labor force participation by 8% in 1940. The decrease in this gender gap is driven by a higher propensity of daughters to work. In contrast, parental exposure to female co–workers does not affect the labor supply decisions of sons. The results hold using a stringent set of fixed effects restricting the identifying variation to within-city and within-department differences in exposure. The results are comparable using pre-existing differences in office occupational structure as the identifying source of variation in an instrumental variable (IV) design.

Our main results are robust to a wide range of falsification and robustness checks. Leveraging the high-frequency nature of the personnel records, we find that the office-level exposure effects are only driven by those civil servants contemporaneously exposed to the increase in female co–workers. In contrast, we do not find any effects on those who worked in the same (to-be exposed) offices but left prior to WWI. We also find that the exposure effects are strongest for teenage children of incumbent civil servants, consistent with existing findings that experiences during "formative years" can shape lifetime economic behavior (Olivetti et al., 2020). We do not find clear exposure effects for older children of civil servants, who are more likely to have left home and already formed a set of rigid preferences and beliefs.

To ensure that our results are not driven by sample selection, we show that our results are not driven by any particular linking method. Furthermore, the point estimates remain virtually identical after reweighting the matched sample to be balanced across a wide range of baseline characteristics. We also rule out that our results are driven by factors correlated with the increase in female co–workers such as the expansion of federal employment or the increase in the share of clerical positions. The intergenerational spillover effects on female labor force participation of daughters of civil servants are solely driven by the parental exposure to greater female representation at their workplace.

After establishing the key result, we study if parental exposure to female co-workers also affected women's marriage, fertility, and schooling outcomes – determinants of female labor supply. Consistent with the fact that female work during our study period was dominated by unmarried women (Goldin, 1991), daughters of exposed parents are less likely to be married and more likely to have fewer children. We also test for changes in human capital, and find weak evidence that parental exposure to female co–workers increased the education of daughters. In a mediation exercise, these observable differences in marriage status, fertility, and schooling cannot fully explain the increased labor force participation of exposed daughters. Parental exposure to female co–workers thus increases labor force participation above and beyond affecting the proximate determinants of labor force participation – consistent with more favorable attitudes towards female work.

Having documented the effects of exposure to female workers on the daughters of incumbent civil servants, we next introduce a simple framework to guide the discussion of channels through which these intergenerational effects may operate. The reduced form effects of greater female representation may reflect improved perceptions about the acceptability of female work – i.e., a change in gender norms. Alternatively, the greater propensity of daughters to work may also reflect the parental transmission of better information about female employment prospects. While it is difficult to fully disentangle both channels, a series of complementary tests indicate that the spillover effects operate beyond only changing awareness of women's employment prospects. First, we find that the spillover effects are not only concentrated in federal government work but also extend to female work in the private sector. Similarly, we observe an increase in female labor force participation even in non-clerical work. Rather than only transmitting information about specific job opportunities for women in the federal

government or as clerical workers, the broader private sector spillover is consistent with parental exposure to female co–workers shifting views about female work in the broader economy. Overall, parental exposure to female co–workers closes the gender earnings gap by 12%, reflecting the move of women not only into the labor force but also into higher-earning positions.

Second, we can probe further if the results reflect greater anticipated *demand* for female workers or changing beliefs about female work. While our reliance on within-city variation ensures that the results are not driven by cross-city differences in labor market conditions, we can rely on "movers" to disentangle the role of exposure to female workers from local labor market conditions correlated with the office-level shock. By 1940, 67% of the civil servant children lived in a city other than their parent's city of residence at the onset WWI, allowing us to separate the role of geography from the role of exposure. Even when comparing children of civil servants who ended up moving to the same city by 1940 (and thus face the same new labor market conditions and institutional environment), variation in their parental exposure to female co–workers at the *origin* city–department remains highly predictive of greater labor force participation. The evidence from the "epidemiological approach" confirms that the parental exposure effects are consistent with improved gender attitudes.

Third, we complement our results on daughters' labor force participation with a revealed preference measure for incumbent men's attitudes: the propensity of unmarried men to marry a working spouse. Consistent with exposure to female co–workers shifting attitudes towards working women, we find that men exposed to a large number of female co–workers are more likely to marry a working spouse. This effect is driven by marriages formed *after* WWI and not mechanically by marriages among co–workers in the government. While none of these individual results can fully separate the role of information from norms, we argue that the intergenerational spillover effects we uncover reflect, at least partly, the vertical transmission of gender norms.

In the last part of the paper, we conclude the analysis by asking whether the increased entry of female workers into the federal government had impacts beyond the vertical transmission through the children of incumbent co–workers we identified. Moving to the local labor market level, we ask whether cities experiencing an increase in female representation in the government also see overall increases in aggregate female labor force participation. We do this by leveraging cross-city variation in female exposure before and after WWI using census data in a difference-in-differences design. While wartime exposure is uncorrelated with female labor force participation before the war, cities with greater female exposure during WWI see a persistent increase in overall female labor force participation. Mirroring our individual-level results, the exposure effects only affect aggregate female labor force participation, leaving male labor force participation unaffected. While this increase is concentrated in the public sector, we also find economically and statistically significant increases in female employment in the private sector – suggestive evidence for broader spillover effects. Finally, we show that the exposure effects on female labor force participation are concentrated in neighborhoods where female wartime workers resided. Leveraging newly digitized data on Women's Clubs from membership directories, we also find that

cities with greater increases in female civil servants saw faster membership growth. The combined findings are consistent with social spillovers through which attitudes and information about women in the workforce are influenced horizontally. Our combined results thus suggest that the sudden war-time increase in female representation in the federal government not only increased female labor force participation among daughters of male co–workers, but as well as women in the broader communities in which those workers reside.

Related literature. Our results have important implications for several strands of the literature. First, we contribute to a growing literature on how increased gender representation can change attitudes (Beaman et al., 2009; Bertrand et al., 2019; Dahl et al., 2021; Miller et al., 2022), especially in male-dominated contexts. Relatedly, by focusing on the implications of exposure to female co–workers among incumbent male workers, our paper relates to an emerging empirical literature that studies how "contact" between groups can help reduce non-egalitarian attitudes and outcomes (Allport, 1954; Schindler and Westcott, 2021; Lowe, 2021).[2] Documenting intergenerational spillovers on labor force participation decisions of children of exposed civil servants, our paper is – to the best of our knowledge – the first to provide evidence for the long-run effects of workplace contact. Our work also resonates with the literature on gender norms (Fortin, 2005; Bertrand, 2011; Olivetti et al., 2020; Miho et al., 2023). This literature has attributed pre-existing differences in gender attitudes and beliefs across groups to cultural determinants that are transmitted across generations (Fernández et al., 2004; Guiso et al., 2006; Alesina et al., 2013; Giuliano, 2021). Our results provide evidence on how the parental exposure to female co–workers can help close gender gaps in labor force participation – a key outcome through which differences in attitudes and norms towards female work may manifest. We provide evidence consistent with both the intergenerational, *vertical* transmission through exposed co–workers, as well as the *horizontal* transmission across neighborhoods in which female war-time civil servants resided (Bisin and Verdier, 2001).

Second, our paper contributes to research on the determinants and historical evolution of gender gaps in economic outcomes, particularly within the labor market (Goldin, 1990, 2006; Bailey, 2006; Bertrand, 2020). By documenting the rise of female clerks within the federal government, we complement existing research on the rise of the clerical sector as one of the gateways to white-collar work in the labor force during the late 19th and early 20th centuries (Costa, 2000; England and Boyer, 2009). We add to this rich literature by documenting how the labor demand shock brought about by WWI contributed to the rise of female clerks. We also contribute to a literature that documents how historical periods of military mobilization affect female labor force participation (Goldin, 1991). A large literature examines the impact of mobilization during World War II on women in the U.S. (Schweitzer, 1980; Acemoglu et al., 2004; Doepke et al., 2015). Our study is – to our knowledge – the first to document how the *World War I* affected the labor market outcomes of American women. This period, termed the "first phase" of women's entry into the labor force by Goldin (2006), marked the transition of women

[2]We note at the outset that given the historical setting and the nature of our data, it is difficult to directly observe contact between individuals in our context. Similar to other studies that document contact effects in historical settings, we infer contact based on geographic proximity of individuals (Schindler and Westcott, 2021; Billings et al., 2021).

from unpaid domestic roles to formal employment. WWI also altered the structure of American goverment, as the scale of federal operations for war mobilization led to many new agencies and a half-million new bureaucrats (Rockoff, 2004). WWI thus had transformative potential for both the federal government and women in the labor market (as suggested in Figure A.I), making the study of this episode as a potential critical juncture particularly interesting. Among the few papers studying WWI are Gay (2023) and Boehnke and Gay (2022), who study how WWI affected women's aggregate labor force participation in France. Our study, in contrast, provides *individual*-level evidence using granular *office*-level variation, which allows us to compare differences in exposure to female co–workers within the same local labor market.

Finally, we contribute to research on the drivers and consequences of diversity within bureaucracies. Much of the existing literature on bureaucratic personnel focuses on selection (Dal Bó et al., 2013) and incentives for improved performance (Khan et al., 2015). Our study relates to recent studies examining the causal factors underlying the demographic composition of the bureaucracy (McCrary, 2007; Moreira and Perez, 2022). To our knowledge, we are among the first to document causal factors leading to the drastic improvement in women's representation within the federal government that occurred during the 20th century. Our paper also relates to a small but growing body of research that investigates the impacts of public sector diversity (Miller and Segal, 2018; Xu, 2023).While these papers study diversity's effects on the performance of public sector organizations, our paper shows that the benefits of increased gender representation extend beyond the workplace, and affect the households of incumbent workers exposed to greater gender diversity.

2 Background and Data

2.1 Context: women and war mobilization in the early 20th century

The United States entered WWI in April 1917. American war engagement was substantial, not only involving 1.2 million American soldiers, but also requiring a nationwide economic mobilization (Keene, 2002).[3] To support its troops and its allies in the war effort, the U.S. provided manpower and resources for war-related government production activities, which included building shipyards, producing munitions, and other industrial activities (Zieger, 2000; Rockoff, 2004). Federal officials were exempt from military service as the federal government took on a central role in mobilizing and coordinating the war.

WWI in turn significantly altered the size and scope of the federal bureaucracy. The U.S. added more than half a million new civilian government workers, mainly during the period 1916-1918 (Rockoff, 2004). Existing departments expanded and new agencies were created to coordinate and administer war-related activities, such as the Food and Fuel Administrations, the War Trade Board, the War Industries Board, and the Bureau of War

[3] In his joint address to Congress leading to the Declaration of War against Germany, President Woodrow Wilson predicted that the war would "involve the organization and mobilization of all the material resources of the country" (April 2, 1917).

Risk Insurance (Cuff, 1980). The first year of the war alone saw 100,000 new government jobs open up, all of which needed to be filled quickly. Thousands of artisans and laborers were needed to meet the demands of a growing bureaucracy (Osborn, 1917). Administrative growth also led to the creation of many new white-collar jobs within the government such as clerks, typists, and secretaries. Across the country, the number of civil servants increased dramatically. For instance, the number of civilian employees in Washington, D.C. more than doubled between 1917 and 1920 but the federal apparatus extended its reach far beyond the capital city to coordinate activities across the nation (Skowronek, 1982). Nationwide, the civilian naval workforce increased from less than 21,000 to more than 100,000, and in government ordnance plants from 11,000 to about 40,000; other branches were likewise greatly augmented (Morgan, 1919). The United States Civil Service Commission – as the agency responsible for personnel administration – was in turn responsible for allocating workers during this era of sudden expansion (Osborn, 1917).

Prior to the war, federal departments retained discretion in making positions available to female employees. Several departments – such as the War and Navy departments – refused to consider women for many jobs within the federal service (Carpenter, 1942).[4] The rapidly expanding federal government provided new opportunities for female workers (Dumenil, 2017). While women had often been explicitly excluded from many positions for decades during the late 19th and early 20th centuries, wartime demands led government bureaus to open up civil service exams to women for the first time. Women who could do the work found themselves readily accepted even by those departments that had been hesitant to employ them. Women made particularly important inroads into clerical jobs. In the 1918 Annual Report of the Civil Service Commission, U.S. Chief Examiner George R. Wales writes of the effect of WWI: "The most notable change in Government personnel brought about by the war is in the employment of women... many of the examinations for technical and scientific positions which in past years have been limited to men may now be taken by women; and the departments are appointing women to these positions. Among the general examinations which war conditions have opened to women are elevator conductor, messenger, junior chemist, computer, bookkeeper, and minor positions in the Ordnance Service at Large" (United States Civil Service Commission, 1918).

The increased presence of women in government work occurred against the backdrop of ongoing changes within the broader labor market. When the U.S. entered WWI, clerical work was on its way to becoming an "archetypal paid job" for women entering the labor force (England and Boyer, 2009). The share of female clerical employment increased from 2.5% in 1870 to 52.5% in 1930 (Goldin, 1984). On the demand side, technological shocks like the emergence of the typewriter and organizational changes related to accounting and bookkeeping allowed more women to compete for clerical jobs in office work (Rotella, 1981; Strom, 1989; Feigenbaum and Gross, 2020).[5] On the supply side, access to secondary school increased the number of young

[4] While the Pendleton Act introduced rule-based exams, there remained discretion in making exams eligible for women.

[5] Other institutional features – such as so-called "marriage bars" that limited the hiring of married women – were instituted in the 19th century, and limited clerical demand to primarily single women (Goldin, 1988).

women who were qualified for clerical work (Goldin, 1990).[6] Our study focuses on how the war-driven demand for clerical work contributed to the increase in female labor force participation.

2.2 Data sources

We combine multiple sources of administrative and personnel data to document the changing patterns of female work in the public sector. In this section, we briefly describe the main sources of this data.

Official Registers of the United States. To study how the war shapes the composition of the civil service, in particular within city-level offices, we rely on the personnel records published in the U.S. Official Registers. The Registers are biennial rosters of the entire workforce of the U.S. federal government, and provide detailed information on a government worker's job title (occupation), salary, department, birthplace, and work location (city).[7] We build on the prior digitization effort by Aneja and Xu (2022), enriching the data to include worker's sex (details discussed in Section 2.3), appointment locations (counties), and harmonized work locations (cities). The latter allows us to identify the department–city of each civil servant, enabling us to construct a granular measure of exposure to female wartime entrants. Figure A.II provides a sample of the records in the Official Registers. For our main analysis, we restrict the sample period to Woodrow Wilson's term (1913–1921), yielding a dataset of 1,041,521 individual-year observations.

Our dataset offers several advantages over data used in previous analyses of gender and labor markets during the pre-1940 period. First, we observe detailed information about workers at a greater frequency than data in other recent studies, as the Registers were published every other year (Withrow, 2021; Eriksson et al., 2022). These data thus increase our confidence that sudden changes in economic outcomes are due to WWI, rather than other time-varying confounding factors, such as the Great Migration or the 1918 influenza epidemic. Second, we observe the salary of each civil servant. Standard data sources for this period – such as the decennial population censuses – typically offer coarser measures of earnings that are imputed from occupational data (Sacerdote, 2005; Collins and Wanamaker, 2021). As discussed in Aneja and Xu (2022), this can be an important advantage, given the potential for discrimination within occupations.

Full-count U.S. decennial Censuses. We make use of the full-count U.S. decennial population Censuses (1900-1940) in all parts of our analysis (Ruggles et al., 2019). To drill down on the changing gender composition of the federal workforce, we first use the 1910 Census. This allows us to study the traits of those who entered as civil servants, based on characteristics before entering the service. In the main part of the paper studying the effects of exposure to female co–workers, we use census data to identify both the children and spouses of serving civil

[6] WWI occurred during America's "High School Movement," a period during which female educational attainment increased from an average of 8 years (for the 1890 birth cohort) to nearly 11 years (for the 1925 cohort) (Goldin and Katz, 2008).

[7] The Official Register was initially compiled by the Department of the Interior, and later by the Census Bureau. Temporary employees who have served for less than six months are not included. In 1923, the Official Register was not published due to federal pressure to reduce costs. The Register resumed in an annual form in 1925, but in a substantially reduced form that covered only senior-level officials within the government.

servants, as well as to examine the adulthood outcomes of the former as a function of parental exposure to female co–workers.[8] For the latter part of our main analysis, we use the rich socioeconomic outcomes that the Census Bureau included in the 1940 Census. Finally, we also use census data to study the aggregate effects of exposure on city-level female labor force participation, as well as spatial heterogeneity within a city.

2.3 Identifying female civil servants

The Official Registers do not uniformly report the pronouns of female workers, making it difficult to consistently identify the sex of civil servants. To overcome this limitation, we use first names to predict a worker's likely sex. We predict an individual to be female if more than half of the individuals sharing the same first name in the 1910 Census self-report to be female, $\Pr(\text{Female}|\text{First name}) > 0.5$. This allows us to impute sex for 87.9% of our sample. For the remainder, the first name is either only recorded with initials (92.6%) or cannot be found in the census, likely due to transcription errors or unique spelling.

To validate our measure, we also matched civil servants directly to the 1910 Census based on their full name and birth state, following Aneja and Xu (2022). While the exact census-linked matches are more accurate, they are only available for the subset (19.8%) of civil servants that could be successfully linked based on name and birth state. When comparing our imputed sex measure with the census-linked measure, both measures coincide with a correlation of 0.97. This gives us confidence in the validity of our imputed measure of sex.[9]

3 Descriptive evidence – the rise of women during World War I

3.1 Expansion of federal employment and female representation

Our main objective in this project is to demonstrate how WWI reduced gender disparities in the labor market by exposing federal civil servants to qualified working women. Before testing the effects of exposure to female workers, we first demonstrate quantitatively how wartime economic mobilization suddenly, substantially, and persistently increased female representation within government work.

We begin this section by showing the sharp increase in the hiring of civil servants that occurred at the onset of the war. Using our high-frequency personnel sample from the Official Registers - the universe of (non-postal) government workers - Figure I, Panel (a) shows that the size of the federal government employment was relatively stable in size between 1907 and 1917. This period of stability was followed by a near *doubling* in total federal employment between 1917 and 1919 (from around 150,000 to over 300,000 workers) after the U.S. entered the conflict in April 1917 (vertical solid line). Notably, this increase in federal employment occurred

[8] We describe the linking of civil servant parents to children in Section 4.3.

[9] When available, we impute the census-linked sex measure to civil servants with missing predicted measures (0.7% of the entire sample).

nationwide, and the growth was not restricted to Washington, D.C. When we omit D.C. from the time series, we still see rapid bureaucratic expansion at the point at which the U.S. entered the war (dashed line).

The raw data presented here also highlights that the growth of the bureaucracy was remarkably persistent. As Figure I highlights, although the size of the federal employment declines slightly from 1919 to 1921, it remains significantly higher than the pre-1917 level even several years after the war concludes. This persistence is evident when using data from the decennial censuses, which allow us to explore federal employment through 1930 (light-gray squares in Figure I, Panel (a)).[10] The pattern is qualitatively similar, showing a gradual expansion in federal employment over time, a sharp increase in the aftermath of WWI, and a subsequent contraction towards a higher employment level. The comparison of the census measures with those derived from the personnel records also highlights the value of using the Official Register data: not only does it deliver a higher frequency, but it allows us to link the rapid expansion of federal employment to the entry of the U.S. into WWI. Moreover, it also reveals a significant under-counting of federal employment by the U.S. Census Bureau.

