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# Immigration Restrictions and the Wages of Low-Skilled Labor: Evidence from the 1920s*

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**Abstract**

This paper examines how the U.S. immigration restrictions of the 1920s affected the wages of low-skilled workers using newly digitized annual wage data from 1910 to 1930. Exploiting variation across local labor markets, we find that wages for low-skilled workers rose faster in areas more affected by the restrictions. These wage effects emerged early in the 1920s and persisted throughout the decade across manufacturing, construction, and agricultural sectors. Our findings help explain previously documented internal migration patterns and demonstrate how reduced immigration affected labor markets through both direct supply effects and subsequent adjustment mechanisms.

**JEL Codes:** J15, J31, K37, N32, N42, R23

**Keywords:** immigration policy, labor markets, wages, low-skilled labor

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# 1 Introduction

The era of mass migration into the United States, characterized by an "open borders" immigration policy, ended with the onset of World War I in Europe and the passage of restrictive immigration laws in 1921 and 1924. These laws, which imposed differential immigration quotas on different countries of origin, substantially reduced immigration from Europe to the United States and remained in effect for several decades. This paper explores the impact of this shift in immigration policy and the resulting reduction in immigration on the wages of low-skilled labor by analyzing various sources of wage data collected in the 1910-1929 period, data that we have hand-collected and digitized. Our identification strategy is based on the fact that the quotas of the 1920s restricted immigration from some countries far more than others and on the known tendency for immigrants to the U.S. to settle in areas in which many of their countrymen have previously settled (Altonji and Card, 1991; Card, 2001). Thus, labor markets which in the early 20th century had a high percentage of past immigrants from these more restricted countries would have been more "exposed" to disruptions caused by the reduction of immigration.

The question of how the immigration restrictions of the 1920s affected wages in the U.S. has been explored in several recent studies (Abramitzky et al., 2023; Price et al., 2020; Tabellini, 2020), some of which use an identification strategy similar to ours. Among these papers, Abramitzky et al. (2023) is the most closely related to our own. Using decadal census data from 1900 through 1930, they find that the foreign-born percentage of the population declined more in those areas more exposed to the quotas, while migration into those areas increased, both from other regions of the U.S. and countries not subject to the quotas. In rural areas more exposed to the quotas, farmers were more likely to buy tractors and switch to less labor-intensive crops. Abramitzky et al. (2023) interpreted these results as evidence that the quotas led wages to rise more quickly in more exposed areas. However, their analyses of wage data from the manufacturing censuses and of wages imputed to individual workers based on characteristics reported in decennial population census data showed no significant effect on wages.

Recent studies have highlighted the significance of time horizons and adjustment dynamics when evaluating the effects of immigration on the labor market (Dustmann et al., 2017; Monras, 2020). Monras (2020) illustrates that the impact of increased Mexican immigration on local labor market outcomes in the United States during the 1990s varied notably in the short term versus the long term. Specifically, the arrival of Mexican immigrants during the Peso crisis of 1995 had substantial short-term effects on wages in high-immigration areas compared to low-immigration areas. However, these effects decreased over time due to factors such as internal migration, local technologies, and the housing market. Building on this research, a potential explanation for the surprising result in Abramitzky et al. (2023) is that analyzing occupation-based wage data from decennial census years instead of annual frequency can only provide a partial understanding of wage dynamics due to missing out on important within-occupation wage changes and short-run responses to reduced immigration. For example, by 1929, the migration documented in their study may have largely closed the wage differentials opened by the differential impact across local labor markets of the war and the quotas.[1]

This paper helps reconcile the puzzle in Abramitzky et al. (2023) – *why would internal migration to areas more exposed to the quotas increase if wages did not increase in those places?* We do so by digitizing high-frequency (annual) wage data collected between 1910 and 1930. These data allow us to look at the short- and medium-term movements of wages during the intercensal years in many different industries and areas. They also report wages for narrowly defined occupations, including the lowest-paid "laborer" occupations in the various sectors. The ability to focus on the wages of the lowest-skilled workers is essential, as the war and quota laws disproportionately reduced the flow of low-skilled labor the U.S.[2] We find compelling evidence that during the 1920s, the wages of low-skilled laborers were higher in labor markets and industries that were more likely to have been exposed to immigration disruptions, thus explaining the migration patterns found by Abramitzky et al. (2023). More generally, our study provides a rare examination of wage dynamics that result from a significant reduction, as opposed to a significant increase, in immigration, as the two types of events may not have symmetric effects on labor markets.

Our most comprehensive source of wage data is a series of industry surveys conducted by the Bureau of Labor Statistics (BLS) between 1910 and 1929. Each survey reports wages, hours, and employment in a particular mining or manufacturing industry, with data reported for detailed occupations within the industry and in different cities or states. Many of the industries were surveyed repeatedly over the period. We also analyze the wages of construction laborers using information from annual surveys conducted by the BLS of union wage scales in construction in various cities between 1910 and 1929. Last, we examine farm

[1] Tabellini (2020), using an identification strategy like that of Abramitzky et al. (2023) and data from the population censuses, also reports no significant effect of the war and quota-related reductions in immigration on manufacturing wages. A number of earlier studies, including Xie (2017); Hatton and Williamson (1995); Goldin (1994), find a negative effect of immigrant flows on wages in the early 20th century.

[2] By contrast, the census of manufacturing wage variable used in previous studies is an average wage for production workers of all skill levels, while Abramitzky et al. (2023) imputed wage variable was based on three-digit occupation, age, and state of residence. The average value of this variable in an area would be almost unaffected by changes in the wages of low-skilled laborers in the region since such workers were almost all assigned to the same 3-digit occupational category.

laborers' wages using the government surveys of wages for farm labor conducted annually from 1910 to 1929 by the United States Department of Agriculture (USDA). We complement our findings by collecting annual wage data available post-1920 to examine whether it is consistent with our primary analysis. First, we use reports of the wages for construction laborers in over 50 cities based on surveys of local building contractors, conducted by the trade journal American Contractor from 1920 to 1928. Second, we use a BLS survey of wages paid in 1928 to workers hired to clean and repair streets in over 2,500 cities and towns.

As noted above, we follow Abramitzky et al. (2023) in arguing that since the quotas of the 1920s restricted immigration from some countries far more than others, labor markets with a high percentage of past immigrants from these more restricted countries would have been more affected by the reduction of immigration. We use measures similar to those created by Abramitzky et al. (2023) to classify local labor markets as more or less exposed to the legislated immigration quotas of the 1920s based on the historical country-of-origin composition of their immigrant population. Because some of our wage data pertain to specific industries at the state level, we have developed modified versions of these measures by defining "relevant labor markets" for particular industries based on the geographical distribution of the industry's employment within the state. Although these exposure measures are based on the design of the legislated immigration quotas of the 1920s, we present evidence that they serve as good proxies for the combined impact on local labor markets of the disruptions in immigration induced by World War I and subsequent changes in immigration policy in the 1920s.

Using these exposure measures, we estimate difference-in-differences models in which we define the treatment as the relevant labor market's exposure to immigration disruptions, conditional on the initial foreign-born share of that labor market and additional controls. Following recent developments in the continuous difference-in-differences literature, we follow Callaway et al. (2024) and assume the "strong parallel trends" assumption as the identifying assumption of our estimation model. That is, we require that the evolution of the foreign-born share and wages of relevant labor markets with lower exposure measures (lower proportions of immigrants from restricted regions) parallels the evolution of these outcomes in relevant labor markets with higher exposure, absent the disruptions in immigration. In practice, this identification assumption restricts treatment heterogeneity and thus justifies comparing labor markets with varying levels of exposure to immigration disruptions.

Our findings show that during the 1920s, the relative wages of low-skilled workers increased faster in labor markets and industries more exposed to immigration restrictions. This finding is consistent across sectors and regions, from manufacturing and construction industries concentrated in urban areas to agriculture, more concentrated in rural areas.

Moreover, our event study analysis suggests that relative wages in more exposed labor markets and industries increased early in the 1920s and remained elevated yet relatively stable throughout the decade, indicating that travel disruptions during World War I were another contributing factor in suppressing labor supply and putting upward pressure on wages. This is consistent with the previous findings in the literature that the 1920s saw internal migration toward areas most affected by the quotas. Our findings also show that wages of low-skilled workers remained elevated in high-immigration areas relative to low-immigration regions. In contrast, other studies in the literature showed that occupational-based wage measures were not different on average between labor markets. This finding is consistent with heterogeneous effects of immigration on workers by skill-type.

We note that our analysis has several caveats. One is that the wage variable is typically an average value for a group of workers in an area-industry cell. Thus, our finding that low-skilled wages rose in areas with more significant declines in foreign-born workers could be due to a composition effect driven by wage discrimination against immigrants. Section 6 of the paper discusses the potential seriousness of this problem. Also, it should be noted that the wage dynamics we document reflect the effects of both the reductions in the flows of low-skilled immigrant labor due to the quotas and any movements of low-skilled labor in response to emerging wage differentials, both in the form of internal migration and of increased immigration from countries for which the quotas did not bind. Therefore, our wage estimates should be interpreted as indications of wage gains in more exposed labor markets relative to less exposed labor markets. More generally, we argue that we are estimating the medium-term effect on wages of the end of mass immigration to the U.S., including any migration that occurred in response to initial short-run wage movements. As we would expect there to be migration responses to any change in immigration that significantly altered inter-area wage differentials, these estimates are more informative about other episodes involving changing immigration levels to which we may wish to generalize than estimates of what would have happened to wages in the 1920s had there been no migration responses to initial wage movements.

Our paper complements the large body of research documenting the wide-ranging social and demographic effects of the end of mass immigration to the U.S. This includes Greenwood and Ward (2015), Massey (2016), and Ward (2017), which examine how the quotas of the 1920s changed the skill selection and probability of return migration for European immigrants, while Collins (1997) and Xie (2017) have studied the relationship between the border closure and the advent of the Great Black Migration. More recent studies show that immigration quotas reduced scientific discovery and patentable ideas (Yoon and Doran, 2020; Moser et al., 2025) and had a small (but detectable) effect on dampening the spread of communicable diseases (Ager et al., 2024). Areas that experienced falling immigration after the border closure also became more receptive to redistribution (Tabellini, 2020).

This study is connected to the broader literature on the impact of immigration on wages. Previous studies have produced conflicting results. Borjas et al. (1997) and Borjas (2003) found evidence that immigration has a significant negative effect on the wages of native workers using national variation in immigration and its effects across education and experience groups. On the other hand, Altonji and Card (1991), Card (2001), and Ottaviano and Peri (2012) reported only minor effects using spatial variation in immigration across local labor markets. Monras (2020) reconciled these findings by demonstrating that increased immigration from Mexico to the U.S. in 1995 had significant short-term effects on wages that dissipated over time due to internal migration and relocation of workers. Our study adds to this literature by examining the impact of immigration restrictions on the wages of low-skilled workers, using cross-sectional variation in exposure to immigration based on immigration networks, along with plausibly exogenous reductions in immigration due to WWI and quota restrictions. We also use high-frequency and actual wage data from the period to document both the short-term and longer-term effects of immigration restrictions on the wages of low-skilled workers.

Numerous papers have utilized natural experiments to evaluate the impact of immigration on labor market outcomes. Some of the notable works in this area include those by Card (1990), Hunt (1992), Angrist and Kugler (2003), Cohen-Goldner and Paserman (2011), Glitz (2012), Borjas (2017), Borjas and Monras (2017), and Dustmann et al. (2017). Borjas (2017) particularly emphasizes the significance of analyzing the wage effects of immigration on workers who are the most similar to the immigrants themselves, and we specifically examine the wages of low-skilled laborers. Dustmann et al. (2017) and Cohen-Goldner and Paserman (2011) consider both local labor markets and internal migration in the adjustment process following an increase in immigration to Germany, providing additional insights into the short-term and longer-term wage adjustments and dynamics for our setting. It should be noted, however, that while almost all of the previous literature focuses on the wage effects of either a sudden increase in the number of immigrants or a longer-term increase in the immigration rate, our study is looking at the effect on wages of a significant longer-term reduction in the immigration rate due to restrictions on immigration, thus throwing light on an understudied aspect of the immigration-wage relationship.[3]

The rest of the paper proceeds as follows. Section 2 documents how the size and nature of immigration flows to the U.S. changed from 1910 through 1930. We also describe some of

[3] An exception is Clemens et al. (2018) which looks at the impact of the US agricultural sector of the reduction of Mexican guest workers caused by the end of the Bracero program in the 1960s.

the perceptions of and reactions to these changes by economists and other interested parties. Section 3 describes the the various datasets that we use to analyze the effect of immigration reductions on workers' wages, along with how we measure exposure to the border closure policy restrictions. Section 4 presents our empirical strategy and estimation framework, Section 5 presents the main findings, and Section 6 provides a composition shift channel analysis. Section 7 concludes.

## 2 Historical Background

During the five years 1910-1914, immigration to the U.S. averaged over 1 million per year, with the overwhelming majority of this being what was called at the time the "new immigration," that is, immigration originating from southern, central, and eastern Europe.[4] During the first full year of World War I, the number of immigrants to the U.S. dropped to 327,000, and over the 1915-1918 period, it averaged about 250,000 per year. Immigration numbers rebounded following the war but were still below half of pre-war levels, averaging about 450,000 per year in 1919-1921.

In May of 1921, the Emergency Immigration Act went into effect. The Act established quotas on immigration based on country of origin, with the number of immigrants allowed annually from a country limited to 3 percent of the number of immigrants from that country living in the United States as of the 1910 Census. The Immigration Act of 1924 lowered the rate for calculating a country's quota to 2 percent and changed the base for the quota calculation to the number of immigrants from that country living in the United States as of the 1890 Census, which disproportionately tightened the quotas for the southern and eastern European countries. Neither law placed quotas on immigration from the Western hemisphere nations.

The laws led to a drastic change in the composition of the immigration flows in terms of the nation of origin of the immigrants. This is shown in panel (a) of Figure 1, which follows Abramitzky et al. (2023) in dividing immigrants to the U.S. into three categories: those from the "high restriction" countries of Asia and of Southern, Eastern and Central Europe (countries for which the quotas introduced in the 1920s were generally binding); those from "low restriction" countries of Northern and Western Europe (for which quotas often did not bind); and those from western hemisphere nations, which were not restricted by the Acts.[5] As the figure shows, immigration during the pre-war period (1910-1914) was dominated

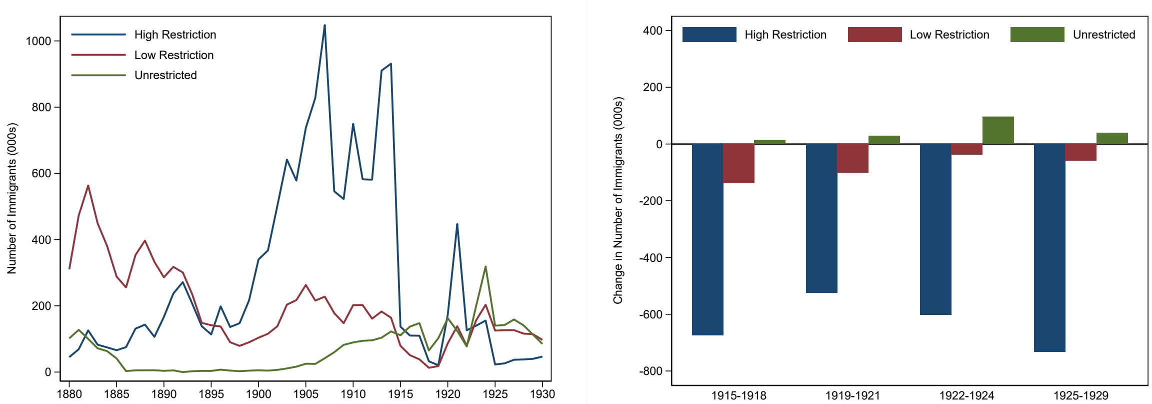
[4] This contrasted with the "old immigration," from northwestern Europe and the British Isles. The new immigration was argued by many to be the source of serious economic, social, and political problems for the U.S. (Leonard (2016), chapter 9 discusses these arguments).

