The Impact of Home-Based Child Care Provider Unionization on the Cost, Type, and Availability of Subsidized Child Care in Illinois

Author(s): Todd Grindal, Martin R. West, John B. Willett and Hirokazu Yoshikawa

Source: *Journal of Policy Analysis and Management*, Vol. 34, No. 4 (FALL 2015), pp. 853-880
Published by: Wiley on behalf of Association for Public Policy Analysis and Management
Stable URL: https://www.jstor.org/stable/43866395

The Impact of Home-Based Child Care Provider Unionization on the Cost, Type, and Availability of Subsidized Child Care in Illinois

Todd Grindal
Martin R. West
John B. Willett
Hirokazu Yoshikawa

Abstract

In February 2005, Illinois became the first U.S. state to grant home-based child care providers (HBCPs) the right to form a labor union in order to bargain collectively with the state government. This policy inspired similar efforts across the country and represents a potentially important direction for child care policy. To date, the implications of labor unions for the cost, type, and availability of subsidized child care have not been evaluated empirically. In this study, we examine the impact of granting Illinois HBCPs the right to form a labor union on (a) the type of child care (licensed vs. license-exempt/home-based vs. center-based) used by subsidy-receiving Illinois infants and toddlers; (b) the per-child cost of subsidized child care for infants and toddlers; and (c) the percentage of Illinois infants and toddlers who use child care subsidies. To conduct these analyses, we combine data from the Current Population Survey with Child Care and Development Fund administrative records on U.S. infants and toddlers whose families received child care subsidies during the period from 2002 to 2008. We use both a traditional difference-in-differences as well as a comparative case study with a "synthetic" control group approach. The synthetic control group approach improves on traditional comparative case studies by providing a transparent, empirical approach for constructing the counterfactual, documenting comparison units' contribution to the synthetically created control group and detailing the degree to which the synthetic control group is, or is not, similar to the treated unit on preintervention measures of the outcome as well as on other selected characteristics. We find that subsidy-receiving Illinois infants and toddlers spent an average of between 6.4 and 7 percentage points more hours in licensed care settings, as compared to license-exempt settings, in the three years following child care unionization. We also find that between 0.7 and 1.1 percentage points fewer Illinois infants and toddlers used child care subsidies following unionization. © 2015 by the Association for Public Policy Analysis and Management.

INTRODUCTION

In February 2005, Illinois became the first state in the United States to grant home-based child care providers (HBCPs) the right to form a labor union and bargain collectively with the state government over wages and working conditions. The implementation of the Illinois policy inspired similar efforts across the country (Blank, Campbell, & Entmacher, 2010). To date, unions representing home-based child care providers have been authorized to organize in more than a dozen states.

Home-based child care is a widely used form of nonparental care provided to low-income children under age 3 years in the United States (Matthews & Lim, 2010).

Journal of Policy Analysis and Management, Vol. 34, No. 4, 853–880 (2015)

Published by Wiley Periodicals, Inc. View this article online at wileyonlinelibrary.com/journal/pam
DOI:10.1002/pam.21840

Unfortunately, despite its ubiquity, it also ranks below formal center-based child care on nearly every indicator of program quality (Fuller, Kagan, & Loeb, 2004; Gormley, 2007). This is cause for concern, as children enrolled in low-quality programs tend to have worse outcomes, in both the short and long term, when compared to children enrolled in higher quality programs (Shonkoff & Phillips, 2000).

In other industries, labor unions have assisted workers' efforts to negotiate for higher wages and provided a stronger voice in determining workplace practices (Freeman & Medoff, 1979). The unionization of teachers, public safety workers, and other municipal employees has also, in many instances, expanded the size of the public sector labor force (Freeman, 1984). Though similar to these other public sector unions in many ways, child care worker unions operate within a unique institutional and market setting that makes their likely impact difficult to anticipate. To our knowledge, there have been no formal evaluations of the impact of the formation of labor unions on critical outcomes for either the child care workers themselves or their charges (Porter et al., 2010).

In the current study, we investigated the causal impact of Illinois HBCPs' first collective bargaining agreement, ratified in January 2006, on several policy-relevant outcomes:

(a) The type of settings (licensed vs. license-exempt/home-based vs. center-based) attended by Illinois infants and toddlers whose parents receive child care subsidies.
(b) The per-child cost of subsidized child care for Illinois infants and toddlers.
(c) The percentage of Illinois infants and toddlers whose parents use child care subsidies.

We analyzed nationally representative administrative data from the Administration for Children and Families (ACF) along with information from the Current Population Survey (CPS). In our preferred analyses, we applied an innovative strategy for conducting comparative case studies developed by Abadie, Diamond, and Hainmueller (2010) that involved the analytic creation of a "synthetic" control group of states that did not implement child care unionization, and against which we could compare average values of the critical outcomes that had been obtained in Illinois, pre- and postunionization. We also analyzed the same outcomes using a more conventional difference-in-differences approach.

We found that unionization had several important consequences in Illinois. First, relative to the synthetic control group, subsidy-receiving Illinois infants and toddlers spent an average of between 6.4 and 7 percentage points more total care hours being cared for in licensed settings, as compared to license-exempt settings, in the three years following child care unionization. Second, we did not find consistent evidence that unionization led to a change in the percentage of child care hours provided in unionized home-based as compared to nonunionized center-based settings following unionization. Third, after unionization between 0.7 and 1.1 percentage points fewer Illinois infants and toddlers received child care subsidies. Finally, we found some evidence that home-based child care workers received more money per child, per month following unionization but that this difference was not consistent across all analyses. We found generally similar effects of unionization when we conducted the analyses using a difference-in-differences approach.

BACKGROUND

Out-of-home child care is a common experience for infants and toddlers in the United States. Approximately 6 million (49 percent) children under the age of three years regularly receive care from someone other than a parent and spend, on average, more than 30 hours each week in these settings (Iruka & Carver, 2006). This is important because it is widely accepted that children's child care experiences exert a meaningful influence on their physical and intellectual development (Belsky et al., 2007) and have lasting effects on their academic achievement (Werner, 1989), economic productivity (Heckman, 2006), and physical and mental health (Shonkoff, Boyce, & McEwen, 2009). However, extant research cautions that only high-quality programs, which provide language-rich environments, responsive interactions, and a variety of stimulating activities and materials, have yielded positive results for children and society (Shonkoff & Phillips, 2000).

