The Consequences of Hometown Regiment what happened in hometown when soldiers had gone?

By KENTARO ASAI AND RYO KAMBAYASHI *

War sometimes caused a large gender imbalance in certain cohorts and in certain areas. Although this gender imbalance may have changed the trajectory of economic development, the controversy is still inconclusive, because the market economy has a strong restoring force. We intend to contribute to this literature by introducing the Japanese experience during the Second World War. Japan lost more than two million soldiers eternally between 1938 and 1945. Furthermore, since the Imperial Japanese Army organized its main force as hometown regiment, the loss of young males is concentrated in certain cohorts of certain geographical areas. By exploiting the variation of changes in gender balance cohort-by-prefecture, we examined the loss of young males may affect the postwar industrial structure. What we have found so far is that the reduction in the gender ratio may have led to slower industrialization, although it is only to a limited degree in terms of quantity.

I. War damage and economic recovery

In this article, we empirically examine how permanent loss of human capital relates to economic recovery. In particular, war affects the industrial structure not only through damage to physical capital stocks, but also through the tremendous loss of human capital. In our case, in addition to the fact that the defended country lost two million soldiers due to the unexpected mismatch of military strategies between offense and defense, the casualties on the battlefield were concentrated in males who were from certain areas and cohorts in the country due to the institutions of *hometown regiments*. Taking advantage of the difference in gender imbalances between geographical areas and cohorts as a consequence of war, we examine the effects of permanent reduction 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 only limited to a small degree in terms of quantity and had gradually disappeared after about fifteen years from the end of the war. These findings imply that average technology may have been augmented by gender during high-speed economic growth after the war. In the medium term, technological change and internal migration may have canceled the permanent loss of human capital in certain geographical areas.

* Asai: Paris School of Economics, 48 Boulevard Jourdan, Paris 75014, kentaro.asai@psemail.eu. Kambayashi: Institute of Economic Research Hitotsubashi University, 2-1 Naka Kunitachi Tokyo, kambayas@ier.hit-u.ac.jp. Acknowledgments

II. Brief literature review

To examine the stability of the market economy, the literature has taken the loss of resources during the war as an exogenous negative shock to the economy. Many researchers have tried to answer questions such as whether the economy returns to the expected trajectory from the prewar era or to what extent the speed of recovery depends on the magnitude of damage. For example, Craft (1995) proposed the interpretation 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 controversy is still inconclusive, since the OLS regression from postwar 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 postwar 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]

Other literature extended the study to use cross-sectional (regional) variation within a country to control for the long-term effect of unobservable regional factors. Naturally, the focus of the articles in this strand moved from aggregated economic growth to the efficiency of the reallocation process after the war. Davis and Weinstein (2002) studied the resilience of the distribution of the size of the city using the destruction of the physical capital by air attacks in Japan as an instrument of changes in the density of the population of the region during the war. Their conclusion is that the size of the city had been recovered 15 years after the war, which implies the strong resilience of the market economy. Brakman, Garrentsen and Schramm (2004) is in the same line, while they used 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 accumulated additional evidence about the resilience of the market economy. On the other hand, other analysis such as Bianchi and Giorcelli (2019) implies that not only the market mechanism, but also the policy intervention may have been needed for recovery. They studied bombing destruction in Italy and found the direct effect of the Marshall Plan on the reconstruction of the Italian economy.

Although physical capital could be reproduced in markets or by governments, damage to human capital may have permanently affected the economy. In the case of the United States, because most mobilized soldiers could return home, human capital in society may have been affected during the war only temporarily. Historians such as Chafe (1972) argued that the work experience of females, although it was temporal during the war, permanently changed their skills and attitudes towards the labor market, resulting in

[1] More concretely, 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 in section 3.2).

a change in the labor supply after the war. Goldin (1991) statistically examined this view, based on the retrospective work histories of more than 4,000 women collected by Gladys Palmer. She reported that the transition matrix on employment status between before and after the war suggests that wartime experiences did *not* increase women's employment. On the other hand, Acemoglu, Autor and Lyle (2004) found that wartime mobilization shifted female labor supply after the war. They define the mobilization rate as the fraction of registered males between the ages of 18 and 44 who were drafted or enlisted for the entire period of war; that is, a single value for each state. Rather than using the mobilization rate as an instrumental variable as in Davis and Weinstein (2002), they used it as an explanatory variable with the interaction of the postwar dummy. Because mobilization rates are only cross-sectionally varied, the OLS estimator is easily contaminated by state-specific unobservable factors, which requires remedying the issue of endogeneity. [2] Goldin and Olivetti (2013) also shows other evidence, using an additional wave of population censuses in the United States. Their conclusion is gray because they found that the shift in labor supply occurred only in the more educated group.[3] Studies in the United States indicate that heterogeneous responses had an already important role in the female labor supply after the Second World War due to a sufficiently industrialized economy. More recently, Boehnke and Gay (2020) extended the literature by introducing the French experience during the First World War, in which the industries had not matured enough to differentiate workers. Furthermore, because more than a million men lost their lives, wartime damage to human capital was persistent in the case. Taking advantage of the difference in casualties between regions (départements 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 that it is important to identify the causes and phases of social development of the destruction in the war. This finding, in turn, suggests the importance of the economic mechanism on the demand side behind it. 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.

[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] Actually, 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 longer-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 a large number of male convicts had been sent. However, as Qian (2008) and Carranza (2014) pointed out, it is generally 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. More specifically, gender imbalance can affect long-term supply conditions through marriage and other demographic changes, such as in Abramitzky, Delavande and Vasconcelos (2011), Grossbard (2015), Brainerd (2017), and Ogasawara and Komura (2018).

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 article contributes to the literature by providing additional evidence on the economic effects of wartime damage in a cleaner environment. First, because wartime damage is measured as the decline in the proportion of male relative to female (gender imbalance) as in Boehnke and Gay (2020) and Cardoso and Morin (2018), it was a *permanent* and *large* loss of gender-specific human capital for the economy that is different from the cases of the United States. Second, the gender imbalances are exogenously diversified not only in regions but also in cohorts in this article, because the variation depends on the unique context of history, as discussed in the next section. Wartime damages can be unrelated to the unobservable regional factors such as potential economic structure, once after controlling for region fixed effects.

III. Hometown Infantry Regiments and the War in Pacific Ocean

During the Second World War, the Imperial Japanese Army (IJA) took the *hometown regimental system* that organized each of the infantry regiments by collecting males from certain cohorts, typically aged 20-21, who lived in certain geographical areas. IJA dispatched regiments to defend the islands scattered in the Pacific Ocean, and the defensive force of the islands sometimes lost all of its war potential, facing the Allies' fierce counterattack. On the other hand, IJA could not expect that the Allies would pass through some islands without any combat. This fact indicates that, in Japan, certain geographical areas lost men massively in certain cohorts, but other areas and cohorts did not always lose them.

A. Construction of Imperial Japanese Army

First Difference: CONSCRIPTION SYSTEM. — The Japanese government had the mandatory draft system before 1945. Each male had to be examined for conscription, usually be-

[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. Clark, Doyle and Stancanelli (2020) also suggests that the impact of the Boston Marathon bombing negatively affects the welfare. In historical context, Costa and Kahn (2003) and Costa and Kahn (2007) argued the effect of the Civil War in the United States.

tween April and August of the year of age 20.[6] In 1944 and 1945, the threshold lowered to age 19. Therefore, in 1944, the examination covered two cohorts at the same time. To define a priority for entering the military service, the examination classified male into five categories; 'Kou ', 'Otsu ', 'Hei ', 'Tei ' and 'Bo '. Only the first two categories ('Kou' or 'Otsu') were regarded as appropriate to the active service. The last two categories indicate unsuitability for the military service; namely, disabled ('Tei') or sick ('Bo'). It should be noted that the evaluation was based only on his physical and health condition, independent of the economic condition of his household or other personal socioeconomic status. Cognitive skills and / or economic conditions of the participant were not considered. There has always been an argument that the member of wealthy family or the elite of society could evade conscription; however, the evidence is at most anecdotes up to the 1920s when the conscription rule actually had exclusions. After 1927, the major revision of the conscription institution, there remains no evidence of bias. As results, the yearly shares of the first category 'Kou' were stable at between 30% and 40%, at least since 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, after receiving the number of its assignment, chose the male from those that were *most recently* classified as 'Kou' at first, then from those classified as 'Otsu'.