[5] The regions and countries assigned to the three categories can be seen in Table A1.

by the new immigration. During the war (1915-1918), immigration from all countries fell considerably, but the reduction was largest from the countries that would later be subject to binding quotas. This was also true in the years between the war and the imposition of the quotas, despite the rebound of immigration from high-restriction countries to nearly a third of its pre-war level. The imposition of quotas in 1921 reduced the new immigration again, while the unrestricted immigration from the western hemisphere (mainly from Canada and Mexico) rose above its pre-war level, and immigration from the low-restriction countries came close to pre-war levels. Finally, immigration from the unrestricted countries remained above pre-war levels following the 1924 Act, while immigration from the high-restriction countries dropped below 50,000 per year.

Figure 1. Immigration Flows by Quota Restriction Category

(a) Annual Flows (b) Change in Mean Annual Flows




Source: Historical Statistics of the United States: Colonial Times to 1957, Series C-88-116. Series C 88-114. "Immigrants, by Country: 1820 to 1957".

Note: See Table A1 for a list of countries/regions and their classification. Change is defined relative to 1910-1914 period.

Panel (b) of Figure 1 shows the absolute change in the annual number of immigrants from the three types of countries – high restriction, low restriction, and unrestricted – from 1915 to 1929, relative to annual averages during the five pre-war years. The figure shows that it is not unreasonable to argue that the "treatment" we identify by our approach – a significant and permanent reduction in the number of immigrants from the restricted countries – commenced in 1915 and remained fairly steady throughout the 1920s. Compared to pre-war flows, the average annual "loss" of new immigrants from the restricted countries to the areas where they would have settled was over half a million per year in all the subperiods shown in the Figure. The changes in flows from the low-restriction and unrestricted countries were small by comparison. In a sense, the quotas of the 1920s simply wrote into law and thus perpetuated a new immigration regime created by the war.[6]

Economists quickly became aware of the impact of the war on immigration to the U.S. and discussed the effect of the decline in immigration on labor markets. Warne (1915) observed that the war had almost cut off immigration, especially from countries like Italy, which had been sending immigrants in large numbers just before the war. Emmet (1917) and Gormly (1918) described the positive impact of war-related reductions in immigration on the fortunes of labor unions in New York's garment industry. Charles Barnes, an official of New York State's Employment Bureau, spoke to the 1918 meeting of the AEA about employers' widespread complaints of "labor shortages." Most such shortages, he believed, resulted from the low wages and poor conditions the complaining employers offered. However, in the case of jobs requiring "laborers of strong physique," he opined that "...there would seem to be a good reason to believe that there is an actual shortage in this line. Immigration of Huns, Poles, and Slavs has practically ceased. Many Greek and Italian reservists returned to their countries soon after the outbreak of the war. We have depended largely upon these races for our laborers, and very few native-born Americans go into this field" (Barnes, 1918).

In the immediate post-war years, most commentators believed that the reduction of immigration would ultimately lead to higher wages. Boris Emmet, an economist and expert on labor issues in the garment trade, discussed in a series of articles an issue of relevance to this paper – the differing effect of the war-related reductions in immigration on different labor markets (See, e.g., Emmet (1917, 1918)).

During the years following the legislative imposition of quotas, there was a broad consensus among economists that limitations on immigration were indeed placing upward pressure on wages. (see, e.g., Hansen (1923, 1925), Berridge (1923), and Douglas (1926)).[7] The accumulating immigration statistics were making obvious another important change in immigration to the U.S. – a reduction in the proportion of low-skilled workers among the admitted immigrants (BLS, 1927). The disproportionate reduction in the supply of low-skilled workers was even more marked if one looked at net migration figures.[8]

Thus, it was widely believed by economists and others during the 1920s that the reductions in immigration that began in 1915 raised wages, particularly the wages of low-skilled

[6]The idea that a single new immigration regime began in 1915 with the war and was maintained by the quota legislation appears in the literature of the 1920s. Discussions of the reduction in the labor force due to immigration commonly refer to both the war and the legislation (e.g..(Douglas, 1926; Wolman, 1929; Jerome, 1934). Baker (1925) speaks of 1915-1920 as "six years without immigration"; Hansen (1923) speaks of the "scarcity of labor resulting from the restricted immigration of the last eight years.", and Slichter (1929) discusses the effects of the reduction of immigration "since 1915".

[7]See also, for example, Holmes (1924) and Soule et al. (1926). Slichter (1929) is a dissenting opinion

[8]Modern analyses of this phenomenon are provided by Greenwood and Ward (2015), Massey (2016), and Ward (2017).

workers. We present evidence supporting this belief by analyzing wage data collected during 1910-1929 for four types of low-skilled workers – laborers in mining and manufacturing industries, construction laborers, municipal workers hired to clean and repair streets, and farm laborers. For each type of labor, we rely on different sources of wage data, necessitating different empirical procedures for each of the four analyses. However, in all four, we rely on the logic underlying the identification strategy employed in Abramitzky et al. (2023). It is based on the observation that in the decades before the war, the new immigrants from the countries of southern, eastern, and central Europe tended to cluster in certain regions and communities. The assumption is that labor markets drawing their workers from such areas before 1915 were the labor markets that were more exposed to the immigration disruptions, that is, more likely to be affected by negative labor supply shocks when overall immigration was reduced by the war and then the quotas. An empirical implication of this logic is that if the reductions in immigration led to higher wages, wages for low-skilled labor would have been relatively higher in labor markets with greater exposure in the years following the war.

# 3 Wages, Labor Markets, and Exposure Measures

## 3.1 Manufacturing and Mining Industries

Early in the 20th century, the Bureau of Labor Statistics (BLS) began conducting periodic surveys of wages, hours, and employment in various industries, examining payrolls from samples of employers in an industry to create standardized measures of employment, hours, and wages for a variety of narrowly defined occupations in the industry. Average values of these variables would be reported for various regions, states, or cities.[9] The detailed descriptions provided by the BLS of data collection procedures and the defined variables support the conclusion that these data are accurate even by modern standards. We draw on surveys conducted from 1910 through 1929.

For most industries, surveys report data for a low-wage occupation with the title "laborer"; for the remaining industries, we chose the average wage for a large occupation at or near the bottom of the industry wage scale as the wage for low-skilled labor in that industry. We construct a sample that includes wages for 21 low-skilled occupations in 19 industries.[10] The sample consists of more than 1200 wage observations, each corresponding to a specific combination of industry-jurisdiction-year, thus denoted $w_{ijt}$, where the jurisdiction can be a

[9]For example, there were frequent surveys of sawmills in over 20 states – between 1910 and 1930, the BLS surveyed the industry in 1910, 1911, 1912, 1913, 1915, 1919, 1921, 1923, 1925, and 1928.

[10]Each industry in the data has one occupation except the coal industry, for which we include the wages of three low-skilled occupations: pick miners, laborers working inside the mine, and laborers working outside of the mine.

state or a city, depending on the industry.[11]

We classify labor markets as more or less exposed to policy-induced disruptions in immigration based on the historical country of origin composition of their immigrant population. To do this, we first define the "Relevant Labor Market" (denoted RLM in figures and tables) underlying each observation $w_{ijt}$ as the area that supplied low-skilled labor to the establishments surveyed by the BLS in industry $i$ and jurisdiction $j$ in time $t$. In the teens and twenties, low-skilled workers were more likely to get to work on foot or by trolley if they lived in a larger city, making the county where an establishment is located the best empirically feasible approximation to the relevant labor market for that establishment. For most of the wage observations, however, jurisdiction $j$ is a state, which creates a problem because we do not know the locations within a state of the establishments surveyed by the BLS.

To address that issue, when wages are reported at the state level, we use census data to estimate the probability that a particular county was part of the relevant labor market of one of the BLS-surveyed establishments contributing data underlying the mean wage observation in an industry. The early 20th-century censuses recorded the industry of each working person surveyed. A county in state $j$ in which no people were working in industry $i$ was clearly not a part of the relevant labor market underlying the wage observation for industry $i$ in that state. In contrast, a county in which fifty percent of the state's total employment in industry $i$ resided was likely a part of the relevant labor market for one or more of the surveyed establishments in that state. Based on that logic, when a wage observation is reported at the state level, we include counties in state $j$ as part of the relevant labor market with a probability weight based on the percentage of the state's total employment in industry $i$ found in that county. We use the full count census of 1910 to construct these weights, which we denote $\alpha_{ijc}$.[12]

Next, we create county-specific exposure measures. For this, we use a set of "quota intensity" measures created by Abramitzky et al. (2023) of the extent to which the legislated quotas of 1921 and 1924 restricted immigration to the U.S. from each of 18 sending regions during the 1920s. Each measure is a ratio defined as the difference between an estimate of the unrestricted flow from a region in the absence of quotas and the quota slots for that region, normalized by the unrestricted flow. The unrestricted flow from a region is estimated using a model that predicts immigration in the 1920s based on historical immigration flows for several pre-1915 decades. This ratio will be zero if the number of allocated slots for a region is greater than or equal to the estimate of the unrestricted flow from that region. It will be one if the

[11] Table A2 lists and describes the industries and associated occupations in the sample.

[12] Column 1 of Table A2 shows the 3-digit industry codes from the 1950 coding system used to create the weights for the industries in our sample for which the data is reported by at the state level.

quota is set equal to zero. The average ratio for the nine "high restriction" countries/regions is 0.925, the average ratio for the nine "low restriction" countries/regions is 0.07, and the average ratio is zero by definition for the four quota-exempted countries/regions.[13]

The quota exposure measure for a county then weights the foreign-born proportion of the county's male working-age (16-65) labor force from each sending region by the quota intensity ratio for that region:

$$q_c = \sum_r \frac{FB_{rc1910}}{LF_{c1910}} \times QI_r \tag{1}$$

where $q_c$ is the quota exposure measure for county $c$, $FB_{rc1910}$ is the number of foreign-born males from region $r$ in the working-age labor force in county $c$ in the 1910 census, $LF_{c1910}$ is the working-age male labor force in the county in the 1910 census, and $QI_r$ is the quota intensity ratio for region $r$. The quota exposure measure for county $c$ would be larger the larger the share of its 1910 populations born in regions with higher quota intensity levels $QI$.[14]

For wage observations reported at the state level, our quota exposure measure for an industry's relevant labor market is the weighted average of the $q_c$ values for the counties in that state, using the $\alpha_{ijc}$ industry employment weights:[15]

$$QE_{ij} = \sum_{c \in j} \alpha_{ijc} \times q_c \tag{2}$$

Column 1 of Table 1 shows the mean exposure measure for the overall sample and each of the 21 industries in the analysis. All the industries in our dataset were exposed to immigration restrictions to some extent, with the mean exposure at 11 percent. However, there is a considerable variation in mean exposure by industry, ranging from a high of 27 percent in iron mining to a low of 6 percent for sawmill workers. Table 1 also provides information from the BLS data on the mean (nominal) hourly wages of low-skilled workers by industry before and after the border closure policies of the 1920s. The mean hourly wage for pre-1920 observations was 26 cents per hour. The mean hourly wage was 46 cents for the post-1920 observations, implying a 67 percent increase in real wages. Table 1 also shows

[13]See Table A1 and Abramitzky et al. (2023) for details on the model used to estimate unrestricted flows.

[14]Abramitzky et al. (2023) also constructed measures of how World War I restricted immigration flows from various sending regions, analogous to $QI_r$ described above (see Table A1). We use them to calculate a "war exposure" measure $w_c$ for each county in a similar manner, which is highly correlated to $q_c$ (correlation coefficient of 0.87), and show that our analysis is robust to controlling for WWI restrictions in Section 5.4.3.

[15]If, as in the meat packing and men's clothing industries, for example, the jurisdiction is a city, the quota exposure measure for the observation is simply the quota exposure measure for the county in which the city is located, that is, $QE_{ij} = q_c$ as defined in Equation 1.

Table 1. Border Closure Exposure Measure and Mean Hourly Wages by Industry

| | Mean Exposure (1) | Pre-1920 (2) | Post-1920 (3) | % Change (4) |
|---|---|---|---|---|
| Overall | 0.11 | $0.26 | $0.43 | 67% |
| Above Median Exposure | 0.17 | $0.31 | $0.53 | 71% |
| Below Median Exposure | 0.04 | $0.21 | $0.33 | 59% |
| **A. Manufacturing and Mining:** | | | | |
| Auto | 0.16 | $0.43 | $0.51 | 19% |
| Boxes | 0.18 | $0.34 | $0.45 | 31% |
| Cement | 0.10 | | $0.40 | |
| Coal In | 0.11 | $0.54 | $0.65 | 20% |
| Coal Miner | 0.11 | $0.77 | $0.71 | -7% |
| Coal Out | 0.11 | $0.47 | $0.57 | 22% |
| Cotton | 0.09 | $0.14 | $0.35 | 139% |
| Foundry | 0.17 | $0.43 | $0.43 | 0% |
| Furniture | 0.11 | $0.23 | $0.37 | 62% |
| Hoisery | 0.09 | $0.20 | $0.38 | 91% |
| Iron Mine | 0.27 | | $0.45 | |
| Machinery Shop | 0.16 | $0.423 | $0.419 | -1% |
| Meat | 0.24 | | $0.45 | |
| Men Clothing | 0.19 | $0.31 | $0.85 | 177% |
| Metal Out | 0.14 | | $0.51 | |
| Mill Work | 0.12 | $0.20 | | |
| Quarry | 0.12 | | $0.40 | |
| Railcar | 0.10 | $0.17 | | |
| Sawmill | 0.06 | $0.19 | $0.30 | 59% |
| Shoes | 0.13 | $0.30 | $0.47 | 57% |
| Wool | 0.16 | $0.26 | $0.46 | 79% |
| Observations | | | 2,975 | |
| Number of RLMs | | | 128 | |
| Number of Occupations | | | 23 | |
| Number of RLM-Occupation Pairs | | | 402 | |

Notes: Column 1 shows the mean policy exposure measure level by industry. Columns 2 and 3 show the pre- and post-1920 mean hourly earnings for the different industries in the sample. Column 4 shows the percent change in mean hourly earnings between the periods.

that industries with higher values of the exposure measure tended to experience higher wage growth. Industries with above median exposure experienced a 71 percent increase in wages relative to pre-1920 wages, compared to a 59 percent increase for industries with below median exposure.

### 3.2 Construction

Throughout the 1920s, the trade journal *American Contractor* surveyed construction contractors in many cities, asking for the wages paid in their city to workers in various construction trades, including "laborers" who generally have the lowest reported wages. We use data from over 50 cities and reports for Oct. 1920, Oct. 1921, May 1922, Oct. 1922, June 1925, Oct. 1925, May 1926, Sept. 1926, April 1927, July 1927, May 1928, and Oct. 1928. When there is more than one observation for a city in a given calendar year, we use the highest wage for that year so that each observation corresponds to a unique city-year pair. The exposure measure we use for this sample is $q_c$ for the county in which the city of the wage report is located (see Equation 1).

*American Contractor* did not begin publishing wage surveys before 1920; thus, the surveys provide no information on geographic wage differentials for construction laborers in the pre-war period. For that reason, we use a second data source on construction wages. From 1910 to 1929, the BLS did annual surveys of labor union leaders in various cities, asking for "union scale of wages and hours of labor." Construction trade unions were among those surveyed, and union officials in many cities reported a union wage for "building laborers." These data are less reliable than the BLS industry surveys as indicators of actual wages paid for several reasons. First, the scales are established minimums, and it was not unusual for unionized workers to be paid more than this minimum. Second, not all construction workers in a surveyed city would be paid according to the reported scale, although the BLS only reported scales for cities in which over 50% of the workers in an occupation were receiving at least the wage stated in the union scale (United States Bureau of Labor Statistics, 1913).

