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Evaluating the Labor Supply and Migration Effects of the Interstate Medical Licensure Compact

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Abstract

A growing body of literature suggests that occupational licensing distorts economic behavior through various mechanisms, such as inflating prices and inhibiting interstate mobility. To combat some of these ill-effects, policymakers have turned to interstate compacts as a way to promote uniformity in licensing requirements across states and facilitate license portability. Despite the development of interstate compacts for numerous licensed occupations and professions, evidence of their efficacy in the literature is thin. Based on data from over 70,000 physicians from 2012 to 2018, I construct a difference-in-differences model to estimate the effects of the adoption of the Interstate Medical Licensure Compact (IMLC) on labor supply and interstate mobility. The results suggest that labor supply and migration/commuting did not significantly change following IMLC adoption.

1. Introduction

The prevalence of occupational licensing has grown at a prolific rate in recent decades. In 1950, the U.S. licensed less than 5% of workers (The Council of State Governments, 1952). Today, that number has swelled to over 25% (U.S. Bureau of Labor Statistics, 2018). The increasing ubiquity of occupational licensing in U.S. labor markets has, correspondingly, captured the attention of policymakers due to its potential distortionary impact on labor supply and interstate migration.

The U.S. generally regulates labor markets in three ways: registration, certification, or licensing. Registration requires individuals to formally add their names to an official roster maintained by a governmental entity. Certification is a wholly voluntary process by which a governmental entity or, more commonly, a private organization "certifies" the competency of individuals wishing to obtain a "right to title," such as a Chartered Financial Analyst or ACSM Certified Personal Trainer®. Licensing requires would-be providers of services and goods to obtain a license from a government-sanctioned licensing agency before they can legally engage in a profession. The process of acquiring licensure usually involves tests, training, and/or fulfillment of relevant experience and education requirements to the satisfaction of a state licensing board or division.

In short, registration is legally required but not meritocratic; individuals cannot be excluded on the basis of professional competence. Certification is meritocratic but not legally required. Occupational licensing, on the other hand, is both legally required and meritocratic. For this reason, many consider licensing to be the most stringent form of labor market regulation.

Certainly, a sizable portion of the growth in occupational licensing in recent years is attributable to the restructuring of the U.S. economy towards more service-oriented industries

that, by their very nature, tend to be more tightly regulated, such as healthcare and education (U.S. Bureau of Labor Statistics, 2018). However, changes in the composition of the workforce only account for slightly more than one-third of the uptick in licensing. An increase in the number of licensed occupations explains the rest, signifying a widening regulatory scope of licensing (U.S. Department of the Treasury Office of Economic Policy, Council of Economic Advisers, and the Department of Labor, 2015).

While the intent of occupational licensing is to safeguard consumers and certify product and service quality, researchers have identified it as a significant source of economic distortion, such as muted geographic mobility (Johnson & Kleiner, 2017). Since occupational licenses are issued at the state level, licensed workers wishing to move or practice across state lines may encounter high costs and redundant red tape associated with obtaining additional licenses, such as fees, exams, background checks, and training. The resulting labor market rigidity may diminish wages, employment and overall economic output (Friedman, 1962).

The healthcare industry may feel these inefficiencies the strongest. New technologies and healthcare delivery modalities, such as telemedicine, bring opportunities for improved consumer access, but due to the nature of state-level licensing, both providers and patients may only realize these opportunities within state borders. To combat some of these ill-effects, policymakers have turned to interstate compacts as a way to promote license portability and facilitate interstate practice. Although the mechanics of each occupational licensing interstate compact are unique, their shared purpose is to create reciprocal agreements between states that allow licensed workers to more easily practice across state lines.

Despite the development of compacts for numerous health professions such as physical therapy, emergency medical services, nursing, and others, there is almost no evidence of their

efficacy in the literature. In the only existing study of occupational licensing interstate compacts, DePasquale and Stange (2016) found that the Nurse Licensure Compact did not have an effect on nurse mobility or labor supply. In this paper, I provide new evidence on the question of the efficacy of occupational licensure interstate compacts by examining the effects of the Interstate Medical Licensure Compact (IMLC) on interstate mobility and economic indicators for physicians. The motivation for my inquiry is multifaceted. The medical licensure compact's utilization of a compact model that is fundamentally different from the Nurse Licensure Compact uniquely positions it to foster interstate migration. This feature makes it an interesting subject to study because researchers have not examined the impacts of this compact model in the literature. Moreover, with the federal government seeking to dedicate funding for the development of additional licensure compacts over the next four fiscal years[1], more research into their impacts is needed.

Using American Community Survey data on physicians from 2012-2018, I compare labor supply and interstate mobility measures for physicians in compact states to those in non-compact states, controlling for observable time-variant and latent time-invariant factors through a difference-in-differences approach. I provide suggestive evidence that the identifying assumptions may be credible through an event-study analysis that tests for differential pre-trends between states that adopted IMLC over my sample and those that did not. My results are consistent with DePasquale and Stange (2016) and find that IMLC adoption did not have a marked impact across outcome variables.

I organize the rest of this paper as follows. Section 2 provides background information on the theory of licensing, the IMLC, and reviews related literature on licensing. Section 3 describes

[1] National Defense Authorization Act for Fiscal Year 2020, Pub. L. 116 - 92

my data. Section 4 delineates my empirical strategy and limitations. Section 5 details my results, and Section 6 presents a summary and concluding remarks.

2. Background

a. Policy Analysis and Theory of Occupational Licensing

The first fundamental theorem of welfare economics states that complete markets – that is, markets in which prices reflect all social costs and benefits – efficiently allocate goods under information symmetry and perfect competition (Hammond, 1998). Advocates of occupational licensing often contend that licensing rectifies violations of this theorem.

Information asymmetry between buyers and sellers is perhaps the most commonly cited violation of this theorem. Transaction costs can be prohibitive for consumers. Buyers may not have the time, know-how, or motivation to adequately assess the professional competence of sellers, which can engender threats to public health/safety and stifle demand (Arrow, 1971). Therefore, proponents of licensing maintain that it corrects information asymmetries by creating a legally binding threshold of minimum competency for anyone seeking to bring a good or service to market. Licensing may have positive supply-side effects as well. When sellers face competition from purveyors of cheap, low-quality substitutes, which may be more common in an unregulated market, incentives to improve or maintain product/service quality may diminish (Akerlof, 1970; Shapiro, 1986). These diminutions are especially pronounced in markets for goods or services that entail large transaction costs. Thus, licensing may encourage human capital investment and foster elevated product/service quality in the long run by increasing the financial return to quality improvements.

Correction for externalities is another common justification for occupational licensing. If I choose to do business with an untutored and negligent mechanic, I certainly put myself at risk of

having more frequent and disastrous traffic incidents, which might not necessarily be of concern to a social planner if I am willing to bear that hazard, but I also gratuitously assign that risk to anyone else who may be on the road in my presence. Therefore, government intervention may be justified in restricting the market for mechanics to only those that possess some socially optimal level of expertise or skill.

While the aforementioned arguments in favor of licensing are bound up with notions of economic efficiency and social desirability, some detractors of licensing claim its origins are not so innocuous. The economics of collective decision-making provides us a framework with which to analyze the case against licensing. Representative government tends to undertake projects that are economically inefficient when the benefits of such projects accrue to a small, politically influential group and the costs are widely dispersed among the public (Gwartney et al. 2014). This is known in public choice theory as the special-interest effect and can be reasonably extended to occupational licensing.

The social costs of licensing largely stem from reduced market competition, which often leads to higher prices, curtailed innovation, and lower employment. For members of the general public, these costs are not particularly salient. The average voter has little incentive to care about the economic distortions of licensing. Although the undesirable economic effects of licensing may be large in the aggregate, they are relegated to insignificance at the individual level – a blip among the slew of causal factors that determine prices, innovation, and employment. However, professional associations and current licensees (in this case, the “special interests”) may stand to profit enormously from regulating competitors out of the market and fortifying their market share. To this end, the pro-licensing groups may provide campaign contributions and volunteers, in addition to an influential and devoted voting bloc, to the amenable politician.

Arguably, one of the consequences of these perverse incentives is that the gatekeepers (i.e., licensing board members) are often market participants themselves. For instance, Kentucky Revised Statutes requires the Kentucky Medical Board of Licensure to comprise eight currently licensed physicians and only three public members. One might riposte that market participants have the appropriate knowledge of the industry and are therefore most qualified to decide who ought to receive a license, but it is clear that if board members succumb to self-interest and are willing to use the powers of the state to regulate competitors out of the market, little recourse remains. The fact that licensing boards have (and often exercise) the ability to burden market entrants with grandfathered regulations to which they themselves to do not have to adhere is suggestive evidence that this hypothesis may carry water.

b. Related Literature

The empirical literature supports a special-interest effect. Past research has largely focused on the relationship between licensing and wages, with most studies finding a positive link between wages and licensing (Kleiner & Vorotnikov, 2017; Thornton & Timmons, 2013; Weeden, 2002). These results are consistent with a reduction in labor supply but are also expected with quality improvements. Kleiner & Krueger (2013) found that licensing is associated with approximately 14% higher wages after controlling for education, training, and experience, suggesting the wage premium associated with licensing may be a function of reduced competition. The authors found that the wage effects related to voluntary certification were much smaller, providing further evidence of rent-seeking. Blair & Chung (2018) employed a border discontinuity design to identify the causal effects of licensing on labor supply. By comparing counties that border another state with different licensing requirements, the authors concluded that licensing reduced labor supply an average of 17%-27%.

