World War I and Female Labor Force Participation
The Case of England and Wales*

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Abstract

Combining district-level measures of World War I military fatalities with full-count census microdata, we show that wartime mortality increased postwar female labor force participation in England and Wales. Women shifted away from traditionally female occupations, while military deaths increased the shares of single and widowed women, whose participation rates were higher. Yet the overall effect of military mortality on postwar female labor force participation remained modest. A harmonized comparison with France suggests that stronger internal migration in England and Wales rapidly dissipated local marriage- and labor-market imbalances, while more generous pensions for war widows weakened their labor supply response.

Keywords World War I, Female labor force participation, England and Wales

JEL codes J16, J21, N33, N34

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1. Introduction

Wars often leave profound and enduring legacies for women's economic and social roles (Munroe et al., 2023; Anderson and Sviatschi, 2025). During both World Wars, mass military mobilization drew unprecedented numbers of women into jobs previously held by men, while military deaths altered postwar marriage and labor markets. A large literature, particularly on World War II in the United States, has examined how wartime mobilization reshaped women's postwar employment (e.g., Goldin, 1991; Acemoglu, Autor and Lyle, 2004; Rose, 2018; Shatnawi and Fishback, 2018).[1] A more recent literature instead focuses on the distinct consequences of military deaths. For instance, Boehnke and Gay (2022) find that World War I military fatalities in France generated a large and persistent increase in female labor force participation, driven primarily by labor supply responses among single women and war widows.[2]

World War I in England and Wales presents an important puzzle. The massive wartime expansion of female employment proved largely temporary, receding rapidly with demobilization (Braybon, 1981; Thom, 1998). Yet the war also left a permanent demographic imprint: more than 600,000 men from England and Wales never returned (Winter, 1985). Despite these losses, female labor force participation showed no aggregate increase by 1921 (Humphries, 1995; Hatton and Bailey, 2001). Did women's labor force participation nevertheless increase in areas that experienced greater military fatalities? More broadly, under what conditions do wartime demographic shocks translate into postwar changes in female labor force participation?

We combine the full-count 1911 and 1921 censuses of England and Wales with a difference-in-differences design that exploits spatial variation in military death rates across nearly two thousand local government districts. We find that military deaths increased female labor force participation, but only modestly: a 5 percentage point higher military death rate—one standard deviation across districts—raised participation by 0.5 percentage points, equivalent to 1.5 percent of the prewar mean. Military deaths affected both sides of the labor market. On the demand side, women shifted away from traditionally female occupations, consistent with employers substituting women for scarce male workers. On the supply side, the marital composition shifted toward single and widowed women, who participated in the labor force at far higher rates than married women.

Why was the overall effect of military deaths on female labor force participation so limited? We investigate this question by comparing England and Wales with France,

[1] See Ferrara (2025) for a review of this literature.

[2] The consequences of wartime fatalities for women's postwar employment have also been documented for World War II in the United States (Brodeur and Kattan, 2022) and Germany (Braun and Stuhler, 2024).

harmonizing the units of aggregation, definitions of the outcome and treatment variables, and estimation strategies across settings. Under this common framework, female labor force participation was an order of magnitude more responsive to military death rates in France than in England and Wales. Two mechanisms help explain this difference. First, internal migration responded more to local military deaths in England and Wales than in France, helping marriage- and labor-market imbalances dissipate more rapidly. Second, the labor supply response of widows to military deaths was large in France but essentially null in England and Wales, where pensions for war widows were considerably more generous.

This article makes three contributions. First, we construct new data on the local incidence of World War I mortality in England and Wales. These data complement recent work exploiting spatial variation in British military fatalities to study civic capital, innovation, and political behavior (Carozzi, Pinchbeck and Repetto, 2026; Repetto et al., 2026; Tchaouchev, 2026), while addressing the measurement challenges created by incomplete information on soldiers' places of residence in archival records. They also contribute to a broader literature using individual-level military records to construct spatially disaggregated measures of World War I mortality, including Fornasin (2017) for Italy, De Juan et al. (2024) for Germany, and Gay and Grosjean (2023) for France.[3]

Second, we extend the growing literature on the effects of wartime mortality on postwar female labor force participation by providing evidence on the case of England and Wales (Boehnke and Gay, 2022; Brodeur and Kattan, 2022; Braun and Stuhler, 2024). Our findings further speak to a longstanding historiographical debate over whether World War I was a watershed for British women's work, with recent scholarship arguing that wartime changes in women's employment were largely temporary (Braybon, 1981; Thom, 1998; Grayzel, 2002). Our results qualify this interpretation: although the wartime expansion of female employment was largely reversed, areas that suffered greater military losses experienced a persistent, if modest, increase in female labor force participation.

Third, our comparative analysis speaks to a broader literature on the external validity of causal estimates across contexts (Egami and Hartman, 2023; Slough and Tyson, 2025). By harmonizing the research design across settings, we identify broader conditions shaping the effect of wartime mortality on female labor, showing that cross-country heterogeneity depends on the extent to which internal migration and social insurance absorb the demographic shock.

[3]Our data contribution also extends to the outcome variable: we construct a corrected measure of female labor force participation based on Schürer and Wakelam's (2025) 1921 census microdata, as their proposed labor force status variable materially understates female labor force participation in that year.

The remainder of the article proceeds as follows. Section 2 provides historical background on women's employment before, during, and after the war. Section 3 describes our data and measurement of female labor force participation and military death rates. Section 4 presents our empirical strategy and results. Section 5 places the case of England and Wales in comparative perspective using France as a benchmark. Section 6 concludes.

2. Women, Work, and World War I

2.1. Women's Work before the War

On the eve of World War I, women's labor force participation in England and Wales had been stagnant since the late nineteenth century, with the aggregate participation rate among women aged 10 and over remaining at 38 percent (Humphries, 1995, pp. 78–80; Hatton and Bailey, 2001, pp. 87–88). This stagnation occurred within a labor market characterized by a pronounced gender division of labor, as female employment was highly concentrated by both marital status and occupation. In 1911, single and widowed women accounted for 86 percent of the female labor force, despite comprising only 55 percent of women aged 10 and over. In addition, domestic service, textiles, and clothing together represented 66 percent of female employment.[4] The institutional environment further reinforced these occupational boundaries: trade-union restrictions, protective labor legislation, and marriage bars all constrained women's access to employment (Humphries, 1995, pp. 83–85).

2.2. Wartime Mobilization and the Demand for Female Labor

The war radically increased the demand for female labor through two simultaneous shocks. First, mass military mobilization dramatically reduced the supply of male workers: by July 1918, 46 percent of the prewar male workforce had been enlisted (Dewey, 1984, pp. 204–5). Second, industrial warfare sharply increased labor demand, particularly following the creation of the Ministry of Munitions in May 1915 (Lloyd-Jones and Lewis, 2016, pp. 174–198). Together, these forces created severe labor shortages and produced an unprecedented expansion and reallocation of female employment: over the course of the war, the number of working women rose from 6 to 7.3 million, while 2 million women moved into jobs previously held by men (Kirkaldy, 1921, Table 1, p. 2).[5]

[4]These figures are calculated from data reported in the *Census of England and Wales, 1911*, Volume 10, *Occupations and Industries*, Part 1 (London, 1914), Table 8, pp. 75–88.

[5]The estimate of total female employment is based on figures reported in the War Cabinet Committee on Women in Industry's (1919) *Women in Industry* report (pp. 80–81).

Over this period, women moved into occupations from which they had previously been largely excluded, often by leaving traditionally female occupations (Braybon, 1981, pp. 44–50; Thom, 1998, pp. 34–38). This shift was particularly pronounced in metalworking, where female employment in private firms rose from 170,000 to 597,000 over the course of the war, with a further 247,000 women employed in government arsenals (Downs, 1995, p. 41). Yet the expansion of female employment extended well beyond war production: by the end of the war, 477,000 women had replaced men in banking, finance, and commerce, compared with 415,000 in the metal and chemical industries (Kirkaldy, 1921, Table 1, p. 2). Nevertheless, much of this expansion was explicitly understood as temporary. Under the Treasury Agreement of March 1915 and other subsequent agreements, trade unions accepted the wartime substitution of women in exchange for the commitment that prewar employment practices would be restored after the war and that returning servicemen would recover their former positions (Braybon, 1981, pp. 50–57; Thom, 1998, pp. 56–59).

2.3. Demobilization and the Postwar Labor Market

The wartime expansion of female employment reversed rapidly after the Armistice as war production contracted and demobilized servicemen returned (Braybon, 1981, pp. 173–215; Downs, 1995, pp. 192–202). By July 1920, female employment had fallen by more than 800,000 relative to November 1918 (Kirkaldy, 1921, Table 1, p. 2). Public policy reinforced this reversal. In particular, the Restoration of Pre-War Practices Act of 1919 facilitated the return of servicemen to their prewar jobs (Rubin, 1989). At the same time, women who refused employment deemed suitable by labor exchanges—often in traditionally female occupations—risked losing unemployment benefits (Braybon, 1981, pp. 180–82; Thom, 1998, pp. 191–93). More broadly, many of the institutional barriers that had constrained women's employment before the war were restored in the 1920s (Braybon, 1981, pp. 216–28). Taken together, these patterns suggest that, by many aggregate measures, World War I was less a watershed than a temporary parenthesis for women's labor: by 1921, female labor force participation had largely returned to its prewar trajectory, while the occupational division of labor remained strongly gendered (Braybon, 1981, pp. 216–28; Thom, 1998, pp. 201–8; Grayzel, 2002, pp. 106–9).

