NBER WORKING PAPER SERIES

ANALYZING THE INFLUENCE OF OCCUPATIONAL LICENSING DURATION ON LABOR MARKET OUTCOMES

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Working Paper 22810
http://www.nber.org/papers/w22810

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
November 2016

We greatly appreciate the comments of Joan Gieseke, Hwikwon Ham, and Mindy Marks on a previous version of the paper. We especially thank participants in seminars at the American Economic Association annual meetings, Collegio Carlo Alberto, Labor and Employment Relations Association annual meetings, London School of Economics, Society of Labor Economists annual meetings, University of Illinois, Urbana-Champaign, University of Minnesota, Twin-Cities, and the W.E. Upjohn Institute for Employment Research for comments. We thank the Smith Richardson Foundation and the Kauffman Foundation for their financial support of our research, but the views expressed here are those of the authors and do not necessarily reflect the views or policies of the Smith Richardson Foundation or the Kauffman Foundation. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 22810
November 2016
JEL No. J08,J3,J38,J44,J8,J88,K0,K2,L12,L38,L51,L84,L88

ABSTRACT

We analyze the labor market influence of the duration of occupational licensing statutes for 12 major universally licensed occupations over a 73 year period. These occupations comprise the vast majority of workers in these regulated occupations in the United States. Time from the start of state occupational licensing statutes (i.e., licensing duration) may matter in influencing labor market outcomes. Adding to or raising the entry barriers is likely easier once an occupation is established and has gained influence in a political jurisdiction. States often enact grandfather clauses and ratchet up requirements that protect existing workers and increase entry costs to new entrants. We provide among the first estimates of potential economic rents to grandfathering. We find that duration years of occupational licensing are positively associated with wages for continuing and grandfathered workers. The estimates show a positive relationship of duration with hours worked, but we find moderately negative results for participation in the labor market. The universally licensed occupations, however, exhibit heterogeneity in outcomes. Consequently, unlike some other labor market public policies, such as minimum wages or direct unemployment insurance benefits, occupational licensing would likely influence labor market outcomes when measured over a longer period of time.

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Introduction

Occupational licensure is the legal process by which governments (mostly the U.S. states but also local governments and the federal government) identify the legal qualifications that are required to practice a trade or profession, after which time only licensed practitioners are allowed by law to receive pay for doing the work in the occupation. This form of labor market regulation has rapidly become one of the most significant factors affecting labor markets in the United States and other industrialized countries (Kleiner, 2015). Over the past several decades, the share of U.S. workers holding an occupational license has grown sharply. For example, during the 2012–2013 state legislative sessions, at least seven new occupations were licensed in at least one state—occupations ranging from scrap metal recyclers in Louisiana to body artists in the District of Columbia.[1] U.S. government estimates suggest that over 1,100 occupations are regulated to some extent in at least one state, but fewer than 60 are regulated in all 50 states, showing substantial differences in which occupations states choose to regulate (Department of the Treasury Office of Economic Policy, Council of Economic Advisers, and Department of Labor, 2015).

The time from the passage of occupational licensing laws may be important in analyzing regulation's influence on the labor market. One rationale is that states often enact grandfather clauses that protect existing workers by allowing them to practice either when licensing laws are passed or after the enactment of new regulations, even though they may not meet the current requirements. We provide among the first estimates of the labor market returns to grandfathering. On the other hand, new entrants must have higher entry standards than the existing members of

[1] These data are from a LexisNexis search of statutes passed during the legislative session.

the occupation.[2] We therefore anticipate that individuals who are grandfathered would have incentives to encourage licensing and their continued participation in the occupation at pre-licensing levels of education and training. They would likely obtain economic rents by limiting supply and increasing the demand for the higher quality service. In the labor market, the process of older, lesser trained workers leaving the workforce or moving to other occupations and newer workers with higher entry requirements entering the field takes many years or even decades as the process works its way through the labor market, resulting in potentially higher wages. We examine duration over a 73 year period to determine the influence of occupational licensing on key labor market outcomes.

In examining the influence of occupational licensing duration on the labor market, we initially review the literature of other studies of duration effects on labor market outcomes and show that our study is the first comprehensive examination of the issue using more than one occupation and it implements a longer period of time. More important, we also present evidence that goes beyond analyzing wage determination to examine hours worked and participation in the regulated occupation for large numbers of workers. Consistent with other findings, we show that occupational licensing raises wages in the regulated occupations and that the duration of state licenses is also associated with higher wages. We find this to be the case across a number of robustness tests, and it is especially the case for grandfathered workers. In addition, the estimates show that the duration of state licensing is associated with an increase in yearly hours worked by those in the occupation by almost 4 percent, but that participation in the occupation in the labor market declines slightly over the first 10 years after the licensing laws are implemented.

[2] A model of grandfathering presented by Shavell (2007) assumes that if the best standard in period 1 exceeds the level of risk that would be appropriate for the expected harm, grandfathering may be desirable. If in period 2 the known harm is below a threshold, grandfathering is optimal—parties who engaged in the activity in period 1 can maintain their period 1 level of risk in period 2, but parties who enter the activity in period 2 should take the new conventionally optimal precaution for the known harm, and they have certainty of the outcome in the second period.

However, the labor market outcomes for the occupations we examine exhibit heterogeneity. We implement several sensitivity tests to examine the robustness of our estimates for labor market outcomes. These results are largely consistent with a monopoly model of regulation that shows gains to those in the regulated occupation through higher wages and more hours worked, but which may restrict entrants into the occupation in the long run.

Reviewing Duration in the Labor Market for Licensed Occupations

The duration of occupational statutes has been identified in previous studies as a factor that may raise wages (Law and Marks, 2009, Timmons and Thorton, 2013). In both studies, the authors examined one occupation and focused on wage determination. Our study expands on these studies by examining 12 universally licensed occupations (i.e., licensed in all states), some of which have been regulated in all states for over 100 years and others that just became universally licensed during the past decade. The number of workers in these occupations represents about 60 percent of all universally licensed workers in the United States in 2013 from our estimates using the American Community Survey. These occupations were chosen because the date of initial licensure was available, there were sufficient observations in the census for statistical analysis, and that the vast majority of workers must obtain a license in order to work (Gittleman and Kleiner, 2016). Also, the states that licensed these occupations regulated them at different times, allowing for a difference in difference estimation strategy.

The Growth and Wage Effects of Occupational Licensing

Occupational licensing has grown to be one of the largest institutions in the U.S. labor market (Kleiner and Krueger, 2013). To illustrate, funeral attendants are licensed in nine states and florists in only one state. Estimates from national surveys find that the wages of unlicensed workers are 8 to 15 percent lower than those of licensed workers with similar levels of education, training, and experience (Kleiner, 2006, Kleiner and Krueger, 2013, Gittleman, Klee, and Kleiner, 2015). More specifically, Kleiner and Krueger (2013) find that licensing at the state level confers a wage premium of around 17 percent, and the combination of state and either federal or local licensing has an estimate effect of around 25 percent. Local licenses by themselves are not associated with higher wages, and certification has a smaller effect on wages using data from the Survey of Income and Program Participation (Gittleman, Klee, and Kleiner, 2015).

The use of other data and methods finds that the wage premium from licensing is more modest and is sometimes estimated as zero. For example, Gittleman, Kleiner, and Klee (2016) find that the wages of licensed workers are around 7.8 to 11 percent higher on average, controlling for detailed occupation, and that licensing also confers better employment opportunities and health and pension benefits. Unlike the minimum wage or unemployment insurance which requires all employers that are covered by the law to pay the new wage or transfer payment immediately, occupational licensing allows individuals who are working in the occupation, but do not meet the current licensing requirements, to continue working. This practice is called "grandfathering." In addition, the regulated occupation generally has the ability to ratchet up the requirements—that is, raise the requirements for initial entry or movement into the occupation from other political jurisdictions with minimal constraints from policy makers (Wheelan, 1999). Again, individuals who do not meet the current requirements are allowed to keep working with permission from the government. In our analysis, we examine how time from initial licensure, which we call duration, influences key labor market outcomes such as wages, hours worked, and participation in the workforce.

