FEDERAL RESERVE BANK OF MINNEAPOLIS
Community Development and Engagement

WORKING PAPER

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May 2023

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Keywords: Occupational licensing, Immigrants, Labor market outcomes

90 Hennepin Avenue, Minneapolis, MN 55401
minneapolisfed.org/community-development-and-engagement/working-papers

Occupational licensing as a barrier to entry for immigrants*

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May 5, 2023

Abstract

Affecting about one-fifth of U.S. workers, occupational licensing is a core labor market institution. However, despite considerable policy focus on the uneven burden of licensure across groups, relatively little is known about the differential impacts of licensure policies by nativity and race/ethnicity. We explore demographic disparities in licensure rates, using variation in licensure within states and occupations to estimate its effects on employment. We find that licensure reduces foreign-born employment in a state-occupation pair by nearly 20 percent relative to native-born employment. Similar effects are evident for Asian, Black, and Latino workers overall, but the effects on Asian and Latino employment are driven largely by foreign-born workers. Wage premiums for immigrants are correspondingly larger than for native-born workers, consistent with the interpretation that licensure requirements constitute a disproportionate barrier to the labor supply of immigrants.

*All findings, views, and conclusions expressed in this paper are those of the authors and not necessarily those of the Federal Reserve Bank of Minneapolis or anyone else in the Federal Reserve System and should not be construed to represent any official USDA or U.S. government determination or policy. This research was supported in part by the U.S. Department of Agriculture, ERS.

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1 Introduction

Occupational licensure requirements are an important determinant of access to occupations and, in turn, labor market opportunity. The classic normative justification for licensure is that the exclusion of unqualified individuals from an occupation can generate public health and safety benefits (Shapiro 1986). Whether or not this is true in any given case, licensing may also serve as a barrier to entry, generating licensure wage premiums and shifting employment out of the licensed sector. In this paper, we tackle the related question of whether licensure requirements have implications for the demographic composition of licensed employment: does the barrier to entry have larger effects on some groups than others? In particular, does licensure disproportionately reduce foreign-born employment relative to any reductions in native-born employment?

Native-born workers are substantially more likely to be licensed, at 20.3%, compared with foreign-born workers (13.9%). A few explanations could contribute to these licensure disparities. First, factors driving occupational choices including differences in worker preferences, educational investments, employment networks, and labor market discrimination indirectly affect licensure rates by generating occupational segregation. Some, but not all, of these worker differences relate to a worker's fitness for licensed practice. Second, licensure disparities could exist because of aspects of licensure requirements that are especially difficult for immigrants to navigate (but that do not relate to fitness for practice). For example, higher levels of native-born wealth could make it easier for those workers to pay licensing fees. Inadequate recognition of credentials obtained abroad could also disproportionately burden immigrants (Council of Economic Advisers, U.S. Department of Labor, and U.S. Department of the Treasury 2015; Council of State Governments 2022). Those two examples are useful for distinguishing how groups of workers may respond differently to the same barriers (e.g., licensure fees that are equal for everyone) or may actually face different barriers (e.g, foreign-trained licensure applicants whose credentials are not accepted on par with U.S.-trained applicants).

To shed light on nativity disparities in licensure, and to better understand licensure policy, we estimate the effects of licensure policy on foreign-born employment in a given state and occupation. We exploit variation in licensing rates across state-occupation pairs to explain variation in the foreign-born employment of those state-occupations. We find that state-occupations with higher licensing shares tend to have lower employment (and lower shares) of foreign-born workers. Highly licensed state-occupations see an 18% reduction in employment of foreign-born workers above and beyond any licensure effects on native-born employment.[1]

Workers who immigrated to the U.S. as adults often have professional credentials and training obtained abroad—an important difference from those who immigrate as children. When we implement the same specification for those who immigrated to the U.S. as children, we estimate a quantitatively similar relationship between licensure policy and the foreign-born employment share. This would seem to suggest that inadequate recognition of foreign-born worker training (by licensing authorities) cannot be the only explanation for the negative effects we estimate. The same likely holds true for the role of English proficiency (which would be less of a barrier for child arrivals).[2] Other potential explanations for our findings include lower levels of financial resources and time available to foreign-born workers for the payment of fees and completion of required training.

Consistent with Cassidy and Dacass [2021], we also estimate substantially larger wage premiums for foreign-born workers than for native-born workers. Under the standard interpretation of licensure wage premiums, and in conjunction with our finding of negative employment effects, this implies that licensure is acting as a larger barrier to entry for foreign-born workers than for native-born workers.

Finally, we explore the employment effects of licensing for Asian, Black, and Latino workers, largely in order to disentangle the overlapping roles of nativity and race. In our sample, 82.6% of foreign-born workers are people of color, and 41.5% of workers of color are foreign-born.[3] As with immigrants, we find a negative effect of licensure on the employment of workers of color in highly licensed state-occupations, ranging from -21 to -26%. However, when we conducted the same analysis for *native-born* workers only, the results were substantially different. While the effect for Black workers' employment was largely unchanged, coefficients for Asian and Latino workers' employment became smaller and statistically insignificant, implying that immigrant-specific barriers are driving the licensure effects for those groups.

2 Data

We use monthly, person-level Current Population Survey (CPS) microdata—obtained from IPUMS-CPS (Flood, King, Rodgers, Ruggles, Warren, and Westberry 2022)—spanning Jan-

[1] Throughout, we characterize estimates in percent terms corresponding to the log point estimates provided in the tables. Where our estimates are relatively large, those percent estimates differ somewhat from their log point equivalents, i.e., percent changes are not well-approximated by log points in those instances.

[2] However, Cassidy and Dacass [2021] find that controlling for English proficiency is an important factor in explaining lower rates of licensure among immigrants.

[3] These calculations are based on data for January 2016–December 2022 in the CPS.

uary 2016 through December 2022. These data provide demographic and labor market outcomes as well as information about the immigration/nativity status of an individual, including years since immigration and age at entry to the United States. The CPS also provides information on whether individuals have an occupational credential issued by a local, state, or federal government and whether that credential is required for one's job. In order to consider a worker as licensed, we require that both be true.

Throughout the analysis, we drop observations with imputed values of licensed status, imputed values of nativity, and (in wage premium regressions) imputed earnings values. Final person weights are used to calculate our summary statistics. Earning weights are used whenever calculating summary statistics of wages. Samples are limited to employed workers 16 years of age and older.

Hourly wages are calculated following the "NBER definition." When hourly wages are provided directly by respondents, those are used, but where only weekly earnings are provided, we divide those earnings by usual weekly hours to calculate the hourly wage.

Table 1 provides summary statistics for our sample showing how licensed and unlicensed workers differ. Licensed workers are more likely to be female, White, and highly educated, and they have higher average wages.

An ideal dataset to study the effects of licensure policy would contain comprehensive policy information on licensure requirements by state and occupation. In the absence of this information, we create a measure for licensure requirements using the share of workers in a state-occupation who report being licensed in the CPS.[4] One complication with this approach is the imperfect mapping between Census-defined occupations (which are units created for economic measurement) and state-defined occupations (which are units created for regulatory purposes). While our policy variable does not correspond exactly to the boundaries of legal occupations, it does correspond precisely to occupation definitions used for our dependent variables.