3. Data and Measurement

We now describe the construction of our measures of female labor force participation (Section 3.1) and military death rates (Section 3.2). While our unit of observation is the individual, we measure exposure to wartime mortality across local government districts

(hereafter, districts). England and Wales were divided into nearly two thousand such districts, whose relatively small size makes them a useful approximation to local labor markets.[6]

3.1. Female Labor Force Participation (1911–21)

Measuring female labor force participation We capture women's labor force participation using the 1911 and 1921 population censuses of England and Wales. Historical censuses have long been criticized for understating women's economic activity, particularly irregular employment and home-based activities among married women (Folbre, 1991; Humphries and Sarasúa, 2012). Recent research, however, suggests that the underenumeration of women's *regular* employment was relatively limited in the historical censuses of England and Wales, even among married women (Anderson, 1999; Higgs and Wilkinson, 2016; You, 2020*a*). Moreover, by the 1911 census, many of the systematic sources of underenumeration identified in earlier censuses had been addressed (Hatton and Bailey, 2001).[7] The 1911 and 1921 censuses thus provide an appropriate basis for measuring female labor force participation.

Our main analysis draws on the full-count original household schedules of the 1911 and 1921 censuses of England and Wales, digitized by FindMyPast and processed by the Integrated Census Microdata (I-CeM) project (Schürer and Higgs, 2025; Schürer and Wakelam, 2025). To capture labor status, we rely on I-CeM's `INACTIVE` variable, which records whether an individual was a labor force participant at the time of the census. We classify an individual as active when this variable equals `0` (working) or `3` (unemployed), and as inactive otherwise. The `INACTIVE` variable is derived from a combination of the original occupational string (`OCC`) recorded in household schedules and I-CeM's standardized occupational code (`OCCODE`) variable (Schürer, Penkova and Shi, 2015). For 1911, I-CeM processed occupations using the census's original Hollerith coding system. By contrast, the processing of occupations for the 1921 census required a series of additional and complex steps, as the logic of occupational classification differed markedly in that census compared to 1911.[8] These procedures, however, introduced a number of inaccuracies into the 1921 `INACTIVE` variable, leading to a pronounced underestimation of

[6] Appendix D describes our methodology for constructing 1,721 consistent local government districts across the 1911 and 1921 censuses.

[7] The instructions for reporting women's occupations were also broadly consistent across the 1911 and 1921 censuses. Appendix Figure A1 reproduces the original household-schedule instructions in both years.

[8] To ensure comparability across the 1911 and 1921 censuses, I-CeM created an additional variable (`HOLLERNOS`) that maps 1921 occupational data onto the `OCCODE` scheme used for 1911. See `https://www.campop.geog.cam.ac.uk/research/projects/icem/construction.html`.

women's labor force participation relative to aggregate data published in the 1921 census report. We document these issues in Appendix B and implement a series of corrections to the 1921 `INACTIVE` variable so as to match the published aggregate statistics.

Female labor force participation rates (1911–21) In Table 1, we report the marital status distribution among women aged 20 to 54 in 1911 and 1921 (Panel A) together with their average labor force participation rates (Panel B).[9,10] Despite the major demographic disruption caused by the war, women's aggregate marital-status composition changed relatively little between 1911 and 1921—consistent with the rapid postwar adjustment of marriage patterns documented by Winter (1985, pp. 249–78). Indeed, the share of single women fell by 2 percentage points, while the shares of married and widowed women each rose by 1 percentage point. The war likewise had limited aggregate implications for women's labor force participation, which fell by 2 percentage points from 35 to 33 percent.[11] We display the spatial distribution of female labor force participation rates across districts in 1911 and 1921 in Panels a and b of Figure 1. Panel c displays the corresponding district-level changes between 1911 and 1921.

Table 1. Female Marital Statuses and Labor Force Participation Rates (1911–21)

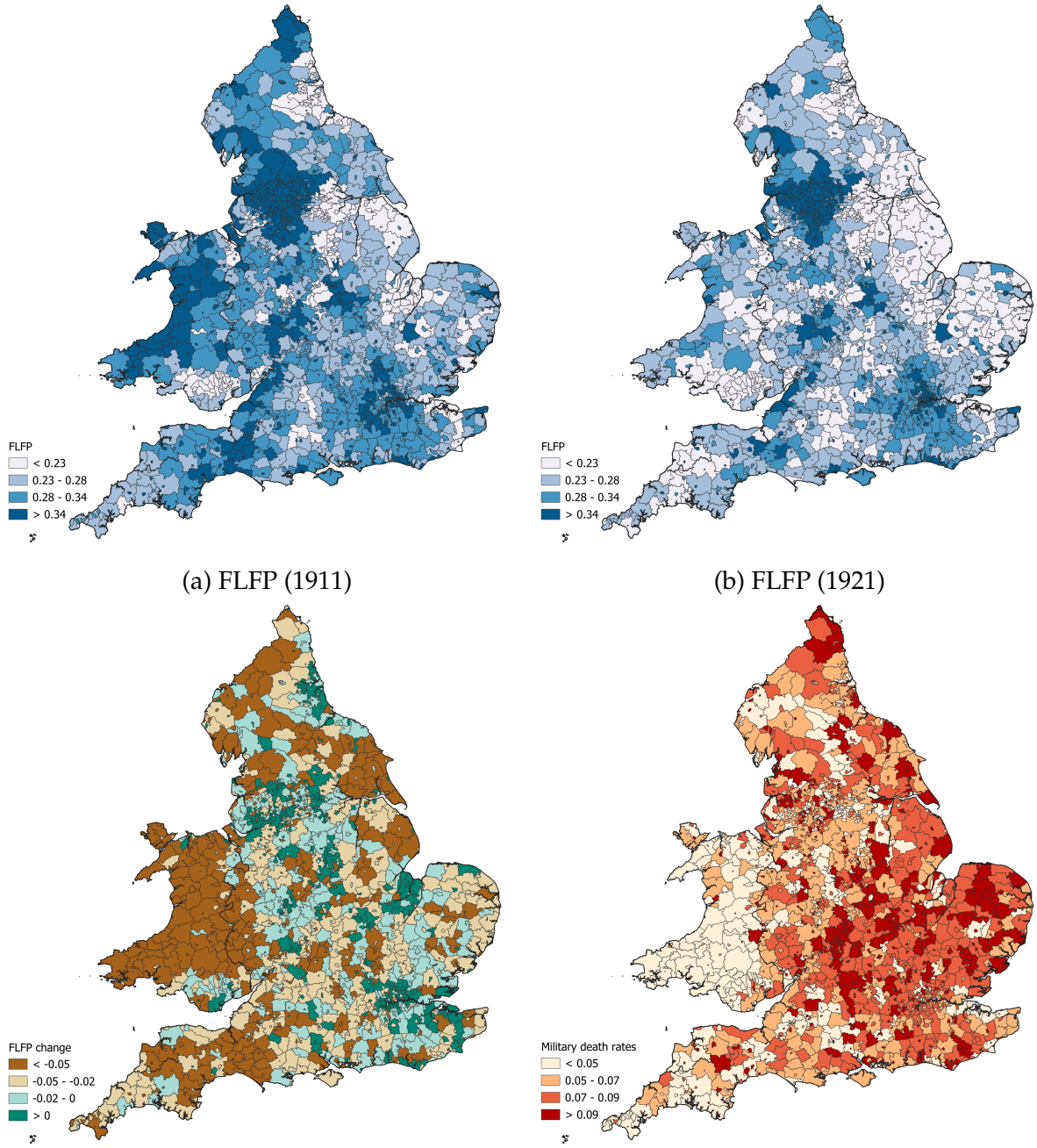
	A. Marital statuses (%)			B. Labor force participation rates (%)			
Marital status	Single	Married	Widowed	All	Single	Married	Widowed
1911	36.2	59.1	4.6	34.5	72.7	9.3	55.8
1921	33.8	60.4	5.7	32.6	72.8	9.2	42.2
Difference	−2.4	1.3	1.1	−1.8	0.1	−0.2	−13.6

Notes. Panel A provides the marital status distribution in percent among women aged 20 to 54. Panel B provides average female labor force participation rates in percent across marital statuses among women aged 20 to 54. Data are based on I-CeM's full-count household schedules of the 1911 and 1921 censuses of England and Wales (Schürer and Higgs, 2025; Schürer and Wakelam, 2025). *Married* only includes married women whose spouse is not explicitly noted as absent (`MAR == 2`). Labor force participants include women who are categorized as working (`INACTIVE == 0`) or unemployed (`INACTIVE == 3`) based on the corrected `INACTIVE` variable (see Appendix B).

[9] To ensure comparability across marital-status groups, we restrict the sample to women who were single, married with a spouse present, or widowed based on I-CeM's `MAR` variable. Married women whose spouse was absent accounted for 5.2 percent of women aged 20 to 54 in 1911 and 5.6 percent in 1921. Women with an unknown marital status made up 0.4 percent of this age group in both years. In addition, there were 308 divorced women in 1911 and 5,548 in 1921. We show in Appendix C that marital status figures derived from household schedules closely match those provided in census reports.

[10] We focus on the core working-age female population to abstract from education and retirement decisions and because household schedules appear less reliable for younger and older women (see Appendix B).

[11] Appendix Tables A1 and A2 provide data similar to Table 1 tabulated by age group. Complete summary statistics by census year both at the individual and district levels are provided in Appendix Table A3.



(a) FLFP (1911)

(b) FLFP (1921)

(c) Change in FLFP (1911–21)

(d) Military Death Rates (1914–18)

Figure 1. Female Labor Force Participation and Military Death Rates (1911–21)

Notes. This figure maps female labor force participation (FLFP) across local government districts in 1911 (Panel a) and 1921 (Panel b), the district-level change in female labor force participation from 1911 to 1921 (Panel c), and military death rates (Panel d). To facilitate comparison over time, Panels a and b use identical class breaks, defined by quartiles of the 1921 distribution. Class breaks in Panels c and d are based on quartiles of the respective variable distributions. Shapefiles are adapted from Southall, Aucott and Burton (2024). See Appendix D for details on the construction of consistent district boundaries.

