THREE ESSAYS ON LABOR MARKET REGULATION IN THE AMERICAN CONSTRUCTION INDUSTRY

By

Matthew Phillip Hinkel

A DISSERTATION

Submitted to
Michigan State University
in partial fulfillment of the requirements
for the degree of

Human Resources and Labor Relations—Doctor of Philosophy

ABSTRACT

THREE ESSAYS ON LABOR MARKET REGULATION IN THE AMERICAN CONSTRUCTION INDUSTRY

By

Matthew Phillip Hinkel

This three-article dissertation focuses on labor market regulation in the American construction industry. The United States faces two parallel crises: one with affordable housing supply, and one with maintaining residential construction labor standards. Historically, issues with labor standards have been addressed on public works through prevailing wage requirements. Labor standards—while good for workers—may increase construction costs; higher costs, in turn, negatively impact low-income families by reducing supplies of affordable housing. In Chapter 1 of this dissertation, I re-examine whether this tradeoff exists and, if so, its implications. I estimate that prevailing wage requirements add, at most, 6% to the costs of affordable housing construction. The implicit baseline for this paper is the current practices in the residential construction industry, including the cost advantages realized by contractors engaging in illegal and undesirable practices. An alternative baseline would be the cost of building affordable housing for contractors who abide by labor standards, classify their workers correctly and pay the required amounts in social insurance and taxes.

Informal employment, defined as the illegal misclassification of employees as independent contractors or employment of workers using cash-only payments, has long been rampant in the American construction industry. These actions rob workers of legally earned benefits, defund social programs, and undermine the competitiveness of law-abiding contractors. While enforcing labor laws has proved difficult, prevailing wage laws may make states abler to strengthen enforcement and limit informality. Under penalty of law, these regulations require

employers to submit weekly certified payrolls to government agencies on public works projects, which increases governmental oversight. In Chapter 3 of this dissertation, I use state-level data from 2010-2019 to examine the relationship between prevailing wage laws and informal construction employment. State prevailing wage laws, even those of weak and average strength, are associated with significant reductions in informality.

Lastly, Chapter 2 of this dissertation focuses on occupational licensing requirements in construction. Over time, disagreements have persisted over the effects of occupational licensing on markets and the appropriate role of government in the regulation of occupations. In Chapter 2, I exploit state variation in occupational licensing laws to examine labor market outcomes of occupational licensing in construction. Data on licensing comes from 2016-2019 Current Population Survey (CPS) data as well as a new 2019 data set on licensing requirements for the three primary construction occupations that require licensing in certain states: electricians, plumbers, and operating engineers. Consistent with prior literature, results suggest the presence of occupational licensing is associated with an 8.3 to 14.8 percentage point increase in earnings for electricians, plumbers, and operating engineers. Employment results are more mixed; while these results suggest occupational licensing is associated with a 1.2 to 1.3 percentage point increase in the proportion of workers employed as electricians, plumbers, or operating engineers, effects on the level of employment in these occupations were not statistically significant. In supplemental analyses I explore possible competing explanations for these employment findings.

With all my love, this dissertation is dedicated to Lauren.
In many ways, this degree is yours, too. Thank you for everything.

ACKNOWLEDGEMENTS

I would first like to thank the members of my dissertation committee, Professors Dale Belman, Peter Berg, Russell Ormiston, and Amanda Chuan, for all of their guidance and support throughout this process. As a first-generation graduate student who was 22 years old when I started the doctoral program, my learning curve has undoubtedly been steep. I could not have made it to this point without such a supportive, patient committee. I will forever be grateful to all of you.

Second, I would like to thank the School of Human Resources and Labor Relations for admitting me to the program. I am so appreciative of the opportunity you gave me, and I intend to represent the HRLR school well as I move into the next stage of my career.

TABLE OF CONTENTS

LIST OF TABLES

INTRODUCTION

This three-article dissertation focuses on labor market regulation in the American construction industry. The existence of multi-person firms suggests the existence of positive transaction costs (e.g., imperfect information, incomplete contracts), which leads to the conclusion that labor markets are "inherently and always and everywhere imperfect" (Kaufman 2007, p. 781) and need regulation.

However, recent shifts in the broader U.S. economy have made the task of regulating labor markets increasingly complex and challenging. These shifts are encapsulated by the notion of workplace "fissuring" (Weil 2014). Weil argues that "the large corporation of days of yore came with distinctive borders around its perimeter, with most employment located inside firm walls. The large business of today looks more like a small solar system, with a lead firm at its center and smaller workplaces orbiting around it" (Weil 2014, p. 43). This fissuring results from a heightened focus by firms on their core competencies and away from those activities not central to their profitability. This broad shift in the structure of the employment relationship stems from two interrelated changes. First, there has been a shift in views of capital markets. Capital markets no longer automatically view layoffs as a sign of weakness; rather, they increasingly view layoffs as routine churning in the labor market and as a signal that businesses are adequately streamlining their focus toward core competencies (Weil 2014). Second, technological change has significantly altered the economy in the last thirty years. Technological change has created new ways of designing and monitoring work, even work that takes place outside of the corporation's walls (Weil 2014).

Few sectors of the U.S. economy have felt the effects of fissuring more than construction. Informal employment, defined as the illegal misclassification of employees as independent contractors or employment of workers using cash-only payments, is rampant in the American construction industry (e.g., Ormiston, Belman, Brockman, & Hinkel 2020; Ormiston, Belman, & Erlich 2020; Juravich, Ormiston, & Belman 2021). Erlich (2020) argues that this stems from construction operating as a large gig economy: "Construction, in particular, was 'gig' long before the term *innovation economy* entered popular consciousness. Building projects start and finish, workers move from employer to employer, location to location, and the breaks in employment and pay caused by inclement weather all add to the inherently insecure nature of the trades" (Erlich 2020, p. 2). That is, labor supply and demand in construction operate on a *project-by-project* basis, with each construction project functioning as a separate "gig" for a highly mobile construction labor force. This has made informality a part of the business model of much of the construction industry.

These problems are particularly strong in specific parts of the industry. Recent work on residential construction (Ormiston, Belman, Brockman, & Hinkel 2020) paints a picture of a sector riddled with low-road labor and employment practices. The residential construction market is highly competitive, with small, labor-intensive firms all competing to complete limited-duration construction projects at the lowest cost possible, key to getting their project bid accepted. This has resulted in the prevalence of illegal practices such as wage theft, workers' compensation and tax fraud, and the exploitation of undocumented labor. Informality has become institutionalized via the usage of labor brokers, check cashing stores, and shell companies that create a tall order for regulators attempting to track illegal behavior in the industry (Ormiston et al. 2020). Enforcement is difficult because as resources available to regulators and tax agencies have declined, contractors (and workers) have gone to great lengths to hide this behavior from these agencies. As these illegal practices expand, law-abiding construction firms face two difficult choices: if they continue following the rules, they risk being shut out of the market and losing construction project bids as other firms reduce labor costs. If they choose to follow suit and operate informally, while enforcement is difficult, they may be caught by regulators. For those who elect to evade labor and employment law, their choice can be profitable: drywall contractors who misclassify employees as independent contractors or employ them off-the-books reduce their labor costs at least 30% (Ormiston et al. 2020). This is because misclassified and off-the-books workers lose access to unemployment insurance, workers' compensation, minimum wages, overtime payments, anti-discrimination defenses, and the right to unionize (Erlich 2020).

This incentive to reduce labor costs does not only come from the structure of the industry. It also stems from the structure of the bidding process for public works projects. Historically, public works construction was often plagued by various forms of corruption between contractors and government officials. To counter this, legislatures have required public bodies, in most instances, to award projects to contractors who offer the *lowest qualified bid*. This low-bid requirement contrasts with the private sector, where owners can consider other factors (e.g., contractors' reputation and experience) alongside the bid price in determining the winning bid. This low-bid requirement also incentivizes contractors to find ways, both legal and illegal, to reduce labor costs for public projects; the easiest way to lower bid prices is paying labor less. This creates a problem for local contractors bidding on public projects: outside competitors enter and submit bids reliant on lower labor costs. As a result, these outside contractors have the lowest bids and win projects, thereby shutting out local contractors from the market.

How can these labor and employment issues be solved? One such regulation that offers a potential solution is a prevailing wage law. Prevailing wages exist for federal projects via the federal Davis-Bacon Act and for state projects via "Little Davis-Bacon" state laws. The Wage and Hour Division (WHD) of the U.S. Department of Labor determines mandated prevailing wage rates by trade, area, and type of construction (e.g., residential, building, highway) by surveying private construction employers where federal projects are scheduled and averaging the surveyed wages. Prevailing wage laws require that employees on public works receive the same hourly compensation paid for similar work on similar private projects. These laws can therefore be conceptualized as minimum wages applied to blue-collar construction workers on public projects, determined by occupation, type of construction, and locality. On public projects with low-bid requirements, prevailing wage laws compel contractors to compete on engineering, project management, and the skills of their labor force, rather than labor costs.

This regulation has a few potential benefits. First, by helping to correct market distortions caused by low-bid requirements on public projects, prevailing wage laws can preserve local labor standards by protecting local contractors (and workers) from being undercut by low-wage contractors. Second, prevailing wage laws also feature inherent enforcement mechanisms that may help minimize illegal behavior in the industry. Under prevailing wage requirements, contractors are required to follow numerous rules (under penalty of law) that involve increased oversight of public projects. This includes the requirement that construction contractors file a weekly certified payroll, which must include lists of employees' names, Social Security numbers, work classification, daily hours worked, and other information. The penalties for violating prevailing wage standards include make whole remedies, fines, and incarceration. Although some payroll fraud undoubtedly takes place on prevailing wage sites, Sinyai and Galeas's recent survey of construction in the District of Columbia reports that evidence of wage theft (i.e., payment by personal checks or cash, and the failure to pay overtime) are not reported by those working on public projects (Sinyai and Galeas, 2021). Thus, prevailing wage laws may offer a solution to labor market problems that stem from both the structure of the industry as well as the structure of the bidding process for public projects.

Though prevailing wage laws may help resolve these issues, the question arises whether these regulations simultaneously increase public project costs, and if so, by how much. This possibility is particularly critical in the affordable housing context, where the U.S. faces a well-known affordable housing supply crisis. The effect of prevailing wage laws on affordable housing construction costs has been the subject of much scholarly debate, with prior literature suggesting that prevailing wage laws increase these costs by 5 to 37%. Research which reconciles this wide variation in cost effects and provides robust estimates will aid policy decisions about affordable housing programs. The magnitude by which prevailing wage laws raise affordable housing construction costs affects governments' ability to build additional affordable housing units, a step central to resolving the affordable housing shortage. In sum, there are two parallel crises affecting the construction industry, and residential construction in particular: (a) the affordable housing shortage, and (b) deteriorating residential construction labor standards. The second crisis might be alleviated via prevailing wage laws on affordable housing construction, but would this come at the price of raising the cost and reducing the supply of new affordable housing? Chapters 1 and 3 explore these dual crises in greater detail and the extent of this potential tradeoff in effects of prevailing wage laws in U.S. construction.

A related issue the construction industry (and the broader economy) faces is declining unionization. Construction union density has steadily declined from 18% in 2000 to only 13% in 2021 (Bureau of Labor Statistics 2022). While unionization has long been in a precipitous decline, occupational licensing, defined as a state-issued credential that must be obtained in order to legally work for pay in certain occupations (Blair & Chung 2019), is one of the fastest-growing institutions in the United States labor market. About 25 percent of the U.S. workforce is required to obtain a license to work in their given occupation; in 1950, that figure stood at 5 percent (Kleiner & Krueger 2010, 2013). The simultaneous decline of unionization and rise of occupational regulation has led scholars to suggest that these two institutions may be acting as substitutes for worker protection in the U.S. (Kleiner & Krueger 2013). That is, it is thought that occupational licensing has increasingly replaced unionization as a means of protecting the wages and employment of skilled workers over the last few decades.

Occupational licensing, also known as the right to practice, is the most restrictive of three forms of occupational regulation in the U.S.; the two less restrictive forms are registration and certification (Kleiner 2000). No prior licensing studies have focused on construction; recent estimates suggest 18.1% of workers in construction and extraction occupations are either licensed or certified (Bureau of Labor Statistics 2019). In construction, licensing requirements are primarily concentrated in three high-skilled occupations: electricians, plumbers, and operating engineers (hereafter referred to as “licensed occupations”). Operating engineers are responsible for operating and maintaining heavy equipment and machinery on construction sites. Licensing requirements are therefore largely clustered in a specific high-skilled portion of the industry’s labor force, leaving less-skilled segments (e.g., carpenters, construction laborers, painters) relatively free of licensing requirements. The declining reach and impact of unionization in the industry and the higher prevalence of licensing in high-skilled construction occupations make studying occupational licensing in construction an important matter for public policy. Further, unlike industries focused on in prior licensing studies, construction is unique in largely not requiring college degrees for entry; a high school diploma is standard, even for the high-skilled occupations that feature licensing. Examining labor market outcomes of occupational licensing is all the more important in a context where workers typically do not have as many other ways to readily signal work quality.

Though occupational regulation has historically been among the most examined institutions in labor economics, disagreements have persisted over its effect on markets and the appropriate role of government in the regulation of occupations (Kleiner 2000). Neoclassical economists have primarily viewed licensing as rent-seeking behavior that interferes with free markets and creates inefficiencies and rents only shared by those who obtain licenses (Friedman 1962; Friedman & Kuznets 1945). Other theory has focused on human capital investments. According to this theory, licensing increases incentives for workers to invest in their human capital and increase their skill levels, given the higher earnings and employment opportunities available to workers who do so (Shapiro 1986). This raises the question as to whether occupational licensing primarily functions as a rent-seeking device to reduce competition with qualified workers.

The most common view in the literature is that licensing acts primarily as a rent-seeking device and a barrier to entry that restricts labor supply to a licensed occupation, thereby driving up the price of labor. Workers with lower skill levels may self-select into non-licensed occupations, leaving licensed occupations with higher average skill levels and employers competing for smaller supplies of workers (Kleiner 2000). This lower labor supply lowers employment but raises wages (Rottenberg 1980; Kleiner 2000). Other literature has presented a different view: by serving as a credible signal of worker quality and training, licensing may increase demand for services by resolving informational asymmetries between workers and consumers in markets where work quality may be difficult to ascertain. This can be especially salient in contexts where low-quality work poses great health and safety risks. To the extent that licensing signals quality and resolves asymmetric information, this increases labor demand, thereby increasing both employment and wages (Leland 1979; Carollo 2020).

While both of the above theoretical frameworks predict positive earnings effects of licensing, employment in licensed occupations could increase or decrease depending on the relative importance of these competing supply and demand effects (Carollo 2020). In construction, while licensing can lower employment in licensed occupations via barriers to entry, it can also signal higher levels of worker training, providing greater consumer protections from negative effects of poor workmanship. In Chapter 2, I follow prior literature by analyzing associations between licensing and earnings as well as employment. Additionally, I add to the literature by exploring competing alternative explanations for the employment results and analyzing potential antecedents of occupational licensing in construction. While the estimates presented in Chapter 2 are not causal, this paper adds to the literature by studying occupational licensing in a construction context that follows the broader economic pattern of declining unionization while also having comparably fewer ways for workers to signal quality to employers and consumers.

CHAPTER 1: SHOULD PREVAILING WAGES PREVAIL? RE-EXAMINING THE EFFECT OF PREVAILING WAGE LAWS ON AFFORDABLE HOUSING CONSTRUCTION COSTS

Abstract

The United States faces two parallel crises: one with affordable housing supply, and one with maintaining residential construction labor standards. Historically, issues with labor standards have been addressed on public works through prevailing wage requirements. Labor standards—while good for workers—may increase construction costs; higher costs, in turn, negatively impact low-income families by reducing supplies of affordable housing. In this paper, we re-examine whether this tradeoff exists and, if so, its implications. We estimate that prevailing wage requirements add, at most, 6% to the costs of affordable housing construction.

Introduction

Policy makers often face tough tradeoffs when designing policies that work at cross purposes. The potential tension between providing affordable housing to low-income families and regulating construction labor standards is a case in point. Labor standards—while good for workers—may increase construction costs; higher costs, in turn, negatively impact low-income families by reducing supplies of affordable housing. In this paper, we analyze whether this tradeoff exists and, if so, its implications.

Affordable housing units are residential housing that, through tax subsidies, vouchers, or other programs, are available at a lower rental cost than for comparable unsubsidized units. Affordable housing can be single family houses or units in multi-family buildings, apartment buildings or mixed-use developments. The availability of affordable housing for U.S. low-income families is in crisis, as low-income families near or below the poverty line face a burdensome shortage of affordable housing (Aurand, Emmanuel, Yentel, Errico, and Pang 2018).[1] Through tenant-based rental assistance, public housing, project-based rental assistance, and homeless assistance grants, federal and state programs seek to alleviate this shortage. Success in expanding the stock of affordable housing stems, in part, from keeping the cost of new projects low while maintaining quality.

The construction labor market is likewise in crisis. In 2017, construction and extraction workers had the second highest rate of working poor in the US—65 percent above the overall labor market average and behind only farming, fishing, and forestry occupations.[2] In particular, in 2010, 22.4% of Hispanics working in construction, 30% of the construction labor force, were below the poverty line. This was 172 percent higher than the overall rate of working poor in the US labor market and 52 percent higher than the rate of working poor among Hispanics in the overall labor market.

The construction labor market crisis is characterized by low-wage segments (particularly residential construction), payroll fraud (particularly in the nonunion segment), workplace accidents (particularly among Hispanics and immigrants) and skilled labor shortages.[3] Just as government affordable housing programs are key policies aimed at alleviating the affordable housing shortage, prevailing wage policies are aimed at preserving labor standards, promoting

[1] Only 37 available affordable housing units are available per 100 U.S. low-income households. The shortage is especially pronounced in the Western U.S. (i.e., Washington, Oregon, California, Nevada, and Arizona), where only 25 affordable housing units are available for every 100 low-income households (The National Low-Income Housing Coalition, Aurand, Emmanuel, Yentel, Errico, & Pang, 2018).

[2] As construction is more volatile than the overall economy, poverty rates among the working poor in construction rise faster in economic downturns than in the overall economy. In 2010, the rate of working poor in construction was 90 percent higher than the overall economy (https://www.bls.gov/opub/reports/working-poor/archive/workingpoor_2010.pdf table 4).

[3] In 2018 construction accounted for 19 percent of all workplace fatalities (https://www.bls.gov/iif/oshcfoi1.htm#2018). In 2015, Hispanic construction workers accounted for 35 percent of all construction fatalities (https://www.bls.gov/iif/oshwc/cfoi/hispanic-or-latino-fatal-injuries.htm).

training and collective bargaining, and mitigating workplace dangers rooted in inexperience, casual labor relations, and inadequate safety and skills training. The question arises whether affordable housing policy and public construction labor standards work at cross purposes by increasing the cost of public construction.

The effect of prevailing wage laws (PWLs hereafter) on the outcomes of construction projects (e.g., costs, safety, training, employment of under-represented minorities) have been matters of public controversy and scholarly debate since the passage of the first PWL in Kansas in 1891 (Philips 2005). PWLs require that employees on public works receive the same hourly compensation paid for similar work on similar private projects. The federal prevailing wage for construction is determined by the Wage and Hour Division of the U.S. Department of Labor by trade, type of project and location. At present, the federal Davis-Bacon Act and the "little Davis-Bacon Acts" in 27 states establish prevailing wage rates on public construction projects. Ongoing controversies and focused political action have resulted in the repeals of 6 state laws in the last decade; further efforts can be expected in the next several years. Opponents of these laws argue they are raids by special interests, construction unions and signatory employers, on the public treasury that prevent citizens from receiving the full value of their taxes. Supporters argue PWLs offset the market distortions caused by low bid requirements on public works. Prevailing wage laws level the playing field in terms of worker compensation and compel contractors to compete on engineering, project management, and the skills of their labor force, rather than labor costs.

The body of peer-reviewed empirical literature finds that PWLs are associated with increased training of the construction labor force and better safety performance, but do not increase the costs of constructing schools or road maintenance (e.g., Azari-Rad, Philips, and Prus 2002, 2003; Bilginsoy 2005; Duncan, Philips, and Prus 2014; Duncan 2015). In contrast, evidence on the effect of PWLs on affordable housing broadly support the view that PWLs increase costs; however, estimates of the magnitude of their effects vary widely between studies, with estimates ranging from 5 to 37% (Dunn, Quigley, and Rosenthal 2005; Littlehale 2017; Palm and Niemeier 2018). Research which reconciles these varying cost effects will aid policy decisions about affordable housing programs. The magnitude by which PWLs raise affordable housing construction costs affects governments' ability to build additional affordable housing units, a step central to resolving the affordable housing shortage.

The current study advances DQR (2005) and Littlehale (2017) with two contributions. First, DQR introduced the use of instrumental variables into the affordable housing literature. The prevailing wage may be endogenous because of correlations between high construction wages, construction costs and use of the prevailing wage. It may also be the result of developers choosing to participate in California state affordable housing programs that offer advantages but also come with a prevailing wage requirement. Internal evidence from DQR's work, however, suggests that their instruments did not meet contemporary relevance requirements for unbiased IV estimates.

Second, we use this exploration of endogeneity to obtain robust estimates of the effects of PWLs on the costs of affordable housing construction. Using data from a state of California affordable housing survey, we instrument the first stage with information on projects' use of programs with prevailing wage requirements. We systematically test our instruments for relevance, discuss the exclusion restriction for each instrument, and consider variants of our models. We use this same data to examine whether DQR's specification meets these criteria.

Using data collected by a California state survey of new affordable housing construction in 2001–2011, we find endogeneity is sensitive to the instruments in the first stage. Our preferred construction cost model, in which the prevailing wage is endogenous, uses two instruments.[4] Whether OLS or IV estimation is appropriate depends on specific exclusion restrictions, restrictions not subject to direct statistical testing. We also replicate DQR's specification and preferred instruments using our more recent data and demonstrate their instruments were too weak to reliably estimate the second stage cost model. We conclude that the weakness of these instruments is likely the cause of their very large prevailing wage cost estimates.

We estimate both OLS and IV models of the prevailing wage cost effects. Our OLS models indicate the PW raises costs by 5–6%. The PW is endogenous in our preferred IV model; it has a negative, non-statistically significant cost effect of -6.3%, and our instruments pass a test for relevance and the endogeneity of the prevailing wage. It is, however, estimated with limited precision; a 95% confidence interval includes our OLS estimates. There is then evidence that PWLs increase affordable housing construction costs by 0-6%.

Background

The federal Davis-Bacon Act of 1931 requires contractors and subcontractors working on federally funded or assisted projects exceeding $2,000 to pay their laborers no less than the wages and fringe benefits that prevail for laborers and mechanics performing similar work on similar projects in the same geographic area (U.S. Department of Labor n.d.). The Wage and Hour Division (WHD) of the U.S. Department of Labor determines mandated PW rates by trade, area, and type of construction (e.g., residential, building, highway) by surveying private construction employers where federal projects are scheduled and averaging the surveyed wages. PWLs are designed to set the compensation paid on public projects to that paid on private projects for similar work in each locality (U.S. Department of Labor 2015). PWLs can then be

[4] Concern over the validity of our Density Bonus instrument is addressed by estimating a second IV model in which Density Bonus is treated as a control. The remaining instrument remains valid, but the PW fails an endogeneity test.

conceptualized as minimum wages applied to blue-collar construction workers on public projects, determined by occupation, type of construction, and locality.

Twenty-seven states and numerous cities have "little Davis-Bacon" laws applying to state and local public works and publicly financed construction, some substantially predating the federal act. The stated goals of these regional laws differ, ranging from the preservation of local labor standards to ensuring a workforce that is safer and supporting apprenticeship training to provide a more skilled workforce (Duncan and Ormiston 2018).[5] Those who oppose these laws view them as a raid by unions and organized contractors on the public till, arguing that they interfere with the operation of free markets and create inefficiencies and rents only shared by the beneficiaries of these laws.

With the support of the American Legislative Exchange Council (ALEC), legislatures in Indiana, West Virginia, Kentucky, Michigan, Wisconsin, and Arkansas have recently debated and repealed their PWLs (ALEC 2020; Duncan and Ormiston 2018). These repeal campaigns have driven the growth in peer-reviewed literature examining PWLs. Duncan and Ormiston (2018) report that in addition to increasing craft workers' wages, improving workplace safety, and increasing access to registered apprenticeship programs, the body of the literature finds that PWLs do not raise costs of school construction and highway maintenance projects (see Duncan and Ormiston 2018 for a thorough literature review).[6]

Affordable Housing: A Break from the Consensus

[5] States' methods of determining a prevailing wage varies considerably. Some use the federal prevailing wage, others use a survey to determine the wage, and others adopt the collectively bargained wage.

[6] The repeal campaigns have also been accompanied by non-peer reviewed literature on the effects of prevailing wage on construction costs that assert high-cost effects. This literature has recently become aware of the large estimates of DQR (2005), and DQR has been cited in these non-peer-reviewed studies. Re-examining DQR's estimates is therefore more important to founding policy decisions on sound research.

The effect of PWLs on the costs of affordable housing construction has been addressed in three peer-reviewed studies (DQR 2005; Littlehale 2017; Palm and Niemeier 2018). California is the only state that covers both public and private affordable housing projects in its state law (DQR 2005).[7] Each find that PWLs result in higher project costs for affordable housing, with estimates ranging from 5 to 37%.

All three studies use detailed California state surveys of affordable housing construction. DQR (2005) analyzed the construction of 205 affordable housing projects in California between 1997 and 2002. They estimated OLS models in which the prevailing wage was treated as exogenous and IV models in which the prevailing wage was treated as endogenous. The OLS models found that PWLs increased the costs of affordable housing projects by 9-11% while the IV models estimated a 19-37% increase in costs (DQR 2005).[8] [9]

Two more recent studies, Littlehale (2017) and Palm and Niemeier (2018) use OLS to examine the affordable housing issue. Using the most recent version of the California affordable housing study, Littlehale (2017) examined 321 affordable housing projects in California built from 2001 to 2011. Using a more fully specified OLS model, he reports a PW effect of 5-7%, statistically significant but smaller than DQR. The most recent study (Palm and Niemeier 2018)

[7] California's expanded its PWL to include privately as well as publicly funded affordable housing projects in 2001 (DQR 2005). Private sources are primarily incentivized by the federal Low-Income Housing Tax Credit (LIHTC) (National Housing Law Project 2017). This law does not include a prevailing wage requirement.
[8] In DQR (2005), the OLS models regressed the natural log of total project costs on explanatory variables including the number of units in a project, the fraction of units meeting affordability guidelines, and structure type (e.g., two-story, townhome, single-family). The authors preferred estimate found that the PW increased costs by 9.7 percent.
[9] In the first stage of their 2SLS model, two models were run: the first model included all project characteristic control variables from the OLS models, except for the number of units in each project. The second model added in the number of units as an additional control. First-stage instruments included the fraction of yes votes on various California propositions in the 1990s, the fraction of voters registered as Democrats, and the percentage of the population over 40 years old (DQR 2005).

examined housing projects built between 2008 and 2016 in California's four largest metropolitan areas. The authors report statistically significant cost effects of 15-16%.[10]

In sum, while PWLs have been shown to not increase costs of school construction and highway maintenance projects, they apparently increase costs of building affordable housing in California. This is plausible, both because California has one of the strictest PWLs (Duncan and Ormiston 2018; Belman, Ormiston, Petty, and Hinkel 2020) and because residential contractors have fewer opportunities to offset the effects of PWLs. PWLs are then unlikely to affect affordable housing construction in a fashion like other sectors of construction. Improving worker productivity, quality of work, safety outcomes or speed of completion is more difficult in affordable housing construction, because of the less educated and less skilled workforce used by residential contractors and their lesser use of sophisticated capital (Blankenau and Cassou 2011). While the gains available via PWLs are more readily available to school and highway construction contractors, they are more likely to increase labor and total project costs in affordable housing. Affordable housing then provides a strong test of the cost effects of the prevailing wage under unfavorable conditions.

Research which reconciles these varying PWL cost effects can aid policy decisions about affordable housing programs.[11] The magnitude by which PW requirements raise the costs of affordable housing affects governments' ability to build additional affordable housing units, a

[10] Palm and Niemeier (2018) estimate OLS models of affordable housing projects for four large metropolitan areas from 2008 to 2016. They find PWLs increase construction costs by 15 to 16%. Their estimates are not comparable to other estimates, as their sample was limited to San Francisco, San Diego, Sacramento, and Los Angeles and do not incorporate controls for the year in which construction was begun. Later in this paper, we show that the omission of year controls greatly increases the magnitude of the prevailing wage effect.

[11] The importance of this issue is reflected in several non-peer-reviewed studies claiming that PW requirements significantly raise the cost of affordable housing construction. For example, the New York's Empire Center for Public Policy suggested that PW inflates costs by as much as 25% (McMahon & Gardner 2017). A study published by the New York Center for Urban Real Estate suggested PW requirements raise costs by 25-30% (Vitullo-Martin 2012). These studies suffer methodological flaws (see Duncan & Ormiston 2018 and Ormiston, Belman, & Hinkel 2018), but influence public discourse on PW laws.

step central to resolving the affordable housing shortage. Robust estimates are needed to determine how much affordable housing is sacrificed to PW requirements. In addition, residential construction is growing rapidly, and with it, employment. Employment rose each year from 2011 to 2016, with 125,000 more workers employed in residential construction in 2016 than 2011 (Quarterly Census of Employment and Wages). Policies that directly affect this industry, such as prevailing wages, now affect a larger number of workers.

Issues with the structure of the residential construction industry also complicate assessing the impact of PWLs on affordable housing project costs. Illegal, low-road labor practices (e.g., cash-only payments, wage theft, worker misclassification, workers' compensation and tax fraud) are rampant in residential construction; up to one fifth of the national construction workforce is misclassified as independent contractors or paid off-the-books, where workers are paid under the table in cash (Ormiston, Belman, Brockman, and Hinkel 2020; Ormiston, Belman, and Erlich 2020; Juravich, Ormiston, and Belman 2021). These profitable illegal practices have become institutionalized in the residential sector. Research finds that drywall contractors who misclassify employees as independent contractors or employ them off-the-books reduce their labor costs at least 30% (Ormiston *et al*. 2020). These practices affect not only the workers and their families but also the public at large, as medical services that should be covered by workers compensation become burdens for public hospitals. This reduction in public tax receipts is taken on by the public and the evasion of mandated employment programs makes it more difficult for law abiding firms to operate profitably, driving them from those segments of construction where this behavior is endemic.