The above conscription rule implies that a man's probability of having military service largely depends on the number of mobilizations at his age of 20 (or 19). And the number of mobilizations increased rapidly, especially after 1943. A rough estimate shows that the numbers of drafted male directly after the examination were 170K in 1937, 320K-360K in between 1938 and 1943, 1,000K in 1944, and 500K in 1945. Therefore, the chance to be assigned to the active military service was concentrated in certain cohorts; especially, 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 between 1938 and 1943, and around 70% of the cohort was drafted in 1944 and 1945, while the share was about 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 to 4K males of the age of 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 consist of 160 infantry regiments.

Generally, the regimental system has two principles. First, it functions as an independent front-line military unit. The regiment is responsible for recruiting new soldiers, equipping weapons, and training the troops. Second, it also works as an administrative

[6] 尾山 (1943), p.4. Current age on the last day of November in the previous year. The exact time of examination was varied between 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

[8] Usually 1-99 numbered regiment

[9] Usually 100 to 299 numbered regiment

unit in peacetime. Once drafted, the majority of Japanese male belonged to a certain regiment and their military service record was registered in the regiment.[10] Even after retiring from active service in IJA, they maintained their relationship with the registered regiment and sometimes were called back for the ceremony. Furthermore, because the regiment organization was based on geographical territory, it was closely accompanied by the community of hometown. This is why the regiment was often called *hometown regiment* in Japan.

The regimental system was once the major principle of organizing a modern army all over the world, at least 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 basically inherited to *Wehrmacht*.[11] The British Army also had a long tradition of the regimental system for more than hundreds of years. Edward Cardwell and Hugh Childer had reorganized it by 1881, clarifying its territorial principle.[12] The Third Republic of France reorganized its regimental system by the law of July 24, 1873. They divided the continental territory into 18 self-contained Army Corps Regions. Boehnke and Gay (2020) discussed French Army was organized by the territorial regimental system at the beginning of the First World War. Until after the First World War, a standing strength of the United States Army had been organized only as territorial regiments. In other words, divisions and brigades were the temporal organizations during the wartime, and these were dissolved when it returned to the peace time.[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 (e.g. the Napoleonic War, the American Civil War), one of the most essential elements of war-fighting was to keep the group morale / discipline of infantry.[15] Second, it may prevent the Army from carrying out coup d'État. For example, the fragmentation of the regimental system prevented the British Army from becoming the political unstabilizing factor.[16]

However, the territorial regimental system started decaying since the middle of the First World War. First, without conscription, the regiment could not afford the necessary capacity for wartime 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 recover its strength by the supports only from its territory, because of the tremendous loss of soldiers, horses, and supplies as the case of France.[18] Third, the infantry tactics in the battlefield changed from mass charge to small-group fire and movement. Small-group tactics utilize several types of branch even at the level of regiment (artillery, engineers,

[10] On the other hand, the Imperial Japanese Navy had a centralized registration system.

[11] Bartov (1991) pp.30-35. In this book, he used the word of "primary group" rather than *Kantonsystem* and he discusses 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.

[17] French (2005).

[18] Bourlet (2010).

medic, telegraph, machine guns, etc.) that require the soldiers for specific skills and experiences. The territory of the regiment was too small to retain a sufficient number of soldiers for every branch. As results, 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, at least during the First World War.[19]

On the other hand, IJA maintained the hometown regimental system even at around 1940. In addition, the infantry regiment remained the main military power in IJA through the Second World War. Poorly mechanized, IJA increased infantry branch up to about 50% of total strength, whereas the United States Army declined 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 three million soldiers over from the North East of China, Continental China, Indo-china, Thailand, Burma, Malaysia, Java, Borneo, Philippine to New Guinea. Because the intensities of fighting depended on the confronting opponents, the causalities of troops were varied by their garrison. A governmental report by the Ministry of Welfare in 1963 estimates the ratio of the number of KIA/MIA to the number of soldiers at the end of the Second World War by the theatre.[20] They are varied from 0.07 in the North-East China to 3.92 in Philippine, whereas the overall average counts 0.27.[21] Assuming that there was no troops that moved from their garrison until the end of war, the probability to be killed or missing was 21% on average in the Japanese military force. The most severe theatre was Philippine that counts 80%; they were high enough at 70% in the Central Pacific islands, 67% in Burma and 61% in the South Pacific islands. On the other hand, it was 7% in the North-East of China. These ratios are consistent with the fact that, the high-intensity fighting was concentrated in the islands in Pacific Ocean, since the counterattack by the Allies was mainly made by the United States Force.

Some may argue that IJA allocated the regiment that were from the area with a particular economic attribute (e.g. agricultural area) to the face of the United States Force. However, in addition to the fact that there is no evidence that IJA intentionally allocated

[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 numbers for other theatre is estimated as is since 1941.

[21] 212.2M/788.9M.

the regiments based on the attribute of hometown, to our best knowledge, IJA even could not exactly anticipate the severeness of a particular battle field because of the strategy that the United States Force took.

THE PACIFIC ISLANDS FRONT AND LEAPFROGGING. — During the Second World War, the Japanese force occupied the South, the South West and the Central Pacific islands to intervene the communication and transportation between the United States and Australia. The first objective of strike back by the United States Force was to secure the approaches to Japan by expelling the Japanese force from these islands. For this purpose, they took 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]

As a result, among the Pacific islands, some islands were hit directly by the United States Force but some islands were passed through. The targets of attack were not precisely anticipated by IJA and often substantial troops got isolated. 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. — An example is the battles in the Palau Islands. Japan began colonizing the Palau Islands as a result of the Treaty of Versailles in 1919. The Japanese government established one of their local branches in Koror, and provides the public services to nearby islands.[23] At the beginning of Japanese domination, the population of islands sums up to 54,358 at the end of 1923, which doubled to 113,562 at the end of June 1939.[24]

On April 1944, anticipating the counter attack by the United States Force, IJA dispatched the 14th division (14D) to defend the Palau Islands. The main strength of 14D was constructed from 3 infantry regiments and each infantry regiment had 3 battalions in IJA; the 2nd 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). 2iR and III/15iR were located on Peleliu Island and I/59iR was located on Angaur Island. The remaining 4 battalions of 15iR and 59iR defended the main island of Palau; Koror island. 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 having occupied the Mariana Islands on July 1944, the United States Force set the next target as the Philippines. Since the Palau Islands are

[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.

located in the middle between the Mariana Islands and the Philippines, the United States Force had to secure there 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 finally succeeded in expelling the Japanese force from the island. As a result, 2iR, III/15iR, and II/15iR (that counterlanded from Palau as reinforcement) were completely destroyed. Within 6,822 soldiers, only 190 survived. US 81st Infantry Division invaded Angaur Island two days later from the landing of 1st Marine Division. After 31 days, almost all of 1,194 soldiers of I/59iR (and 6 soldiers of Navy) lost their life in the island other than 59 as POWs. However, the United States Force left the main island of Palau and Yap. At the end of the War, about 25,000 soldiers from the mainland of Palau and 5,500 soldiers from Yap returned to Japan (in total of 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 POW relative to KIA/MIA. Usually, when an army is defeated on the ground, a substantial number of soldiers is captured as POW by the opponents; some may surrender and some wounded are left behind. For example, the German 6th Army during the Second World War, after losing 147,200 soldiers, leaves 90,000 as POW when General Paulus surrendered at the end of January 1943 at Stalinglad; the ratio of KIA to POW was 1.6 at one of the most dreadful battlefields in Europe. On the other hand, in the battle of Peleliu and Angaur Isalands, these ratios were 42.5 and 19.2, respectively.

The small number of POW of the Japanese force may have been due to mental training by their military code. On Jan 1941, the Japanese force revised their military code and stressed the culture of shame. Since they defined surrender as shame, even after they exhausted arms, that their family should be punished, 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, this is a general mechanism reinforced by the hometown regimental system in the battle field.[26]

As a result, 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 male massively, relative to other cohorts as much as relative to other prefecture. More concretely, the battle in Palau resulted in the loss of almost the entire infantry regiment of Ibaraki prefecture, while Gunma prefecture and Tochigi prefecture lost two-thirds and one-third, respectively. When we approximate the numbers of missing men are 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 single battle in Palau islands, and this variation reflects to the decline in the gender ratio of each prefecture.[27]

[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

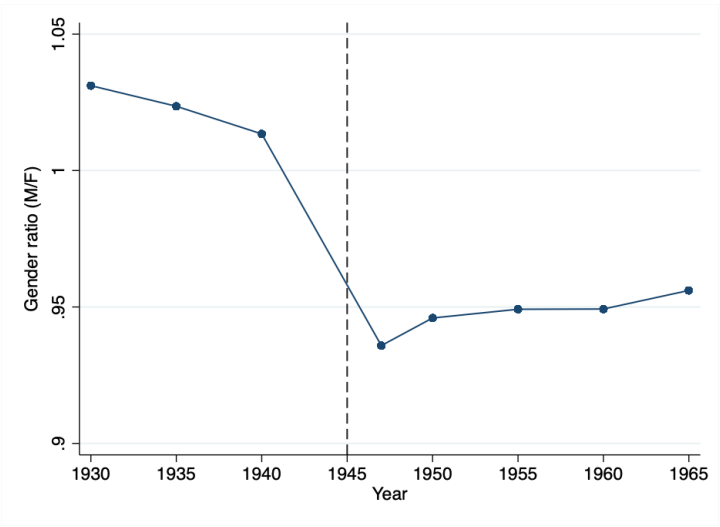
[26] Benedict (1946). Costa and Kahn (2003) discussed about it during the Civil War in the United States.