3.2. Military Death Rates

3.2.1. Constructing Military Death Rates

We define the military death rate of a district as the number of military fatalities among men who resided in that district at the time of enlistment divided by the district's number of men potentially liable for military service. Below, we describe the sources we use to construct the numerator and denominator of our baseline measure of military death rates.

Military fatality records We collect information on military fatalities from England and Wales from the *Soldiers Died in the Great War* (SDGW) military roll, which comprises 661,827 unique entries of deceased soldiers of the British Regular Army and Territorial Force. SDGW records were published in 1921 as 80 regiment- and corps-specific volumes.[12] They report each soldier's first and last name, date of death, military rank, regiment, and service number, as well as places of birth, enlistment, and—most importantly for our purpose—residence at enlistment.[13]

Despite their scale, SDGW records do not encompass the universe of British military fatalities—as a benchmark, we use the estimate of 744,702 British armed forces fatalities reported in Parliament in 1921.[14] Indeed, SDGW records do not include the 41,472 records of deceased officers of the Regular Army and Territorial Force, whose entries were published in the *Officers Died in the Great War* (ODGW) military roll. They also omit the 34,874 records of deceased Royal Navy personnel, whose entries were published in the *Royal Navy and Royal Marine War Graves Roll* (RNWGR).[15] We do not include these records in the numerator of our baseline measure of military death rates because neither source reports places of residence.[16]

Conversely, SDGW records may include soldiers who were listed as missing—and thus presumed dead—when the roll was compiled in 1921, but who in fact survived the war. To ensure that our measure captures soldiers who actually died during the conflict, we link SDGW records to those of the Commonwealth War Graves Commission (CWGC).

[12] We draw on the digitized SDGW volumes available online through Ancestry at `https://www.ancestry.co.uk/search/collections/1543`. Appendix Figure A2 provides an example entry from both the original source and its corresponding Ancestry record.

[13] Panel A of Appendix Table A4 provides field-level availability across SDGW records.

[14] Appendix F.1 evaluates this estimate against alternative estimates.

[15] SDGW records further omit the 8,590 records of deceased Royal Naval Division personnel and the 6,166 records of deceased Air Force personnel. See Appendix F.1 for details.

[16] Both military rolls were published around the same time as the SDGW and are now available online via Ancestry: ODGW records are bundled with the SDGW collection and RNWGR records are available at `https://www.ancestry.co.uk/search/collections/1963`. Panels B and C of Appendix Table A4 provide field-level availability across both types of records.

Established in 1917, the CWGC commemorates Commonwealth service personnel who died in the war and maintains their graves. CWGC records can be regarded as highly reliable as they have been continuously updated over time and verified against multiple administrative sources.[17] CWGC records report each serviceman's first and last names, service number, rank, regiment, unit, medals and honors, date of death, age, burial location, and an "additional information" field that sometimes lists next-of-kin names and addresses—typically those of parents or spouses. Importantly, these records do not report servicemen's places of residence, explaining why we primarily rely on SDGW records. Using a 25-step linking algorithm described in Appendix E, we match 659,412 SDGW records to a CWGC record—a linking rate of 99.6 percent. This leaves 2,415 SDGW records unmatched, likely soldiers presumed dead who survived the war.[18] Taking 744,702 as the total number of British armed forces fatalities, our reliance on SDGW–CWGC records implies that 88.5 percent of British military fatalities are accounted for in our data.

Coverage of military fatalities The resulting SDGW–CWGC linked records report a precise place of residence at enlistment in 48 percent of cases.[19] Using these location strings, we geocode 99.5 percent of them using Google's Geocoding API and map each resulting set of coordinates to a consistent 1911–21 district. In total, 273,803 SDGW–CWGC records are geocoded in England and Wales, accounting for 86 percent of all geocoded records. Nearly all the remaining 14 percent are geocoded in Ireland or Scotland.[20] Because the place of residence at enlistment is missing from many SDGW–CWGC records, the observed spatial distribution of military fatalities may not accurately reflect their true spatial distribution. Reassuringly, this restricted sample is broadly representative of the full sample along observable soldier characteristics—their birth dates, death dates, and military ranks.[21]

[17] CWGC records include all members of Commonwealth military forces and certain auxiliary organizations who died while in service, as well as those who died after discharge if their death was attributable to wartime service and occurred between August 4, 1914, and August 13, 1921. The CWGC database is publicly accessible at `https://www.cwgc.org/find-records/find-war-dead`. An example CWGC record is shown in Appendix Figure A2. Panel D of Appendix Table A4 provides field-level availability across CWGC records.

[18] This process also enables us to impute some missing values in SDGW records and to augment them with additional information from CWGC records (see Appendix Table A5).

[19] Appendix Table A5 reports the field-level availability of variables in SDGW–CWGC records. These records contain a precise place of enlistment for 99.5 percent of cases and a precise place of birth for 90 percent.

[20] Appendix Table A6 reports geocoding rates among SDGW–CWGC records. The figure of 86 percent of records geocoded in England and Wales ($\simeq$ 1 – 691,820/4,970,902) is consistent with the share of enlisted men from Scotland and Ireland reported in the War Office's (1921, p. 9) *General Annual Reports*.

[21] Appendix Table A7 examines potential selection on observable characteristics at three stages. Panel A compares the full sample of SDGW–CWGC records with the subset for which soldiers' place of residence is observed. Panel B compares this subset with the records that we are able to geocode. Panel C compares

To account for this partial coverage, we scale the resulting district-level counts of military fatalities by a correction factor. Given that about 100,000 British military fatalities were from Scotland (Watt, 2019, pp. 83–87) and 30,000 from Ireland (Fitzpatrick, 2015, p. 645), we estimate that about 615,000 of the 744,702 British military fatalities were among men residing in England and Wales at the outbreak of the war.[22] This implies a scaling factor of 2.245 ($\simeq$ 614,702/273,803). However, using a countrywide scaling factor implicitly assumes that the residence field is missing at a uniform rate across England and Wales. In practice, the rate at which this field is missing varies widely over space, reflecting the regimental structure of SDGW volumes: in some regimental volumes, places of residence are almost entirely absent, whereas in others, they are almost always reported (see Appendix Figure G1). We thus correct for the underreporting of soldiers' places of residence by scaling records reporting a place of residence within each enlistment place by the inverse of that enlistment place's residence reporting rate. District-level fatality totals are then obtained by summing these enlistment-place adjusted counts across all enlistment places.[23]

Number of men liable for military service The denominator of our measure of military death rates is the number of men potentially liable for military service over the course of the war in a given district. Recruitment into the British armed forces proceeded in two phases: voluntary enlistment until 1916, followed by conscription (Winter, 1985, pp. 25–39).

The expansion of the Regular Army initially relied on voluntary enlistment among men aged 19 to 30, with the upper age limit gradually raised to 40 by May 1915 (Simkins, 1988).[24] After a recruiting boom in the summer of 1914, voluntary enlistment declined sharply from October onward, as the needs of wartime industry constrained the supply of new recruits. In an effort to revive voluntary recruitment, the Derby Scheme was introduced in October 1915, requiring available men aged 18 to 41 to attest their willingness to serve if called up. By December 1915, however, nearly half of eligible single men had failed

geocoded records with those geocoded in England and Wales. While we find small differences in birth and death dates across samples—and none in military rank—these differences are negligible, averaging less than one month for birth dates and ten days for death dates.

[22] See Appendix F.2 for an assessment of the accuracy of this estimate.

[23] See Appendix G.1 for details on the enlistment-place adjustment procedure. We show that district-level residence reporting rates are uncorrelated with our adjusted measure of military death rates but strongly positively correlated with the unadjusted measure. We also find no evidence that residence reporting rates are correlated with either female labor force participation in 1911 or its change between 1911 and 1921.

[24] Appendix Table A8 summarizes changes in enlistment rules over time, including eligible ages and corresponding birth cohorts. Appendix Figure A3 displays monthly enlistment figures by enlistment regime.

to attest. As a result, conscription was introduced in March 1916.[25] It initially applied to single men and childless widowers aged 18 to 40 unless they qualified for exemption, such as employment in a reserved occupation (Dewey, 1984). It was soon extended to married men and, by April 1918, to all men aged 18 to 50. This final age range corresponds to men aged 11 to 43 in 1911. We therefore use men in this age group to define the population potentially liable for military service over the course of the war, which forms the denominator of our measure of military death rates.[26] This denominator captures potential legal liability rather than effective availability, as recruitment was constrained by occupational exemptions, family circumstances, and medical fitness (see Section 3.2.2).[27]

In total, nearly 5 million men enlisted in the British Army over the course of the war, including 4.3 million from England and Wales (Winter, 1985, p. 28). This amounted to 45 percent of the male population potentially liable for military service. About half enlisted voluntarily, while the remainder were conscripted.[28]

Baseline military death rates Under our definition of military death rates, 6.5 percent of men residing in England and Wales at the outbreak of the war and eligible for military service died during the conflict ($\simeq$ 614,702/9,523,330). We display the spatial distribution of military death rates across districts in Panel d of Figure 1. At this level of aggregation, they averaged 7.7 percent with a cross-district standard deviation of 4.9 percentage points.[29]

Alternative measures of military death rates We construct several alternative measures of military death rates that could, in principle, be used instead of our baseline measure. Each of these measures attempts to address the issue of partial coverage of the numerator—the number of military fatalities—in a different way. The first increases the number

[25] Appendix Table A9 lists the British Military Service Acts enacted between 1916 and 1918 along with their main provisions.

[26] Appendix Figure A4 and Appendix Table A11 document the distribution of military fatalities by age in 1911 based on SDGW–CWGC records geocoded in England and Wales in which the year of birth could be recovered from age at death. More than 98 percent of these fatalities fall within the 11–43 age range.

[27] We use this age-based criterion because a direct measure of the population potentially liable for military service is not feasible: two-thirds of World War I soldier service records were destroyed in September 1940 when an incendiary bomb struck the War Office Record Store in London. The surviving files later became known as the First World War "burnt documents."