The prevalence of these practices in residential construction has economic and legal sources. Economically, the residential sector closely resembles a highly competitive market, with small, labor-intensive firms all competing to complete limited-duration construction projects at the lowest possible cost, key to getting their bid accepted. The easiest way for contractors to reduce costs (and lower bid prices) is to lower labor costs; contractors are therefore incentivized to misclassify their workers or employ them off-the-books. Misclassified and off-the-books workers lose access to unemployment insurance, workers' compensation, minimum wages, and overtime payments, allowing firms to avoid these costs and lower their bid prices (Erlich 2020). On the legal side, the lack of communication and coordination between agencies responsible for regulating the industry, the dwindling enforcement resources available to them, and the development of institutions such as labor brokers, check cashing stores, and shell companies support these practices (Ormiston *et al*. 2020; Juravich *et al.* 2021).

It is far more difficult to engage in this kind of egregious behavior on PW projects. Under Davis-Bacon, employers are required, under penalty of law, to file a weekly certified payroll which lists all employees' names, Social Security numbers, work classification, daily hours worked, rates of pay, payments of Social Security and tax withholdings, with deductions for benefit programs, total deductions, and net wages for the week. Further, under PWLs, contractors are not permitted to "hire' individuals as independent contractors (Wage and Hour Division). The penalties for violating prevailing wage standards include make whole remedies, fines, and incarceration. Although some payroll fraud undoubtedly takes place on prevailing wage sites, Sinyai and Galeas's recent survey of construction in the District of Columbia reports that evidence of wage theft (i.e., payment by personal checks or cash, and the failure to pay overtime) are not reported by those working on public projects (Sinyai and Galeas, 2021). In contrast, 44 to 58% of those employed on private projects report wage theft. At least part of the cost increases attributed to PWLs laws in affordable housing construction may be associated with employers being less able to evade labor and employment law.[12]

The current study also provides insight into the controversies over the purpose and channels through which PWLs act. As discussed in the introduction, opponents view these laws as a raid on the public treasury by special interests. Supporters argue that it corrects a distortion in the public bidding process that compels the public body to deviate from private sector practices in construction markets. Going back to Boss Tweed and Tammany Hall and before, there have been issues with corruption on public works projects. Public officials would arrange overly generous payments to contractors in return for kickbacks and other shameful practices. Since the progressive era, legislatures have countered this by requiring public bodies to accept the lowest qualified bid.[13] This low-bid requirement contrasts with the private sector, where owners can consider other factors (e.g., contractors' reputation and experience) along with bid price in determining the winning bid.[14] The low bid requirement places great pressure on contractors to develop practices, including the gaming of laws, to reduce their bid price; if they do not, they risk losing the bid. Although better engineering techniques and a better-trained workforce lower costs, the easiest way to lower bid prices, particularly for less able contractors, is to pay labor less. Such contractors are motivated to search for pools of the least expensive vulnerable labor they believe can complete the work. In the presence of low bid requirements, prevailing wage laws level the playing field in terms of worker compensation and compel

[12] Enforcement of California's requirements is more comprehensive than the federal Davis-Bacon Act, as building trades unions are directly involved in monitoring projects and informing the state of likely violations (discussion with Littlehale, June 2020).
[13] Bids are considered qualified if the contractor can obtain a bond for the project, a modest hurdle in most instances.
[14] An increasing exception to this low-bid requirement are responsible bidder ordinances (RBOs). RBOs outline sets of criteria a contractor must meet to be awarded contracts for public projects (e.g., past compliance with environmental and safety regulations, prevailing wage compliance, funding of apprenticeship progams). Rather than awarding projects to the lowest bidder, public bodies under RBOs award projects to the lowest *responsible* bidder. These need not be the lowest overall bids, and in many cases, they are not.

contractors to compete on engineering, project management, and the skills of their labor force, rather than labor costs.

Empirical Contribution

The current study advances DQR (2005) and Littlehale (2017) with two contributions. First, DQR introduced the use of instrumental variables into the affordable housing literature. We build on their work by applying contemporary methods and systematically testing the reliability of their IV estimates, the validity of their instruments, and the endogeneity of PWLs. Second, we develop robust estimates of the effect of PWLs on the costs of building affordable housing. The difference between DQR's OLS and IV estimates are very large and imply different policy responses.

The issue is which set of estimates, OLS or IV, provides policy makers with the best, least biased, and lowest variance estimates of the effect of PWLs. If, as contended by DQR, the prevailing wage is endogenous, OLS estimates of prevailing wage effects will be biased. If the prevailing wage variable incorporates observed and unobserved variables that are correlated with the dependent variable in the cost equation, there will be omitted variable bias in the estimate of the prevailing wage on construction costs. DQR's estimates suggest that correcting for omitted variable bias with IV methods increases the estimated effect of the prevailing wage from 10% (OLS) to 37% (IV), nearly a factor of four.

The application of IV methods comes with econometric requirements that, if not met, inflate estimated effects. As Littlehale (2017) pointed out, of the 17 instruments used by DQR, only 2, the fraction of yes votes on California Propositions 160 and 167 in 1992, were significantly related to the requirement that projects require PWLs; these two barely cleared a

10% significance test. The relevance requirement for IV estimation (i.e., that each instrument is statistically significant, *and* the instruments are sufficiently jointly significant) was not met.[15]

Likewise, the exclusion restriction for instruments was not adequately addressed by DQR (2005). All first-stage instruments are required not to have a direct causal effect on the dependent variable, project costs. The instruments can only impact project costs via the explanatory variable of interest, prevailing wages. DQR assert, but do not explain, that their instruments have no direct effect on construction costs (DQR 2005, p. 150). The authors also do not test for PW endogeneity, providing no evidence that their IV estimates are needed or are improvements over OLS. Taken together, their instruments are weak, possibly causing large inconsistencies and bias in IV estimates (Wooldridge 2013). Thus, DQR's weak instrument problem is likely the main reason why their IV estimates were nearly four times greater than their OLS estimates, as weak instruments can inflate IV estimates to unreasonably high levels.

That DQR's IV estimates of prevailing wage cost effects were unreliable and outside a reasonable range is supported by much of their range (19-37%) exceeding typical total labor costs for affordable housing. Using 2007 Economic Census of Construction data, Littlehale (2017) found that total payroll for California contractors specializing in apartment construction work (where most affordable housing work takes place) accounted for 23% of the net value of the work (i.e., exclusive of land costs). That is, in 2007, labor costs were 23% of net project costs—project costs exclusive of land costs—for apartment work in California.[16] DQR found

---

[15] The joint F statistic for instruments was less than 3, well below the value of 11.5 recommended with more than 15 first-stage instruments (Stock, Wright, & Yogo 2002).

[16] We also did similar calculations using 2017 Economic Census data to see if this result was robust and held over time. In 2017, for both the state of California and the entire United States, labor costs as a percentage of the net value of all construction work (NAICS=23) was 23%. This is notably identical to the percentage found by Littlehale (2017) for 2007 apartment work, suggesting that the 2007 result has remained stable over time.

that PWLs increase affordable housing project costs by as much as 37%; the notion that PWLs can impact these costs more than all labor costs combined is implausible.

Given the omission of two-stage methods and exploration of prevailing wage endogeneity by Littlehale (2017) and Palm and Niemeier (2018), the methodological defects of DQR's IV estimates, the availability of newer data, and the persistent shortage of affordable housing, we reconsider the effect of PWLs on the costs of building affordable housing and examine whether PWLs are endogenous.

Methods

Data and Sample

Building on DQR's work, we estimate both OLS and IV models with the survey of affordable housing construction commissioned by the State of California (CAAHCS; State of California 2014). This database was developed from projects that submitted applications to the California Tax Credit Allocation Committee (TCAC), *including* those for which developers sought federal housing tax credits. That is, it is a mix of state and federally funded projects; projects which receive federal funding are excluded from also receiving state funding. This survey was used by Littlehale (2017); a prior version was used by DQR. It includes 356 Low-Income Housing Tax Credit (LIHTC)-financed affordable housing projects built between 2001 and 2011. 233 of these projects were built under PW requirements; 133 were not. Including project built in 153 cities, it captures regional variation in construction methods and conditions. This sample was representative of the larger population of California affordable housing projects. In a comparison of the sample with the universe of the 995 affordable housing projects that begun between 2001 and 2011 across various dimensions (i.e., project type, type of tax credit awarded, project location), the sample was statistically like the larger universe, suggesting a representative sample (CAAHCS; State of California 2014).[17]

DQR (2005) recognized the possible endogeneity of the prevailing wage:

> "If projects located in higher-cost areas (for example, in highly urbanized areas) were more likely to be required to pay prevailing wages…then simple ordinary least squares regression models would falsely attribute these higher costs to the payment of prevailing wages" (DQR 2005, p. 149; Littlehale 2017, p. 124).

If the prevailing wage is endogenous, if it is significantly correlated with an error term which includes observable or unobservable omitted variables, estimates of cost effects associated with those omitted variables would be incorrectly attributed to the prevailing wage. Thus, a statistically significant OLS correlation between PWLs and project costs would not be indicative of a *causal* effect. To ensure that our estimates isolate this "true" effect on project costs, we follow DQR and explore potential endogeneity.

Measures

Our variables, their definitions, means, and standard deviations are in Table 1. Following prior work, the dependent variable, log project cost, is the natural log of total affordable housing project development costs exclusive of land costs. It includes both hard construction costs (e.g., labor costs, material costs) and soft construction costs (e.g., fees, financing).[18] The explanatory variable of central interest is prevailing wage, an indicator with a value of one if a project was subject to state PW requirements, zero otherwise.

---

[17] Our data was originally compiled and graciously provided by Blue Sky Consulting Group, with whom we signed a limited non-disclosure agreement.

[18] A 2019 study on the cost of residential housing in Northern California (Garcia 2019) suggests that hard costs accounted for approximately two-thirds of total development costs, while soft costs (including land costs) accounted for approximately one-third. Prevailing wage requirements directly impact hard costs, since they affect labor costs; any impact on soft costs is indirect. As our dependent project cost variable excludes land costs, which are unaffected by prevailing wages, our estimated effect is an *upper bound* estimate of the effect of prevailing wages on total project costs.

Our OLS model is

(1) $$\log y_i = \beta_0 + \beta_1 PW_{iry} + \boldsymbol{\beta_2 V_{iry}} + \boldsymbol{\beta_3 F_{iry}} + \boldsymbol{\beta_4 X_{iry}} + \mu_{iry},$$

where y_i is project cost, PW_{iry} is the prevailing wage indicator variable, and V_{iry}, F_{iry} and X_{iry} are groups of cost drivers. Variables are indexed by construction project (i), region (r), and year (y). Project specific drivers, $\boldsymbol{V_{iry}}$, include developer characteristics, fees associated with state permitting and impact fees, and project characteristics. Developer type and log developer employees capture whether the developer is the government, nonprofit, or another, and the number of workers employed by the developer[19]. Meetings 4 plus, permit and impact fees, and log permit and impact fees per square foot control for effects of state regulations such as obtaining building permits and requirements for community meetings; should increase total costs. For example, higher fees for obtaining building permits, all else equal, will raise construction costs (Littlehale 2017). Site and structure variables include log building area, parking-to-rentable building, square feet non-residential and non-parking, stories 4 plus, and residential square feet per unit. These capture cost effects specifically associated with the project.

The second group of controls involves drivers of affordable housing construction costs common across adjacent projects ($\boldsymbol{F_{iry}}$): average specialty contractor pay and project snow[20]. Construction projects share the specialty contractor workforce and share these common cost elements;[21] snow depths influence costs though their effect on project design and on windows of time for completion of various stages of construction (Littlehale 2017). Additional controls include log lenders, as a project's costs are affected by the number of loans that finance it;

[19] Developer size, measured by the number of employees, reflects developer capabilities and experience.
[20] Project snow is an indicator variable with a value of one if a project took place in a city with an average snow depth of 1 foot or more, and zero otherwise. Littlehale (2017) used an average city snow depth of one foot in January of 2017 as the threshold. (Littlehale 2017).
[21] Specialty contractors install electrical systems, water systems, prepare land and foundations as well as the carry out the architectural work of carpentry and steel erection including framing and drywall.

duration 24 plus, as longer durations are associated with greater costs; and log architect and engineer (A&E) cost per square foot. A&E costs are included in the models as a proxy for project complexity; more complex projects are more costly and require more time from architects and engineers. Lastly, the vector $\boldsymbol{X_{iry}}$ includes dummies for the (a) the year in which an affordable housing project began and (b) the region in which the project took place.

Our IV model is:

$$PW_{iry} = \beta_0 + \beta_1 CDBG\ Funding_{iry} + \beta_2 Density\ Bonus_{iry} + \boldsymbol{\beta_3 V_{iry}} + \boldsymbol{\beta_4 F_{iry}} + \boldsymbol{\beta_5 X_{iry}} + \mu_{iry}, \tag{2}$$

$$\log y_{iry} = \beta_0 + \beta_1 \widehat{PW_{iry}} + \boldsymbol{\beta_2 V_{iry}} + \boldsymbol{\beta_3 F_{iry}} + \boldsymbol{\beta_4 X_{iry}} + \mu_{iry}, \tag{3}$$

where equation (2) is the first stage prediction of the PW variable and equation (3) predicts the cost of affordable housing projects using the instrumented $\widehat{PW}$. Variables in equation (3) follow our discussion of the OLS model (equation 1), except for the use of the instrumented PW. Equation (2) predicts the use of prevailing wage requirements on a project with CDBG funding and density bonus serving as instruments, with $\boldsymbol{V_{iry}}$, $\boldsymbol{F_{iry}}$, and $\boldsymbol{X_{iry}}$ included as in the OLS model.

Our instruments, community development block grant (CDBG) funding and density bonus, meet the exclusion restriction (that an instrument does not affect costs except through the prevailing wage) and the relevance requirement (that an instrument must be partially correlated with the prevailing wage). The CDBG program is a state of California grant program that, among other purposes, is designed to partially fund the construction of affordable housing units. Any non-entitlement jurisdiction is eligible for CDBG funding, but double dipping is not allowed; a jurisdiction is only eligible for California CDBG funding if it does *not* also participate in the equivalent federal CDBG program.[22] Of note is that the federal CDBG program does not carry a prevailing wage requirement; federal projects are included in our data. Funding is geared toward smaller and more rural communities that otherwise may lack access to other resources, and the likelihood of receiving funding is independent of the expected costs of a project. Consistent with the exclusion restriction, California CDBG grants are not more likely to be awarded to large projects with higher costs.[23] Consistent with the relevance requirement, projects making use of CDBG funds are subject to prevailing wage requirements in most cases (California Department of Housing and Community Development 2018). CDBG funds should only impact costs via the explanatory variable of interest, prevailing wage, making it a valid instrument.[24]

California's density bonus law, like the CDBG program, is designed to spur the development of a range of projects, including affordable housing. It allows housing developers to obtain more favorable development regulations in exchange for offering housing for specific groups on a site or donating land for that purpose. It allows for up to a 35% increase in housing unit densities on a given site. The relevance criterion is met, as the relaxation of prevailing wage requirements is *not* among the regulations that developers can negotiate (Goetz and Sakai 2017).

The exclusion restriction should also be met, as eligibility is independent of project size and costs. Eligibility is based on how housing units are utilized and the specific vulnerable populations they are geared toward. Further, density bonuses are inherently exogenous; as Goetz

[22] Non-entitlement jurisdictions are defined as either (1) cities with populations under 50,000 that are not principal cities of Metropolitan Statistical Areas or (2) counties with populations under 200,000 (California Department of Housing and Urban Development 2018).
[23] Grant awards for housing projects vary. In 2017-2018, the city of Pacific Grove, CA was awarded $390,854 in state CDBG funding for a housing rehabilitation project, whereas the city of Arcata, CA was awarded $1,767,442 for a similar, much larger project (California Department of Housing and Community Development 2018).
[24] Public funds such as the CDBG program do *not* trigger prevailing wage mandates if the grants account for less than 5% of the total project budget. Thus, there are many cases where affordable housing projects can receive CDBG funding and *not* be subject to prevailing wages, as evidenced by a correlation between prevailing wages and CDBG funding of 0.15 found in the data. Thus, while they are positively related, it is not one-to-one.

and Sakai (2017) point out, developers who meet the requirements of the density bonus law are "entitled to receive the density bonus and other benefits as a matter of right" (Goetz & Sakai 2017, p. 2). California cities and counties are *required* by law to grant density bonuses to projects which meet *at least* one of the following guidelines: (1) at least 5% of units are reserved for very low-income residents, (2) at least 10% of units are reserved for low-income residents, and (3) at least 10% of units are geared toward foster youth, disabled veterans, or homeless persons, with subsidized monthly rents that reflect very low-income levels (Goetz and Sakai 2017). The cost impacts of density bonuses will occur only via the prevailing wage requirement, making it a second valid instrument.

We initially replicate Littlehale's OLS model and then estimate the parsimonious model, given by equation (1, OLS), where we eliminate control variables that are highly collinear. Next, we conduct the first-stage regression of prevailing wage on the proposed instruments, including all controls. This is given by equation (2, IV). If the two instruments are sufficiently jointly significant, ($F > 10$, Stock and Yogo 2005; Wooldridge 2013), we estimate the second stage regression, equation (3, IV). We regress log project cost on all proposed controls, instrumenting for the prevailing wage. Lastly, following this IV estimation, we test for endogeneity.

Results

Our results are provided in Tables 2, 3 and 4. Table 2 provides variants on the OLS model, Table 3 provides our preferred IV estimates, and Table 4 investigates sources of the differences between our work and that of DQR (2005). Our OLS estimates indicate that PW requirements increase affordable housing project costs by 5-6%. The IV estimate in Table 3 indicates that the prevailing wage is endogenous but does not have a causal effect on construction costs. This estimate is imprecise; its 95% confidence interval includes the OLS estimates. The reconciliation of our estimates with DQR suggest their very large estimates of the cost effects of PW are a product of invalid instruments as well as their specification.

Model 1 of Table 2 replicates Littlehale's 5% effect. Model 2 is a more parsimonious version dropping highly collinear controls. Our estimates are again like Littlehale: PWLs are associated with a statically significant 5.8% increase in affordable housing project costs. Model 3 is a replication of DQR's work and is discussed in Section 7.

Table 3 provides our IV estimate. The left-hand column presents the first-stage regression given by equation (2), which includes the OLS controls from Model 2 of Table 2. Both instruments are both statistically significant at the 1% level; their joint F statistic is 14.80, above the required threshold (Stock and Yogo 2005). The relevance requirement is also met.

The right-hand column of Table 3 presents the results of the second-stage regression, with CDBG funding and density bonus instrumenting prevailing wage. A χ^2 test for endogeneity weakly rejects the null of no effect; this estimate is then consistent and asymptotically more efficient than OLS (χ^2 =2.847; p = 0.092). With two excluded variables, the system is overidentified; a χ^2 test for overidentification does not reject the validity of the additional instrument (χ^2 = 0.309, p = 0.579). We do not find evidence that the overidentified system was problematic, further supporting our model. The IV estimate of PW (β = -.063) is opposite signed to the OLS estimates and is not statistically significant. The IV estimate is, however, imprecise, with a 95 percent confidence interval reaching from 9.6 to –22.7%; OLS estimates of 5-6% are within this interval. Our tests of significance, of endogeneity and of overidentification indicate the non-statistically significant IV estimate is unbiased and therefore preferred to the OLS estimate.

Any conclusions are complicated as the finding of endogeneity is a product of the use of both instruments; it is not supported when only the CDBG or Density Bonus are the lone instrument. We test the robustness of our IV results by estimating two single-instrument models: (a) an IV model with CDBG funding as the only instrument and Density Bonus as a control; and (b) a model with Density as the instrument and CDBG funding as a control. In model (a), the CDBG variable remains highly significant as an instrument, with an F-statistic of 26.32 . The point estimate for the $\widehat{PW}$ is -3.15, which is non-significant and similar to the two-instrument point estimate of -6.3. The χ^2 test for endogeneity, however, does not reject the null (χ^2 = 1.056; p = 0.30). In model (b), the Density Bonus is the lone instrument, its F-statistic is 6.5, well short of commonly used thresholds. The weakness of this specification is also reflected in the small magnitude of the prevailing wage coefficient (-.13) in the second stage estimate and the failure to reject the null of the exogeneity of the prevailing wage (χ^2 = 2.459; p = 0.12). In these instances, the prevailing wage is not endogenous. With these specifications, our OLS estimates of 5% to 6% are unbiased and reflect the true impact of PW on affordable housing costs.

Where does this leave our estimates? Although the endogeneity of the prevailing wage is sensitive to the instruments, this does not invalidate the two-instrument model. That model passes the standard tests for our instruments, for endogeneity, and for overidentification. As both instruments measure the impact of California state housing policies, it is not unreasonable to view them as reinforcing one another or measuring an unmeasured latent effect. Although our IV estimate of the effect of prevailing wage on the cost of affordable housing is imprecise and includes the OLS estimate within its 95% confidence interval, it meets the fundamental criteria for judging IV models.

Our analyses have produced two key findings. First, PW requirements are associated with OLS estimates of a 5-6% increase in affordable housing construction costs. These findings, as expected, are consistent with Littlehale (2017). Second, when we apply a two-stage, two-instrument IV framework, PW requirements have no significant effect on affordable housing construction costs. Although this estimate is imprecise and endogeneity is sensitive to specification, this paper is the first in the literature to suggest PWLs may have no causal effect. Taken together, we find that prevailing wages have, at most, a modest effect on costs of building affordable housing.

Reconciling Current and Prior Research

What are the sources of the differences between DQR's estimates and ours? Are these associated with differences in the data, specification, samples, or estimation techniques? We investigate this by supplementing our data with variables needed to reproduce DQR's model and systematically move from their preferred specification to ours.[25] Most notably, we replicate DQR's measures of the outcomes of ballot propositions with similar measures for fraction of yes votes for eight California ballot propositions from 2002-2010, obtained from the office of the California Secretary of State. Although not all DQR's controls could be reproduced, our version is reasonably similar and produces similar, if slightly larger estimates of the PW effect.[26] Our OLS replication of DQR (Model 3, Table 2) finds a prevailing wage effect of 17.8%, twice DQR's 2005 estimate of 9.7%, and three times our preferred estimates.

[25] DQR's original data set has been lost, precluding replication of their work.
[26] Our model includes controls for housing type (*Non-Targeted, Senior, SRO, Special Needs*), building characteristics (*Townhouse, Single Family, Three Stories, Number of Bedrooms*), and DQR's preferred regional controls; measures of affordability, the occupancy date, the fraction of financing provided from public sources, and whether the project was built on an inner-city infill site could not be included.

Table 4 summarizes the essential elements of our replication and exploration of DQR's IV estimates, providing PW estimates and standard errors, estimates of instruments and their standard errors, F-tests of the significance of the instruments and tests of PW endogeneity and significance.[27] Model 1 replicates DQR with their preferred instruments and regional controls; Model 2 is similar to Model 1, except that it shifts to our preferred geographic controls; Model 3 adds year indicators; Model 4 adds our preferred instruments; Model 5, our preferred model, drops DQR's instruments.[28]

In Model 1, the replication of DQR, the second-stage PW estimate is larger in magnitude than DQR ($\beta = 0.463$) and is statistically significant, but the endogeneity of PW is rejected.[29] The instruments do not meet the relevance criteria. Only two of the eight instruments are statistically significant; their joint F-statistic is far from significant. Rejection of the relevance criteria is sufficient to reject the accuracy of the IV estimate; Model 1's results are consistent with our concerns about DQR's original work.[30]

Model 2 replaces DQR's regional variables with ours. The estimate of the PW effect is even larger ($\beta = 0.564$) and statistically significant. Model 2 continues to fail the relevance criteria, as only three first stage instruments are significant; a group F-test on the instruments

[27] Complete estimates are available from the corresponding author.

[28] It was not possible to include DQR's measures of union density or the proportion of employment in highly unionized industries and occupations in the first stage of the model (see DQR footnote 14 and Table 4) as they were highly collinear. We recreated these variables using current data from Hirsch and MacPherson (2003) but too collinear to estimate; inclusion resulted in other variables and observations being dropped. When union density was required in our preferred model, five other variables and 43 observations were dropped.

[29] The eight instruments, used to replicate DQR, variables that only enter the first stage model, were the percentage of yes votes on eight California propositions between 2001 and 2010 (Propositions 40, 45, 46, 77, 99, 1A, 1B, and 1C), the percentage of voters registered as Democrats, the percentage of the population over 40 years old, and the percentage of housing units that are owner-occupied. Additional measures were obtained from the supplemental California Secretary of State data.

[30] The larger magnitude of the PW effect in these intermediate estimates may be due to the omission of the union density and highly unionized industry and occupation variables discussed previously.

does not reject the null. This is likely the source of the large PW coefficient estimates in the second stage.

Model 3 adds year effects. The PW coefficient is smaller ($\beta = 0.379$) but remains highly significant. The measure of endogeneity is likewise highly significant. The model continues to fail the relevance criteria, as only three of the instruments are statistically significant and the instruments fail the F-test of group significance.

Model 4 adds CDBG funding and density bonus, the instruments in our preferred model. Both instruments are highly statistically significant, a group test of the significance for the instruments rejects the null and the PW is significant in a test of endogeneity. The PW coefficient is smaller in magnitude than in prior models, but remains large, positive, and statistically significant ($\beta = 0.261$).[31]

Finally, Model 5 summarizes the results from Table 3 for our two-stage model, the model that drops DQR's instruments and only includes our preferred instruments. This model meets each of the criteria for assessing a two-stage model: both instruments are individually significant and reject a joint test of the null; the PW is endogenous. The PW variable is, however, not statistically significant in the second-stage equation ($\beta = -0.063$). Our estimates, supported by consistent application of appropriate statistical tests, are notably smaller than those reported by DQR. Further, the only intermediate version of the model that meets the test standards is the model including our instruments. Table 4 estimates suggest that the omission of project year fixed effects and weak instruments inflated DQR's estimate of effects of prevailing wages on affordable housing construction costs.

[31] Robustness tests indicate that the significant DQR instruments are only significant in the presence of the full set of DQR instruments and the CDBG Funding and Density Bonus variables. If the non-significant instruments are dropped, neither the "% Yes on Prop 1B" or "% Registered Democrats" remain significant. Therefore, the success of Model 4 of Table 4 is due in main to our instruments.

Discussion

There is a dual crisis of (a) a lack of affordable housing in many parts of the United States and (b) poor conditions of work and life for segments of the construction labor force. The latter might be alleviated by the application of prevailing wage laws to publicly supported affordable housing construction, but would this come at the price of raising the cost and reducing the supply of new affordable housing?

We approach this issue by answering two questions. The first (and narrowest) was to systematically apply contemporary instrumental variable (IV) techniques to determine whether the application of the prevailing wage is endogenous to the costs of affordable housing and determine whether DQR's oft cited IV estimates were reliable. We have determined that whether prevailing wage laws are endogenous is sensitive to the instruments adopted for the model. We have also determined that DQR's IV methods, applied to current data, produce unreliable estimates of the effect of the prevailing wage. Their IV estimates of 37% are several orders of magnitude greater than those obtained from an IV model which explicitly follows current practices and applies appropriate statistical tests to the IV estimates. Our current work does not provide the last word on the endogeneity of the prevailing wage but does emphasize the need for thoughtful work in the specification and testing of models in which PWLs are treated as endogenous.

The second purpose was to obtain reliable, robust estimates of the magnitude of the effect of prevailing wage laws on the costs of constructing affordable housing. Building on our IV estimate, which was negative but not statistically significant, we find that the prevailing wage does not affect the cost of affordable housing. Our IV estimate is, however, estimated with limited precision; the 95% confidence interval for the effect of the prevailing wage on costs ranges from - 22% to + 9.6%. Littlehale's (2017) and our OLS estimates of 5 to 6% are within this interval. Despite the non-significant IV estimate, we cannot rule out that the prevailing wage might raise the costs of affordable housing, albeit more modestly than estimated in some prior research. Given some evidence that the prevailing wage is not endogenous, a range from the IV estimate of 0% and the OLS estimates of 5% to 6% is reasonable.

Our estimated range suggests that the prevailing wage requirement has a moderate cost effect; even an upper limit of 6% is well below that of most prior research. Our OLS estimates find that, relative to the costs of construction under current industry practices, there may be an efficiency cost associated with the prevailing wage requirement in affordable housing construction in California. The implicit baseline for this research has been the current practices in the residential construction industry, including the cost advantages realized by contractors engaging in illegal and undesirable practices. An alternative baseline would be the cost of building affordable housing for contractors who abide by labor standards, classify their workers correctly and pay the required amounts in social insurance and taxes. Computation of this baseline, and of its effect on the net costs of prevailing wage requirements for affordable housing construction, awaits additional research.

Conclusion

This study has considered several methodologies for estimating the effect of prevailing wage requirements on the cost of affordable housing and has obtained reliable estimates of those effects. We conclude that California's prevailing wage requirement adds as much as 6% to construction costs of affordable housing but possibly less. Although this cost increase might, as suggested by Fenn et al. (2018), be partially balanced by the reduction in social costs including providing medical services to low-income construction workers injured during their work or, as suggested by Juravich, Ormiston, and Belman (2021), social gains including increased tax revenues, this research does not speak to that issue.

California is the only state that extends prevailing wage requirements to projects supported by state tax concessions. As such, it is currently the only research setting for examining the effects of prevailing wage requirements on affordable housing construction costs. If additional states follow California's lead, this will provide additional opportunities to study the effect of the prevailing wage on the costs of affordable housing.

