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

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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)

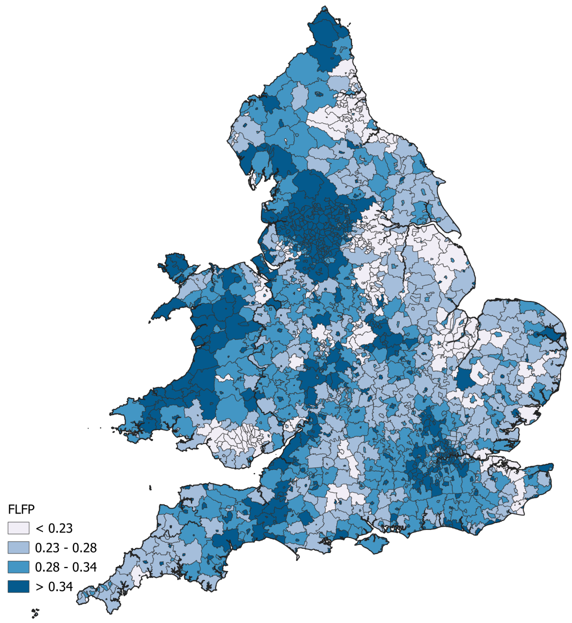
	A. Marital statuses (%)			B. Labor force participation rates (%)			
Marital status	Single	Married	Widowed	All	Single	Married	Widowed
1911	36.2	59.1	4.6	34.5	72.7	9.3	55.8
1921	33.8	60.4	5.7	32.6	72.8	9.2	42.2
Difference	−2.4	1.3	1.1	−1.8	0.1	−0.2	−13.6

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).

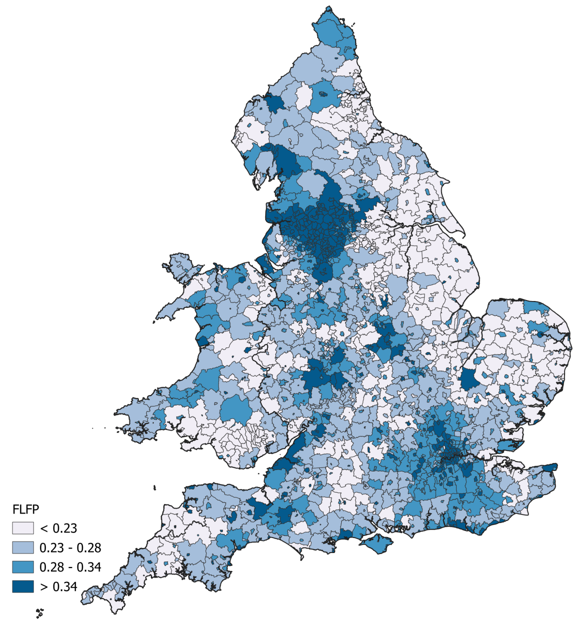
[9] To ensure comparability across marital-status groups, we restrict the sample to women who were single, married with a spouse present, or widowed based on I-CeM's `MAR` variable. Married women whose spouse was absent accounted for 5.2 percent of women aged 20 to 54 in 1911 and 5.6 percent in 1921. 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).

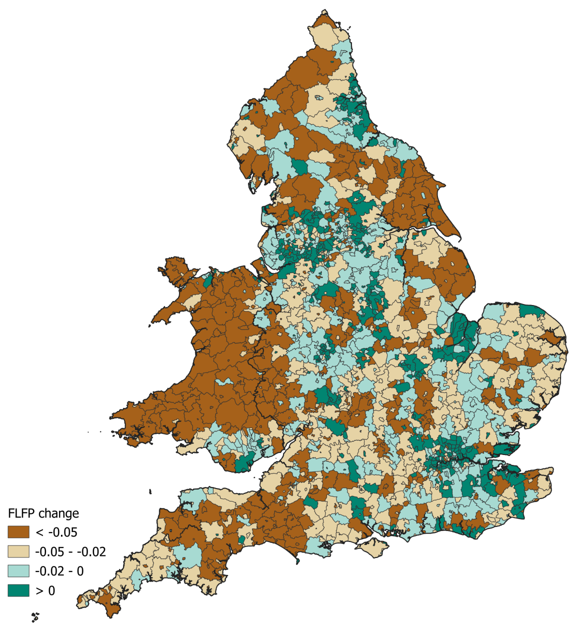
[11] Appendix Tables A1 and A2 provide data similar to Table 1 tabulated by age group. Complete summary statistics by census year both at the individual and district levels are provided in Appendix Table A3.



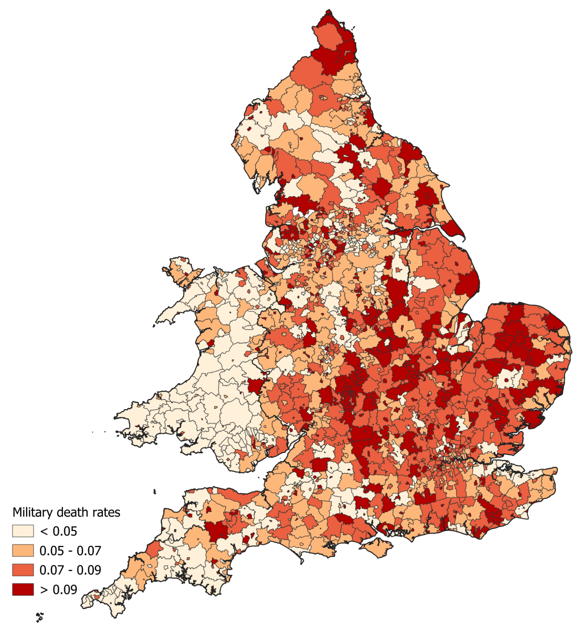
(a) FLFP (1911)



(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:

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

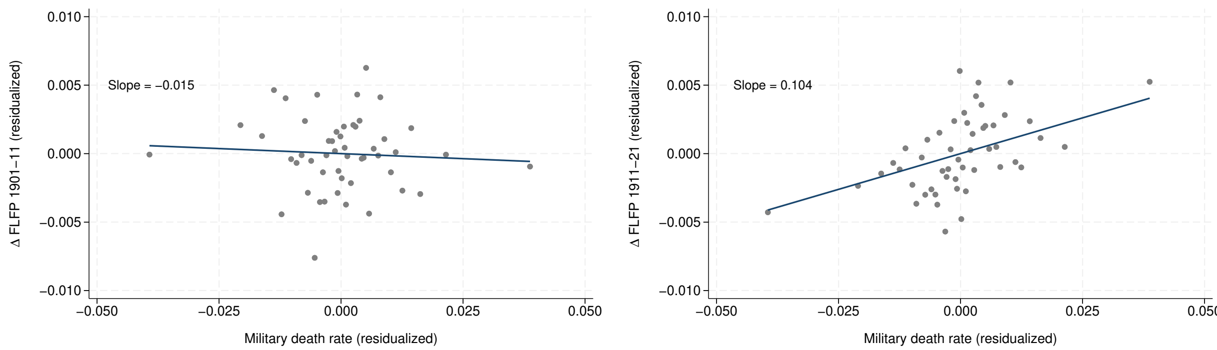
where Y_{iadt} is an indicator equal to one if woman i of age a residing in district d is classified as participating in the labor force in year t, and zero otherwise. The variable $\textit{Death rate}_d$ denotes the military death rate in district d, while $\textit{Post}_t$ is an indicator equal to 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 $\textit{Post}_t$. 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).

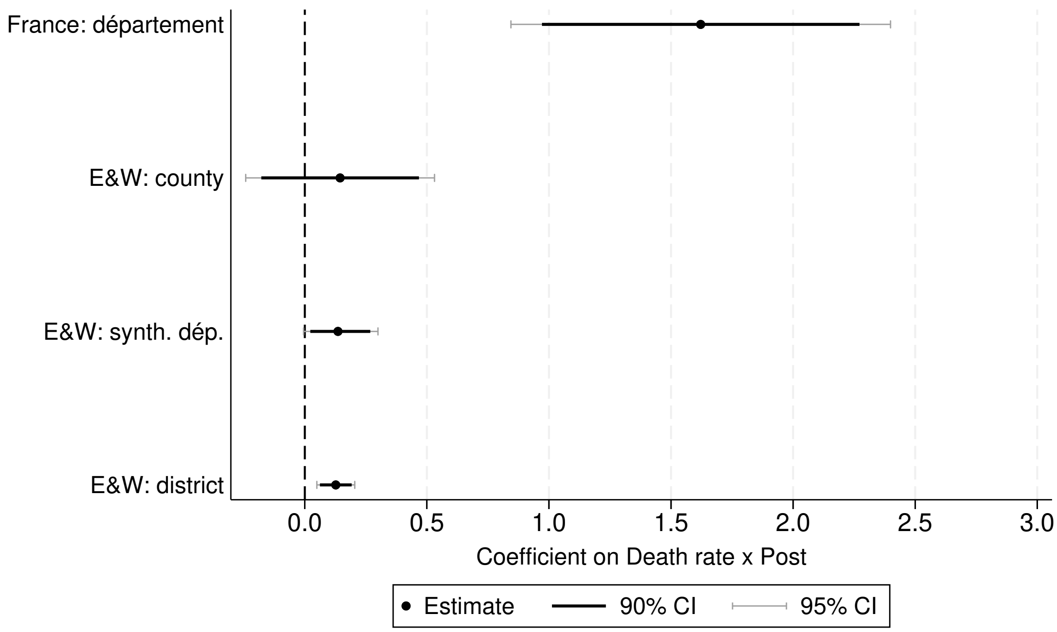


Figure 3. Military Death Rates and Female Labor Force Participation
England and Wales versus France

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

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September 2026

Supplementary Online Appendix

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A. Supplementary Tables and Figures

Table A1. Female Marital Statuses by Age Groups (1911–21)

Age group	20–54	20–29	30–39	40–49	50–54
	A. Single				
1911	36.2	61.8	25.5	18.3	15.9
1921	33.8	59.9	24.5	18.2	16.6
Difference	–2.4	–1.9	–1.0	–0.1	0.6
	B. Married				
1911	59.1	37.7	71.6	73.0	67.1
1921	60.4	38.8	70.3	73.4	68.7
Difference	1.3	1.1	–1.3	0.4	1.6
	C. Widowed				
1911	4.6	0.5	2.9	8.7	17.0
1921	5.7	1.2	5.2	8.5	14.8
Difference	1.1	0.7	2.3	–0.2	–2.3

Notes. This table provides the distribution of marital statuses in percent among women aged 20 to 54 based on the 1911 and 1921 census I-CeM household schedules (Schürer and Higgs, 2025; Schürer and Wakelam, 2025). *Married* only includes married women whose spouse is noted present in the household schedules.

Table A2. Female Labor Force Participation Rates by Age Groups (1911–21)

Age group	20–54	20–29	30–39	40–49	50–54
	A. All marital statuses				
1911	34.5	51.2	26.2	23.5	23.2
1921	32.6	51.2	25.8	21.6	20.5
Difference	–1.8	–0.1	–0.4	–1.9	–2.7
	B. Single				
1911	72.7	75.9	70.2	65.2	59.6
1921	72.8	77.4	70.5	63.0	56.5
Difference	0.1	1.6	0.3	–2.2	–3.1
	C. Married				
1911	9.3	10.7	8.9	9.0	8.7
1921	9.2	10.8	8.9	8.8	8.1
Difference	–0.2	0.0	–0.0	–0.2	–0.5
	D. Widowed				
1911	55.8	64.3	67.2	57.2	46.3
1921	42.2	46.9	43.4	43.7	37.5
Difference	–13.6	–17.4	–23.8	–13.5	–8.8

Notes. This table provides average female labor force participation rates in percent among single, married, and widowed women aged 20 to 54 based on the 1911 and 1921 census I-CeM household schedules (Schürer and Higgs, 2025; Schürer and Wakelam, 2025). 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). *Married* only includes married women whose spouse is noted present in the household schedules.

Table A3. Summary Statistics

	1911					1921				
	Mean	S.d.	Min.	Max.	Obs.	Mean	S.d.	Min.	Max.	Obs.
A. Individual level										
Military death rate						0.066	0.028	0.000	0.732	9,401,537
Active	0.345	0.475	0	1	8,612,461	0.326	0.469	0	1	9,401,537
Single	0.362	0.481	0	1	8,612,461	0.338	0.473	0	1	9,401,537
Married	0.591	0.492	0	1	8,612,461	0.604	0.489	0	1	9,401,537
Widowed	0.046	0.210	0	1	8,612,461	0.057	0.232	0	1	9,401,537
Active | single	0.727	0.445	0	1	3,120,799	0.728	0.445	0	1	3,181,621
Active | married	0.093	0.291	0	1	5,093,468	0.092	0.289	0	1	5,681,522
Active | widowed	0.558	0.497	0	1	398,194	0.422	0.494	0	1	538,394
Age	34.6	9.7	20	54	8,612,461	35.5	9.9	20	54	9,401,537
Children 16–	1.13	1.62	0	22	8,612,461	0.98	1.46	0	18	9,401,537
B. District level										
Military death rate						0.077	0.049	0.000	0.732	1,721
Active	0.317	0.101	0.084	0.672	1,721	0.291	0.096	0.080	0.711	1,721
Share single	0.370	0.087	0.113	0.690	1,721	0.347	0.077	0.118	0.600	1,721
Share married	0.589	0.091	0.255	0.867	1,721	0.601	0.080	0.345	0.848	1,721
Share widowed	0.041	0.011	0.000	0.111	1,721	0.051	0.011	0.013	0.132	1,721
Share active | single	0.673	0.102	0.304	0.935	1,721	0.650	0.121	0.300	0.924	1,721
Share active | married	0.076	0.073	0.000	0.541	1,721	0.077	0.068	0.000	0.584	1,721
Share active | widowed	0.534	0.128	0.000	1.000	1,721	0.367	0.115	0.000	1.000	1,721
Age	34.98	0.72	31.85	37.68	1,721	35.84	0.67	32.51	37.85	1,721
Children 16–	1.14	0.28	0.38	2.16	1,721	0.98	0.26	0.26	1.90	1,721

Notes. This table reports summary statistics for the sample of single, married with spouse present, and widowed women aged 20 to 54 that could be matched to a local government district based on the 1911 and 1921 census I-CeM household schedules (Schürer and Higgs, 2025; Schürer and Wakelam, 2025). Summary statistics are provided at the individual level in Panel A and at the district level in Panel B. 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). *Married* only includes married women whose spouse is noted present in the household schedules. *S.d.* denotes standard deviation; *Max.*, maximum; *Min.*, minimum; and *Obs.*, observations.

Table A4. Field Availability in Military Records

		Military-records database							
		A. SDGW		B. ODGW		C. RNWGR		D. CWGC	
	Field	Count	Share	Count	Share	Count	Share	Count	Share
Vital information	Name	661,827	1.00	41,472	1.00	34,804	1.00	835,740	1.00
	Birth year	0	0.00	0	0.00	34,856	1.00	479,079	0.57
	Death date	661,827	1.00	41,472	1.00	34,830	1.00	835,740	1.00
Military information	Service number	661,330	1.00	0	0.00	0	0.00	771,359	0.92
	Rank	659,567	1.00	41,472	1.00	0	0.00	835,737	1.00
	Military unit	661,827	1.00	41,472	1.00	34,874	1.00	835,740	1.00
Location information	Birth	590,350	0.89	0	0.00	34,843	1.00	44,477	0.05
	Enlistment	658,505	0.99	0	0.00	0	0.00	0	0.00
	Residence: own	319,367	0.48	0	0.00	0	0.00	0	0.00
	Residence: parents	0	0.00	0	0.00	0	0.00	369,498	0.44
	Residence: spouse	0	0.00	0	0.00	0	0.00	125,811	0.15
	Residence: sibling	0	0.00	0	0.00	0	0.00	7,298	0.01
	Residence: child	0	0.00	0	0.00	0	0.00	513	0.00
	Burial	0	0.00	0	0.00	0	0.00	835,740	1.00
	All records	*661,827*	*1.00*	*41,472*	*1.00*	*34,874*	*1.00*	*835,740*	*1.00*

Notes. This table reports field-level data availability across four military-records databases: the *Soldiers Died in the Great War* (SDGW), the *Officers Died in the Great War* (ODGW), the *Royal Navy and Royal Marine War Graves Roll* (RNWGR), and the *Commonwealth War Graves Commission* (CWGC). The *Name* field combines first and last names. The *Birth year* field is inferred from the age at death reported in CWGC records. Field availability for military units and honors listed on CWGC records is omitted. Except for burial location, all location information in CWGC records comes from the *Additional information* field.

Table A5. Field Availability in Linked SDGW–CWGC Records

	Field	Count	Share
Vital information	Name	659,412	1.00
	Birth year	367,416	0.56
	Death date	659,412	1.00
Military information	Service number	659,084	1.00
	Rank	659,412	1.00
	Regiment	659,412	1.00
Location information	Birth	591,204	0.90
	Enlistment	656,139	1.00
	Residence: own	318,187	0.48
	Residence: parents	291,708	0.44
	Residence: spouse	93,688	0.14
	Residence: sibling	6,530	0.01
	Residence: child	420	0.00
	Residence: any	502,184	0.76
	Burial	659,412	1.00
	All records	*659,412*	*1.00*

Notes. This table reports field-level data availability across the linked SDGW-CWGC records. The *Name* field combines first and last names. The *Birth year* field is inferred from the age at death reported in CWGC records. Field availability for military units and honors listed on CWGC records is omitted. Except for burial location, all location information in CWGC records comes from the *Additional information* field. The *Residence: any* field combines all residence location fields.

Table A6. SDGW-CWGC Records Geocoding Rates

	Non-missing		Geocoded		In E&W	
Location field	Count	Share	Count	Share	Count	Share
Residence	318,187	0.483	316,722	0.480	273,803	0.415
Enlistment	656,139	0.995	654,740	0.993	551,008	0.836
Birth	591,204	0.897	586,973	0.890	483,219	0.733

Notes. This table reports geocoding rates among the 659,412 combined SDGW-CWGC records for residence at enlistment, enlistment, and birth places.

Table A7. Selection of SDGW–CWGC Records Geocoded in England and Wales

Characteristics	Full sample Obs.	Full sample Mean	Restricted sample Obs.	Restricted sample Mean	Difference
A. Selection into non-missing place of residence					
Birth year	367,416	1,891.23	178,771	1,891.15	−0.080***
		(6.64)		(6.70)	[0.012]
Death year	659,412	1,917.14	318,187	1,917.12	−0.014***
		(1.15)		(1.15)	[0.001]
Private	659,412	0.813	318,187	0.814	0.001*
		(0.390)		(0.389)	[0.000]
B. Selection into geocoded | non-missing place of residence					
Birth year	178,771	1,891.15	177,930	1,891.15	0.001***
		(6.70)		(6.70)	[0.001]
Death year	318,187	1,917.12	316,722	1,917.12	0.000***
		(1.15)		(1.15)	[0.000]
Private	318,187	0.814	316,722	0.814	−0.000*
		(0.389)		(0.389)	[0.000]
C. Selection into England and Wales | geocoded					
Birth year	177,930	1,891.15	154,759	1,891.19	0.044***
		(6.70)		(6.70)	[0.001]
Death year	316,722	1,917.12	273,803	1,917.15	0.029***
		(1.15)		(1.14)	[0.000]
Private	316,722	0.814	273,803	0.813	−0.000*
		(0.389)		(0.390)	[0.000]

Notes. This table reports means for selected characteristics among the matched SDGW-CWGC records, comparing the full sample with progressively restricted subsamples. In Panel A, the restricted sample includes records with a non-missing place of residence at enlistment. In Panel B, it includes records that we are able to geocode, conditional on a non-missing residence place. In Panel C, it includes records located in England and Wales, conditional on being geocoded. The last column reports differences in means across the full and restricted samples. Bootstrap standard errors based on 1,000 replications are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A8. Enlistment Rules and Eligible Cohorts (1914–18)

Enlistment regime	Effective period	Statutory ages	Implied eligible cohorts: At implementation	Implied eligible cohorts: Ever eligible	Age range in 1911
Voluntary	08/1914 to 09/1914	19–30	1884–1895	1884–1895	16–27
	09/1914 to 10/1914	19–35	1879–1895	1879–1895	16–32
	10/1914 to 05/1915	19–38	1876–1895	1876–1896	15–35
	05/1915 to 11/1918	19–40	1875–1896	1875–1899	12–36
Conscription	03/1916 to 06/1916	18–40*	1876–1898	1876–1898	13–35
	07/1916 to 03/1918	18–40	1876–1898	1876–1900	11–35
	04/1918 to 11/1918	18–50	1868–1900	1868–1900	11–43
Extreme bounds				1868–1900	11–43

Notes. This table reports the main wartime enlistment regimes into the British Army together with their effective periods, statutory age limits, and implied eligible birth cohorts. Voluntary enlistment rules are from Simkins (1988). Conscription rules are based on the successive Military Service Acts (see Table A9). Because legal eligibility depended on exact ages and birthdays, the cohort ranges shown here are approximate calendar-year translations of those rules. *At implementation* denotes the cohorts eligible when each regime took effect. *Ever eligible* denotes all cohorts exposed to that regime during its period of operation. *Conscription restricted to unmarried men and childless widowers.

Table A9. Military Service Acts (1916–18)

Title	Date: Passed	Date: Effective	Eligibility rules: Residency	Eligibility rules: Age	Eligibility rules: Status
Military Service Act, 1916	1916-01-27	1916-03-02	Great Britain	18–40*	Unmarried†
Military Service Act (No. 2), 1916	1916-05-25	1916-06-24	Great Britain	18–40‡	All
Military Service (Review of Exceptions) Act, 1917	1917-04-05	1917-04-05	—	—	—
Military Service (Conventions with Allied States) Act, 1917	1917-07-10	1917-07-10	Any	—	—
Military Service Act, 1918	1918-02-06	1918-02-06	—	—	—
Military Service (No. 2) Act, 1918	1918-04-18	1918-04-18	—	18–50	—

Notes. This table lists the British Military Service Acts enacted between 1916 and 1918. *Passed* gives the date of royal assent of the Act. *Effective* gives the date on which the provisions of the Act came into operation. *Eligibility rules* report the residency, age, and status conditions specified or modified by each Act. The upper bound of each age range refers to men who had attained that bound but not the next age—e.g., *18–40* denotes men aged at least 18 and under 41. Dashes indicate that the Act did not itself alter that dimension of eligibility. Military Service Acts are published in the *Public and General Acts*, *5–9 Geo 5* (London, Eyre and Spottiswoode, 1916–18). *Age as of August 15, 1915. †Also includes childless widowers. Both status as of November 2, 1915. ‡Age as of the appointed date, i.e., June 24, 1916.

Table A10. Enlistments and Enlistment Rates (1914–18)

	Enlistments			Share of men 11–43 in 1911*		
	Voluntary	Conscription†	Total	Voluntary	Conscription†	Total
England and Wales	2,237,497	2,041,585	4,279,082	23.5	21.4	44.9
England	2,092,242	1,913,916	4,006,158	23.7	21.6	45.3
Wales	145,255	127,669	272,924	21.4	18.8	40.2
Scotland	320,589	237,029	557,618	25.6	19.0	44.6
Ireland	117,063	17,139	134,202	9.7	1.4	11.1
United Kingdom	2,675,149	2,295,753	4,970,902	22.3	19.2	41.5

Notes. This table reports the number of enlistments in the British Regular Army and Territorial Force from August 1914 to November 1918, and enlistments as a percentage of the male population aged 11 to 43 in 1911. Enlistment data are drawn from the War Office's (1921, p. 9) *General Annual Reports*. Countries refer to place of enlistment. Male population figures for England and Wales are based on the *Census of England and Wales, 1911. Summary Tables* (London, 1915) table 29, p. 90. Because the published tables do not separately report the England–Wales breakdown by age group, we apportion the male population aged 11 to 43 between England and Wales using the corresponding shares in I-CeM (Schürer and Higgs, 2025). This procedure implies that, among men aged 11 to 43, 92.8 percent resided in England in 1911 and 7.13 percent in Wales. Male population figures for Scotland are based on the *Census of Scotland, 1911. Report on the Twelfth Decennial Census of Scotland*. Volume II (Edinburgh, 1913), table 23, pp. 224–25. Male population figures for Ireland are based on the *Census of Ireland, 1911. General Report, with Tables and Appendix* (Dublin, 1913), table 60, p. 74. *For Ireland, the denominator is the male population aged 10 to 44 in 1911. †Includes all enlistments from November 1917 to November 1918 and men attested under the Derby Scheme.

Table A11. Cumulative Distribution of Military Fatalities across Ages in 1911

A. Fatalities from E&W			B. All fatalities		
Perc.	Count	Age	Perc.	Count	Age
1.6	2,487	11.0	1.5	5,501	11.0
25.0	38,780	14.5	25.0	91,862	14.5
50.0	77,419	18.3	50.0	183,738	18.2
75.0	116,069	24.0	75.0	275,550	23.8
99.6	154,187	43.0	99.7	366,158	43.0

Notes. This table reports the age in 1911 at which the cumulative distribution of military fatalities reaches selected percentiles (*Perc.*), together with the corresponding cumulative counts. In Panel A, the underlying sample consists of the 154,759 matched SDGW-CWGC records whose place of residence at enlistment was geocoded in England and Wales for which the year of birth could be recovered from the reported age at death in CWGC records. In Panel B, the underlying sample consists of the 367,416 matched SDGW-CWGC records for which the year of birth could be recovered.

Table A12. Summary Statistics: 1911 Local Government District Characteristics

	Mean	S.d.	Min.	Max.	Obs.
Baseline control variables					
Share in protected occupations	0.607	0.178	0.095	0.954	1,721
Mean age	32.51	0.75	26.04	35.68	1,721
Share unmarried	0.487	0.060	0.343	0.858	1,721
Share with dependent children	0.404	0.052	0.111	0.554	1,721
Infant mortality rate	0.109	0.049	0.000	1.000	1,721
Rural district	0.385	0.487	0.000	1.000	1,721
Share born outside county	0.383	0.189	0.050	1.000	1,721
Log population	9.117	1.163	4.762	13.643	1,721
Coastal district	0.174	0.379	0.000	1.000	1,721
Welsh district	0.105	0.306	0.000	1.000	1,721
Additional control variables					
Share in military occupations	0.007	0.025	0.000	0.627	1,721
Presence of a war factory	0.444	0.497	0.000	1.000	1,721
Presence of an infantry depot	0.027	0.161	0.000	1.000	1,721

Notes. This table reports summary statistics for the 1911 local government district characteristics used as control variables interacted with the *Post* indicator in the empirical analysis. See the main text for variable definitions. *S.d.* denotes standard deviation; *Min.*, minimum; *Max.*, maximum; and *Obs.*, observations.

