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The Effects of Licensing on the Wages of Radiologic Technologists

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Abstract We analyze the effect that state licensing of radiologic technologists (RTs) has had upon RT wages with a unique dataset that allows us to control for place of work and job specialization. Using OLS and several measures of licensing, we find evidence that RTs working in states with licensing statutes earn as much as 3.3% more than RTs working in states without licensing. When we control for endogeneity using instrumental variables (IV) estimation, our estimate of the licensing premium doubles (6.9%). Our results provide further support for existing theories of the effects of occupational licensing on the wages of practitioners.

Keywords Occupational licensing · Radiologic technologists · Wages · Labor markets · Job regulation

Introduction

Occupational licensing in the USA directly affects more workers than either minimum wage legislation or unionization (Kleiner 2000, 2006). Prior research on the magnitude of the effects of occupational licensing upon the wages and numbers of practitioners has generally produced mixed results, however. In this study we add to the existing literature by examining the impact that state regulation of radiologic technologists (RTs) has had upon RT wages. Radiologic technologists take x-rays, administer nuclear medicine for diagnostic purposes, and also operate other devices that produce diagnostic images (mammograms and MRIs, for example).

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We use micro-level data from a recent wage survey conducted by the American Society of Radiologic Technologists (ASRT 2001). The survey data enable us to control for differences in both place of work (such as hospitals and physician offices) and job specialization (radiation therapy and nuclear medicine, for example) among the RTs in the sample. We estimate a basic wage function that uses four different measures of the strictness of regulation in each state. To control for possible endogeneity between wages and the strictness of regulation, we use the number of licensing board members as an instrument for our licensing variables.

We begin with a review of the existing literature examining the impact of occupational licensing upon the wages of practitioners. We then proceed with an overview and brief regulatory history of the radiologic technologist profession. After formulating our basic model and describing the ASRT data, we present our empirical results.

Prior Research on the Effects of Licensing

There are two divergent views on the effects of occupational licensing in the economics literature. One view argues that licensing is primarily a means for professionals to keep wages high by restricting entry into the profession, thus reducing consumer welfare. The modern origins for this argument can be attributed to Milton Friedman (1962). We label this view the "private interest" hypothesis. Both Friedman and Kuznets (1945) and Stigler (1971) provided some early evidence supporting the private interest hypothesis.

A second view concedes that occupational licensing increases the wages of professionals, but argues that licensing serves as a means of solving an asymmetric information problem. Consumers have less information than practitioners, and licensing protects consumers from poor service. We label this view the "public interest" hypothesis. Leland (1979) and Shapiro (1986) both develop models that support the public interest hypothesis that licensing improves the quality of services delivered to consumers.

Both hypotheses suggest that licensing should increase the wages of workers in the licensed occupations. Dozens of empirical studies have attempted to test this theory but have produced conflicting evidence on the impact of licensure upon wages. For occupations requiring graduate education, most studies find that licensing increases the professional's wages. For example, Shephard (1978), Kleiner (2000), and Kleiner and Kudrle (2000) all find evidence that stricter licensing standards increase the wages of dentists. Kleiner (2000) and Tenn (2001) detect similar findings for lawyers. Kugler and Sauer (2005) also find similar results when examining physician wages in Israel. And Benham and Benham (1975) and Feldman and Begun (1985) find evidence that stricter licensing regulation increases the price of optometry services. The increase in wages resulting from licensing has been found to range from as little as 1% to as much as 340%.

Evidence is more mixed for workers in occupations with less stringent education requirements. For example, White (1978) finds that licensing increases the wages of clinical lab personnel only in the two states that require applicants to be college graduates. In a later study, White (1980) finds evidence that licensing increased the wages of registered nurses in 1950, but he is unable to detect similar results in 1960 and 1970. Kleiner (2000) finds no evidence that licensing improves the wages of barbers and cosmetologists relative to other unlicensed occupations requiring similar levels of training. However, Adams et al. (2002) offer evidence that stricter licensing increases the price of cosmetology services.

Similar mixed findings exist for studies that have examined licensure's impact upon the number of practitioners. Both the private and public interest hypotheses predict that licensing should reduce the number of practitioners in the respective profession. For example, Kleiner and Kudrle (2000) estimate that a 10% decrease in a state dentistry licensing exam's pass rate reduces the number of dentists per capita in the state by 2%. Jackson (2006) and Carpenter and Stephenson (2006) find similar results for CPAs. For the barbering occupation, however, Thornton and Weintraub (1979) find little evidence that stricter licensing requirements have had an impact on a state's number of barbers.

To summarize, most studies examining licensing's impact upon occupations requiring graduate education appear to support the private interest hypothesis. The evidence is mixed, however, for those occupations with lower education requirements. Our study tests the impact that licensing has had upon radiologic technologist wages. RTs are not required to obtain a bachelor's degree; instead, states with licensure statutes generally require that RTs complete 2 years of vocational training beyond high school. To the best of our knowledge, no prior study has examined the economic impact of licensing of the RT profession. One reason, no doubt, is the fact that regulation of the RT profession is a fairly recent phenomenon.