Having demonstrated empirically the wartime expansion of the civilian federal government, we next use the data to document the change in the gender composition of the bureaucracy. Figure I, Panel (b) shows the share of female civil servants over time. The solid line shows the share of female civil servants using our preferred data, based on the probabilistic measure of a civil servant's sex. The dashed line plots the share of female civil servants using the census-linked subsample, with an exact measure of (self-reported) sex. With both measures of sex, we observe a constant share of female workers from 1913 to 1917, and a jump in the female share coinciding with the entry of the U.S. into WWI. The increase in the female share is economically sizable, reflecting an increase of 13 p.p. Compared to the average pre-war share of female civil servants, this increase corresponds to a doubling of the female share among federal workers within only two years.[11]

Figure II breaks down the expansion of the overall federal employment by clerical vs. non-clerical positions. The expansion of the federal workforce was driven almost exclusively by the increase of the clerical workforce to meet the administrative demands of war. While panel (a) shows an expansion of federal government employment across both clerical and non-clerical occupations, the increase is much more pronounced for clerical positions. Panel (b) breaks down the gender composition by clerical vs. non-clerical positions. Consistent with the historical literature (Dumenil, 2017), the expansion of the female workforce was driven almost exclusively by clerks. The share of women in the federal clerical workforce jumped from around 30% to almost 70% within two years, remaining high even after the war.

[10] Federal workers are identified using the industry code (`IND1950=916`) provided by IPUMS.

[11] The graph also suggests a potential benefit of using the imputed measure. While imputed and census-linked measures track each other well pre-war, the census-linked measure begins significantly undercounting female employment during the rapid expansion of federal employment starting in 1919. As Appendix Figure A.III shows, the rapid influx of new workers coincides with a slight decline in census match rates. With men typically matched at a higher rate, the decline in match rate is likely to be a driver of the undercounting observed in the census-linked measure. Our imputed measure, in contrast, does not suffer from this type of selection bias.

3.2 Characterizing women's wartime service

While we primarily aim to understand how the nation's first large-scale expansion of women into government service can persistently reduce gender gaps in labor market outcomes, the rapid change in gender composition raises an intermediate question: did WWI change the type of women who became civil servants? Answering this question is important as it allows us to better characterize the nature of exposure to female co–workers. Unfortunately, the limited biographical information in the Official Registers imposes data limitations. To overcome these, we match the personnel records to the full-count census. From the censuses, we obtain information about workers' socioeconomic backgrounds, including age, labor force status, marital status, and education.

To study how WWI changed selection, we restrict our personnel panel to entrants into the civil service.[12] We then examine how gender gaps in pre-determined individual characteristics differ for entrants before and after the war.[13] Table I, Panel (a) presents descriptive evidence of differential changes in selection into civil service using census-based characteristics. On average, female civil servants tend to enter government service younger than their male counterparts, particularly during and after WWI. These female entrants are also more likely to be single and white. Interestingly, female civil servants entering during and after WWI are also less likely to have been in the labor force in 1910, prior to entering the war.[14] This suggests that WWI helped disproportionately mobilize female entrants from a previously untapped pool of labor.[15]

We also test for differences in human capital. Unfortunately, pre-1940 censuses contain only a coarse measure of human capital: literacy. As such, we follow Aneja and Xu (2022) by linking the Registers forward to the 1940 Decennial Census, which contains a detailed measure of educational attainment – the years of schooling. Table I, Panel (a) shows that female entrants to the civil service have significantly higher levels of education than their male counterparts, particularly during and after WWI.

Next, we study another aspect of selection: the places that civil servants come from. We use information on the county of appointment in the Registers to examine geographic background characteristics of federal personnel. Mean differences for these characteristics are reported in Table I, Panel (b). While female civil servants tend to come from more urban areas, this difference is muted during WWI. We also investigate whether individuals

[12] We define entrants as individuals that are observed for the first time in the data.

[13] To study changes in selection, we estimate the following regression for individual i who enters in year $t = T(i)$:

$$y_i = \beta_0 \text{Female}_i + \beta_1 \text{WWI}_{T(i)} + \beta_2 \text{Female}_i \times \text{WWI}_{T(i)} + \varepsilon_i \quad (1)$$

Here, y_i is the individual-level characteristic of the civil servant, $\text{Female}_i = 1$ if the individual is female, and 0 otherwise. We define the indicator variable $\text{WWI}_{T(i)} = \mathbb{1}[t \geq 1919]$, corresponding to the first year after the U.S. entry into WWI for which we have personnel records. ε_i is the error term.

[14] This difference is not mechanically driven by the fact that female entrants during and after WWI are younger. The coefficient remains of comparable magnitude when conditioning on entry-age fixed effects.

[15] We also study the origin industries from which *previously working* wartime entrants (i.e., those in the labor force, but outside the government in 1910) came (Table A.I). Consistent with increased manufacturing in response to war, fewer manufacturing workers enter the civil service through WWI. We then focus on a common sector for female employment in the 1910s – professional workers, mainly comprised of teachers and clerks (Goldin, 1990). We see that female civil servants are more likely than men to come from professional service jobs, and more so during and after WWI. Interestingly, this selection is driven by female teachers.

were more likely to move for civil service jobs. We define a civil servant to be from the *same state* if the state of the first working position is the same as the state of appointment. While the gender difference before WWI is small, female entrants during and after WWI are less likely to hail from the same state.[16]

Overall, the descriptive evidence allows us to characterize the large number of women who joined the civil service during and after the war: they are younger, more likely to be single, and less likely to have previously been part of the labor force than their male counterparts. At the same time, they are more likely to have moved from rural areas or from a different state.[17] Most importantly, female entrants during and after WWI were significantly more educated, suggesting that WWI helped mobilize a previously untapped and positively selected pool of talent. We now study how their influx into the federal government affected exposed co–workers and future generations.

4 Empirical strategy

4.1 Measuring exposure to female civil servants

In this section, we discuss our estimation approach for answering our central question: did the large influx of female civil servants affect the intergenerational gender gap in labor force participation? Our identifying source of variation stems from the differential intensity of the war-driven increase in female civil servants across cities and departments. This variation allows us to compare the next-generation outcomes for those in offices that were more and less affected by the sudden influx of women. To measure variation in exposure to female civil servants, we compute exposure intensity for each city-department pair:

$$\Delta\text{Exposure}_{jk} = \left(n^f_{jk1919}/(n^m_{jk19}+n^f_{jk19})\right) - \left(n^f_{jk15}/(n^m_{jk15}+n^f_{jk15})\right) \tag{2}$$

where j denotes the city and k the department. n^f_{jk19} denotes the number of female civil servants in the office jk in 1919, while n^m_{jk19} denotes the corresponding number of male civil servants. With American involvement in WWI lasting from April 1917 to November 1918, and the Official Registers recording personnel as of July 1st, we use the change between 1915 and 1919 to fully capture changes that occurred during the war. Equation 2 thus captures the change in the share of female civil servants across offices between both years.[18]

We restrict our main analysis to cities with at least 20 civil servants in a given office in 1915. This ensures

[16] Results shown in Table I, Panel (b) hold if we restrict the sample to civil servants linked to the 1910 decennial Census and we consider their 1910 county of residence rather than their county of appointment as in the personnel records.

[17] While the focus of our study is identifying the long-run consequences of women's increased presence rather than the determinants of this compositional change, Table A.II further considers a range of additional factors potentially related to selection into government service, ranging from the role of military mobilization (Acemoglu et al., 2004; Goldin and Olivetti, 2013), patriotism (Kang and Rockoff, 2015; Caprettini and Voth, 2022) and the 1918 Influenza pandemic.

[18] Since our empirical analysis focuses on incumbent parents and Equation 2 is constructed based on the *subsequent* change in gender composition, there is no need to rely on leave-out measures to compute the exposure variation.

that our sample focuses on administrative offices with economically meaningful federal staffing, excluding dispersed federal presence driven by agricultural extension workers, lighthouse keepers, or park rangers. Since our empirical strategy will rely on both city and department fixed effects, we further limit the sample to cities that have at least two federal government departments. This leaves us with 70 cities and 8 unique federal departments in our sample.[19] Figure III plots the distribution of the exposure measure across offices. The figure shows significant department-by-city variation in the changing gender composition of government offices, with an interquartile range of 10 p.p. Variation in exposure in our main analysis sample is very comparable to the variation in the full sample, reducing concerns about sample selection.

4.2 Estimating exposure effects

To assess the extent to which exposure to female co–workers helps close the intergenerational gender gap in labor force participation among incumbent civil servants, we estimate the following regression:

$$y_i = \beta_1 \text{Female}_i + \beta_2 \Delta \text{ Exposure}_{J(i)K(i)} + \beta_3 \Delta \text{ Exposure}_{J(i)K(i)} \times \text{Female}_i + \theta_{J(i)} + \tau_{K(i)} + \gamma' x_i + \varepsilon_i \quad (3)$$

Equation 3 relates the labor force participation outcome y_i for child i to the intensity of the parental exposure to female co–workers. Parental exposure $\Delta\text{Exposure}_{J(i)K(i)}$ (defined in Equation 2) varies at the office–level, with $j = J(i)$ and $k = K(i)$ denoting the city and department where the child i's parent worked at time of exposure in 1915, respectively. Female_i is a dummy variable that is 1 for daughters and 0 for sons.

The coefficient β_1 captures the (female-male) gender gap in labor force participation, while β_2 captures the extent to which parental exposure to female co–workers affects the labor force participation decision of sons. Since the aim of this study is to understand how an increase in workplace diversity can help decrease gender inequality, our main coefficient of interest is β_3. This coefficient captures the extent to which parental exposure to female co–workers affects the children's gender gap in labor force participation. In line with the gender-specific transmission of spillovers, we hypothesize that parental exposure to female co–workers differentially increases the likelihood of daughters to work later in life, i.e. $\beta_3 > 0$. x_i is a vector of additional controls that will be discussed when relevant, and ε_i is the error term. We cluster the standard errors at the office-level $J(i)K(i)$, coinciding with the level at which the parental exposure to female co–workers varies.[20]

Identification. Our setting offers a few important advantages to allay concerns over identification. To begin with, the focus on the gender gap β_3 helps difference out common shocks to male and females correlated with exposure. This, for example, helps allay concerns that observed increases in female employment capture general improvements in the employment prospects of *both* men and women that are correlated with the greater entry of women into the federal government. In addition, the distinct empirical advantage of our setting is that parental

[19] Figure A.IV shows the spatial distribution of cities in our sample. Cities with federal departments are spread throughout the country.
[20] In robustness checks, we also show that our results hold up using alternative clustering strategies (Table A.III).

exposure to female co–workers varies at the granular office-level, providing variation in exposure even within the same city (across departments) and the same department (across cities). We can thus introduce both city fixed effects θ_j and department fixed effects τ_k to allay concerns over time-invariant confounders that vary at the city- or department-level. This addresses empirical concerns that differences in exposure to female co–workers might be correlated with other unobserved factors that influence whether certain locations or departments are more likely to hire women. For example, urban areas may have a greater supply of female workers and thus see greater entry of female federal workers. These locations may have more progressive gender attitudes to begin with, making it hard to causally attribute differences in outcomes to differences in parental exposure.

Identification of the effect of parental exposure to female co-workers relies on the standard conditional independence assumption – namely, that the residual variation in female exposure is exogenous after partialing out city and department fixed effects. While we cannot directly test this assumption, we provide evidence in support of it by examining how our treatment variable relates to pre-treatment observable characteristics. Table A.IV and Figure A.V show that exposure to female co–workers is, conditional on the city and department fixed effects, uncorrelated with an office's pre-war share of female co–workers. In Table A.V, columns 1–2, we also show that – conditional on city and department fixed effects – exposure to female co–workers is not strongly correlated with any baseline characteristics, consistent with the quasi-random nature of this pre-determined variation.[21]

4.3 Linking parental exposure to children

The focus on intergenerational gender gaps in labor force participation requires (i) identifying the children of civil servants and (ii) tracking them into adulthood. The focus on both daughters and sons further complicates tracking children into adulthood as women traditionally changed their names after marriage. To implement our linkages, we rely on a range of "best practices" from the census linking literature to maximize the match rates of children over time.

Identifying the children of civil servants. To identify the children of civil servants, we restrict the sample of potential parents to civil servants serving in 1915 who can be linked to either the 1900, 1910, or 1920 Censuses. In each of the three census rounds, we define a "potential child" as a household member who (i) shares the same last name of the civil servant, (ii) is at least 18 years younger than the potential parent, and (iii) is aged 18 or younger in that census year. We restrict the sample to children born in or before 1917, the year the US entered WWI. These restrictions yield a sample of 29,290 civil service parents and 75,719 unique children.

Tracking the children of civil servants into adulthood. We match the children identified in the previous step to the 1940 Census to observe labor force participation decisions in adulthood.[22] We combine several linking

[21] When restricting to the variation used in our baseline specification - which relies on both city and department fixed effects interacted by gender, we fail to reject the null hypothesis in our F-test that all coefficients of the excluded variables are equal to zero as indicated by the p-value reported in Table A.V, column 2.

[22] The children in earlier census years are too young for labor market differences to emerge. In 1930, the mean age of children is only 22.

approaches to maximize the linkage rate. First, we leverage genealogical data from Census Tree (Buckles et al., 2023). This crosswalk relies on crowd-sourced genealogies and marriage certificates, providing information on last name changes of women. This approach allows us to match 42,889 children to the 1940 Census. Second, we utilize the IPUMS Multigenerational Longitudinal Panel (Ruggles et al., 2024), which employs a two-step linking process, where first-step high-quality matches for men serve as anchors to link other household members, including women, over time. Additionally, we incorporate a crosswalk developed by Ruggles et al. (2024), which links Social Security application files (SS-5) to full-count censuses. Specifically, we first match children from the 1900, 1910, and 1920 Censuses to their corresponding Social Security identifiers and subsequently link these identifiers to the 1940 Census. Combining the two crosswalks provided by Ruggles et al. (2024) yields a sample of 28,312 children matched to the 1940 Census.

These linkage procedures offer the crucial advantage of tracking women women over time, by employing diverse methodologies to address the challenge of surname changes upon marriage. We complement them with the crosswalk created by the Census Linking Project (Abramitzky et al., 2020), which uses automated name linkage for men, whose last names remain constant over time. This well-tested approach allows us to match 13,346 sons over time. Finally, we rely on our direct matching approach (Aneja and Xu, 2022), linking 6,187 children based on their exact full name and birth state.

Pooling the matches across all linkage methods, we track a total of 45,104 children to the 1940 Census. To reduce the likelihood of false matches, we conservatively drop from our sample children whose civil servant parents are aged 65 or above in 1917 and parents linked to more than 10 children across census rounds. This leaves us with a sample of 42,325 children (56%) of 21,947 civil service parents (75%).[23] As before, we restrict the analysis sample to larger cities (i.e., locations with at least two departments and offices of at least 20 employees in 1915 and 1919, see Section 4.1). To focus on the younger children who are more likely to be affected by exposure to female co–workers, we further limit the sample to those who are younger than 20 in 1917. This results in a final analysis sample of 13,487 children. Appendix Table B1 summarizes the construction of the sample.

5 Intergenerational effects on labor force participation

Table II reports our main results based on Equation 3. Each column presents a separate regression, with increasingly stringent fixed effects and controls. The dependent variable throughout is labor force participation (or "LFP"), a dummy equals to 1 if the individual is in the labor force in 1940 and 0 otherwise.[24]

In short, we consistently observe that parental exposure to female co–workers increases the LFP of daughters

[23] These match rates appear high as they are conditional on parents being matched to the census in the first place.

[24] We measure labor force participation using the IPUMS variable `LABFORCE`. The point estimates remain almost identical when removing farming and (`IND1950=105`) and home production (`IND` $\geq$ `950`).

relative to sons. Relative to sons, daughters whose parents were working in offices that experienced greater increases in the share of female co–workers are more likely to be working in adulthood. On average, daughters of civil servants are 48 percentage points (p.p.) less likely than sons to be in the labor force in 1940 (column 1, row 1).[25] The gender gap in LFP, however, shrinks with greater parental exposure to female co–workers. For the baseline specification (column 1), a 1 SD increase in parental exposure to women increases the relative likelihood that a daughter works by 2 p.p. As indicated by the level coefficient of $\Delta\text{Exposure}_{jk}$, exposure to female co–workers has no statistically discernible impact on the LFP of sons. Thus, the decline in the gender LFP gap is driven by a greater propensity for daughters of exposed civil service parents to work, consistent with a gender-specific transmission of views toward women in the workplace.

The remaining columns gradually make this baseline specification more demanding to demonstrate the robustness of the results. In column 2, we report results where we include office (i.e., city–by–department) fixed effects to account for unobserved differences in labor force participation within federal departments of a given city. These fixed effects fully absorb the parental exposure variation in levels, but still allow us to identify the differential effect of parental exposure to female coworkers by the gender of the child. Similarly, we add both city–by–female fixed effects and department–by–female fixed effects. This set of fully gender-interacted fixed effects restricts the identifying variation in exposure tightly to comparisons within cities and departments. The upshot is that our core finding is strengthened: daughters are 4.5 p.p. more likely to work as adults given a 1 SD increase in the share of female civil servants during WWI.

Column 3 makes this specification yet more demanding by allowing the work-age profile to flexibly differ by the child's gender. In column 4, we introduce the full set of female-interacted individual-level controls for a child's race, number of siblings, the civil servant parent's gender, age, (log) salary in 1915, and whether the parent held a clerical position. Since the exposure variation is largely uncorrelated with individual characteristics (Table A.V), the inclusion of these stringent controls leaves the estimates nearly unchanged.[26] Overall, our preferred estimates relying solely on within-city and within-department variation in parental exposure suggest that a 1 SD increase in parental exposure to female co–workers reduces the gender gap in LFP by 4.1–4.6 p.p. This corresponds to a decline of the mean labor force participation gap by 8.6–9.6% in the sample.

5.1 Heterogeneity and placebo tests

Treatment intensity. Our main result shows that parents working in offices exposed to a larger influx of female workers during WWI saw greater increases in their daughters' propensity to work in adulthood. We now explore

[25] For the sake of clarity, we note that an increase in female LFP relative to men is equivalent to a reduction in the gender LFP *gap*. Unless otherwise noted, subsequent discussions of a change in female LFP (corresponding to β_3 in Equation 3) should be interpreted as indicating a change in female LFP relative to male LFP, or a change in the female-male LFP gap.

[26] In Table A.VI, we further demonstrate the robustness of our main results by including family fixed effects, which restrict the comparison to siblings of different gender originating from the same parental household. The results are – as expected given the smaller sample and restrictive identifying variation – slightly noisier but remain economically and statistically significant.

several sources of heterogeneity to complement our main result. Models of multiple equilibria emphasize the role of tipping points, predicting discontinuities in the relationship between parental exposure to female co–workers and the effects on children's LFP gap. To investigate this possibility, we flexibly estimate our baseline specification (Table II, column 4) by intensity bins. As Figure IV shows, we do not find any discernible effects of exposure for small changes in the share of co–workers (below 15 p.p.). From increases of 15 p.p. and upwards, however, we detect a gradual increase in the exposure effects. While this pattern is consistent with norm-based tipping points, we cannot statistically reject a linear treatment effect in line with our specification in Equation 3.

Placebo parental exposure for civil service leavers. Our preferred interpretation is that parental exposure to female co–workers led to an increased labor force participation of daughters. An alternative interpretation is that female wartime workers were more likely to select into more female-friendly offices in the first place. This endogeneity would make it difficult to distinguish the effect of parental exposure to female co–workers from the fact that employees in more female-friendly offices might generally transmit more egalitarian views about work to their children. While our reliance on within-city and within-department variation mitigates concerns over selection *across* cities and departments, there could still be selection at the *office*-level (i.e., city–by–department). Even if our exposure shock is uncorrelated with an office's pre-war female share and observable worker traits (Table A.IV and Table A.V), selection on unobservable characteristics may still exist.