Cities drop in and out of the survey over the years, and it seems likely that one reason for this is changes in the extent of unionization in various cities. Table A4 lists the cities for which our construction worker data are available by source and coverage. Overall, data on wages in the construction industry is available for 77 different cities, with 58 cities represented in the *American Contractor* data from 1920 to 1928, 45 cities in the union scale data from 1910 to 1929, and 35 cities included in both surveys.

The local labor markets in our construction wage data had varying exposure to immigration disruptions. Column 1 of Table 2 shows that the overall mean exposure measure is 11 percent for the cities in the *American Contractor* sample. The average exposure for cities with a higher-than-median exposure measure is 19 percent, while the average for cities below the median is 4 percent. The corresponding measures for the union scale sample are 13 percent, 20 percent, and 6 percent.

Table 2 also provides information on mean hourly wages before and after the border closure policies of the 1920s. Panel A shows wage data from the *American Contractor* sample. The mean wage was 61 cents in 1920 (column 2) and 49 cents in 1928 (column 3), implying a 21 percent decrease in hourly wages. The data also reveals the relationship between the exposure measure and wage growth. When splitting the sample into cities below and above median exposure, we find that cities with above-median exposure experienced a 16 percent drop in wages in this period, compared to a drop of approximately 26 percent for cities with below-median exposure. We find a similar pattern in Panel B, which shows summary statistics from the union wage scale data before and after 1920 (columns 2-4 in Panel B). We find that cities with above-median exposure to the immigration disruptions experienced more considerable union wage scale growth, consistent with increased bargaining power for labor unions due to reduced labor supply.

Table 2. Construction Workers Wage Data - Summary Statistics

| | **Mean Exposure (1)** | **(Pre-)1920 (2)** | **Post-1920 (3)** | **% Change (4)** |
|---|---|---|---|---|
| **A. The American Contractor Surveys (1920-1928)** | | | | |
| Overall Sample (61 Cities) | 0.11 | \$0.61 | \$0.49 | -21% |
| Below Median Exposure (31 cities) | 0.04 | \$0.54 | \$0.40 | -26% |
| Above Median Exposure (30 cities) | 0.19 | \$0.67 | \$0.56 | -16% |
| **B. Construction Union Survey Data (1910-1929)** | | | | |
| Overall Sample (45 Cities) | 0.13 | \$0.40 | \$0.65 | 39% |
| Below Median Exposure (23 cities) | 0.06 | \$0.39 | \$0.61 | 35% |
| Above Median Exposure (22 cities) | 0.20 | \$0.40 | \$0.70 | 43% |

Notes: In Panel A, Columns 1-4 show the mean policy exposure, the 1920 and post-1920 mean wages for construction laborers, and the percent change in wages between 1920 and post-1920 wages for the cities in the American Contractor sample. In Panel B, columns 1-4 show the mean policy exposure, the pre- and post-1920 construction union wage scale, and the percent change in the union wage scale before and after 1920.

### 3.3 Agriculture

Beginning early in the 20th century, the United States Department of Agriculture collected statistics on the average wages of agricultural laborers by state (United States Bureau of Agricultural Economics, 1943). We use data on average daily and monthly wages for each of the 48 states from 1910 to 1929, and we proxy for hourly wages by dividing daily wages by 10 hours.

To define the state's quota exposure variable, we use a population that closely approximates the labor pool from which farmers in the state hire workers. So, rather than defining the variable over the state's entire labor force, we define it based on those in the state who were classified as "farm laborers" (1950 Census occupation code 820) in the 1910 full count census. Thus, the quota exposure variable for state $s$ is defined as

$$QE_s = \sum_c \frac{FBFW_{cs1910}}{FW_{s1910}} \times q_c \tag{3}$$

where $FBFW_{cs1910}$ is the number of farm workers from county $c$ working in state $s$ in 1910, $FW_{s1910}$ is the total number of farm workers in state $s$ in 1910, and $q_c$ is the exposure measure for county $c$ (see Equation 1). The rationale for defining the quota exposure variable using the population of farm workers is that the ethnic composition of farm workers in a state can be quite different from that of the state's agriculture labor force. A possible concern with this approach is that some low-skilled workers doing something other than farm labor could be part of the relevant labor market for farm workers, even in the short run. Based on this definition, the mean exposure in the agriculture sample is 7%, and the mean wage of farm laborers was \$.19 and \$.23 in pre- and post-1920, respectively, a 22% increase.

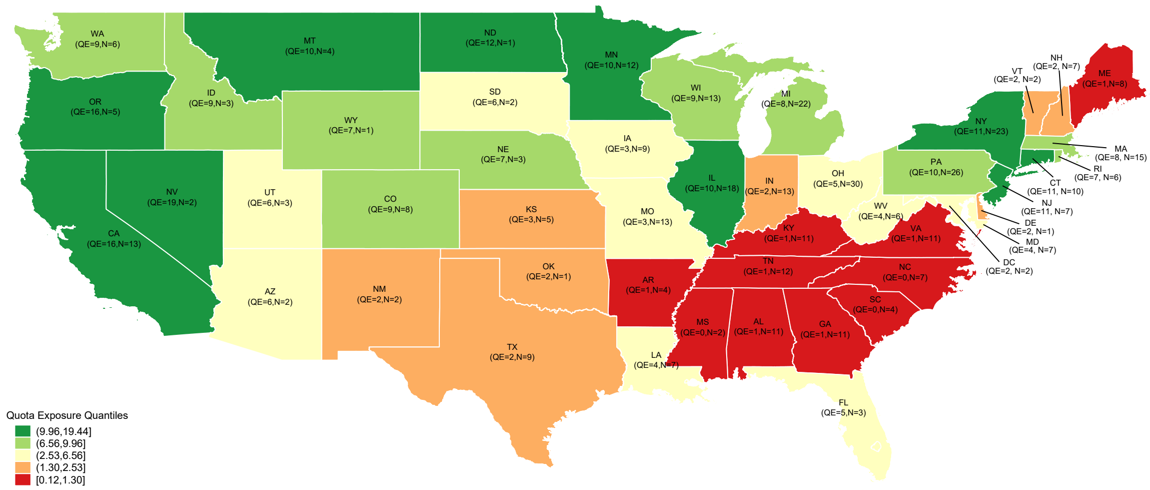
### 3.4 Pooled Sample

We pool RLM-industry pairs from the BLS Manufacturing & Mining, American Contractor Survey, BLS Survey of Construction Unions, and USDA survey on farm labor wages into one dataset to generalize our analysis. Our pooled sample has 403 RLM-industry pairs combined with 249, 61, 45, and 48 for the manufacturing & mining, American Contractor, construction union, and farm labor samples, respectively. Throughout the analysis, we control for dataset fixed effects and time trends to account for any differences in the samples and data collection procedures. We also present results for each dataset separately.

Table A3 closely examines our RLM by industry quota exposure measure. It shows the mean exposure measure, the 1910 foreign-born share in population, and the 1910 composition of the foreign-born population by country of origin. Column 1 shows these statistics for the full sample of 403 RLM by industry pairs. The mean exposure is 0.109, and the mean 1910 foreign-born share is 0.17. Columns 2-5 of Table A3 show the same information but by quartiles of the quota exposure measure. It is noticeable that higher exposed RLM by industry pairs have increased share of foreign-born from high-restricted countries, in particular from Central Europe, Italy, and Russia, while having a lower share of foreign-born from low restriction countries.

Figure 2 provides information regarding the geographic distribution of observations in our dataset. It summarizes the RLM-industry pairs used in our analysis at the state level by indicating the mean exposure measure for RLM-industry pairs and the number of RLM-industry pairs in our analysis for each state. In particular, each state is assigned a different color based on its exposure measure quartile, with darker green states having a higher exposure measure while darker red states have a lower exposure measure.

Figure 2. Geographic Distribution of Labor Market-Industry Pairs and Exposure




Notes: The map plots the quartiles of mean exposure across the relevant labor market (RLM) by industry pairs within a state. Each state has a label indicating its mean RLM-industry exposure measure $QE$ and the number of RLM-industry pairs in the state $N$. The datasets used to prepare the map include the pooled sample that includes the BLS manufacturing and mining, American Contractor, BLS construction union, and USDA farm labor wages.

The observations are distributed very similarly to the overall distribution of the population in 1910, with the Midwest region having the highest number of RLM-industry pairs with 129 (36%), followed by the Northeast and the South with 95 and 93 (26%), respectively, and the West with 38 overall (12%). The Northeast and West have higher quota exposure, while the South and parts of the Midwest have lower exposure, consistent with the geographic distribution of the foreign-born population in the United States in the early 20th Century.

We supplement our data using the 1900-1940 Population Censuses for details on employment, foreign-born share, and other demographics at the RLM-industry level (Ruggles et al., 2024). We calculate mean hourly wages for low-skilled workers in RLM-industry pairs in 1940, the first year wage data was collected, to assess the long-term wage impacts of border closure policy. Additionally, we gather socioeconomic data for that period from the Historical, Demographic, Economic, and Social Data of the United States through ICPSR (Haines et al., 2010).

### 3.5 Municipal Street Laborers

In 1928, the BLS surveyed officials of 2600 municipalities regarding the wage rates of the low-skilled workers hired to clean and repair streets. The survey specifically asked for the wage paid to entry-level workers with no skill or experience. (BLS, 1929). In most cases, a municipality would report one wage rate. When more than one wage rate is reported, we use the lowest wage, and when a range is reported, we use the bottom of the range. Our exposure variable for these wage data was $q_c$ for the county in which the city is located. Towns and villages that overlapped or were very close to a county line in 1930 were excluded from the sample, leaving 2,361 cities and towns with usable wage data.

Table 3. Municipal Street Laborers Workers Wage Data - Summary Statistics

| | **Mean Exposure** (1) | **Wage (1928)** (2) | **Cities** (3) | **Mean Size** (4) | **Largest City** (5) |
|---|---|---|---|---|---|
| Full Sample | 0.10 | $0.43 | 2,273 | 25,254 | New York, NY (6.1m) |
| **Exposure Measure Quartile:** | | | | | |
| Quartile 1 (Bottom) | 0.01 | $0.31 | 569 | 9,248 | Decatur, IL (157k) |
| Quartile 2 | 0.04 | $0.40 | 573 | 19,721 | Davenport, IA (552k) |
| Quartile 3 | 0.12 | $0.49 | 567 | 24,019 | St. Louis, MO (848k) |
| Quartile 4 (Top) | 0.25 | $0.52 | 564 | 48,264 | New York, NY (6.1m) |

Notes: Columns 1-5 show the mean border closure exposure measure, mean hourly municipal laborer wage in 1928, number of cities/towns, mean city population, and the name of the largest city for the full sample and by its exposure quartile of cities and towns, respectively.

Table 3 provides information on the mean quota exposure measure and mean hourly street laborer wage in 1928 for the sample of cities and towns, in column 1 for the full sample, and in columns 2-5 for the policy exposure quartiles. The average exposure measure for the full sample is 10 percent, and the mean hourly wage is 43 cents. The scope of municipalities in the sample is large: the average population is 25,254 but ranges from 2,500 to 6.1 million (New York City). Columns 2-5 show that cities with a higher exposure measure also had higher wages for street laborers. For example, a city in the bottom quartile of exposure, such as Decatur, IL, had a mean quota exposure measure of 1 percent and an average hourly wage of 31 cents compared to a city in the top quartile (such as New York City), that had mean quota exposure of 25 percent and an average hourly wage of 52 cents.

## 4 Empirical Strategy

Our empirical analysis begins by confirming that relevant labor markets with higher exposure to immigration restrictions lost more immigrant workers during the 1920s. Following that, we show evidence that the drop in immigration led to a relatively larger increase in short-term wages for low-skilled occupations in labor markets and industries that historically relied more heavily on immigrant labor from regions that the changes in immigration policy would later restrict. To do so, we estimate various versions of the following equations:

$$FB_{ijt} = \delta_{ij} + \gamma_t + \beta(QE_{ij} \times Post_t) + \Theta(X_{ij1910} \times t) + \varepsilon_{ijt} \quad (4)$$

$$ln(w_{ijt}) = \delta_{ij} + \gamma_t + \beta(QE_{ij} \times Post_t) + \pi_{FB}(FB_{ij1910} \times t) + \Theta(X_{ij1910} \times t) + \varepsilon_{ijt} \quad (5)$$

where $j$ indexes jurisdiction, $FB_{ijt}$ is the foreign-born share of the working-age male workforce in the relevant labor market of industry $i$ in jurisdiction $j$ in year $t$, $ln(w_{ijt})$ is the corresponding log mean hourly wage rate for low-skilled workers in the relevant labor market of industry $i$ in jurisdiction $j$ in year $t$.[16] The primary variable of interest is the interaction between the quota exposure measure $QE_{ij}$ and the post-1920 indicator ($Post_t$) representing the period when immigration restrictions went into effect. The coefficient of interest $\beta$ is identified by comparing relevant labor markets with different shares of workers from restricted regions before and after that point. General correlations between wages and ethnic composition of the $ij$ labor force unrelated to the changes in immigration regime are absorbed by the labor market by industry fixed effects $\delta_{ij}$, while general trends in low-skilled wages are captured by the time fixed effects $\gamma_t$. In addition, we also include a vector $X$ which includes industry- and dataset-specific indicators, in addition to 1910 labor market by industry covariates selected using a Lasso estimation procedure (see below), and interact them with a linear time trend.

Along with our main estimation Equations (4) and (5), we also perform an event-study type analysis by estimating the following specifications:

$$ln(w_{ijt}) = \delta_{ij} + \gamma_t + \sum_k \beta_k(QE_{ij} \times I_k) + \pi_{FB}(FB_{ij1910} \times t) + \Theta(X_{ij1910} \times t) + \varepsilon_{ijt} \quad (6)$$

where $I_k$ are indicators that receive the value 1 if the year of the observation is $k$, and zero otherwise.[17]

Relevant labor markets can be more exposed to the border closure policy because they have a higher foreign-born share of workers or a larger share of their foreign-born population from the restricted regions. In our specification, we interact the initial (1910) foreign-born share of the relevant labor market's population with a linear time trend to control for differential trends by initial foreign-born share, thereby identifying the effect of quota exposure from differences in the composition of the immigrant population.[18]

Given that our treatment variable is continuous, we follow Callaway et al. (2024) and assume the "strong parallel trends" as our identifying assumption instead of the standard parallel trends assumption. That is, we require that conditional on controls for industry and initial foreign-born share, the evolution of the foreign-born share and wages of RLM-industry

[16]When the outcome variable is the foreign-born share, we follow Abramitzky et al. (2023) and do not control for the initial foreign-born share in the relevant labor market.

[17]The years in our analysis vary by industry and region and range from 1910 to 1930.

[18]Areas with different foreign-born shares in 1910 might have different wage trends if, for example, immigrants were more drawn to areas with more robust economies.

pairs with lower exposure measures (lower proportions of immigrants from restricted regions) parallels the evolution of these outcomes in RLM-industry pairs with higher exposure, absent the restrictions on immigration. This identification assumption restricts treatment heterogeneity and thus justifies comparing RLM-industry pairs with varying levels of exposure to immigration disruptions.[19] We provide two pieces of evidence to support this assumption: (1) a Lasso procedure to search for other correlates of the quota exposure measure and (2) an event study analysis that tests for the existence of pre-trends in the evolution of wages by interacting the quota exposure measure with year dummies.

To determine which RLM-industry level controls to include in our estimation, Table A5 considers the relationship between RLM-industry exposure and a series of economic and demographic controls from the 1910 census for the different datasets. The variables selected by a Lasso procedure for the pooled sample are the 1910 foreign-born population share, the log of the total population, and the share of workers in agriculture. Throughout the analysis, we control for these covariates by interacting them with a linear time trend.