The Radiologic Technologist Profession

RTs generally specialize in a particular diagnostic imaging area. The three most common areas are radiography, nuclear medicine, and radiation therapy. The most common, radiography, involves the use of x-rays to produce black and white anatomical images. Nuclear medicine technologists administer radiopharmaceuticals to patients to take pictures of internal organs and analyze their current functionality. Radiation therapists administer radiation to cancer patients. Formal training programs in radiologic technology are one to 4 years in length, with most programs lasting 2 years.

Voluntary certification has existed for RTs since the early 1920s. The current body of accreditation is the American Registry of Radiologic Technologists (ARRT). Candidates for certification must abide by the ARRT Standard of Ethics, must complete an ARRT-accredited education program covering the specialties they are seeking, and must pass an exam in their area of specialization. To maintain certification, an RT is required to complete 24 credit hours of ARRT-approved continuing education every 2 years.

Government regulation of RTs is a much more recent phenomenon. Those states regulating the RT profession generally do so by adopting a licensure statute. A licensure statute makes it illegal to practice without first obtaining a license. The first state to require licensing for RTs was New York in 1965 (ASRT 2004). By 1980, ten states had enacted licensing legislation for the profession. Today, 35 states require all

RTs to be licensed.[1] Four states require licensing for a limited number of specializations, such as mammography.[2] At this time, the states of Kansas and North Dakota are both in the process of establishing licensure requirements. The remaining nine states and the District of Columbia have no form of regulation.

Licensing requirements vary based upon the field of specialization. Generally, states require that applicants complete 2 years of education at an accredited hospital-based program or a 2-to-4-year educational program at an academic institution. Table 1 provides a brief overview of state regulation of RTs. In the table, "full licensure" denotes a state that requires all RTs (regardless of specialization) to obtain licensing.

Basic Model

When micro-level data are available, many studies analyzing the impact of state licensure on wages of professionals employ human capital models (e.g. Kleiner and Kudrle 2000; Tenn 2001). This article utilizes a similar framework. The following equation (Eq. 1) will be estimated:

$$\begin{aligned} (1) \quad \ln(\text{wage}) &= \alpha + \beta_1(\text{age}) + \beta_2(\text{age}^2) + \beta_3(\text{experience}) + \beta_4\left(\text{experience}^2\right) \\ &\quad + \beta_5(R) + \beta_5(S) + \beta_6(V) + \sum_{i=1}^{k}\phi_i X_i + \sum_{i=1}^{l}\lambda_i Y_i + \sum_{i=1}^{m}\eta_i Z_i + \varepsilon \end{aligned} \tag{1}$$

where R is a variable measuring the level of job training, S is a state's gross state product per capita, V is a measure of the strictness of regulation, the X_i are k place-of-work dummies (hospital or outpatient clinic, for example), the Y_{i} are l individual characteristic dummy variables (e.g., gender, marital status, and union affiliation), and the Z_i are m job specialization dummy variables (radiography and nuclear medicine, for example). We explain the rationale for each of the variables below.

Age and experience should have a positive effect upon RT wages, but the effect should decline as age and experience increase (hence their squared values in the wages function). Human capital theory also predicts that individuals with more job training should have higher wages. To measure job training, we include a dummy variable for RTs who are certified in more than one field.

In all professions, wage and salary levels vary from state to state. For the RT profession, some states may have a higher demand for RT services for a variety of reasons (for example, different income levels, tastes and preferences, etc.). In their

[1] State regulation of the profession increased significantly soon after the passage of the Consumer–Patient Radiation Health and Safety Act. This federal legislation, introduced by former West Virginia Senator Jennings Randolph, was meant to establish uniform standards for state regulation of RTs. Although the legislation was accepted into law in 1981, no state was required to comply with the new federal standards. Since the early 1990's RTs have lobbied for the passage of new legislation making state compliance with the Consumer–Patient Radiation Health and Safety Act mandatory.

[2] In our sample, there are only a small number (sometimes none) of RTs in states with partial licensing that are practicing in the particular field that requires licensing. We attempted to control for this possibility, but we found that it did not have a substantial effect upon our results.