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

C. *Its consequences in Japan: decline in gender ratio*

During the Second World War, approximately two 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. This fact is mirrored in the 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 to 64 years dropped from 1.014 to 0.939 between 1940 and 1947.[28] Importantly, the gender ratio was permanently imbalanced, even though the population continued to grow, helped by the postwar baby boom.

Figure 1. : Evolution of gender ratio at national level



Note: The figure shows the gender ratio (defined as 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.

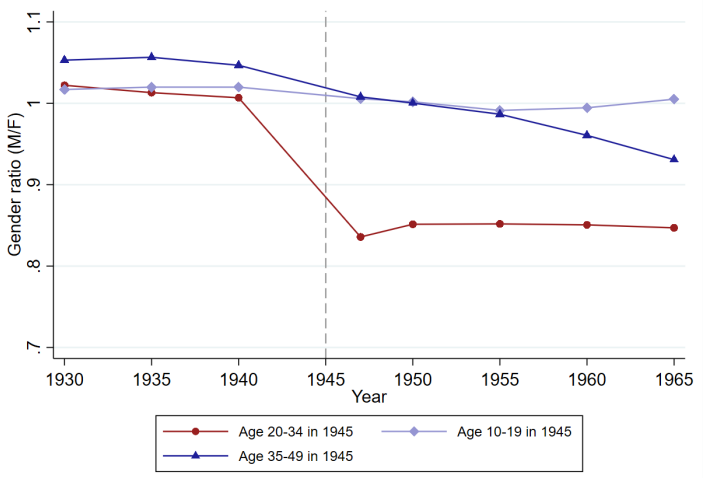
We interpret this decrease in the gender ratio as a gender-specific loss of human capital. Since a comprehensive list of war casualties is not available in Japan, we are unable to directly use death counts by cohort-prefecture, and instead approximate them by the gender ratio changes. Nevertheless, it 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. On the other hand, one caveat of using gender ratio is that when it is

[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.

defined area by area, it is also affected by social migration between areas. However, the gender ratio change seems to approximate the death of male solders well. First, the estimated decrease in the young male population using the cross-prefecture variation of the gender ratio between 1935 and 1947 is 1.75 million, very close to the estimated 2 million soldier deaths from the official statistics.[29] Furthermore, in a case study in Iwate prefecture, where the local authority created the list of war casualties, we find that roughly 70%-80% of the change in the gender ratio is consistent with the death of male soldiers. Note that casualties from air raids could also lead to a human capital loss; however, they were much fewer in number (approximately 300,000) and gender neutral.

The erosion of the gender ratio was particularly strong for the younger cohorts most heavily recruited for the war, as described in Section III.A. Figure 2 shows the evolution of the gender ratio across 3 different birth cohorts, which were, respectively, 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 to 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 were hardly affected, consistent with the fact that they were young enough or old enough to avoid military recruitment.

Figure 2. : Evolution of gender ratio, by cohort

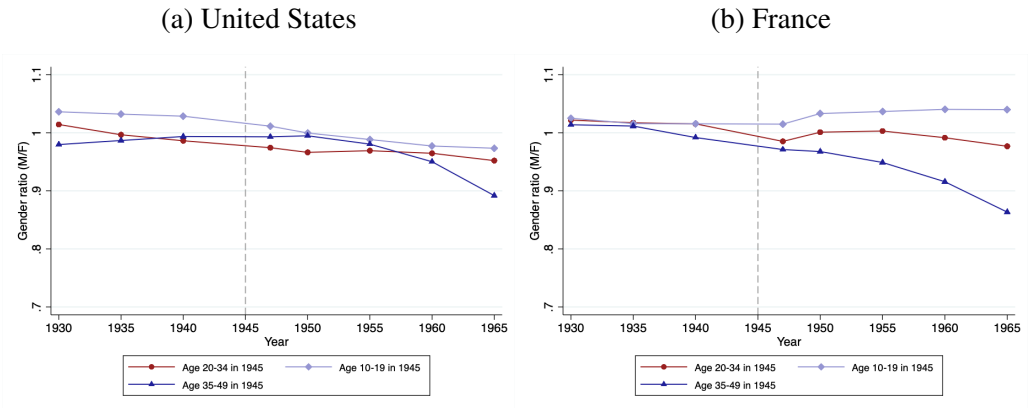


Note: The figure displays the evolution of the gender ratio of three cohorts groups, those aged 20-34 in 1945 (treated cohorts), and aged 10-19 and 35-34 in 1945 (control cohorts), between 1930 and 1965. The figure indicates that the decline in the gender ratio was concentrated among the young cohorts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

[29] We calculated the estimate decline in young male population in the following way: first, for each prefecture, we estimate the decline in male population of the treated cohorts (aged 20-34 in 1945) by multiplying population in 1935 of these cohorts by their change in the gender ratio, and divide this by 2. Then we aggregate these estimates for all prefectures. This aggregate estimate of male disappearance can be significantly different from the estimated number of solder deaths (from the official source) if there is systematic migration between large and small prefectures. Our estimate is 1.75 million, fairly close to official statistics, which confirms that the change in gender ratio approximates the death rate relatively well.

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 include military overseas. As is clear from the first panel, 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 prewar level by 1950, as detained soldiers and evacuees return. The figure suggests 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.

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



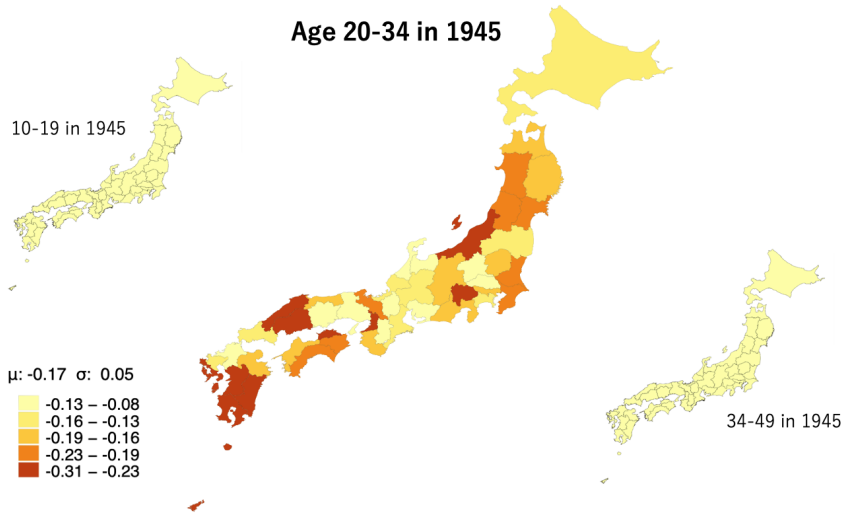
Note: This figure repeats Figure 2 for the United States and France, displaying the evolution of the gender ratio of three cohorts groups, those aged 10-19, 20-34, and 35-34 years in 1945, calculated based on the population census of each country. The figure for the United States include military overseas, 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.

A key ingredient to this paper is the substantial geographical variation in the magnitude of gender ratio changes. This is a combined result of the hometown regiment system of the IJA and the US military strategy, as explained in Section III.B. 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 changes is not systematically related to factors such as urbanity. For example, Tokyo's gender ratio decline was 0.16 (27th out 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 varies

[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.

from Kagawa prefecture's 0.24 (5th) to Ehime prefecture's 0.17 (21st). Again, the figure confirms that the gender ratio change is a unique experience of the treated cohorts. If the change in the gender ratio is entirely due to prefecture-specific factors such as migration, we would expect a similar gender ratio changes for the other cohorts. The two side figures on the changes in the gender ratio of the two control groups confirm that this is not the case.

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 size of 1935-47 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.

