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Is Occupational Licensing a Barrier to Interstate Migration?†

By Janna E. Johnson and Morris M. Kleiner*

Occupational licensure may limit the interstate movement of workers because it adds to the cost of moving between states. We analyze the interstate migration of 22 licensed occupations, proxying for the difficulty of the regulations by comparing state-specific licensed occupations to those with national licensing exams. Our empirical strategy also uses individuals who move a long distance, removing the influence of occupation characteristics and self-selection of migration-averse individuals into licensed occupations. Our estimates show that occupational licensing reduces interstate migration, but the magnitude of the effect can only account for a small part of the overall decline in recent decades. (*JEL* J44, J61, R23)

Occupational licensing has become one of the most significant forms of labor market regulation in the United States (US Department of the Treasury Office of Economic Policy 2015). Recent estimates suggest that about 25 percent of the workforce requires a license to work; in 1950, that figure was only 5 percent (US Department of Labor, Bureau of Labor Statistics 2016; Kleiner and Krueger 2010, 2013). Proponents of occupational licensing contend that it protects consumers, ensuring high service quality and protecting the public from harm by ensuring that all service providers have attained a government-mandated minimum qualification level. Previous work has shown that requiring such qualifications restricts entry into these occupations and increases their earnings (Kleiner 2006, 2013),

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but little work has been done to examine the potential of occupational licensing to limit the geographic mobility of individuals. As the majority of licenses are granted at the state level, licensing may limit the ability of workers to move between states, affecting their capacity to take advantage of job opportunities in other places (Roback 1943). Our research is consistent with the more general analysis of labor market barriers that could be erected by both firms and occupational associations (Krueger and Posner 2018, Krueger and Ashenfelter 2018). We provide new detailed, comprehensive evidence of the influence of occupational licensing on the interstate migration of licensed workers across a variety of occupations.

Economists have long recognized the ability of workers to move to different labor markets without restriction as being fundamental to the efficient functioning of those markets (Smith 1776, Friedman 1962). As most occupational licenses are granted at the state level,[1] the cost of attaining licensure in another state can often be significant, even for those already licensed in another state. Despite the growing importance of occupational licensing, the existing literature investigating the link between occupational licensing and geographic mobility is limited and descriptive in nature (Mulholland and Young 2016), likely because of the lack of available information on changes in licensing requirements that could be used in a traditional causal framework. In our case, this information would ideally be the requirements for currently licensed individuals to obtain licensure in another state, which are often lower than the requirements for initial licensure and referred to as licensing "reciprocity" or "endorsement." Several characteristics of the administration of occupational licensing make accurate identification of these state requirements extremely difficult to obtain. Within each state, the licensing of each occupation is overseen by a licensing board or agency specific to a particular occupation or small group of related occupations. The boards or agencies have little incentive to maintain records of historic changes in reciprocity requirements. Perhaps most importantly, in many occupations and states, licensing reciprocity requirements are determined on a case-by-case basis, where individuals currently licensed in another state must work with the licensing board to determine what they must do to transfer their license. These requirements are often based on characteristics such as the individual's current state of licensure and years of experience in the occupation. Given the idiosyncratic nature of boards and their record keeping for these requirements, it is difficult to determine the current requirements for licensure reciprocity, much less what they were in previous years—information required for a traditional causal analysis of the effect of occupational licensing on interstate migration.

Without measurable exogenous variation in relicensure requirements, we pursue an alternative strategy to estimate the effect of occupational licensure on the ability of workers to move between states. We analyze the geographic mobility of 22 occupations universally licensed in every state, and proxy for the difficulty of transferring this license between states using whether the exam(s) required for licensure

[1] A White House report (US Department of the Treasury 2015) estimated that over 1,100 occupations are licensed in at least one state and 60 are licensed in every state, with more than two-thirds of the growth of licensed workers due to the regulation of new occupations.

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in an occupation are specific to a state (such as the bar for lawyers or the pharmacy jurisprudence exams), or are national exams with a single passing standard (such as those for nurses, physicians, and social workers). We define our "treatment" group to be members of state-specific licensed occupations as they are likely to face the highest relicensure costs, potentially including taking additional licensure exams. This treatment group has two natural comparison groups: members of quasi-national licensed occupations (who also must transfer their license when they move between states, albeit at a likely lower cost), and members of all other occupations (who face no such relicensure costs to interstate migration).

Two characteristics of state-specific licensed occupations, unrelated to relicensure costs, could lead to licensed individuals having lower interstate migration rates relative to these comparison groups, negatively biasing estimates of the effect of licensing on interstate migration. First, members of these occupations could have a lower taste for long-distance migration due to the self-selection of risk-averse individuals into these occupations, which often provide a clear career path and stable employment. Second, many state-specific licensed occupations—such as barbers/cosmetologists and real estate brokers—involve the development of a local clientele or network. Moving to another labor market, as often occurs with an interstate move, would result in the elimination of this "local capital," increasing the cost of making such a move for licensed individuals. To address these sources of negative bias, our empirical strategy exploits the detailed migration information in the American Community Survey to limit our sample to individuals who moved 50 or more miles in the previous year. Using only these long-distance migrants allows for estimation of a causal effect of occupational licensure on interstate migration subject to two identifying assumptions: all moves of 50 or more miles result in the loss of local capital, the cost of which is the same for state-specific and quasi-national licensed occupations, and conditioning on observable characteristics removes any other differences in the returns or costs of long-distance migration between these occupation groups.

Our preferred estimates, using only individuals who move 50 or more miles and reside outside their state of birth, show that individuals in state-specific occupations move between states at a 7 percent lower rate compared to members of quasi-national licensed occupations. This difference is much closer to zero than that which uses all individuals regardless of migration status: −58 percent. We subject our results to tests for remaining unobservable bias developed by Altonji, Elder, and Taber (2005) and Oster (2017), which show that this bias must be very large to explain the negative difference. Our findings are also robust to changing the definition of a long-distance move and alternative standard error specifications. Results are nearly identical using data from the Annual Social and Economic Supplement of the Current Population Survey (CPS ASEC). We proceed to examine each state-specific occupation individually, finding variation in the existence of significantly lower interstate migration rates across the occupations but surprisingly little heterogeneity in the size of the effect for those occupations with reduced migration: point estimates for nearly all affected occupations are approximately −10 to −7 percent.

While the limiting effect of licensing on interstate migration is substantial for affected occupations, it is unlikely that the large increase in occupational licensing

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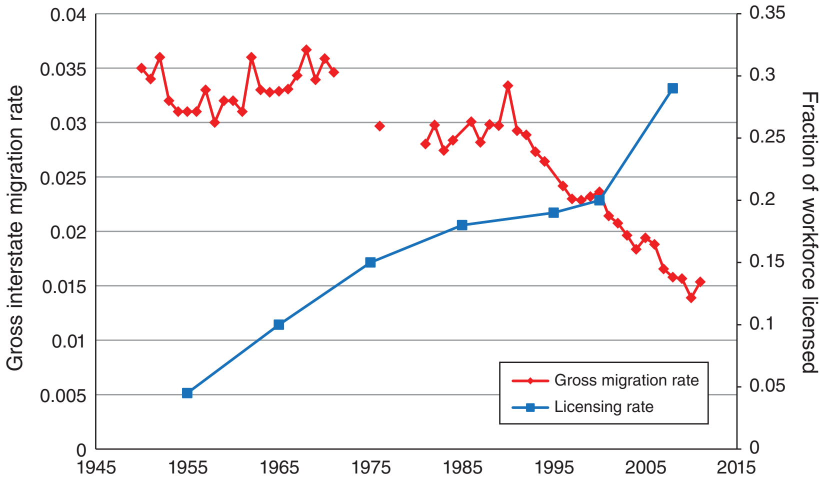

Figure 1. Interstate Migration Rates and Occupational Licensure, 1950–2008

Sources: Gross interstate migration rates from Kaplan and Schulhofer-Wohl (2012). Occupational licensure rates from Kleiner and Krueger (2013).

over the last few decades explains much of the decline in interstate migration experienced over the same period, as shown in Figure 1. If occupational licensing decreases interstate migration by 7 percent on average, as our results indicate, the increase in occupational licensing since 1980 can account for only 2.5 percent of the decline in interstate migration since that time.

Previous work on occupational licensing and interstate migration is quite limited. Several earlier studies showed reduced interstate migration for members of some licensed occupations.[2] More recent work considers the migration of two universally licensed occupations: lawyers and nurses. Tenn (2001) examines the links between the interstate migration of lawyers and their wages, finding that wages are higher in states with lower migration rates. In contrast, DePasquale and Stange (2016) show that the adoption of the Nurse Licensure Compact, which enables registered nurses to practice across state lines without obtaining additional licensure, does not affect the labor supply or the geographic mobility of nurses.