To test the assumption that log mean hourly wages followed a similar pre-1920 trend in RLM-industry pairs with different levels of exposure to immigration restrictions, we restrict our sample to the years 1910-1920.[20] We first regress log mean hourly wages on the exposure measure, a linear trend, and the interaction between the trend and the exposure measure (Table A6, column 1). We then estimate nonlinear pre-policy trends by replacing the linear trend with dummy variables for 1911 to 1920 (Table A6, column 2). In neither case do we find significant evidence of a trend in real wages related to the exposure measure prior to 1921. We then repeat these tests for each dataset separately.[21]

## 5 Results

### 5.1 Foreign-Born Share

Table 4 confirms that RLM-Industry pairs with higher quota exposure measures experienced more significant declines in their foreign-born employment share following the imposition of

[19] The "standard" parallel trends assumption would require the evolution of the foreign-born share and wages of RLM-industry pairs with different proportions of immigrants from restricted regions to evolve similarly regardless of their exposure level.

[20] There is some ambiguity about when the "post" period should begin. As discussed above, the war significantly reduced immigration starting in 1915. We follow the existing literature and choose 1920 as the beginning of the "post" period. Our event study specifications check our decision regarding the beginning of the post-period. Moreover, we also test directly for the impact of the war on wages in Section 5.4.3.

[21] We use year-pair dummies rather than individual year dummies to increase the precision of the estimates since our sample is an unbalanced panel with different industries appearing in different years. Combining consecutive years alleviates this measurement issue to some extent. Estimates with one-year dummies are less precise yet deliver qualitatively similar results and are available upon request.

immigration restrictions in the 1920s. Column 1 shows our estimates for the pooled sample. We estimate a statistically significant coefficient of -0.260 on the interaction between RLM-industry exposure and a post-1920 indicator, implying that the average RLM-industry pair in our sample (exposure measure of 0.11) experienced a 2.9 percentage points drop in foreign-born employment share. We estimate another specification where the outcome variable is the employment share of workers born in quota-restricted countries (see Table A1). The DID coefficient for the pooled sample is -0.312, slightly larger than the DID coefficient for the overall foreign-born share, as immigrants from these countries were more likely to be impacted by the border closure restrictions.

Table 4. Exposure to Border Closure Policy and Foreign-Born Employment Shares

| **Sample:** | **Pooled** | **Manufacturing & Mining** | **American Contractor** | **Construction Unions** | **Agriculture** |
|---|---|---|---|---|---|
| | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** |
| **A. All Foreign-Born:** | | | | | |
| Quota Exposure X Post 1920 | -0.260*** | -0.255*** | -0.150** | -0.146** | -0.112 |
| | (0.0387) | (0.0526) | (0.0627) | (0.0614) | (0.150) |
| Pre-1920 Dependent Mean | 0.280 | 0.320 | 0.208 | 0.292 | 0.155 |
| **B. Quota Restricted Countries:** | | | | | |
| Quota Exposure X Post 1920 | -0.312*** | -0.365*** | -0.144** | -0.120** | -0.0904 |
| | (0.0380) | (0.0505) | (0.0579) | (0.0503) | (0.103) |
| Pre-1920 Dependent Mean | 0.243 | 0.277 | 0.186 | 0.255 | 0.129 |
| Quota Exposure - Mean | 0.11 | 0.11 | 0.11 | 0.13 | 0.07 |
| Quota Exposure - SD | 0.08 | 0.08 | 0.10 | 0.09 | 0.06 |
| RLM-Industry Pairs | 403 | 249 | 61 | 45 | 48 |
| Observations | 1,209 | 747 | 183 | 135 | 144 |

Notes: The table shows estimates of the interaction coefficient between the quota exposure measure $QE$ and the post-policy change indicator. Column 1 includes results for the pooled sample, while columns 2-5 provide results by data source. The Post variable indicates the period after 1920, marking the start of immigration restrictions. Panel A uses the overall foreign-born employment share in the Relevant Labor Market (RLM) by industry pair as the outcome variable, while Panel B focuses on workers from quota-restricted countries. All specifications incorporate year and RLM-industry fixed effects, dataset- and industry-specific linear time trends, RLM-industry level controls listed in Table A5 with a linear time trend interaction. Each RLM-Industry pair has data from the 1910, 1920, and 1930 censuses, with robust standard errors clustered at the RLM by industry level. Significance levels are indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Columns 2-5 of Table 4 replicate the analysis for the four different datasets used to construct the pooled sample. All of the estimates are negative across the the different datasets, ranging from -.255 in the manufacturing and mining sample to around -.15 in the construction samples and to an insignificant -.112 in the agriculture sample. The estimated effects are more significant for the manufacturing & and mining samples compared to other samples because the sample size is larger, these industries tend to have the highest share of foreign-born workers in our analysis, and their size is relatively smaller compared to the construction and agriculture sectors. For agriculture, the effect is negative but insignificant, potentially due to lower relative exposure of the sector to immigration.

### 5.2 Wages

Table 5 documents that RLM-industry pairs that were more exposed to immigration restrictions experienced more significant increases in hourly wages of low-skilled workers in the 1920s, consistent with a scenario in which the emergence of labor shortages in those labor markets led to wage increases in the short term. Our preferred specification in column 1 of Table 5 implies that the average RLM-industry pair in our sample experienced a 5.9 ($= 0.586 \times 0.1$ where the outcome variable is mean log hourly wages) percent increase in mean hourly wages post-1920 relative to a pre-policy mean hourly wage of 26 cents. In addition, each additional standard deviation increase in exposure (0.08) implied an additional 4.7 percent increase in mean hourly wages.

Table 5. Exposure to Border Closure Policy and the Wages of Low-Skilled Workers

| **Sample:** | **Pooled** (1) | **Manufacturing & Mining** (2) | **American Contractor** (3) | **Construction Union** (4) | **Agriculture** (5) |
|---|---|---|---|---|---|
| **A. Log Mean Hourly Wage:** | | | | | |
| Quota Exposure X Post 1920 | 0.586*** | 0.577*** | 0.574* | 1.019*** | 0.288 |
| | (0.150) | (0.216) | (0.309) | (0.317) | (0.241) |
| Pre-1920 Mean Hourly Wage ($) | 0.258 | 0.245 | 0.614 | 0.397 | 0.186 |
| Quota Exposure - Mean | 0.10 | 0.11 | 0.13 | 0.15 | 0.07 |
| Quota Exposure - SD | 0.08 | 0.08 | 0.10 | 0.10 | 0.06 |
| RLM-Industry Pairs | 402 | 248 | 61 | 45 | 48 |
| Observations | 2,975 | 1,234 | 294 | 487 | 960 |

Notes: The table presents estimates of the interaction coefficient between quota exposure measure $QE$ and the post-policy change indicator. The outcome variable is the RLM-industry log mean hourly wage for low-skilled workers. Column 1 presents results for the pooled sample, and columns 2-5 present results for each data source separately. The Post variable is an indicator for post-1920, the last year before immigration restrictions started. All specifications include year and RLM-industry fixed effects, in addition to dataset- and industry-specific linear time trends and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each RLM-industry pair has a varying number of observations from 1910-1930. Robust standard errors, clustered at the RLM by industry, level in parenthesis.. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Figure 3 shows the corresponding event study estimates of the effect of exposure to immigration restrictions on log mean hourly wages for the pooled sample. The pre-1920 estimates confirm the results of Table A6 that RLM-industry pairs with different exposure to immigration disruptions did not experience different log mean hourly wage trends in the pre-1920 years. The positive coefficients in the post-1920 years are consistent with the hypothesis that hourly wages for low-skilled laborers grew faster in markets more affected by reductions in immigration. The event-study estimates suggest that from 1923-24 and throughout the decade, a one percentage point increase in immigration exposure led to about 0.8 percent increase in mean hourly wages.

Taken together, the pooled sample results from Table 4 and Table 5 indicate that a 2.9 percentage point reduction in the foreign-born employment share was associated with a 5.9 percent increase in the mean hourly wage of low-skilled workers in the RLM-industry pair. This "elasticity" estimate of about -2 is in line with earlier studies of immigration and wages in the early 20th century, and at the high end of reported estimates for the late 20th century US. Goldin (1994), looking at the impact of immigration on the wages of urban laborers in the 1898-1907 period, concluded that "a 1-percentage-point increase in the fraction of the city's population that was foreign-born decreased wages by about 1.5 to 3 percent", while Hatton and Williamson (1995), working with time series data and a real wage index covering 1890-1913, reported an estimated wage effect of increased immigration over the period that was "very close to Goldin's". Borjas (2003), regressing log average wages for education-experience cells on percent foreign-born using data from the late 20th century obtained coefficients of -.9 for high school dropouts and -2.07 for high school graduates wages. Looking at the effect of Mexican immigration in response to the Peso crisis in the 1990s, Monras (2020) reports short-run wage elasticity estimates in the -0.7 to -1.4 range, smaller than ours in absolute value. It should be remembered, though, that these studies were looking at periods of increasing immigration, while ours looks at a period of declining immigration, and the movements of wages in the two scenarios may not be symmetric.

Columns 2-5 of Table 5 show the wage estimates for the four datasets used in our analysis. The DID estimates for the manufacturing & mining (column 2) and American Contractor (Column 3) samples are similar in magnitude to the pooled sample estimate. The American Contractor estimate is significant only at a 10 percent confidence level, perhaps due to the smaller sample size, our demanding specification, and the fact that it is a sample with only one rather than multiple industries, which makes capturing time trends of wages less precise in such a case (the same is true for the construction union and agriculture samples).

The estimate for the construction union sample is an outsized 1.019 (Column 4), suggesting that the mean city in our sample (exposure measure of 0.15) experienced a 15 percent increase in union wage scales for construction laborers due to immigration restrictions. Since union wage scales are not actual wages but rather a benchmark, and since not all construction workers were unionized, these estimates should be interpreted with caution.[22]

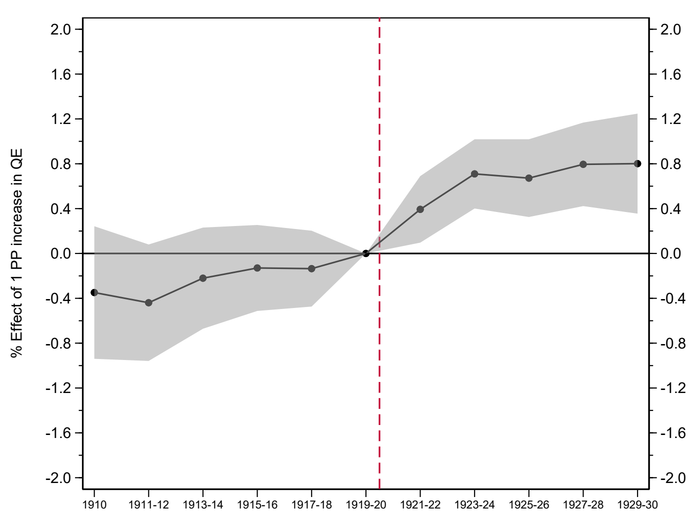
Last, the estimate for the agriculture sample (Column 5, 0.288) is positive yet insignificant

[22]It is reasonable to assume that unionized workers had better working conditions and wages compared to non-unionized workers and that the labor shortages that resulted due to the immigration restrictions have increased the bargaining power of unions, allowing them to increase their wages much more rapidly compared to non-unionized workers.

and smaller in magnitude compared to the pooled sample, suggesting a more muted effects of immigration restrictions on the wages of farm workers during these years, consistent with the result from Table 4 which shows a small and insignificant decline in foreign-born employment share in that sector.

Figure 4 shows the four different event study estimates of the effect of immigration restriction on log mean hourly wages for each dataset included in the pooled sample. The event study estimates are noisier and less precise due to the smaller sample sizes and fewer industries used in each analysis to capture time and industry-specific trends. Nevertheless, all of the post-1920 estimates in the event studies are positive, and the majority are significant at the 5 percent confidence level, consistent with our pooled sample estimates.

Figure 3. Event-Study Estimates of Exposure to Border Closure Policy on Log Mean Wage




Notes: The figure plots the event-study coefficients of the interaction between exposure to border closure policy measure and year from Equation 6. All specifications include year and RLM-industry fixed effects, in addition to dataset- and industry-specific linear time trends and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each RLM-industry pair has a varying number of observations from 1910-1930. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM by industry level.

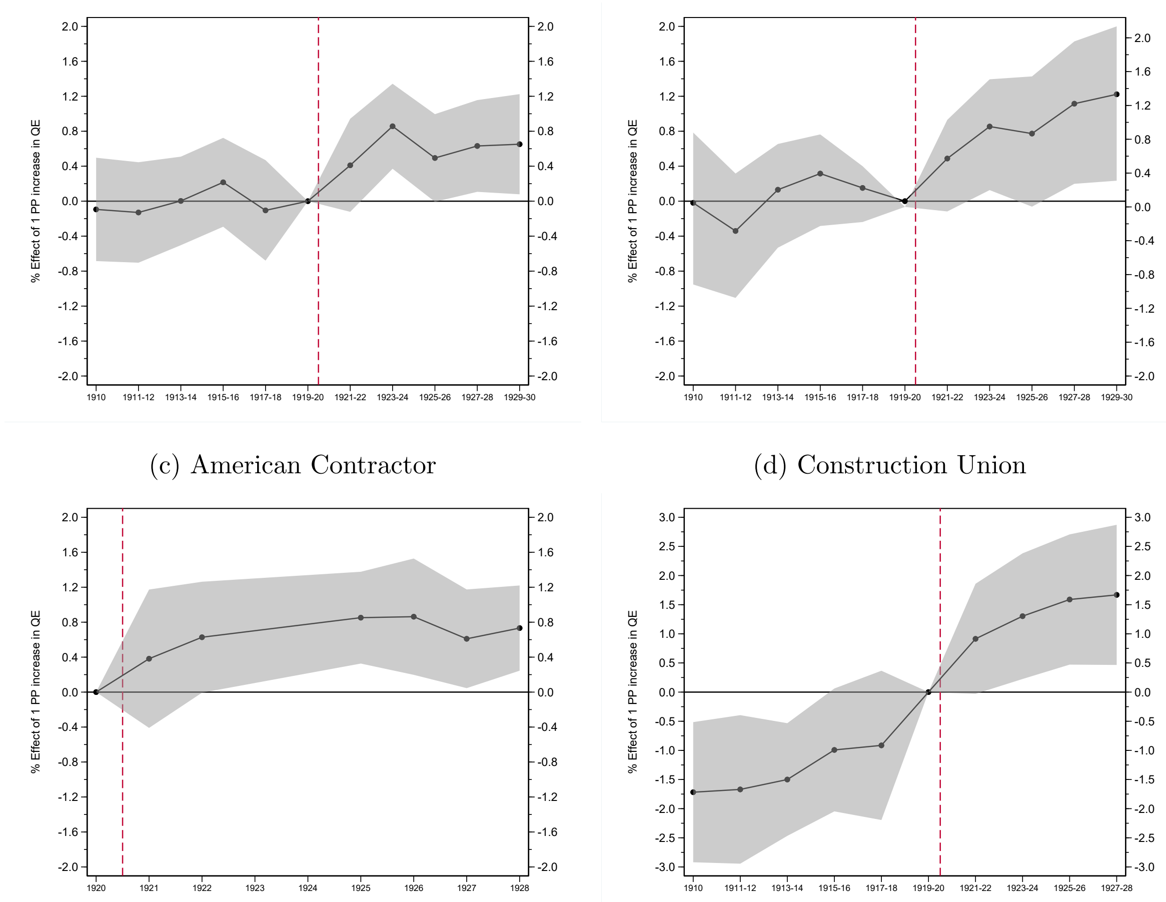
Panel (a) of Figure 4 shows the event study estimates for the manufacturing & mining sample. While the pre-1920 coefficients are noisy, they are all statistically insignificant and close to zero, consistent with no evidence of pre-trends in mean hourly wages growth prior to the inception of the restrictions of immigration.[23] In contrast, all but the 1921-22 coefficients are positive and significant at the 95 percent confidence level.

