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Doctors without Borders? Relicensing Requirements and Negative Selection in the Market for Physicians

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Relicensing requirements for professionals who move across borders are widespread. In this article, we measure the effects of occupational licensing by exploiting an immigrant physician retraining assignment rule. Instrumental variables and quantile treatment effects estimates indicate large returns to acquiring an occupational license and negative selection into licensing status. We also develop a model of optimal license acquisition that, together with the empirical results, suggests that stricter relicensing requirements may lead not only to practitioner rents but also to lower average quality of service in the market for physicians.

I. Introduction

Restricted entry into an occupation through occupational licensing requirements is a widespread phenomenon. At least 18% of the workforce in the United States is affected by occupational licensing, exceeding both

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minimum wage coverage and membership in unions. Occupational licensing is also widespread in many other countries. In the European Union, for example, occupational entry restrictions are thought to substantially affect incentives for internal migration and labor mobility among member states. Licensing requirements not only restrict entry into professional occupations, such as medicine, dentistry, and law, but also entry into less skilled occupations, such as haircutting and cosmetology (see Kleiner 2000).

In the traditional theory of economic regulation, occupational licensing is generally thought of as an institution that allows practitioners to capture monopoly rents (Friedman and Kuznets 1945; Stigler 1971). Licensing is regarded as a tool used by practitioners to restrict labor supply and drive up the price of labor. More recent theoretical analyses of occupational licensing, however, focus on the conditions under which occupational licensing can be socially beneficial. Licensing may improve the average quality of service offered by practitioners when the entry of less competent practitioners is prevented or when less competent practitioners are forced to increase their investments in human capital (see, e.g., Shapiro 1986). The social loss due to excess wages may be outweighed by the social gains from a higher quality of service.

The theoretical ambiguity over the net social benefits of occupational licensing is accompanied by an inconclusive empirical literature. Empirical studies usually find higher mean earnings for individuals in regulated occupations, holding observed human capital levels constant (see, e.g., Muzondo and Pazderka 1980; Kleiner 2000; and Kleiner and Kudrle 2000). But the data used in these studies generally do not permit identification of the causal effect of entry restrictions on earnings and thus inference on the presence of rents. Entry restrictions and self-selection into the regulated occupation are often confounded. Reliable evidence that licensing improves the average quality of service offered by practitioners is even more rare. The main difficulty in measuring quality effects is in obtaining accurate measures of practitioner quality.[1]

In this article, the effects of occupational licensing are measured using data on the labor market outcomes of immigrant physicians in Israel. The data allow us to identify the returns to an occupational license free of biases due to nonrandom selection into licensing status. Identification of the returns is possible for several reasons. First, the immigrant physicians in the data are a relatively homogeneous group of individuals in terms of

[1] Kleiner and Kudrle (2000) indirectly measure quality effects by comparing the dental health of individuals from different states that vary in their entry requirements for dentists. Angrist and Guryan (2003) examine the effect of teacher certification requirements on teacher quality as measured by educational background. Both studies find small quality effects.

education and experience. Second, many immigrant physicians did not get relicensed and/or obtain employment in their original profession. This is important, since identification of the returns to an occupational license may be difficult for the simple reason that there may be very little variation in licensing status among individuals with similar education and training levels (Pashigian 1980). Third, and perhaps most important, the Israel Ministry of Health assigned immigrant physicians to one of two different relicensing tracks, depending on previous experience. Track assignment is not directly based on unobservables, and the two relicensing tracks are inherently different in terms of the likelihood of acquiring a license. The form of identification in this article is, therefore, unique relative to standard studies of occupational licensing.

The Israel Ministry of Health's assignment rule places immigrant physicians on either a retraining track that requires the passing of a general medical knowledge licensing exam (the exam track) or a retraining track that grants an exemption to the exam and issues a temporary general practitioner license for 6 months (the observation track). The temporary license allows the practice of medicine under the observation of native physicians. At the end of the 6-month period, immigrant physicians on the observation track receive a permanent license with near certainty. Retraining track assignment for almost all of the immigrants in the sample follows a "20-year rule": immigrants with more than 20 years of previous physician experience are assigned to the observation track; those with less than 20 years have to pass the relicensing exam.

It is important to note that the physician relicensing regime in Israel is not unique. In the United States, for example, physicians who migrate across states must also be relicensed, and, in many states, exemption from a relicensing exam is a function of previous physician experience. The 10-year rule of state licensing boards requires migrant physicians who have not passed a national board exam within 10 years to take a Special Purpose Examination (SPEX), which tests their general medical knowledge. Physicians who have passed a national board exam within 10 years are exempt from the SPEX and are granted a license.[2] Since other countries use relicensing regimes that are similar to that used in Israel, especially for immigrants or internal migrants, the results in this study may be quite relevant in other contexts. According to ordinary least squares (OLS) estimates of the returns to acquiring a medical license, immigrants who get relicensed have mean monthly earnings that are between 90% and

[2] The states that require the SPEX for migrant physicians who have not passed a national board test within 10 years or who are not board certified are Alabama, California, Illinois, Louisiana, Maryland, Minnesota, Mississippi, Montana, North Carolina, Nevada, New York, Oregon, South Carolina, and Texas. See http://www.ama-assn.org for details.

114% higher than their unlicensed counterparts. However, OLS estimates are biased to the extent that licensing status is related to potential outcomes without a license. Instrumental variables estimates that exploit the retraining assignment rule and isolate the returns to a license among individuals who would not have obtained a license had they not been assigned to the observation track yield an increase in mean monthly earnings that is much higher than that estimated by OLS. Instrumental variables (IV) estimates of the returns to a license range between 180% and 340%.[3] The large IV estimates, compared to OLS estimates, are suggestive of both the presence of rents and negative selection into licensing status. By negative selection, we mean that immigrant physicians who acquire a license have lower intrinsic earnings potential (in the absence of a license) than those who do not.

In order to give an economic interpretation to the OLS and IV estimates and to address the effects of relicensing on the average quality of service, we also develop a model of optimal license acquisition. The theory is linked to the empirical work by showing what OLS and IV are estimating according to the model. The model implies that when IV estimates exceed OLS estimates, stricter relicensing requirements may lead to lower average quality of service in the licensed occupation. The policy implication of the findings is that lowering the costs to immigrant physicians of acquiring a medical license may raise average physician quality.

As a robustness check on the IV results, the returns to an occupational license are also estimated using a quantile treatment effects (QTE) model (Abadie, Angrist, and Imbens 2002). The QTE estimates also account for the endogeneity of licensing status and are less sensitive than standard IV techniques to the inclusion of high-earnings outliers and zero earnings for the unemployed. The QTE estimates show that the returns to an occupational license are largest at the upper quantiles of the earnings distribution. Conventional quantile regression estimates produce the largest returns at the lower quantiles. The QTE estimates of the counterfactual distribution of earnings without a license indicate negative selection into licensing status at all examined quantiles.

The rest of the article is organized as follows. The next section briefly describes the institutional setting of immigrant physician relicensing in Israel. Section III formulates the model of optimal license acquisition. Section IV describes the data, reports OLS estimates, and presents a graphical analysis of the correlation in the discontinuities in licensing and earnings outcomes that forms the basis of identification. Section V

[3] Note that we estimate the returns to license acquisition as opposed to the returns to physician practice. We focus on license acquisition rather than physician practice because we have only one instrument (track assignment), and the channel from institutional design to license acquisition is more direct.

outlines the estimation strategy. Section VI reports reduced-form, IV, and QTE estimates of the returns to a license. Section VII summarizes and concludes.

