Discussion Paper Series A No.743

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What Happened in Hometown When the Soldiers Never Returned?

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June 2023

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The Consequences of Hometown Regiment What Happened in Hometown When the Soldiers Never Returned?*

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Sometimes, war results in a large gender imbalance in certain cohorts and areas that changes the path of economic development. However, there is ambiguity around this notion because the market economy has a strong restoring force. This study contributes to the existing literature by presenting the Japanese experience during the Second World War. Japan lost approximately 2 million soldiers during 1938-1945. Furthermore, the loss of young males concentrated in certain cohorts of certain geographical areas owing to hometown regiment system. By exploiting the variation of changes in gender balance cohort-by-prefecture, we examined the effect of the loss of young males on the post-war industrial structure. We observed that the reduction in the gender ratio may have led to slower industrialization, although to a limited extent quantitatively.

I. War Damage and Economic Recovery

War affects the industrial structure not only through damaging the physical capital stocks but also through the tremendous loss of human capital. In Japan's case, besides the fact that it lost 2 million soldiers due to the unexpected mismatch of military strategies between offense and defense, most casualties were males who were concentrated in certain areas and cohorts in the country due to the institution of hometown regiment. Applying the difference in gender imbalances between geographical areas and cohorts because of war, this study empirically examines how permanent loss of human capital relates to economic recovery. The regression results show that the permanent loss of males may have led to slower industrialization and a tentative increase in agriculture. However, quantitatively, such

* We thank Yusuke Inoue and Masayuki Ikeda for their encouragements. The authors thank Akiko Suwa-Eisenmann, Jérôme Bourdieu, Lionel Kesztenbaum, Kensuke Teshima, Junichi Yamasaki, David Weinstein, and Edward L. Glaeser for their constructive feedbacks. We are also grateful for the helpful comments and discussions from the participants of seminars at the NBER-Japan Project Meeting Fall 2022, Asian and Australian Society of Labour Economists 2022 conference, the University of Tokyo, Tohoku University, Nihon University, Japan Economic Association Meeting, Paris School of Economics, Ecole des Hautes Etudes en Sciences Sociales, and Hitosubashi University. All errors are our own. This study was financially supported by KAKENHI 19K21688, and 19H00592.

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slow-down effects were limited and gradually disappeared about 15 years after the war. These findings imply that gender may have augmented average technology during high-speed economic growth after the war. In the medium term, technological change and internal migration may have neutralized the permanent loss of human capital in certain geographical areas.

II. Brief Literature Review

To examine the stability of the market economy, the literature has considered the loss of resources during the war as an exogenous negative shock to the economy. Many researchers have examined questions such as whether the economy returns to the expected trajectory from the pre-war era or to what extent the speed of recovery depends on the magnitude of damage. For example, Craft (1995) interpreted that the "Golden Age" of European economic growth between 1950 and 1973 was a catch-up stage from the aftermath of war. More recently, Milionis and Vonyo (2015) also found the resilience of economic growth, using data from 57 countries between 1950 and 1980, developed by the Madison project.

However, the research is still inconclusive since the OLS regression from post-war variables on wartime loss easily suffered from bias due to unobservable country effects. For example, considering military or political strategy, wartime loss can be determined by factors that also characterized the post-war economy, such as the development of industry. Milionis and Vonyo (2015) proposed the instrumental variable approach to exploit exogenous variation in wartime loss using the differential proximity of individual countries in the Second World War. [1]

Some literature extended the study using cross-sectional (regional) variation within a country to control for the long-term effect of unobservable regional factors. These studies focused on the efficiency of the reallocation process after the war than aggregated economic growth. Davis and Weinstein (2002) studied the resilience of the distribution of the city size using the destruction of physical capital due to air attacks in Japan as an instrument for the changes in the population density of a region during the war. They noted that the city size recovered 15 years after the war, indicating the strong resilience of the market economy. Brakman, Garrentsen and Schramm (2004) corroborated this, using the German bombing experience to compare the differences between West and East Germany. Their finding that the reversion in the distribution of city sizes existed only in West Germany provided additional evidence about the resilience of the market economy. Conversely, Giorcelli and Bianchi (2021) argued that not only the market mechanism but also the policy intervention may have contributed toward recovery. They studied bombing destruction in Italy and found the direct effect of the Marshall Plan on the reconstruction of the Italian economy.

[1] Specifically, the distance between each country's capital and the closest location of a major battle during the Second World War is used as an instrument for the country's post-war output gap (see the geographical IVs in economic history, section 3.2).

Although physical capital can be restored through markets or governments, damage to human capital permanently affects the economy. In the United States, as most mobilized soldiers returned home, human capital in society was temporarily affected during the war. Historians such as Chafe (1972) argued that the work experience of females, although temporal during the war, permanently changed their skills and attitudes toward the labor market, resulting in a change in the labor supply after the war. Goldin (1991) statistically examined this view based on the retrospective work histories of about 4,000 women collected by Gladys Palmer. She reported that the transition matrix on employment status between pre- and post-war suggests that wartime experiences did not increase women's employment. However, Acemoglu, Autor and Lyle (2004) found that wartime mobilization shifted female labor supply after the war. They defined the mobilization rate as the fraction of registered males between 18 and 44 years who were recruited for the entire war period, in other words, a single value for each state. They used it as an explanatory variable with the interaction of the postwar dummy.

As mobilization rates vary cross-sectionally, the OLS estimator is easily impacted by state-specific unobservable factors, which require remedying the issue of endogeneity.[2] Further, Goldin and Olivetti (2013) showed other evidence, using population censuses in the United States, and found that the shift in labor supply occurred only in the educated group.[3] Studies in the United States indicate that heterogeneous responses played an important role in the female labor supply after the Second World War owing to a sufficiently industrialized economy. Boehnke and Gay (2022) extended the literature by introducing the French experience during the First World War, where the industries had not matured enough to differentiate workers. Furthermore, the loss of more than a million men during the war led to persistent wartime damage. Using the difference in casualties among regions (drtements in France), they showed that the gender imbalance, created by the permanent loss of soldiers, increased the female labor participation. [4]

Research conducted in the United States and France indicated the importance of identifying the causes and phases of social development after the war, suggesting

[2] They used two strategies to address the endogeneity. First, they used age and ethnic structure at the state level in 1940 as instruments for mobilization rates (Table 7). The other strategy is to decompose the mobilization rates into the economy-related component and the economy-independent component. The control related to the economy is imputed by (i) the fractions of men who were farmers and were non-white in 1940 and the average schooling of men in 1940 (Table 5) and (ii) the share of the industry in 1940 (Table 6).

[3] Goldin (1991) has already found heterogeneity among occupations.

[4] Their findings are also consistent with other literature that argues demographical gender imbalance increased female participation in labor markets in the long run. For example, Teso (2019) examined the change in gender balance in sub-Saharan Africa due to slavery trade. Grosjean and Khattar (2019) discussed the case of Australia, where most male convicts were sent. However, as pointed by Qian (2008) and Carranza (2014), it is difficult to identify shocks on gender balance from the change in the structure of the labor market because such shocks are gradual and continue in the long run. Specifically, gender imbalance can affect long-term supply conditions through marriage and other demographic changes (Abramitzky, Delavande and Vasconcelos, 2011; Grossbard, 2015; Brainerd, 2017; Ogasawara and Komura, 2022).

the importance of the demand side economic mechanism. Acemoglu, Autor and Lyle (2004) interpreted the mobilization as a shifter in labor supply and estimated the demand elasticity for males with respect to the female supply. They found that females are imperfect substitutes, except for high-school graduate males. This is consistent with the heterogeneous responses of female employment. Rose (2018) modified the mobilization data to capture missing men more precisely and found that the demand shift was the primary driver of the increase in female employment. Interestingly, he did not find strong effects on manufacturing, but a weak overall effect masks slightly faster growth in durable manufacturing employment and declines in jobs in non-durable goods industries like apparel and textiles, particularly for white women. Shatnawi and Fishback (2018) also estimated the Goldin-Katz type demand curve shift, for females in the United States, by using business data at the establishment level during the war. Cardoso and Morin (2018) is particular for highlighting the role of the employer. They used cohort-region-level data from Portugal after the 1980s and examined the effects of changes in gender ratio on labor market outcomes, such as employment. As an instrument for the gender ratio, they used the cohort-region variation about war casualties between 1961 and 1974 in colonies of Portugal. They found that the increase in female employment accompanied with narrowing pay gap. They interpret it as the change in employers' social norm on female employment.[5]

This study contributes to the literature by providing additional evidence on the economic effects of wartime damage in a broader perspective. We examine the impact of the male soldier loss on the industrial structure, which has not been extensively investigated in the literature. We provide comprehensive analyses on how the labor market absorbs the negative shock to the gender composition of the local labor market, by examining how this change in the demand side is related to the labor force participation of women. Moreover, as wartime damage indicated the decline in the proportion of males relative to females (gender imbalance), it was a permanent and large loss of gender-specific human capital for the economy, which is different from the cases of the United States. Lastly, the gender imbalances varied not only in regions but also in cohorts due the historical context. Therefore, we effectively control for the potential regional confounders, such as pre-war economic structure, by taking the difference between cohorts.

III. Hometown Infantry Regiments and the War in Pacific Ocean

During the Second World War, the Imperial Japanese Army (IJA) implemented the hometown regimental system that organized each infantry regiment by recruiting males from certain cohorts, e.g. aged 20-21, who lived in certain geographical

[5] The consequences of wartime experience are also examined at the individual level, mainly in the interest of health and welfare economics. For example, Cesur, Sabia and Tekin (2013) discussed that military combat on the ground during the war on terrorism leads to a higher probability of suicidal ideation and PTSD. In historical context, Costa and Kahn (2003) and Costa and Kahn (2007) argued the effect of the Civil War in the United States.

areas. IJA dispatched regiments to defend the islands scattered in the Pacific Ocean, and the defensive force of the islands sometimes lost all its war potential, facing the Allies' fierce counterattack. However, IJA did not expect that the Allies would pass through some islands without any combat. This fact indicated that, in Japan, certain geographical areas lost men massively in certain cohorts, but other areas and cohorts did not.

A. Construction of Imperial Japanese Army

First Difference: conscription system. — The Japanese government had the mandatory draft system before 1945. Each male (20 years) was examined for conscription, usually between April and August.[6] In 1944/45, the threshold was lowered to 19 years. Therefore, in 1944, the examination covered two cohorts at the same time. For entering the military service, the examination classified males into five categories: "Kou," "Otsu," "Hei," "Tei," and "Bo." Only the first two categories (Kou or Otsu) were appropriate for the active service. The last two categories indicated unsuitability for the military service, namely disabled (Tei) or sick (Bo). The evaluation was solely based on physical and health conditions and not on the cognitive skills and economic conditions of the participants. It was argued that members of wealthy families or society's elite could evade conscription; however, after the major revision of the conscription institution in 1927, there remains no evidence of bias. Thus, the yearly shares of the first category "Kou" were stable between 30% and 40%, toward the latter half of the 1920s. The actual number of mobilizations was calculated from the military operation plan determined by the General Staff Office in Tokyo.[7] Each regiment chose males among those classified as "Kou" at first and then from "Otsu."

The above conscription rule implies that a man's probability of having military service largely depends on the number of mobilizations at the age of 20 (or 19). And the number of mobilizations increased rapidly, especially after 1943. A rough estimate shows that the number of drafted males directly after the examination was 170K in 1937, 320K-360K during 1938-1943, 1,000K in 1944, and 500K in 1945. Therefore, the chance for the active military service concentrated in certain cohorts: age 20 in 1943, age 19 and 20 in 1944, and age 19 in 1945. Based on the above estimate, given that the male population was around 700K for one cohort at age 20, almost half of the cohort was drafted at age 20 during 1938-1943, and around 70% of the cohort was drafted in 1944/45, which was approximately 25% in 1937.

Second Difference: two principles of the hometown regimental system. — In IJA, an infantry regiment was constructed at the prefecture level, choosing 3K-4K males

[6] Oyama (1943), p.4. Current age on the last day of November in the previous year. The exact time of examination varied among prefectures. The prefectures in Japan correspond to states in the United States, provinces in Canada, and regions in France.

[7] The procedure of mobilization was determined by The Ordinance for Mobilization Plan for the Army.

aged 20-21. [8] When IJA exhausted the generation, they organized a backup regiment recalling men over 22 years of age.[9] On December 1941, IJA mobilized 51 divisions that comprised 160 infantry regiments.

Generally, the regimental system has two principles. First, it functions as an independent frontline military unit responsible for recruiting new soldiers, equipping weapons, and training the troops. Second, it works as an administrative unit in peacetime. Once drafted, most Japanese males belonged to a certain regiment and their military service record was registered in the regiment (heiseki).[10] Even after retiring from active service in IJA, they maintained their relationship with the registered regiment and were invited for ceremony sometimes. Furthermore, as the regiment organization was based on geographical territory, the hometown community closely accompanied it. Therefore, the regiment was called hometown regiment (kyodo rentai) in Japan.

The regimental system was the major principle of organizing a modern army all over the world, until the beginning of the First World War. The famous example is Kantonsystem in Prussia. After the elaboration by Gerhard von Scharnhorst, then by Hans von Seeckt, it was inherited to Wehrmacht.[11] The British Army also had a long tradition of the regimental system for more than 100 years. Edward Cardwell and Hugh Childer reorganized it by 1881, clarifying its territorial principle.[12] The Third Republic of France reorganized its regimental system by the law of 24 July 1873. They divided the continental territory into 18 self-contained Army Corps Regions. Boehnke and Gay (2022) discussed that French Army was organized by the territorial regimental system at the beginning of the First World War. Until the First World War, a standing strength of the United States Army was organized only as territorial regiments. Specifically, divisions and brigades were the temporal organizations during the wartime and were dissolved when the war ended.[13]

The territorial regimental system has several advantages. First, it develops l'esprit de corps by recruiting members of the same community.[14] Especially, when the main infantry tactics in the battlefield was the bayonet charge (for example, the Napoleonic War, the American Civil War), the most essential element of war fighting was to keep the group morale/discipline of infantry.[15] Second, it may prevent the Army from carrying out coup d'Etat. For example, the fragmentation of the regimental system prevented the British Army from being the politically destabilized factor.[16]

[8] Usually 1-99 numbered regiment.
[9] Usually 100 to 299 numbered regiment.
[10] However, the Imperial Japanese Navy had a centralized registration system.
[11] Bartov (1991) pp.30-35. In this book, he used the word "primary group" rather than Kantonsystem and discussed how the primary group was destroyed after 1942, especially in Chapter 3.
[12] French (2005) Ch.1.
[13] McGrath (1947) pp.45-47.
[14] Costa and Kahn (2003)
[15] French (2005) Ch.1.
[16] Strachan (1997) pp.195-214.

However, the territorial regimental system started disintegrating since the middle of the First World War. First, without conscription, the regiment could not afford the necessary wartime capacity only from its territory, as was the case for Britain and the United States.[17] Second, even with conscription, non-negligible regiments were unable to renew their strength by support only from its territory because of the tremendous loss of soldiers, horses, and supplies like France. [18] Third, the infantry tactics in the battlefield changed from mass charge to small-group fire and movement. Small-group tactics utilized several divisions (artillery, engineers, medic, telegraph, machine guns, and so on) even at the level of regiment that required the soldiers with specific skills and experiences. The territory of the regiment was too small to retain enough soldiers for each division. Therefore, the United States, Britain, France, and the Soviets abolished or substantially changed the territorial regiment system as their primary organizational scheme of the army. They introduced the centralized system at the national level during the First World War. [19]

In contrast to the global trend, IJA maintained the hometown regimental system even around 1940. Additionally, the infantry regiment remained the main military power in IJA through the Second World War. Poorly mechanized, IJA increased infantry branch to approximately 50% of total strength; whereas the United States Army decreased the share to 20%. In particular, in the last year of war, about 70% to 80% of newly drafted soldiers were assigned to infantry under the hometown regimental system.

B. Heterogeneous causalities among theatres

With the conscription system, the hometown regimental system of IJA may have lost a certain number of males of certain prefectures if the regiment is completely destroyed. Unfortunately, the particular situation around IJA during the Second World War actually concentrated the losses to certain regiments; due to the heterogeneous opponents due to long front line and the leapfrogging strategy taken by the United States Forces.

Long front line. — At the end of the Second World War, IJA scattered about 3 million soldiers from the North-East China, Continental China, Indo-China, Thailand, Burma, Malaysia, Java, Borneo, Philippine to New Guinea. As the intensity of fighting depended on the confronting opponents, the casualties of troops varied by their garrison. A report by the Ministry of Welfare in 1963 estimated the ratio of the number of killed in action (KIA)/missing in action (MIA) to the number of soldiers at the end of the Second World War by the theatre.[20] They varied from 0.07 in the North-East China to 3.92 in Philippines,

[17] French (2005).
[18] Bourlet (2010).
[19] King (2013) Ch.6.
[20] Total of the Army and the Navy. Estimated KIA/MIA is the accumulated number since 1937 for Continental China, while the number for other theatre is estimated as is since 1941.

whereas the overall average accounted 0.27.[21] The probability of being killed or missing was on average 21% in the Japanese military force, assuming that troops did not move from their garrison until the end of war. The most severe theatre was Philippines accounting for 80%, followed by 70% in the central Pacific Islands, 67% in Burma, 61% in the south Pacific Islands, and 7% in the North-East of China. These ratios indicate that the high-intensity fighting concentrated in the islands in the Pacific Ocean since the counterattack by the Allies was mainly by the United States Force.