IV. Data and sample for analysis

DATA. — Our main outcome of interest is the industrial share of employment, which should be computed for each prefecture-age-industry-gender-year combination, to take the advantage of our identification source. To obtain these variables, we use published tables of Population Census between 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

[31] The microdata of Population Census in Japan is available only for those after 1990, because the Japanese government abolished them for those before 1990.

[32] For the other years, they are only reported by prefecture-gender or by age-gender.

1955, it is available decennially for 1920, 1930 and 1940, the definitions of employment status used in the prewar census were different from the the post-war census, making it difficult to make meaningful comparisons before and after the war. For this reason, we restrict to the years 1955-1980 for our main regression estimation, and prewar employment information is only carefully used for supplementary analyses and robustness checks. We digitized these information for 46 prefectures (excluding Okinawa) and 11 age groups (age 10-14, 15-19,..., 60-64). We use 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 (i.e., 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 also calculated for each prefecture-age combination in the same vein.

SAMPLE AND DESCRIPTIVE STATISTICS. — Our estimation sample is an unbalanced panel of prefecture-cohort pairs from 1955 to 1980. We follow the 8 birth cohorts that were aged 10-49 in 1945. Among these birth cohorts, the three cohorts whose age was between 20-34 correspond to our "treated groups", whereas the other 5 cohorts aged between 10-19 or 35-49 are termed "control groups". The cohorts leave the sample after they pass 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 one hundred thousand, with the gender ratio of 0.90 in 1995-1980 without any population weighting. Mean employment share (without population weighting) is 0.38 for agriculture, 0.24 for manufacturing, and 0.38 for service. Employment rate is 0.75 on average, whilst 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, the table also provides the prefecture-level information. Average population size across prefectures grew by approximately 50% between 1955 and 1980. The periods that is well known as "high speed growth era" is characterized by rapid industrialization. On average, the agricultural share of employment dropped from 0.46 to 0.14, while the manufacturing share increased from 0.20 from 0.33. The employment share of the service sector also increased from 0.33 to 0.53. The impact of the loss of young men is, therefore, examined in this specific context, that is, 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 prewar and postwar years.[33] When a cohort

[33] We use 1935 as the prewar reference year, instead of 1940, because Japan was already in the Second Sino-Japanese

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 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 instrumented 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: that is, the permanent loss of male soldiers would reduce the employment share of manufacturing, which uses male labor in production relatively more intensively, hindering postwar industrialization. In theory, however, it is ambiguous whether the loss of young men would alter the industrial structure or not. The Rybczynski theorem from the international trade literature predicts that, based on the classical Heckscher-Ohlin model, the economy responds to an exogenous increase in endowment of one factor by increasing the output of the sector that uses that factor relatively more intensively (Rybczynski, 1955). On the other hand, 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 (e.g., 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 at play in the context of immigrant influx, but found a larger role of within-sector adjustments. The decomposition method in the international trade literature lead some to find a similar conclusion (e.g., Blum (2010)). Overall, to what extent the change in factor supply is absorbed through bewteen- and within-sector adjustments depend on parameters such as substitutability between production factors, and economic contexts considered.

Applying the tentative conclusion of literature, this article hypothesizes 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

War from 1937 and military deployment has already commenced.

realized in a very short period of time, rather than gradually.[34] Second, though the role of the demand side on female labor supply was acknowledged in the related literature, as summarized in Section II, the analysis in a unified empirical framework and context has not been implemented. Lastly, our context also 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 (e.g. due to 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. In Section III, we discussed why the death rates of soldiers were likely to be random across prefectures from the historical and institutional perspectives. However, there are a few concerns with respect the plausibility of the exogeneity assumption. First, the variation in the gender ratio might be correlated with the prewar industrial structure. For example, conscription used physical examination to assess health conditions and physical strength (i.e., 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 gender-specific migration, a part of the variation may come from the pattern of cross-prefecture mobility 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 prewar prefecture-level characteristics. In particular, we study how they are correlated with the agricultural share of employment.[35] Changes in gender ratio are not aggregated at the prefecture level to have more statistical power. We do not include other outcomes such as female labor force participation, since they are highly correlated with agricultural employment share.

Table 2 show 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

[34] Remember that the majority of the soldiers death occurred in the last few years of the war.

[35] Our digitized data only 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 again that the employment concept of the prewar census is not the same as that of the postwar census, requiring caution in interpretation.

the postwar periods, this result raises the concern that any detected negative effect of the gender ratio change on postwar industrial development merely captures the permanent differences in industrial compositions across prefectures.

However, column (2) shows that adding urban prefecture dummies and population size makes the coefficient on agricultural share virtually zero. 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 remain very small and indeed slightly positive. The obtained R^2 across the models are small, indicating that a large share of changes in the gender ratio changes are not correlated with aggregate observable characteristics.

The simple regression results above support the exogeneity of the gender ratio changes with respect to industrial compositions, once we add appropriate controls. Of course, this is only partial evidence, and we are not able to perfectly control for prefecture-level unobservables. This motivates our estimation strategy to take a second difference between cohorts, on top of the cross-prefecture differences. As will be outlined below, our regression specification includes prefecture fixed effects, which controls for differences in prewar industry characteristics and prefecture-specific unobservables, such as gender norms.

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

	Change in gender ratio 1935-47 (treated cohorts)		
	(1)	(2)	(3)
Agriculture employment share in 1930	-0.049	-0.000	0.007
	(0.052)	(0.095)	(0.111)
Urban prefecture dummy		0.050	0.045
		(0.040)	(0.044)
Log of prime age population in 1935		-0.023	-0.013
		(0.020)	(0.024)
Height-weight ratio (1939 examination)			-1.277
			(1.636)
Log of air raid casualties			-0.005
			(0.005)
Observations	138	138	138
R^2	0.007	0.021	0.031

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 prewar census is not directly comparable to the postwar 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 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 by taking advantage of the variations between prefectures and between 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 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 uni-dimensional fixed effects; $Cohort_c \times Time_t$ and $Pref_j \times Time_t$ represent respectively the cohort-year fixed effects and the 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 the control cohorts, between 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, which we study separately. 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 (e.g., capital destruction) are absorbed by prefecture fixed effects. Time fixed effects control for economic conditions in a given 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 prewar local industrial composition of employment, when that is also different across cohorts (or across ages). To give an example, young people may have worked temporarily in factories located in a neighboring prefecture, or they 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.[36] We deal with this issue in twofold. 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.

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 becomes more negative for a larger loss of men, the signs of the coefficients must be flipped to interpret the effect of the gender specific human capital loss.

The first three columns study our main outcomes of interest, 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.[37] 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 postwar 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. In the period 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, respectively (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.[38]

[36] The direction of the bias is not clear a priori, however. 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 take between cohort differences, 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 prewar period in agricultural intensive prefecture compares to the corresponding difference in the manufacturing intensive sector —- it is difficult to say whether this difference in the differences is positive or negative.

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

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

B. *Labor force participation and employment*

Column (5) indicates that the skewed gender ratio has led to a higher female labor force participation rate, consistent with the previous literature. To put into perspective, 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, 2020; Cardoso and Morin, 2018).

There are several potential explanations for a smaller effect that we find for women's labor force participation. First, labor supply could have been constrained by child care duties in a high fertility postwar period — total fertility rate exceeded 4 before 1949, which gradually decreased until it stabilized at slightly above 2 in the 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 (2018) that showed a significant association between imbalanced sex ratio and higher fertility in the postwar Japan. Second reason might relate to the gender norm. The new constitution was enacted in 1947, guaranteeing the political, economic or social equality between men and women. While the society was moving toward a progressive gender norm, it is unclear if the institutional barriers and normative biases were dissolved in practice, especially relative to the economic participation. [39] Third, 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 men (Boehnke and Gay, 2020). It is possible that the labor supply decision was coordinated within household, whereby other members of the household supply labor instead of the women in the treat cohorts. In this case, our diff-in-diff will be attenuated because of this spillover effect across cohorts.[40] This interpretation is consistent with the supplementary results in Section VII.C where we find a larger and comparable 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.[41]

In contrast, there is no significant association with the male labor force participation

[39] 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 postwar period, but not labor market participation in 1950. This is consistent with the existence of various barriers for women's integration to the labor market.

[40] 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).

[41] 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, had there not been the change in the demand side, the estimated effect on the female labor force participation could have been even smaller. In addition, military pension for bereaved family does not explain the lower effect, because it did not exist until 1952 and is not generous enough to negatively affect labor supply decisions on the extensive margin.

rate as shown in columns (5) and the estimated coefficient is very small. Column (4) indicates that the employment rate has improved when the gender ratio is reduced, consistent with the idea of an improvement in labor market tightness due to the scarcity of male labor, although the coefficient is not statistically significant.