II. Background

The recent mass immigration wave to Israel from the former Soviet Union contained an unusually large number of physicians. Between October 1989 and August 1993, approximately 12,500 Soviet physicians arrived in Israel, nearly doubling the potential supply in the market for physicians. Even before the arrival of immigrant physicians, Israel had one of the highest physician-to-population ratios in the world. The number of doctors per 100,000 Israelis in 1989 was 285. In the same year in the United States, there were only 216 doctors per 100,000 Americans. However, the demand for medical services substantially increased in Israel, as the population grew by 10% during the immigration period.

According to Israeli immigration law, physicians who are licensed to practice medicine in a foreign country and who have their foreign medical credentials recognized by the Israel Ministry of Health must pass a relicensing exam in order to legally practice medicine. However, immigrant physicians who practiced clinical medicine and who have substantial previous physician experience are exempt from the relicensing exam. Until November 1992, the cutoff number of years required for exemption from the licensing exam was 20; the cutoff was subsequently lowered to 14. The lowering of the cutoff was abrupt and not publicized beforehand. Approximately 96% of the Soviet immigrant physicians in our data were subject to the 20-year rule rather than the 14-year rule.

Immigrant physicians who are granted an exemption from the licensing exam must work under observation for 6 months in designated public hospitals or community clinics. During the 6-month observation period, immigrants receive a salary and a small stipend from the Ministry of Absorption. At the end of the 6-month period, these immigrants receive, with near certainty, a permanent license in general medicine. Immigrants who are assigned to the exam track cannot switch to the "easier" observation track. A relicensing track assignment is determined by the government and is binding.

Immigrant physicians who are assigned to the exam track are eligible, but not required, to participate in a government-sponsored examination preparation course. Over 90% of immigrants who are referred to the licensing exam choose to participate in a preparation course. Preparation courses last 6 months, are offered twice a year, and are held in public hospitals throughout the country. In order to be accepted into a preparation course, it is necessary to successfully complete a medical terminology language course that also lasts 6 months. Immigrant physicians who participate in the preparation course receive minimal income support from the Ministry of Absorption. A permanent license in general medicine is acquired after passing the exam.

Upon successful completion of the relicensing requirements, all immigrant physicians, independent of previous experience, must request to be recognized as specialists in order to practice medicine in their former specialty. The Ministry of Health denies an overwhelming majority of these requests. Immigrant physicians whose requests are denied must fulfill a postlicensing residency requirement that includes successful completion of two specialty exams. The residency requirement can last a number of years, depending upon medical specialty. The status of specialist is not required for performing rounds in hospitals or treating patients in residential communities as a general practitioner. Only a small percentage of immigrants were in specialist residency at the time of the immigrant physician survey.

III. The Model

The decision to acquire an occupational license can be described within a general model of optimal license acquisition. The model assumes a continuum of workers of skill type η, where η is drawn from a distribution $F(\cdot)$ with support $[\underline{\eta}, \bar{\eta}]$. Individuals live for two periods and have a subjective discount rate r. In the first period, individuals choose whether to invest in acquiring a license or not. In the second period, all individuals work. Acquisition of a license in the first period involves psychological, out-of-pocket, and opportunity costs. The monetary equivalent of psychological costs includes the effort expended in studying for exams and/or meeting other licensing requirements. The out-of-pocket costs include tuition, fees, and commuting expenses. The opportunity costs are the forgone wages the individual could have earned in the unlicensed occupation in the first period. While out-of-pocket and psychological costs are likely to be relatively lower for more skilled (higher η) individuals, the opportunity costs of acquiring a license are likely to be relatively higher. The assumption that these costs vary among individuals of different types generates selection into the licensed occupation.[4]

The out-of-pocket and psychological costs of acquiring a license are specified as C/η, reflecting the assumption that it is relatively easier for more highly skilled individuals to study for the licensing exam and that there is a positive correlation between skill level and pretraining assets. Highly skilled immigrants are likely to be less liquidity constrained. Production functions in the licensed and unlicensed sectors, $Y_L(\eta, N_L)$ and

[4] The license acquisition model developed here is similar to the general model of training participation and earnings described in Heckman, LaLonde, and Smith (1999).

$Y_U(\eta, N_U)$, and the marginal products of labor in both sectors are assumed to be increasing in η, thus generating higher opportunity costs to acquire a license for more highly skilled individuals. It is also assumed that there are diminishing returns with respect to employment, N_L and N_U, in both occupations. The wages in the licensed and unlicensed occupations, w_L and w_U, are defined as the marginal products of labor in the two occupations, respectively. The wages w_L and w_U are, therefore, increasing in η and decreasing in N_i, $i = L, U$. In what follows, the arguments of w_L and w_U will be suppressed for convenience.

Individuals seek to maximize lifetime income by choosing whether or not to acquire a license. Individuals choose to acquire a license and work in the licensed occupation in the second period, rather than work in the unlicensed occupation in both periods, when

$$\frac{w_L - w_U}{(1 + r)} \geq \frac{C}{\eta} + w_U. \tag{1}$$

Equation (1) states that a license will be acquired when the discounted increase in earnings in the second period is greater than or equal to the sum of all costs in the first period.[5]

The decision rule can also be expressed in terms of the maximum C that an individual of type η is willing to incur to acquire a license. This maximum cost component is denoted as $C_{\max}(\eta)$. The maximum cost component $C_{\max}(\eta)$ is found by equating lifetime income in the two occupations,

$$C_{\max}(\eta) = \frac{\eta[w_L - (2 + r)w_U]}{(1 + r)}. \tag{2}$$

An individual of type η chooses to work in the licensed occupation if $C_{\max}(\eta) \geq C$ and chooses to work in the unlicensed occupation if $C_{\max}(\eta) < C$.

The maximum cost component $C_{\max}(\eta)$ can be shown to vary with η in the following way:

$$\frac{\partial C_{\max}(\eta)}{\partial \eta} = \frac{[w_L - (2 + r)w_U]}{(1 + r)} + \frac{\eta[(\partial w_L/\partial \eta) - (2 + r)(\partial w_U/\partial \eta)]}{(1 + r)}. \tag{3}$$

Note that each of the two terms after the equals sign in equation (3) can be positive or negative, even if $w_L > w_U$ and $\partial w_L/\partial \eta > \partial w_U/\partial \eta$ at every η. If $\partial C_{\max}(\eta)/\partial \eta > 0$, then highly skilled individuals are willing to pay more for a license, and there is positive selection into licensing status. In

[5] Uncertainty both in successfully meeting licensing requirements or in obtaining a job in the licensed occupation after completing the licensing requirements can be thought of as factors that reduce the discounted wage premium. Adding uncertainty to the model does not change anything of substance.

this case, individuals with $\eta \in [\underline{\eta}, \eta^*]$ work in the unlicensed occupation, and individuals with $\eta \in [\eta^*, \bar{\eta}]$ work in the licensed occupation. The variable η^* is the point in the skill distribution where $C_{\max}(\eta) = C$. If $\partial C_{\max}(\eta)/\partial\eta < 0$, then highly skilled individuals are willing to pay less for a license, and there is negative selection into licensing status. In this latter case, individuals with $\eta \in [\underline{\eta}, \eta^*]$ work in the licensed occupation, and individuals with $\eta \in [\eta^*, \bar{\eta}]$ work in the unlicensed occupation.[6]

The theory outlined above can be linked to the empirical work by noting that OLS yields a regression-adjusted estimate of

$$\begin{aligned} E[w_L|C_{\max}(\eta) > C] - E[w_U|C_{\max}(\eta) < C] \\ = E[w_L - w_U|C_{\max}(\eta) > C] \\ + \{E[w_U|C_{\max}(\eta) > C] - E[w_U|C_{\max}(\eta) < C]\}, \end{aligned} \tag{4}$$

where $E[w_L|C_{\max}(\eta) > C]$ and $E[w_U|C_{\max}(\eta) < C]$ are the mean wages, according to the model, in the licensed and unlicensed occupations, respectively. The first term after the equals sign in equation (4) is the effect of treatment on the treated, and the second term is selection bias. Ordinary least squares clearly does not identify the causal effect of acquiring a license.