The availability of high-frequency personnel records allows us to conduct a placebo test to rule out the role of time-invariant selection on the office. In Figure V, we report the estimated parental exposure effects ($\Delta\text{Exposure}_{jk} \times \text{Female}_i$) for different exit cohorts. The "main sample" covers the set of civil servants that were serving in 1915 and last observed in the same year or thereafter.[27] The "placebo sample" covers those civil servants who were serving in the eventually exposed offices but left before the WWI shock was realized. Consistent with actual parental exposure driving our observed results, we find the intergenerational effects only for those civil servants who left after being exposed to the shock. These intergenerational effects are entirely absent for civil servants who selected into the same (to-be exposed) offices but left before the actual exposure. The absence of "pre-trends" helps rule out that our results are driven by time-invariant selection on the office.

Parental exposure by children's age. Motivated by work showing that experiences during "formative years" – childhood and early adulthood – can shape later-life economic behavior (Malmendier and Nagel, 2011; Fuchs-Schündeln and Schündeln, 2015; Roth and Wohlfart, 2018), we also extend the main analysis by studying how the parental exposure effects vary as a function of the child's age. If the intergenerational effects of exposure to female co–workers reflect the transmission of attitudes or information via parents, we expect to observe a greater transmission of workplace norms and information concerning women's work to those children who are

[27] Since the personnel records are biennial, an individual last observed in 1915 could have exited between July 1st 1915 and July 1st 1917. Notice that the reference date (July 1) is after the US entry into WWI (April 4 1917).

still malleable and likely to be living in their parent's household at the time of exposure.

We explore heterogeneity by the child's age at the time of exposure by also including children that are 20 years and older in 1940, and then flexibly estimating our baseline specification (Table II, column 4) for children for five different age bins. Figure VI presents the results, plotting the $\Delta\text{Exposure}_{jk} \times \text{Female}_i$ coefficient for each age group. The exposure effects are concentrated among children who were teenagers at the time of exposure, coinciding with ages at which children are arguably more malleable in terms of their preferences and views. In contrast, we do not find clear exposure effects for older children of civil servants, though the point estimates are less precisely-estimated given the smaller sample sizes. These children are more likely to both have left the home, and to have more rigid preferences and beliefs. These suggestive findings are thus consistent with the literature on the lasting effects of events during "formative years."

Heterogeneity by parental gender. We also consider whether the exposure effect varies by parental gender and occupation. In Table A.VII, we examine the impact of exposure to the WWI-related influx of female co-workers for male and female incumbent civil servants separately (column 1 presents the main effect for reference). We find that the effect of exposure to female workers on daughters' LFP is precise and sizable for incumbent fathers (column 3). This effect is noteworthy given that much existing research highlights vertical transmission through mothers (Fernández et al., 2004; Fogli and Veldkamp, 2011; Morrill and Morrill, 2013; Fernández, 2013). While we find that the effect for mothers is indeed positive (column 2), the effect is imprecise, likely due in part to the small sample of incumbent women working in government. These results thus provide evidence that fathers can also be a conduit for closing intergenerational gender gaps.

5.2 Robustness checks

Selection bias through imperfect census-linking. Although we rely on "best practices" in census-linking, there remains a concern that the results may reflect sample selection bias. For example, if "positively" selected daughters are more likely to be matched, working, and exposed to female co–workers through their parents, the resulting matched sample may be biased towards smaller gender gaps. Although the match rates are comparable to the related literature, the scope for selection is sufficiently large to potentially affect the results.

In our context, sample selection will only affect our main results if the match rate is correlated with our office-level exposure. As Table A.VIII shows, assuringly, the propensity of a civil servant child to be linked to the 1940 Census is – conditional on the city and department fixed effects – uncorrelated with the parental exposure to female co–workers. Nonetheless, differences in the baseline traits of matched and unmatched children (Table A.IX, column 2) motivate us to probe further. Specifically, we use inverse probability weighting (IPW) to ensure that the matched and unmatched samples are statistically balanced in terms of parental city, parental department, the share of initial clerical workers, and whether the child is female or not (Table A.IX, column 3).

Table A.X shows the reweighted point estimates for the intergenerational exposure effects of Table II. As the table shows, the results remain very comparable. Finally, we attempt to also allay concerns over unobservables that IPW cannot balance on. Table A.XI presents a bounding exercise where we assume that all unmatched individuals are either not working (column 1–2) or working (column 3–4) – directly testing for the extent to which this type of selection bias can change our results. Again, the point estimates remain comparable.

Finally, our results are also not driven by any single linking strategy. Genealogy-based linked census data, for example, tend to result in higher match rates for unmarried women, women with children – likely contributors to online genealogies, and women of higher socioeconomic status. Given the difficulty of linking women from their childhood households to their married households, there may be concerns that each of the linking methods induce different types of selection. To that end, we conduct a robustness check assessing the extent to which our main estimates change when dropping one linking method at a time. As Figure A.VI shows, the point estimates remain strikingly stable, leading us to conclude that our results are unlikely to be driven by sample bias due to selective census-linking, or any single census linking approach.

Expansion in federal employment and clerical positions. A limitation of the observational setting is that exposure to a greater share of female co–workers is invariably correlated with an increase in an office's workforce. While making it difficult to disentangle both factors, this "joint treatment" is natural. In a public sector setting where exit is limited, shifting the composition of the workforce goes hand in glove with the expansion of employment. We provide a range of tests to ensure that our results are indeed driven by the increase in female co–workers and not an expansion of the overall federal employment or clerical workforce. Table A.XII, column 2 provides a mediation exercise by directly controlling for the increase in workforce. The exposure effect remains significant and of comparable magnitude.[28] Since the increase in female federal employment is largely driven by the increase in clerks, there is a related concern that our results reflect the exposure to clerks rather than female co–workers. To assess this possibility, Table A.XII, column 3 asks whether daughters of parents exposed to a larger increase in the share of clerks see comparable improvements in labor force participation. While exposure to female co–workers closes the children's gender gap in labor force participation, exposure to more clerks does not. Finally, since female war-time entrants are younger and more educated (Table I), we also assess the extent to which changes in the average office workforce age and education explains – at least in part – our results. As columns 4–5 show, however, the results remain comparable, suggesting that it is indeed the exposure to female civil servants that drives the intergenerational effects we uncover.

Outlier departments and cities. We also assessed the extent to which our results are driven by any particular departments or cities. In Figure A.VII, we report the main results dropping one department at a time. As expected, the exposure effects are somewhat smaller when dropping the department that saw the largest increase in female workforce – the Department of War – but still remain statistically significant at conventional levels. By

[28] The result also holds when controlling for the increase non-parametrically using size bins.

and large, the coefficients are stable, indicating that the exposure effects occur more broadly across all arms of the federal government. In Figure A.VIII, we repeat the exercise now demonstrating the robustness of the point estimates when dropping one city at a time. Once again, as expected, the effect size declines when dropping the city with the greatest female expansion – the District of Columbia. The exposure effects, however, are otherwise remarkably stable and statistically significant throughout. Outliers are thus unlikely to drive our results.

Instrumental variable strategy. An advantage in our empirical setting is that the personnel records allow us to measure exposure to female civil servants at the *office*-level, thus allowing for tighter comparisons within the same governmental department and city. Nonetheless, concerns may remain over whether the remaining within–city and within–department variation is truly exogenous. We thus complement our approach by leveraging an arguably pre-determined source of variation that is predictive of incumbents' exposure to female workers. We rely on the historical fact that the expansion in the female federal workforce is primarily driven by the increased demand for clerical workers in departments that were heavily affected by wartime demand (Figure II). Women were already entering the private sector workforce – and clerical jobs in particular – in the years before WWI (Goldin, 2006). As discussed in Section 2.1, though, the war opened up many new job opportunities for women. As war efforts increased, so did the demand for greater administrative capacity, especially in war-related departments, such as the War and Navy departments. Offices with a greater pre-existing share of clerks were much more likely to see an expansion in the clerical workforce, even more so when the department was war-related (Figure A.IX). This occupational variation allows us to use the interaction of the existing share of the clerical workforce in 1915 and the presence of a war-related department as an instrumental variable.[29]

Our IV specification relies on the interaction between the share of clerical workers in 1915 and a dummy for being a war-related department (i.e., the Departments of War and Navy) as an instrument. More formally, we instrument $\Delta\text{Exposure}_{jk} \times \text{Female}_i$ with the triple interaction $\text{Share of clerks 1915}_{jk} \times \text{War-related dept}_k \times \text{Female}_i$, controlling for $\text{Share of clerks 1915}_{jk} \times \text{Female}_i$ and including the female-interacted department and city fixed effects. As we show in Table A.V, columns 3–4, this instrument appears – conditional on the lower-order interaction and fixed effects – not strongly correlated with any baseline characteristics of interest, consistent with the quasi-random nature of this pre-determined institutional variation.[30]

Table A.XIV reports the reduced form and IV estimates (columns 2-3). Consistent with our main specification, we find that parental exposure to female co–workers closes the children's gender gap in labor force participation.

[29] We can strengthen the intuition of using clerical and war department-related exposure as an instrument in other ways. Figure A.IX breaks down the overall variation in exposure by offices that (i) have an above/below the median share of clerical workers in 1915, and (ii) whether these offices belong to the Departments of War or Navy. As the figure shows, offices with an above-median initial share of clerical workers are more likely to see an expansion in female workers. Similarly, war-related departments with an above-median share of clerical workers see the greatest increase in shares of female workers. Table A.XIII summarizes the visual evidence in regression form, relating the exposure variation (Equation 2) to initial clerk share and being a war-related department. While both factors predict the increase in female workers (column 1), the interaction of both features has the greatest predictive power (columns 2–4) – consistent with Figure A.IX.

[30] Like the OLS balance test using our preferred fixed effects specification (Table A.V, columns 2), we fail to reject the null hypothesis in our F-test that all coefficients of the excluded variables are equal to zero (Table A.V, column 4).

This complementary research design leveraging a more explicit pre-determined source of variation to predict the increase in female co–workers further bolsters our main result. For the remainder of the study, we will continue to rely on our main OLS specification, and report IV estimates of all major results in the Appendix.

5.3 Marriage, fertility, and schooling outcomes

We consider if parental exposure to female co–workers affects other determinants of labor force participation. We first explore the impact of parental exposure on children's future marriages and childbearing – outcomes that are strongly tied to LFP (Gray, 1998; Goldin, 2006; Blau and Winkler, 2018). Women's employment decisions were historically related to marriage decisions. Female workers in our study period are mainly unmarried women, raising the question whether the higher LFP of exposed daughters goes along with a lower propensity to marry. The results are reported in Table III. As column 1 shows, parental exposure to female co–workers does reduce the propensity of daughters (relative to sons) to be married (in 1940). A 1 SD increase in parental exposure to female co–workers reduces the relative likelihood of daughters being married by 3.7 p.p. Similarly, daughters of exposed civil servants tend to also have relatively fewer children by 1940.[31]

We also examine whether parental exposure to female co–workers affects their children's educational attainment. Existing research documents a gender gap in parental human capital investments (Baker and Milligan, 2016; Dizon-Ross, 2019). Guided by this literature, we study whether parental exposure to female co–workers (who are also higher qualified) affects their children's gender gap in schooling (column 3). While parental exposure to female co–workers increases the relative years of education for daughters, the estimated coefficient is small and statistically insignificant.[32] Finally, we conduct a mediation exercise and ask whether the observable differences in marriage, fertility, and education we uncovered can fully account for the difference in labor force participation we observe. We do so by including a dummy for marital status and fixed effects for each number of children and education. As expected, the estimated parental exposure effects is somewhat smaller but still economically large and statistically significant (Table III, column 4). This suggests that parental exposure to female co–workers increases labor force participation above and beyond affecting the proximate determinants of labor force participation – consistent with changing norms to female work.

6 Mechanisms of transmission

Having documented the effects of parental exposure to female civil servants on daughters, we now use a simple framework to guide the discussion of channels through which the intergenerational effects may operate.

[31] The IV results are reported in Table A.XV.

[32] We do, however, find modest evidence of increased educational attainment at higher levels of schooling. To explore this, we use the same specification to regress different levels of education on $\text{Exposure}_{jk} \times \text{Female}_i$. As Figure A.X shows, we see that parental exposure to female co–workers disproportionately increases the schooling levels of daughters (vs sons) at the college level. This increase is driven by the lower propensity to have only completed middle school or high school, reflecting a shift towards higher levels of completed schooling.

Consider a setting where an agent i with group trait $j = \{m, f\}$ – male or female – decides to work. There exists uncertainty over the wage w_i and the extent to which working is penalized (or rewarded) by society, as captured by ν_j, which is common across individuals of the same group. The agent will thus decide to work if the expected utility from working exceeds the certain utility derived from exercising the outside option $\underline{w_i}$:

$$\mathbf{E}[w_i|x_i] - \mathbf{E}[\nu_j|x_i] > \underline{w_i} \tag{4}$$

We follow Fernández (2013) and conceptualize ν_j as capturing the gender-specific societal norm about work. For example, $\nu_f > 0$ can be interpreted as a social sanction that is imposed on working women, reducing their utility from work. Since there is uncertainty over the size of the social sanction, the agent relies on a private, individual-specific signal captured by x_i to form an expectation. Theories of intergenerational transmission highlight mechanisms such as the transmission of norms from parents to daughters (Fernández et al., 2004; Fernández, 2011) or through social interactions (Bisin and Verdier, 2001; Fogli and Veldkamp, 2011). In our setting, this corresponds to the parental exposure to female co–workers, capturing the vertical transmission of norms that was present at the turn of the 19th century.[33]

As Equation 4 implies, the increase in female LFP may reflect improved perceptions about the acceptability of female work, captured by a reduction in the expected sanction from working $\mathbf{E}[\nu_j|x_i]$ – a change in gender norms. Alternatively, the greater propensity for daughters of exposed parents to work could reflect better information about female employment prospects, captured by an increase in $\mathbf{E}[w_i|x_i]$. While it is difficult to fully disentangle both channels, we provide a series of complementary tests to show that the spillover effects are likely to operate beyond only changing the awareness of women's employment prospects.

6.1 Sectoral choice, occupation and income

We first examine if parental exposure effects manifest specifically as work within the federal government, or whether exposure leads to general labor market participation. If parental exposure to female co–workers predominantly captures information about female employment in government, one may expect the LFP effects to be limited to federal employment. As Table IV shows, however, the parental exposure effects extend to the private sector (columns 1–2).[34] In terms of point magnitude, 1 SD increase in parental exposure to female co–workers has the same effect on the intergenerational LFP gap within the public and private sectors, though the relative magnitude is – given the smaller baseline gap – larger for federal employment.

Similarly, if information were the only channel at play, we may expect increases in women's LFP to be limited

[33] Goldin (1980), for instance, concludes that "[p]arents were a large factor in the intergenerational transmission of cultural norms" regarding the decision of single women to work. This channel also directly relates to the literature on cultural transmission, where beliefs and values are transmitted "from generation to generation" (Guiso et al., 2006).

[34] The IV results are presented in Table A.XVI.

to clerical occupations, regardless of whether in the public or private sector.[35] We thus restrict the sample to children who are in the labor force in adulthood and test this hypothesis in Table IV, column 3. While women are generally more well-represented in clerical occupations, we find no effect of our measure of parental exposure on such occupational choice. Both the private sector and clerical effects are thus consistent with parental exposure to female co–workers increasing the overall relative propensity of daughters to work.

To further shed light on how parental exposure to female co–workers affects children's gender gaps in occupational choice, column 4 focuses on gender representation within the chosen occupations. The dependent variable % Male job is constructed as the share of men in the 1910 Census who work in a given occupation-industry. This allows us to examine whether daughters of civil servants are more likely than their counterparts to work in previously male-dominated jobs. While daughters of civil servants are less likely than their male counterparts to select into occupation-industries with pre-existing high male shares (based on the mean of the dependent variable), increased parental exposure to female civil servants helps close that employment gap. In the last column of Table IV, we ask if parental exposure to female co–workers affects the gender pay gap, measured as the total pre-tax wage and salary income in 1940. A 1 SD increase in parental exposure to female co–workers closes the gender pay gap by 12%. The combined evidence suggests that daughters of exposed parents are not only more likely to work, but also to enter higher-paying and previously male-dominated positions.

6.2 Movers and parental exposure at origin city

Despite our reliance on within-city variation in exposure to female co–workers to identify the effects on daughters of incumbent civil servants, there may remain concerns that the greater LFP reflects sustained demand for female workers by the government. While the private sector spillovers help further allay such concerns and suggest broader changes in gender norms, we can rely on "movers" to explicitly disentangle the role of norms from local labor market conditions correlated with the office-level exposure shock.

The early half of the 20th century was a period of substantial internal migration in the United States (Hall and Ruggles, 2004; Rosenbloom and Sundstrom, 2004). By 1940, 67% of the civil servant children in our sample lived in a city other than their parent's city of residence at the onset of WWI. Leveraging this widespread mobility, we can adopt an "epidemiological approach" (Fernández, 2011) and ask whether the parental exposure effects still hold for the subset of children who no longer reside in the original city of parental exposure to the increased representation of women in federal employment. We report the results in Table V. In columns 1–2, we first test whether parental exposure to female co–workers predicts the likelihood of daughters (relative to sons) moving to another city or remaining in the parental city of exposure. While parental exposure decreases the likelihood of moving (column 1), there is no statistically significant relationship once conditioning on the individual's current city – the destination city for movers – as measured in 1940 using fixed effects (column 2).

[35] This is true to the extent that the expansion of the female government workforce was mainly driven by clerks, as we show in Section 3.1.

In columns 3–4, we return to our main outcome (labor force participation) but restrict the sample to movers to test whether the effect of the exposure shock is still present. Even when comparing children of civil servants who ended up moving to the same city by 1940 (and thus face the same labor market conditions and institutional environment), variation in their parental exposure to female co–workers at the *origin* city–department remains highly predictive of greater labor force participation (column 3).

An empirical shortcoming of the "epidemiological approach" is that migration decisions are endogenous. While the destination fixed effects allay concerns over sorting in *levels*, migration decisions may still be origin-destination specific. In Table V, column 4, we go one step further by including origin-destination fixed effects. Even when comparing children of civil servants who moved from the same origin to the same destination city, those with parents in departments with higher exposure to female co–workers during WWI still see higher labor force participation. Consistent with the literature on cultural transmission, the evidence is strongly suggestive of our parental exposure effects reflecting the transmission of norms.

6.3 Direct effects of workplace contact

As a final test in our examination of mechanisms underlying the reduced LFP gap, we provide further evidence that changing gender attitudes are at least one plausible channel through which female LFP is changing. To do so, we complement our results using a direct measure that is an outcome measured for the affected civil servants, rather than their children. We focus in particular on the marriage choices of civil servants differentially exposed to female co–workers, and we ask whether exposed male civil servants were more likely to marry working women. While the early 20th century witnessed an increase in the labor force participation of married women, norms around the social acceptability of married work, as well as the formal institutions enshrining these norms, persisted.[36] In one nationwide survey of professional women conducted by the Bureau of Vocational Information just after WWI, female respondents predominantly reported their primary responsibilities to be that of homemakers, and that marriage and career were contradictory roles. Moreover, the Bureau's survey provided substantial evidence that men's expectations made marriage and career incompatible for women (Drachman, 1989). Similarly, in a 1936 Gallup survey, 72% of respondents disapproved of "a married woman earning money in business or industry if she has a husband capable of supporting her." A large contemporary literature has likewise documented how female labor force decisions affect marriage formation and their durability (Bertrand et al., 2015; Folke and Rickne, 2020).