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 1947-55	-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 1935-47 (first row). Column (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. Column (4)-(6) explores 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 in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

C. Change in gender mix within sector

As discussed in Section V.A, the change in the composition of human capital could also 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 a relatively abundant factor, that is, female labor. Even without the change in technology, hiring of female labor may intensity if there is sufficiently high substitutability between two genders in a given 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. It is clear that in all sectors, share of women in employment is larger when when more men are lost. 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 benchmark sector female shares in the prewar period based on the same employment definition. Nevertheless, the coefficients suggest that the gender ratio change explains larger share of female share variation within-sector than the variation in sector share of employment studied in Table 3.[42]

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 1947-55	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 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 in male labor. In the case of Japan, due to the loss of men during the war, the economy has shifted toward sectors that use relatively more female labor (agriculture and service) than the male-intensive sector (manufacturing). At the same

[42] For example, for manufacturing sector, a reduction in the gender ratio by 0.37 leads to a higher female share in 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.

time, within all sectors, the presence of women has 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 female labor force participation to increase.

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 the 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.

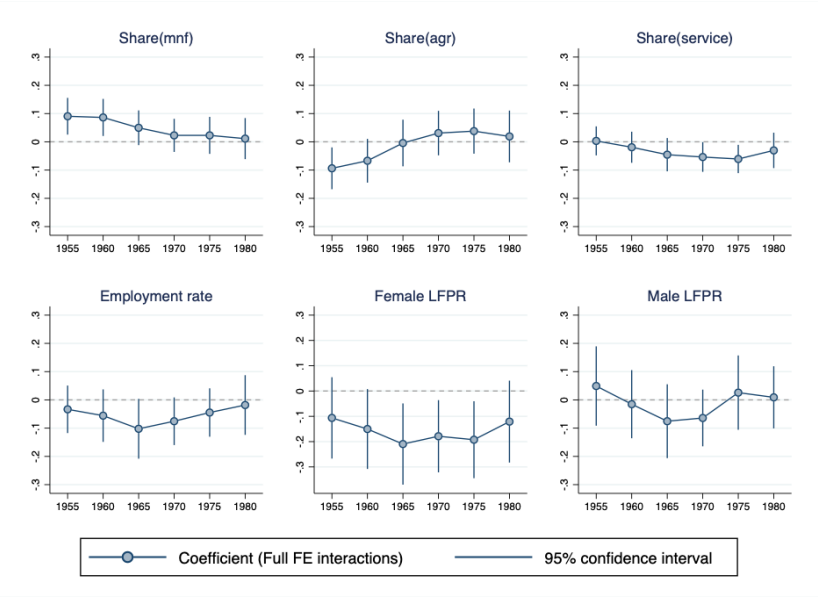
Figure 5 shows the estimated time-varying coefficients and their confidence interval for different outcomes. The top panels indicate that it was only until 1960 that the decline in the gender ratio had a negative impact on the employment share of the manufacturing sector, which mirrors the higher share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

The bottom-middle panel shows the time-varying effects on female labor force participation. Looking at the sign of the effects, the impact of permanent male loss had always been to increase participation of women and it was long-lived even 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 in age 20-34 in 1945, so that women from these cohorts were still constrained by housework and child care early in the postwar 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 trans-

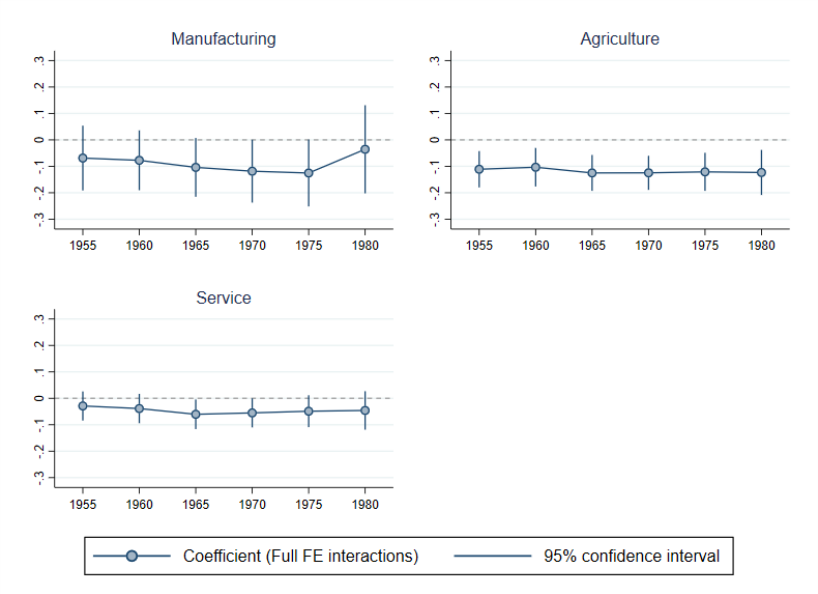
Figure 5. : Long-run effect on industrial share and employment outcomes



Note: The figure provides the long-term effects of the gender ratio changes on industry share of employment (top 3 panels), and employment and labor force participation (bottom 3 panels). These are estimated using the interaction of year dummies with the change in gender ratio in the specification of Equation 3, that 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.

formed 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 case of Japan where many soldiers did not return, the loss of men increased female labor force participation permanently in the postwar period. Moreover, in the short-run, the economy has 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, that 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-specific age effect matters?*

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. In particular, 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. To see weather such prefecture-age effects bias our estimates, we run a placebo test using the data from 1930, the closest prewar census for which labor force participation is available at prefecture-cohort level.[43] 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.

B. *Alternative definition of gender ratio*

We defined $\Delta Ratio_{cj}$ as the change in the gender ratio from 1935 to 1947 of the same *cohort*. A weakness of this definition is that the age of the cohort at the reference year (i.e. 1935) matters. As discussed above, it is possible that $\Delta Ratio_{cj}$ changes due to prefecture-age specific migration, which could correlate with prefecture cohort unobservables. In addition to the placebo test we provided, we perform 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 compare it to the gender ratio in 1935 of the same *age group*. Formally, for each cohort-prefecture pair, we construct:

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

where $a(c)$ stands for 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 $Ratio_{a(c)j,1947}$ is the gender ratio of a cohort c in 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

[43] In the prewar census, definition of employment was different and it was a closer concept to the postwar labor force participation. We use the prewar 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 of are regressed on some of the labor force participation rate in 1930 of the same age groups. Standard errors are clustered at prefecture level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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 exactly the same for both (Equation 2). Standard errors are clustered at prefecture-year level. Standard errors in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

cohort c in 1947.[44] 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 6 compares the coefficients of the two alternative gender ratio estimated from the base specification with full fixed effects. In general, the signs and the qualitative implications are similar. Using the alternative gender ratio change, the effect on the manufacturing share in employment and agriculture are larger and statistically significant at 1% level, while the coefficient on the service sector becomes null.[45] 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.

C. *Control for other cohort*

In our main regression framework, it is assumed 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. Insofar as there 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 own-cohort effect becomes much larger for all the outcomes considered, while the signs are maintained. For example, the manufacturing coefficient is estimated to be 0.521 and significant at 1% level. The magnitude is relatively large: the decline in the gender ratio led a prefecture cohort to have a 5.4 point lower share in manufacturing employment and 3.77 point and 1.4 point higher for the agriculture and service sector, respectively. The coefficient on female labor force participation are also higher.[46]

[44] For example, for a cohort who was aged 30-34 in 1945, this alternative gender ratio change is calculated by the difference between the gender ratio in 1947 when they are aged 32-36, and the gender ratio in 1935 of the *other* cohort who was 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.

[45] 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.

[46] The effect of other cohorts is even larger. It is because of a much smaller variation due to the pooling 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

The increase in own-cohort effect is in fact more linked to the reduced coefficient of log of population in 1935 (which declined from -0.034 to -0.008 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 tells the own-cohort was possibly even larger than originally estimated.

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 calculate simply by the gender ratio change of all other cohorts combined in the same prefecture j excluding own cohort c. The change The estimated coefficient of the change in the gender ratio on own cohort (first row) is much larger than before, suggesting that that our previous estimates are 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 in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

cohorts (recall that we have 3 treated and 5 untreated cohorts), whose gender ratio change might be correlated with the unobservables.