Our study proceeds as follows. Section I presents a simple theoretical framework relating occupational licensing and geographic mobility, which motivates our identifying assumptions. Section II describes the data. Section III outlines the empirical strategy, and Section IV presents our results and robustness tests. In Section V, we summarize, conclude, and present directions for future research.

[2] Holen (1965) showed that dentists and lawyers have limited between-state mobility relative to physicians in the 1950 census. Pashigian (1979) considered the interstate migration of multiple universally licensed occupations, and occupations with little reciprocity between states had lower interstate mobility between 1965 and 1970. In a study of 14 universally licensed occupations using the 1970 census, Kleiner, Gay, and Greene (1982) found a negative relationship between licensing "restrictiveness" (based on state exam and experience requirements) and interstate mobility, as well as a positive correlation between expanded reciprocity and interstate migration rates.

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I. Occupational Licensing and Interstate Migration: A Simple Model

In this section, we present a simple partial equilibrium model that illustrates how occupational licensing can affect interstate migration rates. The model follows the framework of migration models first developed by Sjaastad (1962), which describe the migration decision in terms of relative utility in origin and destination areas. It is an adaptation of the model relating crime rates and labor market conditions in Dix-Carnerio, Soares, and Ulyssea (2016). The model serves mainly as a guide to our empirical investigation and does not intend to be a complete or new theoretical model of the relationship between occupational licensing and interstate migration.

Individuals decide between migrating to a new labor market or staying in their current market.[3] We assume that the individual cares about labor market earnings and knows with perfect certainty the earnings she would receive in her current market (w_s) as well as the potential destination (w_m).[4] Migrating to a new labor market incurs cost c. In addition, individuals are subject to idiosyncratic preference shocks $\varepsilon_{i,m}$ and $\varepsilon_{i,s}$, which tilt preferences toward migrating or staying.

The utilities of migrating and staying are given, respectively, by the following expressions:

$$U_{i,m} = w_m - c + \nu\varepsilon_{i,m},$$

$$U_{i,s} = w_s + \nu\varepsilon_{i,s}.$$

The preference shocks follows standard Gumbel distributions, are independent of each other, and are identically distributed across groups. The scale parameter $\nu > 0$ determines the dispersion of these preference shocks. The migration rate is given by the fraction of individuals who choose to migrate, or $\Pr(U_{i,m} > U_{i,s})$. Following the derivation in Dix-Carniero, Soares, and Ulyssea (2016) and using the properties of Gumbel distributions, we can write the following expression for the (log) migration rate:[5]

$$\log(MR) = \frac{1}{\nu}(V_m - c), \tag{1}$$

where $V_m = w_m - w_s$ is the earnings return to migration. Equation (1) shows that the migration rate is increasing in the relative gain to migration V_m and decreasing

[3] Individuals who move within the same labor market, such as between neighborhoods in a city, are considered as staying in that market, as such moves are likely not accompanied by a change in job.

[4] Adding uncertainty in these earnings does not change the implications of the model for the relationship between occupational licensing and interstate migration.

[5] The derivation of equation (1) is as follows:

$$MR = \Pr(U_{i,m} > U_{i,s}) = \frac{e^{\frac{1}{\nu}(w_m - c)}}{e^{\frac{1}{\nu}(w_m - c)} - e^{\frac{1}{\nu}(w_s)}}$$

$$\Rightarrow \frac{MR}{1 - MR} = \exp\left\{\frac{1}{\nu}(w_m - c - w_s)\right\}$$

$$\Rightarrow \log(MR) \approx \log\left(\frac{MR}{1 - MR}\right) = \frac{1}{\nu}(w_m - w_s - c).$$

The approximation in the last line follows if $MR \ll 1$, which is typically the case.

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in the cost c. The higher the relative earnings gain in the destination, the more individuals choose to migrate; the higher the cost of migration, the more individuals choose to stay put.

Now divide individuals into two different types of occupations: licensed and unlicensed. The cost of moving between labor markets for members of unlicensed occupations is c^U, which includes one-time moving costs, such as transportation to the destination, finding a home and job, and other setup costs incurred at the destination, as well as so-called "psychic costs," such as being farther from family and finding new friends. Those in licensed occupations pay cost c^L, which contains the same components as c^U but is assumed to be larger due to the self-selection of individuals with high risk aversion (who are also assumed to have a strong migration aversion) into licensed occupations.[6] In addition, members of licensed occupations also incur additional costs with a between-labor-market move: c^{LC}, the cost of losing the local capital in their origin labor market and creating it again in the destination, and c^R, the cost of relicensure in their destination. Note that $c^R = 0$ for moves within the same state. If we assume that earnings returns to migration do not vary across occupation type (i.e., $V_m^L = V_m^U$), the difference in (log) *interstate* migration rates between licensed and unlicensed occupations can be written as

$$\text{(2)} \qquad \log\left(MR^L\right) - \log\left(MR^U\right) = -\frac{1}{\nu}\left(c^R + c^{LC} + c^L - c^U\right).$$

The difference in (log) interstate migration rates between members of licensed and unlicensed occupations is a function of the cost of local capital loss c^{LC}, the cost of relicensure c^R, and the difference in migration costs between licensed and unlicensed occupations $c^L - c^U$, which we term the "risk aversion" cost to migration. If c^{LC} and $c^L - c^U$ are sufficiently large relative to c^R, comparing the interstate migration rates of licensed and unlicensed occupations as in equation (2) will yield negatively biased estimates of the effect of relicensure costs $\left(c^R\right)$ on interstate migration.

Rather than considering the fraction of the population who moves between states in a given year, as in equation (2), we now limit the sample to include only those individuals who move a long distance. This restriction removes those who do not move at all or move within their own labor market. As a small fraction of the population moves a long distance within a year,[7] the costs of such migration must be quite high relative to the potential gains. Those individuals who do make such a move are presumed to face lower costs to migration than those who do not (i.e., long-distance migrants must be drawn from the lower end of the distribution of c^U and c^L), as we have assumed that the gains to migration do not vary. As described, c^L is larger than c^U due to the selection of high risk- (and migration-) averse individuals into licensed occupations. By limiting our sample to long-distance migrants in both occupation types, we have removed those with the highest risk aversion (and highest $c^U$and c^L) from the sample, making the average risk aversion costs of migration

[6] Lang and Palacios (2018) provide evidence that one large licensed occupation, teachers, are self-selected to be highly risk averse, at least when it comes to wages.

[7] Using 50 miles to define a long-distance move, the long-distance migration rate in our analysis sample is 3.5 percent for the full population.

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more equal in the subsample of long-distance migrants than in the full population, which includes short- and long-distance migrants as well as non-migrants. If we are willing to assume that among those who move a long distance, the average cost of migration faced by licensed and unlicensed occupations is equal, we remove the $c^L - c^U$ term from equation (2), which now measures the difference in the likelihood that a long-distance move crosses state lines between licensed and unlicensed occupations.

Conditioning on long-distance moves does not help us with local capital costs, c^{LC}, as even among long-distance movers, only those in licensed occupations incur this extra cost to migration. Now we further limit the sample to members of licensed occupations who move a long distance, dividing them into two types: state-specific licensed occupations and quasi-national licensed occupations, each with associated relicensure costs $c^{R,SS}$ and $c^{R,QN}$, respectively.[8] To summarize, under the following assumptions (i) V_m does not vary by occupation type, (ii) conditioning on long-distance movers equalizes other costs of such migration across occupations, removing differences due to selection into occupations due to risk aversion, and (iii) state-specific and quasi-national licensed occupations all incur local capital costs when moving a long distance, we can write the following expression for the difference in interstate migration rates among long-distance migrants in state-specific and quasi-national licensed occupations:

$$\log\left(MR^{SS}\right) - \log\left(MR^{QN}\right) \;=\; -\frac{1}{\nu}\left(c^{R,SS} - c^{R,QN}\right). \tag{3}$$

If $c^{R,SS} > c^{R,QN}$, the model predicts lower interstate migration rates for state-specific licensed occupations relative to quasi-national licensed occupations, and the difference in their migration rates is purely a function of their relative relicensure costs. We can therefore identify the causal effect of relicensure costs on interstate migration by comparing interstate migration rates among long-distance movers in state-specific licensed occupations and quasi-national licensed occupations under the identifying assumptions listed above, plus the condition that we can distinguish long-distance from short-distance moves.

