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The Prevalence and Effects of Occupational Licensing

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Abstract

Our study provides the first national analysis of the labour market implications of workers who are licensed by any agency of the government in the USA. Using a specially designed Gallup survey of a nationally representative sample of Americans, we provide an analysis of the influence of this form of occupational regulation. We find that 29 per cent of the workforce is required to hold a licence, which is a higher percentage than that found in other studies that rely on state-level occupational licensing data or single states. Workers who have higher levels of education are more likely to work in jobs that require a licence. Union workers and government employees are more likely to have a licence requirement than are non-union or private sector employees. Our multivariate estimates suggest that licensing has about the same quantitative impact on wages as do unions — that is about 15 per cent — and that being both licensed and in a union can increase wages by more than 24 per cent. However, unlike unions which reduce variance in wages, licensing does not significantly reduce wage dispersion for individuals in licensed jobs.

1. Introduction

One of the fastest-growing yet least understood institutions in the US labour market is occupational licensing. The movement to a service-oriented economy from manufacturing, where unions and contracts were prominent, created a demand for a 'web of rules' of the workplace that licensing may have provided (Dunlop 1958). While unions have declined, occupational licensing has grown over the last 50 years (Kleiner 2006).

Occupational regulation in the USA generally takes three forms. The least restrictive form is registration, in which individuals file their names, addresses and qualifications with a government agency before practising their

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occupation. The registration process may include posting a bond or filing a fee. In contrast, certification permits any person to perform the relevant tasks, but the government — or sometimes a private, non-profit agency — administers an examination and certifies those who have achieved the level of skill and knowledge for certification. For example, travel agents and car mechanics are generally certified but not licensed. The toughest form of regulation is licensure; this form of regulation is often referred to as 'the right to practice'. Under licensure laws, working in an occupation for compensation without first meeting government standards is illegal. In 1992, the Council of State Governments estimated that more than 800 occupations were licensed in at least one state, and more than 1,100 occupations were licensed, certified or registered (Brinegar and Schmitt 1992).

In this article we use newly available data from a national survey conducted by the Gallup Organization on our behalf to analyse the influence of occupational licensing in the labour market. To determine if a worker is in a licensed position, the survey asks the question: 'Does your job require a license by a federal, state or local government agency?'

This study is the first attempt to gather US data by asking individuals whether a government licence is required to do their work. We find that about 29 per cent of the workforce is required to obtain a licence from either the federal, state or local government to work for pay. The information from the survey shows that licensing is more prevalent among more educated workers, minorities, union members and government workers. The impact of licensing on wages, as judged from a cross-sectional regression, is similar to the influence of unions on wages. Unlike the estimates of the impact of unions, which reduces the variation of wages, licensing results show that there is little effect on wage variations. However, workers who are licensed say they are more competent in doing their jobs, in contrast to what we find for union members.

2. Why license?

The simplest theory of occupational licensing emphasizes the administrative procedural role of licensing. It perceives a costless supply of unbiased, capable gatekeepers and enforcers. Friedman (1962) questioned the role of the government and professional associations as unbiased gatekeepers and enforcers. Instead, he viewed licensing's entry restrictions as creating undesirable monopoly rents through greater barriers to entry. Friedman argued that licensing systems are almost always run by and for incumbents, so that gatekeepers and enforcers are self-interested. Their vested interests lead them to not only create monopoly rents through restrictions on entry but also to limit complaints and disciplinary procedures against most incumbents.

A further modification of both of these perspectives presented above is that licensing requirements can take the form of an unspecified fixed cost of entry that is controlled by the licensing authority and is similar to the unbiased government gatekeeper (Shapiro 1986). The skill and quality of the licensed worker affects the relative cost of producing high-quality services, and licensing takes the more specific form of a minimal human capital requirement. In practice the fixed costs would be requirements that entrants and incumbents take job-specific training programmes, pass an exam or have long-term residency requirements. These models resemble the ones above in predicting, typically, that both the average quality and the average prices or earnings from the services within the regulated industry will rise as licensing requirements are implemented or tightened, resulting in benefits for those who want higher quality, but at a cost to those who are in lower-quality service markets (Shapiro 1986).