Table A13. The Prewar Determinants of Military Death Rates
Unstandardized Coefficients

Outcome:	Military death rate						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Share in protected occupations	−0.061***				−0.071***	−0.069***	−0.071***
	[0.006]				[0.009]	[0.009]	[0.013]
Mean age		0.009***			0.005**	0.004**	0.005**
		[0.002]			[0.002]	[0.002]	[0.002]
Share unmarried		−0.280***			−0.084	−0.036	−0.106
		[0.050]			[0.064]	[0.077]	[0.079]
Share with dependent children		−0.262***			0.024	0.085	−0.022
		[0.059]			[0.072]	[0.089]	[0.092]
Infant mortality rate			−0.110***		−0.037	−0.039	−0.018
			[0.029]		[0.031]	[0.031]	[0.032]
Rural district				−0.011***	−0.001	−0.001	−0.003
				[0.002]	[0.003]	[0.003]	[0.003]
Share born outside county				0.001	−0.019**	−0.018**	−0.033***
				[0.007]	[0.008]	[0.007]	[0.009]
Log population				−0.009***	−0.009***	−0.009***	−0.009***
				[0.001]	[0.001]	[0.001]	[0.001]
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						0.024	0.019
						[0.015]	[0.017]
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. Robust standard errors are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A14. The Prewar Determinants of Military Death Rates
Additional Control Variables

Outcome:	Military death rate						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Share in military occupations	−0.004***				−0.003***	−0.003***	−0.002**
	[0.001]				[0.001]	[0.001]	[0.001]
Presence of a war factory		−0.001			0.008***	0.008***	0.007***
		[0.002]			[0.003]	[0.003]	[0.003]
Presence of an infantry depot			−0.009		−0.006	−0.006	−0.007
			[0.006]		[0.006]	[0.006]	[0.006]
Share in protected occupations				−0.013***	−0.013***	−0.013***	−0.013***
				[0.002]	[0.002]	[0.002]	[0.002]
Mean age				0.003**	0.002	0.002	0.003
				[0.002]	[0.002]	[0.002]	[0.002]
Share unmarried				−0.005	−0.005	−0.002	−0.006
				[0.004]	[0.004]	[0.005]	[0.005]
Share with dependent children				0.001	0.001	0.004	−0.001
				[0.004]	[0.004]	[0.005]	[0.005]
Infant mortality rate				−0.002	−0.002	−0.002	−0.001
				[0.001]	[0.001]	[0.001]	[0.002]
Rural district				−0.001	−0.000	−0.000	−0.003
				[0.003]	[0.003]	[0.003]	[0.004]
Share born outside county				−0.004**	−0.004**	−0.003**	−0.006***
				[0.001]	[0.001]	[0.001]	[0.002]
Log population				−0.010***	−0.012***	−0.012***	−0.012***
				[0.002]	[0.002]	[0.002]	[0.002]
Coastal district				−0.017***	−0.016***	−0.015***	−0.014***
				[0.002]	[0.002]	[0.002]	[0.003]
Welsh district				−0.017***	−0.017***	−0.018***	0.019
				[0.003]	[0.003]	[0.003]	[0.020]
FLFP						0.002	0.002
						[0.002]	[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	0.006	0.000	0.001	0.172	0.180	0.181	0.259

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. Information on war factories is drawn from the *1918 Directory of Manufacturers in Engineering and Allied Trades*. Information on infantry depots is drawn from the *List of commands, military districts and regimental areas* in the August 1916 Army Council *Registration and Recruiting* instructions (102–8), held at the National Archives under record number `MH 47/142/1`. Shares are in percent. Independent variables—except for indicator variables—are standardized. Robust standard errors are reported in brackets.
*** $p \leq 0.01$, ** $p \leq 0.05$, * $p \leq 0.10$.

Table A15. The Prewar Determinants of Military Death Rates
LASSO-Selected Controls

Outcome:	Military death rate				
	(1)	(2)	(3)	(4)	(5)
Share of coal and shale miners	-0.003^{***}		-0.002^{**}	-0.001	-0.002
	[0.001]		[0.001]	[0.001]	[0.001]
Share of bricklayers	0.008^{***}		0.006^{***}	0.006^{***}	0.004^{**}
	[0.001]		[0.001]	[0.001]	[0.002]
Share of butchers	0.007^{***}		0.006^{***}	0.006^{***}	0.006^{***}
	[0.002]		[0.002]	[0.002]	[0.002]
Share of postmen	0.007^{***}		0.008^{***}	0.007^{***}	0.011^{***}
	[0.002]		[0.002]	[0.002]	[0.002]
Share in navy	-0.003^{***}		-0.002^{***}	-0.002^{***}	-0.000
	[0.001]		[0.001]	[0.001]	[0.000]
Share of tramway conductors	-0.004^{***}		-0.002^{*}	-0.002^{**}	-0.003^{***}
	[0.001]		[0.001]	[0.001]	[0.001]
Share of harbour and dock workers	-0.002^{***}		-0.000	-0.000	-0.000
	[0.000]		[0.000]	[0.000]	[0.000]
Share of watch and clockmakers	0.003^{**}		0.003^{**}	0.003^{**}	0.003^{*}
	[0.001]		[0.001]	[0.001]	[0.001]
Share of bakers	0.003^{*}		0.004^{**}	0.003^{**}	0.003^{**}
	[0.002]		[0.002]	[0.002]	[0.002]
Share with no occupation provided	-0.004^{***}		-0.004^{***}	-0.004^{***}	-0.003^{***}
	[0.001]		[0.001]	[0.001]	[0.001]
Log population		-0.011^{***}	-0.006^{***}	-0.006^{***}	-0.007^{***}
		[0.001]	[0.002]	[0.002]	[0.002]
Coastal district		-0.014^{***}	-0.013^{***}	-0.013^{***}	-0.010^{***}
		[0.002]	[0.002]	[0.002]	[0.002]
Welsh district		-0.029^{***}	-0.025^{***}	-0.025^{***}	0.014
		[0.003]	[0.003]	[0.003]	[0.023]
FLFP				0.003^{***}	0.003^{*}
				[0.001]	[0.001]
County FE	No	No	No	No	Yes
Observations	1,721	1,721	1,721	1,721	1,721
R^2	0.200	0.089	0.247	0.249	0.311

Notes. This table reports coefficient estimates from regressing the military death rate on 1911 local government district characteristics selected using LASSO with the plug-in selection method. The candidate set comprises 640 variables. In addition to baseline and additional controls used in Table A14, these cover the detailed age and marital-status distribution among men aged 18 to 50, major HISCO occupational groups, and all 558 occupation or industry Hollerith categories in the 1911 census employing at least 100 men nationally. The unit of observation is a consistent 1911–21 district. Shares are expressed as proportions. Independent variables—except for indicator variables—are standardized. Robust standard errors are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A16. The Impact of Military Death Rates on Female Labor Force Participation
LASSO-Selected Controls

Outcome:	Female labor force participation					MLFP
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate × Post	0.130***	0.071***	0.075***		0.087***	−0.001
	[0.025]	[0.025]	[0.023]		[0.026]	[0.039]
Adj. death rate × Post				0.176***		
				[0.033]		
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	0.107	0.107	0.107	0.107	0.519	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	0.024	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 14 prewar district-level LASSO-selected characteristics included in Column 5 of Table A15, 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.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A17. The Impact of Military Death Rates on Female Labor Force Participation
Robustness to District-Level Sample Restrictions

Outcome:	Female labor force participation							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Death rate × Post	0.104^{***}	0.094^{***}	0.097^{***}	0.094^{***}	0.098^{***}	0.097^{***}	0.156^{***}	0.149^{***}
	[0.024]	[0.024]	[0.032]	[0.025]	[0.034]	[0.025]	[0.041]	[0.045]
Excluded districts	None	Wales	Rural	Coast	Wales, rural, and coast	Bottom	Top	Bottom and top
Observations	18,013,998	16,899,114	12,743,577	14,384,962	10,199,575	16,880,715	17,552,049	16,418,766
Districts	1,721	1,541	1,058	1,421	818	1,548	1,549	1,376
Within R^2	0.108	0.110	0.113	0.112	0.119	0.108	0.107	0.108
1911 outcome mean	0.345	0.351	0.367	0.352	0.382	0.349	0.344	0.349
Treatment s.d.	0.028	0.028	0.027	0.029	0.027	0.027	0.021	0.018

Notes. This table reports coefficient estimates from Equation 1 using alternative district-level sample restrictions. Column 1 reports the baseline specification. Columns 2–4 exclude local government districts located in Wales, rural districts, and coastal districts, respectively. Column 5 excludes all three groups. Columns 6 and 7 exclude districts in the bottom and top deciles of the military death-rate distribution, respectively. Column 8 excludes districts in both the bottom and top deciles. All specifications use the sample of women aged 20 to 54 in the 1911 and 1921 censuses and include age and 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.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A18. The Impact of Military Death Rates on Female Labor Force Participation
Robustness to Alternative Outcome Measures

Outcome:	Active		Employed	
	(1)	(2)	(3)	(4)
Death rate × Post	0.104***	0.120***	0.111***	0.101***
	[0.024]	[0.026]	[0.031]	[0.025]
Corrected measure	Yes	No	Yes	No
Observations	18,013,998	18,013,998	18,013,998	18,013,998
Districts	1,721	1,721	1,721	1,721
Within R^2	0.108	0.109	0.101	0.108
1911 outcome mean	0.345	0.357	0.344	0.344
Treatment s.d.	0.028	0.028	0.028	0.028

Notes. This table reports coefficient estimates from Equation 1 using alternative measures of female labor force participation. Columns 1 and 2 use an indicator for whether a woman was economically active, while Columns 3 and 4 use an indicator for whether she was employed. Columns 1 and 3 use the corrected versions of these measures, while Columns 2 and 4 use the corresponding uncorrected versions. All specifications use the sample of women aged 20 to 54 in the 1911 and 1921 censuses 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.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A19. The Impact of Military Death Rates on Female Labor Force Participation
Robustness to Alternative Sample Definitions

Outcome:	Female labor force participation					
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate × Post	0.104***	0.098***	0.097***	0.098***	0.076***	0.076***
	[0.024]	[0.024]	[0.023]	[0.023]	[0.020]	[0.021]
Age range	20–54	20–54	20–64	14–59	10+	12+
Known marital status	Yes	No	Yes	Yes	Yes	Yes
Observations	18,013,998	19,130,514	20,589,997	23,582,553	29,150,135	28,437,912
Districts	1,721	1,721	1,721	1,721	1,721	1,721
Within R^2	0.108	0.098	0.105	0.155	0.189	0.180
1911 outcome mean	0.345	0.345	0.330	0.389	0.331	0.348
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 alternative sample definitions. Column 1 reports the baseline specification, which includes women aged 20 to 54 with a precise marital-status classification. Column 2 includes all women aged 20 to 54, irrespective of whether marital status is precisely classified. Columns 3–6 use the following age ranges: ages 20–64, following Joshi, Layard and Owen's (1985, table 1, p. S151) definition; ages 14–59, following Hakim's (1980, table 2, p. 559) and Hatton's (2014, table 4.5, p. 110) definition; ages 10 and over, following the 1911 census report definition; and ages 12 and over, following the 1921 census report definition. 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.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A20. The Impact of Military Death Rates on Female Labor Force Participation
Robustness to Alternative Measures of Military Death Rates

Outcome:	Female labor force participation				
	(1)	(2)	(3)	(4)	(5)
Death rate × Post	0.104***	0.082***	0.124***	0.048*	0.073**
	[0.024]	[0.023]	[0.032]	[0.028]	[0.030]
Measure	Enlistment adjusted	Raw measure	Linked compl.	Enlistment compl.	Birth compl.
Observations	18,013,998	18,013,998	18,013,998	18,013,998	18,013,998
Districts	1,721	1,721	1,721	1,721	1,721
Within R^2	0.108	0.108	0.108	0.108	0.108
1911 outcome mean	0.345	0.345	0.345	0.345	0.345
Treatment s.d.	0.028	0.035	0.025	0.038	0.025

Notes. This table reports coefficient estimates from Equation 1 using alternative measures of military death rates. Column 1 reports the baseline specification, which uses the enlistment-adjusted military death rate. Column 2 uses the raw military death rate, which is based only on residence information. Columns 3–5 use the linked-complemented, enlistment-complemented, and birth-complemented military death rates, respectively. These measures are described in Appendix G. All specifications use the sample of women aged 20 to 54 in the 1911 and 1921 censuses 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.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A21. The Impact of Military Death Rates on the Male Share of the Population

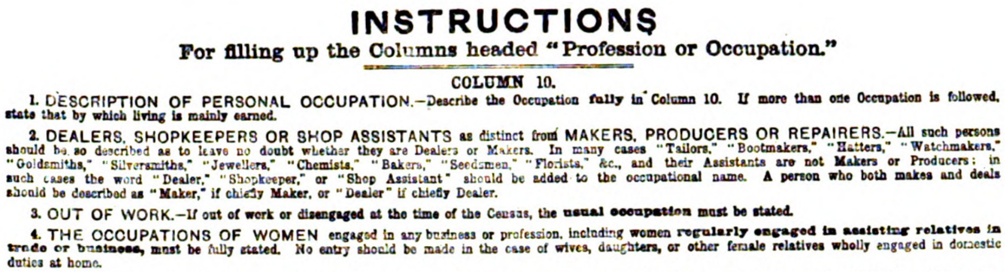
Outcome:	Male share of the population					
	(1)	(2)	(3)	(4)	(5)	(6)
Own-district death rate × Post	−0.046***	−0.040***	−0.042***	−0.042***	−0.042***	−0.041***
	[0.012]	[0.012]	[0.012]	[0.012]	[0.013]	[0.012]
Nearby-district death rate, 0–5 km × Post		−0.114***	−0.105***	−0.113***	−0.117***	−0.109***
		[0.035]	[0.037]	[0.036]	[0.037]	[0.037]
Nearby-district death rate, 5–10 km × Post			−0.014	−0.014	−0.016	−0.019
			[0.022]	[0.023]	[0.024]	[0.024]
Nearby-district death rate, 10–15 km × Post				0.004	0.001	0.001
				[0.031]	[0.032]	[0.033]
Nearby-district death rate, 15–20 km × Post					0.024	0.025
					[0.040]	[0.040]
Nearby-district death rate, 20–25 km × Post						−0.003
						[0.030]
Observations	3,442	3,442	3,404	3,384	3,378	3,368
Districts	1,721	1,721	1,702	1,692	1,689	1,684
Within R^2	0.567	0.572	0.573	0.576	0.577	0.583
1911 outcome mean	0.481	0.481	0.481	0.481	0.481	0.481
Treatment s.d.	0.028	0.028	0.028	0.028	0.028	0.028

Notes. This table reports coefficient estimates from regressions of the male share of the population on military death rates. The unit of observation is a consistent 1911–21 local government district. The outcome is the number of men divided by the total number of men and women aged 20 to 54. Column 1 includes the military death rate in the district. Columns 2–6 sequentially add the population-weighted average military death rates in nearby districts within the indicated distance bands. All specifications include district fixed effects, a postwar indicator, and the 11 prewar district-level characteristics included in Column 6 of Table 2, each interacted with the *Post* indicator. Regressions are weighted by the district's 1911 population aged 20 to 54. Standard errors are clustered at the district level and reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

Table A22. The Impact of Military Death Rates on County-Level Migration

Outcome:		Male share within		Population share of	
	Male share	Stayers	Leavers	Male stayers	Male leavers
	(1)	(2)	(3)	(4)	(5)
Death rate × Post	−0.142*	−0.291***	0.197	−0.329*	0.186
	[0.080]	[0.071]	[0.138]	[0.178]	[0.171]
Observations	108	108	108	108	108
Counties	54	54	54	54	54
Within R^2	0.914	0.955	0.805	0.639	0.726
1911 outcome mean	0.484	0.490	0.477	0.285	0.198
Treatment s.d.	0.012	0.012	0.012	0.012	0.012

Notes. This table reports coefficient estimates from regressions of measures of the male share of the population on county-level military death rates. The unit of observation is a county. Column 1 uses the number of men divided by the total number of men and women as the outcome. Columns 2 and 3 report the male share among stayers and leavers, respectively. Columns 4 and 5 use the numbers of male stayers and male leavers, respectively, divided by the county's total male and female population. Stayers are individuals residing in their county of birth, while leavers are individuals residing outside their county of birth. All outcomes are constructed using men and women aged 20 to 54. All specifications include county fixed effects, a postwar indicator, and population-weighted county averages of the 11 prewar district-level characteristics included in Column 6 of Table 2, each interacted with the *Post* indicator. Regressions are weighted by the county's 1911 population aged 20 to 54. Note that we exclude the county of London because birthplace is recorded using historical counties, whereas the administrative county of London was created from parts of Middlesex, Kent, Essex, and Surrey and therefore has no corresponding county of birth. Standard errors are clustered at the county level and reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.



INSTRUCTIONS

For filling up the Columns headed "Profession or Occupation."

COLUMN 10.

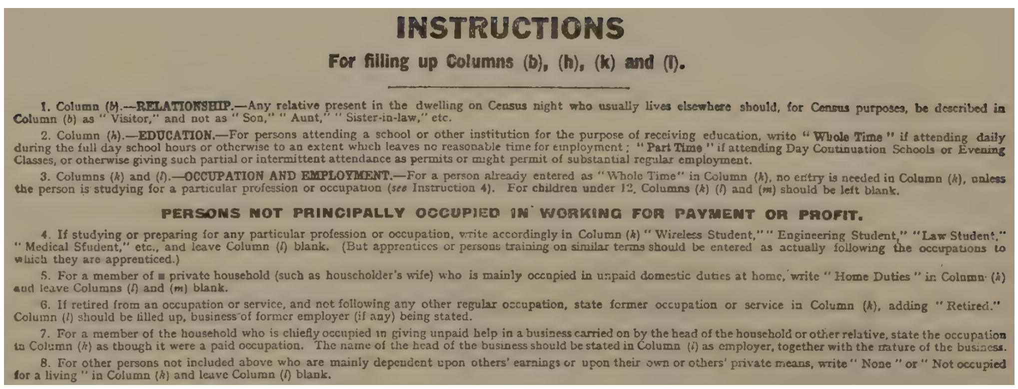
1. DESCRIPTION OF PERSONAL OCCUPATION.—Describe the Occupation fully in Column 10. If more than one Occupation is followed, state that by which living is mainly earned.

2. DEALERS, SHOPKEEPERS OR SHOP ASSISTANTS as distinct from MAKERS, PRODUCERS OR REPAIRERS.—All such persons should be so described as to leave no doubt whether they are Dealers or Makers. In many cases "Tailors," "Bootmakers," "Hatters," "Watchmakers," "Goldsmiths," "Silversmiths," "Jewellers," "Chemists," "Bakers," "Seedsmen," "Florists," &c., and their Assistants are not Makers or Producers: in such cases the word "Dealer," "Shopkeeper," or "Shop Assistant" should be added to the occupational name. A person who both makes and deals should be described as "Maker," if chiefly Maker, or "Dealer" if chiefly Dealer.

3. OUT OF WORK.—If out of work or disengaged at the time of the Census, the usual occupation must be stated.

4. THE OCCUPATIONS OF WOMEN engaged in any business or profession, including women regularly engaged in assisting relatives in trade or business, must be fully stated. No entry should be made in the case of wives, daughters, or other female relatives wholly engaged in domestic duties at home.

(a) 1911 Census



INSTRUCTIONS

For filling up Columns (b), (h), (k) and (l).

1. Column (b).—RELATIONSHIP.—Any relative present in the dwelling on Census night who usually lives elsewhere should, for Census purposes, be described in Column (b) as "Visitor," and not as "Son," "Aunt," "Sister-in-law," etc.

2. Column (h).—EDUCATION.—For persons attending a school or other institution for the purpose of receiving education, write "Whole Time" if attending daily during the full day school hours or otherwise to an extent which leaves no reasonable time for employment; "Part Time" if attending Day Continuation Schools or Evening Classes, or otherwise giving such partial or intermittent attendance as permits or might permit of substantial regular employment.

3. Columns (k) and (l).—OCCUPATION AND EMPLOYMENT.—For a person already entered as "Whole Time" in Column (h), no entry is needed in Column (k), unless the person is studying for a particular profession or occupation (see Instruction 4). For children under 12, Columns (k) (l) and (m) should be left blank.

PERSONS NOT PRINCIPALLY OCCUPIED IN WORKING FOR PAYMENT OR PROFIT.

4. If studying or preparing for any particular profession or occupation, write accordingly in Column (k) "Wireless Student," "Engineering Student," "Law Student," "Medical Student," etc., and leave Column (l) blank. (But apprentices or persons training on similar terms should be entered as actually following the occupations to which they are apprenticed.)

5. For a member of a private household (such as householder's wife) who is mainly occupied in unpaid domestic duties at home, write "Home Duties" in Column (k) and leave Columns (l) and (m) blank.

6. If retired from an occupation or service, and not following any other regular occupation, state former occupation or service in Column (k), adding "Retired." Column (l) should be filled up, business of former employer (if any) being stated.

7. For a member of the household who is chiefly occupied in giving unpaid help in a business carried on by the head of the household or other relative, state the occupation in Column (k) as though it were a paid occupation. The name of the head of the business should be stated in Column (l) as employer, together with the nature of the business.

8. For other persons not included above who are mainly dependent upon others' earnings or upon their own or others' private means, write "None" or "Not occupied for a living" in Column (k) and leave Column (l) blank.

(b) 1921 Census

Figure A1. Instructions for Reporting Women's Occupations in Household Schedules

Notes. This figure reproduces the instructions for reporting women's occupations in the household schedules of the 1911 and 1921 censuses. Panel a reproduces instruction 4 from the *Census of England and Wales, 1911. General Report, with Appendices* (London, 1917), Appendix A, p. 257. Panel b reproduces instruction 5 from the *Census of England and Wales, 1921. General Report with Appendices* (London, 1927), Appendix B, p. 202.

Abbs, George, b. Thorpe-le-Soken, Essex, e. Harwich (Thorpe-le-Soken), 9722, Pte., d. of w., Gallipoli, 20/5/15.

(a) SDGW Record

George Abbs

in the UK, Soldiers Died in the Great War, 1914-1919

Add or update information

Report a problem

Detail Source

Name	George Abbs
Birth Place	Thorpe-le-soken, Essex
Residence	Thorpe-le-soken
Death Date	20 May 1915
Death Place	Gallipoli
Enlistment Place	Harwich
Rank	Private
Regiment	Essex Regiment
Battalion	1st Battalion
Regimental Number	9722
Type of Casualty	Died of wounds
Theatre of War	Balkan Theatre

(b) Ancestry Record

PRIVATE
GEORGE ABBS

Service Number: 9722

Regiment & Unit/Ship

Essex Regiment
1st Bn.

Date of Death

Died 20 May 1915
Age 23 years old

Buried or commemorated at

HELLES MEMORIAL
Panel 146 to 151 or 229 to 233.
Turkey (including Gallipoli)

COMMONWEALTH WAR GRAVES

We are in the process of adding an image of this casualty's point of commemoration

Country of Service	United Kingdom
Additional Info	Son of Mr. G. A. and Mrs. E. C. Abbs, of Abbey St., Thorpe-le-Soken, Clacton-on-Sea, Essex.

(c) CWGC Record

Figure A2. Record of Private George Abbs

Notes. This figure reproduces the record of Private George Abbs (Essex Regiment), who died on May 20, 1915, as it appears in three sources: the original SDGW list in Panel a (*Soldiers Died in the Great War, 1914–19.* Part 48. The Essex Regiment, p. A2, J.B. Hayward & Son, Polstead, Suffolk, 1989), the corresponding Ancestry entry in Panel b (`https://www.ancestry.co.uk/search/collections/1543/records/217414`), and the corresponding CWGC fatality record in Panel c (`https://www.cwgc.org/find-records/find-war-dead/casualty-details/697910`).

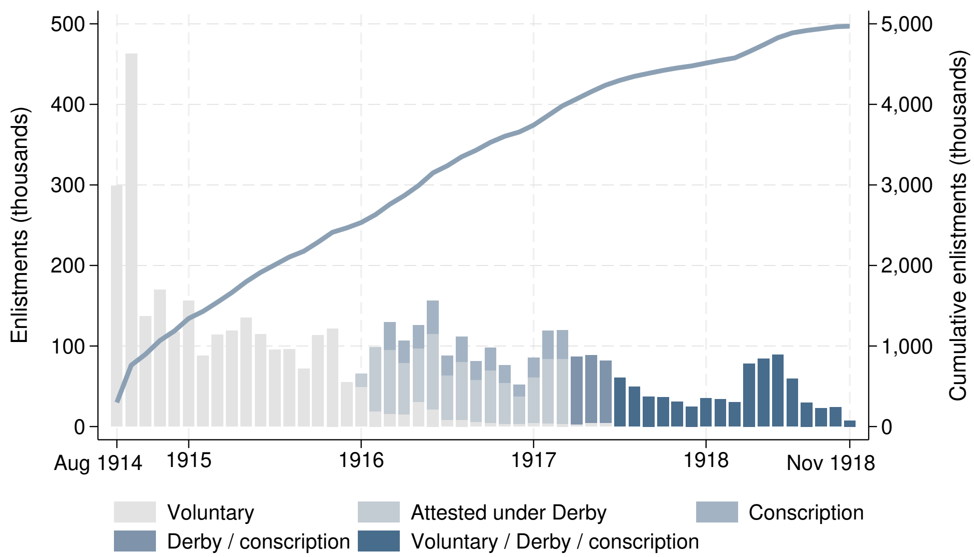


Figure A3. Monthly Enlistments into the British Army (1914–18)

Notes. This figure displays monthly and cumulative enlistments into the British Regular Army and Territorial Force from August 4, 1914, to November 11, 1918, based on table 1B of the War Office's (1921, p. 60) *General Annual Reports*. Bars distinguish voluntary enlistments, enlistments following attestation under the Derby Scheme, and enlistments under conscription. The original source, however, does not provide a consistent monthly decomposition across these categories for the entire period. Voluntary enlistments are separately reported only through June 1917, while Derby Scheme enlistments are separately identified only in January and February 1916. For March 1916 to March 1917, the monthly split between Derby Scheme enlistments and conscription enlistments is imputed from their aggregate ratio over the period. From April to June 1917, the source no longer provides a separate breakdown for these two categories. From July 1917 to November 1918, the source combines all three categories.

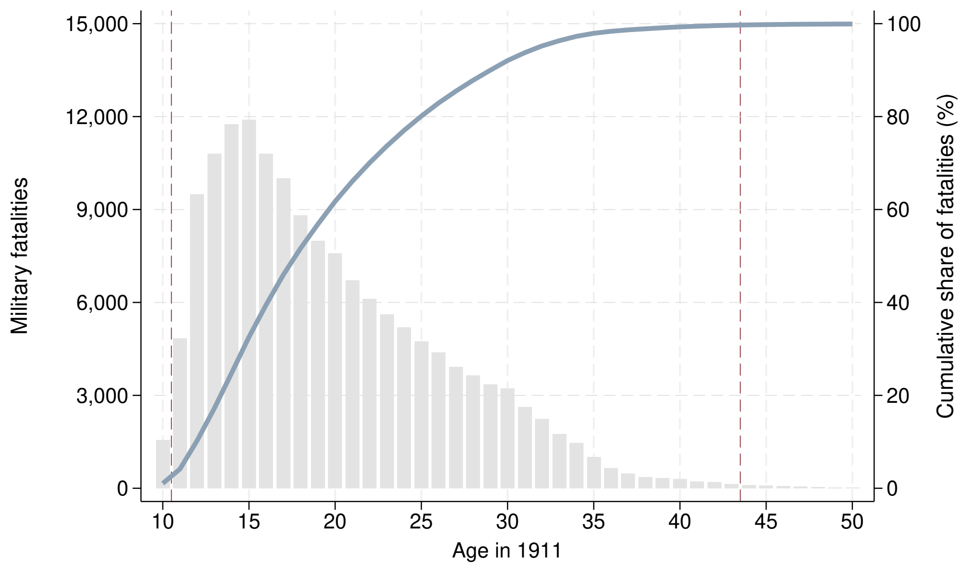


Figure A4. Distribution of Military Fatalities across Ages in 1911

Notes. This figure displays the distribution of military fatalities by age in 1911 together with the cumulative share of fatalities. The underlying sample is restricted to matched SDGW–CWGC records geocoded in England and Wales for which the year of birth could be recovered from the reported age at death in CWGC records, for a total of 154,759 records. Red vertical lines at ages 11 and 43 indicate the lower and upper bounds of the population denominator used in calculating the military death rate. These bounds contain 98.6 percent of military records located in England and Wales for which the year of birth could be recovered. The figure omits 45 fatalities aged 8 to 9 (0.03 percent of all records) and 129 aged 51 to 64 (0.08 percent).

