For Online Publication

Supplemental Appendix to

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

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A Data Sources, Sample Construction, and Exposure Measures

This appendix provides additional detail on the historical sources, data construction, and exposure measures used in the analysis.

A.1 Quota Intensity and War Restrictions

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). Notes: The quota intensity measure for a region equals the predicted unrestricted 1920s inflow minus the number of quota slots, divided by the predicted unrestricted inflow. The WWI restriction measure is constructed in a similar way using wartime disruptions.

Exposure measures to immigration restrictions used in the analysis are constructed using quota intensity indices QI_r that capture how strongly the 1920s quota system restricted immigration from different sending regions. These indices come from Abramitzky et al. (2023), who estimate counterfactual "unrestricted" immigration flows for each region using a pre-1914 fitted polynomial on immigration flows from each sending region and then compare them to the number of quota slots allocated under the new legislation.

Table A1 reports the quota intensity and World War I (WWI) restriction indices by broad

sending region. The quota intensity for region r is defined as:

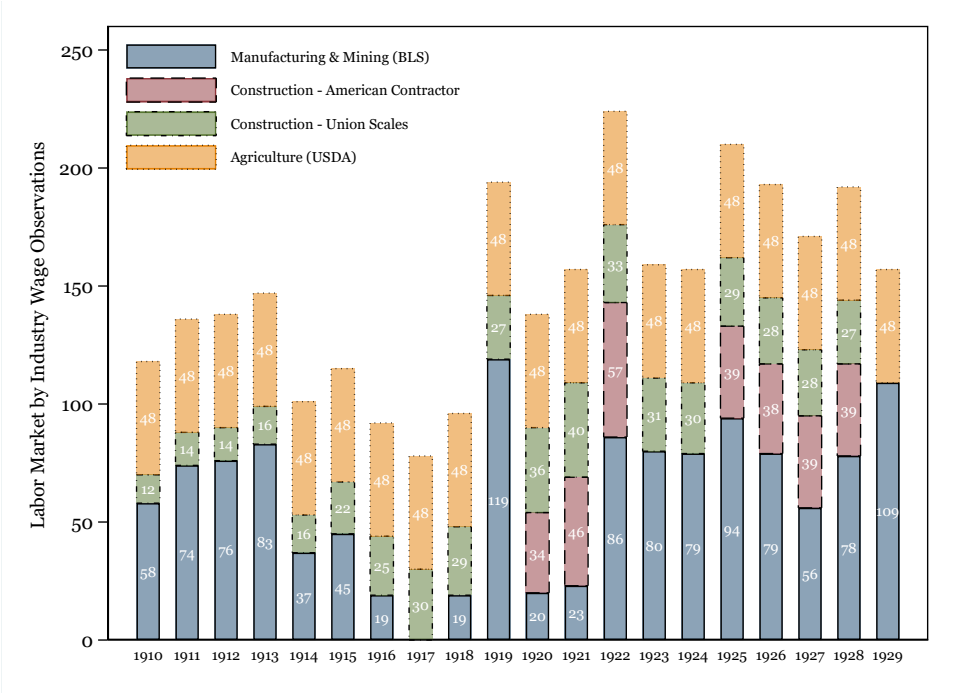
$$QI_r = \frac{\text{unrestricted flow}_r - \text{quota slots}_r}{\text{unrestricted flow}_r},$$

so that $QI_r = 0$ when quotas do not bind and $QI_r = 1$ when flows are fully shut down.

We similarly construct a WWI exposure measure w_c using the WWI restriction indices in column 2 of Table A1. This measure is highly correlated with q_c (correlation coefficient 0.87) and is used in robustness checks that control explicitly for wartime disruptions to migration flows.

A.2 Archival Sources and Coverage

Figure A1. Wage Observations by Labor Market–Industry Dataset, 1910–1929



Notes: The figure reports the number of labor market–industry wage observations per year by dataset. Manufacturing & mining (BLS) are from employer payrolls; construction (American Contractor) from contractor reports of city-level wages; construction (BLS union scales) from negotiated minimums; agriculture (USDA) from state daily/monthly wages converted to hourly (10 hrs/day).

We assemble annual wage data for low-skilled workers in the United States between 1910 and 1929 from four main historical sources: (i) industry-specific surveys of manufacturing and mining wages conducted by the U.S. Bureau of Labor Statistics (BLS), (ii) wage reports from the trade journal *American Contractor*, (iii) BLS surveys of construction trade union wage scales, and (iv) surveys of agricultural wages conducted by the U.S. Department of Agriculture (USDA). These sources provide repeated observations on narrowly defined low-skilled occupations—typically labeled "laborers"—for a variety of local labor markets and industries.

The basic description of what we call our pooled sample is found in the text. To show how the

number of wage observations in this sample evolves over time and across sources, Figure A1 plots the annual number of wage observations by dataset for 1910–1929. Coverage is relatively sparse before World War I, increases during the 1910s, and peaks in the mid- to late 1920s, reflecting the expansion of BLS reporting and the introduction of the *American Contractor* survey.

Manufacturing and Mining Wage Surveys. The BLS industry surveys report mean wages for detailed occupations within individual manufacturing and mining industries. The unit of observation is either a state–industry or a city–industry cell, depending on how BLS summarized the underlying establishment-level returns in a given year.

We focus on the lowest-paid occupation reported in each industry, which is usually labelled "laborer". For a small number of industries where no explicit "laborer" occupation appears, we instead select a large occupation at the bottom of the wage distribution (for example, doffers in cotton textiles). In coal mining we retain three distinct low-skilled occupations—pick miners, inside-mine laborers, and outside-mine laborers. Table A2 lists the industries, occupations, and years covered in the manufacturing and mining sample, together with information on whether semi-skilled wage data are also available for the industry.

Construction Wages: Contractors and Unions. Our data on the construction sector combines information from the journal *American Contractor* with BLS surveys of union wage scales in construction. From *American Contractor* we use reports for October 1920, October 1921, May 1922, October 1922, June 1925, October 1925, May 1926, September 1926, April 1927, July 1927, May 1928, and October 1928. When a city appears more than once in a calendar year, we keep the highest reported wage so that each city–year cell corresponds to a single wage. This yields 294 wage observations in 61 cities. The BLS union wage scale surveys report information for cities where at least half of workers in an occupation were covered by a union agreement. Included are the hourly union scales for "building laborers." These union scales represent minimum pay rates and may understate actual earnings when workers receive premia above the contractual minimum. Nonetheless, they provide broad geographic and temporal coverage of low-skilled construction wages. Table A3 lists the cities covered by each construction source and flags whether a city appears in the wage survey, in the union scale data, or in both. Many large cities appear in both sources, while a set of smaller towns appear only in the contractor survey. For both construction datasets, wages are reported at the city level. Therefore, the RLM for each observation is defined as the county in which the city is located, and exposure is measured using the county-level quota exposure index q_c defined in Equation 7 in the main text.

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	1919/22/25/28	Yes
Boxes: laborers	457	17	11	1919/25	
Cement: laborers	317	4	4	1929	
Coal: In-mine laborers	216	48	10	1919/22/24/26/29	Yes
Coal: pick miners	216	43	9	1919/22/24/26/29	Yes
Coal: Outside of mine laborers	216	48	10	1919/22/24/26/29	Yes
Cotton Textiles: doffers	439	99	12	1911–14/16/18/22/24/26/28	Yes
Foundries: laborers	336–338, 346–348	123	28	1919/23/25/27/29	Yes
Furniture: laborers	309	41	17	1915/19/29	
Hosiery & Underwear: "other occ."	436, 446, 449	73	16	1910–14/22/26/28	
Iron Mines: laborers, outside mine	206	3	3	1924	
Machine Shops: laborers	356–358, 367	123	28	1919/23/25/27/29	Yes
Meat Packing: laborers (maint. & repair)	406	5	1	1921/23/25/27/29	Yes
Men's clothing: basters (coat)	448	59	7	1911–14/19/22/24/26/28	Yes
Metal Mining: laborers (outside mine)	206	8	8	1924	
Mill Work (Wood): laborers	307–308	66	12	1910–13/15/19	
Quarries: laborers	236	4	4	1929	
Railcar Manufacture: laborers	379	63	17	1910–13	
Sawmills: laborers	307–308	211	23	1910–13/15/19/21/23/25/28	Yes
Shoes: cutters, trimming hand	488	81	12	1910–14/20/22/24/26/28	Yes
Woolen Textile: dyehouse laborers	439	87	9	1910–14/16/18/20/22/24/26/28	Yes
B. Construction:					
American Contractor Dataset	246	294	61	1920–22, 25–28	Yes
BLS Construction Union Wage Scales	246	487	45	1910–28	
C. Agriculture:					
USDA Farm Laborer Wage Dataset	105	960	48	1910–29	
Observations				2,973	
Number of Relevant Labor Markets (RLMs)				125	
Number of Industries				18	
Number of RLM–Industry Pairs				399	

Notes: The table lists industries and occupations in the main panel, grouped by dataset. Column (4) reports years with wage observations (with subsequent years abbreviated). The last rows summarize the size of the pooled panel.

Table A3. American Contractor Wage Survey and Construction Union Wage Scales: City 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: "Wage Survey" refers to the *American Contractor* construction wage survey (1920–22, 1925–28); "Union Wage" refers to the BLS survey of construction trade union wage scales (1910–28). YF = city observed in all contractor survey years; YP = city observed in some contractor survey years; N = no observations in that source.

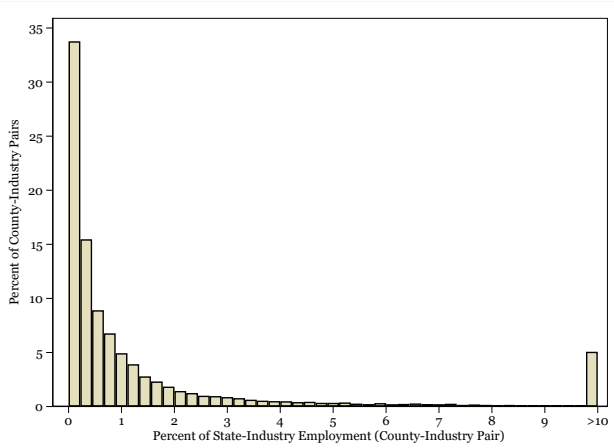
Agricultural Wages. Agricultural wages come from USDA publications that report average daily and monthly wages for farm laborers without board by state (USDA,1943). We use all 48 contiguous states and convert daily wages into hourly rates by assuming a 10-hour workday, which is consistent with contemporary accounts of farm work hours. The resulting series span 1910–1929 and form a balanced panel of 960 state–year observations.

Municipal Laborers (1928). The BLS municipal wage survey in 1928 collected information on the hourly wage paid to entry-level street workers—typically unskilled laborers employed to clean or repair streets—in more than 2,600 municipalities. In most cases, a single wage is reported per city; when multiple rates appear or a range is given, we use the lowest wage to capture the pay of the least experienced workers.

1940 Census Wages and Demographics. We also use the 1900–1940 Population Censuses to construct measures of long-run wages and local labor market characteristics at the RLM–industry level (Ruggles et al., 2024). For 1940 we compute hourly wages for low-skilled workers in each RLM–industry pair by dividing annual earnings by weeks and hours worked. We also use information on foreign-born shares and other demographic and socioeconomic variables from ICPSR's Historical, Demographic, Economic, and Social Data of the United States (Haines et al., 2010).

A.3 Relevant Labor Markets (RLMs) and Exposure to Immigration Restrictions

Figure A2. Distribution of County–Industry Employment Shares



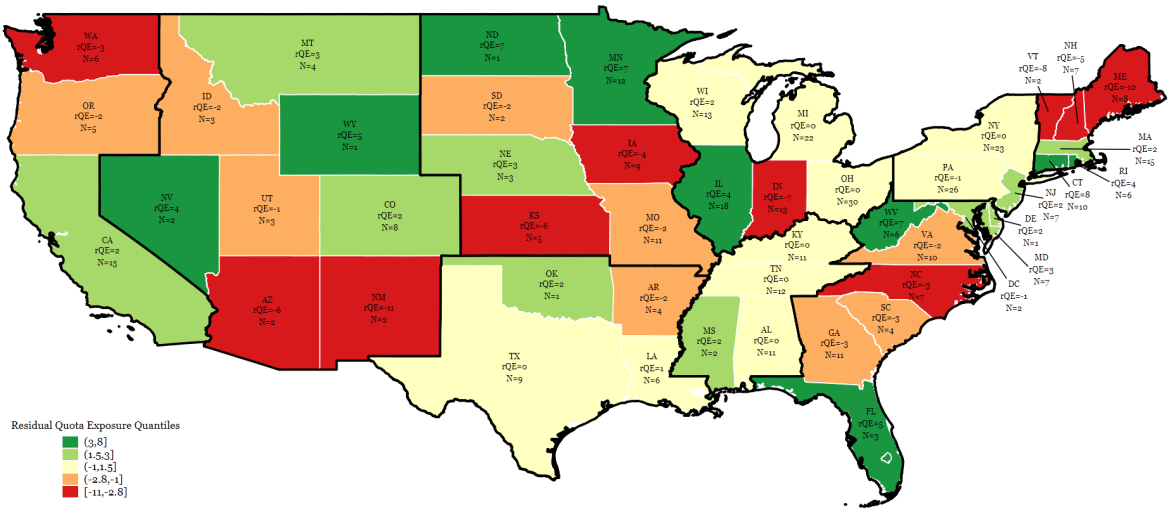
Notes: The figure shows the distribution of county–industry 1910 employment shares across all state–industry pairs in the sample. The x-axis reports the percent of total state–industry employment accounted for by each county–industry pair, while the y-axis shows the percent of such pairs in each bin. Most county–industry pairs represent less than 1% of total state–industry employment.

A central empirical challenge is to map each wage observation onto the correct local labor market — the geographic area that supplied the low-skilled workers whose wages the observation reflects. As described in the text, we do this by defining the *Relevant Labor Market* (RLM) for a given wage observation w_{ijt} as the geographic area that supplies low-skilled labor to the establishments that underlie the BLS mean wage for industry i in jurisdiction j at time t. Where the BLS reports wages at the city level, the RLM is the county in which that city is located; where BLS reports are at the state–industry level, the RLM aggregates county conditions using a set of industry employment weights described in the main text.

Figure A2 provides empirical justification for this choice: the figure plots the within-state distribution of the 1910 industry employment across counties for the state-industry pairs in our sample and shows the geographic distribution employment was highly uneven: a small number of counties typically accounted for a large share of total industry employment.

To illustrate the variation underlying the exposure measure, Figure A3 maps residual quota exposure after removing industry and census-division fixed effects. This residualized measure isolates differences across RLM–industry pairs that are not explained by broad industry composition or by census-division-level differences. The figure shows that meaningful geographic variation remains after netting out these components, with some states containing RLM–industry pairs that are more or less exposed than predicted by their industry mix and census division. This residual variation is the cross-sectional variation in exposure used in the regression analysis after accounting for the main industrial and regional fixed effects.

Figure A3. Residual Geographic Variation in Quota Exposure



Notes: This figure maps residual quota exposure after removing industry and census-division fixed effects from the RLM–industry exposure measure. State shading reflects the mean residualized exposure across RLM–industry pairs located in each state, measured in percentage points. State labels report mean residual quota exposure, rQE, and the number of RLM–industry pairs, N. Negative values indicate that RLM–industry pairs in the state are less exposed than predicted by their industry mix and census division; positive values indicate greater residual exposure. Bold lines delineate census divisions.

Table A4. Residual Quota Exposure and Geographic Coverage by Data Source

	Raw quota exposure		Residual quota exposure			Geographic coverage			
Sample	Mean QE	SD QE	SD rQE	P10 rQE	P90 rQE	RLMs	States	Divisions	Pairs
Manufacturing & Mining	11.30	8.10	4.96	-5.91	6.89	248	42	9	248
American Contractor	11.41	9.79	8.58	-8.81	14.89	58	27	8	58
Construction Union	12.84	8.75	6.76	-7.60	10.12	45	28	9	45
Agriculture	6.60	5.95	4.12	-5.47	5.95	48	48	9	48
Pooled Sample	10.93	8.36	5.73	-6.89	6.98	399	49	9	399

Notes: The table reports raw and residual quota exposure by data source. Mean QE and SD QE are based on the raw RLM–industry quota exposure measure and are reported in percentage points. Residual quota exposure, rQE, is obtained by residualizing the RLM–industry exposure measure on industry, census-division, and dataset fixed effects in the pooled sample, and is reported in percentage points. The table reports the standard deviation and 10th and 90th percentiles of rQE separately by data source. Geographic coverage reports the number of relevant labor markets, states, census divisions, and RLM–industry pairs represented in each sample.

Table A4 complements Figure A3 by summarizing the amount of exposure variation that remains within each wage sample after removing broad industrial, regional, and source-level differences. The table reports the mean and dispersion of raw quota exposure, the dispersion of residual quota exposure after residualizing on industry, census-division, and dataset fixed effects, and the geographic coverage of each source. This exercise shows that residual exposure variation is present in each data source and is not confined to a narrow set of states or census divisions.