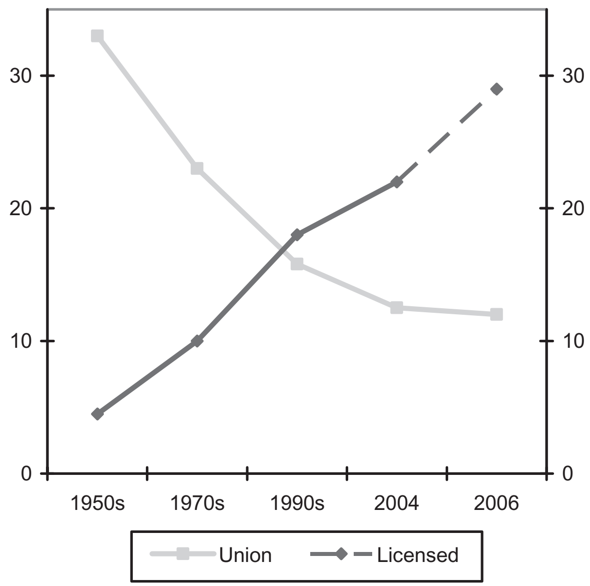
3. Growth of regulation

During the early 1950s, less than 5 per cent of the US workforce was covered by licensing laws at the state level (Council of State Governments 1952). That grew to almost 18 per cent by the 1980s — with an even larger number if federal, city and county occupational licensing is included. By 2000, the percent of the workforce in occupations licensed by states was at least 20 per cent, according to data gathered from the Department of Labor and the 2000 Census (US Department of Labor 2005; see America's Career InfoNet).

As employment shifted from manufacturing to service industries, which typically have lower union representation, the members of the occupations established a formal set of standards that governed members of the occupation. For a professional association, obtaining licensing legislation meant raising funds from members to lobby the state legislature, particularly the chairs of appropriate committees. In addition, the occupation association often solicits volunteers from its membership to work on legislative campaigns. With both financial contributions and volunteers, the occupational association has a significant ability to influence legislation, especially when opposition to regulatory legislation is absent or minimal (Wheelan 1998).

Most prior studies gathered data to estimate the number of individuals who were in licensed occupations at the state level, where most licensing occurs in the USA (Kleiner 2006). However, these estimates understate the regulation that also occurs at the federal and local levels. For this reason the Gallup survey provides a more comprehensive assessment of the coverage of occupational licensing. In Figure 1 we show trends in the growth of occupational licensing and unionization from 1950 to 2006.[1] Licensing data for earlier periods are only available at the state/occupational level; the data gathered through the Gallup survey are denoted with a dashed line in the figure. Despite possible problems in both data series, it is clear that occupational licensing is rising and unionization is declining. By 2006, 29 per cent of workers said they were required to have a government-issued licence to do their job compared with only 12 per cent who said they were union members.

FIGURE 1
Comparisons in the Time-Trends of Two Labor Market Institutions: Licensing and Unionization.[a]



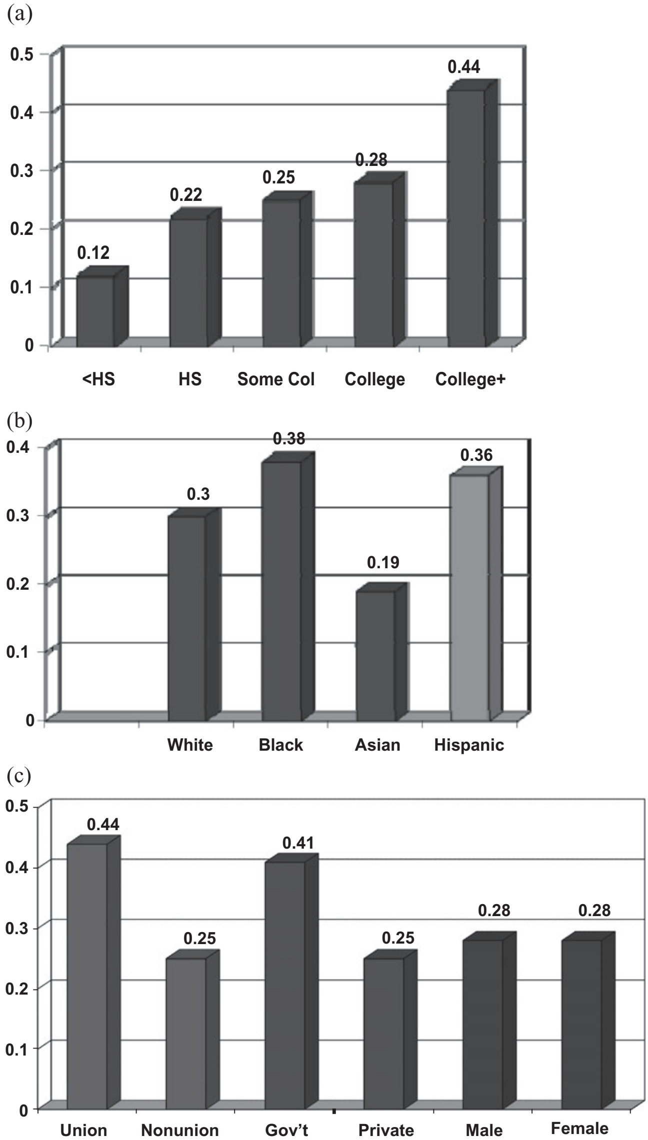
[a] Dashed line shows the value from state estimates of licensing to the Gallup survey results.