An important component of occupational licensing's effects is the degree to which it deters interstate migration, as a lack of labor market fluidity may be partially driving the aforementioned distortions. Johnson & Kleiner (2017) found that individuals in occupations with state-specific licensing exam requirements moved between states at a 36% lower rate than members of occupations with a national licensing exam. The results imply that implementing a national licensing policy would considerably increase interstate migration. This result was foreshadowed by Brüggemann, Bloomfield, Christensen, & Leuz (2015) who found that harmonization of occupational licensing requirements markedly increased labor migration in the EU.

c. The Interstate Medical Licensure Compact

Professional organizations, nonprofits, state licensing boards, and policymakers have turned to interstate compacts as a potential tool to mitigate some of the distortionary effects of occupational licensure. In April 2013, the Federation of State Medical Boards (FSMB) adopted a resolution to begin development of an interstate compact to expedite medical licensure and facilitate multi-state practice. A drafting team composed of state medical board executives, administrators and attorneys finalized the IMLC language in 2014, with activation contingent upon adoption in seven states. The interstate commission, which oversees the administration of the compact, was seated in 2015, and the compact began processing applications in April 2017. Today, the compact has been adopted by 29 states, the District of Columbia, and the Territory of Guam.

The IMLC model permits qualifying physicians to obtain an "expedited license" from any compact state. The expedited license is not tied to residency and physicians are permitted to both practice and live in any compacts state in which they possess an expedited license. This process

is carried out as follows. First, a licensed physician must designate a compact state as his or her "principal state of licensure."[2] The physician must then apply for an expedited license from the medical board of the physician's state of principal licensure. The board will then issue a "letter of verification," provided the physician meets the requirements outlined in the compact, to the Interstate Commission. This letter confirms to member states that the applicant is qualified and thus entitled to a license through the interstate compact. The physician must then register through the Interstate Commission, pay applicable licensing fees,[3] and request a license from the state(s) in which the applicant would like to practice.

The Nurse Licensure Compact (NLC) model, on the other hand, issues a single "multistate license" to applicants that permits nurses to practice in any compact state insofar as state of residence *remains fixed*. By changing state of residence, nurses forgo their "multistate" license and are required to apply for a standard license (or new multistate license) in their new state of residence. This distinction highlights the fundamental difference between the NLC and the IMLC. In short, the NLC issues a single license that permits nurses to work in any compact state provided their state of residence is static, whereas the IMLC provides applicants with a pathway to an expedited license in any compact state. The former facilitates cross-state practice, while the latter facilitates interstate migration.

Rather than focusing legislative efforts on consolidating state-level licensing administrations into a national regime, it appears interstate compacts are the channel through which policymakers have opted to address some of these concerns associated with occupational licensing (The Council of State Governments, 2019). In recent years, state licensing boards,

[2] This is the state in which the physician holds a full and unrestricted license and is where the physician resides; the state where at least 25% of the practice of medicine occurs; the location of the physician's employer; or the state designated as the state of residence for federal income tax purposes.
[3] I provide a list of IMLC licensure fees in Table A1 in the Appendix.

professional organizations and other nonprofits have come together to develop compacts for numerous occupations and professions, including psychology, physical therapy, nursing, and audiology/speech-language pathology. Furthermore, many other professions, such as occupational therapy, dentistry, and teaching have taken preliminary steps to begin the development of an interstate compact. The federal government is positioned to provide financial support to these efforts. The FY 2020 National Defense Authorization Act provides up to $4 million per year (until September 30, 2024) to facilitate the development of occupational licensing interstate compacts. This paper's contribution to the literature is providing an empirical evaluation of a growing policy tool whose influence is felt by numerous occupations and professions and for which existing literature is thin. Physicians are also an important group to consider, as physician shortages are expected to continue to grow in the near future (Dall et al., 2019). Furthermore, expanding healthcare access for underserved communities is an important goal for policymakers that is purportedly achieved through interstate compacts, and up to this point, the empirical impacts of these efforts have largely gone unexamined in the literature.

3. Data

For my empirical analysis, I use public microdata from the American Community Survey (ACS) published by IPUMS-USA (Ruggles, et al. 2019). The ACS is the only public use microdata survey of which I am aware that collects detailed information on both migration, occupation, and labor supply measures. The sample includes 72,175 physicians over the period of 2012 to 2018. The sample begins in 2012 because it is the earliest year in which the ACS collects occupation information for physicians and physician assistants. The sample covers up to 2018 since it is the most recent year for which data are available. The data contain detailed information on migration, travel time to work, and state of employment. I also include economic

and demographic variables such as labor force status, race, sex, age, education, income, and number of children in my analysis.[4] I restrict the sample to physicians and physician assistants who had worked within the previous five years (excluding new workers).[5] The Interstate Medical Licensure Compact Commission provides data on the date in which states adopt and enact the compact.[6]

There are 3,142 counties in the U.S., about a third of which border at least one other state.[7] However, the ACS only collects the county of residence for households residing in counties that (*i*) are coterminous with a single State Economic Area, county group, or Public Use Microdata Area; or (*ii*) contain multiple State Economic Areas, county groups, or Public Use Microdata Areas, none of which extend into other counties. Therefore, I am only able to identify 589 counties over the entire sample period, 209 of which border at least one state. Across the full sample period, individuals residing in border counties account for approximately 22% of my total sample.

Table 1 presents descriptive statistics for the full sample of 72,175 physicians for 2012-2018. I categorize the sample by whether respondents resided in a border county and a compact/non-compact state. Across all sample years, around 93% of the sample were in the labor force, and 92% were employed, working 47 hours per week on average and with mean earnings of $194,000. Only 32% of the sample were female, and 74% of respondents were white. Approximately 4% of physicians in the sample did not work in the same state in which they

[4] Income is measured in real 2018 dollars.
[5] The ACS codes the occupation of unemployed persons as their most recent occupation.
[6] I provide a full list of IMLC implementation dates for all states that are currently part of the compact in Table A2 in the Appendix.
[7] See Curtis, E. Mark, and Ryan A. Decker (2018).

lived. The average commute time (from home to work) was 22 minutes. As expected, physicians in border counties were more likely to work in a different state.

The raw data hint at the potential impacts of the IMLC on outcome variables. Physicians in compact states had higher average incomes, were more likely to have moved between states in the previous year, and had lower commute times than their counterparts in non-compact states. However, the data also signify the need for a more sophisticated analysis in order to approach causal identification, as these baseline differences need not be attributable to the IMLC. Physicians in compact states were also slightly older, more likely to be domestic-born, and white.

4. Empirical Strategy

The IMLC is not randomly assigned, and thus an obvious challenge to estimating the causal effects of its adoption is the presence of selection effects. States that choose to enact the IMLC may be fundamentally different than those that do not, and if those differences are correlated with outcome variables, then the estimates of such a comparison would be biased. For instance, physicians in states that adopt the IMLC may be inherently more mobile or have stronger attachments to the labor force than physicians in non-compact states.

a. Difference-in-Differences Analysis

To control for potential endogenous factors, I employ a difference-in-differences model that accounts for any unobservable, time-invariant state selection effects, aggregate time trends, and time-varying demographic variables. Equation 1 formalizes this approach:

$$y_{ist} = \beta_0 + \beta_1 Compact_{st} + \beta_x X_{ist} + \alpha_s + \gamma_t + \varepsilon_{ist} \quad (1)$$

where y_{ist} are outcome variables (labor force participation, employment status, usual hours worked, income, works in a different state, works in a different compact state, commute time, and lives in a different state) for individual i in state s during year t; $Compact_{st}$ is a dummy variable equal to one during years in which a state is an active member of the compact (equal to zero for non-compact states in all periods and for compact states prior to adoption); coefficient β_1 is the difference-in-differences estimate; X_{ist} is a vector of individual covariates.[8] State and year fixed effects are accounted for by α_s and γ_t. To account for the complex survey design of the ACS, replicate weights at the household and person levels are used to generate empirically derived standard error estimates for all specifications. Considering that the IMLC did not go into effect until 2017, I define the treatment period as 2017 for all states that enacted the compact legislation in my sample prior to 2017. Due to limitations of my data, I am unable to observe if physicians actually obtain a compact license. Therefore, β_1 captures the intention-to-treat (ITT) effect.[9]

b. Triple Difference Specification

Estimates from the difference-in-differences approach outlined above may be biased if IMLC adoption is contemporaneous with shocks that correlate with physician labor supply and migration trends. For example, states may enact the IMLC in anticipation of dwindling physician supply or growing demand. Identifying a second control group within compact states that potentially share these time-varying selection effects can overcome this threat. In their study of the NLC, DePasquale & Stange (2016) construct a triple difference model using non-nurse

[8] These include age, sex, race, education, and number of children.