II. Data

For our empirical analysis, we rely on the American Community Survey (ACS) as available through IPUMS-USA (Ruggles et al. 2019). As the largest nationally representative survey that contains detailed migration and occupation measures, the ACS is the existing dataset most suited to studying the relationship between licensing coverage and migration. We use the ACS from 2005 to 2017 for our main analyses, as the detailed migration information we use is first available in 2005. Since we focus on the migration of currently employed and employable individuals, we limit our sample to those aged 18 to 64. The data available through the ACS have information on occupational licensing coverage but not if the individual attained a license (Gittleman

[8] We still must limit the sample to long-distance movers because there may be differences in migration aversion across the two types of licensed occupations, as there are between licensed and unlicensed occupations.

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Table 1—Universally Licensed Occupations Identifiable in the ACS

State-specific licensed occupations		Quasi-national licensed occupations	
Occupation name	ACS code(s)	Occupation name	ACS code(s)
Elementary/secondary teacher	2300, 2310, 2320, 2330, 2340	Nurse (RN/LPN)	3130, 3255, 3256, 3258, 3500
Lawyer	2100, 2105	Physician	3060
Barber/cosmetologist	4500, 4510	Social worker	2010
Real estate broker/sales agent	4920	Occupational and physical therapist	3150, 3160
Electrician	6350, 6355	Psychologist	1820
Insurance agent	4810	Physician assistant	3110
Pharmacist	3050		
EMT/paramedic	3400		
Dental hygienist	3310		
Dentist	3010		
Real estate appraiser/assessor	810		
Veterinarian	3250		
Pest control worker	4240		
Chiropractor	3000		
Optometrist	3040		
Podiatrist	3120		

Notes: Codes listed are 2003–2017 ACS codes. Teacher sample also conditional on industry code 7860 (elementary and secondary schools). State-specific licensed occupations have state licensing exams of varying content and difficulty; quasi-national licensed occupations are licensed at the state level, all requiring passage of a national exam for licensure. For more details see online Appendix Table A1.

and Kleiner 2016). However, for the occupations we study, licensing coverage and attainment are largely indistinguishable; to be a legally practicing member of these occupations, one must hold a license.[9]

The 22 licensed occupations we examine are shown in Table 1. We chose these occupations based on the following criteria: (i) they were uniquely identifiable using ACS occupation codes, (ii) they were universally licensed in all states, and (iii) entry into the occupation requires licensure, so all members of an occupation must be licensed. These occupations cover a wide variety of skill and income levels—from barbers and cosmetologists, electricians, and pest control workers to lawyers, physicians, and dentists—and represent a range of industries. Some occupations with largely similar tasks, such as occupational and physical therapists, were merged by combining two or more ACS occupational categories to increase the sample size.

As shown in Table 1, we divide these occupations into two groups: "state-specific" and "quasi-national" occupations. As discussed, requirements for attaining licensure vary substantially both across occupations and across states within occupations. Since we are interested in the relationship between state licensing requirements and interstate migration, we require a way to distinguish occupations by their potential ability to transfer an existing license from state to state. Given the lack of information on the exact requirements to do so, we base this classification on a licensing requirement common to all 22 occupations: the passage of a licensing exam. The exam(s) required for licensure take two forms: (i) a common national exam with

[9] This contrasts with other occupations such as accounting and engineering, where individuals can work as a practicing accountant or engineer without holding a license but are restricted in the tasks they can perform (see Hur, Kleiner, and Wang 2018).

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a single passing standard or (ii) an exam with varying content and difficulty from state to state. As well as being a concrete method of classification, licensing exam requirements are much easier to identify than other requirements that potentially vary from state to state, such as additional training and practical experience requirements or board discretion in granting licenses. We ignore such requirements in our classification.[10] Occupations for which at least one licensing exam varies between states are "state-specific" licensed occupations, and those with only national exams (but with licenses granted at the state level) are "quasi-national" licensed occupations. Some occupations, such as pharmacists and veterinarians, have both a national licensing exam as well as a state-specific exam that tests either clinical skills or knowledge of relevant state laws (often called a "jurisprudence" exam). Since the passage of this state exam is required for licensure, we have placed these occupations in the state-specific category. Details on the exam requirements for licensure for each occupation are shown in online Appendix Table A1. The 22 occupations we analyze make up 11 percent of the US labor force, with the state-specific licensed occupations accounting for 7 percent and the quasi-national licensed occupations accounting for 4 percent.[11]

Our empirical strategy relies on the ability to distinguish long-distance from short-distance moves. Starting in 2005, the ACS provides location of current residence, as well as last year's residence for movers, at geography below the state level—the Public Use Microdata Area of migration (MIGPUMA), which is a unit of approximately 100,000 or more residents. MIGPUMAs roughly correspond to a county for densely populated areas and a larger area for more rural areas, and all MIGPUMAs are contained within a single state.[12] For movers between MIGPUMAs, we measure move distance as the straight-line distance between the centroids of the current and former MIGPUMA of residence.

In a robustness analysis, we use the Annual Social and Economic Supplement of the Current Population Survey (CPS ASEC), formerly known as the March CPS. The CPS ASEC allows for identification of the same 22 licensed occupations as the ACS and measures migration in the previous year. However, the CPS ASEC does not contain information on sub-state place of residence last year for movers, meaning we cannot measure move distance as we can in the ACS. It also does not contain state of birth information, a key control variable in our ACS analysis. However, the CPS ASEC reports an individual's occupation in the prior year, information not contained in the ACS. We therefore use the CPS ASEC to show that our results are robust to conditioning on prior year's occupation, as described in Section IVD.

[10] Completion of these additional requirements is often waived for existing license holders seeking to transfer existing licenses between states, another reason to exclude these requirements from our classification scheme. This waiver process is usually done on an individual basis at the discretion of the state licensing board. However, it is unlikely that such a waiver would cover passage of a state-specific exam, such as the pharmacy jurisprudence exam.

[11] In contrast to the 11 percent of the workforce we analyze, the percent of the workforce that belonged to a union was 10.5 percent in 2018 (BLS 2019).

[12] Some MIGPUMAs combine two or more Public Use Microdata Areas (PUMAs) of residence, so the two do not perfectly correspond. For more information, see IPUMS-USA, MIGPUMA1, "Description," https://usa.ipums.org/usa-action/variables/MIGPUMA1#description_section.

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III. Empirical Strategy

Without a source of exogenous variation in licensing requirements, we turn to an alternative strategy to estimate the effect of occupational licensure on interstate migration. Our strategy exploits the current and former place of residence information in the ACS to form a measure of move distance for migrants. Based on our classification of occupations described in the previous section, we use the 16 state-specific licensed occupations as our treatment group, as occupations in this category are likely to face the highest barriers to move their license from one state to another. Our primary estimating equation takes the following form:

$$(4) \quad movedbtstates_{isrt} = \beta state\_specific_{isrt} + X_{isrt}\gamma + \alpha_s \times \eta_t + \theta_r + \varepsilon_{isrt},$$

where $movedbtstates_{isrt}$ is an indicator for individual i born in state r residing in state s in year t moving between states in the previous year; $state\_specific_{isrt}$ is equal to one if individual i's current occupation is one of the 16 state-specific licensed occupations in Table 1; X_{isrt} is a vector of observable individual controls, $\alpha_s \times \eta_t$ are state-year fixed effects (defined based on last year's state of residence), θ_r are state of birth fixed effects, and ε_{isrt} is a conventional error term.[13] We limit the sample to individuals who moved at least 50 miles in the last year, so the coefficient of interest β measures the likelihood that a 50-mile move made by individuals in state-specific licensed occupations over the previous year crossed state lines relative to a comparison group.

Our licensing classification scheme shown in Table 1 generates two possible comparison groups for state-specific licensed occupations: (i) the six quasi-national licensed occupations and (ii) members of all other occupations. The quasi-national licensed occupations are universally licensed in every state, just like the state-specific treatment group. However, given that their licensing exam is a national exam, individuals in quasi-national licensed occupations are likely to face lower relicensure costs than those in state-specific licensed occupations, as they do not have to retake a licensing exam when moving between states. As all individuals who move between states in the sample must transfer their license to continue working in their occupation, using the quasi-national licensed occupations as the comparison group in equation (4) implies β estimates the effect of the more costly license transfer faced by state-specific licensed occupations relative to that faced by members of quasi-national licensed occupations.

The coefficient β using the other natural comparison group—all other occupations—captures the relative effect of facing a costly licensure transfer process to not facing any licensure barriers to moving between states. (We exclude the six quasi-licensed occupations from the sample when we use this comparison group.)