Child Care Subsidies

The U.S. federal government's Child Care and Development Fund (CCDF) provides low-income parents of children under age 14, and parents of children with disabilities, with financial assistance with which to purchase child care. States receive CCDF funds as a block grant that they may supplement by designating other state funds or by transferring up to 30 percent of funds from their Temporary Assistance for Needy Families block grant, which also covers child care as a work support. States are subject to federal requirements regarding how their federal CCDF funds are allocated, but are, by and large, permitted to determine the rules regarding program eligibility, application procedures, levels of assistance, and permissible subsidy use (Greenberg, Lombardi, & Schumacher, 2000). This flexibility leads to state-to-state differences in the proportion of families within a state who receive subsidies, as well as in the value of the subsidies (Rigby, Ryan, & Brooks-Gunn, 2007).

In most states, subsidy-receiving families are permitted to spend their CCDF vouchers in a range of child care settings, including paying for care provided in private homes, for-profit centers, and community-based organizations (Gennetian et al., 2004). Following Porter and colleagues (2010), we define the term "home-based child care" to encompass (a) care provided by a nonparental relative in the child's, or the caregiver's, home; (b) care provided by a nonrelative in the caregiver's home; and (c) care provided in the child's home by babysitters, neighbors, friends, and other nonrelatives. A substantial percentage of low-income families nationwide use home-based child care due to its typically more flexible hours, lower cost, and greater availability in low-income neighborhoods (Laughlin, 2013; National Institute of Child Health and Human Development, 2000).

It is therefore troubling that home-based programs typically lag behind center-based programs on indicators of program quality. One reason for these differences is that home-based providers are often exempt from many of the regulations and oversight that govern the care of young children in center-based settings. Although center-based child care programs generally receive annual inspections and must follow regulations regarding teacher education and instructional content, some states allow home-based providers serving five or fewer nonrelative children to operate without a license or a state inspection. Many states require providers who are otherwise license-exempt to submit to some form of regulation in order to serve subsidy-receiving children, but some states do not even require these providers to submit to criminal background checks. In Illinois, the focal state in this study, home-based providers are permitted to care for up to three nonrelative children before being subject to licensing (225 IL CS 10/part 377). License-exempt Illinois providers who wish to care for subsidy-receiving children must submit to state criminal and Child Protective Services background checks.

Prior research indicates that home-based child care is typically rated as less stimulating cognitively and less safe physically than comparable center-based care (Fuller, Kagan, & Loeb, 2004; National Association of Child Care Resource and Referral

Agencies, 2011). Children in home-based settings spend more time watching television (Christakis & Garrison, 2009), and less time in goal-directed academic or social activities (Layzer & Goodson, 2006) than otherwise similar children in center-based programs. Compared to workers in center-based programs, home-based providers earn lower wages, have less access to affordable health insurance, and are less likely to have completed any formal schooling beyond high school or to have had specific training for the care of young children (Burton et al., 2002).

The Impact of Labor Unions on the Cost and Size of the Public Sector Workforce

Prior work suggests that workers are motivated to join unions out of a desire to increase their levels of compensation, improve job security, promote fairness, and secure roles in workplace decisionmaking (Freeman & Rogers, 1999; McClendon, Wheeler, & Weikle, 1998). Freeman and Medoff (1979) argue that, as a result, labor unions function as instruments of both worker monopoly and worker voice. A union's monopoly functions include raising wages, limiting management authority, and improving the working conditions of members. A union's voice functions, on the other hand, bring the perspectives of workers to policy discussions to reduce worker exit, improve efficiency, and benefit the enterprise as a whole.

In the private sector, market competition may limit the extent of unions' monopoly activities. If a private sector union demands inordinate wages and benefits, the resulting costs may compel firms to cease operation or relocate production to an area that restricts collective bargaining. Union-bargained wage increases must therefore be accompanied by increased productivity from existing workers. As a result, private sector unions typically lead firms to increase per-employee compensation but not to increase the size of the workforce (Freeman & Rogers, 1999).

Limitations on a union's monopoly powers are less clear in the public sector, where unions can use political strategies to pursue increased wages and employment simultaneously (Norcross, 2010). Branchflower and Bryson (2004) found that between 1996 and 2001, public sector workers who were members of a union earned wages that were, on average, 15 percent higher than those of observably similar public sector workers who were not unionized. Valletta (1989) found that local governments with unionized workers had, on average, larger municipal workforces. Similarly, Zax and Ichniowski (1988) found that public sector workers who engaged in collective bargaining had higher average wages than similar nonunion workers, and were employed in greater numbers. They also found that these higher wage and per capita employment levels for unionized workers were offset by lower numbers of full-time employment positions among public sector workers, within the same municipality, who were not covered by a collective bargaining agreement.

That said, there are arguably some important constraints on the power of public sector unions. For example, recent work by Lewin, Keefe, and Kochan (2012) suggests that public sector unions are not, on balance, more powerful than private sector unions. This is consistent with work by Freeman (1986), which maintains that local citizens have the capacity to counter public sector unions' demands for higher wages by relocating to communities with lower public sector wages. However, Brueckner and Neumark (2011) found that this limitation of public sector union monopoly power is not present in communities with high levels of amenities such as warm weather or coastal location, or in large, dense urban areas.

Institutional Setting of Home-Based Child Care

The unique mixed-market setting within which HBCP unions operate makes it unclear whether unionization will yield wage and employment effects similar to those of other public or private sector unions. Unlike most union members, home-based providers are independent businesspeople working out of their homes. As self-employed individuals who do not share a physical workspace, home-based providers appear to lack a common management entity with which to negotiate—something ordinarily seen as essential for collective bargaining. However, the state child care subsidy system links these otherwise independent individuals together. The state provides families with vouchers they can use to purchase child care from providers of their choosing. Families combine these subsidies with personal funds to compensate their chosen provider. This framework, according to those seeking to form child care worker unions, makes the state the home-based providers' de facto employer and makes those providers who accept child care subsidy vouchers public sector workers (Smith, 2007).

The first HBCP collective-bargaining agreement, ratified by Illinois home-based providers in January 2006 and valid for three years, contained four key provisions. First, it provided an approximately 35 percent increase in rates paid to home-based providers to care for the children of subsidy recipients. Second, the agreement provided additional financial incentives for those home-based providers who agreed to participate in the state's new quality rating and improvement system. Third, the agreement provided $27 million in funding to help HBCPs access health insurance. Finally, it provided an additional $18 million in subsidy rate increases for nonunionized center-based child care programs.

The Illinois home-based provider union differed from similar unions in other parts of the country during the study period in some important ways. Illinois allowed automatic union enrollment and dues payments of all home-based providers who care for subsidy-receiving children, while Wisconsin, Kansas, Oregon, and New Mexico did not. The Illinois law provided home-based providers with strong collective-bargaining rights, while the laws in Iowa, Kansas, and Wisconsin permitted workers the right only to "meet and confer" with the state. In Illinois, collective-bargaining agreements between home-based providers and the state were self-executing, whereas agreements in other union states, such as Kansas, Washington, Minnesota, New York, Ohio, and Wisconsin, were subject to approval by the state legislature (Blank, Campbell, & Entmacher, 2010).