3 Demographic disparities in licensure rates

3.1 Nativity disparities

Foreign-born workers are 6.4 percentage points less likely to be licensed than native-born workers (20.3% of whom are licensed), as shown in Figure 1. Moreover, foreign-born workers

[4]When our dependent variable is employment of immigrants, we use native-born workers only to calculate our licensing measures. When our dependent variable is employment of a specific racial group, we use White workers only to calculate our licensing measure. This avoids issues whereby measurement error could produce spurious associations between those policy measures and our dependent variables.

of color are 2.9 percentage points less likely to be licensed than native-born workers of color. After adjusting for education, age, and gender (but not race), we find that foreign-born workers overall are (again) 6.4 percentage points less likely to be licensed than native-born workers, such that a given foreign-born worker is 31% less likely to be licensed than a similar native-born worker. This is broadly consistent with the findings of Cassidy and Dacass [2021], who find that foreign-born workers are 5.0 percentage points less likely to be licensed (after adjusting for their own set of controls).

Disparities in licensure rates between foreign-born and native-born workers increase with education. Only 26.1% of foreign-born workers with more than a bachelor's degree are licensed, well short of the 45.3% licensure rate of native-born workers with the same education level. (See Figure 2).

What explains these differences in licensure rates? Licensure policy may be responsible, acting through several possible channels. The monetary and opportunity costs of obtaining a license may be more of a barrier for foreign-born workers than for native-born workers. Low levels of wealth (Cobb-Clark and Hildebrand 2006), family commitments, and similar factors could all contribute to lower licensure rates. Relatedly, foreign-born workers could have extra difficulty understanding and navigating licensure pathways, in some cases because of language barriers. Another possibility is that foreign-born workers with some or all of the relevant training still face difficulties becoming licensed. Training, education, and credentials obtained abroad are often not recognized by state licensing authorities (Council of Economic Advisers, U.S. Department of Labor, and U.S. Department of the Treasury 2015). For example, a dental hygienist trained abroad is barred in some states from licensed practice even if they hold equivalent experience and passing exam scores (Little Hoover Commission 2016).

Differences in licensure rates could also be due to factors correlated with licensure but not caused by it. Foreign-born workers may simply have different work preferences, education, and employment experiences in their countries of origin that lead them to disproportionately enter certain occupations. Furthermore, the priorities of the U.S. immigration system (e.g., the H-1B visa rules) influence the occupational mix of foreign-born workers. One result is that highly educated foreign-born workers are especially likely to be in computer and mathematical occupations, like computer scientists or programmers, most of which are unlicensed. While they make up 18.3% of highly educated workers (those with a bachelor's degree or more) overall, 34.8% of highly educated workers in computer and mathematical occupations are foreign-born.

Figure 3 shows the occupations where immigrants work, in descending order of median wage. The largest share of licensed foreign-born workers—about one-quarter—are healthcare practitioners, which includes dentists, respiratory therapists, licensed practical nurses, and many other occupations. There are also many foreign-born workers licensed in personal care and service occupations, like the manicurists studied in Federman et al. [2006]. Those authors examined variation across states in the licensing rules that applied to manicurists, finding that English proficiency requirements and high required training hours were barriers to Vietnamese immigrants and a cause of lower overall manicurist employment.

Our identification strategy, described later in the paper, will accommodate occupation and state patterns in immigration that correlate with licensure rates—at the state or occupation levels—but are not attributable to licensure policy.

3.2 Racial disparities

Because 82.6% of foreign-born workers in our sample are people of color, and 41.5% of workers of color are foreign-born, it is important to examine racial disparities alongside nativity disparities.[5] Rates of licensing across racial and ethnic groups in the United States are indeed substantially different, with members of all non-White groups identifiable in Current Population Survey data being less likely to be licensed than White workers. Latino workers (11.8% of whom are licensed) in particular are much less likely to be licensed than White workers (21.9%). Asian, Black, and American Indian and Alaska Native workers are also less likely than White workers to be licensed, by 5.1, 5.0, and 4.5 percentage points, respectively.

As with nativity disparities, the racial gaps in licensing persist after adjusting for differences in education. Figure 4 plots shares of licensed and unlicensed workers, within a given educational group, by race and ethnicity. By contrast to previous figures, Figure 4 shows the share of all licensed (or unlicensed, in separate bars) workers who are members of a given group. Within the group of workers who have a high school degree or less, White workers constitute 52.2% of all unlicensed workers but 63.7% of licensed workers. By contrast, Latino workers within the same education group make up a smaller share of licensed workers (17.5%) than of unlicensed workers (29.0%). At the other end of the education distribution, Asian workers with more than a four-year college degree constitute 16.1% of all unlicensed workers in that education group but only 7.4% of licensed workers.

[5] This section draws from Boesch, Lim, and Nunn [2022b] and Boesch, Kokodoko, and Nunn [2022a].

4 Empirical strategy

4.1 Licensure and foreign-born employment

To move beyond the descriptive analyses presented above, we now take a different approach that exploits variation in how licensing is applied across states and occupations. The core question we address is the extent to which licensure policy affects the ability of foreign-born workers to enter a given occupation within a state. Specifically, we seek to explain how the licensure status of an occupation in a state influences the number—and implicitly the share, because we condition on native-born employment—of foreign-born workers within a given state-occupation.

To proxy for the licensing policy of an occupation in a given state, we estimate the share of native-born workers in that state and occupation who report a government-issued license that is required for their jobs. Because individual licensing data have only become available relatively recently, we pool observations from 2016 to 2022. In our sample, the licensed share of a state-occupation follows the distribution (not weighted by state-occupation employment) shown in figure 6.[6] Inspecting that histogram, we see natural cutoffs at 10%, 33%, and 66%. In particular, many state-occupations have only a very small share of workers who report being licensed—potentially due to measurement error in workers' responses. We consider these state-occupations to be unlicensed.

State-occupations with higher licensed shares often reflect a mix of legally distinct occupations that do not align perfectly with detailed Census occupation codes, generating licensed shares well above 0% but well below 100%. In addition, there is likely a degree of measurement error in workers' responses to the CPS licensing questions. Our strategy—using bins of state-occupation licensed share rather than a continuous measure—tries to accommodate this error by flexibly allowing for different effects depending on licensure share, with a focus on comparisons between the least- and most-licensed cells.