[13] The vector X_{ist} consists of controls for education (dummies for high school completion, some college, 4 years of college, and more than 4 years of college, with less than high school the excluded group), age (dummies for 5-year age groups 20–24, 25–29, . . . , 60–64, with ages 18–19 the excluded group), income (quartile dummies), race (dummies for non-Hispanic white, non-Hispanic black, Hispanic white, and other), marital status (dummies for married, divorced, widowed, single), employment status (dummies for employed, unemployed, not in labor force), citizenship status, and number of children (dummies for 0, 1, 2, 3, 4+).

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However, this group does contain some licensed occupations: we could not cleanly identify some universally licensed occupations in the ACS occupation codes, as well as occupations that are licensed in some states but not others. Therefore, as part of the comparison group is subject to relicensure costs, estimates using unlicensed occupations as a comparison are a lower bound of the effect of state-specific licensing requirements on interstate migration.

Estimates of β in equation (4) constitute a causal effect of state-specific licensure on interstate migration under three assumptions: (i) using only individuals who move 50 or more miles equalizes average risk aversion costs of migration across treatment and comparison groups, (ii) all moves of 50 or more miles result in the destruction of "local capital" for licensed occupations, and (iii) conditioning on moving such a distance as well as our included controls removes all other sources of unobserved selection bias correlated with entering a state-specific licensed occupation and moving between states.

The first assumption is necessary as individuals who select into state-specific licensed occupations may have higher levels of risk aversion than others. Most of these occupations have clear career paths with a high probability of employment once licensure is attained and a strong likelihood of continued stable employment, as labor demand for these occupations is relatively unaffected by macroeconomic conditions (particularly for those in the medical field) (Gittleman, Klee, and Kleiner 2018). These characteristics could lead to members of state-specific occupations having higher average risk aversion than those in other occupations. If those with higher risk aversion are also less likely to migrate between labor markets, this self-selection may account for the lower geographic mobility of members of state-specific licensed occupations. This difference in risk aversion may be mitigated by our use of only individuals who move 50 or more miles, as this requirement excludes individuals so risk averse they are not willing to make such a move at all, potentially equalizing average risk aversion across treatment and comparison groups. However, as many of the quasi-national occupations also have stable employment and clear career paths (physicians, nurses, etc.), the first assumption may be more likely to be satisfied using the quasi-national licensed occupations as the comparison group for state-specific licensed occupations rather than members of all other occupations.[14]

The second assumption—that moves of 50 or more miles are between-labor-market moves and therefore result in the loss of local capital—addresses the fact that success in many state-specific licensed occupations involves developing and maintaining a local reputation or clientele network (or both), such as that for lawyers, real estate agents and appraisers, barbers/cosmetologists, veterinarians, and so on.[15] A move between states would result in the loss of this valuable local capital if such a move also involved changing geographic labor markets. Therefore, reduced interstate

[14] We acknowledge the assumption that conditioning on 50-mile migrants equalizes risk aversion across occupation groups is strong. However, we believe this restriction minimizes the differences in risk aversion between state-specific and quasi-national licensed occupations, and results using this sample certainly present a less-biased estimate of the effect of occupational licensing on interstate migration than using all individuals.

[15] Elementary and secondary teachers do not have a local clientele or network like lawyers and real estate agents, but often tenure and other benefits are specific to seniority in each school district, which could also strongly deter teachers from moving out of their local area.

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mobility of members of state-specific licensed occupations could be a result of long-distance moves destroying local capital and not due to occupational licensing. However, some of the quasi-national occupations, such as physicians, occupational and physical therapists, and psychologists, may also incur a loss of local capital with a long-distance move, particularly if the individual is self-employed.[16] Some occupations in the all other occupations comparison group may also have a local capital component, but we believe the first identifying assumption is more likely to be satisfied using the quasi-national comparison group rather than all other occupations. However, assuming all 50-mile moves result in changing labor markets, using only individuals who move such a distance ensures that all migrants have incurred the cost of local capital destruction if it applies to their occupation.[17]

Satisfying the third assumption entails accounting for all potential sources of bias correlated with being in a state-specific licensed occupation and interstate migration, including differences in the returns and costs of migration. We include a rich set of observable characteristics as well as state-year fixed effects in our specifications to control for such differences. However, there is another potential unobserved source of an increased cost of interstate migration for those in state-specific licensed occupations relative to quasi-national occupations: a strong affinity for, and therefore a reluctance to move out of, one's home state. Consider a young person born in state r deciding which occupation she wants to enter. She knows she wants to remain in state r, whether because she really likes living in r or for some other reason, such as all her family live in r and she prefers to stay near her family. Given the occupational opportunities in r, she may be more likely to enter a state-specific licensed occupation, as she knows she can get a job as a teacher/pharmacist/real estate agent/electrician in her home state no matter if r is Iowa or New York. Such a mechanism could make members of state-specific licensed occupations less likely to move out of their home state than members of other occupations. This selection could also be enforced through the state-specific nature of training to enter such occupations: if an individual born in state r wants to become a teacher, she likely attends the state university in r and receives training designed to satisfy the specific licensing requirements for teachers in state r, which could be quite different than the licensing requirements for teachers in other states. In contrast, if this individual attends the same state university to become a nurse, that training is designed to satisfy the nurse licensure requirements in state r, which are quite similar to those in other states, as they are all based on the national nurse licensing exam. The interstate mobility of a teacher born and trained in state r may therefore be more limited than that of a nurse born and trained in state r. While this may also be considered an "effect" of state-specific licensing on interstate migration, we control for this selection by repeating our analysis using only those individuals residing outside their state of birth in the previous year (state of birth is the only measure of "home state" available in the ACS). Estimates using this sample are less likely to be affected by an increased preference among state-specific licensed

[16] In our ACS sample, about 25 percent of physicians and 33 percent of psychologists are self-employed, compared to 65 percent and 74 percent of (state-specific licensed) dentists and chiropractors, respectively.

[17] We show our results are robust to changes in the 50-mile move definition in Section IVD.

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occupations to remain in their home state than those using all individuals, regardless of birth state.

In sum, our preferred specification compares the likelihood that a move of 50 or more miles was between states for state-specific licensed occupations relative to quasi-national licensed occupations, using only those migrants who were living outside their state of birth last year. We believe this specification is most likely to satisfy the assumptions necessary for this estimated difference to be the causal effect of relicensure requirements on interstate migration.

IV. Results

We begin this section by discussing descriptive differences in migration rates and other characteristics between our treatment and comparison groups. We proceed to estimate the effect of state-specific licensing requirements on interstate migration, and show the sensitivity of our main specification to the inclusion of different control variables. We then provide evidence of the robustness of the results to changing definitions of long-distance migration and different standard error estimation procedures. We also perform a supplementary analysis using the CPS ASEC. Finally, we present our occupation-specific results.

A. *Descriptive Differences in Migration Rates*

Table 2 shows selected descriptive statistics for the full ACS sample as well as our treatment group (state-specific licensed occupations) and two comparison groups (quasi-national licensed occupations and all other occupations).[18] Fifteen percent of the population moves any distance within a year—most of them local, within-MIGPUMA moves. The other moves are split approximately equally between moves between states and between-MIGPUMA moves within a state.[19] Focusing on the migration differences between the state-specific and quasi-national licensed occupations, state-specific licensed occupations are much less likely to move between states: both licensed occupation groups move at similar rates at local and mid-distance levels within a state (within MIGPUMA and between-MIGPUMA/within state), but only 2.0 percent of individuals in state-specific occupations moved between states in the past year, while 2.9 percent of quasi-national occupations did. This pattern persists when we consider the move distance of migrants. Migrants in state-specific licensed occupations are less likely than quasi-national occupations to move 50 or more miles (24.4 percent versus 28.7 percent). Most notably, individuals in state-specific occupations who move 50 or more miles are much less likely to be interstate migrants than members of quasi-national licensed occupations (61 percent versus 71 percent). However, state-specific licensed individuals are much more likely to reside in their state of birth, potentially also a factor in their reduced tendency to move between

[18] Online Appendix Table B1 shows full descriptive statistics for all our control variables for these four groups, as well as for each of the 16 state-specific licensed and 6 quasi-national licensed occupations separately.

[19] All MIGPUMAs are contained within a single state, so a move between states is also a move between MIGPUMAs.