APPENDIX

Table 1.1: Definitions and Descriptive Statistics

Variables	Definitions	Means	Standard Deviations
Log project cost	Natural log of total project cost (excluding land)	16.52	0.60
Prevailing wage	=1 if project required payment of prevailing wages, =0 if not	0.63	0.48
CDBG funding	=1 if a project received funding from the Community Development Block Grant program, =0 if not	0.06	0.25
Density bonus	=1 if a project was incentivized by California's density bonus law, =0 if not	0.29	0.45
Nonprofit developer	=1 if a project's developer was nonprofit, =0 otherwise	0.44	0.50
Government developer	=1 if a project's developer was a government developer, =0 otherwise	0.04	0.19
Other developer	=1 if a project's developer was a private, for-profit developer, =0 otherwise	0.02	0.15
Developer general contractor	=1 if the developer was an in-house general contractor, =0 if not	0.46	0.50
Log developer employees	Natural log of each developer's number of employees	3.88	1.33
Meetings 4 plus	=1 if more than 4 community meetings occurred during a project, =0 if not	0.36	0.48
Permit and impact fees	Sum of permit fees and impact fees	0.94	0.23
Log permit and impact Fees/SF	Natural log of the sum of permit and impact fees, per square foot	2.17	0.93
Funding-redevelopment	=1 if a project was a redevelopment project, =0 otherwise	0.33	0.47
9% tax credit	=1 if a project's developer received a 9% tax credit, =0 if not	0.52	0.50
Log lenders	Natural log of the total number of loans that finances a project	1.26	0.54
Log building area	Natural log of gross building area in square feet, excluding parking	11.15	0.62

Table 1.1 (cont'd)

Structure includes parking	=1 if an affordable housing structure included parking, =0 if not	0.38	0.49
Parking-to-rentable building	Ratio of area of structured parking to gross area of building, minus parking	0.13	0.21
SF non-residential: non-parking	Percentage of total square feet that is non-residential, minus parking	0.11	0.11
Stories 4 plus	=1 if more than 4 stories, =0 otherwise	0.24	0.43
Residential SF per unit (100s)	Gross area of parking, community, common, and commercial divided by total units (100s of SF)	9.26	2.54
Log site in acres	Natural log of the size of an affordable housing project site, in acres	0.97	0.94
Includes non-residential area	=1 if an affordable housing site included a non-residential area, =0 if not	0.96	0.20
Log A&E cost per SF	Natural log of the sum of architect and engineer fees divided by building area, excluding parking	1.98	0.80
Duration 24 plus	=1 if a project took at least 24 months (2 years) to complete, =0 otherwise	0.37	0.48
Average specialty contractor pays	Average pay for specialty trade contractors, by county, as a percentage of state average in project year	1.00	0.17
Log fair market rent (2-bedroom)	Natural log of the HUD-established going rental rate for a 2-bedroom affordable housing unit	6.96	0.31
Project snow	=1 if a project located in a city with an average snow depth of one foot or more, =0 otherwise	0.01	0.12

Table 1.2: Ordinary Least Squares Estimates (Dependent Variable: Log Project Cost)

Variables	Model 1	Model 2	Model 3
Prevailing wage	**0.050****	**0.058****	**0.178*****
	(0.024)	(0.024)	(0.048)
Log units			**0.775*****
			(0.061)
Nonprofit developer	**0.032**	**0.046****	**0.019**
	(0.022)	(0.022)	(0.037)
Government developer	**0.130***	**0.155*****	**0.066**
	(0.052)	(0.048)	(0.088)
Other developer	**-0.003**	**-0.008**	
	(0.041)	(0.041)	
Log developer employees	**-0.022*****	**-0.025*****	
	(0.006)	(0.006)	
Developer general contractor	**(-0.038) ***		
	(0.021)		
Meetings 4 plus	**0.064*****	**0.071*****	
	(0.019)	(0.019)	
Permit & impact fees	**-0.154*****	**-0.174*****	
	(0.050)	(0.049)	
Log permit & impact fees per SF	**0.042*****	**0.048*****	
	(0.014)	(0.014)	
Funding-redevelopment	**0.041****		
	(0.019)		
9% tax credit	**-0.030**		
	(0.021)		
Log building area	**0.877*****	**0.919*****	
	(0.020)	(0.017)	
Structure includes parking	**-0.014**		
	(0.027)		
Parking-to-rentable building	**0.371*****	**0.372*****	
	(0.062)	(0.050)	
SF non-residential: non-parking	**-0.533*****	**-0.597*****	
	(0.135)	(0.128)	
Stories 4 plus	**0.054**	**0.023**	
	(0.034)	(0.031)	
Residential SF per unit	**-0.029*****	**-0.032*****	
	(0.004)	(0.004)	
Log site in acres	**0.026**		
	(0.021)		
Includes non-residential area	**0.024**		
	(0.039)		

Table 1.2 (cont'd)

Average specialty contractor pay	**0.245***	**0.421*****	
	(0.130)	(0.111)	
Log fair market rent (2-bedroom)	**0.212*****		
	(0.070)		
Project snow	**0.190*****	**0.197*****	
	(0.055)	(0.060)	
Log lenders	**0.026**	**0.041****	**0.027*****
	(0.020)	(0.019)	(0.006)
Duration 24 plus	**0.033***	**0.028**	
	(0.018)	(0.018)	
Log A&E cost per SF	**0.148*****	**0.162*****	
	(0.017)	(0.018)	
Housing type: non-targeted			**-0.047**
			(0.108)
Housing type: senior			**-0.280*****
			(0.056)
Housing type: SRO			**-0.594*****
			(0.085)
Housing type: special needs			**-0.234****
			(0.088)
Number of bedrooms			**0.000**
			(0.000)
Mitigation			**0.138*****
			(0.048)
Structure: townhouse			**0.085**
			(0.064)
Structure: single family			**0.157**
			(0.114)
Structure: three stories			**0.101*****
			(0.040)
Constant	**4.72*****	**5.47*****	**12.90*****
	(0.462)	(0.218)	(0.196)
N	321	321	279
R^2	0.958	0.954	0.836
F	246	254	54.50

Estimated coefficients in bold. Robust standard errors in parentheses.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 1.3: Two Stage Least Squares Model

Variables	First-Stage Coefficients	Instrumented Coefficients (Project Cost)
Prevailing wage		***-0.063***
		(0.084)
CDBG funding	***0.378******	
	(0.073)	
Density bonus	***0.140*****	
	(0.055)	
Nonprofit developer	**0.053**	**0.065*****
	(0.064)	(0.025)
Government developer	**0.433*****	**0.210*****
	(0.089)	(0.058)
Other developer	**0.349****	**0.056**
	(0.138)	(0.057)
Log developer employees	**-0.043****	**-0.029*****
	(0.019)	(0.008)
Meetings 4 plus	**0.059**	**0.078*****
	(0.057)	(0.020)
Permit & impact fees	**0.158**	**-0.136****
	(0.143)	(0.057)
Log permit & impact fees	**-0.053**	**0.041****
per SF	(0.037)	(0.017)
Log building area	**-0.083***	**0.913*****
	(0.048)	(0.021)
Parking-to-rentable	**0.148**	**0.419*****
building	(0.150)	(0.056)
SF non-residential: non-	**-0.470**	**-0.673*****
parking	(0.290)	(0.138)
Stories 4 plus	**0.204*****	**0.039**
	(0.069)	(0.037)
Residential SF per unit	**0.022***	**-0.029*****
	(0.012)	(0.005)
Average specialty	**-0.018**	**0.321****
contractor pay	(0.442)	(0.130)
Project snow	**0.376****	**0.244*****
	(0.175)	(0.066)
Log lenders	**0.166*****	**0.059*****
	(0.055)	(0.023)
Duration 24 plus	**-0.030**	**0.017**
	(0.055)	(0.020)
Log A&E cost per SF	**0.075**	**0.194*****
	(0.054)	(0.018)
Constant	**0.873**	**5.57*****
	(0.686)	(0.263)

Table 1.3 (cont'd)

Project year fixed effects	**Yes**	**Yes**
Project region fixed effects	**Yes**	**Yes**
N	281	281
R^2	0.506	0.944
F	15.42	5869

Estimated coefficients are in bold. Robust standard errors are in parentheses.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 1.4: Supplemental Analysis

Variables	Model 1	Model 2	Model 3	Model 4	Model 5
Prevailing wage	**0.463****	**0.564*****	**0.379*****	**0.261*****	**-0.063**
	(0.191)	(0.137)	(0.114)	(0.071)	(0.084)
First-stage joint F statistic	**1.220**	**1.750**	**1.510**	**4.440*****	**14.800*****
% Yes, vote on prop. 40	**3.560**	**1.570**	**2.210**	**1.530**	
water, air, and parks, 2002	(2.190)	(1.770)	(1.850)	(1.780)	
% Yes, vote on prop. 45	**-6.410****	**-1.130**	**-1.270**	**-0.809**	
term limits, 2002	(3.230)	(2.820)	(2.740)	(2.730)	
% Yes, vote on prop. 46	**5.370**	**-4.400**	**-3.660**	**-1.750**	
housing/emergency shelter, 2002	(4.450)	(3.130)	(3.110)	(3.030)	
% Yes, vote on prop. 77	**-1.510**	**0.401**	**-0.175**	**-0.299**	
redistricting, 2005	(2.070)	(1.430)	(1.420)	(1.420)	
% yes, vote on prop. 1B	**-1.940**	**-3.910****	**-3.930****	**-3.830****	
highways and air quality, 2006	(2.350)	(1.750)	(1.740)	(1.620)	
% Yes, vote on prop. 1C	**-5.310**	**7.420***	**6.440**	**4.760**	
housing/emer. shelter, 2006	(3.990)	(3.940)	(3.930)	(3.660)	
%Yes, vote on prop. 99	**0.105**	**-1.610**	**-1.700**	**-1.110**	
eminent domain, 2008	(2.150)	(1.530)	(1.480)	(1.470)	
% Yes, vote on prop. 1A	**0.823**	**3.540***	**3.670****	**3.260***	
state taxes, 2009	(2.220)	(1.850)	(1.770)	(1.750)	
percent of voters registered	**-0.564**	**-4.420****	**-4.760****	**-4.480****	
as democrats	(2.150)	(2.220)	(2.230)	(2.240)	
Percent of population over 40	**-4.160****	**0.067**	**-0.202**	**-0.719**	
years old	(1.960)	(2.240)	(2.170)	(2.030)	
Percent of housing units	**1.110**	**-2.190**	**-1.560**	**-1.110**	
owner-occupied	(1.440)	(1.740)	(1.720)	(1.630)	
CDBG funding				**0.326*****	**0.378*****
				(0.079)	(0.073)
Density bonus				**0.235*****	**0.140****
				(0.059)	(0.055)
Endogeneity statistic	**2.390**	**8.839*****	**6.110****	**3.844***	**2.847***
N	279	279	279	279	281
First-stage R^2	0.508	0.461	0.482	0.523	0.506
Second-stage R^2	0.808	0.781	0.901	0.918	0.944

Estimated coefficients are in bold. Robust standard errors are in parentheses.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Because of missing data, inclusion of the CBDG Funding and Density Bonus variables in Models 4 and 5 reduced the number of observations from 315 to 279 and 281, respectively. As a robustness check, Models 1 – 3 were run with 315 observations. The reduction in the number of observations had little effect on the signs, magnitudes, or significance of the PW results. These results are available from the corresponding author upon request.

REFERENCES

REFERENCES

American Legislative Exchange Council. (2020). Accessed at https://www.alec.org/model-policy/prevailing-wage-repeal-act.

Aurand, A., Emmanuel, D., Yentel, D., Errico, E. and Pang, M. (2018). 'The gap: A shortage of affordable homes.' Washington, D.C.: National Low-Income Housing Coalition.

Azari-Rad, H., Philips, P. and Prus, M. (2002). 'Making hay when it rains: The effect of prevailing wage regulations, scale economies, seasonal, cyclical and local business patterns on school construction costs.' *Journal of Education Finance,* 27: 997-1012.

Azari-Rad, H., Philips, P. and Prus, M. (2003). 'State prevailing wage laws and school construction costs.' *Industrial Relations,* 42: 445-57.

Belman, D., Ormiston, R., Petty, R. and Hinkel, M. (2020). 'The effect of state prevailing wage laws on African American employment in construction.' Unpublished working paper.

Belman, D. and Philips, P. (2005). 'Prevailing wage laws and minority employment in construction." In H. Azari-Rad, P. Philips, and M. Prus (eds.), *The Economics of Prevailing Wage Laws*, pp. 101-122. Aldershot, UK; Burlington, VT: Ashgate.

Bilginsoy, C. (2005). 'Wage regulation and training: The impact of state prevailing wage laws on apprenticeship.' In H. Azari-Rad, P. Philips and M. Prus (eds.), *The Economics of Prevailing Wage Laws*, pp. 149-168. Aldershot, UK; Burlington, VT: Ashgate.

Blankenau, W. and Cassou, S. (2011). 'Industry differences in the elasticity of substitution and rate of biased technological change between skilled and unskilled labor. *Applied Economics,* 43: 3129-42.

Bureau of Labor Statistics, U.S. Department of Labor. (2011-2016). *Quarterly Census of Employment and Wages.* Washington, D.C.

Bureau of Labor Statistics, U.S. Department of Labor. (2010). 'Working poor 2010.' https://www.bls.gov/opub/reports/working-poor/archive/workingpoor_2010.pdf table 4

Bureau of Labor Statistics, U.S. Department of Labor. (2018). 'Census of fatal occupational injuries.' https://www.bls.gov/iif/oshcfoi1.htm#2018

California Department of Housing and Community Development. (2018). 'Community development block grant program. http://www.hcd.ca.gov/grants-funding/active-funding/cdbg.shtml#guidelines.

Duncan, K. (2015). 'The effect of federal Davis-Bacon and disadvantaged business enterprise regulations on highway maintenance costs.' *ILR Review,* 68: 212-37.

Duncan, K. and Ormiston, R. (2018). ‘What does the research tell us about prevailing wage laws?’ *Labor Studies Journal*, 44(2): 139-160.

Duncan, K., Philips, P. and Prus, M. (2014). ‘Prevailing wage regulations and school construction costs: Cumulative evidence from British Columbia.’ *Industrial Relations: A Journal of Economy and Society*, 53: 593-616.

Dunn, S., Quigley, J. and Rosenthal, L. (2005). ‘The effects of prevailing wage regulations on the cost of low-income housing.’ *ILR Review*, 59(1): 141-157.

Erlich, M. (2020). ‘Misclassification in construction: The original gig economy.’ *ILR Review*, forthcoming.

Fenn, A., Li, Z., Pleites, G., Zorigtbaatar, C. and Philips, P. (2018). ‘The effect of prevailing wage repeals on construction income and benefits.’ *Public Works Management & Policy*, 23(4): 346-364.

Garcia, D. (2019). 'Making it pencil: The math behind housing development.' Berkeley, CA: Terner Center for Housing Innovation.

Goetz, J. and Sakai, T. (2017). ‘Guide to the California density bonus law.’ California: Meyers Nave.

Hirsch, B.T. and Macpherson, D.A. (2003). ‘Union membership and coverage database from the Current Population Survey: Note.’ *ILR Review*, 56(2): 349-354 (updated annually at unionstats.com).

Juravich, T., Ablavsky, E., and Williams, J. (2015). ‘The epidemic of wage theft in residential construction in Massachusetts.’ UMass-Amherst Labor Center Working Paper Series.

Juravich, T., Ormiston, R., and Belman, D. (2021), ‘The social and economic costs of misclassification, wage theft and tax fraud in the Massachusetts residential construction industry.’ Amherst, MA: UMass Amherst Labor Center.

Littlehale, S. (2017). ‘Revisiting the costs of developing new subsidized housing: The relative impact of construction wage standards and nonprofit development.’ *Berkeley Planning Journal*, 29(1): 103-130.

McMahon, E.J., and Gardner, K. (2017). ‘Prevailing waste: New York’s costly public works pay mandate.’ Albany, NY: Empire Center for Public Policy.

National Housing Law Project. (2017). ‘Low-income housing tax credits.’ https://www.nhlp.org/resource-center/low-income-housing-tax-credits/.

Ormiston, R., Belman, D. and Hinkel, M. (2018). 'New York's prevailing wage law: A cost-benefit analysis.' Washington D.C.: Economic Policy Institute.

Ormiston, R., Belman, D., Brockman, J. and Hinkel, M. (2020). 'Rebuilding residential construction.' In P. Osterman (ed.), *Creating Good Jobs: An Industry-Based Strategy*, pp. 75-113. MIT Press.

Ormiston, R., Belman, D., and Erlich, M. (2020). 'Payroll fraud in New York's construction industry: Estimating its prevalence, severity and economic costs.' Institute for Construction Research and Training.

Palm, M., and Niemeier, D. (2018). 'Does placing affordable housing near rail raise development costs? Evidence from California's four largest metropolitan planning organizations." *Housing Policy Debate*, 28: 180-98.

Philips, P., Mangum, G., Waitzman, N. and Yeagle, A. (1995). 'Losing ground: Lessons from the repeal of nine "Little Davis-Bacon" acts.' Economics Department, University of Utah.

Sinyai, C., and Galeas, E. (2021). 'The underground economy and wage theft in Washington D.C.'s commercial construction sector.' The Catholic Labor Network: Washington D.C.

State of California. (2014). 'Affordable housing cost study.' http://www.treasurer.ca.gov/ctcac/study.asp.

Stock, J., and Yogo, M. (2005). 'Asymptotic distributions of instrumental variables statistics with many instruments." *Identification and inference for econometric models: Essays in honor of Thomas Rothenberg,* pp. 109-120.

Stock, J., Wright, J. and Yogo, M. (2002). 'A survey of weak instruments and weak identification in generalized method of moments.' *Journal of Business & Economic Statistics*, 20(4): 518-529.

United States Department of Labor. n.d. 'Davis-Bacon and related acts.' https://www.dol.gov/whd/govcontracts/dbra.htm.

United States Department of Labor. (2015). 'Introduction to labor standards statutes coverage.' https://www.dol.gov/whd/recovery/pwrb/Tab3.pdf.

United States Department of Labor Wage and Hour Division. 'Instructions for completing payroll form, WH-347.' https://www.dol.gov/agencies/whd/forms/wh347.

Vitullo-Martin, J. (2012). 'The complex worlds of New York prevailing wage.' New York: Center for Urban Real Estate.

Wooldridge, J. (2013). *Introductory Econometrics: A Modern Approach*. Mason, OH: South-Western.

CHAPTER 2: LABOR MARKET OUTCOMES OF OCCUPATIONAL LICENSING IN CONSTRUCTION

Abstract

Over time, disagreements have persisted over the effects of occupational licensing on markets and the appropriate role of government in the regulation of occupations. In this paper, I exploit state variation in occupational licensing laws to examine labor market outcomes of occupational licensing in construction. Data on licensing comes from 2016-2019 Current Population Survey (CPS) data as well as a new 2019 data set on licensing requirements for the three primary construction occupations that require licensing in certain states: electricians, plumbers, and operating engineers. Consistent with prior literature, results suggest the presence of occupational licensing is associated with an 8.3 to 14.8 percentage point increase in earnings for electricians, plumbers, and operating engineers. Employment results are more mixed; while these results suggest occupational licensing is associated with a 1.2 to 1.3 percentage point increase in the proportion of workers employed as electricians, plumbers, or operating engineers, effects on the level of employment in these occupations were not statistically significant. In supplemental analyses I explore possible competing explanations for these employment findings.

Introduction

An occupational license is a state-issued credential that must be obtained in order to legally work for pay in certain occupations (Blair & Chung 2019). Occupational licensing, also known as the right to practice, is the most restrictive of three forms of occupational regulation in the U.S.; the two less restrictive forms are registration and certification (Kleiner 2000).[32] The study of occupational regulation has a long and distinguished history in economics and has undergone a resurgence in the last 20 years. This renewed focus is merited by occupational licensing's current status as one of the fastest-growing institutions in the United States labor market. While unionization has long been in a precipitous decline, occupational licensing has grown substantially in recent decades. About 25 percent of the U.S. workforce is required to obtain a license to work in their given occupation; in 1950, that figure stood at 5 percent (Kleiner & Krueger 2010, 2013). The simultaneous decline of unionization and rise of occupational regulation has led scholars to suggest that these two institutions may be acting as substitutes for worker protection in the U.S. (Kleiner & Krueger 2013).

In this paper, I exploit state variation in licensing requirements to examine how licensing requirements relate to employment and earnings in one of the largest U.S. sectors: the construction industry. While construction union density was only 13% in 2021, 15.6% of workers in construction and extraction occupations were licensed in 2021 (Bureau of Labor Statistics 2022). In construction, licensing requirements are primarily concentrated in three high-skilled occupations: electricians, plumbers, and operating engineers.[33] Geographic heterogeneity exists in construction licensing requirements because these three occupations require licensing in some states but not others. Although geographic heterogeneity exists, licensing requirements are largely clustered in a specific high-skilled portion of the industry's labor force, leaving less-skilled segments relatively free of licensing requirements.

[32] More specifically, Kleiner (2000) defines occupational licensing as "a process where entry into an occupation requires the permission of the government, and the state requires some demonstration of a minimum degree of competency" (Kleiner 2000, p. 191).
[33] Operating engineers are responsible for operating and maintaining heavy equipment and machinery on construction sites.

Though occupational regulation has historically been among the most examined institutions in labor economics, disagreements have persisted over its effect on markets and the appropriate role of government in the regulation of occupations (Kleiner 2000). Neoclassical economists have primarily viewed licensing as rent-seeking behavior that interferes with free markets and creates inefficiencies and rents only shared by those who obtain licenses (Friedman 1962; Friedman & Kuznets 1945). Other theory has focused on human capital investments. This suggests that licensing increases incentives for workers to invest in their human capital and increase their skill levels, given the higher earnings and employment opportunities available to workers who do so (Shapiro 1986). By signaling worker quality and training to the market, occupational licensing may reduce informational asymmetries between workers and firms, thereby boosting employment and earnings of licensed workers.

The most common view in the literature is that licensing acts primarily as a rent-seeking device and a barrier to entry that restricts labor supply to a licensed occupation, thereby driving up the price of labor. Workers with lower skill levels may self-select into non-licensed occupations, leaving licensed occupations with higher average skill levels and employers competing for smaller supplies of workers (Kleiner 2000). This lowers employment but raises wages (Rottenberg 1980; Kleiner 2000). Empirical licensing literature has largely supported this framework, finding positive wage and earnings effects (e.g., Kleiner & Krueger 2013; Thornton & Timmons 2013; Koumenta et al. 2014; Gittleman et al. 2018). and negative employment and labor supply effects (e.g., Kleiner 2006; Kleiner & Park 2010; Hall et al. 2018; Blair & Chung 2019; Yelowitz & Ingram 2021). No prior licensing studies have focused on construction. Unlike industries focused on in prior licensing studies, construction often does not require college degrees for entry; a high school diploma is standard, even for the high-skilled occupations that feature licensing.[34] Examining labor market outcomes of occupational licensing is all the more important in a context where workers typically do not have as many other ways to readily signal work quality.

Other literature has suggested that licensure creates a greater incentive for workers to invest in more occupation-specific human capital because they will be abler to recoup the full returns to their investment (Shapiro 1986). This had led to a different view: by serving as a credible signal of worker quality and training, licensing may increase demand for services by resolving informational asymmetries between workers and firms in markets where work quality may be difficult to ascertain.[35] This can be especially salient in contexts where low-quality work poses great health and safety risks. To the extent that licensing signals quality and resolves asymmetric information, this increases labor demand, thereby increasing both employment and wages (Leland 1979; Carollo 2020). While both of the above theoretical frameworks predict positive earnings effects of licensing, employment could increase or decrease depending on the relative importance of these competing supply and demand effects (Carollo 2020). In construction, while licensing can lower employment via barriers to entry, it can also boost employment by signaling higher levels of worker training to firms, thereby providing greater consumer protections from negative effects of poor workmanship.

This paper examines how licensing requirements in construction relate to employment and earnings. As the main source of licensing variation is at the state level, the unit of analysis for this study is at the state level. Data on licensing comes from 2016-2019 Current Population

[34] Some lower-skilled construction occupations have no formal education requirement and do not require a high school diploma. For example, no formal education is required to become a construction laborer or helper. An exception is helpers of high-skilled occupations (e.g., helpers of plumbers, helpers of electricians), which typically require a high school diploma (Bureau of Labor Statistics).
[35] Demand for services can also remain the same, depending on the elasticity of demand for construction work.

Survey (CPS) data as well as a new 2019 data set on licensing requirements for the three primary construction occupations that require licensing in certain states: electricians, plumbers, and operating engineers. This new data set was constructed based on correspondence with various construction labor groups: the United Association of Journeymen and Apprentices of the Plumbing and Pipefitting Industry (UA), the National Electrical Contractors Association (NECA), the International Union of Operating Engineers (IUOE), and the Pennsylvania Foundation for Fair Contracting (PAFFC). Consistent with prior literature, results of this study suggest occupational licensing is associated with higher earnings for electricians, plumbers, and operating engineers. Results also suggest the presence of occupational licensing is associated with higher proportions of workers employed as electricians, plumbers, or operating engineers; however, effects of licensing on the level of employment in these occupations were not statistically significant. Taken together, these results suggest that licensing may increase labor demand by resolving informational asymmetries. In supplemental analyses I explore competing explanations for these employment results.

Background and Literature Review

Law and Kim (2005) offer a historical perspective on the appearance and increased prevalence of occupational regulation in the United States during the Progressive Era. Law and Kim (2005) argue that when occupational regulation was introduced, it was designed to improve markets by improving consumers' abilities to judge service quality, rather than restrict entry or reduce competition. Sellers of services are often better informed than buyers about product quality. These informational asymmetries lead to the classic lemons problem, where lower quality goods drive out higher quality goods from the market (Law & Kim 2005).

This same logic can extend to labor markets, where asymmetric information can exist between workers and firms. Workers tend to be better informed than firms about the quality of their work; however, when licensing is present, this information gap is reduced because firms have a credible signal of work quality. Aside from higher earnings, this reduction in asymmetric information can improve markets by boosting the accuracy of work quality judgments on the part of firms (Law and Kim 2005). This is relevant in the construction context because when consumers need plumbing work, they coordinate with plumbing firms; similarly, consumers in need of electrical work coordinate with electrical firms. In construction, firms use licensing to gauge whether workers have achieved the requisite level of training to properly complete the work. Thus, in this industry, occupational licensing may be sought out as a means of improving work quality, reducing workplace accidents, and curbing the effects of asymmetric information.

While licensing may act as a signal of quality and help correct market distortions resulting from informational asymmetries, it can also present a barrier to entry and restrict labor supply (Rottenberg 1980; Kleiner 2000). In the presence of occupational licensing, regulators screen entrants to a given occupation, and entry is limited for those whose skills suggest lower-quality work.[36] Workers who obtain a license and enter the occupation become incumbents; regulators continue to monitor those incumbents and penalize those who perform below established licensure standards. These penalties can range from fines to revocation of the license (Kleiner 2000; Kleiner & Krueger 2010, 2013; Gittleman & Kleiner 2016). This screening process and labor supply reduction then results in a higher price of labor (Rottenberg 1980; Kleiner 2000).

[36] While entry is limited initially, workers can go on to obtain more skills and enter later.

Empirical findings in the literature have largely supported this barrier-to-entry framework and the idea that occupational licensing is primarily rent-seeking. Kleiner and Krueger (2010) use a specially-designed Gallup survey of U.S. workers on occupational regulation and find that 29 percent of the workforce is required to be licensed.[37] The wage effect, as expected, is positive: occupational licensing is associated with an increase in wages of 15 percentage points (Kleiner & Krueger 2010). In a follow-up study, Kleiner and Krueger (2013) use novel national survey data collected by Westat that asked detailed questions on occupational regulation as well as questions on the labor market status of workers. They find that licensing is associated with an 18-percentage-point increase in wages (Kleiner & Krueger 2013).[38] In an industry-specific study, Thornton and Timmons (2013) estimate a licensing earnings premium of 16.2 percent for massage therapists.

Gittleman and Kleiner (2016) compare the economic returns to licensing with the returns to unionization using data collected on licensing statutes linked to the National Longitudinal Survey of Youth from 1979 to 2010.[39] Gittleman and Kleiner (2016) find that occupational licensing is associated with higher wages, but that those returns are lower than those of unionization. Conversely, in Canada, Zhang (2019) finds a licensing premium higher than that of unionization, with a licensing premium of 15.5 percent. Kleiner and Vorotnikov (2017) also find a positive wage premium, suggesting that licensing raises wages by about 11 percent after controlling for human capital and other observable characteristics. Gittleman, Klee, and Kleiner

[37] Kleiner and Krueger (2010) also find that licensing is more prevalent among more educated workers, minorities, union members, and government workers.
[38] Kleiner and Krueger (2013) find that an identical proportion of workers are required to be licensed in this survey as in their 2010 survey: 29 percent of employees were licensed, and 35 percent of employees were either licensed or certified. Additionally, another 3 percent stated that all who worked in their given job would eventually be required to be certified or licensed (Kleiner & Krueger 2013).
[39] This comparison is warranted because even with its decrease in prevalence over time, unionization still covers a nontrivial portion of the U.S. labor force.

(2018) use the 2008 panel of the Survey of Income and Program Participation (SIPP) and find that after controlling for observable heterogeneity (e.g., occupational status), the licensing wage premium is about 7.5 percent (Gittleman, Klee, & Kleiner 2018). In sum, occupational regulation appears to raise earnings across numerous industries, leading to this study's first hypothesis:

Hypothesis 1: *Electricians, plumbers, and operating engineers will have higher earnings in states in which these occupations are licensed than in states without occupational licensing requirements.*

Literature on the effects of licensing on employment and discrimination is more nascent but has grown in recent years.[40] Prior work on employment and labor supply broadly suggests negative effects. Examining librarians, respiratory therapists, and dieticians and nutritionists, Kleiner (2006) shows that controlling for state characteristics, states that did not feature licensing requirements for these occupations experienced 20 percent faster employment growth than states where licensing requirements were imposed. Kleiner and Soltas (2019) find that shifting an occupation in a state from entirely unlicensed to entirely licensed increases average wages by 15% but reduces employment by 29% (Kleiner & Soltas 2019). Yelowitz and Ingram (2021) focus on emergency medical technicians (EMTs) by examining how licensing affects the choice to become an EMT. In so doing, they exploit the demand-side shock from the Affordable Care Act (ACA), which increased demand for EMTs. While the ACA demand shock increased entry by 18 percentage points, occupational licensing requirements reduced entry by a similar magnitude (Yelowitz & Ingram 2021).

[40] Kleiner and Soltas (2018) show that the effect of licensing on employment is a sufficient statistic for the welfare consequences of licensing, adding importance to studying the effect of occupational licensing on employment.

Blair and Chung (2019) employ a boundary discontinuity design and find that licensing reduces labor supply by an average of 17 to 27 percent across multiple industries. In exploiting state variation in licensing, the authors attempt to strengthen the identification assumption by restricting their sample to "border counties", counties that border one another but are in different states.[41] The identification assumption is that counties located next to each other but in different states should have similar characteristics except for licensing requirements that differ across state boundaries (Blair & Chung 2019). In construction, licensing requirements often hold across state boundaries due to reciprocity agreements. These agreements allow workers who previously passed an examination and obtained a license in a state to remain licensed in another state *without* an additional examination (National Center for Construction Education and Research). That is, under reciprocity, a state will accept another state's license in place of their own and allow the worker to remain licensed without having to take an additional test, provided that the worker is in good standing. For example, for electrical licenses, the bordering states of New Hampshire and Massachusetts have reciprocity agreements with one another; so long as a worker is in good standing, New Hampshire electrical licenses are accepted in Massachusetts and vice versa. Similarly, for plumbing licenses, the bordering states of California, Arizona and Nevada have reciprocal agreements with one another (National Center for Construction Education and Research).[42] Thus, the border discontinuity approach and its results may not hold in the construction context.