To gauge how the exposure measure relates to the underlying composition of immigrant populations across local labor markets, we compute the 1910 foreign-born population share and its breakdown by origin region for each RLM and then group RLM–industry pairs into terciles of the exposure distribution. Table A5 summarizes these patterns. High-exposure RLM–industry pairs had both higher overall foreign-born shares and a much larger fraction of immigrants originating from high-restriction countries in Southern and Eastern Europe. In contrast, low-exposure markets had smaller foreign-born shares and immigrants more heavily drawn from low- or non-restriction origins such as Germany, Ireland, Scandinavia, and Canada.

Table A5. Composition of Foreign-Born Population by Exposure to Immigration Restrictions

Quota Exposure:	**Full Sample** (1)	**Bottom Tercile** (2)	**Middle Tercile** (3)	**Top Tercile** (4)
Number of RLM-Industry Pairs	399	133	133	133
Quota Exposure Measure (RLM-Industry):	0.109	0.022	0.101	0.204
Foreign-Born Share in Population (1910):	0.171	0.055	0.177	0.282
A. High-Restriction Countries:	34%	28%	31%	43%
Asia	2.3%	2.9%	1.6%	2.3%
Central Europe	9.8%	5.2%	10.5%	13.5%
Eastern Europe	0.4%	0.3%	0.5%	0.4%
Greece	1.2%	1.8%	1.0%	0.8%
Italy	8.6%	6.7%	7.3%	11.7%
Portugal	0.3%	0.1%	0.1%	0.7%
Rest of World	1.2%	1.4%	1.0%	1.0%
Russia	10.1%	9.1%	9.1%	12.2%
Spain	0.2%	0.2%	0.1%	0.3%
B. Low-Restriction Countries:	53%	57%	56%	47%
Germany	20.9%	22.9%	24.5%	15.2%
Ireland	8.9%	8.4%	7.2%	11.2%
Scandinavia	7.3%	4.9%	9.5%	7.6%
United Kingdom	12.2%	15.9%	10.1%	10.5%
Western Europe	3.9%	4.7%	4.5%	2.5%
C. Non-Restriction Countries:	13%	16%	13%	10%
Canada	10.7%	11.7%	11.5%	8.8%
Caribbean	0.4%	0.5%	0.1%	0.5%
Latin America	0.1%	0.1%	0.0%	0.0%
Mexico	1.7%	3.2%	1.2%	0.7%

Notes: The table reports the composition of the foreign-born population across RLM–industry pairs in 1910, grouped by terciles of the quota exposure measure QE_{ij}. For each group we show the mean quota exposure, the overall foreign-born share in the population, and the distribution of foreign-born by origin region. Percentages in Panels A–C sum to 100 within each column and refer to shares of the foreign-born population.

B Theoretical Framework and Derivations

This appendix presents more details on the formal model used to interpret the empirical results in the main text. The framework is partial equilibrium: local markets take the rental rate of capital and product prices as given, and capital and labor responses are modeled at the industry–labor market-cell level. A full general-equilibrium model with endogenous national factor prices, product markets, and complete inter-regional migration would require additional assumptions that extend beyond the scope of our empirical analysis. Because our empirical strategy exploits variation in exposure across industries and local labor markets within a national economy, the partial-equilibrium framework provides an appropriate basis for interpreting the wage effects we estimate.

B.1 Economic Environment

Consider local labor market–industry cells indexed by (i, j). Output is produced using capital K_{ij}, low-skilled labor $L_{1,ij}$, and skilled labor $L_{2,ij}$:

$$Q_{ij} = \left[(1-\alpha_{ij})K_{ij}^{\delta_{ij}} + \alpha_{ij}L_{ij}^{\delta_{ij}}\right]^{1/\delta_{ij}}, \qquad L_{ij} = \left(\theta_{1,ij}L_{1,ij}^{\beta_{ij}} + \theta_{2,ij}L_{2,ij}^{\beta_{ij}}\right)^{1/\beta_{ij}}, \tag{1}$$

where industry parameters $\alpha_i \in (0,1)$ and $\theta_{1,i}, \theta_{2,i} > 0$ govern factor shares and skill weights. The elasticities of substitution are

$$\sigma_{KL,ij} = \frac{1}{1-\delta_{ij}}, \qquad \sigma_{12,ij} = \frac{1}{1-\beta_{ij}}.$$

Firms are perfectly competitive and maximize profits

$$\Pi_{ij} = Q_{ij} - w_{1,ij}L_{1,ij} - w_{2,ij}L_{2,ij} - rK_{ij}, \tag{2}$$

implying standard factor-pricing first-order conditions:

$$w_{k,ij} = \frac{\partial Q_{ij}}{\partial L_{k,ij}}, \quad k = 1, 2, \qquad r = \frac{\partial Q_{ij}}{\partial K_{ij}}. \tag{3}$$

We treat the rental rate of capital r as fixed to reflect access to the national capital market.

B.2 Cost Shares and Log-Linearized Factor Demand

Let total factor payments be

$$Y_{ij} = w_{1,ij}L_{1,ij} + w_{2,ij}L_{2,ij} + rK_{ij}. \tag{4}$$

Define cost shares:

$$S_{K,ij} = \frac{rK_{ij}}{Y_{ij}}, \qquad S_{L,ij} = 1 - S_{K,ij}, \qquad S_{1,ij} = \frac{w_{1,ij}L_{1,ij}}{w_{1,ij}L_{1,ij} + w_{2,ij}L_{2,ij}}, \quad S_{2,ij} = S_{L,ij} - S_{1,ij}. \tag{5}$$

A log-linear approximation to the first-order conditions around the initial equilibrium yields

$$d\ln w_{k,ij} = \varepsilon_{k1,ij}\, d\ln L_{1,ij} + \varepsilon_{k2,ij}\, d\ln L_{2,ij} + \varepsilon_{kK,ij}\, d\ln K_{ij}, \quad k = 1, 2, \tag{6}$$

where the coefficients depend on $\sigma_{KL,ij}, \sigma_{12,ij}$ and the cost shares $S_{K,ij}, S_{L,ij}, S_{1,ij}, S_{2,ij}$. These parameters vary across industries, generating heterogeneous wage responses to a common shock.

B.3 Short-Run Effects (Fixed Capital)

In the short run, the capital stock is fixed:

$$d\ln K_{ij} = 0. \tag{7}$$

Differentiation of (3) yields, for low-skilled wages,

$$\begin{aligned} d\ln w_{1,ij} = {} & \frac{1}{\sigma_{KL,ij}\sigma_{12,ij}} \left[\sigma_{12,ij} \Big((S_{L,ij} - 1) \tfrac{S_{1,ij}}{S_{L,ij}} \Big) + \sigma_{KL,i} \Big(\tfrac{S_{1,ij}}{S_{L,ij}} - 1 \Big) \right] d\ln L_{1,ij} \\ & + \frac{1}{\sigma_{KL,ij}\sigma_{12,ij}} \left[\sigma_{12,ij} \Big((S_{L,ij} - 1) \tfrac{S_{2,ij}}{S_{L,ij}} \Big) + \sigma_{KL,ij} \tfrac{S_{2,ij}}{S_{L,ij}} \right] d\ln L_{2,ij}, \end{aligned} \tag{8}$$

and an analogous expression for $d\ln w_{2,ij}$.

The relative wage response is independent of capital–labor substitution:

$$d\ln w_{1,ij} - d\ln w_{2,ij} = \frac{1}{\sigma_{12,ij}} \big(d\ln L_{2,ij} - d\ln L_{1,ij} \big). \tag{9}$$

A larger proportional decline in low-skilled labor ($d\ln L_{1,ij} < d\ln L_{2,ij}$) implies an unambiguous increase in $w_{1,ij}/w_{2,ij}$, with magnitude governed by $\sigma_{12,ij}$.

B.4 Medium-Run Effects: Partial Adjustment of Capital

Over the medium run, firms adjust capital in response to changes in labor supply. We model partial adjustment as

$$d\ln K_{ij}^{(m)} = \gamma_{ij}\, d\ln K_{ij}^{*}, \qquad 0 \le \gamma_{ij} \le 1, \tag{10}$$

where $d\ln K_{ij}^{*}$ is the long-run desired change in capital, and γ_{ij} measures the degree of adjustment. Industry differences in capital intensity, replacement costs, and the feasibility of shifting technologies imply heterogeneity in γ_{ij}.

Substituting (10) into the log-linear demand system yields the own-wage elasticity in the medium run:

$$\left. \frac{\partial \ln w_{k,ij}}{\partial \ln L_{k,ij}} \right|_m = \frac{1}{\sigma_{KL,ij}\sigma_{12,ij}} \left[\sigma_{12,ij} \left((S_{K,ij}\gamma_{ij} + S_{L,ij} - 1) \frac{S_{k,ij}}{S_{L,ij}} \right) + \sigma_{KL,ij} \left(\frac{S_{k,ij}}{S_{L,ij}} - 1 \right) \right], \quad k = 1, 2. \tag{11}$$

When $\gamma_{ij} = 0$, this reduces to the short-run elasticity. As γ_{ij} increases, the term $(S_{K,ij}\gamma_{ij} + S_{L,ij} - 1)$ increases, making the own-wage elasticity less negative and attenuating the wage response to a given contraction in labor supply. Industries with slow capital adjustment (small γ_{ij}) thus exhibit larger and more persistent wage increases.

B.5 Labor Mobility

Local labor supply may respond to wage changes through internal migration or labor-force entry. We model labor-supply responses as

$$d \ln L_{k,ij} = \pi_{k,ij} + \rho_{k,ij}\, d \ln w_{k,ij}, \qquad k = 1, 2, \tag{12}$$

where $\pi_{k,ij} < 0$ represents the exogenous decline in labor supply due to immigration restrictions, and $\rho_{k,ij}$ is the semi-elasticity of local labor supply with respect to wages.

To make the interaction between demand and supply explicit, it is useful to write the short-run demand system in a compact form. Let

$$d \ln w_{1,ij} = a_{11,ij}\, d \ln L_{1,ij} + a_{12,ij}\, d \ln L_{2,ij}, \qquad d \ln w_{2,ij} = a_{21,ij}\, d \ln L_{1,ij} + a_{22,ij}\, d \ln L_{2,ij}, \tag{13}$$

where the $a_{kl,ij}$ are the relevant elasticities implied by (8) and its analogue for $w_{2,ij}$.[1]

Case 1: Only Low-Skilled Labor Is Mobile. To illustrate the mechanism clearly, consider first the case in which only low-skilled labor responds to wages (i.e., $\rho_{1,ij} > 0$ and $\rho_{2,ij} = 0$). Then:

$$d \ln L_{1,ij} = \pi_{1,ij} + \rho_{1,ij}\, d \ln w_{1,ij}, \qquad d \ln L_{2,ij} = \pi_{2,ij}. \tag{14}$$

Substituting (14) into the first equation of (13) gives

$$d \ln w_{1,ij} = a_{11,ij}(\pi_{1,ij} + \rho_{1,ij} d \ln w_{1,ij}) + a_{12,ij}\pi_{2,ij}.$$

Rearranging,

$$(1 - a_{11,ij}\rho_{1,ij})\, d \ln w_{1,ij} = a_{11,ij}\pi_{1,ij} + a_{12,ij}\pi_{2,ij}.$$

Hence the equilibrium wage response for low-skilled workers is

$$d \ln w_{1,ij} = \frac{a_{11,ij}\pi_{1,ij} + a_{12,ij}\pi_{2,ij}}{1 - a_{11,ij}\rho_{1,ij}}. \tag{15}$$

The denominator $(1 - a_{11,ij}\rho_{1,ij})$ captures the amplifying or dampening role of labor mobility. Since $a_{11,ij} < 0$ (downward-sloping labor demand) and $\rho_{1,ij} \geq 0$, the denominator is larger than one in

[1] For example, $a_{11,ij}$ is the short-run own-wage elasticity of low-skilled labor demand with respect to low-skilled labor in cell (i, j), and $a_{12,ij}$ is the cross-elasticity with respect to skilled labor. Their exact expressions in terms of $\sigma_{KL,ij}, \sigma_{12,ij}$ and cost shares follow from (8).

absolute value, implying

$$|d \ln w_{1,ij}| \leq |a_{11,ij}\pi_{1,ij} + a_{12,ij}\pi_{2,ij}|$$

with strict inequality when $\rho_{1,ij} > 0$. When labor is immobile ($\rho_{1,ij} = 0$), (15) collapses to the pure demand-side response. As $\rho_{1,ij}$ increases, migration and labor-force entry offset the initial contraction, reducing the local wage effect.

The response of skilled wages follows from the second equation of (13):

$$\begin{aligned} d \ln w_{2,ij} &= a_{21,ij}(\pi_{1,ij} + \rho_{1,ij} d \ln w_{1,ij}) + a_{22,ij}\pi_{2,ij} \\ &= a_{21,ij}\pi_{1,ij} + a_{22,ij}\pi_{2,ij} + a_{21,ij}\rho_{1,ij}\frac{a_{11,ij}\pi_{1,ij} + a_{12,ij}\pi_{2,ij}}{1 - a_{11,ij}\rho_{1,ij}}. \end{aligned} \tag{16}$$

Thus both $d \ln w_{1,ij}$ and $d \ln w_{2,ij}$ are attenuated by labor mobility, but the magnitude of the effect depends on the interaction of own- and cross-price elasticities.

Case 2: Both Skill Groups Are Mobile. In the more general case, let $d \ln L_{ij}$ and $d \ln w_{ij}$ denote the 2×1 vectors of quantity and wage changes. Combining the demand system (13) and supply responses (12) yields

$$d \ln w_{ij} = A_{ij}\, d \ln L_{ij}, \qquad d \ln L_{ij} = \pi_{ij} + R_{ij}\, d \ln w_{ij}, \tag{17}$$

where

$$A_{ij} = \begin{bmatrix} a_{11,ij} & a_{12,ij} \\ a_{21,ij} & a_{22,ij} \end{bmatrix}, \qquad R_{ij} = \begin{bmatrix} \rho_{1,ij} & 0 \\ 0 & \rho_{2,ij} \end{bmatrix}, \qquad \pi_{ij} = \begin{bmatrix} \pi_{1,ij} \\ \pi_{2,ij} \end{bmatrix}.$$

Substituting and solving,

$$d \ln w_{ij} = A_{ij}\big(\pi_{ij} + R_{ij} d \ln w_{ij}\big) \quad \Rightarrow \quad (I - A_{ij}R_{ij})\, d \ln w_{ij} = A_{ij}\pi_{ij}. \tag{18}$$

Provided $(I - A_{ij}R_{ij})$ is nonsingular, the equilibrium wage response is

$$d \ln w_{ij} = (I - A_{ij}R_{ij})^{-1} A_{ij}\pi_{ij}. \tag{19}$$

This expression makes clear that labor mobility attenuates wage responses: when $R_{ij} = 0$ (no mobility), $d \ln w_{ij} = A_{ij}\pi_{ij}$; as R_{ij} increases, the matrix $(I - A_{ij}R_{ij})^{-1}$ moves wage responses closer to zero by making local labor supply more elastic.

Historical evidence indicates that labor mobility among low-skilled workers in the early twentieth century was limited, implying relatively small $\rho_{k,ij}$ and thus limited attenuation of the wage effects. This is consistent with the persistent wage differentials we document and with the internal migration patterns shown in Figure 14 in the main text and Table C5.

B.6 Long-Run Equilibrium

In the long run, capital and labor have fully adjusted. Capital stocks reach their desired levels and internal migration arbitrages most purely local wage differences. However, the immigration quotas permanently reduce the relative supply of low-skilled labor, which affects the long-run skill premium.