[28] Appendix Table A10 reports enlistments in the Regular Army and Territorial Force from August 1914 to November 1918, together with enlistment rates relative to the male population aged 11 to 43 in 1911.

[29] Spatial patterns in this map should be interpreted with caution as 63 percent of districts were classified as urban (see Appendix Table D1), and their small geographic area makes them difficult to discern visually. It nevertheless reveals a clear Welsh–English contrast, with military death rates generally lower in Wales than in England. Indeed, enlistment rates were markedly lower in Wales (Appendix Table A10), as the formation of the Welsh Army Corps in 1914 was delayed and politically contentious (Simkins, 1988, pp. 96–99).

of records with an observed residence at enlistment by matching SDGW–CWGC records to the 1911 census. The second uses the place of enlistment when the place of residence at enlistment is missing. The third uses the place of birth when the place of residence at enlistment is missing. We describe the construction methodology of these measures in detail in Appendix G.2 and show that all are subject to serious biases, due to migration for the enlistment- and birth-place complemented measures and to selection for the linked-place complemented measure.

3.2.2. Determinants of Military Death Rates

We now explore the systematic determinants of military death rates, which may be correlated with changes in female labor force participation. They can be decomposed into two components: the proportion of men liable for military service who enlisted in the armed forces and the rate of wartime mortality among those who enlisted. Although partly overlapping, the sources of variation in these two components differed.

Military enlistment Enlistment rates were a primary source of cross-district variation in military death rates, with only 45 percent of eligible men enlisting nationwide. A first determinant of a district's enlistment rate was its demographic structure, as eligibility during the voluntary enlistment period was governed primarily by statutory age limits. Conscription preserved age as a central criterion, but also made family circumstances an explicit margin of eligibility. The baseline health of a district's population also mattered, since men in poor health could be classified as unfit and outright excluded from military service (Winter, 1985, pp. 48–64; McDermott, 2011, pp. 180–97).

Another key determinant of a district's enlistment rate was its occupational structure due to the system of protected occupations (Lloyd-Jones and Lewis, 2016, pp. 313–55). Initially, protection from military service was uneven and mainly concerned railway and munitions workers. From the summer of 1915 onward, these sector-specific protections evolved into a more systematic regime of certified occupations, which protected workers deemed essential to the war economy (Dewey, 1984, pp. 213–17; Winter, 1985, pp. 39–48). For instance, the November 1916 list of certified occupations contained more than 600 occupations, covering up to 57 percent of the working-age male population.[30] As a result, enlistment rates varied widely across sectors: 35 percent of male workers in

[30] We calculate the figure of 57 percent by matching the 1916 list of certified occupations to I-CeM's `OCCODE`, `HOLLEROCC`, and `HOLLERIND` variables among the male population aged 18 to 50 in the 1911 census. This estimate should be interpreted as an upper bound, since it does not account for occupation-specific age and marital-status criteria for protection from military service. This list is held at the National Archives under record number `MH 47/142/1`.

agriculture and 38 percent in transport enlisted, compared with 63 percent in commerce, where protection was much weaker (Dewey, 1984, pp. 216–19).

The application of these rules was administered locally by military service tribunals. Established in each district in 1916, these tribunals adjudicated individual claims for exemption from conscription (McDermott, 2011, pp. 11–35). Although the government had issued thorough instructions on exemption procedures, tribunals retained considerable discretion in granting exemptions. By April 1917, 1.8 million men had been covered by the protected-occupation scheme, while a further 780,000 had been granted exemptions on other grounds (War Office, 1921, p. 10). Taken together, men protected or exempted from immediate call-up accounted for more than one-quarter of the population potentially liable for military service.

Wartime mortality The second component of military death rates was wartime mortality among enlisted men. The determinants of enlistment discussed above likely also shaped this margin, since they affected not only whether men served, but also when and where they served. Indeed, using a random sample of 2,406 British servicemen born in the 1890s, Bailey, Hatton and Inwood (2023) show that the single most important predictor of death during the war was the date of enlistment: the earlier a man enlisted, the longer his exposure to wartime service, and hence the higher his probability of being killed. They also show that infantrymen faced markedly higher probabilities of death than men serving in other branches—medical fitness was an essential criterion for eligibility for front-line service (Winter, 1985, pp. 48–64; McDermott, 2011, pp. 180–97).

Beyond these factors, the territorial organization of military recruitment constitutes another potential channel through which districts may have differed in wartime mortality. Following the Cardwell (1868–74) and Childers (1881) reforms, infantry regiments were tied to regimental areas typically formed along county lines and drew a large share of their recruits locally (Spiers, 1996, pp. 188–89). The formation of Pals Battalions in 1914 further reinforced this territorial logic of recruitment, as men were encouraged to enlist and form battalions with those from the same towns and districts (Winter, 1985, pp. 30–33; Simkins, 1996, pp. 240–42). As a result, battlefield losses could translate into spatially concentrated mortality when a locally recruited unit experienced heavy casualties. Nevertheless, the territorial link between recruitment and units weakened over the course of the war as casualty replacements and transfers progressively diluted the local composition of battalions (Simkins, 1996, pp. 251–52; Hine, 2018).[31]

[31] For instance, Hine (2018, pp. 54, 287) finds that 82 percent of the 229 Cheshire Regiment 1st Battalion soldiers who died between August 4 and November 30, 1914 had enlisted in the Cheshire regimental area, compared

Control variables To alleviate concerns that military death rates and changes in female labor force participation may both be driven by prewar local labor-market conditions, we capture the systematic part of the variation in military death rates across districts using a parsimonious specification that relates district-level military death rates to the determinants discussed above. We measure the extent to which the system of protected occupations limited military death rates—by reducing enlistment rates—through the share of men aged 18 to 50 in the 1911 census who were employed in occupations covered by the 1916 list of certified occupations.[32] Next, we account for the local demographic structure using three district-level measures: mean age, the share unmarried, and the share with children under 16 in their household—all calculated among men aged 18 to 50 in the 1911 census. Finally, we proxy for baseline health conditions using the district's infant mortality rate.[33]

We further account for district characteristics that may shape our measure of military death rates. Because incomplete coverage of the numerator may be related to rurality and internal migration patterns (see Appendix G), we include an indicator for whether the district was rural, the share of the male population born outside their county of residence, and the logarithm of the district population. We also include a coastal indicator to account for the omission of deceased Royal Navy personnel from our baseline measure. Finally, we include a Wales indicator to capture the systematically lower military death rates in Welsh districts (see Footnote 29).[34]

We report the results of this analysis in Table 2, where we standardize independent variables—except for indicator variables—to facilitate comparison of coefficient magnitudes across specifications.[35] Focusing on Column 5, we find that districts with a higher share of men in protected occupations had lower military death rates, though the magnitude of this correlation remains modest.[36] By contrast, demographic factors played a limited role once other factors are accounted for. Military death rates were also relatively lower in more populated, coastal, and Welsh districts, as well as in districts with a larger

with only 39 percent of the 241 soldiers who died between July 1 and November 30, 1918.

[32] We focus on the age group 18–50 because this corresponds to the widest age range defining the male population potentially liable for military service over the course of the war (see Section 3.2). This implicitly assumes that the occupational structure observed in 1911 would have remained broadly similar in 1914–18 absent the war.

[33] Our district-level measure of infant mortality is the number of deaths among infants under 1 per million births in 1911. It corresponds to Southall and Mooney's (2025) `i_rate` variable divided by 1,000.

[34] Appendix Table A12 provides summary statistics for these 1911 district characteristics.

[35] We report unstandardized coefficients in Appendix Table A13.

[36] The coefficient of −0.013 implies that a 1 standard deviation increase in the share of men in protected occupations was associated with a 1.3 percentage point lower military death rate, compared with a cross-district standard deviation in military death rates of 4.9 percentage points.

Table 2. The Prewar Determinants of Military Death Rates

Outcome:	Military death rate						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Share in protected occupations	−0.011***				−0.013***	−0.012***	−0.013***
	[0.001]				[0.002]	[0.002]	[0.002]
Mean age		0.007***			0.003**	0.003**	0.003**
		[0.001]			[0.002]	[0.002]	[0.002]
Share unmarried		−0.017***			−0.005	−0.002	−0.006
		[0.003]			[0.004]	[0.005]	[0.005]
Share with dependent children		−0.014***			0.001	0.004	−0.001
		[0.003]			[0.004]	[0.005]	[0.005]
Infant mortality rate			−0.005***		−0.002	−0.002	−0.001
			[0.001]		[0.001]	[0.002]	[0.002]
Rural district				−0.011***	−0.001	−0.001	−0.003
				[0.002]	[0.003]	[0.003]	[0.003]
Share born outside county				0.000	−0.004**	−0.003**	−0.006***
				[0.001]	[0.001]	[0.001]	[0.002]
Log population				−0.011***	−0.010***	−0.010***	−0.011***
				[0.002]	[0.002]	[0.002]	[0.002]
Coastal district				−0.013***	−0.017***	−0.016***	−0.014***
				[0.002]	[0.002]	[0.002]	[0.003]
Welsh district				−0.028***	−0.017***	−0.018***	0.019
				[0.003]	[0.003]	[0.003]	[0.020]
FLFP						0.002	0.002
						[0.001]	[0.002]
County FE	No	No	No	No	No	No	Yes
Observations	1,721	1,721	1,721	1,721	1,721	1,721	1,721
R^2	0.049	0.057	0.012	0.101	0.172	0.173	0.253