4. The Gallup survey

The survey we used for our analysis was a national survey conducted by the Gallup Organization from May to August of 2006. The random digit dial phone survey began with the Bureau of Labor Statistics' (BLS) American Time Use Survey (ATUS) and included Current Population Survey (CPS) questions on demographics, industry, occupation, earnings, and education. The response rate for the entire survey was 37 per cent, which is reasonable for a private survey of this kind.[2] The total number of respondents in the survey was 3,982, and 2,037 of these individuals were employed in the reference week. Because of missing data the totals for the analysis of wages and perceived competence is generally lower. Because we dropped non-respondents from our estimates we only include individuals who gave answers to the questions. We examined the individual responses for those who stated that they needed a licence to do their work and found few surprises. When we examined individual responses for accountants, teachers and barbers, they typically stated they needed a licence to do their work. Similarly, factory labourers, carpenters and economists stated they did not need a licence.

In Figure 2 we show the distribution of licensed occupations by education, race, union status, public or private sector and gender from the Gallup

FIGURE 2
Occupational Licensing by Race, Gender, Education, Occupation and Industry.
(a) Education. (b) Licensing by Race. (c) Licensing by Union Status, Industry and Gender.



survey. The results indicate that licensing increases with education: more than 40 per cent of those with post-college education are required to have a licence compared with 11 per cent for those with less than a high school education. Both African Americans and Hispanics have a higher percentage of licences than do whites or Asians. In panel (c) of Figure 2 the results show that union members are more likely to be licensed, reflecting the large number of teachers and nurses who tend to be both union members and licensed. Government workers are more likely to have a license than non-government workers, and there is no difference in the licensing rate by gender.

We estimated correlations for key institutional variables among the variables in Figure 2 in order to further examine the stability of the relationships. We found that there is a 0.14 simple correlation between licensing and unionization, and a correlation of 0.14 between licensing and education. Even deleting the two largest unionized and licensed occupations, teachers and nurses, from the Gallup sample only drops the simple correlation between unionization and licensing to 0.12. Further, there is a 0.15 correlation between licensing and employment in government, and a simple correlation of 0.4 between government employment and unionization. All of these relationships for unionization are similar to national results estimated through the CPS (see Freeman and Kleiner 1999). As a further check on the consistency of the Gallup sample with other national surveys, we found that the estimates were within the plus or minus 3 per cent standard of error usually given within Gallup surveys. For example, the Gallup survey showed that 15.1 per cent belonged to a union compared with 12.5 per cent in the CPS for 2006.

5. Multivariate estimates

In Table 1 we summarize initial results of the impact of licensing on wages. Specifically, we augment a standard human capital earnings equation to include a dummy variable measuring whether a licence is required for the worker's job. We regard these estimates as mainly descriptive as licensed workers may differ from unlicensed workers in unobserved ways, even after we condition on education and occupation. If a licensed dummy is added to a standard wage equation, having a licence is associated with approximately 15 per cent higher hourly earnings ($p < 0.001$). The cross-sectional effect of licensing is remarkably similar to the estimated effect of belonging to a union (see Lewis 1986) and greater than an additional year of schooling.

Of course, jobs that require a licence may also require higher skill and thus may not be directly comparable, even conditional on education, experience, and other covariates variables in the regression. We can partially adjust for unobserved factors by controlling for occupation dummies. Within some occupations (e.g. electricians), identical jobs are licensed in some states and not in others. In an additional specification in column 2, we first controlled for one-digit occupational dummies to further test for the robustness of our initial results, and found no change in the coefficient value or its significance. Next in column 3 we add 306 detailed occupation dummies. We continued to find an estimate of the licensing variable of around 15 per cent, and the estimate continued to be statistically significant. As a further sensitivity test of our results, we dropped the two largest individual occupational groups,

TABLE 1
Multivariate Estimates of the Influence of Occupational Licensing on Hourly Wages with and without Occupation-Specific Controls