The key object is the labor nest in (1):

$$L_{ij} = \left(\theta_{1,ij} L_{1,ij}^{\beta_{ij}} + \theta_{2,ij} L_{2,ij}^{\beta_{ij}}\right)^{1/\beta_{ij}}.$$

Because Q_{ij} depends on $L_{1,ij}$ and $L_{2,ij}$ only through L_{ij}, the ratio of wages equals the ratio of marginal products of the labor aggregate:

$$\frac{w_{1,ij}}{w_{2,ij}} = \frac{\partial Q_{ij}/\partial L_{1,ij}}{\partial Q_{ij}/\partial L_{2,ij}} = \frac{\partial L_{ij}/\partial L_{1,ij}}{\partial L_{ij}/\partial L_{2,ij}}.$$

Differentiating L_{ij},

$$\frac{\partial L_{ij}}{\partial L_{1,ij}} = \theta_{1,ij} L_{1,ij}^{\beta_{ij}-1} L_{ij}^{1-\beta_{ij}}, \qquad \frac{\partial L_{ij}}{\partial L_{2,ij}} = \theta_{2,ij} L_{2,ij}^{\beta_{ij}-1} L_{ij}^{1-\beta_{ij}}.$$

Thus

$$\frac{w_{1,ij}}{w_{2,ij}} = \frac{\theta_{1,ij} L_{1,ij}^{\beta_{ij}-1}}{\theta_{2,ij} L_{2,ij}^{\beta_{ij}-1}} = \left(\frac{\theta_{1,ij}}{\theta_{2,ij}}\right)\left(\frac{L_{2,ij}}{L_{1,ij}}\right)^{1-\beta_{ij}}. \tag{20}$$

Using the relationship $\sigma_{12,ij} = 1/(1-\beta_{ij})$, we have $1-\beta_{ij} = 1/\sigma_{12,ij}$, so

$$\frac{w_{1,ij}}{w_{2,ij}} = \left(\frac{\theta_{1,ij}}{\theta_{2,ij}}\right)\left(\frac{L_{2,ij}}{L_{1,ij}}\right)^{1/\sigma_{12,ij}}. \tag{21}$$

Let $(\bar{w}_{1,ij}, \bar{w}_{2,ij})$ and $(\bar{L}_{1,ij}, \bar{L}_{2,ij})$ denote long-run equilibrium wages and quantities. Then

$$\frac{\bar{w}_{1,ij}}{\bar{w}_{2,ij}} = \left(\frac{\theta_{1,ij}}{\theta_{2,ij}}\right)\left(\frac{\bar{L}_{2,ij}}{\bar{L}_{1,ij}}\right)^{1/\sigma_{12,ij}}. \tag{22}$$

This expression shows that, even after full adjustment of capital and labor, the steady-state skill premium depends on the long-run relative supplies of low-skilled and skilled labor. If immigration restrictions permanently reduce $\bar{L}_{1,ij}$ relative to $\bar{L}_{2,ij}$, the ratio $(\bar{L}_{2,ij}/\bar{L}_{1,ij})$ rises, and the skill premium $\bar{w}_{2,ij}/\bar{w}_{1,ij}$ falls correspondingly.[2]

[2]The exact magnitude depends on the elasticity of substitution $\sigma_{12,ij}$. A higher $\sigma_{12,ij}$ (greater substitutability) implies a weaker effect of relative quantities on relative wages.

B.7 Industry Heterogeneity

All elasticities depend on $\sigma_{KL,ij}, \sigma_{12,ij}$ and the cost shares $(S_{K,ij}, S_{L,ij}, S_{1,ij}, S_{2,ij})$, which vary by industry and region. Two implications are central for our empirical results:

1. Sectors with high capital intensity and large fixed-capital stocks face higher adjustment costs and lower γ_{ij}, generating larger and more persistent wage effects (e.g., manufacturing and mining).

2. Sectors with modular capital or readily available mechanization paths adjust faster (higher γ_{ij}), yielding more muted wage responses (e.g., certain agricultural and light-industrial activities).

B.8 Summary

The model yields three core results:

1. In the short run, a disproportionate contraction in low-skilled labor raises low-skilled wages and compresses the skill premium, with the magnitude governed by $\sigma_{12,ij}$.

2. In the medium run, capital adjustment and limited labor mobility attenuate wage responses, with slower capital adjustment (lower γ_{ij}) producing larger and more persistent wage increases in some industries.

3. In the long run, the permanent reduction in low-skilled labor leaves a lasting imprint on the wage structure through (22), even after full adjustment of capital and labor.

C Additional Results

Additional Results - Immigration and Employment.

Table C1. Exposure to Border Closure Policy and Employment - by Skill Level

Sample: Pooled	**Overall** (1)	**Laborers** (2)	**Other Low-Skill** (3)	**Medium-Skill** (4)	**High-Skill** (5)
A. Recent Immigrants Share:					
Quota Exposure X Post 1920	-0.495***	-0.798***	-0.381***	-0.222***	-0.134***
	(0.0469)	(0.0702)	(0.0340)	(0.0227)	(0.0303)
Pre-1920 Dependent Mean	0.09	0.15	0.07	0.05	0.03
B. Log Overall Employment:					
Quota Exposure X Post 1920	-0.304	-0.720**	-0.804**	-0.460	-0.117
	(0.254)	(0.321)	(0.328)	(0.291)	(0.257)
Pre-1920 Dependent Mean	6,021	1,732	1,214	2,639	388
Quota Exposure - Mean	0.11	0.11	0.11	0.11	0.11
Quota Exposure - SD	0.08	0.08	0.08	0.08	0.08
Number of RLMs	125	125	125	125	125
Number of Industries	18	18	18	18	18
Number of RLM-Industry Pairs	399	399	399	399	399
Observations (RLM-Industry Pair)	1,197	1,197	1,197	1,197	1,197

Notes: This table reports coefficients on the interaction between the quota exposure measure and a post-1920 indicator for the pooled sample. Columns 1–5 report results for different occupation categories within each RLM-industry pair: overall employment in column 1; laborers in column 2; other low-skill workers, including service workers, farmers, and operatives, in column 3; medium-skill workers, including crafts, clerical workers, and sales workers, in column 4; and high-skill workers, including managers and professionals, in column 5. Panel A shows results for the share of recent immigrant arrivals from quota-restricted countries, defined as those who arrived within 10 years. Panel B shows results for log total employment. All outcomes are measured at the Relevant Labor Market (RLM) by industry level. All specifications include RLM-industry, industry-year, and division-year fixed effects. Each RLM-industry pair appears in the 1910, 1920, and 1930 censuses. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. Significance levels are indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table C2. Exposure to Border Closure Policy and Laborers Employment - by Subgroup

Sample:	Pooled (1)	Manufacturing & Mining (2)	American Contractor (3)	Construction Union (4)	Agriculture (5)
1. (Log) Overall Employment (Laborers):					
Quota Exposure X Post 1920	-0.769**	-0.469	-1.654**	-0.916	0.004
	(0.333)	(0.477)	(0.675)	(0.596)	(0.842)
Pre-1920 Dependent Mean	1,732	1,409	2,210	3,148	1,490
2. Recent Immigrants - Quota Restricted Countries:					
Quota Exposure X Post 1920	-0.792***	-1.016***	-0.632***	-0.406**	-0.650***
	(0.0693)	(0.0968)	(0.130)	(0.165)	(0.0542)
Pre-1920 Dependent Mean	0.15	0.17	0.14	0.16	0.05
3. Earlier Immigrants - Quota Restricted Countries:					
Quota Exposure X Post 1920	0.252***	0.315***	0.174	0.252	0.0695
	(0.0599)	(0.0975)	(0.109)	(0.156)	(0.0908)
Pre-1920 Dependent Mean	0.17	0.18	0.17	0.23	0.08
4. Immigrants - Non-Quota Countries:					
Quota Exposure X Post 1920	0.0925***	0.162***	0.0165	0.0321	-0.0324
	(0.0310)	(0.0576)	(0.0232)	(0.0431)	(0.142)
Pre-1920 Dependent Mean	0.04	0.05	0.01	0.03	0.03
5. Native Born - Whites:					
Quota Exposure X Post 1920	0.319***	0.447***	0.289	-0.0828	0.550**
	(0.0739)	(0.0885)	(0.180)	(0.151)	(0.220)
Pre-1920 Dependent Mean	0.47	0.46	0.41	0.37	0.68
6. Native Born - Non-Whites:					
Quota Exposure X Post 1920	0.128***	0.0917*	0.152	0.205	0.0631
	(0.0492)	(0.0532)	(0.124)	(0.139)	(0.0430)
Pre-1920 Dependent Mean	0.17	0.14	0.27	0.21	0.17
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
Number of RLMs	125	48	58	45	48
Number of Industries	18	16	1	1	1
Number of RLM-Industry Pairs	399	248	58	45	48
Observations (RLM-Industry Pair)	1,197	744	174	135	144

Notes: This table reports coefficients on the interaction between the quota exposure measure and a post-1920 indicator. Column (1) uses the pooled sample; columns (2)–(5) present results by data source. The outcome variable is the log number of laborers, defined using OCC1950 codes 810–979, employed within each RLM-industry pair in row (1) and the employment share of a group in rows 2–6. Results are reported for six population groups, in the following order: overall laborer employment, recent immigrant arrivals from quota-restricted countries, earlier immigrants from quota-restricted countries, immigrants from non-quota countries, native-born whites, and native-born non-whites. All outcomes are measured at the Relevant Labor Market (RLM) by industry level. Specifications include RLM-industry, industry-year, and division-year fixed effects. Each RLM-industry pair appears in the 1910, 1920, and 1930 censuses. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. Significance levels are indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Additional Results - Semi-Skilled Workers.

Table C3. Exposure to Border Closure Policy and Construction Workers' Wages

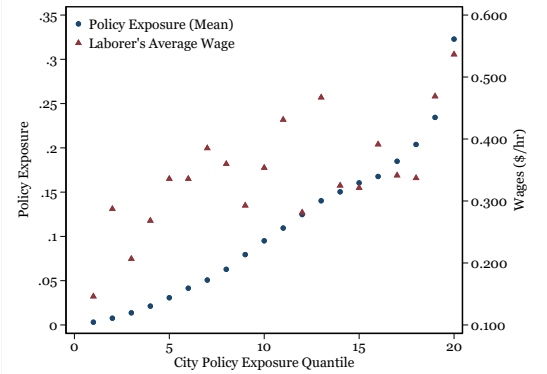
Sample:	**American Contractor**		
Occupation (Construction Industry):	**Laborers**	**Hod-Carriers**	**Bricklayers**
	(1)	**(2)**	**(3)**
A. Foreign Born Share:			
1. *All Foreign-Born*			
Quota Exposure X Post 1920	-0.443***	-0.425	-0.0396
	(0.159)	(0.278)	(0.0834)
Pre-1920 Dependent Mean	0.32	0.05	0.28
2. *Quota Countries*			
Quota Exposure X Post 1920	-0.461***	-0.144	-0.0587
	(0.152)	(0.113)	(0.0894)
Pre-1920 Dependent Mean	0.30	0.04	0.26
B. Log Mean Hourly Wage:			
Quota Exposure X Post 1920	0.628**	0.108	0.0746
	(0.274)	(0.261)	(0.152)
1920 Mean Hourly Wage	0.614	0.757	1.234
Quota Exposure - Mean	0.13	0.13	0.13
Quota Exposure - SD	0.10	0.10	0.10
Number of Cities	54	49	54
Number of States	26	26	26
Observations (City-Year Pairs)	288	279	296

Notes: This table presents estimation results using the American Contractor survey sample, which reports mean wages of various construction occupations during the 1920s. Column 1 shows estimates for construction laborers, column 2 for hod-carriers, and column 3 for bricklayers. The outcomes in Panel A are the foreign-born share, overall and from quota-restricted countries, for the occupation in the city. The outcomes in Panel B are the log mean hourly wage for workers in a given occupation in a city. The table presents coefficients on the interaction between the quota exposure measure and the post-1920 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 are at the city-year level and are obtained from the American Contractor surveys of construction contractors in the 1920s. Panel A specifications include city and division-year fixed effects. Panel B specifications include city, year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the city level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

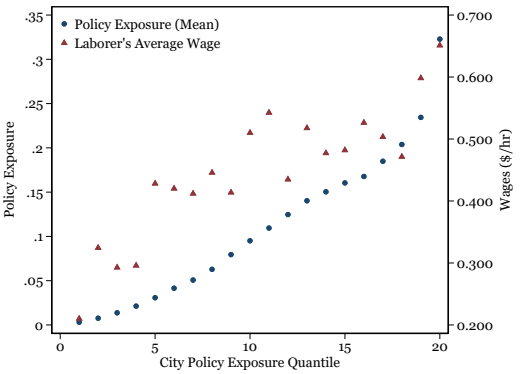
Additional Results - Municipal Street Laborers.

Figure C1. Cross-Sectional Exposure to Border Closure Policy and Wages - Pooled Sample

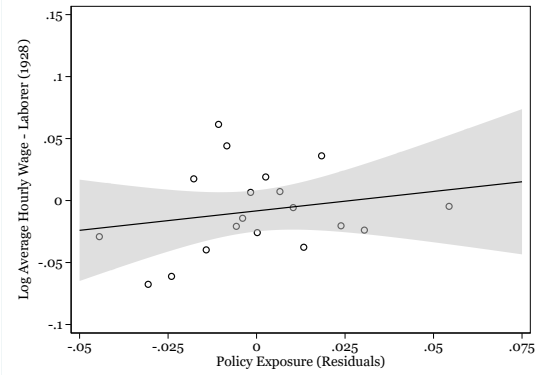
(a) Unadjusted: Pre-1920



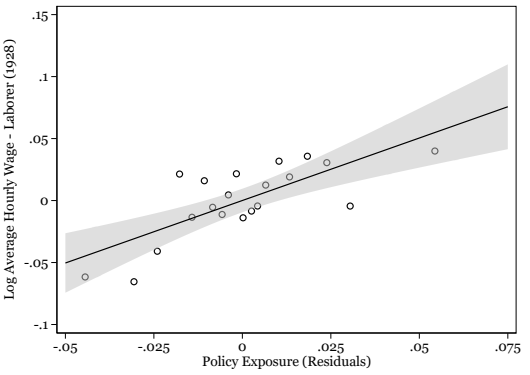
(b) Unadjusted: Post-1920



(c) Residualized: Pre-1920



(d) Residualized: Post-1920



Notes: Panels (a) and (b) plot the relationship between RLM-industry policy exposure quantile and policy exposure mean (blue) and mean hourly wage pre- and post-1920 (red triangles), respectively, for each quantile. Each of the 399 RLM-industry pairs from the pooled sample is assigned into quantiles. Panels (c) and (d) plot the relationship between policy exposure and log mean hourly wage pre- and post-1920 using a binned scatter plot. We exclude outliers with wage residuals above .2 or below -.1. Each RLM-industry policy exposure and wages are adjusted for the set of controls listed in Table D7, 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.

Table C4. 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	2,295	25,216	New York, NY (6.1m)
Exposure Measure Quartile:					
Quartile 1 (Bottom)	0.01	\$0.31	576	9,257	Decatur, IL (157k)
Quartile 2	0.04	\$0.40	574	20,087	Davenport, IA (552k)
Quartile 3	0.12	\$0.49	574	23,868	St. Louis, MO (848k)
Quartile 4 (Top)	0.25	\$0.52	571	47,827	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.

Additional Results - Mechanisms

Table C5. Composition Change and Internal Migration by Quota Exposure

Sample:	DID Estimate (1)	Std. Error (2)	Pre-1920 Mean (3)
A. Entries (not in sample in T and in sample in T+10):			
Overall	0.021	(0.034)	0.884
Remained Laborer	0.048	(0.045)	0.288
Remained Laborer & Move to Higher QE	-0.030	(0.040)	0.157
Became Laborer	-0.063	(0.045)	0.300
Enter LF	0.049	(0.048)	0.298
B. Exits (in Sample in T but not in T+10):			
Overall	-0.105***	(0.029)	0.887
Remained Laborer	-0.067	(0.043)	0.265
Remained Laborer & Move to Higher QE	-0.004	(0.039)	0.106
Change Occupation from Laborer	0.007	(0.057)	0.527
Exit LF	-0.006	(0.055)	0.097
RLM-industry Pairs		399	
Observations		798	

Notes: This table presents difference-in-differences estimates of the interaction between quota exposure QE and a post-1920 indicator. Column 1 shows the DID estimate, column 2 shows robust standard errors clustered at the RLM-industry level, and column 3 shows pre-1920 means. Panel A examines entries, defined as individuals not in the sample at T but in the sample at $T + 10$: overall entries, remained laborer, remained laborer and moved to higher-QE RLM-industry pairs, became laborer, and entered the labor force. Panel B examines exits, defined as individuals in the sample at T but not in the sample at $T + 10$: overall exits, remained laborer, remained laborer and moved to higher-QE RLM-industry pairs, changed occupation from laborer, and exited the labor force. All specifications include RLM-industry, industry-year, and division-year fixed effects. The sample includes 399 RLM-industry pairs observed from 1910 to 1930. Robust standard errors are reported in parentheses. Significance levels are indicated as *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

D Robustness

This appendix presents the full set of robustness analyses summarized in Section 6.4 in the main text. We examine whether the main results are sensitive to pre-trends, alternative geography-time and industry-time controls, the influence of individual data sources, alternative definitions of quota exposure, additional covariates, aggregation levels, and controls for wartime disruptions. Across these exercises, the estimated post-1920 wage effects remain positive and statistically significant. The results are robust to specifications with non-overlapping treatment periods, alternative geographic units, and different frequencies of industry-time controls; they also remain stable when excluding each data source one at a time. In addition, the estimates are similar when exposure is defined at the state level, when using binary or 1900-based exposure measures, and when aggregating the data to broader units such as RLMs or region–industry cells. Adding pre-1910 demographic and industrial covariates has little effect on the pooled sample estimates, suggesting that the results are not explained by pre-existing differences across labor markets or industries. Finally, controlling directly for World War I immigration exposure does not alter the quota-exposure estimates, supporting the interpretation that the wage responses reflect the 1920s immigration restrictions rather than temporary wartime disruptions.

Pre-Trends. To assess whether log mean hourly wages exhibited similar pre-1920 trends across RLM–industry pairs with different exposure levels, we restrict the sample to 1910–1920 and regress wages on the exposure measure, a linear time trend, and their interaction (Table D1, col. 1). We then allow for nonlinear pre-policy dynamics by replacing the linear trend with year dummies for 1911–1920 (Table D1, col. 2). In neither specification do we find significant evidence that pre-1921 wage trends varied with exposure. Repeating these tests separately by dataset yields the same conclusion.