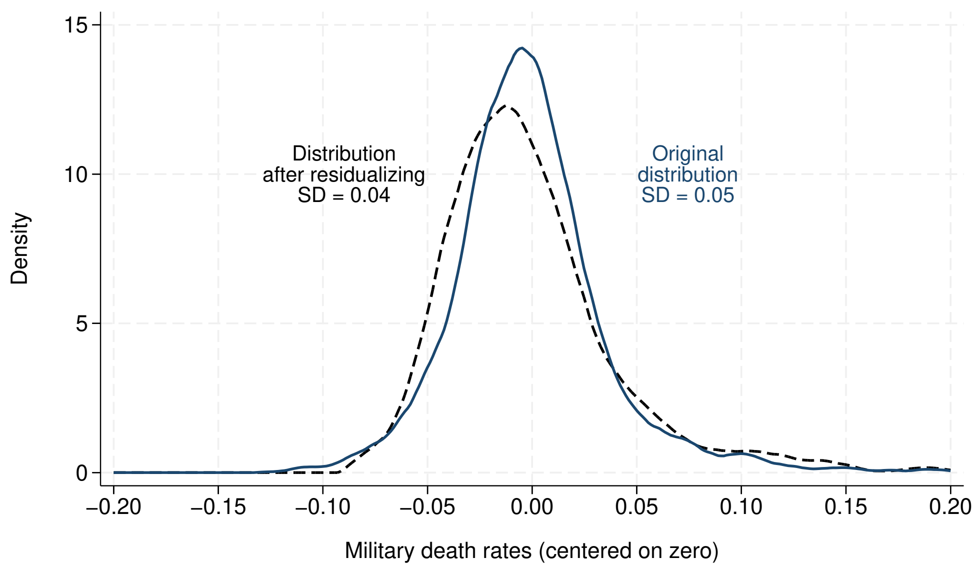


Figure A5. Raw and Residual Distributions of Military Death Rates

Notes. This figure displays the raw and residual distributions of military death rates across the 1,721 consistent local government districts. Densities are estimated using the `kdensity` Stata command. *SD* denotes the standard deviation. This figure is inspired by Mummolo and Peterson (2018, figure 1, p. 831).

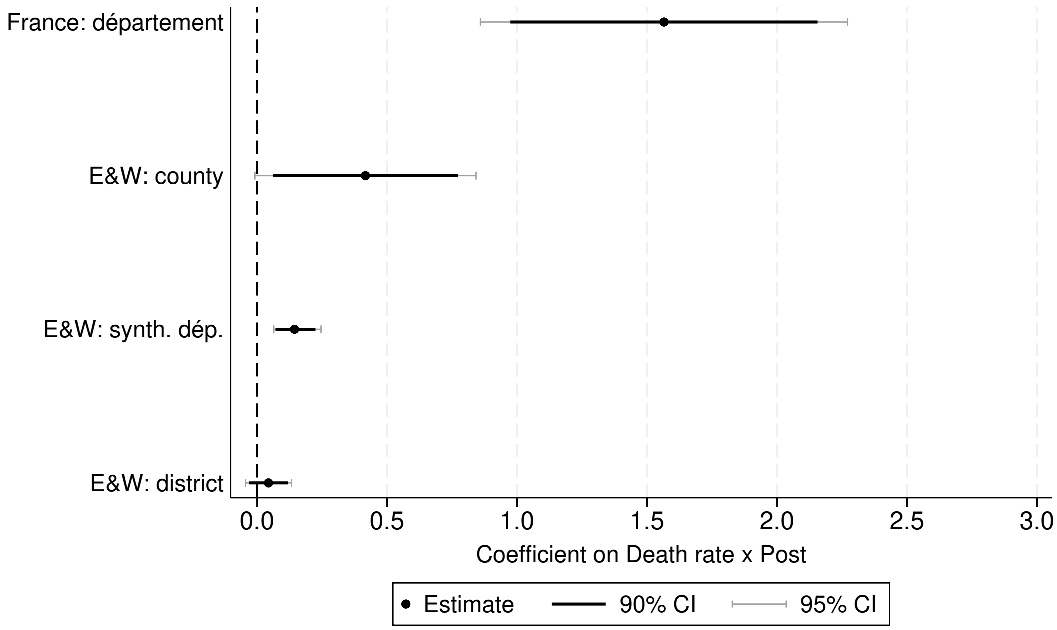


Figure A6. Military Death Rates and Female Labor Force Participation
England and Wales versus France: No Controls

Notes. This figure reports estimates of the effect of military death rates on female labor force participation between 1911 and 1921 without controls. 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.

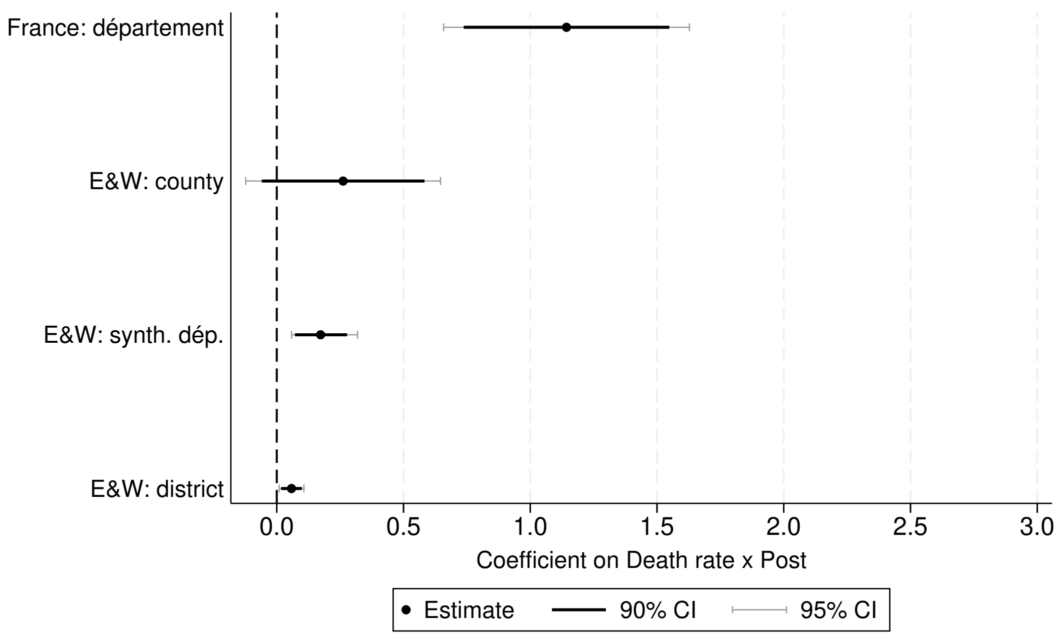


Figure A7. Military Death Rates and Female Labor Force Participation
England and Wales versus France: No Population Weights

Notes. This figure reports estimates of the effect of military death rates on female labor force participation between 1911 and 1921 without population weights. 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.

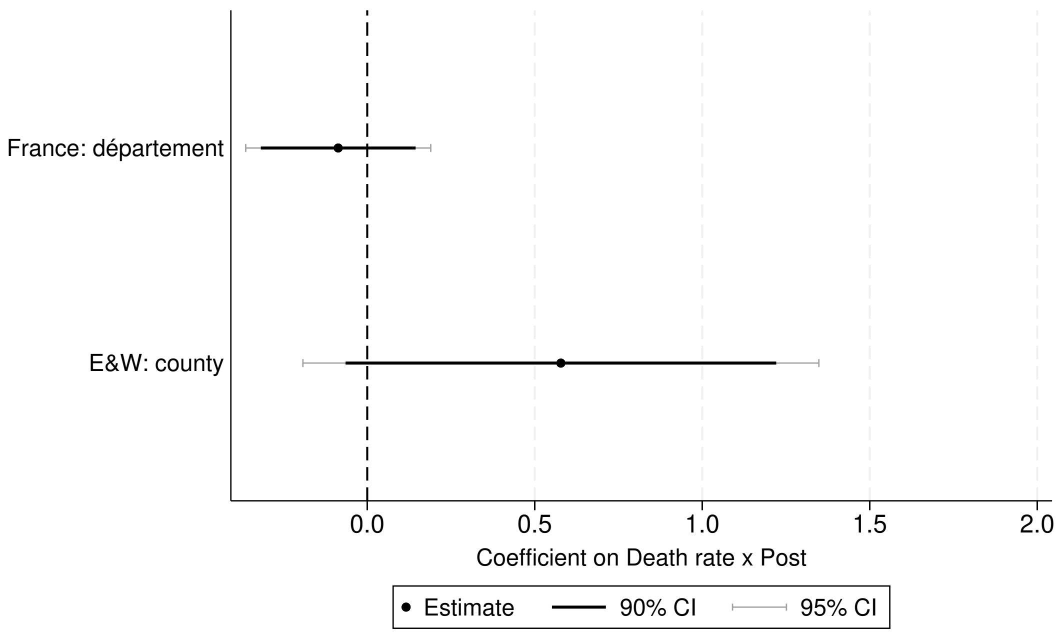


Figure A8. Military Death Rates and Male Internal Migration
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the share of male residents born outside their county or département of residence. For England and Wales, the estimate is reported at the administrative-county level and compares 1911 and 1921. For France, the estimate is reported at the département level and compares 1906 and 1921. Bars report 90 and 95 percent confidence intervals based on standard errors clustered at the corresponding spatial level.

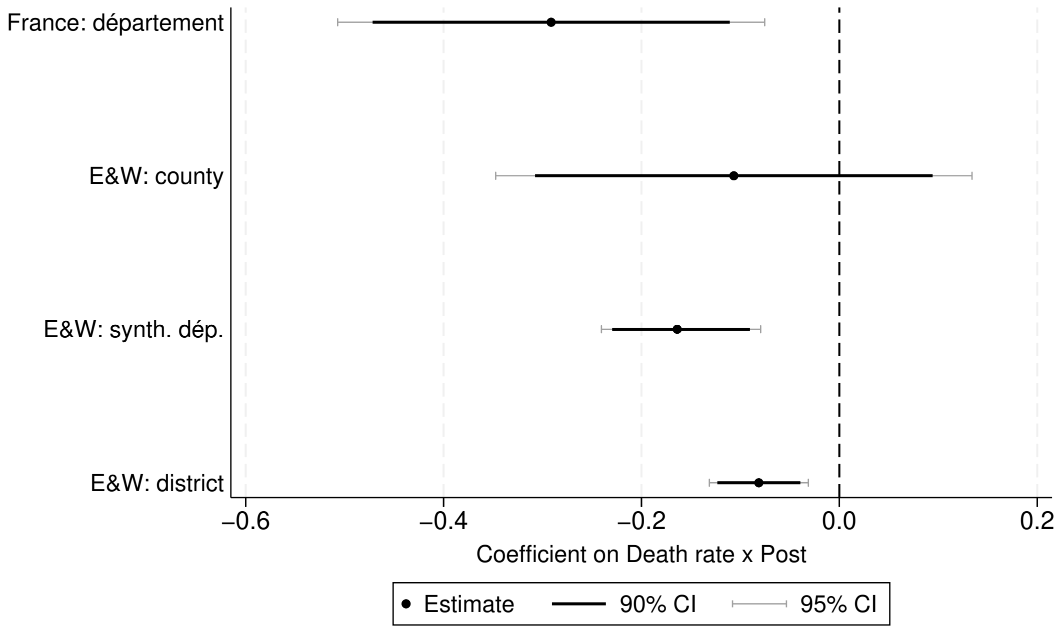


Figure A9. Military Death Rates and the Male Share of the Population
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the male share of the population among individuals aged 20 to 49 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.

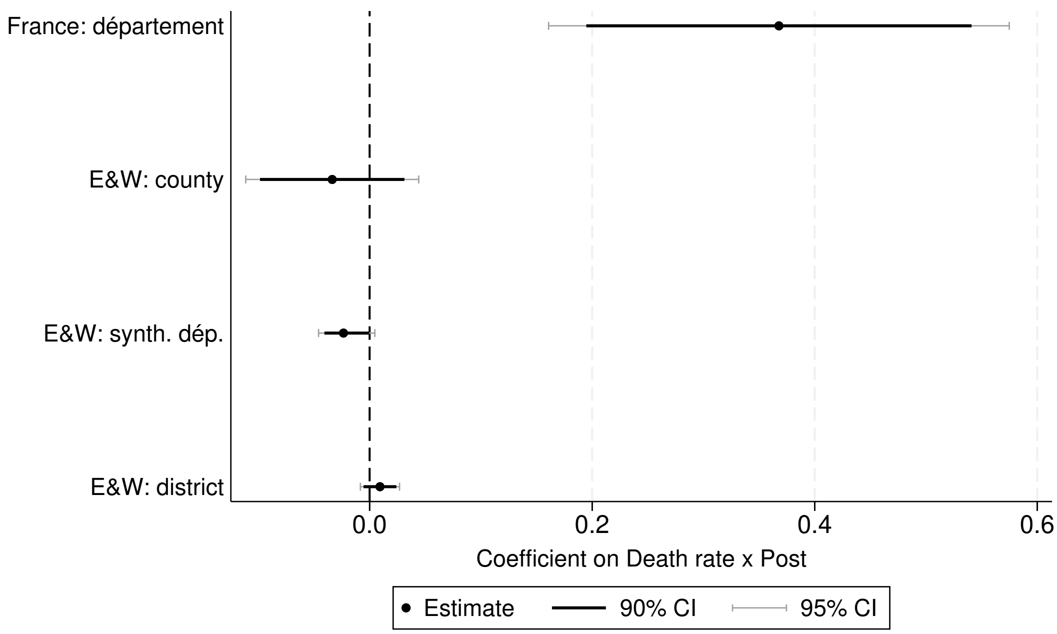


Figure A10. Military Death Rates and the Share of Active Widows
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the share of women aged 15 and over who are both widowed and active in the labor force. For England and Wales, estimates are reported at the local government district, synthetic-département, and administrative-county levels and compare 1911 with 1921. For France, the estimate is reported at the département level and compares 1901 with 1926, the census years for which labor force participation by widow status is available. 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.

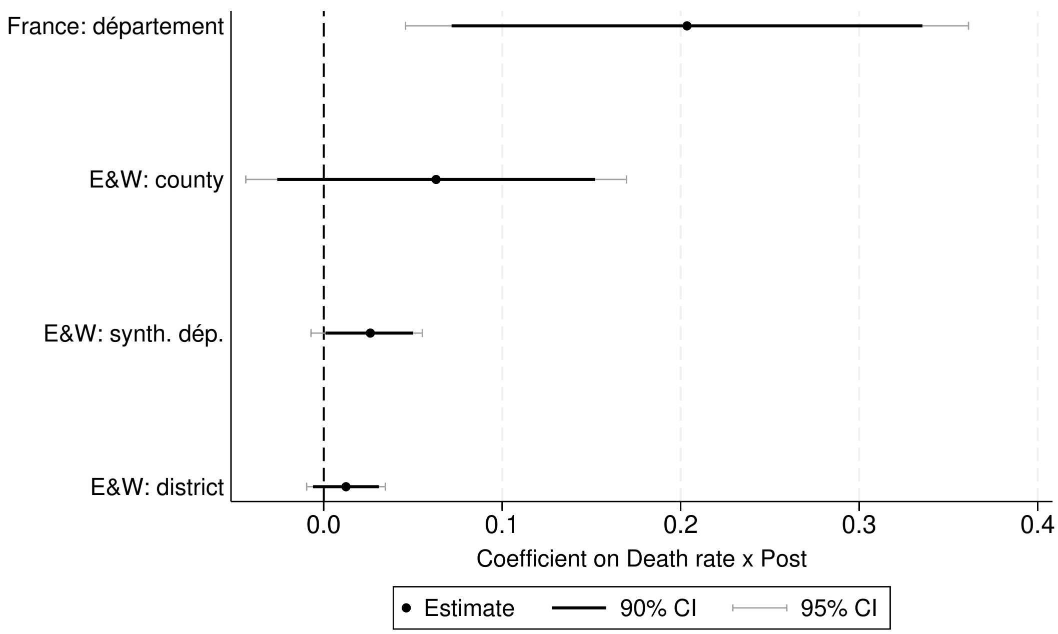


Figure A11. Military Death Rates and Widowhood
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the share of women aged 15 and over who are widowed. For England and Wales, estimates are reported at the local government district, synthetic-département, and administrative-county levels and compare 1911 with 1921. For France, the estimate is reported at the département level and compares 1901 with 1926, the census years used for the corresponding widow analysis. 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.

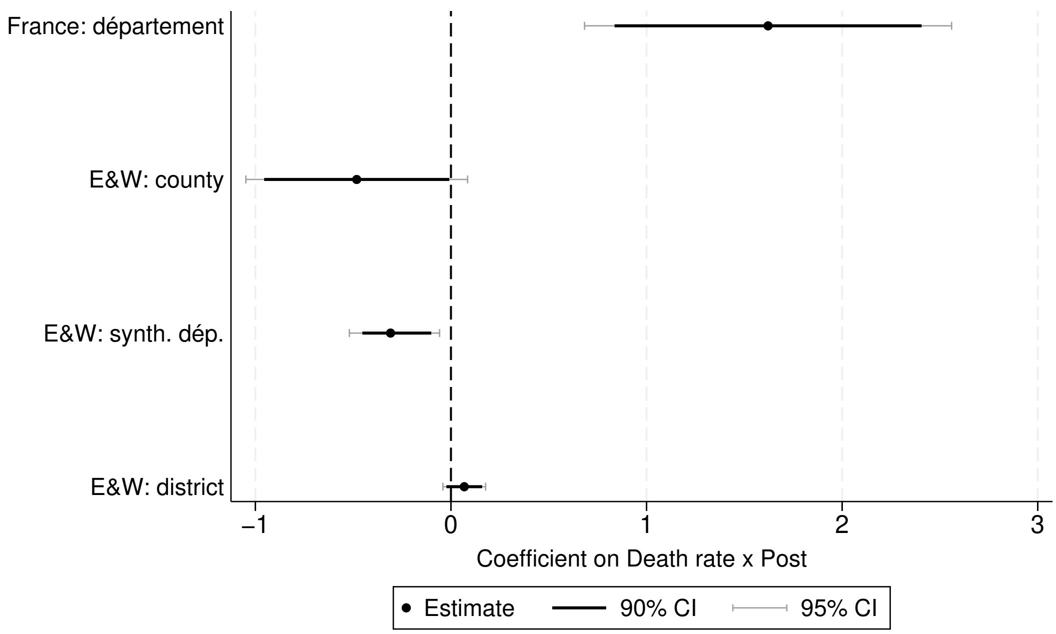


Figure A12. Military Death Rates and Female Labor Force Participation among Widows
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on labor force participation among widowed women aged 15 and over. For England and Wales, estimates are reported at the local government district, synthetic-département, and administrative-county levels and compare 1911 with 1921. For France, the estimate is reported at the département level and compares 1901 with 1926, the census years for which labor force participation among widows is available. 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.

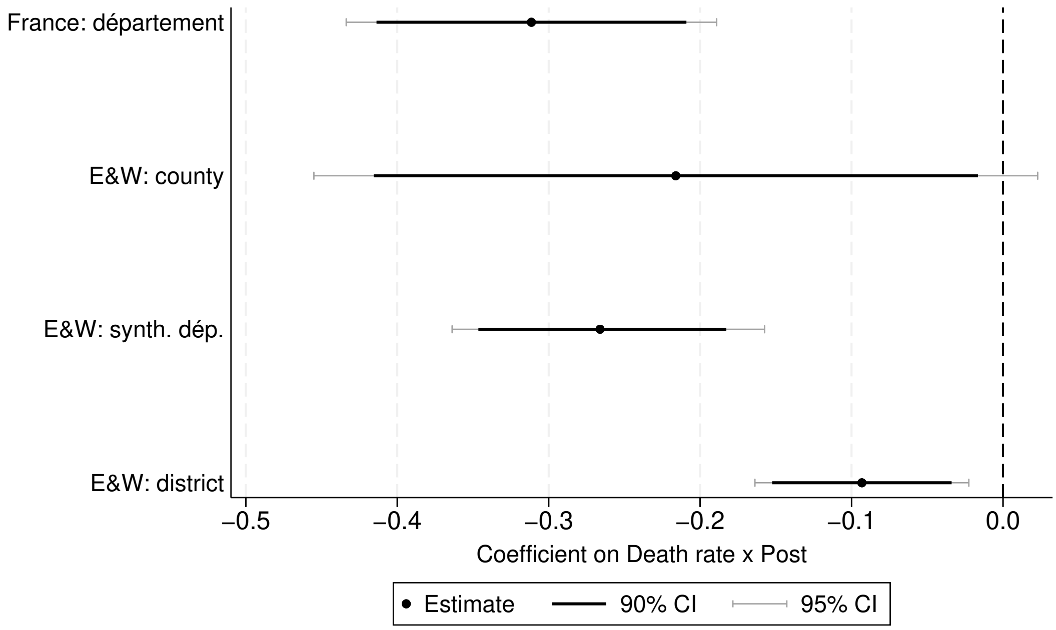


Figure A13. Military Death Rates and Marriage
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the share of women aged 15 and over who are married 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.

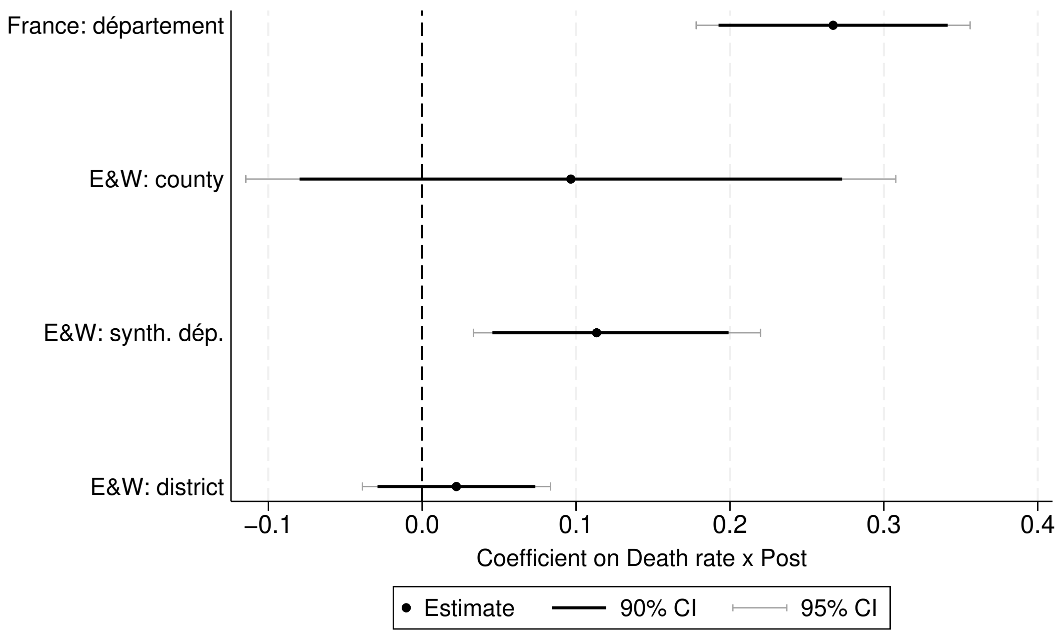


Figure A14. Military Death Rates and Singlehood
England and Wales versus France

Notes. This figure reports estimates of the effect of military death rates on the share of women aged 15 and over who are single 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.

B. Female Labor Force Participation in the Censuses

This section compares female labor force participation rates derived from census reports with those reconstructed from household schedules. Census reports provide a valid benchmark because, from 1911 onward, published tables were compiled directly from household schedules—rather than from census enumerators' books—using Hollerith punch-card machines.[1] Section B.1 focuses on the 1911 census, and Section B.2, on the 1921 census.

B.1. Female Labor Force Participation in the 1911 Census

Table B1 compares female labor force participation rates based on the 1911 census, contrasting figures published in the census report with those derived from household schedules.[2] We classify women as labor force participants if they are recorded as working or unemployed by I-CeM's original `INACTIVE` variable.[3] The table reports figures for all women aged 10 and over—the population used in the census report—as well as for women aged 20 to 64, which can be reconstructed from the age groups available in the report—this estimate also appears in Joshi, Layard and Owen (1985, table 1, p. S151). We additionally provide labor force participation rates by marital status, combining widowed and divorced women as in the census report—divorced women represent only 0.03 percent

[1] In this respect, the *General Report* of the 1911 census states the following: "In one respect this routine differed materially from that followed at previous censuses. At these the enumerator was required to copy the replies to the questions on the schedule into an 'enumeration book,' in which form the information was very much more compactly arranged, blank spaces being eliminated. It was from these enumeration books that the census tables were prepared. This procedure [. . .] led, despite the checking provided for in the regulations, to many copyists' errors, some of a purely accidental nature, and others due to mistaken emendations from misinterpretation of the replies. It was, therefore, decided to omit this copying process in 1911, and to tabulate from the schedules themselves as received from the public" (*Census of England and Wales*, 1911. *General Report with Appendices* [London, 1917], p. 10). It further notes that "[o]wing to extension of the subjects of inquiry, and the demand for greater detail in the published results, the increased work consequently thrown upon the Census Office necessitated some alteration in the system of tabulation adopted at previous Censuses. After investigating the possibilities of the mechanical systems applicable to such an undertaking it was decided to adopt the Hollerith system" (Appendix B, p. 259).

[2] Census report data are from the *Census of England and Wales, 1911.* Volume 10, *Occupations and Industries*, part 1 (London, 1914), table 8, p. 75. Household schedule data are from I-CeM (Schürer and Higgs, 2025).

[3] One exception concerns the measurement of unemployment. In the 1911 census, most women classified as `INACTIVE == 3` explicitly report no occupation—e.g., 51 percent hold the occupational string `NONE`, 17 percent `NO OCCUPATION`, and 15 percent `NIL`. Only 0.5 percent are explicitly labeled `UNEMPLOYED`, consistent with the 1921 definition of unemployment, i.e., active and seeking work. To align our measures with this definition, we recode all women with `INACTIVE == 3` as unoccupied (`INACTIVE == 7`) and classify as unemployed those whose occupational strings contain `UNEMP`, `OUT OF WORK`, or close variants. This remains a residual category in 1911, as "[n]o inquiry as to unemployment was made at this or at any previous census, and persons temporarily out of work on census day were on this occasion instructed in the schedule to state their usual occupation" (*Census of England and Wales, 1911.* Volume 10, part 1, p. vii).

of this combined category.

Table B1. Female Labor Force Participation Rates in 1911 (%)
Broad Comparisons between Census Reports and Household Schedules

	Age group		Marital status		
Group	10+	20–64	Single	Married	Widowed
Census report	32.5	32.5	54.5	10.3	30.1
Household schedules	33.1	33.1	54.4	11.1	33.1
Difference	−0.6	−0.6	0.1	−0.8	−3.0

Notes. Census report: data are from the *Census of England and Wales, 1911*. Volume 10. *Occupations and Industries*, part 1 (London, 1914), table 8, p. 75. *Household schedules*: data are from I-CeM (Schürer and Higgs, 2025) and based on the original `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 10 and over, while *Married* and *Widowed*, the female population aged 15 and over.