[9] Depending on the degree to which physicians in compact states utilize the compact, my estimates may be smaller than the expected effect on physicians who actually obtain a compact license. However, the relevant research question is the effect of occupational licensing interstate compacts *as a policy tool*. Compacts cannot force physicians to obtain a compact license, so ITT effects are the estimates of interest if we seek to evaluate compacts on policy grounds.

healthcare workers as a second control group. They argue that time-varying selection effects can effectively be purged from their estimates to the extent that these selection effects equally impact nurses and non-nurse healthcare workers within compact states.

I explored the possibility of a triple difference specification in this study. Ultimately, I did not opt to employ this strategy here. For one, nurses, who are now part of a compact (i.e., treated), comprise a significant proportion of this potential control group (i.e., non-physician healthcare workers). What's more, state boards do not license some non-physician healthcare workers at all, casting doubt on the credibility of this group as a counterfactual for physicians. In an attempt to end-run this issue by identifying a better analog, I considered the prospect of using physician assistants as a narrower and more precise second control group. However, a series of simple line graphs showed that trends between physicians and physician assistants across outcome variables over the sample period were not parallel, signifying a clear dissimilarity and lack of comparability between the two groups.[10]

c. Event-Study

One of the key assumptions of the standard difference-in-differences model is that the control group serves as a compelling counterfactual for the treatment group. That is, the control group can reasonably serve as a proxy for how the treatment group would have trended across outcome variables in the absence of the treatment. In this case, I assume that, conditional on covariates and state/year fixed effects, states that did not enact the IMLC exhibit a trend equivalent to what IMLC states would have followed "but for" IMLC adoption. While fundamentally untestable, this assumption can be made more credible if one is able to demonstrate that the treatment and control groups exhibit parallel trends in the pre-intervention

[10] These graphs are available in the Appendix.

period. I attempt to establish suggestive evidence of this assumption by dint of an event-study. This analysis is formalized by Equation 2:

$$y_{ist} = \beta_0 + \beta_1 IMLC_{st} * \theta_{t-5} + \beta_2 IMLC_{st} * \theta_{t-4} + \beta_3 IMLC_{st} * \theta_{t-3} + \beta_4 IMLC_{st} * \theta_{t-2} + \beta_6 IMLC_{st} * \theta_t + \beta_6 X_{ist} + \alpha_s + \gamma_t + \varepsilon_{ist} \quad (2)$$

where y_{ist} is the outcome for individual i in state s in year t. $IMLC_{st}$ is a dummy variable that is equal to 1 if a state joined the compact during the sample period. X_{ist} is a vector of individual covariates. State and year fixed effects are accounted for by α_s and γ_t. θ_t is a dummy variable indicating each unit's timing relative to the treatment year. The omitted category is the year just before compact implementation.

d. Border County Analysis

Given that one of the primary goals of the IMLC is to facilitate license portability across state boundaries, it is important to consider the possibility that the effect of IMLC adoption is different for residents who live in close proximity to neighboring states. Residents of border counties have much smaller costs associated with working across state lines. This is evident in the sample. Across the sample period, about 10% of physicians living in border counties worked in a different state from which they lived, whereas this was true for only about 2% of physicians in non-border counties. This trend need not necessarily hold for migration, at least for the border sample, as the ability to more easily become permitted to commute to a nearby state for work may offset any increase in the incentive to move.

To explore this potential relationship, I supplement my primary state-level inquiry with a more refined border county analysis. For this portion of my paper, I reorient my state-level difference-in-differences approach to the county-level. Adapting Equation 1 slightly, I restrict my sample to only those individuals who reside in border counties identified in the ACS and

modify my panel variable from the state to the county. For this analysis, I define treatment as (*i*) living in a state that has enacted the IMLC and (*ii*) living in a county that borders at least one other state that is also part of the compact. I utilize two counterfactual groups for this analysis. In my county fixed effects specification, I use the experience of all border counties in untreated states as the counterfactual. I can probe the robustness of this approach by employing state fixed effects, comparing outcomes for individuals in treated border counties with those in untreated border counties within the same state. The latter being the more valid approach if political, cultural, and institutional differences between states influence outcome variables.

e. Limitations

The IMLC was not activated until April of 2017, yet the ACS data used here describe the population of interest over the entire year of 2017. Consequently, my treatment indicator captures the policy change imperfectly, as I code treatment as having occurred in 2017. This measurement error introduces bias into my estimates, understating the IMLC's "true" effect.[11] Furthermore, the policy likely takes more than a single year to be fully effectuated due to application processing time, lagged awareness among potential applicants, etc.

While I control for time-invariant selection effects (e.g., states with more mobile physicians may be more likely to adopt the compact) in my specifications, I fail to control for unobserved time-varying selection effects. DePasquale & Stange (2016) attempt to mitigate this concern in their study of the NLC by constructing a triple difference estimator, using non-nurse healthcare workers in treated states as an additional control group. I attempted to build on this approach by constructing a triple difference model with a more refined second control group

[11] In cases where the expected direction of IMLC's effect is positive, the bias is negative. In cases where the expected direction is negative, the bias is positive.

(physician assistants), but a trend graph demonstrated a clear lack of comparability between the two groups. Therefore, time-varying selection effects remain a threat to validity.

The external validity of my border county analysis is inherently limited, given that some border counties are not observable in the ACS data. Likewise, the sample does not include new market entrants. Accordingly, my results must be interpreted as conditional on being in the workforce for at least a year. This excludes workers who might have a higher propensity to migrate for work. Further, the ACS' primary place of work measure does not capture all of the relevant dynamics of interstate practice. Ideally, I would like to know how much work physicians conduct outside of their state of residence. However, I only observe a proxy measure for this: the state in which a physician *primarily* works. If the IMLC catalyzes additional out of state work but does not change how a physician interprets his or her "primary state of work," my estimates will be understated. Lastly, states may adopt other policies in concert with the IMLC. If these policies also affect outcome variables, my estimates will be biased.

5. Results and Discussion

a. Main Results

I present my main results from the full sample in Table 2. Columns (1)-(4) report the estimated impact of the IMLC on labor supply/economic measures. The IMLC did not have an observable effect on labor force participation, employment, or usual hours worked. Point estimates are not particularly precise. I can only rule out moderate positive effects for labor force participation (95% CI = -0.009 to 0.014), employment (95% CI = -0.007 to 0.017), and hours worked (95% CI = -0.670 to 1.302). Within compact states, IMLC adoption was associated with

a $11,238 increase in income, all else constant (95% CI = $4,314.76 to $18,161.14).[12] This was significant at the 1% level. The IMLC's significant positive relationship with income hints at the possibility of emboldened bargaining power afforded to physicians in compact states.

Columns (5)-(8) report the estimates for migration and commuting outcome variables. Within compact states, the impacts of IMLC adoption on working in a different state (95% CI = -0.004 to 0.013), living in a different state (95% CI = -0.010 to 0.015), working in a compact state (95% CI = -0.003 to 0.007), and travel time to work (95% CI = -0.666 to 1.324) were not statistically different from zero. These null effects are consistent with DePasquale & Stange (2016). Interestingly, I find no evidence that the expedited licensure compact model had a stronger impact on migration (i.e., the living location) or commuting (i.e., working out of state) than the mutual recognition compact model of the NLC.

While IMLC enactment did not appear to have a significant relationship with migration and commuting variables in the full sample, it is noteworthy that this was not the case for income. Even if the IMLC did not spur migration or interstate practice, the mere possibility of working across state lines may have increased bargaining power for physicians in compact states, allowing them to negotiate higher pay. Previous literature has argued that licensing is associated with higher wages because of decreased mobility and reduced labor supply. However, this result suggests that, at least for physicians in compact states, inhibited mobility may have actually reduced wages. Unfortunately, labor market shocks that may correlate with IMLC adoption (e.g., growing demand for physicians or declining supply) could also reasonably be behind this result, and therefore I caution this interpretation.

b. Border County Results

[12] Income of unemployed persons is coded as zero. I test the robustness of this result by excluding the sample to observations with incomes above zero and obtain nearly identical point estimates.

Table 3 presents results from my border county analysis. Surprisingly, I find no effect of the IMLC on labor supply measures. Again, point estimates are not particularly precise, and I am unable to rule out small positive effects for labor force status (95% CI = -0.022 to 0.018), employment (95% CI = -0.024 to 0.017), and hours worked (95% CI = -2.895 to 0.852). Within border counties, IMLC adoption was associated with a 4-point decrease in the proportion of physicians who worked in a different state, on average (95% CI = -0.071 to -0.025). This result is somewhat unexpected but does not hold when I include state fixed effects. IMLC adoption did not have a significant relationship with other migration and commuting measures. Estimates permit ruling out only moderate positive effects for working in a different compact state (95% CI = -0.008 to 0.014), living in a different state (95% CI = -0.026 to 0.016), and travel time to work (95% CI = -2.816 to 0.786).