[41] Blair and Chung (2019) also include boundary fixed effects.
[42] It should be noted that in some cases, the term "reciprocity" can be somewhat of a misnomer; there are cases where reciprocity agreements do not go both ways. For example, West Virginia accepts electrical licenses from all 50 states in place of their own, yet only one state (Virginia) accepts West Virginia electrical licenses (National Center for Construction Education and Research).

The same can be said for recent work that suggests licensing creates barriers to geographic mobility. Using American Community Survey (ACS) data, Johnson and Kleiner (2020) find that occupational licensing is associated with significant reductions in interstate migration across multiple occupations, including pharmacists, dentists, and lawyers. While licensing may restrict the flow of incumbent workers in other industries, reciprocity agreements can increase the flow of incumbent construction workers across state boundaries. The highly mobile nature of the construction labor force is supported and potentially enhanced by the ability for licensed workers to move between states with reciprocal agreements and maintain their prior license.[43]

Prior work has also studied the effects of licensing on discrimination. Licensing may reduce representation of underrepresented groups because members of these groups may find it costlier to obtain licenses or because licensing represents a deliberate effort to exclude these groups. On the other hand, to the extent that licensing acts as a solution for informational asymmetries and a signal of quality, it may increase representation of underrepresented groups in licensed occupations. This is because licensing can reduce uncertainty about worker quality that might otherwise lead to statistical discrimination over observable characteristics like race or gender (Law & Marks 2009).

Studies focusing on funeral directors (Cathles et al. 2010) and Uber (Hall et al. 2018) find that licensing requirements reduce female employment relative to men. Additionally, Gomez et al. (2015) find that while occupational licensing raises immigrant wages more than non-immigrant wages, the authors also find that the probability of employment in a licensed occupation is significantly lower for immigrants (Gomez et al. 2015). Law and Marks (2009) use

[43] Additionally, it is common for union locals to issue permission to members to travel and work in other states, further enhancing the mobility of the construction labor force.

longitudinal data across multiple industries from 1870 to 1960 and find that licensing does not reduce labor supply for female or black workers. Across all industries, Law and Marks (2009) do not find evidence that licensing impacts labor supply for female or black workers. Of relevance to construction, the authors include plumbers, a licensed construction occupation, in their analysis. While licensing did not affect representation of black plumbers, representation of female plumbers increased more in states with licensing requirements than states without licensing requirements (Law & Marks 2009). Women may be particularly at risk of discrimination in the construction context due to the heavily male-dominated nature of the industry. These results suggest that the signal of quality licensing provides can boost female representation in an otherwise male-dominated context and alleviate discrimination.

Recent work on occupational licensing for dental assistants (Xia 2021) finds further evidence that licensing alleviates statistical discrimination, so long as the license reveals information about the worker’s productivity. Blair and Chung (2018) find wage effects consistent with this story: when licensing is present, result suggest that firms rely less on observable characteristics, such as race and gender, in wage determination. Licensed minorities and women experience smaller racial and gender wage gaps than their unlicensed counterparts (Blair & Chung 2018). Similarly, in addition to finding wage effects of 5 to 10 percentage points, Law and Marks (2017) find that licensing equalizes wages within occupations, with minority wages rising faster than nonminority wages.

In sum, the primary view of occupational licensing is that it is primarily rent-seeking and restricts competition via barriers to entry. In turn, this shifts the labor supply curve leftward, thereby lowering employment but increasing wages. These predictions have been supported by empirical research that has found negative labor supply and employment effects (Kleiner 2006;

Blair & Chung 2019; Kleiner & Soltas 2019; Yelowitz & Ingram 2021) as well as lower female employment relative to men and lower immigrant employment resulting from licensing (Cathles et al. 2010; Gomez et al. 2015; Hall et al. 2018). Other literature has suggested licensing may increase demand for services in markets where work quality may be difficult to ascertain by resolving informational asymmetries, resulting in negligible or positive employment effects. This idea may be particularly applicable in contexts where low-quality work poses great health and safety risks (Leland 1979; Carollo 2020). This theory has been partially supported by evidence that licensing alleviates wage and employment discrimination by reducing information gaps and lessening firms' tendencies to base wage and employment decisions on observable characteristics (Law & Marks 2009, 2017; Blair & Chung 2018; Xia 2021). Although this theory has received partial support, the idea that occupational licensing is primarily rent-seeking and restricts competition has received more support in the literature. This leads to the following hypothesis:

Hypothesis 2: *Employment of electricians, plumbers, and operating engineers will be lower in states with licensing requirements than in states without licensing requirements.*

Methods

Data and Sample

As the main source of licensing variation is at the state level, the unit of analysis for this study is at the state level.[44] As there is limited year-to-year variation in licensing requirements

[44] A limitation of this study is that while the vast majority of licensing variation is at the state level, there are cases where licensing is determined locally, rather than the state level. There are cases where a state does not have licensing requirements for an occupation, but a major locality within that state does. An example is St. Louis, MO: while Missouri does not feature licensing requirements for the three licensed occupations in this study, St. Louis

across states, the data used in this study can best be described as a pooled cross section. Data for all dependent and control variables used in this study come from 2016-2019 American Community Survey (ACS) data from the Integrated Public Use Microdata Series (IPUMS) provided by the University of Minnesota (Ruggles et al. 2022).[45] Since the unit of analysis is at the state level, all individual-level ACS data was aggregated to the state level. The initial state-year sample size was 200; however, there were some states that featured low numbers of construction workers that resulted in wide and unreliable year-to-year variation in ACS construction employment estimates. This wide variation was assumed to have resulted from sampling error. The ACS, in some cases, may only survey small samples of construction workers. As this study only focuses on a single industry (and specific occupations within that industry), it is imperative to restrict the sample to states that feature enough annual within-occupation observations. Five states were excluded from the analysis due to low numbers of construction workers in the sample: these states were Alaska, Delaware, North Dakota, Rhode Island, and Wyoming. as a result, 45 states were used in this analysis over 4 years, bringing the final state-year sample to 180.[46] Prior to aggregating data to the state level, I follow prior

---

does. However, the single-industry focus of this study precluded the usage of metropolitan statistical area (MSA)-level variation in licensing; many MSAs did not have sufficient numbers of construction workers to reliably estimate effects on earnings or employment. For the purposes of this study, if an occupation is not classified as licensed in a given state and year, it is classified as unlicensed for the *entire* state, including localities where there may be separate licensing requirements at the local level.

[45] The ACS provides a far superior sample size to that of the Current Population Survey (CPS). The ACS is the largest annual household survey in the United States, surveying over 3.5 million households annually, while the CPS only surveys 60,000 households each month. As mentioned later in this section, the CPS was used for job certification data that proxies for licensing requirements; this is because the ACS does not include this data. Aside from the certification data, the larger sample size provided in the ACS is more suitable for state-aggregated data in this study given its focus on only one industry.

[46] As a robustness check, models in this study were re-estimated with all 50 states included and a total state-year sample of 200. This did not meaningfully change the results of this study; therefore, results were robust to the sample reduction. In these models, the presence of occupational licensing for all three occupations was associated with an 8.6 to 11 percentage point increase in earnings of electricians, plumbers, and operating engineers; this range was nearly identical to the main results presented in Table 2. Similarly, employment results did not meaningfully change; the presence of occupational licensing was associated with a 1.3 to 1.5 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers.

literature and only include working-age adults who are not in school and do not live in group quarters (e.g., jails). This is done to ensure that effects of licensing are captured for individuals who are actually in the workforce.

Dependent and Independent Variables

Hypothesis 1 stated that electricians, plumbers, and operating engineers will have higher earnings in states with occupational licensing requirements than in states without licensing requirements. For this hypothesis, the dependent variable is y_{st}, defined as the natural logarithm of average earnings of electricians, plumbers, and operating engineers indexed by state (s) and year (t).[47] The earnings dependent variable as well as the dependent variables used for employment (explained below) do not depend on whether these occupations feature licensing requirements in a given state and year. Rather, all three occupations are included in this study's dependent variables in every state, based on the fact that these are the three occupations in U.S. construction that tend to feature licensing requirements. To mitigate effects of outliers, prior to aggregating earnings to the state-year level and calculating the natural log, earnings were top- and bottom-coded based on the top and bottom 2 percent of the distribution.

Hypothesis 2 predicted that employment of electricians, plumbers, and operating engineers will be lower in states with licensing requirements than in states without licensing requirements. For this hypothesis, two different dependent variables were used. The first dependent variable is *Proportion Employed*$_{st,}$ defined as the proportion of construction workers that are employed as plumbers, electricians, or operating engineers, also indexed by state and year. The rationale for this dependent variable is that increases (or decreases) in employment in

[47] ACS occupational codes do not distinguish between different types of employment within a given construction occupation. For example, the data does not distinguish between helpers of electricians, journeyworkers, and master electricians; the same is true for plumbers. Thus, the results of this study cannot speak to further employment variation that exists within each construction occupation.

licensed occupations should be reflected by higher (or lower) proportions of workers employed as electricians, plumbers, or operating engineers. However, it is plausible that electricians, plumbers, and operating engineers may be more involved in certain types of construction projects than others. For example, workers in these licensed occupations may be more involved in non-residential projects than residential projects; as a result, higher proportions of electricians, plumbers, and operating engineers in a given state and year could partially reflect higher proportions of certain types of projects. This leads to the second dependent variable for Hypothesis 2: $Employment_{st}$, defined as the natural log of total employment of electricians, plumbers, and operating engineers, indexed by state and year. While the first dependent variable is proportional, this dependent variable measures the total level of employment in these three occupations.

For the main independent variables of interest, I use two distinct datasets for two different measures of licensing in this study. For the first definition of licensing, I use 2016-2019 Current Population Survey (CPS) data provided by IPUMS (Ruggles et al. 2022). I focus on the CPS variable JOBCERT, an indicator with a value of 1 if a worker states that certification is required for them to do their job, and 0 otherwise.[48] The year 2016 was the first year that JOBCERT was included in the CPS, preventing the usage of prior years in my analysis. The underlying assumption in using this variable to proxy for licensing is that certification requirements should be highly correlated with licensing requirements across states, making certification requirements a useful proxy for licensing requirements.[49] Rather than the 50-50 rule

[48] In using the JOBCERT variable, I incorporated its data quality flag, QJOBCERT, to account for cases where respondents do not answer the certification question and the Census Bureau indicates an answer in their place. I only include observations where the value of QJOBCERT equals zero; these are observations in which respondents answered the question and for which there was no change in value.

[49] Other studies in the literature (e.g., Blair & Chung 2019) base their licensing variable on a 50-50 rule: an occupation is defined as a licensed occupation if at least 50 percent of workers in a state-occupation pair report having a license or certification. For example, Blair and Chung (2019) use the CPS variable "STATECERT", an

used in some prior studies, this study employs a 70-30 rule in classifying occupations as licensed in a state-year cell: an occupation is defined as a licensed occupation if at least 70 percent of workers in a given state and year report that certification is required for their respective occupation in the CPS. The justification for this higher bar is that if an occupation does indeed require certification in a given state, it should be reflected in the data via a high level of consensus among workers in that occupation.[50]

For the second definition of licensing, this study introduces a new 2019 data set on licensing requirements for electricians, plumbers, and operating engineers. This data set was constructed based on correspondence with union contacts in each of these occupations. The United Association of Journeymen and Apprentices of the Plumbing and Pipefitting Industry (UA) graciously provided results of an internal UA survey that provided a comprehensive list of states that required licensure of plumbers. Similarly, contacts at the National Electrical Contractors Association (NECA), the International Union of Operating Engineers (IUOE), and the Pennsylvania Foundation for Fair Contracting (PAFFC) provided information on state licensing requirements for electricians and operating engineers. This resulted in a comprehensive data set for 2019 on state licensing requirements for plumbers, electricians, and operating engineers. This results in two different measures of licensing requirements used in this study: the first measure is based on 2016-2019 job certification data from the CPS that proxies for

---

indicator with a value of 1 if a worker possesses a government-issued professional certification or license, and 0 otherwise. If at least 50 percent of workers in a state-occupation pair respond in the affirmative, the occupation is classified as a "licensed occupation" in a given state (Blair & Chung 2019). While the STATECERT variable may deal with licensing more directly, it is still limited by the fact that it combines licensure with certification and does not distinguish between those two forms of occupational regulation. Under the assumption that certification requirements are highly correlated with licensing requirements across states, the JOBCERT variable used in this paper should be a better proxy for licensing requirements because it captures whether workers are *required* to be certified to do their job, rather than whether or not the workers have a license or certification.

[50] As a robustness check, separate models were estimated with licensing definitions based on the 50-50 rule used in prior work. As explained further later in the paper, results were robust to this alternative definition and did not meaningfully change.

licensing, and the second measure is based on a new data set from 2019 that includes more detailed information on licensing requirements in all three occupations by state. While the second measure is limited by its one-year state sample size of 45, both of these licensing measures are included for two reasons. First, even with the smaller sample size, the second definition is arguably more robust, since it is based on a data set that specifically focuses on state licensing requirements in construction. Second, including both measures of licensing in this study allows for a comparison between the two measures and can be informative with respect to whether certification requirements may indeed be a useful proxy for licensing requirements.

For the first set of earnings and employment models that tests Hypotheses 1 and 2, the independent variable of interest is $License\text{-}70_{st}$, an indicator with a value of 1 if all three occupations are licensed in a given state (s) and year (t) and 0 otherwise, based on the 70-30 rule for certification requirements that proxies for licensing requirements. For the second set of earnings and employment models, the independent variable of interest is $License_{s}$. This variable is not indexed by year because it is based on this study's detailed 2019 licensing data collected via correspondence with building trades unions. It is an indicator with a value of 1 if all three occupations are licensed in a given state in 2019, and 0 otherwise. The third and fourth sets of earnings and employment models involve separate licensing indicator variables for electricians, plumbers, and operating engineers. For the third set of models, the indicator variables are indexed by state and year and use the 70-30 certification requirement rule; for the fourth and final set of models, the indicator variables are only indexed by state and are based on the 2019 licensing data only.

Controls

First, demographic controls used in this study come from 2016-2019 ACS data. These variables are also aggregated to the state level and featured in the models as proportions. Though the variables are aggregated to the state level, the vector X_{st} follows prior licensing literature by including demographic controls related to gender, race, ethnicity, educational attainment, and age. For gender, I control for the proportion of construction workers that are female. In a male-dominated industry such as construction, higher proportions of female workers may be associated with lower average earnings and employment as a result of possible gender pay and employment gaps. For race and ethnicity, I control for proportions of construction workers that are black and Hispanic. Higher proportions of black and/or Hispanic workers may also be associated with lower average earnings and employment. In aggregating age to the state level, I separate construction workers into age groups, resulting in proportions of construction workers that are (a) 23 to 29 years old; (b) 30 to 36 years old; (c) 37 to 43 years old; (d) 44 to 50 years old; (e) 51 to 57 years old; and (f) 58 to 64 years old, with 16 to 22 years old as the reference group. This is done to capture the effect of each particular age group as well as to examine how associations between age and earnings as well as employment might change over the course of construction workers' careers. Lastly, rather than controlling for average educational attainment, I include two educational controls: the proportion of construction workers with a high school diploma, as well as the proportion of construction workers who are college-educated.

The vector Y_{st} is a vector of construction industry controls, most of which come from 2016-2019 Quarterly Census of Employment and Wages (QCEW) data. First, I control for the annual average weekly wage and employment "location quotients." Location quotients are designed to control for the fact that some areas of the country, on average, have higher values of a given variable than others. In this case, the weekly wage location quotient controls for some areas of the country having higher (or lower) average wages or employment for construction workers than others (e.g., higher or lower costs of living, larger or smaller construction industries). This location quotient variable is defined as follows: a location quotient of 1 means the area has a weekly construction wage or employment value in line with the U.S. average; a location quotient greater than 1 means the area has a value greater than the U.S. average; and a location quotient less than 1 means the area has a value less than the U.S. average. Second, I control for average firm size, defined as annual average employment divided by the number of establishments (i.e., firms). I also control for private construction union density (Hirsch & Macpherson 20003), as prior literature (e.g., Kleiner & Krueger 2010, 2013; Gittleman & Kleiner 2016) finds union earnings premiums that are independent of licensing premiums.

Models

Hypothesis 1 predicted that construction workers in licensed occupations (i.e., electricians, plumbers, and operating engineers) will have higher earnings in states with occupational licensing requirements than in states without occupational licensing requirements. As mentioned previously, dependent variable of interest for Hypothesis 1 is *y*, defined as the natural logarithm of average earnings of construction workers in licensed occupations and indexed by state (s) and year (t). The OLS earnings equations estimated in this study are as follows:

(1) $y_{st} = \beta_0 + \beta_1 License\text{-}70_{st} + \boldsymbol{\beta_2 X_{st}} + \boldsymbol{\beta_3 Y_{st}} + \phi_t + \mu_{st}$

(2) $y_s = \beta_0 + \beta_1 License_s + \boldsymbol{\beta_2 X_s} + \boldsymbol{\beta_3 Y_s} + \mu_s$

(3) $y_{st} = \beta_0 + \beta_1 Electrical\ Licensing\text{-}70_{st} + \beta_2 Plumbing\ Licensing\text{-}70_{st} + \beta_3 Operating\ Engineer\ Licensing\text{-}70_{st} + \boldsymbol{\beta_4 X_{st}} + \boldsymbol{\beta_5 Y_{st}} + \phi_t + \mu_{st}$

(4) $y_s = \beta_0 + \beta_1 Electrical\ Licensing_s + \beta_2 Plumbing\ Licensing_s + \beta_3 Operating\ Engineer\ Licensing_s + \boldsymbol{\beta_4 X_s} + \boldsymbol{\beta_5 Y_s} + \mu_s$

where X is the vector of demographic controls, Y is the vector of construction industry controls, ϕ_t denotes year fixed effects, and μ is the error term. Earnings equations (1) and (3) utilize licensing variables based on the 70-30 certification rule and feature variables indexed by state and year; earnings equations (2) and (4) utilize licensing variables based on 2019 data only and feature variables that are only indexed by state.

Hypothesis 2 predicted that employment in licensed occupations will be higher in states with licensing requirements than in states without licensing requirements. As mentioned previously, the dependent variables of interest for Hypothesis 2 are *Proportion Employed* and *Employed*. *Proportion Employed* is measured as the proportion of construction workers that are employed as plumbers, electricians, or operating engineers; *Employed* is measured as the natural log of total employment of electricians, plumbers, and operating engineers. The OLS employment equations estimated in this study are as follows:

(5) $Proportion\ Employed_{st} = \beta_0 + \beta_1 License\text{-}70_{st} + \boldsymbol{\beta_2 X_{st}} + \boldsymbol{\beta_3 Y'_{st}} + \phi_t + \mu_{st}$

(6) $Proportion\ Employed_s = \beta_0 + \beta_1 License_s + \boldsymbol{\beta_2 X_s} + \boldsymbol{\beta_3 Y'_s} + \mu_s$

(7) $Proportion\ Employed_{st} = \beta_0 + \beta_1 Electrical\ Licensing\text{-}70_{st} + \beta_2 Plumbing\ Licensing\text{-}70_{st} + \beta_3 Operating\ Engineer\ Licensing\text{-}70_{st} + \boldsymbol{\beta_4 X_{st}} + \boldsymbol{\beta_5 Y'_{st}} + \phi_t + \mu_{st}$

(8) $Proportion\ Employed_s = \beta_0 + \beta_1 Electrical\ Licensing_s + \beta_2 Plumbing\ Licensing_s + \beta_3 Operating\ Engineer\ Licensing_s + \boldsymbol{\beta_4 X_s} + \boldsymbol{\beta_5 Y'_s} + \mu_s$

(9) $Employed_{st} = \beta_0 + \beta_1 License\text{-}70_{st} + \boldsymbol{\beta_2 X_{st}} + \boldsymbol{\beta_3 Y'_{st}} + \phi_t + \mu_{st}$

(10) $Employed_s = \beta_0 + \beta_1 License_s + \boldsymbol{\beta_2 X_s} + \boldsymbol{\beta_3 Y'_s} + \mu_s$

(11) $Employed_{st} = \beta_0 + \beta_1 Electrical\ Licensing\text{-}70_{st} + \beta_2 Plumbing\ Licensing\text{-}70_{st} + \beta_3 Operating\ Engineer\ Licensing\text{-}70_{st} + \boldsymbol{\beta_4 X_{st}} + \boldsymbol{\beta_5 Y'_{st}} + \phi_t + \mu_{st}$

(12) $Employed_{s} = \beta_0 + \beta_1 Electrical\ Licensing_{s} + \beta_2 Plumbing\ Licensing_{s} + \beta_3 Operating\ Engineer\ Licensing_{s} + \boldsymbol{\beta_4 X_{s}} + \boldsymbol{\beta_5 Y'_{s}} + \mu_{s}$

where the vector X is the same as before and the vector Y' is a vector of construction industry controls specific to employment.

The vector Y' retains average firm size and private construction union density from the vector Y in the earnings equations, but it replaces the weekly wage location quotient with a location quotient related to annual construction employment that controls for some areas of the country having higher (or lower) construction employment than others (e.g., construction industries that are larger or smaller than other areas). In the first set of employment equations (equations 5-8), equations (5) and (7) utilize licensing variables based on the 70-30 certification rule and feature variables indexed by state and year; equations (6) and (8) utilize licensing variables based on 2019 data only and feature variables that are only indexed by state. Likewise, in the second set of employment equations (equations 9-12), equations (9) and (11) utilize licensing variables based on the 70-30 certification rule and feature variables indexed by state and year; equations (10) and (12) utilize licensing variables based on 2019 data only and feature variables that are only indexed by state.

Results

All standard errors in this study are clustered by state. Table 1 presents summary statistics for variables used in this analysis, and Table 2 presents a detailed summary of licensing requirements by state and occupation. In Table 2, the first three columns focus on the 70-30 rule and list the number of years in the period 2016-2019 in which each respective occupation required licensing in each state based on this rule. The last three columns focus on the 2019 data set and list whether the occupation was licensed in a given state. An annual average of 25.5 percent of states required licensing for all three occupations based on the 70-30 certification rule used in the CPS, nearly identical to the 24 percent of the 2019 state sample that required licensing for all three occupations based on the building trades union data set.[51] Based on the 70-30 rule, 13 states required licensing for all three occupations in 2016; 12 states required licensing for all three occupations in 2017; in 2018 and 2019, those numbers of states were 13 and 14, respectively. In 2019, based on the new data, 11 states required licensing for all three occupations. The average state-year proportion of construction workers who were female was 8.6 percent, reflecting the male-dominated nature of the industry.[52] While the average black proportion was only 5.4 percent, the average Hispanic proportion was 20.1 percent, reflecting the fact that construction is one of the industries with the most Hispanic workers (Siniavskaia 2018).

Table 3 presents results for earnings equations (1) and (3), which focus on the 70-30 definition of licensing. Models 1-3 focus on equation (1), while Models 4-6 focus on equation (3). All sets of models in Tables 3 and 4 proceed in the same fashion: the first model includes the licensing independent variable(s) of interest and the demographic controls from the ACS; the second model adds construction industry controls; and the third model adds private construction union density as a control. In Models 1-3 of Table 3, results suggest that the presence of occupational licensing for all three occupations (electricians, plumbers, and operating engineers)

---

[51] Further, *License-70* and *License* were highly positively correlated, suggesting that certification requirements at the state level may serve as a useful proxy for state licensing requirements.

[52] This female share of the construction workforce is similar to estimates from the Center for Construction Research and Training (CPWR), which suggest that women comprise approximately 9.2% of the construction workforce. The CPWR estimate (as well as this study's estimate) include clerical and support work, which features the highest proportion of women in the industry.

is associated with a statistically significant 8.3 to 11.6 percentage point increase in earnings of electricians, plumbers, and operating engineers[53] These findings are in line with the positive licensing earnings premium established by prior literature and within the range of earnings effects found in prior work (5 to 18 percent). In most specifications, higher proportions of workers who are black are associated with lower earnings for electricians, plumbers, and operating engineers. Additionally, in most specifications, higher proportions of workers who are between the ages of 30 and 43 as well as between the ages of 51 and 57 are associated with higher earnings for electricians, plumbers, and operating engineers, while coefficients for younger workers (ages 23-29) and older workers (ages 58-64) were not statistically significant. As expected, this suggests that construction workers have higher earnings in their prime earning years. Lastly, a 1% increase in private construction union density is associated with approximately a 1.9 percentage point increase in earnings. This is consistent with prior literature (e.g., Kleiner & Krueger 2010, 2013; Gittleman & Kleiner 2016) that finds positive union earnings premiums independent of licensing premiums.

Models 4-6 of Table 3 focus on equation (3), where separate indicator variables are used for electrical licensing, plumbing licensing, and operating engineer licensing. Across all three specifications, results suggest that operating engineer licensing may be the main driver of positive associations between licensing and earnings in Models 1-3: the presence of licensing for operating engineers is associated with a statistically significant 7.4 to 12 percentage point increase in earnings across all three occupations. The coefficients for electrical and plumbing

[53] These models were also estimated with licensing definitions based on the 50-50 rule. This did not meaningfully change the results; the presence of occupational licensing for all three occupations in a given state and year was associated with a statistically significant 8 to 12.9 percentage point increase in earnings of electricians, plumbers, and operating engineers. Additionally, in models where separate indicators were used for electrical licensing, plumbing licensing, and operating engineer licensing, results were robust to the usage of the 50-50 rule; the presence of operating engineer licensing was associated with a statistically significant 7.4 to 12.5 percentage point increase in earnings of electricians, plumbers, and operating engineers.

licensing were not statistically significant. This may be because operating engineer licensing requirements are less common than plumbing and electrical licensing. While an annual average of 87 and 85 percent of states required licensing for electricians and plumbers respectively, only an average of 45 percent of states required licensing for operating engineers. Similar respective proportions were also found in the 2019 licensing data set. With a vast majority of states requiring licensing for electricians and plumbers (and those two types of licensing requirements being highly positively correlated), the results in Models 4-6 suggest that the positive association between licensing and earnings may be largely driven by areas where operating engineers are also required to be licensed in addition to electricians and plumbers.

Table 4 presents results for earnings equations (2) and (4), which define licensing based on the 2019 building trades union data set. Similar to Table 3, Models 1-3 focus on equation (2), while Models 4-6 focus on equation (4). In Models 1 and 2, results suggest that the presence of occupational licensing for all three occupations (electricians, plumbers, and operating engineers) is associated with a statistically significant 12.4 to 14.8 percentage point increase in earnings of electricians, plumbers, and operating engineers, similar to albeit slightly higher than the associations found in Table 3. While the coefficient on licensing remains positive, the statistical significance vanishes in Model 3, suggesting that the results in Table 4 are more sensitive to specification. Models 4-6, which focus on earnings equation (4), follow a similar pattern. While the operating engineer coefficients in Models 4 and 5 suggest that operating engineer licensing requirements are associated with an 11.5 to 13.6 percentage point increase in earnings, the statistical significance vanishes in Model 6. The increased sensitivity to specification in Table 4 likely results from the significantly smaller sample size. Even so, the results of Table 4, as was the case with Table 3, are in line with the range of earnings effects found in prior work (5 to 18 percent). Overall, across Tables 3 and 4, results suggest that occupational licensing requirements are associated with an 8.3 to 14.8 percentage point increase in earnings,, with results also suggesting that operating engineer licensing requirements may be the key driver of these positive associations.

Though these earnings results are in line with the consensus of prior studies, they cannot provide clear evidence in favor of or against either theory of licensing presented in this paper, for two reasons. First, as the models were estimated using OLS, the findings represent associations. There may be other factors that could possibly explain these results, such as (a) states with more robust construction industries having more licensing requirements and/or (b) licensing serving as a stronger barrier to entry in states with higher union densities. Building trades unions may want barriers to entry in place that protect their members from lower-skilled competition, thereby ensuring higher wages for their members. It therefore cannot be said with certainty that these in fact represent causal impacts of licensing on earnings of construction workers in licensed occupations. Second, as mentioned in the introduction, both theoretical frameworks presented in this paper predict positive earnings effects of licensing. Findings of positive associations between licensing and earnings of electricians, plumbers, and operating engineers therefore cannot disentangle these two theories. Even so, as expected, the licensing earnings premium seems to be significantly positive, and the magnitudes fall within the range established by prior literature.

Table 5 presents results for employment equations (5) and (7). This table proceeds the same as before: Models 1-3 focus on equation (5), while Models 4-6 focus on equation (7) and include separate indicators for electrical, plumbing, and operating engineer licensing requirements. Licensing is defined by the 70-30 rule, and the employment dependent variable is

Proportion Employed, measured as the proportion of construction workers employed as plumbers, electricians, or operating engineers. Across Models 1-3, results suggest the presence of occupational licensing for all three occupations is associated with a statistically significant 1.2 to 1.3 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers.[54] Further, these results are similar to the earnings results in that operating engineer licensing requirements seem to be the main driver of the positive associations. Across Models 4-6, operating engineer licensing requirements are associated with a 1.4 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers, while the coefficients for electrical and plumbing licensing were not statistically significant.[55] As was the case with earnings, positive associations between licensing and employment may be largely driven by areas where operating engineers are also required to be licensed in addition to electricians and plumbers, since electrical and plumbing licensing requirements are present in a vast majority of states and are correlated. Overall, these employment findings seem to support the quality signal and human capital framework: the signal

[54] These models were also re-estimated with licensing defined using the 50-50 rule. Employment results were also robust to this change; the presence of licensing requirements for all three occupations in a given state and year was associated with a 1.8 to 2 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers. Additionally, in models where separate indicators were used for electrical licensing, plumbing licensing, and operating engineer licensing, results were robust to the usage of the 50-50 rule; the presence of operating engineer licensing was associated with a statistically significant 1.9 to 2.2 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers.

[55] As the dependent variable here is a proportion, these same models were also estimated using a generalized linear model (GLM) specifying a binomial family with a logit link, a variant of logistic regression (Papke & Wooldridge 1996). This accounts for the possibility that the errors of this proportional dependent variable are not normally distributed, which can be an issue with proportional dependent variables. Compared to logistic regression, GLM estimation is attractive because it accounts for possible non-normal error distributions without the need to logit transform the dependent variable. Positive, statistically significant associations between licensing and employment (and between operating engineer licensing and employment) were also found in these models; results are available from the author upon request.

of quality licensing provides and the resulting positive demand effects may outweigh negative supply effects resulting from barriers to entry.[56]

Table 6 presents results for employment equations (9) and (11). Models 1-3 focus on equation (9), while Models 4-6 focus on equation (11). Licensing remains defined by the 70-30 rule, but the employment dependent variable is now *Employed*, measured as the natural log of total employment of electricians, plumbers, and operating engineers. This dependent variable is used to account for the possibility that the positive licensing-employment associations found in Table 5 may not reflect increases in licensed occupation employment but may instead reflect decreases in *unlicensed* occupation employment. No significant associations were found in Table 5 between occupational licensing requirements and employment of electricians, plumbers, and operating engineers.