Table D1. 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	-0.004		-0.044		0.015		0.000	
	(0.033)		(0.055)		(0.035)		(0.020)	
Quota Exposure X Year 1911-12		-0.073		0.202		-0.057		-0.378**
		(0.168)		(0.298)		(0.127)		(0.146)
Quota Exposure X Year 1913-14		0.010		0.197		0.443		0.039
		(0.176)		(0.339)		(0.393)		(0.106)
Quota Exposure X Year 1915-16		0.213		0.481		0.250		0.169
		(0.206)		(0.297)		(0.481)		(0.172)
Quota Exposure X Year 1917-18		-0.053		-0.374		0.582		-0.053
		(0.276)		(0.475)		(0.461)		(0.164)
Quota Exposure X Year 1919-20		-0.377		-0.781		1.041**		-0.261
		(0.302)		(0.495)		(0.421)		(0.205)
R-Squared	0.587	0.663	0.636	0.753	0.537	0.695	0.620	0.707
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. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Geography-Time Controls. To assess whether the results are sensitive to the geography-time controls, Table D2 reports estimates under specifications that vary both the geographic unit and the period structure. The specifications range from broader census-region controls to narrower state controls, and include non-overlapping periods such as post-1920 fixed effects and four-period fixed effects that separate 1915–1919 from 1920–1924. Across these alternatives, the effects on recent immigrant shares and wage remain large, statistically significant, and similar in magnitude.

Figure D1 presents the corresponding event-study estimates. Across most specifications, the pre-1920 coefficients are close to zero, while the post-1920 coefficients become positive after the quota restrictions and remain elevated through the 1920s. This pattern is similar when using no geography-time controls, region-by-post-1920 controls, division-by-post-1920 controls, and division-by-four-period controls.

Panels (e) and (f) illustrate the limits of increasingly saturated geography-time controls in our setting. In panel (e), state-by-period controls preserve a positive post-1920 wage response and

do not generate statistically significant pre-1920 estimates, but the event-study path is noisier and suggests some pre-period movement. This is not surprising: within-state exposure variation is limited, especially once the specification also includes industry-by-year fixed effects. Panel (f) shows a more demanding specification with both industry-by-year and division-by-year fixed effects. This saturated specification absorbs much of the policy-relevant variation in exposure and leaves identification to come from a narrow residual comparison within division–industry–year cells. As a result, the coefficients are shifted downward and become less precise. We therefore prefer the baseline specification, which controls flexibly for industry-by-year shocks and broad regional trends while retaining meaningful within-division variation in quota exposure.

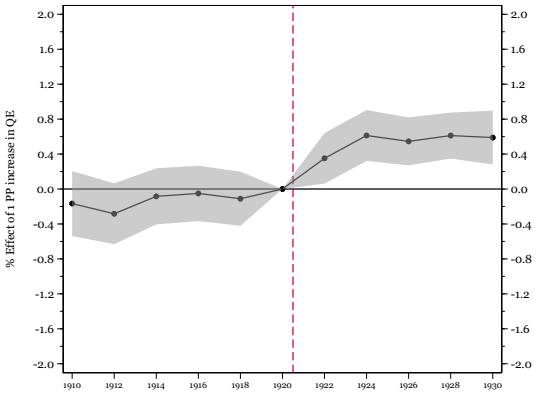
Table D2. Robustness to Alternative Geography-Time Controls

Geography-Time Controls:	**Baseline (Division X 3 Periods)**	**No Geography-Time Controls**	**Region X Post-1920**	**Region X 3 Periods**	**State X 3 Periods**	**Division X 4 Periods**	**Division X Year**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
A. Recent Immigrants Share:							
Quota Exposure X Post 1920	-0.496***	-0.511***	-0.468***	-0.468***	-0.518***	–	–
	(0.0471)	(0.0334)	(0.0455)	(0.0455)	(0.0651)	–	–
Pre-1920 Dependent Mean	0.09	0.09	0.09	0.09	0.09	–	–
Quota Exposure - Mean	0.11	0.11	0.11	0.11	0.11	–	–
Quota Exposure - SD	0.08	0.08	0.08	0.08	0.08	–	–
Number of Panel Units	399	399	399	399	399	–	–
Observations	1,197	1,197	1,197	1,197	1,197	–	–
B. Log Mean Hourly Wage:							
Quota Exposure X Post 1920	0.532***	0.616***	0.534***	0.584***	0.576***	0.449***	0.414***
	(0.128)	(0.140)	(0.148)	(0.129)	(0.126)	(0.131)	(0.146)
Pre-1920 Dependent Mean	0.258	0.258	0.258	0.258	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.100	0.100	0.100	0.100	0.100	0.100
Quota Exposure - SD	0.082	0.082	0.082	0.082	0.082	0.082	0.082
Number of Panel Units	399	399	399	399	399	399	399
Observations	2,973	2,973	2,973	2,973	2,973	2,973	2,973

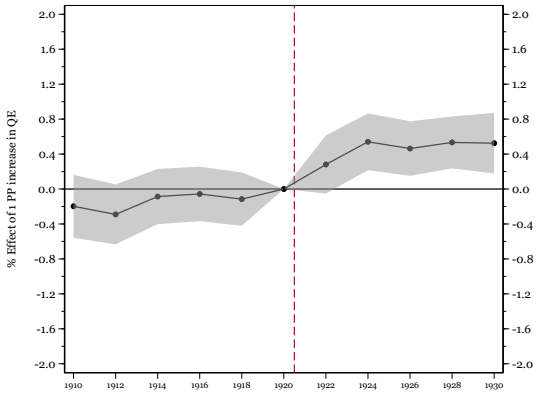
Notes: This table reports the coefficients of the interaction between quota exposure QE and the post-1920 indicator under alternative geography-time controls. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses the RLM-industry log mean hourly wage for low-skilled workers. Column (1) reports the baseline specification. All specifications include RLM-industry and industry-year fixed effects. Columns differ only in the geography-time controls included. For Panel A, which uses decennial census outcomes observed in 1910, 1920, and 1930, geography-by-period controls correspond to geography-by-year fixed effects. For Panel B, three-period controls are defined using 1910–1914, 1915–1924, and 1925–1930. Four-period controls are defined using 1910–1914, 1915–1919, 1920–1924, and 1925–1930. Columns (6) and (7) are reported only for the annual wage outcome in Panel B. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D1. Event-Study Estimates Under Alternative Geography-Time Controls

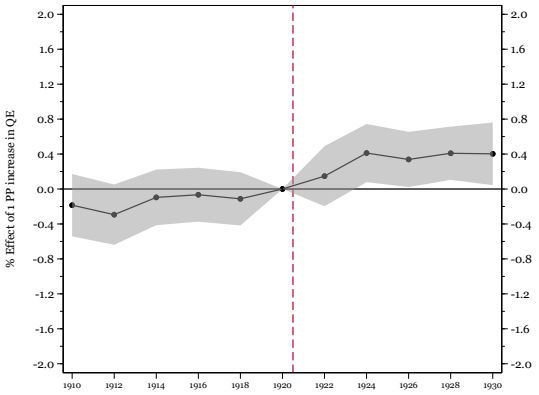
(a) No geography-time controls



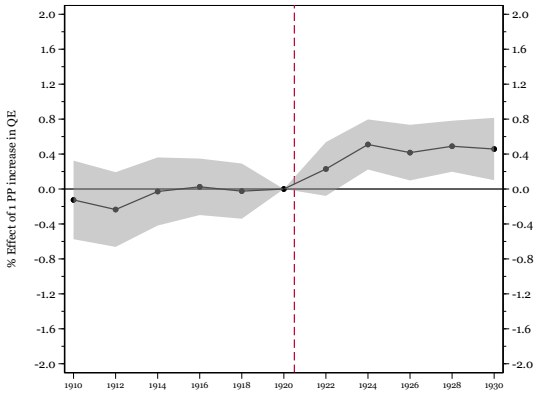
(b) Region × Post-1920



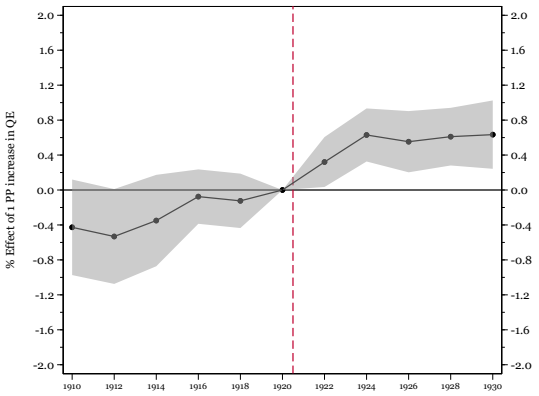
(c) Division × Post-1920



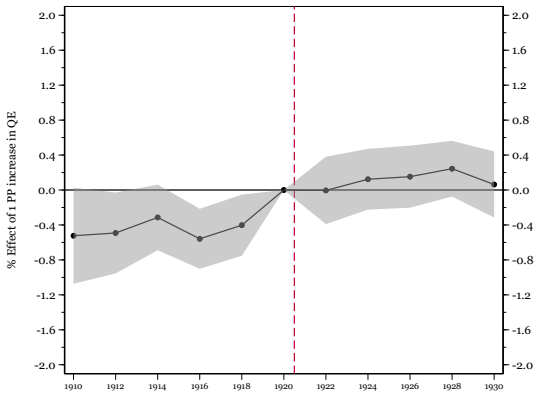
(d) Division × 4 Periods



(e) State × 3 Periods



(f) Division × Year



Notes: The figures plot event-study coefficients of the interaction between exposure to the border-closure policy and year-pair indicators for the pooled sample under alternative geography-time controls. All specifications include RLM-industry, industry-year, and dataset-year fixed effects. The specifications differ only in the geography-time controls listed in each panel title. The omitted period is 1919–1920. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM-industry level.

Industry-Time Controls. To assess whether the results are sensitive to the treatment of industry-specific time shocks, Table D3 reports estimates under alternative industry-time controls. The baseline specification includes industry-by-year fixed effects, which absorb high-frequency shocks common to each industry in a given year. We compare this specification to models with no industry-time controls and to lower-frequency controls that allow industry effects to vary by the post-1920 period, by the baseline three-period structure, or by four non-overlapping periods. Across these alternatives, the estimated effects on recent immigrant shares remain large and statistically significant, and the wage estimates remain positive, statistically significant, and similar in magnitude.

Figure D2 presents the corresponding event-study estimates. The figure moves from the most flexible specification, with industry-by-year fixed effects, to progressively lower-frequency controls: industry-by-four-period, industry-by-three-period, industry-by-post-1920, and no industry-time controls. Overall, the event studies are broadly consistent with the baseline specification. In most panels, the pre-1920 coefficients are close to zero, while the post-1920 coefficients rise after the quota restrictions and remain elevated through the 1920s.

The event studies also show why some control for industry-specific time variation is important in this setting. When we omit industry-time controls entirely, the estimates become noisier and the pre-period pattern is less clean, suggesting some differential wage movements before the quota restrictions. This is not surprising given the structure of the data: the composition of industries varies over time, industries differ substantially in their exposure to the quotas, and industries also experienced different wage dynamics during the 1910s and 1920s. Controlling for industry-time shocks therefore helps ensure that the estimated wage effects are identified from within-industry differences in quota exposure across RLMs, rather than from differential trends across industries.

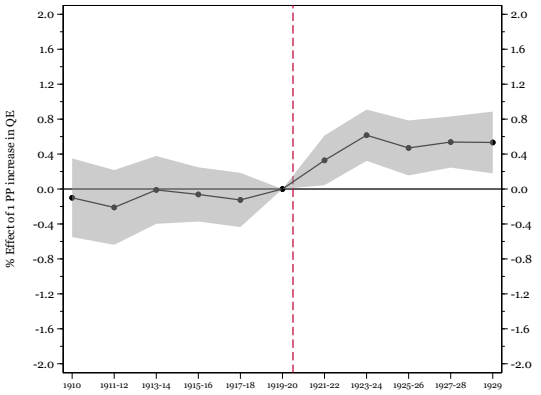
Table D3. Robustness to Alternative Industry-Time Controls

Industry-Time Controls:	**Baseline Industry X Year**	**No Industry-Time Controls**	**Industry X Post-1920**	**Industry X 3 Periods**	**Industry X 4 Periods**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
A. Recent Immigrants Share:					
Quota Exposure X Post 1920	-0.496***	-0.530***	-0.496***	–	–
	(0.0471)	(0.0578)	(0.0467)	–	–
Pre-1920 Dependent Mean	0.09	0.09	0.09	–	–
Quota Exposure - Mean	0.11	0.11	0.11	–	–
Quota Exposure - SD	0.08	0.08	0.08	–	–
Number of Panel Units	399	399	399	–	–
Observations	1,197	1,197	1,197	–	–
B. Log Mean Hourly Wage:					
Quota Exposure X Post 1920	0.532***	0.657***	0.507***	0.554***	0.535***
	(0.128)	(0.135)	(0.127)	(0.125)	(0.127)
Pre-1920 Dependent Mean	0.258	0.258	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.100	0.100	0.100	0.100
Quota Exposure - SD	0.082	0.082	0.082	0.082	0.082
Number of Panel Units	399	399	399	399	399
Observations	2,973	2,973	2,973	2,973	2,973

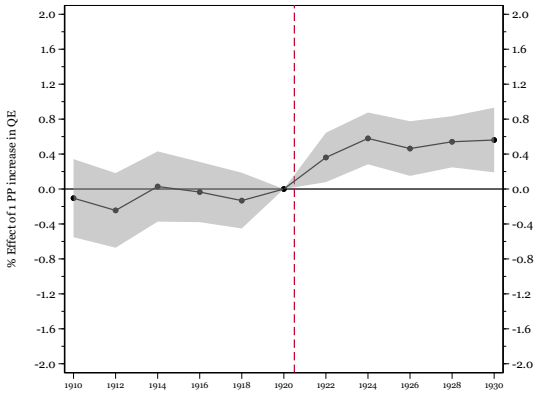
Notes: This table reports the coefficients of the interaction between quota exposure QE and the post-1920 indicator under alternative industry-time controls. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses the RLM-industry log mean hourly wage for low-skilled workers. Column (1) reports the baseline specification. All specifications include RLM-industry, dataset-year, and division-by-period fixed effects. Columns differ only in the industry-time controls included. For Panel A, which uses decennial census outcomes observed in 1910, 1920, and 1930, industry-by-period controls correspond to industry-by-year fixed effects; columns (4) and (5) are therefore reported only for the annual wage outcome in Panel B. Three-period controls are defined using 1910–1914, 1915–1924, and 1925–1930. Four-period controls are defined using 1910–1914, 1915–1919, 1920–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D2. Event-Study Estimates Under Alternative Industry-Time Controls

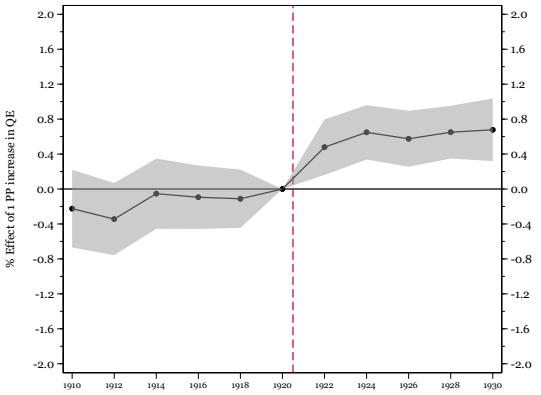
(a) Industry × year



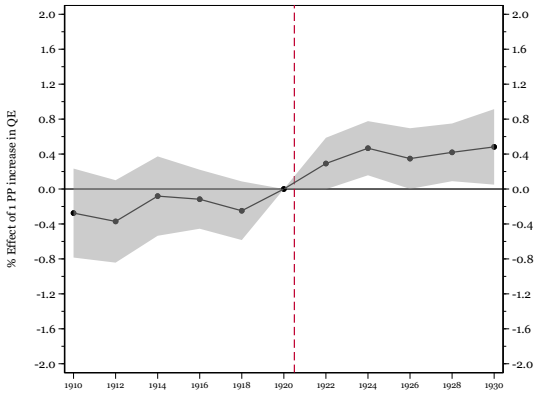
(b) Industry × 4 periods



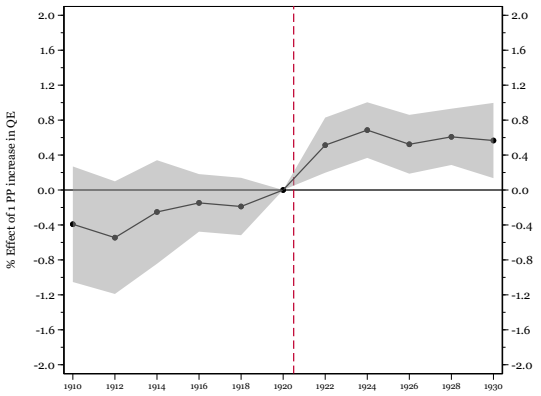
(c) Industry × 3 periods



(d) Industry × post-1920



(e) No industry-time controls



Notes: The figures plot event-study coefficients of the interaction between exposure to the border-closure policy and year-pair indicators for the pooled sample under alternative industry-time controls. All specifications include RLM-industry, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. The specifications differ only in the industry-time controls listed in each panel title. The omitted period is 1919–1920. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM-industry level.