[23]The only coefficient that is slightly elevated during the pre-1920 period is the 1915-16 coefficient, which might capture a WWI effect. However, it is smaller than all post-1920 coefficients. We show that our analysis is robust to controlling for WWI restrictions in Section 5.4.3.

Figure 4. Dataset-Specific Event-Study Estimates

(a) Manufacturing & Mining (b) Farm Labor

(c) American Contractor (d) Construction Union




Notes: The figures plot the event-study coefficients of the interaction between exposure to border closure policy measure and year from Equation 6 for each of the four datasets used in our analysis separately. All specifications include year and RLM by industry fixed effects, in addition to dataset- and industry-specific linear time trends and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each RLM-industry pair has a varying number of observations from 1910-1930. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM by industry level.

Panel (b) of Figure 4 shows the event study estimates for the agriculture sample. The pre-1920 coefficients hover around zero and are statistically insignificant. The positive and increasing coefficients in the post-1920 period reveal that farm labor wages increased gradually with time and became significant only in the second half of the decade in states with higher exposure of their agriculture sector to immigration, explaining why our estimate in Column 5 of Table 5 is small and insignificant. These results support the conclusion that immigration restrictions did lead to rising wages of farm labor, which provides one causal explanation for Abramitzky et al. (2023) finding of a positive relationship between a rural area's quota exposure and both the level of farm mechanization and share of farm acreage planted with less labor-intensive crops in the 1920s.

Panel (c) of Figure 4 shows the event study estimates for the log mean hourly wages of construction workers in the American Contractor sample. Although we cannot test for the existence of pre-trends using this dataset since it began in 1920, we can confirm that wages of construction laborers increased faster in cities with higher exposure, an estimated effect that became statistically significant in the second half of the decade, with the 1930 estimate suggesting about 0.7 percent increase in mean hourly wages for each one percentage point increase in the city's construction sector exposure to immigration restrictions.

Last, panel (d) of Figure 4 shows the event study estimates on construction union scale wages for laborers. The 1910-14 coefficients are negative and statistically significant, yet increasing over time, suggesting that the parallel trends assumption might not hold for this sample. The increasing pre-trend we observe for the construction union sample is consistent with Medici (2023), which shows that immigration positively affected the emergence of organized labor in the United States during these years. Nevertheless, the 1915-16 and 1917-18 estimates are statistically insignificant but still negative. With that in mind, the post-1920 estimates are positive and significant and increasing over time, reaching an estimated 1.6 percent increase in union wage scales for each one percentage point increase in exposure to immigration restrictions.

## 5.3 Additional Results

### 5.3.1 Semi-Skilled Workers

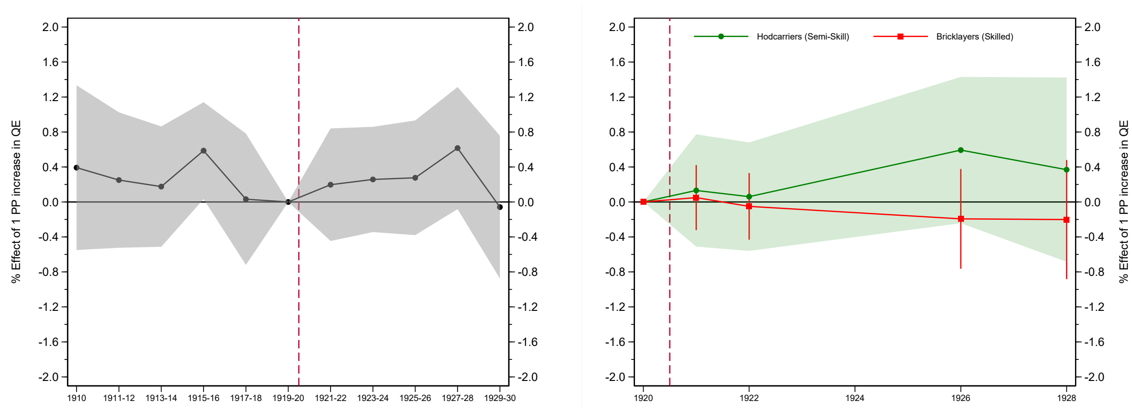
As part of our data collection and digitization process, we have also digitized the wages of semi-skilled operatives using the BLS survey data on manufacturing & mining industries. Column 5 of Table A2 indicates which industries provided information regarding the mean wages of semi-skilled operatives. For 12 of the 21 industries, we identified and recorded the average wage of a large occupation that was classed as a "semi-skilled operative" occupation, yielding a sample of 165 RLM industry pairs. Semi-skilled operatives had higher wages than laborers and were less likely to be recent immigrants. Therefore, we would expect wage pressures due to a shortage of immigrant workers to be less crucial for these occupations and the estimated effects of immigration restrictions to be more muted.

Panel (a) of Figure 5 presents the event-study estimates of exposure to border closure policy on log mean hourly wages for semi-skilled operatives in manufacturing & mining industries. The pre-and post-1920 coefficients are insignificant and remain similar in magnitude, suggesting no effect of exposure to border closure on the wages of semi-skilled workers, consistent with our hypothesis.

We also digitized data from the American Contractor surveys on the wages of hod-carries and bricklayers in different cities, considered semi-skilled and high-skilled occupations in the construction industry, respectively. Panel (b) of Figure 5 shows the event-study estimates for the (log) wages of hod-carriers in green and bricklayers in red. The estimates of both occupations are insignificant throughout the decade, suggesting that the wages of more skilled workers in the construction sector were less affected by the restrictions on immigration, in keeping with economists' expectations at the time.[24]

Figure 5. Exposure to Border Closure Policy and the Wages of Semi-Skilled Workers

(a) Manufacturing & Mining (b) American Contractor




Notes: The figures plot the event-study coefficients of the interaction between exposure to border closure policy measure and year from Equation 6 for the log mean hourly wages of semi-skilled operatives in the manufacturing and mining sample in panel (a) and for hod-carriers (semi-skilled) and bricklayers (skilled) from the American Contractor dataset in panel (b). All specifications include year and RLM-industry fixed effects, in addition to dataset- and industry-specific linear time trends and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each RLM-industry pair has a varying number of observations from 1910-1930. The gray area in panel (a), the green area in panel (b), and the red spikes in panel (b) show 95% confidence intervals. Standard errors are clustered at the RLM by industry level.

### 5.3.2 Municipal Street Laborers

Panel (a) of Figure 6 shows the relationship between quota exposure and mean hourly wages for municipal street laborers in 1928 by exposure quantile. The vertical axis on the left demonstrates that the exposure measure increases exponentially by quantiles. The vertical axis on the right shows that hourly wages of street laborers increase linearly in quota exposure quantile, suggesting that cities and towns whose labor supply was more affected by immigration restrictions paid higher wages to their lowest-skilled employees in the late 1920s.

Panel (b) of Figure 6 further documents that cities and towns more exposed to immigration restrictions had higher hourly wages for street laborers. The graph depicts the partial relationship between exposure and hourly wages of street laborers after controlling for a large set of city-level variables listed in Table A5, along with state fixed effects. The figure sug-

[24]Table A7 provides the DID estimates for the foreign-born shares and (log) mean hourly wages for laborers, hod-carriers, and bricklayers in the American Contractor dataset.

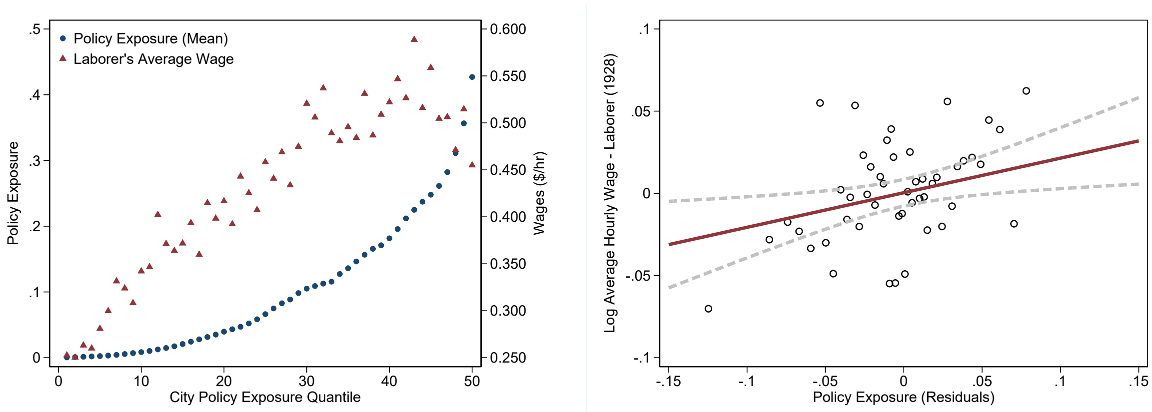
gests that a percentage point increase in a town's exposure is associated with a 0.66 percent larger mean hourly wage, consistent with our DID estimates.

### 5.3.3 Long-Run Effects

Our results indicate that labor markets and industries with increased exposure to immigration had experienced a short-term relative drop in immigration inflows and a subsequent increase in the wages of low-skilled workers. However, Abramitzky et al. (2023) and other studies from the period relying on occupational-level measures of income found no significant wage effect in areas with higher exposure to immigration. As we noted in the introduction, two important reasons for the discrepancy between our results and these studies are that we are using actual wage data and focusing on a population more likely to be impacted by a reduction in immigration - low-skilled laborers.

Figure 6. Exposure to Border Closure Policy and Municipal Street Laborers' Wages

(a) Unadjusted (b) Residualized




Notes: Panel (a) plots the relationship between city policy exposure quantile and policy exposure mean (blue) and municipal street laborers' mean hourly wage in 1928 (red triangles) for each quantile, respectively. Each of the 2,361 cities in towns from the municipal street laborers sample is assigned into quantiles. Panel (b) plots the relationship between policy exposure and log mean hourly wage in 1928 for municipal street laborers. Each city or town's policy exposure and wages are adjusted for the set of controls listed in Table A5, and state fixed effects. The red line presents the fitted linear relationship between the residuals of policy exposure and log mean hourly wages. The dashed gray lines are 95% confidence intervals.

Nevertheless, several studies highlight the importance of time horizon adjustment dynamics when evaluating the effects of immigration on the labor market. In particular, Monras (2020) shows that adjustment factors such as internal migration and local technologies might play an important role in mitigating the short-term effects of an immigration shock, while Abramitzky et al. (2023) show that the immigration restrictions in the 1920s did induce internal migration from low- to high-exposure areas and also increased adoption of capital-intensive technologies and cops in agriculture. Clemens et al. (2018) show that eliminating the Bracero "guest worker" program for agricultural labor induced farmers to adopt labor-saving technologies and crops.

Table 6. Long-Run Effects of Exposure to Border Closure Policy on Low-Skilled Wages

| Sample: | Pooled (1) | Manufacturing & Mining (2) | American Contractor (3) | Construction Union (4) | Agriculture (5) |
|---|---|---|---|---|---|
| **A. Log Total RLM-Industry Employment:** | | | | | |
| Quota Exposure X (Year > 1920 & Year < 1940) | -0.921** | -1.446* | -0.305 | -0.151 | 0.357 |
| | (0.463) | (0.803) | (0.390) | (0.383) | (0.905) |
| Quota Exposure X (Year = 1940) | -1.744** | -2.454 | -0.536 | -0.819 | -1.116 |
| | (0.860) | (1.493) | (0.503) | (0.722) | (1.441) |
| Pre-1920 Dependent Mean | 3,745 | 827 | 9,738 | 14,851 | 852 |
| Observations (RLM-Industry Pair) | 1,612 | 996 | 244 | 180 | 192 |
| **B. Log Mean Hourly Wage (RLM-Industry)** | | | | | |
| Quota Exposure X (Year > 1920 & Year < 1940) | 0.613*** | 0.548*** | 0.622** | 0.721** | 0.477* |
| | (0.146) | (0.174) | (0.239) | (0.289) | (0.239) |
| Quota Exposure X (Year = 1940) | 0.538** | 0.718** | -0.0605 | 0.607 | 0.0725 |
| | (0.242) | (0.291) | (0.492) | (0.677) | (0.610) |
| Pre-1920 Mean Hourly Wage ($) | 0.308 | 0.325 | 0.544 | 0.420 | 0.190 |
| Quota Exposure - Mean | 0.10 | 0.11 | 0.13 | 0.15 | 0.07 |
| Quota Exposure - SD | 0.08 | 0.08 | 0.10 | 0.10 | 0.06 |
| Observations | 2,975 | 1,234 | 294 | 487 | 960 |

Notes: This table presents estimates of the interaction coefficients between quota exposure measure $QE$ and the post-1920 and 1940 indicators. Column 1 presents results for the pooled sample, and columns 2-5 present results for each data source separately. In panel A, the outcome variable is the log total RLM-industry employment of low-skilled workers. In panel B, the outcome variable is the log mean hourly wages for RLM-industry pair. All specifications include year and RLM-industry fixed effects, in addition to dataset- and industry-specific linear time trends and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each RLM-industry pair has a varying number of observations from 1910-1930. Robust standard errors, clustered at the RLM by industry, level in parenthesis.. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

To investigate the long-run effects of immigration restrictions, we use the 1940 Full Count Population Census to estimate the impact of the 1920s restrictions on wages and low-skilled employment for the industries we investigate in our analysis. We begin by examining what happened to labor supply. Panel A of Table 6 shows the DID estimates of the impact of the quota exposure measure on log total RLM by industry employment of low-skilled workers. The pooled sample estimates in column 1 show that in the 1920s, the average RLM by industry low-skilled workforce dropped by 9.2 percent and that this decline increased to 17.4 percent by 1940. That is, even though internal migration and immigration from non-restricted countries increased to higher exposure areas as Abramitzky et al. (2023) shows, this was not enough to substitute the missing immigrants from restricted countries. Columns 2-5 show that the negative effect of labor supply occurred across all of our datasets, but the estimates are imprecise.

Panel B of Table 6 shows the DID estimates for log mean hourly wages of low-skilled workers. For the pooled sample, the 1920s DID coefficient is .613, and it drops slightly to .538 in 1940, indicating that the higher relative wages for low-skilled workers persisted over time as low-skilled labor remained in shortage. We do see the most substantial effect in the mining and manufacturing dataset, where labor shortages by 1940 seem to be the worst, while in construction and agriculture, the elevated wage growth seems to have moderated despite the decrease in labor supply, potentially due to adoption of new technologies and organization of labor in these industries. However, these estimates should be taken with caution as they come after the Great Depression and as the impact of World War II was taking effect across the U.S., and also because wages and labor markets in the 1940 Census are not precisely comparable with our wage data despite our best attempts.

## 5.4 Robustness Checks

### 5.4.1 Defining A Labor Market Based on Geography Rather than Industry

One concern about our decision to define a labor market using geography (state or city) and industry is that it implies that geographical differences in exposure to immigration reductions generated the same wage dynamics as inter-industry differences in exposure within the same geographic area. However, it seems reasonable to suppose that if highly exposed industries began to raise wages in response to labor shortages, low-skilled workers in less exposed industries in the same geographical area could have switched industries reasonably easily since there is no significant skill constraint that bars low-skilled workers in one industry from switching to work in another industry. Put another way, our specification ignores the possibility that wage equalization across industries in a given geographic region (state or city) could occur more quickly than wage equalization across geographic regions with different levels of exposure to immigration restrictions.

If low-skilled workers within a geographical labor market moved fairly quickly to take advantage of emerging inter-industry wage differentials, a specification that ignores inter-industry differences in wages and exposure within a geographic area and relies only on geographic differences should lead to estimated wage effects of quota exposure that are larger than those reported in Table 5. Based on this reasoning, we collapse our RLM-industry pairs to the RLM (jurisdiction) level by constructing an exposure measure for the jurisdiction that assumes that there is no variation between industries in exposure to immigration restrictions and compute the (employment-weighted) mean wage across all industries in a given year for the RLM. Our findings in Table A8 show that higher exposure of an RLM to immigration restrictions reduced the overall foreign-born share in that RLM and increased mean real wages for low-skilled workers, although the wage effect is marginally insignificant. However, the wage estimate is smaller than the one reported in Table 5, suggesting that wage equalization across industries within an RLM in response to industry-specific labor supply shocks caused by immigration restrictions was not a considerably important phenomenon.