Some argue that IJA allocated the regiment with a particular economic attribute (for example, agricultural area) to face the United States Force. However, reportedly, there is no evidence regarding IJA intentionally allocating the regiments based on the hometown attribute; even IJA could not anticipate the severity of a particular battlefield because of the strategy adopted by the United States Force.

the Pacific islands front and leapfrogging. — During the Second World War, the Japanese forces occupied the south, southwest, and central Pacific Islands to impede the communication and transportation between the United States and Australia. The first objective of the United States Force was to secure the approaches to Japan by expelling the Japanese forces from these islands for which they adopted a strategy called "Leapfrogging". Samuel Morison explained it as "instead of invading every island which held a Japanese garrison, we bypassed the strongest concentrations such as Rabaul, Truk, and Wewak; landed amphibious forces on beaches comparatively free of the enemy; built an airfield; and, using our sea supremacy to seal off the bypassed enemy garrisons, left them to 'wither on the vine'." [22]

Thus, among the Pacific Islands, some islands were directly hit by the United States Force, while some islands were bypassed. IJA did not anticipate the targets of attack, thereby isolating substantial troops. Thus, without naval and air supremacy, isolated troops were neutralized until the end of the war, keeping their soldiers away from severe warfighting.

Example of Palau. — Japan began colonizing the Palau Islands after the Treaty of Versailles in 1919. The Japanese government established one of their local branches in Koror and provided public services to nearby islands.[23] At the beginning of Japanese domination, the total population of islands was approximately 54,358 in 1923, which doubled to 113,562 by June 1939.[24]

On April 1944, anticipating counter attack by the United States Force, IJA dispatched the 14th division (14D) to defend the Palau Islands. The 14D constituted

[21] 212.2M/788.9M.
[22] Morison (1968) p.7.
[23] South Sea Agency
[24] The average annual growth rate was approximately 4.7% for 16 years, while it was 2.6% in Taiwan from 1925 to 1940 and 1.3% in the mainland Japan from 1925 to 1940.

three infantry regiments, and each infantry regiment had three battalions; the second infantry regiment (2iR; Ibaraki, capacity 3,166), the 15th infantry regiment (15iR; Gunma, capacity 3,964), and the 59th infantry regiment (59iR; Tochigi, capacity 3,166). The location of 2iR and III/15iR and I/59iR was Peleliu Island and Angaur Island, respectively. The remaining four battalions of 15iR and 59iR defended the Koror Island (main island of Palau). For Yap Island, 14D located the 49th independent mixed brigade (49MBs; 8 independent infantry battalions, capacity 5,591). Lastly, the 53rd independent mixed brigade (53MBs; 6 independent infantry battalions, capacity 4,263) was assigned to the mainland of Palau. After occupying the Mariana Islands on July 1944, the next target of the United States Force was the Philippines. As the Palau Islands were located between the Mariana Islands and Philippines, the United States Force had to occupy it to serve as the staging base to recapture the Philippines.[25]

On 15 September 1944, the US 1st Marine Division began landing on Peleliu Island. After 74 days, the United States Force successfully expelled the Japanese forces from the island destroying 2iR, III/15iR and II/15iR (which counter landed from Palau as reinforcement). Of 6,822 soldiers, only 190 survived. The US 81st Infantry Division invaded Angaur Island 2 days after the landing of 1st Marine Division. After 31 days, approximately 1,194 soldiers of I/59iR (and 6 soldiers of Navy) lost their lives and 59 were captured as prisoners of war (POWs). However, the United States Force left the main island of Palau and Yap. After the war, about 25,000 soldiers from the mainland of Palau and 5,500 soldiers from Yap returned to Japan (both the Army and Navy). KIA in these areas was estimated at 770.

The Breaking Jewel. — The particularity of the Pacific theater was the small number of POWs relative to KIA/MIA. Generally, if defeated in a war, a substantial number of soldiers were captured as POWs by the opponents; some surrender and the wounded are left behind. For example, the German 6th Army during the Second World War, after losing 147,200 soldiers, left 90,000 as POWs when General Paulus surrendered at the end of January 1943 at Stalinglad; the ratio of KIA to POWs was 1.6 at one of the most dreadful battlefields in Europe. Contrastingly, this ratio was 42.5 and 19.2 in the battle of Peleliu and Angaur Islands, respectively.

The small number of Japanese POWs may have been due to the mental training by their military code. On January 1941, the Japanese force revised their military code and stressed the culture of shame. Since they defined surrender as shame, even after exhausting their arms, and punishment to their families, Japanese soldiers avoided surrender under the hometown regimental system. While Ruth Benedict, an anthropologist who worked for United States Office of War Information, contrasts it as "shame culture" to "guilt culture" in the United States,

[25] The distance between Saipan Islands in Mariana and Leyte Gulf in Philippine is more than 2,000 km. The air-cover by P-51 from Saipan cannot reach the Philippines directly.

this is a general mechanism reinforced by the hometown regimental system in the battle field.[26]

Therefore, the hometown regiment that faced the counterattack by the United States Force in the Pacific Islands was likely to lose its soldiers, which means that specific prefecture lost a specific cohort of males massively, relative to other cohorts and prefectures. Ibaraki prefecture lost the entire infantry regiment in the battle of Palau, while neighboring Gunma prefecture and Tochigi prefecture lost two-thirds and one-third, respectively. Approximately, the number of missing men was 3K, 2K, and 1K in Ibaraki, Gunma, and Tochigi, respectively. They lost 21.2%, 17.4%, and 10.1% of one cohort of male in a single battle in Palau islands, and this variation indicated the decline in the gender ratio of each prefecture.[27]

C. Its consequences in Japan: decline in gender ratio

During the Second World War, approximately 2 million Japanese soldiers died. This loss was massive not only from the humanitarian perspective but also from the economic viewpoint, as the economy lost nearly 5% of the working population at that time, representing a large decline in the male-female ratio. Figure 1 shows that the gender ratio (number of men divided by number of women) of the population aged 15-64 years dropped from 1.014 to 0.939 between 1940 and 1947.[28] Notably, the gender ratio was permanently imbalanced, even though the population continued to grow, led by the post-war baby boom.

We interpret this decrease in the gender ratio as a gender-specific loss of human capital. Since a comprehensive list of war casualties is not available in Japan, we cannot directly use death counts by cohort-prefecture and instead approximate them by the gender ratio changes. Nevertheless, it is natural to use the gender ratio change in our reduced-form regressions, as we are interested in how the death of soldiers affected the economic outcomes through the change in the composition of local human capital, that is gender ratio. One caveat of using gender ratio is that when it is defined area by area, it is also affected by social migration between areas. However, the change in gender ratio reasonably approximates the death of male soldiers. First, the decrease in the total young male population between 1935 and 1947 imputed from the cross-prefecture variation of the gender ratio is 1.75 million. This is close to the estimated 2 million soldier deaths from the official statistics.[29] Second, in a case study in Iwate prefecture, where the local authority created the list of war casualties, we find that roughly 70%-80%

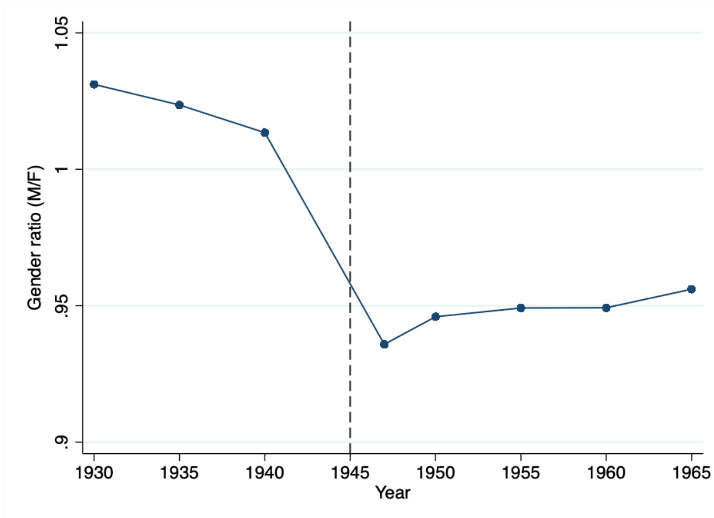
[26] Benedict (1946). Costa and Kahn (2003) discussed it during the Civil War in the United States.

[27] According to the number of examinations for conscription, which is the number of males aged 20, average cohort sizes of males are 14155.5, 11500.8, and 9858.5 in Ibaraki, Gunma, and Tochigi, respectively, between 1940 and 1943.

[28] A small recovery in the gender ratio between 1947 and 1950 is due to the return of Japanese migrants overseas such as from the Korean Peninsula and mainland China, commonly known as Hikiage-sha. Ministry of Health, Labour and Welfare estimates the number of these returnees to be more than 3 million, in addition to the return of approximately 3 million soldiers.

[29] We calculated the estimate decline in young male population in the following way: First, for each prefecture, we estimated the decline in male population of the treated cohorts (aged 20-34 in 1945) by

Figure 1. : Evolution of gender ratio at national level



Note: The figure shows the gender ratio (defined as the number of men divided by number of women) of population aged 15-64 in Japan, calculated from the published Census reports. The gender ratio declined from 1.014 to 0.939 between 1940 and 1947. A small recovery in the gender ratio between 1947 and 1950 is due to the return of Japanese migrants overseas.

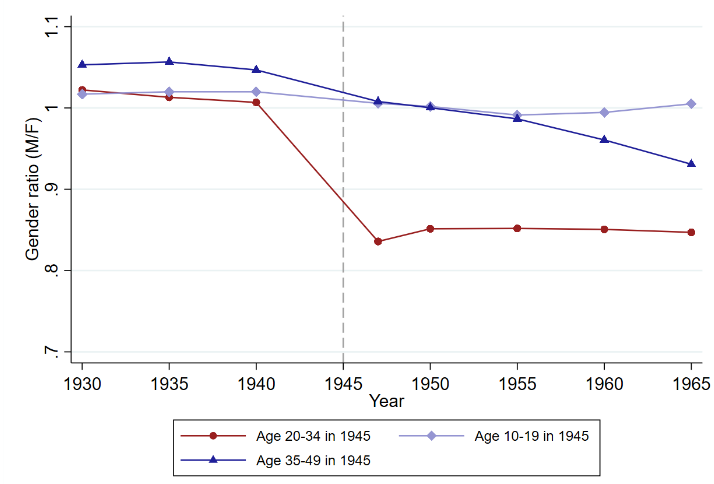
of the change in the gender ratio is consistent with the death of male soldiers. Third, while casualties from air raids could also lead to a human capital loss, they were fewer in number (approximately 0.3 million) and gender neutral. Fourth, we will check the robustness of our approximation later, by using a new prefecture-level approximation constructed from Image Information Retrieval System for the Code-Number List of Adjudication Notice for Condolence Payments.

The erosion of the gender ratio was particularly strong for the younger cohorts heavily recruited for the war. Figure 2 shows the evolution of the gender ratio across three different birth cohorts (aged 10-19, 20-34, and 35-49) in 1945. The gender ratio sharply drops for the cohorts aged 20-34 in 1945 between 1940 and 1947: the decline as large as -0.17, going from 1.01 to 0.84. In contrast, the gender ratio of the two other cohorts was hardly affected, as they were either young or old enough to avoid military recruitment.

To compare the experience of Japan with the victorious countries of the Second World War, Figure 3 shows the evolution of the gender ratio by birth cohort for the United States and France. Note that the figure for the United States includes

multiplying population in 1935 of these cohorts by their change in the gender ratio and divided by 2. Then, we aggregated these estimates for all prefectures. This aggregate estimate of male disappearance can be significantly different from the estimated number of soldier deaths (from the official source) if there is systematic migration between large and small prefectures. Our estimate is 1.75 million, close to official statistics, which confirms that the change in gender ratio approximates the death rate relatively well.

Figure 2. : Evolution of gender ratio, by cohort



Note: The figure displays the evolution of the gender ratio of three cohorts groups, those aged 20-34 in 1945 (treated cohorts), and aged 10-19 and 35-34 in 1945 (control cohorts), between 1930 and 1965. The figure indicates that the decline in the gender ratio concentrated among the young cohorts heavily recruited for the war.

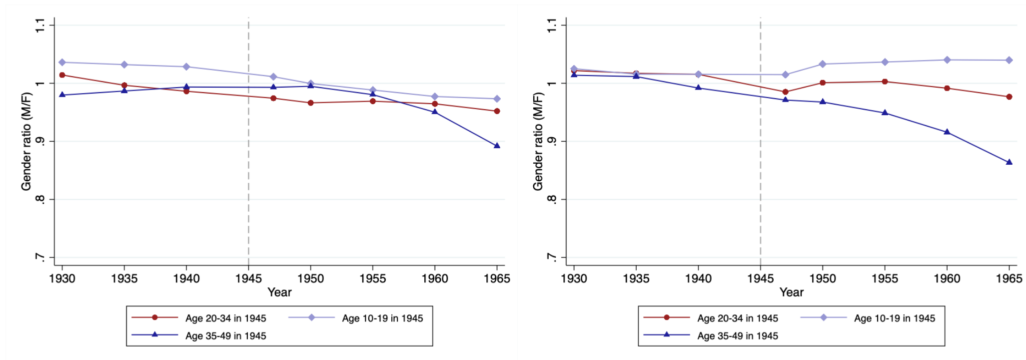
military overseas. The first panel indicates there is no discernible decrease in the gender ratio for any of the cohorts in the United States. In France, the gender ratio drops temporarily for the cohorts aged 20-34 in 1945, but it recovers almost to the pre-war level by 1950, as detained soldiers and evacuees return. The figure demonstrates that a large and permanent decline in the gender ratio is a particular experience of countries that lost in the Second World War, including Japan.

We exploit a substantial geographical variation in the magnitude of gender ratio changes. This is a combined result of the hometown regiment system of the IJA and US military strategy. Figure 4 provides the changes in the gender ratio between 1935 and 1947 by prefecture for the three cohort groups. Among the treated cohorts who were aged 20-34 in 1945, on average across 46 prefectures, the gender ratio declined by 0.17, with a standard deviation of 0.05. The largest decline is 0.31 in Nagasaki prefecture, while the smallest is 0.08 in Hyogo prefecture. [30] Note that the variation in the gender ratio change is not systematically related to factors such as urbanity. For example, Tokyo's gender ratio decline was 0.16 (27th of 46 prefectures), while it was 0.23 (7th) for Osaka. Furthermore, neighboring prefectures also had different experience; within Shikoku islands, the change in gender ratio varied from Kagawa prefecture's 0.24 (5th) to Ehime prefecture's 0.17 (21st). Thus, Figure 4 confirms that the gender ratio change is a

[30] The variation does not change much when measured by percentage changes that take into account the initial differences in the gender ratio in 1935.

Figure 3. : The evolution of the gender ratio in victorious countries

(a) United States (b) France



Note: This figure repeats Figure 2 for the United States and France, displaying the evolution of the gender ratio of three cohort groups, those aged 10-19, 20-34, and 35-34 years in 1945, calculated based on the population census of each country. The figure for the United States includes military overseas. Thus, there is no temporary decline in the gender ratio during the war. Compared to Japan, there is virtually no permanent and large decline in the gender ratio.

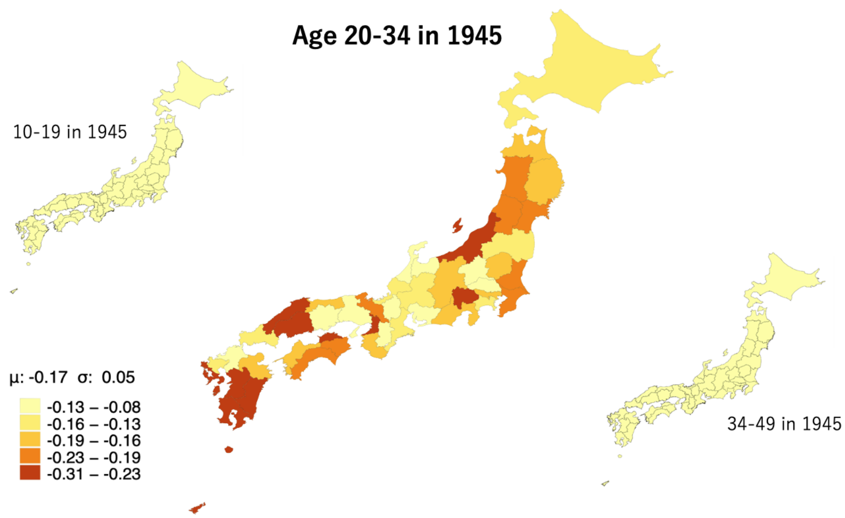
unique experience of the treated cohorts. If the change in the gender ratio is entirely due to prefecture-specific factors such as migration, we can expect a similar trend for the other cohorts. The two figures (on either side) on the changes in the gender ratio of the two control groups confirm that this is not the case.

IV. Data and Sample for Analysis

Data. — Our main objective is to examine the impact of human capital loss on industrial share of employment, which should be computed for each prefecture-age-industry-gender-year combination. To obtain these variables, we use published tables of population census from 1920 to 1980, available for every 5 years except for 1945, which was replaced by the concise version in 1947. [31] While population census report headcounts by prefecture-age-gender combination, employment status, and worker counts are consistently reported only from 1955, by 1-digit sector.[32] Although, prior to 1955, it is available decennially for 1920, 1930, and 1940, the definitions of employment status used in the pre-war census were different from the post-war census, making it difficult to make meaningful pre- and post-war comparisons. For this reason, we restricted the years 1955-1980 for

[31] The microdata of population census in Japan is available only for those after 1990.
[32] For the other years, they are reported by prefecture-gender or by age-gender.