The Role of Different Institutions on Wage Determination and Labor Market Outcomes

A helpful analogy of the influence of institutions in the labor market can be drawn from unions. When unions first organize a firm or establishment, the wage increases are generally small (Freeman and Kleiner, 1990, DiNardo and Lee, 2004, Lee and Mas, 2012). However, cross-sectional estimates of the impact of unions are between 15 to 20 percent (Hirsch and Macpherson, 2013). The additional cost of having a union worker is approximately $40,500 over the course of that worker's employment with the firm (Lee and Mas, 2012). Moreover, unions appear to raise the wages and benefits with a statistically significant effect the longer they are in an establishment (Freeman and Kleiner, 1990). We examine whether these wage outcomes may also be the case for occupational licensing.

Unions may raise wages through collective bargaining and withholding their labor services through concerted activities to gain wages and benefits. On the other hand, occupational licensing could raise wages by choosing the right set of regulations to restrict supply and limit the tasks of unlicensed workers, and thus enhance demand by signaling and education that they are providing a higher quality service (Friedman, 1962, Spence, 1973). In a manner similar to unions, the institutional mechanism and design that occupational licensing uses also takes time to implement and the full effects may only reach fruition over several decades of strengthening these rules (Hurwicz, 1973).

Background on Grandfathering and Ratcheting Requirements

Initially, the influence of licensing duration on labor market outcomes was identified in a National Bureau of Economic Research volume published in 1945 by Milton Friedman and Simon Kuznets (Friedman and Kuznets, 1945). They noted that in 1911, the American Medical Association, through the implementation of the Flexner Report, ratcheted up requirements for becoming a doctor through tougher admissions requirements, length of education in medical school, and limits on the number of new openings for medical education (Beck, 2004). While increasing the requirements for graduation from medical school and pushing for tougher licensing, the Flexner Report did not require currently working doctors to meet the same higher requirements; this was a classic case of grandfathering (Beck, 2004). Friedman and Kuznets went on to examine the influence of the regulations more than 20 years later in the late 1930s, and they found that doctors were able to raise their wages by greater than 17 percent more than dentists, who did not substantially change their requirements. This example illustrates how an occupation can raise wages that involved rents to those who were in the occupation and how entry requirements for an occupation were raised for just new entrants.

More recent estimates of the influence of the length of licensing statutes on wage determination include results for massage therapists, nurses, lawyers, and barbers (Law and Marks, 2009, Pagliero, 2010, Timmons and Thornton, 2010, Timmons and Thornton, 2013). The main results suggest that for specific occupations such as massage therapists and barbers, the length of time that a licensing statute has been in place enhances the earnings of these practitioners, but little evidence of the influence of duration was found for nurses (Law and Marks, 2013). However, the estimates are limited to these occupations over a relatively short time period. Our estimates expand upon and provide evidence beyond simply the wage determination of the effects of licensing duration on labor market outcomes.

Although not explicitly addressed, the process occurs by allowing current practitioners to avoid the explicit general and specific education requirements, internships, tests, continuing education mandates, and good moral character investigations, assuming that they were in good standing prior to the new licensing laws. To the extent that these requirements raise marginal productivity, they may also raise wages. Also, it takes many years for the individuals who did not meet these requirements to leave the occupation or retire, and as a result, the educational quality of the new entrants is higher, and they dominate the current members of the occupation only after many years. Moreover, the longer the occupation is licensed, the greater the ability of the members of the occupation to lobby the legislature and licensing boards to ratchet up requirements for entry within the occupation for those who might enter from unregulated states or occupations. For example, accountants increased the years of university schooling from four to five years in the 1990s in order to attain a Certified Public Accountant (CPA) license (Carpenter and Stephenson, 2006). In addition, physical therapists raised their education requirements from a bachelor's degree in the 1990s to a doctor of physical therapy license by 2016 through 2018 (Cai and Kleiner, 2016). In both cases, the national professional association promoted these enhanced or ratcheted-up requirements through the state boards of licensing or the state legislature. Although the policies may have enhanced the educational quality of the new workers, they could have also reduced access to the occupation by practitioners and consumers and limited the supply of labor to the occupation.

A Licensing Model with Duration

The model uses a framework in which the work of one occupation or individual cannot legally be done without the inputs of the other occupation. The focus of the model serves as a basis to inform and develop hypotheses about the empirical work, rather than as a fully specified general equilibrium model of production of services under regulation. The model uses a modified standard production function:

$$Q_{\mathrm{pt}} = HH = f(P(z),K)_{\mathrm{t}} \qquad (1)$$

$$Q_{\mathrm{nt}} = HL = f(P(z), N(z),K)_{\mathrm{t}}, \qquad (2)$$

where Q_{pt} is the output produced by the licensed practitioner over time t, which we will refer to as "high-skilled services (*HH*)." Q_{nt} is the output produced by the unlicensed worker (NP) over time t, which we will refer to as "low-skilled services (*HL*)." $P(z)$ represents the licensed worker's labor, recognizing that output relies on their decision of personal input, and $N(z)$ represents the NP's labor, recognizing that output relies on their decision of personal input. The variable K represents the quantities of capital inputs used in a service production function (Reinhardt, 1972).

By law, however, the technology needed for NPs to produce *HL* is tied to supervision of entry by the licensed practitioner. Nevertheless, within a profit function, the NP's wage is tied to the decisions of the licensed practitioner to add the labor input and technology mix to the high-skilled provider, *HH*. Regulation acts like a shifter of both the supply and demand curves with long time lags for full implementation over time t because of grandfathering and the ratcheting up of skills for regulated practitioners. In the model, time is the proxy for grandfathering licensed workers and capturing the work of unregulated or lesser regulated practitioners. For example, a licensed engineer or architect can restrict the work of an unlicensed interior designer, reducing earnings, hours worked, or the number of workers who may choose that occupation over time (Kleiner, 2013). Regulated practitioners, who are generally in control of the production of these services, can allocate relatively low-skilled work to unlicensed workers while taking on higher-skilled and value-added services for themselves and thus increasing their hours worked and earnings, but still restricting the number of workers who may enter the occupation. These central theoretical issues raised in these models are empirical questions that the rest of the paper examines.

The Rationale for Grandfathering and Ratcheting

In the labor market, the process of older, lesser trained workers leaving the workforce or moving to other occupations and newer workers with higher entry requirements entering the field takes many years or decades as the process works its way through the labor market, resulting in potentially higher wages. An illustration of the process over three periods is shown in Figure 1. The figure shows the evolution of grandfathered participants over time and how they diminish by leaving the occupation, through retirement or death. By the end of the period, only individuals who have gone through the licensing process are in the occupation. However, the process may limit the supply of labor in the long run by increasing entry and mobility requirements, and may allow those licensed in the occupation to gain economic benefits by limiting employment growth. In addition, occupations could also ratchet up the requirements for already licensed occupations. Therefore, licensing duration—the time from the implementation of occupational licensing legislation—may matter. It may take years for the full effects of occupational licensing to be realized in the labor market, and for the analyst to observe these changes on wages, hours, and employment. A similar effect of regulation would occur when the occupation ratchets up the requirements for entry, such as the increases in education that occurred in accounting and physical therapy. The licensing requirements went from four to five years in accounting and the addition of an advanced degree for licensure in physical therapy.

A further implication of the role of time for occupational licensing is that it captures the work of unregulated workers and tasks as exemplified in the *North Carolina State Board of Dental Examiners v. Federal Trade Commission* Supreme Court case (2015). Moreover, legal cases involving the Institute for Justice challenged cosmetologists capturing the work of hair braiding for their occupation. In addition, veterinarians have tried to legally capture the work of farmhands who do teeth filing for horses, suggesting that only trained veterinarians can do these tasks for farm animals. In all of these cases, the number of individuals in the regulated occupations would grow as unlicensed workers declined and as the tasks were legally mandated by regulated workers, as presented in our theory overview.