Variables	*(1)* *Log wage*	*(2)* *Log wage*	*(3)* *Log wage*
Licensed	0.15***	0.15***	0.16***
	(0.03)	(0.04)	(0.05)
Experience	0.03***	0.03***	0.03***
	(0.004)	(0.004)	(0.005)
Experience^2/100	–0.0435***	–0.0388***	–0.0398***
	(0.008)	(0.008)	(0.009)
Education	0.11***	0.10***	0.08***
	(0.006)	(0.007)	(0.008)
Female	–0.29***	–0.26***	–0.23***
	(0.03)	(0.03)	(0.04)
Union	0.12***	0.13***	0.14***
	(0.04)	(0.04)	(0.04)
Government	–0.08**	–0.08**	–0.01
	(0.04)	(0.04)	(0.04)
Constant	1.28***	1.38***	0.92***
	(0.07)	(0.09)	(0.14)
Observations	1614	1614	1614
R^2	0.36	0.40	0.55
One-digit occupation controls	No	Yes	No
306 three-digit occupation controls	No	No	Yes

Note: Robust standard errors in parentheses.
*** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$.

teachers and nurses, from our analysis, which reduced the sample size to 1,519, and found the results were still about a 15 per cent wage effect for the remaining licensed occupations. The resilience of the licensing wage effect to occupation controls suggests that the cross-sectional estimate is not severely biased by omitted variables associated with licensing at the occupation level.

In Table 2 we show the influence of both unionization and licensing on wage determination and their interactions. In column 1 we estimate a standard union wage equation with controls usually found in the CPS. The estimates in column 2 show that with the same controls as in Table 1 along with one digit occupational controls, including licensing, the wage equation reduces the union impact on wages from 0.14 to 0.13. In column 3 we expand the estimation procedure to include the union and licensing variables and their interactions. We find that individuals who are both in a union and are licensed earn a wage premium approximately 24.6 per cent higher than those individuals who are neither in a union nor licensed (Halvorsen and Palmquist 1980). The interaction term for licensing and unionization is negative and significant, which is suggestive that being both licensed and in a union may result in diminishing returns to wages that is consistent with some potential substitution between these two institutions.

Does licensing reduce variations in wages similar to the impact of unions? To examine the role of each labour market institution on the wage structure

TABLE 2
Influence of Unionization and Licensing on Wage Determination

Variables	*(1)* *Log wage*	*(2)* *Log wage*	*(3)* *Log wage*
Union	0.14***	0.13***	0.19***
	(0.04)	(0.04)	(0.05)
Licensed		0.15***	0.18***
		(0.04)	(0.04)
Union × licence			0.15**
			(0.08)
Experience	0.03***	0.03***	0.03***
	(0.00)	(0.00)	(0.00)
Experience^2	–0.04***	–0.04***	–0.04***
	(0.01)	(0.01)	(0.01)
Education	0.10***	0.10***	0.10***
	(0.01)	(0.01)	(0.01)
Female	–0.26***	–0.26***	–0.26***
	(0.03)	(0.03)	(0.03)
Government	–0.07*	–0.08**	–0.08**
	(0.04)	(0.04)	(0.04)
Constant	1.37***	1.38***	1.38***
	(0.09)	(0.09)	(0.09)
Observations	1614	1614	1614
R^2	0.387	0.396	0.397
10 one-digit occupation controls	Yes	Yes	Yes

within large occupational categories, we grouped workers into three large occupational categories: professional and technical workers, managers and administrators, other occupations, and all in the Gallup sample. Large occupational groups are required to have sufficient observations for a meaningful analysis of wage dispersion in similar types of jobs. Within these categories, we compared the residual wage variability from a modified model presented in Table 1, column 1 between licensed and unlicensed positions, and between union and non-union workers using a similar estimation procedure to the one that estimates the impact of union wage policies on the wage structure within establishments (see Freeman 1982).[3] We examine the mean within category squared residual of the log of wages in licensed occupations and in unionized ones controlling for human capital characteristics. We then present the difference in the coefficient value of the mean squared error in the next column. The results in Table 3 show that the dispersion of wages of union members is lower than for non-union members since the p-values are different from zero with a value of 0.002. Similarly, the measure of dispersion of wages among licensed jobs is about the same or only slightly smaller than the unregulated ones, and the p-value for difference in the mean squared error is not significant.