D. Attention to prefecture-level variation

Our mains specification uses industry employment share at cohort-prefecture level to take advantage of 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 assume cohort-level segmentation of the labor market. However, if the labor market is not perfect segmented by cohort and there are between-cohort substitutions, our estimate will be attenuated due 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 labor stock at the prefecture level. This motivates a specification that exploits only the cross-prefecture variation in the gender ratio change. 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 (dummies for large prefectures - Tokyo, Kanagawa, Osaka, Aichi, Kyoto, and Hyogo). Moreover, we also add the prefecture-level outcome in 1930 (of working population) to control for the permanent prefecture characteristics. We also 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 significant. They are larger than the previous estimates, consistent with the idea that our main specification (with cohort-prefecture variation) underestimated the effect of the gender ratio changes. On the other hand, for employment outcomes, the cross-prefecture variation provides a very different picture. In particular, the female labor force participation rate is significantly lower in prefectures with a larger male loss.

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)
Change in the gender ratio (treated cohorts)	0.192***	-0.196***	-0.085***	0.134***	0.192***	0.033**
	(0.055)	(0.050)	(0.032)	(0.042)	(0.068)	(0.016)
Bomb casualty per capita	0.365	-0.653**	0.331**	0.630***	1.080***	0.095
	(0.273)	(0.284)	(0.156)	(0.188)	(0.283)	(0.113)
Destroyed building per capita	0.253**	-0.098	-0.201**	-0.049	-0.119	0.013
	(0.104)	(0.149)	(0.096)	(0.098)	(0.160)	(0.043)
Height weight ratio	-2.607***	5.019***	-1.293***	0.657*	1.308**	0.227
	(0.695)	(0.642)	(0.341)	(0.367)	(0.599)	(0.197)
ln(pop) change 1935-47	0.129**	-0.198***	0.121***	-0.122***	-0.263***	-0.011
	(0.055)	(0.053)	(0.033)	(0.043)	(0.071)	(0.017)
Gender ratio in 1935	0.107*	-0.184***	-0.077	-0.182***	-0.235***	-0.017
	(0.057)	(0.070)	(0.049)	(0.050)	(0.082)	(0.017)
ln(pop) in 1935	0.020	-0.163***	0.075***	-0.019	-0.078	0.013
	(0.027)	(0.040)	(0.028)	(0.040)	(0.065)	(0.013)
ln(worker)	-0.017	0.160***	-0.076***			
	(0.026)	(0.040)	(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
R-squared	0.834	0.933	0.856	0.787	0.771	0.624
Observations	2484	2484	2484	2484	2484	2484

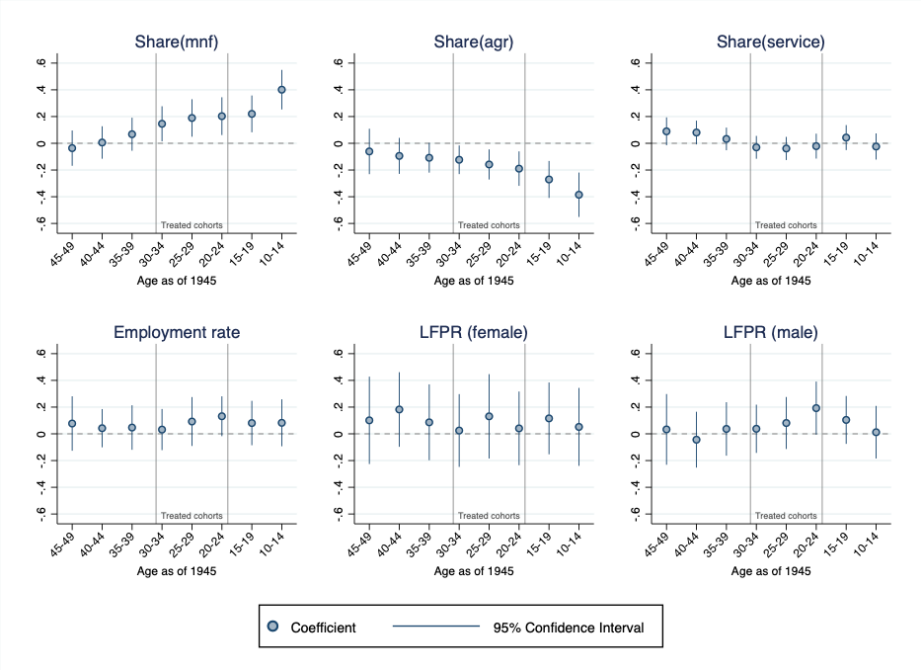
Note: The table provides the regression results where we only exploit cross-prefecture differences in the gender ratio change (of treated cohorts who were aged 20-34 in 1945). While maintaining the same specification as Equation 2, prefecture FE and prefecture-year FE cannot be included. To minimize the endogeneity of the gender ratio change, we include a number of 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 in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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 to 1955 and 1960 to examine the immediate effect of the gender imbalance and highlight the mechanism of between-cohort substitution or complementarity.

Figure 7 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.

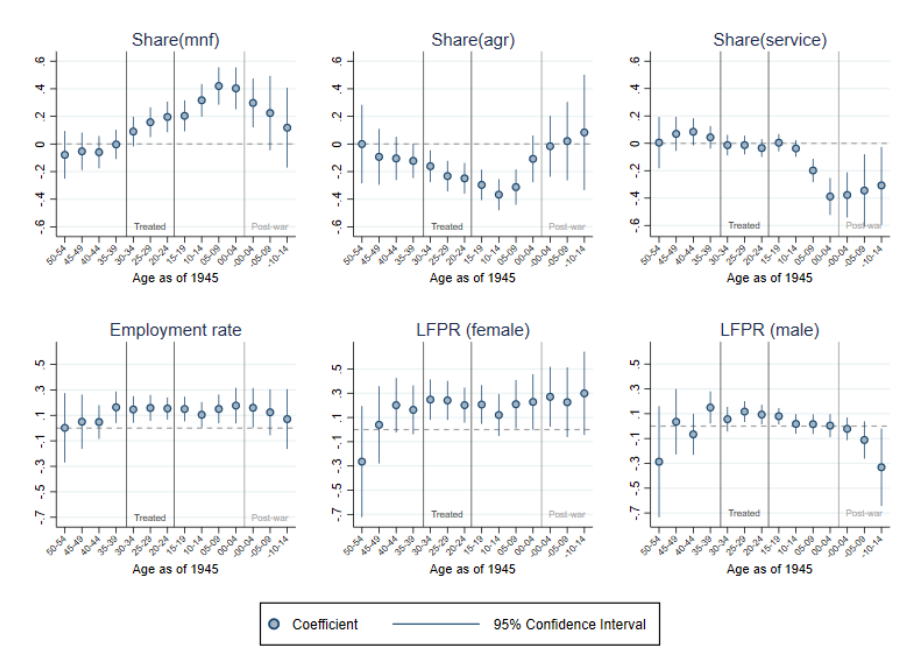
Figure 7. : Coefficient by cohorts, until 1960



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.

An interesting pattern emerges when we analyze even the younger cohort and study the entire 1955-1980 period, shown in Figure 8. It demonstrates that the effect is starting to plateau or decrease as we go done to the younger cohorts, indicative of the catch-up process as the younger and large cohorts enter the labor market. In the case of agriculture, the sign even reverses (although not precisely estimated).

Figure 8. : Coefficient by cohorts, all years 1955-1980



Note: The figure extends Figure 7 by adding cohorts that were not born in 1945 in the sample.

VIII. Conclusion

In this paper, we examined how the permanent loss of human capital is related to economic recovery. In particular, war affects the industrial structure not only through damage to physical capital stocks but also through the tremendous loss of human capital. In the case of Japan in the Second World War, in addition to the fact that the country lost two million soldiers due to the unexpected mismatch of military strategies between offense and defense, the casualties on the battlefield were concentrated in males who were from certain areas and cohorts in the country due to the institutions of *hometown regiments*. Taking advantage of the difference in gender imbalances between geographical areas and cohorts as a consequence of war, we examined the effects of permanent reduction of men 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 only limited to a small degree in terms of quantity and had gradually disappeared after about fifteen years from the end of the war. These findings imply that average technology may have been enhanced by gender during high-speed economic growth after the war. In the medium term, technological change and internal migration may have canceled the permanent loss of human capital in certain geographical areas. We also found 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. Consistent with these findings, the permanent loss of men increased female labor force participation, the effect that persisted over 30 years since the end of the war. The paper highlights the demand-side pulling factors that led to the integration of women into the labor market when men are lost during the war.