Excluding Individual Data Sources. We next assess whether the pooled results are driven by any single data source. This is especially important for two reasons. First, the construction union sample exhibits some evidence of differential pre-trends in the dataset-specific event study in Figure 9, panel (c), in the main text. Second, the American Contractor sample contains mostly post-1920 observations, limiting its usefulness for assessing pre-quota wage trends. Table D4 therefore reports estimates that sequentially exclude each of the four data sources from the pooled sample.

The results are highly stable. The estimated effects on recent immigrant shares remain negative, large, and statistically significant in every leave-one-out specification. The wage estimates also remain positive, statistically significant, and similar in magnitude regardless of which data source is excluded. Importantly, excluding the construction union sample leaves the pooled wage estimate essentially unchanged, indicating that the main result is not driven by the dataset with the least clean pre-period pattern. Excluding the American Contractor sample also produces a similar estimate, showing that the results do not depend on the data source with limited pre-1920 coverage. More broadly, the leave-one-out estimates indicate that no single source is responsible for the pooled wage effect.

Figure D3 presents the corresponding event-study estimates. Across the leave-one-out samples, the dynamic patterns remain broadly similar to the baseline: the pre-1920 coefficients are generally close to zero and insignificant, while the post-1920 coefficients become positive and remain elevated through the 1920s. This holds when excluding each data source in turn, including the construction union data and the American Contractor data.

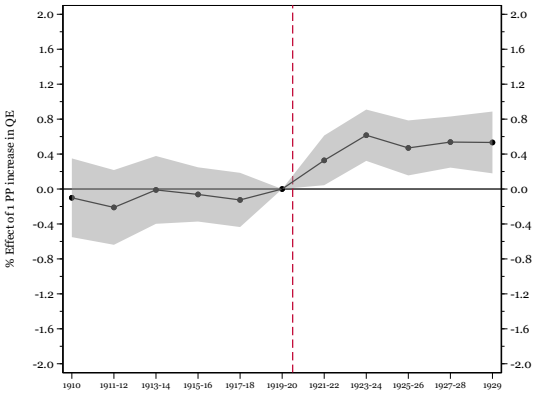
Table D4. Robustness to Excluding Individual Data Sources

Sample:	Pooled Sample (1)	Excluding Manufacturing & Mining (2)	Excluding American Contractor (3)	Excluding Construction Union (4)	Excluding Agriculture (5)
A. Recent Immigrants Share:					
Quota Exposure X Post 1920	-0.496***	-0.291***	-0.591***	-0.547***	-0.492***
	(0.0470)	(0.0334)	(0.0598)	(0.0529)	(0.0498)
Pre-1920 Dependent Mean	0.09	0.06	0.09	0.09	0.10
Quota Exposure - Mean	0.11	0.10	0.11	0.11	0.12
Quota Exposure - SD	0.08	0.09	0.08	0.08	0.08
Number of Panel Units	399	151	341	354	351
Observations	1,197	453	1,023	1,062	1,053
R-Squared	0.892	0.901	0.896	0.894	0.891
B. Log Mean Hourly Wage:					
Quota Exposure X Post 1920	0.532***	0.529***	0.507***	0.501***	0.565***
	(0.128)	(0.166)	(0.139)	(0.132)	(0.138)
Pre-1920 Dependent Mean	0.258	0.267	0.249	0.228	0.304
Quota Exposure - Mean	0.100	0.094	0.099	0.088	0.121
Quota Exposure - SD	0.082	0.084	0.082	0.074	0.087
Number of Panel Units	399	151	341	354	351
Observations	2,973	1,739	2,681	2,486	2,013
R-Squared	0.979	0.979	0.981	0.982	0.971

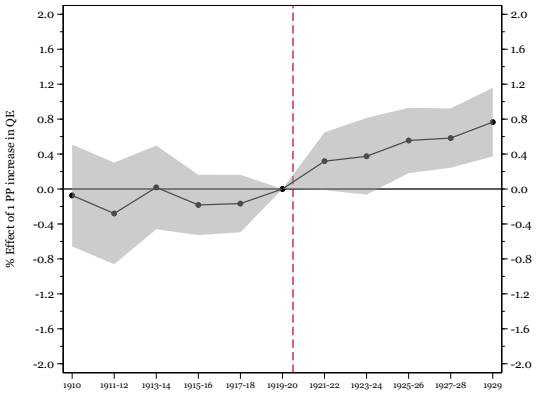
Notes: This table reports leave-one-data-source-out estimates of the interaction between quota exposure QE and the post-1920 indicator. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses the RLM-industry log mean hourly wage for low-skilled workers. Column (1) reports the pooled-sample baseline specification. Columns (2)–(5) sequentially exclude one data source from the pooled sample. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D3. Event-Study Estimates Excluding Individual Data Sources

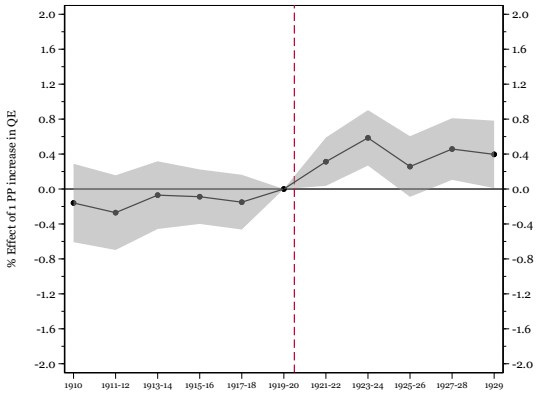
(a) Pooled sample (Baseline)



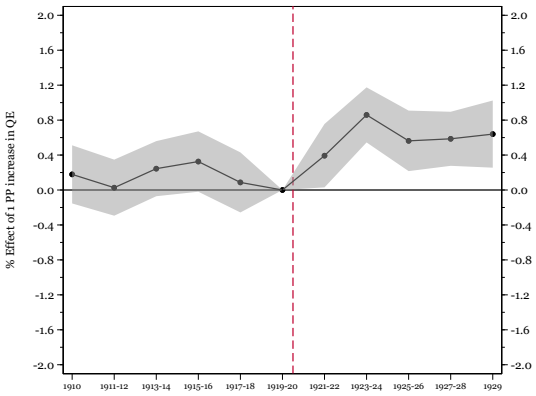
(b) Excluding Manufacturing & Mining



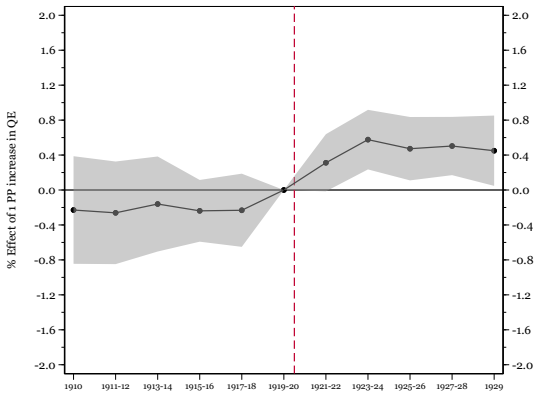
(c) Excluding American Contractor



(d) Excluding Construction Union



(e) Excluding agriculture



Notes: The figures plot event-study coefficients of the interaction between exposure to the border-closure policy and year-pair indicators for the pooled sample and for samples that sequentially exclude one data source. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. The omitted period is 1919–1920. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM-industry level.

Alternative Exposure Measures. We test the robustness of our findings to alternative definitions of the quota exposure measure in Table D5. Column 1 reports our baseline specification, where exposure is defined at the RLM–industry level by applying county–industry employment shares within each state to the county-level exposure measure as defined in Equation 8 in the main text. To address possible concerns with use of country employment shares in creating RLM exposure measures, we construct a state-level exposure measure that ignores the RLM definition and aggregates exposure at the state level. Formally, this is the analogous of Equation 7 in the main text,

$$q_s = \sum_r \frac{FB_{rs1910}}{LF_{s1910}} \times QI_r \tag{23}$$

where FB_{rs1910} and LF_{s1910} denote the number of foreign-born and total working-age (16–65) male workers in state s in 1910, respectively. For the construction datasets, we retain the county-level exposure q_c, as these measures are directly aligned with the underlying data.

Column (2) of Table D5 reports estimates using this state-based exposure. The results for both recent immigrant shares and wages remain statistically significant and similar in magnitude to our preferred RLM–industry specification. Furthermore, Table D6 and Figure D4 shows that event study estimates using a state-based exposure measure produces very similar results to our preferred RLM-based exposure measure event studies. This confirms that our results are not driven by the additional variation created when assigning distinct exposure measures at the geography–industry level rather than at the geographic level alone. In column (3), we further modify the country-specific quota intensity variable QI_r, replacing the continuous counterfactual measure with a binary indicator equal to one for high-restriction countries and zero otherwise. The results remain robust.

Finally, columns (4)–(6) of Table D5 we use the 1900 Census instead of 1910 to construct both the RLM-based and state-based measures (including the binary specification). The results remain virtually unchanged.

Table D5. Robustness to Alternative Quota Exposure Measures

Exposure Measure:	1910 Census			1900 Census		
	Baseline (RLM-Industry) (1)	State Based (2)	Binary Exposure (3)	RLM-Industry Based (4)	State Based (5)	Binary Exposure (6)
A. Recent Immigrants Share:						
Quota Exposure X Post 1920	-0.496***	-0.352***	-0.478***	-0.512***	-0.424***	-0.506***
	(0.0470)	(0.0338)	(0.0467)	(0.0737)	(0.0662)	(0.0754)
Pre-1920 Dependent Mean	0.09	0.08	0.09	0.09	0.08	0.09
Quota Exposure - Mean	0.11	0.10	0.11	0.06	0.06	0.06
Quota Exposure - SD	0.08	0.08	0.09	0.05	0.05	0.05
Number of Panel Units	399	399	399	397	399	397
Observations	1,197	1,197	1,197	1,191	1,197	1,191
R-Squared	0.892	0.886	0.891	0.880	0.881	0.880
B. Log Mean Hourly Wage:						
Quota Exposure X Post 1920	0.532***	0.571***	0.579***	0.509**	0.600***	0.661***
	(0.128)	(0.118)	(0.112)	(0.216)	(0.213)	(0.212)
Pre-1920 Dependent Mean	0.258	0.258	0.258	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.105	0.104	0.059	0.062	0.058
Quota Exposure - SD	0.082	0.081	0.083	0.054	0.053	0.052
Number of Panel Units	399	399	399	397	399	399
Observations	2,973	2,973	2,973	2,965	2,973	2,973
R-Squared	0.979	0.980	0.980	0.979	0.979	0.979

Notes: This table reports the coefficients of the interaction between quota exposure measures and the post-1920 indicator. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses log mean hourly wages for low-skilled workers. The post-1920 indicator equals one in years after 1920, the last census year before the quota policy took effect. Columns (1)–(3) construct quota exposure using the 1910 Census, while columns (4)–(6) use the 1900 Census. Columns (1) and (4) use continuous RLM-industry quota exposure; columns (2) and (5) use continuous state-based quota exposure; and columns (3) and (6) use binary exposure indicators. All specifications include RLM-industry, industry-year, and dataset-year fixed effects. Panel A also includes division-year fixed effects, while panel B includes division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

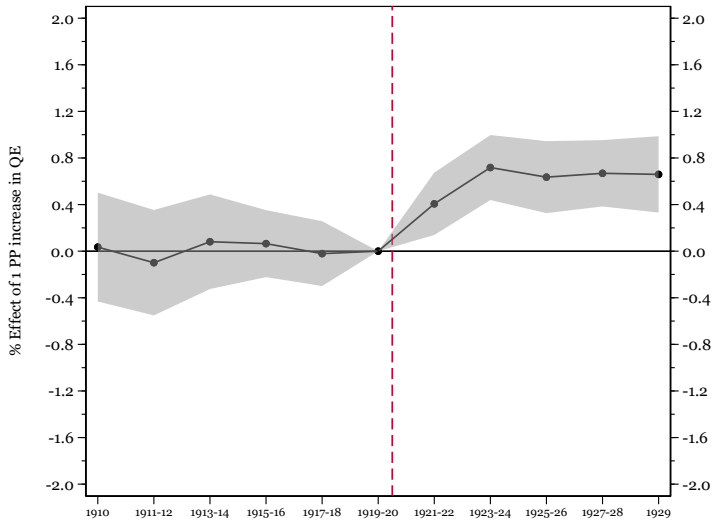
Table D6. Robustness to Alternative Quota Exposure Measures – Event-Study Estimates

Exposure Measure:	1910 Census			1900 Census		
	Baseline (RLM-Industry) (1)	State Based (2)	Binary Exposure (3)	RLM-Industry Based (4)	State Based (5)	Binary Exposure (6)
Quota Exposure X Year 1910	-0.100	0.0350	0.0223	-0.212	-0.0211	-0.0248
	(0.229)	(0.238)	(0.236)	(0.297)	(0.319)	(0.335)
Quota Exposure X Year 1911–12	-0.211	-0.0990	-0.102	-0.287	-0.143	-0.125
	(0.218)	(0.230)	(0.228)	(0.287)	(0.316)	(0.330)
Quota Exposure X Year 1913–14	-0.0102	0.0812	0.0715	-0.0930	0.0497	0.0621
	(0.197)	(0.207)	(0.204)	(0.255)	(0.280)	(0.288)
Quota Exposure X Year 1915–16	-0.0620	0.0648	0.0692	-0.167	-0.0134	0.00572
	(0.158)	(0.146)	(0.145)	(0.243)	(0.227)	(0.235)
Quota Exposure X Year 1917–18	-0.125	-0.0211	-0.0253	-0.219	-0.0780	-0.0822
	(0.158)	(0.142)	(0.139)	(0.230)	(0.212)	(0.217)
Quota Exposure X Year 1921–22	0.328**	0.406***	0.410***	0.311	0.454**	0.500**
	(0.145)	(0.136)	(0.132)	(0.204)	(0.201)	(0.200)
Quota Exposure X Year 1923–24	0.616***	0.718***	0.724***	0.431*	0.605**	0.674***
	(0.149)	(0.142)	(0.136)	(0.242)	(0.247)	(0.245)
Quota Exposure X Year 1925–26	0.470***	0.635***	0.654***	0.327	0.572**	0.669***
	(0.160)	(0.157)	(0.152)	(0.244)	(0.248)	(0.245)
Quota Exposure X Year 1927–28	0.537***	0.669***	0.689***	0.498**	0.703***	0.808***
	(0.149)	(0.145)	(0.141)	(0.231)	(0.236)	(0.232)
Quota Exposure X Year 1929	0.533***	0.659***	0.672***	0.501*	0.752***	0.853***
	(0.180)	(0.167)	(0.163)	(0.268)	(0.258)	(0.259)
Pre-1920 Dependent Mean	0.258	0.258	0.258	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.105	0.104	0.059	0.062	0.058
Quota Exposure - SD	0.082	0.081	0.083	0.054	0.053	0.052
Number of Panel Units	399	399	399	397	399	399
Observations	2,973	2,973	2,973	2,965	2,973	2,973
R-Squared	0.976	0.976	0.976	0.975	0.975	0.975

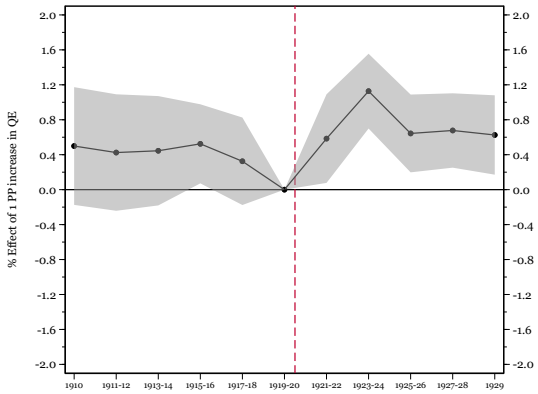
Notes: This table reports event-study coefficients from specifications that interact quota exposure measures with year-pair indicators. The outcome is the RLM-industry log mean hourly wage for low-skilled workers. The omitted period is 1919–1920. Columns (1)–(3) construct quota exposure using the 1910 Census, while columns (4)–(6) use the 1900 Census. Columns (1) and (4) use continuous RLM-industry quota exposure; columns (2) and (5) use continuous state-based quota exposure; and columns (3) and (6) use binary exposure indicators. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D4. Event-Study Estimates Using the State-Based Quota Exposure Measure

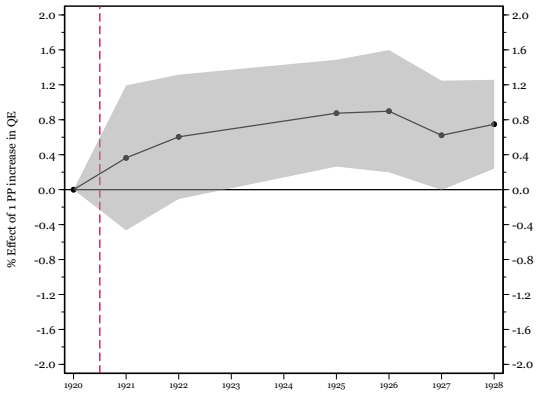
(a) Pooled sample



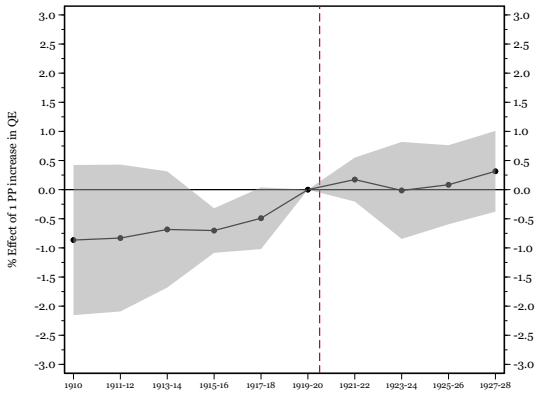
(b) Manufacturing & mining



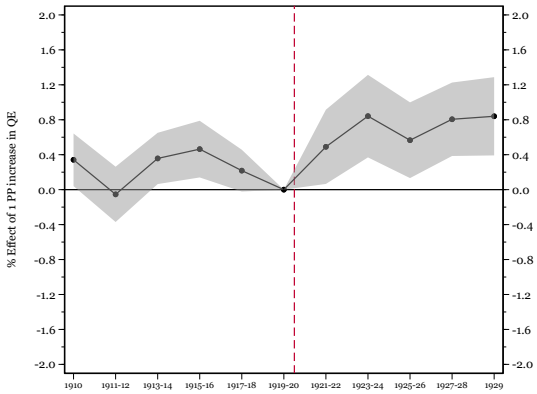
(c) American Contractor



(d) Construction Union



(e) Farm labor



Notes: The figures plot event-study coefficients of the interaction between the state-based quota exposure measure and year-pair indicators. The outcome is the RLM-industry log mean hourly wage for low-skilled workers. The omitted period is 1919–1920. All specifications include RLM-industry, industry-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. The pooled specification also includes dataset-year fixed effects. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM-industry level.