Figure 4. : Geographical variation in gender ratio changes



Note: The figure shows the variation in the gender ratio changes across 46 prefectures. The map in the center classifies prefectures by the size of 1935-1947 change in gender ratio of the treated cohorts aged 20-34 in 1945, where the threshold values are based on quantiles. The mean change in the gender ratio of the treated cohorts is 0.17, with the standard deviation of 0.05. The side figures supplement the gender ratio changes of control cohorts aged 10-19 and 34-49 in 1945, respectively, where the same threshold values are applied.

our main regression estimation, and pre-war employment information is only used for supplementary analyses and robustness checks. We digitized this information for 46 prefectures (excluding Okinawa) and 11 age groups (10-14, 15-19,..., 60-64). We use the terms "age" and "age group" interchangeably.

We aggregate 1-digit industries into three large industry categories for simplicity: agriculture, manufacturing, and service. Agriculture combines 1-digit sectors of agriculture, forestry, and fishery. Manufacturing consists of 1-digit sectors of mining, construction, and manufacturing. Lastly, service includes all other sectors (in other words, retail and wholesale, finance, real estate and insurance, transport, communication and utility, service, government service, and others). In Appendix, we also provide the results based on 1-digit industry classifications. Employment rate and labor force participation rate by gender are calculated for each prefecture-age combination similarly.

Sample and descriptive statistics. — Our estimation sample is an unbalanced panel of prefecture-cohort pairs from 1955 to 1980. We follow the eight birth cohorts that were aged 10-49 in 1945. Among these birth cohorts, the three cohorts whose age was between 20 and 34 in 1945 correspond to our "treated groups," whereas the other five cohorts aged between 10-19 and 35-49 in 1945 are termed "control groups." The cohorts leave the sample after the age 60-64.

Table 1 provides summary statistics of our sample. It first shows the mean and standard deviation of our estimation sample in which prefecture-cohort is the unit of observation. On average, prefecture-cohort population size is slightly above 100K, with the gender ratio of 0.90 in 1955-1980 without any population weighting. Mean employment share (without population weighting) is 0.38 for agriculture, 0.24 for manufacturing, and remaining 0.38 for service. Employment rate is 0.75 on average, while mean labor force participation rate is 0.58 and 0.95 for women and men, respectively.

To give a sense of the industrial characteristics of our studied period, Table 1 also provides the prefecture-level information. Average population size across prefectures grew by approximately 50% between 1955 and 1980. The period, well known as "high speed growth era," is characterized by rapid industrialization. On average, the agricultural share of employment dropped from 0.46 in 1955 to 0.14 in 1980, while that of the manufacturing share and service sector increased from 0.20 to 0.33 and 0.33 to 0.53, respectively. Therefore, the impact of the loss of young males is examined in the specific context of whether the industrialization was partially impeded by the loss of young male labor.

Table 1—: Descriptive Statistics

	Prefecture-cohort	Prefecture	
	1955-1980	1955	1980
Population (in thousand)	123 (112)	1,002 (756)	1,521 (1,428)
Gender ratio	0.90 (0.09)	0.92 (0.04)	0.94 (0.04)
Industry Share of Employment			
Agriculture	0.38 (0.16)	0.46 (0.14)	0.14 (0.07)
Manufacturing	0.24 (0.09)	0.20 (0.07)	0.33 (0.06)
Service	0.38 (0.09)	0.33 (0.07)	0.53 (0.05)
Employment and Participation			
Employment rate	0.75 (0.09)	0.76 (0.05)	0.76 (0.04)
Labor force participation rate (female)	0.58 (0.13)	0.58 (0.10)	0.58 (0.07)
Labor force participation rate (male)	0.95 (0.06)	0.93 (0.02)	0.92 (0.02)
Observations	1748	46	46

Note: The table provides summary statistics of our sample. It first summarizes sample means and standard deviations of population, gender ratio, and outcome variables with prefecture-cohort as a unit, pooled over the period of 1955-1980 (second column). The last two columns provide prefecture-level information to characterize the developmental transitions of our studied period. No population weight is applied in computing means and standard deviations.

V. Empirical Strategy

A. Conceptual framework

Variable definition and baseline econometric model. — We use the change in the gender ratio as a measure of a gender-specific human capital loss. For each prefecture-cohort combination, we define the gender ratio as:

$$Ratio_{cjt} = \frac{M_{cjt}}{F_{cjt}}$$

where, M_{cjt} and F_{cjt} are, respectively, the number of men and women for a cohort c, in a prefecture j, in year t. The value is less than 1 when there are fewer men than women. For each prefecture-cohort pair, we calculate the change in the gender ratio before and after the war:

$$\Delta Ratio_{cj} = Ratio_{cj,1947} - Ratio_{cj,1935}$$

where we take 1935 and 1947 as reference pre-war and post-war years. [33] When a cohort in a prefecture loses more men from the war, $\Delta Ratio_{cj}$ becomes more negative.

We are interested in the effect of the change in the gender ratio on own cohort outcomes:

$$Y_{cjt} = \alpha + \beta \cdot \Delta Ratio_{cj} + \varepsilon_{cjt} \tag{1}$$

where Y_{cjt} is a post-war outcome ($t \geq 1955$) and ε_{cjt} is the error term. Given the exogeneity of $\Delta Ratio_{cj}$ (between c and between j), the estimate β provides the average impact of the change in the gender ratio during the war on the economic outcome of the own cohort after the war. This is a difference-in-differences estimator that compares the outcomes of treated cohorts and control cohorts, between prefectures that lost more vs. less men in the treated cohorts. Note that Equation 1 is essentially a reduced form of the structural relationship where the outcomes in 1955-1980 are determined by the contemporaneous gender ratio, which is influenced by the past change in the gender ratio during the war.

Impact on industrial structure. — Focusing on the employment shares of industry as outcomes, we empirically discuss the economic mechanism behind Equation 1. Specifically, the permanent loss of male soldiers would reduce the employment

[33] We used 1935 as the pre-war reference year, instead of 1940, because Japan was already in the Second Sino-Japanese War from 1937 and military deployment had already commenced.

share of manufacturing, which uses male labor in production relatively more intensively, hindering post-war industrialization. In theory, however, it is ambiguous whether the loss of young men would alter the industrial structure. The Rybczynski theorem from the international trade literature predicts that, based on the classical Heckscher-Ohlin model, the economy responds to an exogenous decline in endowment of one factor by decreasing the output of the sector that uses that factor relatively more intensively (Rybczynski, 1955). Nevertheless, labor market may also absorb the labor supply shock through the change in production technology or input mix within sector especially in the long run, theorized as directed technological change (Acemoglu, 1998, 2002) or optimized production factor choices (Beaudry and Green, 2005).

These theoretical predictions have been empirically tested. For example, Dustmann and Glitz (2015) confirmed that both mechanisms of between- and within-sector adjustments were effective in the context of immigrant influx, but found a larger role of within-sector adjustments. The decomposition method in the international trade literature led some to a similar conclusion (Blum, 2010).[34] Overall, the extent of the change in factor supply absorbed through between- and within-sector adjustments depends on parameters such as substitutability between production factors and considered economic contexts.

Applying the tentative conclusion of literature, this study hypothesizes that the gender comparative advantage plays a key role in our context, encompassing both endowment changes (total female and male population) in the trade literature, and labor supply changes (total female and male workforce) in the labor literature. Our study departs from the previous literature in some respects. First, the large and permanent supply shock was realized in a very short period, rather than gradually.[35] Second, although the role of the demand side on female labor supply was acknowledged in the related literature, the analysis in a unified empirical framework and context has not been implemented. Finally, our context offers an interesting case study in which the change in factor endowment arises from an absolute decrease in one labor input, in contrast to the large body of work that studies the increases of the educated workforce or the influx of immigrants. The economy's response might be asymmetric between gaining and losing of production input, if there is strong inertia or recovering force of the market (for example, invested fixed cost).

B. Exogeneity of the change in the gender ratio

The crucial assumption to identify the causal effect of gender-biased loss of human capital on industrial composition is that the observed changes in the gender ratio were exogenous. Further, we already discussed why the death rates of soldiers were likely to be random across prefectures from the historical and

[34] We thank Kensuke Teshima for informing us of the literature on decomposition methods.
[35] Remember that most soldiers died in the last few years of the war.

institutional perspectives. However, there are a few concerns relative to the plausibility of the exogeneity assumption. First, the variation in the gender ratio might be correlated with the pre-war industrial structure. For example, conscription used physical examination to assess health conditions and physical strength (height-weight ratio) in ranking one's ability as a soldier. If male workers in the agricultural sector were more likely to pass the examination, then prefectures that lost more men could have been relatively more agricultural-intensive even before the war. Second, since the gender ratio change is affected by age-gender-specific migration, a part of the variation may come from the pattern of cross-prefecture mobility in certain ages related to work or education, which could also be correlated with industrial and economic characteristics of the prefecture.

To assess randomness of the changes in the gender ratio, we regress the changes in the gender ratio of the treated cohorts (aged 20-24, 25-29, and 30-34 in 1945) on observed pre-war prefecture-level characteristics. We study how they are correlated with the agricultural share of employment. [36] We do not aggregate the changes in gender ratio at the prefecture level to have more statistical power. Moreover, we do not include other outcomes such as female labor force participation since they are highly correlated with agricultural employment share.

Table 2 shows the results. Column (1) includes only the share of agricultural employment at the prefecture level in 1930 as a regressor. There is a modest negative relationship, indicating that prefectures with an initially higher agricultural share experienced a greater decrease in the gender ratio, although the coefficient is not significant. Since our hypothesis is that male loss during the war led to a higher share in agriculture in the post-war periods, this result raises the concern that any detected negative effect of the gender ratio change on post-war industrial development merely captures the permanent differences in industrial compositions across prefectures.

Column (2) shows that adding urban prefecture dummies and population size makes the coefficient on agricultural share smaller. Similarly, column (3) further adds war-related variables related to recruitment (average height-weight ratio of examined individuals in 1939 conscription) and war damage (log of number of air raid casualties). The coefficient on the agriculture share is even more reduced. In all cases, the estimated coefficients are far from statistically significant. Importantly, the obtained R^2 across the models are small, indicating that a large share of changes in the gender ratio are not correlated with pre-war observable characteristics of prefectures.

The simple regression results above support the exogeneity of the gender ratio changes relative to industrial compositions. Of course, this is only partial evidence, and we are not able to perfectly control for prefecture-level unobserv-

[36] Our digitized data have narrowly defined (1-digit) agriculture share and manufacturing share in employment, but no other sectors. We use agriculture in the explanatory variable because 1-digit agriculture represents more employment share in the broad agriculture sector (defined in Section IV) than 1-digit manufacturing does for the broad manufacturing sector. Note that the employment concept of the pre-war census is not the same as that of the post-war census, requiring caution in interpretation.

ables. This motivates our estimation strategy to take a second difference among cohorts, on top of the cross-prefecture differences. Our regression specification includes prefecture fixed effects, which controls for differences in pre-war industry characteristics and prefecture-specific unobservables, such as gender norms.

Table 2—: $\Delta Ratio_{cj}$ and pre-war characteristics

	Change in gender ratio 1935-47 (treated cohorts)		
	(1)	(2)	(3)
Agriculture employment share in 1930	-0.059	-0.040	-0.031
	(0.050)	(0.088)	(0.109)
Urban prefecture dummy		0.039	0.035
		(0.038)	(0.042)
Log of prime age population in 1935		-0.024	-0.012
		(0.020)	(0.024)
Height-weight ratio (1939 examination)			-0.958
			(1.767)
Log of air raid casualties			-0.005
			(0.005)
Observations	138	138	138
R^2	0.010	0.022	0.032

Note: The table provides the regression result where the gender ratio changes of the treated cohorts (aged 20-24, 25-29, and 30-34 in 1945) are regressed on the observed pre-war prefecture-level characteristics. Outcome variables have cohort-prefecture variations, but explanatory variables vary only across prefectures. Note that employment concept in the pre-war census is not directly comparable to the post-war census and interpretation requires caution. Urban prefecture dummy takes 1 for Tokyo, Osaka, Hokkaido, Hyogo, Aichi, Fukuoka, Kanagawa, and Kyoto and 0 otherwise. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

C. Estimation Strategy

We identify the effect of the gender ratio change on industrial composition and labor market outcomes using the variations among prefectures and cohorts. Our baseline specification takes the following form:

$$
\begin{aligned}
Y_{cjt} = \alpha + \beta \cdot \Delta Ratio_{cj} + \gamma_1 X_{cjt} + \gamma_2 X_{cj} \\
+ Cohort_c + Pref_j + Time_t \\
+ (Cohort_c \times Time_t) + (Pref_j \times Time_t) + \varepsilon_{cjt}
\end{aligned}
\tag{2}
$$

where Y_{cjt} is the outcome of interest of cohort c, in a prefecture j, in year t; $\Delta Ratio_{cj}$ is the gender ratio change from 1935 to 1947; X_{cjt} represents contemporaneous cohort-prefecture controls; X_{cj} is time-invariant cohort-prefecture controls; $Cohort_c$, $Pref_j$, and $Time_t$ are the unidimensional fixed effects; $Cohort_c \times Time_t$ and $Pref_j \times Time_t$ represent, respectively, the cohort-year fixed effects and prefecture-year fixed effects; ε_{cjt} is the error term. The coefficient of interest is β. As explained earlier, it is a difference-in-differences estimator that captures the difference in average outcomes over 1955-1980 between the treated cohorts and control cohorts and among prefectures that differ in the size of the gender ratio change of the treated cohorts.

With respect to X_{cjt}, we include only log of population (or log of number of workers where appropriate). We do not include contemporaneous gender ratio or migration because we consider these factors as mechanisms that magnify or mitigate the effect of human capital loss in the long run. For time-constant prefecture-cohort controls X_{cj}, we add the gender ratio and the log of population in 1935. Moreover, to distinguish the effect of male soldier loss (gender-specific) from general population loss (gender-neutral), such as from the air raids, we also include the change in log of population between 1935 and 1947.

Prefecture-specific factors, such as location specificity, pre-war industrial compositions, and other war damages (for example, capital destruction) are absorbed by prefecture fixed effects. Time fixed effects control for economic conditions during a year, and cohort fixed effects control for cohort-common factors. Prefecture-time fixed effects deal with differential trends in outcomes across prefectures or time-specific shocks that are specific to a prefecture. Lastly, with cohort-time fixed effects, the specification controls for factors such as life cycle.

The only source of potential endogeneity is the correlation of $\Delta Ratio_{cj}$ with prefecture-cohort unobservables. Among others, most concerning is cohort-specific (or age-specific) cross-prefecture migration between 1935 and 1947. This can bias our results if it is correlated with pre-war local industrial composition of employment, when it is different across cohorts (or across ages). For example, young people may have worked temporarily in factories located in a neighboring prefecture or went to a school in larger prefectures just before the war. If such migration is gender-biased, $\Delta Ratio_{cj}$ could correlate with the error term. [37] We deal with this issue in two ways. First, we control for the initial gender ratio in 1935 in the regression to capture cohort-prefecture specific factors influencing the cross-prefecture migration. Second, we perform two robustness checks: one test uses 1930 outcomes as a placebo, and the other uses age-fixed gender ratio change as an alternative measure of $\Delta Ratio_{cj}$ that are robust to prefecture-age specific migration.

[37] The direction of the bias is not clear a priori. Suppose that the gender ratio change of the treated cohorts is more negative for agricultural intensive sector because of temporary migration of young men. However, since we consider differences among cohorts, the direction of the bias depends on how the difference between treated and control cohorts in agricultural employment share (or industrial employment share in general) in the pre-war period in agricultural intensive prefecture compared to the corresponding difference in the manufacturing intensive sector —- it is difficult to say whether this difference in the differences is positive or negative.

VI. Results

A. Industrial composition

Table 3 shows the estimated impact of the gender-specific human capital loss on industrial composition and employment outcomes identified in the specification of Equation 2. Since the gender ratio change is negative for a larger loss of men, the coefficient signs must be transposed to interpret the effect of the gender specific human capital loss.

Table 3—: Effect of the gender ratio change on industry composition and employment

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.052*	-0.019	-0.032	-0.057	-0.159***	-0.013
	(0.027)	(0.033)	(0.021)	(0.037)	(0.061)	(0.053)
ln(pop) change 1935-47	-0.029**	-0.001	0.030**	0.083***	0.146***	0.071**
	(0.015)	(0.019)	(0.013)	(0.024)	(0.042)	(0.028)
Gender ratio in 1935	0.091***	-0.085**	-0.007	-0.042	-0.142**	-0.043
	(0.030)	(0.034)	(0.023)	(0.040)	(0.065)	(0.052)
ln(pop) in 1935	-0.032**	-0.028	0.060***	0.101***	0.184***	0.077***
	(0.015)	(0.020)	(0.012)	(0.022)	(0.043)	(0.023)
ln(worker)	-0.037***	0.133***	-0.096***			
	(0.011)	(0.015)	(0.008)			
ln(pop)				0.037**	-0.025	0.018
				(0.016)	(0.032)	(0.019)
Mean outcome 1955-1980	0.241	0.382	0.377	0.747	0.583	0.948
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.924	0.948	0.922	0.610	0.671	0.320
Observations	1748	1748	1748	1748	1748	1748

Note: The table provides our main results on the effect of the gender-specific human capital loss on post-war economic outcomes, corresponding to the estimated coefficients on the gender ratio change 1935-1947 (first row). Columns (1)-(3) suggests that the loss of men led to lower employment share of manufacturing, and higher share of agriculture and service, although imprecisely estimated. Columns (4)-(6) explore the effect on labor force participation and employment rate. The results show that the female labor force participation significantly increased. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

The first three columns study industrial employment shares. Column (1) shows that the decrease in the gender ratio is associated with a lower share of employment in the manufacturing sector between 1955 and 1980 and is statistically significant at 10% level. On the contrary, columns (2) and (3) indicate that the employment share of agriculture and service sectors increased, although the estimates are imprecise individually. [38] The signs of the coefficients are consistent with the hypothesis that the permanent loss of young men led to a lower employment share of manufacturing sector, that is, lower level of industrial development in the post-war period.