Motivated by the historical context and the related literature, we use a man's propensity to marry a working spouse as an implicit measure of favorable attitudes toward women's work, and thus another way of capturing gender norms.[37] Focusing on male civil servants, we test whether exposure to female co-workers increases the

[36] "Marriage bars," or formal rules limiting the hiring of married women, were adopted by firms and local school boards throughout the first half of the 20th century (Goldin, 1988).

[37] In addition to being a direct test of changing gender norms, this approach allows us to expand the analysis to the subset of civil servants

propensity for their spouses to work. We track the spouses of men who were working in the federal government in 1915 across the census rounds 1900-1940. We identify female spouses for 56,378 male civil servants corresponding to a pooled sample of 94,092 spouses across 1900-1940 census rounds.[38] We then ask in a difference-in-differences setting whether exposure to female co-workers increases the spouses' propensity to work post-WWI (1920 and 1940) relative to the pre-war period (1900 and 1910).

Table VI shows the results, providing evidence consistent with exposure to female co–workers shifting attitudes towards working women. On average, we find that a 1 SD increase in exposure to female co–workers increases the propensity of a male civil servant's wife to work by 4.3 p.p. after WWI (column 1). Reassuringly, this increase is driven by marriages formed *after* the actual exposure to female co–workers (column 3). In contrast, we do not find that the eventual exposure to female co–workers is predictive of having a working spouse for those who married *prior* to the realization of the exposure shock during WWI (column 2). Finally, we exclude from the sample marriages among civil servants to ensure that our results are not exclusively driven by the greater likelihood of exposed men marrying their female co–workers. As column 4 shows, the results are primarily driven by exposed men marrying spouses who end up working outside the public sector. The evidence based on revealed–preference marriage patterns is thus collectively consistent with an interpretation that exposure to female co–workers shifted the gender norms of incumbent civil servants.[39]

7 Horizontal transmission and aggregate implications

We conclude our analysis by exploring the aggregate effects of the WWI-induced shock to women's participation in the federal service. While the previous section provided evidence consistent with the intergenerational transmission of changed attitudes toward women's work, the role of vertical transmission in generating rapid aggregate changes to female LFP is likely limited. This raises the question of whether the large increase in the female federal workforce had additional, contemporaneous spillovers.

To study whether the increased entry of female civil servants had impacts beyond the vertical transmission documented in Section 5, we aggregate our measure of office-level exposure to the city-level:

$$\Delta\text{Exposure}_j = \left(N^f_{j1919}/(N^m_{j19}+N^f_{j19})\right) - \left(N^f_{j15}/(N^m_{j15}+N^f_{j15})\right) \tag{5}$$

where $\Delta\text{Exposure}_j$ measures the city-level change in the share of female federal workers and $N^f_{j1919} = \Sigma_k n^f_{jk1919}$ is the total city-level number of federal government workers in 1919, computed by summing across all depart-

without children. In Appendix C, we validate this alternative measure of gender norms relying on current data from the World Value Survey.

[38] Appendix Table B2 summarizes the construction of the sample.

[39] Relatedly, we also investigated whether parental exposure to female co–workers increases the likelihood of sons to marry women in the workforce. In contrast to the direct effects of workplace contact (i.e., effects on incumbent civil servants), we do not find that *indirect* exposure through parents has similar effects (Table A.XVII). This suggests that the intergenerational effects we uncover operate primarily through the daughters.

ments k (see Equation 2). Figure VII summarizes the cross-city variation in the exposure measure, documenting substantial variation in the change in female shares. We leverage this cross-city variation to assess whether the overall presence of women affected female labor force participation beyond affected households.

7.1 Cross-city level variation in exposure

Moving from the individual level to the city level, we ask whether cities experiencing an increase in female representation in the government also see overall increases in female labor force participation.[40] We thus use this aggregate measure of female exposure in a difference-in-difference (DID) framework to compare female LFP in more versus less-exposed cities, before and after WWI:

$$y_{jt} = \beta_1 \Delta\text{Exposure}_j \times \text{Post WWI}_t + \theta_j + \tau_t + \varepsilon_{jt} \tag{6}$$

Here y_{jt} is the outcome in city j in year $t = \{1900, ..., 1940\}$. $\Delta\text{Exposure}_j$ captures the city-level change in the share of female federal workers, as defined by Equation 5. Post WWI_t is 1 for census years after 1910 and 0 before. θ_j are city fixed effects and τ_t are census year fixed effects. ε_{jt} is the error term, with standard errors clustered at the city level. Note that given the construction of our treatment variable, $\Delta\text{Exposure}_j$, we estimate the continuous version of the standard DID estimator (Callaway et al., 2024).

Our primary estimate of interest is β_1, which captures the aggregate-level exposure effect. Along the conceptual framework introduced in Section 6, this estimate would capture both the effects of the vertical transmission as well as any additional horizontal transmission. While moving to the city level allows us to explore impacts beyond vertical transmission, exploiting cross-city variation in exposure raises different potential identification concerns. Unlike our previous results, which rely on within-city and within-department variation in exposure, using cross-city variation may raise concerns about local labor market shocks that are correlated with exposure. To ameliorate concerns, we conduct a few important supplementary analyses. First, we estimate more flexible versions of Equation 6 where we include time-interacted controls for (log) city size as well as federal employment levels. Second, we test for the presence of pre-trends using a standard event study design.

Table VII shows our results for aggregate city-level effects of the increased presence of women in federal work. On average, we find that a 1 SD increase in exposure at the city-level increases female labor force participation by 0.9–1.1 p.p. (columns 1–2). The point estimates remain comparable when introducing controls to hold constant changes in the federal workforce size and differences in city sizes (column 2). These robustness tests suggest that the effects are unlikely to reflect changes driven by the expansion of federal employment or the fact that larger cities see a general increase in labor force participation after WWI. Finally, in column 4, we do not observe similar effects on male labor force participation, suggesting that the results indeed reflect a gender-

[40] We focus on cities because they are arguably a relevant characterization of the local labor market.

specific shock as opposed to generic city-level employment shocks. Figure VIII, panel (a) presents the flexible difference-in-difference estimates. As the figure shows, the impact on female labor force participation persists until 1940. In contrast, city-level exposure to female federal workers is uncorrelated with female LFP prior to WWI. For comparison, panel (a) also shows the corresponding estimates for male LFP, which reveals a flat, statistically insignificant effects of female exposure over time.

The magnitude of the aggregate effects we document is notable. In practical terms, the 0.9 p.p. increase in city-level female labor force participation in response to a 1 SD increase in exposure – corresponding to an increase in the city-level share of female federal government workers of 7.6 p.p. – is economically large. A back-of-the-envelope calculation suggests that, on average, each additional female war-time civil service entrant leads to 2.4 additional women entering the workforce – a sizable multiplier effect.[41]

We also exploit heterogeneity in the public sector size across cities to better understand the mechanisms of transmission. If the aggregate effects we observe are indeed driven by the increase in the share of the female federal workforce, we expect the horizontal transmission to be stronger in cities where the federal government has a larger footprint. To operationalize this, we split the sample into cities with above and below-median shares of federal employment. As column 3 of Table VII shows, we indeed find that the city-level exposure effects are larger in cities with a high federal employment share. Once again, this pattern is only restricted to female labor force participation (columns 3 and 5).

Finally, we break down female labor force participation for the public and the private sector. As expected, the largest effects are concentrated in the public sector. While part of the short-run effect is mechanical, the persistence and increasing magnitudes well into 1940 are striking (Figure VIII, panel (b)).[42] We also find that exposure to female federal workers increases female labor force participation in the *private* sector. The magnitude of this spillover is smaller but remains persistent throughout 1940. Mirroring the individual-level spillover effects on private sector employment, these results provide suggestive evidence for a wider effect of female exposure at the local labor market level, consistent with a horizontal transmission channel.

7.2 Channels of horizontal transmission

The effects uncovered at the city-level suggest the presence of private sector spillovers, raising the question of underlying mechanisms through which an increase in the female workforce in the public sector can also increase female participation in the private sector labor force. While identifying the exact channel of spillover is beyond the scope of this paper, we provide a few pieces of evidence for one plausible channel – social spillovers. A

[41] On average, an increase in the city-level share of female government workers by 7.6 p.p. corresponds to 96 additional female civil servants (0.076×1266 – the mean number of civil servants in 1915). An increase in the city-level female labor force participation by 0.9 p.p. corresponds to 324 additional female workers ($0.009 \times 35{,}997$ – the mean number of women in a city). Taking the ratio, we obtain the multiplier of $324/96 - 1 = 2.4$.

[42] The corresponding regression table can be found in Table A.XVIII.

large literature has highlighted the role of social learning in disseminating information and changing norms (Fogli and Veldkamp, 2011; Bursztyn et al., 2020; Miho et al., 2023). Such social spillover effects are often spatially concentrated. Moreover, social spillovers may also reflect underlying local social networks such as church communities, schools, clubs, and local neighborhoods.

Heterogeneity by neighborhood. Research suggests that economically consequential social interactions often occur at the level of one's neighborhood (Bayer et al., 2008; McCartney and Shah, 2022). Guided by this literature on the role of proximity for social interactions, we first ask if the gains in female LFP within a given city are concentrated in neighborhoods where female wartime civil servants resided. To implement this test, we restrict the sample to those residing in the cities of interest in 1920, differentiating by whether female wartime civil servants resided in the same census enumeration district (a proxy for neighborhood) or not. We then rely on our census–linking approach (Section 4.3) to track the cohorts of interest through the census rounds 1900–1940.

Figure IX summarizes the results.[43] The figure reports the increase in female labor force participation in response to a 1 SD city-level increase in female exposure for each census round, broken down by areas with and without female wartime entrants. As the breakdown reveals, the gains in female labor force participation are almost entirely concentrated in census enumeration districts where female wartime civil servants were residing. In contrast, neighborhoods without female war–time entrants do not see a comparable increase in female labor force participation following a similar-sized exposure shock. The presence of this spatial spillover is consistent with social learning as a potential mechanism through which horizontal transmission might operate.

Growth in women's clubs. Another way we can further examine the role of social learning is by looking at whether the influx of female civil servants during WWI changed women's civic participation.[44] We do so by considering whether women's large-scale emergence within the civil service affected the network of voluntary associations, particularly for women. As far back as the early 19th century, the U.S. was noted for the proclivity of its citizens to join voluntary associations as a way to participate in the social and political affairs of a community (Skocpol et al., 2000). Moreover, the early 20th century witnessed nationwide growth in federated women's organizations, due in part to the fact that clubs were the primary form of social inclusion and political activity in the absence of the formal franchise (Scott, 1992). Given that women's clubs were perhaps the primary form of political activity for women during our period, we study whether they functioned as potential conduits through which social spillovers might operate.

To test for this possibility, we rely on the *Official Register and Directory of the Women's Clubs in America* (Winslow, 1922). These novel data provide the most comprehensive list of women's clubs in the U.S., providing

[43] Once again, the corresponding regressions as shown in Table A.XIX.

[44] Examining standard outcomes like voting is challenging in our setting because women were almost universally excluded from the franchise until 1920 (Arnsbarger, 2024).

information about clubs and their membership at the city-level. We digitize every second year of the directories between 1913–1921 and compute the aggregate membership numbers for cities in our main sample. Consistent with our hypothesis, we find that cities with a greater exposure to female civil servants during WWI experienced faster growth in women's club memberships (Table A.XX). This pattern, however, only appears after WWI, coinciding with the rise in female labor force participation (Figure A.XI). While this evidence is suggestive, it nevertheless is consistent with women that are exposed to the prospect of female civil service becoming more engaged in socioeconomic life.

8 Conclusion

In this paper, we demonstrate that WWI was a pivotal juncture in the history of women in the labor market. We document both the rise of women working in the federal government as well as the transformation of gender norms over generations. Newly-digitized personnel records allow for a close examination of how the war mobilization effort transformed American bureaucracy. We first show that after America entered the war, the share of female civil servants more than doubled within two years – driven by the increase in clerical jobs. These war-time entrants were more likely to be young, unmarried, and highly qualified (relative to their male counterparts). Importantly, they were also less likely to have been in the workforce, suggesting that WWI helped draw in a previously untapped source of labor. Moreover, the change in composition persisted: the government continued on a path towards greater gender diversity for the next half-century.

We then consider how the sudden increased presence of female public sector workers generated intergenerational effects on female labor market outcomes. Leveraging city-by-department variation in the exposure of incumbent civil servants to greater gender diversity, we find strong evidence for the intergenerational transmission of attitudes towards women's work, as captured in the reduced gender gap in LFP of incumbent workers' children. Additional outcomes bolster our interpretation of a change in gender norms within more exposed government offices. Within-city variation and a complementary "movers" design further allow us to disentangle the effects of geography from changes in norms brought about by the sweeping increase in female representation across different offices of the U.S. government. Finally, we complement our evidence on the vertical transmission channel with aggregate evidence consistent with a horizontal transmission channel that acts as a social multiplier to magnify the effects of increased gender representation on broader changes in the labor market. In addition to contributing to opening the black box of female employment and clerical work during WWI, our evidence on the significance of "critical junctures" has broader policy implications. Our study offers a novel contribution to the understanding of self-reinforcing mechanisms through which temporary shocks can give rise to persistent changes at both the organizational and local labor market level. The paper highlights the role of personnel policies – and in particular public employment – in reducing labor market inequalities.

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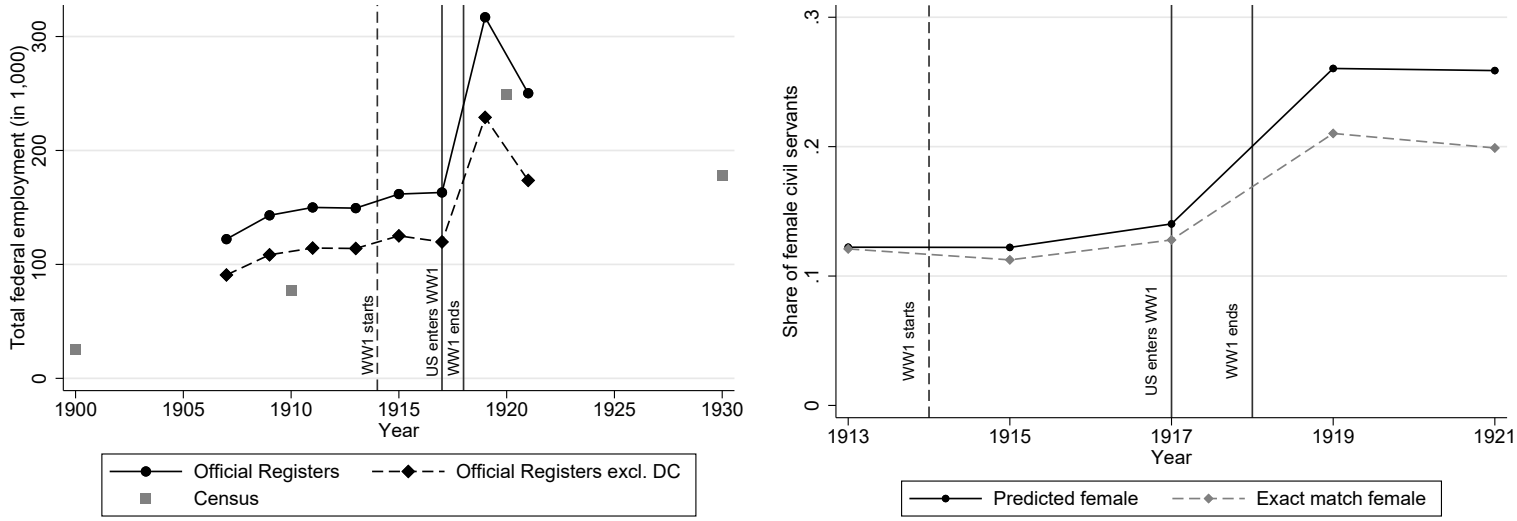
Figures and Tables

Figure I: Expansion of the federal workforce and increase in female share

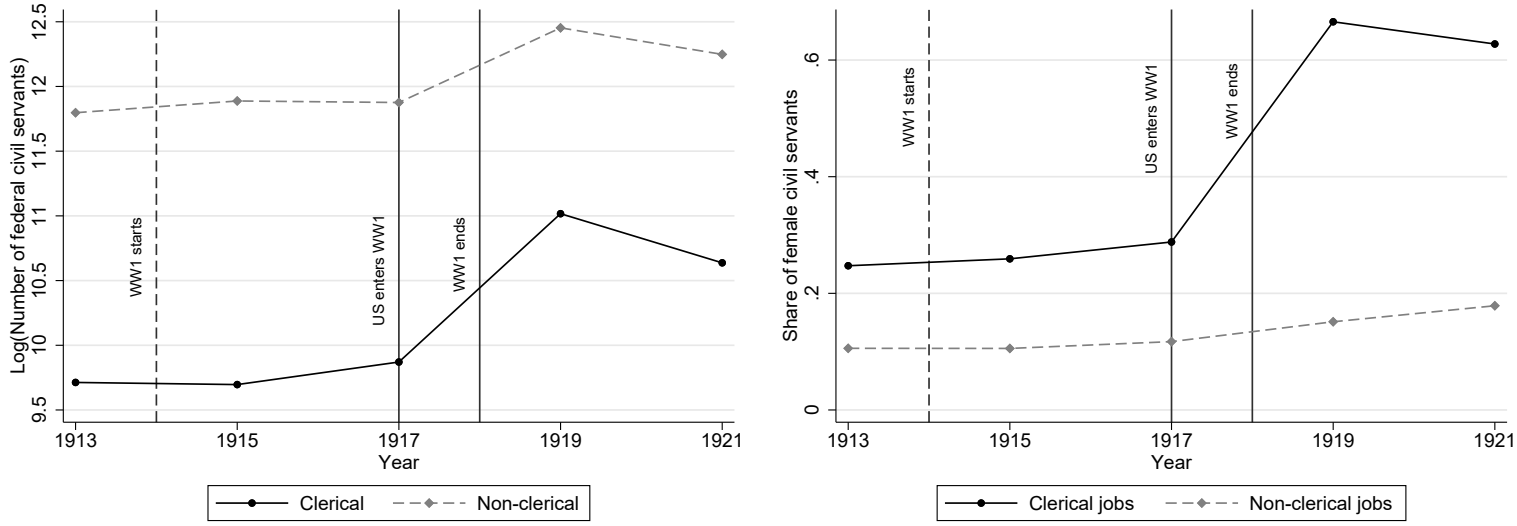


(a) Total federal employment

(b) Share of female civil servants

Notes: Panel (a): The figure shows the number (in thousands) of federal workers over time. The solid (resp., dashed) line reports the number computed using the Official Register data on the entire (resp., dropping District of Columbia) sample. The light-gray squares report the number derived from the decennial censuses. Panel (b): The figure shows the share of female civil servants over time. The measure is derived from the Official Register data, using our imputed gender measure (solid line) and census-linked gender measure (dashed line). See Section 2.3 for a detailed description of the procedure used to identify female civil servants. Note that US enters WWI in April 1917, while the Official Registers record personnel as of July 1st in each year.