Next, we estimate the effect of licensure in a state-occupation on the employment of foreign-born workers, holding constant native-born employment. The baseline specification we implement is as follows:

$$ln(E^{fb}_{s,o}) = ln(E^{nb}_{s,o}) \cdot \beta_{nb} + lic^{10-33}_{s,o} \cdot \beta_{10-33} + lic^{33-66}_{s,o} \cdot \beta_{33-66} + lic^{66-100}_{s,o} \cdot \beta_{66-100} + \eta_s + \eta_o + \epsilon_{s,o} \tag{1}$$

where E is a population count of employed workers who are either unlicensed or licensed

[6] Note that we do not include state and occupation combinations with fewer than 100 observations in our sample; this removes about 60% of observations.

and either foreign-born or native-born. In other words, $E^{fb}_{s,o} = E^{u,fb}_{s,o} + E^{l,fb}_{s,o}$ and $E^{nb}_{s,o} = E^{u,nb}_{s,o} + E^{l,nb}_{s,o}$. s denotes state and o denotes occupation, and the unit of observation is a state-occupation pair. In the specification above, the dependent variable is the natural log of the total number of foreign-born workers employed in a state-occupation, and the first covariate is the natural log of the total number of native-born workers in that same state-occupation.

lic indicators are equal to 1 when the state-occupation licensed share falls within a given range, and η_s and η_o are state and occupation fixed effects, respectively. Licensing indicators are defined using only native-born workers. For example, $lic_{66-100} = (\frac{E^{l,nb}}{E^{l,nb}+E^{u,nb}} > 0.66)$. The coefficient β_{66-100} can be interpreted as the log point effect of licensure on foreign-born individuals' employment, conditional on native-born workers' employment.

State-occupations with less than or equal to 10% licensed share are the omitted category. In order to limit the influence of very small state-occupation cells, we omit cells with fewer than 100 total survey respondents (regardless of nativity). We also exclude cells with zero employment of foreign-born workers (all cells have nonzero native-born employment). Accordingly, our estimates should be interpreted as reflecting the effect of licensure on state-occupations with positive foreign-born employment.

In addition to total foreign-born employment, we consider several related dependent variables: the number of workers who arrived in the U.S. as adults (i.e., those 18 or older when they immigrated), the number of workers who arrived in the U.S. as children (i.e., those who were 17 or younger when they immigrated), and the number of second-generation immigrants (i.e., native-born workers with at least one foreign-born parent). To provide consistency across these three specifications, we exclude any state-occupation cells for which adult-arrival foreign-born employment is zero, child-arrival foreign-born employment is zero, or second-generation employment is zero, allowing implementation of our log-linear regressions. This set of restrictions gives us a common sample across our child-arrival-only, adult-arrival-only, and second-generation specifications. For child-arrival and adult-arrival specifications, we control for total log native-born employment. For the second-generation specification, we control for the log of native-born employment less second-generation employment.

For each specification, we implement a version that is unweighted (i.e., each state-occupation cell is equally weighted) and a version that weights by the sum of individual-level observations associated with a state-occupation. Because our analysis is at a level of observation that is aggregated beyond the individual level, it is appropriate to place more emphasis on the state-occupation cells with greater representation in our sample. However, differences in the unweighted and WLS results can indicate misspecification (e.g., heterogenous treatment effects), and we report both sets of results.

4.2 Licensure and employment by race

We take an analogous approach to estimating effects of licensure on employment of different racial and ethnic groups. Rather than log employment of immigrants, log employment in a given racial/ethnic group is the dependent variable. And rather than log employment of native-born workers, log employment of White workers is used as a covariate. As above, the coefficient β_{66-100} can be interpreted as the log point effect of licensure on a given group's employment, conditional on White employment.

Specifically, we implement the following regression:

$$ln(E^r_{s,o}) = ln(E^w_{s,o}) \cdot \beta_w + lic^{10-33}_{s,o} \cdot \beta_{10-33} + lic^{33-66}_{s,o} \cdot \beta_{33-66} + lic^{66-100}_{s,o} \cdot \beta_{66-100} + \eta_s + \eta_o + \epsilon_{s,o} \tag{2}$$

where r indexes racial/ethnic groups: Asian, Black, Latino, or White.[7]

4.3 Identification

The identifying variation in all the above regressions—whether focused on nativity or race—comes entirely from differences in the licensed share of different state-occupations. For example, the licensed share of massage therapists in Wisconsin is 76.9%, such that they fall into the 66%–100% license-share bin. In Minnesota, by contrast, the licensed share of massage therapists is 39.7%.

It is important to note that, because of the inclusion of state and occupation fixed effects in all specifications, differences in foreign-born workers' preferences over location or type of work and preparation for the labor market will not be conflated with the effect of licensure, if those differences manifest at the state or occupation levels. In other words, the fact that some states (or, separately, occupations) have higher or lower shares of immigrants does not threaten our identification. Any self-selection on the part of immigrants into states based on licensure policy, will contribute to the effect we estimate and is one mechanism through which licensure policy can affect foreign-born employment. An implication is that universally licensed occupations like physicians do not contribute to identification.

Another important point is that our approach identifies the effect of licensure on foreign-born employment *above and beyond* any effects on native-born employment. In other words,

[7] Groups are defined to be mutually exclusive, with our use of "Latino" defined as any individual who identifies as Hispanic as their ethnicity in the CPS. For example, someone identifying as both Latino and Asian will be coded as Latino. Those identifying as members of multiple races (e.g., "Asian" and "White") are assigned to the Other category, for which we do not report results. Because of insufficient sample size, results for American Indian and Alaska Native respondents are also not reported.

we do not attempt to identify the total change in foreign-born employment, but rather the change relative to any effect of licensure on native-born employment. If, for example, licensure reduces both native- and foreign-born employment by the same amount, our approach would yield an estimated licensure effect of zero. We believe this is a strength of our approach, because it weakens the necessary identification assumptions. If we were not controlling for native-born employment, it would be necessary to assume away the possibility that some unobserved (perhaps historical) factor affected total employment in a state-occupation (conditional on state and occupation fixed effects) as well as that state-occupation's licensed status. By controlling for native-born employment, the necessary assumption is weakened such that any omitted factor could affect total employment in a state-occupation, but—more narrowly—could not affect the ratio of foreign- and native-born employment.

There are two principal threats to identification. The first is simply measurement error associated with our proxy for licensed status of a state-occupation. Workers may be confused about their licensed status, assigned by the Census Bureau to the incorrect occupation, or have a license that is not legally required for the occupation they're currently working in. (The relevant CPS question is about whether a credential is required for the respondent's job, which may be an informal employer requirement rather than a legal restriction.) The second threat is the potential endogeneity of policy variation. If state legislatures are more likely to license an occupation for reasons related to the foreign-born employment share, then our estimate of the effect of licensure will be biased. For example, if foreign-born workers are less effective in lobbying state legislatures to become licensed—and this differential effectiveness is quantitatively important to the pattern of licensure across states and occupations—our estimates could erroneously suggest that licensure reduces the foreign-born share of workers.

4.4 Wage premium from licensure

The method above is meant to directly assess whether licensure affects the composition of employment in an occupation, and in particular whether licensure constitutes a disproportionate barrier to employment for foreign-born workers. Here we implement a complementary, indirect method of assessing whether this is the case. In line with the previous literature—from Kleiner and Krueger [2013] to Cassidy and Dacass [2021]—we estimate the wage premiums associated with licensure of native- and foreign-born workers.