However, the magnitude of the effects is modest: the decline in the gender ratio by 0.1, corresponding to approximately 10% of men disappearing from the local labor market for a fixed number of women, is associated with a 0.52-point decrease in manufacturing employment share, and a corresponding increase in employment share in agriculture and service by 0.19 and 0.32 points, respectively. During 1955-1980, the mean share of employment of the prefecture cohort is 0.38 in manufacturing, 0.24 in agriculture, and 0.38 in service (see Table 1). The largest decline in the gender ratio at the cohort-prefecture level was around 0.37, indicating that the decline in the gender ratio does not account for much share of the cross-cohort-prefecture variation in the sector composition of employment. [39]

B. Labor force participation and employment

Column (5) in Table 3 indicates that the skewed gender ratio has led to a higher female labor force participation rate, consistent with the previous literature. Labor force participation rate is 1.6-point higher for women from a cohort in a prefecture that experienced 0.1 decline in the gender ratio, relative to those whose gender ratio was unaltered. This effect is much smaller than estimates from the previous literature that also exploited a permanent loss of men, finding an approximately 3-4 percentage points increase in female labor force participation rate in response to a gender ratio difference of 0.1 (Boehnke and Gay, 2022; Cardoso and Morin, 2018).

There are several potential explanations for a smaller effect on women's labor force participation. First, limited labor supply was possibly due to child care duties in a high-fertility, post-war period — total fertility rate exceeded 4 before 1949, which gradually decreased until it stabilized at slightly above 2 during late 1950s. Since women in the treated cohorts were aged 20-34 in 1945, our sample period overlaps the stage of the traditional life cycle of women when child care duties could prevent them from supplying labor. This is consistent with Ogasawara and Komura (2022) that showed a significant association between imbalanced sex ratio and higher fertility in the post-war Japan. Second reason may relate to the conservative gender norm. Although the new constitution enacted in 1947 guaranteed the political, economic, or social equality between men and women, and the society was moving toward a progressive gender norm, it is unclear if the institutional barriers and normative biases were dissolved in practice, especially

[38] Notice that the three coefficients add up to 0 as they represent shares.

[39] For example, a simple computation indicates that the most treated prefecture-cohort had 1.9-point lower manufacturing employment share compared to non-treatment. This is only one-fifth of the standard deviation in manufacturing employment share.

relating to the economic participation.[40] The smaller effects on the female labor force participation might be due to the econometric specification that takes cross-cohort differences. Income shock is one of the supply-side mechanisms of increased female labor force supply in response to the permanent loss of men (Boehnke and Gay, 2022). Possibly, the labor supply decision was undertaken within household, whereby other members of the household supplied labor instead of the women in the treated cohorts. In this case, our diff-in-diff estimate attenuated because of spillover effect across cohorts. [41] This interpretation is consistent with the supplementary results in the following section, where we find a larger effect on female labor force participation in the specification that controls for the change in the gender ratio of the other cohorts, and the change in the gender ratio of the other cohorts in the same prefecture increase labor force participation of the own cohort. [42]

In contrast, there is no significant association with the male labor force participation rate as shown in column (5), and the estimated coefficient is very small. Column (4) indicates that the employment rate increased when the gender ratio reduced, possibly reflecting an improvement in labor market tightness due to the scarcity of male labor, although the coefficient is not statistically significant.

C. Change in gender mix within sector

The change in the composition of human capital can be absorbed by the change in gender mix within sector. Faced with the shortage of male labor in the local labor market, employers may adjust their production technology or input mix to accommodate female labor. Even without the change in technology, hiring female labor may intensify if there is sufficiently high substitutability between two genders in each sector. To test these hypotheses, we study the impact of the gender ratio change on female share in each sector, using the same specification in Equation 2.

Table 4 summarizes the results. In each sector, the female share is higher when there is a larger loss of males due to the war. For example, female share in manufacturing is 0.95 points higher in cohort-prefecture where the gender ratio declined by 0.1 during the war. Unfortunately, it is difficult to compare the magnitude of the effect across sectors because we do not have a benchmark female shares

[40] Interestingly, Okuyama (2021) finds that the GHQ's radio program promoting Japanese women's social, political, and economic rights increased women's political participation in the post-war period, but not labor market participation in 1950. This is consistent with the existence of various barriers for women's economic participation.

[41] The post-war period had a relatively high share of multi-generation household (over one-third of all households in 1955, although it decreased to one-fifth by 1980).

[42] Labor demand side mechanisms would rather facilitate female labor force participation since the industry moved toward more female intensive sectors and, as discussed later, all sectors increased female share in their labor input. Hence, the change in the demand side led to an increased estimated effect on the female labor force participation. In addition, military pension for bereaved family does not explain the lower effect because it did not exist until 1952 and was not generous enough to negatively affect labor supply decisions on the extensive margin.

in each sector in the pre-war period based on the same employment definition. Nevertheless, the coefficients suggest that the gender ratio change explains large share of female share variation within-sector than the variation in sector share of employment studied in Table 3. [43]

Table 4—: Effect of the gender ratio change on within-sector gender composition

	Female Share in Employment		
	Manufacturing	Agriculture	Service
	(1)	(2)	(3)
Gender ratio change 1935-47	-0.095*	-0.130***	-0.050**
	(0.050)	(0.025)	(0.022)
ln(pop) change 1935-47	0.014	0.098***	0.065***
	(0.030)	(0.019)	(0.015)
Gender ratio in 1935	-0.077	-0.129***	-0.079***
	(0.052)	(0.028)	(0.025)
ln(pop) in 1935	0.062**	0.099***	0.073***
	(0.030)	(0.018)	(0.015)
ln(worker)	0.024*	0.019*	0.006
	(0.014)	(0.011)	(0.009)
Mean outcome 1955-1980	0.261	0.524	0.369
3-way fixed effect	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes
Cohort-year FE	Yes	Yes	Yes
R-squared	0.909	0.946	0.949
Observations	1748	1748	1748

Note: The table provides the effect of the gender-specific human capital loss on female share within each sector. Negative and significant coefficients on the first row indicates that all sectors used female worker more intensively when they lost young male working force. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Overall, the results provide new empirical evidence on the demand-side factors that facilitate the integration of women into the labor market when there is a negative shock to the availability of male labor. In the case of Japan, due to the loss of men during the war, the economy shifted toward sectors that used relatively more female labor (agriculture and service) than the male-intensive sector (manufacturing). Meanwhile, within all sectors, the presence of women increased. These changes in the demand structure should have increased the economic return of entering the labor market for women. However, the estimated increase in female labor force participation rate was rather small, indicating the existence of obstacles and counteracting factors that prevented women from supplying labor.

[43] For example, for manufacturing sector, a reduction in the gender ratio by 0.37 leads to a higher female share by 3.5 points, two fifths of the standard deviation of the female share in manufacturing in our sample. This is almost twice what the change in gender ratios explains in terms of variation in the manufacturing share of employment (one-fifth of the standard deviation) in Table 3.

D. Long-run effects

It is unclear whether the impact of gender-specific human capital loss is permanent or temporary. On the one hand, the initial "shock" can determine the path of the local industrial development and, therefore, the effect of the male human capital loss persists, or even propagates, over time. However, the shock can also be transitory and neutralized in the long run through, for example, cross-prefecture migrations and entry of new birth cohorts into the local labor market.

We study these alternative hypotheses by looking at the time-varying effects of the change in gender ratio in the following specification:

$$\begin{aligned} Y_{cjt} = {} & \alpha + \beta \cdot (\Delta Ratio_{cj} \times I_t) + \gamma_1 X_{cjt} + \gamma_2 X_{cj} \\ & + Cohort_c + Pref_j + Time_t \\ & + (Cohort_c \times Time_t) + (Pref_j \times Time_t) + \varepsilon_{cjt} \end{aligned} \tag{3}$$

where I_t is an indicator for time periods.

Figure 5 shows the estimated time-varying coefficients and their confidence interval for different outcomes. The top panels indicate that it was only until 1960 that the decline in the gender ratio had a negative impact on the employment share of the manufacturing sector, resulting in a higher share of the agriculture sector. However, the effects gradually disappear afterward. These results suggest that the loss of male workers hampered industrialization initially, but this negative impact attenuated subsequently in the catching up process.

The bottom middle panel shows the time-varying effects on female labor force participation. The impact of permanent male loss had always been an increased female labor force participation that lasted after 30 years since the end of the Second World War. The coefficient becomes large enough to be significant only after 1965. This pattern is consistent with the life cycle: the treated cohorts were aged 20-34 in 1945, so women from these cohorts were constrained by housework and child care early in the post-war period.

Figure 6 shows the time-varying effects of the gender ratio on within-sector female shares. In all three sectors, the increase in female share has been a persistent phenomenon, although it is not always significant. Focusing on manufacturing and agriculture sectors, this is in stark contrast to the results found in Figure 5, that is, even though the impact of the gender ratio on the industrial structure has disappeared in the long run, more intensive use of female labor within sectors has remained in the long run.

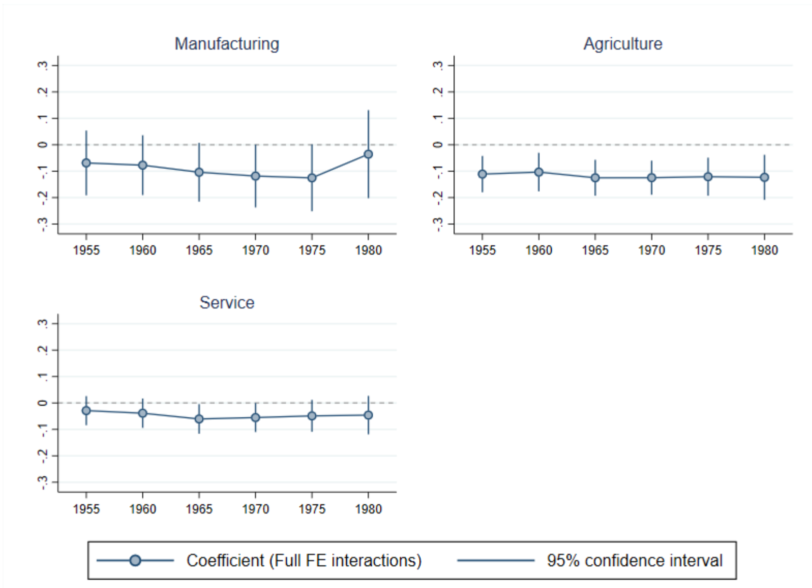
The controversy remains in the United States on how the Second World War transformed women's economic life: specifically, if the temporary absence of men during the war have had a positive and persistent impact on women's labor force even after the war, and what are the economic areas (in terms of occupations and industries) in which female representation changed. We find that, in the

Figure 5. : Long-run effect on industrial share and employment outcomes

Note: The figure provides the long-term effects of the gender ratio changes on industry share of employment (top three panels) and employment and labor force participation (bottom three panels). These are estimated using the interaction of year dummies with the change in gender ratio in the specification of Equation 3, which include full set of fixed effects. For each outcome, we provide the coefficients and 95% confidence interval. The figure shows that the declined gender ratio led to lower (higher) employment share of manufacturing (agriculture) up to 1960, after which the effect decays to null. The effect on service sector is gradual. While somewhat imprecisely estimated, the effect on female labor force participation is persistent over time.

case of Japan where many soldiers did not return, the loss of men increased female labor force participation permanently in the post-war period. Moreover, in the short-run, the economy shifted toward the sectors that used female labor more intensively, as seen in the tentative increase in agriculture sector and decline in manufacturing sector. However, in all sectors, female representation has increased and the effect persisted in the long run, even during the era of massive industrialization and high-speed economic growth.

Figure 6. : Long-run effects on within-sector female share



Note: The figure provides the long-run effects of the gender ratio changes on female share in manufacturing, agriculture, and service sector. These are estimated using the interaction of year dummies with the change in gender ratio in the specification of Equation 3, which include full set of fixed effect. For each outcome, we provide the coefficients and 95% confidence interval. The figure shows that increased female share was relatively a persistent phenomenon.

VII. Discussion

A. Prefecture-age-specific effect does not matter

Our identification assumption is that the change in the gender ratio is not systematically correlated with prefecture-cohort specific unobservables, once added controls and fixed effects. One major threat is the prefecture-specific age effect. People migrate across prefectures, and such migration can be prefecture-age specific and gender-biased. For example, young men in rural prefectures might temporarily relocate to urban prefectures for jobs or schools. In this case, some of the variation in the change in the gender ratio between 1935 and 1947 could be endogenous to prefecture-age specific migration.

Placebo test. — To see whether such prefecture-age effects bias our estimates, we run a placebo test using the data from 1930, the closest pre-war census for which labor force participation is available at prefecture-cohort level. [44] We run the same regression specified in Equation 2 on the outcomes in 1930 of those whose age corresponds to the age in 1945 of our sample cohorts who were aged 20-49 in 1945.

Table 5 shows that there is no statistically significant association between the change in gender ratio between 1935 and 1947, and labor force participation in 1930. While the signs of the coefficient for female labor force participation is negative, the coefficient is not as large as estimated in our main regression results. Although this placebo test remains only suggestive, the results are consistent with the exogeneity assumption of the gender ratio changes.

Alternative definition of gender ratio. — As discussed above, it is possible that $\Delta Ratio_{cj}$ changed due to prefecture-age specific migration, which could correlate with prefecture cohort unobservables, because we defined $\Delta Ratio_{cj}$ as the change in the gender ratio from 1935 to 1947 of the same cohort, and a weakness of this definition is that the age of the cohort at the reference year (1935) matters. In addition to the placebo test, we performed robustness checks by introducing an alternative measure of the gender ratio change robust to the prefecture-age specific unobservables. Instead of comparing the gender ratio in 1947 to that in 1935 of the same cohort, we compared it with the gender ratio in 1935 of the same age group. Formally, for each cohort-prefecture pair, we construct:

$$\Delta Ratio_{cj}^{al} = Ratio_{a(c)j,1947} - Ratio_{a(c')j,1935}$$

where $a(c)$ refers to the age group of a cohort c and $a(c')$ is the age group of another cohort c'. The first term of the RHS is the same as before, that is,

[44] In the pre-war census, definition of employment was different and it was a closer concept to the post-war labor force participation. We use the pre-war employment rate as the labor force participation rate in these placebo tests.

Table 5—: Results of placebo test using 1930 outcomes

	Outcome in 1930		
	Total LFPR	Female LFPR	Male LFPR
	(1)	(2)	(3)
Gender ratio change 1935-47	-0.023	-0.076	-0.014
	(0.052)	(0.108)	(0.025)
ln(pop) change 1935-47	0.083***	0.244***	-0.063
	(0.027)	(0.064)	(0.039)
Gender ratio in 1935	-0.005	0.003	-0.022
	(0.051)	(0.108)	(0.021)
ln(pop) in 1935	0.048**	0.065	-0.000
	(0.023)	(0.051)	(0.018)
ln(pop)	0.009	-0.012	0.024
	(0.014)	(0.026)	(0.016)
Prefecture FE	Yes	Yes	Yes
Cohort FE	Yes	Yes	Yes
Year fixed FE	No	No	No
R-squared	0.963	0.966	0.888
Observations	276	276	276

Note: The table shows the results of placebo test where the gender ratio changes are regressed on some of the labor force participation rate in 1930 of the same age groups. Standard errors are clustered at prefecture level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

$Ratio_{a(c)j,1947}$ is the gender ratio of a cohort c in a prefecture j in 1947. The second term, however, is now the gender ratio of another cohort c' whose age group in 1935 is the same as that of the cohort c in 1947.[45] Since this alternative measure fixes age rather than cohort, it is not contaminated by prefecture-age specific factors affecting the gender ratio, based on the assumption that these factors were constant between 1935 and 1947.

Table 6compares the coefficients of the two alternative gender ratios estimated from the base specification with full fixed effects. In general, the signs and qualitative implications are similar. Using the alternative gender ratio change, the effect on the manufacturing share in employment and agriculture is large and statistically significant at 1% level, while the coefficient on the service sector becomes null.[46] With respect to employment outcomes, the effect of female labor force participation becomes smaller, but the coefficient is still negative and remains significant at 10% level.

Table 6—: Comparison of coefficients with the alternative gender ratio change

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Coefficients from the model with:						
Change in gender ratio (own-cohort)	0.052*	-0.019	-0.032	-0.057	-0.159***	-0.013
	(0.027)	(0.033)	(0.021)	(0.037)	(0.061)	(0.053)
Change in gender ratio (age-fixed)	0.088***	-0.088***	-0.000	-0.035	-0.079*	-0.020
	(0.016)	(0.019)	(0.010)	(0.23)	(0.042)	(0.024)

Note: The table provides the estimated effects of the gender-specific human capital loss when the alternative measure of the gender ratio change is used, which is robust to prefecture-age specific unobservable. The first row provides the coefficients from our main results using the gender ratio change fixing the cohort. In the second row, we provide the estimated coefficients on the alternative measure of the gender ratio change that fix age rather than cohort. Specification is the same for both (Equation 2). Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B. Control for other cohort

In our main regression framework, we assume that $\Delta Ratio_{cj}$ independently affects own cohort outcome. However, through indirect mechanisms such as substitution, the gender ratio of other cohorts may affect own cohort outcome. If there

[45] For example, for a cohort who was aged 30-34 in 1945, this alternative gender ratio change was calculated by the difference between the gender ratio in 1947 when they were aged 32-36, and the gender ratio in 1935 of the other cohort aged 32-36 in 1935. Previously, the comparison was made against the gender ratio of the own cohort in 1935, when this cohort was at age 20-24.