Overall, the comparison shows that figures derived from household schedules closely match those in the census report—the gap in overall female labor force participation is under one percentage point. The match is especially close for single women. As expected from the literature (e.g., Folbre, 1991; Humphries and Sarasúa, 2012; Burnette, 2025; Mol, 2026), the census report undercounts married women's participation, though only modestly (see, e.g., McKay, 1998; Anderson, 1999; Goose, 2007; Shaw-Taylor, 2007; McGeevor, 2014; Higgs and Wilkinson, 2016; You, 2020). By contrast, discrepancies are larger for widowed women, reaching about three percentage points.

We examine these patterns in greater detail in Table B2, which compares female labor force participation rates from both sources by age group and marital status, using all age categories published in the census report. For all marital statuses combined (Panel A), discrepancies are most pronounced among women aged 55 and over, where the census report undercounts participation by two to three percentage points. This undercount is evident across all marital groups but is particularly marked among single and widowed women (Panels B and D, respectively). Larger differences also appear for married and widowed women under 20 (Panels C and D), though these ages constitute less than 0.5 percent of their respective marital groups.

On the basis of these comparisons, we take the 1911 household schedule data at face value in our analysis—though see Footnote 3. Moreover, the discrepancies observed among the youngest and oldest women motivate restricting our main analysis to the core working-age population, defined here as ages 20 to 54.

Table B2. Female Labor Force Participation Rates in 1911 (%)
Detailed Comparisons between Census Reports and Household Schedules

Age group	10+/15+	10–14	15–19	20–24	25–34	35–44	45–54	55–64	65–74	75+
					A. All marital statuses					
Census report	32.5	10.4	68.8	62.0	33.8	24.1	23.0	20.4	13.8	5.7
Household schedules	33.1	11.0	68.0	61.3	34.0	24.7	24.3	22.7	17.0	8.7
Difference	−0.6	−0.6	0.8	0.7	−0.2	−0.6	−1.3	−2.3	−3.2	−3.0
					B. Single					
Census report	54.5	10.4	69.5	77.7	74.0	66.1	58.9	46.2	26.0	9.4
Household schedules	54.4	10.9	68.5	76.3	73.5	67.7	62.0	50.8	31.8	14.7
Difference	−0.0	−0.5	1.0	1.4	0.5	−1.6	−3.1	−4.6	−5.8	−5.3
					C. Married					
Census report	10.3		*13.4*	12.9	10.6	10.6	10.5	8.8	5.7	2.3
Household schedules	11.1		*20.8*	14.1	11.1	11.0	11.4	10.3	8.0	5.0
Difference	−0.8		*−7.4*	−1.2	−0.5	−0.4	−0.9	−1.5	−2.3	−2.7
					D. Widowed					
Census report	30.1		*39.3*	*59.2*	66.4	62.3	47.0	32.1	17.0	5.8
Household schedules	33.1		*55.6*	*59.5*	66.8	63.8	49.7	35.8	20.5	8.6
Difference	−3.0		*−16.3*	*−0.3*	−0.4	−1.5	−2.7	−3.7	−3.5	−2.8

Notes. Census report: data are from the *Census of England and Wales, 1911*. Volume 10. *Occupations and Industries*, part 1 (London, 1914), table 8, p. 75. *Household schedules*: data are from I-CeM (Schürer and Higgs, 2025) and based on the original `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 10 and over, while *Married* and *Widowed*, the female population aged 15 and over. Figures are italicized when the relevant age group accounts for less than 0.5 percent of the entire marital group.

B.2. Female Labor Force Participation in the 1921 Census

B.2.1. Comparisons with the Original I-CeM Measures

Table B3 compares female labor force participation rates based on the 1921 census, contrasting figures published in the census report with those derived from household schedules.[4] As for 1911, we classify women as labor force participants if they are recorded as working or unemployed by I-CeM's original `INACTIVE` variable. The table reports figures for all women aged 12 and over—the population used in the census report—as well as for women aged 14 to 59, which can be reconstructed from the age groups available in the report—this estimate also appears in Hakim (1980, table 2, p. 559) and Hatton (2014, table 4.5, p. 110).[5] We additionally provide labor force participation rates by marital status, combining widowed and divorced women as in the census report—divorced women represent only 0.4 percent of this combined category.

Table B3. Female Labor Force Participation Rates in 1921 (%)
Broad Comparisons between Census Reports and Household Schedules
Original I-CeM Measures

	Age group		Marital status		
Group	12+	15–59	Single	Married	Widowed
Census report	32.3	36.9	60.9	9.1	26.1
Household schedules	26.6	30.5	50.6	7.8	19.6
Difference	5.7	6.4	10.3	1.3	6.5

Notes. Census report: data are from the *Census of England and Wales, 1921. Occupation Tables* (London, 1924), table 4, p. 54. *Household schedules*: data are from I-CeM (Schürer and Wakelam, 2025) and based on the original `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 12 and over, while *Married* and *Widowed*, the female population aged 14 and over.

Overall, the comparison indicates that figures derived from household schedules using I-CeM's original `INACTIVE` variable *do not* align with those in the census report—the gap in overall female labor force participation rate is close to six percentage points. The discrepancy is especially large for single women, at about ten percentage points. Moreover, contrary to expectations from the literature (e.g., Folbre, 1991; Humphries and Sarasúa, 2012; Burnette, 2025; Mol, 2026), I-CeM's original measure implies that the census report substantially *overcounts* women's activity, though less so for married women.

[4]Census report data are from the *Census of England and Wales, 1921. Occupation Tables* (London, 1924), table 4, p. 54. Household schedule data are from I-CeM (Schürer and Wakelam, 2025).

[5]Both studies denote this group as ages *15*–59, though strictly speaking the census report combines ages 14 and 15.

Table B4. Female Labor Force Participation Rates in 1921 (%)
Detailed Comparisons between Census Reports and Household Schedules
Original I-CeM Measures

Age group	12+/14+	12–13	14–15	16–17	18–19	20–24	25–34	35–44	45–54	55–59	60–64	65–69	70+
						A. All marital statuses							
Census report	32.3	3.9	44.8	70.9	76.3	62.2	33.5	22.9	21.0	20.1	18.3	15.3	6.5
Household schedules	26.6	3.7	38.8	60.5	64.1	52.4	27.9	18.5	16.3	15.2	13.5	11.1	5.1
Difference	5.7	0.2	6.0	10.4	12.2	9.8	5.6	4.4	4.7	4.9	4.8	4.2	1.4
						B. Single							
Census report	60.9	3.9	44.8	71.0	78.9	80.5	76.3	68.2	60.4	52.9	43.9	33.8	13.1
Household schedules	50.6	3.7	38.8	60.7	66.3	67.9	63.1	54.2	45.8	39.3	32.3	24.9	11.8
Difference	10.3	0.2	6.0	10.3	12.6	12.6	13.2	14.0	14.6	13.6	11.6	8.9	1.3
						C. Married							
Census report	9.1		*40.3*	*17.8*	*15.0*	13.2	9.9	9.3	8.8	8.0	7.0	5.7	2.8
Household schedules	7.8		*22.1*	*17.8*	*14.5*	12.1	8.8	8.0	7.3	6.4	5.5	4.5	2.3
Difference	1.3		*18.2*	*0.0*	*0.5*	1.1	1.1	1.3	1.5	1.6	1.5	1.2	0.5
						D. Widowed							
Census report	26.1		*33.3*	*43.2*	*44.5*	*50.4*	46.8	45.8	41.1	33.1	26.2	19.3	6.5
Household schedules	19.6		*17.1*	*45.5*	*18.8*	*36.9*	36.7	35.0	30.8	24.5	19.2	13.9	4.9
Difference	6.5		*16.2*	*−2.3*	*25.7*	*13.5*	10.1	10.8	10.3	8.6	7.0	5.4	1.6

Notes. Census report: data are from the *Census of England and Wales, 1921. Occupation Tables* (London, 1924), table 4, p. 54. *Household schedules*: data are from I-CeM (Schürer and Wakelam, 2025) and based on the original `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 12 and over, while *Married* and *Widowed*, the female population aged 14 and over. Figures are italicized when the relevant age group accounts for less than 0.5 percent of the entire marital group.

We examine these patterns in greater detail in Table B4, which compares female labor force participation rates from both sources by age group and marital status, using all age categories published in the census report. For all marital statuses combined (Panel A), discrepancies are most pronounced among younger women aged 16 to 24, where the census report overcounts participation by 10 to 12 percentage points. This overcount is evident across all marital groups but is especially marked among single and widowed women (Panels B and D, respectively). By contrast, discrepancies for married women are small, consistently below two percentage points (Panel C).

On the basis of these comparisons, we do not take the 1921 household schedule data at face value in our analysis. The classification logic of occupations in 1921 differed substantially from that of 1911. To ensure comparability, I-CeM constructed an additional variable (`HOLLERNOS`) that maps the 1921 census to the `OCCODE` scheme developed for 1911, which underlies the `INACTIVE` variable.[6] These transformations, however, introduced several inaccuracies into the 1921 `INACTIVE` variable, which we address next.

B.2.2. Corrections to I-CeM's 1921 `INACTIVE` Variable

We correct I-CeM's 1921 `INACTIVE` variable by leveraging the raw occupation (`OCC`) and employer (`EMPLOYER`) strings recorded in the 1921 household schedules.[7] We then implement ten types of corrections for the female population aged 10 and over.[8]

1. **Erroneously classified as `INACTIVE == 1` (Retired)**

   We reclassify as *occupied* those women originally coded as retired whose occupation or employer strings do not contain `RETIRED` or close variants.[9] This adjustment moves about 212,000 women into the occupied category—75 percent of those originally classified as retired in 1921—including 93,000 clerks, 83,000 cooks, and 20,000 nurses.[10]

2. **Erroneously classified as `INACTIVE == 2` (Late/formerly occupied)**

[6]For details on these coding procedures, see the Cambridge Group for the History of Population and Social Structure: https://www.campop.geog.cam.ac.uk/research/projects/icem/construction.html (accessed November 2025). See also the discussion of limitations in the coding of occupations in the 1911 and 1921 censuses: https://www.campop.geog.cam.ac.uk/research/projects/icem/limitations.html (accessed November 2025).

[7]The `INDUSTRY` and `EMPLOY` strings are not available for the 1921 census. For details, see https://www.campop.geog.cam.ac.uk/research/projects/icem/occupation.html (accessed November 2025).

[8]Individuals classified as `INACTIVE == 3` (Unemployed, not currently working but with specified occupation), `INACTIVE == 6` (Living on own means, independent), or `INACTIVE == 9` (Other without a specified occupation) are correctly classified.

[9]Specifically, we require that these strings contain none of the following 18 terms: `RETIRED`, `RETRIED`, `RETIERD`, `RETIRD`, `RETD`, `RETIRE`, `REITRED`, `RETIERED`, `RETIED`, `RETRED`, `RETAIRED`, `RITIRED`, `RETTIRED`, `PENSIONER`, `PREVIOUSLY`, `FORMERLY`, `PAST WORK`.

[10]A small number of those reclassified were in fact unemployed. We address this in our final correction.

We reclassify as *occupied* those women originally coded as late or formerly occupied whose occupation or employer strings do not contain the terms `FORMERLY`, `PREVIOUSLY`, `LATE`, or close variants.[11] This adjustment moves about 146,000 women into the occupied category—83 percent of those originally classified as late or formerly occupied in 1921—including 29,000 domestic servants, 20,000 teachers, 14,000 barmaids, and 13,000 cashiers.[12]

3. **Erroneously classified as `INACTIVE == 4` (Pensioner, annuitant, or superannuated)**

   We reclassify as *occupied* those women originally coded as pensioner, annuitant, or superannuated whose occupation or employer strings do not contain the terms `PENSION`, `ANNUIT`, `RETIRED`, `FORMERLY`, or close variants, or who worked in pension administration.[13] This adjustment moves about 8,000 women into the occupied category—23 percent of those originally classified as pensioner, annuitant, or superannuated in 1921—nearly all of whom are performing artists.

4. **Erroneously classified as `INACTIVE == 5` (Pauper, in receipt of alms)**

   We reclassify as *occupied* those women originally coded as pauper whose occupation or employer strings do not contain the terms `ASSIST`, `HELP`, `POOR`, or close variants.[14] This adjustment moves 355 women to the occupied category—20 percent of those initially classified as pauper in 1921.

5. **Erroneously classified as `INACTIVE == 7` (Others with description suggesting not currently working) with `OCCODE < 772`**

   We reclassify as *occupied* those women originally coded as not working in this category (`INACTIVE == 7`) whose occupation or employer strings suggest an occupation *and* whose occupational code also indicates an occupation (`OCCODE < 772`).[15] We do

[11] Specifically, we require that these strings include none of the following 24 terms: `FORMER`, `FORMLY`, `FORMLEY`, `FOMERLY`, `FORMALY`, `FORMELRY`, `FORMELY`, `FOERMERLY`, `FOEMERLY`, `FOMERLEY`, `FOREMERLY`, `FORMARLY`, `FORMARLEY`, `FOREMARLY`, `FORMELEY`, `FORMLERY`, `PREVIOUS`, `PREVIOUSLY`, as well as `RETIRED`, `LATE `, `LATE,`, `LATE)`, and `LATE-`; while excluding spurious matches such as `CHOCOLATE`, `SLATE`, `PLATE`, `PERFORMER`, and close variants.

[12] Again, a small share of those reclassified were in fact unemployed. We address this in our final correction.

[13] Specifically, we require that these strings include none of the following 27 terms: `PENSION`, `PENSON`, `O A P`, `OAP`, `ANNUIT`, `ANNUAT`, `OLD AGE`, `RETIRED`, `RETRIED`, `FORMERLY`, `FORMLEY`, as well as `UNEMP`, `NOT OCCUP`, `OUT OF WORK`, `OUT OF EMP`, `NOT OUT AT WORK`, `NOT WORK`, `NO WORK`, `NOT EMP`, `NO EMP`, `UNMPLOYED`, `DISABLED`, `RESTING`, `DISENGAG`, `DISCUGAG`, `DIS ENGAGED`, and `DESENGAGED`; while excluding spurious matches such as `MINISTRY`, `COMMISSION`, `COMMITTEE`, `SOAP`, `ROAP`, and close variants.

[14] Specifically, we require that these strings include none of the following 31 terms: `ASSIST`, `HELP`, `POOR`, `PAUP`, `PUPER`, `PATIENT`, `PATENT`, `HOUSE`, `HOME`, `HOOK`, `RETIRED`, `SPINISTER`, `GUEST`, `CRIPPLE`, `BLIND`, `DEPENDANT`, `DEPENDENT`, `DEPENDING`, `HOLIDAY`, `CHARITY`, `GONE`, `NOT OCCUPIED`, `NONE`, `CRIP`, `INVALID`, `NO OCC`, `ALLOW`, `PENSION`, `PERSION`, and `OUT OF WORK`.

[15] Women classified as `INACTIVE == 7` and with `OCCODE >= 772` are correctly classified as not occupied.

not, however, consider as occupied women whose occupation or employer strings contain the terms `HOME`, `HOUSEHOLD`, `HELP`, `ASSISTING`, or related terms.[16] This adjustment moves about 101,000 women to the occupied category—59 percent of those originally in this category with an occupational code `OCCODE < 772` in 1921—nearly all of whom are domestic servants.

6. **Erroneously classified as `INACTIVE == 999999` with `OCCODE < 772`**

   We reclassify as *occupied* those women originally coded as not working in this category whose occupation or employer strings suggest an occupation *and* whose occupational code also indicates an occupation (`OCCODE < 772`).[17] Here, we do not consider as occupied women whose occupation or employer strings contain the terms `DUTIES`, `WIFE`, `BLIND`, or related terms.[18] This adjustment moves about 130,000 women to the occupied category—91 percent of those originally in this category with an occupational code `OCCODE < 772` in 1921—including 19,000 confectioners and 6,000 domestic servants.

7. **Erroneously classified as `INACTIVE == 999999` with `OCCODE == 999999`**

   Occupation strings are uninformative in these cases, as they typically indicate household rather than employment status. We therefore rely on the employer string, classifying women as *occupied* whenever an employer is listed, except when the employer field clearly denotes non-employment, such as `HOME DUTIES` and 300 other strings. Using this approach, we move about 109,000 women to the occupied category—4 percent of those originally classified as inactive in this group with a missing occupational code in 1921.

8. **Erroneously classified as `INACTIVE == 0` (Working)**

   We reclassify women who are erroneously coded as *occupied*—we do so for both censuses, though errors are marginal in the 1911 census. We implement corrections across four categories: retired, formerly occupied, unemployed, and pensioners. Overall, these procedures move about 6,000 women out of the occupied category in 1911 and 277,000 in 1921.

[16] Specifically, we require that these strings include none of the following 25 terms: `HOME`, `HOUSE W`, `HOUSEW`, `HOSUE W`, `HOUSEHOLD`, `NO WORK`, `NOT EMP`, `OUT OF`, `UNEMPLOYED`, `NO OCC`, `NOT OCC`, `INVALID`, `RETIRED`, `FORMERLY`, `PREVIOUSLY`, `INMATE`, `STUD`, `HELP`, `ASSIST`, `ASSIT`, `DUTIE`, `DUTY`, `WIFE`, `WIDOW`, and `WIDDOW`.

[17] Women classified as `INACTIVE == 999999` and with `OCCODE >= 772 & != 999999` are correctly classified as not occupied.

[18] Specifically, we require that these strings include none of the following 24 terms: `DUTIE`, `DUTY`, `WIFE`, `WIDOW`, `PARTNER`, `SINGLE`, `COMPANION`, `BLIND`, `UNFIT`, `SICK`, `INVALID`, `TRAIN`, `STUD`, `OUT OF`, `UNEMP`, `NOT WORK`, `NO OCC`, `NOT OCC`, `NOT EMP`, `NO EMP`, `RETIRED`, `PREVIOUSLY`, `FORMERLY`, and `OLD AGE`.

i. **Retired (`INACTIVE == 1`)**

We reclassify as *retired* those women originally coded as working whose occupation or employer strings contain the term `RETIRED` or close variants.[19] This adjustment moves about 300 women to the retired category in the 1911 census and 5,000 in the 1921 census.

ii. **Late/formerly occupied (`INACTIVE == 2`)**

We reclassify as *late or formerly occupied* those women originally coded as occupied whose occupation or employer strings contain the terms `FORMER`, `PREVIOUS`, `LATE`, or close variants.[20] This adjustment moves about 1,000 women to the late or formerly occupied category in the 1911 census and 2,000 in the 1921 census.

iii. **Unemployed (`INACTIVE == 3`)**

We reclassify as *unemployed* those women originally coded as occupied whose occupation or employer strings contain the terms `UNEMP`, `OUT OF`, or close variants.[21] This adjustment moves about 3,000 women to the unemployed category in the 1911 census and 263,000 in the 1921 census.

iv. **Pensioner (`INACTIVE == 4`)**

We reclassify as *pensioner* those women originally coded as occupied whose occupation or employer strings contain the terms `PENSION`, `OAP`, `ANNUIT`, or close variants.[22] This adjustment moves about 1,000 women to the pensioner category in both the 1911 and 1921 censuses.

9. **Consistency with the 1911 occupational classification**

Women with certain occupational strings were systematically classified as occupied in the I-CeM data in the 1911 census, but not in the 1921 census—these occupations relate to domestic services such as housekeepers, housemaids, and domestic servants. To ensure comparability across the two censuses, we reclassify approximately 101,000 women with these occupational strings as *occupied* in 1921.

[19] Specifically, we require that these strings include one of the following 14 terms: `RETIRED`, `RETRIED`, `RETIERD`, `RETIRD`, `RETD`, `RETIRE`, `REITRED`, `RETIERED`, `RETIED`, `RETRED`, `RITIRED`, `RETTIRED`, or `RETAIRED`.

[20] Specifically, we require that these strings include one of the following 24 terms: `FORMER`, `FORMLY`, `FORMLEY`, `FOMERLY`, `FORMALY`, `FORMELRY`, `FORMELY`, `FOERMERLY`, `FOEMERLY`, `FOMERLEY`, `FOREMERLY`, `FORMARLY`, `FORMARLEY`, `FOREMARLY`, `FORMELEY`, `FORMLERY`, `PREVIOUS`, `PREVIOUSLY`, `LATE `, `LATE,`, `LATE)`, or `LATE-`; while excluding spurious matches such as `CHOCOLATE`, `SLATE`, `PLATE`, `PERFORMER`, and close variants.

[21] Specifically, we require that these strings include one of the following 7 terms: `OUT OF`, `UNEMP`, `NO WORK`, `NOT WORK`, `UN EMP`, `NOT EMP`, or `ENEMP`.

[22] Specifically, we require that these strings include one of the following 7 terms: `PENSION`, `PENSON`, `O A P`, `OAP`, `ANNUIT`, `ANNUAT`, or `OLD AGE`; while excluding spurious matches such as `MINISTRY`, `COMMISSION`, `COMMITTEE`, `SOAP`, `ROAP`, and close variants.

10. **Final check using employers' addresses and surnames**

    Finally, for remaining women in 1921 not classified as working, retired, formerly occupied, pensioned, or unemployed, but who nonetheless have an occupational code that indicates an occupation (`OCCODE < 772`) *and* have an employer listed, we reclassify them as *occupied* if their employer string does not include the woman's own surname or address, and does not clearly denote non-employment, using the same 300 strings noted in step 7. This adjustment moves about 23,000 women into the occupied category—31 percent of women in this case—including 14,000 domestic servants.

B.2.3. Comparisons with the Corrected I-CeM Measures

Table B5 compares female labor force participation rates based on the 1921 census, contrasting figures published in the census report with those derived from the corrected I-CeM measures. Overall, the corrected figures align much more closely with the census report than the original I-CeM variable reported in Table B3—the gap in overall female participation falls from six percentage points to less than half a percentage point. The corrections bring the measure especially close for married and widowed women, while discrepancies remain slightly larger for single women—under two percentage points, compared with eleven under the original measure. The detailed patterns reported in Table B6 confirm the superiority of our corrected measure.

Table B5. Female Labor Force Participation Rates in 1921 (%)
Broad Comparisons between Census Reports and Household Schedules
Corrected I-CeM Measures

	Age group		Marital status		
Group	12+	15–59	Single	Married	Widowed
Census report	32.3	36.9	60.9	9.1	26.1
Household schedules	32.0	36.4	59.4	10.1	25.6
Difference	0.3	0.5	1.5	−1.0	0.5

Notes. Census report: data are from the *Census of England and Wales, 1921. Occupation Tables* (London, 1924), table 4, p. 54. *Household schedules*: data are from I-CeM (Schürer and Wakelam, 2025) and based on the corrected `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 12 and over, while *Married* and *Widowed*, the female population aged 14 and over.

Table B6. Female Labor Force Participation Rates in 1921 (%)
Detailed Comparisons between Census Reports and Household Schedules
Corrected I-CeM Measures

Age group	12+/14+	12–13	14–15	16–17	18–19	20–24	25–34	35–44	45–54	55–59	60–64	65–69	70+
						A. All marital statuses							
Census report	32.3	3.9	44.8	70.9	76.3	62.2	33.5	22.9	21.0	20.1	18.3	15.3	6.5
Household schedules	32.0	4.5	43.9	68.8	73.5	60.5	33.3	23.2	21.3	20.3	18.3	15.4	7.3
Difference	0.3	−0.6	0.9	2.1	2.8	1.7	0.2	−0.3	−0.3	−0.2	−0.0	−0.1	−0.8
						B. Single							
Census report	60.9	3.9	44.8	71.0	78.9	80.5	76.3	68.2	60.4	52.9	43.9	33.8	13.1
Household schedules	59.4	4.5	43.9	69.0	76.0	78.2	74.5	66.8	59.1	51.6	43.1	33.6	15.8
Difference	1.5	−0.6	0.9	2.0	2.9	2.3	1.8	1.4	1.3	1.3	0.8	0.2	−2.7
						C. Married							
Census report	9.1		*40.3*	*17.8*	*15.0*	13.2	9.9	9.3	8.8	8.0	7.0	5.7	2.8
Household schedules	10.1		*27.3*	*20.9*	*17.2*	14.4	10.8	10.2	9.9	9.0	7.9	6.6	3.8
Difference	−1.0		*13.0*	*−3.1*	−2.2	−1.2	−0.9	−0.9	−1.1	−1.0	−0.9	−0.9	−1.0
						D. Widowed							
Census report	26.1		*33.3*	*43.2*	*44.5*	*50.4*	46.8	45.8	41.1	33.1	26.2	19.3	6.5
Household schedules	25.6		*24.4*	*63.6*	*24.1*	*45.5*	45.0	44.0	40.0	32.3	25.5	18.8	7.0
Difference	0.5		*8.9*	*−20.4*	*20.4*	*4.9*	1.8	1.8	1.1	0.8	0.7	0.5	−0.5

Notes. *Census report*: data are from the *Census of England and Wales, 1921. Occupation Tables* (London, 1924), table 4, p. 54. *Household schedules*: data are from I-CeM (Schürer and Wakelam, 2025) and based on the corrected `INACTIVE` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. *Single* concerns the female population aged 12 and over, while *Married* and *Widowed*, the female population aged 14 and over. Figures are italicized when the relevant age group accounts for less than 0.5 percent of the entire marital group.

C. Female Marital Statuses in the Censuses

This section compares the distribution of female marital statuses based on census reports and household schedules. Section C.1 focuses on the 1911 census, and Section C.2, on the 1921 census.

C.1. Female Marital Statuses in the 1911 Census

Table C1 compares the distribution of marital statuses among women aged 20 to 54—our sample of interest—based on the 1911 census, contrasting figures published in the census report with those derived from household schedules.[23] The 1911 census report tabulates the distribution of women by age across three marital groups: unmarried, married, and widowed. By contrast, I-CeM's `MAR` variable—derived from the transcribed marital status field on the household schedule (`COND`)—distinguishes six categories: single (`1`), married (`2`), married, spouse absent (`3`), widowed (`4`), divorced (`5`), and not recorded or unknown (`9`). To align these two sources, we recode household schedule data by pooling married women regardless of the status of their spouse, excluding women whose marital status is unknown, and classifying divorced women as widowed as in the census report—divorced women represent only 0.1 percent of this combined category.

Table C1. Female Marital Statuses in 1911 (%)
Comparisons between Census Reports and Household Schedules

Marital status	Unmarried	Married	Widowed
Census report	34.6	60.9	4.5
Household schedules	34.3	61.3	4.4
Difference	0.2	–0.4	0.2

Notes. Census report: data are from the *Census of England and Wales, 1911. Summary Tables* (London, 1915), table 29, pp. 90–91. *Household schedules*: data are from I-CeM (Schürer and Higgs, 2025) and based on the original `MAR` variable. *Married* includes married women whose spouse is noted absent in the household schedules. *Widowed* includes widowed and divorced women. The sample comprises women aged 20 to 54 and excludes women whose marital status is unknown in the household schedules.

The comparison shows that figures derived from household schedules closely match those in the census report, although the gap for married women is 0.4 percentage points. This discrepancy likely reflects the nontrivial share of women whose marital status is unknown in household schedule data—they represent 0.4 percent of women aged 20 to

[23] Census report data are from the *Census of England and Wales, 1911. Summary Tables* (London, 1915), table 29, pp. 90–91. Household schedule data are from I-CeM (Schürer and Higgs, 2025).

54.[24]

C.2. Female Marital Statuses in the 1921 Census

Table C2 compares the distribution of marital statuses among women aged 20 to 54 based on the 1921 census, contrasting figures published in the census report with those derived from household schedules.[25] The 1921 census report tabulates the distribution of women by age across four marital groups: single, married, widowed, and divorced. To align this classification with I-CeM's, we recode the household schedule data by pooling married women regardless of the status of their spouse and excluding women whose marital status is unknown.