Inclusion of Covariates. To assess the robustness of our results to the inclusion of economic and demographic controls, we examine the relationship between RLM–industry exposure and a broad set of 1910 covariates across datasets. As shown in Table D7, a Lasso procedure identifies three key predictors of quota exposure in the pooled sample: the 1910 log of total population, the share of workers in agriculture, and the share of workers in mining. Following Abramitzky et al. (2023), we also include the 1910 foreign-born share directly in our wage regressions to account for potential "scale effects" of immigration restrictions and isolate the impact driven by differences in immigrant composition.

Table D8 and Table D9 report estimates from specifications that add all Lasso-selected 1910 controls (and the foreign-born share for wage regressions) interacted either with a linear time trend or with a post-1920 indicator.[3] The pooled difference-in-differences estimates in Table D8 show that adding these controls leaves the main results unchanged. For wage regressions in Panel B, the inclusion of covariates does not affect the pooled estimates materially.

Table D9 presents the corresponding event-study estimates with covariates included. The pre-1920 coefficients remain statistically insignificant and the post-1920 effects stay positive and similar in magnitude to the baseline. Overall, these results indicate that our findings are robust to controlling for pre-existing demographic and economic differences across local labor markets, and that the estimated effects are not driven by omitted variation in 1910 characteristics.

[3]Interacting 1910 controls with full year fixed effects is too demanding given the limited number of observations, particularly in the event-study regressions. Allowing post-1920 trend changes based on these covariates provides a more parsimonious way to capture differential dynamics.

Table D7. Lasso Selection of 1910 Characteristics Predicting Quota Exposure

Sample:	Pooled	Manufacturing & Mining	American Contractor	Construction Union	Agriculture
	(1)	(2)	(3)	(4)	(5)
A. Demographics:					
Log Total Population	x	x	-	x	-
Share Urban Population	-	-	-	-	-
Share Black Population	-	-	-	-	-
Literacy Rate	-	-	-	x	-
B. Occupation:					
Share Workers in Blue-Collar Occupations (Manufacturing)	-	-	-	-	-
Share Workers in Farming Occupations (Agriculture)	x	x	x	x	x
Share Workers Holding White Collar Occupation	-	-	-	-	-
C. Industry:					
Share Workers in Mining	x	-	x	-	-
Share Workers in Construction	-	-	-	-	-
Share Workers in Transportation	-	-	-	-	-
Share Workers in Wholesale/Retail	-	-	-	-	x
Share Workers in Services	-	-	-	-	-
Share Workers in Public Administration	-	-	-	-	-
D. Farming:					
Log Average Farm Value	-	-	-	-	x
Log Value of Farm Output per Acre	-	-	-	-	-
Share Owner Operated Farms	-	-	-	-	-
Share Farmland Cultivated	-	-	-	-	-
Share Wheat in Cultivated Farmland	-	-	-	x	-
Share Cotton in Cultivated Farmland	-	-	-	-	-
Share Hay/Corn in Cultivated Farmland	-	-	-	-	-
Quota Exposure - Mean	0.109	0.113	0.111	0.129	0.066
Quota Exposure - SD	0.083	0.081	0.096	0.087	0.059
Number of Panel Units	399	248	58	45	48

Notes: Columns 1–5 show the variables selected by a lasso procedure in cross-sectional specifications where the dependent variable is the RLM-industry quota exposure measure and the potential explanatory variables are 1910 RLM-industry characteristics. Division fixed effects are included as unpenalized controls in all specifications. For the pooled sample, dataset and industry fixed effects are also included as unpenalized controls. For the manufacturing-and-mining sample, industry fixed effects are also included as unpenalized controls. Controls marked with an "x" are selected by the lasso procedure.

Table D8. Robustness to Inclusion of Covariates

Sample:	Pooled (1)	Manufacturing & Mining (2)	American Contractor (3)	Construction Union (4)	Agriculture (5)
A. Recent Immigrants Share:					
Baseline Specification	-0.496***	-0.720***	-0.285***	-0.246***	-0.438***
	(0.0470)	(0.0707)	(0.0474)	(0.0685)	(0.0491)
Covariates: Lasso Selected Controls (Linear Trends)	-0.432***	-0.631***	-0.244***	-0.146	-0.254***
	(0.0520)	(0.0835)	(0.0583)	(0.116)	(0.0871)
Covariates: Lasso Selected Controls × Post-1920	-0.512***	-0.754***	-0.301***	-0.340***	-0.414***
	(0.0545)	(0.0914)	(0.0609)	(0.0904)	(0.0709)
Pre-1920 Dependent Mean	0.09	0.11	0.06	0.08	0.03
Observations	1,197	744	174	135	144
R-Squared – Baseline	0.892	0.901	0.888	0.909	0.926
R-Squared – Linear Trends	0.894	0.903	0.890	0.916	0.942
R-Squared – Post-1920 Interactions	0.892	0.901	0.889	0.912	0.927
B. Log Mean Hourly Wage:					
Baseline Specification	0.532***	0.475***	0.628**	0.545**	0.403
	(0.128)	(0.159)	(0.275)	(0.224)	(0.256)
Covariates: 1910 Foreign-Born Share + Lasso Selected Controls (Linear Trends)	0.570***	0.491**	0.561*	0.560**	0.283
	(0.135)	(0.229)	(0.323)	(0.247)	(0.289)
Covariates: 1910 Foreign-Born Share + Lasso Selected Controls × Post-1920	0.681***	-0.0202	0.645	1.601***	0.217
	(0.227)	(0.340)	(0.821)	(0.568)	(0.410)
Pre-1920 Dependent Mean	0.258	0.245	0.614	0.397	0.186
Observations	2,973	1,234	292	487	960
R-Squared – Baseline	0.979	0.981	0.905	0.932	0.975
R-Squared – Linear Trends	0.980	0.981	0.906	0.935	0.976
R-Squared – Post-1920 Interactions	0.980	0.982	0.906	0.938	0.978
RLM-Industry Quota Exposure - Mean	9.95	10.69	12.84	15.18	6.60
RLM-Industry Quota Exposure - SD	8.21	7.86	9.53	9.74	5.89
Number of RLMs	125	48	58	45	48
Number of Industries	18	16	1	1	1
Number of RLM-Industry Pairs	399	248	58	45	48

Notes: This table reports robustness checks for the baseline specifications by including additional covariates. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses log mean hourly wages for low-skilled workers. Coefficients correspond to the interaction between quota exposure QE and the post-1920 indicator. Column (1) reports estimates for the pooled sample, and columns (2)–(5) report estimates separately for manufacturing and mining, American Contractor, Construction Union, and agriculture. The first row of each panel reports the baseline specification. In Panel A, the next two rows add all lasso-selected controls listed in Table D7, allowing the controls to enter either as linear time trends or interacted with the post-1920 indicator. In Panel B, the next two rows also add the 1910 foreign-born share. All specifications include RLM-industry fixed effects. Panel A includes industry-year and division-year fixed effects, while panel B includes industry-year and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. The pooled specifications also include dataset-year fixed effects. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table D9. Robustness to Inclusion of Covariates – Event-Study Estimates

Covariates Included:	**Baseline Specification**	**1910 Foreign-Born Share + Lasso Controls Linear Trends**	**1910 Foreign-Born Share + Lasso Controls Post-1920 Interactions**
	(1)	**(2)**	**(3)**
Quota Exposure X Year 1910	-0.100	-0.506	0.00395
	(0.229)	(0.337)	(0.227)
Quota Exposure X Year 1911–12	-0.211	-0.536*	-0.125
	(0.218)	(0.296)	(0.221)
Quota Exposure X Year 1913–14	-0.0102	-0.262	0.0422
	(0.197)	(0.252)	(0.202)
Quota Exposure X Year 1915–16	-0.0620	-0.149	0.0690
	(0.158)	(0.190)	(0.152)
Quota Exposure X Year 1917–18	-0.125	-0.134	-0.0269
	(0.158)	(0.154)	(0.144)
Quota Exposure X Year 1921–22	0.328**	0.512***	0.651***
	(0.145)	(0.141)	(0.230)
Quota Exposure X Year 1923–24	0.616***	0.938***	0.951***
	(0.149)	(0.156)	(0.212)
Quota Exposure X Year 1925–26	0.470***	0.961***	0.892***
	(0.160)	(0.188)	(0.217)
Quota Exposure X Year 1927–28	0.537***	1.100***	0.927***
	(0.149)	(0.204)	(0.213)
Quota Exposure X Year 1929	0.533***	1.135***	0.875***
	(0.180)	(0.257)	(0.222)
Pre-1920 Dependent Mean	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.105	0.105
Quota Exposure - SD	0.082	0.081	0.081
Number of Panel Units	399	399	399
Observations	2,973	2,973	2,973
R-Squared	0.976	0.976	0.977

Notes: This table reports event-study estimates of the interaction between quota exposure and year-pair indicators, using log mean hourly wages for low-skilled workers as the dependent variable. The omitted period is 1919–1920. Column (1) reports the baseline specification. Columns (2) and (3) add the 1910 foreign-born share and all lasso-selected controls listed in Table D7, allowing the controls to enter either as linear time trends or interacted with the post-1920 indicator. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Aggregation of Labor Markets. A potential concern with our definition of a labor market as a geography–industry (RLM–industry) pair is that it assumes both geographic and inter-industry differences in exposure to immigration restrictions generated similar wage dynamics. Yet, low-skilled workers may have been able to move easily between industries within the same geographic

area, implying that wage equalization across industries within a region could occur more rapidly than wage equalization across regions with differing exposure levels. In this case, our baseline specification—which treats each geography–industry cell as a separate labor market—could overstate within-region variation in exposure and understate the degree of adjustment through worker mobility.

To evaluate this possibility, we collapse the RLM–industry data to the geographic labor market (jurisdiction) level. Specifically, we construct an exposure measure for each RLM assuming no within-region variation across industries and compute the employment-weighted mean wage across all industries in that RLM and year. As reported in Table D10, higher exposure to immigration restrictions is associated with a decline in the recent-immigrant share and an increase in mean real wages for low-skilled workers. Both effects are smaller than in our baseline estimates (Tables 2 and 4 in the main text), suggesting that within-region wage equalization across industries was limited and that our main specification does not overstate inter-industry adjustment.

Column (3) of Table D10 aggregates the data to the region–industry level. This specification collapses across local labor markets within the same census region and industry. The resulting estimates remain consistent with the main results: higher-exposure region–industry cells experience larger declines in immigrant shares and larger increases in wages after 1920. This indicates that the estimated effects are not driven by idiosyncratic variation across individual RLMs, but are also present when the analysis is conducted at a more aggregated region–industry level.

Taken together, these aggregation exercises confirm that our main results are not sensitive to the level of analysis. Whether exposure is defined at the RLM–industry, RLM, or region–industry level, more exposed markets consistently exhibit lower immigrant employment shares and higher wages.

Table D10. Robustness of Estimates to Aggregation Level

Aggregation:	**RLM-Industry (Baseline)** **(1)**	**RLM** **(2)**	**Region-Industry** **(3)**
A. Recent Immigrants Share:			
Quota Exposure X Post 1920	-0.496***	-0.394***	-0.755***
	(0.0470)	(0.0577)	(0.151)
Pre-1920 Dependent Mean	0.089	0.070	0.130
Observations	1,197	375	174
R-Squared	0.892	0.862	0.928
B. Log Mean Hourly Wage:			
Quota Exposure X Post 1920	0.280**	0.242**	0.224***
	(0.135)	(0.104)	(0.0713)
Pre-1920 Dependent Mean	0.337	0.348	0.528
Observations	843	299	122
R-Squared	0.991	0.968	0.998
Quota Exposure - Mean	0.106	0.116	0.228
Quota Exposure - SD	0.084	0.134	0.185
Number of Panel Units	399	125	58

Notes: This table reports coefficients on the interaction between quota exposure and a post-1920 indicator for different levels of sample aggregation. Column (1) reports the baseline RLM-industry specification, column (2) aggregates observations to the RLM level, and column (3) aggregates observations to the region-industry level. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable. RLM-industry pairs are employment-weighted to construct measures at the RLM or region-industry level. Panel B uses log mean hourly wage as the dependent variable. RLM-industry pairs are first employment-weighted, and the data are then collapsed to the decadal level by assigning 1910–1914 to 1910, 1915–1924 to 1920, and 1925–1930 to 1930. All specifications include panel-unit fixed effects. Column (1) includes industry-year, division-year, and dataset-year fixed effects. Column (2) includes year and division-year fixed effects. Column (3) includes industry-year and region-year fixed effects. Robust standard errors are clustered at the panel-unit level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

We next examine whether the results are sensitive to overlapping wage observations across cells or sources. These overlaps arise in two ways. First, within the manufacturing and mining sample, some RLM–industry–year cells contain multiple wage reports because related occupations or wage series are reported separately. As shown in Table A2, this occurs in parts of the BLS manufacturing and mining data where multiple low-skilled occupations or closely related wage series are available within the same harmonized industry. For example, coal mining includes multiple low-skilled occupations, and some related manufacturing series map to the same RLM–industry exposure cell. Second, the two construction sources may cover the same city in the same year. Table A3 documents the coverage of the *American Contractor* data, which overlaps in some city-years with

the BLS Construction Union data.

In the baseline analysis, we retain these observations because they represent distinct wage reports rather than duplicate records. Nevertheless, to assess whether repeated coverage of the same local labor market and year affects the estimates, we re-estimate the main specifications after collapsing overlapping observations by taking mean wages within the relevant cells.