When a wage premium is larger, a standard interpretation of that fact is that the barrier to entry to the licensed sector is larger. This is not the only interpretation: importantly, the regression might be misspecified, such that unobservable differences between licensed and unlicensed workers are actually responsible for some of the wage difference. Nonetheless, we estimate wage premiums and present them as part of the overall picture of licensure and foreign-born workers. Wage premiums are especially informative in conjunction with estimates of employment effects: for example, higher wages for licensed foreign-born workers—but lower employment—are consistent with licensure policy causing a leftward shift in the foreign-born labor supply curve in the licensed sector.

Typically, researchers estimate a linear regression to calculate an average wage premium. But licensing, whether of native- or foreign-born workers, can pose different challenges for workers at different points in the wage distribution. We therefore implement conditional quantile regression to assess whether the premium varies by wage level.

Specifically, we implement the following quantile regression at the 25th, 50th, and 75th percentiles using individual-level data, separately for native-born and foreign-born workers:

$$\begin{aligned} Wage_i = lic_i \cdot \beta_{lic} + age_i \cdot \beta_{age} + female_i \cdot \beta_{female} + \Sigma^r race_i^r \cdot \beta_{race}^r \\ + union_i \cdot \beta_{union} + \Sigma^p region_i^p \cdot \beta_{region}^p + \Sigma_n educ_i^n \cdot \beta_{educ}^n + \epsilon_i \end{aligned} \quad (3)$$

where i indexes individual workers, r indexes racial/ethnic groups, p indexes Census regions, n indexes education groups, *union* is an indicator for union membership, and *region* is Census region.

5 Results

5.1 Licensure and foreign-born employment

Table 2 provides summary statistics for the aggregated state-occupation cells, describing their characteristics and how many individual observations underlie the state-occupations that are the unit of observation for our core analysis. The weighted median number of immigrants who arrived in the U.S. as adults ("adult arrivals") in a state-occupation cell is 34,000 and the weighted median number of immigrants who arrived in the U.S. as children ("child arrivals") in a state-occupation cell is 20,700.

Next, Table 3 shows results for our core analysis of how licensure affects total foreign-born employment. We estimate both unweighted and weighted least-squares specifications, with state-occupation sample counts as weights. All standard errors are heteroskedasticity-robust.[8]

Our preferred specifications are those that weight observations based on the number

[8] Because we aggregate the data to the level of policy variation, i.e., the state-occupation, it is not necessary to cluster standard errors.

of state-occupation individual-level observations. Focusing on the weighted least squares (WLS) specification and the most-licensed indicator variable β_{66-100}, we find a substantial negative effect of licensure on foreign-born employment in a given state-occupation: an 18% reduction above and beyond any reduction in native-born employment.[9] Because we control for the natural log of native-born employment, our coefficient of interest can be interpreted as a change in the ratio of foreign-born to native-born employment.

One advantage of our licensure rate bin approach is that it is straightforward to see that the coefficient magnitude rises monotonically as a state-occupation's licensure rate increases. In every specification of Table 3, effects become more negative as the licensure rate rises.

Next we turn to equivalent estimates of licensure effects for immigrants who arrived in the U.S. as adults or as children, and for children of immigrants. See Table 4 for results. Higher licensing shares are again associated with progressively more negative effects, and the coefficient on licensure at a rate above 66% (relative to the omitted category of licensure below 10%) is -24%–27% for adult arrivals and -27%–30% for child arrivals (ranges are between ordinary least squares, or OLS, and WLS estimates).[10] Turning to second-generation immigrants, we find a 21%–25% reduction in employment (relative to "third-plus-generation" employment).

Though generally similar, point estimates in the OLS and WLS specifications are not identical. Following Gary Solon and Wooldridge [2015], we interpret this as potential evidence of misspecification. Specifically, there may be different-sized effects of licensure in different state-occupations, such that changing the weights on state-occupation pairs can have implications for our estimates. This would not be surprising, given the heterogeneous nature of licensing-policy barriers. While we slightly prefer our WLS specification because it places more emphasis on the state-occupation cells about which our survey data are more precise (due to larger sample sizes), we believe that both the OLS and WLS estimates are informative about the range of underlying disemployment effects in different occupations and states.

Comparisons of the adult and child arrival estimates can be informative about the mechanisms by which licensure is reducing foreign-born employment share. Adult arrivals may have obtained training abroad that could (in principle) qualify them for an occupational license, whereas child arrivals cannot have done so. Comparing the results for adult and child

[9]As noted above, tables report log point coefficients. In the text we have generally converted those log point estimates to percent effects.

[10]It may be surprising that the coefficients for immigrants who arrived as children and as adults are sometimes larger in magnitude than the coefficients on the effect of licensure for total immigrant employment. In unreported analysis, we confirmed that this was not entirely due to the difference in samples. We believe it is, rather, due to shifting relationships across specifications between our controls and the different dependent variables.

arrivals can therefore help to distinguish alternative mechanisms for any effect of licensure on employment composition.

The pattern we observe, with child arrivals having equal or (if anything) larger effects of licensure, suggests that inadequate U.S. recognition of credentials and training obtained abroad is not the only driver of our lower estimated foreign-born share in the licensed sector. As discussed above, remaining possibilities include the disproportionate difficulty that immigrants may have in paying the explicit costs (e.g., licensing fee) and opportunity costs (e.g., foregone wages during training) of obtaining a license, as well as potentially limited understanding of licensure pathways. Substantial negative licensure effects for second-generation immigrants also lead us to prefer this set of explanations.

5.2 Licensure and employment by race

Turning now to effects on employment for different racial and ethnic groups, we again find negative effects of licensure on employment. Results are shown in Table 5. The largest employment reductions are for Asian workers (-26%), with Black (-21%) and Latino (-22%) workers experiencing somewhat smaller effects. These magnitudes are roughly comparable to those experienced by immigrants.

A natural question that arises is whether effects on foreign-born workers can be distinguished from effects on Asian, Black, and Latino workers. Are the estimated effects on foreign-born employment entirely a function of race and ethnicity, rather than nativity? To investigate the differential effects, we restrict the sample to native-born workers only and otherwise conduct the same analysis as immediately above. Results are shown in Table 6.

We find that, for Asian and Latino workers, effects of licensure are much smaller when restricting to native-born workers. For Latino workers in particular, estimated effects fall to -6%–9% and are no longer statistically significant. We infer from this that licensure effects on foreign-born workers are driving the overall effects in Table 5, at least in the case of Latino workers, and to a lesser extent for Asian workers. For Black workers, effects of licensure on native-born and foreign-born employment are similar.

5.3 Wage premium from licensure

Figure 5 shows that foreign-born workers have substantially larger wage premiums, i.e., the wage differences (conditional on observable factors) between licensed and unlicensed workers are larger for foreign-born workers than for native-born workers. Results come from a quantile regression that conditions on age, gender, educational attainment groups, race, union membership, and census region.

Interestingly, the gap is smallest at the 25th percentile. Whereas the wage premium for native-born workers is monotonically falling in the wage quantile, the premium for foreign-born workers is highest and roughly the same at the 25th and the 50th percentiles (12.7%). By contrast, the premium for native-born workers at the 50th percentile is 9.8%.