[46] Given that some young male workers (typically sons other than the eldest who did not take over the family agriculture business) moved to neighboring prefectures to work temporarily in factories located in neighboring prefectures, it seems important to control for the age-prefecture effect.

are correlations between the gender ratio change across cohorts within a prefecture, this might bias the estimate of the own cohort effect Furthermore, cross-cohort effect provides some insights in itself. We control for $\Delta Ratio_{c'j}$ (change in gender ratio of other combined cohorts) in the base specification:

$$\begin{aligned}
Y_{cjt} = \alpha + \beta \cdot \Delta Ratio_{cj} + \theta \cdot \Delta Ratio_{c'j} + \gamma_1 X_{cjt} + \gamma_2 X_{cj} \\
+ Cohort_c + Pref_j + Time_t \\
+ (Cohort_c \times Time_t) + (Pref_j \times Time_t) + \varepsilon_{cjt}
\end{aligned} \tag{4}$$

with $\Delta Ratio_{c'j}$ representing the gender ratio change of all other cohorts c' in the same prefecture.

Table 7 shows the results. The coefficients of the own-cohort gender ratio change have same signs as the main results, but the magnitude is relatively large: the 0.1 decline in the gender ratio led a prefecture cohort to have a 5.2-point lower share in manufacturing employment and 3.8-point and 1.4-point higher for the agriculture and service sector, respectively. The coefficient on female labor force participation are also higher.[47]

The increase in own-cohort effect is linked to the reduced coefficient of log of population in 1935 (which declined from -0.034 to -0.009 in case of manufacturing). For its average value of 11.7, it translates to a 30 ppt difference in manufacturing share, while it is 26 ppt for the other cohorts' gender ratio, whose mean value is -0.08. Overall, controlling for the correlation across cohort does not alter the sign of the coefficient, while the magnitude implies that the own-cohort was possibly even larger than originally estimated.

C. Attention to prefecture-level variation

Our main specification uses industry employment share at cohort-prefecture level to capitalize on the cohort-prefecture variation in the change in the gender ratio. To interpret the estimated coefficient as the effect on the local industrial structure, we implicitly assumed cohort-level segmentation of the labor market. However, if the labor market is not perfectly segmented by cohort and there are between-cohort substitutions, our estimate will be attenuated owing to cross-cohort spillovers within prefecture; in this case, our main specification only provides the lower bound. Here, we estimate the effect of young male loss under the assumption that the local industrial structure is determined by the total labor stock at the prefecture level. This motivates a specification that exploits only the cross-prefecture variation in the gender ratio change.

[47] The effect of other cohorts is even larger because of a much smaller variation due to the consolidation of different cohorts: while mean and standard deviation of the own-cohort change in the gender ratio is -0.083 and 0.093, they are -0.083 and 0.029, respectively, for the other-cohort's change. Moreover, part of the variation comes from the untreated cohorts (as we have three treated and five untreated cohorts), whose gender ratio change might be correlated with the unobservables.

Table 7—: Results of the regression controlling for the rest of cohorts

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.521***	-0.377***	-0.144**	-0.179*	-0.296*	-0.063
	(0.056)	(0.079)	(0.056)	(0.092)	(0.161)	(0.116)
Gender ratio change 1935-47 (rest)	2.805***	-2.135***	-0.669**	-0.737	-0.819	-0.304
	(0.278)	(0.407)	(0.269)	(0.534)	(0.915)	(0.613)
ln(pop) change 1935-47	-0.004	-0.021	0.024*	0.076***	0.138***	0.068**
	(0.014)	(0.019)	(0.013)	(0.023)	(0.043)	(0.030)
Gender ratio in 1935	0.147***	-0.127***	-0.020	-0.056	-0.158**	-0.049
	(0.031)	(0.036)	(0.026)	(0.039)	(0.064)	(0.053)
ln(pop) in 1935	-0.009	-0.046**	0.054***	0.095***	0.177***	0.075***
	(0.014)	(0.019)	(0.013)	(0.022)	(0.043)	(0.025)
ln(worker)	-0.039***	0.135***	-0.096***			
	(0.010)	(0.014)	(0.008)			
ln(pop)				0.038**	-0.023	0.018
				(0.016)	(0.032)	(0.019)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.971	0.977	0.965	0.913	0.900	0.632
Observations	1748	1748	1748	1748	1748	1748

Note: The table provides the results of the specification where the gender ratio change of the rest of the cohorts is included as a control. The gender ratio change of the rest of the cohort is calculated by the gender ratio change of all other cohorts combined in the same prefecture excluding own cohort. The estimated coefficient of the change in the gender ratio on own cohort (first row) is much larger than before, suggesting that our previous estimates attenuated due to cross-cohort spillovers. The estimated effect of the gender ratio change of the other cohorts is large because the size of the gender ratio change is much smaller compared to own-cohort gender ratio changes. All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Specifically, we use the gender ratio change of the treated cohorts (those aged 20-34 in 1945), which varies only across prefectures:

$$\begin{aligned} Y_{cjt} = {} & \alpha + \beta \cdot \Delta Ratio_j + \gamma_1 X_{cjt} + \gamma_2 X_{cj} + \delta Z_j \\ & + Cohort_c + Time_t \\ & + (Cohort_c \times Time_t) + \varepsilon_{jct} \end{aligned} \tag{5}$$

where $\Delta Ratio_j$ is the change in the gender ratio of those aged 20-34 years in each prefecture (and therefore has only a subscript j). In this specification, we cannot include prefecture fixed effects. Our approach to minimize the endogeneity of $\Delta Ratio_j$ is to add appropriate prefecture-level controls, represented by Z_j. Since our treatment variable is now at the prefecture level, we also include cohorts who were less than 9 years old or not born in 1945 in our estimation sample.

Regarding prefecture controls Z_j, we include variables related to military recruitment (average height-weight ratio of the male examined for conscription in 1939), other mechanisms of the war effect (number of bomb casualties per capita, number of destroyed buildings per capita, and prefecture-level log of population change between 1935 and 1947), and urbanity (large prefectures—Tokyo, Kanagawa, Osaka, Aichi, Kyoto, and Hyogo). Moreover, we add the prefecture-level outcome in 1930 (of working population) to control for the permanent prefecture characteristics. We include time-varying prefecture controls, such as age structure (share of those aged below or 19, 20s, 30s, 40s, 50s, and 60 or above, respectively, and contemporaneous log of population or workers). We maintain cohort-prefecture controls such as pre-war controls population and gender ratio in 1935.

Table 8 shows the results. The coefficients on the sector employment share have expected signs and are statistically significant. They are larger than the previous estimates from the regressions that exploit both cross-prefecture and cross-cohort variations (in Table 3). For example, the manufacturing share in employment is 2.6-point higher when the gender ratio declines by 0.1, larger than the effect of 0.5 point from the previous model. This is consistent with the idea that our main specification underestimated the effect of the gender ratio changes due to cross-cohort substitution within the prefecture.

However, the gender ratio changes (of the treated cohorts) at the prefecture level may still correlate with prefecture-specific unobservables because we cannot include prefecture fixed effects in this specification. This concern is reflected in the (wrong) signs of the coefficients on the war-related variables — while we expect negative signs because air strikes and building destruction hinder the development of manufacturing, they enter with positive signs. These signs may rather reflect the endogeneity that bombing was strategic, targeting prefectures with high manufacturing intensities.

If such endogeneity is present, in particular if the gender ratio declined more (less) in prefectures that were already intensive in agriculture and service (manufacturing), the estimated effects of the gender ratio on the employment share in Table 8 provides only an upper bound (in absolute terms). Since both the lower-bound estimates (from Table 3) and the upper-bound estimates (from Table 8) have the same signs, our overall conclusion remains unchanged that the loss of male soldiers negatively affected the industrial development of the prefecture in the post-war period.

On the other hand, the effect on employment and labor force participation is inconclusive. The coefficients in Table 8 are drastically different (both in terms of signs and magnitudes) from the regression results using cohort-prefecture variations. Notably, the gender ratio decline is associated with lower female labor force participation rate.

The direction of the endogeneity bias deserves some discussion. it is opposite: the bias is downward (more negative) for agricultural employment share, whereas it is upward (positive) for female labor force participation. One interpretation is that the direction of bias is different before and after the war. For example, there might be an unobservable related to potential frontier or productivity of the areas, which also correlates with the gender ratio changes. This unobservable is positively correlated with the agriculture sector share (as the largest source of economic income) and predicts higher female labor force participation (due to higher gains from labor market participation) before the war. However, after the war, the correlation reverses for the agricultural sector share as manufacturing becomes the most productive sector. In contrast, the relationship remains unchanged for the participation as the economic motive still give positive incentives for labor market participation. Testing this interpretation requires the measure of potential frontier or productivity of each prefecture, but we do not have such data. Another and more indirect way is to use the prefecture-cohort level outcomes to compare the role of fixed effects before and after the way. We have not digitized the prefecture-cohort outcomes that are available for the 1930 census.

D. Robustness check using prefecture-level death counts

New Data and Artificial Gender Ratio. — As pointed out in III.C, the Japanese government has been reluctant to count and publicize the number of casualties during the Second World War. Recently, the Ministry provided a new prefecture-level table that was constructed from Image Information Retrieval System for the Code-Number List of Adjudication Notice for Condolence Payments. This data is a mechanical extraction, by prefecture, of the number of condolence payment documents listed in the Image Information Retrieval System, which is used by the Ministry of Health, Labor and Welfare as an index to search for original condolence payment documents in its archives. Therefore, the reported numbers are not the numbers of condolence payments by prefecture itself, as some may be missing or duplicated. In addition, the recipients of condolence payments do

Table 8—: Prefecture-level treatment, with pre-war outcome control

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change (treated cohorts)	0.258***	-0.179***	-0.085**	0.134***	0.192***	0.033**
	(0.055)	(0.050)	(0.036)	(0.042)	(0.068)	(0.016)
Bomb casualty per capita	0.524*	-0.485*	-0.082	0.630***	1.080***	0.095
	(0.297)	(0.276)	(0.163)	(0.188)	(0.283)	(0.113)
Destroyed buildings per capita	0.190*	-0.217	-0.071	-0.049	-0.119	0.013
	(0.110)	(0.152)	(0.104)	(0.098)	(0.160)	(0.043)
Height weight ratio	-3.999***	4.200***	0.048	0.657*	1.308**	0.227
	(0.747)	(0.709)	(0.374)	(0.367)	(0.599)	(0.197)
ln(pop) change 1935-47	0.042	-0.221***	0.193***	-0.122***	-0.263***	-0.011
	(0.054)	(0.053)	(0.034)	(0.043)	(0.071)	(0.017)
Gender ratio in 1935	0.090	-0.270***	0.072	-0.182***	-0.235***	-0.017
	(0.063)	(0.068)	(0.049)	(0.050)	(0.082)	(0.017)
ln(pop) in 1935	0.011	-0.163***	0.100***	-0.019	-0.078	0.013
	(0.028)	(0.041)	(0.028)	(0.040)	(0.065)	(0.013)
ln(worker)	-0.016	0.159***	-0.088***			
	(0.028)	(0.041)	(0.028)			
ln(pop)				0.005	0.052	-0.014
				(0.040)	(0.065)	(0.013)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.834	0.933	0.856	0.787	0.771	0.624
Observations	2484	2484	2484	2484	2484	2484

Note: The table provides the regression results where we only exploit cross-prefecture differences in the gender ratio change (of treated cohorts who were aged 20-34 in 1945). While maintaining the same specification as 2, prefecture FE and prefecture-year FE cannot be included. To minimize the endogeneity of the gender ratio change, we include various prefecture-level controls. In columns (1)-(3), the estimated effects on the industrial compositions have expected signs, while the magnitude is larger than our main results, because this specification does not take within-prefecture cross-cohort differences. Column (4) shows a surprising negative effect of male worker loss on female labor force participation when using cross-prefecture variations. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

not necessarily coincide with the war dead according to the eligibility, defined by the Law for the Relief of Bereaved Families of War Victims. Regardless of these limitations, the table can be a direct approximation of the number of deaths by prefecture [48].

Using the newly counted number of death, it is possible to construct a measure of gender ratio changes that is more robust to the cross-prefecture migration. We newly define artificial gender ratio, that is:

$$\Delta Ratio_j^{Art} = \frac{M_{j,1935} - Death_j}{F_{j,1947}} - \frac{M_{j,1935}}{F_{j,1935}}$$

where male population size changes only through soldier deaths (ignoring the cross-prefecture migration) between 1935 and 1947. We continue to use the treated cohorts to construct this measure for each prefecture.

Figure 7 shows the correlation between this artificial gender ratio change and the original gender ratio change. Tokyo and Osaka have exceptionally high (and positive) artificial gender ratio changes because the population of the treated cohorts was already male-skewed in 1935, and there was a large outmigration in particular for men between 1935 and 1947. As a result, $\Delta Ratio_j^{Art}$ is inflated for these two prefectures, because it ignores this outmigration in the numerator. Apart from these two outliers, Figure 7 show a positive correlation between the two measures, confirming that the gender ratio change in population census does largely capture the male soldier deaths.

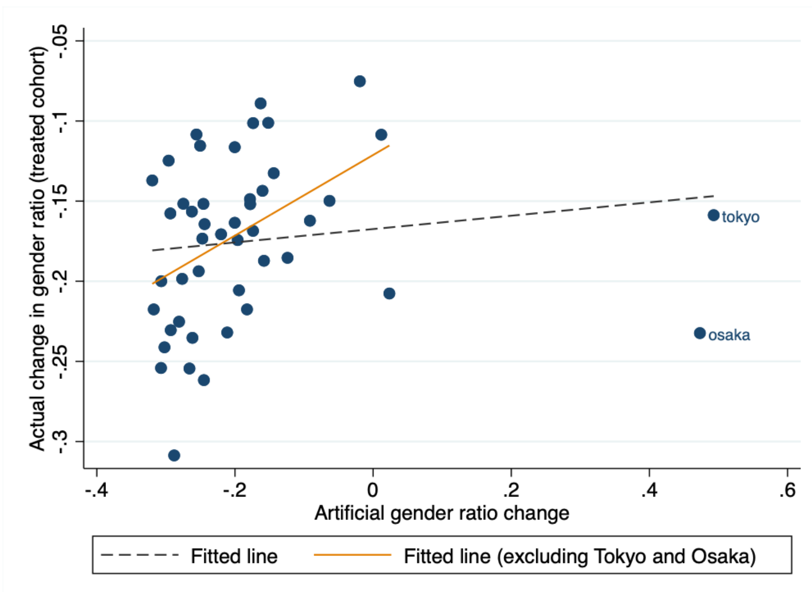
Table 9 replicates Table 8 using the artificial gender ratio, but we add a dummy variable that takes 1 for Tokyo and Osaka. The signs of the coefficients are consistent with the results in Table 8. However, the coefficients are smaller for the industry employment shares. This confirms that the previous estimates using prefecture-level gender ratio change may indeed suffer an omitted variable bias, that is related to the movement of male population between 1935 and 1947. The size of the coefficients here is closer to one from the coefficients in Table 3, that are robust to unobservable prefecture characteristics. In the appendix, we provide the result using the artificial gender ratio change that also ignores female migration (Table B7) and controlling for the population changes due to migration for men and women separately (Table B8).

Death rate. — It is also possible to use death rate by prefecture directly in the estimation. This resembles the approach taken by Boehnke and Gay (2022) and Cardoso and Morin (2018).

We impute death rate for each prefecture by dividing death count by number of men in the treated cohorts when they were aged 10-14. We do not use number of men in the treated cohorts in 1935 in the denominator. This is because our death count is based on the prefecture of registration and population size at the

[48] Due to the contract with the Ministry, we cannot publish the table by ourselves.

Figure 7. : Correlation between actual and artificial gender ratio change



Note: The figure shows the correlation between actual gender ratio change of the treated cohort and artificial gender ratio change.

Table 9—: Regression with artificial gender ratio change

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Artificial gender ratio change	0.121*	-0.016	-0.080*	0.201***	0.302***	0.053***
	(0.061)	(0.059)	(0.041)	(0.046)	(0.075)	(0.016)
Bomb casualty per capita	-0.129	0.075	-0.067	0.555***	1.051***	0.064
	(0.318)	(0.300)	(0.182)	(0.193)	(0.283)	(0.118)
Destroyed building per capita	0.339***	-0.298*	-0.121	0.034	-0.002	0.036
	(0.117)	(0.155)	(0.102)	(0.089)	(0.145)	(0.040)
Height weight ratio	-4.657***	4.208***	0.258	-0.158	0.181	-0.033
	(0.805)	(0.728)	(0.419)	(0.389)	(0.642)	(0.219)
ln(pop) change 1935-47	0.019	-0.150**	0.106*	-0.033	-0.103	0.009
	(0.060)	(0.074)	(0.061)	(0.060)	(0.099)	(0.023)
Gender ratio in 1935	0.023	-0.188***	0.090*	-0.187***	-0.244***	-0.017
	(0.058)	(0.066)	(0.048)	(0.046)	(0.078)	(0.016)
ln(pop) in 1935	0.028	-0.167***	0.100***	-0.050	-0.125*	0.005
	(0.030)	(0.045)	(0.030)	(0.041)	(0.066)	(0.014)
ln(worker)	-0.036	0.163***	-0.086***			
	(0.030)	(0.045)	(0.030)			
ln(pop)				0.030	0.089	-0.007
				(0.040)	(0.065)	(0.013)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.854	0.940	0.919	0.794	0.776	0.625
Observations	2484	2484	2484	2484	2484	2484

Note: The table provides the regression results for artificial gender ratio change, controlling for 1930 outcomes. In addition to the variables listed in the table, the prefecture-level controls include age share of population (10s, 20s, 30s, 40s, 50s and 60s), dummy for large prefecture, dummy for Tokyo and Osaka. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

age of 10-14 is closer to the registered population of the treated cohorts, rather than the current population size in 1935 when they were aged 20-34. In fact, in our data, we observe that population size (and gender ratio) starts to diverge across prefectures from the age group of 15-19. Death rate is therefore imprecise and endogenous to migration if we simply use population size in 1935 in the denominator.[49] For example, death rate of large prefectures becomes lower (as receiving prefectures of young men) and that of small prefectures becomes higher (as sending prefectures). Our imputed death rate is on average 0.2, with the highest of 0.26 for Saga and the lowest of 0.12 for Hokkaido.