Taken together, the results in Tables 5 and 6 suggest that employment effects of licensing in construction may be negligible or slightly positive. This seems to contradict most prior occupational licensing literature that has found negative employment and labor supply effects of licensing. Employment could increase or decrease depending on the relative importance of competing supply and demand effects of licensing (Carollo 2020). While licensing can lower employment in licensed occupations via barriers to entry, it can also signal higher levels of worker training, thereby reducing asymmetric information between workers and consumers and increasing labor demand. Findings of negligible to slightly positive associations between occupational licensing and employment suggest that these supply and demand effects likely counterbalance one another in construction, with labor demand effects potentially slightly

[56] Earnings and employment results were robust to alternative classification rules. For example, alternative models were run with licensing defined based on an 80-20 rule (a higher bar); employment and earnings results held.

outweighing labor supply effects. This is plausible, given the health and safety risks poor construction work poses to consumers.

However, there are other possible explanations that warrant further exploration. First, it may be the case that states with more robust construction industries (i.e., higher total construction employment) are the ones with more licensing requirements. This could potentially be an alternative explanation of the positive associations found in Table 4. A second alternative explanation is connected to union density: it could be that states with higher union densities are also the states where licensing serves as a greater barrier to entry. This could derive from unions in these high-skill occupations wanting barriers to entry in place that protect their members from lower-skilled competition, thereby increasing employment for their members. To analyze the first alternative explanation (i.e., that states with larger and more robust construction industries also have more licensing requirements), I conduct supplemental analyses with two new dependent variables: (a) the proportion of *all* workers (i.e., construction and non-construction) employed as electricians, plumbers, and operating engineers by state and year, and (b) the proportion of all workers employed in *unlicensed* construction occupations (i.e., all occupations except electricians, plumbers, and operating engineers), by state and year.[57] If it is indeed the case that states with larger construction industries also have more licensing requirements, positive associations between licensing and employment should be found for *both* sets of construction workers and should not only be concentrated in occupations that may require licensing.

[57] For this analysis, it was not possible to use the same type of dependent variable as was used in Tables 4 and 5. For these new licensed and unlicensed worker proportions, rather than the denominator being comprised of only construction workers as before, it was necessary to change the denominator to include all workers. A denominator made up only of construction workers would result in mechanically opposite coefficients between the two sets of models, because the proportion of construction workers in unlicensed occupations would simply be 1 – the proportion in licensed occupations. It was thus necessary to measure these proportions out of all workers in a given state and year.

Tables 7 and 8 present these results. In Table 7, the dependent variable is the proportion of all workers (i.e., construction and non-construction workers) employed as electricians, plumbers, and operating engineers. The models proceed in the same way as before, and all variables remain aggregated to the state level. While the coefficients are naturally much smaller as a result of the larger dependent variable denominators, results in Table 7 suggest that the presence of occupational licensing is associated with a 0.1 percentage point increase in the proportion of the total state-year workforce employed as electricians, plumbers, and operating engineers. Licensing requirements are then associated with increases in the respective proportions of construction workers and *all* workers employed as electricians, plumbers, and operating engineers. These results must be compared with results in Table 8 to assess the alternative explanation that states with larger construction industries (i.e., higher total employment) also have more licensing requirements.

In Table 8, the dependent variable is the proportion of all workers employed in *unlicensed* construction occupations (i.e., all occupations except electricians, plumbers, and operating engineers). This is done to compare effects on unlicensed occupations against effects on electricians, plumbers, and operating engineers. The licensing coefficients in Table 7 are negative and not statistically significant; I do not find positive associations between occupational licensing and employment in unlicensed occupations as a share of the total state-year workforce. Taken together, results in Tables 7 and 8 suggest that licensing requirements are associated with higher proportions of workers employed as electricians, plumbers, and operating engineers but do not have a significant association with employment in unlicensed occupations. While the results of this supplemental analysis do not provide direct support for the alternative explanation, it also cannot be entirely ruled out; nonsignificant impacts on employment in occupations other than electricians, plumbers, and operating engineers may still be consistent with licensing acting as a barrier to entry. As electricians, plumbers, and operating engineers are the only occupations in construction that are significantly directly impacted by licensing requirements, it is conceivable that employment effects of licensing acting as a barrier to entry would not be felt in other occupations.[58]

A second alternative explanation is that states with higher union densities are also the states where licensing serves as a greater barrier to entry. This could result from unions in these high-skill occupations wanting barriers to entry in place that protect their members from lower-skilled competition. To analyze this possible alternative explanation, the final analysis in this paper examines union density as a moderator of the association between licensing and employment in licensed occupations. Table 9 presents this analysis. As a starting point, Model 1 presents the employment model also found in Model 3 of Table 4. In Models 2 and 3, I separate the sample into two parts: Model 2 includes state-year cells that have union densities that are below the mean (0.13), whereas Model 3 includes state-year cells that are above the mean union density. This is done to compare the respective strengths of licensing-employment associations between states where construction unions are weaker (Model 2) and states where construction unions have a stronger presence (Model 3). Lastly, in Model 4, I include an interaction term of licensing and union density.

In Model 2, where construction unions are weaker, occupational licensing requirements for all three licensed occupations are associated with a 1.8 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers.

[58] Given the usage of proportional dependent variables in Tables 6 and 7, additional models were estimated using GLM estimation, as was the case for the models presented in Table 4. Results are available from the author upon request; results in Tables 6 and 7 were robust in these specifications.

Conversely, in Model 3 (where construction unions are stronger), this positive association vanishes and is no longer statistically significant. These differing coefficients between Models 2 and 3 were found to be statistically significant. However, even with this difference, neither of these coefficients suggest negative impacts of licensing that would otherwise be consistent with licensing acting as a barrier to entry. Further, in Model 4, the coefficient on the interaction term is not statistically significant; a significantly negative interaction would be consistent with the alterative explanation. Taken together, the results in Table 9 do not provide evidence in support of the alternative explanation that alternative explanation that states with higher union densities are also the states where licensing serves as a greater barrier to entry. Across Models 2 and 3 of Table 9, neither licensing coefficient is significantly negative, nor is the interaction term in Model 4 significantly negative.

Discussion and Conclusion

In sum, the presence of occupational licensing for all three construction occupations that require licensing in certain states (electricians, plumbers, and operating engineers) is associated with a statistically significant 8.3 to 14.8 percentage point increase in earnings of electricians, plumbers, and operating engineers. These findings are in line with the positive licensing earnings premium established by prior literature and within the range of earnings effects found in prior work (5 to 18 percent). For employment, the evidence is more mixed: the presence of occupational licensing for all three occupations is associated with a statistically significant 1.2 to 1.3 percentage point increase in the proportion of construction workers employed as electricians, plumbers, or operating engineers. However, these associations vanish when using an alternative dependent variable centered on the level of employment of electricians, plumbers, and operating engineers. Overall, results suggest occupational licensing is associated with higher earnings and negligible to slightly higher employment for electricians, plumbers, and operating engineers. This suggests that labor demand effects may outweigh supply effects in the construction context. Yet as suggested in supplemental analyses, the alternative explanation that states with more robust construction industries (i.e., higher total construction employment) are the ones with more licensing requirements cannot be ruled out.

What explains the differing results that suggest negligible to slightly positive employment effects of licensing? This could result from changes in the proportional employment dependent variable primarily reflecting fewer workers in occupations that do not require licensing, rather than reflecting more workers employed as electricians, plumbers, or operating engineers. In this case, lower proportions of construction workers in other occupations naturally raises the proportion of construction workers employed as electricians, plumbers, and operating engineers. However, these effects on levels of employment in other occupations would not be reflected in the additional employment dependent variable that explicitly focuses on levels of employment of electricians, plumbers, and operating engineers. This could explain why licensing may raise the *proportion* of construction workers employed as electricians, plumbers, or operating engineers but not raise the *level* of employment in these three occupations.

The literature has presented two different views of occupational licensing. The most common theory is that licensing primarily represents rent-seeking behavior, acting as a barrier to entry and restricting supplies of labor to a licensed occupation. This shifts the labor supply curve leftward and increases the price of labor (Rottenberg 1980; Kleiner 2000). Other literature has suggested that licensure creates a greater incentive for workers to invest in more occupation-specific human capital (Shapiro 1986). Licensing may increase demand for services and shift the labor demand curve rightward by resolving informational asymmetries between workers and firms, which may be particularly important in contexts where low-quality work poses significant health and safety risks (Leland 1979; Carollo 2020). While both frameworks predict positive earnings effects of licensing, employment could increase or decrease depending on the relative importance of these competing supply and demand effects (Carollo 2020). The risks posed by construction work to consumers necessitate high levels of training, and licensure may signal to construction firms that workers have indeed achieved the requisite level of training to properly complete the work. This can help correct market distortions resulting from informational asymmetries and increase labor demand in a context where work quality may otherwise be difficult to ascertain.

However, these results do not constitute causal effects; the positive associations between licensing and employment may also still reflect licensing primarily acting as a barrier to entry and limiting competition. Supplemental analysis cannot rule out the alternative explanation that states with more robust construction industries also feature more licensing requirements. Though this study examines labor market outcomes of occupational licensing in a construction context not yet studied in prior work, the descriptive nature of these results invites future research to build on this work by establishing causal impacts of licensing on earnings and employment in construction. This would shed more light on which of the two competing effects is stronger in the construction context.

Like all studies, this study has limitations. First, using certification requirements as proxies for licensing requirements in this study assumes that certification requirements are highly correlated with licensing requirements across states. Much of the literature on occupational licensing has utilized data sources, such as the Survey of Income and Program Participation (SIPP), that combine licensing requirements with certification requirements and do not provide a mechanism to disentangle the two. Other literature has defined licensing based on data on the proportion of workers in an occupation who have a license or certification, which also combines licensure with certification and does not speak to whether licensure or certification is required in a particular state. In addition to the new construction licensing data this study introduces, the certification requirement data used in this paper should be a better proxy for licensing requirements than other data used in prior literature because it captures whether workers are *required* to follow occupational regulation to do their job, rather than whether or not the workers have a license or certification. Indeed, the two distinct measures of licensing in this study were positively correlated and produced similar results, suggesting that this proxy may be useful. However, this rests on the assumption that certification and licensing requirements are indeed highly correlated. Future work can improve on this by obtaining comprehensive data specific to construction licensing requirements across states that runs over multiple years. This would result in a more reliable measure of licensing requirements as well as a complete picture of how licensing requirements may vary over time.

Second, all dependent variables used in this study constitute aggregate earnings and employment outcomes of three distinct construction occupations. These occupations were grouped together because they represent a high-skilled portion of the construction workforce that differs from the rest of the workforce with respect to licensing requirements. As previously mentioned, electricians, plumbers, and operating engineers are similar in that they are the only three occupations in construction that tend to feature licensing requirements in certain states. These aggregated dependent variables are designed to capture the impact of licensing on this high-skilled portion of the construction workforce as a whole. However, there are differences between these occupations across certain dimensions; for example, while electricians and plumbers most commonly train via apprenticeships, operating engineers are more likely to learn on-the-job than go through apprenticeship programs (Bureau of Labor Statistics).

Further, while these three construction occupations are the only occupations that tend to require licensing across certain states, they differ with respect to the number of states in which these occupations are required to be licensed. As Tables 1 and 2 indicate, while electricians and plumbers were required to be licensed in over 80% of this study's state-year cells (thereby making licensing requirements for these two occupations highly positively correlated), less than half required operating engineers to be licensed, making operating engineer licensing less common. This could explain why licensing effects on earnings and employment were concentrated among operating engineers; the similarly high extent of licensing for electricians and plumbers likely precluded detection of effects for those two occupations. Even so, findings of a negligible to slightly positive effect of licensing on employment contradict most prior licensing literature and suggest that demand effects of licensing may outweigh supply effects in construction. This is plausible, given the health and safety risks posed by poor construction work and the signal of quality licensing can provide to construction firms.

Lastly, the results of this study cannot speak to further variation that exists within licensing requirements in construction. There are multiple dimensions of licensing that can affect entry into licensed occupations. The first dimension is the qualifications needed to obtain a license, particularly passing an examination and obtaining work experience in the industry. The second dimension is whether all workers must be licensed in order to work in their respective occupation. In some states, construction licensing requirements are more stringent and do not allow unlicensed workers to work in a licensed occupation. However, in other states, workers in licensed occupations can work on projects without a license so long as a licensed contractor supervises them. Lastly, other states only require workers to have their work inspected by a licensed worker. There may be significant variation across states along this dimension, and future research is strongly encouraged to explore how these additional sources of heterogeneity may affect labor market outcomes of licensed workers.

It is plausible that the licensing earnings premium in construction would be lower in areas where only contractors are required to be licensed, due to licensing not presenting as much of a barrier to entry for workers and the resulting increased competition. It may also be the case that employment effects of licensing would lessen in states where only contractors are required to be licensed. Additionally, as mentioned previously, licensing requirements in construction also vary based on reciprocity agreements. These agreements allow workers who previously passed an examination and obtained a license in a state to remain licensed in another state *without* an additional examination (National Center for Construction Education and Research). Given recent work suggesting that licensing inhibits geographic mobility (Johnson & Kleiner 2020), it would be informative to explore whether this holds in the construction context given the nature of reciprocity agreements and how this might impact earnings and employment. These areas can be explored via collection of comprehensive construction-specific licensing data over multiple years that specifies which states are in reciprocal agreements and also whether only contractors or all workers are required to be licensed.

APPENDIX

Table 2.1: Descriptive Statistics

Variables	Means	Standard Deviations
Log average earnings of workers in licensed occupations	10.79	0.16
Proportion of construction workers employed in licensed occupations	0.15	0.02
Log total employment in licensed occupations	9.89	0.83
License-70	0.26	0.44
Electrical licensing-70	0.87	0.34
Plumbing licensing-70	0.85	0.36
Operating engineer licensing-70	0.45	0.50
License	1.85	1.01
Percent female	0.09	0.02
Percent black	0.05	0.05
Percent Hispanic	0.20	0.16
Percent high school	0.51	0.05
Percent college	0.10	0.02
Percent age 23-29	0.15	0.02
Percent age 30-36	0.18	0.02
Percent age 37-43	0.18	0.02
Percent age 44-50	0.17	0.02
Percent age 51-57	0.17	0.02
Percent age 58-64	0.11	0.02
Weekly wage location quotient	1.04	0.12
Employment location quotient	1.03	0.22
Percent employed residential	0.23	0.02
Percent state and local	0.50	0.12
Average firm size	8.20	2.22
Private construction union density	0.14	0.10

Table 2.2: Licensing Requirements by State and Occupation

State	Electricians: 70-30 Rule	Plumbers: 70-30 Rule	Operating Engineers: 70-30 Rule	Electricians: 2019 Definition	Plumbers: 2019 Definition	Operating Engineers: 2019 Definition
Alabama	4 years	4 years	0 years	Yes	Yes	No
Arizona	3 years	3 years	1 year	Yes	Yes	No
Arkansas	4 years	4 years	0 years	Yes	Yes	No
California	4 years	3 years	3 years	Yes	Yes	Yes
Colorado	4 years	4 years	2 years	Yes	Yes	No
Connecticut	4 years	4 years	3 years	Yes	Yes	Yes
Florida	4 years	3 years	1 year	Yes	Yes	No
Georgia	4 years	4 years	1 year	Yes	Yes	No
Hawaii	4 years	3 years	4 years	Yes	Yes	Yes
Idaho	2 years	2 years	1 years	No	No	No
Illinois	4 years	4 years	2 years	Yes	Yes	No
Indiana	1 years	2 years	0 years	No	No	No
Iowa	2 years	2 years	1 year	No	No	No
Kansas	1 year	2 years	2 years	No	No	No
Kentucky	4 years	4 years	0 years	Yes	Yes	No
Louisiana	4 years	4 years	0 years	Yes	Yes	No
Maine	4 years	4 years	2 years	Yes	Yes	No
Maryland	4 years	3 years	4 years	Yes	Yes	Yes
Massachusetts	4 years	4 years	4 years	Yes	Yes	Yes
Michigan	4 years	4 years	1 year	Yes	Yes	No
Minnesota	4 years	4 years	3 years	Yes	Yes	Yes
Mississippi	4 years	4 years	0 years	Yes	Yes	No
Missouri	2 years	2 years	1 year	No	No	No
Montana	1 year	2 years	4 years	No	No	Yes
Nebraska	4 years	2 years	2 years	Yes	No	No
Nevada	4 years	4 years	4 years	Yes	Yes	Yes
New Hampshire	2 years	1 year	1 year	No	No	No

Table 2.2 (cont'd)

New Jersey	4 years	4 years	4 years	Yes	Yes	Yes
New Mexico	4 years	4 years	4 years	Yes	Yes	Yes
New York	2 years	3 years	4 years	No	Yes	Yes
North Carolina	4 years	3 years	2 years	Yes	Yes	No
Ohio	4 years	4 years	1 year	Yes	Yes	No
Oklahoma	4 years	4 years	0 years	Yes	Yes	No
Oregon	4 years	4 years	3 years	Yes	Yes	No
Pennsylvania	3 years	3 years	4 years	No	No	Yes
South Carolina	2 years	3 years	2 years	No	Yes	No
South Dakota	3 years	2 years	1 year	No	No	No
Tennessee	4 years	4 years	0 years	Yes	Yes	No
Texas	4 years	4 years	2 years	Yes	Yes	No
Utah	4 years	4 years	3 years	Yes	Yes	Yes
Vermont	4 years	4 years	2 years	Yes	Yes	No
Virginia	4 years	3 years	1 year	Yes	Yes	No
Washington	4 years	3 years	3 years	Yes	Yes	Yes
West Virginia	2 years	2 years	4 years	No	No	Yes
Wisconsin	4 years	4 years	1 year	Yes	Yes	No

Table 2.3: Earnings Results: 70-30 Rule

Variables	Demographics	Industry Controls	Union Density	Demographics	Industry Controls	Union Density
License-70	**0.116*** **	**0.109*** **	**0.083*** **			
	(0.032)	(0.034)	(0.022)			
Electrical licensing-70				**-0.037**	**-0.021**	**-0.014**
				(0.068)	(0.060)	(0.044)
Plumbing licensing-70				**0.024**	**0.048**	**0.045**
				(0.042)	(0.037)	(0.033)
Operating engineer licensing-70				**0.120*** **	**0.105*** **	**0.074*** **
				(0.036)	(0.036)	(0.024)
Percent female	**0.483**	**-0.307**	**-0.087**	**0.457**	**-0.257**	**-0.049**
	(0.869)	(0.872)	(0.545)	(0.863)	(0.857)	(0.545)
Percent black	**-0.818*** **	**-0.776*** **	**-0.252**	**-0.873*** **	**-0.834*** **	**-0.294**
	(0.292)	(0.255)	(0.221)	(0.301)	(0.261)	(0.225)
Percent Hispanic	**-0.210**	**-0.260**	**-0.144**	**-0.196**	**-0.253**	**-0.134**
	(0.129)	(0.186)	(0.113)	(0.135)	(0.193)	(0.115)
Percent high school	**0.371**	**0.383**	**0.055**	**0.419**	**0.448**	**0.126**
	(0.424)	(0.410)	(0.312)	(0.433)	(0.417)	(0.310)
Percent college	**1.706*** **	**1.884*** **	**0.808**	**1.590** **	**1.845** **	**0.821**
	(0.633)	(0.669)	(0.543)	(0.673)	(0.706)	(0.556)
Percent age 23-29	**1.057**	**0.533**	**-0.685**	**1.005**	**0.519**	**-0.686**
	(1.143)	(0.979)	(0.839)	(1.135)	(0.968)	(0.829)
Percent age 30-36	**3.303*** **	**3.034*** **	**1.006**	**3.157*** **	**2.951*** **	**0.968**
	(0.862)	(0.876)	(0.638)	(0.843)	(0.840)	(0.611)
Percent age 37-43	**2.443** **	**2.202* **	**-0.111**	**2.143** **	**2.232* **	**-0.067**
	(1.159)	(1.114)	(0.807)	(1.176)	(1.120)	(0.811)
Percent age 44-50	**1.836* **	**1.380**	**-0.933**	**1.721**	**1.353**	**-0.924**
	(1.007)	(0.981)	(0.602)	(1.027)	(0.997)	(0.603)
Percent age 51-57	**2.717*** **	**2.647*** **	**0.909**	**2.589*** **	**2.566*** **	**0.860**
	(0.689)	(0.700)	(0.622)	(0.710)	(0.715)	(0.631)

Table 2.3 (cont'd)

Percent age 58-64	**1.264**	**1.156**	**-0.142**	**1.320**	**1.244**	**-0.062**
	(1.271)	(1.122)	(0.827)	(1.258)	(1.101)	(0.829)
Weekly wage location quotient		**-0.036**	**-0.350****		**-0.068**	**-0.378****
		(0.269)	(0.146)		(0.276)	(0.158)
Percent employed residential		**1.080**	**0.158**		**1.021**	**0.134**
		(0.645)	(0.506)		(0.652)	(0.524)
Percent state and local		**0.052**	**0.021**		**0.063**	**0.037**
		(0.132)	(0.100)		(0.133)	(0.100)
Average firm size		**0.017**	**0.008**		**0.018**	**0.009**
		(0.013)	(0.008)		(0.013)	(0.008)
Private construction union density			**1.088*****			**1.086*****
			(0.125)			(0.131)
Constant	**8.343*****	**8.288*****	**10.744*****	**8.417*****	**8.285*****	
	(0.830)	(0.853)	(0.625)	(0.837)	(0.844)	
N	180	180	180	180	180	180
R^2	0.446	0.482	0.693	0.449	0.482	0.691

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.4: Earnings Results: 2019 Definition

Variables	Demographics	Industry Controls	Union Density	Demographics	Industry Controls	Union Density
License	**0.124*** **	**0.148** **	**0.060**			
	(0.047)	(0.078)	(0.052)			
Electrical licensing				**0.067**	**0.103**	**0.044**
				(0.098)	(0.083)	(0.051)
Plumbing licensing				**-0.067**	**-0.117**	**-0.033**
				(0.089)	(0.088)	(0.065)
Operating engineer licensing				**0.115** **	**0.136** **	**0.042**
				(0.045)	(0.073)	(0.045)
Percent female	**0.427**	**0.182**	**0.333**	**0.723**	**0.585**	**0.420**
	(1.569)	(1.892)	(1.252)	(1.545)	(1.887)	(1.295)
Percent black	**-0.642**	**-0.663**	**-0.059**	**-0.603**	**-0.556**	**-0.047**
	(0.402)	(0.508)	(0.405)	(0.391)	(0.522)	(0.490)
Percent Hispanic	**-0.215**	**-0.270**	**-0.114**	**-0.224**	**-0.343**	**-0.126**
	(0.182)	(0.279)	(0.241)	(0.183)	(0.329)	(0.261)
Percent high school	**0.420**	**0.290**	**-0.120**	**0.281**	**0.116**	**0.129**
	(0.577)	(0.794)	(0.564)	(0.601)	(0.924)	(0.636)
Percent college	**1.288**	**0.963**	**-0.089**	**1.278**	**0.864**	**-0.046**
	(1.043)	(1.580)	(0.990)	(1.031)	(1.585)	(1.036)
Percent age 23-29	**3.294**	**2.980**	**1.001**	**2.884**	**2.375**	**0.758**
	(2.588)	(2.700)	(2.170)	(2.619)	(2.700)	(2.201)
Percent age 30-36	**4.050** **	**4.190** **	**2.663**	**3.884***	**4.258***	**2.704**
	(1.971)	(1.969)	(1.547)	(2.168)	(2.118)	(1.723)
Percent age 37-43	**2.191***	**1.948**	**0.640**	**2.702** **	**2.666** **	**0.956**
	(1.270)	(1.251)	(1.086)	(1.336)	(1.297)	(1.083)
Percent age 44-50	**2.244**	**2.793**	**1.320**	**1.630**	**2.479**	**1.033**
	(2.236)	(2.740)	(2.014)	(2.470)	(2.855)	(2.085)
Percent age 51-57	**3.142*** **	**2.989** **	**0.864**	**3.535*** **	**3.542*** **	**1.127**
	(1.060)	(1.153)	(1.279)	(1.166)	(1.190)	(1.427)

Table 2.4 (cont'd)

Percent age 58-64	**2.153**	**2.013**	**2.171**	**1.605**	**1.486**	**2.033**
	(2.254)	(2.560)	(2.065)	(2.473)	(2.759)	(2.014)
Weekly wage location quotient		**-0.013**	**-0.375**		**-0.169**	**-0.419**
		(0.358)	(0.265)		(0.383)	(0.298)
Percent employed residential		**0.964**	**-1.038**		**1.133**	**-0.976**
		(0.839)	(0.757)		(0.963)	(0.827)
Percent state and local		**-0.232**	**-0.059**		**-0.237**	**-0.028**
		(0.330)	(0.233)		(0.338)	(0.241)
Average firm size		**0.003**	**-0.009**		**0.007**	**-0.007**
		(0.021)	(0.016)		(0.021)	(0.017)
Private construction union density			**1.352*** **			**1.350*** **
			(0.225)			(0.231)
Constant	**7.825*** **	**7.854*** **	**10.139*** **	**7.966*** **	**7.995*** **	**10.123*** **
	(1.296)	(1.308)	(1.183)	(1.427)	(1.409)	(1.289)
N	45	45	45	45	45	45
R^2	0.441	0.466	0.727	0.451	0.484	0.725

Estimated coefficients in bold. Robust standard errors in parentheses.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.5: Employment Results: 70-30 Rule

Variables	Demographics	Industry Controls	Union Density	Demographics	Industry Controls	Union Density
License-70	**0.013*** **	**0.012** **	**0.012** **			
	(0.004)	(0.005)	(0.005)			
Electrical licensing-70				**-0.004**	**-0.006**	**-0.006**
				(0.006)	(0.006)	(0.006)
Plumbing licensing-70				**0.003**	**0.003**	**0.003**
				(0.004)	(0.004)	(0.004)
Operating engineer licensing-70				**0.014*** **	**0.014** **	**0.014** **
				(0.004)	(0.005)	(0.005)
Percent female	**-0.173**	**-0.216**	**-0.214**	**-0.176**	**-0.212**	**-0.209**
	(0.151)	(0.182)	(0.178)	(0.149)	(0.181)	(0.176)
Percent black	**-0.072***	**-0.059***	**-0.061**	**-0.077** **	**-0.066***	**-0.070***
	(0.037)	(0.034)	(0.039)	(0.035)	(0.033)	(0.039)
Percent Hispanic	**-0.051*** **	**-0.053*** **	**-0.054** **	**-0.050*** **	**-0.052*** **	**-0.053** **
	(0.016)	(0.019)	(0.021)	(0.016)	(0.019)	(0.020)
Percent high school	**-0.057**	**-0.056**	**-0.055**	**-0.054**	**-0.053**	**-0.050**
	(0.057)	(0.060)	(0.061)	(0.057)	(0.060)	(0.061)
Percent college	**-0.325*** **	**-0.344*** **	**-0.340*** **	**-0.342*** **	**-0.361*** **	**-0.354*** **
	(0.097)	(0.013)	(0.115)	(0.101)	(0.106)	(0.116)
Percent age 23-29	**-0.230**	**-0.198**	**-0.194**	**-0.238**	**-0.197**	**-0.190**
	(1.167)	(0.175)	(0.171)	(0.167)	(0.176)	(0.172)
Percent age 30-36	**-0.191**	**-0.147**	**-0.140**	**-0.211**	**-0.162**	**-0.150**
	(0.145)	(0.162)	(0.167)	(0.145)	(0.162)	(0.165)
Percent age 37-43	**-0.267**	**-0.200**	**-0.193**	**-0.273**	**-0.205**	**-0.192**
	(0.191)	(0.201)	(0.209)	(0.190)	(0.201)	(0.208)
Percent age 44-50	**-0.132**	**-0.046**	**-0.040**	**-0.149**	**-0.058**	**-0.048**
	(0.139)	(0.177)	(0.184)	(0.138)	(0.178)	(0.184)
Percent age 51-57	**-0.303*** **	**-0.223***	**-0.220***	**-0.319*** **	**-0.240***	**-0.235***
	(0.103)	(0.126)	(0.130)	(0.104)	(0.128)	(0.131)

Table 2.5 (cont'd)

Percent age 58-64	**0.100**	**0.123**	**-0.127**	**0.104**	**0.133**	**0.140**
	(0.197)	(0.179)	(0.185)	(0.193)	(0.174)	(0.180)
Employment location quotient		**0.021***	**0.020***		**0.021***	**0.019**
		(0.012)	(0.111)		(0.012)	(0.011)
Percent employed residential		**-0.014**	**-0.008**		**-0.029**	**-0.020**
		(0.082)	(0.092)		(0.080)	(0.091)
Percent state and local		**-0.001**	**-0.001**		**-0.002**	**-0.001**
		(0.021)	(0.021)		(0.022)	(0.022)
Average firm size		**-0.001**	**-0.001**		**-0.001**	**-0.001**
		(0.001)	(0.001)		(0.001)	(0.001)
Private construction union density			**-0.004**			
			(0.028)			
Constant	**0.408*****	**0.349*****	**0.342*****	**0.419*****	**0.364****	**0.352****
	(0.118)	(0.141)	(0.148)	(0.116)	(0.142)	(0.148)
N	180	180	180	180	180	180
R^2	0.380	0.398	0.398	0.387	0.405	0.405

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.6: Employment Results: 70-30 Rule