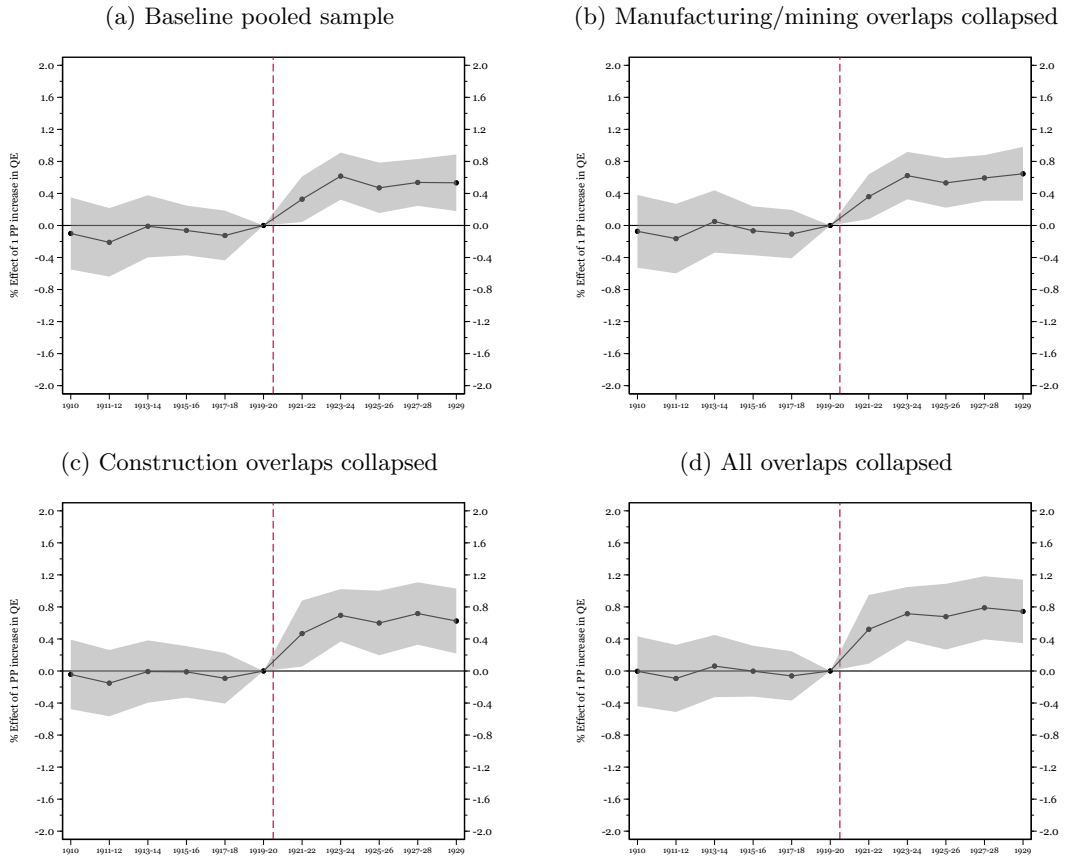
Table D11 reports the resulting DID estimates. Column (1) reproduces the baseline specification. Column (2) collapses overlapping observations in manufacturing and mining within repeated RLM–industry–year cells. Column (3) collapses overlapping observations across the two construction samples within city-year cells. Column (4) applies both adjustments. The estimates for both recent immigrant shares and log wages remain close to the baseline estimates and statistically significant. The corresponding wage event studies in Figure D5 show the same dynamic pattern: pre-1920 coefficients remain close to zero, while post-1920 coefficients rise in more exposed markets.

Table D11. Robustness to Collapsing Overlapping Wage Observations

		Overlapping Wage Reports Collapsed		
Sample Adjustment:	**Baseline**	**Manufacturing & Mining**	**Construction**	**All**
	(1)	**(2)**	**(3)**	**(4)**
A. Recent Immigrants Share:				
Quota Exposure X Post 1920	-0.496***	-0.474***	-0.528***	-0.506***
	(0.0470)	(0.0494)	(0.0506)	(0.0535)
Pre-1920 Dependent Mean	0.09	0.09	0.09	0.09
Observations	1,197	1,098	1,122	1,023
B. Log Mean Hourly Wage:				
Quota Exposure X Post 1920	0.532***	0.557***	0.602***	0.633***
	(0.128)	(0.130)	(0.162)	(0.167)
Pre-1920 Dependent Mean	0.26	0.26	0.25	0.25
Quota Exposure - Mean	0.10	0.10	0.10	0.10
Quota Exposure - SD	0.08	0.08	0.08	0.08
Number of Panel Units	399	366	374	341
Observations	2,973	2,793	2,871	2,691

Notes: This table reports the coefficients of the interaction between quota exposure and the post-1920 indicator. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses log mean hourly wages for low-skilled workers. Column (1) reports the baseline specification. Column (2) collapses overlapping observations in the manufacturing and mining sample by averaging observations within repeated RLM–industry–year cells. Column (3) collapses overlapping observations across the two construction samples by averaging observations within overlapping city-year cells. Column (4) applies both adjustments. The post-1920 indicator equals one in years after 1920, the last census year before the quota policy took effect. All specifications include RLM-industry, industry-year, and dataset-year fixed effects. Panel A also includes division-year fixed effects, while Panel B includes division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D5. Event-Study Robustness to Collapsing Overlapping Wage Observations - Pooled Sample



Notes: The figures plot event-study coefficients of the interaction between quota exposure and year-pair indicators. The outcome is the RLM-industry log mean hourly wage for low-skilled workers. Panel (a) reports the baseline pooled sample. Panel (b) collapses overlapping observations in the manufacturing and mining sample by averaging wages within repeated RLM–industry–year cells, while retaining the pooled-sample structure. Panel (c) collapses overlapping observations across the two construction samples by averaging wages within overlapping city-year cells, while retaining the pooled-sample structure. Panel (d) applies both adjustments. The omitted period is 1919–1920. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. The gray area shows 95% confidence intervals. Standard errors are clustered at the RLM-industry level.

World War I. Because World War I preceded the introduction of the immigration quotas, disentangling the effects of wartime disruptions from those of the subsequent policy change is a central identification concern. Regions that relied heavily on European immigrants experienced large inflows before 1915, sharp wartime declines, and further reductions following the implementation of the quotas. This pattern raises the possibility that the observed wage and employment effects reflect temporary wartime shocks rather than the long-term consequences of the 1920s immigration restrictions.

To assess this possibility, we first turn to the event-study estimates presented earlier. These results show no systematic increase in wages during the war years. The pre-1920 coefficients are flat and statistically insignificant, indicating that wages in high- and low-exposure markets evolved similarly before the policy change. Wage effects emerge only after 1920, when the quota system came into force, and remain persistently positive throughout the decade. The absence of

wage movements during the war period reinforces that our findings capture the causal impact of immigration restrictions rather than temporary disruptions in labor demand or supply related to wartime production and mobilization.

Table D12. Robustness to World War I Exposure

Exposure Measures:	**QE Exposure** (1)	**WWI Exposure** (2)	**QE + WWI Exposure** (3)
A. Recent Immigrants Share:			
Quota Exposure X Post 1920	-0.496***		-0.494***
	(0.0470)		(0.106)
WWI Exposure X Post 1920		-0.323***	-0.0009
		(0.0297)	(0.0604)
Pre-1920 Dependent Mean		0.089	
Quota Exposure - Mean		0.109	
Quota Exposure - SD		0.084	
WWI Exposure - Mean		0.183	
WWI Exposure - SD		0.124	
Observations	1,197	1,197	1,197
R-Squared	0.892	0.888	0.892
B. Log Mean Hourly Wage:			
Quota Exposure X Post 1920	0.532***		0.639***
	(0.128)		(0.124)
WWI Exposure X Post 1920		0.170	-0.132
		(0.104)	(0.0936)
Pre-1920 Dependent Mean		0.258	
Quota Exposure - Mean		0.100	
Quota Exposure - SD		0.082	
WWI Exposure - Mean		0.169	
WWI Exposure - SD		0.124	
Observations	2,973	2,973	2,973
R-Squared	0.979	0.979	0.980

Notes: This table reports coefficients on the interaction between exposure measures and the post-1920 indicator. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses log mean hourly wages for low-skilled workers. Column (1) includes the baseline quota exposure measure, column (2) replaces quota exposure with the World War I exposure measure, and column (3) includes both exposure measures. All specifications include RLM-industry, industry-year, and dataset-year fixed effects. Panel A includes division-year fixed effects, while panel B includes division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

We further test this interpretation using a more direct approach. Following Abramitzky et al. (2023), we construct a World War I exposure measure that captures the extent to which immigration

flows were restricted by the war. We then re-estimate our main specification including both the quota and WWI exposure measures:

$$y_{ijt} = \delta_{ij} + \gamma_{it} + \beta(QE_{ij} \times 1(t \geq 1920)_t) + \theta(WE_{ij} \times 1(t \geq 1920)_t) + \varepsilon_{ijt} \tag{24}$$

Table D12 reports the results for the pooled sample. Column (1) presents the baseline estimates, Column (2) replaces quota exposure with the WWI exposure measure, and Column (3) includes both exposure measures. When entered on its own, WWI exposure is associated with a decline in recent immigrant shares, but the coefficient is smaller than the corresponding quota-exposure estimate. Once both measures are included, the quota-exposure coefficient remains large and statistically significant, while the WWI-exposure coefficient falls essentially to zero. This indicates that the post-1920 decline in recent immigrant shares is captured primarily by quota exposure rather than by wartime exposure.

The wage results tell a similar story. WWI exposure alone is positively but imprecisely related to post-1920 wage growth. When both measures are included, the quota-exposure coefficient becomes slightly larger and remains statistically significant, while the WWI-exposure coefficient turns negative and remains statistically insignificant. Thus, conditional on quota exposure, WWI exposure has no systematic relationship with post-1920 wage growth. Taken together with the event-study evidence showing no wage increase during the war years, these results suggest that the wage and employment responses we document reflect the 1920s immigration restrictions rather than temporary wartime disruptions.

E Sample Selection and Composition Appendix

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; see Angrist et al. (1996)). 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 E1 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 2 in the main text for reference. In columns 2-4, we show our 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 E1, 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.

Table E1. Testing for Potential Bias from Potential SUTVA Violation

		Exclude Bottom X% Quota Exposure:			Exclude Top X% Out-Migration Rate:		
Sample:	**Baseline**	**X = 1**	**X = 5**	**X = 10**	**X = 5**	**X = 10**	**Top State Within Division**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
A. Recent Immigrants Share:							
Quota Exposure X Post 1920	-0.496***	-0.498***	-0.496***	-0.490***	-0.479***	-0.449***	-0.510***
	(0.0470)	(0.0471)	(0.0475)	(0.0479)	(0.0446)	(0.0387)	(0.0394)
Pre-1920 Dependent Mean	0.089	28.473	29.655	31.196	27.732	26.927	25.702
Quota Exposure - Mean	0.109	0.110	0.115	0.121	0.108	0.106	0.099
Quota Exposure - SD	0.084	0.083	0.082	0.080	0.084	0.084	0.079
Number of RLMs	125	123	118	117	119	112	102
Number of RLM-Industry Pairs	399	395	379	359	384	366	312
Observations	1,197	1,185	1,137	1,077	1,152	1,098	936
R-Squared	0.892	0.891	0.890	0.889	0.894	0.899	0.897
B. Log Mean Hourly Wage:							
Quota Exposure X Post 1920	0.532***	0.522***	0.548***	0.533***	0.575***	0.563***	0.475***
	(0.128)	(0.129)	(0.133)	(0.141)	(0.127)	(0.124)	(0.146)
Pre-1920 Dependent Mean	0.258	0.260	0.273	0.278	0.256	0.253	0.253
Quota Exposure - Mean	0.100	0.101	0.110	0.116	0.099	0.098	0.091
Quota Exposure - SD	0.082	0.082	0.080	0.078	0.083	0.085	0.079
Number of RLMs	125	123	118	117	119	112	102
Number of RLM-Industry Pairs	399	395	379	359	384	366	312
Observations	2,973	2,920	2,723	2,588	2,844	2,697	2,341
R-Squared	0.979	0.979	0.976	0.976	0.980	0.980	0.981

Notes: This table reports coefficients on the interaction between quota exposure QE and the post-1920 indicator for the pooled sample. Panel A uses the share of recent immigrants from quota-restricted countries as the dependent variable, while Panel B uses log mean hourly wages for low-skilled workers. Columns (2)–(4) exclude RLMs in the bottom 1, 5, and 10 percent of the quota-exposure distribution, respectively. Columns (5)–(7) use the Abramitzky et al. (2023) out-migration rate to high-exposure areas, aggregated to the state level, and exclude states in the top 5 percent, top 10 percent, and top state within each division of out-migration to high-exposure areas, respectively. All specifications include RLM-industry fixed effects. All specifications include RLM-industry, industry-year, and dataset-year fixed effects. Panel A includes division-year fixed effects, while panel B includes division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Sample Selection. We also examine whether changes in the composition of our analysis sample from the pre- to post-1920 period might bias our estimated wage effects. Specifically, we assess selection into the wage sample—that is, the likelihood of having a wage observation in our data—by comparing demographic characteristics of laborers in RLM–industry pairs included in our wage sample to those of laborers in RLM–industry pairs not observed in the wage data. These comparisons are constructed using the full-count population censuses for 1910, 1920, and 1930, which provide comprehensive information on worker demographics and occupational structure. We then compare differences in these characteristics across high- and low-exposure RLM–industry pairs to evaluate

whether the sample composition changed differentially following the immigration restrictions.

Table E2 examines differences in demographic composition between the wage sample and the broader population of laborers, as well as between high- and low-exposure RLM–industry pairs, using data from the 1910, 1920, and 1930 full-count censuses. Columns (1)–(3) report pre-1920 means for laborers in RLM–industry pairs included in the wage sample, those excluded from the sample, and the adjusted differences between the two groups after controlling for industry and state fixed effects. On average, RLM–industry pairs included in our wage analysis are slightly more exposed to immigration restrictions than those excluded, with a mean quota exposure of 0.11 versus 0.09. Before 1920, wage-sample laborers were more likely to be foreign-born (recent immigrant share of 0.16 vs. 0.13 and early immigrant share of 0.16 vs. 0.14, adjusted differences of 3 and 2.1 percentage points, respectively) and more likely to reside in urban areas (urban household rate of 0.61 vs. 0.55, an adjusted difference of 7.4 percentage points). These pre-1920 differences indicate that the wage-sample RLM–industry pairs were selected primarily along foreign-born and urban characteristics.

Columns (3)–(5) of Table E2 report difference-in-differences (DID) estimates that control for both industry and state fixed effects, isolating changes in sample composition from broader sectoral or geographic trends. Column (4) captures the overall post-1920 trend in workforce composition across all laborers, while column (5) presents the DID estimates comparing how the composition of workers in the wage sample evolved relative to that overall trend. The results indicate a 2.2 percentage-point decline in the immigrant share among wage-sample observations, suggesting that immigration restrictions disproportionately affected in-sample RLM–industry pairs relative to out-of-sample ones. At the same time, we observe a 1.8 percentage-point increase in the share of Black workers and a 1.9 percentage-point decline in the share of White workers, while the share of young workers aged 16–25 declined by about 1.7 percentage points. Lastly, while the initial difference in urban rates between in-sample and out-of-sample observations is significant, this gap remains stable post-1920, suggesting no composition change driven by urban status.

We next compare high- versus low-exposure RLM–industry pairs within the wage sample in columns (6)–(10) of Table E2, allowing us to assess whether composition shifted differentially across RLM–industry pairs with varying exposure to immigration restrictions. By construction, the difference in mean quota exposure between these groups is large (0.17 vs. 0.04): high-exposure RLM–industries are those above the median exposure value, while low-exposure pairs fall below it. Pre-1920 differences confirm that these groups were distinct even before the quota laws. High-exposure RLM–industry pairs had substantially higher recent and early immigrant shares (0.26 and 0.25, respectively, versus 0.06 and 0.07 in low-exposure cells) and were more urban (0.73 vs. 0.49). These contrasts reflect the pre-existing concentration of foreign-born labor in industrial and urban centers most affected by the quotas.

Post-1920, the DID estimates reveal sharp compositional adjustments. The share of recent immigrants fell dramatically in high-exposure areas—by 14.3 percentage points—while the Black share increased by 2.6 percentage points and the average age rose by 0.83 years. These changes are

consistent with a shift away from foreign-born workers toward older, native-born, and more racially diverse labor forces in regions most affected by the restrictions. The rise in Black worker shares is less concerning from the standpoint of bias, as Black workers typically earned lower wages during this period—implying that this shift would attenuate, rather than inflate, our estimated wage effects. Similarly, the modest increase in mean age is unlikely to meaningfully affect wages,. Consequently, the primary source of potential mechanical bias remains the sharp decline in immigrant shares in high-exposure labor markets.

Table E2. Changes in Sample Composition and Demographic Characteristics Across Wage Sample and Exposure Groups

Sample Split:	**Wage Sample vs. Not in Sample**					**High vs. Low Exposure**				
	Pre-1920 Means		**DID Results**			**Pre-1920 Means**		**DID Results**		
	Not in-Sample	**In-Sample**	**In-Sample Indicator**	**Post-1920 Indicator**	**DID Estimate**	**Low-Exposure**	**High-Exposure**	**High-Exposure Indicator**	**Post-1920 Indicator**	**DID Estimate**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**
Mean Quota Exposure	0.09	0.11				0.04	0.17			
Recent Immigrant (< 10 Years in U.S.)	0.13	0.16	0.030***	-0.093***	-0.022**	0.06	0.26	0.146***	-0.044***	-0.143***
			(0.007)	(0.010)	(0.009)			(0.013)	(0.007)	(0.010)
Early Immigrant (10+ Years in U.S.)	0.14	0.16	0.021***	0.006	0.012*	0.07	0.25	0.064***	-0.005	0.047***
			(0.007)	(0.006)	(0.007)			(0.013)	(0.004)	(0.010)
Black Indicator	0.19	0.16	0.003	-0.009	0.018***	0.27	0.05	-0.002	-0.004	0.026***
			(0.007)	(0.006)	(0.006)			(0.016)	(0.008)	(0.008)
White Indicator	0.81	0.84	-0.002	0.009	-0.019***	0.73	0.95	-0.003	0.004	-0.026***
			(0.008)	(0.006)	(0.006)			(0.017)	(0.008)	(0.008)
Urban Household	0.55	0.61	0.074***	0.020**	-0.016	0.49	0.73	0.106***	0.015	-0.021
			(0.016)	(0.010)	(0.012)			(0.017)	(0.011)	(0.013)
Age	33.8	33.2	-0.199	1.084***	0.369*	32.2	34.1	-0.210	1.038***	0.830***
			(0.145)	(0.188)	(0.206)			(0.236)	(0.133)	(0.169)
Age 16 to 25 Indicator	0.33	0.35	0.002	-0.009	-0.017***	0.40	0.31	-0.023***	-0.029***	0.006
			(0.004)	(0.005)	(0.006)			(0.008)	(0.005)	(0.006)
Single Indicator	0.44	0.47	-0.001	-0.071***	0.008	0.47	0.47	0.008	-0.064***	0.003
			(0.005)	(0.007)	(0.007)			(0.010)	(0.006)	(0.009)
Head of Household Indicator	0.49	0.44	-0.005	0.065***	0.007	0.45	0.43	-0.040***	0.054***	0.036***
			(0.006)	(0.007)	(0.007)			(0.010)	(0.007)	(0.008)

Notes: Columns (1)–(2) report pre-1920 means for laborers in region–industry cells that are included in the wage sample analysis versus laborers in region-industry cells that are not included in the wage analysis sample. Columns (3)–(5) present coefficients from a difference-in-differences regression that adjusts for state and industry fixed effects. Columns (6)–(7) report pre-1920 means for laborers in region–industry cells with below-median (low exposure) versus above-median (high exposure) values of the quota exposure measure. Columns (8)–(10) present coefficients from a difference-in-differences regression that adjusts for state and industry fixed effects. Each region–industry pair appears in the 1910, 1920, and 1930 U.S. censuses. The sample is restricted to laborers. "Recent immigrant" refers to individuals in the U.S. for fewer than 10 years, while "Early immigrant" refers to those in the U.S. for 10 or more years. Urban status, household headship, and other demographic indicators are coded at the individual level from census data. Robust standard errors, clustered at the state level, in parentheses. *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$.