6 Discussion

Occupational licensure shifts employment out of the licensed sector and into the unlicensed sector (Blair and Chung 2018). While it is possible that this shift occurs proportionately for each demographic group, this seems unlikely on its face even before looking at the data, given the multitude of experiences that people have with licensing and the different kinds of resources they bring to the application process.

Leveraging licensing-policy variation across states and occupations, we present evidence that licensure reduces foreign-born employment substantially more than it does native-born employment. We also estimate larger wage premiums for licensed foreign-born workers than for licensed native-born workers. Together, these findings imply that licensure policy is limiting foreign-born labor supply in the licensed sector, diminishing employment, and raising wages (relative to the unlicensed sector).

However, we do not find evidence that licensure effects are larger for those who arrived in the U.S. as adults than for those who arrived as children. We infer that licensure may be reducing foreign-born employment through the monetary and opportunity cost of entry, or other factors that apply to both adult and child arrivals, and not exclusively through inadequate recognition of credentials and training obtained abroad.

We further show that workers of color—especially Latinos—are less likely to be licensed than observably similar White workers, and that (for Asian and Latino workers) this is driven by licensure effects on immigrants. Gaps between native- and foreign-born licensure rates meaningfully contribute to racial disparities in occupational licensure. When we limit the sample to native-born workers and adjust for education, age, and gender, the licensure rate gap for Latino workers is 2.1 percentage points (rather than 4.3 percentage points in the entire sample) and the gap for Asian workers is 5.5 percentage points (rather than 9.0 percentage points).[11]

Further research should improve upon the analysis of this paper in at least two important respects. First, a comprehensive licensure policy dataset, directly taken from statute and

[11]By contrast, the licensure rate gap for native-born Black workers is actually larger (3.7 percentage points) than for the entire sample (2.8 percentage points), but the number of foreign-born Black workers is considerably smaller than their Asian or Latino counterparts.

rulemaking, would constitute a more accurate assessment of licensure policy and (if well-matched to occupation classifications in census data) an improvement over the approach in this paper. Second, a complementary approach of identifying licensure effects from changes in state policy over time would be helpful in addressing concerns about the potential endogeneity of state-occupation variation in contemporary licensure policy.

7 Acknowledgement

The authors are grateful to Nick Carollo, Ed Timmons, and LERA Annual Meeting participants for insightful feedback. However, any errors remain the authors' responsibility.

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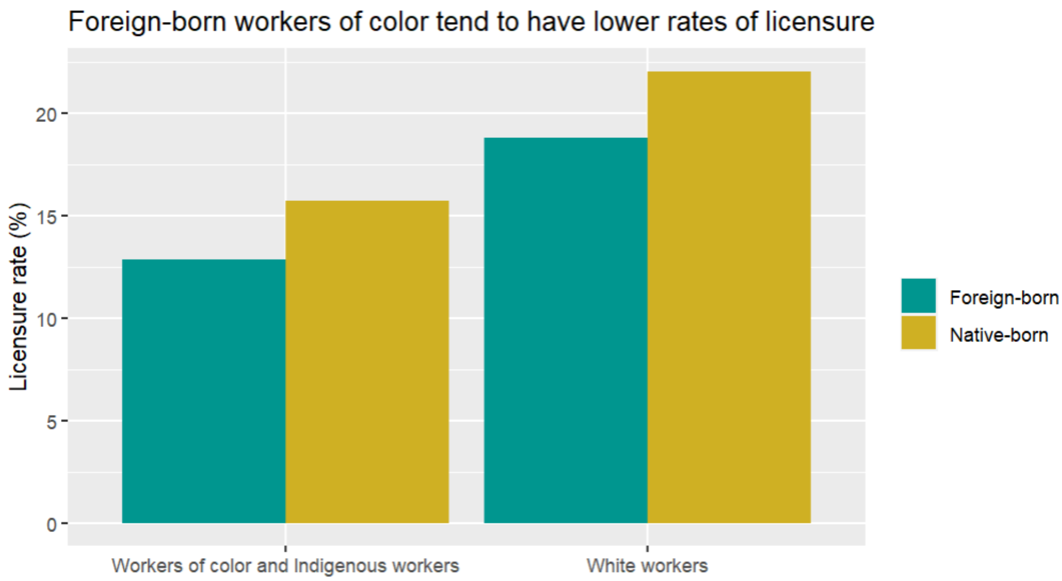
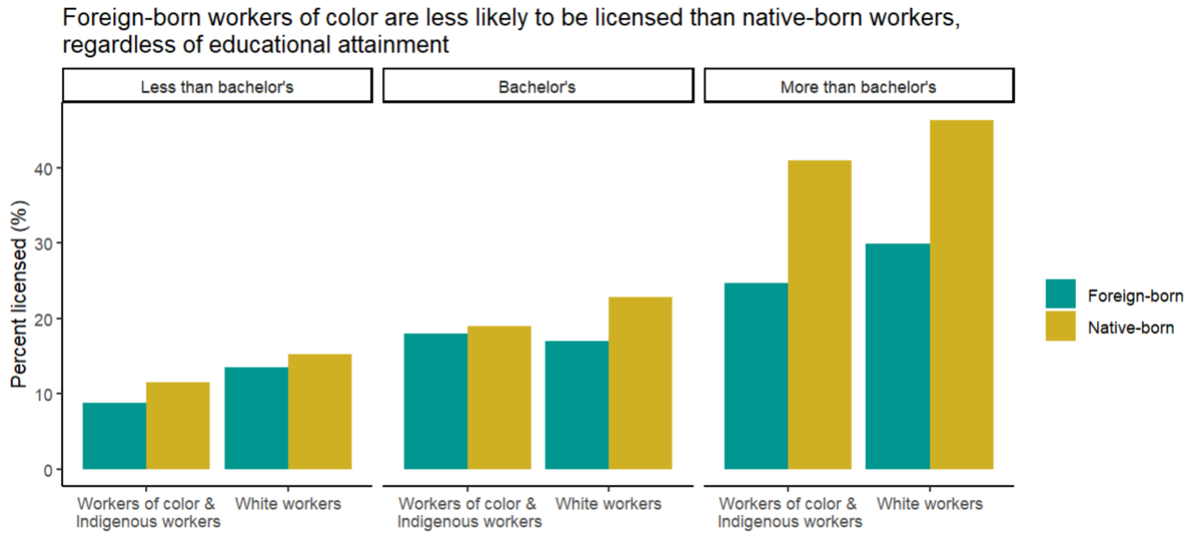
Figure 1: Licensure rates by nativity and race



Source: Authors' calculations using data from the Bureau of Labor Statistics Current Population Survey (CPS) January 2016–December 2022 (accessed via IPUMS-CPS).
Note: Sample includes individuals 16 and older who were employed or employed but not at work last week. Licensed individuals are defined as those who indicated having a professional certification or industry license required for their job that was issued by a government entity. Individuals whose nativity is coded as "Unknown" are excluded from the sample. Categorizations are mutually exclusive.