The Empirical Model

We gathered statutory information for each occupation by year for each state that passed a licensure law from several different legal data sources. In order to calculate the duration of licensure for all states, we used a couple of different resources. Our major source of data used a Council of State Governments (1952) report to obtain information by year for each state listing their first licensing legislation for the major universally licensed occupations in our analysis. From this source alone, we were able to obtain full information for almost 60 percent of the major universally licensed occupations in the United States. We also used the LexisNexis legal resource database to obtain the remaining statutory information.[3]

In order to develop a model with a sufficient time line to analyze how duration may influence labor market outcomes, we use all available data from the census and the American Community Survey (ACS) for a 73-year time period from 1940 to 2013. We begin with 1940 since that was the first year wage data was added to the Census. We include in our sample individuals who worked in 12 major universally licensed occupations that had more than 174 million workers and which represented more than 10 percent of the U.S. workforce and about 60 percent of all universally licensed occupations. The sample includes both blue- and white-collar occupations and ones that are high, middle, and low income. One set of controls are individuals

[3] For the additional remaining information on attorneys for 19 states, we contacted the Supreme Court library and Board of Examiners. We managed to obtain responses for 7 states: Arkansas, Delaware, Illinois, Indiana, Maryland, Michigan, and Minnesota. We replaced the average duration with missing values on attorneys for 12 states: Connecticut, Georgia, Missouri, Montana, Nebraska, Nevada, New Hampshire, New Jersey, New Mexico, South Carolina, Utah, and West Virginia, and we denoted with dummy variables in our statistical analysis for completeness (Little and Rubin, 1987).

who worked in occupations that were unlicensed during the period of analysis and those in service occupations. We limit the analysis to those occupations that have sufficient number of state and year observations in the census and were licensed in all states by 2013, the end of our period of analysis.

An illustration of the licensed occupations that are in the sample and the time line are presented in Figure 2 for the period 1800 to 2013 (Meyer and Osborne, 2005). The sample includes individuals who either were in one of the major universally licensed occupations when they became regulated or were in an unlicensed occupation for the period. Our analysis is limited because we can only include individuals who are covered by licensing statutes, but some may not have attained a license (Gittleman and Kleiner, 2016). Also, we cannot cover the same individuals over their careers as in smaller data sets such as the National Longitudinal Survey of Youth (NLSY). However, we define individuals who worked in major licensed occupations as a treatment group after they were required to have a license, and individuals who worked in all other unlicensed occupations or were unlicensed prior to their state passing a licensure statute as a comparison or control group.

Next, we include standard human capital variables from the census and for more recent years from the ACS, such as gender, age, education, and potential experience. In order to generate a reliable sample for our analysis, we dropped those individuals whose education is "below 12th grade without a diploma" for dentists, lawyers, accountants, and pharmacists. Also, we dropped those individual whose education level was "below high school diploma" for nurses. For barbers, we screened for those with at most a high school diploma. In addition, individuals older than 65 or younger than 16 and whose years of potential experience are estimated to be below zero were also deleted.[4]

Finally, hourly real earnings were determined by dividing the annual earnings including profits and dividends from work, by annual hours worked, adjusted by the 2014 consumer price index (CPI). Annual hours worked were calculated by multiplying the usual working hours by the number of weeks for the past 12 months. We eliminated from the sample those with more than 60 hours of work in a week as a response error or coding mistake. In addition to these restrictions, the original sample was trimmed by excluding individuals with wages below the federal minimum wage level in that year. The resulting sample consists of over 1 billion observations from 1940 to 2013 using the census and the ACS sample.[5] In Table 1 we show the means and standard deviations of the licensed and unlicensed workers in our sample with wage data adjusted by the 2014 CPI to standardize our results.

Implementing Descriptive and Causal Estimates

The following sections present both descriptive and causal estimates of our empirical models, initially by using a kernel estimation approach for a descriptive approach and then by using a difference-in-difference causal model that takes into account the different times that each of the occupations in our model initially became licensed in each of the 50 U.S. states over the time period of the analysis.

Nonparametric Kernel Estimation

[4]We also used individuals with graduate school education for dentists, lawyers, and physicians, and the results were similar. These estimates are available from the authors.
[5] We show the number of observations by year in Appendix Table 1.

In order to provide basic descriptive data for our analysis for more recent and older licensed occupations, we used a nonparametric kernel estimation procedure. The estimates for the effects of occupational licensing duration on wage determination use the kernel estimation procedure, and they are shown in Figure 3.[6] The estimates are a form of data smoothing. We try to visually inspect whether the longer duration of licensing exhibits wider variance in the earnings distribution. The black line represents the wage distribution for unlicensed workers, the pink line is the distribution for licensed workers with a licensing duration of less than 10 years, and the blue line is the distribution for licensed workers with licensing durations of more than 10 years. Our nonparametric kernel estimation results suggest that it takes at least 10 years to fully realize the economic effects of occupational licensing on wage determination.

Empirically Modeling Duration Effects

In order to more fully empirically model the influence of occupational licensing on wage determination, hours worked, and participation in the labor market, we use a basic difference-in-difference approach. Since each of the states implemented their licensing statute at different times, we are able to develop an estimate of causal inference for the influence of duration on labor market outcomes. We would expect the relationship to initially move slowly as both new more skilled workers enter, and as fewer less skilled grandfathered workers continue in the occupation. When grandfathered workers retire or leave the occupation, wages would then increase more rapidly. Further, when workers' representatives are more fully in control of the supply of labor by ratcheting up requirements, this would also result in wages increasing.

[6] The kernel estimation procedure develops and uses an autoregressive approach to predict observed outcomes and is a theoretical method to provide basic forecasts of observed phenomena.

To causally link occupational licensing and labor market outcomes, we employ a difference-in-difference (DID) strategy using data on changes in state licensing requirements for the 12 occupations in our sample. Such changes affect the ability of someone to work in a licensed occupation in a particular state without needing to fulfill additional licensing requirements. For estimation purposes, our model shown in equation (3) would take the following general form:

$$Y_{istk} = \beta_0 + \boldsymbol{\delta} Duration_{its} + X_{it}\beta + \eta_s + \alpha_t + \theta_k + \varepsilon_{it} \qquad (3)$$

where Y_{istk} is an indicator of a labor market outcome such as earnings or hours worked, $Duration_{its}$ is duration of an initial occupational licensing statute, and δ is the DID estimate of the effect of the change on the DID strategy using data on changes in state licensing requirements for the 12 occupations in our sample. The variable X_{it} is individual characteristics (education level, male, race, potential experience), α_t includes year fixed effects, η_s includes state fixed effects, and θ_k includes the size of the occupation in the industry.

$$Y_{jst} = \beta_0 + \boldsymbol{\delta} Duration_{jts} + Income_{st} + \mu_s + \alpha_t + \varepsilon_{it} \quad (4)$$

In equation (4) we show our model of licensed worker participation in the workforce. The variable Y_{jst} is the number of licensed workers in occupation j per 10,000 population, and $Income_{st}$ is per capita mean income in state s in year t. The other variables have the same definitions as in equation (3). Such changes could affect the ability of someone to be able to work in a licensed occupation in that state. We use the DID model by exploiting changes in state licensing laws and requirements over time in each of the tables presented in the rest of the paper. Our sources of identification are the changers in states that adopted occupational licensing laws over time relative to the non-adopters, individuals who were licensed in the same occupation in comparison to those who did not achieve licensure, and any individual who was licensed relative to those who were not licensed. In order to focus only on changers during the period of analysis, we develop separate estimates for occupations that were licensed during the period 1940–2013. In order to focus on changers in licensing, we also examine by discrete time periods the influence of becoming a licensed occupation on the participation rate in the occupation (Law and Marks, 2013). However, we also present estimates of all 12 occupations in our sample, many of whom were initially regulated prior to 1940. Moreover, since we do not assume a linear relationship between licensing adoption and its labor market effects, consequently we also present nonlinear estimates in our tables.

In Table 2 we show the influence of duration of the passage of a licensing statute and licensing on earnings using clustered standard errors. We show both a linear and quadratic specification in the table. In addition, we show in panel A estimates using all the occupations in our sample. In panel B, we show estimates for only those occupations that changed their licensing status over the period of our analysis. In the first column, we show the influence of the duration of the passage of a licensing law on wage determination with no occupation controls as a benchmark for our other specifications. The estimates suggest that for every 10 years that an occupation is licensed, wages increase by a statistically significant 4 percent. Moreover, in column (3), we see that becoming licensed raises earnings by almost 15 percent, which is at the midrange of estimates shown by Kleiner and Krueger (2010, 2013). We estimate our models using two-digit occupation controls, but we do not introduce more detailed occupational controls because they would result in identification taking place largely through individuals who were in the occupation but were not licensed, in comparison with those who were regulated and licensed. Since our objective is to examine the influence of the change in laws over various time periods, adding detailed occupation controls would not be an appropriate strategy for identification. The estimates across these specifications or groups of occupations were that older regulated occupations, as opposed to those who were regulated more recently, found relatively small differences across the specifications as a robustness test of our estimates. As a robustness test of our estimates across these specifications or groups of occupations, we found relatively small differences across the specifications between older regulated occupations and those who were regulated more recently. As a further test of the strength of our findings we also developed a generalized propensity score (GPS) based on race, sex, experience, education, and year and found that the results are similar to those presented in Table 2[7].

