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Occupational Licensing and Immigrants

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Abstract

This study examines the incidence and impact of occupational licensing on immigrants using two sources of data: the Current Population Survey and the Survey of Income and Program Participation. We find that immigrants are significantly less likely to have a license than similar natives and that this gap is largest for men, workers in the highest education level, and nonnaturalized immigrants. The licensing rate increases with years since migration and shows large variation by immigrants' region of origin. A lack of English proficiency reduces the probability that an immigrant has a license. The wage premium from having a license is much larger for immigrants than natives, though this may in part reflect licensing status proxying for English-language ability.

1. Introduction

Occupational licensing has become an increasingly prevalent aspect of the US labor market over the past several decades. Approximately 29 percent of US workers are required to have an occupational license; as a result, occupational licensing directly impacts more workers than either the minimum wage or unionization (Kleiner and Krueger 2010; Kleiner 2000). Immigrants may be particularly impacted by occupational licensing for several reasons: acquiring a license takes time, and newly arrived immigrants are essentially new entrants in the US labor market; some licenses have a residency and/or citizenship requirement; a license may require educational credentials that may not transfer easily (if at all) between an immigrant's source country and the United States; and most licenses require taking an exam, typically in English, and thus have at least a de facto English-ability requirement. A report on occupational licensing from the Obama White House, recognizing the potentially important connection between immi-

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grants and occupational licensing, says, “Our licensure system can also prevent immigrants who have considerable training and work experience abroad from applying their skills in the U.S., since often they do not meet the relevant licensing requirements” (US Department of the Treasury 2015, p. 8).

A worker with a license typically enjoys a wage premium that averages between 10 and 15 percent (Kleiner and Krueger 2013). Natives and immigrants may differ in the types of occupational licenses they obtain and may therefore differ in the wage premiums they obtain by being licensed. In the presence of such a wage premium, any difficulty acquiring an occupational license by immigrants may contribute to their well-known wage gap with natives. Furthermore, there is evidence that more recent immigrant cohorts to the United States have experienced slower earnings assimilation relative to earlier cohorts (Borjas 2015), though this evidence has been challenged in the literature (Villarreal and Tamborini 2018). The increasing prevalence of occupational licensing may help to explain this apparent slowdown in assimilation if immigrants are being partly excluded from licensed occupations.

An occupational license, being a costly investment, serves as a labor market signal, similar to educational attainment in the classic Spence signaling model (Blair and Chung 2018). Employers in the United States likely have less accurate information about immigrants compared with otherwise similar natives, in part because assessing educational quality may be harder for education acquired abroad. Thus, immigrants may turn to occupational licenses to help alleviate the asymmetric-information problems that are more severe for them than for natives. Thus, we might expect higher licensing rates for immigrants than natives.

Using data from the Current Population Survey (CPS) and the 2008 Survey of Income and Program Participation (SIPP), we examine whether, and to what extent, immigrants are less likely to hold an occupational license compared with otherwise similar natives. We also explore whether an immigrant’s source country is related to licensing probability. We then test whether the wage premium from having a license differs between natives and immigrants. We use the CPS because of its much larger sample size and the inclusion of year-of-immigration and source country variables. The SIPP, while a smaller sample than the CPS, contains an English-proficiency measure and additional details about the type of license the worker holds.

Both the CPS and the SIPP allow us to directly measure the occupational licensing attainment of workers. In contrast, because of a lack of direct licensing-attainment data, much of the literature has relied on licensing coverage, that is, whether an occupation in a given state is licensed.¹ By directly measuring licensing attainment, we can estimate the relationships between licensing status and wages and hours worked, where these relationships are identified from variation in who reports having a license in an occupation in a state and licensing rates by occupation across states. As a robustness check, we replicate our main results

¹ See, for example, Gittleman and Kleiner (2016), which uses the 1979 National Longitudinal Survey of Youth and a database of occupations that require a license by state maintained by the US Department of Labor.

but utilize the imputed licensing status based on occupation code and state of residence used in Gittleman and Kleiner (2016). A more thorough discussion of these results is in Section OA2 in the Online Appendix. In brief, though our main results and the results based on the imputed licensing status overlap in some places, there are substantial differences. In addition, though the imputed licensing status is somewhat predictive of self-reported licensing status, the correlation is only .24. These results suggest that our measures of self-reported licensing status and imputed licensing status are meaningfully different, a finding that implies that additional scrutiny of the self-reported licensing measure and the imputed measure is warranted.

The results from both the CPS and the SIPP show that immigrants are 30–35 percent less likely than otherwise similar natives to have an occupational license. This licensing gap is much larger for men (45–50 percent) than women (15–20 percent). Immigrants' licensing rates increase with years since migration (consistent with labor market assimilation) and English-language proficiency. However, we also find that most of the characteristics of the licenses held by natives and immigrants are the same; for example, they are equally likely to have taken an exam to receive a license.

The licensing rate rises with education more rapidly for natives than for immigrants. Thus, we find that the native-immigrant licensing gap is much larger for workers with more than a bachelor's degree compared with other education levels. In fact, for women with a bachelor's degree or less, there is only a modest gap in licensing rates between natives and immigrants.

We also find large variations in the licensing rate by immigrants' region of origin. Immigrants from three regions in particular—the Caribbean, Southeast Asia, and Africa—have high licensing rates, especially among women. These results are due largely to immigrants from these regions clustering into occupations, notably nursing and cosmetology, that are typically licensed.

We find an overall wage premium from having a license of around 9 percent in the CPS and 7 percent in the SIPP, and in both the CPS and the SIPP women have a much higher licensing wage premium than men. The licensing premium is 7 percentage points higher for immigrants than natives in the CPS. However, controlling for English ability in the SIPP eliminates the positive immigrant and license interaction term, which suggests that licensing status may be serving—at least in part—as a proxy for English-language ability (which is not measured in the CPS). We find that the wage gap between natives and immigrants with an occupational license is much smaller than the wage gap between natives and immigrants without a license.²

Section 2 provides additional background regarding the regulation of occupations and discussion of some of the existing occupational licensing literature. Section 3 describes the data used in the analysis, while Section 4 discusses the results. Section 5 concludes. Analysis of the relationship between citizenship status and

² This result mirrors the results by race from Blair and Chung (2018), where the black-white and male-female wage gaps are lower among licensed than nonlicensed workers.

occupational licensing is reported in the Appendix, together with a discussion of the most common occupations of licensed workers.

2. Background and Existing Literature

Among regulated occupations, there are typically three levels of regulation in the United States. From least to most restrictive, they are registration, certification, and licensure (Kleiner and Krueger 2010). Registration simply requires individuals to file their name and address with a government authority before practicing their occupation. Certification is voluntary and involves the recognition by a government body or some other agency that an individual has achieved a certain level of skill. Only certified individuals can legally use the title “certified” in a professional capacity; however, noncertified individuals are still able to work for pay in an occupation. An occupational license is a legal right to practice an occupation in a given state, and an individual who lacks a license cannot work in the occupation. In this paper, we focus on the most restrictive case (licensure), but we also replicate our results for the less restrictive case of certification.