Table 1 Summary of state regulation of radiologic technologists

State	Full licensure (1=yes, 0=no)	Year enacted[a]	CE hours per year	Number of licensing board members	Number of RTs on board	% RTs on board
Alabama	0	No statute	0	0	0	0
Alaska	0	No statute	0	0	0	0
Arizona	1	1977	12	10	4	40
Arkansas	1	1999	6	10	4	40
California	1	1969	0	11	2	18
Colorado	0	No statute	0	0	0	0
Connecticut	1	1993	0	0	0	0
Delaware	1	1989	0	15	0	0
District of Columbia	0	No statute	0	0	0	0
Florida	1	1979	6	15	3	20
Georgia	0	No statute	0	0	0	0
Hawaii	1	1974	12	10	6	60
Idaho	0	No statute	0	0	0	0
Illinois	1	1990	12	14	3	21
Indiana	1	1982	0	9	3	33
Iowa	1	1987	12	10	4	40
Kansas	0	No statute	0	0	0	0
Kentucky	1	1978	12	10	2	20
Louisiana	1	1984	12	11	4	36
Maine	1	1984	12	9	4	44
Maryland	1	1992	12	8	3	38
Massachusetts	1	1987	10	9	4	44
Michigan	0	No statute	0	0	0	0
Minnesota	1	1997	0	0	0	0
Mississippi	1	1996	12	10	4	40
Missouri	0	No statute	0	0	0	0
Montana	1	1977	6	7	3	43
Nebraska	1	1987	12	9	0	0
Nevada	0	No statute	0	0	0	0
New Hampshire	0	No statute	0	0	0	0
New Jersey	1	1968	0	15	3	20
New Mexico	1	1983	10	11	4	36
New York	1	1965	0	9	3	33
North Carolina	0	No statute	0	0	0	0
North Dakota	0	No statute	0	0	0	0
Ohio	1	1995	6	13	0	0
Oklahoma	0	No statute	0	0	0	0
Oregon	1	1979	12	7	5	71
Pennsylvania	1	1987	0	0	0	0
Rhode Island	1	1994	12	7	3	43
South Carolina	1	1999	12	13	5	38
South Dakota	0	No statute	0	0	0	0
Tennessee	1	1985	10	0	0	0
Texas	1	1987	12	11	3	27
Utah	1	1989	8	7	4	57
Vermont	1	1984	12	5	2	40
Virginia	1	1997	12	5	3	60
Washington[b]	1	1991	0	N/A	N/A	100
West Virginia	1	1977	12	9	3	33
Wisconsin	0	No statute	0	0	0	0
Wyoming	1	1985	0	5	3	60

All data from *Legislative Guidebook for State and Federal Legislation*, ASRT 2004. Data were confirmed (and adjusted in some cases) by contacting each state directly.

[a] The states of Colorado, Michigan, and Nevada require licensing only for RTs practicing mammography. The state of Wisconsin requires only RTs practicing radiation therapy or computerized tomography to obtain a license.

[b] All members of the licensing board in Washington are RTs, but the number of members varies from year to year based upon the governor's appointments.

study examining the effect of federal transfer programs on corruption, Fisman and Gatti (2002) use state gross domestic product per capita to control for such differences across states. We likewise include state GDP per capita in our wage regression. We hypothesize that the coefficient will have a positive sign—higher GDP per capita will result in higher demand for RT services and thus higher RT wages.[3] Place of work is also likely to have some influence on RT wages. For example, RTs working for larger employers (e.g., hospitals) will be paid higher wages than RTs working in smaller establishments (e.g., physicians' offices). To control for these differences, we include place-of-work dummy variables. Among RTs there are also differences in pay based upon their area of specialization. RTs specializing in radiation therapy typically earn much more than RTs specializing in radiography, for example. To control for any differences in wages resulting from differences in the RT's field of specialization, we include a dummy for each area of specialization.

The individual characteristic dummies are included to control for some common findings in the empirical literature. In the majority of empirical human capital studies, females are generally found to earn less than males (Blau and Kahn 2000). The effect of marriage on wages depends upon the sex of the worker—married males are generally found to earn more than unmarried males while the reverse is generally true for married females (Korenman and Neumark 1991, 1992). Hence, the sign of the coefficient of the married variable is indeterminate. In addition, workers covered by collective bargaining agreements are generally found to earn more than workers with no collective bargaining (see, for example, Blanchflower and Bryson 2004).

With respect to the strictness of regulation, our major variable of interest, there are a number of potential measures.[4] The most rudimentary is a dummy variable equal to 1 if the state requires licensure for RTs and 0 if the state does not require licensing. There may be some lag, however, in the impact of licensure upon RT wages, and to control for this we include as an alternate specification the number of years that a state has possessed a licensing statute.[5]

Another way of measuring the effects of licensing is to include variables measuring specific regulations from the statute.[6] One possible measure is based on the *composition* of a state's licensing board. Graddy and Nichol (1989) argue that if a licensing board has more public members, the state will likely have less stringent regulation. Public members of a licensing board will be more interested in protecting the public interest and preventing the passage of frivolous licensing requirements. Adams (1996) suggests that board *size* may be positively related to the strictness of

[3] We also calculated OLS regressions excluding this variable. Our results were very similar.

[4] Each of our proposed measures is meant to measure the same thing—the extent or strictness of licensing. We therefore include each variable separately to avoid potential multicollinearity problems.

[5] States without statutes are coded as zero.