I present results from my border county model with state fixed effects in Table 4. When comparing treated border counties to untreated border counties within the same state, I find similar null effects for labor supply measures. Estimates are slightly more imprecise with state fixed effects included. I can only rule out large positive effects for labor force status (95% CI = -0.019 to 0.023), employment (95% CI = -0.021 to 0.024), and hours worked (95% CI = -2.033 to 1.770). Estimates for migration measures were not statistically different from zero, and I can rule out modest positive effects for working in a different state (95% CI = -0.030 to 0.016), working in a different compact state (95% CI = -0.014 to 0.011), living in a different state (95% CI = -0.030 to 0.015), and travel time to work (95% CI = -1.819 to 1.680). These null effects are striking, considering that this subset of my sample likely faces the lowest non-licensure related costs to interstate practice and should thus have greater incentive to utilize license portability options that the compact affords.

c. Event-Study

My difference-in-differences analysis requires that I assume that outcome variables in non-compact states exhibit trends that compact states would have followed in the absence of compact adoption. Figures 1 and 2 present estimates from my event-study for labor supply and migration/commuting outcome variables, respectively, to test if compact and non-compact states exhibited similar trends prior to compact adoption. Figure 1 shows that labor supply measures for physicians in compact and non-compact states appear to be trending similarly in the pre-period. Note that trends for my income measure are slightly negative and near zero in the pre-period and display a marked increase following IMLC adoption, providing further evidence that the IMLC put upward pressure on wages. Figure 2 shows additional compelling evidence of parallel trends in the pre-period for migration/commuting measures, although living in a different state appears to have been trending differently prior to compact adoption.

6. Conclusion

As the economic distortions that stem from occupational licensing continue to capture the interest of policymakers, the efficacy of proposed solutions to reduce inefficiencies are of growing importance. Interstate compacts have become a popular policy tool, yet we know very little about their effectiveness. DePasquale & Stange (2016) provided the first empirical analysis of occupational licensing interstate compacts by studying the impacts of the Nurse Licensure Compact. However, since the NLC's development, a large number of professions and occupations have either developed (or are in the process of developing) additional compacts that differ in how they function. This paper seeks to contribute to the literature by evaluating the Interstate Medical Licensure Compact, a policy tool with aims that are similar to those of the NLC but with important functional differences that have gone unevaluated.

Using data on 72,175 physicians over the period of 2012 to 2018, I estimate the impacts of the IMLC on various labor market and interstate mobility measures. Using a difference-in-differences model that compares outcome variables for physicians in compact states to those in non-compact states, I find that IMLC adoption did not have an observable effect on the proportion of physicians in the labor force or employed in compact states. Similarly, I find no impact of the compact on migration or commuting patterns. Nevertheless, IMLC adoption was positively associated with income, consistent with the possibility of improved bargaining power for physicians in compact states. Contrary to my expectation, the evidence from this study suggests that the IMLC's expedited licensure model is not better situated to foster interstate mobility than the mutual recognition model of the NLC.

It should be noted that this study is limited in several ways. First, the treatment I study is still in its nascent stages. It is possible that the compact's effects are unlikely to materialize until at least a few years after adoption. Anecdotal evidence from conversations with the Executive Director of the IMLC, Marschall Smith, suggests rates of applications were relatively tepid until the compact reached its second year of activation. I also cannot rule out the possibility that labor market shocks are biasing my estimates, although I am unable to provide empirical evidence that this is the case, and my results are consistent with previous literature. If my assumptions hold, one possibility is that occupational licensing compacts do not sufficiently streamline the licensing process so as to yield robust effects. Even with the compact in place, licensees are subject to a host of costs that may perpetuate the disincentives of state-level licensing.[13] Another possibility is that physicians may be a particularly immobile profession, and thus relatively impervious to the license portability afforded by the IMLC. However, the null effects of the NLC

[13] See Table A1 in the Appendix for the fees associated with obtaining a license through the IMLC.

on migration measures for nurses diminishes the plausibility of this interpretation, at least for healthcare workers.

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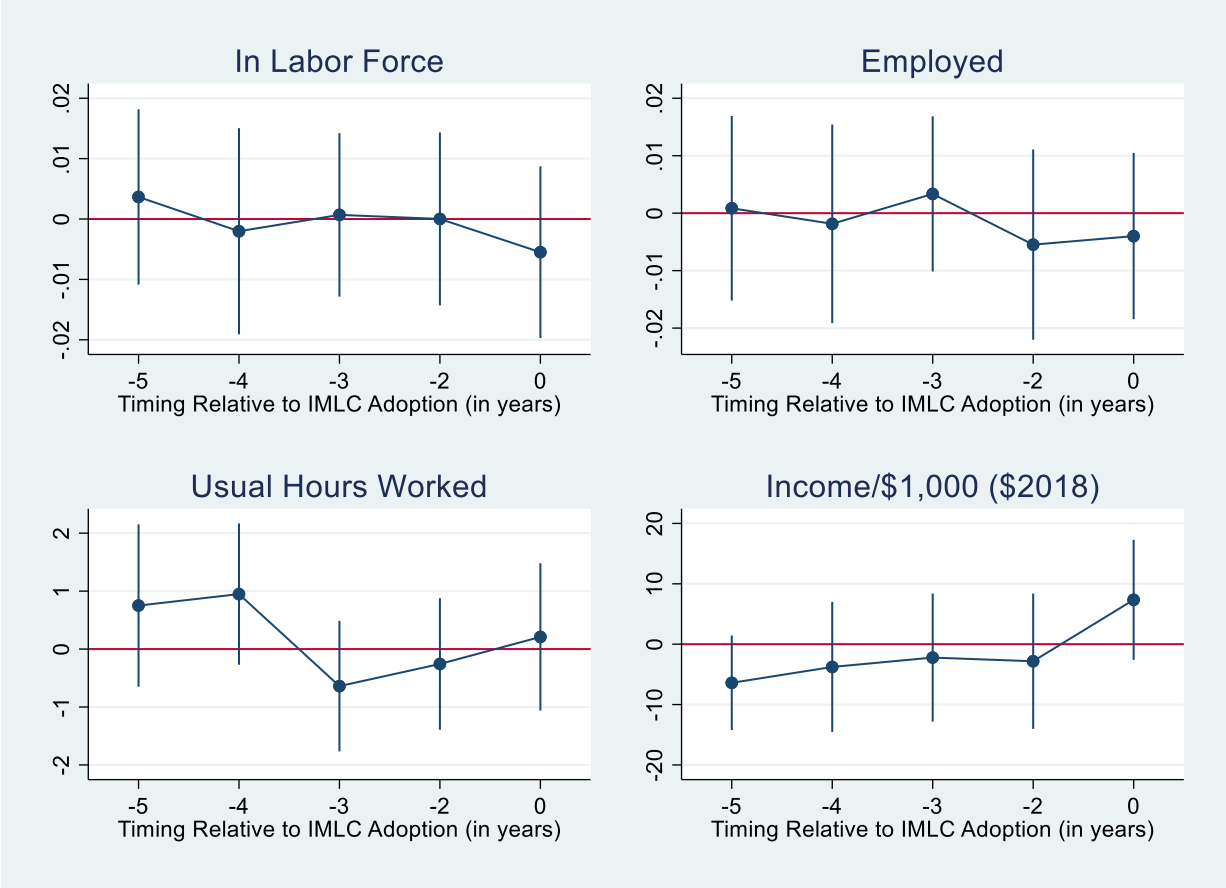
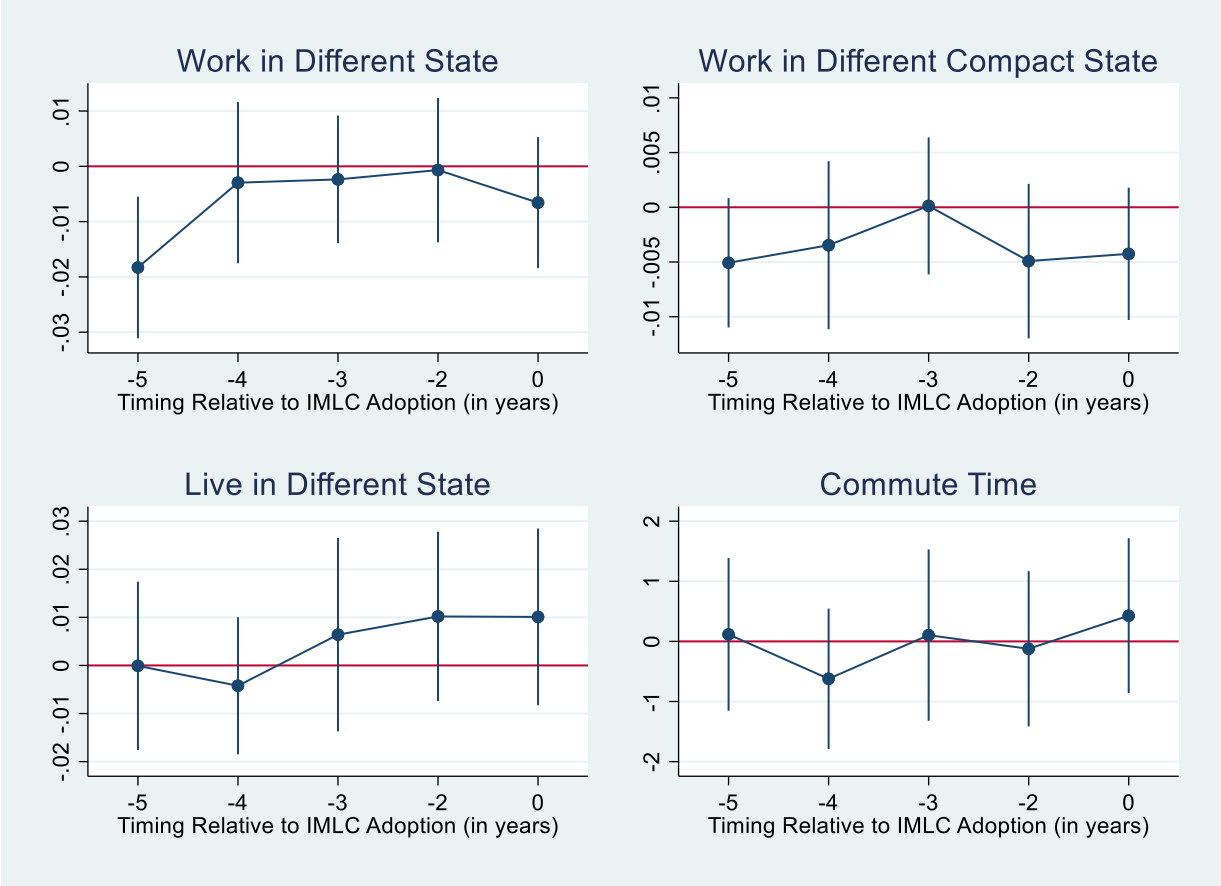
Figures