Most studies of occupational licensing and immigrants focus on high-skilled occupations such as doctors and how their credentials transfer (or not) internationally. Kugler and Sauer (2005) study relicensing of immigrant physicians in Israel and find a substantial increase in earnings for those who acquire a license. McDonald, Warman, and Worswick (2015) analyze international medical graduates in the United States and Canada and find that they are more likely to work as physicians in the United States (where employers’ nomination of immigrants is prevalent) than in Canada (which uses a points-based immigration system). Peterson, Pandya, and Leblang (2014) find evidence that in states where physicians have greater control over the licensing board, there are more stringent requirements for migrant physicians to become licensed, and thus there are fewer new migrant physicians in those states. This is consistent with incumbent workers using occupational licensing to limit competition from high-skilled migrant workers.

Federman, Harrington, and Krynski (2006) explore the occupational licensing of manicurists and Vietnamese immigrants. Though all states require manicurists to be licensed, they vary in the level of English ability required: some states do not require an exam to become a licensed manicurist, while some offer the exam in Vietnamese. In addition, some states allow the transfer of a license from another state. In states with stricter (de facto) English-ability requirements, Vietnamese immigrants are less likely to become manicurists, which suggests that an important link between English ability and licensing may exist in other occupations.

To our knowledge, there is no study that uses large-scale, representative data to explore the relationship between immigration and occupational licensing in the United States. Gomez et al. (2015) investigate the occupational licensing of immigrants to Canada and find that immigrants are less likely to have licenses than natives and that immigrant men may benefit more than natives in terms of wages from being licensed. In addition to studying a different country, Gomez et al.

(2015) impute licensing status on the basis of occupation code alone; they do not observe if a worker is licensed and therefore do not assess licensing attainment. They also are not able to control for English- or French-language proficiency. Tani (2018), using longitudinal data on immigrants in Australia, finds that occupational licensing improves the wages of licensed workers, though because of the occupational licensing hurdle many immigrants do not continue to work in their premigration occupations after moving, which leads to skill wastage.

Examining occupational licensing in Canada, Girard and Smith (2013) find that immigrants are less likely than similar natives to work in regulated occupations, and this lower probability is associated with immigrants' source country. The authors provide evidence to suggest that immigrants educated in Asia, Latin America, and the Caribbean are less likely to enter a regulated occupation than either the native born or immigrants from the United States and Europe. Interestingly, our results indicate that immigrants in the United States from the Caribbean are much more likely to be licensed than other immigrants, largely because of their prevalence in nursing and cosmetology.

Studies of the effects of occupational licensing on wages in the United States, with standard labor market controls, indicate that licensing is associated with 7–18 percent higher wages. Using a relatively small database on occupational licensing from the Princeton Data Improvement Initiative, Kleiner and Krueger (2013) report that licensing is associated with an approximately 18 percent higher hourly wage. Kleiner and Vortnikov (2017) find that occupational licensing raises wages by about 11 percent after controlling for human capital and other observable characteristics, while Gittleman, Klee, and Kleiner (2018) report a 7.5 percent increase in hourly wage when a government-issued license is required for a job. Pizzola and Tabarrok (2017), exploiting a natural experiment to perform a causal analysis, find a licensing premium of 11–12 percent in the Colorado funeral service industry. Our estimates of the occupational licensing wage premium are closer to the lower bound of the estimated range.

3. Data

3.1. *Current Population Survey*

The first data set used in our analysis is the CPS, drawn from the Integrated Public Use Microdata Series (IPUMS) (Flood et al. 2017). The CPS is a monthly representative data set in the United States that includes interviews with each household for 4 consecutive months, followed by an 8-month period in which the household is not interviewed; the household is then interviewed for an additional 4 months before exiting the survey permanently. Month 5 refers to the fifth month an individual household is in the sample, which because of the 8-month break occurs 1 year after the first month. During months 4 and 8 when a household is about to be rotated out of the survey, it is part of the outgoing rotation group (ORG) and is asked more detailed income questions than during the other 3 months.

Beginning in 2015, in months 1 and 5 respondents are asked three questions

related to occupational certification and licensing: “Do you have a currently active professional certification or a state or industry license?” “Were any of your certifications or licenses issued by the federal, state, or local government?” “Is your certification or license required for your job?” Issues with the implementation of the survey led to the information from the third question being unavailable for 2015, but it is available for 2016 onward.

We say that individuals have a credential if they answer yes to the first question and that they are licensed if they answer yes to all three questions. Our main results use the strictest category—licensed—following Kleiner and Soltas (2019), which also uses the CPS licensing data. Thus, for our main results, we drop data from 2015 since the third question is not available in that survey year. In the Online Appendix, we replicate our main results with two alternative definitions: credentials and nonrequired license, that is, where workers answer yes to the first two questions but may answer yes or no to the third question. Since these alternative definitions do not make use of the third question, we are able to include observations from 2015 in the analysis to increase our sample size. Overall, our results are not qualitatively sensitive to these alternative definitions.

We utilize the IPUMS-created variable *cpsidp* to link individuals across survey months (Flood and Pacas 2017). Only very limited income information is available in months 1 and 5 of the monthly basic CPS when data on an individual’s credential are collected. So to explore the relationship between occupational licenses, wages, and hours worked, we link these months with the ORG months (months 4 and 8).

The CPS contains two important immigrant-related variables not included in the 2008 SIPP: year of immigration and birthplace. Knowing the year of immigration allows us to calculate an immigrant’s years since migration and age at migration. An immigrant is defined as someone born outside the United States.³

Our data set includes CPS basic monthly observations from January 2016 to January 2019. We use two estimation samples from the CPS—one for exploring the probability of having a license and another for exploring the relationship between licensing and wages and hours worked. All samples include respondents who are in the labor force, currently working, not in the armed forces, not self-employed, and between ages 18 and 64.⁴ We exclude childhood immigrants, defined as immigrants who arrived in the United States at age 17 or younger, from all samples.⁵ Omitting childhood immigrants, which we are unable to do

³ Individuals from Puerto Rico, Guam, the Northern Mariana Islands, American Samoa, and the US Virgin Islands are considered born in the United States and are thus natives.

⁴ We drop self-employed respondents from all samples in the Current Population Survey and the Survey of Income and Program Participation because the outgoing rotation group sample does not contain earnings data for the self-employed.

⁵ There are a number of compelling reasons to focus on adult immigrants: childhood immigrants are much more likely to be proficient in English than adult immigrants, the effect of years since migration is likely to vary greatly between adult and childhood immigrants, a childhood immigrant will acquire nearly all of his or her work experience in the United States, and childhood immigrants will likely have at least some exposure to the US schooling system. Results that include childhood immigrants do not differ qualitatively from those presented here and are available from the authors on request.

with the SIPP samples since age at migration cannot be calculated, does make the SIPP and CPS immigrant samples different, so care is needed when comparing the CPS and SIPP results.