In Table 3 we present estimates of the influence of duration of occupational licensing on hours worked per year using clustered standard errors. Using an approach similar to the one shown in Table 2, we begin by estimating the influence of duration with no occupation controls and a simple linear relationship. We also show estimates of the two panels for all the occupations in our sample and the ones that experienced changes. We find that duration is associated with increases in work hours. Also, the estimates of increasing the hours worked due to becoming licensed is more than 76 hours per year for those who were licensed more recently. The resulting increase in hours worked per year could be due to the substitution effect of wage increases dominating the income effect for the occupations evaluated in our sample. Again, we show the influence of becoming licensed using the occupations that were regulated during the period of analysis and those who were licensed during earlier periods.

[7] We also estimate a propensity matching approach using the nearest neighbor method to implement further robustness checks of the estimates, and find similar results for wages and hours worked. In order to remove selection bias caused by endogenous selection into the occupations, we implement the generalize propensity score approach (Hirano and Imbens, 2004). We calculate the GPS based on race, sex, experience, education and year. Appendix Table 3 contains the estimation results for the dose-response function. Standard errors are bootstrap standard errors from 100 replications.

As an additional robustness check, we implemented a two-stage procedure that uses the state as the unit of observation rather than individual characteristics.[8] The model we implement is applied to occupation-specific log wages and is a two-way fixed effects version of the standard cross-sectional human capital wage equation. We estimate the earnings equations using two different approaches. In the first approach, we estimated the model using the full micro-level data set and estimated standard errors that are robust to heteroskedasticity and clustering at the state level. In the second approach, we aggregated the data to the level of state $\times$ year cells using the two-stage procedure described in Hanushek (1974), Amemiya (1978), and Conley and Taber (2011). In the first stage, we regressed individual-level outcomes on individual covariates and a full set of state $\times$ time fixed effects. The coefficients on the state $\times$ time fixed effects represent state $\times$ time cell means that have been purged of the variation associated with the within-cell variation in the covariates. In the second stage, the covariate-adjusted cell means are regressed on the policy variables, state fixed effects, and year fixed effects as described earlier. Standard errors are again constructed to allow for heteroskedasticity and clustering at the state level. The estimates using this procedure are shown in Table 4 for wages and hours worked. The estimates in Table 4 using the two-stage process are similar to those using individual observations for both wage determination and hours worked.

We also provide basic approximations of the potential rents that occupational licensing offer to individuals who are grandfathered to show the potential incentives for these individuals to promote this type of regulation. To develop these estimates, we use those individuals whose expected tenure in an occupation occurred during the period that the occupation initially became licensed. To illustrate, if an occupational therapist had 10 years of experience and licensing

[8] In Appendix Tables 4 and 5, we show basic aggregates of the state-level estimates.

occurred after she had been in the occupation for 5 years, then that person would be considered grandfathered. In this example, individuals whose tenure was less than 5 years would be considered to have entered the occupation after licensing, and would be a new entrant who started after the initial regulation of the occupation. The estimates used the propensity matched sample of those unlicensed individuals who were most like the licensed sample based on race, sex, experience, education, and year of the observation. The results in Panel A of Table 5 show that individuals, who are grandfathered, gained almost 1 percent per year in earnings, and their overall earnings are about 5 percent higher than their unlicensed control group.

In panel B of the table we show the earnings effect of grandfathering relative to new entrants. In this case new entrants make about 12 percent more than grandfathered workers. However, using an Oaxaca decomposition, human capital differences widens the wage gap to 15 % because new entrants require substantially higher human capital to obtain occupational licensure (Oaxaca, 1973). But we also find that the unexplained part, which is the potential rents to grandfathered workers, is able to explain about 3 percent in the wage gap in favor of grandfathered workers. Therefore, occupational licensing offers potential rents to individuals who are grandfathered into occupational licensing relative to unlicensed workers or new licensed entrants with similar observable covariates.

To the extent that an increase in hours worked could reflect a reduction in the number of practitioners, we next turn our attention to Table 6, which focuses on labor market participation. Perhaps one of the most speculated yet little researched areas of occupational licensing focuses on the role of the regulated institution on the labor supply of regulated practitioners (Law and Marks, 2013). In Table 6 we estimate the influence of the duration of an occupational license statute on the number of practitioners in the occupation per 10,000 in the population, again using different ways of categorizing the occupations in our sample from all, changers and non-changers. Unlike our previous analysis, we use the state as the major unit of observation, since the statutes are largely influencing state-level observations of labor market participation.[9] Our estimates show the influence of the passage of the licensing statute by 3-year intervals following the passage of the law. Throughout the first 10 years following the passage of the statute, the number and the ratio of individuals who worked in licensed occupation both decrease. The estimates show that 4 to 6 years following the passage of a licensing statute, the number of licensed workers per 10,000 population decreases by a statistically significant seven practitioners. Different results for each panel imply that the occupations that are most likely to be licensed were licensed in an earlier period, and ratcheting up the requirements results in fewer workers. The overall effect of licensing is to reduce the relative number of workers as a consequence of becoming licensed. These declines in the number of workers may be the reason for wages going up and the number of hours worked by practitioners increasing, as shown in Tables 2 and 3. Consistent with the theoretical section, occupational licensing statutes can provide the basis for regulated practitioners reallocating work toward members in their own occupation at the expense of other workers in the occupation. A recent illustration of the issue was the ***North Carolina State Board of Dental Examiners v. Federal Trade Commission*** (2015). The result of these practices could be the higher observed pay.

In order to provide a further robustness check on the estimates shown in Tables 2–4, in Figure 4 we show the slope of the hourly wage curve before and after the change to licensure. Wages in the regulated occupations are relatively flat before the introduction of licensing, but decrease slightly immediately following regulation, perhaps because there is a surge in the

[9] Appendix Tables 4 and 5 show the state-level analysis for log hourly earnings and total worked hours per year as a basis of sensitivity and robustness analysis.

number of people who want to enter the occupation now that licensing has been introduced and lower skilled grandfathered practitioners can continue to work, but the regulations have little bite initially. After the start of initial regulation, hourly wages increase as the duration of licensing grows, also perhaps due to workers who were grandfathered leaving the occupation and the ability of the lobbyists for the occupations to ratchet up the legal requirements for entry. The statistically significant different slopes of the wage lines before and after the change to licensure presented in Figure 4, along with the similarities between the 12 universally licensed and never licensed groups shown in Table 1[10], suggest that the DID approach of identification has validity for our empirical model.[11]

In Table 7 we present estimates for a wide variety of occupations, each of which may have experienced different economic and institutional environments on the road to becoming licensed. The table summarizes the influence of duration and licensing for each of the occupations in our sample separately. The estimates suggest that wages are generally similar but that dentists have much higher wages and lower hours worked as a consequence of occupational licensing. All the occupations have greater participation except for teachers whose participation declined significantly, and since they make up about half of the total number of workers in our sample, this resulted in a negative effect on workforce participation for the combined sample.

From a policy perspective, our estimates are consistent with the decision of the U.S. Supreme Court in holding that dentists were operating as monopolists in reallocating tasks to themselves when the North Carolina State Board of Dental Examiners outlawed the work of

[10] Appendix Table 2 reinforces the validity of our DID approach by showing the similarities between occupations that changed their regulation status (i.e. treatment group) and occupations that did not changed their regulation status (i.e. control group).

[11] The estimates in Figure 4 show only occupations that changed regulatory status over the period we examined from 1940 to 2013. We also charted all the individual occupations in our sample, and they showed similar plots; these results are available from the authors.