Figure 2: Licensure rates by nativity and education



Source: Authors' calculations using data from the Bureau of Labor Statistics Current Population Survey (CPS) January 2016–December 2022 (accessed via IPUMS-CPS).
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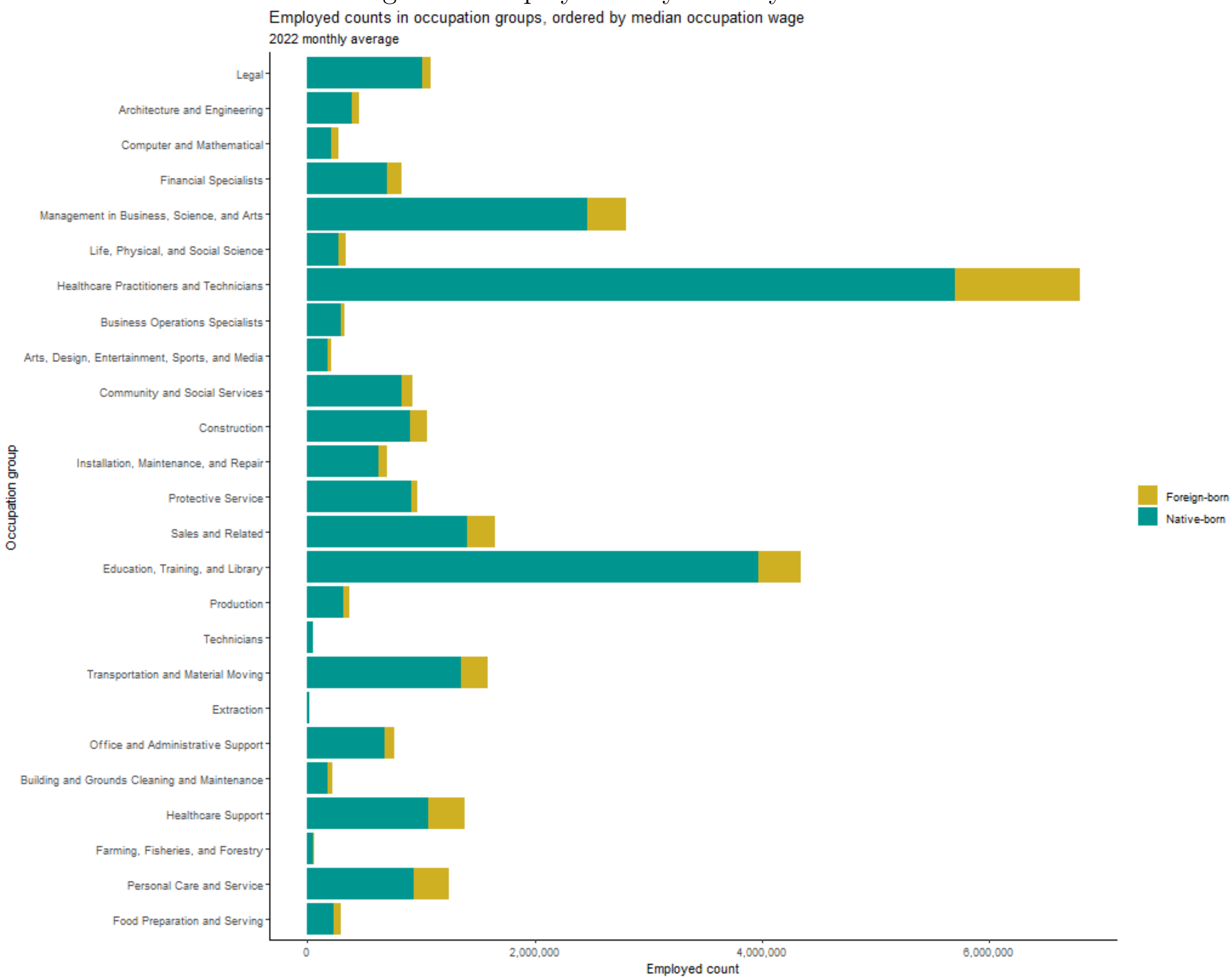
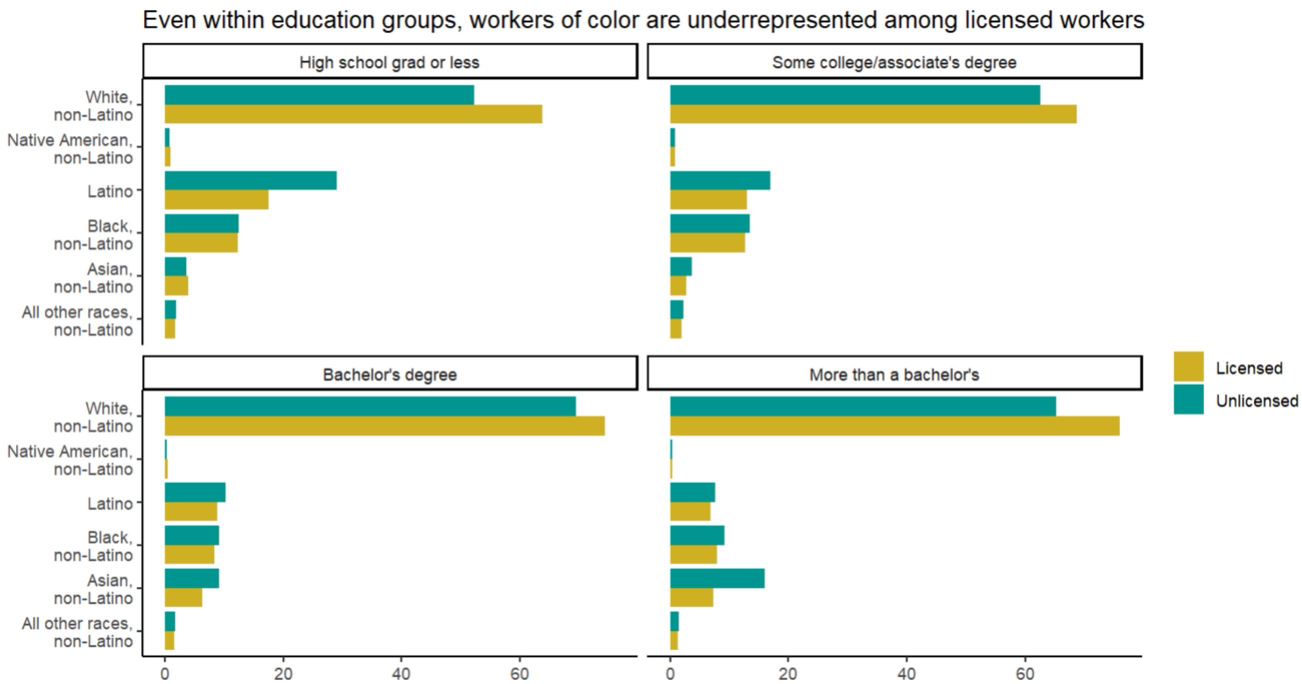
Figure 3: Employment by nativity