The probability sample includes only months 1 and 5 since those are the months in which licensing questions are asked. We have 428,455 person-months of observations from 335,557 individuals. Our wage and hours-worked sample uses the ORG. We use the licensing status in month 1 (5) and assign it to the ORG sample in month 4 (8). We use the directly reported hourly wage for workers paid by the hour and calculate the hourly wage of salary workers on the basis of their weekly earnings and usual hours worked per week. We include workers with hourly wages between \$5 and \$100 and drop observations with an imputed wage. This sample includes 218,437 observations from 181,192 individuals.

Summary statistics of the probability sample are shown in Table 1. Immigrants are much less likely than natives to have a credential, license, or certification. Workers with licenses are on average older, more educated, more likely to be married, and disproportionately white. With regard to natives and immigrants with or without a license, licensed natives are much more educated than unlicensed natives, while the difference in education level between licensed and unlicensed immigrants is less dramatic. The number of years since migration is greater for licensed than unlicensed immigrants, consistent with immigrants assimilating over time in terms of licensing status.

3.2. *Survey of Income and Program Participation*

The 2008 panel of the SIPP is a nationally representative longitudinal survey of the United States. Every 4 months respondents answer a core group of questions about the preceding 4 months. The responses provide detailed monthly information about demographics, employment situations, wages, and a variety of other characteristics. Occupational licensing information is derived from the Wave 13 Professional Certifications, Licenses, and Educational Certificates topical module collected between September and December 2012.

A respondent is said to have a credential if he or she responds yes to the following question: “Do you have a professional certification or state or industry license?” Respondents are also asked “Who awarded this certification or license?” and “Is this certification or license required for your current or most recent job?” If the certificate or license was awarded by the federal, state, or local government, and if it is required for the respondent’s job, we refer to the credential as a license; otherwise, we refer to it as a certificate.⁶ Thus, as with the CPS, we use the strictest licensing definition.

The SIPP licensing and certification module includes several additional questions about the characteristics of the individual’s license not included in the CPS, such as “Did you take courses or training to earn this certification or license?”

⁶ This distinction between certificate and license follows definitions 1 and 2 from Gittleman, Klee, and Kleiner (2018).

Table 1
Summary Statistics: Current Population Survey Sample

	All		License		No License	
	Native	Immigrant	Native	Immigrant	Native	Immigrant
Female	.492	.435				
Qualification:						
Credential	.254	.163	1.000	1.000	.070	.042
License	.198	.127	1.000	1.000	.000	.000
Certificate	.056	.036	.000	.000	.070	.042
Education:						
High school dropout	.045	.222	.016	.062	.052	.246
High school graduate	.270	.243	.138	.157	.302	.255
Some college	.315	.152	.281	.198	.323	.146
BA	.245	.205	.280	.290	.237	.193
More than BA	.125	.178	.285	.293	.085	.161
Race or ethnicity:						
White	.714	.158	.780	.204	.697	.151
Black	.133	.118	.104	.215	.140	.104
Asian	.020	.284	.018	.337	.021	.276
Other	.133	.440	.099	.245	.141	.469
Hispanic	.113	.451	.081	.252	.121	.479
Other:						
Age	39.850	43.957	42.170	45.658	39.277	43.710
Married	.505	.695	.623	.725	.476	.691
Children	.783	1.189	.967	1.181	.738	1.191
Years since migration		15.051		16.548		14.833
N	382,161	46,294	78,510	5,858	303,651	40,436

Note. Values are means. BA = bachelor of arts.

These additional licensing questions allow us to go into more depth regarding the differences in occupational licensing between natives and immigrants. Importantly, unlike the CPS, the SIPP contains a self-reported measure of English-speaking proficiency.

Our sample includes workers ages 18–64. We use a single observation per worker from wave 13, which is either month 4 (in 93 percent of cases) or the nearest month for those not employed in month 4. As in the CPS, our sample includes individuals who are currently employed, not in the armed forces, not self-employed, and have hourly wages between \$5 and \$100; we drop observations with an imputed wage. An immigrant is defined as anyone born outside the United States. At 21,269 observations, the SIPP sample is much smaller than the CPS probability and wage samples. Unlike the CPS, we have only one SIPP sample that we use to explore licensing probability, wage premium, and hours worked.

Descriptive statistics of our SIPP sample are shown in Table 2. The credential rate is higher in the SIPP than the CPS, though a smaller fraction of the credentials are licenses in the SIPP than the CPS, which leads to an overall lower licensing rate of natives in the SIPP (17.3 percent) than the CPS (19.8 percent). Other-

Table 2
Summary Statistics: Survey of Income and Program Participation Sample

	All		License		No License	
	Native	Immigrant	Native	Immigrant	Native	Immigrant
Female	.497	.440				
Qualification:						
Credential	.293	.184	1.000	1.000	.145	.090
License	.173	.103	1.000	1.000	.000	.000
Certificate	.119	.081	.000	.000	.145	.090
Education:						
High school dropout	.036	.219	.010	.493	.041	.235
High school graduate	.235	.238	.113	.097	.261	.255
Some college	.386	.241	.346	.322	.394	.232
BA	.229	.189	.284	.282	.217	.178
More than BA	.115	.113	.247	.221	.088	.100
Race or ethnicity:						
White	.837	.644	.861	.501	.832	.661
Black	.112	.099	.092	.177	.117	.090
Asian	.014	.217	.016	.293	.014	.208
Other	.036	.040	.031	.030	.037	.042
Hispanic	.102	.468	.077	.252	.108	.493
Other:						
Age	40.706	41.392	42.719	42.991	40.283	41.209
Married	.563	.674	.658	.691	.543	.672
Children	.738	1.080	.795	.891	.726	1.102
Union	.135	.105	.263	.225	.108	.091
Paid by hour	.533	.593	.392	.394	.563	.616
Government worker	.184	.099	.354	.239	.149	.083
Speaks English:						
Very well		.601		.827		.575
Not well		.399		.173		.425
N	17,582	3,158	3,080	326	14,502	2,832

Note. Values are means. BA = bachelor of arts.

wise, very similar patterns are observed in the SIPP and CPS: licensed workers are older, more educated, and less likely to be immigrants and have higher wages. Licensed immigrants have much greater English proficiency than unlicensed immigrants, which is consistent with a lack of language proficiency imposing a barrier to obtaining a license, though this, of course, does not control for other characteristics such as age or level of education.

In Table 3, we take advantage of the richer certificate information provided in the SIPP than the CPS to see if immigrants and natives acquire different types of licenses. For workers with licenses, we show the fractions of natives and immigrants that are required to take a course or training, have to demonstrate skills or take an exam, have to fulfill continuing education requirements, and acquire their license for work or personal reasons. We detect no major differences between natives and immigrants in any of these characteristics. While these are

Table 3
Licensing Requirements: Survey of Income and Program Participation Sample

	All		Men		Women	
	Native	Immigrant	Native	Immigrant	Native	Immigrant
Course or training	.942	.937	.921	.895	.957	.961
Skills or exam	.923	.916	.922	.910	.923	.920
Continuing education	.768	.757	.721	.702	.802	.790
Work-related license	.990	.992	.990	.992	.990	.991
Personal license	.021	.010	.021	.015	.020	.007
<i>N</i>	3,080	326	1,211	119	1,869	207

Note. Values are means.

only a small number of descriptives, the results suggest that the types of licenses acquired by natives and immigrants may not differ much in their characteristics.