The literature on public sector labor unions and the contents of the January 2006 Illinois HBCP collective-bargaining agreement suggest two hypotheses regarding how home-based provider labor unions might affect the child care used by Illinois subsidy-receiving families. The first is a "professionalization" hypothesis. Exemplified by the theoretical work of Freeman and Medoff (1979), this hypothesis suggests that unions will help child care become a respected profession and thereby improve services for children. There is some prior empirical evidence to support this hypothesis. A survey of unionized providers in Washington conducted by Burris (2012) found that union-provided training sessions increased providers' self-reported knowledge and skills and allowed them to establish professional connections with other providers. Similar research focused on New Jersey home-based providers conducted by Houser, Nisbet, and White (2012) also indicated that providers believed that union-provided information was helpful and that the union was focused broadly on improving the quality of child care.

Another observable way unionization might help to professionalize home-based child care would be to encourage license-exempt providers, as well as people who enter the field following unionization, to submit to the regulations necessary to obtain licensure. All else being equal, this would produce an increase in the percentage of HBCPs who are licensed, rather than license-exempt, following unionization. However, enhancing the professional status of home-based providers could also encourage some center-based child care workers to provide home-based, rather than center-based, services. Were such a phenomenon to occur, it could obscure any changes in the percentage of licensed care used by subsidy-receiving families. In this study, we therefore compared the overall percentage of care hours during which the children of subsidy recipients received licensed infant/toddler care in both center- and home-based settings. Given prior research indicating that licensed care is, on average, of higher quality than license-exempt care, an increase in the percentage of care experienced in licensed home-based settings should be interpreted as improving the overall quality of the child care used by subsidy-receiving families, and thus represent a positive effect of unionization.

We term the second hypothesis concerning how unionization might affect the child care used by subsidy-receiving families to be a "rent-seeking" hypothesis (Hoxby, 1996). In economics, individuals or firms are said to seek rents when they try to use political action to secure benefits beyond what they would receive in a competitive marketplace. A rent-seeking home-based provider union would therefore be expected to manipulate political and regulatory processes to increase payments to unionized providers, and increase the proportion of the subsidized market served by unionized providers.

In this context, rent-seeking behavior might be manifest in two ways. First, home-based care provider unions could influence bureaucratic and regulatory processes to steer more subsidy-receiving families to choose home-based, as opposed to center-based, care for their young children. Larger numbers of children enrolling in unionized home-based care, as opposed to nonunionized center-based care, would generate increased income both for union members and, via compulsory membership dues, for the union itself. If this were true, we would expect to observe an increase in the percentage of infant and toddler care provided in home-based settings, as compared to center-based settings, following unionization.[1]

The second observable way HBCP unions might engage in rent-seeking would be to use their political power to increase the number of families who receive child care vouchers. Increasing the number of families who receive child care vouchers would provide new clients for child care providers that would likely, in turn, increase income for HBCPs and their union. In many communities, the numbers of families who are eligible to receive child care subsidies exceeds the number of subsidies available (Schulman & Blank, 2012). Prior work has indicated that receipt of a child care subsidy allows families to purchase higher quality care for their young children (Ryan et al., 2011). Receipt of child care subsidies has also been shown to facilitate parent employment and reduce economic hardship (Tekin, 2007). Therefore, although any increase in the percentage of infants and toddlers who received child care subsidies would represent an increased expense for states (and a corresponding benefit for labor unions and their members), it might also be interpreted as a positive effect of unionization.

The professionalization and rent-seeking hypotheses are not mutually exclusive, and it is possible for a given effect of unionization to be interpreted as supporting both theories. Were per-child payments to increase due to unionization, for example, the increased income to both HBCPs and the union as a whole could represent rents to the union and its members. However, it is also plausible that increasing provider wages could lead to improvements in the quality of care (Folbre, 2006). A study of early efforts to unionize HBCPs by Brooks (2005) suggested that unionization may improve the quality of care because the workplace issues that child care workers

[1] Although prior research has documented that home-based programs lag behind center-based programs on nearly all quality indicators, it is possible that the union-provided training and support services would help to lower the well-documented quality gap between center- and home-based child care programs. It is therefore not clear whether an increase in the percentage of subsidy-receiving children enrolled in home-based settings should be viewed as a negative effect of unionization.

experience—long hours, low wages, and no benefits—are directly connected to the quality of care they are able to provide. Higher compensation levels could therefore encourage more highly educated and better prepared workers to become or remain HBCPs and thus help to professionalize the HBCP workforce.

In this study, we investigated whether the impacts of HBCP unionization were consistent with the professionalization and rent-seeking hypotheses. Our specific research questions (RQs) were as follows:

Professionalization hypothesis:

RQ1: Did the Illinois home-based child care provider collective bargaining agreement increase the percentage of child care provided to Illinois subsidy-receiving infants and toddlers in licensed rather than license-exempt child care programs?

Rent-seeking hypothesis:

RQ2: Did the Illinois home-based child care provider collective bargaining agreement increase the percentage of child care provided to Illinois subsidy-receiving infants and toddlers in home-based, rather than center-based, child care programs?

Professionalization and rent-seeking hypotheses:

RQ3: Did the Illinois home-based child care provider collective bargaining agreement increase the percentage of Illinois infants and toddlers who received child care subsidies?

RQ4: Did the Illinois home-based child care provider collective bargaining agreement increase the dollar amount of payments to home-based providers, per subsidy-receiving child?

RESEARCH DESIGN

We employed a comparative case-study approach with a synthetic control group design as proposed by Abadie, Diamond, and Hainmueller (2010) as our primary method for addressing our research questions. Specifically, we compared the average values of the selected outcomes in Illinois in each of the months after January 2006—the date when the first contract between Illinois HBCPs and the state was ratified—to the corresponding average monthly values of the same outcome for a "synthetic" pre-2006 "Illinois." We constructed the monthly values of outcomes for this latter counterfactual condition by pooling data from states that did not permit child care workers to form unions, and weighting their contributions to generate a single "control group" state that was similar to Illinois during the pretreatment period. We checked the robustness of the findings produced by the synthetic control analyses using a difference-in-differences approach.

Data Sources

We drew our principal data from ACF state administrative records for subsidy-receiving families, for the years 2002 through 2008. These family- and child-level data were reported monthly on representative samples of families who received child care subsidies; they contain indicators of child age, and the type and amount of care used in a given month. We aggregated these family- and child-level indicators to the state level, prior to our principal analyses.

We also drew data from the monthly administrations of the CPS for the same time period, in order to incorporate additional information on the social, economic, demographic, and housing characteristics in the constituent states. The CPS provides monthly state-level data on a representative sample of individuals in each state. We obtained these data through the Integrated Public Use Microdata Series (King et al., 2010).