Variables	Demographics	Industry Controls	Union Density	Demographics	Industry Controls	Union Density
License-70	**-0.085**	**0.030**	**-0.039**			
	(0.216)	(0.204)	(0.189)			
Electrical licensing-70				**-0.248**	**-0.044**	**-0.058**
				(0.256)	(0.230)	(0.209)
Plumbing licensing-70				**-0.177**	**-0.173**	**-0.205**
				(0.202)	(0.192)	(0.193)
Operating engineer licensing-70				**-0.140**	**-0.011**	**-0.096**
				(0.257)	(0.240)	(0.221)
Percent female	**0.483**	**3.463**	**2.569**	**-0.020**	**3.238**	**2.183**
	(4.945)	(4.761)	(4.370)	(4.770)	(4.168)	(4.171)
Percent black	**4.515*** **	**3.115***	**4.268****	**4.535*****	**3.067***	**4.346****
	(1.647)	(1.554)	(1.586)	(1.636)	(1.603)	(1.626)
Percent Hispanic	**2.100***	**1.975***	**2.280***	**2.161***	**1.988**	**2.319**
	(1.056)	(1.169)	(1.138)	(1.077)	(1.184)	(1.143)
Percent high school	**1.288**	**0.622**	**-0.105**	**1.616**	**0.736**	**-0.006**
	(2.276)	(2.143)	(2.130)	(2.350)	(2.149)	(2.135)
Percent college	**3.873**	**5.177**	**3.259**	**4.169**	**5.182**	**3.248**
	(4.140)	(3.797)	(4.245)	(4.270)	(3.910)	(4.318)
Percent age 23-29	**4.334**	**0.571**	**-1.279**	**4.925**	**0.850**	**-1.016**
	(5.963)	(4.916)	(5.570)	(5.863)	(4.921)	(5.606)
Percent age 30-36	**7.194**	**3.583**	**0.554**	**7.707**	**3.783**	**0.706**
	(5.090)	(4.988)	(4.448)	(4.690)	(4.736)	(4.325)
Percent age 37-43	**7.697***	**2.741**	**-0.510**	**8.192****	**2.968**	**-0.441**
	(3.974)	(4.078)	(3.004)	(3.786)	(3.944)	(2.880)
Percent age 44-50	**12.214****	**5.279**	**2.472**	**12.659*****	**5.519**	**2.663**
	(4.859)	(4.645)	(3.867)	(4.562)	(4.488)	(3.712)
Percent age 51-57	**4.457**	**-1.663**	**-3.158**	**4.475**	**-1.462**	**-2.965**
	(3.386)	(3.365)	(2.927)	(3.363)	(3.341)	(2.904)

Table 2.6 (cont'd)

Percent age 58-64	**-0.832**	**-2.237**	**-4.100**	**-0.279**	**-2.015**	**-3.965**
	(6.014)	(5.180)	(5.394)	(5.878)	(5.061)	(5.326)
Employment location quotient		**-1.713*** **	**-1.301** **		**-1.690*** **	**-1.241** **
		(0.517)	(0.618)		(0.518)	(0.610)
Percent employed residential		**1.246**	**-1.142**		**1.102**	**-1.388**
		(4.730)	(5.335)		(4.678)	(5.287)
Percent state and local		**-0.780**	**-0.906**		**-0.779**	**-0.899**
		(0.653)	(0.642)		(0.638)	(0.625)
Average firm size		**0.107**	**0.068**		**0.105**	**0.063**
		(0.078)	(0.081)		(0.078)	(0.082)
Private construction union density						**2.078** **
						(1.013)
Constant	**2.056**	**7.530***	**10.606** **	**1.955**	**7.519***	**10.698*** **
	(3.841)	(4.326)	(4.196)	(3.484)	(4.009)	(3.933)
N	180	180	180	180	180	180
R^2	0.382	0.506	0.529	0.390	0.508	0.534

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.7: Supplemental Analysis: Employment in Licensed Occupations as Share of Total Workforce

Variables	Demographics	Industry Controls	Union Density
License-70	**0.001****	**0.001***	**0.001***
	(0.000)	(0.000)	(0.000)
Percent female	**0.000**	**-0.020**	**-0.018**
	(0.012)	(0.015)	(0.015)
Percent black	**-0.007***	**-0.004**	**-0.006***
	(0.004)	(0.003)	(0.003)
Percent Hispanic	**-0.001**	**-0.003***	**-0.003****
	(0.001)	(0.001)	(0.001)
Percent high school	**-0.007**	**-0.008***	**-0.007**
	(0.006)	(0.005)	(0.005)
Percent college	**-0.022****	**-0.025*****	**-0.022****
	(0.009)	(0.008)	(0.008)
Percent age 23-29	**-0.022**	**-0.015**	**-0.012**
	(0.016)	(0.015)	(0.014)
Percent age 30-36	**-0.032*****	**-0.020**	**-0.015**
	(0.012)	(0.013)	(0.013)
Percent age 37-43	**-0.058*****	**-0.038****	**-0.033***
	(0.020)	(0.018)	(0.018)
Percent age 44-50	**-0.044*****	**-0.020**	**-0.015**
	(0.015)	(0.015)	(0.016)
Percent age 51-57	**-0.045*****	**-0.023***	**-0.016**
	(0.013)	(0.011)	(0.011)
Percent age 58-64	**-0.009**	**-0.018**	**0.003**
	(0.021)	(0.011)	(0.017)
Employment location quotient		**0.007*****	**0.006*****
		(0.001)	(0.001)
Percent employed residential		**-0.002**	**0.001**
		(0.007)	(0.008)

Table 2.7 (cont'd)

Percent state and local		**0.001**	**0.001**
		(0.002)	(0.002)
Average firm size		**0.000**	**0.000**
		(0.000)	(0.000)
Private construction union density			**-0.003**
			(0.002)
Constant	**0.052*** **	**0.033** **	**0.028** **
	(0.013)	(0.012)	(0.013)
N	180	180	180
R^2	0.356	0.544	0.551

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.8: Supplemental Analysis: Employment in Unlicensed Occupations as Share of Total Workforce

Variables	Demographics	Industry Controls	Union Density
License-70	**-0.001**	**-0.001**	**-0.001**
	(0.001)	(0.001)	(0.001)
Percent female	**0.028**	**-0.015**	**-0.013**
	(0.026)	(0.027)	(0.026)
Percent black	**-0.012***	**-0.008**	**-0.011**
	(0.007)	(0.007)	(0.008)
Percent Hispanic	**0.016*****	**0.011****	**0.011****
	(0.005)	(0.005)	(0.005)
Percent high school	**-0.001**	**-0.003**	**-0.001**
	(0.011)	(0.010)	(0.010)
Percent college	**0.036***	**0.036****	**0.041****
	(0.018)	(0.016)	(0.017)
Percent age 23-29	**0.002**	**0.002**	**0.007**
	(0.032)	(0.025)	(0.024)
Percent age 30-36	**-0.034**	**-0.019**	**-0.011**
	(0.033)	(0.027)	(0.026)
Percent age 37-43	**-0.089*****	**-0.064****	**-0.055****
	(0.030)	(0.028)	(0.027)
Percent age 44-50	**-0.093*****	**-0.064****	**-0.057****
	(0.032)	(0.028)	(0.023)
Percent age 51-57	**-0.034****	**0.005**	**0.009**
	(0.016)	(0.016)	(0.014)
Percent age 58-64	**-0.047**	**-0.033**	**-0.028**
	(0.032)	(0.031)	(0.031)
Employment location quotient		**0.010*****	**0.009*****
		(0.002)	(0.003)
Percent employed residential		**0.010**	**0.016**
		(0.019)	(0.020)

Table 2.8 (cont'd)

Percent state and local		**-0.001**	**0.000**
		(0.003)	(0.003)
Average firm size		**0.000**	**0.000**
		(0.000)	(0.000)
Private construction union density			**-0.005**
			(0.005)
Lagged employment			
Constant	**0.071*** **	**0.043***	**0.035***
	(0.024)	(0.021)	(0.021)
N	180	180	180
R^2	0.372	0.489	0.495

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.9: Supplemental Analysis: Interaction of Licensing and Union Density

Variables	Full Base Model	Below Mean	Above Mean	Interaction
License-70	**0.012****	**0.018****	**0.003**	**0.018****
	(0.005)	(0.007)	(0.004)	(0.008)
Private construction union density	**-0.004**			**0.006**
	(0.028)			(0.026)
License x union density				**-0.035**
				(0.031)
Percent female	**-0.214**	**-0.153**	**-0.358**	**-0.222**
	(0.178)	(0.235)	(0.226)	(0.179)
Percent black	**-0.061**	**-0.024**	**-0.107**	**-0.059**
	(0.039)	(0.044)	(0.137)	(0.038)
Percent Hispanic	**-0.054****	**-0.028**	**-0.072*****	**-0.055*****
	(0.021)	(0.020)	(0.024)	(0.020)
Percent high school	**-0.055**	**-0.053**	**-0.074**	**-0.050**
	(0.061)	(0.082)	(0.090)	(0.060)
Percent college	**-0.340*****	**-0.393****	**-0.037**	**-0.338*****
	(0.115)	(0.145)	(0.144)	(0.114)
Percent age 23-29	**-0.194**	**0.027**	**-0.302**	**-0.186**
	(0.171)	(0.190)	(0.257)	(0.172)
Percent age 30-36	**-0.140**	**0.118**	**-0.251**	**-0.116**
	(0.167)	(0.198)	(0.221)	(0.162)
Percent age 37-43	**-0.193**	**-0.073**	**-0.203**	**-0.185**
	(0.209)	(0.297)	(0.163)	(0.207)
Percent age 44-50	**-0.040**	**-0.033**	**0.275**	**-0.030**
	(0.184)	(0.224)	(0.204)	(0.182)
Percent age 51-57	**-0.220***	**-0.321**	**-0.094**	**-0.209**
	(0.130)	(0.190)	(0.151)	(0.129)

Table 2.9 (cont'd)

Percent age 58-64	**0.127**	**0.447**	**-0.071**	**0.141**
	(0.185)	(0.237)	(0.196)	(0.189)
Employment location quotient	**0.020***	**-0.002**	**0.066****	**0.020***
	(0.011)	(0.014)	(0.028)	(0.011)
Percent residential	**-0.008**	**0.345****	**-0.232****	**0.006**
	(0.092)	(0.125)	(0.103)	(0.098)
Percent state and local	**-0.001**	**-0.026**	**0.012**	**0.000**
	(0.021)	(0.025)	(0.022)	(0.021)
Average firm size	**-0.001**	**-0.002***	**-0.002**	**-0.001**
	(0.001)	(0.001)	(0.002)	(0.001)
Constant	**0.342****	**0.183**	**0.330****	**0.323****
	(0.148)	(0.215)	(0.142)	(0.146)
N	180	100	80	180
R^2	0.398	0.485	0.600	0.402

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 2.10: Earnings and Employment Results: 50-50 Rule

Variables	Demographics	Industry Controls	Union Density	Demographics	Industry Controls	Union Density
License-70	**0.125*** **	**0.129*** **	**0.080*** **	**0.018*** **	**0.019** **	**0.021*** **
	(0.035)	(0.035)	(0.025)	(0.004)	(0.004)	(0.005)
Percent female	**0.447**	**0.032**	**0.068**	**-0.180**	**-0.230**	**-0.191**
	(0.862)	(0.915)	(0.589)	(0.139)	(0.167)	(0.161)
Percent black	**-0.833*** **	**-0.768*** **	**-0.278**	**-0.071** **	**-0.070** **	**-0.090** **
	(0.291)	(0.256)	(0.226)	(0.032)	(0.031)	(0.036)
Percent Hispanic	**-0.245** **	**-0.193**	**-0.105**	**-0.060*** **	**-0.070*** **	**-0.074*** **
	(0.120)	(0.142)	(0.100)	(0.013)	(0.015)	(0.015)
Percent high school	**0.334**	**0.284**	**0.056**	**-0.075**	**-0.089* **	**-0.077**
	(0.403)	(0.414)	(0.325)	(0.051)	(0.052)	(0.054)
Percent college	**1.350** **	**1.220* **	**0.528**	**-0.394*** **	**-0.407*** **	**-0.395*** **
	(0.656)	(0.690)	(0.512)	(0.088)	(0.091)	(0.096)
Percent age 23-29	**0.871**	**0.360**	**-0.732**	**-0.265**	**-0.269**	**-0.230**
	(1.183)	(1.150)	(0.949)	(0.168)	(0.176)	(0.168)
Percent age 30-36	**2.983*** **	**2.725*** **	**0.920**	**-0.251**	**-0.223**	**-0.166**
	(0.857)	(0.898)	(0.628)	(0.155)	(0.171)	(0.171)
Percent age 37-43	**2.405** **	**2.200* **	**0.011**	**-0.287**	**-0.241**	**-0.174**
	(1.183)	(1.167)	(0.869)	(0.196)	(0.204)	(0.203)
Percent age 44-50	**1.543**	**1.205**	**-0.919**	**-0.193**	**-0.140**	**-0.082**
	(1.121)	(1.084)	(0.645)	(0.142)	(0.175)	(0.184)
Percent age 51-57	**2.565*** **	**2.431*** **	**0.866**	**-0.332*** **	**-0.268** **	**-0.245* **
	(0.759)	(0.780)	(0.681)	(0.110)	(0.128)	(0.132)
Percent age 58-64	**1.091**	**0.827**	**-0.242**	**0.062**	**0.092**	**0.117**
	(1.280)	(1.283)	(0.904)	(0.208)	(0.194)	(0.187)
Weekly wage location quotient		**0.149**	**-0.245**			
		(0.195)	(0.150)			
Employment location quotient					**0.015**	**0.007**
					(0.011)	(0.010)

Table 2.10 (cont'd)

Percent employed residential		**0.880**	**0.091**		**-0.010**	**0.031**
		(0.623)	(0.496)		(0.074)	(0.081)
Percent state and local		**-0.122**	**-0.067**		**-0.016**	**-0.018**
		(0.127)	(0.112)		(0.021)	(0.021)
Average firm size		**0.015**	**0.006**		**-0.001**	**-0.001**
		(0.013)	(0.007)		(0.001)	(0.001)
Private construction union density			**1.052*** **			**-0.037**
			(0.129)			(0.026)
Constant	**8.585*** **	**8.614*** **	**10.787*** **	**0.462*** **	**0.437*** **	**0.388** **
	(0.870)	(0.903)	(0.666)	(0.125)	(0.142)	(0.146)
N	180	180	180	180	180	180
R^2	0.475	0.498	0.689	0.437	0.450	0.461

Estimated coefficients in bold. Robust standard errors in parentheses. Year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

REFERENCES

REFERENCES

Blair, Peter Q., and Bobby W. Chung. 2018. "Job market signaling through occupational licensing." No. w24791. National Bureau of Economic Research.

Blair, Peter Q., and Bobby W. Chung. 2019. How much of barrier to entry is occupational licensing? *British Journal of Industrial Relations 57*(4): 919-943.

Bureau of Labor Statistics, U.S. Department of Labor. 2011-2016. *Quarterly Census of Employment and Wages.* Washington, D.C.

Bureau of Labor Statistics. 2019. Occupational outlook handbook. Accessed at https://www.bls.gov/ooh/.

Carollo, Nicholas A. 2020. "The impact of occupational licensing on earnings and employment: evidence from state-level policy changes." Working paper, available at: https://www.ncarollo.net.

Cathles, Alison, David E. Harrington, and Kathy Krynski. 2010. The gender gap in funeral directors: Burying women with ready-to-embalm laws? *British Journal of Industrial Relations 48*(4): 688-705.

Friedman, Milton. 1962. "Occupational Licensure." In *Capitalism and Freedom.* Chicago, IL: University of Chicago Press.

Friedman, Milton, and Simon Kuznets. 1945. "Income from independent professional practice." National Bureau of Economic Research.

Gittleman, Maury, and Morris M. Kleiner. 2016. Wage effects of unionization and occupational licensing coverage in the United States. *Industrial and Labor Relations Review 69*(1): 142-172.

Gittleman, Maury, Mark A. Klee, and Morris M. Kleiner. 2018. Analyzing the labor market outcomes of occupational licensing. *Industrial Relations: A Journal of Economy and Society 57*(1): 57-100.

Hall, Jonathan, Jason Hicks, Morris M. Kleiner, and Rob Solomon. 2018. Occupational licensing of Uber drivers. *Institute for Theoretical Economics.*

Hirsch, Barry T., and David A. Macpherson. 2003. Union membership and coverage database from the Current Population Survey: Note. *Industrial and Labor Relations Review* 56(2): 349-354 (updated annually at unionstats.com).

Johnson, Janna E., and Morris M. Kleiner. 2020. Is occupational licensing a barrier to interstate migration? *American Economic Journal: Economic Policy 12*(3): 347-73.

Kleiner, Morris M. 2000. Occupational licensing. *Journal of Economic Perspectives 14*(4): 189-202.

Kleiner, Morris M. 2006. *Licensing occupations: Ensuring quality or restricting competition?* Kalamazoo, MI: WE Upjohn Institute.

Kleiner, Morris M., and Alan B. Krueger. 2010. The prevalence and effects of occupational licensing. *British Journal of Industrial Relations 48*(4): 676-687.

Kleiner, Morris M., and Alan B. Krueger. 2013. Analyzing the extent and influence of occupational licensing on the labor market. *Journal of Labor Economics 31*(S1): S173-S202.

Kleiner, Morris M., and Kyoung Won Park. 2010. "Battles among licensed occupations: Analyzing government regulations on labor market outcomes for dentists and hygienists." No. w16560. National Bureau of Economic Research.

Kleiner, Morris M., and Evan J. Soltas. 2019. A welfare analysis of occupational licensing in U.S. states. No. w26383. National Bureau of Economic Research.

Kleiner, Morris M., and Evgeny Vorotnikov. 2017. Analyzing occupational licensing among the states. *Journal of Regulatory Economics 52*(2): 132-158.

Koumenta, Maria, Amy Humphris, Morris Kleiner, and Mario Pagliero. 2014. "Occupational regulation in the EU and UK: Prevalence and labour market impacts." Final report, department for business, innovation and skills, school of business and management, Queen Mary University of London, London.

Law, Marc T., and Sukkoo Kim. 2005. Specialization and regulation: The rise of professionals and the emergence of occupational licensing regulation. *The Journal of Economic History 65*(3): 723-756.

Law, Marc T., and Mindy S. Marks. 2009. Effects of occupational licensing laws on minorities: Evidence from the progressive era. *The Journal of Law and Economics 52*(2): 351-366.

Law, Marc T., and Mindy S. Marks. 2017. The labor-market effects of occupational licensing laws in nursing. *Industrial Relations: A Journey of Economy and Society 56*(4): 640-661.

Leland, Hayne E. 1979. Quacks, lemons, and licensing: A theory of minimum quality standards. *Journal of Political Economy 87*(6):*1328*-1346.

Rottenberg, Simon. 1980. *Occupational licensure and regulation*. American Enterprise Institute.

Ruggles, Steven, Sarah Flood, Ronald Goeken, Josiah Grover, Erin Meyer, Jose Pacas and Matthew Sobek. 2020. *IPUMS USA: Version 10.0* [dataset]. Minneapolis, MN: IPUMS, 2020. https://doi.org/10.18128/D010.V10.0

Shapiro, Carl. 1986. Investment, moral hazard, and occupational licensing. *The Review of Economic Studies* 53(5): 843-862.

Siniavskaia, Natalia. 2018. Immigrant workers in the construction labor force. *National Association of Home Builders (NAHB).*

Thornton, Robert J., and Edward J. Timmons. 2013. Licensing one of the world's oldest professions: Massage. *The Journal of Law and Economics 56*(2): 371-388.

Xia, Xing. 2021. Barrier to entry or signal of quality? The effects of occupational licensing on minority dental assistants." *Labour Economics 71*: 102027.

Yelowitz, Aaron, and Samuel J. Ingram. "How does occupational licensing affect entry into the medical field? An examination of emergency medical technicians." *Southern Economic Journal* (2021).

Zhang, Tingting. 2019. Effects of occupational licensing and unions on labour market earnings in Canada. *British Journal of Industrial Relations 57*(4): 791-817.

CHAPTER 3: THE EFFECT OF PREVAILNIG WAGE LAWS ON INFORMAL CONSTRUCTION EMPLOYMENT

Abstract

Informal employment, defined as the illegal misclassification of employees as independent contractors or employment of workers using cash-only payments, has long been rampant in the American construction industry. These actions rob workers of legally earned benefits, defund social programs, and undermine the competitiveness of law-abiding contractors. While enforcing labor laws has proved difficult, one way a state may be able to strengthen enforcement—and limit informality—is a prevailing wage law. Under penalty of law, these regulations require employers to submit weekly certified payrolls to government agencies on public works projects, which increases governmental oversight. This study uses state-level data from 2010-2019 to examine the relationship between prevailing wage laws and informal construction employment. State prevailing wage laws, even those of weak and average strength, are associated with significant reductions in informality.

Introduction

Informal employment, defined as (a) the illegal misclassification of employees as independent contractors using 1099-MISC forms or (b) off-the-books employment of workers using cash-only payments, is rampant in the American construction industry (e.g., Ormiston, Belman, Brockman, & Hinkel 2020; Ormiston, Belman, & Erlich 2020; Juravich, Ormiston, & Belman 2021).[59] This is partly due to the fact that as an industry, construction tends to operate as

[59] Informal firms are those who choose to evade established government regulations and are not registered with tax authorities; the totality of these informal firms comprises the informal sector (Wells 2007; Ulyssea 2018). One set of

a gig economy (Erlich 2020). Misclassification and off-the-books employment are linked to the nature of the construction industry because like other gig work, construction features a short-term employment structure. Contractors hire workers for a specific project, but only for the duration of that project; once the project is complete, workers move to other projects and often work for other employers. Labor supply and demand in construction then operate on a *project-by-project* basis, with each construction project functioning as a separate "gig" for a highly mobile construction labor force.[60] While this looser employment relationship can offer greater fluidity and flexibility for workers who are not beholden to a particular employer, it also has helped incentivize a business strategy built on purposeful exploitation of construction workers (Erlich 2020).

Alongside its short-term and decentralized employment structure, informality has become institutionalized in the construction industry because of minimal enforcement of regulations. Enforcement is difficult because contractors (and workers) have gone to great lengths to hide this behavior from regulators and tax agencies as resources available to these agencies have declined. A common practice is that contractors hire and pay workers off-the-books to avoid paying taxes. Additionally, contractors have increasingly utilized labor brokers to hide illegal behavior. When contractors hire labor brokers (who in turn provide workers for a construction project), it looks as though subcontractors are being hired; however, as the hiring contractors oversee the workers, labor brokers operate off-the-books to illegally reduce labor costs. Recent work suggests these illegal practices are profitable in residential construction, a highly competitive market with small,

regulations firms may elect to evade are labor market regulations (e.g., labor legislation, health and safety rules), which define the terms and conditions of employment (Wells 2007).

[60] The idea of construction operating as a gig economy is supported by the prevalence of independent contracting in the industry" while independent contractors only comprise 7% of the total U.S. workforce, about 20% of all U.S. independent contractors work in construction (Erlich 2020; U.S. BLS 2018).

labor-intensive firms all competing to complete limited-duration construction projects at the lowest cost possible (Ormiston, Belman, Brockman, & Hinkel 2020). As these illegal practices expand, law-abiding construction firms face two difficult choices. If they continue following the rules, they risk being shut out of the market and losing construction project bids as other firms reduce labor costs. If they choose to follow suit and operate informally, while enforcement is difficult, they may be caught by regulators.

Prior work suggests reduced enforcement increases informal employment (e.g., Friedman et al. 2000; Dabla-Norris, Gradstein, & Inchauste 2008; Almeida & Carneiro 2006). But in construction, one way a state may be able to strengthen its enforcement system—and limit the degree of informality—is a prevailing wage law. On the federal level, the Davis-Bacon Act regulates federal construction projects by determining mandated “prevailing wage” rates by construction trade, area, and project type by surveying private construction employers where projects are scheduled. In addition, under Davis-Bacon, employers are required to follow numerous rules, under penalty of law, that involve increased federal oversight of public projects. This includes the requirement that construction contractors file a weekly certified payroll, which must include lists of employees’ names, Social Security numbers, work classification, daily hours worked, and other information (Wage and Hour Division). On the state level, 27 states currently feature their own prevailing wage laws, known as “little Davis-Bacon" laws, that apply regulations like Davis-Bacon on state-funded construction projects.

This study uses state-level, cross-sectional variation during the period 2010-2019 to examine the relationship between state prevailing wage laws and informal construction employment. Results suggest that state prevailing wage laws are associated with significant reductions in informality, with additional results suggesting that this association holds for prevailing wage laws of average and even weak strength. These results are in line with this paper's theoretical argument that via certified payroll requirements, prevailing wage laws allow states to strengthen their enforcement systems and curb the degree of informality. This benefits workers, law-abiding firms, and taxpayers alike.

This paper is structured as follows. First, the next section summarizes relevant informal employment literature, both within construction and outside of construction. Second, this study integrates recent theoretical work centered on firm informal employment decisions to develop a framework that yields hypotheses about the effect of prevailing wage laws on informal construction employment. Third, this study's methodology of measuring informal employment is presented, along with detailed descriptions of estimation strategies and statistical models. Lastly, results are presented and discussed.

Literature Review

Studying informality is important because it exerts undeniable influences on markets, workers, and governments. It distorts markets primarily because it creates an uneven playing field between firms (Ulyssea 2018; Carre & Wilson 2004). In the presence of informality, high-road firms who adhere to labor and employment law have higher costs than low-road firms who are successfully able to evade labor and employment law with illegal labor practices. This leaves high-road firms at a substantial competitive disadvantage: if firms can reduce their labor costs so significantly via these practices and avoid detection, other firms risk being shut out of the market if they fail to follow suit (Ormiston, Belman, Brockman, & Hinkel 2020; Ulyssea 2018).

Informality also affects workers and governments. Misclassified and off-the-books workers lose access to unemployment insurance, workers' compensation, minimum wages, overtime payments, anti-discrimination defenses, and the right to unionize (Harris & Krueger

2015; Erlich 2020). Misclassification and off-the-books employment also impact federal and state tax revenues. Ormiston, Belman, and Erlich (2020) estimate that due to payroll fraud in construction, under conservative income assumptions, federal and state governments face hundreds of millions of dollars in budget shortfalls annually. Employment relationships in informal labor markets are largely controlled by employers, as the balance of power heavily favors employers who can evade labor and employment law with practices that take place in society's shadows. As a result, workers endure revoked labor law protections and governments face shortfalls in tax revenues.

Different theories have attempted to explain the persistence of informal employment in the United States (Weil 2014; Friedman, Johnson, Kaufmann, & Zoido-Lobaton 2000). First, Friedman et al. (2000) suggest informal employment is centered on taxes, where firms choose to operate informally because they wish to keep more of their profits for themselves. Friedman et al. (2000) argue that weak labor market institutions (e.g., limited enforcement of regulations) fail to curb informal employment and allow it to continue, incentivizing profit-maximizing firms to operate informally and avoid taxes.

Second, Weil (2014) offers a more nuanced perspective, suggesting that economic restructuring and workplace "fissuring" in the last few decades has demanded small, flexible units of production that created incentives for more informal employment, such as subcontracting. The author details the increased prominence of subcontracting in recent decades, as firms have shifted their focus toward their core competencies and away from those activities not central to their profitability. Weil (2014) notes that "the large corporation of days of yore came with distinctive borders around its perimeter, with most employment located inside firm walls. The large business of today looks more like a small solar system, with a lead firm at its center and smaller workplaces orbiting around it" (Weil 2014, p. 43). In construction, this shift away from core competencies is related to the issue of liability. If a general contractor hires a subcontractor and does not employ their own workers, the general contractor is not liable for the practices of the subcontractor, which benefits and protects the general contractor. Additionally, in some trades, subcontractors will hire their own subcontractors, creating a multi-layered system of subcontracting that is better able to hide illegal behavior. This happens because the tasks that are most often misclassified or employed off-the-books are simpler tasks. Lesser-skilled trades in construction are likely to have larger labor supplies because of fewer barriers to entry; thus, lesser-skilled trades are more likely to feature subcontractors who accept cash-only payments.

Prior empirical work has supported the notion that labor market regulations and institutions affect the degree of informal employment (e.g., Johnson, Kaufmann, & Zoido-Lobaton 1998; Friedman et al. 2000; Dabla-Norris, Gradstein, & Inchauste 2008; Almeida & Carneiro 2006; Enste 2010). In considering the impact of regulation, it is important to distinguish between regulatory *quantity* and *quality*. Regulatory quantity refers to the number of regulations in place; regulatory quality refers to how well-developed, efficient, and enforced the regulations are (Enste 2010; Dabla-Norris et al. 2008). Both factors can exert independent (and often conflicting) effects on informal employment. For example, if more regulations are present, but they are low quality and not properly enforced, firms may decide that their optimal decision is to operate informally, evade those regulations, and thereby increase their profits (Johnson et al. 1997; Enste 2010). Johnson et al. (1998) find support for the prediction: countries with more labor market regulations have significantly larger informal sectors, as measured by total GDP.

Friedman et al. (2000) obtain a similar finding. Using 1990s data on tax rates, the legal environment, and the size of informal economies in 69 countries (including developing and transition countries), the authors find that higher regulation (i.e., higher regulatory *quantity*) is associated with larger informal sectors. This finding indicates support for the authors' theoretical explanation that more regulations, with insufficient enforcement, entice profit-maximizing firms to operate informally (Friedman et al. 2000). Lastly, Enste (2010) analyzes the impact of regulation on the degree of informality in 25 OECD countries, including the United States. Enste (2010) finds that in these 25 countries, higher regulatory quantity significantly increases the size of a country's informal sector.

Regarding regulatory quality, Dabla-Norris et al. (2008) used the World Business Environment Survey (WBES), a survey of over 4,000 firms in 41 countries, most of which being developed and transition countries. Their findings reinforce their theoretical prediction, based in part on Friedman et al. (2000): regulatory quality is a primary determinant of informality in both developed and transition countries. Specifically, Dabla-Norris et al. (2008) find that more efficient and developed legal institutions in a given country significantly reduce its incidence of informality, even after controlling for other country-specific institutions and firm-level variables.[61]

Almeida and Carneiro (2006) analyze the impact of regulatory enforcement on informal employment in Brazil. Brazil simultaneously has (a) one of the most regulated labor markets in the world and (b) a relatively decentralized and ineffective enforcement structure.[62] This combination of high regulation and low enforcement provided an optimal research setting to examine this relationship (Almeida & Carneiro 2006). Using firm-level data of 1,641 Brazilian

[61] In measuring legal quality, Dabla-Norris et al. (2008) use a "rule of law" variable, an index measured from 0 to 10 that proxies for the quality of legal institutions and level of legal enforcement in a country. This index includes measures for (a) perceptions of both violent and nonviolent crime, (b) the effectiveness and predictability of a country's judiciary, and (c)the enforceability of contracts, with higher values denoting higher quality of the rule of law.