Table C2. Female Marital Statuses in 1921 (%)
Comparisons between Census Reports and Household Schedules

Marital status	Single	Married	Widowed	Divorced
Census report	32.2	62.3	5.4	0.1
Household schedules	31.9	62.6	5.4	0.1
Difference	0.3	−0.3	0.0	0.0

Notes. Census report: data are from the *Census of England and Wales, 1921. General Report with Appendices* (London, 1927), table 39, p. 81. *Household schedules*: data are from I-CeM (Schürer and Wakelam, 2025) and based on the original `MAR` variable. *Married* includes married women whose spouse is noted absent in the household schedules. The sample comprises women aged 20 to 54 and excludes women whose marital status is unknown in the household schedules.

The comparison shows that figures derived from household schedules again closely match those in the census report, with a gap for single and married women of 0.3 percentage points. This discrepancy likely reflects the nontrivial share of women whose marital status is unknown in household schedule data—they represent 0.4 percent of women aged 20 to 54.[26]

[24]Assigning all women whose marital status is unknown to the unmarried category would eliminate the remaining gap between figures from the census report and household schedules.

[25]Census report data are from the *Census of England and Wales, 1921. General Report with Appendices* (London, 1927), table 39, p. 81. Household schedule data are from I-CeM (Schürer and Wakelam, 2025).

[26]Assigning all women whose marital status is unknown to the single category would again eliminate the remaining gap between figures from the census report and household schedules.

D. Consistent 1911–21 Local Government Districts

D.1. Local Government Districts as the Unit of Aggregation

To quantify the local exposure to the war's demographic shock, we compute military death rates at the level of local government districts (hereafter, districts). We choose this level of aggregation for four reasons (Gregory, Dorling and Southall, 2001, pp. 298–99). First, it aligns our measure with labor statistics provided in the 1911 and 1921 census reports, which both use districts as their level of aggregation—although the 1911 census does not report such statistics separately for rural districts, as well as municipal boroughs and urban districts with fewer than 5,000 inhabitants. Second, districts are the closest historical analogue to current travel-to-work areas: cities were rarely split across multiple districts, so these districts tend to coincide with internally consistent local labor markets. Third, districts were explicitly classified as rural or urban, which yields greater within-district homogeneity of economic activity than alternative units such as Registration or Poor Law districts, which were not explicitly divided along rural or urban lines. Finally, during the war, appeals for exemption from conscription were heard by tribunals established at the district level, generating variation across districts in the likelihood that men served—and, in turn, in their probability of dying in the conflict.

Local government districts were known as Sanitary districts until 1894. These units comprised several types of local authorities with slightly different powers. The vast majority were designated either urban or rural, but the system also included county, municipal, or metropolitan boroughs—urban districts which retained additional municipal privileges. In practice, their powers related primarily to sanitation and transportation.[27]

D.2. The Making of Consistent 1911–21 Local Government Districts

We obtain the geometries of the 1911 and 1921 local government districts from the *Great Britain Historical Database* (Southall, Aucott and Burton, 2024).[28] These shapefiles identify districts consistently over time through a unique `g_unit` identifier, which tracks the same district across years even when its name, boundaries, or administrative status changed. According to these shapefiles, England and Wales included 1,836 districts in 1911 and 1,824 in 1921, of which 191 (10 percent) were located in Wales.[29] The vast majority

[27] For additional details on the different levels of aggregation reported in the censuses of England and Wales, see, e.g., Gregory and Southall (2000) and Gregory, Dorling and Southall (2001). For a contemporary account, see Part II of the 1894 Local Government Act.

[28] Specifically, we use the raw shapefiles `ukds_ew1911_lgdistricts2` and `ukds_ew1921_lgdistricts2`.

[29] Both shapefiles include a unit assimilated to a local government district that corresponds to "[a]reas of common that were a combination of land held by several different parishes and which have never been

Table D1. Administrative Statuses of Local Government Districts (1911–21)

	1911		1921	
District administrative status	Count	Share	Count	Share
Rural district	675	0.368	673	0.369
Urban district	808	0.440	792	0.434
County borough	75	0.041	82	0.045
Municipal borough	248	0.135	247	0.135
Metropolitan borough	28	0.015	28	0.015
London county corporate	1	0.001	1	0.001
No administrative status	1	0.001	1	0.001
Total	1,836	1.000	1,824	1.000

Notes. This table reports the administrative statuses of local government districts in 1911 and 1921 according to Southall, Aucott and Burton's (2024) shapefiles. *No administrative status*: see Footnote 29 for details.

of districts were rural or urban districts, accounting for 37 and 44 percent of all districts, respectively (Table D1). The remaining units were municipal boroughs (14 percent), county boroughs (4 percent), or metropolitan boroughs (2 percent). London formed a separate county corporate. Figure D1 displays the spatial distribution of administrative statuses of districts in 1911—the 1921 distribution is similar.

The geography of local government districts was broadly stable between 1911 and 1921. In total, 1,798 districts can be linked across the two shapefiles using their `g_unit` identifier, representing 98 percent of 1911 districts and 99 percent of 1921 districts. Only 38 districts present in 1911 are no longer observed in 1921, while 26 observed in 1921 were created after 1911. As shown in Table D2, among districts that existed both in 1911 and 1921, 1,627 (90 percent) did not experience change of any type, while 171 (10 percent) experienced a change of some kind (name, status, or boundary). Of direct interest to us, 160 (9 percent) experienced some boundary change—while 1,638 did not (= 1,798 − 160). Combining these with the ones that disappeared or were created over the period results in our creation of 83 "super local government districts" whose boundaries remained constant between 1911 and 1921. We thus base our analysis on a consistent set of 1,721 1911–21 districts (= 1,638 + 83).

recognized as separate administrative units" (Southall, Aucott and Burton, 2024, Study Guides for 1911 and 1921, p. 2). This unit in North Riding is recorded as "intermixed lands (common)."

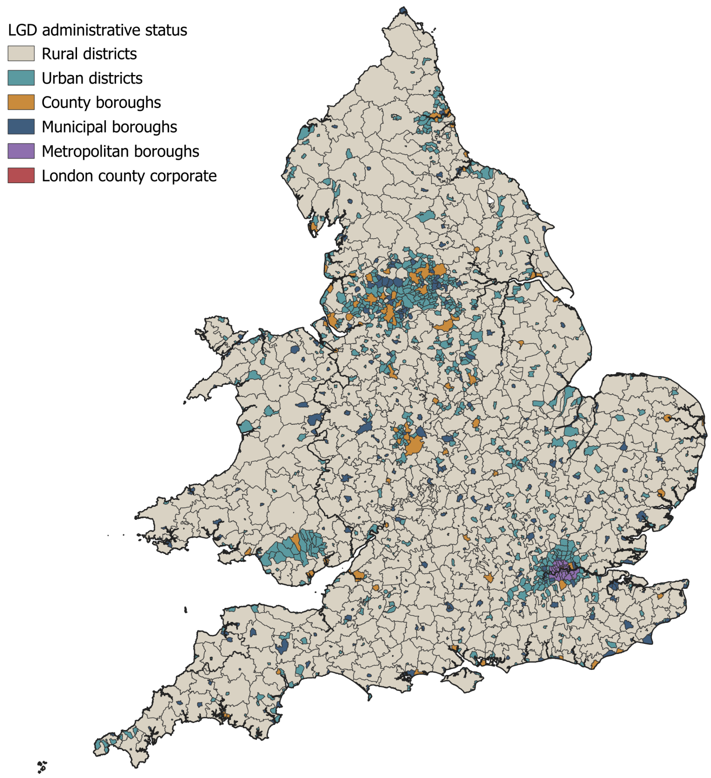


Figure D1. Administrative Statuses of Local Government Districts (1911)

Notes. This figure maps the administrative statuses of local government districts in 1911. The base shapefile is from Southall, Aucott and Burton (2024).

Table D2. Types of Changes to Local Government Districts (1911–21)

	Count	Share
Name change	6	0.003
Administrative status change	16	0.009
Boundary change	160	0.089
Any change	171	0.095
No change	1,627	0.905
Total	1,798	1.000

Notes. This table reports the types of changes to local government districts among those that existed both in 1911 and 1921, according to Southall, Aucott and Burton's (2024) shapefiles.

E. Linking Military Fatality Records to CWGC Records

E.1. Linking SDGW Records to CWGC Records

The numerator of our baseline measure of military death rates relies on SDGW records. However, these records may include soldiers who were listed as missing—and thus presumed dead—when the roll was compiled but who ultimately survived.[30] To ensure that our measure captures soldiers who actually died during the conflict, we link the 661,827 SDGW records to the 835,740 records of the Commonwealth War Graves Commission (CWGC).

The linking proceeds through 25 linking steps based on different combinations of the following fields: last name, first name, initials, date of death, service number, and regiment.[31] At each step, we restrict both datasets to records with non-missing values for all linking fields used in that step and retain only those records that are unique with respect to that combination of fields. We then perform an exact match on the selected fields, remove matched records from both datasets, and carry the remaining unmatched observations forward to the next step. Throughout, we give priority to matches based on name variables.

The results of the linking process are reported in Table E1. Starting from 661,827 SDGW records, we match 659,412 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. Most matches are obtained in the early stages of the procedure, where richer combinations of identifying variables are available, while later steps yield smaller but still non-negligible numbers of additional matches.

[30] The preface to SDGW volumes states that "[t]he rolls of missing, and missing believed killed, contain the names of those whose cases were or are held in abeyance in deference to the wishes of their next of kin." In addition, the Army Record Offices compiled SDGW records during a period of downsizing and heavy administrative pressure, as they were simultaneously processing mass demobilization files, supplying documentation for war-related pension claims, and issuing medals (Nash, 2018*b*). On the verification process of SDGW records against battalion ledgers by the Army Record Offices, see Nash (2018*a*).

[31] The raw CWGC records include a field for first-name initials, whereas the raw SDGW records do not. We therefore construct two initials variables for the SDGW records: a "simple" initials variable, defined as the first letter of the first given name, and a "full" initials variable, defined as the first letters of all first names. We then implement the linking in four substeps. First, we match the raw initials field in the CWGC records to the "simple" initials variable in the SDGW records. Next, we match it to the corresponding "full" initials variable. We then repeat the same two substeps using analogous "simple" and "full" initials variables constructed from the first-name field in the CWGC records, since the two CWGC fields do not always coincide.

Table E1. SDGW–CWGC Records Linking Steps

	Linking fields						Matches		Cumulative	
Step	Last name	First name	Initials	Death date	Service number	Regiment	Count	Share	Count	Share
1	X	X		X	X	X	355,299	0.537	355,299	0.537
2	X		X	X	X	X	201,950	0.305	557,249	0.842
3	X			X	X	X	7,560	0.011	564,809	0.853
4	X	X		X		X	20,704	0.031	585,513	0.885
5	X		X	X		X	19,947	0.030	605,460	0.915
6	X			X		X	1,067	0.002	606,527	0.916
7	X	X			X	X	10,746	0.016	617,273	0.933
8	X		X		X	X	7,850	0.012	625,123	0.945
9	X				X	X	334	0.001	625,457	0.945
10	X	X		X	X		1,028	0.002	626,485	0.947
11	X		X	X	X		858	0.001	627,343	0.948
12	X			X	X		74	0.000	627,417	0.948
13	X	X		X			2,056	0.003	629,473	0.951
14	X		X	X			3,096	0.005	632,569	0.956
15	X			X			724	0.001	633,293	0.957
16	X	X			X		40	0.000	633,333	0.957
17	X		X		X		42	0.000	633,375	0.957
18	X				X		8	0.000	633,383	0.957
19	X	X				X	1,023	0.002	634,406	0.959
20	X		X			X	1,429	0.002	635,835	0.961
21	X					X	1,293	0.002	637,128	0.963
22				X	X	X	20,453	0.031	657,581	0.994
23					X	X	901	0.001	658,482	0.995
24				X	X		24	0.000	658,506	0.995
25				X		X	906	0.001	659,412	0.996

Notes. This table reports matches between SDGW and CWGC records across successive linking steps, starting from an initial set of 661,827 SDGW records.

E.2. Linking RNWGR Records to CWGC Records

We also link RNWGR records to CWGC records. As with the SDGW–CWGC linking, this serves both to exclude servicemen who were reported missing but did not die and to facilitate subsequent linkage to the 1911 census using information available only in CWGC records, such as parents' names. Because RNWGR does not report places of residence at enlistment, these records are excluded from our baseline analysis. They often report place of birth, however, and many can be linked to the 1911 census, enabling us to incorporate them in robustness checks that construct military death rates using either place of birth or place of residence in 1911.[32]

The linking of RNWGR to CWGC records follows the same sequence of steps as the SDGW–CWGC linking described above, with two modifications: steps relying on service numbers are omitted because this field is unavailable in RNWGR, and ship replaces regiment as the corresponding military-unit field. The results of this linking process are reported in Table E2. Of the 34,874 RNWGR records, we match 33,457 to a CWGC record—a linking rate of 96.0 percent.

Table E2. RNWGR–CWGC Records Linking Steps

	Linking fields					Matches		Cumulative	
Step	Last name	First name	Initials	Death date	Ship	Count	Share	Count	Share
1	X	X		X	X	16,074	0.461	16,074	0.461
2	X		X	X	X	4,237	0.122	20,311	0.583
3	X			X	X	140	0.004	20,451	0.587
4	X	X		X		8,762	0.251	29,213	0.838
5	X		X	X		3,230	0.093	32,443	0.931
6	X			X		122	0.003	32,565	0.934
7	X	X			X	410	0.012	32,975	0.946
8	X		X		X	91	0.003	33,066	0.949
9	X				X	8	0.000	33,074	0.949
10				X	X	383	0.011	33,457	0.960

Notes. This table reports matches between RNWGR and CWGC records across successive linking steps, starting from an initial set of 34,874 RNWGR records.

[32] See Appendix G.2 for a detailed description of the construction of these alternative death rates.

F. Estimating the Number of British Armed Forces Fatalities in England and Wales

F.1. Number of British Armed Forces Fatalities

The total number of military fatalities in the British armed forces remains contested. For instance, Winter (1976, table 1, p. 540) lists more than 30 distinct estimates drawn from different sources, ranging from 550,000 to 1,184,000 deaths. Among these, Winter (1976, p. 541) argues that "the most reliable treatment of British war casualties is the [1921] *General Annual Reports of the British Army* for the war years" as "figures therein were checked against the official regimental lists of war dead." On this basis, he concludes that the War Office's (1921) *General Annual Reports* provides the "best estimate of total deaths of British servicemen in the First World War [at] 722,785."[33]

In practice, we show in Table F1 that Winter's aggregate of 722,785 corresponds to the sum of the 6,166 Air Force fatalities (Jones, 1937, p. 160), the 34,654 Royal Navy and 8,590 Royal Naval Division fatalities (Newbolt, 1931, p. 434), and the 673,375 fatalities of the Regular Army and Territorial Force reported in the *General Annual Reports* (War Office, 1921, table 1G, pp. 71–72). While the figures reported by Jones (1937) and Newbolt (1931) appear reliable—for instance, the RNWGR contain 34,874 Royal Navy records—the total of 673,375 for the Regular Army and Territorial Force seems an underestimate. Indeed, our matched SDGW–CWGC records amount to 700,884 fatalities once ODGW–CWGC records are included (= 659,412 + 41,472).

By contrast, the "official number of British war dead" published by Parliament in 1921 (Watt, 2019, p. 77; see House of Commons Debates, *141*, May 4, 1921, cols. 1033–4) implies a total of 695,292 for the Regular Army and Territorial Force after subtracting fatalities of the Air Force, the Royal Navy, and the Royal Naval Division—much closer to our total (= 744,702 – 6,166 – 34,654 – 8,590). In line with Watt (2019, p. 78), we therefore treat 744,702 as the most credible aggregate estimate of British military fatalities.[34]

[33] Winter (1976, pp. 541–42) notes that this total includes "over 41,000 army officers and men [who] were killed between 1 October 1918 and 30 September 1919," as well as 10,000 fatalities among British soldiers attributed to the 1918–19 influenza pandemic.

[34] After subtracting fatalities in the Royal Naval Division, the War Office's (1922, table ia, p. 237) *Statistics of the Military Effort* reports 693,820 fatalities for the Regular Army and Territorial Force, which is likewise close to our benchmark. We nonetheless prefer the *Hansard* estimate, as it is closer to that benchmark of 700,884 fatalities by 1,472. By contrast, Carozzi, Pinchbeck and Repetto (2026, Appendix A.3, p. 4) and Tchaouchev (2026, p. 20) cite the *Statistics* total of 702,410 as a benchmark for the total number of British fatalities, although that figure excludes the 6,166 Air Force fatalities and the 34,654 Royal Navy fatalities.

Table F1. Military Fatalities by Source and Coverage

Source	Branch coverage	Officers	Other ranks	Total
A. Official aggregate benchmarks				
General Annual Reports	Army + TF	37,484	635,891	673,375
Hansard	British armed forces	43,480	701,222	744,702
	Army + TF	*35,964*	*659,328*	*695,292*
Statistics of the Military Effort	Army + TF + RND	37,452	664,958	702,410
	Army + TF	*36,989*	*656,831*	*693,820*
B. Service-specific official counts				
Jones (1937)	Air Force	4,579	1,587	6,166
Newbolt (1931)	Royal Navy	2,474	32,180	34,654
	RND	463	8,127	8,590
C. Individual-level records				
SDGW–CWGC	Army + TF		659,412	659,412
ODGW–CWGC	Army + TF	41,472		41,472
RNWGR–CWGC	Royal Navy			34,457

Notes. This table reports aggregate military fatalities across sources and branches of the British military. *Army* refers to the Old and New Regular Armies, *TF*, to the Territorial Force, and *RND*, to the Royal Naval Division. *British armed forces* comprises the Regular Army, Territorial Force, Air Force, Royal Navy, and Royal Naval Division. *SDGW–CWGC* refers to records from Soldiers Died in the Great War, *ODGW–CWGC*, to records from Officers Died in the Great War, and *RNWGR–CWGC*, to the Royal Navy and Royal Marine War Graves Roll, all matched to the Commonwealth War Graves Commission database. Italicized figures are derived by subtracting the relevant Air Force, Royal Navy, and Royal Naval Division totals from broader aggregate counts. Sources for the official aggregate benchmarks in Panel A are the War Office's (1921, pp. 62–72) *General Annual Reports of the British Army*, House of Commons Debates (Hansard, *141*, 1921, cols. 1033–34), and the War Office's (1922, p. 237) *Statistics of the Military Effort of the British Empire During the Great War*.

F.2. Number of Fatalities Residing in England and Wales

Given that about 100,000 of all British military fatalities were from Scotland (Watt, 2019, pp. 83–87) and 30,000 were from Ireland (Fitzpatrick, 2015, p. 645), we estimate that 614,702 of the 744,702 British military fatalities were among men residing in England and Wales at the outbreak of the war. This figure is necessarily approximate. For instance, it is lower than Winter's (1976, p. 547) estimate of 621,595 fatalities for England and Wales. Indeed, Watt (2019, pp. 87–98) shows that Winter's (1976) assumptions that 14 percent of British servicemen came from Scotland and Ireland *and* that they died at similar rates are inaccurate and likely lead to an underestimate. Conversely, Watt's (2019, pp. 83–87) figure of 100,500 Scottish fatalities is likely an overestimate because it counts servicemen born in Scotland as opposed to residing there at the outbreak of the war. Likewise, Fitzpatrick (2015, p. 645) points out that "[t]he true number of deaths attributable to war service, excluding those of 'Irish' servicemen enlisted outside Ireland, probably exceeded 30,000."

G. Measuring Military Death Rates

This appendix describes the construction of different measures of military death rates and explains the rationale for our preferred measure. We first detail how our baseline measure corrects for the underreporting of residence places in SDGW–CWGC records (Section G.1). We then compare this measure with alternatives that complement missing residence information using census-linked records, enlistment places, or birth places, and discuss the biases associated with each of these alternative approaches (Section G.2). Finally, we compare our preferred measure of military death rates with those used in the literature (Section G.3).

G.1. Correcting for the Underreporting of Residence Places

The SDGW–CWGC records we rely on to construct the numerator of our measure of military death rates cover only a subset of deceased British servicemen who resided in England and Wales at the outbreak of the war. First, these records exclude deceased officers of the Regular Army and Territorial Force, as well as deceased military personnel of the Royal Navy, Royal Naval Division, and Air Force. Second, a precise place of residence at enlistment is reported for only 48 percent of SDGW–CWGC records. And third, although we are able to geocode 99.5 percent of records with a reported place of residence, 14 percent of these are located outside of England and Wales—in Ireland or Scotland. As a result, we estimate that the SDGW–CWGC records whose places of residence we geocode in England and Wales account for only 45 percent of deceased British servicemen who resided there at the outbreak of the war ($\simeq$ 273,803/614,702; see Table A6 and Appendix F). In this section, we describe how we correct for both the underreporting of residence places among SDGW–CWGC records and the broader undercoverage of this population.

Enlistment-place adjustment procedure To account for the undercoverage of the universe of deceased British servicemen who resided in England and Wales, we scale the observed count of military fatalities. Applying a countrywide scaling factor of 2.245 ($\simeq$ 614,702/273,803) would implicitly assume that the residence field is missing from SDGW–CWGC records at a uniform rate across England and Wales. In practice, however, the rate at which it is missing varies substantially over space, reflecting the regimental structure of the original SDGW volumes: in some regimental volumes, places of residence are almost entirely absent, whereas in others they are almost always reported. For instance, among the 6,682 deceased soldiers of the Royal Sussex Regiment, only one has a recorded place of residence. By contrast, among the 13,067 deceased soldiers of the

King's Liverpool Regiment, all but 36 have a recorded place of residence. These two cases are not outliers: regiments reporting a place of residence for fewer than 2 percent of their death records account for 13 percent of all SDGW–CWGC records, while those reporting a place of residence for more than 98 percent account for 11 percent. The blue curve of Figure G1 displays the full distribution of missing residence rates across regiments, with each regiment weighted by its relative share of SDGW–CWGC records.

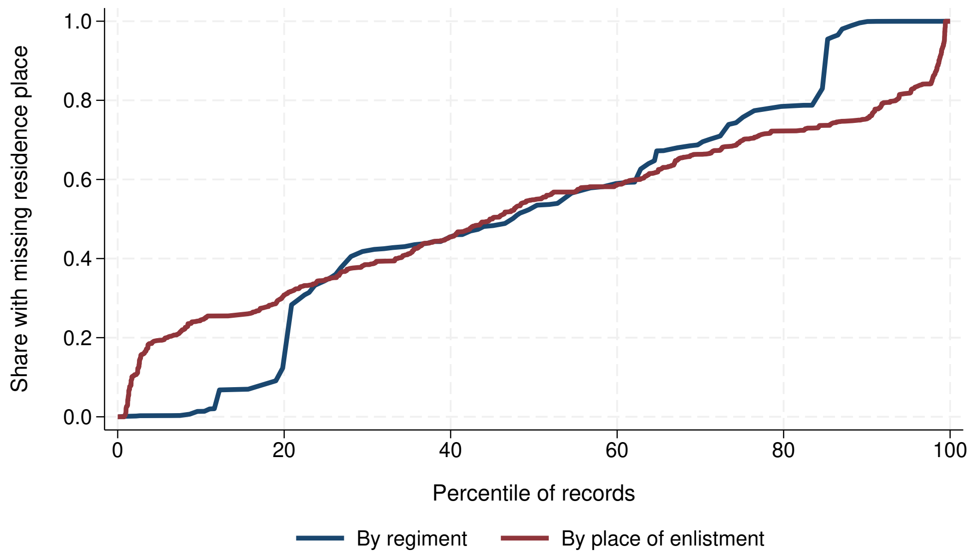


Figure G1. Distribution of SDGW–CWGC Records with a Missing Place of Residence

Notes. This figure displays the distribution of the share of SDGW–CWGC records with a missing place of residence across 108 regiments and 18,740 places of enlistment. Units are ordered by their missing residence rate. We weight each regiment and place of enlistment by its relative share of SDGW–CWGC records.

Given the territorial structure of military recruitment (see Section 3.2.2 in the main text), one way to address this variation in missing residence rates is to scale SDGW–CWGC records with a reported place of residence by the inverse of their regiment-specific residence reporting rate. Although this approach is valid in principle (under the assumption that the likelihood of reporting residence is independent of place of residence given regiment), it is imprecise in practice because a non-negligible number of SDGW–CWGC records are associated with regiments that report places of residence for a very small share of their records. For instance, under this procedure, all 6,682 deceased soldiers of the Royal Sussex Regiment would effectively be assigned to the place of residence of the single record in this regiment for which that information is reported.[35]

[35] In addition, the residence field is missing entirely for 15 regiments, representing 14 percent of all regiments and 0.6 percent of all SDGW–CWGC records.

We therefore apply this scaling procedure at the enlistment-place level, which is reported for 99.5 percent of SDGW–CWGC records (see Table A6). The validity of this method rests on the relatively weak assumption that the likelihood of reporting residence is independent of place of residence given enlistment place. Further, the red curve of Figure G1 shows that missing residence rates are far less concentrated on the right-hand side of the distribution when calculated at the enlistment-place level, making it a more stable basis for scaling observed SDGW–CWGC records with a reported place of residence. For instance, enlistment places reporting a place of residence for fewer than 2 percent of their death records account for 0.6 percent of all SDGW–CWGC records, as a given enlistment place may be associated with multiple regiments.[36,37] We thus estimate the true number of deceased soldiers of the Regular Army and Territorial Force for each local government district by first weighting the records reporting a place of residence within each enlistment place by the inverse of that enlistment place's residence reporting rate, and then summing the adjusted fatality counts across all enlistment places.[38] Applying this method to the 273,803 SDGW–CWGC records geocoded in England and Wales yields an adjusted fatality total of 542,821—this procedure therefore imputes places of residence for an additional 269,018 records (see Table G1).

Coverage of the enlistment-place adjusted fatality count While our enlistment-place adjustment procedure dramatically improves the number of records geocoded at a place of residence, it remains incomplete in two respects: first, with respect to the universe of SDGW–CWGC records of soldiers who resided in England and Wales, and second, with respect to military fatalities whose records are not included in SDGW volumes.

The first aspect encompasses two sets of cases. First, our enlistment-place adjustment procedure does not account for the 3,671 SDGW–CWGC records with no recorded place of residence whose enlistment place reports no residence information for any of its death records. Of these, 2,550 have an enlistment place geocoded in England and Wales.[39]

[36] In addition, the residence field is missing entirely for 2,186 enlistment places, representing 30 percent of all enlistment places and 0.6 percent of all SDGW–CWGC records.

[37] An additional advantage of applying the adjustment procedure at the enlistment-place rather than the regiment level is that all enlistment places can be linked to a specific district, whereas several regiments were not tied to a particular territory, such as the Corps of Royal Engineers.

[38] Consider the following simple example. Suppose that 100 records enlisted at place A, of which 20 report district 1 as their place of residence, 30 report district 2, and 50 have no recorded place of residence. Since the residence reporting rate for enlistment place A is 50 percent, we estimate the total number of deceased soldiers who enlisted at place A and resided in district 1 to 40 (= $20 \times 100/50$). Likewise, we estimate the total number of deceased soldiers who enlisted at place A and resided in district 2 to 60 (= $30 \times 100/50$).