Analytic Sample

The ACF data sets contain monthly information on representative samples of families who received CCDF subsidies from January 2002 through September 2008. However, prior to generating the state-level monthly aggregate values of those variables that became the target of our subsequent data analyses, we limited our sample in two important ways. First, we excluded 12 states that enacted similar child care unionization laws during the study period. Because these states allowed some form of unionization, we could not consider them pure nonunion contributors to any synthetic control group condition. On the other hand, the substantial interstate differences in the scope and content of the laws enacted by each state also made it difficult to conceive of them as enacting replications of the Illinois law and becoming members of a potential treatment group.

We further limited our sample to include information on children who were between birth and 35 months only. We limited ourselves to this younger sample for three reasons. First, older school-aged children (five years and older) spend the majority of their days in free school-based settings, and receive child care subsidies only for care in home-based settings as a secondary form of care. Second, three- and four-year-old children, in some cases, also had access to free school-based care. In fact, during the period following unionization, Illinois as well as many of the nonunion comparison states expanded the availability of publicly funded center-based prekindergarten programs to three- and four-year-old children (Barnett et al., 2010). This increased availability could have altered the type and cost of CCDF care used by families of three- and four-year-old children, making it difficult to isolate a pure effect of unionization. Finally, we excluded two states, New Mexico and New Hampshire, for which there were substantial amounts of missing data on our outcomes of interest.

Measures

We organized our data in a "state-month" data set, where each row contains the values of the following outcomes, predictors, and covariates, aggregated to the state level for that month.

Outcomes:

- *Percent of hours in licensed care:* A continuous variable that recorded the percentage of total care hours that subsidy-receiving infants and toddlers (between birth and age 35 months) spent in licensed, as opposed to unlicensed, settings (both home-based and center-based), in the state in that month.
- *Percent of hours in home-based care*: A continuous variable that recorded the percentage of total care hours subsidy-receiving infants and toddlers (between birth and age 35 months) spent in a home-based setting, on average, in the state in that month.
- *Percent subsidy*: A continuous variable that recorded the percentage of children (between birth and age 35 months) in the state who received child care subsidies in that month.
- *Average monthly cost per child:* A continuous variable that recorded the average per-child amount (in U.S. dollars) paid to providers as compensation for caring for subsidy-receiving children in the state each month during the reporting period. This figure includes both contributions made by the state subsidy payment and any additional parent co-payments.

Question Predictors:

- *Treatment*: A dichotomous predictor to distinguish Illinois from the states used to constitute the synthetic comparison state (1 = Illinois, 0 = otherwise).
- *Time:* An ordinal predictor to record the month (and year) during which each cross-section of data was recorded (1 = January 2002 through 81 = September 2008). Waves 1 through 48 constitute the preintervention period, while waves 49 through 81 represent the postintervention period.
- In addition, we included a vector of selected state-level family, macroeconomic, data collection, and policy characteristics as covariates in our creation of the synthetic control group.

The family characteristics were the percentage of state residents who had graduated from college; the percentage of parents who were married; the percentage of state residents who identified as African American, Hispanic, or white; the percentage of state residents who were not U.S. citizens; and the percentage of state residents who lived in an urban setting. Macroeconomic characteristics used as covariates were the state median annual family income, the state unemployment rate, and the percentage of state residents who were members of a union. We also included as covariates an indicator of whether the state reported data on a sample, rather than the full population, of subsidy recipients, as well as a measure of the portion of the state's subsidy funds that were provided by the federal government. We selected these covariates used in the creation process based on theory and prior research describing the characteristics of recipients and the cost, type, or amount of child care used by, or available to, low-income families. For example, prior work has documented associations between families' child care preferences and parent race (Fuller, Holloway, & Liang, 1996), parent immigration status (Yoshikawa, 2011), family income (Early & Burchinal, 2002), parent marital status (Rose & Elicker, 2010), and mother's level of education (Hirshberg, Huang, & Fuller, 2005). There is considerably less research on how state-level characteristics are related to use of child care, but a 2008 review by Lippman and colleagues suggested that state-level aggregates of the above-cited characteristics, along with state policy and macroeconomic conditions, are key factors in explaining cross-state variation in the type, cost, and amount of child care that low-income families use. A detailed definition of these covariates is provided in the Appendix to this paper.[2]

DATA ANALYSES

Often, comparative case studies are used to evaluate the causal impact of large-scale interventions, such as policy changes or natural disasters, on aggregate units such as cities, states, or countries. With such research designs, the researcher compares the observed values of the outcome for a unit that experienced the intervention to the values of the same outcome observed in a unit or group of units that did not experience the intervention. In such a design, outcome data on the untreated unit(s) observed during the treatment period are assumed to provide a valid estimate of the values of the same outcomes that would have been observed for the treated unit(s) if they had not experienced treatment (i.e., the counterfactual).

The synthetic control group approach developed by Abadie, Diamond, and Hainmueller (2010) improves on the comparative case study method by providing a transparent analytic approach for constructing the counterfactual empirically. Rather

[2] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://onlinelibrary.wiley.com.

than choosing a single comparison unit (in this case, a state) to represent the counterfactual condition, the synthetic control group approach uses information on pretreatment outcomes and selected covariates to create an empirically weighted combination of potential comparison states (referred to as the "donor pool"), in a way that best approximates the treatment state, prior to treatment. These same weights are then applied in data on postintervention outcomes from the donor pool states, to estimate the values that relevant posttreatment outcomes would have had under the counterfactual condition. The weights, which are applied in the donor pool, show the degree to which each potential comparison unit contributes to the synthetically created control group, and thereby make explicit the degree to which the synthetic control group is, or is not, similar to the treated unit on preintervention characteristics hypothesized to be associated with the dependent variable.

The estimation of inferential statistics and testing of hypotheses is also conducted empirically, by conducting multiple "placebo" studies, to estimate the effects of a "pseudo-treatment" on an untreated state, compared to a corresponding synthetically generated control state, and then enumerating the empirical probability (*p*-value) that the actual observed treatment effect (or an effect that is larger than it) could have been produced by an accident of sampling from a population in which there was, by definition, no treatment effect. A limitation of this approach is that statistical power is limited by the number of available placebo states, making it difficult to reach the conventional 0.05 critical value.

We applied this approach to address our research questions in four steps. We document each of these steps briefly below. In the findings section that follows, we provide a detailed example of the process.