Guided by these findings, we conduct formal tests of sample selection and potential selection bias in our wage estimates, reported in Table E3. These exercises assess whether the initial differences and compositional changes we document in Table E2 translate into systematic differences in the likelihood of being included in the wage sample or in a high-exposure RLM–industry pair, and

whether such differences bias the estimated effect of immigration restrictions on wages. We focus on two complementary approaches: (i) an inverse probability weighting (IPW) correction that re-weights the wage sample to match the full population of laborers, and (ii) a pseudo-Heckman selection model that explicitly models selection into the wage sample and into high-exposure.

The first set of estimates, shown in columns (2) and (5) of Table E3, apply inverse probability weights based on the estimated probability that a laborer in a given RLM–industry cell is observed in the wage sample. These probabilities are estimated using a random-forest model on the 1910–1930 full-count censuses, where the outcome is a binary indicator for being a laborer in a sample RLM–industry pair in year t (1910, 1920, or 1930), and zero otherwise. The predictors include the 1910 RLM–industry characteristics listed in Table D7, along with year, industry, state, and dataset fixed effects. Reweighting the wage data using these predicted inclusion probabilities leaves the estimated wage effects virtually unchanged, suggesting that sample selection—particularly differences in urbanization and immigrant presence—does not materially affect our treatment estimates.

Columns (3) and (7) of Table E3 report estimates from a pseudo-Heckman selection correction model, designed to account for potential unobserved factors driving inclusion in the wage sample. We refer to it as "pseudo" because, unlike a standard Heckman two-step model estimated at the individual level, our first-stage selection equation is estimated using individual-level data from the full-count censuses, while the second-stage wage regressions are estimated at the RLM–industry–year level. To bridge these levels of analysis, we aggregate the individual-level inverse Mills ratio (IMR) to its mean value within each RLM–industry–year cell before including it in the wage equation.

In the first stage, we estimate a probit model of selection into the wage sample at the RLM–industry level:

$$S^*_{prijt} = \beta \times QE_{ij} + \gamma^p \times X^p_{pijt} + \gamma^{RLM} \times X^{RLM}_{ij1910} + \delta_t + \pi_r + \omega_i + \varepsilon_{ijt} \tag{25}$$

where S^*_{prijt} is an indicator equal to one if individual p in census region r in RLM-industry pair ij in year t is a laborer in the wage sample and zero otherwise, QE_{ij} is RLM-industry ij exposure measure, X^p is a set of individual-level controls, X^{RLM} includes the 1910 RLM-industry characteristics listed in Table D7, and δ_t, π_r, and ω_i are year, region, and industry fixed effects, respectively.[4]

From this model, we compute the RLM-industry aggregated inverse Mills ratio (IMR), defined as

$$\bar{\lambda}_{ijt} = \frac{1}{N_{ijt}} \sum_{p \in ijt} \frac{\phi(\hat{S}_{prijt})}{\Phi(\hat{S}_{prijt})} \tag{26}$$

where N_{ij} is the number of individuals working as laborers in RLM-industry pair ij in year t, $\hat{S}_{prijt}$ is the fitted value from the probit model in Equation 25, $\phi(\cdot)$ and $\Phi(\cdot)$ denote the standard normal probability density function and cumulative distribution function, respectively. Intuitively, the inverse Mills ratio measures the expected value of the error term in the outcome equation

[4]Individual level controls include age, race, nativity, marital status, urban status, head of household indicator, currently in school indicator, and group quarters indicator.

conditional on being observed in the sample—capturing the extent to which unobserved factors jointly affect selection and outcomes. A significant coefficient on the IMR in the second stage would thus indicate that selection into the observed wage sample is non-random and correlated with unobserved determinants of wages.

Columns (3) and (7) of Table E3 report estimates from the pseudo-Heckman correction for selection into the wage sample. The IMR coefficient is negative and marginally significant in the DID specification, but smaller and statistically insignificant in the event-study specification. This provides at most modest evidence of negative selection into the wage sample: RLM–industry cells more likely to appear in the wage data tend to have slightly lower unobserved wage components. Importantly, accounting for this selection has little effect on the estimated treatment effects. The DID coefficient on quota exposure remains nearly unchanged, and the event-study coefficients continue to show no differential pre-trends and a persistent post-1920 wage increase in more exposed RLM–industry pairs.

Columns (4) and (8) of Table E3 extend the pseudo-Heckman framework to examine selection into high-exposure RLM–industry pairs, using only individuals observed in the wage sample. The first-stage probit models the probability that a laborer belongs to a high-exposure RLM–industry pair (above the median quota exposure value),

$$H^*_{prijt} = \gamma^p \times X^p_{pijt} + \delta_t + \pi_r + \omega_i + \varepsilon_{ijt} \tag{27}$$

where H^*_{prijt} is an indicator equal to 1 if individual p is a laborer in RLM-industry pair ij with above median exposure in year t and zero if he is a laborer in an RLM-industry pair in year t with below median exposure. As before, we compute an individual-level inverse mills ratio and aggregate it to the RLM-industry-year level by averaging the individual-level observations.

As before, we aggregate the resulting IMR to the RLM–industry–year level and include it in the wage regressions. The estimated IMR coefficients provide only limited evidence of selection into the high-exposure sample. The coefficient is negative in the DID specification but statistically insignificant, while the event-study specification yields a larger negative and statistically significant coefficient. This suggests that RLM–industry cells more likely to be classified as highly exposed may have somewhat lower unobserved wage components. However, accounting for this selection has little effect on the main estimates. The DID coefficient on quota exposure remains stable, and the event-study estimates continue to show no systematic differential pre-trends and a persistent post-1920 wage increase in more exposed RLM–industry pairs.

Overall, the results provide little evidence that selection materially affects our estimated wage effects. Although some specifications indicate mild negative selection into the wage sample or into high-exposure markets, the correction terms do not meaningfully change the main treatment estimates. The DID coefficients remain stable, and the event-study estimates continue to show no systematic differential pre-trends and a persistent post-1920 wage increase in more exposed RLM–industry pairs. If anything, the negative selection patterns suggest that our estimates may slightly understate the wage response to the immigration restrictions.

Table E3. Robustness of Wage Effects to Sample Selection Models

Estimation Model:	Pooled DID				Event Study			
	Baseline	IPW	Pseudo Heckman Selection to Sample	Pseudo Heckman Selection to Exposure	Baseline	IPW	Pseudo Heckman Selection to Sample	Pseudo Heckman Selection to Exposure
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Quota Exposure X Post 1920	0.532***	0.535***	0.544***	0.592***				
	(0.128)	(0.128)	(0.128)	(0.129)				
Inverse-Mills Ratio			-0.239*				-0.0731	
			(0.133)				(0.137)	
Inverse-Mills Ratio: Exposure Selection				-0.429				-0.981***
				(0.308)				(0.318)
Quota Exposure X Year 1910					-0.100	-0.0935	-0.109	-0.270
					(0.229)	(0.228)	(0.233)	(0.230)
Quota Exposure X Year 1911–12					-0.211	-0.205	-0.219	-0.337
					(0.218)	(0.217)	(0.220)	(0.218)
Quota Exposure X Year 1913–14					-0.0102	-0.00441	-0.0158	-0.0830
					(0.197)	(0.197)	(0.199)	(0.196)
Quota Exposure X Year 1915–16					-0.0620	-0.0570	-0.0655	-0.163
					(0.158)	(0.159)	(0.158)	(0.156)
Quota Exposure X Year 1917–18					-0.125	-0.122	-0.124	-0.166
					(0.158)	(0.158)	(0.158)	(0.154)
Quota Exposure X Year 1921–22					0.328**	0.338**	0.329**	0.385***
					(0.145)	(0.146)	(0.145)	(0.146)
Quota Exposure X Year 1923–24					0.616***	0.621***	0.617***	0.735***
					(0.149)	(0.149)	(0.150)	(0.150)
Quota Exposure X Year 1925–26					0.470***	0.474***	0.474***	0.523***
					(0.160)	(0.160)	(0.161)	(0.160)
Quota Exposure X Year 1927–28					0.537***	0.539***	0.542***	0.653***
					(0.149)	(0.148)	(0.150)	(0.156)
Quota Exposure X Year 1929					0.533***	0.538***	0.536***	0.704***
					(0.180)	(0.179)	(0.181)	(0.193)
Pre-1920 Dependent Mean	0.258	0.258	0.258	0.258	0.258	0.258	0.258	0.258
Quota Exposure - Mean	0.100	0.100	0.100	0.100	0.100	0.100	0.100	0.100
Quota Exposure - SD	0.082	0.082	0.082	0.082	0.082	0.082	0.082	0.082
Number of RLMs	125	125	125	125	125	125	125	125
Number of Industries	23	23	23	23	23	23	23	23
Number of RLM-Industry Pairs	399	399	399	399	399	399	399	399
Observations	2,973	2,973	2,973	2,973	2,973	2,973	2,973	2,973
R-Squared	0.979	0.979	0.980	0.980	0.976	0.976	0.976	0.976

Notes: This table reports robustness checks addressing potential non-random selection into the wage sample. The dependent variable is log mean hourly wages for low-skilled workers at the RLM-industry level. Columns (1)–(4) report pooled difference-in-differences estimates, where the coefficient on quota exposure interacted with the post-1920 indicator captures the average post-policy wage effect. Columns (5)–(8) report event-study estimates using year-pair indicators, with 1919–1920 as the omitted period. Columns (1) and (5) report the baseline specification. Columns (2) and (6) reweight observations using inverse probability weights constructed from first-stage estimates of the probability of observing wage data in an RLM-industry cell. Columns (3) and (7) add an inverse-Mills ratio from a pseudo-Heckman correction for selection into the wage sample. Columns (4) and (8) add an inverse-Mills ratio from a pseudo-Heckman correction for selection into exposure. All specifications include RLM-industry, industry-year, dataset-year, and division-by-period fixed effects, where periods are defined as 1910–1914, 1915–1924, and 1925–1930. Robust standard errors, clustered at the RLM-industry level, are shown in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Immigrant Wage Penalty. In this section, we describe in more detail our assessment of the extent to which the wage effects we observe are driven by the falling immigrant share among laborers in the more exposed RLM-industry pairs, following the imposition of quotas.

As described in the text, we produce estimates of the native immigrant wage differential from three data sources. The first is the report of the US Immigration Commission or Dillingham Commission, established by Congress in 1907 to conduct an investigation into "the subject of im-

migration". 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 (USIC,1911). To measure the immigrant-native wage differential for each industry, we use the reported industry-wide ratio of the average wage for foreign-born male employees to the average wage for native-born workers. 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.

One caveat of the Dillingham Commission wage data is that 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. Thus, we complement our analysis with data from the 1915 Iowa State Census Project, which provides one of the only individual-level earnings datasets available before 1920 (Goldin and Katz, 2000). The census covers a representative sample of Iowa's population and includes information on age, nativity, education, occupation, and annual income. Although confined to a single state, the Iowa data allow us to estimate the immigrant wage penalty using pre-1920 individual-level wage observations.

After restricting the sample to men aged 16–65 working as laborers with reported annual earnings, we estimate regressions of the form:

$$ln(w_{ic}) = \beta immig_i + \Gamma X_i + \gamma_c + \varepsilon_{ic} \tag{28}$$

where $ln(w_{ij})$ is the log annual earnings of individual i residing in county c, $immig_i$ is an indicator equal to one if individual i is foreign-born arriving at the US within the last 10 years (recent immigrant), X_i is a set of additional individual-level covariates, and γ_c are county fixed-effects.[5] Although industry information is unavailable, we estimate separate models for farm laborers and coal-mine laborers, allowing partial differentiation across broad sectors.

Our Iowa estimates indicate an immigrant wage penalty of 7.1 percentage points on average. The adjusted penalties for farm laborers and coal miners are slightly larger, around 9 percentage points. Overall, these Iowa results align closely with the nationwide estimates from the Dillingham Commission and provide a valuable pre-1920 benchmark confirming that immigrants systematically earned lower wages than natives.

Despite that, neither the Dillingham Commission nor the Iowa State Census report wage differentials for all the industries covered in our analysis, so we also estimate the immigrant wage penalty using the 1940 Full Count Population Census, the first Census to report actual wages. De-

[5]The additional covariates include an indicator for early-immigrant status (more than 10 years in the US), age and age-squared, indicators for marital status (single, married, divorced, widowed, missing), race indicators (white, black, other), and education attainment categories (no education, common school. grammar school, high school, and college or more).

spite being reported many years after the restrictions went into effect, using the full count Census allows us to estimate wage penalties for all the industries in our data. 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:

$$ln(w_{ijc}) = \sum_j \beta_j immig_{ijc} + \Gamma X_i + \gamma j + \delta_c + \varepsilon_{ij} \tag{29}$$

where $ln(w_{ijc})$ is the log hourly wage of individual i in industry j in county c, β_j are industry indicators, $immig_{ijc}$ is a recent immigrant indicator (arrived within the last 5 years), X_i is a set of additional individual-level covariates, γ_j are industry fixed effects, and δ_c are county fixed effects.[6]

Our estimates range from 5 percent in the agriculture sector to as large as 70 percent in the shoes industry. The aggregated estimate from the 1940 Census is 9 percent.[7]

Assessing the Composition Effects Channel The main text describes how we use the approach of Ager et al. (2024) to evaluate the extent to which the observed post-1920 wage increases might reflect mechanical composition effects arising from the sharp decline in immigrant shares.

Panel (a) of Figure 13 in the main text 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, suggesting that composition shifts are not the primary explanation for the estimated wage effects in most RLM-industry pairs. Table E4 presents the table analogue of panel (a) of Figure 13 in the main text when assuming a constant 10 percent immigrant wage penalty. In the median RLM–industry pair, composition effects explain 6.2 to 9.3 percent of total wage changes, depending on whether we use recent-immigrant or overall foreign-born shares. Over two-thirds of pairs exhibit effects below 10 percent, and more than 90 percent fall below 25 percent.

[6]The additional covariates include an indicator for early-immigrant status (more than 5 years in the US), age and age-squared, indicator single status, head of household status, urban household, farm household, race indicators (white, black, hispanic, other), and five education attainment categories.

[7]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 5.9% in that specification.

Table E4. Composition Effects Estimates and Distribution Across RLM-Industry Pairs

	DID Estimate	DID – Wage	Comp. Effect (%, Median)	Composition Effect			
				$< 10\%$	$< 15\%$	$< 25\%$	$\geq 25\%$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Recent Immigrants Share – Quota Countries	-0.496***	0.532***	7.19	198	232	279	25
	(0.0470)	(0.128)		{65%}	{76%}	{92%}	{8%}
Foreign-Born Share	-0.307***	0.532***	10.70	146	191	258	46
	(0.0428)	(0.128)		{48%}	{63%}	{85%}	{15%}

Notes: This table reports calculations of the share of the predicted increase in hourly wages due to the immigration quotas that is accounted for directly by changes in the workforce share of recent immigrants from quota-restricted countries or foreign-born workers in general. Changes in wages are calculated from 1910–1920, before the quotas, to 1930, the first post-quota year with census data on foreign-born shares. Column (1) reports the estimated effect of quota exposure on immigrant shares. Column (2) reports the estimated effect of quota exposure on log mean hourly wages. Column (3) reports the median share of the wage effect accounted for by composition. Columns (4)–(6) report the number of RLM-industry pairs, with shares in curly brackets, where composition accounts for less than 10%, 15%, and 25% of the wage effect, respectively. Column (7) reports the number and share of RLM-industry pairs where composition accounts for 25% or more of the wage effect. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

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