[62] In the early 2000s, approximately 40 percent of total Brazilian employment was informal (Almeida & Carneiro 2006).

manufacturing firms in 306 cities from the 2002 World Bank Investment Climate Survey, combined with 2002 administrative data on enforcement of regulation at the city level, the authors find a significant negative relationship: firms employ significantly smaller amounts of informal employment in areas where regulatory enforcement is stricter (Almeida & Carneiro 2006). Further, the authors find evidence of a negative causal effect of enforcement on informal employment, by instrumenting for enforcement with (1) measures of access to labor inspectors to firms and (2) measures of general law enforcement in a firm's given location (Almeida & Carneiro 2006). Thus, Dabla-Norris et al. (2008) and Almeida and Carneiro (2006) find that higher regulatory quality lowers informal employment.

In sum, prior literature provides theoretical justification and empirical support for the idea that labor market regulations and institutions, both within and outside the United States, significantly affect informal employment. Further, previous research has distinguished between regulatory quantity and quality and found that each factor exerts independent effects on the degree of informality. While higher regulatory quantity leads to increases informality, higher regulatory quality, with stronger enforcement of labor and employment law, lowers informality.

It is important to distinguish between these findings related to regulatory quantity and quality and how prevailing wage laws act. While prevailing wage laws are an additional form of labor market regulation in construction (i.e., an increase in quantity), inherent enforcement mechanisms in prevailing wage laws imply that these laws also act as an increase in regulatory quality. On public projects subject to prevailing wage requirements, contractors are required to follow numerous rules that involve increased governmental oversight. This includes the requirement that construction contractors file a weekly certified payroll with federal and state governments, which must include lists of employees' names, Social Security numbers, work classification, daily hours worked, and other information (Wage and Hour Division). While prior research on regulatory *quantity* suggests that prevailing wage laws, an additional form of regulation in construction, could increase informal employment, research on regulatory *quality* suggests that increases in efficiency and enforcement of regulations (e.g., certified payroll requirements on prevailing wage projects) significantly lower informal employment. This makes assessing the impact of prevailing wage laws on informal employment more complex.

Theory and Hypothesis Development: Prevailing Wage Laws

Duncan and Ormiston (2018) summarize the empirical literature on prevailing wage laws; research has examined the effect of this legislation on costs, training, safety and other factors. But this paper adds to this literature by addressing and examining an unaddressed question: the effect of prevailing wage laws on informal construction employment. The incentive to reduce labor costs in construction does not only result from the short-term and decentralized nature of the employment relationship in the industry. It also stems from the structure of the bidding process for public works projects. Historically, public works construction was often plagued by corruption between contractors and government officials in the form of kickbacks and other shameful practices. Federal, state, and local legislatures have countered this by requiring public bodies, in most instances, to accept the lowest qualified bid.[63,64] This low bid requirement contrasts with the private sector, where owners are allowed to consider other factors

[63] There are exceptions to this low-bid requirement; a notable example is a responsible bidder ordinance (RBO). RBOs outline sets of criteria a contractor must meet in order to be awarded contracts for public projects (e.g., past compliance with environmental and safety regulations, prevailing wage compliance). Rather than awarding projects to the lowest bidder, public bodies under RBOs award projects to the lowest *responsible* bidder. These need not be the lowest overall bids, and in many cases, they are not (Indiana-Illinois-Iowa Foundation for Fair Contracting 2018).

[64] In order for a bid to be considered "qualified," contractors must obtain a bond for the project. This is a modest hurdle in most instances (Hinkel & Belman 2021).

(e.g., contractors' reputation and experience) along with the bid price in determining the winning bid.

Thus, on public projects, contractors are incentivized to find ways to reduce their bid price; if they do not, they risk losing the bid. Although better engineering techniques and a better-trained workforce lower costs, the most straightforward approach is to pay labor less. In the presence of low bid requirements, prevailing wage laws level the playing field in terms of worker compensation and instead compel contractors to compete on engineering, project management, and the skills of their labor force, rather than labor costs (Hinkel & Belman 2021). As prevailing wage laws make labor costs a smaller source of competition between firms, the question arises whether these laws also impact informality. To address this question, this study integrates and builds upon recent theoretical work by Cuff, Mongrain, and Roberts (2020) and Ulyssea (2018) on evasion of labor market regulations by firms.

Consider a model where construction firms are heterogeneous, and θ, defined as the productivity of each individual firm, is a function of k (capital) and λ (labor skill). Assume each construction firm, on a given project, produces a homogenous good. Assume product and labor markets are both competitive, and formal and informal workers each supply one unit of labor, ℓ, at an identical opportunity cost (i.e., wage) of ω (Ulyssea 2018; Cuff et al. 2020). Define the output of a given construction project, y, as a function of θ and ℓ. Project output is then given by $y(\theta, \ell) = \theta q(\ell)$, where the function q is assumed to be increasing, concave, and twice continuously differentiable (Ulyssea 2018).

Assume that firms face two distinct choices for each project: (1) to hire all informal workers, or (2) to hire all formal workers. If firms choose informality, assume they pay no payroll or revenue taxes (Ulyssea 2018). While informal firms can avoid these costs, they face a probability of detection, $\rho > 0$, by government regulators. If detected, informal firms face a punishment imposed by regulators, c. Possible punishments for violations of prevailing wage standards include make-whole remedies, fines, and/or incarceration (Hinkel & Belman 2021). Ulyssea (2018) defines the expected cost of detection in the form of a general labor distortion but never explicitly defines this distortion.

However, this can be conceptualized in the form of an expected value, as a function of the probability of detection (ρ), the punishment (c), and Ω, defined as the sum of the wages of all workers on a project. As such, define the expected cost of detection as $E(D)_p = \rho c\Omega$, indexed by construction project (p).[65] That is, the expected cost of detection is the probability of being caught (ρ) times the overall penalty ($c\Omega$). In this equation, c is an unknown multiple of the sum of wages, based on the punishment levied by regulators if the contractor is caught. If a contractor is required to simply pay back wages as a penalty, then $c = 1$ (i.e., the penalty is equal to the sum of wages). If a contractor is required to pay back wages plus an additional proportional penalty, then $c > 1$. Lastly, there are instances where contractors are able to pay less than full back wages as a penalty (Juravich et al. 2021); in these cases, $c < 1$. Define Π^I as the profit function of informal construction firms, indexed by construction project (p). It is the following (based on Ulyssea 2018):

(1) $$\Pi^I_p(\theta) = \max\{\theta q(\ell) - \rho c\Omega\},$$

where the price of the final good is normalized to 1.

[65] I index this expected detection cost (and each profit function) by construction project because, as mentioned in the introduction, labor supply and demand in construction tends to function on a project-by-project basis (Erlich 2020).

Conversely, if a construction firm chooses to operate formally on a project, it elects to comply with labor market regulations and pay all relevant taxes. As such, it is assumed that formal firms must pay a constant payroll tax on all formal workers, τ_w, and a revenue tax, τ_y. However, given that these firms are abiding by labor and employment law, they face no probability of evasion detection (i.e., $\rho = 0$). It follows then that $E(D)_p = 0$ for formal firms. Following Ulyssea (2018), define Π^F as the profit function of formal construction firms, indexed by construction project (p):

$$\Pi^F{}_p(\theta) = \max\{(1 - \tau_y)\theta q(\ell) - (1 + \tau_w)\Omega\}. \quad (2)$$

Faced with these competing profit functions, construction firms choose whether to evade labor market regulations and operate informally. Assuming that construction firms are profit maximizers, define the evasion decision as follows, based on Cuff et al. (2020):

Evasion Decision: A construction firm with parameters (θ, ω) decides to evade labor market regulations and operate informally if and only if $\Pi^I{}_p(\theta) \geq \Pi^F{}_p(\theta)$.

Next, consider how prevailing wage laws factor into this decision. As previously mentioned, in addition to helping define the terms and conditions of employment in construction, prevailing wage laws feature built-in enforcement mechanisms, such as the requirement to file weekly certified payrolls to government regulators. Thus, on projects subject to prevailing wage requirements, firms who choose to operate informally face increased probabilities of detection, due to higher enforcement. Let PW denote a prevailing wage requirement; let PW = 1 denote a project subject to these requirements and let PW = 0 denote a project not subject to these requirements. It follows that for firms choosing to operate informally, indexed by project:

$$(3) \qquad \rho_p \mid PW = 1 > \rho_p \mid PW = 0,$$

As ρ increases, the expected cost of detection, E(D), also naturally increases. Assuming that $\delta\theta q(\ell) / \delta\rho = 0$ (i.e., that production decisions do not depend on detection probability), all else equal, an increase in ρ lowers the profits of a firm choosing to operate informally, directly following from equation (1):

$$(4) \qquad \delta\Pi^I{}_p(\theta, \omega) / \delta\rho = -c\Omega < 0.$$[66]

Next, note that the profits of a firm choosing to operate formally remain unaffected; for formal firms, the probability of evasion detection is 0. Therefore, for formal firms:

$$(5) \qquad \delta\Pi^F{}_p(\theta, \omega) / \delta\rho = 0.$$

Since prevailing wage requirements increase the probability of evasion detection, they lower profits of firms choosing to operate informally. Conversely, since formal firms face a probability of evasion detection of 0, their profits are unchanged by prevailing wage requirements. Therefore, on projects subject to prevailing wage requirements, construction firms

[66] It should be noted that this assumes production decisions do not depend on the probability of evasion detection.

should be *less likely* to make the decision to evade labor market regulations; informal employment should be lower on prevailing wage projects.

While the Davis-Bacon Act affects construction employers working on federally-funded projects throughout the United States, 27 states have similar laws applying Davis-Bacon-like regulations on state-funded projects. Since construction firms have disincentive to operate informally on state-funded projects, informal employment in states with prevailing wage laws should be lower, due to the increased evasion detection risk.

Hypothesis 1: *States with prevailing wage laws will have lower rates of informal construction employment than states without prevailing wage laws.*

Simply comparing states with prevailing wage laws and states without them does not tell the full story about the effect of prevailing wage laws on informality. As prior literature has done outside of construction (Dabla-Norris et al. 2008; Almeida & Carneiro 2006), it is important to consider the *quality* of such regulations, as state prevailing wage laws differ in strength and reach. Thieblot (1995) developed a methodology for measuring the strength of a state's prevailing wage law, which he then applied in a follow-up study (Thieblot 1999). This methodology is a point system that scores each state's prevailing wage law across the following scales: (1) the minimum contract threshold for coverage under the state law, (2) whether both state and local contracts are covered, (3) the enforced wage rate, (4) the breadth of work and occupations covered by the law and (5) "other factors," including state and local resident preferences, administrative and/or compliance requirements, and penalties for violating the law. The higher the score, the stronger the law (Thieblot 1995, 1999).

This study employs this methodology to measure the strength of each state's prevailing wage law over time. Prevailing wage law strength should also be related to informal employment. For example, lower coverage thresholds lead to more projects being subject to prevailing wage laws' enforcement mechanisms; this should be associated with lower informality. Similarly, if penalties for violation are higher, all else equal, informal employment is disincentivized. In sum, states with stronger laws (e.g., lower coverage thresholds, wider variety of work and occupations covered, higher penalties for violation) will likely feature lower levels of informal employment than states with weak laws.

Hypothesis 2: *States with stronger prevailing wage laws will have lower rates of informal construction employment than states with weaker laws.*

A related question is whether there are certain aspects of prevailing wage laws (e.g., thresholds, coverage, enforcement) that affect informality more than others. Coverage thresholds determine the minimum state-funded construction project costs subject to state prevailing wage requirements. These coverage thresholds vary from state to state. For example, California has a coverage threshold of $2,000, meaning that any state-funded construction project costing at least $2,000 must be covered by prevailing wages. In contrast, Nevada has a coverage threshold of $100,000, meaning that only state-funded projects costing $100,000 or more can be subject to prevailing wages (Wage and Hour Division). As a result, California's prevailing wage law covers proportionately more construction projects than Nevada's. All else equal, lower coverage thresholds should be associated with reduced informality, as prevailing wage protections and enforcement mechanisms cover more construction projects. The same logic follows for breadth of coverage: wider varieties of projects covered by prevailing wage requirements should be associated with lower informality. This is because certified payroll requirements are applied to a wider variety of construction projects, making informality less profitable for firms on these projects.

Hypothesis 3: *States with lower coverage thresholds will have lower rates of informal construction employment than states with higher thresholds.*

Hypothesis 4: *States with more types of projects covered will have lower rates of informal construction employment than states with fewer project types covered.*

Methods

Measuring Informal Employment

To measure informal employment, the current study compares annual household employment data from the American Community Survey (ACS) against payroll records from the Bureau of Economic Analysis (BEA) over the 2010-2019 period, a period of growth in U.S. construction.[67,68] . The unit of analysis is at the state level. While household surveys like the ACS capture all forms of employment (e.g., wage-and-salary employment, self-employment), firm payroll records like the BEA only capture legal wage-and-salary (W-2) employment. Measuring informality presents a significant hurdle: while household surveys commonly used by

[67] The ACS provides a far superior sample size to that of the Current Population Survey (CPS), the other oft-used household survey in labor economics. The ACS is the largest annual household survey in the United States, surveying over 3.5 million households annually, while the CPS only surveys 60,000 households each month. As Ormiston et al. (2020) observe, while this sample size difference is likely not problematic for nationwide estimates of construction informal employment, it can become problematic for state-by-state estimates.
[68] The ACS data was obtained from the Integrated Public Use Microdata Series (IPUMS) provided by the University of Minnesota (Ruggles et al. 2020).

economists (e.g., the American Community Survey, the Current Population Survey) ask many relevant questions about employment and income, they do *not* ask questions about the legality of each worker's employment situation. It is then impossible for economists to directly measure informality with existing data.

As a result, researchers have turned to indirect approaches to measuring informal employment, which involve examining discrepancies between two different measures of economic activity. The current study follows this literature in employing an indirect approach. Bohn and Owens (2012) proposed this strategy to analyze informality across multiple industries by comparing official firm records from the Quarterly Census of Employment and Wages (QCEW) against three worker surveys of labor force participation.[69] In so doing, they use the following equation to estimate informal employment, indexed by state (s) and year (t):

$$(6) \qquad \textit{Informal Employment}_{st} = \textit{Total Workforce}_{st} - \textit{Legal Employment}_{st},$$

where the total workforce is total wage-and-salary employment in state s and year t, as estimated by household surveys, and legal employment is the total number of wage-and-salary jobs in state s and year t, as estimated by firm payroll records.[70]

Given Bohn and Owens' (2012) explicit focus on comparing wage-and-salary employment across these two sources, equation (6) ignores all self-employed workers. As a

[69] It would be possible to use QCEW data for this purpose as Bohn and Owens (2012) did, as the QCEW publishes annual averages in addition to quarterly firm data. However, BEA data is more optimal in this case. Whereas QCEW data is collected by state unemployment insurance agencies and captures between 95 and 97 percent of all legal wage and salary civilian employment in the U.S., the BEA augments this by reviewing additional administrative records, including workers who are not covered by state unemployment insurance programs. The BEA also corrects for QCEW reporting errors. This adjusts estimated wage and salary employment upward, making the BEA a more complete estimate of official legal employment than the QCEW.

[70] As a robustness check, this study's models were estimated using equation (6) (i.e., without the informal self-employment adjustment). Results were robust to the omission of informal self-employment.

result, the above definition of informality is incomplete; any measure of informal employment should also include self-employed workers who are either misclassified or working off-the-books.[71] This is particularly true in construction, where 20% of all U.S. independent contractors work (Erlich 2020). While legitimate independent contractors and other legally self-employed workers certainly exist in construction and should not be counted as informal, excluding the entirety of the self-employed workforce drastically understates the degree of informality in the industry. In light of this, the current study's informal construction employment equation is presented below, indexed by state and year:

$$\text{(7)} \quad \textit{Informal Employment}_{st} = (1.095)(1.02)\textit{Total Workforce}_{st} - \textit{Legal Employment}_{st} + (1.095)(1.02)\textit{Informal Self-Employment}_{st},$$

where informal self-employment is equal to 0.44*self-employment, with self-employment taken from the ACS.

Moving from equation (6) to equation (7) involves a number of steps. Three adjustments are made to equation (6) to arrive at equation (7): (a) the addition of a measure of informal self-employment, using self-employed income underreporting as a proxy; (b) accounting for employment underreporting in household surveys compared to payroll records; and (c) accounting for workers who are in construction as a second job.

The first issue is the need to find a way to distinguish between two groups at the state level: (a) legitimate self-employed workers and (b) self-employed workers who are employed informally. Ormiston, Belman, and Erlich (2020) suggest that income underreporting rates of

[71] Bohn and Owens (2012) implicitly acknowledge the incompleteness of this definition; the authors classify their estimates of informality as lower bounds.

self-employed workers on IRS tax forms be used to proxy for the proportion of self-employment that is informal. While these are imperfect measures, these rates can offer a useful proxy for illegal activity, as a primary goal of firms who operate informally is to conceal payments to workers and evade taxes.[72] The relationship between income underreporting and informal employment arrangements is supported by work from the IRS (2016).[73] Proxying for informal self-employment with self-employment income underreporting is attractive because it relaxes the assumption that every tax filer is operating entirely in accordance with tax law. This approach incorporates workers who may operate *legally* in some work (reporting it to the IRS) but do additional work on the side that goes unreported (Ormiston et al. 2020).[74]

This paper follows Ormiston et al. (2020) and proxies for informal self-employment with income underreporting rates of self-employed workers. Ormiston et al. (2020) use a variety of income underreporting rates as proxies for informal self-employment; these rates range from 23.3% to 64% (Alm & Erard 2016; Johns & Slemrod 2010; Hurst, Li, & Pugsley 2014; IRS 2016). [75,76] While it is not construction-specific, one of these rates (44%) derives from the fact

[72] Ormiston et al. (2020), in advocating for this method of distinguishing between legitimate self-employed workers and informal self-employed workers, note that "to be clear, the decision to report—or not report—income on tax returns is the responsibility of the worker. But employers who rely on cash-only payments...effectively open the door for income underreporting" (Ormiston et al. 2020, p. 28-29).

[73] While only 1% of wage-and-salary income across all industries was underreported on IRS tax forms, 64% of nonfarm proprietor income was underreported to the IRS. That is, for those in legal jobs that feature more detailed W-2 documentation, the underreporting rate was near zero; meanwhile, for non-proprietor income, subject to far less information reporting than W-2 work, the underreporting rate was far higher (IRS 2016).

[74] This includes wage-and-salary employees who do work on the side, as well as sole proprietors who report income documented on 1099-MISC forms but fail to report cash-only payments (Ormiston et al. 2020).

[75] Ormiston et al. (2020) use a range of income underreporting rates as proxies. First, Alm and Erard (2016) use 2001 CPS and 2001 IRS tax data to explore the extent of income underreporting by informal workers in 12 industries, including construction. Using CPS data, the authors estimated that self-employed construction workers earned an aggregate of $53.3 billion in income in 2001. In contrast, using IRS tax data, the authors concluded that only $23.2 billion was actually reported as self-employment income and that an additional $17.7 billion was mistakenly reported as wages on workers' tax filings. This meant that out of $53.3 billion, only $40.9 billion was reported on worker tax returns, resulting in a self-employed construction worker underreporting rate of 23.3 percent.

[76] However, this is only a starting point: Hurst, Li, and Pugsley (2014) note that self-employed workers (across all industries) underreport their income by an average of 25 percent on worker surveys like the CPS. Ormiston, Belman, and Erlich (2020) point out that had this held for the sample in Alm and Erard (2016), the adjusted aggregate CPS income for self-employed construction workers would have been $66.6 billion, rather than the reported $53.3

that the BEA applies a 44% misreporting adjustment in their analysis of sole proprietors and partnerships to account for unreported self-employment net income.

The current study applies this percentage (44%) as a proxy for the proportion of construction self-employment that is informal, where each state-year ACS estimate of construction self-employment is multiplied by 0.44.[77] The result is this study's estimate of informal self-employment in construction, by state and year. Given that income underreporting rates are being used to proxy for informal self-employment, this measure is subject to error and is best regarded as an approximation; it is impossible to directly measure this with existing data. However, higher self-employment income underreporting rates have been found in academic tax literature (57%; Johns & Slemrod 2010) and by the IRS itself (64%; IRS 2016).[78] Given the prevalence of informal employment in construction (Ormiston, Belman, & Erlich 2020; Ormiston, Belman, Brockman, & Hinkel 2020), the assumption that slightly less than half (44%) of all self-employed construction workers are either misclassified or working off-the-books, while an approximation, is assumed to be reasonable if not conservative.

The second overarching issue is the underreporting of employment in household surveys, compared to payroll records. Abraham et al. (2013, 2018, 2020) point out that workers do not report all of their employment in household surveys like the CPS, thereby undercounting total

billion. This would have resulted in an income underreporting rate of 38.6 percent, rather than 23.3 percent (Ormiston et al. 2020). Even this number can be considered conservative, as there exist even higher adjustments for self-employed informal employment. The BEA applies a 44 percent misreporting adjustment in their analysis of sole proprietors and partnerships to account for unreported net income. Also, Johns and Slemrod (2010) used tax data from 2001 to show that 57 percent of nonfarm proprietor income went unreported to the IRS (Johns & Slemrod 2010; Ormiston et al. 2020). Lastly (and most importantly in the context of the current study), using more recent 2008-2010 data, the IRS recently estimated the net misreporting percentage on nonfarm proprietors' income to be 64% (IRS 2016). Ormiston et al. (2020) use all these income underreporting rates as proxies for informal self-employment and present a range of findings.

[77] In this calculation, both *incorporated* and *unincorporated* self-employment are included.

[78] Results were robust to varying proxies, both higher and lower than 44%.

employment.[79] Abraham et al. (2020) analyze employment discrepancies over the period 1996-2015. The authors use records from the Annual Social Economic Supplement to the CPS (CPS-ASEC) linked to tax information for the same individuals from the Detailed Earnings Record (DER) supplied by the Social Security Administration. The authors find that (a) self-employment in payroll records is *45% larger* than self-employment in household surveys and (b) wage-and-salary employment in payroll records is *9.5% larger* than wage-and-salary employment in household surveys. This presents a puzzle for methodology reliant on comparing household surveys with payroll records, as is done in the current study.

This study makes two adjustments in equation (7) to account for employment undercounting in household surveys. First, ACS wage-and-salary employment is adjusted upward by 9.5% to account for wage-and-salary employment underreporting in household surveys (Abraham et al. 2018; Abraham et al. 2020). Second, I adjust this study's measure of informal self-employment upward by the same percentage (9.5%). Although Abraham et al. (2020) found that self-employment in payroll records was 45% higher than household surveys, this finding was across *all* industries, not just construction. Self-employment takes on many different forms (e.g., Uber driving, babysitting), some of which may be underreported at higher rates. To be conservative, this study treats off-the-books construction work the same way as regular employment and adjusts informal self-employment upward by 9.5%, consistent with the total workforce adjustment.

[79] This issue can arise because household surveys collect information from households about work done for pay or profit. As a result, employment questions on these surveys may not always prompt respondents to report work outside of a conventional job or business (Abraham et al. 2018). This issue is exacerbated in the case of the ACS because unlike the CPS, which includes questions about second jobs, the ACS only focuses on the first (i.e., main) job. Given the prevalence of moonlighting in the construction industry (Ormiston et al. 2020), where workers perform construction work outside normal business hours, this suggests that construction employment in household surveys is *undercounted.*

The third and final issue is that the ACS does *not* ask about second jobs, while the CPS does. While the ACS is used in this study for its superior sample size, moonlighting is common in construction, and it must be accounted for. A recent CPS national average suggests that moonlighting increases total construction employment by about 2% (Ormiston et al. 2020). Although this does not account for state-to-state variation in moonlighting, this percentage is used as a general adjustment to account for workers who are in construction as a second job. All of the above adjustments bring this study's measure of informal employment from equation (6) to equation (7). Lastly, in order to express effects on informality in terms of percent change, this study's informal employment dependent variable is converted to a proportion and defined as an informal employment rate, indexed by state (s) and year (t). To calculate this dependent variable, the rate of informal construction employment, I follow Bohn and Owens (2012) in using the following equation:

$$\text{(8)} \qquad \textit{Informal Employment Rate}_{st} = \frac{Informal\ Employment}{Total\ Employment,}$$

where informal employment is the measure of informality presented in equation (7), and where total employment is the summation of (a) wage-and-salary employment, (b) incorporated self-employment, and (c) non-incorporated self-employment in the ACS.

This method for measuring informal employment involves limitations. Using the difference between worker surveys and payroll records to determine the number of jobs unaccounted for within a specific industry assumes that industry codes are (a) 100% compatible across data sources and (b) identified and coded correctly on both workers surveys and payroll records. Regarding assumption (a), these concerns are minimized—if not erased—by the fact that construction is a distinct industry code in both worker surveys and payroll records. This produces a reliable one-to-one comparison between the ACS (Census=770) and payroll records (NAICS=23). As for (b), it is less certain; prior research suggests it is not uncommon for workers' occupations to be miscoded on surveys, and it is conceivable that a similar issue exists for workers' industries.[80] While this issue can affect estimates of informal employment, no known credible assessments exist of their net effect in the construction industry. As a result, following Juravich et al. (2021), this paper assumes the net effect of this issue is zero and the data used in this study are accurate.

Sample

This study uses 2010-2019 state-level data from all 50 states to examine the impact of state prevailing wage laws on informal construction employment. Six states repealed their prevailing wage laws during this period: Arkansas (2017), Indiana (2015), Kentucky (2017), Michigan (2018), Wisconsin (2017), and West Virginia (2016), the first repeals since the 1980s (Wage and Hour Division). At the beginning of the decade, 33 states had prevailing wage laws, and 17 did not. By the end of the decade, 27 states had prevailing wage laws, and 23 did not. The initial sample size for this study was 500 (i.e., 50 states over 10 years). However, due to issues with sampling error, this sample was slightly reduced. While it is expected that payroll records like the BEA are accurate in terms of the number of legal jobs in an industry, ACS statewide estimates are based on samples that may feature substantial error. The ACS, in some cases, may only survey small samples of construction workers in a given state and year. Four states (Montana, North Dakota, Utah, and Wyoming) featured wide and unreliable year-to-year

[80] Juravich, Ormiston, and Belman (2021) suggest that employers' industry classifications may evolve over time without companies updating their industry code when filing payroll records with the state. Further, Juravich et al. (2021) point out that construction employers may be strategic in how they classify their industry code on state and federal forms, as workers' compensation insurance premiums may be higher for some industries than others.

variation in ACS construction employment estimates that resulted in wide variation in estimates of informality; in each of these four states, year-to-year changes in informality of at least 10% (and in some cases, 15%) were found in the data. It was presumed that these issues stemmed from significant sampling error; as a result, these four states were excluded from the study, bringing the sample to 460.[81]

Models

Hypothesis 1 predicted that states with prevailing wage laws will have lower rates of informal construction employment than states without prevailing wage laws. For this hypothesis, the independent variable of interest is an indicator with a value of one if a state (s) had its own prevailing wage law in year (t), and zero otherwise. Using the Wage and Hour Division (WHD) website, which publishes these lists annually, this information is compiled for the years 2010-2019.

To test Hypothesis 1, the model is the following:

$$(9) \qquad IE_{st} = \beta_0 + \beta_1 PW_{st} + \beta_2 NCIE_{st} + \boldsymbol{\beta_3 X_{st}} + \boldsymbol{\beta_4 Y_{st}} + \alpha_s + \phi_t + \mu_{st},$$

where IE_{st} denotes the rate of informal employment in a state in a given year, PW_{st} denotes state prevailing wage requirements, $NCIE_{st}$ is a lagged non-construction informal employment variable, X_{st} is a vector of state-aggregated construction industry controls, Y_{st} is a vector of state

81 Results were robust to this sample reduction; the exclusion of these four states did not alter this study's results. Including all 50 states in this analysis (with a complete 10-year sample size of 500), the presence of prevailing wage laws was associated with a 2% reduction in informal employment, nearly identical to the main result presented in Table 2.

political and legislative controls, and α_s and ϕ_t are state and year fixed effects, respectively. All variables are indexed by state (s) and/or year (t).[82]

An important control variable is the lagged rate of non-construction informal employment (i.e., one year prior). This controls for prior rates of informality in each state in all industries *except* construction. Informality can differ by state for numerous reasons, and the NCIE variable helps control for these non-construction factors (e.g., enforcement, demographics, legal frameworks). The calculations for non-construction informality (and the *rate* of non-construction informality) were created in a similar way to those of equations (7) and (8). With lagged non-construction informal employment as an explanatory variable for construction informality, the other explanatory variables measure the difference between informality in construction and informality in the rest of the economy. However, this approach does not change the interpretation of the prevailing wage variable because prevailing wage laws are unique to the construction industry. That is, even with non-construction informal employment included in the models, the prevailing wage variable still captures construction-specific effects on informality.

Next, X_{st} is a vector of state construction industry controls. The first is state construction union density (Hirsch & Macpherson 2003). Higher union densities may be associated with lower informality as more workers are afforded labor and employment law protections that combat misclassification and off-the-books employment.[83] Additionally, union reps often serve

[82] This study's models do *not* feature state-aggregated controls for race, ethnicity, or gender. Given this study's usage of panel data with state and year fixed effects, this study's models measure the effects of *year-to-year* changes in the explanatory variables on the dependent variable. The racial, ethnic, and gender makeup of a state's construction labor force is not expected to meaningfully change from year to year; this was supported by the data. As such, inclusion of these controls as explanatory variables in prior iterations of these models produced biased, inflated coefficients with high standards errors, and as a result, these controls were dropped from subsequent analyses.

[83] These two controls (non-construction informality and union density) could also be related to whether a state has a prevailing wage law, thereby constituting potential "bad controls" (Angrist & Pischke 2008). As a result, all models used in this paper were re-estimated without these two variables; results were robust to the omission of these two controls.

as advocates for workers in the *non-union* sector who are mistreated and subject to illegal practices, which could lead to lower informality. Next, this study controls for year-to-year construction employment growth, average construction firm size, year-to-year changes in building permits, and the proportion of construction workers employed in building construction. These variables are all indexed by state and year. Higher employment growth, as measured by QCEW payroll records, is included to control for fluctuations in the business cycle. Higher building permit growth could lead to higher informality as more construction projects take place.

Average firm size (as measured by QCEW payroll records) is included because recent work (Juravich et al. 2021) suggests that smaller firms may be more likely to elect to evade regulators and engage in illegal behavior. Lastly, the proportion of construction workers employed in residential construction, again taken from QCEW payroll records, is included because recent work has shown that illegal labor practices are particularly rampant in residential construction (Ormiston, Belman, Brockman, & Hinkel 2020). As a result, all else equal, it is expected that the higher the proportion of workers employed in residential construction in a given state, the higher the rate of construction informality. Lastly, a control for the proportion of annual construction value put in place from public projects is included.[84] Annual construction value put in place, measured in millions of dollars, includes both public and private construction projects. Given that the low-bid requirement only applies to public projects and does not apply to private projects, it stands to reason that informality would be higher on public projects so as to lower labor costs and make project bids as low as possible (Hinkel & Belman 2021).