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Table 2—Selected Descriptive Statistics, 2005–2017 ACS

	Full sample	State-specific licensed occupations	Quasi-national licensed occupations	All other occupations
Moved at all	0.148	0.126	0.134	0.150
Moved within MIGPUMA	0.096	0.078	0.079	0.098
Moved beween MIGPUMA ∣ within state	0.027	0.027	0.026	0.027
Moved between states	0.025	0.020	0.029	0.025
Moved 50+ miles, given moved at all	0.237	0.244	0.287	0.235
Moved between states, given moved 50+ miles	0.652	0.613	0.705	0.652
Lived outside state of birth last year	0.472	0.443	0.509	0.473
Mean years of education	13.45	16.18	16.13	13.17
Education				
Less than high school	9.93	1.05	0.16	10.88
High school graduate	34.48	12.82	5.15	37.02
Some college	25.58	14.76	30.33	26.11
Bachelor's degree	19.41	30.81	32.15	18.18
More than bachelor's degree	10.59	40.56	32.20	7.80
Race				
Non-Hispanic white	67.22	79.78	73.25	66.17
Non-Hispanic black	10.66	6.75	10.47	10.92
Hispanic white	9.25	5.80	4.40	9.66
Other	12.87	7.67	11.88	13.25
Female	51.70	62.80	80.30	49.89
Mean age	41.22	42.81	43.41	41.03
Mean labor income US$(2017)	39,435	62,085	78,563	36,468
Observations	17,953,437	1,166,297	666,335	16,120,805

Notes: Sample includes all individuals aged 18–64 residing in the 50 US states and DC not residing in group quarters with nonimputed values for migration status, education, income, occupation, age, sex, race, citizenship status, marital status, and employment status, excluding those who lived outside the 50 US states and DC in the previous year. Sample also excludes individuals residing in the PUMAs of migration affected by Hurricane Katrina in Lousiana and those residing in PUMA of migration 51000 in Virginia in the current or previous year. Move distance calculated as distance between centroids of current and previous PUMA of migration.

states. Comparing the migration of state-specific licensed occupations to that of members of other occupations yields similar conclusions.

The remaining statistics in Table 2 show that both state-specific and quasi-national licensed occupations have higher education, are less likely to be a member of a minority group, and earn a higher average income than those in other occupations. A higher fraction of members of both licensed occupation groups are female relative to members of other occupations due to the presence of two large female-dominated occupations (teachers and nurses) in each licensed group.

B. *The Effect of State-Specific Licensing Requirements on Interstate Migration*

Estimates of the effect of state-specific licensing requirements on interstate migration are shown in Table 3. Panel A uses quasi-national licensed occupations as the comparison group, and panel B uses members of all other occupations. The dependent variable in all specifications is an indicator for moving between states in

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Table 3—Interstate Migration and Occupational Licensing, 2005–2017 ACS

		Moved between states given moved 50+ miles	
	Moved between states (1)	All individuals (2)	Outside state of birth (3)
Panel A. State-specific versus quasi-national licensed occupations			
State-specific licensed	−0.014	−0.073	−0.051
	(0.001)	(0.009)	(0.007)
Dependent variable mean	0.024	0.652	0.774
Percentage effect	−58.33	−11.20	−6.59
R^2	0.025	0.146	0.126
Observations	1,832,632	55,368	34,271
Panel B. State-specific licensed occupations versus all other occupations			
State-specific licensed	−0.011	−0.084	−0.057
	(0.001)	(0.008)	(0.009)
Dependent variable mean	0.024	0.650	0.774
Percentage effect	−45.83	−12.92	−7.36
R^2	0.016	0.113	0.088
Observations	17,287,102	544,140	330,044

Notes: Sample described in notes to Table 2. All specifications include last year's state of residence × year fixed effects, state of birth fixed effects, and controls for income, race, sex, education, marital status, age, employment status, citizenship status, and number of children. Percentage effects calculated as coefficient/dependent variable mean × 100. Sample in panel A excludes members of quasi-national licensed occupations, and sample in panel B includes only members of licensed occupations. Estimated using OLS and sample weights. Standard errors clustered on last year's state of residence in parentheses.

the past year, and all specifications include the full set of controls in equation (4). As a basis of comparison for estimates using our preferred empirical strategy, column 1 contains results comparing the likelihood of moving between states between treatment and comparison groups using the full ACS sample; individuals who do not move between states in this sample include both non-migrants and within-state migrants. Results show that members of state-specific licensed occupations move between states at a 1.4 percentage point lower rate than members of quasi-nationally licensed occupations and 1.1 percentage points lower than other occupations. Scaling by the mean interstate migration rate in the sample, state-specific licensed occupations move between states at a rate 58 percent or 46 percent lower than members of each comparison group, respectively. However, these differences are likely contaminated with negative selection bias from the local capital-intensive nature and higher risk aversion of state-specific licensed occupations, both of which could lead to a lower observed interstate migration rate for reasons unrelated to relicensure costs.

The remaining columns in Table 3 limit the sample to migrants who moved at least 50 miles in the previous year. Column 2 uses all such migrants, and column 3 uses only those migrants who lived outside their state of birth last year. In all cases, the estimated difference in interstate migration for these long-distance migrants is much closer to zero than that in column 1. Given a move of 50 or more miles, individuals in state-specific licensed occupations are 11 percent less likely to move between states than members of quasi-national licensed occupations, and 13 percent less likely than members of all other occupations. When the sample is limited to those residing outside their state of birth in the previous year, these differences

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shrink to approximately −7 percent using both comparison groups. These estimates constitute a causal effect of relicensure costs on interstate migration for these occupations, subject to the assumptions outlined in Section III.

In the online Appendix, we repeat this analysis using two alternative treatment groups: all licensed occupations combined (all 22 licensed occupations in Table 1), and the 6 quasi-national licensed occupations. Results are shown in online Appendix Table B2. The comparison group in both cases is all other occupations (excluding state-specific licensed occupations for the quasi-national treatment group). Results using the 50-mile migrant sample residing outside their state of birth show a 5 percent reduction in interstate migration for all licensed occupations combined, and a 2 percent reduction for quasi-national occupations. Both estimates are strongly statistically significant. Online Appendix Table B3 repeats the analyses in Table 3 and online Appendix Table B2 using a cell matching estimator (Black 2015), with virtually identical results.

C. *Control Sensitivity and Bounding a Causal Effect*

We use methods first developed by Altonji, Elder, and Taber (2005) (hereafter AET) and further developed by Oster (2017) to examine the sensitivity of our results in Table 3 to observable and unobservable selection bias. AET and Oster's work shows that simply comparing coefficient changes with the addition of control variables is not enough to show that bias from unobserved sources is negligible. Such changes must also be scaled by changes in R^2.

Table 4 repeats the analysis from panel A, columns 2 and 3 of Table 3, progressively adding more controls. Panel A of Table 4 uses the sample of 50-mile migrants; panel B uses only those migrants residing outside their state of birth last year. Column 1 shows the simple bivariate regression of an indicator for moving between states on an indicator for being in a state-specific licensed occupation. (Quasi-national licensed occupations are the comparison group in Table 4. Online Appendix Table B4 repeats the exercise using all other occupations as the comparison group.) Columns 2 through 6 progressively add more control variables, including education, age and sex, all other observable controls, state-year fixed effects, and state of birth fixed effects. Column 5 repeats our main specifications from Table 3. Column 6 adds interactions between the vector of age category dummies and state-year fixed effects; and between the age category dummies and state-of-birth fixed effects. In both panels, the point estimate and standard error of the effect of state-specific licensing changes slightly, but R^2 increases substantially.

AET and Oster's method relies on assumptions about the values of two parameters: the relative importance of selection on observed and unobserved variables (denoted δ), and the hypothetical R^2 from a regression of the outcome on all observed and unobserved variables ($R_{\max}$). The method also assumes that bias from unobservable characteristics moves the coefficient point estimate the same direction as that from observable characteristics.