In contrast to OLS, IV yields a regression-adjusted estimate of the local average treatment effect (LATE) (Angrist, Imbens, and Rubin 1996), which is a causal expression. In this case, LATE is

$$E[w_L - w_U|C_1 > C_{\max}(\eta) > C_0], \tag{5}$$

where C_0 is the lower C incurred when assigned to the observation track and C_1 is the higher C incurred when assigned to the exam track. Assuming that LATE is a good approximation to the effect of treatment on the treated, $E[w_L - w_U|C_1 > C_{\max}(\eta) > C_0] \approx E[w_L - w_U|C_{\max}(\eta) > C]$; the difference between OLS and IV is the selection bias. If IV exceeds OLS, selection bias is negative, and, according to the model, the least-skilled (lowest η, lowest earnings potential) immigrant physicians choose to become licensed physicians in the host country.

IV. The Data

The population of immigrant physicians in this study consists of immigrants who arrived in Israel from the former USSR between October 1989 and June 1992, submitted a request to the Israel Ministry of Health to start the process toward relicensing, and had recognized medical credentials in the former USSR. Of the immigrants who declared at the

[6] A fuller development of the model can be found in an earlier version of the article posted on the Social Science Research Network (SSRN) at http://papers.ssrn.com.

airport on the day of arrival that they were physicians in the former USSR, 27% did not submit their credentials to the Ministry of Health. Of the immigrants who submitted their credentials, 3% did not have their credentials recognized. Among immigrants who had their medical credentials recognized, 3% were not referred to one of the two retraining tracks. These latter immigrants were either required to complete a 1-year internship before being eligible for the exam track or were immediately granted recognition as specialists. The total number of immigrant physicians in this restricted population is 6,754.[7]

Between the months of May and November of 1994, 731 immigrant physicians, who were confirmed by the Israel Ministry of Health to have been physicians in the former USSR, were surveyed in face-to-face interviews by Russian-speaking enumerators, using a questionnaire written in Russian. The survey was conducted under the auspices of the JDC-Brookdale Institute of Jerusalem, the Israel Ministry of Health, and the Israel Ministry of Immigrant Absorption. The random sample of 731 immigrant physicians was stratified by assigned retraining track and geographical region. The goal was to interview 10% of the restricted population. A reserve list of immigrants was prepared, according to the same stratification rules, to substitute for those on the original list who could not be interviewed. In total, 1,002 immigrant physicians were approached for interviewing. In descending order of importance, those on the original list who were not interviewed either were not located, refused to be interviewed, return migrated, or had passed away.

A. Descriptive Statistics

Table 1 displays selected descriptive statistics for the sample by assigned retraining track. Of the 731 immigrant physicians in the sample, only two immigrants did not have a retraining track coded. Of the 414 immigrant physicians assigned to the exam track, according to either the 14-year rule or the 20-year rule, 73% passed the relicensing exam. Immigrants who were assigned to the exam track and did not acquire a license either never took the exam or took the exam and failed. Of the 315 immigrant physicians assigned to the observation track, according to either the 14-year rule or the 20-year rule, 89% worked under observation and acquired a permanent license. The 11% among this latter group who are coded as not having acquired a permanent license reported that they never looked for a place to begin work under observation.

Overall, 80% of the immigrant physicians in the sample acquired a license, but there is a large 16-percentage-point difference between the

[7] Nonsubmitters to the Ministry of Health are much more likely than submitters to be over 55 years of age on arrival. Immigrants who did not have their credentials recognized are younger than those who had their credentials recognized.

Table 1
Descriptive Statistics

Variable	Exam Track	Observation Track
% Licensed	72.71	88.57
% Employed	86.23	65.40
% Physician	58.70	37.14
Monthly earnings (NIS)	2,552	1,703
	(1,811)	(2,011)
Months in Israel	44.31	42.68
	(6.24)	(7.45)
Age on arrival	34.54	53.06
	(5.00)	(7.36)
Previous physician experience	10.32	28.18
	(4.82)	(7.57)
% Male	44.44	44.13
% Married on arrival	84.30	79.36
No. children under age 18 on arrival	1.23	.59
	(.75)	(.76)
% from Russia	46.14	41.59
% from Ukraine	16.67	23.81
% from city > 1,000,000	52.17	53.33
% Advanced medical degree	26.81	25.4
% Former specialist	40.34	85.08
% Former general practitioner	22.95	18.73
% Former pediatrician	16.18	12.70
% Former obstetrics/gynecology	7.49	5.71
% Arrived in 1991–92	77.3	67.62
N	414	315

NOTE.—The table reports means and percentages by assigned retraining track. Standard deviations are in parentheses. Monthly earnings are in 1994 New Israeli Shekels (NIS), where 1 NIS equals US$.33. There are 382 exam track earnings observations and 294 observation track earnings observations (including zeros for the unemployed).

proportion acquiring it on the exam track and the proportion acquiring it on the observation track. The most obvious difference between individuals on the two different tracks is age at time of arrival. Individuals assigned to the exam track are, on average, 18 years younger, have 18 years less physician experience in the former USSR, and have more children under the age of 18 living at home at the time of arrival.

Table 1 also shows that mean monthly earnings (including zeros for the unemployed), the employment rate, and the rate of employment as a physician at the time of the survey are higher among immigrants assigned to the exam track. Note that a considerable proportion of immigrant physicians on the observation track are not employed as physicians. Immigrants who are employed but not working as physicians are, for the most part, working in skilled occupations as post–secondary education teachers, social workers, qualified nurses, optometrists, medical technicians, and paramedics. There is only a small proportion of immigrant physicians working in less skilled occupations, such as unqualified nurse-

Table 2
OLS Estimates of the Returns to a Medical License

	Full Sample				Discontinuity Sample	
Regressors	(1)	(2)	(3)	(4)	(5)	(6)
Licensed	1,279	1,211	1,162	−175	1,254	169
	(140)	(140)	(141)	(138)	(373)	(358)
% Impact	1.095	.979	.904	−.067	1.137	.077
Physician	. . .	. . .	. . .	2,324	. . .	1,885
				(158)		(301)
Experience	. . .	. . .	4.90	56.21	1,111	928
			(25.1)	(23.24)	(547)	(462)
Experience^2	. . .	. . .	−1.35	−2.05	−27.15	−22.80
			(.61)	(.58)	(13.93)	(11.65)
Other regressors	No	Yes	Yes	Yes	Yes	Yes
RMSE	1,876	1,672	1,648	1,390	1,631	1,447
R^2	.0711	.2847	.3073	.5086	.3227	.4702
N	676	676	676	676	181	181

NOTE.—Robust standard errors are in parentheses. RMSE = root mean squared error. Other regressors include dummies for age on arrival, year of arrival, months in Israel, gender, marital status, profession of spouse, number of children, size of last city of residence, republic of origin, advanced medical degrees, previous specialist status, previous type of medical practice, and type of reported earnings (after tax and/or after other deductions). The discontinuity sample in cols. 5 and 6 uses the subsample of observations between 14 and 26 years of previous physician experience.

maids, cleaners in institutions, security guards, and skilled and unskilled workers in industry.[8]

In terms of gender composition, size of last city of residence in the former USSR (more than 1,000,000 inhabitants), continuation of studies in the former USSR toward an advanced medical degree, and the number of months since arrival, the immigrants are quite similar by retraining track. There are only slight differences in marital status on arrival, republic of origin, and type of medical practice in the former USSR. Note that over 95% of the immigrants in the sample arrived at the start of the immigration wave, during the years 1990–91. Overall, the similarities in immigrant characteristics by assigned retraining track, except for characteristics related to age, constitute strong evidence that retraining track assignment was indeed mainly a function of previous experience. The differences in former specialist status by track arise because specialist status is also a function of age in the former Soviet Union.