#### 5.4.2 Possible SUTVA Violations

One of the assumptions required for the DID model to be interpreted as causal is the Stable Unit Treatment Value Assumption (SUTVA). This assumption requires that there should be no changes in the composition of the treatment and control groups and that there is no spillover effect from the treatment of one unit on the outcome of another unit. However, as observed in prior studies by Abramitzky et al. (2023) and Price et al. (2020), the restrictions on immigration in the 1920s resulted in population composition changes in labor markets due to increased internal migration from low- to high-exposure areas. As a result, SUTVA could be violated in our context because the immigration policy restrictions indirectly affect labor markets that experience increased out-migration to more highly exposed labor markets.

We would note that first, we are claiming to identify the effect of immigration restrictions and disruptions on more-exposed markets relative to less-exposed markets, given the structure of interactions across relevant labor markets that determine equilibrium adjustments. Put another way, we are identifying wage dynamics in the 1920s in response both to the labor supply shocks directly caused by the quota laws and the movements of labor in response to the wage changes initially caused by those labor supply shocks. Second, if migration from low- to high-exposed markets pushed wages down in the latter while pulling them up in the former, our estimates would be downward-biased estimates of the initial, direct effect of the immigration reductions on wages.

Nevertheless, we test for this potential bias by conducting two robustness checks: omitting labor markets with low exposure to treatment and labor markets with high rates of out-migration to treated areas. The first test is based on the idea that the least exposed areas are potentially most likely to have experienced high out-migration to high-exposed areas. The second test is based on Abramitzky et al. (2023) calculations of out-migration rates to higher exposed labor markets between 1920 and 1930. They define a labor market as a State Economic Area (SEA), a collection of counties within a state that are economically integrated. For our purposes, we compute the mean out-migration rate to highly exposed areas across SEAs within a given state and use the state's mean out-migration rate as a proxy for the extent of changes in composition in each relevant labor market in our sample. We then exclude from the estimation sample areas that this measure indicates were likely exposed to larger composition changes and spillover effects.

Table A9 provides the results of these tests for our the pooled sample. The outcome in Panel A is the relevant labor market by industry foreign-born share (in each decennial census), and the outcome in Panel B is the log mean hourly wage for each relevant labor market by industry pair (annual observations). Column 1 shows our baseline estimates from column 1 of Table 4 for reference. In columns 2-4, we show our first test results when excluding RLM-Industry pairs at the exposure measure's bottom 1st,5th, and 10th percentiles. The foreign-born share and wage estimates remain unchanged in any of the specifications. In columns 5 and 6 of Table A9, we show the results of excluding RLM-industry pairs with the highest 5 percent or 10 percent of values of our "out-migration to higher exposure markets" rate. In column 7, we exclude the state with the highest out-migration rate within each census division. Again, the coefficients are all statistically significant and are similar to the baseline estimate.

### 5.4.3 World War I

The fact that World War I preceded the change in immigration policy makes disentangling the effect of one from another challenging, which is why we have characterized our estimates as capturing a combination of the effect of WWI and legislated changes in immigration policy. As noted earlier, we have constructed a World War I restriction measure in the spirit of Abramitzky et al. (2023) that aims to capture how much immigration flows were restricted due to the war and the disruption in immigration it caused. The fact that it is highly correlated with our quota exposure measure makes distinguishing the effect of WWI versus that of changes in immigration policy even more complex.

As a result, one might worry that WWI entirely drives our results and that policy changes had no real effect on the wages of low-skilled workers. For example, it might be the case that the direct and varying effects of WWI on U.S. labor markets due to war production or the removal of men from the labor force due to military service were correlated with the effect of reduced immigration due to the later immigrant quotas. While we cannot fully conclude that WWI effects do not play a role in our estimates, we can test how important this role might be. In particular, we assume that the war had a short-term effect on production and labor supply that was limited to the war years, as it is unlikely that these increases in demand and reductions in labor supply persisted for long after the war's end. Thus, we re-estimate our main specification using our war exposure measure interacted with the war years:

$$y_{ijt} = \delta_{ij} + \gamma_t + \beta(QE_{ij} \times 1(t \geq 1920)_t) + \theta(WE_{ij} \times 1(t \geq 1915, t \leq 1918)_t) + \Theta(X_{ij1910} \times t) + \varepsilon_{ijt} \quad (7)$$

Table A10 presents our results for this specification for the pooled sample. Column 1 in those tables presents our preferred estimations, column 2 includes only the WWI exposure measure term, and column 3 includes both the quota and WWI exposure measures, respectively. Across all specifications, the coefficients on WWI exposure are significantly smaller in magnitude than those of the quota exposure, suggesting that the effect on labor markets of the war alone made little contribution to our overall findings.

### 5.4.4 Alternative Exposure Measures

We test whether our results are sensitive to alternative definitions of the quota exposure measure in Table A11. Column 1 shows our baseline specification. In column 2, we alter our country-specific quota intensity measure $QI_r$ in column 1 of Table A1 from a counterfactual that measures the percent of immigrants who did not arrive due to the restrictions to a binary indicator that is equal to one for high-restriction countries or regions listed in Panel A and zero otherwise. In columns 3 and 4, we use the 1900 Census to calculate the county- and industry-specific geographic distribution of immigrants in Equations 1, 2, and 3. All the estimates are significant and similar in magnitude to our baseline estimates.

## 6 Composition Effects

Our wage observations are means from samples that include both native and foreign-born workers. This raises a concern that the relative increase in mean wages for low-skilled labor that we observe in more exposed markets might be due to composition effects. That is, if we assume that in this period, native workers earned more than immigrant workers for the same type of work in the same labor market, the decline in the foreign-born share of workers in the more exposed markets would lead mechanically to an increase in the average wage in those markets, without any increase in the wages paid to either native or foreign-born workers.

Goldin (1994) discusses this issue in her analysis of the impact of immigration on wages in the early 20th century U.S. Using a first-differences regression model, she estimates that the effect of immigration on the mean wage of laborers in cities is generally negative and often substantial. However, she also notes that her results might suffer from the mechanical effect of increasing immigration. She believes this to be unlikely, however, writing "...The difference in wages between immigrants and natives in the same occupation would have to have been extremely high to account for the large negative impact of immigration on wages in general and even for those occupations in which the foreign-born were a large percentage." (Goldin 1994, p. 253). We take Goldin's argument one step further and test it by using our data to approximate the contribution of composition effects to our estimates and by considering how large the wage differential between natives and immigrants would have to have been for our estimates to be explained solely by the mechanical effect of fewer immigrant workers in highly exposed labor markets.

### 6.1 Immigrant Wage Penalty

We begin by estimating a native-immigrant wage differential using two data sources. The first data source we use is the report of the US Immigration Commission or Dillingham Commission. The commission was established by Congress in 1907 to conduct a thorough investigation into "the subject of immigration". One aspect of the commission's investigation was the collection of employment and earnings data in over 20 industries that employed a disproportionate number of immigrants. Information on over half a million immigrant and native-born workers was obtained from employer payroll records of 1907 (United States Immigration Commission, 1911). To measure the immigrant-native wage differential for each industry, we use the reported industry-wide average wage ratio for foreign-born male employees to the comparable earnings figure reported for native-born workers. The red circles in Figure 7 plot the data the commission collected by industry, and the red line shows the mean wage penalty across those industries. The commission reported data for 8 of the 21 industries used in our analysis. According to the commission's data, the immigrant wage penalty in those 8 industries was 9.3% on average, and it ranged from 25% in the construction industry to a wage premium of about 5% in metal mining.

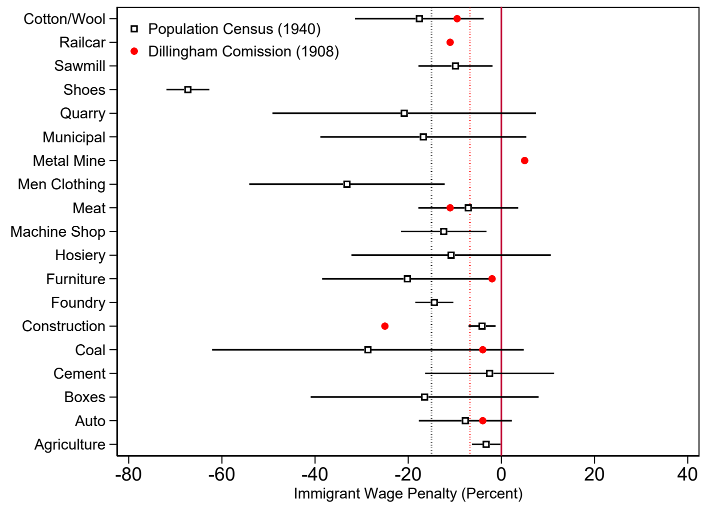
The Dillingham Commission wage data have two caveats. First, they do not report wage differentials for all the industries covered in our analysis. Second, they report industry-wide wage differentials rather than low-skilled labor wage differentials, which might imply an inflated wage penalty since immigrants were more likely to be concentrated in low-skilled and low-wage occupations during this period.

To address the caveats of the Dillingham Commission data, we complement it by estimating the immigrant wage penalty using the 1940 Full Count Population Census, the first Census to report actual wages. To estimate the immigrant wage penalty, we construct a sample of working-age males aged 16-65 with a reported laborer occupation in one of the 21 industries used in our analysis. Then, we estimate the following regression model for the sample of low-skilled workers:

$$ln(w_{ijc}) = \beta_j immig_{ijc} + \gamma_j + \delta_c + \varepsilon_i \tag{8}$$

where $ln(w_{ijc})$ is the log hourly wage of individual $i$ in industry $j$ in county $c$, $\beta_j$ are industry indicators, $immig_{ijc}$ is a recent immigrant indicator (arrived within the last 5 years), $\gamma_j$ are industry fixed effects, and $\delta_c$ are county fixed effects.

Figure 7. Immigrant Wage Penalty Estimates by Industry




Notes: The figure plots the industry-specific immigrant wage penalties reported by the Dillingham Commission (1908) in red, and our 1940 Population Census estimates in black squares. The 1940 population Census estimates are the coefficients on a recent immigrant indicator from a regression where the outcome is an individual log mean hourly wage and additional demographics, in addition to industry and county fixed effects. The vertical bars show 95% confidence intervals.

The black squares in Figure 7 plot the 1940 Census estimates for industries and their respective 95 percent confidence intervals. The estimates range from 5 percent in the agriculture sector to as large as 70 percent in the shoes industry. The black dashed line shows the aggregated estimate from the 1940 Census, which is 9.6 percent.[25]

### 6.2 Assessing the Composition Effects Channel

We assess how much of the estimated wage effects can be attributed to induced changes in the composition of the employment. Our back-of-the-envelope calculations suggest that about 10% of the total effect of immigration restrictions on mean wages in the median RLM-industry pair is due to composition shifts and that such changes account for less than 25% the total effect in 90% of the RLM-industry pairs in our sample.

We take to the following steps to assess the magnitude of the composition effect in our analysis. First, we predict the post-1920 foreign-born share for each RLM-industry pair using each RLM-industry's quota exposure measure and our estimate of the effect of exposure to immigration restrictions on foreign-born employment share in Column 1 of Table 4. Second,

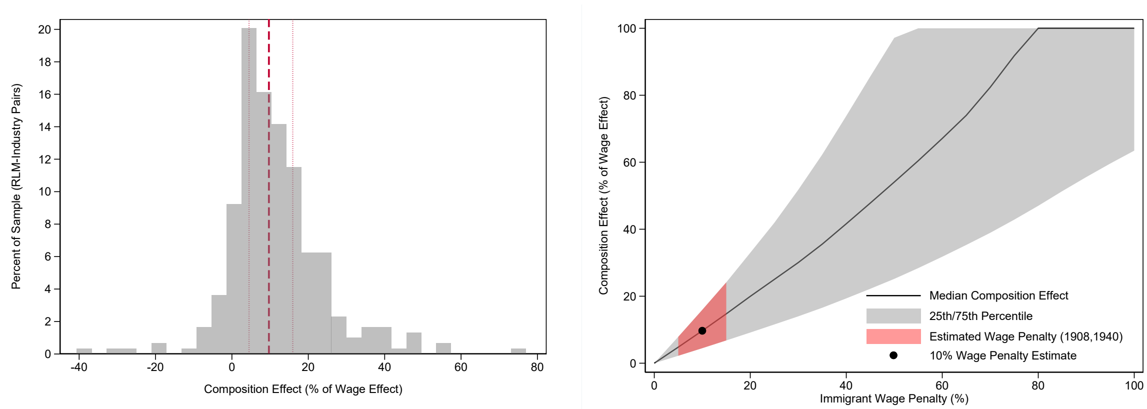
[25]We also have another specification where we control for individual demographics, including age and age-squared, place of birth, marital status, education, and race. The estimated immigrant wage penalty is 6.2% in that specification.

we assume a constant immigrant wage penalty of 10 percent based on our estimates in Figure 7.[26] Third, we use the pre-1920 mean hourly wage and foreign-born employment share at the RLM-industry level to calculate the pre-1920 foreign- and native-born mean hourly wages. Fourth, we calculate the implied post-1920 mean hourly wage for the RLM-industry pair if the only thing that changed was the composition of the population and the pre-1920 hourly wages remained unchanged. Fifth, we compute the percent wage growth for the RLM-industry pair due to composition change only by taking the ratio of the composition implied post-1920 mean hourly wage and the pre-1920 mean hourly wage. Sixth, we compute the implied wage growth due to immigration restrictions by multiplying the RLM-industry quota exposure measure by our estimate in Column 1 of Table 5. Last, we compute the percentage of wage effect due to composition by taking the ratio of the implied growth due to composition alone relative to the overall wage growth due to immigration restrictions for each RLM-industry pair.

Figure 8. Composition Effects Analysis

(a) 10% Wage Penalty Distribution

(b) Varying Wage Penalties




Notes: Panel (a) plots the histogram of the percent contribution of composition shifts on the overall estimated effect of immigration restrictions on log mean hourly wages under the assumption of a constant and uniform 10 percent immigrant wage penalty. The dashed red line shows the value for the median RLM-industry pair, and the thin dotted red lines present the 25th and 75th percentiles. Panel (b) presents the back-of-the-envelope distribution of composition effects for varying immigrant wage penalties. The black line plots the median composition effect for each wage penalty, the gray area shows the 25th and 75th percentiles of the composition effect, and the red area highlights the estimated composition effects that match our estimates of immigrant wage penalties from Figure 7.

Panel (a) of Figure 8 shows the distribution of the percent of total wage effect that is due to composition shift across RLM-industry pairs when assuming a constant 10 percent wage penalty. The thick dashed red line presents the median percent, estimated at slightly less than 10 percent, and the thin dotted red lines show the 25th and 75th percentiles of the composition effect. Most RLM-industry pairs have a very low estimated composition effect,

[26] We relax this assumption by allowing the immigrant wage penalty to vary from 0 to 100 percent. However, we assume the wage penalty remains constant over time and across industries and locations.

suggesting that composition shifts are not the primary explanation for the estimated wage effects in most RLM-industry pairs.

Panel (b) of Figure 8 expands the analysis to varying immigrant wage penalties. The black line presents the median composition effect for varying wage penalties, and the gray area represents the 25th and 75th percentiles. The red area shows the range of estimates consistent with our estimates of immigrant wage penalties from the Dillingham Commission and the 1940 Census. As expected, composition effects are more prominent when the wage penalty is higher. However, our back-of-the-envelope calculations suggest that composition effects explain our findings entirely only when the immigrant wage penalty is above 80 percent, consistent with Goldin (1994)'s argument cited above.