[84] For this variable, only state and local projects are included in public projects; federal projects are excluded due to a lack of data availability.

Y_{st} is a vector of state political and legislative controls. The first is whether a state has a minimum wage higher than the federal rate of $7.25.[85] Higher minimum wages, all else equal, may incentivize firms to elect to operate informally to cut labor costs. The final set of legislative controls pertain to preemption, a strategy where state governments enact laws that *preempt* (i.e., void) local ordinances. The use of preemption by state governments to roll back local regulations has gained significant traction in the last decade, as local ordinances focused on expanding worker protections such as project labor agreements (PLAs), fair scheduling, and paid leave have increasingly been targeted by this strategy. Whereas only 15 cases of state preemption existed prior to 2010, this has jumped to 114 cases since 2010 (Economic Policy Institute 2019). Given the increase in year-to-year variation during the 2010-2019 period, the current study controls for this increasingly common phenomenon. States with more preemption should have higher rates of informality; as preemption occurs, local regulations are rolled back, thereby making it easier for firms to operate informally. Three controls are included, indexed by state and year: PLA preemption, fair scheduling preemption, and paid leave preemption. These variables are all reverse-coded dummy variables equaling 0 if these local ordinances have been preempted at the state level and equaling 1 if they have not.[86]

Hypothesis 2 predicted that states with stronger prevailing wage laws will have lower rates of informal employment than states with weaker laws. For this hypothesis, the independent variable of interest, prevailing wage law strength, is a measure of the strength of each state prevailing wage law, indexed by state and year. This variable is measured using the scoring system first proposed by Thieblot (1995) and used in a follow-up paper (Thieblot 1999) and

[85] The federal rate has remained at $7.25 for the entire period of this study (2010-2019).
[86] Reverse-coding these preemption variables means the coefficients can be interpreted as the effect of *an absence* of preemption. These coefficients, therefore, are expected to be negative: in the absence of preemption, local ordinances are not being rolled back by state law. This should be associated with lower informality.

recent working paper (Belman, Ormiston, Petty, & Hinkel 2020). Prevailing wage law strength is classified into three distinct groups: “strong” (12+ points), “average” (7-11 points) and “weak” (1-6 points), with the reference group being states without prevailing wage laws (i.e., scores of 0). Strong is a dummy variable equaling 1 if a state prevailing wage law has 12+ points, and 0 otherwise; average is a dummy equaling 1 if a state law has 7-11 points, and 0 otherwise; weak is a dummy equaling 1 if a state law has 1-6 points, and 0 otherwise (Belman et al. 2020). To test Hypothesis 2, these variables are all included in the following model:

$$(10) \qquad IE_{st} = \beta_0 + \beta_1 Weak_{st} + \beta_2 Average_{st} + \beta_3 Strong_{st} + \beta_4 NCIE_{st} + \boldsymbol{\beta_5 X_{st}} + \boldsymbol{\beta_6 Y_{st}} + \alpha_s + \phi_t + \mu_{st},$$

which is identical to equation (10), except that the prevailing wage dummy is replaced by the measures of prevailing wage law strength indexed by state and year.

Lastly, to address Hypotheses 3 and 4, each portion of the Thieblot scoring system is separated and included as separate variables in the model. This is to examine whether coverage threshold strength and breadth of coverage significantly impact informality more than other aspects of prevailing wage laws. This model is identical to equation (10) above, except that weak, average, and strong are replaced by the separate portions of the Thieblot scoring system.

Estimation

State and year fixed effects and all explanatory variables were included in all models, and robust standard errors (clustered by state) are estimated.[87] In testing this study’s four hypotheses,

[87] Correlations, while not included in this paper, are available upon request. The lone source of collinearity in the models were between the three types of preemption. The highest correlation between any of these three types was between fair scheduling and paid leave preemption, which was 0.48. While this was moderately high, the mean VIF

each set of models features two estimation strategies. The first (Model 1) uses standard ordinary least squares (OLS) estimation. The second strategy (Model 2) employs a generalized linear model (GLM), specifying a binomial family with a logit link (Papke & Wooldridge 1996). This accounts for the possibility that the errors of this study's proportional dependent variable are not normally distributed, a common issue with proportional dependent variables. Is it also common to employ a logit transformation, where the dependent variable is transformed into a log-odds ratio before being modeled as a linear function with OLS estimation and robust standard errors.[88] GLM estimation is used in this study rather than a logit transformation because it allows the proportional dependent variable to remain in its original form. It is designed to account for non-normal error distributions while not requiring logit transformations. As such, each set of models features two variants, the first using OLS estimation and the second using GLM estimation.

Results

Table 1 presents means and standard deviations for all variables used in this analysis. On average, in a given state and year, approximately 23% of the construction workforce was subject to informal employment (i.e., misclassification or off-the-books employment). This is in line with recent national estimates suggesting that up to one-fifth of the national construction workforce is subject to these illegal practices (Ormiston, Belman, & Erlich 2020), especially when factoring in upward adjustments made to this study's measure of informal employment. Table 1 also highlights that informality impacts construction more than the rest of the U.S. economy: while average construction informality was 23%, average non-construction informality

---

of the main OLS model was 4.63, an acceptable level. Further, fair scheduling and paid leave preemption, while moderately correlated, still focus on different issues. As a result, they were both included.

[88] Even though the proportional dependent variable is naturally bounded between 0 and 1, the log-odds ratio of this proportion can take on any value on the real line as the proportion varies between 0 and 1. As such, one can model its population regression as a linear function following the logit transformation.

in a given state and year was only 3%. Over all 460 state-year cells in the 2010-2019 period, 61% had prevailing wage laws. This reflects the fact that 33 states had prevailing wage laws at the beginning of the decade and only 27 states had these laws at the end of the decade.

Table 2 presents the results for Hypothesis 1, which predicted that states with prevailing wage laws will have lower rates of informal construction employment than states without prevailing wage laws. This hypothesis was supported. According to Model 1 of Table 2, the presence of state prevailing wage laws is associated with a 2.1% reduction in the rate of informal employment in construction, a finding statistically significant at the 99% confidence level. This finding supports the notion that due to their inherent enforcement mechanisms (e.g., certified payroll requirements), prevailing wage laws make it more difficult for construction firms to choose to operate informally and are associated with reduced informal employment.[89] In addition, Model 1 finds that a one-worker increase in average firm size is associated with a 1.1% reduction in construction informality. This is in line with the hypotheses of Juravich et al. (2021), which suggested that smaller firms, particularly those focused on unlicensed trades (e.g., carpentry, painting), may be more likely to engage in illegal behavior (Juravich et al. 2021). Also, a 1% increase in the proportion of construction value put in place from public projects is associated with an 0.05% increase in informality. This supports the notion that informality is

[89] As a further robustness check, at the suggestion of a reviewer, the results of this study were tested with a separate dependent variable: nonemployer establishments. The Census Bureau releases Nonemployer Statistics (NES) data annually; nonemployer establishments are businesses who have no employees on their payroll and are subject to federal income tax. To calculate the proportion of establishments in construction that are nonemployer, by state and year, NES data was supplemented with annual QCEW data, which provided the number of *employer* establishments in construction (i.e., those with at least 1 employee on their payroll). The proportion of nonemployer establishments was then calculated, by state and year, as a fraction of total establishments (i.e., nonemployer and employer establishments combined). The prevailing wage effect on the proportion of nonemployer construction establishments was found to be negative, but not statistically significant. However, this dependent variable does not capture the full extent of misclassification and off-the-books work. While it is a good proxy for firms who operate *exclusively* informally, it does not capture firms that employ both formal and informal workers. As a result, this methodology likely *understates* the extent of misclassification and off-the-books work.

higher on public projects so as to lower labor costs and make project bids as low as possible, key to getting bids accepted (Hinkel & Belman 2021).

The finding that prevailing wage laws are associated with lower construction informality is expected and is in line with prior work suggesting that increases in regulatory quality lower informal employment (e.g., Dabla-Norris et al. 2008; Almeida & Carneiro 2006). Further, the finding of a 2.1% reduction is reasonable. Less than a quarter of the construction workforce is misclassified or employed off-the-books (Ormiston, Belman, & Erlich 2020); thus, a moderate reduction in informality resulting from increases in regulatory quality is expected when informality itself only impacts a distinct minority of the construction workforce. Further, public projects (which prevailing wage laws directly govern) are only a part of total construction activity; construction also features a significant amount of private work. Although firms are required to submit certified payrolls on public projects covered by prevailing wage requirements, misclassification and off-the-books employment still occurs in cases where local regulators may not notice or pursue illegal practices. Thus, the prevailing wage coefficient has the expected sign, and its magnitude is reasonable.

Model 2 of Table 2 presents GLM results. Log odds coefficients, while less intuitive than traditional OLS coefficients, can be interpreted in terms of the likelihood of a particular outcome (in this case, informal employment). While the dependent variable was not logit transformed, GLM coefficients still require log odds interpretations. According to Model 2, the presence of state prevailing wage laws is associated with reduced informality odds of 10.2%, which was statistically significant.[90] In Table 2, the variation in prevailing wage coefficients between Models 1 and 2 exists because the two models estimate different outcomes. While the OLS

[90] The prevailing wage log odds coefficient of -0.108 is interpreted as follows: exp(-0.108) = 0.898; 1-0.898 = 0.102.

model (Model 1) estimates the association between state prevailing wage laws and statewide *rates* of misclassification and off-the-books employment, the GLM model (Model 2) estimates the association between state prevailing wage laws and the *likelihood* of workers being misclassified or employed off-the-books.[91] Though the coefficients necessarily differ, they should have the same sign; indeed, both prevailing wage coefficients are negatively signed. Informality is significantly lower in states with prevailing wage laws than in states without prevailing wage laws, thereby supporting Hypothesis 1. Taken together, while not a panacea, these results suggest that prevailing wage laws can help attenuate the issue of informality that plagues the U.S. construction industry. With one-fifth or more of the industry's labor force either misclassified or working off-the-books, costing workers numerous labor law protections and federal and state governments hundreds of millions of dollars in lost tax revenues (Ormiston, Belman, & Erlich 2020), heightened enforcement from state prevailing wage laws can benefit construction workers and taxpayers alike by making formal employment a more profitable business decision.

Hypothesis 2 focused on prevailing wage law strength and predicted that states with stronger prevailing wage laws will have lower rates of informal construction employment than states with weak laws. Table 3 presents these results; Hypothesis 2 did not receive support. Weak prevailing wage laws are associated with a marginally significant 1.7% reduction in informality; meanwhile, average and strong prevailing wage laws are associated with statistically significant 2.8% and 2.2% reductions, respectively. Following estimation, differences between these coefficients were tested, and they were not found to be statistically significant. The differences

[91] Given the indirect methods used to measure informality in this study, in cases where odds of informality are lower, it cannot be determined whether workers move from informal to formal work in construction or whether workers move from informal work in construction to informal work in another industry.

were also not found to be statistically significant in Models 2 and 3. However, this stems from the result that having a weak or average prevailing wage law is enough to significantly curb informality.

Table 4 presents the results for Hypotheses 3 and 4. Here, the measure of prevailing wage law strength is separated into distinct categories: (1) the minimum contract threshold for coverage under the state law, (2) the enforced wage rate, (3) the breadth of work and occupations covered by the law and (4) "other factors," including state and local resident preferences, administrative and/or compliance requirements, and penalties for violating the law.

Hypothesis 3 predicted that states with lower coverage thresholds will have lower rates of informal construction employment than states with higher thresholds. Following Thieblot (1995, 1999), the threshold strength variable was measured in the following way: a value of 0 denotes a construction contract threshold for prevailing wage coverage of more than $50,000; a value of 1 denotes a coverage threshold between $2,000 and $50,000; and a value of 2 denotes a coverage threshold below $2,000. As the value of this variable increases from 0 to 2 (i.e., as the contract threshold for prevailing wage coverage lessens), more projects are required to adhere to prevailing wage laws, thereby increasing the strength of the threshold. Hypothesis 3 was supported. According to Model 1 of Table 4, a 1-point increase in threshold strength is associated with a statistically significant 1.8% reduction in informality. Further, the GLM coefficient in Model 2 was also negative and statistically significant, suggesting that a 1-point increase in threshold strength is associated with reduced informality odds of 8.7%.[92] As for Hypothesis 4, which predicted that states with more types of projects covered will have lower rates of informal construction employment than states with fewer project types covered, this

---

[92] The interpretation of the GLM coefficient in Model 3 of Table 3 is as follows: exp(-0.091)=0.913; 1-0.913=0.087.

hypothesis did not receive support. Across all models in Table 4, while the coefficients were negative, breadth of coverage was not statistically significant. In sum, these results suggest that coverage thresholds may be the main factors of prevailing wage laws that affect the degree of informality in construction.

Discussion and Conclusion

In sum, these analyses produced two key results. First, state prevailing wage laws are associated with lower informal construction employment. Results indicate that prevailing wage laws are associated with a 2.1% reduction in informality and reduced informality odds of 10.2%. Second, even weak or average prevailing wage laws are associated with lower informality. Based on this study's methodology for estimating informal employment, across all six states that repealed their prevailing wage laws during the 2010-2019 period (Arkansas, Indiana, Kentucky, Michigan, West Virginia, and Wisconsin), an annual average of 192,000 construction workers (22.7% of the workforce) were either misclassified or employed off-the-books prior to repeal. Since the presence of state prevailing wage laws is associated with a 2.1% reduction in informality, it follows that repealing these laws is associated with a 2.1% increase in informality. This translates to an average of approximately 4,030 additional construction workers misclassified or employed off-the-books across these six states in a given year.[93] The results of this study suggest that while prevailing wage laws are an additional form of labor market regulation in construction (thereby increasing regulatory quantity), they also increase regulatory *quality* in the construction context and are associated with reductions in informal employment in the industry. This is consistent with prior research suggesting that increases in regulatory quality

[93] This estimate is based on a state average of approximately 670 additional construction workers misclassified or employed off-the-books in a given year. This is then multiplied by the number of state repeals (6) to arrive at an annual estimate.

significantly reduce informal employment (e.g., Dabla-Norris et al. 2008; Almeida & Carneiro 2006).

It should be noted, however, that these results do not represent causal effects of prevailing wage laws on informality. For example, it is plausible that contractors simply operate legally under prevailing wage requirements, rather than change illegal behavior and move from informal to formal employment. That is, these results cannot establish a causal impact where prevailing wage laws directly lead to changes in contractor behavior. Even so, the results of this study seem to support this paper's theoretical predictions that due to inherent enforcement mechanisms (e.g., certified payroll requirements), prevailing wage laws make it more difficult and less profitable for construction firms to choose to operate informally. This benefits law-abiding construction firms by making formal employment (and conducting business legally) a better business decision.

Historically, prevailing wage laws have been utilized on public works to address issues with labor standards. As mentioned in the introduction, enforcement of labor standards and regulating informality in construction is difficult. Resources available to government regulators and tax agencies have declined over time; further, even with more resources, it is particularly difficult for regulatory agencies to crack down on off-the-books employment due to workers being entirely absent from payrolls and a lack of paper trails. This study's results suggest that prevailing wage laws can boost enforcement and limit the degree of informality without states having to significantly increase investments in enforcement resources. This results in a net gain for taxpayers: as prevailing wage laws increase enforcement on public projects (without a significant increase in taxpayer costs), firms and workers pay more of their income in taxes as informality declines. This boosts state government revenues, thereby leaving them better able to provide vital services to citizens.

These findings also have implications for construction workers. As mentioned in the introduction, misclassified and off-the-books workers lose access to unemployment insurance, workers' compensation, minimum wages, overtime payments, anti-discrimination defenses, and the right to unionize (Harris & Krueger 2015; Erlich 2020). By incentivizing construction firms to play by the rules and limiting the degree of informality, state prevailing wage laws allow more workers to directly benefit from these key labor and employment law protections. While there are plenty of legitimate independent contractors in construction (Erlich 2020), there are also numerous misclassified and off-the-books workers in the industry (Ormiston, Belman, & Erlich 2020; Ormiston, Belman, Brockman, & Hinkel 2020). State prevailing wage laws directly benefit workers who would otherwise be subject to these illegal practices.

This study is not without limitations. First, while including the self-employed in calculations of informality represents an improvement over prior literature, proxying for informal self-employment with income underreporting assumes that income underreporting is indeed an accurate proxy for informal activity. The assumption that 44% of all self-employment activity in construction is informal, while assumed to be reasonable if not conservative, is subject to error; the *actual* proportion of construction self-employment that is informal could feasibly be higher than 44%. Additionally, as previously mentioned, this study's measurement of informality relies on comparing two disparate data sets (worker surveys and payroll records) to determine the number of jobs unaccounted for within a specific industry. This indirect methodology is admittedly an imperfect measure and an approximation.

Second, this methodology cannot capture all types of employment fraud in construction. It cannot capture other types of fraud that take place even as firms seemingly comply with prevailing wage and certified payroll requirements. For example, a firm may agree to pay workers prevailing wages when taking a project bid, but then illegally use third-party payroll companies to circumvent the requirements. In this case, firms pay the required prevailing wage rate to the payroll company and indicate on the certified payroll form that they did so. Then, the payroll company pays workers only a fraction of this prevailing wage rate, leaving the rest for the contractor. From a prevailing wage and certified payroll perspective, the contractor is seemingly conducting business legally; in order for regulators to catch this behavior, they would need to check the payroll company's paystubs *in addition to* the certified payroll form. This behavior cannot be captured by this study's methodology; these cases would be classified by this study's methodology as formal rather than informal because firms look like they are following the law.

Employers may also illegally misrepresent the occupations of their workers on the certified payroll forms. For example, an employer may label a worker who performed electrical work as a "construction laborer," or a worker who performed plumbing work as a "painter" or a "carpenter." This is done to save on labor costs, as prevailing wage rates for laborers, painters, and carpenters are often significantly lower than those of electricians and plumbers. It is unclear how often (or to what extent) this practice goes on; however, conversations with industry experts suggest employers use this mechanism as a way to reduce labor costs, even as they comply with certified payroll requirements on prevailing wage projects. In sum, contractors may choose to respond to prevailing wage and certified payroll requirements by finding other creative ways to illegally reduce labor costs that would be unaccounted for with this study's methodology.

Lastly, any study involving the measurement of informality is limited by how difficult informality is to track. These illegal activities are intentionally designed to be concealed from regulators, as they typically lack paper trails and go unreported to appropriate taxation bureaus. Off-the-books arrangements can be especially difficult to track and measure, as paying workers under the table in cash leaves these workers entirely absent from payrolls and not a part of reports to taxation bureaus. Future research is encouraged to improve the accuracy of informality estimates via the usage of matched administrative data (e.g., the Longitudinal Employer-Household Dynamics, or LEHD, database). For example, if a worker claims to be an employee of a particular firm, but the firm does not list that worker on their payroll records, this could offer clearer evidence of misclassification (Ormiston et al. 2020). While accessibility to this data requires a lengthy approval process, by linking worker surveys to employer payroll files, it presents a promising avenue for researchers interested in studying informality.

APPENDIX

Table 3.1: Descriptive Statistics

Variables	Means	Standard Deviations
Informal employment	0.23	0.06
Non-construction informal employment	0.03	0.03
Prevailing wage	0.61	0.49
Strong prevailing wage law	0.45	0.50
Average prevailing wage law	0.17	0.38
Weak prevailing wage law	0.08	0.27
Coverage threshold	0.68	0.87
Breadth of coverage	1.28	1.65
Setting of prevailing wage rate	3.28	3.03
Other factors	0.53	1.13
Private construction union density	0.14	0.10
Employment growth rate	0.02	0.05
Average firm size	7.75	2.04
Percent employed residential	0.22	0.02
Building permits	0.08	0.17
Percent state and local	0.54	0.12
Minimum wage	0.47	0.50
Fair scheduling preemption	0.93	0.25
Paid leave preemption	0.77	0.42
Project labor agreement preemption	0.73	0.44

Table 3.2: Effect of Prevailing Wage Laws on Informal Construction Employment

Variables	OLS	GLM
Prevailing wage	**-0.021*** **	**-0.108** **
	(0.008)	(0.046)
Non-construction informal employment	**0.060**	**0.403**
	(0.151)	(0.818)
Private construction union density	**-0.023**	**-0.123**
	(0.065)	(0.346)
Employment growth rate	**-0.208*** **	**-1.324** **
	(0.073)	(0.539)
Average firm size	**-0.011* **	**-0.072** **
	(0.005)	(0.033)
Percent employed residential	**0.801**	**5.199**
	(0.538)	(3.597)
Building permits	**0.018**	**0.087**
	(0.014)	(0.069)
Percent state and local	**0.055** **	**0.344** **
	(0.026)	(0.150)
Minimum wage	**0.006**	**0.038**
	(0.009)	(0.049)
Fair scheduling preemption	**-0.010**	**-0.070**
	(0.010)	(0.053)
Paid leave preemption	**-0.014** **	**-0.075** **
	(0.007)	(0.036)
Project labor agreement preemption	**-0.014* **	**-0.075* **
	(0.008)	(0.042)
Constant	**0.161**	**-1.071* **
	(0.138)	(0.903)
N	455	455
R^2	0.706	

Estimated coefficients in bold. Robust standard errors in parentheses. State and year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 3.3: Effect of Prevailing Wage Law Strength on Informal Construction Employment

Variables	OLS	GLM
Strong prevailing wage law	**-0.022***	**-0.123**
	(0.013)	(0.079)
Average prevailing wage law	**-0.028****	**-0.151****
	(0.012)	(0.075)
Weak prevailing wage law	**-0.017***	**-0.085**
	(0.009)	(0.056)
Non-construction informal employment	**0.070**	**0.447**
	(0.153)	(0.820)
Private construction union density	**-0.020**	**-0.107**
	(0.065)	(0.347)
Employment growth rate	**-0.211*****	**-1.342****
	(0.073)	(0.546)
Average firm size	**-0.011***	**-0.070****
	(0.005)	(0.033)
Percent employed residential	**0.830**	**5.368**
	(0.539)	(3.612)
Building permits	**0.018**	**0.086**
	(0.014)	(0.070)
Percent state and local	**0.056****	**0.347****
	(0.026)	(0.152)
Minimum wage	**0.007**	**0.038**
	(0.009)	(0.049)
Fair scheduling preemption	**-0.007**	**-0.052**
	(0.010)	(0.053)
Paid leave preemption	**-0.014****	**-0.077****
	(0.007)	(0.037)
Project labor agreement preemption	**-0.014***	**-0.076***
	(0.009)	(0.046)
Constant	**0.151**	**-1.770***
	(0.140)	(0.913)
N	455	455
R^2	0.707	

Estimated coefficients in bold. Robust standard errors in parentheses. State and year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

Table 3.4: Effect of Prevailing Wage Strength: Individual Components

Variables	OLS	GLM
Coverage threshold	**-0.018****	**-0.091****
	(0.007)	(0.039)
Breadth of coverage	**-0.005**	**-0.022**
	(0.006)	(0.039)
Setting of prevailing wage rate	**-0.000**	**-0.004**
	(0.003)	(0.019)
Other factors	**0.001**	**0.012**
	(0.005)	(0.030)
Non-construction informal employment	**0.065**	**0.424**
	(0.149)	(0.801)
Private construction union density	**-0.028**	**-0.147**
	(0.065)	(0.345)
Employment growth rate	**-0.202*****	**-1.290****
	(0.072)	(0.535)
Average firm size	**-0.012****	**-0.079****
	(0.006)	(0.035)
Percent employed residential	**0.810**	**5.211**
	(0.531)	(3.535)
Building permits	**0.019**	**0.090**
	(0.013)	(0.068)
Percent state and local	**0.061****	**0.376****
	(0.027)	(0.154)
Minimum wage	**0.008**	**0.044**
	(0.009)	(0.049)
Fair scheduling preemption	**-0.011**	**-0.078**
	(0.010)	(0.056)
Paid leave preemption	**-0.015****	**-0.079****
	(0.006)	(0.034)
Project labor agreement preemption	**-0.011**	**-0.062**
	(0.009)	(0.047)
Constant	**0.166**	**-1.672***
	(0.138)	(0.888)
N	455	455
R^2	0.707	

Estimated coefficients in bold. Robust standard errors in parentheses. State and year fixed effects included in all models.
*Statistically significant at the 10% level; ** at the 5% level; *** at the 1% level.

REFERENCES

REFERENCES

Abraham, Katharine G., John Haltiwanger, Kristin Sandusky, and James R. Spletzer. 2013. Exploring differences in employment between household and establishment data. *Journal of Labor Economics 31*(S1): S129-S172.

Abraham, Katharine G., John Haltiwanger, Kristin Sandusky, and James R. Spletzer. 2018. Measuring the gig economy: Current knowledge and open issues. NBER Working Paper.

Abraham, Katharine G., John Haltiwanger, Claire Hou, Kristin Sandusky, & James R. Spletzer. 2020. Reconciling survey and administrative measures of self-employment. *Journal of Labor Economics*, forthcoming.

Alm, James, and Brian Erard. 2016. Using public information to estimate self-employment earnings of informal suppliers. *Public Budgeting & Finance 36*(1): 22-46.

Almeida, Rita, and Pedro Manuel Carneiro. 2006. Enforcement of regulation, informal labour, firm size and firm performance. CEPR.

Belman, Dale, Russell Ormiston, Ryan Petty, and Matthew Hinkel. 2020. The effect of state prevailing wage laws on African American employment in construction. Unpublished working paper.

Bohn, Sarah, and Emily Greene Owens. 2012. Immigration and informal labor. *Industrial Relations 51*(4): 845-873.

Carre, Francoise, and Randall Wilson. 2004. "The social and economic costs of employee misclassification in construction." Accessed at http://citeseerx.ist.psu.edu/viewdoc/download?doi=10.1.1.679.4695&rep=rep1&type=pdf

Cuff, Katherine, Steeve Mongrain, and Joanne Roberts. 2020. The evasion of fiscal and labor regulations: Firm behavior and optimal tax policy. *Journal of Public Economic Theory 22*(1): 69-97.

Dabla-Norris, Era, Mark Gradstein, and Gabriela Inchauste. 2008. What causes firms to hide output? The determinants of informality. *Journal of Development Economics 85*(1-2): 1-27.

Duncan, Kevin, and Russell Ormiston. 2018. What does the research tell us about prevailing wage laws? *Labor Studies Journal* 44(2): 139-160.

Economic Policy Institute. 2019. "Worker rights preemption in the U.S.: A map of the campaign to suppress worker rights in the states." Accessed at https://www.epi.org/preemption-map/.

Enste, Dominik H. 2010. Regulation and shadow economy: Empirical evidence for 25 OECD countries. *Constitutional Political Economy 21*(3): 231-248.

Erlich, Mark. 2020. Misclassification in construction: The original gig economy. *ILR Review,* forthcoming.

Friedman, Eric, Simon Johnson, Daniel Kaufmann, and Pablo Zoido-Lobaton. 2000. Dodging the grabbing hand: The determinants of unofficial activity in 69 countries. *Journal of Public Economics 76*(3): 459-493.

Harris, Seth D., and Alan B. Krueger. 2015. A proposal for modernizing labor laws for twenty-first-century work: The "independent worker". Washington, DC: Brookings.

Hinkel, Matthew, and Dale Belman. 2021. Should prevailing wages prevail? Re-examining the effect of prevailing wage laws on affordable housing construction costs. Revise and resubmit at *British Journal of Industrial Relations.*

Hirsch, Barry T., and David A. Macpherson. 2003. Union membership and coverage database from the Current Population Survey: Note. *ILR Review 56*(2): 349-354.

Hurst, Erik, Geng Li, and Benjamin Pugsley. 2014. Are household surveys like tax forms? Evidence from income underreporting of the self-employed. *Review of Economics and Statistics 96*(1): 19-33.

Internal Revenue Service. 2016. "Federal tax compliance research: Tax gap estimates for tax years 2008-2010." IRS Publication 1415. https://www.irs.gov/pub/irs-soi/p1415.pdf.

Johns, Andrew, and Joel Slemrod. 2010. The distribution of income tax noncompliance. *National Tax Journal 63*(3): 397.

Johnson, Simon, Daniel Kaufmann, and Pablo Zoido-Lobaton. 1998. Regulatory discretion and the unofficial economy. *The American Economic Review 88*(2): 387-392.

Juravich, Tom, Russell Ormiston, and Dale Belman. 2021. The social and economic costs of illegal misclassification, wage theft, and tax fraud in residential construction in Massachusetts. University of Massachusetts-Amherst Labor Center Working Paper Series.

Ormiston, Russell, Dale Belman, and Mark Erlich. 2020. An empirical methodology to estimate the incidence and costs of payroll fraud in the construction industry. *United Brotherhood of Carpenters and Joiners of America (UBC).*

Ormiston, Russell, Dale Belman, Julie Brockman, and Matthew Hinkel. 2020. Rebuilding residential construction. In P. Osterman (Ed.), *Creating Good Jobs: An Industry-Based Strategy,* pp. 75-113. MIT Press.

Papke, Leslie E., and Jeffrey M. Wooldridge. 1996. Econometric methods for fractional response variables with an application to 401 (k) plan participation rates. *Journal of Applied Econometrics 11*(6): 619-632.

Ruggles, Steven, Sarah Flood, Ronald Goeken, Josiah Grover, Erin Meyer, Jose Pacas and Matthew Sobek. 2020. *IPUMS USA: Version 10.0* [dataset]. Minneapolis, MN: IPUMS**,** 2020. https://doi.org/10.18128/D010.V10.0

Thieblot, Armand J. 1995. State prevailing wage laws: An assessment at the start of 1995. Rosslyn, Va.: State Relations Department, Associated Builders and Contractors, 1-39.

Thieblot, Armand J. 1999. Prevailing wage laws and Black employment in the construction industry. *Journal of Labor Research 20:* 155-159.

Ulyssea, Gabriel. 2018. Firms, informality, and development: Theory and evidence from Brazil. *American Economic Review 108*(8): 2015-47.

United States Department of Labor. 2015. Introduction to labor standards statutes coverage. Accessed at https://www.dol.gov/whd/recovery/pwrb/Tab3.pdf.

United States Department of Labor Wage and Hour Division. Instructions for completing payroll form, WH-347. Accessed at https://www.dol.gov/agencies/whd/forms/wh347.

Weil, David. 2014. *The fissured workplace*. Harvard University Press.

Wells, Jill. 2007. Informality in the construction sector in developing countries. *Construction Management and Economics 25*(1): 87-93.