4. Results

We begin by discussing the licensing probability results from the CPS and SIPP in Section 4.1 and then discuss the relationships between licensing and hourly wage and hours worked in Section 4.2. The estimation models are as similar as possible between the two data sets. The two notable differences are that controls for years since migration are included in the CPS, and an English-proficiency control is included in the SIPP.

4.1. Probability of Having an Occupational License

The summary statistics suggest a close connection between educational attainment and licensing status that may differ for immigrants and natives. To demonstrate the relationship between licensing rates and educational attainment, Figure 1 shows the fraction of each education level that reports having a license in the CPS probability sample. The increasing trend of licensing with education for all groups is clear; however, licensing increases with educational attainment more rapidly for natives than immigrants. The native-immigrant licensing gap is substantially larger for the most educated group than any other, particularly for women. In fact, native and immigrant women exhibit little difference in their licensing rates across the other four education levels. Thus, it seems plausible that the labor market impacts of occupational licensing may be more relevant for highly educated immigrants.

We estimate a series of linear probability regressions in which the dependent variable equals one if the individual reports having an occupational license and zero otherwise:

$$\begin{aligned} \text{License}_{it} = & \beta_X \mathbf{X}_{it} + \beta_{\text{Immigrant}} \text{Immigrant}_i + \beta_{\text{YSM}} \text{YSM}_{it} \\ & + \beta_{\text{YSM}^2} \text{YSM}_{it}^2 + \delta_s + \varepsilon_{it}, \end{aligned} \quad (1)$$

where License_{it} equals one if worker i has a license in month t and zero otherwise, $\mathbf{X}_{it}$ is a vector of individual and other characteristics, δ_s is the fixed effect for state

Figure 1. Occupational licensing by education level

of residence s , and Immigrant_i equals one if the worker is an immigrant and zero otherwise. Values for years since migration (YSM_{it} and YSM_{it}^2) equal 0 for natives. We control for age (introduced as a third-order polynomial), educational attainment (five categories), survey year, month of sample (month 1 or 5), month of the year, whether the individual appears once or twice in the sample, and number of children and include racial and ethnic dummy variables (black, Asian, other, and Hispanic) and a married dummy variable.⁷ This estimation is performed on the full sample and separately for men and women.

The results from these estimations are shown in Table 4. As our focus is on immigration and licensing, our primary coefficients of interest are those on the

⁷ Since attrition between months 1 and 5 may not be random, and thus including month 5 may bias our results, we repeated the estimates using data from only month 1. The results are nearly identical.

Table 4
Occupational Licenses: Current Population Survey Sample

	All	Men	Women
Female	.054*** (.0014)		
Immigrant	-.097*** (.0042)	-.088*** (.0050)	-.095*** (.0073)
YSM/10	.036*** (.0054)	.014* (.0064)	.056*** (.0094)
YSM ² /100	-.004** (.0016)	-.000 (.0018)	-.009** (.0027)
High school graduate	.018*** (.0021)	.022*** (.0027)	.018*** (.0034)
Some college	.089*** (.0023)	.075*** (.0030)	.117*** (.0036)
BA	.126*** (.0026)	.079*** (.0033)	.185*** (.0040)
More than BA	.310*** (.0033)	.232*** (.0044)	.396*** (.0048)
License (mean)	.189	.156	.224
N	428,455	217,081	211,374
R ²	.088	.059	.111

Note. Results are from linear probability regressions, with individual-level clustered standard errors in parentheses. The dependent variable equals one if the worker has an occupational license and zero otherwise. All estimations include controls for age (as a third-order polynomial), state of residence, month in sample, month of the year, survey year, number of times the respondent is observed, marital status, and number of children and racial and ethnic dummy variables (black, Asian, other, and Hispanic). The omitted group is high school dropouts. YSM = years since migration; BA = bachelor of arts.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

immigrant dummy and years since migration, though we also show the coefficients for gender and educational attainment. Table 4 also includes the mean of the dependent variable, that is, the fraction of the sample that has an occupational license.

Note the gender gap in licensing rates: controlling for other characteristics, we find that women are 5.4 percentage points more likely than men to be licensed. Consistent with Figure 1, educational attainment has a strong, positive relationship with licensing status, with workers with a bachelor of arts (BA) being 12.6 percentage points more likely than high school dropouts to be licensed, while workers with more than a BA are 31.0 percentage points more likely to be licensed.

The coefficient on the immigrant dummy variable is negative, statistically significant, and economically large in all estimations. Consistent with labor market assimilation, the probability that an immigrant is licensed increases with years since migration: 10 years after migration, the probability increases by 3.6 per-

centage points. In the full sample, we find that 10 years after migration, immigrants are 6.5 ($= 9.7 - 3.6 + .4$) percentage points less likely to have an occupational license; given the overall rate of 18.9 percent, this corresponds to a 34.4 percent lower probability.

The licensing gap between natives and immigrants 10 years after migration is greater for men (7.4 percentage points) than women (4.8 percentage points). Furthermore, the overall licensing rate for women is higher than that for men (22.4 percent versus 15.6 percent), so 10 years after migration, immigrant men and women are 47.4 percent and 21.4 percent, respectively, less likely to be licensed than otherwise similar natives. In addition, the effect of years since migration is weaker for men than women (that is, male immigrants seem to assimilate more slowly in terms of licensing status than female immigrants). Ten years after migration, the licensing probability increases by only 1.4 percentage points for men but by 4.7 percentage points for women, and thus the gender licensing gap for immigrants grows with years since migration.

A prominent phenomenon among immigrant workers in the United States is occupational clustering in which immigrants of the same nationality tend to work in similar occupations, especially when clustered in the same city (Patel and Vella 2013). In the presence of occupational clustering, since certain occupations require a license, we might expect variation in occupational licensing rates by birthplace, such as is the case for Vietnamese manicurists (Federman, Harrington, and Krynski 2006). We explore this idea by dividing immigrants by region of birth.⁸

Figure 2 shows the license rates of each region relative to natives: the raw difference in licensing rate and the adjusted difference that accounts for workers' characteristics. The adjusted differences are computed by adding region of birth to equation (1) and then using the coefficients on the region dummy variables adjusted such that years since migration equals 15, which is approximately the mean of the sample.

Licensing rates are lowest for immigrants from Mexico and Central America for both men and women. Interestingly, and consistent with occupational clustering, we find high licensing rates for immigrants from the Caribbean, Southeast Asia, and Africa. Immigrants from these regions, especially women, tend to work in occupations such as nursing and cosmetology, which usually require occupational licensure. For example, 8.3 percent of immigrant women from Southeast Asia work as miscellaneous personal appearance workers, a category that includes manicurists, while only .1 percent of native women work in that occupation. Among immigrant women from Africa, 13.1 percent are nurses, compared with only 5.5 percent among native women.

Adjusting for workers' characteristics expands the licensing gaps for many regions; the main driver for this difference appears to be educational attainment.