Overall, our analysis suggests that while composition effects have played a limited role in increasing mean wages in areas more affected by the immigration restrictions, they were certainly not the dominant factor behind those increases in most RLM-industry pairs.

## 7 Conclusion

This paper explores the effect on the wages of low-skilled labor from the substantial restrictions to immigration to the United States caused by the "quota laws" of 1921 and 1924. To identify the effect of these immigration restrictions on workers' wages, we rely on the insight of past researchers that the size of the negative labor supply shock experienced by a local labor market as a result of policy-related immigration reductions would have depended on the ethnic composition of the labor force in that market prior to the restrictions. We apply this identification strategy to various sources of historical data on actual wages being paid during the period 1910-1929 to low-skilled laborers, the group whose wages were most likely to have been affected by the disruptions to immigration.

We conducted analyses of wages paid to laborers in the mining and manufacturing industry, wages paid to construction workers, wages of street cleaners and repair workers hired by municipalities, and wages of farm laborers. All four analyses support the same conclusion: during the 1920s, low-skilled workers in labor markets who experienced larger adverse shocks to their labor supply as a result of the disruptions to immigration were being paid higher wages. This result is robust to the inclusion of a variety of pre-war labor market characteristics that might have been correlated with wages or wage growth in the pre-war period. We also showed that this pattern of geographic wage differentials did not exist before the war and that the post-war pattern was not a result of differential trends across labor markets in the pre-war period.

Our finding of higher wages in areas more affected by the reductions in immigration is consistent with the finding of earlier researchers that by 1930, these areas had received more significant inflows of internal migrants and immigrants from countries unaffected by the immigration quotas imposed in the early 1920s. They are also consistent with evidence that farmers in rural areas that lost more immigrant labor due to the quotas were quicker to mechanize and shift to less labor-intensive crops. A matter for future research is whether these increases in the wages of low-skilled labor contributed to the wave of mechanization in manufacturing that was a source of comment and concern among economists and policymakers in the 1920s (Woirol, 2006).

While the results suggest that reduced immigration led to wage increases for low-skilled workers, the extent to which these findings apply to other immigration policy changes or different economic settings remains uncertain. Factors such as internal migration adjustments, technological shifts in labor markets, and differences in industry composition may influence the results' external applicability. Nonetheless, the study provides valuable insights into the economic consequences of immigration restrictions, which could inform contemporary policy debates on immigration and labor market dynamics.

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## For Online Publication: Online Appendix

Table A1. Country and Region-Specific Quota Intensity and War Restriction Variables

| **Country or Region** | **Quota Intensity Measure (1)** | **WWI Restriction Measure (2)** |
|---|---|---|
| **A. High-Restriction Countries:** | | |
| Asia | 0.95 | 0.50 |
| Central Europe | 0.97 | 0.98 |
| Eastern Europe | 0.94 | 0.96 |
| Greece | 0.97 | 0.50 |
| Italy | 0.96 | 0.89 |
| Portugal | 0.95 | 0.41 |
| Rest of World | 0.69 | 0.00 |
| Russia | 0.93 | 0.95 |
| Spain | 0.98 | 0.14 |
| **B. Low-Restriction Countries:** | | |
| Germany | 0.00 | 0.92 |
| Ireland | 0.00 | 0.79 |
| Scandinavia | 0.10 | 0.68 |
| United Kingdom | 0.00 | 0.80 |
| Western Europe | 0.56 | 0.72 |
| **C. No-Restrictions Countries:** | | |
| Canada | 0.00 | 0.00 |
| Caribbean | 0.00 | 0.11 |
| Latin America | 0.00 | 0.00 |
| Mexico | 0.00 | 0.00 |

Source: Abramitzky et al. (2023).

Table A2. Datasets Scope and Coverage

| | 1950 Industry Code | Observations | RLMs | Years Covered | Semi-Skilled Wages |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| **A. Manufacturing and Mining:** | | | | | |
| Auto: laborers | 376 | 28 | 7 | 1919, 1922, 1925, 1928 | Yes |
| Boxes: laborers | 457 | 17 | 11 | 1919, 1925 | |
| Cement: laborers | 317 | 4 | 4 | 1929 | |
| Coal: In-mine laborers | 216 | 48 | 10 | 1919, 1922, 1924, 1926, 1929 | Yes |
| Coal: pick miners | 216 | 43 | 9 | 1919, 1922, 1924, 1926, 1929 | Yes |
| Coal: Outside of mine laborers | 216 | 48 | 10 | 1919, 1922, 1924, 1926, 1929 | Yes |
| Cotton Textiles: doffers | 439 | 99 | 12 | 1911-1914, 1916, 1918, 1922, 1924, 1926, 1928 | Yes |
| Foundries: laborers | 336-338, 346-348 | 123 | 28 | 1919, 1923, 1925, 1927, 1929 | Yes |
| Furniture: laborers | 309 | 41 | 17 | 1915, 1919, 1929 | |
| Hosiery & Underwear: "other occupations" | 436, 446, 449 | 73 | 16 | 1910-1914, 1922, 1926, 1928 | |
| Iron Mines: laborers, outside mine | 206 | 3 | 3 | 1924 | |
| Machine Shops: laborers | 356-358, 367 | 123 | 28 | 1919, 1923, 1925, 1927, 1929 | Yes |
| Meat Packing: laborers (maintenance and repair) | 406 | 5 | 1 | 1921, 1923, 1925, 1927, 1929 | Yes |
| Men's clothing: basters (coat) | 448 | 59 | 7 | 1911-1914, 1919, 1922, 1924, 1926, 1928 | Yes |
| Metal Mining: laborers (outside mine) | 206 | 8 | 8 | 1924 | |
| Mill Work (Wood): laborers | 307-308 | 66 | 12 | 1910-1913, 1915, 1919 | |
| Quarries: laborers | 236 | 4 | 4 | 1929 | |
| Railcar Manufacture: laborers | 379 | 63 | 17 | 1910-1913 | |
| Sawmills: laborers | 307-308 | 211 | 23 | 1910-1913, 1915, 1919, 1921, 1923, 1925, 1928 | Yes |
| Shoes: cutters, trimming hand | 488 | 81 | 12 | 1910-1914, 1920, 1922, 1924, 1926, 1928 | Yes |
| Wollen Textile: dyehouse laborers | 439 | 87 | 9 | 1910-1914, 1916, 1918, 1920, 1922, 1924, 1926, 1928 | Yes |
| **B. Construction:** | | | | | |
| American Contractor Dataset | 246 | 294 | 61 | 1920-1922, 1925-1928 | Yes |
| BLS Construction Union Wage Scales | 246 | 487 | 45 | 1910-1928 | |
| **C. Agriculture:** | | | | | |
| USDA Farm Laborer Wage Dataset | 105 | 960 | 48 | 1910-1929 | |
| Observations | | | | 2,975 | |
| Number of Relevant Labor Markets (RLMs) | | | | 128 | |
| Number of Industries | | | | 23 | |
| Number of RLM-Industry Pairs | | | | 402 | |

Notes: Panels A through C break down industries by data source and the industries reported in them. Column 1 shows the 1950 Census industry codes matched to each industry title. Column 2 shows the number of observations in the sample for each industry. Column 3 shows the number of jurisdictions (RLM - state or city) that have at least one observation for the industry, Column 4 shows the years where data is available for each industry, and Column 5 indicates whether wages of semi-skilled operatives are also reported for the industry.

Table A3. Composition of Foreign-Born Population by Exposure to Immigration Restrictions

| **Quota Exposure Quartile:** | **Full (1)** | **Bottom (2)** | **Second (3)** | **Third (4)** | **Top (5)** |
|---|---|---|---|---|---|
| Number of RLMs | 403 | 103 | 101 | 99 | 100 |
| Quota Exposure Measure (RLM): | 0.109 | 0.016 | 0.066 | 0.137 | 0.220 |
| Foreign-Born Share (1910): | 0.170 | 0.035 | 0.136 | 0.230 | 0.285 |
| **A. High-Restriction Countries:** | **34%** | **30%** | **28%** | **32%** | **46%** |
| Asia | 2.3% | 3.3% | 1.4% | 2.2% | 2.1% |
| Central Europe | 9.8% | 5.3% | 8.9% | 10.3% | 15.0% |
| Eastern Europe | 0.4% | 0.3% | 0.4% | 0.5% | 0.4% |
| Greece | 1.2% | 2.0% | 1.0% | 1.1% | 0.6% |
| Italy | 8.6% | 7.2% | 6.5% | 7.6% | 13.2% |
| Portugal | 0.3% | 0.1% | 0.0% | 0.5% | 0.6% |
| Rest of World | 1.2% | 1.6% | 1.0% | 1.1% | 1.0% |
| Russia | 10.2% | 9.7% | 8.7% | 9.1% | 13.1% |
| Spain | 0.2% | 0.2% | 0.2% | 0.1% | 0.4% |
| **B. Low-Restriction Countries:** | **53%** | **58%** | **56%** | **53%** | **45%** |
| Germany | 20.9% | 23.0% | 25.9% | 19.4% | 15.3% |
| Ireland | 8.9% | 8.2% | 7.8% | 8.7% | 10.9% |
| Scandinavia | 7.3% | 4.1% | 8.3% | 10.2% | 6.6% |
| United Kingdom | 12.1% | 17.4% | 10.0% | 11.0% | 10.1% |
| Western Europe | 3.9% | 5.0% | 3.8% | 4.3% | 2.6% |
| **C. Non-Restriction Countries:** | **13%** | **13%** | **16%** | **14%** | **8%** |
| Canada | 10.6% | 7.9% | 14.1% | 13.6% | 6.9% |
| Caribbean | 0.4% | 0.6% | 0.2% | 0.1% | 0.6% |
| Latin America | 0.1% | 0.1% | 0.0% | 0.0% | 0.0% |
| Mexico | 1.7% | 3.8% | 1.7% | 0.4% | 0.7% |

Notes: The table displays the 1910 Relevant Labor Market (RLM) by the composition of the foreign-born population based on their country of origin and quota exposure quartiles. Column 1 includes data for the full (pooled) sample of 403 different RLMs by industry pairs. Columns 2 through 5 provide information ranging from the lowest quota exposure quartile (column 2) to the highest (column 5). For each sample, we present the number of observations, the mean quota exposure measure, and the share of the foreign-born population in 1910. Additionally, the foreign-born population share is further broken down by country of origin, displaying the percentage of individuals from each country within the overall foreign-born population.

Table A4. American Contractor Construction Contractors Wage Survey and Construction Trade Union Wage Scales - Participating Cities and Coverage

| **City** (1) | **Wage Survey** (2) | **Union Wage** (3) | **City** (4) | **Years Covered** (5) | **Union Wage** (6) |
|---|---|---|---|---|---|
| Akron, OH | YP | N | New Haven, CT | YP | Y |
| Alliance, OH | YP | N | New Orleans, LA | N | Y |
| Atlanta, GA | YF | Y | New York, NY | YF | Y |
| Baltimore, MD | YF | Y | Newark, OH | YP | N |
| Binghamton, NY | YP | N | Norfolk, VA | YF | N |
| Boston, MA | YF | Y | Omaha, NE | YF | Y |
| Bridgeport, CT | N | Y | Peoria, IL | N | N |
| Buffalo, NY | YF | Y | Philadelphia, PA | YF | Y |
| Butte, MT | N | Y | Pittsburgh, PA | YF | Y |
| Chicago, IL | YF | Y | Portland, OR | N | Y |
| Cincinnati, OH | YF | Y | Portland, ME | N | Y |
| Cleveland, OH | YF | Y | Providence, RI | N | Y |
| Columbia, SC | YP | N | Raleigh, NC | YP | N |
| Columbus, OH | YF | Y | Reading, PA | YF | N |
| Dallas, TX | N | Y | Redfield, SD | YP | N |
| Dayton, OH | YF | N | Richmond, IN | YP | N |
| Denver, CO | N | Y | Richmond, VA | YF | N |
| Des Moines, IA | YP | Y | Rochester, NY | YP | Y |
| Detroit, MI | YF | Y | Saginaw, MI | YP | Y |
| Dubuque, IA | YP | N | Salt Lake City, UT | N | Y |
| Duluth, MN | YP | N | San Francisco, CA | YP | Y |
| Erie, PA | YF | Y | Savannah, GA | YP | N |
| Fairmont, WV | YP | N | Scranton, PA | N | Y |
| Fitchburg, MA | YP | N | Seattle, WA | N | Y |
| Flint, MI | YP | N | Sharon, PA | YP | N |
| Grand Rapids, MI | YF | Y | Shreveport, LA | YP | N |
| Greensboro, NC | YP | N | Sioux City, IA | YF | N |
| Houston, TX | N | Y | Spokane, WA | N | Y |
| Indianapolis, IN | YF | N | St. Joseph, MO | YF | N |
| Kansas City, MO | N | Y | St. Louis, MO | YF | Y |
| Kent, OH | YP | N | St. Paul, MN | N | Y |
| Lansing, MI | YP | N | St. Petersburg, FL | YF | N |
| Lima, OH | YP | N | Toledo, OH | YP | Y |
| Little Rock, AR | YP | Y | Warren, OH | YP | N |
| Los Angeles, CA | N | Y | Washington, DC | YF | Y |
| Louisville, KY | YF | Y | Webster City, IA | YP | N |
| Memphis, TN | YF | N | Wichita, KS | N | Y |
| Milwaukee, WI | YF | Y | Youngstown, OH | YF | N |
| Minneapolis, MN | N | Y | | | |

Notes: The table shows a list of the cities where (i) the American Contractor journal surveyed construction contractors regarding construction workers' wages throughout the 1920s and (ii) the BLS did an annual survey of construction trade union leaders where they reported a union wage scale for building laborers. The wage survey years are 1920, 1921, 1922, 1925, 1926, 1927, and 1928. The union wage scale survey years are 1910-1928. In the wage survey columns 2 and 4, "YF" means the city has data for each of the survey years, "YP" means the city has data for some of the survey years, and "N" means the city has no wage data at all. In the union wage scale columns, "Y" implies the city has union wage scale reported in at least one year, and "N" means no union wage scale is reported for the city.

Table A5. Lasso Results for the Relationship Between Exposure to Border Closure Policy and 1910 Region Characteristics

| **Sample:** | **Pooled** | **Manufacturing & Mining** | **American Contractor** | **Construction Union** | **Agriculture** |
|---|---|---|---|---|---|
| | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** |
| **A. Demographics:** | | | | | |
| Foreign-Born Share | x | x | x | x | x |
| Log Total Population | x | x | x | x | x |
| Share Urban Population | - | - | - | - | - |
| Share Black Population | - | - | - | - | - |
| Literacy Rate | - | - | - | - | - |
| **B. Occupation:** | | | | | |
| Share Workers in Blue-Collar Occupations (Manufacturing) | - | - | - | - | - |
| Share Workers in Farming Occupations (Agriculture) | x | x | x | - | - |
| Share Workers Holding White Collar Occupation | - | - | - | - | - |
| **C. Industry:** | | | | | |
| Share Workers in Mining | - | - | - | - | - |
| Share Workers in Construction | - | - | - | - | x |
| Share Workers in Transportation | - | - | - | - | - |
| Share Workers in Wholesale/Retail | - | - | - | - | x |
| Share Workers in Services | - | - | - | - | x |
| Share Workers in Public Administration | - | - | - | - | - |
| **C. Farming:** | | | | | |
| Log Average Farm Value | - | - | - | - | - |
| Log Value of Farm Output per Acre | - | - | - | - | - |
| Share Owner Operated Farms | - | - | - | - | - |
| Share Farmland Cultivated | - | - | - | - | - |
| Share Wheat in Cultivated Farmland | - | - | - | - | - |
| Share Cotton in Cultivated Farmland | - | - | - | - | - |
| Share Hay/Corn in Cultivated Farmland | - | - | - | - | - |
| Region Quota Exposure - Mean | 0.109 | 0.113 | 0.111 | 0.129 | 0.066 |
| Region Quota Exposure - SD | 0.083 | 0.081 | 0.096 | 0.087 | 0.059 |
| Number of Regions | 403 | 249 | 61 | 45 | 48 |

Notes: Columns 1-5 show the variables selected by a lasso procedure of a cross-sectional specification where the dependent variable is the Relevant Labor Market (RLM) by industry exposure measure and the potential explanatory variables are a set of 1910 RLM by industry characteristics. Controls marked with an "x" are chosen by the Lasso specification.