B. OLS Estimates

OLS estimates of the increase in mean monthly earnings due to acquisition of a license are reported in table 2. We concentrate on monthly earnings because there is little overall variation in hours worked among

[8] Among immigrant physicians employed as physicians, 41% work for the government (local and national). The remainder work for health maintenance organizations and other private employers. Only 6% found physician work as a direct continuation of retraining.

these immigrants, and hours of work are poorly measured. Column 1 does not include any other covariates and yields a precisely estimated coefficient on licensed of 1,279 New Israeli Shekel (NIS).[9] In 1994, the year in which earnings are reported, 1 NIS = approximately US$.33. The estimated increase in earnings of 1,279 NIS corresponds to a percentage impact of 109%.[10]

Column 2 adds covariates to the regression but excludes previous physician experience. The other covariates include dummies for age on arrival, year of arrival, months in Israel, gender, marital status, profession of spouse, number of children under age 18 living at home on arrival, size of last city of residence, republic of origin, advanced medical degrees, previous specialist status, previous type of medical practice, and type of reported earnings (after tax and/or after other deductions). The coefficient on licensed in this latter regression is a precisely estimated 1,211, which corresponds to a percentage impact of 98%.

Column 3 adds years of physician experience in the USSR and its square. The coefficient on licensed further decreases to 1,162 but is still precisely estimated. The percentage impact is 90%. Column 4 adds an indicator for being employed as a physician. The physician dummy is an interaction between having acquired a license and being employed as a physician. The coefficient on licensed turns negative, is quite small in magnitude, and is not precisely estimated. The coefficient on the physician dummy, however, is substantial. The OLS results do not indicate a significant return to a license when not employed as a practicing physician. Note that a significant coefficient on licensed in this latter specification would have been suggestive of a signaling value to license acquisition.

Columns 5 and 6 repeat the specifications in columns 3 and 4 for the subsample of immigrants who have previous physician experience between 14 and 26 years. This subsample of immigrants defines the discontinuity sample. Restricting the analysis to the discontinuity sample helps control for differences in unobservables between immigrants of different ages (e.g., differences in the quality of medical education in the former USSR). The discontinuity sample also has a relatively greater degree of variation in assigned retraining track than the full sample.[11]

The estimated coefficient on licensed in the discontinuity sample is

[9] All standard errors are heteroskedasticity robust.

[10] Monthly earnings include zeros for the unemployed. The percentage impact is calculated as the ratio of the coefficient on licensed to the fitted value from the regression with the licensed dummy set to zero and other covariates set to the means among individuals with a license, when other covariates are included.

[11] The experience levels 14 and 26 are the 45th and 75th percentiles, respectively, in the previous physician experience distribution. The experience distribution is skewed to the right with a mean of 16, a standard deviation of 11, and a median of 18.

larger than in the corresponding specification in the full sample. With a quadratic in previous physician experience and no indicator for physician employment, the coefficient on licensed is a precisely estimated 1,254. The percentage impact is 114%. Column 6 reports the results with the physician employment indicator added. The estimated coefficient on licensed is positive but is negligible in magnitude and imprecisely estimated.

Additional OLS results that are not reported in table 2, for the sake of brevity, indicate that within the retraining track the increase in mean monthly earnings due to acquisition of a license is similarly large in magnitude. The percentage impact on earnings for immigrant physicians assigned to the exam track is 99%, with other covariates and previous physician experience and its square included in the regression. The corresponding percentage impact on earnings among physicians assigned to the observation track is somewhat smaller but not significantly different. In addition, the effect of license acquisition on employment probabilities is similar to the effect of track assignment on physician employment probabilities, which is explicitly considered below.

C. Graphical Discontinuity Analysis

The OLS estimates presented in the previous subsection do not exploit the Ministry of Health's retraining assignment rule. The assignment rule can be used to construct instrumental variables estimates of the returns to an occupational license even though the assignment rule is a near deterministic function of years of previous physician experience, and previous physician experience directly affects earnings in the host country. Identification relies on matching discontinuities (or nonlinearities) in the relationship between previous physician experience and licensing status and discontinuities in the relationship between previous physician experience and earnings. The correlation between these discontinuities identifies the causal effect of acquiring a license on earnings as long as it is the assignment rule and not some other mechanism that is generating the discontinuities in licensing outcomes (see Angrist and Krueger 1999).

The discontinuities in licensing and labor market outcomes resulting from the retraining assignment rule are illustrated graphically in figures 1–3. The figures are smoothed by estimating weighted local linear regressions that do not straddle the break points of 14 and 20 years of previous physician experience (see Hahn, Todd, and van der Klaauw 2001). Figure 1 plots the proportion of immigrant physicians assigned to the observation track along with the residuals from a linear probability model that has acquisition of a license as the dependent variable and that excludes previous physician experience as a covariate.

Note that in figure 1 the proportion assigned to the observation track is zero in the range 0–13 years of experience. From 14 to 19 years of

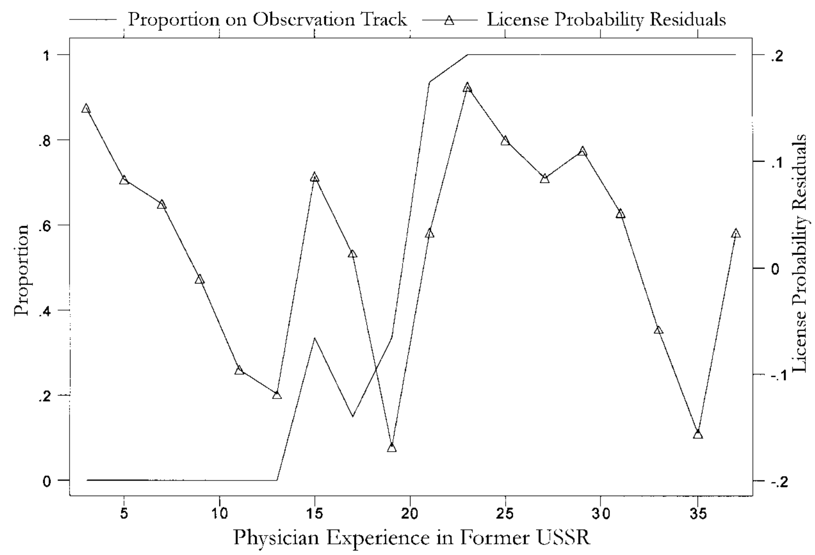

FIG. 1.—Track assignment and license acquisition

experience, the proportion fluctuates between 12% and 33%. This is due to the small number (32 out of 729 immigrants) assigned to the observation track according to the later instituted 14-year rule. At 20 years of experience, the proportion sharply jumps up. At 20 years of experience and more, the proportion fluctuates between 92% and 100%. Immigrant physicians in this latter experience group who are not assigned to the observation track did not previously practice clinical medicine.

Figure 1 also shows that the license probability residuals jump up together with jumps in the proportion assigned to the observation track. License probability residuals remain high after 20 years of experience but drop off rather sharply after 30 years of experience. Immigrant physicians with 30 or more years of experience are, on average, 60 years old. Clearly, the incentive to invest in license acquisition is much weaker the closer one is to retirement age.