"teeth whiteners" (*North Carolina State Board of Dental Examiners v. Federal Trade Commission*, 2015). Our results suggest that occupational licensing works slowly over time as older, less skilled workers retire or move to other occupations and the state boards or legislatures that regulate the professions ratchet up the requirements for entry. Our ability to document these changes shows how important labor market institutions work with deliberate speed to enhance the work and pay arrangements for their members, in contrast to policies such as the minimum wage or changes in unemployment insurance policies, whose influence is more immediate (Kleiner, 2015).

Conclusions

Since the implementation of new occupational licensing statutes takes time to fully carry out, duration from the passage of a statute should matter in influencing labor market outcomes. For example, states often enact grandfather clauses that allow continuing practitioners to continue working without meeting the new requirements, or they ratchet up the requirements for entry, such as education and reciprocity agreements with other states or nations, that protect existing workers. One implication is that new entrants must have higher regulatory standards than those already in the occupation. The process of older, less educated workers leaving and newer workers with higher entry requirements entering the occupation takes time to work its way through the labor market. Our analysis uses a model in which licensed practitioners influence the number and kinds of jobs that they and unregulated workers can do over the long run. We use data for 12 large, diverse licensed occupations covering a 73-year period to examine the labor market effects of initial licensure. Our results that are consistent with theory of regulation for continuing workers and those that were grandfathered show that their wages rise relative to unlicensed individuals and those that were in the regulated workforce following licensure.

Moreover, there are incentives for incumbents in the occupation to raise standards because they can get higher wages. However, the relative number of workers in the occupation declines somewhat. Our study should allow policy analysts and policy makers to develop and implement more informed decisions on the long-run implications of the rapidly growing labor market institution of occupational licensing.

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Figure 1 Potential Influence of the Grandfathering of New Regulations

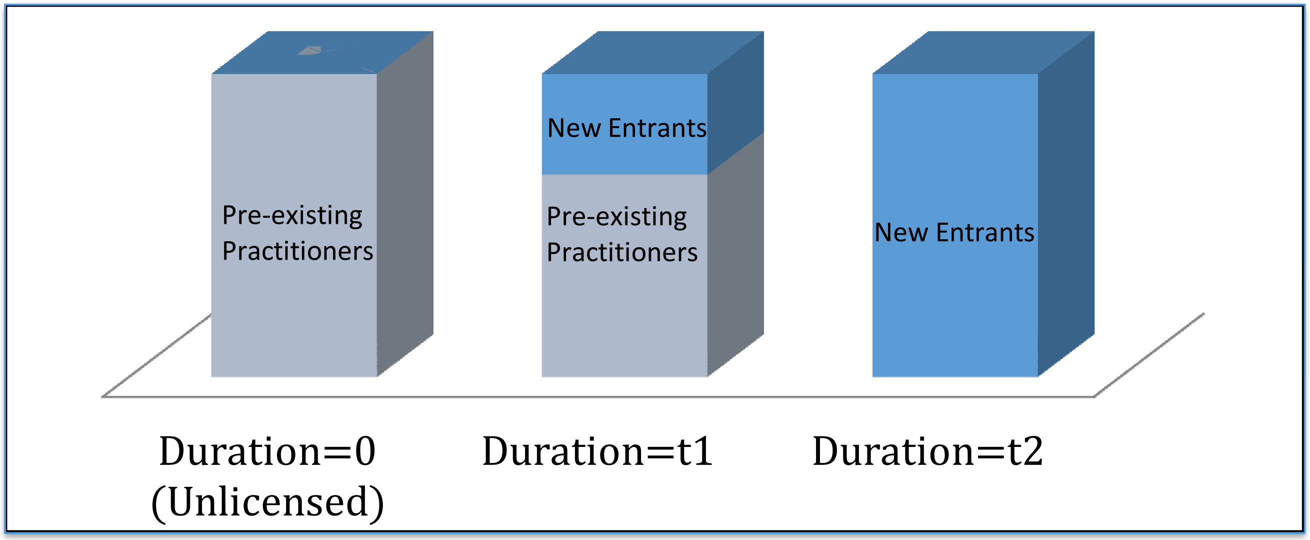
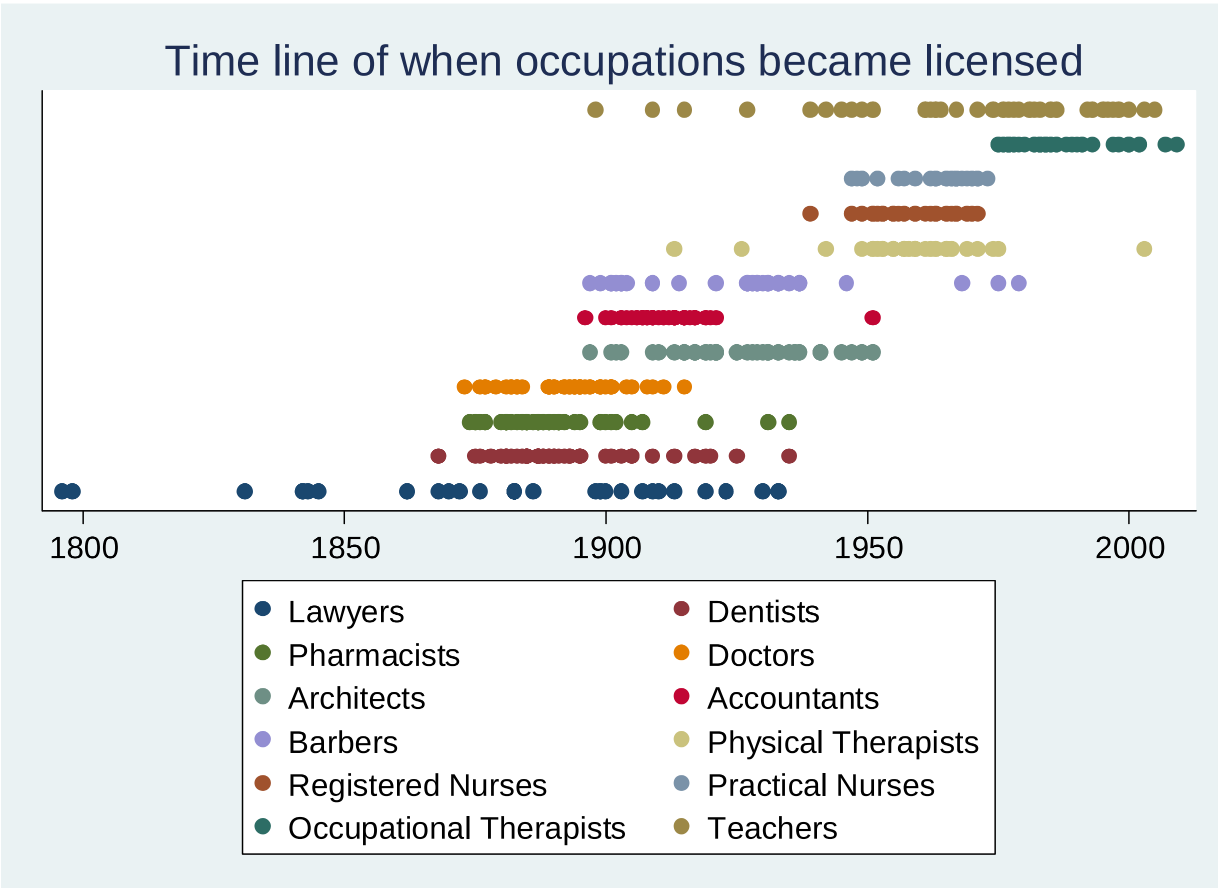


Figure 2



Developed from the authors' examination of the initial implementation of occupational licensing using *Occupational Licensing Legislation in the States* (Council of State Governments, 1952) and LexisNexis legal data services.