Step 1: Generating a Synthetic Control State

Separately for each dependent variable, we generated a weighted combination of states that, when pooled, provided an appropriate synthetic control group for that outcome. This was achieved, for each outcome, by estimating weights, W, that minimized a statistical "distance" criterion, W^*, as follows:

$$W = \sqrt{(X_1 - X_0 W)' V (X_1 - X_0 W)} \tag{1}$$

where vector X_1 contains pretreatment values of selected covariates observed in Illinois, vector X_0 contains values of the same characteristics for donor pool states, and V is a diagonal matrix with non-negative elements chosen to reflect the importance of each of the selected covariates in predicting variation in preintervention values of the dependent variable. Elements along the diagonal of V were determined by regressing values of the preintervention-dependent variable on the corresponding values of the selected covariates, and choosing values that reflected the importance of each covariate in that prediction. Weights W were estimated iteratively to minimize distance criterion W^* and, once obtained, were applied to generate a single weighted composite value of that outcome for the resulting synthetic Illinois, for each month pretreatment.

Step 2: Checking that the Synthetic Control State Replicates Pretreatment Illinois Data Well

We then summarized the degree to which our newly created synthetic Illinois approximated actual Illinois by estimating a root mean-squared error average difference (RMSE) in outcome values between the two during the pretreatment period as follows:

$$RMSE = \sqrt{\frac{\sum_t (\bar{Y}_t^I - \bar{Y}_t^S)^2}{T}} \tag{2}$$

where subscript t enumerates the time period ($t_1, \ldots, t_T$), with t_1 representing the first pretreatment observation period (January 2002), and t_{48} representing December 2005, the last observation before the HBCPs ratified their first collective-bargaining agreement. In the denominator, T is the total number of periods involved in the sum of squares (48, in this case). Thus, $\bar{Y}_t^I$ then represents values for the dependent variable observed in Illinois during time t and $\bar{Y}_t^S$ represents values for the dependent variable observed during time t in the synthetic comparison group, and the difference represents the extent to which the one matched the other. Then, squaring (to eliminate negative signs) and summing the squared differences provides a summary statistic whose magnitude captures the quality of the match between Illinois and synthetic Illinois, pretreatment. We used the estimated value of pretreatment RMSE from equation (2) to determine empirically the particular covariates included in the estimation of distance W^* in equation (1). If the inclusion of the specific covariate helped reduce the magnitude of pretreatment RMSE (and thus improved the capacity of pre-treatment $\bar{Y}_t^S$ to approximate pretreatment $\bar{Y}_t^I$), we retained the covariate in the estimation of W^*. We repeated this process until we had identified that set of covariates that, when combined, produced the smallest possible RMSE between Illinois and synthetic Illinois, in the pretreatment period.

As with other quasi-experimental methods, our ability to make causal claims from this analysis rests on the degree to which the synthetically derived control state represents an appropriate counterfactual for the treatment state. The synthetic control method relies heavily on the pretreatment values of covariates to determine distance W^*, weights W, and thereby the composition of the synthetic counterfactual unit. This makes it critical to first choose an appropriate set of covariates with which to model preintervention $\bar{Y}_t^I$.

Step 3: Estimating the Treatment Effect

To determine whether the treatment then caused differences subsequently between actual and synthetic Illinois, we first inspected a graphical display of monthly values for $\bar{Y}_t^I$ and $\bar{Y}_t^S$ for the entire time period, months 1 through 81 (September 2008), covering the entire pre- and posttreatment periods. If the treatment did indeed have an impact on the value of the respective dependent variable, values for $\bar{Y}_t^I$ and $\bar{Y}_t^S$ should be observably *similar* in the pretreatment period but observably *different* in the posttreatment period. We quantified any differences between actual and synthetic Illinois in each month by subtracting $\bar{Y}_t^S$ from $\bar{Y}_t^I$, and we refer to the resulting difference as $\bar{Y}_t^D$. We then estimated the overall average difference in the outcome between actual and synthetic Illinois before and after unionization by estimating the mean of all values for $\bar{Y}_t^D$ in the pretreatment period, hereafter referred to as $\overline{\bar{Y}_{pre}^D}$, and posttreatment period, hereafter referred to as $\overline{\bar{Y}_{post}^D}$. Finally, we estimated the overall population treatment effect, α, by subtracting the average difference between actual and synthetic Illinois in the posttreatment period $\overline{\bar{Y}_{post}^D}$ from the average difference between actual and synthetic Illinois in the pretreatment period $\overline{\bar{Y}_{pre}^D}$, to form estimate $\bar{\alpha}$. If there was a noticeable treatment effect, we would expect the value of $\bar{\alpha}$ to be large.

Step 4: Conducting Empirical Tests of Inference

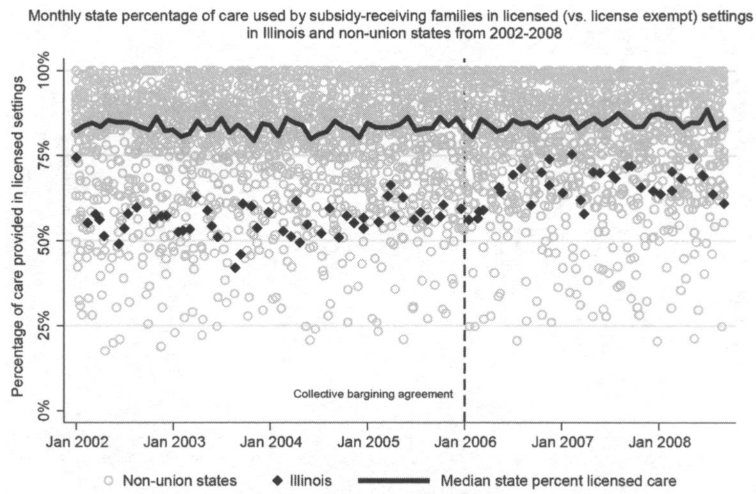
Finally, we tested whether the value of parameter α is different from zero, in the population, by conducting the series of placebo or falsification studies, in order to enumerate an empirical p-value. Specifically, we repeated the analytic process outlined above, treating each state that did not enact child care unionization as if it were the treatment state. Thus, we generated a new synthetic control group, one by one, for each of these nonunion pseudo-treatment or placebo states. The synthetic control unit for each placebo state is again composed of a unique combination of other states (excluding Illinois) that, when combined, produce a reasonable pretreatment match for that specific state. For each placebo state, we then estimated the ratio of pre- and posttreatment root mean-squared errors, and constructed a histogram to display the empirical distribution of these estimated ratios. This histogram then summarized empirically the range of possible pre–posttreatment differences that might have been observed at random, if there had been no effect of the treatment. Finally, we superimposed the magnitude of the actual post/pre-RMSE ratio obtained in the observed comparison of Illinois with its corresponding synthetic control group on the histogram of pseudo-values, and estimated from it the probability that a value this large, or larger, could have been obtained as a pseudo-value, as an idiosyncrasy of sampling from a null population. Thus, we generated an empirical p-value for the test of the null hypothesis that there was no treatment effect in the population.