Figure 2 plots license acquisition residuals and physician employment residuals from separate linear probability models that exclude previous physician experience. The figure shows that employment as a physician follows licensing status quite closely. Figure 3 examines the relationship between monthly earnings residuals and license acquisition residuals. The matching of the discontinuity at 20 years of experience and the change in trend in earnings after 20 years of experience is quite evident.

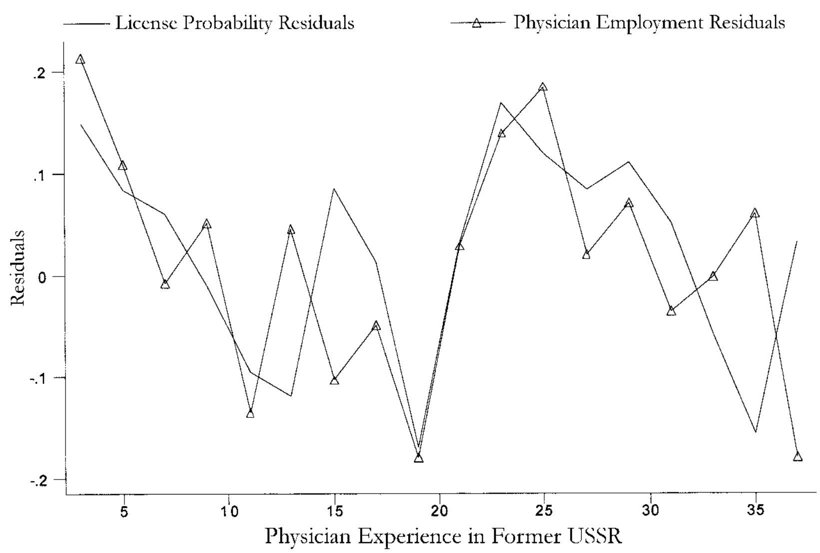

FIG. 2.—License acquisition and physical employment outcomes

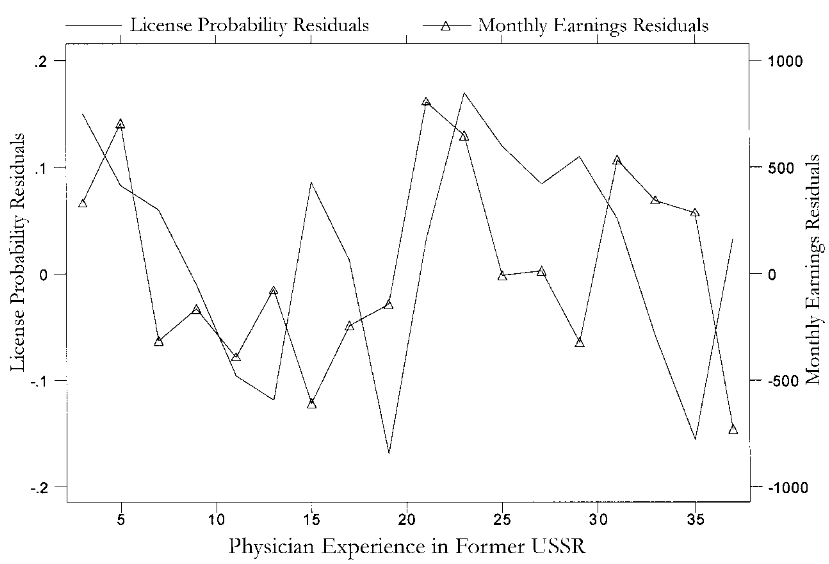

FIG. 3.—License acquisition and monthly earnings

V. Estimation Strategy

The estimation framework that we consider is a linear, constant-effects model that connects the earnings of immigrant i at time t, Y_{it}, with the occupational licensing status of individual i at time t, L_{it}, plus a vector X_i of immigrant characteristics at the time of arrival, and a random error component specific to individuals at time t, ϵ_{it}:

$$Y_{it} = X_i'\beta + t\delta + L_{it}\alpha + \epsilon_{it}. \tag{6}$$

Time in Israel, t, is, like the elements of X_i, widely believed to be exogenous to potential labor market outcomes in Israel among those immigrants who arrived in the first 3 years of the immigration wave. The variable t generically captures changes in language ability, social networks, and knowledge of local institutions, and is also correlated with the size of the last city of residence in the former USSR (included in X_i). Immigrants who arrived earlier generally came from big cities in which there was greater access to information, government offices, and consulates. Measures of English and Hebrew ability at the time of the survey and host country work experience are endogenous (and not included in X_i) but are also strongly correlated with t. In empirical implementation, t is represented by a series of 6-month dummy variables (e.g., a dummy for immigrants who have been in Israel between 1 and 6 months, a dummy for those who have been in Israel between 7 and 12 months, etc.).

The interpretation of equation (6) is that it describes the earnings of immigrants under alternative assignments of licensing status, controlling for any effects of X_i and t. However, since L_{it} is not randomly assigned and is likely to be correlated with potential earnings, ϵ_{it}, OLS estimates of equation (6) do not have a causal interpretation. Instrumental variables estimates of equation (6) do have a causal interpretation as long as it is reasonable to assume that, after controlling for X_i and t, the association between assignment to the observation track and monthly earnings is solely due to the association between observation track assignment and licensing status.

In IV estimation, the first-stage relationship between licensing status, assignment to the observation track, X_i, and t is

$$L_{it} = X_i'\pi_0 + t\pi_1 + TR_i\pi_2 + \xi_{it}, \tag{7}$$

where $TR_i = 1$ indicates assignment to the observation track and $TR_i = 0$ indicates assignment to the exam track. The error term ξ_{it} is defined as the residual from the population regression of L_{it} on X_i and t and the instrument, TR_i. This residual captures other factors that are correlated with licensing status and may be correlated with ϵ_{it}.

The key identifying assumption that underlies estimation using TR_i as an instrument is that any effects of previous physician experience on monthly earnings in Israel are adequately controlled by the smooth functions of previous physician experience included in $X_i'\beta$ and partialed out of TR_i by the inclusion of smooth functions of previous physician experience in $X_i'\pi_0$. If this assumption is reasonable, the discontinuities in earnings with previous physician experience, as depicted in the graphical analysis, can be attributed to the acquisition of an occupational license.[12]

The same discussion above carries through for measuring the effect of working as a physician in place of the effect of acquiring a license. However, since we have only one instrument, we focus on the results that use license acquisition, rather than physician practice, as the sole endogenous variable. We believe that focusing on the licensing outcome is more natural because the channel from institutional design to license acquisition is more direct. Government selection of a retraining track for an immigrant physician, which is not directly based on immigrant unobservables, more closely corresponds to random assignment of a license as opposed to random assignment to a physician job.

VI. Estimation Results

A. Reduced-Form Estimates

Reduced-form estimates of the effect of being assigned to the observation track are reported in table 3. Columns 1–3 of table 3 report coefficients for the first stage described in equation (7), that is, the effect of being assigned to the observation track on licensing status. In all three linear probability models, assignment to the observation track increases the probability of acquiring a license. The estimated coefficient on track, without any other regressors, is .159. Including other regressors increases the estimated coefficient to .231. Adding previous physician experience and its square further increases the estimated coefficient on track to .338. Column 4 reports the same specification as in column 3 in the discontinuity sample. The coefficient on track in this latter specification is .258. In all four specifications, the coefficient on track is precisely estimated.