Under these assumptions, we can measure the influence of omitted variable bias by either placing bounds on the treatment effect or by calculating the relative degree of unobservable selection δ that would result in a treatment effect of zero. (The latter

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TABLE 4—CONTROL SENSITIVITY AND TREATMENT EFFECT BOUNDS, STATE-SPECIFIC LICENSED OCCUPATIONS VERSUS QUASI-NATIONAL LICENSED OCCUPATIONS, 2005–2017 ACS

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. All 50+-mile migrants						
State-licensed occupation	−0.092	−0.082	−0.088	−0.075	−0.073	−0.074
	(0.011)	(0.010)	(0.013)	(0.009)	(0.009)	(0.009)
Dependent variable mean	0.652	0.652	0.652	0.652	0.652	0.652
Percentage effect	−14.11	−12.58	−13.50	−11.50	−11.20	−11.35
Treatment effect bounds and deltas						
$R_{\max} = 0.34$						$(-0.074, -0.066)$, $\delta = 5.38$
$R_{\max} = 0.5$						$(-0.074, -0.046)$, $\delta = 1.90$
$R_{\max} = 0.7$						$(-0.074, -0.006)$, $\delta = 1.05$
$R_{\max} = 1$						$(-0.074, 0.138)$, $\delta = 0.63$
R^2	0.009	0.026	0.037	0.138	0.146	0.258
Observations	55,368	55,368	55,368	55,368	55,368	55,368
Panel B. 50+-mile migrants residing outside state of birth last year						
State-licensed occupation	−0.049	−0.046	−0.058	−0.049	−0.051	−0.054
	(0.007)	(0.006)	(0.009)	(0.007)	(0.007)	(0.008)
Dependent variable mean	0.774	0.774	0.774	0.774	0.774	0.774
Percentage effect	−6.33	−5.94	−7.49	−6.33	−6.59	−6.98
Treatment effect bounds and deltas						
$R_{\max} = 0.37$						$(-0.056, -0.054)$, $\delta = 15.77$
$R_{\max} = 0.5$						$(-0.060, -0.054)$, $\delta = 6.46$
$R_{\max} = 0.7$						$(-0.071, -0.054)$, $\delta = 3.39$
$R_{\max} = 1$						$(-0.132, -0.054)$, $\delta = 1.98$
R^2	0.003	0.011	0.022	0.120	0.126	0.283
Observations	34,271	34,271	34,271	34,271	34,271	34,271
Controls						
Education, age, sex		X	X	X	X	X
Other controls			X	X	X	X
State × year fixed effects				X	X	X
State of birth fixed effects					X	X
Age interacted with fixed effects						X

Notes: Sample described in notes to Table 2 and further limited to members of licensed occupations listed in Table 1. Other control variables include income, race, marital status, employment status, citizenship status, and number of children. Estimated using OLS and sample weights. Treatment effect bounds calculated assuming $\delta = 1$ and stated $R_{\max}$. Deltas calculated assuming treatment effect of 0 and stated $R_{\max}$. Standard errors clustered on last year's state of residence in parentheses.

is AET's original test statistic, which also assumes an $R_{\max}$ of 1.) We report both measures in column 6 of Table 4 under different assumptions about $R_{\max}$. Oster suggests an $R_{\max}$ of 1.3 times the R^2 from the fully controlled model (column 6), which is 0.34 for the specification using all migrants, and 0.37 for that using only those residing outside their state of birth last year. In both specifications, the treatment effect bounds are tight and do not include zero.[20] The controlled coefficient of −0.054 is the *upper bound* of the treatment effect for those outside their state of birth, as this value is *lower* than the uncontrolled estimate of −0.049. Alternatively, we report the value of δ required for the true treatment effect to be zero. Using $R_{\max}$ values of 0.34 and 0.37 the required δ is 5.38 for all migrants and 15.77 for those outside their state

[20]These and all reported bounds are calculated under the assumption of $\delta = 1$, or equal selection on observables and unobservables.

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of birth, respectively. In other words, the remaining selection on unobservable characteristics (such as differences in the returns or costs to interstate migration between state-specific and quasi-national licensed occupations) must be 5 (or 15) times more important than that from the included observable characteristics (and fixed effects) for the causal effect of state-specific licensing on interstate migration to be equal to 0.

As these suggested values for $R_{\max}$ may be too low, we report bounds and δs for three other possible values of $R_{\max}$: 0.5, 0.7, and 1. Using all 50-mile migrants, the bounds do not include zero for $R_{\max}$ values up to 0.7 and, as the lower bound moves downward, never include zero for those residing outside their state of birth. For this group, the unobservable selection must be nearly twice as important as the observable selection for the true effect to be zero, even with an $R_{\max}$ of 1.

Online Appendix Table B4 repeats the exercise using all other occupations as the comparison group. Results are similar to those in Table 4 using only those outside their state of birth: as the controlled coefficient is lower than the uncontrolled coefficient, the controlled coefficient is the upper bound of the treatment effect, since the method assumes unobserved bias changes the point estimate in the same direction as that from observable characteristics. The treatment effect bounds therefore never include zero.

D. *Additional Robustness Tests*

In the online Appendix, we provide several more tests of the robustness of our results. Online Appendix Table B5 repeats the analysis in Table 3 using a 100-mile move, and results are very similar to those using a 50-mile move. The ACS changed the definition of MIGPUMAs in 2012, which could affect our move distance measures. Online Appendix Table B6 shows that results are nearly identical in the earlier period using 2000 census definitions (2005–2011) as the later 2010 definitions (2012–2017). Given the dominance of two large occupations in our state-specific and quasi-national licensed groups—teachers and nurses, respectively—we show results excluding them from the sample in online Appendix Table B7. Estimated effects are nearly identical to the results using the full sample shown in Table 3.

All our specifications estimate standard errors clustering at the state of residence in the previous year. Online Appendix Table B8 provides standard errors for our main results using several alternative estimators. While standard errors increase using clustering at the occupation level and two-way clustering on occupation and state of residence last year, the estimated effect of state-specific licensing on interstate migration for 50-mile movers remains strongly statistically significant in all specifications (p-values are all less than 0.02). The p-values produced using randomization inference permuting an occupation's licensure status (i.e., either state-specific or quasi-national, or state-specific or all other occupations, depending on the comparison group used) within state-year strata are zero for all specifications using 1,000 replications.

One limitation of using the ACS data to study the effect of occupational licensing on interstate migration is the inability to observe an individual's occupation in the previous year. This means we cannot distinguish "continuing" members of an

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occupation, for whom relicensure costs would be relevant when moving between states, from new entrants into the occupation, who would not face relicensure costs but instead initial licensure costs. As we cannot distinguish these two groups, our estimates capture the effect of both relicensure and initial licensure costs, which could affect an individual's decision to migrate in different ways. Fortunately, the Annual Social and Economic Supplement of the Current Population Survey (CPS ASEC, formerly known as the March CPS) *does* record occupation of last year, so we repeat our analysis using this dataset.

While the CPS ASEC (Flood et al. 2018) does allow us to distinguish continuing members of an occupation from new entrants, it has several major disadvantages relative to the ACS that lead us to not use it in our main analysis. First, the CPS ASEC has a much smaller sample size than the ACS—about 100,000 observations per year compared to 1.4 million per year in the ACS. Since our analysis focuses on a small subset of the population, the small sample size of the CPS ASEC limits its usefulness. Second, the CPS ASEC does not allow for the measurement of move distance, as it does not contain information on sub-state place of residence in the prior year. Third, the CPS ASEC does not report state of birth.[21] The lack of detailed sub-state location and state of birth in the CPS ASEC means we cannot implement our empirical strategy using migrants who move 50 or more miles as we can in the ACS. However, the CPS ASEC does allow for the identification of the same 22 state-specific and quasi-national licensed occupations as the ACS.

We provide migration statistics using the 2005–2018 CPS ASEC in online Appendix Table B9. The overall migration rate (12.0 percent) and the interstate migration rate (1.6 percent) are lower than that in the ACS, consistent with previous research (Kaplan and Schulhofer-Wohl 2012).[22] However, the patterns across the different occupation categories are similar to those observed in the ACS (i.e., individuals in state-specific licensed occupations migrate between states at a lower rate than those in quasi-national licensed occupations and all other occupations).

Online Appendix Table B9 also reports the fraction continuing members of an occupation by migration status and licensure group. Continuing members are those reporting the same 3-digit CPS occupation in the current and previous year. Interstate migrants have the lowest fraction continuing members among all occupation types, but this fraction is substantially higher among the two licensed occupation groups relative to all other occupations (82.5 and 87.2 percent for state-specific and quasi-national licensed occupations, respectively, and 66.9 percent for all other occupations). However, licensed occupations tend to be stable, easily defined occupations, while members of unlicensed occupations could potentially switch occupations as defined by the 3-digit occupation code without changing the nature of their job (for example, there are around 30 codes for various types of managers). The fraction continuing members is relatively similar across within-state/between-county movers (the best proxy for long-distance within-state migrants available in

[21] The CPS ASEC does contain a birthplace variable, but it only identifies country of birth and does not identify state of birth for those born in the United States.

[22] The CPS ASEC within-county and between-county/within-state migration rates are not comparable to the ACS within-MIGPUMA and between-MIGPUMA/within-state migration rates, as counties and ACS MIGPUMAs do not necessarily correspond.

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the CPS ASEC) and interstate movers for both groups of licensed occupations (86.7 versus 82.5 percent for state-specific; 90.3 versus 87.2 percent for quasi-national). We therefore find it unlikely that the inability to distinguish continuing members of an occupation from new entrants affects our ACS results substantially.