Figure II: Size and female share of federal workforce, broken down by clerical vs. non-clerical work

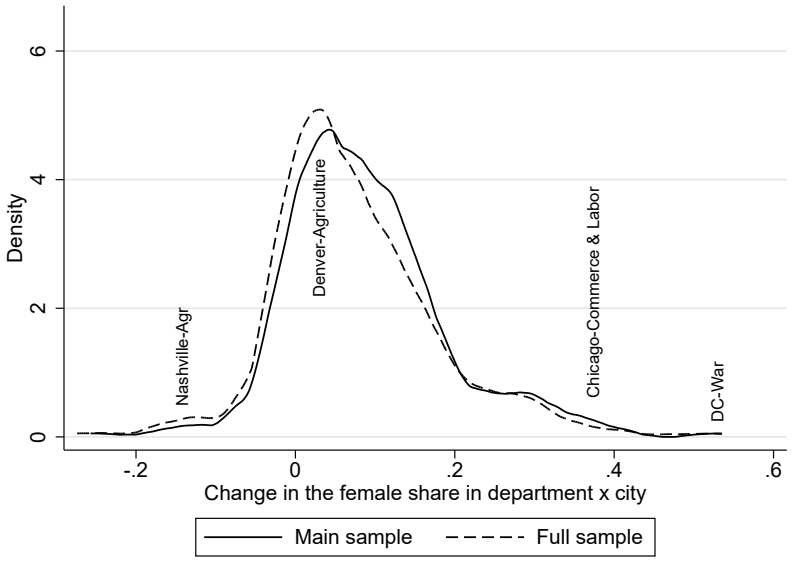


(a) Federal employment by clerk vs. non-clerk

(b) Share of female civil servants by clerk vs. non-clerk

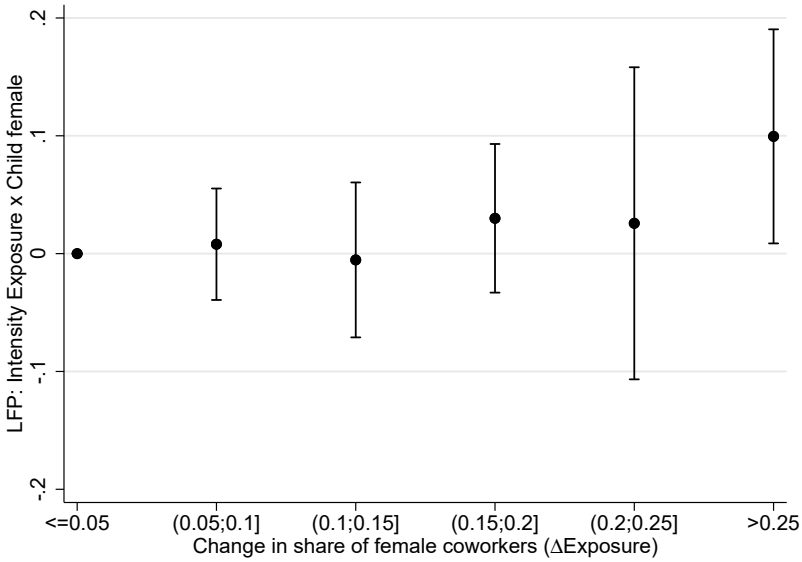
Notes: Panel (a): The figure shows the (log) number of federal civil servants by job type over time. The measure is derived from the personnel records and classifies job types as clerical (solid line) and non-clerical (dashed line). Panel (b): The figure shows the share of female civil servants, broken down by clerical (solid line) vs. non-clerical work (dashed line) over time.

Figure III: Distribution of change in female share of civil servants at the department-city level



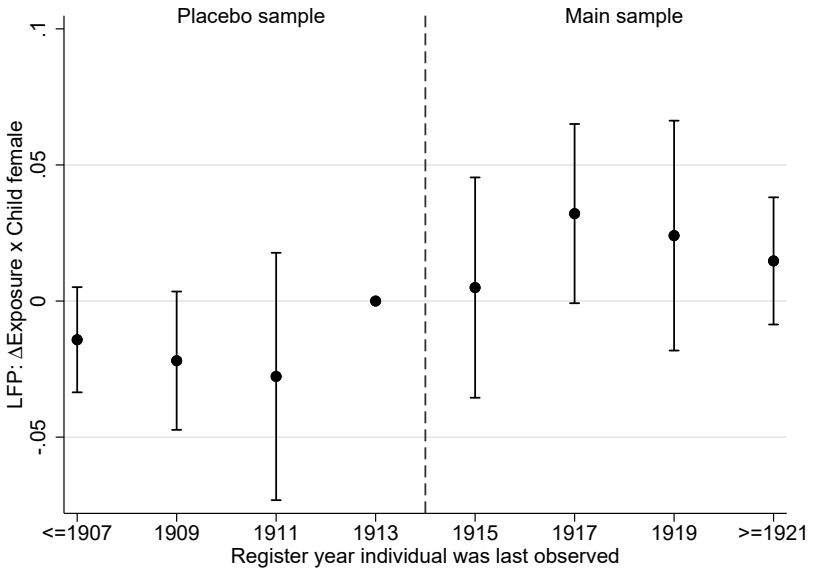
Notes: The figure shows the distribution of the change in the share of female civil servants across offices (i.e., city-department) between 1915 and 1919. The dashed line plots the distribution over the full sample of cities (307) with at least 20 civil servants in a given office in 1915. The solid line plots the distribution over our main sample of cities (70), which restricts to cities with at least two federal departments.

Figure IV: Exposure to female co–workers and children's LFP gender gap by intensity



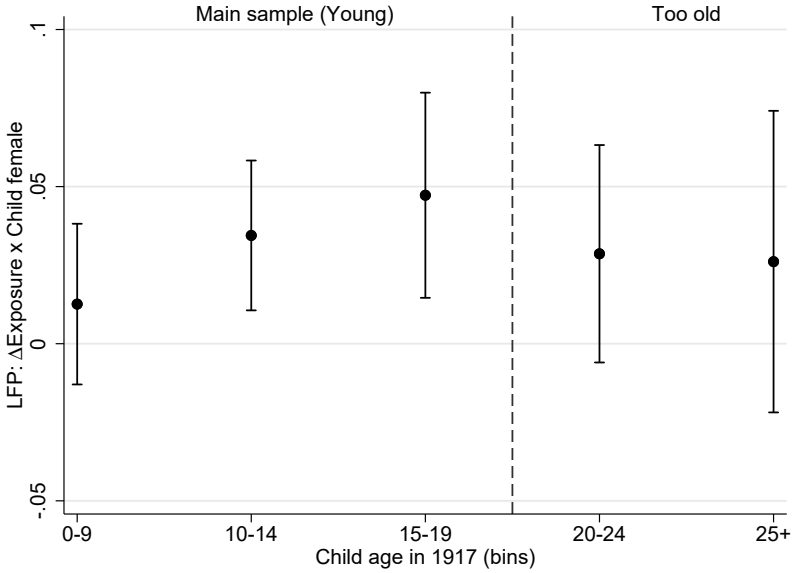
Notes: The figure shows the point estimates of our baseline specification (Table II, column 4), allowing the coefficient on the interaction term ΔExposure × Female to vary by intensity bins. The bins capture changes in the share of female co–workers across offices between 1915 and 1919 (e.g., the bin (0.05;0.1] captures changes in the share of female co–workers between 5 and 10 p.p.). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. The 95% confidence intervals are based on standard errors clustered at the office (i.e. city-department) level.

Figure V: Effect on intergenerational LFP gap by exposed vs. unexposed cohorts



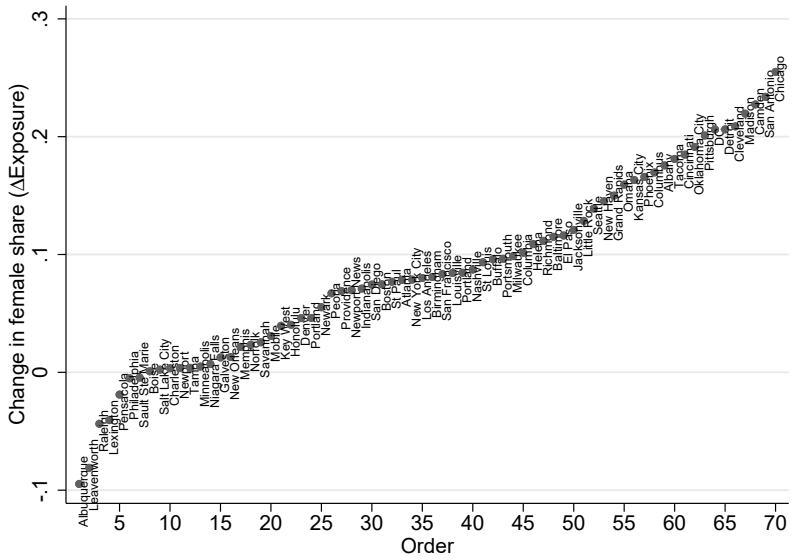
Notes: The figure shows the point estimates of our baseline specification (Table II, column 4), allowing the coefficient on the interaction term ΔExposure × Female to vary by the year the parent exited the civil service. The main sample refers to the sample of actually exposed civil servants used to estimate Table II. The placebo sample refers to those parents who were serving in the to-be exposed offices but left before the realization of the exposure shock during WWI. The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. The 95% confidence intervals are based on standard errors clustered at the office (i.e. city-department) level.

Figure VI: Exposure to female co–workers and children's LFP gender gap by age in 1917



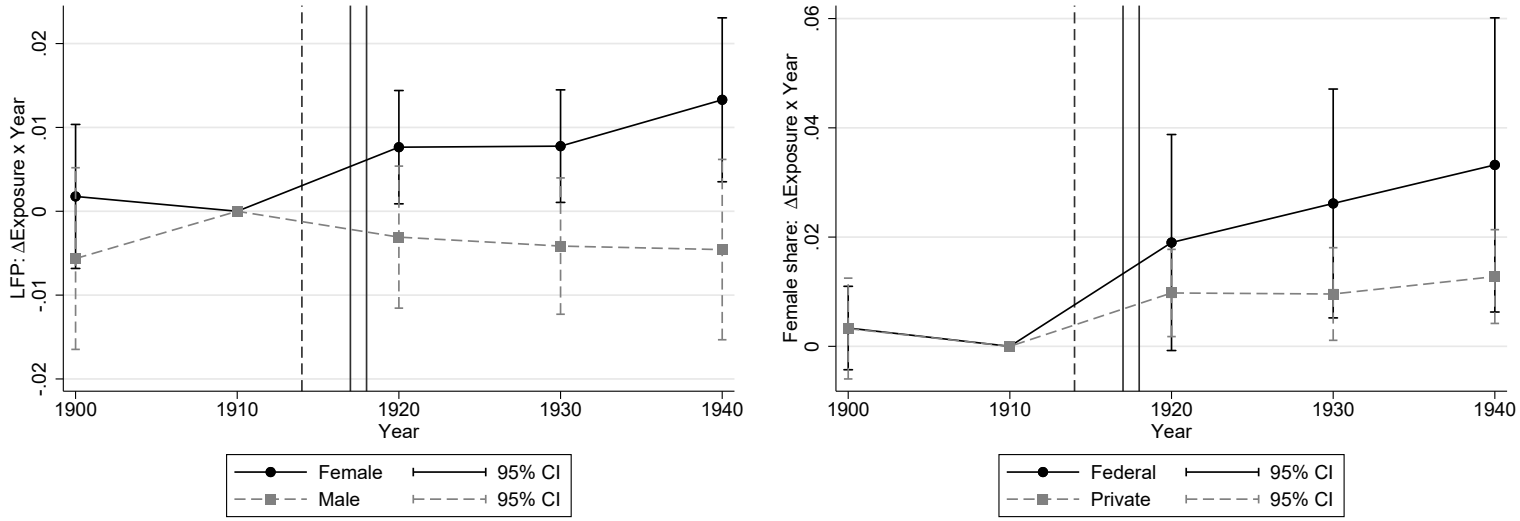
Notes: The figure shows the point estimates of our baseline specification (Table II, column 4), allowing the coefficient on the interaction term ΔExposure × Female to vary according to child's age in 1917. The main sample refers to the sample used to estimate Table II. The "too old" sample comprises those who were 20 years old or above in 1917. The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. The 95% confidence intervals are based on standard errors clustered at the office (i.e. city-department) level.

Figure VII: Cross-city variation in exposure to female wartime civil servants



Notes: The figure plots the ranked distribution of cities according to their change in the share of female civil servants between 1915 and 1919 (see Equation 2). The sample consists of the 70 cities in our main sample - that is, cities with at least 20 civil servants in a given office in 1915 and with at least two federal government departments.

Figure VIII: City-level exposure to female civil servants and labor market outcomes

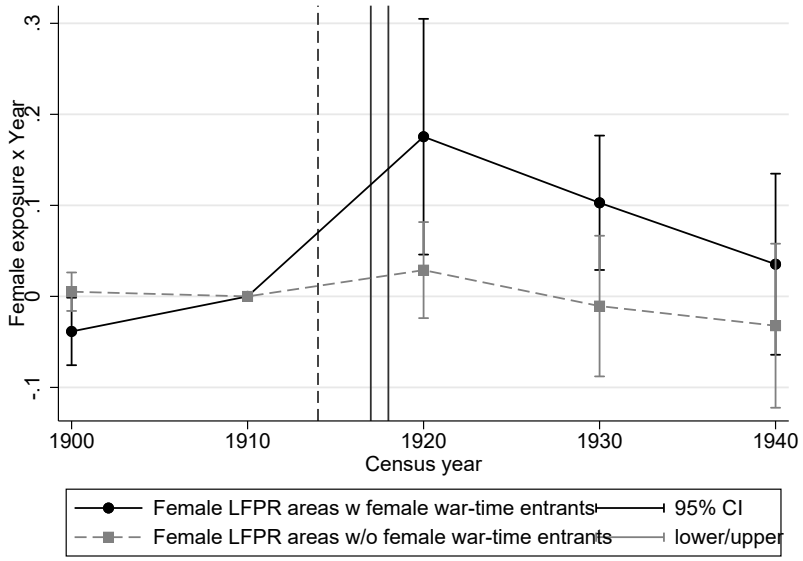


(a) Female – Male LFP

(b) Female share: federal – private

Notes: The figure shows a version of the full-control specification in Table VII, columns 2, allowing the coefficient on ΔExposure to vary by each decade. The unit of analysis is the city-year. The sample consists of a balanced panel of 66 cities in our main sample covering the 1900-1940 period. In panel (a), the outcome variable is the city-year-level share of women (resp. men) in the labor force for the solid (resp. dashed) line. In panel (b), the outcome variable is the city-year-level female share in the federal public (resp. private) sector for the solid (resp. dashed) line. All outcome variables are derived from the decennial censuses. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919.The 95% confidence intervals are based on standard errors clustered at the city-level. The dashed vertical line marks the start of WWI. The solid vertical lines mark the entry of the U.S. and the end of WWI.

Figure IX: City-level exposure to female civil servants and labor market outcomes by neighborhoods



Notes: The figure shows the flexible version of Table A.XIX, columns 3-4, allowing the coefficient on ΔExposure to vary by each decade. The solid line shows the estimates for the sample restricted to women who in 1920 lived in a census enumeration district where female wartime civil servants were living. The dashed line shows the estimates for the sample restricted to women who in 1920 lived in a census enumeration district with no female wartime civil servants. The 95% confidence intervals are based on standard errors clustered at the city-level. The dashed vertical line marks the start of WWI. The solid vertical lines mark the entry of the U.S. and the end of WWI.

Table I: Descriptive statistics of civil servant pre-entry characteristics

	(1) Mean	(2) Female	(3) WWI	(4) Female × WWI	(5) Obs.
Panel A: Individual traits					
Entry age	38.86	-2.448***	4.074***	-0.572**	70,853
		(0.197)	(0.094)	(0.224)	
White	0.915	-0.044***	-0.008***	0.066***	70,853
		(0.006)	(0.002)	(0.007)	
Never married in 1910	0.488	0.063***	-0.007	0.050***	70,853
		(0.009)	(0.004)	(0.011)	
In labor force in 1910	0.803	-0.438***	0.030***	-0.056***	70,853
		(0.009)	(0.003)	(0.010)	
Literate in 1910	0.978	0.002	-0.004***	0.010***	70,853
		(0.003)	(0.001)	(0.003)	
Years of education in 1940	11.72	0.557***	-0.575***	0.777***	42,100
		(0.108)	(0.047)	(0.121)	
Panel B: Geographic traits					
Ln(pop density 1910)	6.325	0.402***	0.455***	-0.359***	340,441
		(0.022)	(0.011)	(0.024)	
Same state	0.728	-0.010***	0.030***	-0.147***	340,441
		(0.004)	(0.002)	(0.004)	

Notes: The table compares the differential changes in selection into civil service between wartime and non-wartime periods and between male and female civil servants. The sample consists of entry cohorts of civil servants into the U.S. federal civil service, from 1913 to 1921. WWI is a dummy equal to 1 from year 1919 on. Each row reports the (difference in) means of a specific characteristic attached to each entrant civil servant. Column 1 reports the mean of each characteristic over the entire sample. Column 2 (resp. 3) reports the female-male (resp. post-pre WWI) difference in means. Column 4 reports post-pre WWI difference of the (female-male) gender gap. Panel A reports individual-level characteristics and is restricted to civil servants linked to decennial censuses. Entry age is derived from civil servants' age as reported in the 1910 Census, dummies for being white, being in the labor force, not being married, and being literate are derived from the 1910 Census. Years of education is derived from the highest year of completed school as reported in the 1940 Census. Panel B sample consists of all entrant civil servants, regardless of whether they are linked to decennial censuses. Population density in 1910 is at the county-of-appointment level. Same state is a dummy equal to 1 if the civil servant's first work location is in the same state as the state of appointment and 0 otherwise. Robust standard errors reported. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table II: Parental exposure to female civil servants and labor force participation in adulthood

	(1)	(2)	(3)	(4)
	Labor force participation			
Dep. var. mean female-male gap	-0.479	-0.479	-0.479	-0.479
Female	-0.477***			
	(0.011)			
Δ Exposure	-0.006			
	(0.005)			
Δ Exposure × Female	0.022***	0.045***	0.046***	0.041***
	(0.008)	(0.009)	(0.009)	(0.009)
City FEs	✓	✓		
Department FEs	✓	✓		
City × Department FEs		✓	✓	✓
City FEs × Female		✓	✓	✓
Department FEs × Female		✓	✓	✓
Age FEs × Female			✓	✓
Controls × Female				✓
Observations	13,487	13,487	13,487	13,487

Notes: The table shows OLS estimated coefficients from Equation 3. The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table III: Parental exposure to female co–workers and determinants of labor force participation

	(1)	(2)	(3)	(4)
	Married	Children	Education	Labor force
Dep. var mean female-male gap	-0.099	0.066	0.287	-0.480
Δ Exposure × Female	-0.037***	-0.058***	0.086	0.033***
	(0.010)	(0.021)	(0.068)	(0.010)
City × Department FEs	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓
Mediation controls (col. 1-3)				✓
Observations	13,407	13,407	13,151	13,072

Notes: The table shows OLS estimated coefficients from Equation 3. The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is married (col. 1), number of children (col. 2), and years of education (col. 3). All outcome variables are derived from the 1940 Census. In column 4, we augment our baseline specification (Table II, column 4) by including a dummy for being married, fixed effects for the number of children and years of education. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table IV: Parental exposure to female co–workers, sectoral choice and income

	(1)	(2)	(3)	(4)	(5)
	Employed in		Working in		
	Federal	Private	Clerical	% Male job	Log(Income)
Dep. var. mean female-male gap	-0.033	-0.446	0.327	-0.439	-0.366
Δ Exposure × Female	0.020***	0.022**	-0.010	0.026**	0.118***
	(0.007)	(0.009)	(0.018)	(0.011)	(0.043)
City × Department FEs	✓	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓	✓
Sample	Full sample		Labor force participation=1		
Observations	13,487	13,487	8,998	8,622	8,998