Figure 1. Event-Study Estimates Leading up to Compact Adoption (with Controls)



Notes: Figure plots coefficients and confidence intervals for estimates for the interaction terms between being in an IMLC state and timing relative to enactment.

Figure 2. Cont'd. Event-Study Estimates Leading up to Compact Adoption (with Controls)



Notes: Figure plots coefficients and confidence intervals for estimates for the interaction terms between being in an IMLC state and timing relative to enactment.

Tables

Table 1. Summary Statistics

	All counties			Border counties only		
	Full Sample	Compact States	Non-compact States	Full Sample	Compact States	Non-compact States
	Mean	Mean	Mean	Mean	Mean	Mean
Number of couples in household	0.84	0.86	0.83	0.82	0.83	0.82
Number of own family members in household	2.86	2.91	2.85	2.83	2.83	2.82
Number of own children in the household	0.93	0.97	0.93	0.92	0.91	0.92
Number of own children under age 5 in household	0.22	0.23	0.22	0.22	0.22	0.22
Age	49.61	50.06	49.58	48.88	49.51	48.84
Usual hours worked per week	47.45	46.32	47.51	48.23	46.57	48.34
Wage and salary income, $2018	194,000	212,000	193,000	192,000	206,000	191,000
Travel time to work	22.01	21.17	22.05	23.16	23.76	23.12
Border county	0.22	0.26	0.21	1	1	1
In labor force	0.93	0.91	0.93	0.94	0.91	0.94
Employed	0.92	0.91	0.92	0.93	0.91	0.93
Work in a different state	0.04	0.04	0.04	0.1	0.07	0.11
Live in a different state	0.05	0.06	0.05	0.05	0.06	0.05
Female	0.35	0.35	0.35	0.37	0.39	0.37
White	0.74	0.79	0.74	0.72	0.71	0.72
Black	0.04	0.03	0.04	0.05	0.04	0.05
American Indian	0	0	0	0	0	0
Chinese	0.04	0.03	0.04	0.04	0.04	0.04
Japanese	0	0	0.01	0	0	0
Other race	0.17	0.15	0.17	0.19	0.21	0.19
College	0.01	0.01	0.01	0.01	0.01	0.01
Postgraduate	0.98	0.99	0.98	0.98	0.99	0.98
American citizen	0.72	0.78	0.72	0.71	0.72	0.71
Naturalized citizen	0.2	0.16	0.2	0.21	0.21	0.21

Non citizen	0.06	0.05	0.06	0.06	0.06	0.06
Bord abroad	0.01	0.01	0.01	0.01	0.01	0.01
Observations	72,175	3,666	68,509	15,687	965	14,722

Note: Sample includes physicians age 16 and over who had worked within the previous five years

Table 2. Regression Estimates - Full Sample

	(1) Labor force status	(2) Employed	(3) Usual hours worked per week	(4) Income	(5) Work in a different state	(6) Work in a different compact state	(7) Live in a different state	(8) Travel time to work
Compact (Treatment)	0.0028	0.0051	0.316	$11,238**	0.0043	0.0018	0.0023	0.3288
	(0.006)	(0.006)	(0.495)	(3478.2)	(0.004)	(0.002)	(0.006)	(0.5)
Constant	1.58***	0.58***	52.73***	-$45,0586***	-0.1003***	-0.0021	0.47***	-4.1
	(0.04)	(0.041)	(4.002)	(32,415.02)	(0.023)	(0.006)	(0.02)	(5.586)
Observations	72,175	72,175	72,175	72,175	72,175	72,175	72,175	72,175

Notes: Standard errors in parentheses. **p<0.01; ***p<0.001

Table 3. Regression Estimates - Border County Sample

	(1) Labor force status	(2) Employed	(3) Usual hours worked per week	(4) Income	(5) Work in a different state	(6) Work in a different compact state	(7) Live in a different state	(8) Travel time to work
Compact (Treatment)	-0.0017	-0.0033	-1.0218	$7,436.07	-0.0480***	0.0030	-0.0048	-1.0147
	(0.010)	(0.010)	(0.941)	(6174.469)	(0.012)	(0.006)	(0.011)	(0.905)
Constant	1.7113***	0.7070***	58.8054***	-$44,0389.71***	-0.08	-0.0177	0.4704***	6.9695
	(0.063)	(0.064)	(9.223)	(42529.206)	(0.072)	(0.013)	(0.041)	(6.071)
Observations	15,687	15,687	15,687	15,687	15,687	15,687	15,687	15,687

Notes: Standard errors in parentheses. ***p<0.001

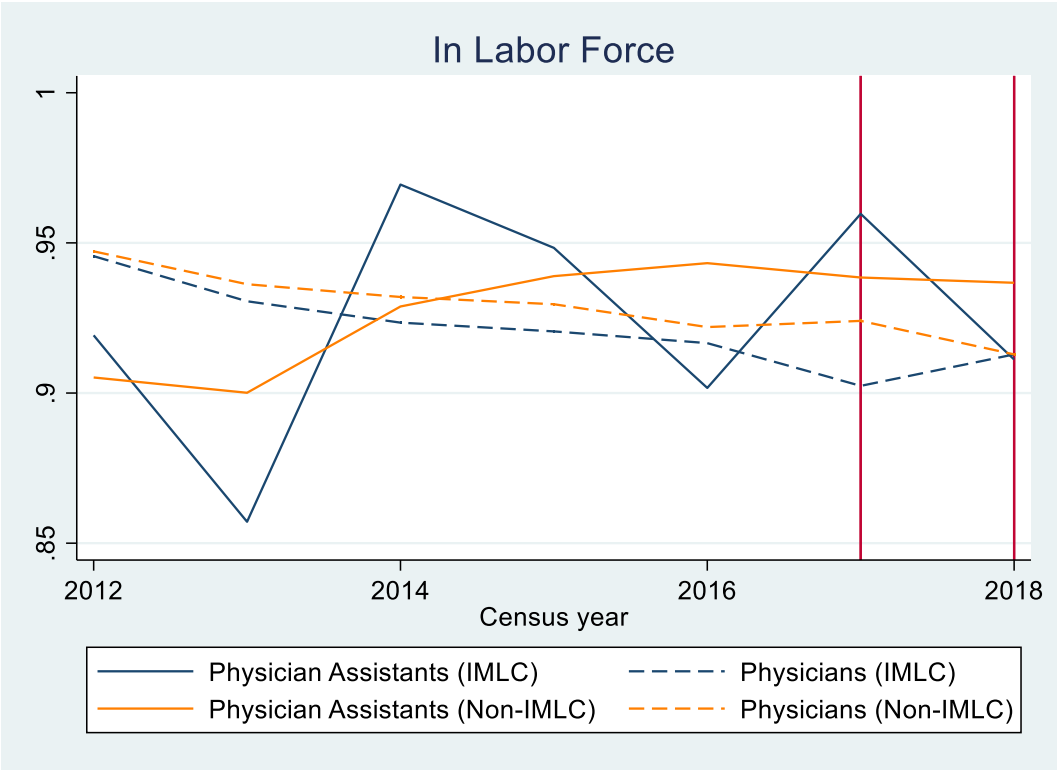
Table 4. Regression Estimates - Border County Sample (with State Fixed Effects)