Table 10 shows the results. Note that the sign of the coefficient must be reversed from the previous tables as male loss is larger for higher death rate. The estimated coefficient for the industry employment shares have expected signs, that is higher death rate of young soldiers is associated with lower (higher) share of manufacturing (agricultural and service) sector share in employment. However, again, higher death rate is negatively related to female labor force participation and other employment outcomes, similarly to the previous results that exploit cross-prefecture variations in the gender ratio.[50]

E. Cross-cohort relation

We extend the analysis based on cross-prefecture variation in the gender ratio change to study the heterogeneous effect on different cohorts. The estimation sample is restricted between 1955 and 1960 to examine the immediate effect of the gender imbalance and highlights the mechanism of between-cohort substitution or complementarity.

Figure 8 provides the cohort-specific treatment effect. For the manufacturing sector employment share, the negative effects are detected for the treated cohort, and even more strongly for the younger cohort, while the coefficient is close to zero for older cohorts. This pattern is similar in the agricultural sector. We do not see a clear pattern of heterogeneity with respect to the employment and participation outcomes.

An interesting pattern emerges when we analyze the younger cohort and study the entire 1955-1980 period, shown in Figure 9. It demonstrates that the effect diminishes for younger cohorts, indicating catch-up process as the young and large cohorts enter the labor market. In the case of agriculture, the sign even reverses (although not precisely estimated).

[49] The reference age for the male population in the denominator can be younger than 10-14. The result is generally unchanged when we use age 5-9. However, we do not use age 0-4 as a reference because it will be contaminated by infant mortality rate, that is likely to be correlated with prefecture-level economic indicators.

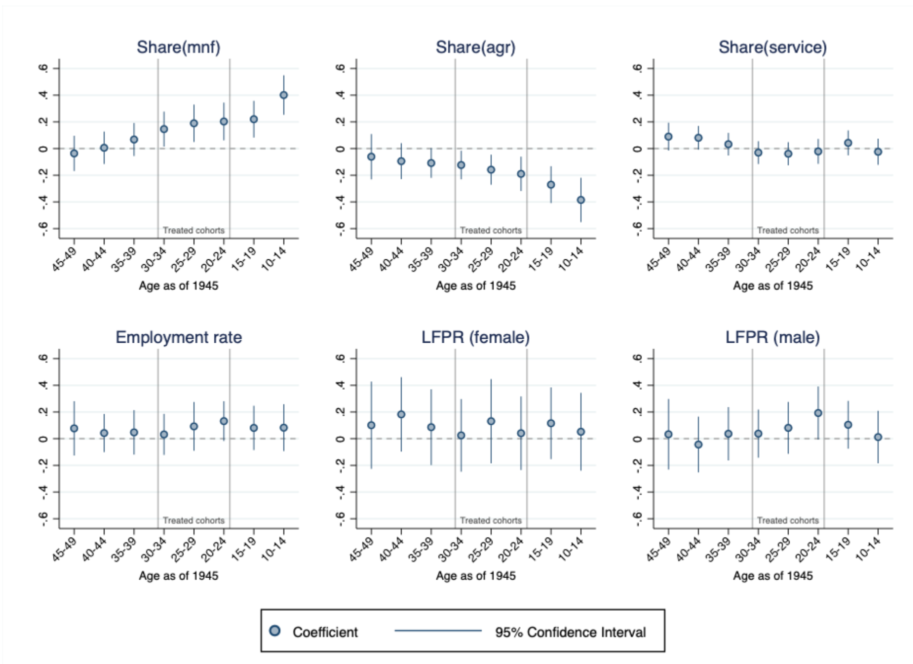
[50] We provide the regression results using the version of death rate based on number of men in the treated cohorts in 1935 in the appendix (Table B9).

Table 10—: Regression using death rate

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate	-0.297***	0.233**	0.028	-0.356***	-0.391***	-0.111***
	(0.109)	(0.110)	(0.064)	(0.073)	(0.116)	(0.035)
Bomb casualty per capita	0.240	-0.186	-0.048	0.727***	1.089***	0.133
	(0.410)	(0.389)	(0.193)	(0.215)	(0.322)	(0.126)
Destroyed building per capita	0.323**	-0.302*	-0.121	-0.032	-0.066	0.016
	(0.128)	(0.165)	(0.103)	(0.085)	(0.145)	(0.042)
Height weight ratio	-4.142***	4.502***	-0.193	-0.474	-0.073	-0.215
	(0.682)	(0.681)	(0.359)	(0.422)	(0.649)	(0.243)
ln(pop) change 1935-47	-0.021	-0.165***	0.134***	-0.180***	-0.284***	-0.040
	(0.054)	(0.063)	(0.050)	(0.051)	(0.079)	(0.025)
Gender ratio in 1935	0.055	-0.201***	0.134***	-0.208***	-0.290***	-0.011
	(0.070)	(0.068)	(0.039)	(0.046)	(0.074)	(0.023)
ln(pop) in 1935	-0.057***	-0.019	0.069***	-0.011	0.017	-0.011*
	(0.020)	(0.031)	(0.020)	(0.016)	(0.027)	(0.006)
ln(worker)	0.048***	0.017	-0.058***			
	(0.018)	(0.028)	(0.019)			
ln(pop)				-0.014	-0.059**	0.007
				(0.015)	(0.025)	(0.005)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.854	0.940	0.919	0.794	0.776	0.625
Observations	2484	2484	2484	2484	2484	2484

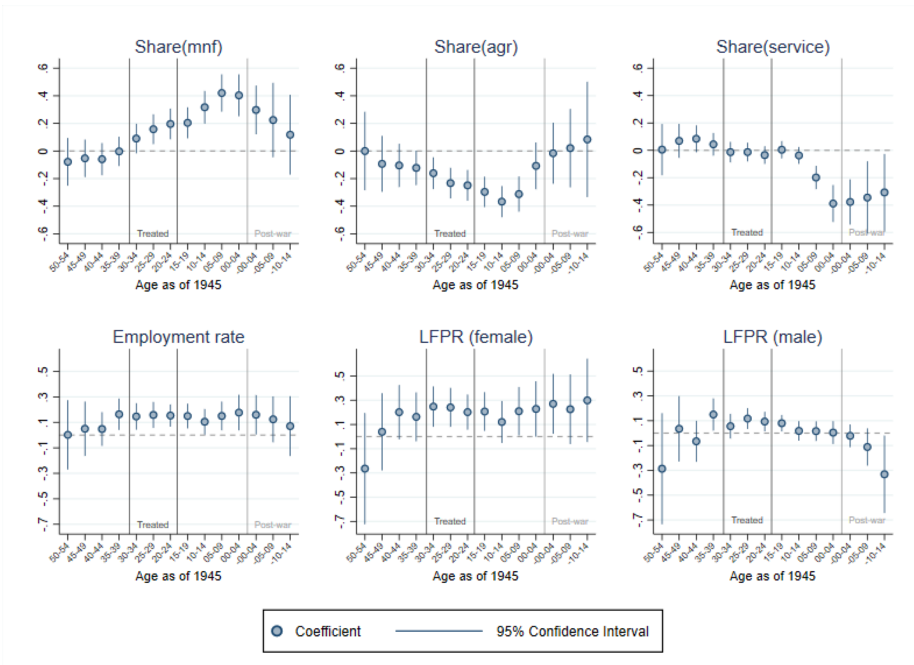
Note: The table provides the regression results for death rate where death rate is calculated based on number of death counts in each prefecture, divided by number of male population of treated cohorts when they were aged 10-14. The regression controls for 1930 outcomes. In addition to the variables listed in the table, the prefecture-level controls include share of young population (aged 20-30) in 1935, dummy for large prefecture, dummy for Tokyo and Osaka. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure 8. : Coefficient by cohorts, until 1960



Note: The figure provides the heterogeneous effect of the gender ratio change on the outcomes across cohorts. The regression is based on Equation 5, but the heterogeneous coefficients are obtained by interacting the cohort dummies with the change in the gender ratio of treated cohorts that varies only across prefectures.

Figure 9. : Coefficient by cohorts, all years 1955-1980



Note: The figure extends Figure 8 by adding cohorts that were not born in 1945 in the sample.

VIII. Conclusion

In this study, we examined how the permanent loss of human capital is related to economic recovery that led to the integration of women into the labor market owing to the loss of men during the war. War affects the industrial structure not only through damage to physical capital stocks but also through the tremendous loss of human capital. In the Second World War, Japan lost 2 million soldiers due to the unexpected mismatch of military strategies between offense and defense, and most casualties were males from certain areas and cohorts in the country owing the institution of hometown regiment.

Using the difference in gender imbalances between geographical areas and cohorts because of the war, we examined the effects of permanent loss of males on the consecutive economic development of areas. The regression results show that the permanent loss of males may have led to slower industrialization and a tentative increase in agriculture. However, such slow-down effects were limited quantitatively and had gradually disappeared after approximately 15 years from the end of the war.

These findings imply that average technology is augmented by gender during high-speed economic growth after the war. In the long term, technological change and internal migration may have neutralized the permanent loss of human capital in certain geographical areas. Further, we find that all sectors increased female share in employment in response to the imbalanced gender ratio, indicating that the within-sector production technology adjustment was one mechanism for the local economy to absorb the negative shock in human capital composition. This effect was persistent even after the impact on the industrial structure has disappeared. Consistently, the permanent loss of men moderately increased female labor force participation rate, that persisted over 30 years since the end of the war, highlighting the demand-side pulling factors that led to the integration of women into the labor market.

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Online Appendix A: Data

Our main data source is census published tables. Prefecture-age-gender data for industry share in employment is only available since 1940. Moreover, a direct comparison of employment is difficult for census waves before and after the war because the pre-war census used different definitions of employment (in other words, Hongyo in 1920 and Yugyo in 1930 and 1940), and coverage also was not the same (for example, not covering self-employees). The definition of employment in 1930 and 1940 is close to labor force participation in the post-war census. For this reason, we use the employment rate in the pre-war periods as a proxy for labor force participation rate. Note that industry classifications are slightly different in the pre-war and post-war census. When we use prefecture-level outcomes, it is of those aged 20-64 when cohort-level information is available (post-1940); otherwise, the values refer to those aged 14 or older, directly taken from the census. Regarding the sector employment share at prefecture level in 1930, we have digitized only for 1-digit manufacturing and agriculture for the moment, and service sector share is extrapolated by one minus the employment shares of these two sectors combined. It will be updated after digitizing all other 1-digit sectors to make the sector grouping comparable to the one used in the main regression.

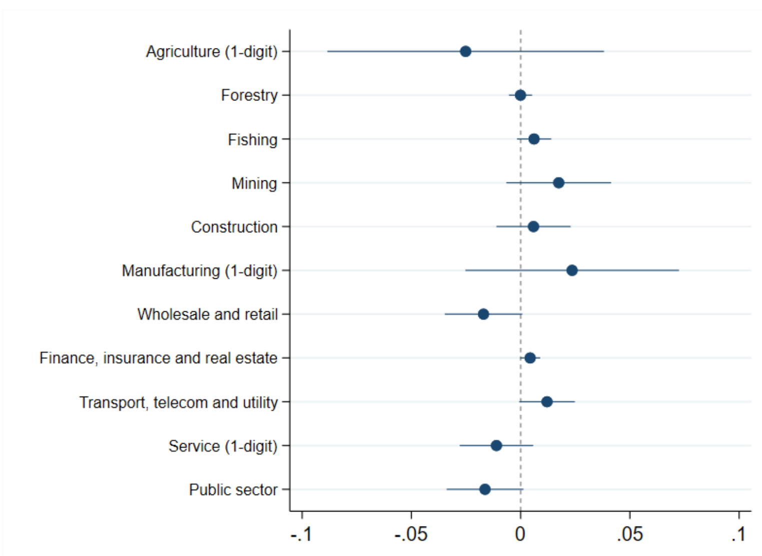
Online Appendix B: Supplementary Regression Results

B1. Results based on 1-digit industry classifications

The Figures B1-B4 provide the main results on the industry employment share and within-sector female share at 1-digit level (without aggregating them into three large sector groups as in the main text). Figure B1 shows that male loss has led to the higher employment share of Agriculture and certain service sectors, namely wholesale and retail, service, and Public Sector. It decreased not only the employment share of all manufacture sector but also some relatively high-paying service industries such as finance, insurance, and real estate, and transport, telecommunication, and utility. The effects are in general short-lived, as confirmed in Figure B2.

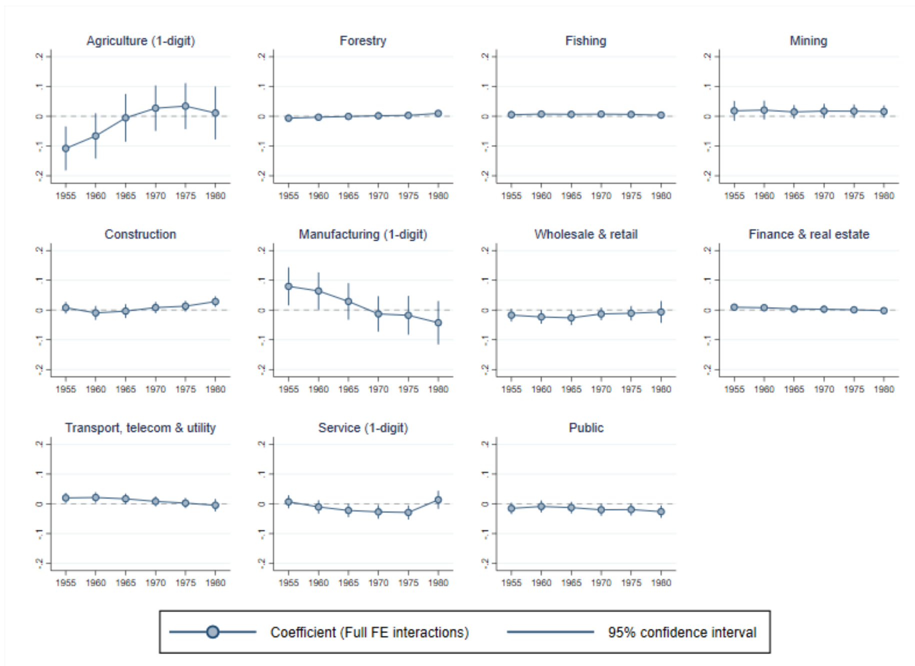
Regarding within-sector female share, Figure B3 shows that female representation has increased in many 1-digit industries in response to the loss of men, although there are some variations. Figure B4 shows the time-varying effects, in general confirming a persistent change in gender composition of workers in sectors where we find an effect.

Figure B1. : Effect on industry share - 1-digit industry



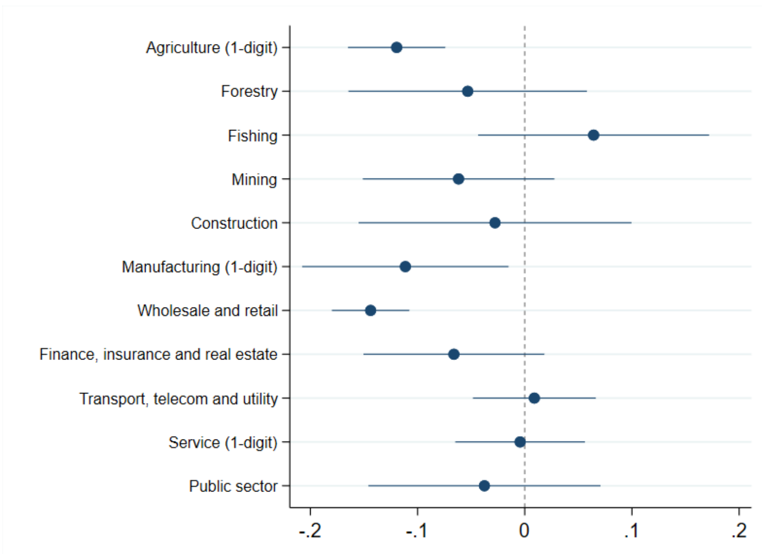
Note: The figure shows the impact of the change in the gender ratio estimated from the specification of Equation 2, where the outcomes are industrial employment shares based on 1-digit sector dis-aggregation.