Notes. This table reports coefficient estimates from regressing the military death rate on 1911 local government district characteristics. The unit of observation is a consistent 1911–21 district. See the main text for variable definitions. Shares are in percent. Independent variables—except for indicator variables—are standardized. Robust standard errors are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

share of men born outside their county of residence. Altogether, these factors explain 17 percent of the cross-district variation in military death rates. Importantly, including prewar female labor force participation rates in Column 6 leaves the explanatory power of the specification virtually unchanged. The coefficient on this regressor is also close to zero and statistically insignificant, further alleviating concerns that military death rates are correlated with prewar local labor-market conditions that also shape subsequent changes in female labor force participation. Including county fixed effects in Column 7 likewise leaves the coefficients on the main regressors broadly unchanged, suggesting that the observed

correlations are not driven by cross-county variation.[37] We therefore use the regressors included in Column 6 as control variables in our baseline difference-in-differences specification described below and rely on the remaining 83 percent of the raw cross-district variation in military death rates for identification.[38]

4. The Impact of World War I on Female Labor Force Participation

4.1. Empirical Strategy

To assess the impact of wartime mortality on women's labor, we use a difference-in-differences strategy that compares changes in female labor force participation between 1911 and 1921 across districts differentially exposed to military death rates. We estimate the following specification on the full-count sample of women aged 20 to 54:

$$Y_{iadt} = \beta \, Death\ rate_d \times Post_t + \boldsymbol{\gamma}' \left(\mathbf{X}_d \times Post_t\right) + \alpha_a + \mu_d + \lambda_t + \varepsilon_{iadt} \,, \tag{1}$$

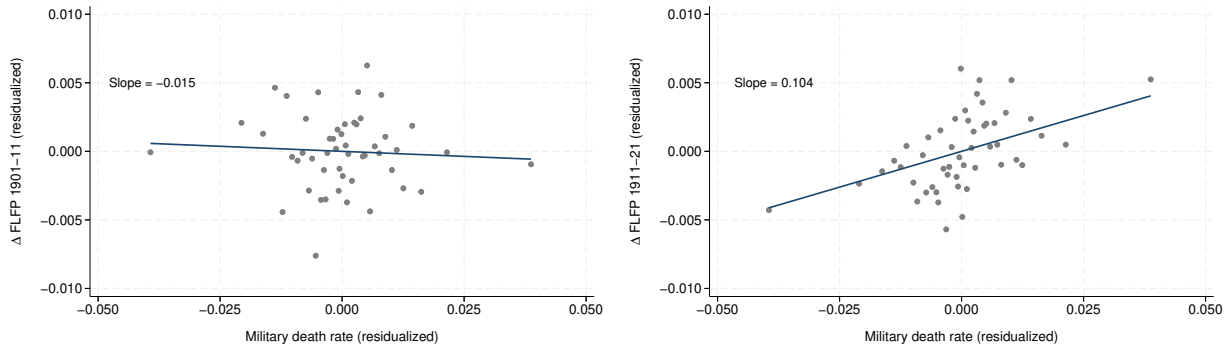
where Y_{iadt} is an indicator equal to one if woman i of age a residing in district d is classified as participating in the labor force in year t, and zero otherwise. The variable $Death\ rate_d$ denotes the military death rate in district d, while $Post_t$ is an indicator equal to one in 1921 and zero in 1911. The vector $\mathbf{X}_d$ contains the 11 prewar district-level controls included in Column 6 of Table 2, each interacted with $Post_t$. Age fixed effects α_a account for differences in labor force participation across ages, district fixed effects μ_d absorb time-invariant differences across districts, and year fixed effects λ_t capture aggregate changes in female labor force participation common to all districts. The coefficient of interest β identifies the differential change in female labor force participation between 1911 and 1921 associated with greater district-level military death rates. To account for the within-district correlation in the error term ε_{iadt}, we cluster standard errors at the district level.

This identification strategy relies on the parallel-trends assumption: absent differential exposure to wartime mortality, districts would have experienced similar changes in female labor force participation. To support the validity of this assumption, we examine whether

[37] Local government districts are nested within 55 administrative counties. We show in Appendix Table A14 that the results are robust to controlling for additional characteristics that may be correlated with military death rates, including the prewar share of men employed in military occupations and indicators for the presence of a war factory or an infantry depot in the district. As an alternative to our theory-guided selection of controls, we also use LASSO to select predictors of military death rates from a set of 640 prewar district characteristics. This procedure selects 14 variables. Our results are robust to using this data-driven approach (Appendix Table A15).

[38] Appendix Figure A5 plots the distributions of raw and residualized military death rates. The residualization leaves the dispersion in military death rates across districts largely unchanged: the standard deviation declines only from 0.05 in the raw data to 0.04 after residualization.

prewar changes in female labor force participation between 1901 and 1911 are predicted by subsequent military death rates. Specifically, we estimate Equation 1 on the full-count sample of women aged 20 to 54 in 1901 and 1911, setting $Post_t$ to one in 1911 and zero in 1901.[39] In this placebo specification, the coefficient on $Death\ rate_d \times Post_t$ captures whether districts that later had greater military death rates were already experiencing differential changes in female labor force participation in the decade preceding the war. We find no evidence of differential prewar trends: the estimated coefficient is close to zero at −0.015 and statistically insignificant (standard error of 0.026), supporting the credibility of the parallel-trends assumption.[40] Panel a of Figure 2 provides a graphical counterpart to this placebo test.



(a) 1901–11 (b) 1911–21

Figure 2. Military Death Rates and Changes in Female Labor Force Participation

Notes. These figures plot the relationship between military death rates and changes in female labor force participation (Δ FLFP) among women aged 20 to 54 between 1901 and 1911 in Panel a and between 1911 and 1921 in Panel b. Both variables are residualized with respect to local government district and age fixed effects, a post-period indicator, and the 11 prewar district-level controls included in Column 6 of Table 2, each interacted with a post-period indicator. The figures group observations into 50 equal-sized bins based on residualized military death rates and plot mean residualized female labor force participation and military death rates within each bin. The solid lines represent OLS fits estimated using the full individual-level sample rather than the binned observations.

[39] We also draw on I-CeM's data for the 1901 census (Schürer and Higgs, 2025). This placebo should be interpreted with caution because the 1901 census data derive from census enumerators' books rather than from household schedules—female labor force information in census enumerators' books has been shown to vary markedly across enumeration districts due to inconsistencies in enumeration practices (You, 2020*b*). Moreover, the methodology for classifying occupations in the 1901 census differed markedly from that used in the 1911 and 1921 censuses because of changes in schedule format and tabulation procedures—see table 63 (*Differences in Classification of Occupations as between 1911 and 1901 and as between 1901 and 1891*) in the *Census of England and Wales, 1911. Summary Tables* (London, 1915), pp. 246–61.

[40] Without controls, the estimated coefficient is −0.052 and marginally statistically significant at the 10 percent level (standard error of 0.027).

4.2. Results

Before turning to the results of the estimation, Panel b of Figure 2 provides a visual representation of the identifying variation by plotting residualized military death rates against changes in female labor force participation between 1911 and 1921. The relationship is clearly positive: the fitted line has a slope of 0.104, which corresponds to the baseline difference-in-differences estimate.

4.2.1. Baseline Estimate

We report the coefficient from estimating Equation 1 in Column 1 of Table 3. Exposure to wartime mortality increased women's labor force participation: the coefficient of 0.104 implies that in districts where the military death rate was 5 percentage points higher—one standard deviation of the treatment variable at the district level—women were 0.5 percentage points more likely to participate in the labor force after the war. This effect is positive and statistically significant at the 1 percent level but modest in magnitude, as it represents 1.5 percent of the prewar mean female labor force participation rate (34.5 percent).

Table 3. The Impact of Military Death Rates on Female Labor Force Participation

Outcome:	Female labor force participation					MLFP
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate × Post	0.104***	0.071***	0.067***		0.054**	−0.001
	[0.024]	[0.025]	[0.024]		[0.023]	[0.037]
Adj. death rate × Post				0.141***		
				[0.035]		
Controls	Yes	No	Yes	Yes	Yes	Yes
County × Post	No	No	Yes	No	No	No
Observations	18,013,998	18,013,998	18,013,998	18,013,998	3,442	16,283,440
Districts	1,721	1,721	1,721	1,721	1,721	1,721
Within R^2	0.108	0.107	0.107	0.108	0.603	0.011
1911 outcome mean	0.345	0.345	0.345	0.345	0.317	0.963
Treatment s.d.	0.028	0.028	0.028	0.023	0.049	0.028

Notes. This table reports coefficient estimates from Equation 1. Columns 1–4 use the sample of women aged 20 to 54, while Column 6 uses the corresponding sample of men. These specifications include age and local government district fixed effects. In Column 5, the unit of observation is a consistent 1911–21 district, so age fixed effects are not included. *Controls* comprise the 11 prewar district-level characteristics included in Column 6 of Table 2, each interacted with the *Post* indicator. *Adj. death rate* is the military death rate adjusted for measurement error, as described in Appendix I. *MLFP* stands for male labor force participation. Standard errors are clustered at the district level and reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

4.2.2. Robustness Tests

Next, we conduct a series of robustness tests that support the validity of the baseline estimate. The estimated effect remains similar—albeit slightly smaller—when we omit the controls (Column 2), suggesting that observable prewar district characteristics do not drive the relationship between military death rates and changes in female labor force participation.[41] Including county-by-year fixed effects—equivalent to county-specific time trends in our two-period setting—and thus relying only on within-county variation for identification also yields a similar estimate (Column 3).

A potential concern is that measurement error in military death rates may attenuate the estimated treatment effect. Reassuringly, correcting for the sampling component of this measurement error using regression calibration yields a similar effect when expressed in standard-deviation units (Column 4, see Appendix I for details). The results are also robust to aggregating the data at the district level (Column 5). Finally, estimating Equation 1 on the sample of men instead of women results in a null effect.

We conduct several additional robustness tests. We first verify that the estimated effect is not driven by particular types of districts: excluding Welsh, rural, or coastal districts produces estimates similar to the baseline, as does excluding districts in the bottom and top deciles of the distribution of military death rates (Appendix Table A17). Next, we show that the results are not sensitive to alternative measures of female labor force participation: using I-CeM's original `INACTIVE` variable rather than our corrected measure yields similar estimates, as does focusing on employment rather than labor force participation (Appendix Table A18). Including women of all marital statuses or using broader age ranges also leaves the results essentially unchanged (Appendix Table A19). By contrast, using alternative measures of military death rates yields markedly different results: although all coefficients remain positive and statistically significant at conventional levels, most are substantially smaller (Appendix Table A20), a pattern likely reflecting greater measurement error (see Appendix G).