⁸ The regions are the English-speaking developed countries Canada, Australia, New Zealand, the United Kingdom, and Ireland; Mexico; Central America; the Caribbean; South America; northern and western Europe; the rest of Europe and Russia; eastern Asia; Southeast Asia; central, southern, and southwestern Asia; the Middle East; and Africa.

Figure 2. Occupational licensing rates and immigrants' origin

For example, female immigrants from English-speaking developed countries have a mean licensing rate that is higher than that for natives, but their adjusted licensing rate is lower largely because of their high levels of education. Immigrants from Mexico, on the other hand, have a smaller adjusted licensing rate gap than their raw gap (though both are negative), mostly because of their low average education levels.

We now turn to a discussion of the licensing probability results using SIPP data. We first estimate a linear probability regression model, as described in equation (1), where the dependent variable equals one if the worker has a license and zero otherwise. Our controls are as similar as possible to those in the CPS probability estimations. In particular, all estimations control for an immigrant dummy variable, state of residence, age (introduced as a third-order polynomial), educational attainment (five categories), racial and ethnic dummy variables (black,

Asian, other, and Hispanic), a married dummy variable, and number of children. Notably, unlike with the CPS data, we cannot control for years since migration. We thus add a dummy variable for ability to speak English very well, which we set equal to zero for natives.

Results are shown in Table 5. Without language-ability controls, immigrants in the full sample are 5.0 percentage points less likely to have a license; given the overall licensing rate of 16.2 percent, this represents a 30.9 percent lower probability. Immigrants in the full sample who are proficient in English are 24.7 percent less likely than natives to have a license, while immigrants who do not speak English well are 42.6 percent less likely to have a license, and this difference is statistically significant at the 5 percent level. Thus, as expected, immigrants who are proficient in English are more likely than other immigrants to report having an occupational license. However, it is important to note that we cannot make a causal claim, since it is possible that some other characteristic for which we do not control is related to English proficiency and the probability of being licensed.

Male immigrants are 47.3 percent less likely and female immigrants are only 17.9 percent less likely to have a license relative to natives. Female immigrants who speak English very well are only 11.7 percent less likely to have a license than native women, while male immigrants who speak English very well are 41.1 percent less likely than male natives to have a license. The values for female and male immigrants who do not speak English well are 31.1 percent and 58.1 percent, respectively. The differences between the licensing rates of immigrants proficient versus not proficient in English are statistically significant at the 10 percent significance level for women, while the p -value for men is slightly above that level, at .115.

Comparing the SIPP and CPS licensing probability results, we find that at the mean number of years since migration, about 15 (in the CPS), the results are similar, with a 30.9 percent lower probability in the SIPP and 27.5 percent lower probability in the CPS. Notably, the SIPP results provide suggestive evidence that language proficiency, or the lack thereof, is an important contributor to the native-immigrant licensing difference, especially for female immigrants.

4.2. Licensing, Wages, and Hours Worked

We turn now to the relationship between occupational licensing and wages and hours worked. We estimate the following regression model:

$$Y_{it} = \beta_X X_{it} + \beta_{\text{License}} \text{License}_{it} + \beta_{\text{Immigrant}} \text{Immigrant}_i + \beta_{\text{License} \times \text{Immigrant}} \text{License}_{it} \times \text{Immigrant}_i + \delta_s + \varepsilon_{it}, \quad (2)$$

where Y_{it} is the labor market outcome of interest, that is, either log wage or hours worked for individual i in month t . The dummy variable License_{it} equals one if the worker has a license in month t , Immigrant_i is a dummy variable that equals one if the worker is an immigrant, X_{it} is a vector of control variables, and δ_s is the

Table 5
Occupational Licenses: Survey of Income and Program Participation Sample

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Female	.055*** (.0052)	.055*** (.0052)				
Immigrant	-.050*** (.0084)	-.069*** (.0115)	-.061*** (.0099)	-.075*** (.0120)	-.035** (.0128)	-.061** (.0200)
Speaks English very well		.029* (.0128)		.022 (.0138)		.038+ (.0219)
High school graduate	.015+ (.0076)	.010 (.0077)	.019+ (.0096)	.015 (.0095)	.015 (.0136)	.008 (.0150)
Some college	.080*** (.0087)	.074*** (.0086)	.067*** (.0100)	.063*** (.0102)	.109*** (.0148)	.101*** (.0157)
BA	.128*** (.0092)	.122*** (.0097)	.064*** (.0113)	.059*** (.0114)	.207*** (.0165)	.198*** (.0180)
More than BA	.261*** (.0142)	.254*** (.0143)	.176*** (.0175)	.172*** (.0176)	.356*** (.0216)	.346*** (.0223)
License (mean)	.162	.162	.129	.129	.196	.196
N	20,740	20,740	10,314	10,314	10,426	10,426
R ²	.071	.071	.047	.048	.093	.094

Note. Results are from linear probability regressions, with standard errors (calculated using balanced repeated replication, with Fay's adjustment factor of .5 and 120 survey-provided replicate weights) in parentheses. All estimations include controls for age (as a third-order polynomial), state of residence, union status, marital status, number of children, and dummy variables for government worker, paid by the hour, and race or ethnicity (black, Asian, other, and Hispanic). The omitted groups are high school dropouts and immigrants who do not speak English well. BA = bachelor of arts.

+ $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

fixed effect for state of residence s . The term $\text{License}_{it} \times \text{Immigrant}_i$ tests if the effect of having a license differs between immigrants and natives.

The baseline estimation includes the same set of controls used in the license probability regressions. In addition, we include a number of job-related controls, including dummy variables for the worker's union status and whether he or she is full-time versus part-time, paid by the hour, and a government worker. We again perform the estimates on the full sample and separately for men and women. For brevity, only the immigrant dummy variable and the occupational licensing variables are shown. Table 6 presents the CPS results for log wage and hours worked, while Table 7 presents the SIPP results.

We begin by discussing the log wage results from the CPS (Table 6). The licensing coefficient is positive and statistically significant in all specifications. In the full sample without interacting licensing status with the immigrant dummy, we find a wage premium of 8.5 log points, which is higher than the results from Gittleman, Klee, and Kleiner (2018) and our results from the SIPP, though it is somewhat low relative to other results in the literature (Kleiner and Krueger 2013; Pizzola and Tabarrok 2017). The licensing premium is much higher for women (12.3 log points) than men (4.5 log points), a finding that has received little attention in the literature.⁹

The immigrant and license interaction terms are positive, statistically significant at the .1 percent level, and economically meaningful in all three samples: the wage premium for immigrants is 15.1 log points ($7.8 + 7.3$) for the full sample, while for natives it is only 7.8 log points. We find that immigrant men earn a much larger licensing premium than native men (12.1 log points versus 3.7 log points), while immigrant women also earn a higher premium than native women (18.0 log points versus 11.7 log points).

It is unclear why immigrants appear to benefit more than natives from holding an occupational license. One possible explanation is that, because an occupational license is a costly investment, it also serves as a signal, and the value of that signal is greater for immigrants than natives. Employers likely have less information about immigrants than natives, for example, about the quality of their educational credentials, and thus a license may be a stronger signal for immigrants than for natives.¹⁰ Alternatively, an occupational license may signal a lower probability of return migration and thus a stronger attachment to the US labor force,

⁹ Using the 2008 SIPP, Blair and Chung (2018) find evidence of a higher return to occupational licensing for women than men, which they argue is due to higher returns to human capital acquired along with a license.