RESULTS

Percent of Hours in Licensed Care

Our first research question asked whether HBCP unionization caused an increase in the percentage of care hours that subsidy-receiving children spent in licensed, as compared to license-exempt, child care settings. Such an increase would be interpreted as supporting the professionalization hypothesis. Illinois subsidy-receiving families were permitted to enroll their children in either licensed or license-exempt child care facilities during the study period. In Illinois, HBCPs who served three or fewer nonrelative children were permitted to forego licensure. This policy was consistent across the period under evaluation in this study.

In Figure 1, we display the percentage of care hours provided in licensed settings to subsidy-receiving children (aged birth to 35 months old) in Illinois and from donor pool states, from January 2002 through October 2008. Values for Illinois are displayed as black diamonds, values for all donor pool states are displayed as gray circles, and median values for the entire analytic sample (including Illinois and all states that did not implement some form of the treatment subsequently) are displayed as a solid black line. During the pretreatment period, the percentage of care hours provided in licensed settings in Illinois typically fell below the monthly median for nonunion states, and ranged from a low of 42.2 percent in August of 2003 to a high of 74.1 percent in January of 2002. In contrast, the percentage of care hours subsidy-receiving infants and toddlers spent in licensed settings in donor pool states ranged from a low of 16.4 percent for Hawaii in May of 2002 to a high of 100 percent in Arkansas, the District of Columbia, North Carolina, and Oklahoma in one or more months during the pretreatment period. States with values of 100 percent in one or more months during the pretreatment period represent a ceiling effect. That is, although these states could potentially contribute to a pretreatment synthetic control group, it would not be possible for the percentage of care provided in licensed settings to increase in these states during the posttreatment period. We therefore exclude these "ceiling" states from analyses of this dependent variable.



Note: State-Average Monthly Percentage of Care Hours Used by Subsidy-Receiving Families in Licensed (vs. License-Exempt) Settings in Illinois (Black Diamond) and Nonunion States (Circle), by Chronological Year, from 2002 through 2008. Median Value of the Percentage of Licensed Care Hours Across All Union and Nonunion States is Shown As a Solid Black Line.

Figure 1. Care Hours Used In Licensed Settings in Illinois and Nonunion States, 2002 through 2008.

Using the process outlined above, in our description of the data analyses, we generated—based on selected pretreatment covariates—a synthetic comparison group to approximate the trajectory of observed values of *Percent Licensed* in Illinois prior to treatment. The synthetic Illinois featured in this first comparison of average values of this dependent variable was a weighted combination of eight states. Hawaii was the largest contributor (37.1 percent), followed by contributions from California (18.1 percent), Montana (15.4 percent), Connecticut (15.2 percent), Texas (5.2 percent), Mississippi (4.3 percent), Missouri (0.7 percent), and Rhode Island (0.1 percent). During the pretreatment period, both actual and synthetic Illinois were quite similar in terms of the variables that were used as covariates in the estimation of the synthetic control group (see Tables A1 and A2).[3]

In Figure 2, we display the percentage of infant and toddler care hours provided in licensed settings in actual (solid black line) and synthetic Illinois (dotted gray line), prior to and following treatment. On average, subsidy-receiving Illinois infants and toddlers spent approximately 0.05 percentage points ($SD = 5.2$) fewer hours in licensed care when compared to children in the synthetic control group prior to the ratification of the collective-bargaining agreement. Following ratification, Illinois infants and toddlers spent an average of 6.4 ($SD = 5.5$) percentage points more hours in licensed settings when compared to children in the synthetic control group.

The 6.4 percentage point average posttreatment difference between Illinois and its synthetic control group was sizable (approximately 1.2 standard deviations of the observed values of percentage of care hours provided in licensed settings in Illinois

[3] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://onlinelibrary.wiley.com.

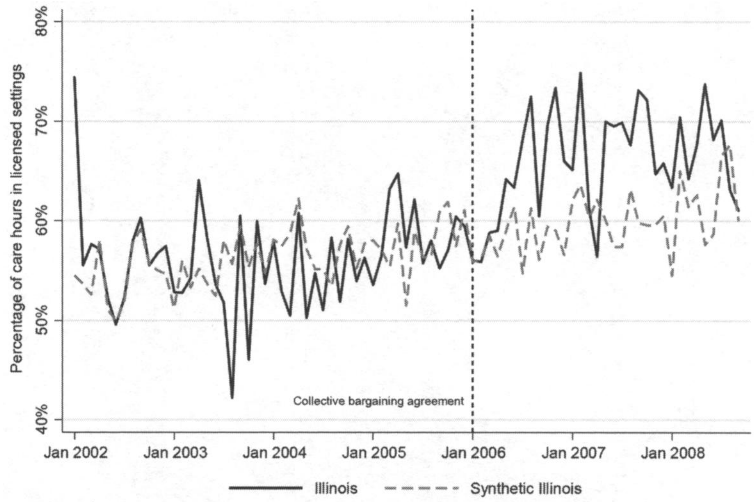


Figure 2. State-Average Percentage of Care Hours in Licensed (vs. License-Exempt) Child Care Programs, Before and After Unionization in Illinois (Solid Line) and Synthetic Illinois (Dashed Line), by Chronological Year, from 2002 through 2008.

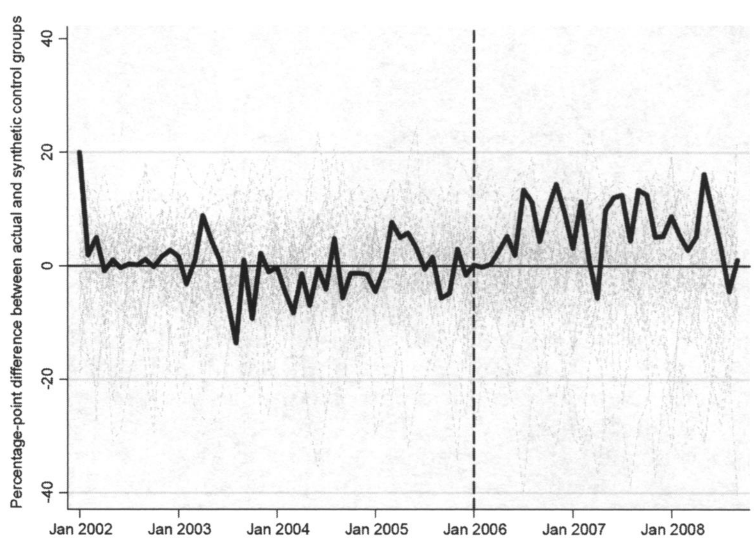


Figure 3. Percentage-Point Differences between Actual and Synthetic Control Groups for Illinois (Solid Black Line) and Placebo States (Dotted Gray Lines), by Chronological Year, from 2002 through 2008.