4.3. Mechanisms

We now turn to the mechanisms that may explain the small but positive effect of military death rates on female labor force participation. We first examine the roles of labor demand (Section 4.3.1) and labor supply (Section 4.3.2), before considering migration as a

[41] The baseline estimate is also virtually unchanged when age fixed effects are omitted, with a coefficient of 0.102 (standard error of 0.025). We also obtain similar results when reproducing the specifications in Table 3 using the 14 prewar district-level control variables selected through the data-driven LASSO procedure described in Footnote 37 (Appendix Table A16).

factor that may have mitigated these effects (Section 4.3.3).

4.3.1. Labor Demand

The war reduced the stock of working-age men, potentially increasing firms' demand for women in occupations previously held by those lost to the war. By this logic, the increase in female labor force participation should be disproportionately concentrated in previously male-dominated occupations, which were most directly affected by male scarcity, as well as in gender-balanced occupations, where male and female workers were likely close substitutes. We test this mechanism by first calculating the female share of workers in each of the 326 HISCO occupational categories observed in the 1911 and 1921 censuses. We then classify occupations into three mutually exclusive groups based on their 1911 composition: those in which more than 75 percent of workers were women ("female-dominated occupations"), those in which women accounted for 25–75 percent of workers ("gender-balanced occupations"), and those in which fewer than 25 percent of workers were women ("male-dominated occupations"). We then estimate Equation 1 using as the outcome an indicator equal to one if a woman participated in the labor force in a given occupational group, and zero otherwise.

Table 4. Military Death Rates and Occupational Substitution Among Women

Sample:	A. All women			B. Active women		
Outcome:	Female-dom. occ.	Gender-bal. occ.	Male-dom. occ.	Female-dom. occ.	Gender-bal. occ.	Male-dom. occ.
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate × Post	0.007	0.063***	0.034**	−0.138**	0.072	0.066
	[0.027]	[0.020]	[0.015]	[0.066]	[0.052]	[0.043]
Observations	18,013,998	18,013,998	18,013,998	6,035,360	6,035,360	6,035,360
Districts	1,721	1,721	1,721	1,721	1,721	1,721
Within R^2	0.030	0.037	0.028	0.022	0.003	0.022
1911 outcome mean	0.172	0.139	0.034	0.500	0.402	0.099
Treatment s.d.	0.028	0.028	0.028	0.028	0.028	0.028

Notes. This table reports coefficient estimates from Equation 1 using indicators for employment in different occupational groups as outcomes. Columns 1–3 use the sample of all women aged 20 to 54, while Columns 4–6 restrict the sample to women active in the labor force. Occupations are classified according to the female share of workers in each HISCO category in the 1911 census. *Female-dominated occupations* are those in which more than 75 percent of workers were women. *Gender-balanced occupations* are those in which women accounted for 25–75 percent of workers. *Male-dominated occupations* are those in which fewer than 25 percent of workers were women. In Columns 1–3, the outcome equals one if a woman was active in the relevant occupational group and zero otherwise. In Columns 4–6, the outcomes indicate the occupational group of active women. All specifications include age and local government district fixed effects, together with the 11 prewar district-level characteristics included in Column 6 of Table 2, each interacted with the *Post* indicator. Standard errors are clustered at the district level and reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

We report the results in Table 4. Panel A uses the full sample, while Panel B restricts the

sample to labor force participants. Because the three occupational groups are mutually exclusive, the coefficients in Panel A sum to the baseline estimate. Taken together, the estimates support labor demand as an important mechanism: gender-balanced occupations account for 61 percent of the baseline effect ($\simeq 0.063/0.104$) and male-dominated occupations for 32 percent ($\simeq 0.034/0.104$), while female-dominated occupations account for only 7 percent ($\simeq 0.007/0.104$), despite employing half of active women before the war. Focusing on labor force participants in Panel B, a 5 percentage point higher military death rate shifted the occupational composition of female employment by 0.7 percentage points away from female-dominated occupations.

4.3.2. Labor Supply

Wartime military mortality may also have increased female labor supply by shifting the marital composition toward single and widowed women, whose baseline participation rates were far higher than those of married women (see Table 1). Panel A of Table 5 supports this mechanism: a 5 percentage point higher military death rate raised postwar shares of single and widowed women by 0.2 and 0.1 percentage points, respectively, while reducing the married share by a corresponding 0.3 percentage points. Holding participation rates by marital status at their prewar levels, these compositional changes imply a 0.2 percentage point increase in female labor force participation or 38 percent of the baseline effect.[42]

Once changes in marital composition are accounted for, the remaining effect operates through changes in labor force participation rates within marital-status groups. Estimates in Panel B imply that a 5 percentage point higher military death rate raised labor force participation by 0.6 percentage points among single women and by 0.2 percentage points among married women—a sizable magnitude relative to married women's prewar labor force participation rate of 9.3 percent. By contrast, the estimated effect among widows is close to zero and statistically insignificant—likely due to the relatively high pensions received by war widows (see Section 5.4).

To decompose the overall effect into contributions from single, married, and widowed women, Panel C uses indicators for being both active and in a given marital-status group. Because the three marital groups are mutually exclusive, the coefficients in this panel sum to the baseline estimate. The results imply that single women accounted for 78 percent of the overall effect ($\simeq 0.081/0.104$), as higher military death rates both increased their share

[42] This share is obtained by comparing the implied increase in female labor force participation from changes in marital composition with the baseline effect of a 5 percentage point increase in the military death rate: $5 \times [(0.045 \times 0.727) - (0.069 \times 0.093) + (0.024 \times 0.558)] / (5 \times 0.104) \simeq 0.198/0.520 \simeq 0.382$.

Table 5. Military Death Rates, Marital Status, and Female Labor Force Participation

Marital status:	Single	Married	Widowed
	(1)	(2)	(3)
A. Marital status			
Death rate × Post	0.045*	−0.069***	0.024***
	[0.023]	[0.022]	[0.007]
Observations	18,013,998	18,013,998	18,013,998
1911 outcome mean	0.362	0.591	0.046
B. FLFP within marital-status groups			
Death rate × Post	0.119***	0.041**	0.012
	[0.036]	[0.020]	[0.069]
Observations	6,302,420	10,774,990	936,587
1911 outcome mean	0.727	0.093	0.558
C. Contribution to female labor force participation			
Death rate × Post	0.081***	0.012	0.011**
	[0.020]	[0.011]	[0.005]
Observations	18,013,998	18,013,998	18,013,998
1911 outcome mean	0.264	0.055	0.026

Notes. This table reports coefficient estimates from Equation 1. Panel A uses indicators for being single, married, or widowed as outcomes. In Panel B, the outcome is female labor force participation, and each column restricts the sample to the corresponding marital-status group. Panel C uses indicators equal to one if a woman was both active in the labor force and belonged to the corresponding marital-status group. All specifications use women aged 20 to 54 and include age and local government district fixed effects, together with the 11 prewar district-level characteristics included in Column 6 of Table 2, each interacted with the *Post* indicator. Standard errors are clustered at the district level and reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

in the female population and labor force participation rate. By contrast, married women accounted for just 12 percent ($\simeq 0.012/0.104$), as their increased labor force participation rate was largely offset by a corresponding decline in their share of the female population. Finally, widowed women accounted for the remaining 11 percent ($\simeq 0.011/0.104$), mostly because they continued to represent a small share of the female population despite the sharp increase in their numbers.

4.3.3. Migration and Spatial Adjustment

The mechanisms we identify may have been mitigated by the migration response to local military deaths. Stronger labor demand in districts with high military death rates may have attracted surviving men from less affected areas, reducing both the demand for women to fill vacancies and pressure on local marriage markets. On the supply side,

comparatively favorable marriage prospects in districts with low military death rates may have induced single and widowed women to migrate in the opposite direction, thereby alleviating marriage-market imbalances in the most affected areas. The local equilibrium response of female labor force participation to military deaths is therefore likely to have been weaker where migration reacted more strongly. Given the high level of internal mobility in England and Wales during this period (Pooley and Turnbull, 1998) and its responsiveness to spatial differences in both labor- and marriage-market conditions (Boyer and Hatton, 1997; Day, 2026), migration may have constituted an especially important mechanism of spatial adjustment.[43]

To assess whether local military deaths induced internal migration, we examine changes in district-level male shares of the population. In the absence of migration—and abstracting from other demographic changes—a marginal 1 percentage point increase in the military death rate would mechanically reduce the male share of the population by 0.25 percentage points, assuming a balanced prewar sex ratio.[44] Estimating Equation 1 at the district level with the male share of the population aged 20 to 54 as the dependent variable yields a coefficient of −0.046 significant at the 1 percent level (Appendix Table A21). The observed local sex-ratio response was therefore only one-fifth of the closed-district benchmark, consistent with internal migration mitigating local demographic imbalances. Sequentially adding population-weighted average military death rates for neighboring districts in successive distance bands, we find that military death rates in districts within 5 kilometers are associated with an additional decline in the focal district's male share. Such highly localized spillovers are consistent with nearby districts competing for a common pool of mobile surviving men. An implication of these local cross-district migrations is that effect sizes should increase as the analysis is aggregated to broader spatial units—moves between neighboring districts are then internalized, thereby reducing migration spillovers. Consistent with this prediction, the county-level estimate on the male share of the population is more than three times as large as the district-level estimate (Appendix Table A22).

[43] Comparing county of birth with county of residence, about half of men and women aged 20 to 54 resided outside their county of birth in both 1911 and 1921. Birthplace is not recorded at the district level, preventing us from constructing an analogous measure at that level.