Figure 3

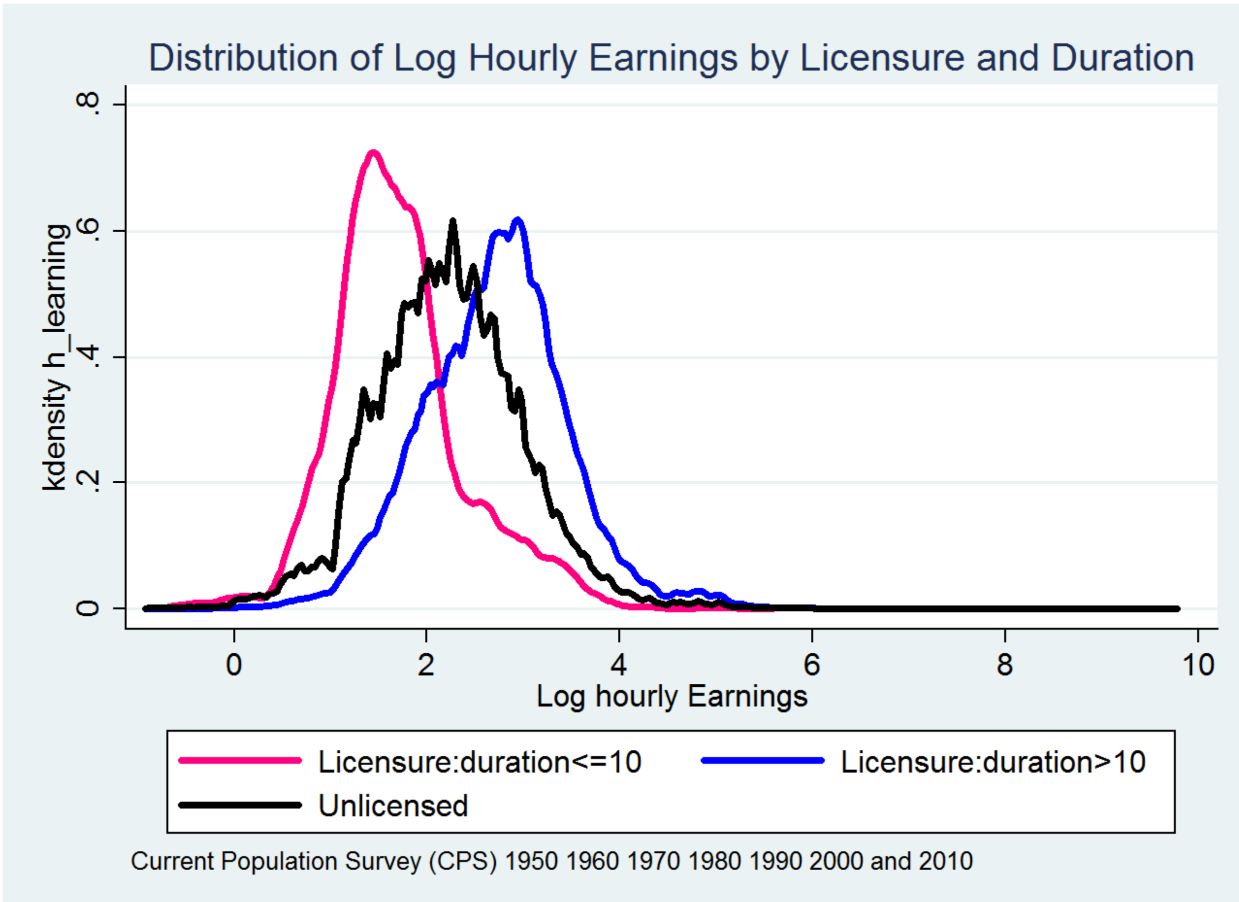
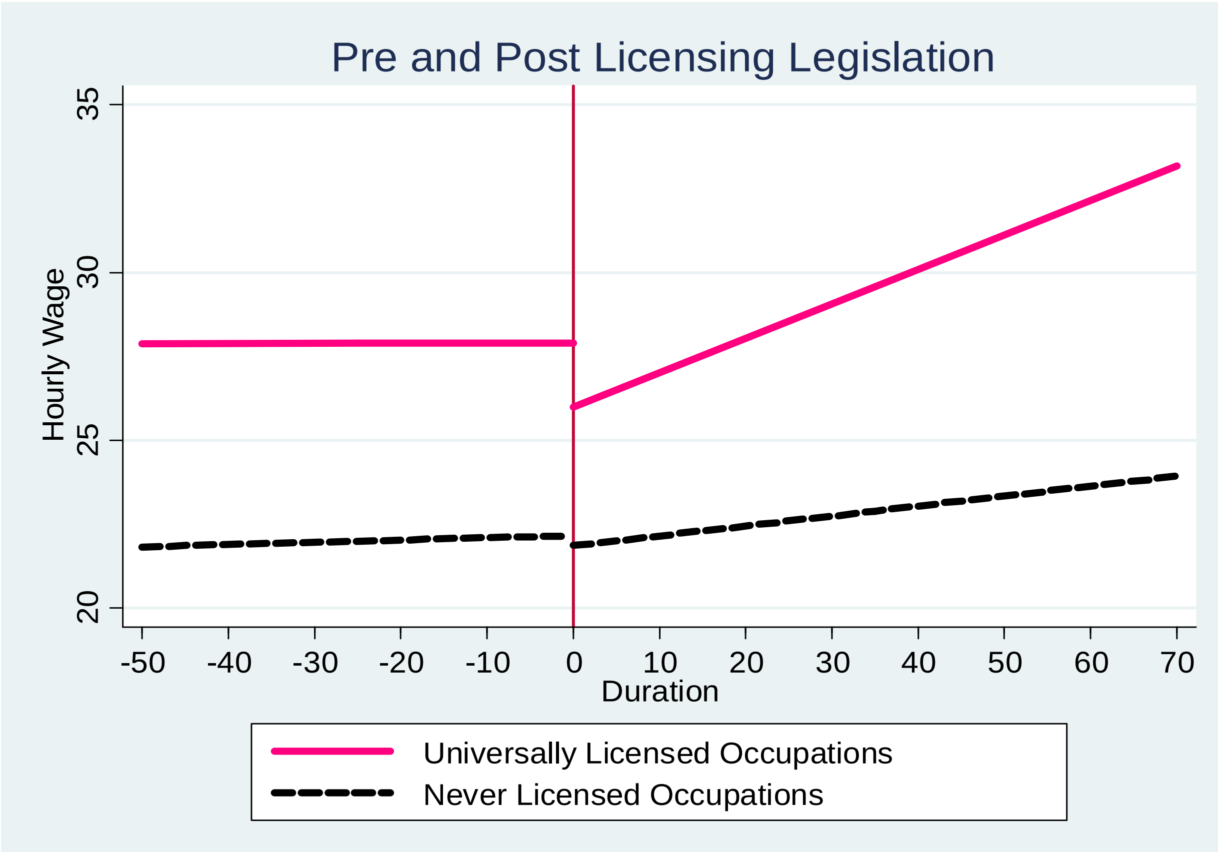

Distribution of Log Hourly Earnings by Licensure and Duration
kdensity h_learning
0
.2
.4
.6
.8
0
2
4
6
8
10
Log hourly Earnings
Licensure:duration<=10
Licensure:duration>10
Unlicensed
Current Population Survey (CPS) 1950 1960 1970 1980 1990 2000 and 2010

Figure 4. Estimated Slopes of Hourly Wages before and after the Implementation of Licensing Statutes*



*Slopes of the lines are statistically different after the implementation of a licensing statute.

Table 1. Means and Standard Deviation of Licensed and Unlicensed Occupations in the ACS sample

VARIABLES	12 Universally Licensed Occupations	Never Licensed Occupations
Age	42.03	39.12
	(11.30)	(12.57)
Male	0.31	0.49
	(0.46)	(0.50)
White	0.83	0.79
	(0.38)	(0.41)
Potential Experience	19.88	19.90
	(11.32)	(12.64)
Years of Education	16.16	13.20
	(1.82)	(2.28)
Total Worked Weeks per year	46.59	46.23
	(9.81)	(11.80)
Average Worked Hours per week	40.92	39.84
	(8.18)	(8.21)
Total Worked Hours per year	1921.72	1,867.99
	(579.73)	(636.66)
Hourly Earnings (2014 CPI)	$37.34	$24.92
	(52.78)	(40.29)
Licensure	0.94	0.00
	(0.24)	(0.00)
Years of Duration	54.27	0.00
	(38.52)	(0.00)
Industry size (% of GDP)	4.14	6.20
	(3.26)	(4.26)
N	173,935,423	842,399,749

Note: Data are weighted using population weights.