Columns 5–7 of table 3 report the effect of being assigned to the observation track on the probability of employment as a physician in linear probability models. Adding previous physician experience yields a statistically significant coefficient of .234. The specification with previous physician experience in the discontinuity sample only, reported in column 8, produces a statistically significant coefficient of .258. The first-stage relationship between physician practice and track assignment is slightly weaker than the first-stage relationship between license acquisition and track assignment. The physician employment regressions in columns 6–8

[12] Card (1999) surveys evidence supporting the smoothness assumption in the relationship between experience and earnings.

Table 3
Reduced Form Estimates of the Effect of Track on Monthly Earnings, License Acquisition, and Physician Employment

	Licensed				Physician Employment				Monthly Earnings			
	Full Sample			Discontinuity Sample	Full Sample			Discontinuity Sample	Full Sample			Discontinuity Sample
Regressors	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Track	.1587	.2305	.3383	.2853	−.2155	.0206	.2336	.2576	−849	42	571	628
	(.0283)	(.0308)	(.0490)	(.0679)	(.0365)	(.0584)	(.0754)	(.1230)	(149)	(207)	(255)	(371)
% Impact	. . .	. . .	. . .	. . .	. . .	. . .	. . .	. . .	.3325	.0094	.1409	.2262
Experience	. . .	. . .	−.0053	−.2198	. . .	. . .	−.0222	−.1889	. . .	. . .	−14.12	765
			(.0075)	(.1009)			(.0077)	(.1525)			(28.01)	(530)
Experience^2	. . .	. . .	−.0002	.0052	. . .	. . .	.0001	.0038	. . .	. . .	−1.38	−19.49
			(.0002)	(.0025)			(.0001)	(.0038)			(.62)	(13.52)
Other regressors	No	Yes	Yes	Yes	No	Yes	Yes	Yes	No	Yes	Yes	Yes
RMSE	.3961	.3908	.3854	.3269	.4891	.4621	.4551	.4579	1,900	1,740	1,705	1,675
R^2	.0380	.0879	.1153	.2260	.0456	.1716	.1987	.2605	.0469	.2257	.2589	.2857
N	729	729	729	203	729	729	729	203	676	676	676	181

Note.—Robust standard errors are in parentheses. RMSE = root mean squared error. Other regressors include dummies for age of arrival, year of arrival, months in Israel, gender, marital status, profession of spouse, number of children, size of last city of residence, republic of origin, advanced medical degrees, previous specialist status, previous type of medical practice, and type of reported earnings (after tax and/or after other deductions). The discontinuity sample in cols. 4, 8, and 12 uses the subsample of observations between 14 and 26 years of previous physician experience. The licensed and physician employment regressions are linear probability models. The physician employment regressions include interaction terms between experience and year of arrival.

also require interaction terms between previous experience and year of arrival that have no explanatory power in columns 2–4.

Columns 9–11 of table 3 report reduced-form effects of being assigned to the observation track on mean monthly earnings. Without controls for other regressors or previous physician experience, there is a negative association between assignment to the observation track and monthly earnings. The track variable is picking up a downward trend in mean earnings with previous physician experience. The association between being assigned to the observation track and monthly earnings becomes significantly positive with the addition of other regressors and previous physician experience and its square. The coefficient on track in this latter specification is 571 with a standard error of 255. The percentage impact is 14%. Column 12 reports the same specification as in column 11 in the discontinuity sample only. The coefficient on track is stronger, 628, but somewhat less precisely estimated. The percentage impact increases to 23%.

As the reduced-form estimates illustrate, IV estimates of the returns to a medical license and the returns to physician practice should differ only by a scale factor.[13] The IV returns to physician practice are, in general, larger than the IV returns to a license, as was found in the OLS results. In what follows, we focus on the returns to a license rather than on the returns to practicing as a physician, for the reasons previously mentioned.

B. Instrumental Variables Estimates

Instrumental variables estimates of the effect of acquiring a license are reported in table 4.[14] Acquisition of a license is instrumented with assigned retraining track. Columns 1–3 of table 4 report the estimated coefficients on licensed without any other regressors, with other regressors but excluding previous physician experience, and with other regressors and linear and quadratic terms for previous physician experience, respectively. The estimated coefficient on licensed with other regressors and linear and quadratic terms for previous physician experience is 1,638 with a standard error of 706. The percentage impact is 182%. Correcting for nonrandom selection in licensing status yields a percentage impact that is approximately double the corresponding percentage impact produced by OLS (90%).

[13] The IV estimates of the return to a license can be calculated from the ratio of the coefficient in the reduced-form earnings regression to the coefficient in the reduced-form license acquisition. The IV estimates of the return to physician practice can be calculated from the ratio of the coefficient in the reduced-form earnings regression to the coefficient in the reduced-form physician employment regression.

[14] The instrumental variables estimates reported in this subsection deviate somewhat from the ratio of the relevant reduced-form estimates in table 3 due to different sample sizes.

Table 4
Two-Stage Least Squares Estimates of the Returns to a Medical License

	Full Sample				Discontinuity Sample
Regressors	(1)	(2)	(3)	(4)	(5)
Licensed	−5,244	178	1,638	1,865	1,886
	(1,516)	(863)	(706)	(864)	(1,043)
% Impact	−.827	.086	1.818	2.617	3.422
Experience	. . .	. . .	−1.24	44.90	1,210
			(25.16)	(76.64)	(619)
Experience^2	. . .	. . .	−1.17	−3.90	−29.84
			(.62)	(4.34)	(15.82)
Experience^3	. . .	. . .	. . .	.0407	. . .
				(.0639)	
Other regressors	No	Yes	Yes	Yes	Yes
RMSE	3,245	1,722	1,659	1,672	1,646
R^2	. . .	.2417	.2982	.2883	.3099
N	676	676	676	676	181

Note.—Robust standard errors are in parentheses. RMSE = root mean squared error. Licensed is instrumented with track. Other regressors include dummies for age on arrival, year of arrival, months in Israel, gender, marital status, profession of spouse, number of children, size of last city of residence, republic of origin, advanced medical degrees, previous specialist status, previous type of medical practice, and type of reported earnings (after tax and/or after other deductions). The discontinuity sample in col. 5 uses the subsample of observations between 14 and 26 years of previous physician experience.

Considering that instrumental variables estimates in a regression discontinuity design may be quite sensitive to the way in which the variable generating the discontinuity is controlled, column 4 of table 4 reports the results of including a third-order polynomial in previous physician experience. The estimated effect of a license in this latter specification is 1,865 with a standard error of 864. The percentage impact grows to 262%.

Column 5 of table 4 reports the results of including linear and quadratic terms for previous physician experience in the discontinuity sample only. The estimated coefficient on licensed further grows to 1,886 but is somewhat less precisely estimated. The percentage impact is 342%. The corresponding percentage impact according to OLS in table 2 is 114%. Narrowing the range of the discontinuity sample produces even larger percentage impacts, but they are less precisely estimated as the sample size increasingly shrinks.

Additional IV results of interest that are not shown in table 4 for the sake of brevity are as follows. First, there are no significant interactions between licensing status and other covariates. In particular, there are no significant interactions between licensing status and time in Israel. Licensing status does not affect the slope of immigrant earnings profiles, only earnings levels. There are also no significant interactions between licensing status and previous physician experience. Second, including separate controls for the small number of immigrant physicians who were assigned to the observation track according to the 14-year rule does not change the results.[15]

The IV estimates in table 4 isolate the increase in mean earnings for immigrant physicians who acquire a license. The counterfactual is an immigrant physician who does not acquire a license and who cannot work as a physician. The parameter being estimated in this analysis is not the increase in mean earnings due to the imposition of licensing requirements on practicing physicians. This latter parameter would also be interesting to estimate, but it is not identified. One reason that it is not identifiable, which was already mentioned earlier, is that both licensing status and physician employment are endogenous and we have only one valid instrument at our disposal.