[44] To see this, let s denote the prewar male share of the population and d the military death rate. Absent migration, the postwar male share would be $s(1-d)/(1-sd)$, with a marginal effect of the death rate of $-s(1-s)/(1-sd)^2$. Assuming a balanced prewar sex ratio ($s = 0.5$), a 1 percentage point increase in the military death rate would reduce the male share by $s(1-s) = 0.25$ percentage points.

5. England and Wales in Comparative Perspective: France as a Benchmark

We evaluate the magnitude of the responsiveness of female labor force participation to military death rates in England and Wales by using the French case as a benchmark. France likewise experienced a postwar increase in female labor force participation, but differed markedly along several demographic and institutional dimensions that may have shaped its labor market response to wartime mortality. This comparison enables us to identify the broader conditions under which wartime demographic shocks may translate into greater subsequent female labor force participation. We begin by reviewing the French case (Section 5.1), before outlining our methodology for the cross-country comparison (Section 5.2). We then show that female labor force participation was substantially more responsive to military death rates in France than in England and Wales (Section 5.3). Finally, we argue this larger response likely reflects cross-country differences along two dimensions: internal migration and pensions to war widows (Section 5.4).

5.1. The French Case

During World War I, France suffered an exceptionally high death toll relative to other belligerent countries. Under its system of universal conscription, most male citizens aged 20 to 48 were mobilized during the war. Of these, 1.3 million died, a military death rate of 13.0 percent. Using census data across 87 *départements*, Boehnke and Gay (2022) find that wartime mortality substantially increased women's labor force participation: in départements with military death rates 3.8 percentage points higher—one standard deviation of the treatment variable—women were 1.6 percentage points more likely to participate in the labor force in 1921, an effect size representing 5.2 percent of the prewar mean female labor force participation rate (31.4 percent).[45]

They further show that this effect was driven primarily by labor supply responses: low pensions pushed war widows into employment, while deteriorating marriage prospects led single women to delay marriage and work longer. By contrast, labor demand mechanisms were secondary, as firms responded to male scarcity mainly by increasing capital rather than substituting female for male labor.

5.2. Cross-Country Comparison Methodology

To make the estimated effects comparable across contexts, we harmonize the data (Section 5.2.1) and estimation strategy (Section 5.2.2) across settings, thereby limiting

[45]This estimate of 0.430 is obtained using the baseline specification in Boehnke and Gay (2022, p. 1221) and restricting the sample to 1911 and 1921.

differences attributable to measurement and research design rather than context (Slough and Tyson, 2025).

5.2.1. Harmonizing the Data

We first harmonize the data along three dimensions: the level of aggregation and the definitions of the outcome and treatment variables. This harmonization is nevertheless constrained by major differences in data availability: whereas individual-level census data are available for England and Wales, the French data are limited to département-level tabulations published in census reports.

Level of aggregation Given the local migration response to the demographic shock in England and Wales and the resulting cross-district spillovers, a meaningful comparison first requires harmonizing the level of aggregation of the analysis. Because of limitations of the French data, this entails aggregating the data for England and Wales to a spatial unit comparable to the French département. However, no close equivalent existed in the administrative hierarchy of England and Wales: the nearest unit was the administrative county, of which there were 55, compared with 87 départements. Because the two countries' populations were broadly similar in 1911—40 million in France and 36 million in England and Wales—counties represent a much coarser level of aggregation than départements. We therefore construct "synthetic English and Welsh départements" (hereafter, synthetic départements) by grouping contiguous districts into 87 units whose population distribution matches that of French départements. Using a stochastic spatial partitioning algorithm described in Appendix J, we generate 200 different mappings of synthetic départements so as to produce a distribution of estimated effects of military death rates on female labor force participation in England and Wales.

Female labor force participation We also harmonize the definition of the outcome variable. Here again, the comparison is constrained by the French data, which only report the number of labor force participants among women aged 15 and over at the département level. Accordingly, we define female labor force participation as the share of women aged 15 and over who are in the labor force, irrespective of marital status. In addition, to address issues in French statistical conventions for the classification of farmers' wives, we exclude women classified as farm owners from the active population (Boehnke and Gay, 2022, pp. 1215–16; Gay, 2025).[46] Under this harmonized definition, the average female

[46] Beginning in 1906, the French censuses systematically classified farmers' wives as economically active, artificially inflating the measured female labor force participation rate from 31.4 to 51.5 percent in 1911. This

labor force participation rate across (synthetic) départements declined from 33.9 percent in 1911 to 30.7 percent in 1921 in England and Wales, whereas it increased from 31.4 to 35.0 percent in France.[47]

Military death rates Finally, we harmonize the definition of the treatment variable, beginning with its numerator—the number of military fatalities. For England and Wales, we use the enlistment-place adjusted number of military fatalities. In France, military records locate servicemen only by place of birth. Because internal migration was relatively limited, assigning military fatalities by département of birth rather than residence is unlikely to introduce serious bias (Boehnke and Gay, 2022, p. 1225).[48] Nonetheless, to address this concern, we approximate a residence-based military death rate by adjusting the spatial allocation of military fatalities using bilateral migration patterns among men across départements in 1911.[49] In addition, because records for England and Wales provide only partial coverage of military fatalities—while French records cover the universe of fatalities—we correct the corresponding military death rate for potential measurement error using regression calibration (see Appendix I).

We further harmonize the denominator of the military death rate. Whereas men potentially liable for military service in England and Wales were aged 11 to 43 in 1911, the corresponding age range in France was 13 to 44 (Boehnke and Gay, 2022, p. 1216). To ensure comparability across settings, we thus use the number of men aged 11 to 44 in 1911 as the denominator. Under this harmonized definition, the average military death rate across (synthetic) départements was 6.4 percent in England and Wales and 12.8 percent in France, with corresponding standard deviations of 1.5 and 1.7 percentage points.

5.2.2. Harmonizing the Estimation Strategy

To compare the effect of military death rates on female labor force participation between 1911 and 1921 across the two settings, we estimate the following specification using 87 (synthetic) départements:

$$Y_{dt} = \beta\ Death\ rate_d \times Post_t + \boldsymbol{\gamma}'\left(\mathbf{X}_d \times Post_t\right) + \mu_d + \lambda_t + \varepsilon_{dt}\ , \tag{2}$$

adjustment has essentially no effect in England and Wales, where only 173,839 women aged 15 and over were classified in farming occupations in the 1911 and 1921 censuses combined.

[47] For England and Wales, the reported mean female labor force participation rate is the median of the means obtained across the 200 synthetic mappings.

[48] In 1911, only 19 percent of French men aged 15 to 44 resided outside their département of birth, compared to about half in England and Wales (Boehnke and Gay, 2022, p. 1216).

[49] The methodology underlying this correction is described in Boehnke and Gay (2022, Appendix E).

where Y_{dt} denotes the female labor force participation rate in (synthetic) département d in year t. The remaining variables and parameters are defined analogously to those in Equation 1. The vector $\mathbf{X}_d$ contains prewar controls that capture the systematic determinants of military death rates that may be correlated with subsequent changes in female labor force participation in each setting. Because the determinants of wartime mortality differed markedly across the two countries, we allow the set of controls to reflect the relevant data generating process in each case rather than impose a common specification. For England and Wales, $\mathbf{X}_d$ comprises the 11 prewar district-level controls included in Column 6 of Table 2, aggregated to the synthetic-département level. For France, it comprises the two controls defined in Boehnke and Gay (2022, pp. 1217–19), i.e., the rural share of the population and the share residing in their département of birth. The resulting sets of control variables therefore differ by design. We cluster standard errors at the (synthetic) département level. In addition, to make the results more directly comparable to the individual-level estimates reported in Section 4, we weight regressions by each (synthetic) département's 1911 population. Finally, to further provide benchmarks for the synthetic-département estimates for England and Wales, we also estimate Equation 2 at the district and county levels, where we aggregate control variables to the corresponding spatial unit and redefine unit fixed effects and clustered standard errors accordingly.

5.3. Comparative Results

Figure 3 reports the coefficients from estimating Equation 2. They imply that a 1 percentage point higher military death rate across French départements was associated with a 1.6 percentage point increase in female labor force participation between 1911 and 1921. By contrast, the responsiveness of female labor force participation to military death rates was less than one-tenth as large at 0.14 percentage points in England and Wales at the synthetic-département level.[50] This gap is even starker when comparing magnitudes: a 1 standard deviation higher military death rate was associated with a 2.9 percentage point increase in female labor force participation in France compared with 0.16 percentage points in England and Wales, or 8.0 versus 0.4 percent of their respective prewar means (Table 6). Overall, the response in France was thus about one order of magnitude larger than in England and Wales, pointing to substantial context-dependent treatment-effect heterogeneity (Egami and Hartman, 2023).[51]

[50] Consistent with the argument in Section 4.3.3, the estimated effect in England and Wales increases with the spatial scale of aggregation—from the local government district at 0.127 to the administrative county at 0.145—as migration-induced spillovers across neighboring districts are progressively internalized.

[51] These results are robust to estimating Equation 2 without controls (Appendix Figure A6) or without population weights (Appendix Figure A7).