Results in online Appendix Table B10 substantiate this belief. The table repeats the analyses in Table 3 and online Appendix Table B2, showing that results using all members of an occupation are very similar to those using all occupations for all four treatment/comparison group combinations. For example, the difference in interstate migration rates between members of state-specific and quasi-national licensed occupations, using only individuals who moved between counties, is −27 percent using all current members of these occupations, and −28 percent using only continuing members of the occupations.[23]

Table 5 repeats the control sensitivity and bounding analysis from Table 4 for these CPS ASEC results. The table uses quasi-national occupations as the comparison group for state-specific licensed occupations, limiting the sample to individuals who moved between counties in the past year. Panel A uses the full migrant sample and panel B uses only continuing occupation members. Columns 1 through 4 progressively add more controls, and column 5 adds interactions between the state-year fixed effects and age category dummies. The addition of these interactions does not substantially change the point estimate of the difference between interstate migration rates of state-specific and quasi-national licensed occupations, but the R^2 increases from approximately 0.3 to 0.7 in both panels of the table. However, given the much smaller sample size of the CPS ASEC (approximately 3,000 long-distance migrants compared to 55,000 in the ACS) and the large number of parameters added to the model through these interactions, overfitting is a concern. The AET/Oster treatment effect bounds do not include zero even with an $R_{\max}$ of one, and a treatment effect of zero requires approximately equal selection on unobservable and observable characteristics.[24]

E. *Occupation-Specific Results*

We proceed to investigate whether the relationship between state-specific licensure and interstate migration varies across licensed occupations. To do so, we use a variation of the specification in equation (4):

$$(5) \qquad movedbtstates_{isrt} = \sum_{j=1}^{J} \beta_j \mathbf{1}\big[occupation = j\big]_{isrt} + X_{isrt}\gamma + \alpha_s \times \eta_t + \theta_r + \varepsilon_{isrt},$$

[23] These estimated differences are larger than the corresponding estimate from the ACS (−11 percent), likely because many between-county migrants in the CPS ASEC do not change jobs as they move between adjacent counties or counties within the same metropolitan area.

[24] Online Appendix Table B11 repeats the bounding exercise using the CPS ASEC and the all other occupations comparison group. Results are similar to those shown in Table 5.

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Table 5—Control Sensitivity and Treatment Effect Bounds, State-Specific Licensed Occupations versus Quasi-national Licensed Occupations, 2005–2018 CPS ASEC

	(1)	(2)	(3)	(4)	(5)
Panel A. All migrants between counties					
State-licensed occupation	−0.114	−0.110	−0.115	−0.111	−0.100
	(0.026)	(0.025)	(0.026)	(0.031)	(0.060)
Dependent variable mean	0.410	0.410	0.410	0.410	0.410
Percentage effect	−27.80	−26.83	−28.05	−27.07	−24.39
Treatment effect bounds					
$R_{\max} = 0.92$					$(-0.100, -0.073)$, $\delta = 1.50$
$R_{\max} = 1$					$(-0.100, -0.035)$, $\delta = 1.10$
R^2	0.013	0.043	0.069	0.303	0.704
Observations	3,579	3,579	3,579	3,579	3,579
Panel B. Continuing members of occupation who migrated between counties					
State-licensed occupation	−0.118	−0.116	−0.116	−0.115	−0.098
	(0.029)	(0.029)	(0.030)	(0.034)	(0.070)
Dependent variable mean	0.401	0.401	0.401	0.401	0.401
Percentage effect	−29.43	−28.93	−28.93	−28.68	−24.44
Treatment effect bounds					
$R_{\max} = 0.95$					$(-0.098, -0.051)$, $\delta = 1.26$
$R_{\max} = 1$					$(-0.098, -0.010)$, $\delta = 1.03$
R^2	0.014	0.043	0.071	0.333	0.730
Observations	3,084	3,084	3,084	3,084	3,084
Controls					
Education, age, sex		X	X	X	X
Other controls			X	X	X
State × year fixed effects				X	X
Age interacted with fixed effects					X

Notes: Sample includes all individuals aged 18–64 residing in the 50 US states and DC not residing in group quarters with nonimputed values for migration status, education, occupation, age, sex, race, marital status, and employment status, excluding those who lived outside the 50 US states and DC in the previous year and noncivilians (i.e., those outside of the universe of the CPS labor force status question). Continuing members of occupation defined as individuals reporting same CPS occupation code as last year and current occupation. Sample only contains members of occupations listed in Table 1. Other control variables include income, race, marital status, employment status, citizenship status, and number of children. Estimated using OLS and sample weights. Treatment effect bounds calculated assuming $\delta = 1$ and stated $R_{\max}$. Deltas calculated assuming treatment effect of 0 and stated $R_{\max}$. Standard errors clustered on last year's state of residence in parentheses.

where $\mathbf{1}[occupation = j]_{isrt}$ indicates individual i's occupation. All other variables are defined as in equation (4). We use two different treatment/comparison group combinations. The first includes indicators for each of the 16 state-specific licensed occupations listed in Table 1, with the 6 quasi-national licensed occupations as the comparison group. The second includes dummies for 21 of the 22 occupations, with one quasi-national licensed occupation as the excluded comparison group.

Results for the first specification are shown in Figure 2.[25] The 16 state-specific licensed occupations are sorted in ascending order by the size of their associated coefficient β_j (converted to percent difference) in the model using all individuals (equivalent to that shown in column 1 of Table 3). In this specification, which does not condition on move distance, 12 of the 16 occupations move at a significantly

[25] Results are shown in tabular form in online Appendix Table B12.

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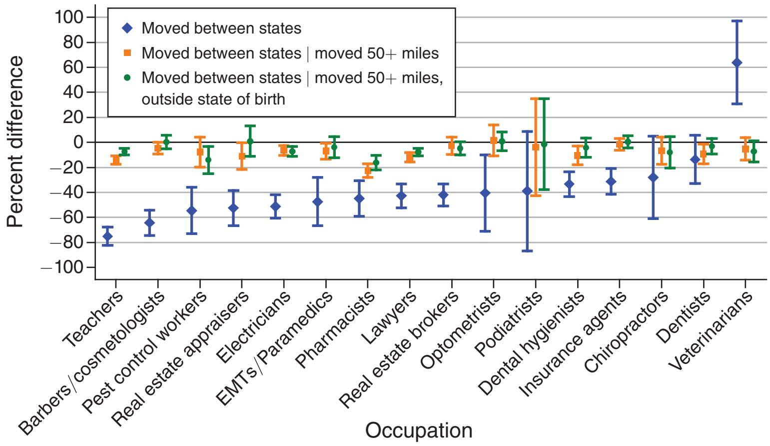

Figure 2. Interstate Migration and Occupational Licensing, Occupation-Specific Results, State-specific Licensed Occupations, 2005–2017 ACS

Notes: Sample described in notes to Table 2. Percent difference in interstate migration likelihood relative to quasi-national licensed occupations and 95 percent confidence intervals shown. Excluded group is quasi-national licensed occupations. Percentage effects calculated as coefficient/dependent variable mean × 100. All specifications include last year's state of residence × year fixed effects, state of birth fixed effects, and controls for income, race, sex, education, marital status, age, employment status, citizenship status, and number of children. Estimated using OLS and sample weights. Standard errors clustered on last year's state of residence. Results shown in tabular form in online Appendix Table B12.

lower rate between states relative to members of quasi-national licensed occupations. However, after conditioning on a 50-mile move, the coefficients only remain statistically significant for 6 of the occupations (teachers, electricians, pharmacists, lawyers, dental hygienists, and dentists). Further limiting the sample to 50-mile movers residing outside their state of birth in the previous year changes the results slightly, and five occupations show significantly lower interstate migration rates (teachers, pest control workers, electricians, pharmacists, and lawyers). For these five occupations, interstate migration rates are between −25 percent (pharmacists) and −7 percent (electricians) lower relative to members of quasi-national licensed occupations. Although not all are statistically significant, point estimates for 12 of the 16 occupations are less than zero. Among the occupations with the largest difference in overall interstate migration, pharmacists have the smallest change in the estimated migration rate difference when limiting the sample to 50-mile migrants, from approximately −44 to −22 percent. Local capital is unlikely to be as important for the career success of pharmacists, who largely work for large firms, as it is for other occupations such as lawyers and teachers (Goldin and Katz 2016).