[6] We chose to use proxies since a number of more direct measures were infeasible to use. One commonly used measure of the strictness of regulation is the pass rate on the licensing exam. For many occupations, each individual state will construct its own independent exam for licensure. This is not true for RTs. The overwhelming majority of states that regulate RTs contract out the examination preparation and grading to the ARRT. Education requirements among states are also very similar. Furthermore there is little variation in reciprocity arrangements across states. If a state has reciprocity, it will recognize the license of another state (under certain conditions). All states have reciprocity, with the exception of Delaware.

licensing in a state. He contends that a larger board will contain more public board members. As a result, there is a greater possibility of public members having an influence on board decision making. Although both of these variables may be related to the strictness of licensing, it would seem that some combination of the two would provide a measure more appropriate for our study. The number of public members alone is not satisfactory since it does not take into account the size of the board. A public member will be much more effective in limiting the stringency of the licensing regulation if there are fewer members of the profession on the board. The size of the board is not totally satisfactory either. A larger board may mean either that there are more public members or that there are more members from the profession. The impact appears to be ambiguous. In this study, therefore, we measure the strictness of a state's licensing regulation by the percentage of licensing board members who are RTs ("RT board density"). The hypothesized sign of the coefficient is positive. RT board density will increase if the number of RT board members increases or if the number of public members decreases. Both effects should increase the stringency of licensing in a state and also, we would expect, the wages of RTs, all other things being equal.

Our last measure of the strictness of a state's licensing statute is the number of continuing education (CE) hours required for RTs. To maintain ARRT certification, RTs are required to complete 24 hours of CE every 2 years. Adams (1996) argues that CE requirements should be positively correlated with the stringency of licensing standards. CE requirements constitute an additional cost for licensed RTs, hence a barrier to entry, and should therefore discourage entry into the occupation. To test this hypothesis, we use the number of hours that the CE requirements of a state are *below* ARRT guidelines. A negative sign would indicate support for the Adams hypothesis.[7]

Data

The ASRT has periodically conducted salary surveys of its members and other RTs. We use data from the 2001 survey for our empirical analysis.[8] The sample includes RTs paid an hourly wage and those who are paid a salary. For the purposes of this study, we focus upon RTs paid an hourly wage.[9] Table 2 provides summary statistics for the ASRT sample. In the table we compare means for the variables in states with licensing to the corresponding means in states without licensing. A standard *t*-test reveals that average RT wages are higher in states with licensing than in states without licensing. Most RTs in the sample are married females whose wages are not covered by collective bargaining agreements and who practice

[7] Of course stiffer continuing education requirements also constitute an additional cost for already licensed RTs. Therefore, it might be argued that the expected sign for this coefficient is indeterminate.

[8] The ASRT also conducted surveys in 1997 and 2004. The 1997 data has several coding problems we were unable to resolve. The ASRT also admitted experiencing distribution problems with the 2004 survey. As a result, we focus solely on the 2001 survey data.

[9] The determinants of wages for salaried employees are likely to differ from the determinants of wages for RTs paid an hourly wage. To avoid any resulting complications, we decided to focus on wage earners in this study.

Table 2 Summary statistics for ASRT survey data

States without licensing (n=2,440)		States with licensing (n=6,960)	
Variable	Mean/%	Variable	Mean/%
Hourly wage	19.78	Hourly wage	20.85
Age	40.6	Age	40.9
Experience	15.4	Experience	15.3
Female	79.8%	Female	78.0%
Married	73.4%	Married	71.9%
Collective bargaining	2.8%	Collective bargaining	8.1%
Two or more credentials	66.3%	Two or more credentials	66.6%
Job specialization			
Radiography	27.8%	Radiography	26.9%
Radiation therapy	15.0%	Radiation therapy	16.8%
Nuclear medicine	5.0%	Nuclear medicine	4.4%
Sonography	4.9%	Sonography	4.7%
Mammography	11.0%	Mammography	10.3%
Cardiovascular interventional	8.0%	Cardiovascular interventional	7.7%
Computed tomography	9.2%	Computed tomography	9.4%
MRI	8.2%	MRI	8.6%
Medical Dosimetry	1.4%	Medical Dosimetry	1.4%
Other Field	9.5%	Other field	9.7%
Place of work			
Physician's office	18.5%	Physician's office	17.4%
Imaging center	8.4%	Imaging center	11.4%
Not-for-profit hospital	49.6%	Not-for-profit hospital	47.6%
For-profit hospital	15.4%	For-profit hospital	16.1%
Mobile unit	2.0%	Mobile unit	1.3%
VA or government hospital	1.2%	VA or government hospital	0.7%
Other workplace	4.8%	Other workplace	5.4%

All data from Radiologic Technologist Wage and Salary Survey, ASRT 2001

radiography.[10] More than half of the RT sample have more than one credential. Approximately 90% of RTs work in non-government hospitals, physician's offices, or imaging centers. Incidentally, no previous empirical study of licensing has used data possessing this level of detail about the professional's place of work.

Basic Model Estimation

To test for the impact of licensure on RT wages, Eq. 1 is estimated for RT wage-earners in the sample. Experience is measured using the number of years that an individual has worked as an RT. For the initial estimation, the most rudimentary measure of regulation is employed: a dummy variable with a value of 1 if the state had licensing legislation for RTs before the sample year and 0 otherwise. Throughout our paper we use standard errors corrected for state-level heteroskedasticity (Moulton 1990).