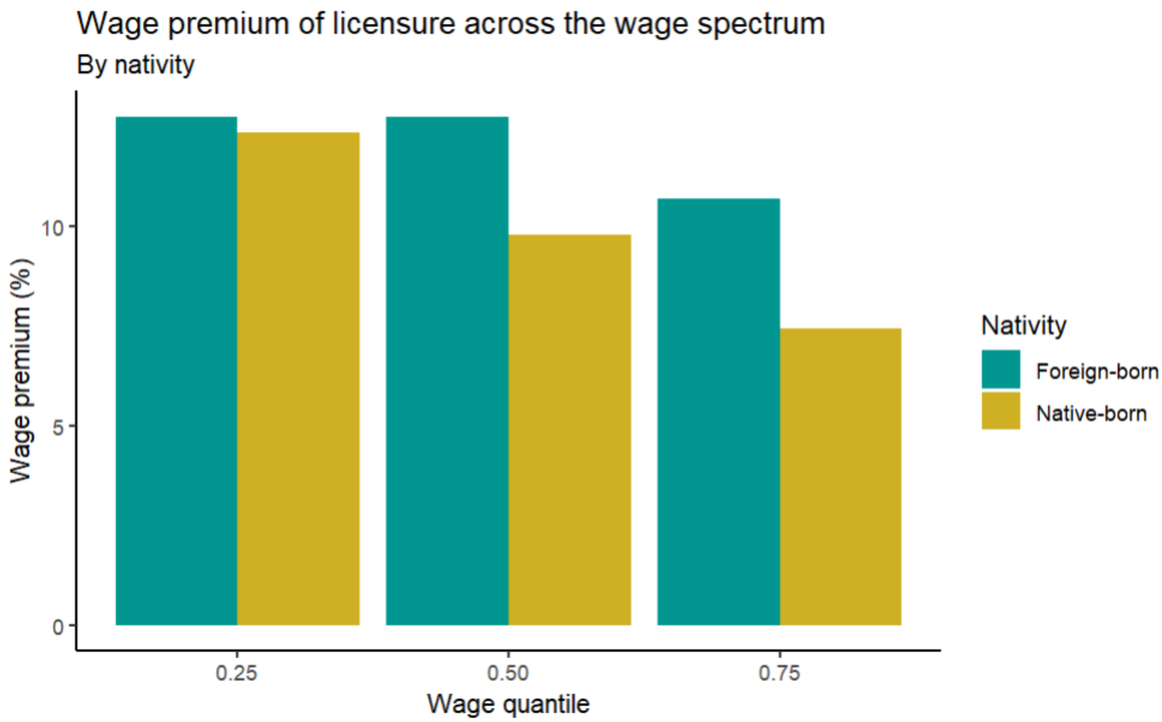
Source: Authors' calculations using data from the 2022 Bureau of Labor Statistics Current Population Survey (CPS).
Note: Sample includes licensed individuals 16 and older who were employed or employed but not at work last week. Licensed individuals are defined as those who indicated having a professional certification or industry license required for their job that was issued by a government entity. Individuals whose nativity is coded as "Unknown" are excluded from the sample. Categories are mutually exclusive. Categorizations are mutually exclusive.

Figure 4: Licensed employment shares by race/ethnicity and education



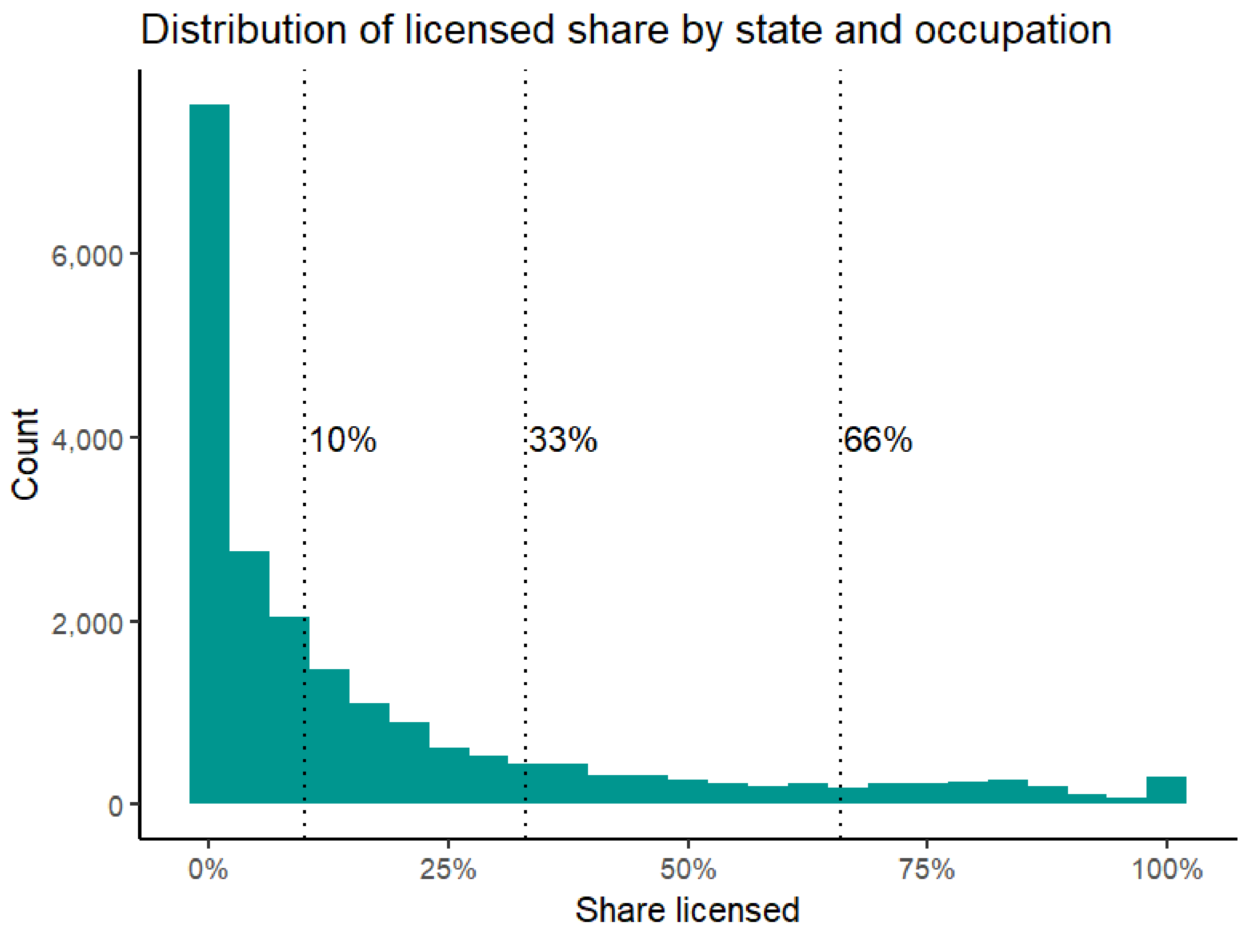
Source: Source: Authors' calculations using data from the Bureau of Labor Statistics Current Population Survey (CPS) January 2016–December 2022 (accessed via IPUMS-CPS).
Note: Sample is restricted to those 16 and older who were employed or employed but not at work last week. Licensed individuals are defined as those who indicated having a professional certification or industry license required for their job that was issued by a government entity. Categorizations are mutually exclusive.