## Table A6. Pre Border Closure Policy Trends in Wages

| **Sample:** | **Pooled Sample** | | **Manufacturing & Mining** | | **Construction Union** | | **Farm Wages** | |
|---|---|---|---|---|---|---|---|---|
| **Trend Type:** | **Linear** | **Non-Linear** | **Linear** | **Non-Linear** | **Linear** | **Non-Linear** | **Linear** | **Non-Linear** |
| | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** | **(6)** | **(7)** | **(8)** |
| Quota Exposure X Linear Trend | 0.0385 | | -0.0399 | | 0.00224 | | -0.00132 | |
| | (0.0297) | | (0.0577) | | (0.0402) | | (0.0202) | |
| Quota Exposure X Year 1911-12 | | -0.0500 | | 0.250 | | -0.0402 | | -0.381** |
| | | (0.137) | | (0.267) | | (0.117) | | (0.146) |
| Quota Exposure X Year 1913-14 | | 0.143 | | 0.403 | | 0.278 | | 0.0316 |
| | | (0.159) | | (0.320) | | (0.289) | | (0.107) |
| Quota Exposure X Year 1915-16 | | 0.257 | | 0.443 | | 0.176 | | 0.157 |
| | | (0.158) | | (0.274) | | (0.417) | | (0.172) |
| Quota Exposure X Year 1917-18 | | 0.00240 | | -0.473 | | 0.494 | | -0.0665 |
| | | (0.198) | | (0.371) | | (0.414) | | (0.165) |
| Quota Exposure X Year 1919-20 | | 0.106 | | -0.609 | | 0.899** | | -0.277 |
| | | (0.264) | | (0.541) | | (0.366) | | (0.208) |
| R-Squared | 0.710 | 0.780 | 0.673 | 0.781 | 0.578 | 0.739 | 0.767 | 0.854 |
| Region Real Wage - Mean | 0.258 | | 0.245 | | 0.397 | | 0.186 | |
| Region Quota Exposure - Mean | 0.100 | | 0.107 | | 0.152 | | 0.066 | |
| Region Quota Exposure - SD | 0.082 | | 0.079 | | 0.098 | | 0.059 | |
| Observations | 1,353 | | 550 | | 241 | | 528 | |

Notes: The table presents the results of two tests for the existence of pre-trends in log mean hourly wages in the pooled sample (columns 1-2), BLS industry surveys of manufacturing and mining industries (columns 3-4), the BLS surveys of construction unions in selected cities (columns 5-6), and USDA survey of agriculture labor wages (columns 7-8). The years included in all specifications are the pre-policy years 1910-1920. The table presents the interaction coefficient for each sample between quota exposure and a linear time trend in one specification and the coefficients between the quota exposure measure and year dummies in another specification. The omitted year is 1910. The specification also includes year dummies and quota exposure measures. Robust standard errors in parenthesis are clustered at the level of the relevant region by industry. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Table A7. Exposure to Border Closure Policy and Construction Workers' Wages

| **Sample:** | **American Contractor** | | |
|---|---|---|---|
| **Occupation (Construction Industry):** | **Laborers** | **Hod-Carriers** | **Bricklayers** |
| | **(1)** | **(2)** | **(3)** |
| **A. Foreign Born Share:** | | | |
| Overall | | | |
| Quota Exposure X Post 1920 | -0.573*** | -0.385 | -0.0119 |
| | (0.174) | (0.288) | (0.0671) |
| Pre-1920 Dependent Mean | 0.30 | 0.04 | 0.27 |
| Pre-1920 Share of Industry Employment | 0.06 | 0.0007 | 0.01 |
| Quota Restricted Countries | | | |
| Quota Exposure X Post 1920 | -0.588*** | -0.127 | -0.0286 |
| | (0.175) | (0.163) | (0.0644) |
| Pre-1920 Dependent Mean | 0.29 | 0.03 | 0.25 |
| **B. Log Mean Hourly Wage:** | | | |
| Quota Exposure X Post 1920 | 0.574* | 0.0709 | -0.0280 |
| | (0.310) | (0.276) | (0.166) |
| 1920 Mean Hourly Wage (City Level) | 0.614 | 0.757 | 1.234 |
| RLM-Industry Quota Exposure - Mean | | 0.12 | |
| RLM-Industry Quota Exposure - SD | | 0.10 | |
| Number of Cities | 61 | 56 | 61 |
| Number of States | 27 | 27 | 27 |
| Observations (City-Year Pairs) | 355 | 336 | 364 |

Notes: The table presents estimation results for the American Contractor survey sample. Column 1 shows estimates for construction laborers (low-skilled), column 2 for hod-carriers (medium-skill), and column 3 for bricklayers (skilled). The outcomes in Panel A are the foreign-born share (overall and from quota-restricted countries) for the occupation in the city. The outcome in Panel B is the (log) mean hourly wage for workers in a given occupation in a city. The table presents the coefficients of the interaction between the quota exposure measure and the post-1920 policy change indicator. Foreign-born shares are calculated as the share of foreign-born workers in the construction industry at the county level using the full count population censuses of 1910,1920, and 1930. Wage data is at the city-year level and is obtained from the American Contractor surveys of construction contractors in the 1920s. Each observation in the regression represents a city-year pair. All specifications include year and city fixed effects, in addition to all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Each city has a varying number of observations from 1920 through 1928. Robust standard errors, clustered at the RLM by industry, level in parenthesis. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Table A8. Robustness of Estimates to Aggregation: RLM Level

| **Aggregation Level:** | **RLM-Industry (1)** | **RLM (2)** |
|---|---|---|
| **A. Foreign-Born Share:** | | |
| Quota Exposure X Post 1920 | -0.260*** | -0.210*** |
| | (0.0387) | (0.0633) |
| Pre-1920 Dependent Mean | 0.280 | 0.252 |
| Quota Exposure - Mean | 0.11 | 0.107 |
| Quota Exposure - SD | 0.08 | 0.086 |
| Number of Panel Units | 403 | 128 |
| Observations | 1,209 | 307 |
| **B. Log Mean Hourly Wage:** | | |
| Quota Exposure X Post 1920 | 0.586*** | 0.454 |
| | (0.150) | (0.277) |
| Pre-1920 Dependent Mean | 0.308 | 0.339 |
| Quota Exposure - Mean | 0.102 | 0.107 |
| Quota Exposure - SD | 0.083 | 0.086 |
| Number of Panel Units | 401 | 128 |
| Observations | 3,371 | 307 |

Notes: This table presents the coefficient of the interaction between quota exposure $QE$ and the post-policy change indicator. The Post variable is an indicator for post-1920, the last year before the border closure policy was enacted. In Panel A, the dependent variable in the specifications is the foreign-born share for the Relevant Labor Market (RLM) by industry in column 1 and for the RLM, aggregated by the list of industries in Table A2, in column 2. In Panel B, the dependent variable in the specifications is the log mean hourly wage for Relevant Labor Market (RLM) by industry in column 1 and for the (employment-weighted) aggregated RLM across the same group of industries. All specifications include year and RLM by industry fixed effects, in addition to dataset- and industry-specific linear time trend and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Robust standard errors, clustered at the RLM by industry, level in parenthesis.. Robust standard errors clustered at the RLM-industry or RLM level in parenthesis. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Table A9. Testing for Potential Bias from Potential SUTVA Violation

| | | Exclude Bottom X% Quota Exposure: | | | Exclude Top X% Outmigration Rate | | |
|---|---|---|---|---|---|---|---|
| **Sample:** | **Baseline** | **X =1** | **X = 5** | **X = 10** | **X = 5** | **X = 10** | **Top state (Division)** |
| | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** | **(6)** | **(7)** |
| **A. Foreign-Born Share:** | | | | | | | |
| Quota Exposure X Post 1920 | -0.260*** | -0.263*** | -0.270*** | -0.267*** | -0.239*** | -0.232*** | -0.324*** |
| | (0.0387) | (0.0392) | (0.0399) | (0.0407) | (0.0408) | (0.0424) | (0.0442) |
| Pre-1920 Dependent Mean | 0.280 | 0.283 | 0.295 | 0.310 | 0.275 | 0.267 | 0.255 |
| Quota Exposure - Mean | 10.876 | 11.011 | 11.456 | 12.046 | 10.735 | 10.525 | 10.003 |
| Quota Exposure - SD | 8.347 | 8.311 | 8.188 | 8.006 | 8.408 | 8.353 | 7.908 |
| Number of Panel Units | 403 | 398 | 382 | 362 | 388 | 370 | 316 |
| Observations | 1,209 | 1,194 | 1,146 | 1,086 | 1,164 | 1,110 | 948 |
| **B. Log Mean Hourly Wage:** | | | | | | | |
| Quota Exposure X Post 1920 | 0.586*** | 0.564*** | 0.576*** | 0.561*** | 0.636*** | 0.598*** | 0.566*** |
| | (0.150) | (0.151) | (0.154) | (0.157) | (0.151) | (0.145) | (0.174) |
| Pre-1920 Dependent Mean | 0.258 | 0.258 | 0.266 | 0.276 | 0.256 | 0.253 | 0.248 |
| Quota Exposure - Mean | 9.961 | 9.968 | 10.523 | 11.255 | 9.894 | 9.785 | 9.140 |
| Quota Exposure - SD | 8.217 | 8.216 | 8.103 | 7.910 | 8.332 | 8.483 | 7.797 |
| Number of Panel Units | 402 | 400 | 390 | 377 | 387 | 369 | 315 |
| Observations | 2,975 | 2,962 | 2,825 | 2,676 | 2,846 | 2,699 | 2,461 |

Notes: The table presents the coefficient of the interaction between quota exposure $QE$ and the post-policy change indicator for the pooled sample. The Post variable is an indicator for post-1920, the last year before immigration restrictions started. In Panel A, the outcome variable in the specifications is the foreign-born employment share for the Relevant Labor Market (RLM) by industry pair. In Panel B, the outcome variable in the specifications is the log mean hourly wage for Relevant Labor Market (RLM) by industry pair. All specifications include year and RLM by industry fixed effects, in addition to dataset- and industry-specific linear time trend and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Columns 2-4 exclude RLM by industry pairs in the bottom 1,5 and 10 percentile of the $QE$ distribution, respectively. Columns 5-7 use Abramitzky et al. (2023)'s out-migration rate to high exposure areas, aggregated to the state level, and exclude the states in the top 5, 10, and highest in division out-migration to high exposure areas rates, respectively. In Panel A, each RLM-industry pair has three observations for the 1910, 1920, and 1930 censuses. In Panel B, each RLM-industry pair has a different number of observations from 1910 to 1930 based on pooled sample data. Robust standard errors, clustered at the RLM by industry level, in parenthesis. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Table A10. Testing Robustness for Controlling for World War I Exposure

| | (1) | (2) | (3) |
|---|---|---|---|
| **A. Foreign-Born Share:** | | | |
| Quota Exposure X Post 1920 | -0.260*** | | -0.400*** |
| | (0.0387) | | (0.0536) |
| WWI Exposure X (Year = 1920) | | -0.00544 | -0.134*** |
| | | (0.0112) | (0.0203) |
| Pre-1920 Dependent Mean | | 0.280 | |
| Quota Exposure - Mean | | 0.109 | |
| Quota Exposure - SD | | 0.083 | |
| WW1 Exposure - Mean | | 0.182 | |
| WW1 Exposure - SD | | 0.124 | |
| Observations | | 1,209 | |
| **B. Log Mean Hourly Wages:** | | | |
| Quota Exposure X Post 1920 | 0.586*** | | 0.634*** |
| | (0.150) | | (0.152) |
| WWI Exposure X (Year $\geq$ 1915 & Year $\leq$ 1918) | | -0.0377 | 0.100 |
| | | (0.0713) | (0.0628) |
| Pre-1920 Dependent Mean | | 0.258 | |
| Quota Exposure - Mean | | 9.961 | |
| Quota Exposure - SD | | 8.217 | |
| WW1 Exposure - Mean | | 16.902 | |
| WW1 Exposure - SD | | 12.454 | |
| Observations | | 2,975 | |

Notes: The table presents the coefficient of the interaction between quota exposure $QE$ and the post-policy change indicator. The Post variable is an indicator for post-1920, the last year before immigration restrictions started. In Panel A, the dependent variable is the foreign-born employment share for the Relevant Labor Market (RLM) by industry pair. In Panel B, the dependent variable is the log mean hourly wage for Relevant Labor Market (RLM) by industry pair. All specifications include year and RLM by industry fixed effects, in addition to dataset- and industry-specific linear time trend and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. Column 1 presents our baseline estimates. Column 2 uses the WWI exposure measure instead of our preferred quota exposure measure and interacts it with the war years (1920 for the decadal census and 1915-1918 for the wage sample). Column 3 includes both quota exposure and WWI exposure measures and their interactions. In Panel A, each RLM-industry pair has three observations for the 1910, 1920, and 1930 censuses. In Panel B, each RLM-industry pair has a different number of observations from 1910 to 1930 based on pooled sample data. Robust standard errors, clustered at the RLM-industry level, in parenthesis. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Table A11. Robustness to Alternative Quota Exposure Measures

| | 1910 Census | | 1900 Census | |
|---|---|---|---|---|
| **Exposure Measure:** | **Baseline** | **Binary Exposure** | **Baseline** | **Binary Exposure** |
| | **(1)** | **(2)** | **(3)** | **(4)** |
| **A. Foreign-Born Share:** | | | | |
| Quota Exposure X Post 1920 | -0.260*** | -0.229*** | -0.333*** | -0.276*** |
| | (0.0387) | (0.0380) | (0.0646) | (0.0650) |
| Pre-1920 Dependent Mean | 0.280 | 0.280 | 0.280 | 0.280 |
| Quota Exposure - Mean | 10.876 | 10.778 | 6.179 | 5.742 |
| Quota Exposure - SD | 8.347 | 8.549 | 5.437 | 5.366 |
| Number of Panel Units | 403 | 403 | 401 | 401 |
| Observations | 1,209 | 1,209 | 1,203 | 1,203 |
| **B. Log Mean Hourly Wage:** | | | | |
| Quota Exposure X Post 1920 | 0.586*** | 0.599*** | 0.535** | 0.582** |
| | (0.150) | (0.145) | (0.226) | (0.232) |
| Pre-1920 Dependent Mean | 0.258 | 0.258 | 0.258 | 0.258 |
| Quota Exposure - Mean | 9.961 | 9.832 | 5.850 | 5.409 |
| Quota Exposure - SD | 8.217 | 8.390 | 5.361 | 5.255 |
| Number of Panel Units | 402 | 402 | 400 | 400 |
| Observations | 2,975 | 2,975 | 2,967 | 2,967 |

Notes: The table presents the coefficient of the interaction of four different quota exposure measures and the post-policy change indicator for the labor market by industry foreign-born share (Panel A) and log mean hourly wage (Panel B). The Post variable is an indicator for post-1920, the last year before immigration restrictions started. All specifications include year and RLM by industry fixed effects, in addition to dataset- and industry-specific linear time trend and all of the RLM-industry level controls listed in Table A5, interacted with a linear time trend. The quota exposure measure in column 1 is from our preferred specification, based on continuous country-level restriction measures and the 1910 foreign-born shares. The exposure measure in column 2 uses binary instead of continuous country-level exposure measures. Columns 3 and 4 use 1900 foreign-born shares instead of 1910 foreign-born shares. Robust standard errors, clustered at the RLM-industry level, in parenthesis. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.