[10] In the occupational licensing literature, there is evidence of a relationship between migration rates and licensing (Pashigian 1980; Tenn 2001). Since the majority of RTs are married females and most likely not the primary wage earners in the family, we do not think that mobility between licensed and non-licensed states will influence our results.

Column (1) of Table 3 contains the results. Most of the variable coefficients are statistically significant and also have the expected signs. For example, we find that RTs working in highly technical fields (such as radiation therapy and medical dosimetry) earn more than RTs working in more basic fields (e.g., radiography and mammography). Our results also suggest that female and married RTs earn less than male and unmarried RTs respectively, and that RTs covered by collective bargaining agreements earn more than those not covered.[11] The main variable of interest to us, though, is the licensure dummy variable. Our results suggest that RTs working in states with licensing statutes earn approximately 3.3% more than RTs working in states without licensing.[12]

As mentioned previously, there is the possibility of a lagged effect of licensure upon the wages of practitioners. To investigate this possibility further, we next include the number of years that a state has had a licensing statute as a variable in our regression. Column (2) of Table 3 contains the results of this estimation. We continue to find evidence that licensing has a positive and statistically significant effect upon RT wages. Our results suggest that a 5-year increase in the length of time a state has a licensing statute in place increases RT wages by approximately 1%.

We next re-estimate Eq. 1 measuring the extent of licensing by using RT board density. We reason that a larger percentage of licensing board members who are RTs will lead to higher RT wages. Column (3) of Table 3 reports the results from this specification of our model. We continue to find evidence that stricter licensing statutes increase RT wages. The size of the effect is rather small, however. The average size of licensing boards for RTs is approximately 6 members. As a result, replacing one non-RT member with an RT member would increase RT wages by approximately 1.7%.

For our final specification, we measure the extent of licensing using the number of hours that the CE requirements of a state are below ARRT guidelines. The results of this specification are reported in Column (4) of Table 3. The sign of the coefficient is negative, suggesting support for the Adams (1996) hypothesis that continuing education requirements are positively correlated with the strictness of licensing; however, the coefficient is statistically insignificant. It does not appear that CE requirements affect RT wages.

In summary, we find some evidence that licensing affects RT wages. Our results suggest that licensing increases wages by as much as 3.3%. The magnitude of our estimate is relatively small but yet comparable to those found in other licensing studies examining professions not requiring a bachelor's degree.

Controlling for Possible Endogeneity

It is possible that our results in the previous section may have been influenced by a simultaneity problem in Eq. 1. In short, RTs might push for licensing regulation (or

[11] Since most RTs are female, the negative sign on the marriage dummy coefficient is consistent with the findings of other studies.

[12] Because the coefficients of variables in semilogarithmic regressions are only approximately equal to percentage changes, in our text discussion we have converted regression coefficients to percentage changes using the conversion formula $e^{\beta}-1$. See Halvorsen and Palmquist (1980) and Thornton and Innes (1989).

Table 3 OLS Estimates of the impact of licensing upon RT log hourly wages[a]

	(1)	(2)	(3)	(4)
Licensure dummy	0.0321 (0.0151)**			
Years licensed		0.0021 (0.0008)***		
RT board density			0.0010 (0.0003)***	
Hours below ARRT requirements				−0.0010 (0.0016)
Age	0.0095 (0.0029)***	0.0091 (0.0030)***	0.0087 (0.0030)***	0.0096 (0.0029)***
Age^2	−0.0001 (0.0000)***	−0.0001 (0.0000)***	−0.0001 (0.0000)***	−0.0001 (0.0000)***
Experience	0.0139 (0.0016)***	0.0141 (0.0016)***	0.0142 (0.0016)***	0.0138 (0.0016)***
$Experience^2$	−0.0002 (0.0000)***	−0.0002 (0.0000)***	−0.0002 (0.0000)***	−0.0002 (0.0000)***
Two or more credentials	0.0239 (0.0051)***	0.0231 (0.0053)***	0.0242 (0.0054)***	0.0238 (0.0051)***
Female	−0.0461 (0.0059)***	−0.0436 (0.0061)***	−0.0448 (0.0061)***	−0.0467 (0.0058)***
Married	−0.0234 (0.0046)***	−0.0224 (0.0043)***	−0.0222 (0.0046)***	−0.0235 (0.0047)***
Collective bargaining	0.0646 (0.0187)***	0.0614 (0.0162)***	0.0583 (0.0210)***	0.0696 (0.0199)***
Radiography	−0.1201 (0.0082)***	−0.1191 (0.0081)***	−0.1210 (0.0081)***	−0.1206 (0.0081)***
Radiation therapy	0.1839 (0.0113)***	0.1852 (0.0113)***	0.1858 (0.0114)***	0.1846 (0.0112)***
Nuclear medicine	0.0999 (0.0145)***	0.1023 (0.0142)***	0.1021 (0.0142)***	0.0994 (0.0145)***
Sonography	0.0860 (0.0133)***	0.0886 (0.0132)***	0.0856 (0.0133)***	0.0857 (0.0134)***
Mammography	−0.0751 (0.0113)***	−0.0736 (0.0114)***	−0.0768 (0.0114)***	−0.0758 (0.0113)***
Cardio. inter.	0.0198 (0.0103)*	0.0220 (0.0105)**	0.0212 (0.0105)**	0.0197 (0.0102)*
CT	−0.0222 (0.0105)**	−0.0204 (0.0103)*	−0.0220 (0.0106)**	−0.0224 (0.0105)**
MRI	0.0644 (0.0109)***	0.0660 (0.0111)***	0.0636 (0.0109)***	0.0639 (0.0109)***
Medical dos.	0.3248 (0.0218)***	0.3277 (0.0218)***	0.3225 (0.0221)***	0.3252 (0.0217)***
Phys. office	−0.1586 (0.0149)***	−0.1587 (0.0151)***	−0.1602 (0.0151)***	−0.1589 (0.0148)***
Imaging center	−0.0828 (0.0156)***	−0.0835 (0.0158)***	−0.0826 (0.0160)***	−0.0810 (0.0156)***
Not-for-profit hospital	−0.0819 (0.0155)***	−0.0815 (0.0155)***	−0.0835 (0.0152)***	−0.0827 (0.0154)***
For-profit hospital	−0.0850 (0.0160)***	−0.0849 (0.0162)***	−0.0856 (0.0158)***	−0.0852 (0.0159)***
Mobile unit	−0.0808 (0.0167)***	−0.0783 (0.0167)***	−0.0809 (0.0169)***	−0.0833 (0.0165)***
Government or VA hospital	−0.1117 (0.0265)***	−0.1117 (0.0268)***	−0.1125 (0.0267)***	−0.1170 (0.0273)***
GDP per capita	0.0000007 (0.0000)***	0.0000007 (0.0000)***	0.0000007 (0.0000)***	0.0000007 (0.0000)***
Constant	2.5149 (0.0882)***	2.5222 (0.0874)***	2.5148 (0.0867)***	2.5351 (0.0873)***
Observations	9,400	9,400	9,400	9,400
R^2	0.46	0.46	0.47	0.46