Figure 5: Licensing wage premium by nativity



Source: Authors' calculations using data from the Bureau of Labor Statistics Current Population Survey (CPS) January 2016–December 2022 (accessed via IPUMS-CPS).
Note: Sample includes individuals 16 and older who were employed or employed but not at work last week. Licensed individuals are defined as those who indicated having a professional certification or industry license required for their job that was issued by a government entity. Individuals whose nativity is coded as "Unknown" are excluded from the sample. Categorizations are mutually exclusive.

Figure 6: Distribution of state-occupation licensed shares



Source: Authors' calculations using data from the Bureau of Labor Statistics Current Population Survey (CPS) January 2016–December 2022 (accessed via IPUMS-CPS).
Note: Sample includes individuals 16 and older who were employed or employed but not at work last week. Licensed individuals are defined as those who indicated having a professional certification or industry license required for their job that was issued by a government entity. Categorizations are mutually exclusive.

Table 1: Individual-level sample statistics

	$\overline{Age}$	Female	White	$\geq$ BA	$\overline{Wage}$	N
Licensed	45	54.7%	71.7%	58.0%	$36.2	883314
Unlicensed	42	45.2%	60.6%	34.2%	$27.3	3539607

Table 2: State-occupation sample statistics

	20th	Median	80th	N
Adult arrivals	4065	34014	171617	7285
Child arrivals	3337	24232	101619	7210
Second-generation	7156	32928	127800	7651
Licensed share	2.91%	9.95%	35.20%	8229

Note: Unit of observation is the state-occupation pair. Percentiles of estimated population-level counts (for adult and child arrivals as well as second-generation immigrants) are calculated for the set of state-occupation pairs with non-zero counts in each instance. Child arrivals are those who immigrated at age 17 or younger. Percentiles are sample-weighted.

Table 3: Effects of licensure on foreign-born employment

	First generation	
β_{10-33}	-0.05	-0.08
	(0.03)	(0.03)
β_{33-66}	-0.09	-0.15
	(0.05)	(0.06)
β_{66-100}	-0.22	-0.20
	(0.08)	(0.09)
Native employment	0.62	0.69
	(0.03)	(0.03)
Occ & State FE	Yes	Yes
WLS	No	Yes
F-test	121.00	260.80
R^2	0.859	0.929
N	7832	7832

Note: Dependent variable is the natural log of employment (zeros are dropped). The unit of observation is the state-occupation pair. Counts of individual workers per state-occupation are used for weighted least squares. Standard errors are heteroskedasticity-robust.

Table 4: Effects of licensure on foreign-born employment

	Adult arrivals		Child arrivals		Second generation	
β_{10-33}	−0.12	−0.11	−0.10	−0.13	−0.03	−0.08
	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)
β_{33-66}	−0.22	−0.20	−0.17	−0.27	−0.08	−0.10
	(0.06)	(0.07)	(0.06)	(0.08)	(0.05)	(0.06)
β_{66-100}	−0.32	−0.27	−0.31	−0.35	−0.29	−0.23
	(0.11)	(0.11)	(0.11)	(0.11)	(0.10)	(0.09)
Native employment	0.59	0.67	0.67	0.71	0.75	0.83
	(0.03)	(0.04)	(0.03)	(0.04)	(0.03)	(0.03)
Occ & State FE	Yes	Yes	Yes	Yes	Yes	Yes
WLS	No	Yes	No	Yes	No	Yes
F-test	76.61	159.8	62.48	136.60	75.93	179.50
R^2	0.820	0.906	0.788	0.891	0.819	0.915
N	6409	6409	6409	6409	6409	6409

Note: Dependent variable is the natural log of employment. The unit of observation is the state-occupation pair. State-occupation observations are dropped whenever either adult-arrival, child-arrival, or second-generation employment is zero. Child arrivals are those who immigrated at age 17 or younger. Counts of individual workers per state-occupation are used for weighted least squares. Standard errors are heteroskedasticity-robust. The native employment results in the second-generation specification are for native-born workers that are not second-generation immigrants.

Table 5: Effects of licensure on employment of different racial groups

	Asian		Black		Latino	
β_{10-33}	-0.08	-0.07	-0.06	-0.04	-0.05	-0.03
	(0.04)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
β_{33-66}	-0.17	-0.18	-0.10	-0.10	-0.09	-0.12
	(0.07)	(0.09)	(0.05)	(0.05)	(0.05)	(0.06)
β_{66-100}	-0.34	-0.30	-0.16	-0.24	-0.24	-0.25
	(0.12)	(0.13)	(0.10)	(0.09)	(0.08)	(0.09)
Native employment	0.52	0.67	0.48	0.60	0.52	0.58
	(0.03)	(0.04)	(0.03)	(0.03)	(0.02)	(0.04)
Occupation & State FE	Yes	Yes	Yes	Yes	Yes	Yes
WLS	No	Yes	No	Yes	No	Yes
F-test	55.50	114.60	96.75	191.20	107.50	236.00
R^2	0.771	0.875	0.838	0.912	0.849	0.925
N	6147	6147	7095	7095	7549	7549

Note: Dependent variable is the natural log of employment (zeros are dropped). The unit of observation is the state-occupation pair. Counts of individual workers per state-occupation are used for weighted least squares. Child arrivals are those who immigrated at age 17 or younger. Standard errors are heteroskedasticity-robust.

Table 6: Effects of licensure on employment of different racial groups, native-born only

	Asian		Black		Latino	
β_{10-33}	−0.03	−0.09	−0.07	−0.07	0.03	0.00
	(0.05)	(0.06)	(0.03)	(0.03)	(0.03)	(0.03)
β_{33-66}	−0.06	−0.00	−0.09	−0.14	0.03	−0.03
	(0.09)	(0.12)	(0.06)	(0.06)	(0.06)	(0.06)
β_{66-100}	−0.18	−0.12	−0.15	−0.28	−0.06	−0.09
	(0.16)	(0.21)	(0.11)	(0.11)	(0.09)	(0.09)
Native employment	0.53	0.73	0.47	0.57	0.57	0.65
	(0.04)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
Occupation & State FE	Yes	Yes	Yes	Yes	Yes	Yes
WLS	No	Yes	No	Yes	No	Yes
F-test	21.80	43.24	89.19	180.30	77.38	179.30
R^2	0.665	0.801	0.832	0.910	0.811	0.909
N	3794	3794	6791	6791	7027	7027

Note: Dependent variable is the natural log of employment (zeros are dropped). The unit of observation is the state-occupation pair. Counts of individual workers per state-occupation are used for weighted least squares. Child arrivals are those who immigrated at age 17 or younger. Standard errors are heteroskedasticity-robust.