If licensing is associated with increases in earnings, does it result in more competent services? Our evidence here depends on subjective self-assessments. We model answers to a question that asked: ‘On a scale of 0 to 6, where 0 means not at all and 6 means very much, how competent did you

TABLE 3
Licensing and Union Wage Dispersion for Broad Occupational Categories and Overall Variations

Occupation category	*Mean within category squared residual*		*Coefficient: (difference in mean squared error license — non-licensed)*	*p-value*
	Licensed	*Unlicensed*		
Managers	0.48	0.37	0.11	0.293
Professional/technical	0.33	0.32	0.01	0.749
Other	0.34	0.38	–0.03	0.756
All	0.36	0.34	0.02	0.567
Occupation category	*Mean within category squared residual*		*Coefficient: (difference in mean squared error union — non-union)*	*p-value*
	Union	*Non-union*		
Managers	0.25	0.41	–0.15	0.35
Professional/technical	0.22	0.35	–0.13	0.017
Other	0.14	0.40	–0.26	0.06
All	0.22	0.37	–0.15	0.002

$N = 1680$.

feel while you were at work yesterday?' Using responses to this question as the dependent variable, we estimate a multivariate equation with standard human capital and labour market variable as controls. Table 4 shows our results using OLS and an ordered probit. We find that individuals who have a licence perceive themselves as being more competent. In contrast, union members perceive themselves as no more competent than other workers. This is consistent with other studies that show the union productivity effect is close to zero.[4] Although a more precise measure of quality would gather information from consumers, these estimates show that self-reported abilities are higher for licensed workers. Moreover, the contrast with union members suggests that the results are not merely the consequence of a wage premium. If the results can be replicated for consumer outcomes, there may be support for the Shapiro (1986) model of greater ability, quality and costs of licensing for some consumers as a consequence of licensing.

6. Conclusions

Our study provides the first national analysis of labour market implications of workers who are licensed by any agency of the government in the USA. Using a specially designed Gallup survey of a nationally representative sample of Americans, we provide a preliminary analysis of the influence of this form of occupational regulation. We find that 29 per cent of the workforce is required to hold a licence, which is a higher percentage than that found in other studies that rely on state-level occupational licensing data or

TABLE 4
OLS and Ordered Probit Estimates of Standardized Competence Measure

Variables	*(1) Competence OLS*	*(2) Competence category probit*	*Marginal effects of an ordered probit estimate*[a]
Licensed	0.157***	0.24***	0.09***
	0.05	(0.07)	(0.02)
Education	0.017*	–0.004	
	(0.009)	(0.01)	
Age	0.007***	0.010***	
	(0.002)	(0.002)	
Union	–0.006	0.06	
	(0.07)	(0.085)	
Government	0.14**	0.19**	
	(0.06)	(0.08)	
Self-employed	0.05	0.038	
	(0.07)	(0.08)	
Non-profit	0.07	0.09	
	(0.08)	(0.093)	
Constant	–0.59***		
	(0.12)		
Observations	1,951	1,951	
R^2	0.025		

[a] Estimates of the ordered probit are from the second highest to the highest value.
Note: Standard errors in parentheses.
*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

single states. Workers who have higher levels of education are more likely to work in jobs that require a licence. This pattern suggests that our results are not spuriously reflecting driver's licenses. Union workers and government employees are more likely to have a licence requirement than are non-union or private sector employees. Our multivariate estimates suggest that licensing has about the same quantitative impact on wages as do unions — that is about 15 per cent — and that being both licensed and in a union can increase wages by more than 24 per cent. If these results hold up to further scrutiny, occupational licensing would be a much more pervasive phenomenon for the distribution of income than labour unions.

With the large and growing number of workers required to obtain an occupational licence, and the apparently large effect of licensing requirements on the labour market, we think it would be prudent for the government to measure the extent of occupational licensing in a manner similar to information that is collected for unions. To help this effort, we are in the process of developing a small number of questions on occupational licensing that can be added to a labour force survey, such as the CPS. These questions would help to answer more fully how much regulation is optimal, the effect of licensing on wages and productivity, and the type of regulation that is best suited for the emerging jobs in the workforce.

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Notes

1. The method used to calculate the percent licensed prior to 2006 first involved gathering the listing of licensed occupations in the BLS Occupation and Employment Survey. This was matched with occupations in the 2000 Census. If no match was obtained the occupation was dropped. From the Census the number working in the licensed occupation in each state was estimated and used to calculate a weighted average of the percent of the workforce in the USA that works in a licensed occupation.
2. Sample weights were developed to adjust for non-response. The pattern of time use closely matched the ATUS. All of the results we present rely on weighted estimates.
3. Our first-stage estimates that generated the residual value used education, experience-squared, public sector, licensed, union and gender, and we then estimated the impact of either licensing or unionization on the squared error. We compared the significance of the values generated in each category using the *p*-value.
4. This is consistent with the Fuchs *et al.* (1998) survey of economists of perceived union effects but lower than Freeman and Medoff's (1984) estimates in their survey of union productivity effects.

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