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DATA APPENDIX

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 prewar census used different definitions of employment (i.e., Hongyo (本業) in 1920 and Yugyo (有業) in 1930 and 1940), and coverage also was not the same (e.g., not covering self-employees). The definition of employment in 1930 and 1940 is close to labor force participation in the postwar census. For this reason, we sometimes use the employment rate in the prewar periods as a proxy for labor force participation rate. Note also that industry classifications are also slightly different in the prewar and postwar census (e.g., 工業 vs 製造業). When we use prefecture-level outcomes, it is of those aged 20-64 when cohort-level information is available (i.e. 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 simply extrapolated by one minus the employment shares of these two sectors combined. It will be updated after digitizing all other 1-digit sectors so as to make the sector grouping comparable to the one used in the main regression.

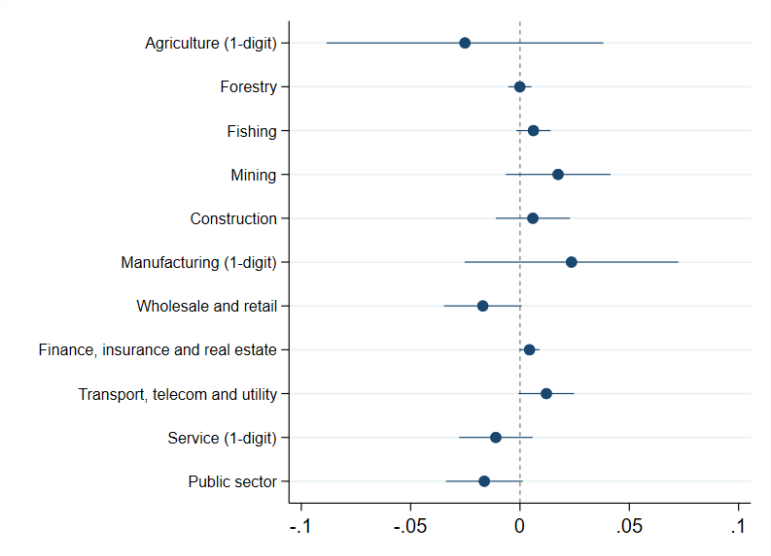
REGRESSION APPENDIX

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.

With regard to 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.

ROLE OF FE INTERACTIONS. — Our main results are base 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 (i.e. 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 signif-

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



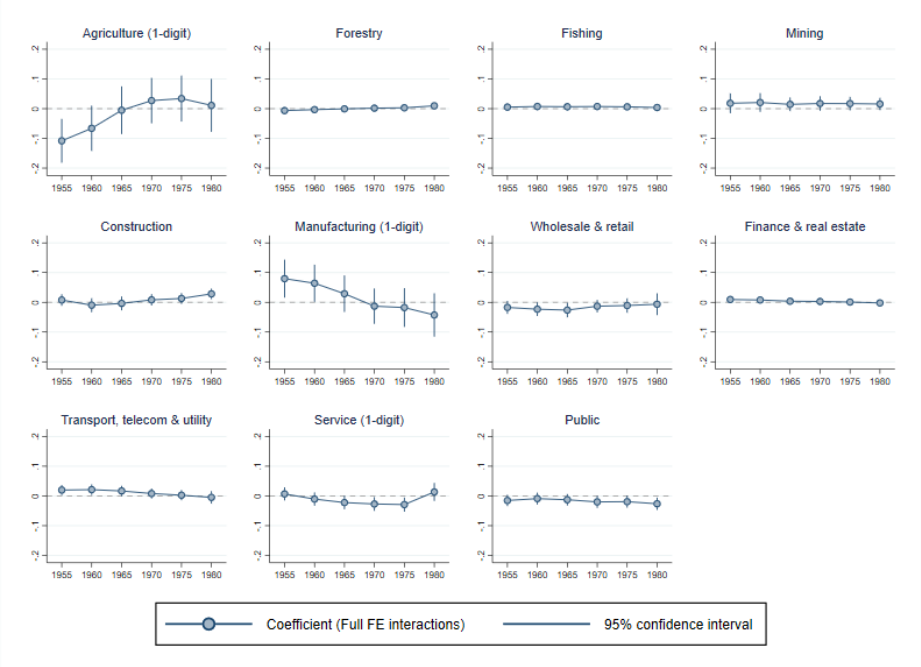
Note: The figure shows the impact of the change in the gender ratio estimated from the specification of Equation 2 where the outcomes are industrial employment shares based on 1-digit sector dis-aggregation.

icantly 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. This is due to the fact that the female labor supply has a strong life cycle component.

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; Table 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.

CONTEMPORANEOUS GENDER RATIO. — The skewed gender due to the war could be alleviated by the migration in the postwar period. We consider that should rather be considered as a mechanism to alleviate the war shock in the long run. In theory, such

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



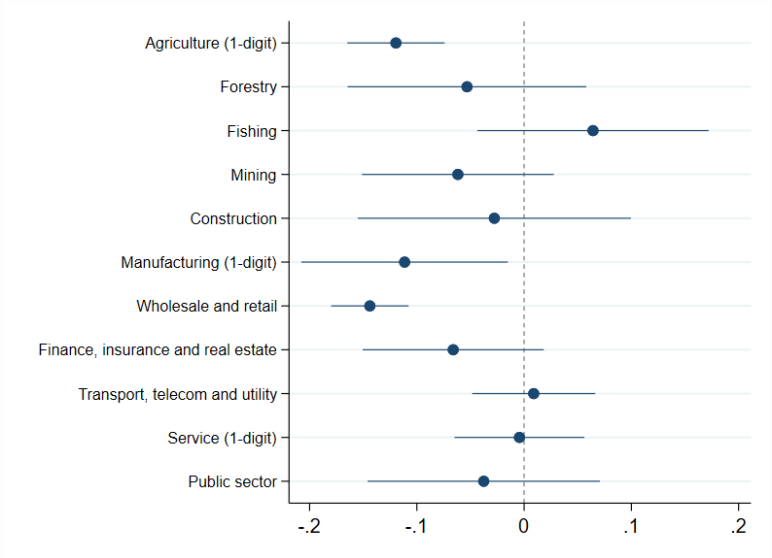
Note: The figure repeats Figure 5 for the industrial employment shares based on 1-digit classification.

effect can be captured by including the contemporaneous gender ratio. We provide the results of the specification that includes contemporaneous gender ratio, in part for the sake of comparison with previous literature. The result is presented in Table B5. Note that the regression suffers a high correlation between the change in the gender ratio between 1935-1947 (our treatment variable) and the contemporaneous gender ratio. This makes the separate identification difficult and in part 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 are consistent for the other outcomes.

ALTERNATIVE DEFINITION OF GENDER RATIO. — In Section VII.B, 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.

CORRELATION ACROSS COHORTS. — Figure B8 provides the correlation matrix across cohorts in terms of the change the gender ratio.

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



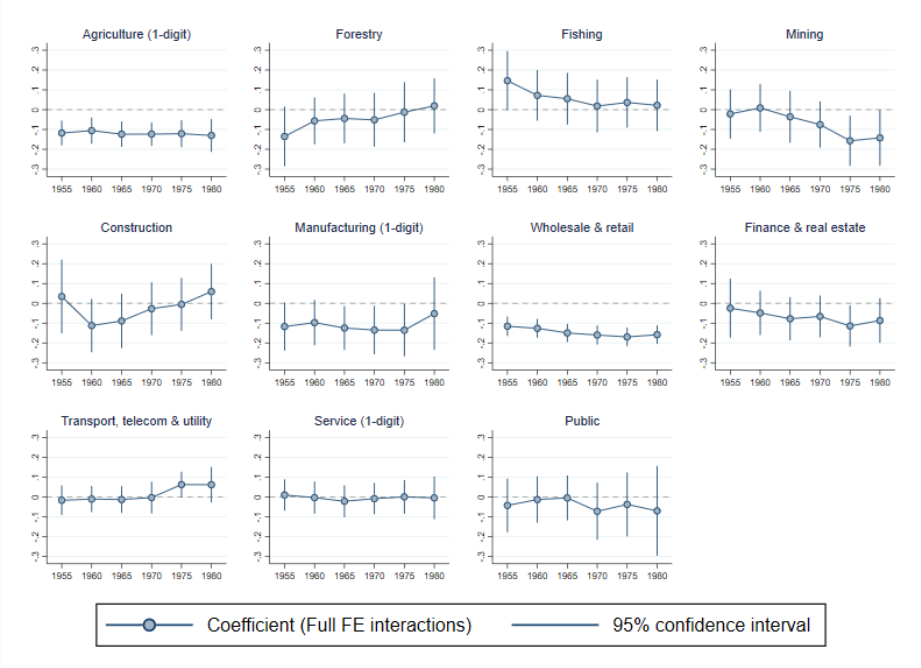
Note: The figure shows the impact of the change in the gender ratio estimated from the specification of Equation 2 where the outcomes are female share wihtin each 1-digit sector.