It is also worth reiterating that the return to acquiring a medical license and the return to practicing as a physician differ. One possible reason why they differ is that there is signaling value to a license when not practicing as a physician (the noncausal evidence in table 2 does not support this hypothesis). Another possibility is that some immigrant physicians fail to use the medical license that they acquire for physician practice, perhaps because they do not receive good physician job offers. Consistent with this latter explanation, the first-stage results in table 3 show that assignment to the observation track has more of an impact on the probability of license acquisition than on the probability of physician employment.

Although we are not estimating the same parameter as in many other papers in the licensing literature, it is still instructive to compare our estimates to those found by others. It is interesting that Kleiner (2000) reports licensing effects on hourly earnings, not taking into account nonrandom selection, that are very similar in magnitude to our OLS estimates presented in table 2. He finds that licensed dentists earn 91% more than unlicensed biological and life scientists, and licensed lawyers earn 94% more than unlicensed economists. Results for licensed physicians are not reported.

Studies that estimate the effect of imposing licensing requirements in a single profession usually find more modest earnings effects than in our study and in Kleiner (2000). Using cross-state variation in the stringency of regulation to become a dentist, Kleiner and Kudrle (2000) find that the hourly earnings of dentists in the most regulated states are only 12% higher than those of dentists in less regulated states. Reduced-form results in Angrist and Guryan (2003) indicate that state licensing requirements raise the salaries of teachers with a BA by between 3% and 5%. Given

[15] Further experimentation with the small number of immigrant physicians subject to the 14-year rule does not produce reliable results. There are also no other data sources available that could be brought into the analysis for this purpose.

that the imposition of licensing requirements increases the probability of license acquisition by about .5, implicit IV estimates, using state licensing requirements as instruments, are twice as large as reduced-form estimates.

Potential Threats to Identification

There are several potential threats to a causal interpretation of the IV results that we have presented in this study. First, since 27% of all immigrants who reported themselves to be physicians on arrival did not submit their credentials to the Ministry of Health, the sample could be biased. However, almost all of these immigrants were over 55 years of age on arrival. It is also unlikely that younger immigrant physicians are more likely to be misclassified by the Ministry of Health as immigrant physicians after the credential verification process and that these immigrant physicians are more likely to choose to become nonphysicians.

Second, there is a possibility that assignment to the exam track and/or failure to pass the exam led to out-migration. Information on the outcomes of these immigrants would be missing from the sample. However, as mentioned in Section II, return migration was one of the least important reasons for nonresponse to the survey. Only 3.7% of those approached to be interviewed were not interviewed because they had out-migrated or were absent from the country for an extended period of time.

Third, physicians in the former USSR may have decided not to immigrate to Israel (or decided to delay immigration) based on knowledge of the assignment rule. However, immigrant physicians were asked on the survey if they were aware of the relicensing process upon arrival to Israel, and only 4% responded that they had knowledge of relicensing procedures. This is consistent with the widely held belief that early arrivals from the former USSR were panic migrants, not economic migrants.

Interpretation

Our interpretation of the large IV estimates, compared to OLS estimates, is that relicensing requirements in the medical profession led to practitioner rents. This is consistent with standard economic theory. In addition, our theoretical model of optimal license acquisition suggests that when IV estimates exceed OLS estimates, the requirement of having to pass a relicensing exam (as opposed to being observed on the job) leads to lower average quality in the market for physicians. When IV estimates exceed OLS estimates, there is negative selection. Negative selection implies lower average quality of service under the assumption that there is a positive association between outside earnings potential, general talent, and physician quality, as specified in our model.

Note that the possibility of a large earnings premium coexisting with negative selection has been raised before in the occupational licensing literature (see Friedman and Kuznets 1945; and, more recently, Angrist and Guryan 2003). The combination of high psychological costs to retraining, lengthy delays in being able to secure employment as a physician, liquidity constraints, and relatively good outside opportunities for the generally talented can induce the most skilled immigrant physicians in Israel to seek employment opportunities as nonphysicians. In favor of this interpretation is the evidence that the average age of immigrant physicians on arrival is 43, that the average number of months from arrival to the first physician job is 36, and that most immigrant physicians who do not acquire a license tend to work in skilled occupations. It is well established that investment activity decreases with age, length of the investment period, and alternative returns. As in the simplest human capital models in the schooling literature, which predict that ability bias is negative (e.g., Griliches 1977), the highest-skilled individuals are the ones who do not invest in license acquisition, because their opportunity costs of acquiring a license are the highest. A large earnings premium to license acquisition is not necessarily enough to cover the costs.

A large return to licensing, such as estimated here, can also be interpreted as arising from high returns to occupation-specific human capital, instead of rents and negative selection, as we have argued. Working in favor of our rent/selection story is the finding that interactions of licensing status with previous physician experience do not yield significant coefficients on the interaction terms. Previous physician experience is generally not significant, even without interactions, and even has small negative returns in several specifications. The main reason for this result is that many immigrant physicians who obtained jobs as physicians were placed at the bottom of the physician job ladder in public hospitals and community clinics, regardless of age and former specialty. For similar reasons, negligible returns to previous work experience have also been estimated for Soviet immigrant engineers, managers, and unskilled workers (see Friedberg 2000; and Weiss, Sauer, and Gotlibovski 2003). Nonetheless, substantial earnings differences remain between those who find work at the bottom of the job hierarchy in the medical profession and those who find work at the bottom of the job hierarchy in alternative occupations.[16]

C. Quantile Regression and Treatment Effect Estimates

The constant-effects IV model relied on to measure the returns to an occupational license has a drawback in that it does not allow for differential effects of acquiring a license at different points in the earnings

[16] It is unlikely that earnings differences between professions are highly confounded with differences in hours worked, given that part-time jobs in Israel are not in abundant supply.

distribution. This can be problematic, since the earnings distribution has a mass point at zero. In general, the effect of acquiring a license on participation may be substantially different from the effect of acquiring a license on conditional-on-positive mean earnings. An alternative estimation strategy that is less sensitive to the inclusion of zero earnings for the unemployed, less demanding than formal sample-selection models, and less sensitive to high-earnings outliers is the quantile treatment effects model (Angrist 2001; and Abadie et al. 2002). The QTE model modifies conventional quantile regression for inclusion of an endogenous binary regressor and provides a good robustness check on the conclusions reached from the standard IV model. The QTE model is briefly outlined in the appendix.

The top panel of table 5 reports quantile regression and QTE estimates of the effect of licensing status on monthly earnings. Licensing effects are measured at the 0.15, 0.25, 0.50, 0.75, and 0.85 quantiles of the monthly earnings distribution. Quantile regression treats licensing status as exogenous and produces the largest percentage impacts of acquiring a license, 93% and 117%, at the 0.15 and 0.25 quantiles, respectively. The percentage impact steadily declines at higher quantiles, falling to 59% at the 0.85 quantile. At each quantile the coefficient on licensed is precisely estimated.

Quantile treatment effects estimates that correct for the endogeneity of licensing status yield substantially different results. The percentage impact of acquiring a license at the 0.15 and 0.25 quantiles is 50% for both, considerably lower than the quantile regression estimates. The percentage impact at the 0.50, 0.75, and 0.85 quantiles is, however, considerably higher than the quantile regression estimates. The QTE model produces the largest percentage impact, 169%, at the 0.85 quantile. At each quantile the coefficient on licensed is precisely estimated.

Note that the highest percentage impact of 169% at the 0.85 quantile is less than the percentage impact estimated in the corresponding specification in the constant-effects model in table 4 (182%). This suggests that the licensing effect on mean earnings in the constant-effects model is relatively more sensitive to high-earnings outliers than the inclusion of zeros for the unemployed.