[39] Adding these records to the adjusted total yields 545,371 (= 542,821 + 2,550). This figure is close to the 551,008 SDGW–CWGC records for which a place of enlistment is reported and geocoded to England and Wales (see Table A6). The small discrepancy of 5,637 records arises because the two figures capture slightly

Table G1. Coverage of Alternative Measures of Military Fatalities

Measure	Base	SDGW–CWGC records Additions	SDGW–CWGC records Total	SDGW–CWGC records Coverage	All fatalities Additions	All fatalities Coverage	Scaling
Raw residence places	273,803	—	273,803	0.500	—	0.445	2.245
Enlistment-place adjustment	273,803	269,018	542,821	0.992	—	0.883	1.132
Linked-place complement	273,803	94,497	368,300	0.673	—	0.599	1.669
+ RNWGR + other CWGC	273,803	94,497	368,300	0.673	37,014	0.659	1.517
Enlistment-place complement	273,803	281,228	555,031	1.014	—	0.903	1.108
Birth-place complement	273,803	255,267	529,070	0.967	—	0.861	1.162
+ RNWGR	273,803	255,267	529,070	0.967	22,579	0.897	1.114

Notes. This table reports the coverage of alternative measures of military fatalities for England and Wales. *Base* denotes the number of SDGW–CWGC records with a place of residence geocoded in England and Wales, 273,803 (see Table A6). *Additions* denote records assigned to England and Wales through each adjustment or complement procedure. *Total* denotes the sum of the base count and additions. *Coverage* in the *SDGW–CWGC* panel is the total count divided by the estimated total number of SDGW–CWGC records for soldiers who resided in England and Wales, i.e., 547,307. By contrast, *Coverage* in the *All fatalities* panel is the total count divided by the estimated total number of military fatalities among men residing in England and Wales, i.e., 614,702. The scaling factor is the inverse of the all-fatality coverage. The difference between 614,702 and 547,307 is 67,395 fatalities whose records are not included in the SDGW volumes.

Second, it also does not account for the 2,732 SDGW–CWGC records that report neither a place of residence nor a place of enlistment. Of these, 1,936 have a place of birth geocoded in England and Wales. Assuming that these soldiers also resided in England and Wales, these figures imply a total of 547,307 SDGW–CWGC records of soldiers who resided in England and Wales (= 542,821+2,550+1,936). As a result, our enlistment-place adjustment procedure yields a coverage rate of 99.2 percent of all SDGW–CWGC records of soldiers who resided in England and Wales—compared to 50.0 percent when relying only on records with a reported place of residence.

The second aspect encompasses four sets of cases, since our enlistment-place adjustment procedure does not account for military fatalities whose records are not included in SDGW volumes—this concerns fatalities among officers of the Regular Army and Territorial Force, as well as among military personnel of the Royal Navy, Royal Naval Division, and Air Force. Subtracting our estimate of SDGW–CWGC records of soldiers who resided

different populations. The figure of 545,371 corresponds to the number of SDGW–CWGC records—among those with a reported place of residence or enlistment—of servicemen who resided in England and Wales. It therefore includes servicemen who enlisted outside England and Wales but resided within it—this is the case for 3,032 records, and our procedure extrapolates from these to estimate a total of 6,953 among all SDGW–CWGC records with a reported place of enlistment. By contrast, the figure of 551,008 includes all SDGW–CWGC records with a place of enlistment geocoded to England and Wales, and therefore also includes servicemen who enlisted in England and Wales but resided elsewhere—this is the case for 6,319 records, and our procedure extrapolates from these to estimate a total of 12,590 such records among all SDGW–CWGC records with a reported place of enlistment. The two quantities can therefore be reconciled as follows: 545,371 − 6,953 = 551,008 − 12,590 (= 538,418, i.e., the number of SDGW–CWGC records of servicemen who enlisted *and* resided in England and Wales).

in England and Wales from our estimate of the total number of military fatalities among men residing in England and Wales yields 67,395 fatalities (= 614,702 – 547,307).[40]

Table G2. Alternative Measures of Military Death Rates

	Mean	S.d.	Min.	Max.	Obs.
Raw military death rate	0.075	0.051	0.000	0.585	1,721
Enlistment-place adjusted	0.078	0.050	0.000	0.732	1,721
Linked-place complemented	0.072	0.039	0.000	0.452	1,721
+ RNWGR + other CWGC	0.072	0.036	0.000	0.414	1,721
Enlistment-place complemented	0.060	0.050	0.000	0.689	1,721
Birth-place complemented	0.071	0.042	0.000	0.613	1,721
+ RNWGR	0.071	0.042	0.000	0.628	1,721

Notes. This table reports summary statistics at the local government district level for alternative measures of military death rates in England and Wales. Each measure is defined as the number of military fatalities assigned to a district divided by the male population aged 11 to 43 in that district in 1911. The *raw military death rate* measure uses only SDGW–CWGC records with a reported place of residence geocoded in England and Wales. The *enlistment-place adjusted* measure is the baseline measure described in Section G.1 and scales observed residence records within enlistment places to account for missing residence information. The *linked-place complemented* measure supplements the raw measure with residence places recovered by linking SDGW–CWGC records—as well as RNWGR records and unlinked CWGC records—to the 1911 census. The *enlistment-place complemented* and *birth-place complemented* measures instead assign records with missing residence information to their geocoded place of enlistment or birth, respectively. *S.d.* denotes standard deviation; *Max.*, maximum; *Min.*, minimum; and *Obs.*, observations.

These figures imply that our enlistment-place adjustment procedure nearly doubles the overall coverage of the numerator used to construct military death rates, from 45 to 88 percent (≃ 542,821/614,702). This corresponds to a countrywide scaling factor of 1.132, substantially lower than the original scaling factor of 2.245. Summary statistics for the resulting military death rate are reported in Table G2, together with alternative measures discussed in Section G.2. Because the numerator of each measure is rescaled to match the same countrywide fatality total of 614,702, the mean of the enlistment-place adjusted measure at 7.8 percent is close to that obtained using only SDGW–CWGC records with a reported place of residence geocoded in England and Wales, at 7.5 percent. Its cross-district standard deviation is also similar, at 5.0 percentage points.

But similarity in aggregate moments does not imply that the two measures allocate military fatalities similarly across districts. Indeed, Panels a and b of Figure G2 suggest that

[40] Table F1 documents 41,472 fatalities among officers of the Regular Army and Territorial Force, 34,457 among military personnel of the Royal Navy, 8,590 among the Royal Naval Division, and 6,166 among the Air Force, for a total of 91,102. The gap of 23,707 (= 91,102 – 67,395) is likely due to a substantial share of these men residing outside England and Wales, but it also reflects the marginal uncertainty surrounding these calculations. As highlighted in Section 3.2.1 in the main text, however, this uncertainty does not materially affect our analyses because our scaling is performed at the countrywide level.

Table G3. Baseline Versus Alternative Measures of Military Death Rates

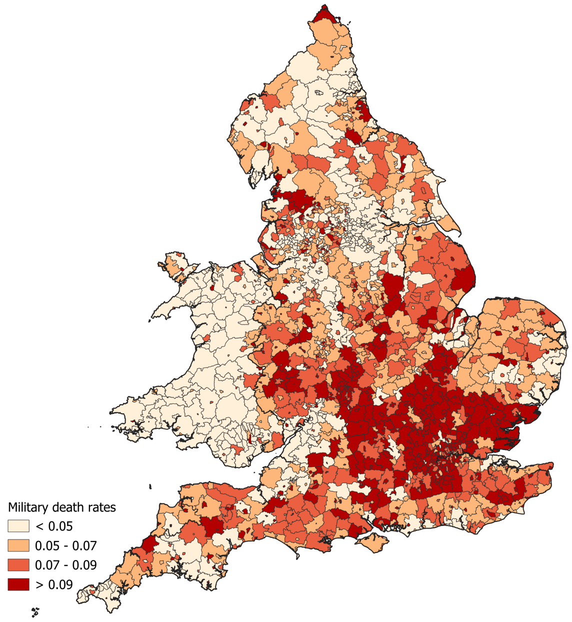
Outcome:	Alternative military death rate			
	Raw measure	Linked compl.	Enlistment compl.	Birth compl.
	(1)	(2)	(3)	(4)
Enlistment-place adjusted military death rate	0.899***	0.703***	0.643***	0.721***
	[0.032]	[0.027]	[0.054]	[0.039]
R^2	0.757	0.802	0.403	0.724
Districts	1,721	1,721	1,721	1,721
Alternative measure raw s.d.	0.051	0.039	0.050	0.042
Alternative measure residual s.d.	0.025	0.017	0.039	0.022
Residual / raw s.d.	0.493	0.445	0.773	0.525

Notes. This table reports coefficient estimates from regressions at the local government district level of each alternative military death rate on our baseline enlistment-place adjusted military death rate. Each column uses a different alternative measure as the dependent variable: the raw measure, the linked-place complemented measure, the enlistment-place complemented measure, and the birth-place complemented measure. Robust standard errors are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

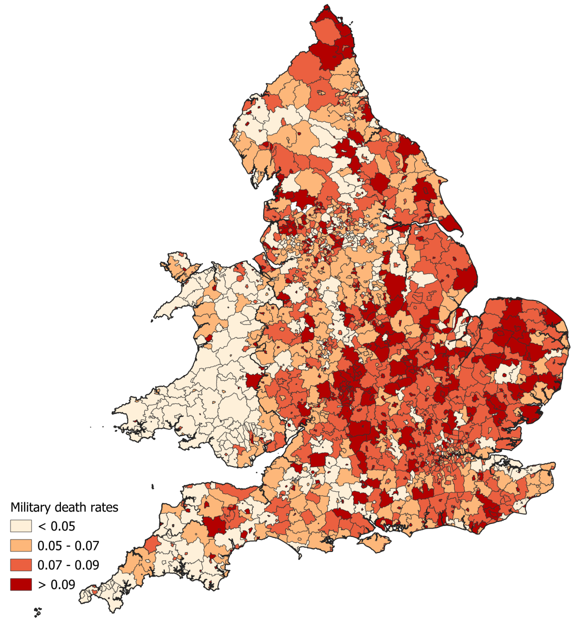
military death rates constructed using the enlistment-place adjustment are substantially less spatially concentrated than those based on the raw measure. Consistent with these patterns and despite a strong positive relationship between the two variables (Panel a of Figure G3), we find that this relationship is far from one-to-one: Column 1 of Table G3 implies that, after projecting the raw measure on the enlistment-place adjusted measure, the residual standard deviation remains about half as large as the raw measure's unconditional standard deviation—equivalently, about one quarter of the cross-district variance in the raw measure is orthogonal to that stemming from our baseline measure.

Enlistment-place adjusted military death rates and residence reporting rates A remaining concern is that spatial variation in residence reporting rates could be correlated with either military death rates or female labor force participation rates, in which case even the enlistment-place adjustment might leave systematic measurement error. Figure G4 displays the district-level distribution of residence reporting rates based on enlistment places.[41] These rates exhibit little spatial correspondence with the distributions of military death rates and female labor force participation rates reported in Figure 1 in the main text. The scatterplot in Panel a of Figure G5 confirms that, after applying the enlistment-place adjustment, residence reporting rates are orthogonal to the resulting measure of

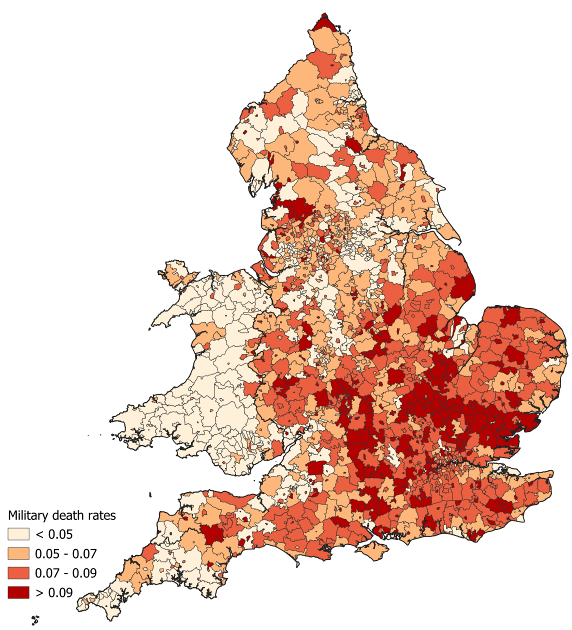
[41] Note that residence reporting rates are undefined for the 145 districts that did not contain any enlistment place. Among these districts, 63 percent were urban districts and 37 percent were rural districts—a composition similar to the overall distribution of administrative statuses across districts (see Table D1).



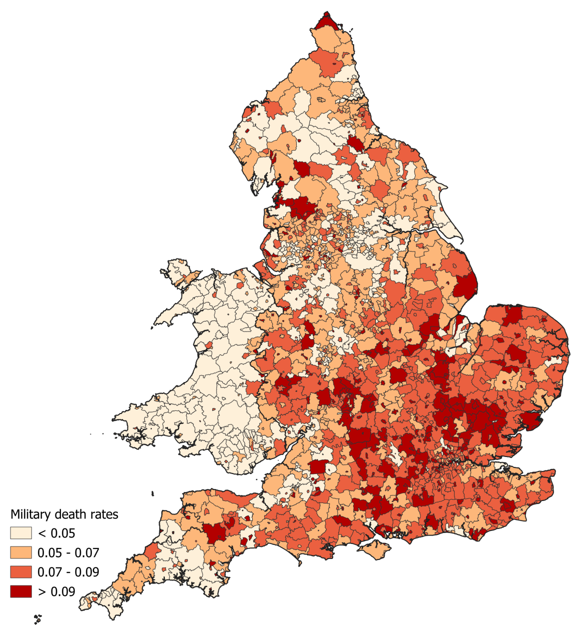
(a) Raw measure



(b) Enlistment-Place Adjusted



(c) Linked-Place Complemented



(d) Linked-Place Complemented + RNWGR + other CWGC

Figure G2. Alternative Measures of Military Death Rates

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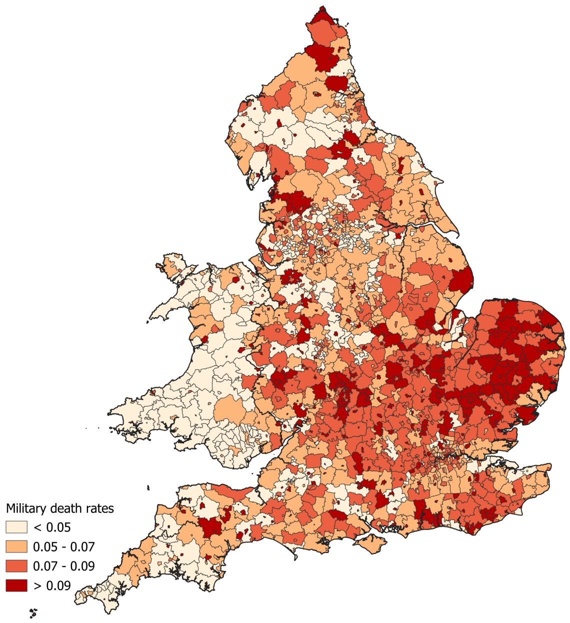
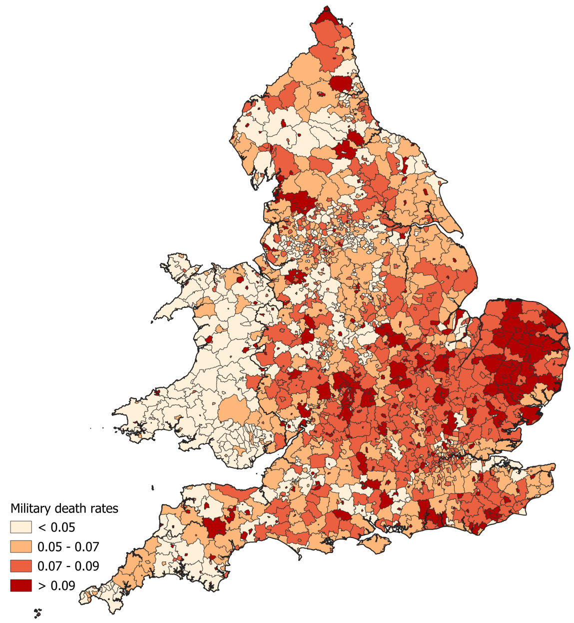
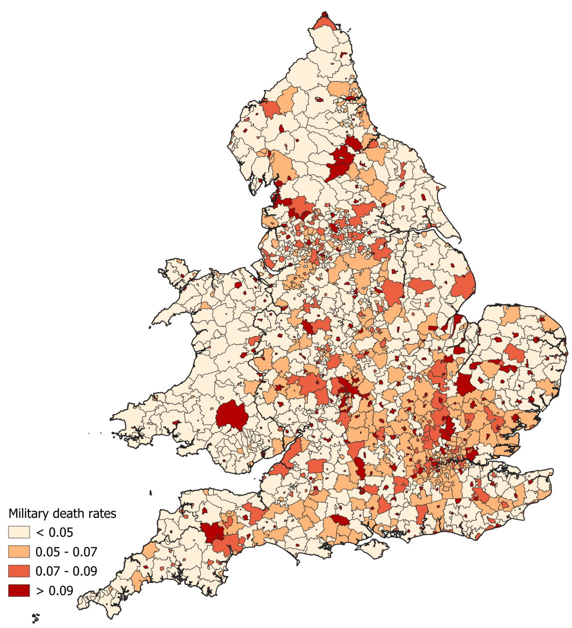


(e) Birth-Place Complemented



(f) Birth-Place Complemented + RNWGR



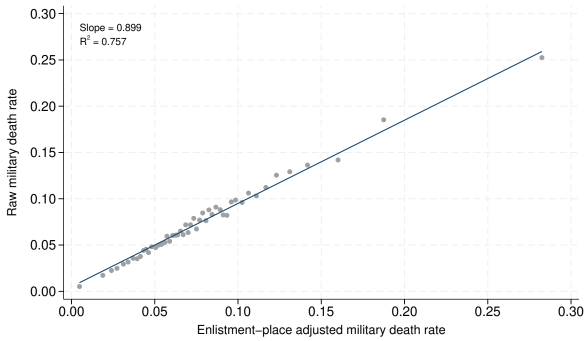
(g) Enlistment-Place Complemented

Figure G2. Alternative Measures of Military Death Rates
Continued

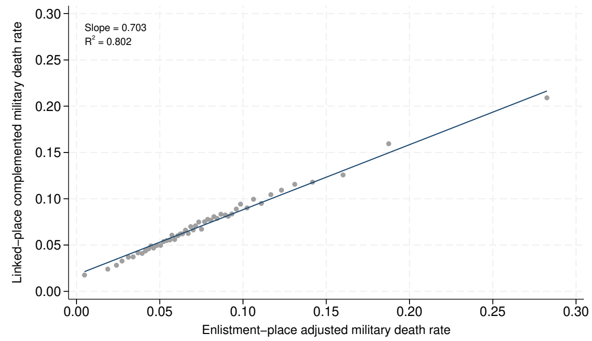
Notes. This figure maps alternative measures of military death rates across local government districts. See the text for definitions. Class breaks are defined by quartiles of the distribution of the enlistment-place adjusted measure. Shapefiles are adapted from Southall, Aucott and Burton (2024).

military death rates. By contrast, Panel b shows that military death rates calculated from the raw distribution of observed places of residence and adjusted by a single countrywide scaling factor of 2.245 are strongly correlated with residence reporting rates. This contrast supports the credibility of our correction procedure.

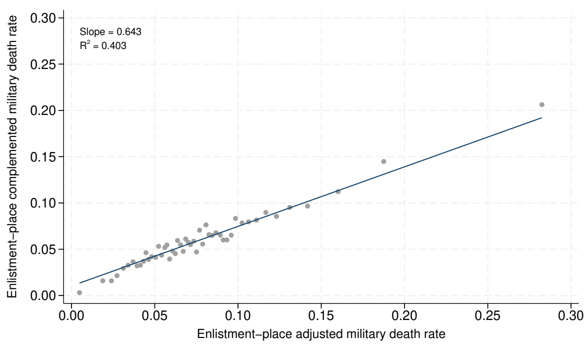
Table G4 provides further evidence on this point by reporting coefficient estimates from district-level regressions of residence reporting rates on alternative measures of military death rates and female labor force participation rates. The enlistment-place adjusted measure is uncorrelated with residence reporting rates, whereas the raw measure is



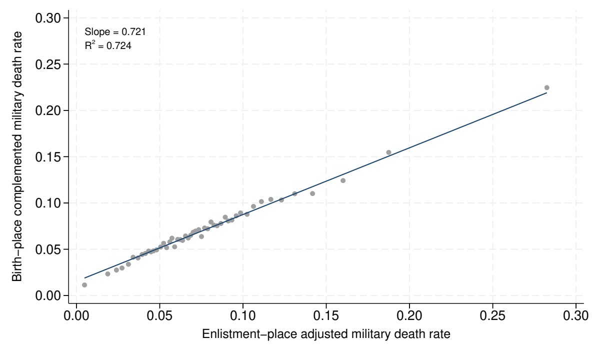
(a) Raw Measure



(b) Linked-Place Complemented



(c) Enlistment-Place Complemented



(d) Birth-Place Complemented

Figure G3. Baseline Versus Alternative Measures of Military Death Rates

Notes. This figure displays the relationship between alternative measures of military death rates and our baseline enlistment-place adjusted measure at the local government district level. Each panel reports a binned scatterplot with 50 quantile bins, where points represent bin means. In addition, solid lines represent linear fits whose slopes correspond to the regression estimates reported in Table G3. Binned scatterplots are generated using Stepner's (2013) `binscatter` Stata command.

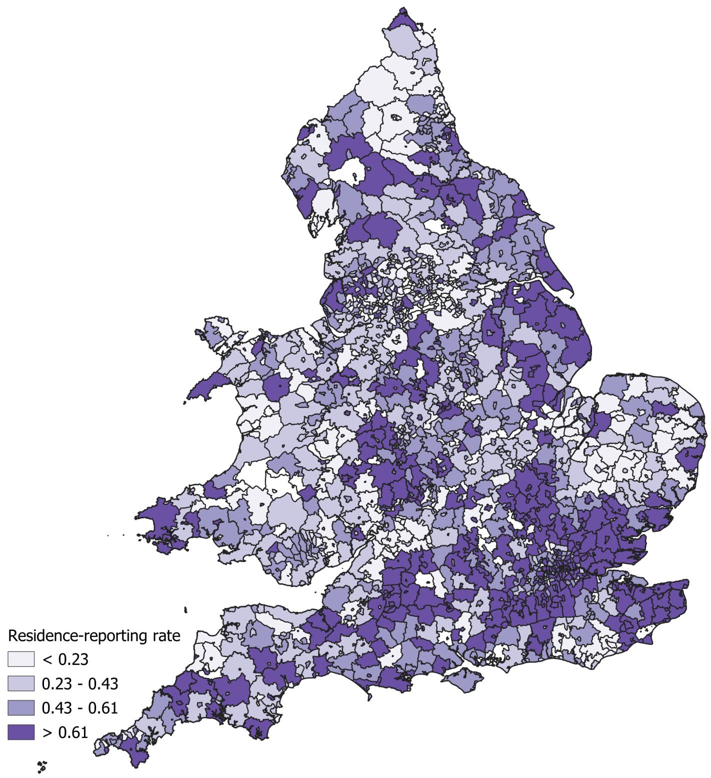


Figure G4. Residence Reporting Rates

Notes. This figure maps residence reporting rates across local government districts based on enlistment places. Specifically, we assign each enlistment place to a district and compute, for each district, the share of SDGW–CWGC records associated with enlistment places for which a place of residence is reported. Class breaks correspond to quartiles of the resulting residence reporting rate distribution. The shapefile is adapted from Southall, Aucott and Burton (2024).

strongly positively correlated.[42] These results therefore suggest that differential residence reporting rates are unlikely to bias our estimates of the relationship between enlistment-place adjusted military death rates and changes in female labor force participation rates.

Formal derivation We now formally demonstrate the validity of our enlistment-place adjustment procedure. Let P_i be an indicator equal to one if record i lists a place of residence, and zero otherwise, and D_{di}, an indicator equal to one if the local government district of residence listed in death record i is d, and zero otherwise. For $p \in \{0, 1\}$, let $S_{e,p}$ denote the set of records that enlisted at place e and for which $P_i = p$, and let $|S_{e,p}|$ be the number of records from enlistment place e. Assuming that P_i is independent of place of

[42]We also find no evidence that residence reporting rates are correlated with either 1911 female labor force participation rates (estimate: −0.014, s.e.: 0.060) or changes in female labor force participation rates between 1911 and 1921 (estimate: 0.119, s.e.: 0.196).

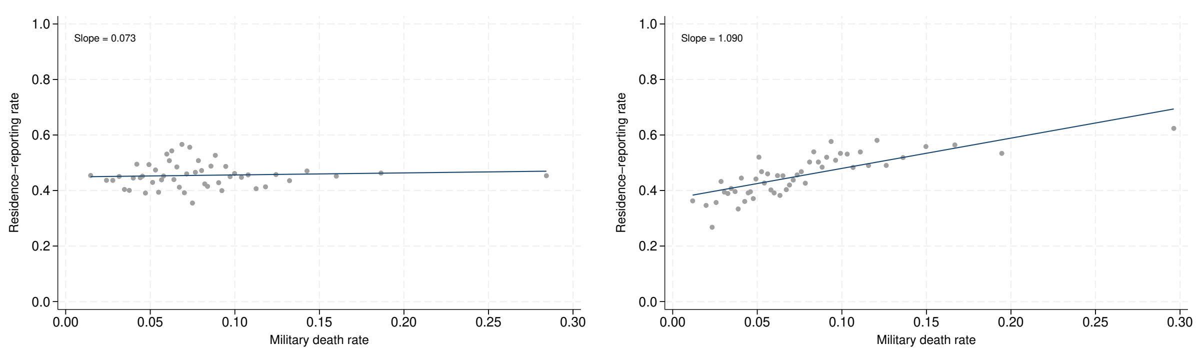
Table G4. Residence Reporting Rates and Military Death Rates

Outcome:	Residence reporting rate				
	(1)	(2)	(3)	(4)	(5)
Military death rate	1.090^{***}	0.073	1.035^{***}	0.358^{***}	0.274^{**}
	[0.135]	[0.129]	[0.161]	[0.117]	[0.127]
Measure	Raw measure	Enlistment adjusted	Linked compl.	Enlistment compl.	Birth compl.
R^2	0.051	0.000	0.026	0.005	0.002
Districts	1,576	1,576	1,576	1,576	1,576

Notes. This table reports coefficient estimates from regressions at the local government district level of residence reporting rates on military death rates. Residence reporting rates are defined as the share of SDGW–CWGC records with enlistment places located in a given district for which a place of residence is reported. The sample excludes the 145 districts with no enlistment places. Each column uses a different measure of military death rates: the raw measure, the enlistment-place adjusted measure, the linked-place complemented measure, the enlistment-place complemented measure, and the birth-place complemented measure. Robust standard errors are reported in brackets.
$^{***}p \leq 0.01$, $^{**}p \leq 0.05$, $^{*}p \leq 0.10$.

residence i given place of enlistment e,

$$\mathbb{E}\left[\sum_{i \in S_{e,1}} \frac{D_{di}}{|S_{e,1}|}\right] = E\left[D_{di} | i \in S_{e,1}\right] = E\left[D_{di} | i \in (S_{e,0} \cup S_{e,1})\right] \tag{G.1}$$



(a) Enlistment-Place Adjusted

(b) Raw Measure

Figure G5. Military Death Rates and Residence Reporting Rates

Notes. This figure displays the relationship between residence reporting rates and military death rates at the local government district level. Panel a uses the enlistment-place adjusted measure of military death rates, which reweights SDGW–CWGC records reporting a place of residence by enlistment-place residence reporting rates. Panel b uses the raw measure of military death rates based on the observed distribution of geocoded SDGW–CWGC records and a single countrywide scaling factor of 2.245. Each panel reports a binned scatterplot with 50 quantile bins, where points represent bin means. In addition, solid lines represent linear fits whose slopes correspond to the regression estimates reported in Columns 1 and 2 of Table G4. Binned scatterplots are generated using Stepner's (2013) `binscatter` Stata command.