	(1) Labor force status	(2) Employed	(3) Usual hours worked per week	(4) Income	(5) Work in a different state	(6) Work in a different compact state	(7) Live in a different state	(8) Travel time to work
Compact (Treatment)	0.0020	0.0014	-0.1315	$9,651.32	-0.0071	-0.0013	-0.0074	-0.0697
	(0.011)	(0.011)	(0.955)	(6389.237)	(0.012)	(0.006)	(0.011)	(0.879)
Constant	1.7058***	0.7010***	63.0869***	-$44,0393.48***	-0.2993*	-0.0147	0.4644***	5.5546
	(0.064)	(0.066)	(9.649)	(48275.845)	(0.115)	(0.021)	(0.058)	(6.342)
Observations	15,687	15,687	15,687	15,687	15,687	15,687	15,687	15,687

Notes: Standard errors in parentheses. ***p<0.001; *p<0.05

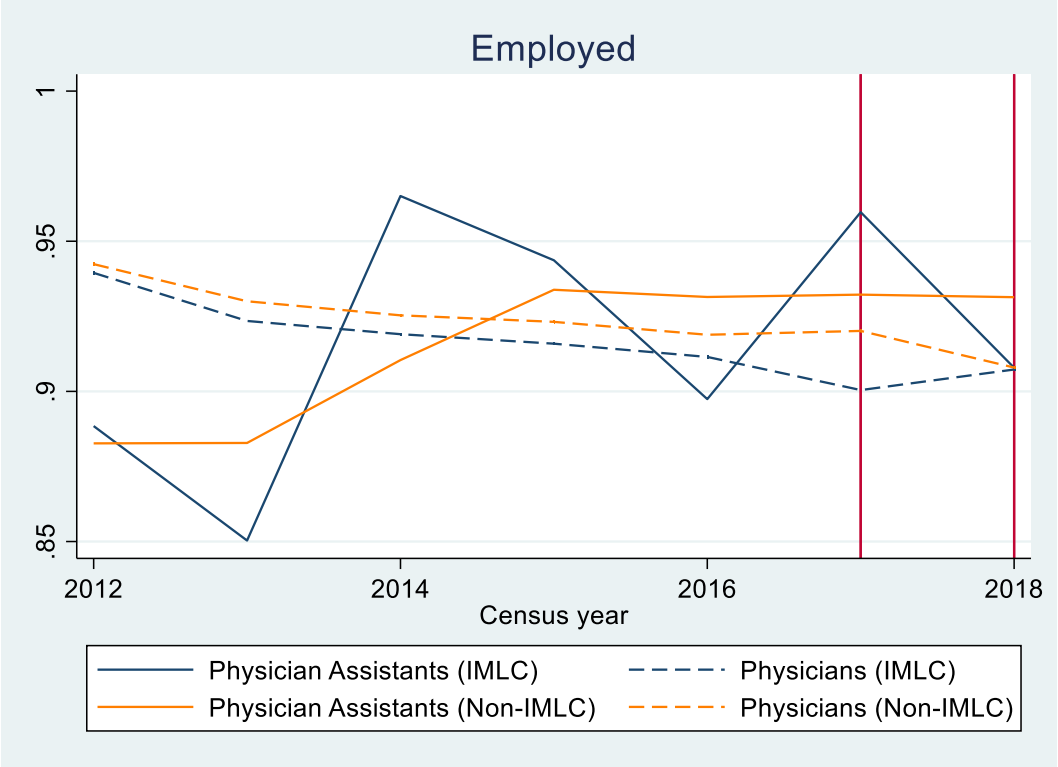
Appendix

Figure A1. Parallel Trends (In Labor Force), Graphical Evidence



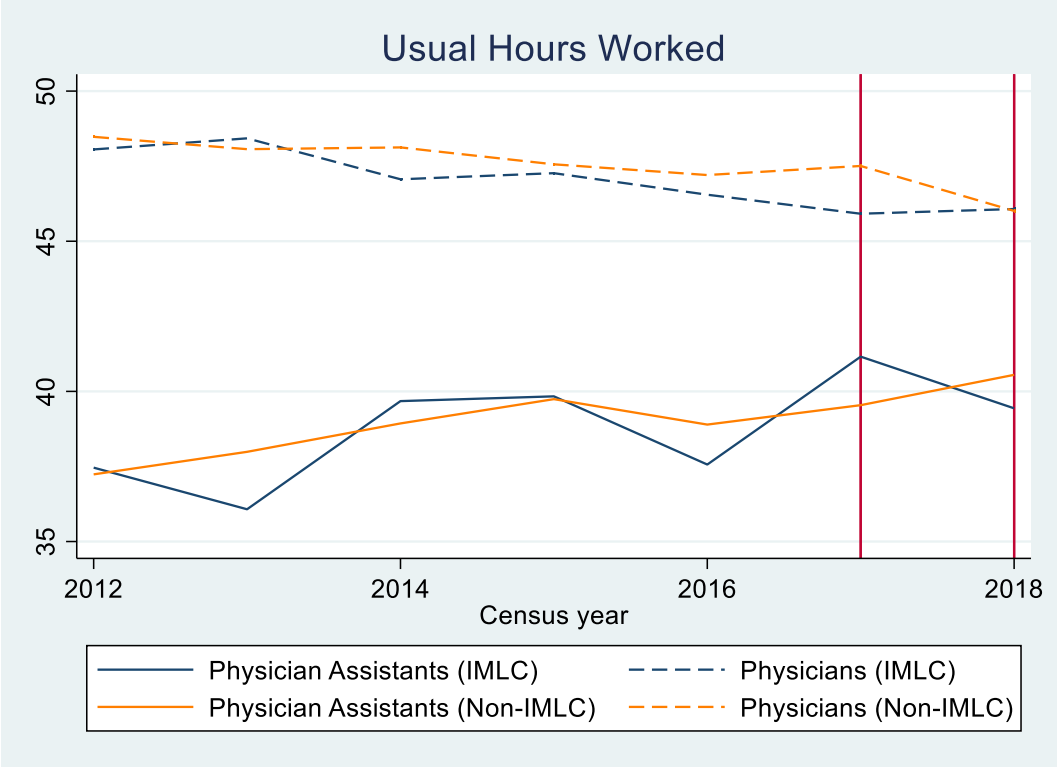
Notes: Graph plots average outcome variable for each year.

Figure A2. Parallel Trends (Employed), Graphical Evidence



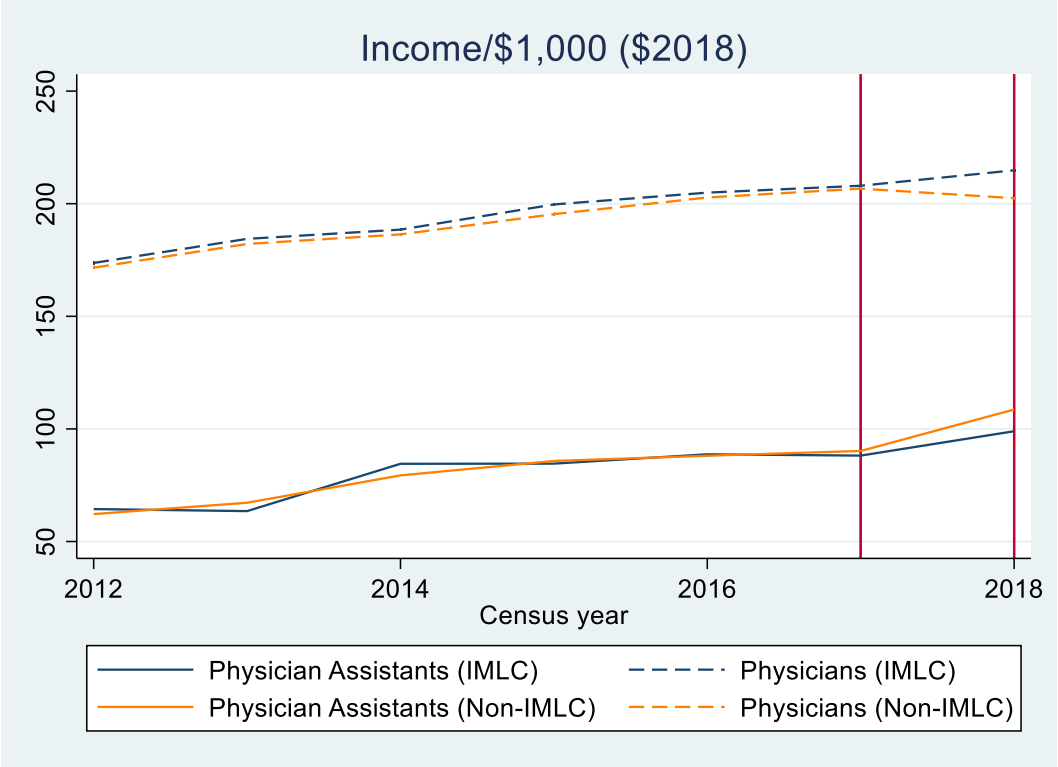
Notes: Graph plots average outcome variable for each year.