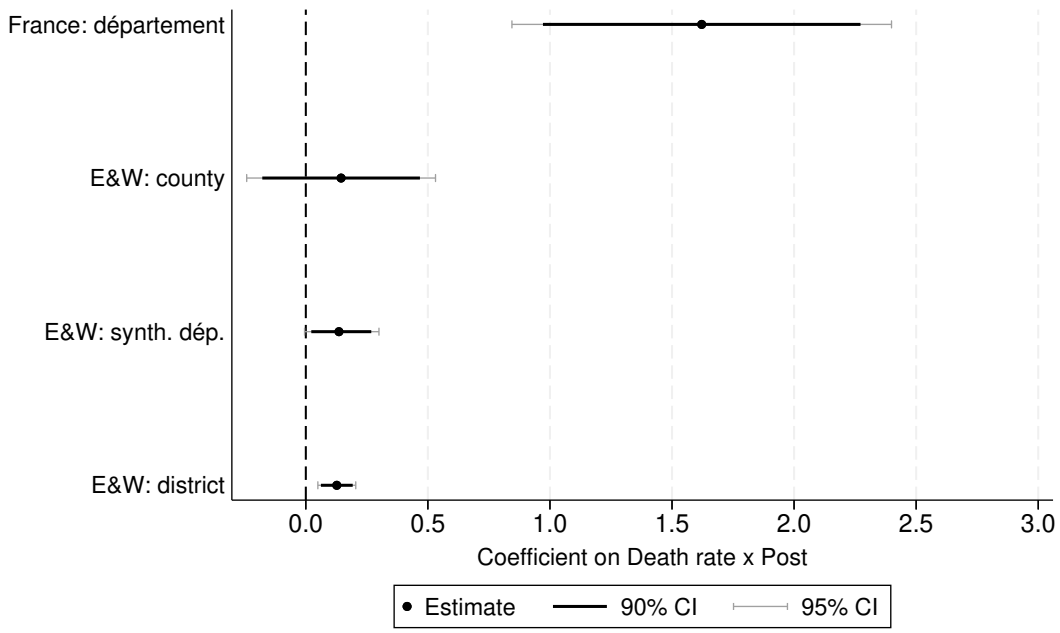


Figure 3. Military Death Rates and Female Labor Force Participation
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on female labor force participation between 1911 and 1921. For England and Wales, estimates are reported at the local government district, synthetic-département, and administrative-county levels. For France, the estimate is reported at the département level. For all but the synthetic-département specifications, bars report 90 and 95 percent confidence intervals based on standard errors clustered at the corresponding spatial level. For the synthetic-département specification, the point estimate is the median coefficient across 200 alternative synthetic mappings and bars report the 5th and 95th percentiles (and 2.5th and 97.5th percentiles) of the resulting coefficient distribution.

Table 6. Magnitudes of Estimated Effects
England and Wales versus France

		Parameters			Magnitudes	
Country	Unit	$\widehat{\beta}$	σ_D	$\overline{Y}$	$\widehat{\beta}\sigma_D$	$\widehat{\beta}\sigma_D/\overline{Y}$
France	Département	1.622	0.018	0.357	2.86	8.02
E&W	County	0.145	0.010	0.361	0.15	0.42
	Synth. dép.	0.136	0.012	0.361	0.16	0.44
	District	0.127	0.021	0.361	0.27	0.74

Notes. This table reports magnitudes of estimated effects of military death rates on female labor force participation across settings and levels of aggregation. The first column reports the estimated coefficient. The second column reports the population-weighted treatment standard deviation. The third column reports the population-weighted mean female labor force participation in 1911. The fourth column reports the effect of a one standard deviation increase in the military death rate in percentage points. The final column divides this effect by mean female labor force participation in 1911 and reports the result in percent.

5.4. Explaining the Anglo-French Gap

We now examine why female labor force participation was far more responsive to military death rates in France than in England and Wales. We focus on two factors that may condition how demographic shocks translate into changes in female labor force participation: internal migration, which shaped the persistence of local demographic imbalances (Section 5.4.1), and pensions for war widows, which conditioned their labor supply response (Section 5.4.2). We then consider several alternative explanations for this difference (Section 5.4.3).

5.4.1. Internal Migration

Internal migration was much more prevalent in England and Wales than in France before the war. Thus, even absent any differential postwar migration response to military deaths, the continuation of prewar migration patterns alone would have dissipated war-induced local demographic imbalances more rapidly in England and Wales than in France.

Yet we find suggestive evidence of a differential postwar migration response between the two countries. Estimating Equation 2 with the share of male residents born outside their county or département of residence as the outcome, we find that, in England and Wales, a 1 percentage point higher county-level military death rate was associated with a 0.58 percentage point increase in this share, while the corresponding estimate for France is close to zero (Appendix Figure A8). This pattern is consistent with internal migration more strongly attenuating local demographic imbalances—and the associated labor- and marriage-market pressures—in England and Wales than in France. This differential migration response is reflected in the greater persistence of local demographic imbalances in France: estimating Equation 2 with the male share of the population aged 20 to 49 as the outcome, we find that a 1 percentage point higher military death rate was associated with a 0.29 percentage point decline in this share across French départements, compared with a 0.16 percentage point decline across synthetic départements in England and Wales (Appendix Figure A9).

5.4.2. Pensions for War Widows

Boehnke and Gay (2022, p. 1232) identify war widows as a central mechanism behind the French response: military deaths not only sharply increased the number of widows, but also induced many of them to enter the labor force to compensate for the loss of their husbands' income. Indeed, French pensions for war widows remained relatively low throughout the 1920s, providing little compensation for this negative income shock.

Boehnke and Gay (2022) accordingly find that widows accounted for nearly half of the overall effect of military death rates on female labor force participation in France.

We obtain a similar result using our comparative specification: estimating Equation 2 with the share of women who were both widowed and active as the outcome, we find that a 1 percentage point increase in the military death rate was associated with a 0.37 percentage point increase in this share in France (Appendix Figure A10). This contribution represents 43 percent of the overall effect of military death rates on female labor force participation.[52] By contrast, this channel was largely absent in England and Wales as the corresponding estimate is close to zero at the synthetic-département level.

Part of this difference reflects the stronger effect of military death rates on widowhood in France: a 1 percentage point increase in the military death rate raised the widow share by 0.20 percentage points in France, compared with only 0.03 percentage points at the synthetic-département level in England and Wales (Appendix Figure A11). This gap may reflect greater spatial adjustment through internal migration in England and Wales, which attenuated local marriage-market imbalances. It may also reflect differences in the marital composition of military recruitment: whereas married men were initially exempt from conscription in England and Wales, French conscription was largely universal.

Quantitatively, however, the larger difference appears in the labor force participation response among widows themselves. Estimating Equation 2 with female labor force participation among widows as the outcome, we find that a 1 percentage point higher military death rate was associated with a 1.6 percentage point increase in labor force participation among widows in France (Appendix Figure A12). By contrast, the corresponding response in England and Wales was slightly negative at the synthetic-département level. Differences in state support provide one plausible explanation for this heterogeneity: French pensions for war widows amounted to only one-quarter of the earnings of a full-time female textile worker in 1921, whereas British pensions amounted to at least 60 percent of comparable earnings and provided more generous supplements for children (see Appendix K). The markedly different contribution of widows to the overall female labor force participation response therefore appears to stem primarily from a much stronger labor supply response among French widows, with the larger increase in widowhood playing a secondary role.

[52] While we find an overall coefficient of 1.622 in the specification reported in Table 6, the analysis of widow participation relies on comparing 1901 to 1926 due to data availability constraints. Re-estimating the effect on female labor force participation over these census years yields a coefficient of 0.864 (standard error of 0.276), implying that active widows accounted for 43 percent of this total effect ($\simeq 0.368/0.864$).

5.4.3. Alternative Explanations

We finally consider two alternative explanations: differences in the postwar marital-status composition of women and in the sectoral composition of female employment. Neither appears sufficient to account for the much larger effect of military death rates on female labor force participation in France.

First, because single and widowed women participated in the labor force at far higher rates than married women, military deaths could have mechanically increased female labor force participation by shifting the postwar marital-status composition of women away from marriage and toward singlehood and widowhood even absent any change in participation within marital-status groups. Differences in the magnitude of this compositional shift between France and England and Wales could therefore potentially account for part of the cross-country gap. Estimating Equation 2 with the shares of married, single, and widowed women as outcomes, we find that higher military death rates reduced the share of married women less in England and Wales than in France, with estimated effects of −0.27 and −0.31, respectively—a difference of 13 percent (Appendix Figure A13). Likewise, higher military death rates increased the share of single women less in England and Wales than in France, with estimated effects of 0.11 and 0.27, respectively—a difference of 59 percent (Appendix Figure A14). Finally, over the 1911–21 period, the widowhood response was also smaller in England and Wales than in France, with estimated effects of 0.03 and 0.10, respectively—a difference of 70 percent.[53] While these compositional responses were smaller in England and Wales than in France, they remain too small to account for the 92 percent gap in the overall female labor force participation response to military death rates. Moreover, the smaller compositional response in England and Wales may itself partly reflect the stronger migration response documented above, as greater male in-migration into high-death areas would have helped offset wartime imbalances in local marriage markets.

A second potential explanation is the stark difference in the sectoral composition of female employment across the two countries. In 1911, more than half of French working women were employed in agriculture, compared with less than 2 percent in England and Wales. If female labor in agriculture responded particularly strongly to military death rates, this difference in sectoral composition could account for part of the cross-country gap. The evidence, however, provides little support for this explanation, as redefining the outcome as participation in non-agricultural employment leaves the estimates for both settings essentially unchanged.

[53] Appendix Figure A11 reports an estimate of 0.20 for France using 1901 and 1926, the census years used for the analysis of widows' labor force participation in Section 5.4.2.

6. Conclusion

World War I produced a massive but largely temporary expansion of female employment in England and Wales. Yet this aggregate reversal concealed a persistent local response to wartime mortality: districts that suffered greater military losses experienced larger postwar increases in female labor force participation, as women shifted away from traditionally female occupations and military deaths shifted the marital composition toward groups with higher baseline participation rates. The magnitude of this effect was nevertheless modest: a 5 percentage point higher military death rate increased female labor force participation by only 0.5 percentage points, or 1.5 percent of the prewar mean. Our comparative analysis with France helps explain why this response was so limited: internal migration adjusted more strongly to local military deaths in England and Wales, dissipating the resulting marriage- and labor-market imbalances, while relatively generous pensions for war widows dampened their labor supply response. Our findings thus suggest that the consequences of wartime demographic shocks for women's work depend not only on the scale of military losses, but also on the spatial and institutional mechanisms through which economies absorb them.

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