during the pretreatment period). However, it is possible that such differences could have been observed by chance, even with no effect of unionization. Thus, as described earlier, we conducted a series of placebo tests to estimate an empirical p-value, in order to test the null hypothesis that this difference was zero, in the population. In Figure 3, we display the differences observed between actual and synthetic control for Illinois (displayed as a solid black line) and each of the placebo states (displayed as dotted gray lines) during the study period. Notice

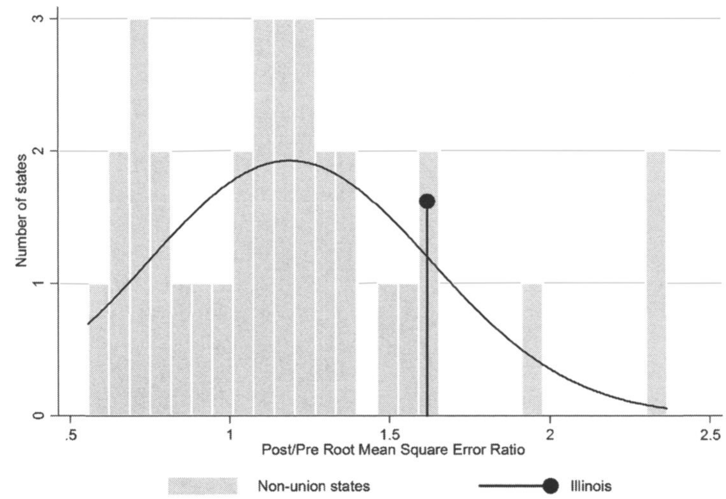


Figure 4. Sample Distribution of the Ratio of Post- to Pretreatment Root Mean-Square Errors for Illinois (Vertical Black Line) and the Placebo Study States (Histogram and Smoothed Envelope).

the considerable scatter in the placebo profiles, and contrast that scatter with the profile obtained in the primary analysis.

For each state, we then estimated the root mean-squared error summarizing the differences in outcome over months, between actual Illinois and its synthetic counterpart separately for the pre- and posttreatment periods. We then estimated the *ratio* of posttreatment RMSE to pretreatment RMSE for each, and display the distribution of these ratios as a histogram in Figure 4. Of the 33 comparisons made, the comparison involving actual versus synthetic Illinois registered the fifth largest RMSE ratio. Thus, enumerating across the figure, there is a 0.12 probability (empirical p-value) that differences between Illinois and its synthetic control were observed by chance, in comparisons in which there were really no such differences. We therefore come close to rejecting this null hypothesis empirically, at conventional levels of Type I error, and being able to declare that actual Illinois displayed a higher percentage of care hours provided in licensed settings, on average, postunionization.[4]

As noted above, values for *Percent Licensed* for Illinois ranged in the pretreatment period from a low of 42.2 percent in August 2003, to a high of 74.1 percent in January 2002. This outlying January 2002 value (the first occasion of observation, for this

[4] It is notable that although, on average, the percentage of care hours provided in licensed settings was higher in Illinois than in its synthetic counterfactual in the post-treatment period, there are two months, April 2007 and August 2008, in which the percentage of care hours provided in licensed settings in Illinois was lower than the corresponding percentage in the synthetic control group. This negative difference was accounted for in estimating the average difference between actual and synthetic Illinois, but the formula for deriving the empirical p-value is based only on the absolute value of the difference. To test the sensitivity of inferences to the presence of these negatively signed differences, we smoothed these month-to-month differences by averaging the monthly values within a four-month window. We then replicated the analyses using these smoothed values (for which the difference between actual and synthetic Illinois was consistently positive in the post-treatment period), and obtained the same empirical p-value of 0.12.

study) was more than two standard deviations larger than the other values observed in Illinois during the pretreatment period. Later, we found similarly extreme values reported in Illinois during this same time period for additional dependent variables *Percent Home-Based* (RQ2) and *Average Monthly Cost per Child* (RQ4). The consistent presence of aberrant values in this time period suggests the possibility that these values may be the result of data entry error in the ACF data files, rather than something unique about the characteristics of subsidized care used in Illinois during January 2002. We therefore tested the sensitivity of our findings to the removal of this data point. That is, we replicated each step of the analyses described above using February 2002, rather than January 2002, as the beginning of the pretreatment period. These additional sensitivity analyses yielded a synthetic control group that provided a better fit to actual Illinois pretreatment, and led to an estimated treatment effect that was larger and for which the empirical p-value was lower. Specifically, removing the potentially aberrant data in January 2002 indicated that unionization caused a 7 percentage point higher share of subsidized infant and toddler care to be provided in licensed settings (compared to 6.4 percentage points in the formal comparison). Placebo tests based on the reduced pretreatment period revealed that we would expect to find an effect of treatment as large as, or greater than, the effect observed in Illinois only 3 percent of the time ($p = 0.03$), if there were in fact no treatment effect. Thus, based on the results from analyses using both the full and truncated pretreatment time series, we conclude that unionization caused an increase in the percentage of care hours Illinois infants and toddlers were cared for in licensed, as compared to license-exempt, settings by between 6.4 and 7 percentage points.

Percent of Hours in Home-Based Care

Our second research question asked whether HBCP unionization caused an increase in the percentage of care hours subsidy-receiving children spent in home-based care, as compared to center-based care. Since all home-based care was provided by persons unionized by the policy of interest, and center-based care was provided by persons who were not unionized by the policy, an increase would be interpreted as supporting the rent-seeking hypothesis. Home-based care was relatively common in Illinois compared to in nonunion states during the pretreatment period. For instance, from 2002 through 2005, between 58 percent and 76 percent of the care hours provided to Illinois subsidy-receiving infants and toddlers were provided in home-based settings. The monthly median percentage of home-based care received by infants and toddlers in donor pool states during the pretreatment period ranged from 30 percent to 43 percent. We display these summaries, as above, in the upper left corner of Figure 5, with the percentages for Illinois displayed as black diamonds, values for donor pool states displayed as gray circles, and median values for the analytic sample displayed as a solid black line.