Figure A3. Parallel Trends (Usual Hours Worked), Graphical Evidence



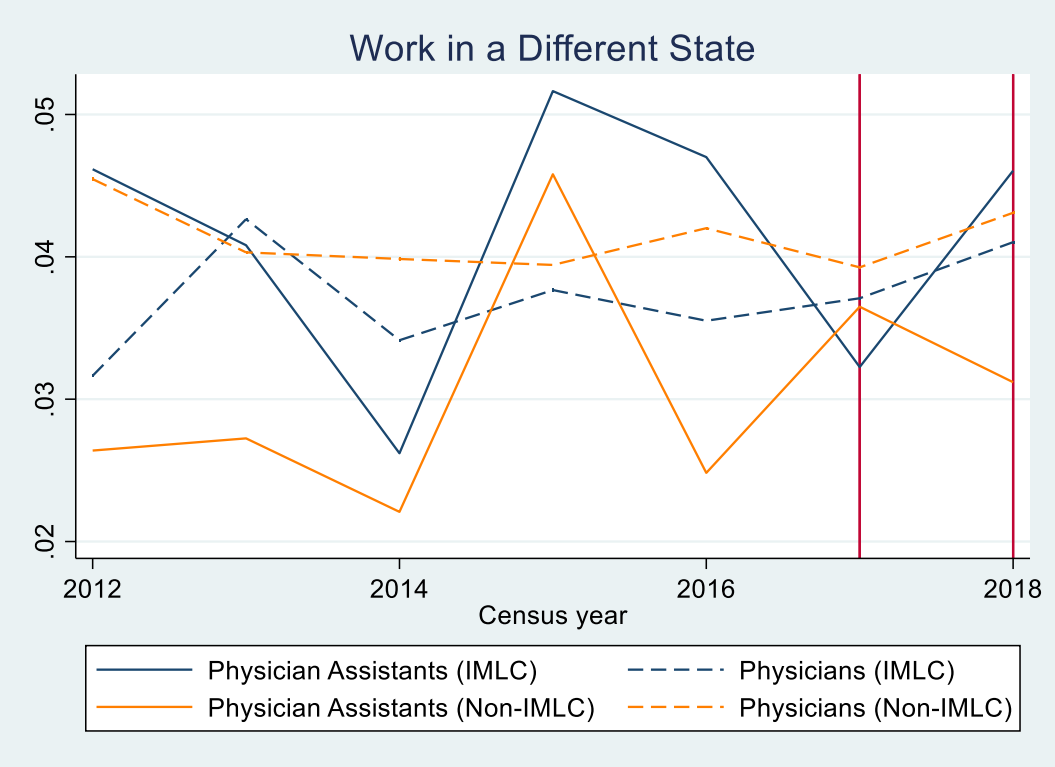
Notes: Graph plots average outcome variable for each year.

Figure A5. Parallel Trends (Income), Graphical Evidence



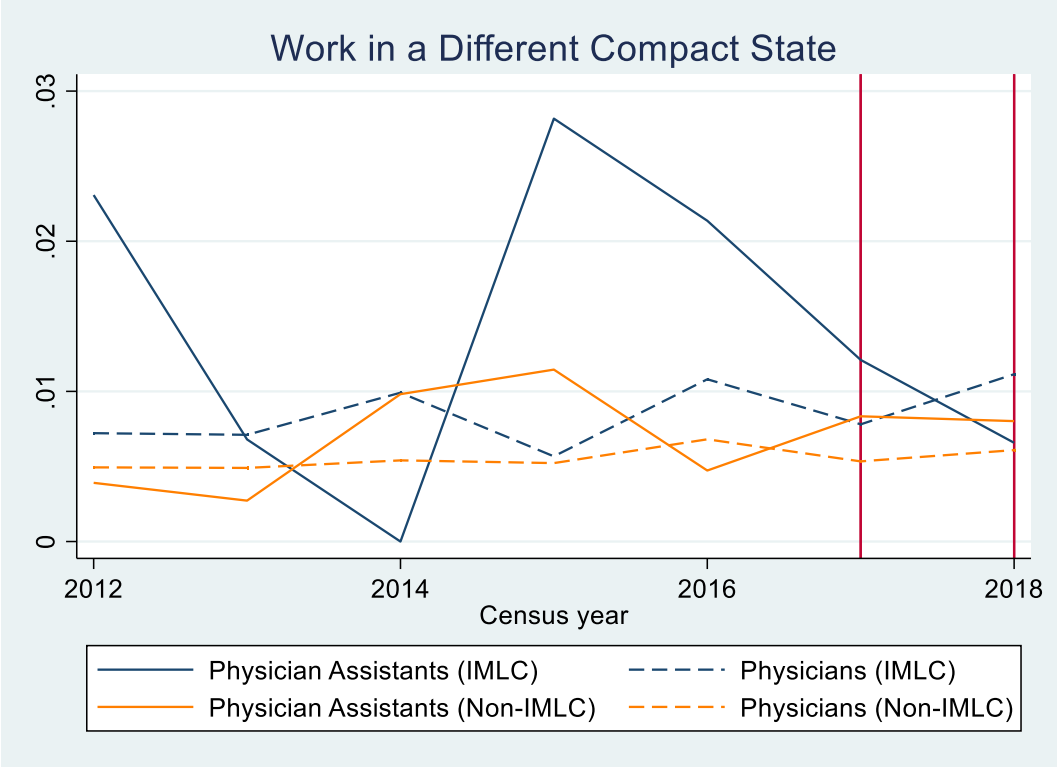
Notes: Graph plots average outcome variable for each year.

Figure A6. Parallel Trends (Work in a Different State), Graphical Evidence



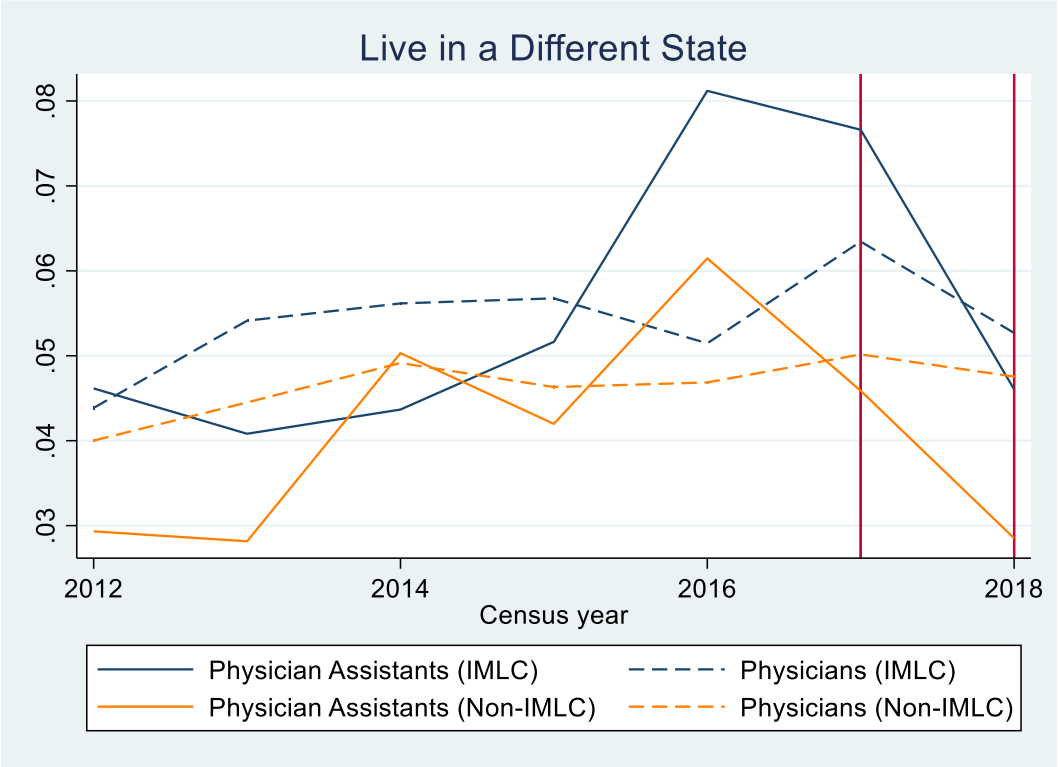
Notes: Graph plots average outcome variable for each year.