Standard errors in parentheses are adjusted for heteroskedasticity resulting from observations from the same state (Moulton 1990).

All data except licensure dummy and GSP per capita from Radiologic Technologist Wage and Salary Survey, ASRT 2001; licensure variables are from ASRT *Legislative Guidebook for State and Federal Legislation*, ASRT 2004. Data were confirmed (and adjusted in some cases) by contacting each state directly. State GDP per capita is obtained from the Bureau of Economic Analysis (BEA).

[a] Fifty states and District of Columbia. Dependent variable is the natural logarithm of the hourly wage.

*Significant at 10% level

**Significant at 5% level

***Significant at 1% level

Table 4 IV Estimates of the impact of licensing upon RT log hourly wages[a]

	(1)	(2)	(3)	(4)
Licensure dummy	0.067 (0.0242)***			
Years licensed		0.003 (0.0010)***		
RT board density			0.0016 (0.0005)***	
Hours below ARRT requirements				−0.006 (0.0024)**
Age	0.0094 (0.0029)***	0.0088 (0.0030)***	0.0082 (0.0029)***	0.0096 (0.0030)***
Age^2	−0.0001 (0.0000)***	−0.0001 (0.0000)***	−0.0001 (0.0000)***	−0.0001 (0.0000)***
Experience	0.014 (0.0016)***	0.0142 (0.0016)***	0.0144 (0.0016)***	0.0141 (0.0016)***
$Experience^2$	−0.0002 (0.0000)***	−0.0002 (0.0000)***	−0.0002 (0.0000)***	−0.0002 (0.0000)***
Two or more credentials	0.0238 (0.0050)***	0.0227 (0.0053)***	0.0244 (0.0054)***	0.023 (0.0050)***
Female	−0.0454 (0.0059)***	−0.0422 (0.0062)***	−0.0436 (0.0061)***	−0.0462 (0.0060)***
Married	−0.0231 (0.0045)***	−0.0219 (0.0041)***	−0.0213 (0.0043)***	−0.0227 (0.0047)***
Collective bargaining	0.0587 (0.0182)***	0.0575 (0.0163)***	0.0512 (0.0190)***	0.0671 (0.0213)***
Radiography	−0.1197 (0.0082)***	−0.1185 (0.0080)***	−0.1213 (0.0081)***	−0.1211 (0.0083)***
Radiation therapy	0.1831 (0.0114)***	0.1855 (0.0113)***	0.1865 (0.0118)***	0.1845 (0.0116)***
Nuclear medicine	0.101 (0.0142)***	0.1038 (0.0140)***	0.1041 (0.0142)***	0.1022 (0.0143)***
Sonography	0.0863 (0.0130)***	0.0899 (0.0133)***	0.0856 (0.0132)***	0.0853 (0.0129)***
Mammography	−0.0743 (0.0114)***	−0.0727 (0.0115)***	−0.0774 (0.0113)***	−0.0758 (0.0112)***
Cardio. inter.	0.0202 (0.0105)*	0.0231 (0.0108)**	0.0222 (0.0108)**	0.0207 (0.0103)**
CT	−0.0221 (0.0105)**	−0.0195 (0.0103)*	−0.0219 (0.0106)**	−0.0233 (0.0110)**
MRI	0.0648 (0.0109)***	0.0669 (0.0113)***	0.0634 (0.0107)***	0.0635 (0.0107)***
Medical dos.	0.3244 (0.0218)***	0.3288 (0.0220)***	0.3208 (0.0225)***	0.3251 (0.0221)***
Phys. office	−0.1584 (0.0148)***	−0.1586 (0.0151)***	−0.1611 (0.0153)***	−0.1589 (0.0146)***
Imaging center	−0.0849 (0.0153)***	−0.0847 (0.0159)***	−0.0837 (0.0161)***	−0.0817 (0.0154)***
Not-for-profit hospital	−0.0814 (0.0153)***	−0.0811 (0.0154)***	−0.0842 (0.0151)***	−0.0836 (0.0150)***
For-profit hospital	−0.085 (0.0158)***	−0.0849 (0.0161)***	−0.0859 (0.0156)***	−0.0861 (0.0155)***
Mobile unit	−0.0779 (0.0165)***	−0.076 (0.0161)***	−0.0795 (0.0170)***	−0.0832 (0.0166)***
Government or VA hospital	−0.1063 (0.0266)***	−0.1094 (0.0265)***	−0.1099 (0.0257)***	−0.1184 (0.0274)***
GDP per capita	0.0000007 (0.0000)***	0.0000007 (0.0000)***	0.0000007 (0.0000)***	0.0000007 (0.0000)***
Constant	2.4951 (0.0901)***	2.5173 (0.0883)***	2.5037 (0.0885)***	2.5451 (0.0854)***
Observations	9,400	9,400	9,400	9,400
R-squared	0.45	0.46	0.46	0.44
Instrument tests				
F-test of excluded instruments (BSIZE)	36.24	41.4	34.25	25.07
Wu–Hausman *F*-test	76.99	25.57	28.72	156.79