Table 2. Effects of Licensing Duration on Log Hourly Earnings

Panel A. All Occupations

VARIABLES	(1) Log Hourly Earnings	(2) Log Hourly Earnings	(3) Log Hourly Earnings	(4) Log Hourly Earnings	(5) Log Hourly Earnings	(6) Log Hourly Earnings
Duration	0.003***	0.003***	0.003***		0.0029***	
	(0.000)	(0.000)	(0.000)		(0.000)	
Duration^2			0.000			
			(0.000)			
Licensure				0.198***		0.111***
				(0.007)		(0.013)
Constant	1.009***	0.862***	0.862***	0.842***	1.463***	1.601***
	(0.028)	(0.027)	(0.027)	(0.028)	(0.014)	(0.254)
Individual Covariates	Yes	Yes	Yes	Yes	Yes	Yes
Industry Size	No	Yes	Yes	Yes	Yes	Yes
2-Digit SOC	No	No	No	No	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	14,564,036	14,564,036	14,564,036	14,564,036	14,564,036	14,564,036
R-squared	0.329	0.338	0.338	0.333	0.390	0.370

Robust Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B. Occupations that changed their regulation status

VARIABLES	(1) Log Hourly Earnings	(2) Log Hourly Earnings	(3) Log Hourly Earnings	(4) Log Hourly Earnings	(5) Log Hourly Earnings	(6) Log Hourly Earnings
Duration	0.0039***	0.0053***		0.0014***	0.0005	
	(0.000)	(0.001)		(0.000)	(0.001)	
Duration^2		-0.0000			0.0000	
		(0.000)			(0.000)	
Licensure			0.1463***			0.0363
			(0.010)			(0.025)
Constant	0.8938***	0.9061***	0.9131***	0.9108***	0.9064***	0.9168***
	(0.029)	(0.036)	(0.030)	(0.030)	(0.032)	(0.030)
Individual Covariates	Yes	Yes	Yes	Yes	Yes	Yes
Industry Size	Yes	Yes	Yes	Yes	Yes	Yes
2-Digit SOC	No	No	No	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	13,801,113	13,801,113	13,801,113	13,801,113	13,801,113	13,801,113
R-squared	0.303	0.303	0.302	0.306	0.306	0.306

Robust Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3. Effects of Licensing Duration on Total Worked Hours per Year Using the ACS

Panel A. All Occupations

VARIABLES	(1) Total Worked Hours per year	(2) Total Worked Hours per year	(3) Total Worked Hours per year	(4) Total Worked Hours per year	(5) Total Worked Hours per year	(6) Total Worked Hours per year
Duration	0.541***	0.526***	-1.693***		0.089	
	(0.061)	(0.048)	(0.332)		(0.092)	
Duration^2			0.022***			
			(0.003)			
Licensure				-15.933***		-6.465
				(5.523)		(9.431)
Constant	851.815***	713.422***	709.993***	685.108***	1,177.785***	1,129.937***
	(19.094)	(23.624)	(24.157)	(23.982)	(16.661)	(25.639)
Individual Covariates	Yes	Yes	Yes	Yes	Yes	Yes
Industry Size	No	Yes	Yes	Yes	Yes	Yes
2-Digit SOC	No	No	No	No	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	14,564,036	14,564,036	14,564,036	14,564,036	14,564,036	14,564,036
R-squared	0.141	0.149	0.150	0.149	0.176	0.189

Robust Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B. Occupations that Changed their Regulation Status

VARIABLES	(1) Total Worked Hours per year	(2) Total Worked Hours per year	(3) Total Worked Hours per year	(4) Total Worked Hours per year	(5) Total Worked Hours per year	(6) Total Worked Hours per year
Duration	-1.4147***	-3.2553***		0.5858*	1.7950**	
	(0.110)	(0.608)		(0.347)	(0.678)	
$Duration^2$		0.0345***			-0.0177**	
		(0.010)			(0.008)	
Licensure			- 72.3954***			76.2312***
			(8.935)			(17.756)
Constant	705.2900***	689.6890***	692.1555***	690.6337***	696.7370***	696.8075***
	(23.753)	(23.558)	(24.102)	(25.333)	(24.887)	(24.595)
Individual Covariates	Yes	Yes	Yes	Yes	Yes	Yes
Industry Size	Yes	Yes	Yes	Yes	Yes	Yes
2-Digit SOC	No	No	No	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	13,801,113	13,801,113	13,801,113	13,801,113	13,801,113	13,801,113
R-squared	0.147	0.148	0.148	0.148	0.148	0.148

Robust Standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 4: Results for Hourly Earnings and Annual Hours Worked Using the Two-Stage Model for Occupations that Changed Regulatory Status

VARIABLES	(1) Hourly Log Earnings	(2) Hourly Log Earnings	(3) Total Hours Worked	(4) Total Hours Worked
Duration	0.0042***		6.5442***	
	(0.001)		(0.685)	
Licensure		0.1665***		112.2298***
		(0.027)		(21.890)
Male	0.2560***	0.2560***	265.0099***	265.0099***
	(0.004)	(0.004)	(3.398)	(3.398)
Black	0.0048	0.0048	-33.5889***	-33.5889***
	(0.005)	(0.005)	(2.445)	(2.445)
White	0.0560***	0.0560***	46.6355***	46.6355***
	(0.004)	(0.004)	(2.124)	(2.124)
Potential Experience	0.0339***	0.0339***	46.2489***	46.2489***
	(0.000)	(0.000)	(0.324)	(0.324)
Potential Experience Squared	-0.5151***	-0.5151***	-874.5550***	-874.5550***
	(0.006)	(0.006)	(7.951)	(7.951)
Years of Education	0.1051***	0.1051***	35.1331***	35.1331***
	(0.002)	(0.002)	(0.512)	(0.512)
SOC 2-digits dummies	Yes	Yes	Yes	Yes
State Dummies	Yes	Yes	Yes	Yes
Year Dummies	Yes	Yes	Yes	Yes
Number of States	48	48	48	48
Number of Individuals	14,527,380	14,527,380	14,527,380	14,527,380

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 5: Estimates of the Wage Effects of Occupational Licensing with Grandfathering

Panel A. Estimates of the Influence of Grandfathering

VARIABLES	Log Hourly Earning	Log Hourly Earnings
Overall Effect of Grandfathering		0.05339+
Duration	0.0076***	
	(0.0001)	
Duration^2	0.0000	
	(0.0000)	
GPS	5.2380***	
	(0.0117)	
GPS^2	-13.1227***	
	(0.0188)	
GPS*Duration	0.0083***	
	(0.0008)	
Constant	2.8949***	
	(0.0018)	
Adjusted R Squared	0.1038	-
Number of Observations	27,863,774	12,472640

+ Estimated with a propensity matched control group

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B: Oaxaca Decomposition Analysis of New Entrants Relative to Grandfathered Workers

VARIABLES	Log Hourly Earnings
New Entrants	3.4057***
	(0.000)
Grandfathered	3.2831***
	(0.000)
Difference	0.1225***
	(0.000)
Explained	0.1505***
	(0.000)
Unexplained	-0.0280***
	(0.000)
Observations	2,320,215

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 6. Effects of Licensing Duration on Labor Market Participation

Panel A. All Occupations

VARIABLES	(1) Participation (per10,000capita)	(2) Participation (per10,000capita)	(3) Participation (per10,000capita)	(4) Participation (per10,000capita)
Duration	-1.0113***	-1.0112***	0.0434	
	(0.108)	(0.108)	(0.452)	
Duration^2			-0.0082**	
			(0.003)	
Licensure				39.7285***
				(10.779)
Constant	226.3771***	123.8382***	96.5994**	-0.0496
	(8.132)	(36.924)	(40.811)	(38.813)
Per Capita Income by State	No	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	11,157	11,157	11,157	11,157
R-squared	0.040	0.041	0.045	0.011

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B. Occupations that Changed Their Regulation Status

VARIABLES	(1) Whether Worked in Licensed Occupation	(2) Ratio of Licensed Workers Relative to Unlicensed Workers in the Service Sector
Duration: 1-3 years	-0.00002	-0.00002
	(0.0001)	(0.0001)
Duration:4-6 years	-0.0007***	-0.0007***
	(0.0001)	(0.0001)
Duration:7-9 years	-0.0003***	-0.0003***
	(0.0001)	(0.0001)
⋮	⋮	⋮
Individual Characteristics	Yes	Yes
% of people over age 65	Yes	Yes
Health Sector	Yes	Yes
Per Capita Income by State	Yes	Yes
State FE	Yes	Yes
Year FE	Yes	Yes
Number of Individuals	28,118,887	28,118,887