Figure B2. : Effect on industry share (long run) - 1-digit industry



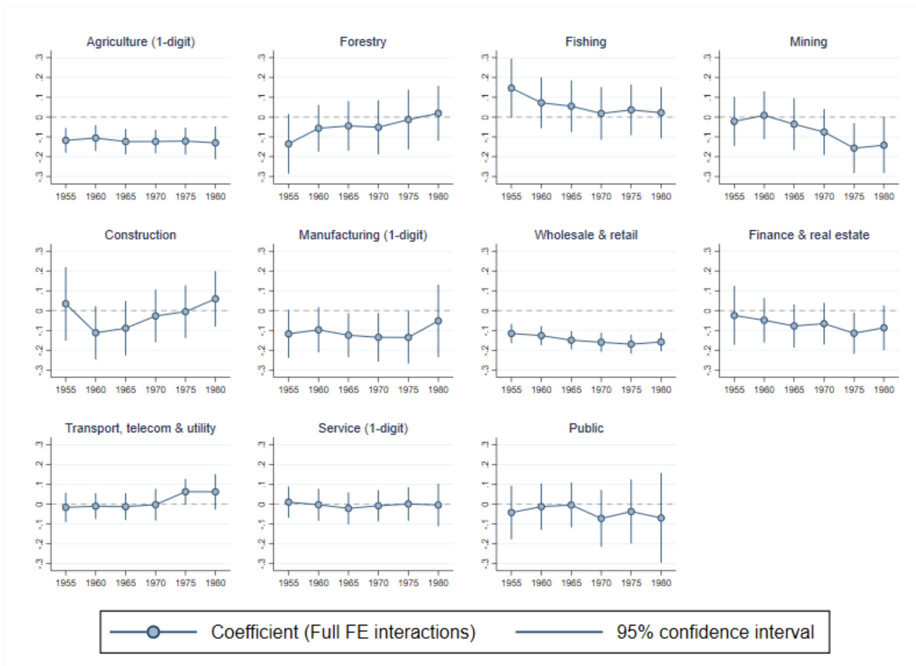
Note: The figure repeats Figure 5 for the industrial employment shares based on 1-digit classification.

Figure B3. : Effect on within sector female share - 1-digit industry



Note: The figure shows the impact of the change in the gender ratio estimated from the specification of Equation 2 where the outcomes are female share wihtin each 1-digit sector.

Figure B4. : Effect on within sector female share (long run) - 1-digit industry



Note: The figure repeats Figure 6 for the industrial employment shares based on 1-digit classification.

B2. Role of FE interactions

Our main results are based on the specification (in Equation 2) which includes interaction fixed effects. Table B1 shows the comparison of the estimated coefficients with the alternative specifications. The first row shows that, without adding any interaction fixed effects (in other words, specification with only one-dimensional fixed effects in year, cohort and prefecture), the estimated effects on industrial share are overestimated, as indicated in columns (1)-(3). This remains similar when the prefecture-year FE is added in the second row. The third row shows that the effect attenuates significantly when we add cohort-year fixed effects, indicating the importance of controlling for the overall differences in employment share across cohorts (or ages). The estimated coefficients are much closer to the main specification that fully interacts fixed effects as shown in the last row. Column (5) shows that controlling for cohort-year specific effect is important in identifying the effect on female labor force participation because the female labor supply has a strong life cycle component. Column (5) shows that controlling for cohort-year specific effect is important in identifying the effect on female labor force participation. This is due to the fact that the female labor supply has a strong life cycle component.

Table B1—: Comparison of coefficients across models

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Coefficient from the model with:						
No FE interactions (Baseline)	0.171***	-0.249***	0.078***	0.135*	0.060	0.147**
	(0.032)	(0.044)	(0.026)	(0.075)	(0.103)	(0.067)
Prefecture-year FE interaction	0.177***	-0.267***	0.091***	0.236***	0.179*	0.186***
	(0.031)	(0.043)	(0.026)	(0.077)	(0.105)	(0.071)
Cohort-year FE interaction	0.075**	-0.051	-0.024	-0.056	-0.161***	0.019
	(0.030)	(0.035)	(0.022)	(0.037)	(0.061)	(0.052)
Both FE interactions	0.052*	-0.019	-0.032	-0.057	-0.159***	-0.013
	(0.027)	(0.033)	(0.021)	(0.037)	(0.061)	(0.053)

Note: The table summarizes the estimated coefficient of the change in the gender ratio with alternative specifications. All regressions are based on Equation 2, but the first row removes the interaction fixed effects. The second and the third rows include interaction fixed effects only along prefecture-year and cohort-year, respectively. The last row is our main results from Table 3. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B3. Time-varying effect regression results

This section presents the complete regression results for time-varying effect. Table B2 summarizes the results corresponding to the main text that includes a full set of fixed effect interactions; we also supplement Table B3 and Table B4 that include interaction fixed effects only along cohort-year and prefecture-year, respectively. We also visually compare the time-varying effects across the regression specifications: Figure B5 compares the results between specifications without interacted fixed effects and with both fully interacted effects; Figure B6 compares the result of the specification with full fixed effect interactions to the results that add only one of the interaction fixed effects. The figure confirms the important role of cohort-year fixed effects in controlling the life cycle effects.

Table B2—: Time-varying effects with two FE interactions (3)

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.091***	-0.094**	0.003	-0.033	-0.106	0.049
	(0.033)	(0.038)	(0.026)	(0.043)	(0.082)	(0.071)
× 1960	-0.004	0.027	-0.022	-0.022	-0.044	-0.064
	(0.028)	(0.030)	(0.027)	(0.047)	(0.088)	(0.061)
× 1965	-0.041	0.090***	-0.049*	-0.069	-0.103	-0.125*
	(0.027)	(0.033)	(0.028)	(0.048)	(0.082)	(0.064)
× 1970	-0.068***	0.125***	-0.057**	-0.042	-0.073	-0.113**
	(0.026)	(0.032)	(0.025)	(0.035)	(0.069)	(0.051)
× 1975	-0.068**	0.132***	-0.064**	-0.011	-0.086	-0.023
	(0.029)	(0.036)	(0.026)	(0.036)	(0.073)	(0.061)
× 1980	-0.079**	0.113**	-0.034	0.015	-0.015	-0.040
	(0.034)	(0.045)	(0.033)	(0.049)	(0.082)	(0.061)
ln(pop) change 1935-47	-0.028*	-0.003	0.031**	0.082***	0.147***	0.071**
	(0.015)	(0.019)	(0.013)	(0.024)	(0.043)	(0.028)
Gender ratio in 1935	0.093***	-0.086**	-0.007	-0.043	-0.144**	-0.044
	(0.030)	(0.034)	(0.023)	(0.040)	(0.065)	(0.052)
ln(pop) in 1935	-0.030**	-0.030	0.060***	0.100***	0.183***	0.076***
	(0.015)	(0.020)	(0.012)	(0.022)	(0.043)	(0.024)
ln(worker)	-0.038***	0.136***	-0.097***			
	(0.011)	(0.015)	(0.008)			
ln(pop)				0.037**	-0.026	0.017
				(0.016)	(0.032)	(0.019)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.968	0.977	0.965	0.913	0.900	0.634
Observations	1748	1748	1748	1748	1748	1748

Note: The table shows the results of the time-varying effects in the specification that includes both prefecture-year and cohort-year interacted fixed effects. The first row corresponds to the effect of the gender ratio in 1955, and the next five rows are the estimated interaction coefficient for subsequent years (representing the difference from 1955 effect). All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B3—: Time-varying effects with cohort-year FE interaction

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.109**	-0.103**	-0.006	-0.043	-0.150*	0.078
	(0.044)	(0.052)	(0.030)	(0.045)	(0.086)	(0.061)
× 1960	-0.000	0.021	-0.020	-0.056	-0.068	-0.084*
	(0.046)	(0.056)	(0.035)	(0.049)	(0.092)	(0.046)
× 1965	-0.019	0.050	-0.031	-0.081*	-0.113	-0.124**
	(0.038)	(0.048)	(0.032)	(0.043)	(0.076)	(0.048)
× 1970	-0.010	0.063	-0.053*	-0.029	-0.041	-0.095**
	(0.038)	(0.048)	(0.030)	(0.038)	(0.069)	(0.040)
× 1975	-0.045	0.082	-0.036	0.033	0.029	-0.004
	(0.047)	(0.059)	(0.035)	(0.041)	(0.081)	(0.051)
× 1980	-0.161***	0.120	0.041	0.101**	0.199*	-0.023
	(0.059)	(0.074)	(0.043)	(0.051)	(0.101)	(0.046)
ln(pop) change 1935-47	-0.057***	0.025	0.032**	0.070***	0.131***	0.053*
	(0.019)	(0.024)	(0.014)	(0.024)	(0.042)	(0.027)
Gender ratio in 1935	0.133***	-0.135***	0.002	-0.062	-0.185***	-0.009
	(0.034)	(0.036)	(0.025)	(0.040)	(0.064)	(0.051)
ln(pop) in 1935	-0.033*	-0.028	0.060***	0.074***	0.139***	0.058**
	(0.019)	(0.025)	(0.013)	(0.021)	(0.041)	(0.023)
ln(worker)	-0.007	0.087***	-0.080***			
	(0.013)	(0.020)	(0.011)			
ln(pop)				0.037**	-0.017	0.039***
				(0.015)	(0.028)	(0.013)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	No	No	No	No	No	No
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.936	0.960	0.940	0.883	0.859	0.608
Observations	1748	1748	1748	1748	1748	1748

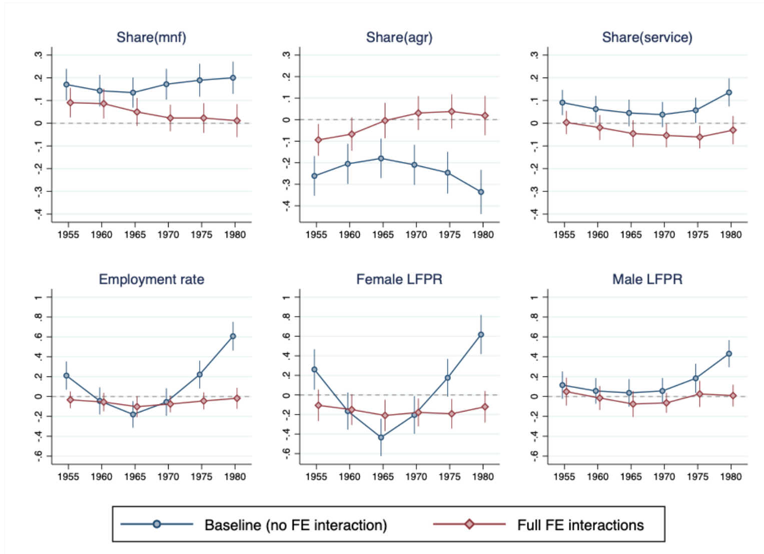
Note: The table shows the results of the time-varying effects in the specification that includes only cohort-year interacted fixed effects. The first row corresponds to the effect of the gender ratio in 1955, and the next five rows are the estimated interaction coefficient for subsequent years (representing the difference from 1955 effect). All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B4—: Time-varying effects with pref-year FE interaction

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.162***	-0.262***	0.100***	0.278***	0.359***	0.124
	(0.031)	(0.043)	(0.027)	(0.072)	(0.103)	(0.076)
×1960	-0.029	0.052***	-0.023	-0.244***	-0.431***	-0.039
	(0.018)	(0.018)	(0.015)	(0.033)	(0.051)	(0.044)
×1965	-0.036**	0.072***	-0.036**	-0.385***	-0.715***	-0.054
	(0.017)	(0.019)	(0.016)	(0.030)	(0.052)	(0.048)
×1970	-0.004	0.039*	-0.035**	-0.261***	-0.483***	-0.039
	(0.018)	(0.023)	(0.016)	(0.028)	(0.044)	(0.040)
×1975	0.032	-0.003	-0.029*	0.018	-0.106**	0.086*
	(0.020)	(0.028)	(0.016)	(0.037)	(0.052)	(0.049)
×1980	0.080***	-0.113***	0.033	0.426***	0.344***	0.362***
	(0.022)	(0.037)	(0.023)	(0.040)	(0.057)	(0.048)
ln(pop) change 1935-47	-0.086***	0.112***	-0.027	-0.019	0.035	-0.002
	(0.020)	(0.032)	(0.018)	(0.041)	(0.063)	(0.043)
Gender ratio in 1935	0.218***	-0.343***	0.125***	0.210***	0.134	0.145**
	(0.034)	(0.043)	(0.026)	(0.074)	(0.099)	(0.070)
ln(pop) in 1935	-0.073***	0.055*	0.018	0.034	0.108**	0.030
	(0.020)	(0.032)	(0.017)	(0.037)	(0.051)	(0.042)
ln(worker)	0.067***	-0.076***	0.010			
	(0.007)	(0.012)	(0.007)			
ln(pop)				0.288***	0.249***	0.203***
				(0.024)	(0.034)	(0.026)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cohort-year FE	No	No	No	No	No	No
R-squared	0.957	0.967	0.950	0.725	0.779	0.391
Observations	1748	1748	1748	1748	1748	1748

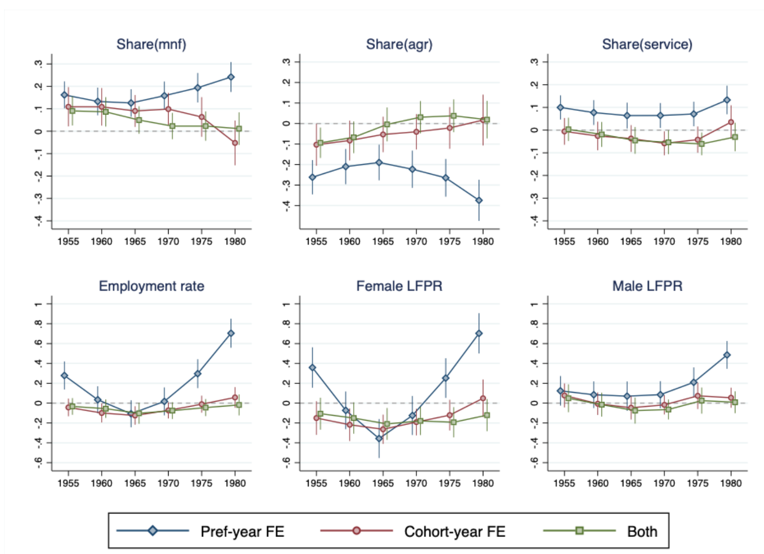
Note: The table shows the results of the time-varying effects in the specification that includes only prefecture-year interacted fixed effects. The first row corresponds to the effect of the gender ratio in 1955, and the next five rows are the estimated interaction coefficient for subsequent years (representing the difference from 1955 effect). All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure B5. : No interaction vs full interactions



Note: The figure repeats Figure 5. It adds the results from the specification that includes no interaction fixed effects.

Figure B6. : No interaction vs full interactions



Note: The figure repeats Figure 5, comparing the main results with two less strict models that add only one of the interacted fixed effects.

B4. Contemporaneous gender ratio

The skewed gender due to the war can be alleviated by the migration in the post-war period. In theory, such effect can be captured by including the contemporaneous gender ratio partially for the sake of comparison with the previous literature The result is presented in Table B5. Note that the regression suffers a high correlation between the change in the gender ratio during 1935-1947 (our treatment variable) and the contemporaneous gender ratio. This makes the separate identification difficult and partially explains why the estimated effect of the change in the gender ratio is close to zero for manufacturing sector, and the sign of the effect has changed for agriculture, although the result is consistent for the other outcomes.

Table B5—: Controlling for contemporaneous gender ratio

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47	0.004	0.044	-0.047**	-0.091**	-0.200***	0.050
	(0.030)	(0.037)	(0.022)	(0.040)	(0.072)	(0.073)
ln(pop) change 1935-47	-0.003	-0.035*	0.039***	0.100***	0.166***	0.040
	(0.016)	(0.021)	(0.013)	(0.023)	(0.040)	(0.032)
Contemporaneous gender ratio	0.102***	-0.134***	0.032*	0.066**	0.079	-0.121
	(0.025)	(0.031)	(0.017)	(0.032)	(0.071)	(0.082)
Gender ratio in 1935	0.028	-0.002	-0.026	-0.086*	-0.195**	0.038
	(0.035)	(0.041)	(0.025)	(0.046)	(0.082)	(0.081)
ln(pop) in 1935	-0.003	-0.065***	0.069***	0.119***	0.205***	0.045
	(0.016)	(0.022)	(0.013)	(0.021)	(0.040)	(0.029)
ln(worker)	-0.055***	0.157***	-0.102***			
	(0.011)	(0.015)	(0.008)			
ln(pop)				0.021	-0.044	0.048*
				(0.017)	(0.030)	(0.027)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.968	0.977	0.964	0.913	0.900	0.634
Observations	1748	1748	1748	1748	1748	1748

Note: The table shows the results of the regression that adds contemporaneous gender ratio in the main specification. All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B5. Alternative definition of gender ratio

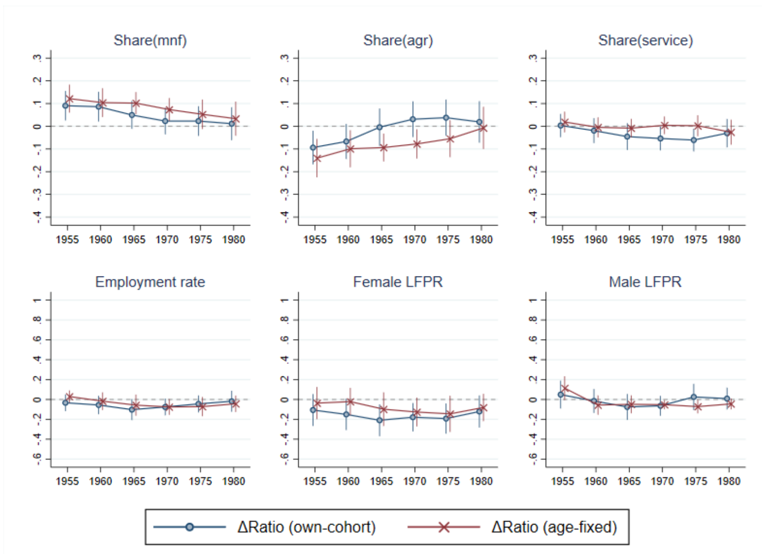
In Section VII.A, we implemented the robustness check using an alternative definition of the gender ratio change that fixed age rather than cohort. Table B6 shows the full results for this age-fixed alternative gender ratio change under the main specification of Equation 2. Figure B7 shows the time-varying when the alternative measure of the gender ratio is used, compared with the main result using the cohort-fixed measure.