Mechanically, the true number of records from enlistment place e that have place of residence d (that is, $N_{e,d}$) is simply the fraction of records from enlistment place e with place of residence d, multiplied by the total number of records from e:

$$N_{e,d} = \mathbf{E}\left[D_{di} | i \in (S_{e,0} \cup S_{e,1})\right] \times \left(|S_{e,0}| + |S_{e,1}|\right) \tag{G.2}$$

Putting these two equations together, it is straightforward to see that one can estimate the true number of records from enlistment place e that have place of residence d by taking the sum of death records from enlistment place e with place of residence d recorded, and multiplying by the adjustment factor $\left(|S_{e,0}| + |S_{e,1}|\right) / |S_{e,1}|$:

$$N_{e,d} = \mathbf{E}\left[\sum_{i \in S_{e,1}} D_{di}\right] \times \frac{|S_{e,0}| + |S_{e,1}|}{|S_{e,1}|} \tag{G.3}$$

The adjustment factor $\left(|S_{e,0}| + |S_{e,1}|\right) / |S_{e,1}|$ is analogous to that used in the baseline measure we use—the inverse of the proportion of records for which place of residence is reported—but it is computed and applied separately within each enlistment place rather than at the country level.

G.2. Alternative Measures of Military Death Rates

In this section, we discuss alternative measures of military death rates that could, in principle, be used instead of our baseline measure—the enlistment-place adjusted measure described in Section G.1. Each of these measures attempts to address the issue of partial coverage of the numerator—the number of military fatalities—in a different way. To avoid compounding multiple adjustment margins and to keep comparisons simple, all alternative measures begin with the same set of 273,803 SDGW–CWGC records for which the place of residence at enlistment is available and geocoded in England and Wales (see Table A6). They then attempt to improve coverage by supplementing missing places of residence with alternative sources of information. The first alternative measure increases the number of records with 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 instead of the place of enlistment.[43]

[43] Note that, by design, all three alternative measures—like the raw and enlistment-place adjusted measures—remain incomplete because they do not capture military fatalities outside the universe of SDGW–CWGC records—although we provide an additional alternative that partly addresses this gap. Hence, these alternative measures all still require scaling by some countrywide adjustment factor (see Table G1).

In what follows, we first describe the construction of each alternative measure (Section G.2.1). For each measure, we then report the number of records it adds relative to the uncorrected baseline of 273,803 SDGW–CWGC records with a reported place of residence geocoded in England and Wales, the resulting coverage of the numerator, and the implied countrywide scaling factor. We then assess how closely these alternative measures relate to our baseline measure of military death rates. Next, we examine the potential biases associated with each of them in detail (Section G.2.2).

G.2.1. Construction of Alternative Measures of Military Death Rates

Linked-place complemented measure The first alternative measure increases the number of records with an observed place of residence by linking SDGW–CWGC records without a reported place of residence to the 1911 census. Using a modified version of Abramitzky et al.'s (2021) ABE-JW deterministic record-linking algorithm described in Appendix H, we identify the place of residence for an additional 94,497 SDGW–CWGC records (Table G1). This raises the SDGW–CWGC coverage rate to 67.3 percent and the overall coverage rate to 60.0 percent, implying a countrywide scaling factor of 1.669.

The resulting military death rate averages 7.2 percent and has a cross-district standard deviation of 3.9 percentage points (Table G2). Its spatial distribution closely resembles that of the raw measure (Panel c of Figure G2). It is also strongly positively correlated with our baseline measure with a slope of 0.70 (Panel b of Figure G3) and one fifth of its cross-district variance orthogonal to that stemming from our baseline measure (Table G3).[44]

Enlistment-place complemented measure The second alternative measure assigns SDGW–CWGC records with no reported place of residence to their listed place of enlistment—this information is available for 99.5 percent of all SDGW–CWGC records (Table A6) and for 99.3 percent of those without a reported place of residence. Since 74 percent of these enlistment places are geocoded in England and Wales, this procedure adds 281,228 SDGW–CWGC records to the 273,803 already localized at their place of residence (Table G1). This procedure raises the SDGW–CWGC coverage rate to 101.4 percent and the overall coverage rate to 90.3 percent, implying a countrywide scaling factor of 1.108.[45]

[44]We construct an additional measure using the same methodology, further incorporating census records matched to RNWGR records of Royal Navy personnel and to CWGC records not matched to SDGW—mainly officers of the Regular Army and Territorial Force and military personnel of the Air Force. This procedure adds 37,014 records, raising the overall coverage rate to 65.9 percent and reducing the scaling factor to 1.517 (Table G1). The resulting distribution of military death rates is broadly similar to that of the baseline linked-place complemented measure (see Table G2 and Panel d of Figure G2).

[45]The SDGW–CWGC coverage rate exceeds 100 percent because many soldiers who enlisted in England and Wales resided elsewhere at the time of enlistment—see Footnote 39.

The resulting military death rate averages 6.0 percent and has a cross-district standard deviation of 5.0 percentage points (Table G2). Its spatial distribution differs sharply from both the raw and baseline measures (Panel g of Figure G2), as this procedure assigns most additional records to urban districts—where enlistment centers were generally located. This concentration in urban districts also helps explain why the average military death rate is lower than for the other measures—the number of men eligible for military service was higher in urban than in rural areas. Nevertheless, this alternative measure remains positively related to our baseline measure, with a slope of 0.64 (Panel c of Figure G3). However, the relationship is substantially weaker as three fifths of its cross-district variance is orthogonal to that stemming from our baseline measure (Table G3).

Birth-place complemented measure The third alternative measure assigns SDGW–CWGC records with no reported place of residence to their listed place of birth—this information is available for 89.7 percent of all SDGW–CWGC records (Table A6) and for 92.3 percent of those without a reported place of residence. Since 72 percent of these birth places are geocoded in England and Wales, this procedure adds 255,267 SDGW–CWGC records to the 273,803 already localized at their place of residence (Table G1). This procedure raises the SDGW–CWGC coverage rate to 96.7 percent and the overall coverage rate to 86.1 percent, implying a countrywide scaling factor of 1.162.

The resulting military death rate averages 7.1 percent and has a cross-district standard deviation of 4.2 percentage points (Table G2). Its spatial distribution somewhat resembles that of our baseline measure (Panel e of Figure G2). It is also strongly positively correlated with our baseline measure with a slope of 0.72 (Panel d of Figure G3) and a little over one quarter of its cross-district variance orthogonal to that stemming from our baseline measure (Table G3).[46]

G.2.2. Biases of Alternative Measures of Military Death Rates

In this section, we assess the potential biases associated with each alternative measure. We first show that complementing missing residence information with enlistment or birth places introduces substantial spatial bias. We then show that the measure complemented with census-linked SDGW–CWGC records is subject to selection into linking and remains vulnerable—as the other alternatives—to differential residence reporting rates

[46]We construct an additional measure using the same methodology but further incorporating RN-WGR records of Royal Navy personnel localized at their place of birth. This procedure adds 22,579 records, raising the overall coverage rate to 89.7 percent and reducing the scaling factor to 1.114 (Table G1). The resulting distribution of military death rates is broadly similar to that of the baseline birth-place complemented measure (see Table G2 and Panel f of Figure G2).

across enlistment places. Together, these results support our baseline measure as the preferred measure for analyzing the relationship between military death rates and changes in female labor force participation.

Biases of the enlistment- and birth-place complemented measures The treatment variable is intended to approximate the local demographic shock caused by the war. This requires assigning deceased servicemen to the places where they would likely have resided after the war had they survived. For this reason, we prefer a measure based on deceased servicemen's place of residence at enlistment—with the caveat that this information is available for only 45 percent of deceased British servicemen who resided in England and Wales at the outbreak of the war. Alternative measures that complement missing residence information with birth or enlistment places each rely on specific assumptions: using birth places assumes that servicemen still resided in the district where they were born, and using enlistment places assumes that they enlisted in the same district in which they resided.

Both assumptions are far from accurate. Table G5 reports the distribution of SDGW–CWGC records across all possible configurations of birth, residence, and enlistment places among the 225,721 records for which all three locations are available.[47] Indeed, only 46.2 percent of deceased soldiers in this sample resided in the district where they were born (= 0.146 + 0.316) and only 25.9 percent resided in the district where they enlisted (= 0.146 + 0.113). By contrast, the most common configuration is that of soldiers who were born in one district, resided in another, and enlisted in a third—this configuration accounts for 37.1 percent of the sample.

Importantly, this mobility is not spatially neutral: the characteristics of residence, birth, and enlistment districts differ systematically across deceased soldiers' mobility status. Panel A of Table G6 compares the characteristics of residence and birth districts separately for soldiers whose residence and birth districts coincide (46 percent of the sample) and for those whose residence and birth districts differ (54 percent). Among non-movers, 31 percent were born and resided in a rural district.[48] By contrast, among movers, 33 percent were born in a rural district, but only 27 percent resided in one. This gap implies that at least 6,800 mobile deceased soldiers—about 6 percent of movers ($\simeq$ 6,800/121,434)—moved from a rural to an urban district. As a result, a measure based on birth places would

[47]This subset is relatively close to the sample with an available place of residence, which includes 273,803 SDGW–CWGC records (Table A6). Table A7 shows little evidence of selection into this sample along observable characteristics.

[48]We classify a district as rural if its administrative status was a rural district, as opposed to an urban district, a county borough, a municipal borough, a metropolitan borough, or the London county corporate (see Table D1).

Table G5. Birth, Residence, and Enlistment Districts in SDGW-CWGC Records

District configuration	Count	Share
Birth = residence = enlistment	32,985	0.146
Birth = residence ≠ enlistment	71,302	0.316
Birth = enlistment ≠ residence	12,254	0.054
Residence = enlistment ≠ birth	25,432	0.113
Birth ≠ residence ≠ enlistment	83,748	0.371
Total	225,721	1.000

Notes. This table classifies SDGW–CWGC records along their birth, residence at enlistment, and enlistment local government districts. Categories are mutually exclusive and exhaustive among the 225,721 records for which information on all three locations is available.

overstate military fatalities in rural areas by at least 6,800 or 3 percent of the full sample ($\simeq$ 6,800/225,721). Given that these birth-place districts experienced larger declines in female labor force participation after the war, assigning military fatalities based on birth places would overstate military death rates in these areas, thus biasing downward their estimated effect on female labor force participation.

Table G6. Characteristics of Residence, Birth, and Enlistment Districts by Mobility Status

A. Birth-place complemented measure				
Sample:	Res. = birth	Res. ≠ birth		
District characteristics:	Res./birth	Res.	Birth	Birth − res.
Rural district	0.312	0.274	0.330	0.056
1911 FLFP	0.343	0.339	0.346	0.007
1911–21 FLFP Δ	−0.015	−0.012	−0.016	−0.004
Number of records	104,287	121,434	121,434	
Share of records	0.462	0.538	0.538	

B. Enlistment-place complemented measure				
Sample:	Res. = enlist.	Res. ≠ enlist.		
District characteristics:	Res./enlist.	Res.	Enlist.	Enlist. − res.
Rural district	0.183	0.330	0.130	−0.200
1911 FLFP	0.358	0.335	0.381	0.046
1911–21 FLFP Δ	−0.011	−0.014	−0.016	−0.003
Number of records	58,417	167,304	167,304	
Share of records	0.259	0.741	0.741	

Notes. This table compares the mean characteristics of local government districts associated with deceased soldiers' residence, birth, and enlistment places. Panel A compares residence and birth districts, separating records for which the two locations coincide from those for which they differ. Panel B compares residence and enlistment districts analogously. All statistics are computed at the SDGW–CWGC record level on the sample of 225,721 records for which information on all three locations is available.

A measure based on enlistment places introduces an analogous spatial bias. Panel B of Table G6 compares the characteristics of residence and enlistment districts separately for soldiers whose residence and enlistment districts coincide (26 percent of the sample) and for those whose residence and enlistment districts differ (74 percent). Among non-movers, 18 percent resided and enlisted in a rural district. By contrast, among movers, 33 percent resided in a rural district, but only 13 percent enlisted in one. This gap implies that at least 33,461 mobile deceased soldiers—about 20 percent of movers ($\simeq$ 33,461/167,304)—traveled from a rural to an urban district to enlist. As a result, a measure based on enlistment places would understate military fatalities in rural areas by at least 33,461 or 15 percent of the full sample ($\simeq$ 33,461/225,721). Given that these enlistment-place districts experienced larger declines in female labor force participation after the war, assigning military fatalities based on enlistment places would overstate military death rates in these areas, thus substantially biasing downward their estimated effect on female labor force participation.

Biases of the linked-place complemented measure The alternative measure that complements residence information based on linking SDGW–CWGC records with the 1911 census is theoretically not subject to the aforementioned biases. Nevertheless, the linking process—described in Appendix H—introduces distinct issues, notably selection into matching and false-positive matches.

The linking procedure may lead to selection because records with more distinctive characteristics are more likely to be matched. Table G7 outlines several variables on which selection appears to occur. For instance, SDGW–CWGC records with a rural place of birth are more likely to have a unique candidate match because rural parishes typically had smaller populations than urban ones. Likewise, SDGW–CWGC records with a later birth year are more likely to have a unique candidate match because younger men were more likely to reside with their parents in 1911, which enables parental names to be used as additional matching information. Linked records are also slightly less likely to hold the rank of Private; one possible explanation for this is that educated men may have been both more likely to be linked (due in part to a decreased likelihood of age rounding when completing the census), and more likely to hold a higher rank.

Finally, linked records have slightly earlier death dates than those in the full CWGC dataset, but those linked and in SDGW–CWGC have slightly later death dates than those in the full SDGW–CWGC dataset. The first difference is due to the fact that records that appear in both the SDGW and CWGC datasets are considerably more likely to be linked than those that appear in the CWGC dataset alone (due to typically having more information to link on from SDGW), and at the same time tend to have slightly earlier death

Table G7. Selection of Death Records Matched to 1911 Census

Characteristics	Full sample Obs.	Full sample Mean	Linked sample Obs.	Linked sample Mean	Difference
A. All CWGC vs. linked CWGC records					
Birth year	479,079	1,890.71	184,209	1,891.89	1.181***
		(7.33)		(6.34)	[0.013]
Death year	835,736	1,917.25	225,238	1,917.23	−0.019***
		(1.24)		(1.19)	[0.002]
Private	835,736	0.699	225,238	0.695	−0.003***
		(0.459)		(0.460)	[0.001]
B. All SDGW-CWGC vs. linked SDGW-CWGC records					
Birth year	367,416	1,891.23	150,144	1,892.01	0.776***
		(6.64)		(6.34)	[0.013]
Death year	659,412	1,917.14	186,624	1,917.18	0.039***
		(1.15)		(1.12)	[0.002]
Private	659,412	0.813	186,624	0.806	−0.007***
		(0.390)		(0.396)	[0.001]
Rural birth district	483,219	0.312	166,182	0.337	0.025***
		(0.463)		(0.473)	[0.001]

Notes. This table reports means for selected characteristics in the full and linked samples. In Panel A, the full sample includes all CWGC records and the linked sample, CWGC records that could be matched to the 1911 census. In Panel B, the full sample includes all SDGW–CWGC records and the linked sample, SDGW–CWGC records that could be matched to the 1911 census. The last column reports differences in means across the full and linked samples. Bootstrap standard errors based on 1,000 replications are reported in brackets.
$^{***}p \le 0.01$, $^{**}p \le 0.05$, $^{*}p \le 0.10$.

dates (likely due to differences in major battle dates between army and navy records). The second difference is likely due to the previously discussed fact that linked men tend to be younger, and thus on average would have enlisted slightly later and in turn missed some of the earlier fighting.

Table G8. Characteristics of Residence and 1911 Census Districts by Census Linking Status

Sample:	Res. = census	Res. ≠ census		
District characteristics:	Res./census	Res.	Census	Census – res.
Rural district	0.321	0.290	0.378	0.088
1911 FLFP	0.339	0.339	0.329	−0.010
1911–21 FLFP Δ	−0.013	−0.015	−0.016	−0.001
Number of records	51,246	35,456	35,456	
Share of records	0.591	0.409	0.409	

Notes. This table compares the mean characteristics of local government districts associated with deceased soldiers' place of residence at time of enlistment (*Res.*), and place of residence in the 1911 census (*Census*), separating records for which the two locations coincide from those for which they differ. All statistics are computed at the SDGW–CWGC record level on the sample of 86,702 records for which information on both locations is available.

The linking procedure may also produce false-positive matches. Bailey et al. (2020) suggest that common matching procedures typically yield 66 to 90 percent of true matches—Abramitzky et al. (2021) provide a more optimistic range of 70 to 95 percent. While our matching procedure is conservative, some false-positive matches are unavoidable. And although the false-positive rate cannot be estimated, we can assess the broad quality of the resulting matches by comparing the characteristics of districts of residence at enlistment listed on SDGW–CWGC records with districts of residence from the 1911 census among records for which both locations are observed—this is the case for 86,702 of the 181,203 SDGW–CWGC records that we are able to link to the 1911 census (48 percent). We find that only 59.1 percent of these records report the same district of residence at enlistment as in the 1911 census. Although part of the remaining 40.9 percent likely reflects geocoding errors or genuine migration between 1911 and enlistment, many are presumably due to false-positive matches.[49] Importantly, these potential mismatches are not spatially neutral: among soldiers whose district of residence at enlistment differs from that recorded in the 1911 census, 37.8 percent resided in a rural district in 1911, compared with only 27.4 percent at enlistment (Table G8). Again, this pattern may partly reflect mi-

[49]Geocoding is particularly challenging in London, given the large number of local government districts in the city—28 metropolitan boroughs in addition to the London county corporate. As a result, only 47.3 percent of linked records geocoded in London report the same district of residence at enlistment as in the 1911 census.

gration toward urban areas between 1911 and enlistment, perhaps to service war factories. Nevertheless, this suggests that the rural bias of the linking procedure itself may play a role as well. Overall, given that these 1911 residence districts experienced slightly larger declines in female labor force participation after the war, this measure is likely to slightly bias downward the effect of military death rates on female labor force participation.

Finally, like the raw measure and the two other alternative measures, the military death rate based on this adjustment remains strongly positively correlated with residence-reporting rates across enlistment places (Columns 3–5 of Table G4). Although these reporting rates are uncorrelated with changes in female labor force participation between 1911 and 1921 (see Footnote 42), the resulting measurement error would still generate attenuation bias. For these reasons, we still prefer the enlistment-place adjusted military death rate as our baseline measure.

G.3. Comparison with Measures of Military Death Rates in the Literature

In this section, we compare our approach with studies of World War I in England and Wales that use spatial variation in military fatalities as their core treatment variable. Two studies fall into this category: Carozzi, Pinchbeck and Repetto (2026, henceforth CPR26) and Tchaouchev (2026, henceforth T26).[50,51] Table G9 summarizes the key characteristics of each measure. Although both studies focus on England and Wales, they use different levels of aggregation: 10,807 (grouped) parishes for CPR26 and 509 parliamentary constituencies for T26, compared with our 1,721 local government districts. For this reason, the comparison below focuses primarily on the construction of the numerator rather than on military death rates per se.[52]

Differences in methodology The first difference concerns the archival sources used to identify military fatalities. We begin with SDGW records and restrict this sample to records that can be linked to CWGC (see Appendix E). By contrast, CPR26 begin with CWGC records and supplement them with location information from SDGW.[53] Because CWGC records include servicemen who died while serving in Commonwealth military

[50] Repetto et al. (2026) also use military death rates as their core treatment variable, but they explicitly rely on CPR26's measure.

[51] We focus on the March 2026 version of T26.

[52] CPR26 use the log number of military fatalities as their main treatment variable but report summary statistics for military death rates (Table 1, p. 13), which we reproduce in Table G9.

[53] CPR26 state that they supplement CWGC records with "information from Forces War Records" (Appendix A.1, p. 1), but the relevant Forces War Records database corresponds to the same SDGW source we use (see `https://uk.forceswarrecords.com/publication/1278/uk-soldiers-died-in-the-great-war-1914-1919`, accessed May 2026). Although ODGW records are included in this source, they do not add location information as they report no birth, residence, or enlistment places (see Panel B of Table A4).

Table G9. Measures of Military Fatalities and Death Rates in the Literature

Study	Numerator: Source	Numerator: Assignment	Numerator: Records	Denominator*
GM26	SDGW–CWGC	Residence	542,821	Men 11–43
CPR26	CWGC (+ FWR/SDGW)	Residence + birth (+ other)	584,241	Population
T26	CWGC	Other	356,690	Men 11–46†

Study	Military death rate: Scaling	Level of aggregation	Mean	S.d.
GM26	Yes	Local government district	0.078	0.050
CPR26	No	(Grouped) Parish	0.010	0.010
T26	No	Parliamentary constituency	0.069	0.030

Notes. This table compares the construction of measures of military fatalities and death rates in GM26, CPR26, and T26. *GM26* refers to the present article, *CPR26*, to Carozzi, Pinchbeck and Repetto (2026), and *T26*, to Tchaouchev (2026). *Source* indicates the archival records used to construct the numerator. *Assignment* indicates the location rule used to allocate records across spatial units. *Other* corresponds to the location information drawn from the "Additional information" field in CWGC records. *Records* reports the number of records effectively retained to compute military death rates. *Scaling* indicates whether the measure is rescaled to match an external estimate of total military fatalities. *Population data are from the 1911 population census. †T26 denominator includes men aged 11 to 46 in 1911 net of those occupied in a scheduled protected occupation.

forces as well as certain auxiliary organizations, CPR26's final sample may thus be broader than ours, including servicemen outside the British Army who were nonetheless born in England and Wales. This difference also applies to T26, whose measure is likewise based on CWGC records.

The second difference concerns how military fatalities are assigned to spatial units. We restrict location information to the SDGW residence field and address missing residence coverage using the adjustment procedure described in Section G.1. CPR26 instead "give priority to place of residence at enlistment in the geolocation" and, "[w]hen this is not available, [they] use place of birth instead" (p. 12).[54] Their baseline measure broadly corresponds to the birth-place complemented measure discussed in Section G.2 and is therefore vulnerable to the spatial biases we document.[55,56] This procedure enables CPR26 to geocode 584,241 records in England and Wales. T26 adopts a different approach, relying exclusively on the location information reported in the "additional information" field of

[54] They further complement their records using "the 'additional information' string included in the CWGC source" (Appendix A.3, p. 4), which contains several types of location information, including the residence of parents or spouses (see Table A4).

[55] CPR26 justify their reliance on birth places because they "expect that both sources would yield very similar figures for deaths because most people reside in the same parish in which they were born" (Appendix A.3, p. 6). Our data do not support this assumption, as only 46 percent of deceased soldiers resided in the same district in which they were born (see Table G5). Since local government districts are much larger than parishes, the corresponding parish-level figure should be lower still.

[56] The downward bias entailed by this measure may help explain why CPR26 find no effect of World War I on female labor force participation—see Column 2 of table B7 (Appendix B, p. 20).

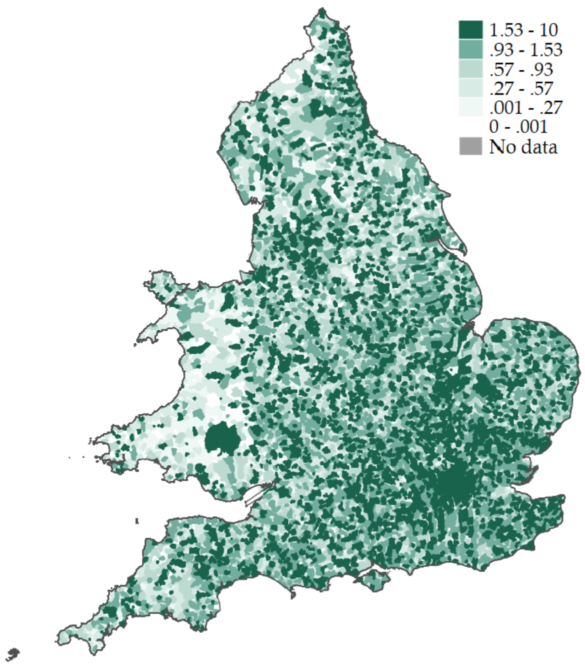
CWGC records. Based on the information reported in Table A4, about two thirds of these locations correspond to the residence of deceased servicemen's parents, one fifth to that of their spouse, and the remainder to their birth place or to the residence of a sibling or child—the interpretation of this measure is thus relatively unclear. This procedure enables T26 to geocode 356,690 records in England and Wales.

A third difference concerns whether the numerator accounts for the underreporting of location information. We address this issue explicitly. First, we define the universe of military fatalities covered by our source records—soldiers of the Regular Army and Territorial Force. Second, we estimate the total number of fatalities within this universe who resided in England and Wales at the outbreak of the war—614,702 fatalities. Finally, we scale the observed fatality counts so that the aggregate numerator matches this total. Neither CPR26 nor T26 explicitly accounts for this issue of limited coverage. Instead, both studies implicitly treat their geocoded records as covering the universe of military fatalities associated with England and Wales.[57]

The final difference concerns the denominator. We use the number of men aged 11 to 43 in 1911 to approximate the male population potentially exposed to military service during the war. CPR26 instead use the total population as the denominator—though their main treatment variable is the log number of fatalities. T26 nets out men employed in scheduled protected occupations. This choice is debatable, since at least half of wartime enlistments occurred voluntarily rather than through conscription, so that many men in protected occupations may nevertheless have enlisted (see Table A10).

Comparing spatial distributions Figure G6 displays the spatial distributions of measures of military death rates proposed by CPR26 and T26. Both differ markedly from our preferred enlistment-place adjusted measure, shown in Panel b of Figure G2. CPR26's measure is (unsurprisingly) closest to Panel e of Figure G2, which reports the birth-complemented measure: both maps exhibit a broad concentration of high military death rates across eastern and southeastern England and the Midlands, together with a relatively diffuse pattern elsewhere. By contrast, T26's measure is much harder to reconcile with any of our proposed measures. At most, it bears a loose resemblance to Panel g of Figure G2, which reports the enlistment-complemented measure.

[57]Note that CPR26 acknowledge that their geolocation procedure covers only 73 percent of the fatalities in their source data and discuss the resulting measurement-error concerns, but they do not rescale the observed counts to correct for this incomplete coverage.



(a) Carozzi et al. (2026)

(b) Tchaouchev (2026)

Figure G6. Alternative Measures of Military Death Rates in the Literature

Notes. This figure maps alternative measures of military death rates proposed in the literature. Panel a corresponds to the military death rate at the parish level proposed by Carozzi, Pinchbeck and Repetto (2026, figure 2, p. 14). Panel b corresponds to the military death rate at the parliamentary constituency level proposed by Tchaouchev (2026, figure 3, p. 11).

H. Linking CWGC Records to the 1911 Census

To maximize the number of military records with an observed place of residence at the outbreak of the war, we link all 835,740 CWGC records—augmented, where possible, with information from SDGW and RNWGR (see Appendix E)—to the 1911 census. We use a modified version of Abramitzky et al.'s (2021) ABE-JW deterministic record-linking algorithm, matching on names, age, place of birth, and parents' names. Starting with the pool of potential draftees in the 1911 census, i.e., men aged 11 to 43, we proceed as follows.

1. **Name normalization**

   We apply Abramitzky et al.'s (2021) name-normalization algorithm, which removes non-alphabetic characters and harmonizes common first names so that, for instance, 'Ben' and 'Benjamin' are treated as a match.