¹⁰ Using Canadian data, Gomez et al. (2015) also find evidence that immigrants enjoy a larger earnings premium from holding licenses compared with natives. This result, coupled with a lower probability of an immigrant being in a regulated occupation, could support proposals to facilitate recognition of immigrants' foreign credentials. Using data on immigrants to Canada, Li and Sweetman (2014) discuss the importance of quality source-country education in determining immigrants' wages. They provide evidence that the average quality of educational outcomes (measured using international test scores) in an immigrant's source county is positively associated with wages in the host country.

Table 6
Wages and Hours Worked: Current Population Survey Sample

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Log hourly wage:						
Immigrant	-.178*** (.0097)	-.185*** (.0097)	-.175*** (.0134)	-.181*** (.0134)	-.180*** (.0137)	-.188*** (.0138)
License	.085*** (.0032)	.078*** (.0033)	.045*** (.0048)	.037*** (.0049)	.123*** (.0043)	.117*** (.0044)
License × Immigrant		.073*** (.0125)		.084*** (.0189)		.063*** (.0166)
N	218,437	218,437	110,627	110,627	107,810	107,810
R ²	.480	.480	.484	.484	.462	.462
Usual hours worked per week:						
Immigrant	-.588*** (.1587)	-.619*** (.1589)	-.737*** (.2177)	-.687** (.2172)	-.251 (.2284)	-.359 (.2297)
License	1.071*** (.0564)	1.042*** (.0584)	1.648*** (.0906)	1.706*** (.0947)	.632*** (.0701)	.551*** (.0721)
License × Immigrant		.321 (.2051)		-.655** (.3175)		.882*** (.2669)
N	218,437	218,437	110,627	110,627	107,810	107,810
R ²	.383	.383	.289	.289	.439	.439

Note. Result are from ordinary least squares estimations with individual-level clustered standard errors in parentheses. All regressions include controls for age (as a third-order polynomial), state of residence, educational attainment (five categories), years since migration, years since migration squared, month in sample, month of the year, survey year, number of times the respondent is observed, marital status, number of children and dummy variables for full-time, government worker, union, paid by the hour, and race or ethnicity (black, Asian, other, and Hispanic).

** $p < .01$.

*** $p < .001$.

Table 7
Wages and Hours Worked: Survey of Income and Program Participation Sample

	All			Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log hourly wage:									
Immigrant	-.103*** (.0143)	-.107*** (.0144)	-.245*** (.0154)	-.125*** (.0197)	-.124*** (.0199)	-.250*** (.0205)	-.078*** (.0178)	-.083*** (.0181)	-.239*** (.0234)
License	.066*** (.0106)	.063*** (.0113)	.066*** (.0114)	.012 (.0140)	.012 (.0145)	.016 (.0145)	.112*** (.0146)	.108*** (.0156)	.112*** (.0156)
License × Immigrant		.027 (.0354)	-.016 (.0360)		-.001 (.0545)	-.042 (.0552)		.030 (.0410)	-.013 (.0396)
Speaks English very well			.215*** (.0180)			.202*** (.0287)			.233*** (.0228)
N	20,740	20,740	20,740	10,314	10,314	10,314	10,426	10,426	10,426
R ²	.441	.441	.445	.451	.451	.456	.422	.422	.427
Usual hours worked per week:									
Immigrant	-.494* (.2238)	-.477* (.2336)	-.527+ (.3048)	-.925** (.2845)	-.992** (.3017)	-1.034** (.3519)	-.038 (.3273)	.125 (.3310)	-.036 (.4662)
License	.824*** (.1848)	.840*** (.1909)	.842*** (.1920)	1.344*** (.2908)	1.263*** (.3081)	1.264*** (.3091)	.559* (.2375)	.683** (.2479)	.687** (.2494)
License × Immigrant		-.145 (.6181)	-.161 (.6293)		.784 (.9106)	.770 (.9201)		-1.041 (.7693)	-1.085 (.7828)
Speaks English very well			.079 (.3701)			.069 (.3554)		.241	.241 (.6375)
N	20,740	20,740	20,740	10,314	10,314	10,314	10,426	10,426	10,426
R ²	.193	.193	.193	.176	.176	.176	.175	.176	.176

Note. Result are from ordinary least squares estimations, with standard errors (calculated using balanced repeated replication, with Fay's adjustment factor of .5 and 120 survey-provided replicate weights) in parentheses. All regressions include controls for age (as a third-order polynomial), state of residence, educational attainment (five categories), union status, marital status, number of children and dummy variables for government worker, paid by the hour, and race or ethnicity (black, Asian, other, and Hispanic). The omitted groups are high school dropouts and immigrants who do not speak English well.

+ $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

which may make employers more likely to train immigrant workers with licenses relative to immigrants without licenses.

Because of the positive interaction term for immigrant status and licensing status, the wage gap between licensed natives and immigrants is lower than that between unlicensed natives and immigrants: for unlicensed workers, the native-immigrant wage gap is 17.8 log points; however, when native and immigrant licensed workers are compared, the wage gap narrows significantly to only 11.2 log points. Thus, immigrants able to acquire an occupational license do relatively well in the labor market compared with native workers. This result mirrors that of Blair and Chung (2018), who find that occupational licensing can help to narrow the black-white and male-female wage gaps.

To estimate the effect of occupational licensing on wages in the SIPP (Table 7), we repeat a log wage regression similar to equation (2) with the same set of controls as in the CPS wage estimations except we cannot control for years since migration. In the full sample, we find a return to holding a license of 6.6 log points. Interacting the license and immigrant controls, we find a positive coefficient, which is consistent with the CPS results: immigrants with a license earn wages that are 9.0 log points higher, while natives with a license earn wages that are 6.3 log points higher. The interaction term, however, is not statistically significant at conventional levels, so we cannot rule out that natives and immigrants receive the same wage premium from being licensed. Nonetheless, adding controls for immigrants' language proficiency for the full sample reduces the interaction term by .043 log point. This result suggests that holding a license may partly serve to proxy for English proficiency, which is consistent with the large, positive returns to English-language proficiency for immigrants found in the literature (Bleakley and Chin 2004).

As in the CPS results, we find a much larger licensing wage premium for women (11.2 log points) than for men (1.2 log points). In fact, we find no evidence of a licensing premium for men. Contrary to the wage results from the CPS, the license and immigrant interaction terms are not statistically significant at conventional levels, though both are positive. Adding English proficiency again results in the coefficient of the interaction term becoming negative for both men and women, consistent with holding a license proxying for English ability.

Our wage results, using both data sets, identify a positive and statistically significant license premium. However, these estimates may not fully capture the relationship between licensing and workers' earnings if the number of hours worked is related to licensing status.

In the CPS results in Table 6, we find that workers with a license work 1.1 more hours per week than those without a license. Interacting licensing status with immigrant status, we find a positive though statistically insignificant effect. We find a much stronger (more positive) relationship between having a license and hours worked for men than women, though the coefficient is positive and statistically significant for both. This larger effect of licensing on hours worked for men than

women should, to some extent at least, offset the much larger wage premium for women than men.