Notes: The table shows OLS estimated coefficients from Equation 3. The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). In columns 3–5, the sample is further restricted to individuals who are in the labor force in 1940. The outcome variable is a dummy equal to 1 if the child is employed in the federal government (col. 1), a dummy equal to 1 if the child is employed in the private sector (col. 2), a dummy equal to 1 if the child is employed in a clerical occupation (col. 3), the share of male workers in the same job (occupation and industry) as the unit of observation (col. 4), (log) working income (col. 5). Outcome variables are derived from the 1940 Census but for the one in column 4 which is derived from the 1910 Census. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table V: Labor force participation of movers and parental exposure at origin city

	(1)	(2)	(3)	(4)
	Moved city		Labor force participation	
Dep. var mean female-male gap	-0.031	-0.030	-0.524	-0.524
Δ Exposure × Female	-0.028***	-0.007	0.033**	0.032**
	(0.010)	(0.006)	(0.013)	(0.014)
City × Department FEs	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓
Current city FEs		✓	✓	
Origin-current City FEs				✓
Sample	Full		Moved county=1	
Observations	13,487	13,293	8,340	8,340

Notes: The table shows OLS estimated coefficients from Equation 3. The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). In columns 1–2, the dependent variable is a dummy that is 1 if the child's city of residence in 1940 is different from the parent's city in 1915 (the original city of exposure). In columns 3–4, the dependent variable is a dummy that is 1 if the child is in the labor force and the sample is restricted to those who have moved. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Current city FEs are fixed effects capturing the city the individual resides in 1940. Origin-current City FEs capture bilateral FEs for the parental city of exposure (the origin) and the individual's current city in 1940 (e.g., DC-Chicago). Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table VI: Exposure to female co-workers and being married to a working spouse

	(1)	(2)	(3)	(4)
	Wife is working			
Mean of dep. var	0.078	0.066	0.093	0.087
Δ Exposure	0.023	0.022		
	(0.021)	(0.021)		
Δ Exposure × Post WWI	0.043*	-0.034	0.106***	0.087***
	(0.022)	(0.040)	(0.028)	(0.027)
Age FEs	✓	✓	✓	✓
Department × Year FEs	✓	✓	✓	✓
City × Year FEs	✓	✓	✓	✓
Sample	Full	Married pre WWI	Married post WWI	
			All	Excl. govt' work
Observations	44,261	24,596	19,656	19,396

Notes: The unit of observation is individual-year. The sample includes spouses (i.e., wives) of men who were working in the federal government in 1915 and identified across the census rounds 1900-1940. The outcome variable is a dummy equal to 1 if the male civil servant's spouse is working. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. Year FEs correspond to census round fixed effects (1900-1940). In column 2, the sample is restricted to those who married prior to WWI (i.e., in census rounds 1900-1910). In columns 3–4, the sample is restricted to only include marriages that were formed after WWI (i.e. in census rounds 1920-1940). In column 4, the sample only includes male civil servants' wives that are not federal workers listed in the Official Registers. Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

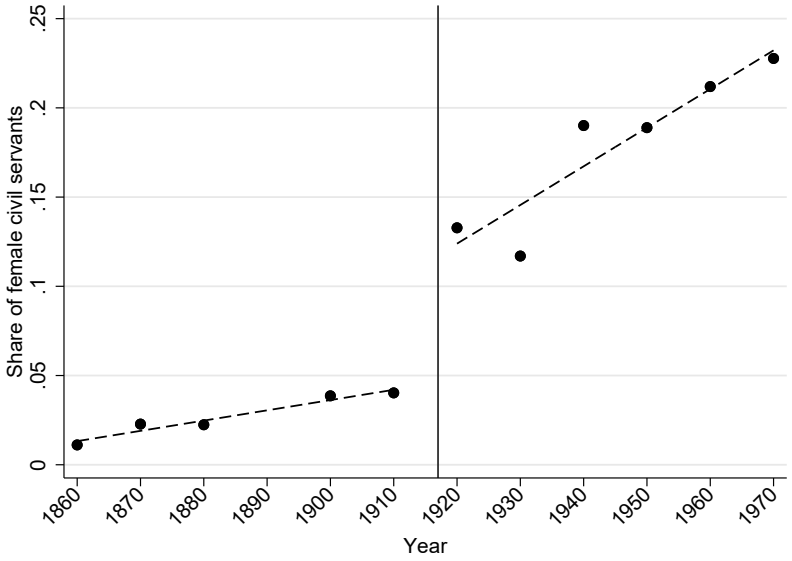
Table VII: Aggregate exposure to female civil servants and city-level labor market outcomes

	(1)	(2)	(3)	(4)	(5)
	Female LFP			Male LFP	
Mean of dep. var	0.239	0.239	0.239	0.646	0.646
Δ Exposure × Post	0.011***	0.009***	0.004	-0.001	-0.003
	(0.003)	(0.003)	(0.004)	(0.004)	(0.006)
Δ Exposure × Post × High federal employment city			0.010*		0.005
			(0.006)		(0.007)
Year FEs	✓	✓	✓	✓	✓
City FEs	✓	✓	✓	✓	✓
Year FEs × City-level controls		✓	✓	✓	✓
Lower oder interactions	-	-	✓	-	✓
Δ Exposure × Post +	-	-	0.014***	-	0.001
Δ Exposure × Post × High federal employment city			(0.004)		(0.004)
Observations	330	330	330	330	330

Notes: The table shows OLS regression estimates from Equation 6. The unit of analysis is the city-year. The sample consists of a balanced panel of (66) cities in our main sample covering the 1900-1940 period. The outcome variable is the city-year-level share of women in the labor force (columns 1-3) and the share of men in the labor force (columns 4–5) as derived from decennial censuses. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. High federal employment city is a dummy that is 1 if the city's share of federal employment is above median (1.3%) and 0 otherwise. City-level controls include: the 1915-1919 change in the (log) federal workforce size and (log) population in 1910. Standard errors are clustered at the city level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

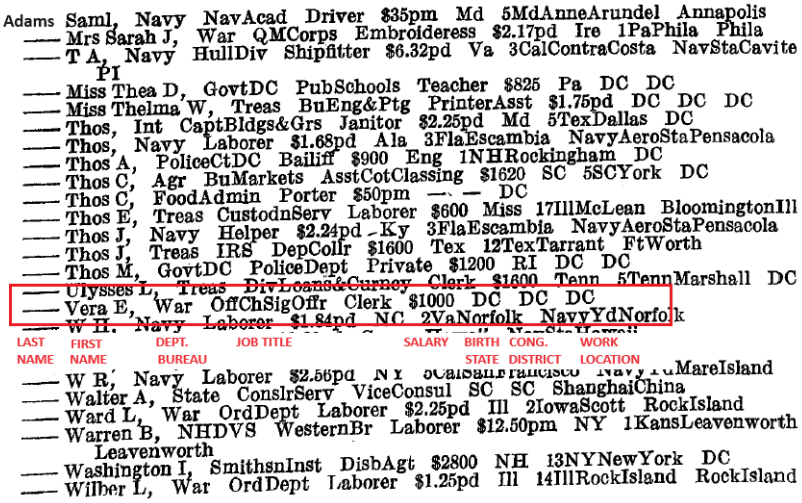
A Online Appendix

Figure A.I: Share of female civil servants over the long-run 1860–1970



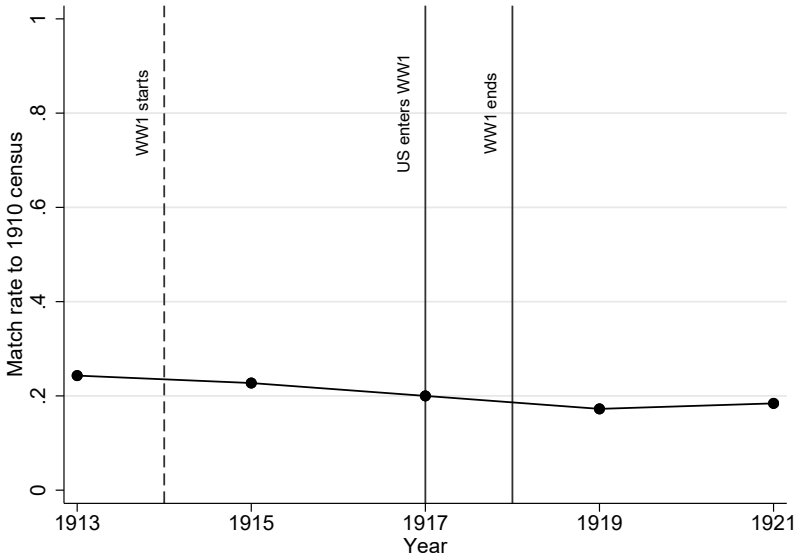
Notes: The figure shows the evolution of the share of female civil servants from 1860 to 1970 as it results from US decennial censuses. The black vertical line represents the US entry in World War I in 1917.

Figure A.II: Example of Official Register records



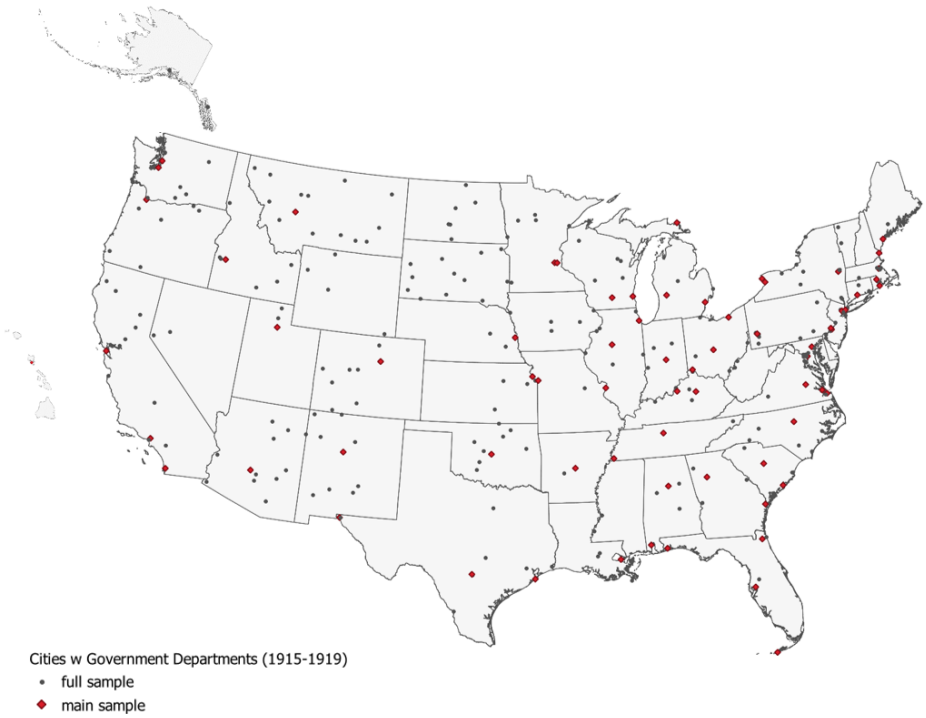
Notes: The figure shows a sample of the Official Registers for the year 1917 (p. 124). See Section 2.2 for a detailed description of the source.

Figure A.III: Census match rates over time



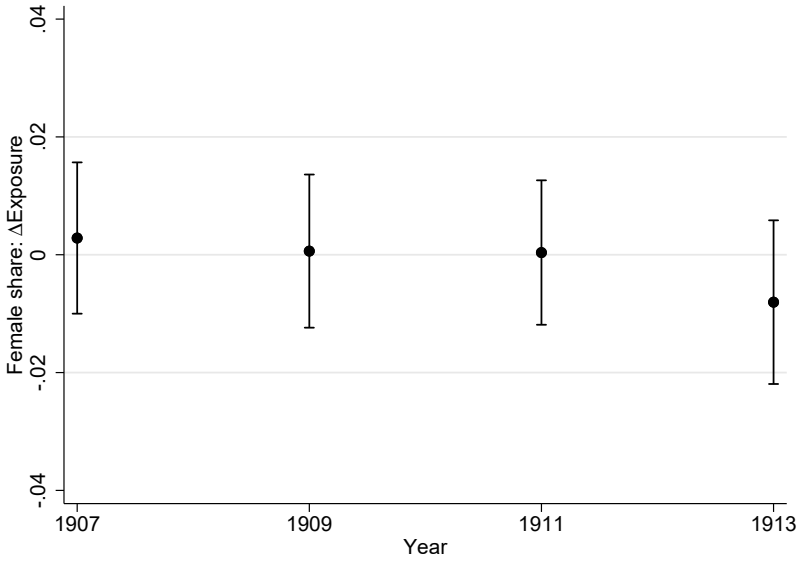
Notes: The figure shows the share of civil servants from the Official Register dataset (covering the 1913-1921 period) who could be linked to the 1910 Census.

Figure A.IV: Spatial distribution of cities with federal departments



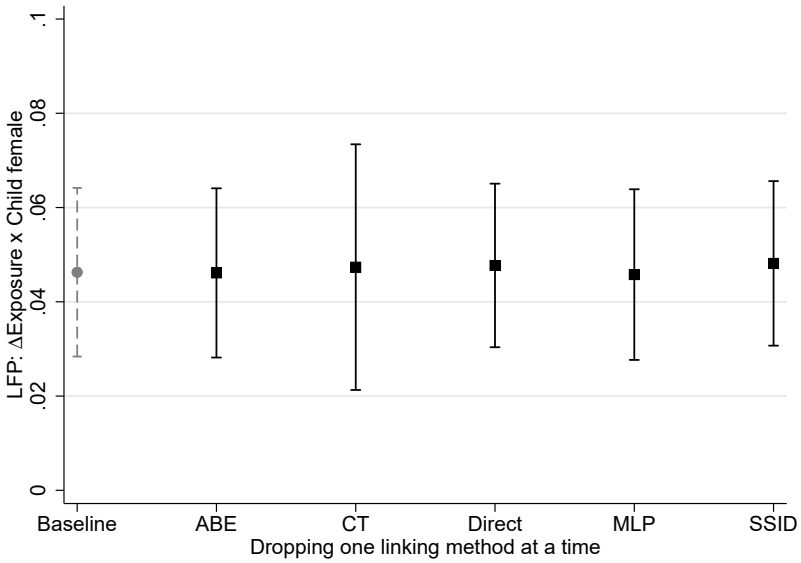
Notes: The figure shows the spatial distribution of cities with federal departments with at least 20 civil servants in a given office in 1915. Main sample (red squares) additionally restricts to cities with at least two departments.

Figure A.V: Pre-war levels of and wartime increase in female share across department-cities



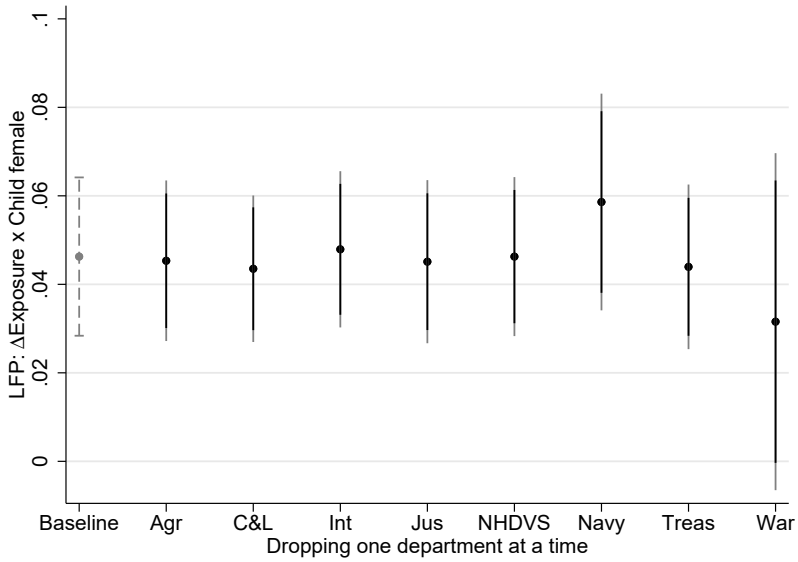
Notes: The figure shows the point estimates of Table A.IV (column 3), allowing the coefficient on ΔExposure to vary by year. The outcome variable is the female share of civil servants across offices in each year of Official Registers between 1907 and 1913. ΔExposure captures the change in the share of female civil servants across offices between 1915 and 1919 (see Equation 2). City and department fixed effects are included in the regression. The 95% confidence intervals are based on standard errors clustered at the office (i.e., city-department) level.

Figure A.VI: Exposure to female civil servants and LFP – dropping one linking method at a time



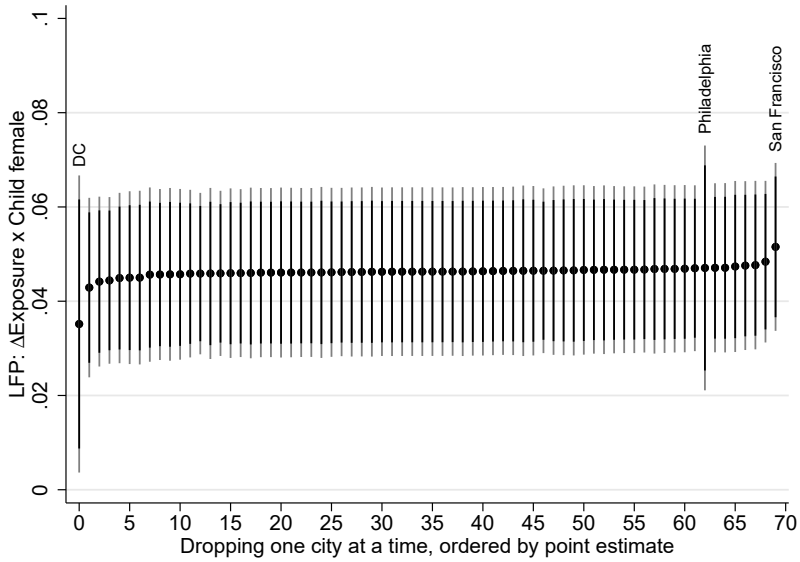
Notes: The figure shows the point estimates relating exposure to female civil servants to labor force participation dropping one linking method (ABE, Census Tree, direct linking, MLP, and SSID) at a time. All point estimates correspond to the specification of Table II, column 3, except that we rely on different matched samples. Baseline reflects the preferred matched sample relying on all matches. Standard errors are clustered at the office (i.e., city-department) level. Reporting 95% confidence intervals.

Figure A.VII: Exposure to female civil servants and LFP – dropping one department at a time



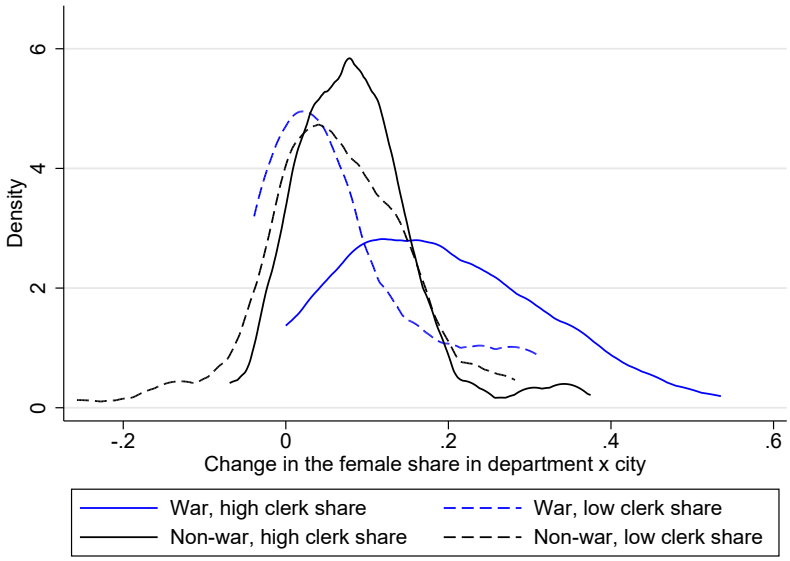
Notes: The figure shows the point estimates relating exposure to female civil servants to labor force participation dropping one department at a time. All point estimates correspond to the specification of Table II, column 3, except that we rely on different subsamples. Baseline includes all departments. Standard errors are clustered at the office (i.e., city-department) level. Reporting 95% (black) and 90% (gray) confidence intervals.