2. **Set of candidate matches**

   For each death record, we construct a set of candidate matches in the 1911 census. Candidate matches are census records that share the same first and last initials as the death record *and* satisfy two additional criteria. First, when the death record reports the county of birth, we restrict matches to individuals born in that county—otherwise, we allow candidates from all birth counties in the 1911 census. Second, when the death record reports the age at death, we restrict matches to individuals whose year of birth is within two years of the implied year of birth—otherwise, we allow candidates from all birth years.

3. **Jaro–Winkler name similarity score**

   For each candidate match, we compute Jaro–Winkler similarity scores for first and last names separately, and for parents' first names when available in both the death and census records. The Jaro–Winkler score is a string-comparison metric that places greater weight on agreement at the beginning of the string and ranges from 0 (no similarity) to 1 (exact match). We then retain only those candidate matches that satisfy *all* the following criteria:

   (a) The first-name Jaro–Winkler score is greater than 0.9.

   (b) The last-name Jaro–Winkler score is greater than 0.9.

   (c) If both parents' first names are provided in both the death and census records, the smallest Jaro–Winkler score is greater than 0.75.

   (d) One of the following conditions is true:

- The middle name Jaro–Winkler score is greater than 0.9.
- Both parents' first-name Jaro–Winkler scores are greater than 0.9.
- We have a birth-parish match.
- We have a birth-county match *and* a birth-year match within one year.

4. **Composite match-quality score**

   We then assign a match-quality score to each surviving candidate match, awarding one point for each of the following conditions:

   (a) The middle-name Jaro–Winkler score is greater than 0.9.

   (b) Both parents' first-name Jaro–Winkler scores are greater than 0.9.

   (c) We have a birth-parish match.

   (d) We have a birth-county match *and* a birth-year match within one year.

5. **Selection of unique best matches among death records**

   For each death record, we designate as the "best candidate" match the one that has the highest match-quality score among surviving candidate matches. If multiple candidate matches share the highest score, we do not assign any "best candidate" match.

6. **Selection of unique best matches among census records**

   We proceed in the same way as above but for each census record.

7. **Selection of best matches**

   Finally, we retain only those matches that are identified as the "best candidate" match among both the death and census records.

Applying this procedure, we uniquely match 225,238 CWGC records to the 1911 census. Of these, 218,217 (97 percent) are matched to a consistent local government district using the 1911 parish of residence.[58] Of these 218,217 death records, 181,203 (83 percent) correspond to an SDGW–CWGC record, 9,898 (5 percent) correspond to an RNWGR-CWGC record, and the remaining 27,116 (12 percent) appear only in CWGC.

[58] Of the remaining 7,021, 6,792 have 1911 place of residence classified as either "Royal Navy At Sea And In Ports Abroad" or "Military Abroad." The remaining 229 have a 1911 residence in one of 37 parishes that we are unable to match to a district.

I. Measurement Error Correction

I.1. Estimating the Measurement-Error Variance

The place of residence at the time of enlistment is available for only about 45 percent of British military fatalities who resided in England and Wales at the outbreak of the war (see Appendix F). As a result, any measure of military death rates using this location information is potentially subject to measurement error. As described in Appendix G, we mitigate this error by exploiting information on places of enlistment. But this procedure does not eliminate measurement error entirely, as district-level averages are estimated from incomplete samples of fatalities rather than from the full population.

To see why, let $p_{e,d}$ denote the proportion of fatalities who enlisted at place e and resided in district d at the time of enlistment. Using the notation from Appendix G,

$$p_{e,d} = \mathbf{E}\left[D_{di} | i \in (S_{e,0} \cup S_{e,1})\right] \tag{I.1}$$

Next, let $\bar{p}_{e,d}$ denote the corresponding proportion among fatalities who enlisted at e and for whom a place of residence is reported:

$$\bar{p}_{e,d} = \frac{\sum_{i \in S_{e,1}} D_{di}}{|S_{e,1}|}. \tag{I.2}$$

Under the assumption that, conditional on place of enlistment, the likelihood of reporting a place of residence is independent of the place of residence itself, Equation G.1 implies that $E(\bar{p}_{e,d}) = p_{e,d}$. Thus, $\bar{p}_{e,d}$ is an unbiased estimator of $p_{e,d}$. Its realized value may nevertheless differ from $p_{e,d}$ because it is estimated from a finite sample. The variance of $\bar{p}_{e,d}$ can be written as:

$$\widehat{\sigma}^2_{\bar{p}_{e,d}} = \frac{\bar{p}_{e,d}\left(1 - \bar{p}_{e,d}\right)}{|S_{e,1}| - 1} \times \frac{|S_{e,0}|}{|S_{e,1}| + |S_{e,0}|} \tag{I.3}$$

We then estimate the number of SDGW–CWGC fatalities who enlisted in e and resided in d by multiplying $\bar{p}_{e,d}$ by the total number of fatalities associated with enlistment place e:

$$\widehat{N}_{e,d} = \bar{p}_{e,d}\left(|S_{e,1}| + |S_{e,0}|\right) \ , \tag{I.4}$$

where $\sigma^2_{N_{e,d}} = \sigma^2_{\bar{p}_{e,d}}\left(|S_{e,1}| + |S_{e,0}|\right)^2$.

Finally, we estimate military death rates associated with place of residence d by summing the estimated number of fatalities residing in d across all enlistment places in the SDGW–CWGC records, denoted E, and dividing by the number of eligible men in d. Let M_d denote this denominator. The estimated military death rate is:

$$\widehat{\text{Death rate}}_d = c\frac{\sum_{e\in E}\widehat{N}_{e,d}}{M_d}\,, \tag{I.5}$$

where c is the scaling factor and $\sigma^2_{dr_d} = \sum_{e\in E}\sum_{f\in E} Cov\left(\widehat{N}_{e,d}, \widehat{N}_{f,d}\right)/M_d^2$.

The sampling variation in the estimated military death rate for each district d constitutes measurement error. Because this variation arises from sampling uncertainty, we estimate its variance using a bootstrap procedure. Specifically, we draw 1,000 bootstrap samples with replacement from the full set of SDGW–CWGC fatality records. In each replication, we then re-estimate the distribution of reported residence places for every enlistment place and use these estimates to reconstruct the district-level fatality counts and military death rates. For each district d, the variance of its estimated military death rate across bootstrap replications thus provides an estimate of $\sigma^2_{dr_d}$. The resulting estimates of measurement-error variance can then be incorporated into a regression-calibration procedure to correct for attenuation bias.

I.2. Correcting for Attenuation Bias Using Regression Calibration

Once the measurement-error variance of an independent variable is known, the resulting attenuation bias in estimated regression coefficients can be addressed using regression calibration. Because this method is well established, we provide only a brief overview of its mechanics and intuition. Carroll et al. (2006) offer a comprehensive treatment, while Sullivan (2001) discusses its application when measurement error is heteroskedastic, as in our setting.

For exposition, consider first the simplified model $Y_i = \beta X_i^* + \epsilon_i$. The researcher does not observe the true regressor X_i^*, but instead observes $X_i = X_i^* + u_i$, where u_i has mean zero and is independent of X_i^*. It follows that the variance of X_i is the sum of the variance of X_i^* and u_i: $\sigma^2_{x_i} = \sigma^2_{x_i^*} + \sigma^2_{u_i}$. Suppose that the latent regressor X_i^* has a common variance across observations (i.e., is homoskedastic with $\sigma^2_{x_i^*} = \sigma^2_{x^*}$), while the measurement-error variance $\sigma^2_{u_i}$ is known. Then, for a sufficiently large sample of size n, the variance of the latent regressor can be estimated as

$$\widehat{\sigma}^2_{x^*} = \frac{1}{n}\sum\left(X_i - \bar{X}\right)^2 - \frac{1}{n}\sum\sigma^2_{u_i}.$$

This estimate can then be used to calculate, for each observation, the share of the variance in the observed regressor attributable to variation in the true regressor rather than to measurement error:

$$r_i = \frac{\hat{\sigma}^2_{x^*}}{\hat{\sigma}^2_{x^*} + \sigma^2_{u_i}} \tag{I.6}$$

This quantity is known as the "reliability ratio." When measurement error is homoskedastic, so that $r_i = r$, regression calibration is equivalent—in the simplified single-regressor setting—to the conventional errors-in-variables correction obtained by dividing the naive OLS coefficient from a regression of Y_i on X_i by r. Regression calibration produces the same estimate, but by replacing X_i with $X_i^{\text{adj}} = \mathbf{E}[X] + r\,(X_i - \mathbf{E}[X]) = \bar{X} + r\left(X_i - \bar{X}\right)$, and regressing Y_i on this adjusted regressor. The method extends to heteroskedastic measurement error by allowing the reliability ratio to vary across observations and using $X_i^{\text{adj}} = \bar{X} + r_i\left(X_i - \bar{X}\right)$ as the regressor (Sullivan, 2001).

The intuition behind this method is straightforward: observations for which X_i is measured with greater uncertainty are adjusted more strongly toward the mean, effectively decreasing their weight in the regression. Equivalently, X_i^{adj} can be interpreted as the best linear predictor of X_i^* based on X_i. That is, if the latent regressor X_i^* were observed and linearly projected onto X_i, the resulting fitted value would be X_i^{adj} (Carroll et al., 2006, p. 45).

Extending this framework to the multivariate case is also straightforward. Suppose that $Y_i = \beta X_i^* + \alpha Z_i + \epsilon_i$. The researcher observes Z_i, but not the true regressor X_i^*. Instead, the observed regressor is $X_i = X_i^* + u_i$, where u_i has mean zero and is independent of X_i^*. In this setting, regression calibration is based on the variation in X_i^* conditional on Z_i, rather than on its unconditional variation:

$$r_{i|Z} = \frac{\hat{\sigma}^2_{x^*|Z}}{\hat{\sigma}^2_{x^*|Z} + \sigma^2_{u_i}} \tag{I.7}$$

where

$$\hat{\sigma}^2_{x^*|Z} = \frac{1}{n}\sum (X_i - \hat{\gamma} Z_i)^2 - \frac{1}{n}\sum \sigma^2_{u_i}$$

and $\hat{\gamma} Z_i$ estimates the linear projection $\mathbf{E}\left[X_i^*|Z_i\right]$. Because the measurement error has mean zero and is independent of Z_i, this conditional expectation is also equal to $\mathbf{E}\left[X_i|Z_i\right]$. The regression-calibrated regressor is therefore $X_i^{\text{adj}} = \mathbf{E}\left[X_i|Z_i\right] + r_{i|Z}\left(X_i - \mathbf{E}\left[X|Z\right]\right) = \hat{\gamma} Z_i + r_{i|Z}\left(X_i - \hat{\gamma} Z_i\right)$.

J. Constructing Synthetic English and Welsh Départements

This section describes the construction of synthetic spatial units formed by grouping contiguous English and Welsh local government districts into 87 units so as to approximate France's 87 prewar départements. We generate 200 such synthetic mappings, each based on a substantially different partition of districts while matching the population distribution of French départements. We first describe the stochastic spatial partitioning algorithm we use to generate these synthetic mappings (Section J.1), and then assess the extent of variation across the resulting mappings (Section J.2).

The motivation for this construction is to enable a meaningful comparison of estimated effects between England and Wales and France. Although the two countries had similar populations in 1911—40 and 36 million, respectively—the French population was distributed across only 87 départements, compared with 1,721 local government districts in England and Wales. This difference matters insofar as individuals migrated after the war in response to local military death rates: migration across units of analysis generates spatial spillovers, whereas migration within larger units is internalized and therefore does not attenuate the estimated effect to the same extent. Thus, a common spatial support is needed for a meaningful comparison across settings. Since the French data are typically only available at the département level, the comparison must be harmonized at that scale by aggregating smaller spatial units in England and Wales into synthetic equivalents of French départements.[59]

J.1. Stochastic Spatial Partitioning Algorithm

Each of the 200 mappings of synthetic départements is generated in three stages: seed selection (Section J.1.1), contiguous regional expansion (Section J.1.2), and boundary refinement (Section J.1.3). In the seed selection stage, 87 districts are chosen as starting points for constructing synthetic départements. Next, in the contiguous regional expansion stage, unassigned districts are added sequentially until all 1,721 have been assigned to a synthetic département. Finally, in the boundary refinement stage, districts along synthetic-département boundaries are reassigned when doing so improves the match to the French population distribution while preserving spatial compactness.

[59]We are not the first to adopt this type of spatial harmonization. For instance, Rozenfeld et al. (2011) and Dingel, Miscio and Davis (2021) apply a similar logic to cross-country comparisons of cities, where differences in the administrative definition of metropolitan areas hinder comparability. They address this problem by constructing synthetic metropolitan areas from smaller spatial units so as to obtain a more consistent geographical basis for comparison.

J.1.1. Seed Selection

In the seed selection stage, 87 of the 1,721 districts are chosen as starting points for constructing synthetic départements. The five most populous districts are always selected to prevent them from being absorbed into already large synthetic départements during the regional expansion stage, which would worsen population balance.[60] The remaining 82 seeds are selected sequentially and randomly, with selection probabilities favoring both spatial dispersion and larger population size. For each draw, every unselected district i is assigned the weight $w_i = (d_i + 1)\sqrt{P_i}$, where d_i is its distance in meters from the nearest previously selected seed and P_i its 1911 population. The probability that an unselected district i is chosen as the next seed is then proportional to its weight: $w_i / \sum_j w_j$, where j is taken over the set of currently unselected districts.

Once all 87 seeds have been selected, each is assigned a population target that will guide the subsequent regional expansion stage—more precisely, the population target assigned to seed r, denoted T_r, equals the population of its matched département P_d, scaled by the ratio of the total population of England and Wales to that of France. The least populous seed is matched to the least populous département, the second least populous seed is matched to the second least populous département, and so on, with the most populous seed matched to the most populous département.

J.1.2. Contiguous Regional Expansion

In the contiguous regional expansion stage, unassigned districts are added sequentially until all 1,721 have been assigned to a synthetic département. Let R_r denote the set of districts assigned to synthetic département r at the beginning of the current expansion round. The population of synthetic département r at that point is therefore $P_r = \sum_{i \in R_r} P_i$. At the start of each expansion round, the 87 synthetic départements are ranked by their proportional population shortfall, $(T_r - P_r)/T_r$. Proceeding from the largest to the smallest shortfall, each synthetic département is then given an opportunity to expand, so that those furthest below their targets select additional districts first.

During synthetic département r's turn, each unassigned district v adjacent to its current boundary is assigned a score S_{rv} based on four criteria.

1. G_{rv}^{pop} measures how assigning district v to synthetic département r would affect its

[60] The five most populous districts were centered on Birmingham (841,483 inhabitants in 1911), Liverpool (754,523), Manchester (713,235), Sheffield (595,769), and Leeds (473,786). For comparison, the mean district population in 1911 was 20,921 and the median was 8,543. The five most populous départements in 1911 were Seine (4,154,042), Nord (1,961,780), Pas-de-Calais (1,068,155), Rhône (915,581), and Seine-Inférieure (877,383). The mean population across départements was 455,198 and the median was 341,205.

population relative to its target, with $G_{rv}^{\text{pop}} = (|T_r - P_r|)/T_r - (|T_r - P_r - P_v|)/T_r$.

2. N_{rv} denotes the number of districts adjacent to v that are already assigned to synthetic département r.

3. D_{rv}^{centroid} denotes the distance between district v and synthetic département r's current centroid, normalized by a national scaling factor.[61]

4. D_{rv}^{seed} denotes the distance between district v and synthetic département r's original seed district, normalized by the same scaling factor.

We then define S_{rv} as:

$$S_{rv} = 6G_{rv}^{\text{pop}} + N_{rv} - 2D_{rv}^{\text{centroid}} - D_{rv}^{\text{seed}} + \epsilon \,, \tag{J.1}$$

where $\epsilon \sim \mathcal{N}\left(0, 0.1^2\right)$. Among the unassigned districts adjacent to synthetic département r, candidates receive higher scores when they bring the synthetic département closer to its population target and share boundaries with more districts already assigned to it. By contrast, a candidate's score decreases with distance from both the synthetic département's current centroid and its original seed. Finally, the random term ϵ introduces stochastic variation, helping generate distinct partitions across mappings.

After scoring all unassigned districts adjacent to synthetic département r, the five highest-scoring candidates—or all candidates when fewer than five are available—are placed into the selection set K_r. One district is then drawn from K_r, with selection probability:

$$\Pr\left(v \text{ is selected } | v \in K_r\right) = \frac{\exp(S_{rv}/0.2)}{\sum_{k \in K_r} \exp(S_{rk}/0.2)}. \tag{J.2}$$

Once a district has been added to synthetic département r, the next synthetic département in the round's ordering follows the same procedure. The process continues until all 87 synthetic départements have had an opportunity to expand. A new expansion round then begins, and the procedure is repeated until every district has been assigned to a synthetic département.

[61] This national scaling factor is defined as $\sqrt{\left(\sum_i ||c_i - \bar{c}||^2\right)/1{,}721}$, where c_i denotes the centroid of district i and $\bar{c}$ denotes the mean centroid across all districts.

J.1.3. Boundary Refinement

In the boundary refinement stage, districts along synthetic-département boundaries are reassigned when doing so improves the match to the French population distribution while preserving spatial compactness. During this stage, boundary districts are examined in random order. For each one, the algorithm evaluates whether it should be reassigned from its current synthetic département to an adjacent one. A reassignment occurs only when both conditions are satisfied: removing the district leaves the donor synthetic département contiguous and the move improves the following objective function:

$$Q(M, D, C) = M + 0.5D + 0.01C. \tag{J.3}$$

In this objective function, M measures the mean squared proportional deviation of synthetic département populations from their targets with $M = \left[\sum_{r=1}^{87} ((P_r - T_r)/T_r)^2\right]/87$. Next, C denotes the share of all district edges—where an edge is a pair of districts sharing a boundary—that are separated by a synthetic-département boundary. This term penalizes irregular boundaries, thereby encouraging smoother and less fragmented synthetic départements. Finally, D measures centroid-based compactness by penalizing synthetic départements whose component districts are spatially dispersed around the synthetic-département centroid. Let A_i denote the area of district i, c_i its centroid, $\bar{c_r}$ the centroid of synthetic département r, and $\bar{c}$ the mean centroid across all districts. The compactness measure D is proportional to the following expression:

$$\frac{\sum_{r=1}^{87} \sum_{i \in r} A_i ||c_i - \bar{c_r}||^2}{\frac{1}{1{,}721} \sum_{i=1}^{1{,}721} A_i ||c_i - \bar{c}||^2} \tag{J.4}$$

Once all boundary districts have been evaluated, a new refinement round begins, since earlier reassignments may have created new boundary districts and thus new candidates. The procedure is repeated until either no district is reassigned in a round or ten refinement rounds have been completed, at which point the algorithm terminates.

J.2. Variation Across Synthetic Mappings

We generate 200 synthetic mappings using the stochastic spatial partitioning algorithm described above, so as to produce a distribution of estimated effects of military death rates on female labor force participation. However, such a distribution would be informative only if the 200 mappings differed substantially from one another—if the algorithm instead produced minor variations on only one or a few underlying partitions, the resulting estimates would be clustered around a small number of values.

We therefore assess the variation across the 200 mappings by comparing all 19,900 pairwise mapping combinations. For each pair of mappings, we calculate the proportion of pairs of districts grouped together in at least one mapping that are grouped together in both. We find a mean and median similarity of only 26 percent. Moreover, even the most similar pair of mappings—displayed in Figure J1—shares only 33 percent of its co-grouped pairs of districts.

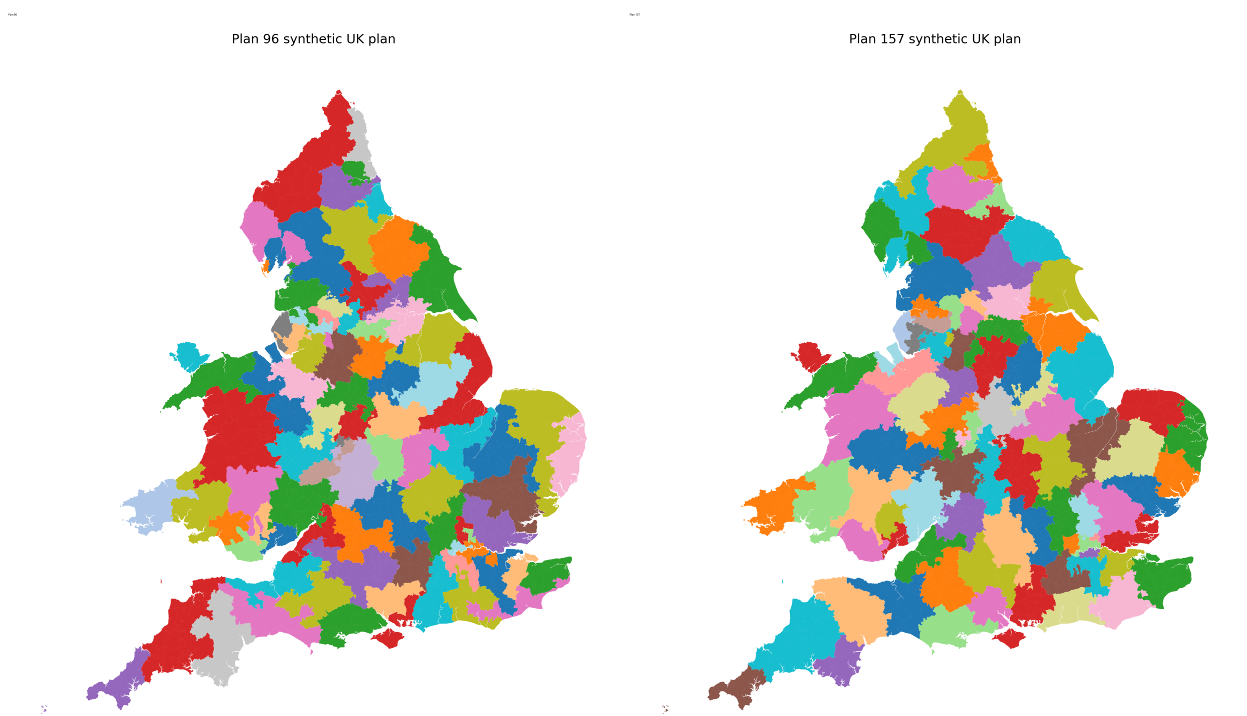


Figure J1. Most Similar Pair of Synthetic Mappings

Notes. This figure displays the most similar pair among the 19,900 pairwise combinations of the 200 synthetic département mappings. Similarity is measured as the proportion of pairs of local government districts grouped together in at least one mapping that are grouped together in both.

K. Pensions for War Widows in Britain and France

This appendix describes and compares the pension systems for war widows in postwar Britain and France. We first compare baseline pension entitlements and benchmark them against female earnings in each country. We then show that several additional features of the British system made British pensions more generous still.

Baseline Pension Entitlements In France, the pension law of March 31, 1919 (art. 19) entitled war widows of regular soldiers to an annual pension of 800 francs.[62] By contrast, from 1921 onward, British war widows under 40 were entitled to a minimum weekly pension of 20 shillings (or £1), while those aged 40 and over received at least 26.67 shillings (26s. 8d.) (Ministry of Pensions, 1921*b*, p. 5).

Given that the franc-to-pound exchange rate fluctuated between 45 and 60 francs per pound in 1921, a direct conversion implies that the French pension for a childless widow amounted to only about 26–34 percent of the British pension for widows under 40. Such a conversion, however, does not adequately capture differences in the extent to which pensions affected widows' labor supply decisions across the two countries. To provide a more meaningful comparison, we therefore benchmark each country's pensions for war widows against the wage a widow could expect to earn from full-time employment.

Pensions Relative to Female Earnings Following Boehnke and Gay (2022), we use average hourly female wages in the textile industry across 186 French cities in 1921 as a proxy for the labor income of a working French woman. Based on this measure, the pension for a childless war widow amounted to about one-quarter of a working woman's income.

To provide a comparable benchmark, we also relate the British pensions for war widows to female textile earnings in 1921. We estimate these earnings in two steps. First, the Ministry of Labour's October 1924 earnings inquiry reports average weekly earnings of 28.6 shillings for women employed in textiles (Bowley, 1937, Table XI, p. 51).[63]

No comparable inquiry exists for 1921. However, the Ministry of Labour's 1928 *Abstract of Labour Statistics* reports changes in female wages relative to their 1914 levels for both 1921 and 1924 in three major segments of the textile industry—Yorkshire piece-workers, Yorkshire time-workers, and Lancashire time-workers—allowing us to backcast the 1924 benchmark to 1921 (Ministry of Labour, 1928, pp. 98–99). Female Lancashire time-workers earned about 19.83 shillings (19s. 10d.) more per week than their unreported prewar wage in 1921, compared with 15.58 shillings (15s. 7d.) more in 1924. Applying this difference

[62]Elements regarding the French case are drawn from Boehnke and Gay (2022, Appendix Section J.3).
[63]For comparison, average weekly earnings for women across all industries were 27.5 shillings.

to the 1924 average weekly wage of 28.6 shillings therefore yields an estimated 1921 wage of 32.85 shillings. The Yorkshire figures produce similar estimates: female piece-workers earned 88 percent more than their prewar wage in 1921 and 63.75 percent more in 1924, implying a 1921 weekly wage of 32.83 shillings ($\simeq 28.6 \times 1.88/1.6375$). Female time-workers earned 110 and 79.75 percent more than their prewar wage in 1921 and 1924, respectively, implying a corresponding estimate of 33.41 shillings ($\simeq 28.6 \times 2.10/1.7975$).

These estimates thus imply that the British pension replaced about 60 percent of the average female textile wage ($\simeq 20/33$), compared with about 25 percent in France.

Additional Sources of Pension Generosity in Britain If anything, the comparison above understates the difference between the British and French pensions for three reasons. First, as already noted, childless British widows aged 40 and over received an even larger pension—26.67 shillings per week, equivalent to about 80 percent of their estimated full-time earnings.

Second, although both countries provided more generous pensions to widows with dependent children, the child supplements were generally larger in Britain. A British widow under 40 received an additional 16.67 shillings per week for her first child under 16, 7.5 shillings for her second, and 6 shillings for each subsequent child—equivalent to increases of 83.4, 37.5, and 30 percent, respectively, relative to the childless pension of 20 shillings. In France, by contrast, each child under 18 increased a widow's pension by 300 francs per year, or 37.5 percent relative to the childless pension of 800 francs (Ministry of Pensions, 1919, pp. 4–5). Given these schedules, British widows with one to eight dependent children received larger proportional increases relative to the childless pension than their French counterparts—only among widows with nine or more children was the proportional increase larger in France.[64]

Third, the British pension amounts described above were only minimum rates. Britain also operated an "Alternative Pension," under which a widow could opt to receive 106.67 percent of her deceased husband's nominal prewar income instead of the standard pension, up to a maximum of 50 shillings per week if she was childless and 66 shillings otherwise (Ministry of Pensions, 1919, p. 4; 1921*b*, p. 5).[65] As of March 1921, 17.7 percent of the 174,215 registered pensions for war widows were calculated under this alternative scheme (Ministry of Pensions, 1921*a*, appendix VI).

[64]For instance, a British widow with five children received $20 + 16.67 + 7.5 + 3 \times 6 = 62.17$ shillings per week, or about 311 percent of the childless pension of 20 shillings. In France, a widow with five children received $800 + 5 \times 300 = 2{,}300$ francs per year, or 287.5 percent of the childless pension of 800 francs.

[65]More precisely, the pension was calculated as two-thirds of 160 percent of the husband's prewar income, with the 160 percent adjustment intended to account for wartime inflation.

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