Standard errors in parentheses are adjusted for heteroskedasticity resulting from observations from the same state.
All data except licensure variables, *BSIZE,* and GSP per capita are from Radiologic Technologist Wage and Salary Survey, ASRT 2001. Licensure variables and BSIZE are from *Legislative Guidebook for State and Federal Legislation*, ASRT 2004. State GDP per capita is obtained from the Bureau of Economic Analysis (BEA).
[a] BSIZE is used as an instrument for each licensing variable. Dependent variable is the natural logarithm of the hourly wage.
*Significant at 10% level
**Significant at 5% level
***Significant at 1% level

stricter regulation if a statute is already in place) to increase wages. To eliminate any possible simultaneity bias, we re-estimate Eq. 1 using instrumental variables (IV).

To perform the IV estimation, we identify a variable to serve as an instrument that is correlated with the probability of a state imposing stricter regulation but not correlated with RT wages. In the first stage, we estimate the following equation:

$$V = \alpha + \beta_1(\mathrm{BSIZE}) + \sum_{i=1}^{k} \lambda_i Y_i + \gamma \tag{2}$$

where V is the appropriate measure of regulation, *BSIZE* is the size—the number of members of any type—of a state's licensing board, and the Y_i are the k exogenous variables from Eq. 1. Several studies have suggested that licensing board size makes an excellent instrument for licensing restrictions (Graddy and Nichol 1989; Adams et al. 2002; and Jang 2000). The size of a licensing board (as opposed to its composition) is correlated with the probability of a state imposing stricter regulation, but is not related to the earnings of practitioners.[13] Following the literature, we hypothesize that the coefficient on BSIZE will be positive. The larger the licensing board, the more likely a state is to have stricter licensing legislation.

To estimate Eqs. 1 and 2, we utilize two-stage least squares (2SLS).[14] To investigate the validity of our instruments, we refer to the F-test of our excluded instrument (BSIZE) at the bottom of Table 4. For each specification, the F-statistic is significantly larger than 10 so we feel confident that our instruments are valid (Staiger and Stock 1997). In addition, we find evidence in all cases that we can reject the null hypothesis that our OLS results are unbiased as a result of endogeneity (Wu 1973). Given these test results, we feel confident that our estimation strategy is correct and that our IV coefficient estimates are unbiased.

We now turn to the results of our IV regression results. Again we estimate one regression equation for each of the four licensing measures used in the previous section. The results are contained in Table 4 in Columns 1–4. For our first three measures of licensing, our estimated coefficients are larger for our IV results than they were for our OLS results. For example, our IV results suggest that RTs working in states with licensing statutes earn as much as 6.9% more than RTs working in states without licensing statutes. One difference between our OLS and IV estimates is the coefficient on the CE requirements variable. The IV estimate provides support for the Adams hypothesis—that is, CE requirements are positively correlated with other licensing restrictions and should therefore increase RT wages. In general, however, our results for all of the other licensing variables are roughly of the same order of magnitude after we control for endogeneity between licensing restrictions and RT wages. In other words, we continue to find evidence that licensing has moderately increased RT wages.