Figure A.VIII: Exposure to female civil servants and LFP – dropping one city at a time



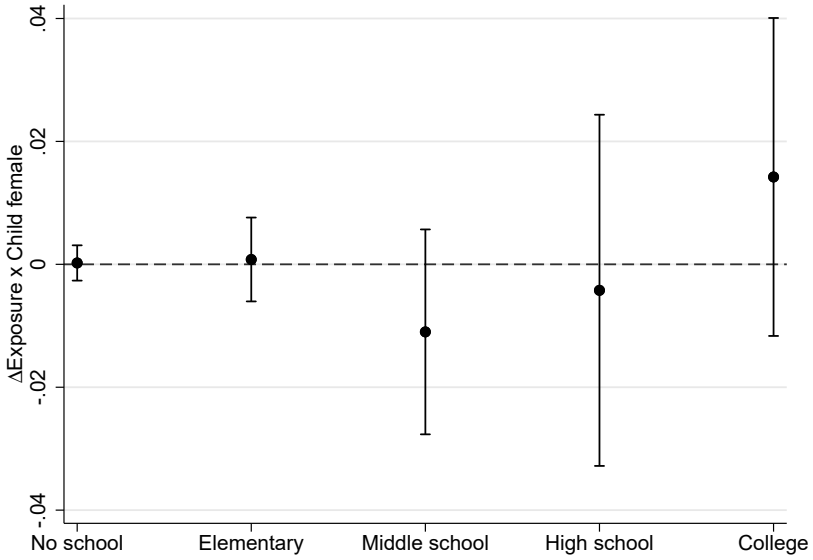
Notes: The figure shows the point estimates relating exposure to female civil servants to labor force participation dropping one city at a time. All point estimates correspond to the specification of Table II, column 3, except that we rely on different subsamples. Baseline includes all cities in sample. Standard errors are clustered at the office (i.e., city-department) level. Reporting 95% (black) and 90% (gray) confidence intervals.

Figure A.IX: Distribution of change in female share of civil servants – by clerical share and war department



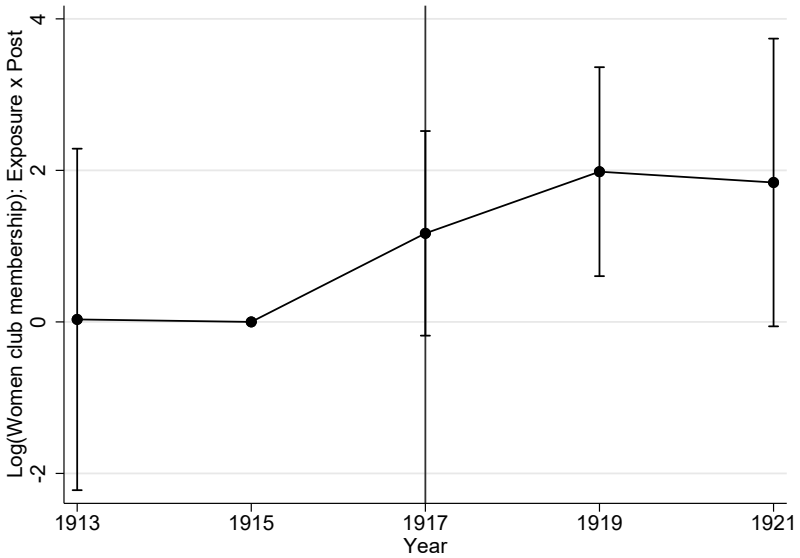
Notes: The figure shows the distribution of the change in the share of female civil servants across offices (i.e., city-department) between 1915 and 1919. Blue (resp., black) lines restrict to offices in war (resp., non-war) departments. Solid (resp., dashed) lines restrict to offices with the share of clerical workers in 1915 above (resp., below) the median.

Figure A.X: Exposure to female civil servants and children's schooling gap



Notes: The figure shows a version of the specification in Table III, column 3, where the the outcome variable is a dummy equal to 1 if the highest year of school as in the 1940 Census corresponds to the level of education indicated on the x-axis and 0 otherwise. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The 95% confidence intervals are based on standard errors clustered at the office (i.e. city-department) level.

Figure A.XI: City-level exposure to female civil servants and women's club membership



Notes: The figure shows a version of the full-control specification in Table VII, columns 2, allowing the coefficient on ΔExposure to vary by each year. The unit of analysis is the city-year. The sample consists of a panel of cities covering the 1913–1921 period, with observations available biennially. The outcome variable is the (log) total number of members in women clubs. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. City-level controls include: the 1915-1919 change in the (log) federal workforce size and (log) population in 1910. Standard errors are clustered at the city level. 95% confidence intervals reported.

Table A.I: Descriptive statistics of pre-entry characteristics - industries

	(1) Mean	(2) Female	(3) WWI	(4) Female × WWI	(5) Obs.
Manufacturing sector	0.184	-0.117***	-0.011**	0.024**	41,616
		(0.009)	(0.004)	(0.011)	
– Light manufacturing	0.037	0.011*	0.004**	0.000	41,616
		(0.007)	(0.002)	(0.008)	
Professional occupations	0.106	0.221***	-0.026***	0.088***	41,616
		(0.015)	(0.003)	(0.017)	
– Education	0.054	0.184***	-0.008***	0.112***	41,616
		(0.013)	(0.002)	(0.015)	

Notes: The table compares the differential changes in selection into civil service between wartime and non-wartime periods and between male and female civil servants. The sample consists of entry cohorts of civil servants into the U.S. federal civil service, from 1913 to 1921 and it is further restricted to only include those who were in the labor force, but outside the government, as reported in the 1910 Census. WWI is a dummy equal to 1 from year 1919 on, corresponding to the first year after the U.S. entry into WWI for which we have personnel records. Each row reports the (difference in) means of a specific characteristic attached to each entrant civil servant. Column 1 reports the mean of each characteristic over the entire sample. Column 2 (resp., 3) reports the female-male (resp., post-pre WWI) difference in means. Column 4 reports post-pre WWI difference of the (female-male) gender gap. Manufacturing sector and Professional occupations are dummy variables equal to 1 if the individual is reported to work in the corresponding industry sector in the 1910 Census, according to the `IND1950` classification provided by IPUMS, and 0 otherwise. Light manufacturing is a dummy equal to 1 if the individual is reported to work in textile, paper, and printing sectors in the 1910 Census, and 0 otherwise. Education is a dummy equal to 1 if the individual is reported to work in educational services in the 1910 Census, and 0 otherwise. Robust standard errors reported. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.II: Descriptive statistics of additional pre-entry characteristics - WWI and 1918 pandemic

	(1) Mean	(2) Female	(3) WWI	(4) Female × WWI	(5) Obs.
Panel A: Patriotism					
Liberty bonds - origin state	0.218	-0.007***	-0.003***	0.009***	290,727
		(0.001)	(0.000)	(0.001)	
Desertion rate - origin state	0.015	-0.001***	-0.001***	0.001***	290,727
		(0.000)	(0.000)	(0.000)	
Enlistment rate - origin county	0.120	-0.003***	0.005***	-0.004***	270,503
		(0.001)	(0.000)	(0.001)	
Panel B: 1918 pandemic					
Influenza excess mortality – origin state	2.647	0.021***	-0.043***	-0.068***	217,269
		(0.005)	(0.002)	(0.006)	
Influenza excess mortality – destination city	588.4	18.147***	-12.481***	-7.688***	228,346
		(0.980)	(0.652)	(1.112)	

Notes: The table compares the differential changes in selection into civil service between wartime and non-wartime periods and between male and female civil servants. The sample consists of entry cohorts of civil servants into the U.S. federal civil service, from 1913 to 1921. WWI is a dummy equal to 1 from year 1919 on, corresponding to the first year after the U.S. entry into WWI for which we have personnel records. Each row reports the (difference in) means of a specific characteristic attached to each entrant civil servant. Column 1 reports the mean of each characteristic over the entire sample. Column 2 (resp., 3) reports the female-male (resp., post-pre WWI) difference in means. Column 4 reports post-pre WWI difference of the (female-male) gender gap. Panel A reports war-related measures, attached to each civil servant according to the state (or county) of appointment. Liberty bonds is the state-level subscription rate of the Fourth Liberty Loan during WWI (Hilt et al., 2022). Desertion rate at the state level and enlistment rate at the county level are taken from Crowder (1920). Panel B reports measures of severity of the 1918 Influenza pandemic, attached to each civil servant according to the state of appointment or the city of first federal job. The state-level measure (available for 27 states) is computed dividing the total number of people died to influenza and pneumonia in the 1918-1919 period by the corresponding number in the 1915-1916 period (data digitized from Rogers, 1920). The city-level measure (available for 46 cities) is taken from Correia et al. (2022) and is defined as the difference between the deaths to influenza and pneumonia that took place in the 1918-1919 period and the number expected in the absence of the pandemic (i.e., median mortality in the 1910-1916 period), reported as a rate per 100,000 inhabitants. Robust standard errors reported. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.III: Exposure and children's LFP gap – alternative clustering

	(1)	(2)	(3)	(4)
	Labor force participation			
Dep. var. mean female-male gap	-0.479	-0.479	-0.479	-0.479
Δ Exposure × Female	0.041***	0.041***	0.041**	0.041**
	(0.0094)	(0.0087)	-	-
p-value	*0.000*	*0.000*	*0.020*	*0.018*
Cluster	City-Dept	City	Dept	Two-way
N. clusters	227	70	8	70 & 8
City × Department FEs	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓
Observations	13,487	13,487	13,487	13,487

Notes: The table shows a version of our OLS baseline specification (Table II, column 4, and reported in column 1) with alternative clustering strategies. Standard errors are clustered at the office (i.e., city-department) level in column 1, at the city level in column 2, at the department level in column 3, and two-way clustered at the city and department levels in column 4. In columns 3 and 4, we use wild cluster bootstrap standard errors because of the small number of department clusters (Cameron et al., 2008). The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.IV: Pre-war office-level female shares and exposure to female civil servants during WWI

	(1)	(2)	(3)	(4)
	Female share, 1907-1913			
Mean dep. var.	0.078	0.078	0.078	0.077
Δ Exposure	-0.002	-0.012*	-0.001	-0.001
	(0.006)	(0.006)	(0.006)	(0.006)
Observations	904	904	904	884
Year FEs	✓	✓	✓	
City FEs		✓	✓	
Department FEs			✓	
City × Year FEs				✓
Department × Year FEs				✓

Notes: The table shows the relationship between our measure of exposure to female co–workers and office's pre-war female share. The unit of observation is office-year. The sample consists of the 70 cities in our main sample - that is, cities with at least 20 civil servants in a given office in 1915 and with at least two federal government departments. The sample pools offices over the 1907-1913 period. The year-by-year estimates are summarized in Figure A.V. The outcome variable is the female share of civil servants across offices in each year of Official Registers between 1907 and 1913. ΔExposure captures the change in the share of female civil servants across offices between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1 (see Equation 2). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.V: Individual-level correlates of key regressor

	(1)	(2)	(3)	(4)
	OLS		IV	
	Δ Exposure × Female		Share clerks × War × Female	
Mean of dep. var	0.001	0.001	-0.079	-0.079
Panel A: Children's traits				
White	0.019	0.008	-0.065***	0.004
	(0.043)	(0.012)	(0.024)	(0.007)
Number of siblings (standardized)	-0.011***	-0.002	-0.006**	-0.001
	(0.004)	(0.002)	(0.003)	(0.001)
Age in 1917 (standardized)	0.002	-0.001	0.004	0.001
	(0.005)	(0.002)	(0.004)	(0.001)
Panel B: Civil servant parent's traits				
Female	-0.046	-0.009	-0.044	-0.002
	(0.045)	(0.006)	(0.034)	(0.004)
Literate	-0.090	-0.028	-0.034	-0.024
	(0.072)	(0.020)	(0.048)	(0.016)
White	-0.055	0.004	0.028	0.002
	(0.038)	(0.013)	(0.026)	(0.008)
Married in 1910	0.005	-0.002	0.014**	0.002
	(0.008)	(0.005)	(0.007)	(0.003)
Age in 1917	0.007*	0.004*	-0.001	-0.000
	(0.004)	(0.002)	(0.004)	(0.002)
Log(salary 1915) (standardized)	0.001	-0.002	-0.002	-0.001
	(0.006)	(0.002)	(0.004)	(0.001)
Tenure in 1915 (standardized)	-0.000	0.002	0.010*	0.001
	(0.007)	(0.001)	(0.005)	(0.001)
Clerk in 1915	0.056**	-0.002	0.039*	-0.004
	(0.027)	(0.004)	(0.022)	(0.002)
Work state in 1915 same as birth state	-0.006	0.002	-0.005	-0.001
	(0.011)	(0.004)	(0.010)	(0.003)
p-value (H_0: All coefficients=0)	0.029	0.413	0.000	0.798
City FEs	✓		✓	
Department FEs	✓		✓	
City FEs × Female		✓		✓
Department FEs × Female		✓		✓
Observations	13,487	13,487	13,487	13,487

Notes: The table reports regressions of ΔExposure × Female (columns 1–2) and Share clerks × War × Female (columns 3–4) on individual-level baseline characteristics of civil servant's children (Panel A) and their civil servant parents (Panel B). Δ Exposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919. Share clerks is the clerical share in 1915 at the office-level. Both variables are standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. War is a dummy equal to 1 if the office is in a war-related department and 0 otherwise. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). Children's number of siblings, and civil servant parents' age in 1917, (log) salary in 1915, and tenure in 1915 are standardized to mean 0 and standard deviation 1. The table also reports the p-value of a joint significance test, with the null hypothesis that all coefficients are jointly 0. Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.VI: Exposure and children's LFP gap – inclusion of family fixed effects

	(1)	(2)	(3)
	Labor force participation		
Dep. var. mean female-male gap	-0.479	-0.486	-0.486
Δ Exposure × Female	0.041***	0.023**	0.021*
	(0.009)	(0.010)	(0.012)
City × Department FEs	✓	✓	✓
City FEs × Female	✓	✓	✓
Department FEs × Female	✓	✓	✓
Age FEs × Female	✓	✓	✓
Controls × Female	✓	✓	✓
Parental household FEs			✓
Observations	13,487	9,043	9,043

Notes: The table shows a version of our OLS baseline specification (Table II, column 4, and reported in column 1), augmented by parental household fixed effects (column 3). Column 2 replicates the baseline specification, restricting the sample to match the one used in the household fixed effects specification. The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.VII: Exposure and children's LFP gap– by parent gender

	(1)	(2)	(3)
	Labor force participation		
Dep. var. mean female-male gap	-0.479	-0.529	-0.476
Δ Exposure × Female	0.041***	0.049	0.040***
	(0.009)	(0.112)	(0.010)
City × Department FEs	✓	✓	✓
City FEs × Female	✓	✓	✓
Department FEs × Female	✓	✓	✓
Age FEs × Female	✓	✓	✓
Controls × Female	✓	✓	✓
Sample	Full	Mothers	Fathers
Observations	13,487	780	12,652

Notes: The table shows a version of our OLS baseline specification (Table II, column 4, and reported in column 1), restricting the sample according to whether the parent working in the civil service in 1915 is female (col. 2) or male (col. 3). The unit of observation is the individual-year. The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.VIII: Exposure to female civil servants and propensity of child to be matched in 1940

	(1)	(2)	(3)	(4)	(5)
	Child is matched in 1940 Census				
Dep. var. mean female-male gap	-0.247	-0.247	-0.247	-0.247	-0.247
Female	-0.247***				
	(0.010)				
Δ Exposure	-0.011**				
	(0.006)				
Δ Exposure × Female	0.003	0.006	0.006	0.002	0.005
	(0.007)	(0.009)	(0.009)	(0.009)	(0.017)
City FEs	✓	✓			
Department FEs	✓	✓			
City × Department FEs		✓	✓	✓	✓
City FEs × Female		✓	✓	✓	✓
Department FEs × Female		✓	✓	✓	✓
Age FEs × Female			✓	✓	✓
Controls × Female				✓	✓
Kleibergen-Paap F-stat	-	-	-	-	46.02
Estimation method	OLS				2SLS
Observations	23,454	23,454	23,454	23,454	23,454

Notes: The table shows OLS estimated coefficients from Equation 3. The unit of observation is the individual-year. The outcome is a dummy equal to 1 if the civil servant's child has been linked to the 1940 Census. The sample includes all identified children of civil servants in 1915 who are younger than 20 in 1917, regardless of whether they are linked to 1940 Census (see Section 4.3 for a detailed description of the sample). ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Columns 1-4 report OLS estimates, whereas Column 5 reports 2SLS estimates (see Section 5.2). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.IX: Descriptive statistics of matched vs. unmatched sample – raw and reweighted

	(1) Overall mean	(2) Diff. means matched vs. unmatched	(3)	(4) Obs.
		Unweighted	Weighted	
Parent female	0.063	0.000	0.000	25,345
		(0.003)	(0.003)	
Parent city: DC	0.363	-0.030***	-0.003	25,345
		(0.006)	(0.006)	
Parent dept: War	0.392	-0.015**	-0.001	25,345
		(0.006)	(0.007)	
Share of clerks	0.157	-0.004**	-0.001	25,345
		(0.002)	(0.002)	
Child female	0.497	-0.251***	-0.005	25,345
		(0.006)	(0.007)	

Notes: The table compares 1940 census-linked and not-linked civil servants' children. The overall sample includes all identified children of civil servants in 1915 who are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample and of the linking procedures). Column 1 reports the mean characteristics of the full sample. Columns 2 and 3 report the difference in means between the linked and not-linked sample, using the raw variables (col. 2) or inverse probability weighting (col. 3). Parent female is a dummy equal to 1 if the child has a civil servant parent who is a female. DC (resp., War) is a dummy equal to 1 if the child has a civil servant parent who works in DC (resp., in a war-related department) in 1915. Share of clerks is the clerical share of the parental office in 1915. Child female is a dummy equal to 1 if the civil servant child in the sample is a female. Robust standard errors reported. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.X: Exposure and children's LFP gap – reweighting based on parental characteristics

	(1)	(2)	(3)	(4)	(5)
	Labor force participation				
Dep. var. mean female-male gap	-0.479	-0.479	-0.479	-0.479	-0.479
Female	-0.474***				
	(0.011)				
Δ Exposure	-0.003				
	(0.006)				
Δ Exposure × Female	0.020**	0.046***	0.047***	0.042***	0.089***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.021)
City FEs	✓	✓			
Department FEs	✓	✓			
City × Department FEs		✓	✓	✓	✓
City FEs × Female		✓	✓	✓	✓
Department FEs × Female		✓	✓	✓	✓
Age FEs × Female			✓	✓	✓
Controls × Female				✓	✓
Kleibergen-Paap F-stat	-	-	-	-	46.52
Estimation method			OLS		IV
Observations	13,487	13,487	13,487	13,487	13,487