The bottom panel of table 5 reports licensing effects on median earnings in the discontinuity sample only. Effects at other quantiles are difficult to identify, given the reduced variation in earnings and smaller sample size. The results indicate a large effect on median earnings, 239%. The corresponding percentage impact when treating licensed as exogenous is 108%. Both effects are precisely estimated. These percentage impacts stand in sharp contrast to the percentage impact on mean earnings in the discontinuity sample (342%). This again suggests greater sensitivity of standard IV to high-earnings outliers than the inclusion of zeros.

The quantile regression and QTE estimates reported in table 5 can be

Table 5
Quantile Regression and Treatment Effects Estimates

	Quantile Regression Estimates					Treatment Effects Estimates				
	.15	.25	.5	.75	.85	.15	.25	.5	.75	.85
A. Full sample:										
Licensed	359	644	997	1,455	1,507	119	257	1,115	1,589	2,774
	(130)	(163)	(143)	(178)	(320)	(4)	(63)	(65)	(267)	(212)
% Impact	.932	1.175	.832	.782	.588	.496	.499	1.466	1.076	1.688
Pseudo R^2	.1626	.2467	.2397	.2116	.2055	.1190	.1744	.2315	.1579	.1748
N	676	676	676	676	676	548	548	548	548	548
B. Discontinuity sample:										
Licensed	. . .	. . .	1,093	. . .	. . .	. . .	. . .	1,793	. . .	. . .
			(520)					(310)		
% Impact	. . .	. . .	1.076	. . .	. . .	. . .	. . .	2.386	. . .	. . .
Pseudo R^2	. . .	. . .	.1897	. . .	. . .	. . .	. . .	.2413	. . .	. . .
N	. . .	. . .	181	. . .	. . .	. . .	. . .	164	. . .	. . .

NOTE.—Other regressors include a quadratic in previous physician experience, dummies for age on arrival, year of arrival, months in Israel, republic of origin, advanced medical degrees, previous specialist status, previous type of medical practice, and type of reported earnings (after tax and/or after other deductions). The discontinuity sample is the subsample of observations between 14 and 26 years of previous physician experience. Bootstrapped standard errors are in parentheses. The bootstrapped standard errors when licensed are treated as endogenous and adjusted for the first-step estimation.

Table 6
Earnings Quantiles without a License

	Quantile				
	.15	.25	.5	.75	.85
A. Full sample:					
Licensed immigrants	239	515	760	1,476	1,643
Unlicensed immigrants	385	548	1,199	1,859	2,563
% Selection bias	−.379	−.060	−.366	−.206	−.359
B. Discontinuity sample:					
Licensed immigrants	. . .	. . .	751	. . .	. . .
Unlicensed immigrants	. . .	. . .	1,016	. . .	. . .
% Selection bias	. . .	. . .	−.261	. . .	. . .

NOTE.—The table reports monthly earnings quantiles without a license for immigrants who acquired a license and for immigrants who did not acquire a license. Earnings without a license for immigrants who acquired a license are calculated using quantile treatment effect estimates. Earnings without a license for immigrants who did not acquire a license are calculated using quantile regression estimates. The discontinuity sample is the subsample of observations between 14 and 26 years of previous physician experience.

used to estimate the marginal distributions of monthly earnings without a license both for immigrants who acquired a license and immigrants who did not acquire a license. Potential earnings without a license for immigrants who acquired a license are obtained by using the QTE coefficients together with the covariate means among those who acquired a license and by setting the licensing status dummy to zero. The counterfactual earnings of all immigrants with a license can be approximated by the counterfactual earnings of compliers (individuals whose treatment status is affected by the instrument), under the assumption that compliers are a random sample of all immigrants with a license. The monthly earnings without a license for immigrants who did not acquire a license are also computed conditional on the mean of the covariates among immigrants who acquired a license, and with the licensing status dummy set to zero, but using the quantile regression coefficients.

The top panel of table 6 shows that licensed immigrants have lower potential earnings without a license than unlicensed immigrants at all examined quantiles. The negative selection bias is greatest in the tails and at the median of the distribution, varying between 36% and 38%. The bottom panel of table 6 shows that negative selection bias is also present at the median of potential earnings in the discontinuity sample.

VII. Conclusion

This study measures the effects of occupational licensing requirements by exploiting an immigrant physician retraining assignment rule. The OLS estimates of the returns to a license among immigrant physicians in Israel range between 90% and 114%. Instrumental variables estimates of the returns range between 180% and 340%. The large IV estimates, compared to OLS estimates, are suggestive of the presence of rents accruing to practitioners and negative selection into licensing status.

In order to give an economic interpretation to the OLS and IV estimates, we develop a model of optimal license acquisition. According to the model, IV estimates that exceed OLS estimates imply that stricter relicensing requirements lead to lower average quality of service in the licensed occupation. The policy implication of the results is that lowering the costs of acquiring a license may raise physician quality.

As a robustness check, the returns to an occupational license are also estimated using a QTE model. The QTE estimates indicate that the returns to an occupational license are largest at the upper quantiles of the earnings distribution. Conventional quantile regression estimates indicate that the returns to a license are largest at the lower quantiles. The QTE estimates also illustrate that standard IV estimates are more sensitive to high-earnings outliers than the inclusion of zero earnings for the unemployed. The QTE estimates of the counterfactual distribution of earnings without a license among those immigrants who obtained a license also provide a rich picture of negative selection into licensing status.

The particular social experiment that we exploit in this article to identify the causal effects of relicensing requirements in the market for physicians may be relevant in other contexts. Other countries use similar licensing regimes for immigrants as well as for internal migrants. Future research could examine the earnings and quality effects of relicensing regimes in other countries and in other professions using a methodology similar to the one employed in this study.

Appendix

The Quantile Treatment Effects Model

The QTE model specifies a linear relationship between earnings and licensing status at each quantile. That is,

$$Q_\theta(Y_i|X_i, L_i, L_{1i} > L_{0i}) = X_i'\beta_\theta + L_i\alpha_\theta, \tag{A1}$$

where L_{1i} denotes licensing status when assigned to the observation track ($TR_i = 1$) and L_{0i} denotes licensing status when assigned to the exam track ($TR_i = 0$). The time in Israel subscript is suppressed for convenience. The coefficient α_θ has a causal interpretation because L_i is independent of potential earnings outcomes conditional on X_i and being a complier ($L_{1i} > L_{0i}$). The proof is found in Abadie et al. (2002).

The parameters of the QTE model are estimated by minimizing the sample analog of

$$E[\kappa_i\rho_\theta(Y_i - X_i'b_\theta - L_ia_\theta)], \tag{A2}$$

where ρ_θ is the check function and κ_i is weights that transform the conventional quantile regression minimand into a problem for compliers only. For computational reasons, κ_i is replaced by an estimate of $E(\kappa_i|X_i, L_i, Y_i)$, where

$$E(\kappa_i|X_i, L_i, Y_i) = 1 - \frac{L_i[1 - E(TR_i|Y_i, L_i, X_i)]}{[1 - E(TR_i|X_i)]} - \frac{(1 - L_i)E(TR_i|Y_i, L_i, X_i)}{E(TR_i|X_i)}. \tag{A3}$$

The first-step estimate of $E(\kappa_i|X_i, L_i, Y_i)$ is obtained by separately estimating $E(TR_i|Y_i, L_i, X_i)$ and $E(TR_i|X_i)$ with a probit. Predicted values of $E(\kappa_i|X_i, L_i, Y_i)$ that are negative are set to zero, leading to a reduced sample size. Standard errors are computed by bootstrapping the first- and second-step estimations 100 times.

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