PREFECTURE LEVEL VARIATION CONTROLLING FOR OTHER COHORTS. — We provide supplementary results to the regression that exploits prefecture-level variation (Table 8). Table B7 controls for the change in the gender ratio of the control cohorts — added separately for those aged 10-19 and 35-49 in 1945.

COHORT SUBSTITUTIONS. — In the main text, we provided the results on the cohort-substitution effect in Figure 7 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 Table B9.

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 B8 show the results. The estimates coefficients have consistent signs, and the impact on agricultural share becomes actually significant.

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.

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 in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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 5 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 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 1947-55	-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

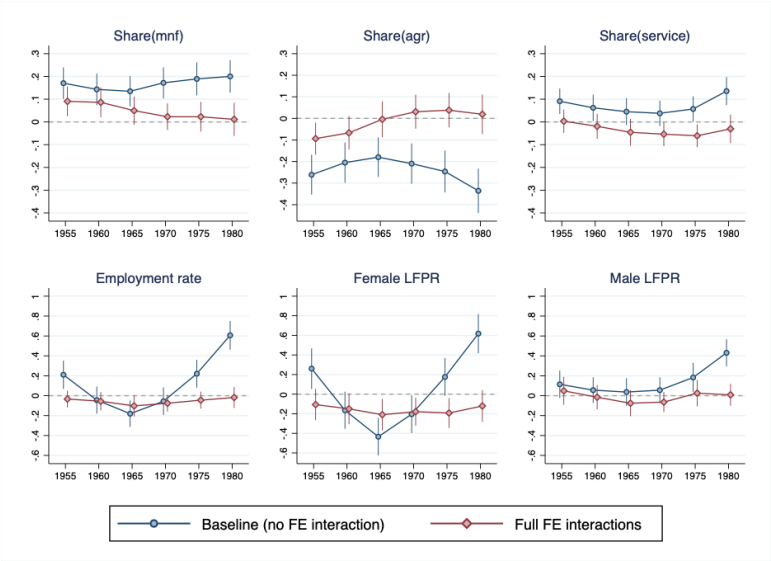
Note: The table shows the results of the time-varying effects in the specification that includes only cohort-year interacted fixed effects. The first row corresponds to the effect of the gender ratio in 1955 and the next 5 rows are the estimated interaction coefficient for subsequent years (representing the difference from 1955 effect). All standard errors are clustered at prefecture-year level. Standard errors 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 1947-55	-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

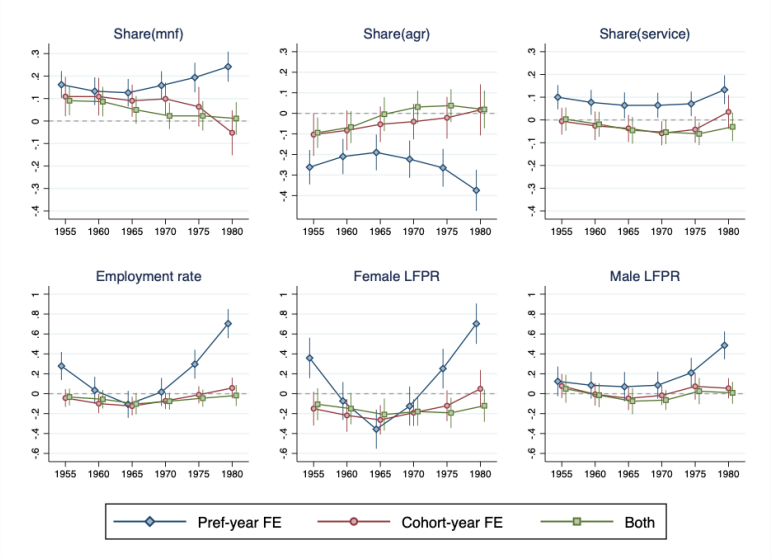
Note: The table shows the results of the time-varying effects in the specification that includes only prefecture-year interacted fixed effects. The first row corresponds to the effect of the gender ratio in 1955 and the next 5 rows are the estimated interaction coefficient for subsequent years (representing the difference from 1955 effect). All standard errors are clustered at prefecture-year level. Standard errors 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.

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

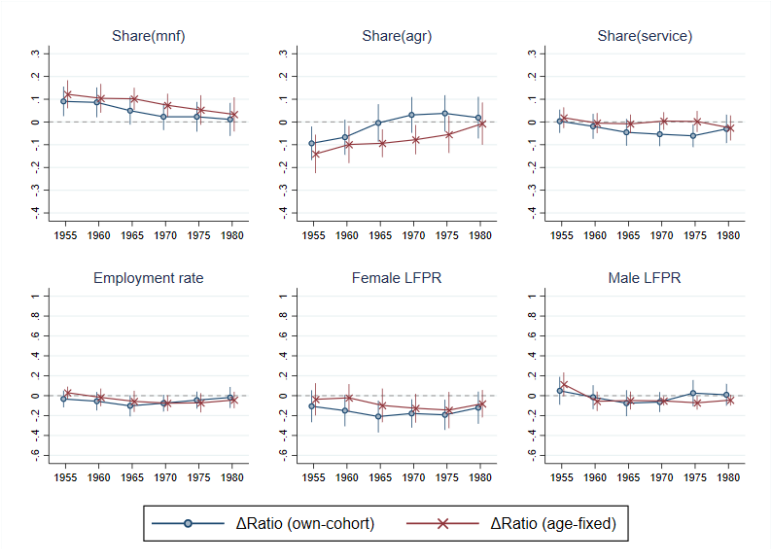
Note: The table shows the results of the regression that adds contemporaneous gender ratio in the main specification. All standard errors are clustered at prefecture-year level. Standard errors in parentheses. $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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, as explained in Section VII.B. All standard errors are clustered at prefecture-year level. Standard errors 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.

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

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

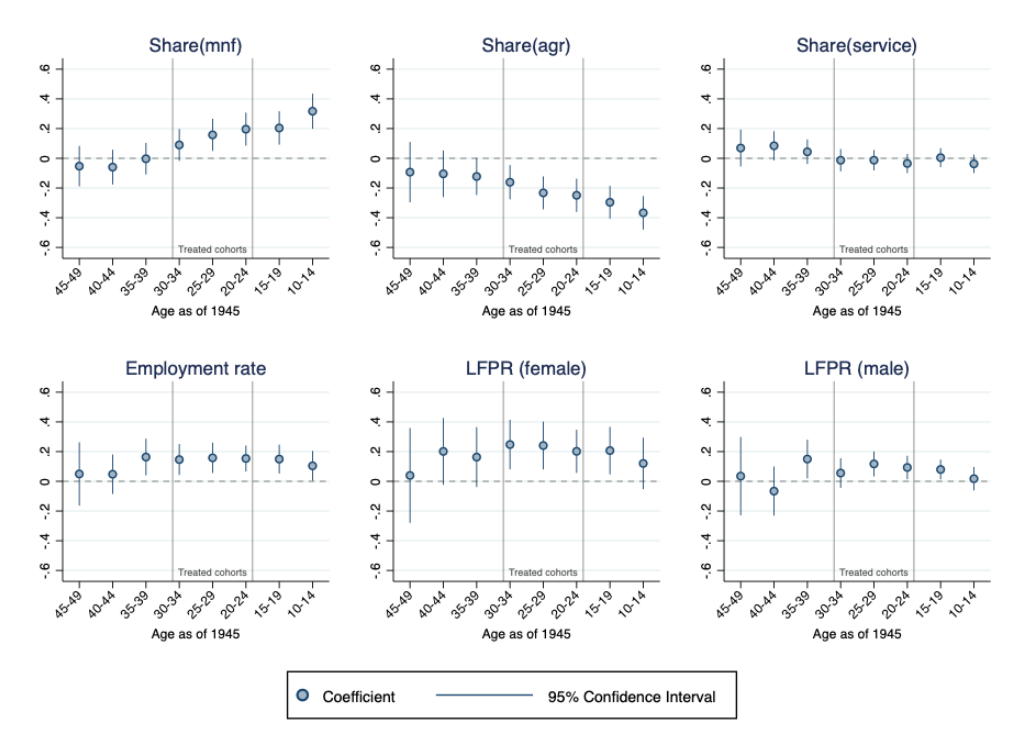


Table B7—: With pre-war outcome control and other cohorts' ΔRatio

	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 in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B8—: Results excluding outlier prefectures (top and bottom 3 manufacturing share)

	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, so as to check 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 in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.