Figure A7. Parallel Trends (Work in a Different Compact State), Graphical Evidence



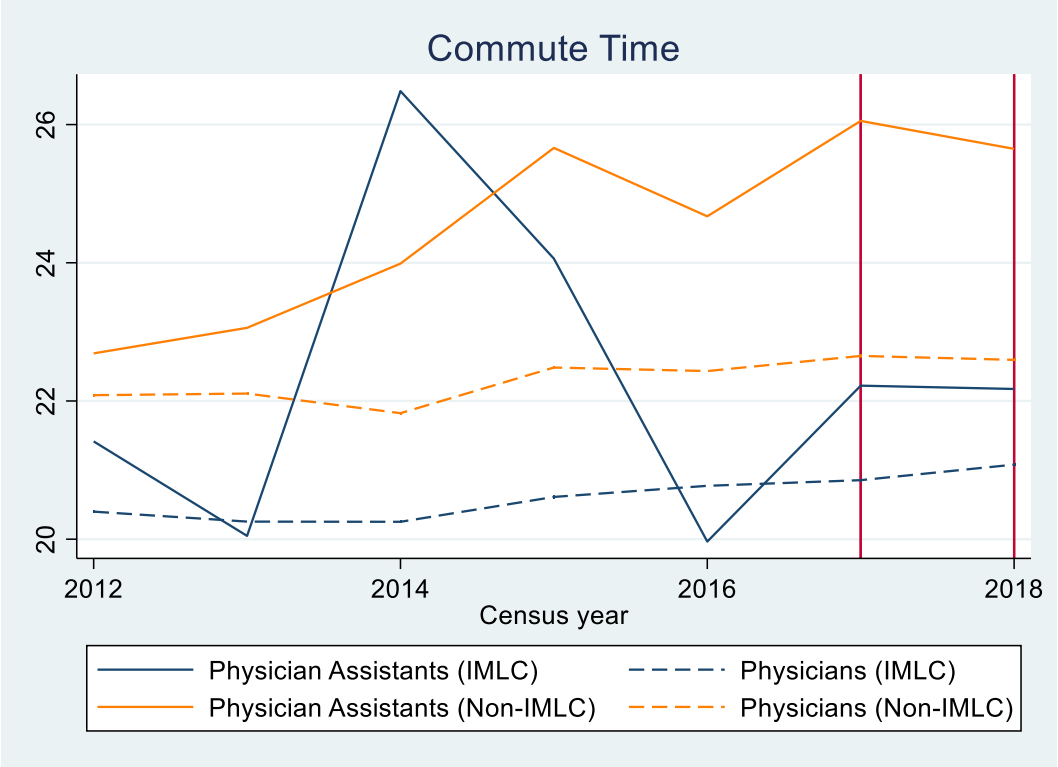
Notes: Graph plots average outcome variable for each year.

Figure A8. Parallel Trends (Live in a Different State), Graphical Evidence



Notes: Graph plots average outcome variable for each year.

Figure A9. Parallel Trends (Commute Time), Graphical Evidence



Notes: Graph plots average outcome variable for each year.

Table A1. IMLC Licensure Costs

State	State Regulatory Authority	Cost	State
ALABAMA	ALABAMA MEDICAL LICENSURE COMMISSION	$75.00	ALABAMA
ARIZONA D.O.	ARIZONA BOARD OF OSTEOPATHIC EXAMINERS	$400.00	ARIZONA D.O.
ARIZONA M.D.	ARIZONA MEDICAL BOARD	$500.00	ARIZONA M.D.
COLORADO	COLORADO MEDICAL BOARD	$390.00	COLORADO
GEORGIA	GEORGIA COMPOSITE MEDICAL BOARD	$500.00	GEORGIA
GUAM	GUAM BOARD OF MEDICAL EXAMINERS	$400.00	GUAM
IDAHO	IDAHO BOARD OF MEDICINE	$251.00	IDAHO
ILLINOIS	ILLINOIS DIVISION OF PROFESSIONAL REGULATION	$500.00	ILLINOIS
IOWA	IOWA BOARD OF MEDICINE	$450.00	IOWA
KANSAS	KANSAS BOARD OF HEALING ARTS	$300.00	KANSAS
MAINE M.D.	MAINE BOARD OF LICENSURE IN MEDICINE	$700.00	MAINE M.D.
MAINE D.O.	MAINE BOARD OF OSTEOPATHIC LICENSURE	$350.00	MAINE D.O.
MARYLAND	MARYLAND BOARD OF PHYSICIANS	$790.00	MARYLAND
MICHIGAN M.D.	MICHIGAN BOARD OF MEDICINE	$361.00	MICHIGAN M.D.
MICHIGAN D.O.	MICHIGAN BOARD OF OSTEOPATHIC MEDICINE AND SURGERY	$361.00	MICHIGAN D.O.
MINNESOTA	MINNESOTA BOARD OF MEDICAL PRACTICE	$392.00	MINNESOTA
MISSISSIPPI	MISSISSIPPI STATE BOARD OF MEDICAL LICENSURE	$600.00	MISSISSIPPI
MONTANA	MONTANA BOARD OF MEDICAL EXAMINERS	$500.00	MONTANA
NEBRASKA	NEBRASKA BOARD OF MEDICINE AND SURGERY	$350.00	NEBRASKA
NEVADA M.D.	NEVADA STATE BOARD OF MEDICAL EXAMINERS	$750.00	NEVADA M.D.
NEVADA D.O.	NEVADA STATE BOARD OF OSTEOPATHIC MEDICINE	$500.00	NEVADA D.O.
NEW HAMPSHIRE	NEW HAMPSHIRE BOARD OF MEDICINE	$300.00	NEW HAMPSHIRE
NORTH DAKOTA	NORTH DAKOTA BOARD OF MEDICINE	$200.00	NORTH DAKOTA
OKLAHOMA M.D.	OKLAHOMA STATE BOARD OF MEDICAL LICENSURE & SUPERVISION	$500.00	OKLAHOMA M.D.
OKLAHOMA D.O.	OKLAHOMA STATE BOARD OF OSTEOPATHIC EXAMINERS	$575.00	OKLAHOMA D.O.
SOUTH DAKOTA	SOUTH DAKOTA BOARD OF MEDICAL AND OSTEOPATHIC EXAMINERS	$400.00	SOUTH DAKOTA
TENNESEE M.D.	TENNESSEE BOARD OF MEDICAL EXAMINERS	$510.00	TENNESEE M.D.
TENNESEE D.O.	TENNESSEE BOARD OF OSTEOPATHIC EXAMINATION	$410.00	TENNESEE D.O.

UTAH D.O.	UTAH OSTEOPATHIC PHYSICIANS & SURGEONS	$200.00	UTAH D.O.
UTAH M.D.	UTAH PHYSICIANS & SURGEONS LICENSING BOARD	$200.00	UTAH M.D.
VERMONT M.D.	VERMONT BOARD OF MEDICAL PRACTICE	$650.00	VERMONT M.D.
VERMONT D.O.	VERMONT BOARD OF OSTEOPATHIC PHYSICIANS	$500.00	VERMONT D.O.
WASHINGTON D.O.	WASHINGTON BOARD OF OSTEOPATHIC PHYSICIAN AND SURGEON	$391.00	WASHINGTON D.O.
WASHINGTON M.D.	WASHINGTON MEDICAL COMMISSION	$491.00	WASHINGTON M.D.
WEST VIRGINIA M.D.	WEST VIRGINIA BOARD OF MEDICINE	$400.00	WEST VIRGINIA M.D.
WEST VIRGINIA D.O.	WEST VIRGINIA BOARD OF OSTEOPATHIC MEDICINE	$100.00	WEST VIRGINIA D.O.
WISCONSIN	WISCONSIN MEDICAL EXAMINING BOARD	$75.00	WISCONSIN
WYOMING	WYOMING BOARD OF MEDICINE	$600.00	WYOMING

Application cost is a $700.00 fee PLUS the cost of a license(s) selected to practice. Source: https://imlcc.org/what-does-it-cost/

Table A2. IMLC Enactment Dates

State	Compact Legislation Effective Date
Alabama	5/19/2015
Arizona	8/9/2016
Colorado	6/8/2016
Idaho	7/1/2015
Illinois	7/20/2015
Iowa	7/2/2015
Kansas	7/1/2016
Maine	6/24/2017
Maryland	7/1/2019
Michigan	3/28/2019
Mississippi	5/16/2016
Montana	10/1/2019
Nebraska	5/23/2017
Nevada	10/1/2015
New Hampshire	7/4/2016
North Dakota	4/9/2019
South Dakota	3/12/2015
Tennessee	1/1/2019
Utah	5/11/2015
Washington	7/23/2017
West Virginia	6/10/2015
Wisconsin	12/16/2015
Wyoming	7/1/2015

Source: https://imlcc.org/