[13] We examined simple correlations between board size, RT wages, and our licensing variables. We found that the correlation coefficient between board size and wages is approximately 0.10 while the correlation coefficient between board size and the licensing variables ranges from 0.54–0.70.

[14] Full results from the first stage estimations are available from the authors upon request.

Conclusion

In this article we have estimated the impact that occupational licensure has had upon the wages of RTs. Using OLS, we find that licensing has increased the wages of RTs by as much as 3.3%. When we control for potential endogeneity bias using IV regression, the upper bound of our estimated effect increases to 6.9%. In general, the magnitude of our results is comparable to those found in other studies of the effects of licensing on earnings in professions with relatively low education requirements.

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References

Adams AF III (1996) Occupational licensing of cosmetologists and midwives: two empirical studies of the effects of regulation. Dissertation, Auburn University

Adams AF III, Ekelund R, Jackson J (2002) Occupational licensing in a 'competitive' labor market: the case of cosmetology. J Labor Res 23:261–278

American Society of Radiologic Technologists (2001) Radiologic technologist wage and salary survey

American Society of Radiologic Technologists (ASRT) (2004) Legislative guidebook for state and federal legislation

Benham L, Benham A (1975) Regulating through the professions: a perspective on information control. J Law Econ 18:421–447

Blanchflower D, Bryson A (2004) What effect do unions have on wages now and would Freeman and Medoff be surprised? J Labor Res 25:383–414

Blau F, Kahn L (2000) Gender differences in pay. J Econ Perspect 14:75–99

Bureau of Labor Statistics (2007) U.S. Department of Labor, Occupational outlook handbook, 2006–07 Edition

Carpenter C, Stephenson EF (2006) The 150-hour rule as a barrier to entering public accountancy. J Labor Res 27:115–126

Feldman R, Begun J (1985) The welfare cost of quality changes due to professional regulation. J Ind Econ 34:17–32

Fisman R, Gatti R (2002) Decentralization and corruption: evidence from U.S. Federal Transfer Programs. Public Choice 113:25–35

Friedman M (1962) Capitalism and freedom. University of Chicago, Chicago

Friedman M, Kuznets S (1945) Income from independent professional practice. NBER, New York

Graddy E, Nichol M (1989) Public members on occupational licensing boards: effects of legislative regulatory reforms. South Econ J 55:610–625

Halvorsen R, Palmquist R (1980) The interpretation of dummy variables in semilogarithmic equations. Am Econ Rev 70:474–475

Jackson R (2006) E. Post-graduate education requirements and entry into the CPA profession. J Labor Res 27:101–114

Jang C (2000) Occupational licensing in the new health care environment: an integrated framework and empirical analysis. Dissertation, University of Southern California

Kleiner M (2000) Occupational licensing. J Econ Perspect 14:189–202

Kleiner M (2006) Licensing occupations: ensuring quality or restricting competition. W.E. Upjohn Institute, Michigan

Kleiner M, Kudrle R (2000) Does regulation affect economic outcomes? The case of dentistry. J Law Econ 18:547–576

Korenman S, Neumark D (1991) Does marriage really make men more productive? J Hum Resour 26:282–307

Korenman S, Neumark D (1992) Marriage, motherhood, and wages. J Hum Resour 27:282–307

Kugler A, Sauer R (2005) Doctors without borders? Relicensing requirements and negative selection in the market for physicians. J Labor Econ 23:437–465

Leland H (1979) Quacks, lemons, and licensing: a theory of minimum quality standards. J Polit Econ 87:1328–1346

Moulton B (1990) An illustration of a pitfall in estimating the effects of aggregate variables on micro units. Rev Econ Stat 72:334–338

Pashigian BP (1980) Has occupational licensing reduced geographic mobility and raised wages? In Rottenberg S (ed) Occupational licensure and regulation. American Enterprise Institute, Washington, DC, pp 299–333

Shapiro C (1986) Investment, moral hazard, and occupational licensing. Rev Econ Stud 53:843–862

Shephard L (1978) Licensing restrictions and the cost of dental care. J Law Econ 21:187–201

Staiger D, Stock J (1997) Instrumental variable regression with weak instruments. Econometrica 65:557–586

Stigler G (1971) The theory of economic regulation. Bell J Econ 2:3–21

Tenn, S (2001) Three essays on the relationship between migration and occupational licensing. Ph.D. Dissertation, University of Chicago

Thornton R, Weintraub A (1979) Licensing in the barbering profession. Ind Labor Relat Rev 32:242–249

Thornton R, Innes J (1989) Interpreting semilogarithmic regression coefficients in labor research. J Labor Res 10:443–447

White W (1978) The impact of occupational licensure of clinical laboratory personnel. J Hum Resour 13:91–102

White W (1980) Mandatory licensing of registered nurses: introduction and impact. In Rottenberg S (ed) Occupational licensure and regulation. American Enterprise Institute, Washington, DC, pp 47–72

Wu D (1973) Alternative tests of independence between stochastic regressors and disturbances. Econometrica 41:733–750