Table B6—: Regression results using alternative definition of the gender ratio change

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Gender ratio change 1935-47 (age-controlled)	0.091***	-0.065***	-0.016***	-0.033	-0.079*	-0.020
	(0.015)	(0.020)	(0.005)	(0.023)	(0.042)	(0.024)
ln(pop) change 1935-47 (age-controlled)	-0.044***	-0.061***	0.031***	0.016	0.007	-0.008
	(0.010)	(0.021)	(0.005)	(0.013)	(0.020)	(0.021)
Gender ratio in 1935	-0.005	-0.040**	0.003	0.019	0.035	-0.024
	(0.014)	(0.016)	(0.004)	(0.021)	(0.037)	(0.022)
ln(pop) in 1935	0.019**	-0.021	0.021***	0.036***	0.072***	0.031**
	(0.009)	(0.016)	(0.006)	(0.013)	(0.023)	(0.016)
ln(worker)	-0.046***	0.076***	-0.060***			
	(0.011)	(0.018)	(0.004)			
ln(pop)				0.052***	0.010	0.025
				(0.015)	(0.030)	(0.018)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Age-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.970	0.978	0.964	0.912	0.899	0.631
Observations	1748	1748	1748	1748	1748	1748

Note: The table shows the full results of the specification where we use the age-fixed version of the gender ratio change. All standard errors are clustered at prefecture-year level. Standard errors are in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure B7. : Time varying effect: own-cohort vs age-fixed $\Delta Ratio$



Note: The figure compares the estimated time-varying effects of the gender ratio change when the gender ratio change calculated is cohort-fixed or age-fixed.

B6. Correlation across cohorts

Figure B8 provides the correlation matrix across cohorts in terms of the change the gender ratio.

Figure B8. : Correlation of $\Delta Ratio$ across cohorts

Matrix of correlations

Variables	189600	190105	190610	191115	191620	192125	192630	193135
189600	1.000							
190105	0.876	1.000						
190610	0.734	0.796	1.000					
191115	0.414	0.385	0.627	1.000				
191620	-0.018	-0.087	0.211	0.521	1.000			
192125	-0.415	-0.279	-0.130	-0.131	0.341	1.000		
192630	-0.176	0.040	0.068	-0.228	-0.254	0.646	1.000	
193135	-0.065	0.093	-0.071	-0.223	-0.556	-0.039	0.425	1.000

Note: Correlation of change in gender ratio 1935-47 across cohorts

B7. Regressions related to the artificial gender ratio and death rate

The artificial gender ratio change defined in Discussion VII.D is subject to the endogenous change in female social movement across prefectures. We can also create another version of the artificial gender ratio change that fixes female population as 1935. By doing so, this gender ratio change in this measure reflects solely the change in male population by death.

$$Ratio_j^{Art2} = \frac{M_{j,1935} - Death_j}{F_{j,1935}} - \frac{M_{j,1935}}{F_{j,1935}}$$

Table B7 shows the results and find relatively similar results. The coefficients for sector employment shares are larger in magnitude compared to Table 9.

Since the artificial gender ratio ignores the (cross-prefecture) mobility, Table B8 controls for the percentage change in male and female population from 1935 to 1947 in the treated cohorts that are due to migration. For men, the change in population due to migration is imputed from the change in population less death count of soldiers. For women, it is simply the population change, since there are no military death for female.

Lastly, Table B9 supplements Table 10 by using the death rate calculated using the treated cohorts' male population size in 1935, instead of fixing ages.

Table B7—: Regression with alternative artificial gender ratio change

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Artificial gender ratio (female pop fixed)	0.217**	-0.090	-0.107*	0.366***	0.507***	0.093***
	(0.086)	(0.094)	(0.060)	(0.061)	(0.103)	(0.028)
Bomb casualty per capita	0.051	-0.038	-0.127	0.833***	1.407***	0.129
	(0.339)	(0.322)	(0.201)	(0.211)	(0.318)	(0.126)
Destroyed building per capita	0.333***	-0.301**	-0.116	0.019	-0.025	0.033
	(0.112)	(0.152)	(0.101)	(0.082)	(0.137)	(0.041)
Height weight ratio	-4.487***	4.324***	0.137	-0.029	0.456	-0.012
	(0.743)	(0.711)	(0.405)	(0.410)	(0.674)	(0.216)
ln(pop) change 1935-47	-0.062	-0.137**	0.160***	-0.172***	-0.307***	-0.028
	(0.045)	(0.057)	(0.051)	(0.053)	(0.087)	(0.021)
Gender ratio in 1935	-0.010	-0.185***	0.108**	-0.237***	-0.320***	-0.028*
	(0.056)	(0.065)	(0.045)	(0.042)	(0.071)	(0.017)
ln(pop) in 1935	0.034	-0.168***	0.092***	-0.039	-0.104*	0.008
	(0.030)	(0.045)	(0.029)	(0.037)	(0.061)	(0.013)
ln(worker)	-0.045	0.166***	-0.077**			
	(0.030)	(0.045)	(0.030)			
ln(pop)				0.015	0.062	-0.012
				(0.036)	(0.061)	(0.012)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.854	0.940	0.919	0.794	0.776	0.625
Observations	2484	2484	2484	2484	2484	2484

Note: The table provides the regression results for alternative artificial gender ratio change that fixes female population. The regression controls for 1930 outcomes. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B8—: Regression with alternative artificial gender ratio change, controlling for movement for men and women

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Artificial gender ratio (female pop fixed)	0.340***	-0.236***	-0.094*	0.332***	0.441***	0.080**
	(0.080)	(0.090)	(0.055)	(0.062)	(0.105)	(0.032)
Percentage change by migration (male)	0.274***	-0.239***	-0.064*	-0.007	-0.026	-0.005
	(0.050)	(0.058)	(0.038)	(0.043)	(0.071)	(0.021)
Percentage change by migration (female)	-0.179**	0.006	0.124**	-0.087	-0.155	-0.030
	(0.070)	(0.065)	(0.048)	(0.060)	(0.095)	(0.022)
Bomb casualty per capita	0.218	-0.307	-0.144	0.797***	1.328***	0.114
	(0.307)	(0.305)	(0.190)	(0.210)	(0.325)	(0.130)
Destroyed building per capita	0.212**	-0.234	-0.081	0.008	-0.040	0.030
	(0.106)	(0.148)	(0.103)	(0.082)	(0.138)	(0.041)
Height weight ratio	-4.812***	3.998***	0.176	-0.181	0.142	-0.071
	(0.806)	(0.719)	(0.396)	(0.419)	(0.690)	(0.224)
ln(pop) change 1935-47	-0.176**	0.068	0.107	-0.096	-0.162	0.001
	(0.080)	(0.095)	(0.074)	(0.074)	(0.118)	(0.035)
Gender ratio in 1935	0.102*	-0.247***	0.057	-0.213***	-0.279***	-0.021
	(0.061)	(0.069)	(0.053)	(0.051)	(0.085)	(0.018)
ln(pop) in 1935	0.031	-0.173***	0.105***	-0.050	-0.123*	0.004
	(0.028)	(0.044)	(0.030)	(0.039)	(0.064)	(0.013)
ln(worker)	-0.035	0.167***	-0.092***			
	(0.029)	(0.044)	(0.030)			
ln(pop)				0.026	0.083	-0.008
				(0.038)	(0.063)	(0.013)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.854	0.940	0.919	0.794	0.776	0.625
Observations	2484	2484	2484	2484	2484	2484

Note: The table provides the regression results for alternative artificial gender ratio change that fixes female population. The regression controls for 1930 outcomes. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B9—: Regression using alternative death rate

	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Death rate	-0.071	0.060	-0.039	-0.258***	-0.316***	-0.065**
	(0.080)	(0.086)	(0.050)	(0.058)	(0.097)	(0.028)
Bomb casualty per capita	-0.105	0.088	0.023	0.470**	0.839***	0.046
	(0.398)	(0.374)	(0.172)	(0.210)	(0.302)	(0.120)
Destroyed building per capita	0.365***	-0.337**	-0.130	0.029	-0.000	0.035
	(0.135)	(0.169)	(0.102)	(0.086)	(0.147)	(0.040)
Height weight ratio	-3.595***	4.129***	-0.393	-0.424	-0.134	-0.098
	(0.666)	(0.670)	(0.366)	(0.419)	(0.664)	(0.235)
ln(pop) change 1935-47	-0.006	-0.177***	0.125**	-0.174***	-0.281***	-0.035
	(0.057)	(0.065)	(0.050)	(0.050)	(0.079)	(0.024)
Gender ratio in 1935	0.026	-0.177**	0.133***	-0.254***	-0.341***	-0.028
	(0.074)	(0.073)	(0.039)	(0.046)	(0.071)	(0.023)
ln(pop) in 1935	-0.062***	-0.017	0.070***	-0.012	0.016	-0.011*
	(0.021)	(0.031)	(0.020)	(0.016)	(0.028)	(0.006)
ln(worker)	0.058***	0.009	-0.061***			
	(0.019)	(0.028)	(0.019)			
ln(pop)				-0.013	-0.059**	0.008
				(0.015)	(0.026)	(0.005)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war outcome (1930 value)	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.854	0.940	0.919	0.794	0.776	0.625
Observations	2484	2484	2484	2484	2484	2484

Note: The table provides the regression results for death rate where death rate is calculated based on number of death counts in each prefecture, divided by number of male population in treated cohorts in 1935 in each prefecture. The regression controls for 1930 outcomes. In addition to the variables listed in the table, the prefecture-level controls include share of young population (aged 20-30) in 1935, dummy for large prefecture, dummy for Tokyo and Osaka. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B8. Prefecture level variation controlling for other cohorts

We provide supplementary results to the regression that exploits prefecture-level variation (Table 8). Table B10 controls for the change in the gender ratio of the control cohorts — added separately for those aged 10-19 and 35-49 in 1945.

Table B10—: With pre-war outcome control and other cohorts' ΔRatio

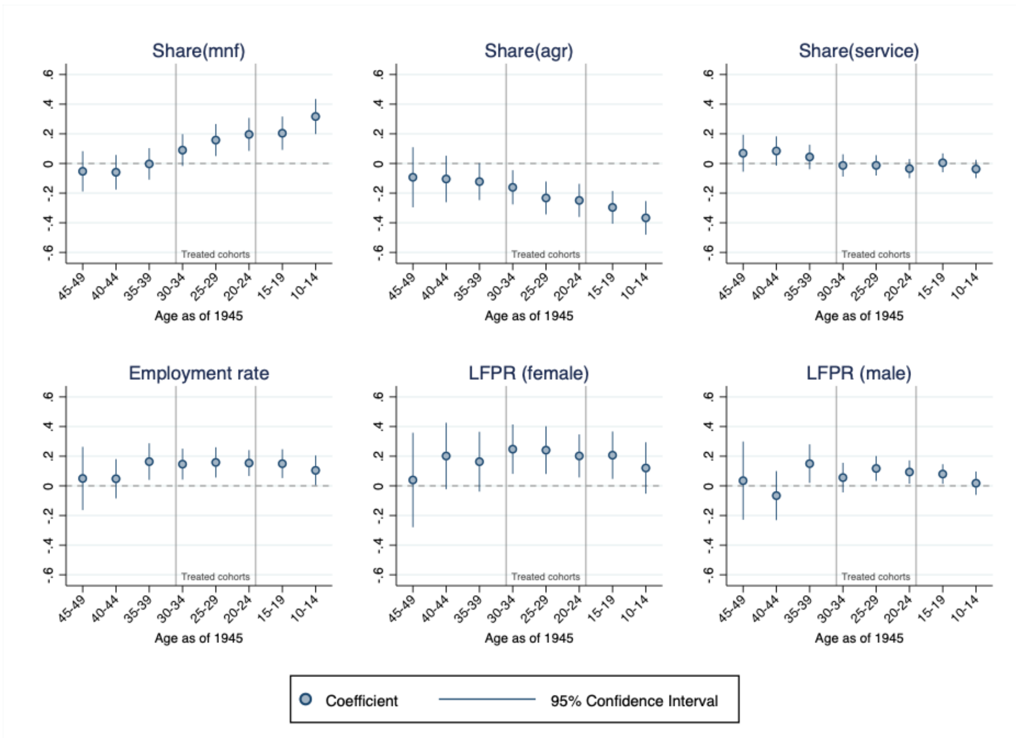
	Industry Employment Share			Employment and Participation		
	Manufacturing	Agriculture	Service	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Delta Ratio (age 20-34 in 1945)	0.176***	-0.203***	-0.040	0.109**	0.173**	0.014
	(0.051)	(0.056)	(0.035)	(0.045)	(0.072)	(0.018)
Delta Ratio (age 10-19 in 1945)	-0.278***	-0.071	0.259***	-0.208**	-0.194	-0.149***
	(0.105)	(0.123)	(0.087)	(0.094)	(0.147)	(0.043)
Delta Ratio (age 35-49 in 1945)	-0.208**	0.157	0.113	0.153*	0.218*	0.082**
	(0.087)	(0.119)	(0.078)	(0.089)	(0.128)	(0.040)
Bomb casualty per capita	0.123	-0.590**	0.501***	0.618***	1.114***	0.058
	(0.244)	(0.296)	(0.157)	(0.196)	(0.302)	(0.114)
Destroyed building per capita	0.386***	-0.162	-0.278***	-0.083	-0.186	0.006
	(0.104)	(0.140)	(0.087)	(0.100)	(0.161)	(0.043)
Height weight ratio	-2.958***	4.884***	-1.329***	0.479	1.054*	0.062
	(0.674)	(0.659)	(0.333)	(0.383)	(0.622)	(0.205)
ln(pop) change 1935-47	0.127**	-0.208***	0.109***	-0.132***	-0.276***	-0.014
	(0.051)	(0.058)	(0.029)	(0.043)	(0.072)	(0.016)
Gender ratio in 1935	0.047	-0.096	-0.017	-0.068	-0.083	0.049
	(0.087)	(0.099)	(0.064)	(0.077)	(0.119)	(0.031)
ln(pop) in 1935	0.020	-0.156***	0.092***	-0.013	-0.065	0.013
	(0.026)	(0.041)	(0.029)	(0.041)	(0.066)	(0.014)
ln(worker)	-0.011	0.153***	-0.098***			
	(0.026)	(0.040)	(0.028)			
ln(pop)				0.001	0.038	-0.013
				(0.040)	(0.066)	(0.013)
Cohort FE and time FE	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	No	No	No	No	No	No
Cohorot-time FE interaction	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
Pre-war prefecture-level controls	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.850	0.933	0.869	0.789	0.772	0.625
Observations	2484	2484	2484	2484	2484	2484

Note: The table extends the results of Table 8 by adding the change in the gender ratio of the control cohorts. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B9. Cohort substitutions

In the main text, we provided the results on the cohort-substitution effect in Figure 8 which used balance panel of the cohorts until 1960. We provide the results using a balance panel of the cohorts for 1955-1980 (with the attrition of the older cohorts later in the sample) in Figure B9.

Figure B9. : Coefficient for selected cohorts (unbalanced panel), all years



B10. Robustness to outliers

In order to make sure that our main results on the impact on the industrial share is not driven by some prefecture, we exclude prefectures with top and bottom 3 prefectures in terms of the 1955 manufacturing share. Table B11 show the results. The estimates coefficients have consistent signs, and the impact on agricultural share becomes actually significant.

Table B11—: Results excluding outlier prefectures

	Outcome					
	Share(mnf)	Share(agr)	Share(service)	Employment rate	Female LFPR	Male LFPR
	(1)	(2)	(3)	(4)	(5)	(6)
Change in gender ratio 1935-47	0.113***	-0.085**	-0.028	-0.032	-0.111*	0.020
	(0.035)	(0.040)	(0.025)	(0.038)	(0.064)	(0.064)
change_lnpop_1935_47	-0.081***	0.070**	0.011	0.107***	0.190***	0.047
	(0.025)	(0.033)	(0.018)	(0.031)	(0.053)	(0.035)
1935 gender ratio	0.171***	-0.156***	-0.014	-0.020	-0.099	-0.009
	(0.039)	(0.040)	(0.027)	(0.038)	(0.064)	(0.060)
lnpop_1935	-0.062***	0.047	0.015	0.128***	0.247***	0.040
	(0.023)	(0.031)	(0.016)	(0.027)	(0.049)	(0.027)
lnworker	0.019	0.034	-0.054***			
	(0.014)	(0.022)	(0.012)			
lnpop				0.012	-0.068**	0.039**
				(0.019)	(0.034)	(0.015)
3-way fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year FE	No	No	No	No	No	No
Cohort-year FE	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.919	0.969	0.961	0.885	0.857	0.591
Observations	1520	1520	1520	1520	1520	1520

Note: The table provides robustness checks of our main results in Table 3 that excludes top three and bottom three prefectures in terms of manufacturing employment share in 1955 to check that our results are not driven by specific prefectures. Excluded prefectures are Osaka, Tokyo, Aichi, Kagoshima, Aomori, and Ibaraki. Standard errors are clustered at prefecture-year level. Standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.