Figure 3 shows the results using all occupations, with occupational and physical therapists as the excluded group.[26] Quasi-national licensed occupations are

[26]We use this quasi-national occupation as the excluded group as it is the largest such occupation with a statistically insignificant difference in interstate migration rates relative to all other occupations in a regression using 50-mile movers and including all controls (coefficient = −0.034, SE = 0.018). Results using nurses (the largest quasi-national occupation, coefficient = −0.028, SE = 0.005) and physician assistants (the smallest

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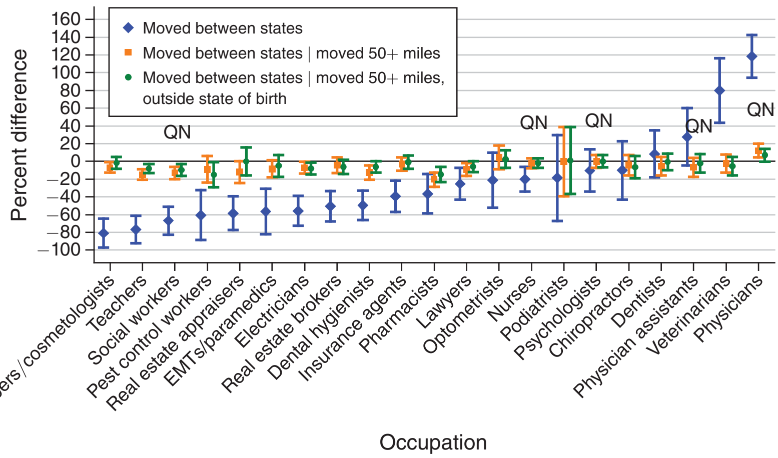

Figure 3. Interstate Migration and Occupational Licensing, Occupation-Specific Results, All Licensed Occupations Relative to Occupational and Physical Therapists, 2005–2017 ACS

Notes: Sample described in notes to Table 2. Percent difference in interstate migration likelihood relative to quasi-national licensed occupations and 95 percent confidence intervals shown. Excluded group is occupational and physical therapists. QN denotes quasi-national licensed occupation. Percentage effects calculated as coefficient/dependent variable mean × 100. All specifications include last year's state of residence × year fixed effects, state of birth fixed effects, and controls for income, race, sex, education, marital status, age, employment status, citizenship status, and number of children. Estimated using OLS and sample weights. Standard errors clustered on last year's state of residence. Results shown in tabular form in online Appendix Table B13.

indicated with a "QN" symbol. Six occupations show significantly lower interstate migration rates relative to occupational and physical therapists in at least one of the two specifications using only 50-mile movers: teachers, social workers, electricians, pharmacists, dental hygienists, and lawyers. One quasi-national licensed occupation moves between states at a lower rate relative to occupational and physical therapists: social workers. Social workers do have a national exam but no system of reciprocity between states, such as the nurse licensure compact, and many states require additional state-specific training courses for licensure.[27]

In sum, the limiting effect of state-specific licensure on interstate migration varies across occupations, and not all such occupations experience such an effect. Pharmacists and pest control workers are the worst affected, with 50-mile movers residing outside their state of birth 16 and 14 percent less likely to move between states relative to quasi-national licensed occupations, respectively. The other occupations with significantly lower relative interstate migration rates (including "quasi-national licensed" social workers) move between states at a rate approximately 7 percent lower.[28]

quasi-national occupation, coefficient = −0.065, SE = 0.027) are shown in online Appendix Figures B1 and B2, respectively, and are similar. Results are shown in tabular form in online Appendix Table B13.

[27] See Social Work License Map, "Frequently Asked Questions," https://socialworklicensemap.com.

[28] We provide our preferred occupation-specific estimates, which use only 50-mile migrants residing outside their state of birth in the previous year, for both specifications in online Appendix Figures B3 and B4. Both re-sort the occupations by effect size and rescale the *y*-axis for ease of interpretation.

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V. Conclusion

This paper shows that the interstate mobility of occupations that face high relicensure costs is reduced relative to occupations that face lower migration costs. Our preferred estimates, using only individuals who move 50 or more miles and reside outside their state of birth, show that individuals in state-specific occupations move between states at a 7 percent lower rate compared to members of quasi-national licensed occupations. This difference is much closer to zero than that which uses all individuals regardless of migration status: −58 percent. Our estimate constitutes a causal effect of licensing on interstate migration, assuming that (i) using only individuals who move 50 or more miles and reside outside their state of birth removes bias from differences in risk aversion across occupations, (ii) all moves of 50 or more miles result in changing labor markets, (iii) the cost of the local capital loss resulting from a between-labor-market move is equivalent between the licensed occupation groups, and (iv) conditioning on our included controls removes all other sources of bias correlated with entering a state-specific licensed occupation (versus the comparison occupations) and moving between states. Using methods developed by Altonji, Elder, and Taber (2005) and Oster (2017), we show that any remaining negative bias caused by unobservable differences must be quite large to explain away our results. The estimates are also robust to different standard error specifications, changing the definition of a long-distance move, and using another independent dataset (the CPS ASEC). Occupation-specific results show heterogeneity in effect size across occupations.

Based on our results, the increase in occupational licensing likely does not explain a substantial amount of the decrease in overall interstate migration rates between 1980 and 2015. To illustrate, assume unlicensed individuals migrated between states at the population average rate of 3 percent per year in 1980, and licensed individuals moved between states at a 7 percent lower rate (i.e., the same relative rate as state-specific licensed occupations), or 2.79 percent per year. Assuming no change in these underlying rates over time, an increase in licensing from 15 to 25 percent of the population (the increase experienced between 1980 and 2015) would result in a decrease in overall population interstate migration rates from 2.97 percent in 1980 (the weighted average of the licensed and unlicensed migration rates) to 2.95 percent in 2015—a decrease of 0.02 percentage points. Work by Molloy et al. (2014) shows that the unexplained decline in interstate migration over this period is 0.8 percentage points (after controlling for observable characteristics). Based on this figure, our estimates imply the increase in occupational licensing explains about 2.5 percent of the total decrease in interstate migration.[29]

While the increase in occupational licensing does not explain the broader trend in interstate migration, our results show that for individuals in affected occupations, the limiting effect of licensing costs on interstate migration can be substantial, up to 20 percent for some occupations. These findings are relevant to policymakers

[29] Using the unadjusted estimate of 46 percent lower migration for licensed occupations increases this amount to 0.14 percentage points, or 17.5 percent of the total 0.8 percentage point decrease in interstate migration rates between 1980 and 2015.

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currently considering revising licensure requirements to reduce the cost of relicensure for individuals licensed in other states. For example, as of April 2019, Arizona recognizes occupational licenses from all other states for nearly all licensed occupations (Office of the Governor of Arizona 2019).

Economists have long held that restrictions on geographic mobility limit the ability of the labor market to operate efficiently. Within this context, occupational licensing provisions that restrict job entry through interstate migration could also be a barrier to economic opportunity and labor market efficiency for affected occupations. Our results may also have legal implications. For example, in 1941, the US Supreme Court ruled against a California statute, making it illegal to restrict indigent individuals from migrating to the state during the Great Depression.[30] The court ruled that the California statute "prevent[ed] a citizen because he was poor from seeking new horizons in other States" (Roback 1943). In this way, limits on occupational entry might withhold the ability to migrate across states from large segments of the population.

Our empirical strategy estimates the causal effect of occupational licensing on interstate migration subject to stronger assumptions than those required by a traditional causal model based on exogenous variation in relicensure requirements. We hope that additional information on such requirements will become available in the future for a more traditional causal analysis of the effect of occupational licensing on geographic mobility.[31]

Our analysis examines the migration of individuals and may therefore miss an additional important effect of relicensure costs incurred as a result of interstate migration. For many, migration is not an individual decision; instead, it is a choice made based on overall household or family well-being. As our analysis is limited to the individuals we observe in an occupation after their move, we do not capture the effect on individuals who are forced out of an occupation or out of the labor force entirely as a result of moving between states. One example is so-called "trailing spouses"—those who move because their partner obtains a better job in another state. If these spouses were in a licensed occupation prior to the move, they may have had to switch careers as a result.[32] The effect of licensure on career changes or labor force exits that were made as a result of household migration is potentially important, and because we cannot identify individuals affected by these phenomena in the ACS, we also leave this analysis for future research.

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[30] Edwards v. People of State of California, 314 US 160 (1941).

[31] In online Appendix C, we investigate the effect of initial adoption of a reciprocity agreement on the interstate migration of lawyers using an event study framework. While there is some evidence that the in-migration of lawyers increases in the year after adoption, the overall results are too noisy to be conclusive.

[32] North Dakota enacted a bill in 2019 that establishes occupational license reciprocity for military spouses for the more than 11,000 military personnel in the state (Associated Press 2019).

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