The immigrant and license interaction term is large, negative, and statistically significant for men, which again would serve to reduce the overall earnings effect of having a license on immigrants versus natives. For women, the interaction term is positive and statistically significant at conventional levels; thus, the larger wage premium for licensed immigrant women understates the difference in the effect of holding a license on earnings between immigrant and native women.

Turning to the SIPP results (Table 7), we find that, as with the CPS result, there is a positive relationship between hours worked and having a license, and this positive relationship is also larger for men than women. Of note, the magnitude of the coefficient on licensing status is lower in the SIPP results than in the CPS results. The interaction term is slightly negative for the full sample but is positive for men and negative for women, though the coefficient is statistically insignificant in all estimations.

Overall, while there are some differences between the SIPP and the CPS results in terms of hours worked, we find evidence that licensed individuals work more hours per week than unlicensed workers. This suggests that the earnings premium enjoyed by licensed workers is greater than their wage premium.

5. Further Discussion and Conclusion

Using data from the CPS and the SIPP, we find that immigrants (especially men) are much less likely than natives to have occupational licenses. The wage premium from being licensed appears to be larger for immigrants than natives; however, part of this difference may be due to the licensing status of immigrants proxying for their English-language ability.

Given the substantial increases in the prevalence of occupational licensing, the immigrant fraction of the US workforce, and the apparent difficulty that immigrants have in acquiring occupational licenses, our results provide what we believe to be important insights into a largely underexplored aspect of the occupational licensing debate. Furthermore, our results suggest that the impact of licenses is heterogeneous in a number of important dimensions, including education level and immigrants' birthplace.¹¹ Understanding how occupational licensing affects different groups of workers, such as men versus women and natives versus immigrants, is important in providing a clear picture of the labor market impacts of occupational licensing.

One important characteristic of immigrant workers in the United States is that they are a more mobile group than natives, especially newer immigrants, and thus help labor market efficiency by making locational choices on the basis of

¹¹ This result is similar to discussions surrounding union wage differentials, which vary by gender and education (Blackburn 2008; Gomez and Lamb 2019) and employment contract (Blanchflower and Bryson 2010). Union membership seems to benefit disadvantaged workers more, which narrows the long-standing earning gaps between workers in traditionally advantaged and disadvantaged groups in society.

wage differences across locations. In the words of Borjas (2001, p. 117), immigrants are “grease on the wheels of the labor market.” Occupational licensing requirements, however, have the opposite effect: Johnson and Kleiner (2017) find evidence that between-state migration rates are significantly lower for workers in occupations with state-specific licensing exam requirements than for workers in other occupations. Since occupational licensing appears to impact immigrants more strongly, licensing requirements may serve to diminish this important role that immigrants play in the economy, thus leading to lower labor market efficiency and greater interstate wage inequality.

Finally, our results have important implications for potential changes to immigration policy, particularly movements toward more skill-based immigration.¹² Skill-based immigration would give preference to immigrants with high levels of education. Our results indicate that it is precisely this group that exhibits the largest licensing attainment gap with natives. Increasing the flow of highly educated immigrants may lead to occupational mismatch if they face difficulty in acquiring licenses needed to work in their premigration occupations. Policy discussions surrounding the movement to a more skill-based immigration system would benefit from carefully considering the issue of how immigrants—especially those who are highly educated—may acquire occupational licenses after migration.

Appendix

Occupational Licensing, Citizenship, and Common Occupations

A1. Occupational Licensing and Citizenship

As documented in Calvo-Friedman (2014), there are numerous examples of occupations with licensure restricted to citizens or to citizens and/or permanent residents. For example, a funeral home director in Massachusetts is required to be a citizen, while a funeral home director in New York must be either a citizen or a permanent resident (Calvo-Friedman 2014, app. 2). Thus, it seems quite plausible that a lack of citizenship may impede an immigrant’s ability to acquire an occupational license.

To test for the relationship between citizenship status and licensing among immigrants, we repeat our estimation of equation (1) but restrict our sample to immigrants and include a control for whether the individual is a citizen. The results are shown in Table A1 for the CPS and Table A2 for the SIPP. As expected, they confirm that, even after controlling for individuals’ characteristics such as educational attainment, immigrants who are citizens are substantially more likely to

¹² Canada and Australia currently use a points-based immigration system that determines eligibility for immigration partly based on criteria such as age, English- (or French-) language ability, and educational attainment. See Antecol, Cobb-Clark, and Trejo (2003) and Sweetman and Warman (2013) for details regarding immigration to Canada and Australia and how these programs compare with immigration to the United States. In the United States, the Reforming American Immigration for Strong Employment Act (S. 354, revised version S. 1720) introduced by Senators Tom Cotton and David Perdue in 2017 includes a points system for potential immigrants to the United States, also in part on the basis of age, educational attainment, and English-speaking ability.

Table A1
Immigrants' Occupational Licenses: Current Population Survey Sample

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Citizen		.058*** (.0044)		.051*** (.0056)		.062*** (.0070)
Female	.063*** (.0037)	.059*** (.0037)				
YSM/10	.051*** (.0060)	.035*** (.0060)	.035*** (.0072)	.022** (.0072)	.064*** (.0103)	.046*** (.0103)
YSM ² /100	-.009*** (.0017)	-.007*** (.0017)	-.005* (.0020)	-.004 (.0020)	-.013*** (.0029)	-.011*** (.0029)
High school graduate	.030*** (.0039)	.022*** (.0039)	.029*** (.0045)	.023*** (.0045)	.038*** (.0069)	.028*** (.0070)
Some college	.097*** (.0058)	.082*** (.0058)	.086*** (.0073)	.073*** (.0073)	.121*** (.0092)	.105*** (.0093)
BA	.122*** (.0059)	.109*** (.0059)	.096*** (.0071)	.085*** (.0071)	.163*** (.0098)	.146*** (.0098)
More than BA	.159*** (.0070)	.149*** (.0069)	.135*** (.0084)	.127*** (.0083)	.206*** (.0118)	.193*** (.0118)
License (mean)	.127	.127	.094	.094	.170	.170
N	46,294	46,294	25,455	25,455	20,839	20,839
R ²	.075	.081	.050	.056	.089	.094

Note. Results are from linear probability regressions, with individual-level clustered standard errors in parentheses. All estimations include controls for age (as a third-order polynomial), state of residence, month in sample, month of the year, survey year, number of times the respondent is observed, marital status, number of children, and racial or ethnic dummy variables (black, Asian, other, and Hispanic). The omitted group is high school dropouts. YSM = years since migration; BA = bachelor of arts.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

Table A2
Immigrants' Occupational Licenses, Survey of Income and Program Participation Sample