Notes: The table shows a version of Table II, by reweighting civil servants' children in the sample (i.e. linked to the 1940 Census) to be comparable to their not-linked counterparts. The reweighting is based on: parental working city in 1915 (whether District of Columbia), parental working department in 1915 (whether war-related), 1915 clerical share (split in 20 bins) of the office of parental exposure, parental gender, and child's gender. The unit of observation is the individual-year. The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Columns 1-4 report OLS estimates, whereas Column 5 reports 2SLS estimates (see Section 5.2). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XI: Exposure and children's LFP gap – bounding potential selection bias

	(1)	(2)	(3)	(4)
	Labor force participation			
Dep. var. mean female-male gap	-0.479	-0.479	-0.479	-0.479
Δ Exposure × Female	0.026***	0.044***	0.024***	0.039***
	(0.007)	(0.014)	(0.006)	(0.013)
City × Department FEs	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓
Impute for missing	LFP=0		LFP=1	
Estimation method	OLS	IV	OLS	IV
Kleibergen-Paap F-stat		46.02		46.02
Observations	23,454	23,454	23,454	23,454

Notes: The table shows a version of Table II, estimated on a larger sample including all identified children of civil servants in 1915 who are younger than 20 in 1917, regardless of whether they are linked to 1940 Census (see Section 4.3 for a detailed description of the sample). The unit of observation is the individual-year. Unlinked children in the estimation sample are imputed not to be in the labor force in columns 1–2; similarly, unlinked children in columns 3–4 are imputed to all be in the labor force. Column 1 (resp., 2) reports coefficients estimated with OLS (resp., 2SLS) strategy. The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XII: Exposure and children's LFP gap – change in other workplace traits

	(1)	(2)	(3)	(4)	(5)
	Labor force participation				
Dep. var mean female-male gap	-0.479	-0.479	-0.479	-0.480	-0.480
Δ Exposure × Female	0.041***	0.052***	0.040***	0.033***	0.037***
	(0.009)	(0.012)	(0.011)	(0.013)	(0.010)
Δ (log) size × Female		-0.027			
		(0.017)			
Δ Clerk share × Female			0.000		
			(0.009)		
Δ Age × Female				0.001	
				(0.012)	
Δ Education × Female					-0.017*
					(0.009)
City × Department FEs	✓	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓	✓
Observations	13,487	13,487	13,487	11,670	11,670

Notes: The table shows a version of our OLS baseline specification (Table II, column 4, and reported in column 1). The unit of observation is the individual-year. The baseline specification is augmented with the (log) 1915-1919 change in workforce (col. 2), the 1915-1915 change in clerical share (col. 3), 1915-1915 change in the average age of civil servants (col. 4), and 1915-1919 change in the average years of schooling of civil servants (col. 5). The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e. city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XIII: Predicting increase in female share of civil servants across department-cities

	(1)	(2)	(3)	(4)
		Δ Exposure		
Mean dep. var.	0.089	0.089	0.089	0.089
War dept	0.064***	0.003	0.014	
	(0.017)	(0.020)	(0.020)	
Share clerks 1915 (%)	0.209***	0.061	-0.001	-0.049
	(0.065)	(0.053)	(0.054)	(0.077)
War dept × Share clerks 1915 (%)		0.567***	0.526***	0.416**
		(0.127)	(0.142)	(0.170)
Observations	233	233	233	233
City FE			✓	✓
Department FE				✓

Notes: The table shows the first stage of our 2SLS regression estimates (see Section 5.2). The unit of observation is city-department (i.e., office). The sample consists of the 70 cities in our main sample - that is, cities with at least 20 civil servants in a given office in 1915 and with at least two federal government departments. The outcome variable, ΔExposure, captures the change in the share of female civil servants across offices between 1915 and 1919 (see Equation 2). War dept is a dummy equal to 1 if the office is in a war-related department. Share clerks is the 1915 share of clerks in the office. Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XIV: Exposure and children's LFP gap – instrumental variable

	(1)	(2) Labor force participation	(3)	(4) Δ Exposure × Female
Dep. var. mean female-male gap	-0.479	-0.479	-0.479	
ΔExposure × Female	0.041***		0.085***	
	(0.009)		(0.021)	
Share clerk × Female		-0.025	-0.032	0.083
		(0.019)	(0.021)	(0.106)
Share clerk × War dept × Female		0.078***		0.913***
		(0.016)		(0.132)
City × Department FEs	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓
Kleibergen-Paap F-stat	-	-	48.13	-
Estimation method	OLS	Reduced	IV	First-stage
Observations	13,487	13,487	13,487	13,487

Notes: The table shows the coefficients from the regression model Equation 3, estimated with OLS and our 2SLS specification (see Section 5.2). The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section Section 4.3 for a detailed description of the sample). The outcome variable is a dummy equal to 1 if the child is in the labor force in 1940 and 0 otherwise (colums 1-3). Column 1 reports our baseline OLS estimates (Table II, column 4). Column 2 reports the reduced-form estimates, Column 3 shows the 2SLS estimates, and Column 4 reports the corresponding first-stage coefficients with instrumented exposure to female workers. ΔExposure is instrumented with the interaction between the office-level share of clerical workers in 1915 and a dummy equal to 1 for being in a war-related department (controlling for lower order interactions). ΔExposure and Share clerks are standardized to mean 0 and standard deviation 1. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XV: Exposure and children's outcome gaps - sociodemographics, IV

	(1) Married	(2) Children	(3) Education	(4) Work
Dep. var mean female-male gap	-0.099	0.066	0.287	-0.480
Δ Exposure $\times$ Female	-0.054**	-0.093**	0.176	0.080***
	(0.023)	(0.044)	(0.144)	(0.021)
City FEs $\times$ Department FEs	✓	✓	✓	✓
City FEs $\times$ Female	✓	✓	✓	✓
Department FEs $\times$ Female	✓	✓	✓	✓
Age FEs $\times$ Female	✓	✓	✓	✓
Controls $\times$ Female	✓	✓	✓	✓
Kleibergen-Paap F-stat	47.26	47.26	48.15	47.19
Observations	13,407	13,407	13,151	13,072

Notes: The table shows a version of Table III, with 2SLS estimation strategy (see Section 5.2). The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). All outcome variables are derived from the 1940 Census. In column 4, we augment our baseline specification (Table II, column 4) by including a dummy for being married, fixed effects for the number of children and years of education as additional mediation controls. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XVI: Exposure and children's outcome gaps - labor market, IV

	(1)	(2)	(3)	(4)	(5)
	Employed in		Working in		
	Federal	Private	Clerical	% Male job	Log(Income)
Dep. var. mean female-male gap	-0.033	-0.446	0.327	-0.439	-0.366
Δ Exposure × Female	0.039***	0.048**	0.042	-0.012	0.091
	(0.013)	(0.022)	(0.029)	(0.021)	(0.077)
City × Department FEs	✓	✓	✓	✓	✓
City FEs × Female	✓	✓	✓	✓	✓
Department FEs × Female	✓	✓	✓	✓	✓
Age FEs × Female	✓	✓	✓	✓	✓
Controls × Female	✓	✓	✓	✓	✓
Kleibergen-Paap F-stat	48.13	48.13	48.32	47.73	48.32
Sample	Full sample		Labor force participation=1		
Observations	13,487	13,487	8,998	8,622	8,998

Notes: The table shows a version of Table IV, with 2SLS estimation strategy (see Section 5.2). The unit of observation is the individual-year. The sample includes all children of civil servants in 1915 who could be linked to the 1940 Census and are younger than 20 in 1917 (see Section 4.3 for a detailed description of the sample). In columns 3–5, the sample is further restricted to individuals who are are in the labor force in 1940. Female is a dummy equal to 1 if the child of the civil servant is female and 0 otherwise. Age is the child's age as in 1917. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table A.XVII: Parental exposure and son's propensity to being married to a working spouse

	(1)	(2)	(3)
	Wife is working		
Mean of dep. var	0.20	0.20	0.20
Δ Exposure	-0.002	-0.003	-0.004
	(0.013)	(0.013)	(0.013)
City FEs	✓	✓	✓
Department FEs	✓	✓	✓
Age FEs		✓	✓
Controls FEs			✓
Sample	Married sons		
Observations	5,433	5,433	5,433

Notes: The table shows OLS regression estimates, restricting our baseline sample (Table II) to married sons. The outcome variable is a dummy equal to 1 if the wife is working in 1940, and 0 otherwise. ΔExposure captures the change in the share of female civil servants across offices (i.e., city-department) between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Controls include: (i) a set of child's characteristics (i.e., a dummy equal to 1 if white, and number of siblings), and (ii) a set of civil servant parent's characteristics (i.e., a dummy equal to 1 if female, age in 1917, a dummy equal to 1 if clerical position in 1915, and (log) salary in 1915). Standard errors are clustered at the city-department level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XVIII: Exposure to female civil servants and city-level labor market outcomes

	(1)	(2)	(3)	(4)
	Female LFP			
	Public sector		Private sector	
Mean of dep. var	0.0996	0.0996	0.277	0.277
Δ Exposure × Post	0.024**	0.008	0.009**	0.008
	(0.010)	(0.013)	(0.004)	(0.006)
Δ Exposure × Post × High federal employment city		0.034*		0.002
		(0.020)		(0.008)
Year FEs	✓	✓	✓	✓
City FEs	✓	✓	✓	✓
Year FEs × City-level controls		✓	✓	✓
Lower oder interactions	-	✓	-	✓
Δ Exposure × Post +	-	0.042***	-	0.010**
Δ Exposure × Post × High federal employment city		(0.015)		(0.004)
Observations	330	330	330	330

Notes: The table shows OLS regression estimates from Equation 6. The unit of analysis is the city-year. The sample consists of a balanced panel of (66) cities in our main sample covering the 1900-1940 period. The outcome variable is the city-year level share of women in the federal government sector (columns 1–2) and the private sector (columns 3–4) as derived from decennial censuses. High federal employment city is a dummy that is 1 if the city's share of federal employment is above median (1.3%) and 0 otherwise. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. City-level controls include: the 1915-1919 change in the (log) federal workforce size and (log) population in 1910. Standard errors are clustered at the city level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XIX: Exposure to female co–workers and neighborhood-level spillovers

	(1)	(2)	(3)	(4)
	Labor force participation			
Mean dep. var.	0.214	0.697	0.214	0.214
Δ City-level exposure × Post	0.086**	-0.021	0.037	
	(0.043)	(0.029)	(0.034)	
Female war-time entrant in neighborhood × Post			-0.002	0.000
			(0.002)	(0.001)
Δ City-level exposure × Post			0.077**	0.041**
× Female war-time entrant in neighborhood			(0.032)	(0.016)
Individual FEs	✓	✓	✓	✓
Year FEs	✓	✓	✓	
Year FEs × City-level controls	✓	✓	✓	
City × Age FEs	✓	✓	✓	✓
City × Year FEs				✓
Sample	Female	Male	Female	
Observations	11,991,020	13,724,485	11,990,765	11,991,276

Notes: The table shows OLS regression estimates from Equation 6. Unit of observation is the individual-year. The sample covers a panel of census-linked individuals present in the main sample of (70) cities in 1920 through the years 1900–1940. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. The outcome variable is a dummy equal to 1 if the individual is in the labor force in 1940 and 0 otherwise. Column 1 and columns 3–4 restrict the sample to women only, while column 2 restricts the sample to men. In column 3 (resp., 4), the sample is restricted to women who in 1920 lived in a census enumeration district with (resp., without) at least one female wartime civil servant. City-level controls includes: the 1915-1919 change in the (log) federal workforce size. Standard errors are clustered at the city-level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.XX: Exposure to female civil servants and city-level membership in women's clubs

	(1)	(2)
	Log(Women club membership)	
Mean of dep. var	6.721	6.721
Δ Exposure × Post	0.129**	0.131**
	(0.057)	(0.060)
Year FEs	✓	✓
City FEs	✓	✓
Year FEs × City-level controls		✓
Observations	250	250

Notes: The table shows OLS regression estimates from Equation 6. The unit of analysis is the city-year. The sample consists of a panel of cities covering the 1913–1921 period, with observations available biennially. The outcome variable is the (log) total number of members in women clubs. ΔExposure is the change in the share of female civil servants across cities between 1915 and 1919 and it is standardized to mean 0 and standard deviation 1. Post is a dummy equal to 1 for the years 1920 and later and 0 otherwise. City-level controls include: the 1915-1919 change in the (log) federal workforce size and (log) population in 1910. Standard errors are clustered at the city level. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

B Appendix – Linking Procedure

Table B1: Identifying children of civil servants in adulthood

	Total	Share female
Civil servants in 1915 (parents)		
(1) Civil servants employed in 1915	161,773	0.122
(2) Matched to 1900-1920 Census based on full name and birth state	67,194	0.123
Children of civil servants between 1900-20		
(3) Children found in households of matched civil servants in (2)	75,719	0.498
(4) Civil servants in (2) that are parents of children in (3)	29,290	0.069
Children of civil servants in 1940		
(5) Children from (3) matched via Census Tree	42,889	0.400
(6) Children from (3) matched via Multigenerational Longitudinal Panel	25,843	0.306
(7) Children from (3) matched via Numident-to-census	2,469	0.562
(8) Children from (3) matched via Census Linking Project	13,346	0.001
(9) Children from (3) matched via full name, birth state, age	6,187	0.195
Pooled sample of civil servants' children in 1940		
(10) Combined sample of unique children from (5)-(9)	45,104	0.397
(11) Data quality check in (10)	42,325	0.397
(12) Civil servant parents of children in (11)	21,947	0.065
(13) Imposing large city restriction in (11)	18,921	0.393
(14) Imposing age restriction (<20) in 1917 in (13)	14,624	0.391
(15) Dropping children with missing data	13,487	0.389

Notes: The table summarizes the main steps of the linking procedures that have been undertaken in this paper. See Section 4.3 for a detailed description. At each census round, a child is identified as a household member who shares the same last name as the civil servant's, is at least 18 years younger than the parent, and is aged 18 or younger in that census year. The sample is always restricted to children born before 1918. See Section 4.1 for details on the restriction to large cities. Census Tree dataset is a cross-census linkage developed by Buckles et al. (2023). Multigenerational Longitudinal Panel and Numindent-to-Census crosswalks are provided by Ruggles et al. (2024). The Census Linking Project provides a crosswalk to link census respondents across time developed by Abramitzky et al. (2020).

Table B2: Identifying spouses of civil servants

	Total	Share female
Civil Servants in 1915		
(1) Civil servants employed in 1915	161,773	0.122
(2) Matched to 1900-1940 Census based on full name and birth state	80,945	0.120
(3) Male civil servants in (2)	71,261	0.000
Spouses of civil servants between 1900-1940		
(4) Spouses found in households of matched civil servants in (3)	56,378	1.000
(5) Imposing large city restriction in (4)	26,228	1.000
(6) 1900-1940 pooled sample of (4)	94,092	1.000
(7) Imposing large city restriction in (6)	44,267	1.000

Notes: The table summarizes the main steps of the linking procedures that have been undertaken in this paper. At each census round, a spouse is identified within the civil servant's household through the `SPLOC` variable provided by IPUMS. See Section 4.1 for details on the restriction to large cities.

C Appendix – Validation of alternative gender norms measure

In Section 6.3, we complement our main results using a direct measure of civil servants' attitudes toward working women. Specifically, we proxy favorable attitudes towards women's labor force participation with men's propensity to marry a working woman. In what follows, we validate this alternative measure of gender norms.

Ideally, one would need to have survey data contemporaneous to the period of our study in which respondents are asked not only about gender norms but also about their marital status and their spouse's labor market participation. To the best of our knowledge, no historical survey collects such information for our period of analysis. We overcome the lack of historical data by relying on data from the World Values Survey (henceforth, WVS). The WVS is a collection of national, individual-level surveys on a wide range of topics, including attitudes towards gender roles and gender equality. The survey also contains information on demographic characteristics, such as gender, age, education, labor market status, and income. The survey has been administered seven times, from 1981 to 2021. Importantly for our purpose, the seventh wave (2017-2022) records the employment status of respondents' spouse. We can then investigate whether men with a working spouse have more favorable attitudes towards women's labor force participation and in general, gender equality in the labor market.

We restrict WVS data from Wave 7 to respondents in the United States with a spouse and with non-missing employment status. These restrictions yield a sample of 1,481 individuals. We construct a comprehensive measure of positive attitudes towards women's labor force participation, using a z-score index over a set of questions previously used in the literature (Alesina et al., 2013; Bertrand et al., 2020).[45] With this index at hand, we investigate whether marital decisions predict variation in attitudes towards women in the labor market. In particular, we regress the gender norms z-score against a dummy variable equal to one if the respondent is working, and a dummy equal to one if the spouse is working.[46] We also include respondent's age, income, and education as additional individual controls.

We first run the regression restricting the sample to married men. Estimated coefficients are reported in Table C2, columns 1 and 2. As expected, being married to a working wife is positively and significantly associated with progressive attitudes on gender roles in the labor market, whereas the male respondent's own employment status shows no significant relationship. We then estimate the same regression separately for married women (Table C2, columns 3 and 4). Estimated coefficients reveal a reversed pattern, where a woman's own employment status is significantly associated with progressive attitudes toward gender roles, while her spouse's employment status has no significant effect. We interpret these results - differential by respondent's gender - as a validation for our alternative measure of gender norms, relying on male civil servants' marital decisions (see Section 6.3).

[45] Table C1 reports the exact wording of the survey questions. All questions are rescaled so that the higher values stand for more positive attitudes towards gender equality.

[46] In our analysis sample, around 70% of male respondents work and 63% of men's spouses work. Whereas, around 60% of female respondents work and 80% of their spouses work.

Table C1: WVS gender norms index: questions

Question	**Answers coded as**
When jobs are scarce, men should have more right to a job than women	1 = strongly agree to 5 = strongly disagree
If a woman earns more money than her husband, it's almost certain to cause problems	1 = strongly agree to 5 = strongly disagree
Being a housewife is just as fulfilling as working for pay	1 = strongly agree to 4 = strongly disagree
On the whole, men make better business executives than women do	1 = strongly agree to 4 = strongly disagree

Notes: The table reports the exact wording and answer scale of the questions in the World Value Survey, Wave 7 (2017-2022) that we aggregate to construct a comprehensive measure of gender norms. All questions are rescaled so that higher values stand for more positive attitudes toward women and gender equality in the labor market.

Table C2: Cvil servants' marital decisions – WVS validation

	(1)	(2)	(3)	(4)
	Gender norms z-score			
Mean dep. var.	-0.021	-0.017	0.055	0.065
Working respondent	0.046	-0.079	0.173***	0.137***
	(0.052)	(0.055)	(0.050)	(0.051)
Working spouse	0.248***	0.209***	0.000	-0.078
	(0.049)	(0.049)	(0.064)	(0.073)
Observations	823	812	658	638
Sample	Men	Men	Women	Women
Individual controls		✓		✓
Mean working repondent	0.720	0.718	0.622	0.622
Mean working spouse	0.633	0.633	0.809	0.810

Notes: The table shows OLS regression estimates. The sample is restricted to WVS Wave 7 respondents in the United States with a spouse. Columns 1 and 2 (resp., columns 3 and 4) restrict the sample to respondents who are married men (resp., women). The outcome variable is a z-score averaged across a set of questions on gender norms in the labor market: higher values indicate more positive attitudes toward women in the labor market. See Table C1 for the exact wording of the questions under scrutiny. Working respondent (resp., working spouse) is a dummy equal to 1 if the respondent (resp., the respondent's spouse) is working (full-time, part-time, or self-employed). Individual controls include: respondent's age, self-reported income and education. Robust standard errors reported. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.