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 7. Heterogeneity of the Influence of Licensing and Its Duration on Wage Determination, Hours Worked and Participation in the Occupation

Summary of the Estimates

Log Hourly Wage

	Occupational Therapists	Physical Therapists	Registered Nurses	Practical Nurses	Teachers	Architects	Barbers	Accountants	Lawyers	Dentists	Pharmacists	Physicians
Duration	0.013***	0.005***	0.007***	0.002***	0.001***	0.002***	-0.003***	0.002***	0.004***	0.006***	0.005***	0.006***
Licensure	0.313***	0.264***	0.320***	0.089***	0.048***	0.131***	-0.178***	0.158***	0.438***	0.733***	0.507***	0.694***

Total Worked Hours

	Occupational Therapists	Physical Therapists	Registered Nurses	Practical Nurses	Teachers	Architects	Barbers	Accountants	Lawyers	Dentists	Pharmacists	Physicians
Duration	-4.092***	-1.193***	-1.520***	-0.450***	-1.607***	0.487***	1.068***	0.7668***	1.208***	-2.120***	-0.556***	1.684***
Licensure	-111.051***	-64.433***	-72.787***	-21.846 ***	-80.412***	44.610***	110.711***	70.7803***	154.158***	-247.025***	-58.480***	201.049***

Labor Market Participation

	Occupational Therapists	Physical Therapists	Registered Nurses	Practical Nurses	Teachers	Architects	Barbers	Accountants	Lawyers	Dentists	Pharmacists	Physicians
Duration	0.0005***	0.0002***	0.0002***	0.0002***	-0.0003***	0.0001***	0.000	0.000***	0.000**	-0.000	0.000***	0.000
Licensure	0.0099***	0.0130***	0.0115***	0.009***	-0.0153***	0.0074***	0.006*	0.010***	0.002	-0.002	0.004***	0.001

*** p<0.01, ** p<0.05, * p<0.1

Appendix Table 1: Number of Observations by year

Year	12 Universally Licensed Occupations		Never Licensed Occupations	Total Observations
	Occupations that changed their regulation status	Occupations that did not change regulation status		
1950	433,144	227,038	4,818,133	5,478,315
1960	1,570,942	664,150	14,828,936	17,064,028
1970	2,331,500	976,950	20,511,350	23,819,800
1980	4,180,640	1,608,500	31,886,760	37,675,900
1990	5,809,132	2,637,992	43,990,679	52,437,803
2000	6,744,760	3,055,109	51,386,716	61,186,585
2001	6,671,902	3,011,143	52,091,756	61,774,801
2002	6,986,823	3,036,777	52,605,593	62,629,193
2003	7,193,663	3,222,882	51,976,359	62,392,904
2004	7,194,203	3,317,048	52,258,247	62,769,498
2005	7,498,915	3,469,741	52,602,529	63,571,185
2006	7,556,357	3,519,840	54,100,647	65,176,844
2007	7,763,758	3,641,144	54,890,858	66,295,760
2008	7,955,422	3,738,645	54,976,604	66,670,671
2009	7,915,266	3,735,327	52,901,321	64,551,914
2010	7,824,206	3,614,249	49,970,894	61,409,349
2011	7,776,931	3,665,235	48,229,650	59,671,816
2012	7,881,929	3,725,257	48,895,429	60,502,615
2013	7,920,655	3,709,090	49,477,288	61,107,033

Note: Data are weighted using population weight.

Appendix Table 2: Means and Standard Deviation of Occupations that changed their regulation status and Occupations that did not change regulation status in the ACS sample

VARIABLES	Occupations that changed their regulation status	Occupations that did not change regulation status
Age	42.06 (11.30)	39.14 (12.56)
Male	0.31 (0.46)	0.48 (0.50)
White	0.83 (0.38)	0.79 (0.41)
Potential Experience	19.88 (11.31)	19.89 (12.63)
Years of Education	16.18 (1.81)	13.23 (2.29)
Total Worked Weeks per year	46.74 (9.74)	46.20 (11.79)
Average Worked Hours per week	41.03 (8.18)	39.83 (8.20)
Total Worked Hours per year	1932.47 (577.95)	1,866.38 (636.07)
Hourly Earnings (2014 CPI)	$37.53 (52.97)	$25.01 (40.43)
Licensure	1.00 (0.00)	0.00 (0.00)
Years of Duration	57.69 (37.04)	0.00 (0.00)
Industry size (% of GDP)	4.19 (3.27)	6.17 (4.25)
N	163,918,377	850,228,619

Note: Data are weighted using population weights.

Appendix Table 3: Robustness check: Estimated Dose Response Function using the Generalized Propensity Score

In order to remove selection bias caused by endogenous selection into the occupations, we implement the generalize propensity score approach (Hirano and Imbens, 2004). We calculate the GPS based on race, sex, experience, education and year.

This appendix table 3 contains the estimation results for the dose-response function. Standard errors are bootstrap standard errors from 100 replications.

VARIABLES	Log Hourly Earning
Duration	0.0076***
	(0.0001)
Duration^2	0.0000
	(0.0000)
GPS	5.2380***
	(0.0117)
GPS^2	-13.1227***
	(0.0188)
GPS*Duration	0.0083***
	(0.0008)
Constant	2.8949***
	(0.0018)
Adjusted R Squared	0.1038
Number of Observations	27,863,774

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Appendix Table 4: State-Level Effects of Licensing Duration on Log Hourly Earnings

Panel A. All Occupations

VARIABLES	(1) Log Hourly Earnings	(2) Log Hourly Earnings	(3) Log Hourly Earnings
Duration	0.0068***	0.0039***	
	(0.000)	(0.001)	
$Duration^2$		0.0000*	
		(0.000)	
Licensure			0.4296***
			(0.027)
Constant	2.8240***	2.8841***	2.7619***
	(0.016)	(0.031)	(0.026)
State FE	YES	YES	YES
Year FE	YES	YES	YES
Observations	11,157	11,157	11,157
R-squared	0.417	0.423	0.212

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B. Occupations that Changed Their Regulation Status

VARIABLES	(1) Log Hourly Earnings	(2) Log Hourly Earnings	(3) Log Hourly Earnings
Duration	0.0052***	0.0110***	
	(0.000)	(0.001)	
$Duration^2$		-0.0001***	
		(0.000)	
Licensure			0.2868***
			(0.013)
Constant	2.7062***	2.7443***	2.6952***
	(0.013)	(0.017)	(0.016)
State FE	YES	YES	YES
Year FE	YES	YES	YES
Observations	5,112	5,112	5,112
R-squared	0.458	0.490	0.509

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Appendix Table 5: State-Level Effects of Licensing Duration on Total Worked Hours per Year

Panel A. All Occupations

VARIABLES	(1) Total Worked Hours per year	(2) Total Worked Hours per year	(3) Total Worked Hours per year
Duration	2.9184***	3.3150***	
	(0.103)	(0.314)	
Duration^2		-0.0031	
		(0.002)	
Licensure			159.1023***
			(14.140)
Constant	1,948.1033***	1,939.7617***	1,940.2907***
	(11.503)	(13.396)	(14.857)
State FE	YES	YES	YES
Year FE	YES	YES	YES
Observations	11,157	11,157	11,157
R-squared	0.212	0.213	0.062

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Panel B. Occupations that Changed Their Regulation Status

VARIABLES	(1) Total Worked Hours per year	(2) Total Worked Hours per year	(3) Total Worked Hours per year
Duration	0.1820	-0.4740	
	(0.138)	(0.376)	
Duration^2		0.0109	
		(0.007)	
Licensure			-23.5850***
			(5.951)
Constant	1,796.5019***	1,792.2151***	1,806.2319***
	(14.667)	(16.259)	(14.196)
State FE	YES	YES	YES
Year FE	YES	YES	YES
Observations	5,112	5,112	5,112
R-squared	0.108	0.109	0.111

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1