	All		Men		Women	
	(1)	(2)	(3)	(4)	(5)	(6)
Citizen	.061*** (.0130)	.054*** (.0132)	.053*** (.0154)	.046** (.0162)	.068** (.0218)	.062** (.0213)
Female	.063*** (.0112)	.063*** (.0112)				
Speaks English very well		.030* (.0141)		.029+ (.0158)		.030 (.0249)
High school graduate	-.015 (.0142)	-.020 (.0143)	.002 (.0151)	-.001 (.0146)	-.030 (.0246)	-.036 (.0266)
Some college	.047 (.0174)	.036 (.0170)	.029 (.0204)	.020 (.0202)	.081 (.0260)	.068 (.0272)
BA	.068 (.0229)	.055 (.0243)	.015 (.0238)	.004 (.0239)	.134 (.0387)	.120 (.0427)
More than BA	.117 (.0288)	.101 (.0300)	.057 (.0273)	.043 (.0287)	.197 (.0498)	.180 (.0524)
License (mean)	.103	.103	.068	.068	.147	.147
N	3,158	3,158	1,694	1,694	1,464	1,464
R ²	.087	.089	.065	.067	.139	.140

Note. Results are from linear probability regressions, with standard errors (calculated using balanced repeated replication, with Fay's adjustment factor of .5 and 120 survey-provided replicate weights) in parentheses. All regressions include controls for age (as a third-order polynomial), state of residence, union status, marital status, number of children and dummy variables for government worker, paid by the hour, and race or ethnicity (black, Asian, other, and Hispanic). The omitted groups are high school dropouts and immigrants who do not speak English well. BA = bachelor of arts.

+ $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

hold occupational licenses than noncitizens. Of course, since we do not address concerns about self-selection into citizen status, care is needed when interpreting our results, and we cannot conclude that naturalization necessarily increases the licensing rate of immigrants. Bratsberg, Ragan, and Nasir (2002), using data from the 1990s, find that naturalization of US immigrants accelerates their wage growth and reduces employment barriers. Our findings suggest that one possible reason for the accelerated wage growth due to naturalization may be that becoming a citizen provides access to licensed occupations not available to noncitizens. In addition, the returns to naturalization may have risen because of the increasing prevalence of occupational licensing in the economy since the 1990s.

A2. Most Common Occupations of Licensed Workers

To give an idea of the type of occupations that licensed workers perform, Table A3 shows the 10 most common occupations reported among licensed workers in the CPS probability sample. Given the apparent relevance of the most educated group of workers in understanding the native-immigrant licensing gap, we repeat this exercise for the most educated workers.

For both natives and immigrants, licensed women are frequently in nursing or a closely related occupation: nearly a fifth of licensed immigrant women are registered nurses, a further 17.1 percent are nursing orderlies, and 3.1 percent are licensed practical nurses, while 4.8 percent are physicians, 1.9 percent are health aides, and 1.7 percent are physical therapists. Thus, licensed immigrant women are well represented in the medical industry. After nurses, hairdressers and cosmetologists are the most common occupational group for licensed immigrant women at 8.6 percent, which is much higher than the rate for native licensed women (2.3 percent).

For licensed workers with the highest education level, native workers are much more likely than immigrants to be teachers: 37 percent of native women with more than a BA are some type of teacher, while only 20 percent of immigrant women are teachers. For educated men, 23 percent of natives but only 11 percent of immigrants are teachers. So while likely not the full story, it seems that the teaching profession may be a large contributor to the significant native-immigrant licensing gap at the highest level of education and suggests that it may be disproportionately unattractive or difficult for immigrants to enter.

While nursing is a common occupation among licensed workers, the rates for men are much lower than those for women, especially for native male workers. Truck driver is a common occupation for both native and immigrant male licensed workers. Licensed immigrant men are much more likely than native men to be physicians.

Among the most educated group, physician is by far the most common occupation for both immigrant men and women. Lawyer is the most common occupation among highly educated licensed male natives, but only 2.5 percent of highly educated licensed immigrant men are lawyers. These patterns suggest that,

Table A3
Most Common Occupations of Licensed Workers: Current Population Survey Sample

Rank	All		Highly Educated	
	Natives	Immigrants	Natives	Immigrants
Men:				
1	Truck, delivery, or tractor driver (8.1)	Physician (8.6)	Lawyer (16.1)	Physician (26.0)
2	Primary schoolteacher (4.8)	Truck, delivery, or tractor driver (7.0)	Physician (12.6)	Subject instructor (high school or college) (7.0)
3	Manager or administrator, N.E.C. (4.2)	Registered nurse (4.6)	Primary schoolteacher (11.0)	Manager or administrator (5.0)
4	Police or detective (4.0)	Manager or administrator, N.E.C. (3.8)	Secondary schoolteacher (8.4)	Primary schoolteacher (3.6)
5	Lawyer (4.0)	Nursing aide, orderly, or attendant (3.3)	Manager in education (4.3)	Computer software developer (3.3)
6	Secondary schoolteacher (3.3)	Taxi cab driver or chauffeur (2.8)	Subject instructor (high school or college) (3.3)	Accountant or auditor (2.8)
7	Physician (3.1)	Electrician (2.7)	Manager or administrator, N.E.C. (2.9)	Registered nurse (2.7)
8	Electrician (2.7)	Subject instructor (high school or college) (2.7)	Pharmacist (2.6)	Civil engineer (2.5)
9	Registered nurse (2.5)	Cook, variously defined (2.2)	Registered nurse (2.0)	Lawyer (2.5)
10	Fire fighter (2.2)	Computer software developer (2.1)	Accountant or auditor (1.9)	Computer systems analyst (2.5)
Women:				
1	Registered nurse (16.7)	Registered nurse (19.6)	Primary schoolteacher (24.0)	Physician (17.4)
2	Primary schoolteacher (14.3)	Nursing aide, orderly, or attendant (17.1)	Registered nurse (8.0)	Registered nurse (11.8)
3	Nursing aide, orderly, or attendant (5.9)	Hairdresser or cosmetologist (8.6)	Secondary schoolteacher (6.4)	Primary schoolteacher (11.0)
4	Secondary schoolteacher (3.5)	Primary schoolteacher (5.4)	Lawyer (6.4)	Subject instructor (high school or college) (5.6)
5	Licensed practical nurse (2.6)	Physician (4.8)	Physician (5.3)	Lawyer (4.3)
6	Hairdresser or cosmetologist (2.3)	Licensed practical nurse (3.1)	Vocational counselor (4.2)	Secondary schoolteacher (3.6)
7	Lawyer (2.1)	Health aide, except nursing (1.9)	Manager in education (3.8)	Pharmacist (3.5)
8	Health aide, except nursing (2.1)	Accountant or auditor (1.8)	Special education teacher (3.4)	Physical therapist (3.1)
9	Teacher, N.E.C. (2.0)	Physical therapist (1.7)	Social worker (3.2)	Accountant or auditor (3.0)
10	Kindergarten or early-childhood teacher (1.9)	Subject instructor (high school or college) (1.7)	Subject instructor (high school or college) (2.9)	Dentist (2.3)

Note. Numbers in parentheses indicate fractions of the licensed workers in that group. Highly educated workers have more than a bachelor's degree. N.E.C. = not elsewhere classified.

though there is certainly overlap in the occupations common to licensed natives and immigrants, there exist significant differences as well, such as a relative lack of immigrant teachers.

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