We then followed the same analytic process as described above for our first outcome, and again generated a synthetic control group to contrast with Illinois, using data on the set of pretreatment characteristics that yielded the best possible pretreatment fit between Illinois and a synthetically produced composite of nonunion states. The resulting synthetic control group for RQ2 comprised four states: Connecticut (42 percent), Hawaii (33 percent), West Virginia (19 percent), and Utah (6 percent). This synthetic Illinois was similar to actual Illinois on some characteristics, such as the percentage of families with young children in which the parents were married and the percentage of workers in the state who were members of a union or covered

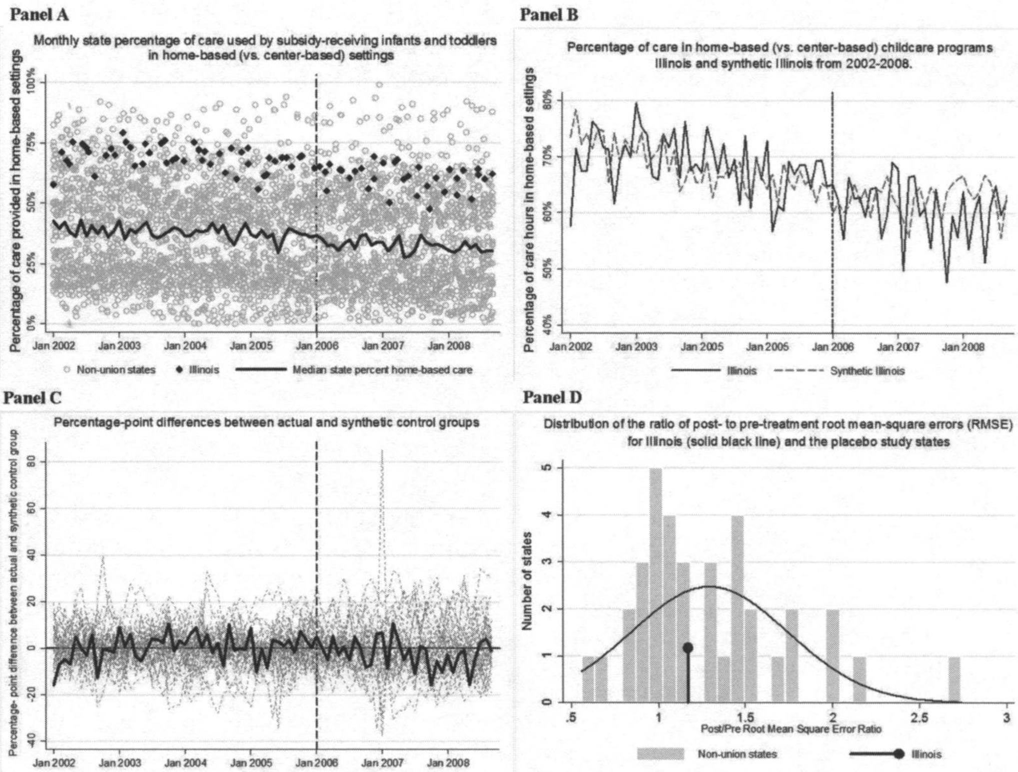


Figure 5. Graphical Summary of the Critical Components of the Synthetic Control group Analysis of the Impact of Unionization on the Percentage of Care Hours Used by Subsidy-Receiving Infants and Toddlers in Unionized Home-Based, Compared to Nonunionized Center-Based, Settings (RQ2).

by a union contract, but notably dissimilar on other characteristics such as median family income and unemployment rate (see Tables A3 and A4).[5]

We display values of *Percent Home-Based* for Illinois and its corresponding synthetic control group in the upper right panel of Figure 5. During the pretreatment period, on average, Illinois infants and toddlers spent 0.5 ($SD = 5.8$) percentage points more time in home-based settings than did the infants and toddlers in the synthetic counterfactual. During the posttreatment period, by contrast, Illinois infants and toddlers spent 2.4 ($SD = 6.4$) percentage points fewer care hours in home-based settings compared to the synthetic counterfactual. Although this average posttreatment difference is sizable, it is important to note that, as evidenced by the relatively large standard deviation, the differences between actual and synthetic Illinois varied substantially from month to month and did not, upon visual inspection of the graph displayed in the bottom right panel of Figure 5, provide compelling and clear evidence of a causal change between the pre- and posttreatment periods.

[5] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://onlinelibrary.wiley.com.

Enumerating summary statistics from placebo testing (displayed in the bottom left panel of Fig. 5) indicates that differences greater than those observed between pre- and posttreatment Illinois and their synthetic controls could be observed by chance 50 percent (18/36) of the time in situations in which there was no intervention. Thus, we could not reject the null hypothesis that there was a causal change in the percentage of care subsidy-receiving infants and toddlers in unionized home-based, as compared to nonunionized center-based, settings postunionization. As with RQ1, we conducted sensitivity analyses in which we omitted the potentially aberrant January 2002 data. Our sensitivity analyses again produced a better pretreatment fit between actual and synthetic Illinois, and lead to a substantially smaller estimate of posttreatment differences, of –0.4 percentage points. Placebo testing once again suggested a high probability (p-value $= 0.62$) that these effects were observed by chance, confirming our inability to reject the corresponding null hypothesis.

Percent Subsidy

For our third research question, we asked whether unionization led to an increase in the percentage of Illinois infants and toddlers (aged birth to 35 months old) who received child care subsidies. States have substantial flexibility to expand or contract the size of their subsidy-receiving populations. An increase in the percentage of children in the population who received subsidies would be interpreted as lending support to the rent-seeking hypothesis. Across the country, a small percentage of infants and toddlers received child care subsidies in a given month during the study period. The percentage of Illinois infants and toddlers who received child care subsidies was similar to the median value for other states in the analytic sample, ranging between a low of 3.2 percent in February 2002 and a high of 5.4 percent in February 2004.

The synthetic comparison group was composed of Rhode Island (39.1 percent), Connecticut (19.6 percent), Oklahoma (6.7 percent), Arizona (4.2 percent), West Virginia (4.0 percent), and small contributions (less than 2 percent) from 30 other states (additional details in Tables A5 and A6).[6] The monthly average percentages of Illinois infants and toddlers who received subsidies were similar in the pretreatment (4.1 percent, $SD = 0.4$) and posttreatment (3.8 percent, $SD = 0.3$) periods. The percentage of infants and toddlers in the synthetic control group who received subsidies by contrast increased sharply from 4.1 percent ($SD = 0.2$) during the pretreatment to 4.5 percent ($SD = 0.3$) in the posttreatment period. Thus, although Illinois served a slightly lower percentage of its infants and toddlers through the subsidy system following unionization, the synthetic control group served a higher percentage of these children during this period. This suggests that unionization led to 0.7 percentage points fewer Illinois children using child care subsidies.

We again conducted a set of placebo tests to estimate the empirical probability that the pre-to-posttreatment differences observed in Illinois might have occurred by chance, when there was really no treatment effect. The results of these analyses, displayed in the bottom right and bottom left of Figure 6, show that three of the 35 placebo comparisons (8 percent) yielded pre-to-post-differences that were greater in magnitude than those observed in Illinois. Thus, the observed treatment effect of 0.7 percentage points fewer Illinois infants and toddlers using child care subsidies, or a value more extreme, could have been observed 8 percent of the time when there was no effect of treatment ($p = 0.08$).

[6] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://onlinelibrary.wiley.com.

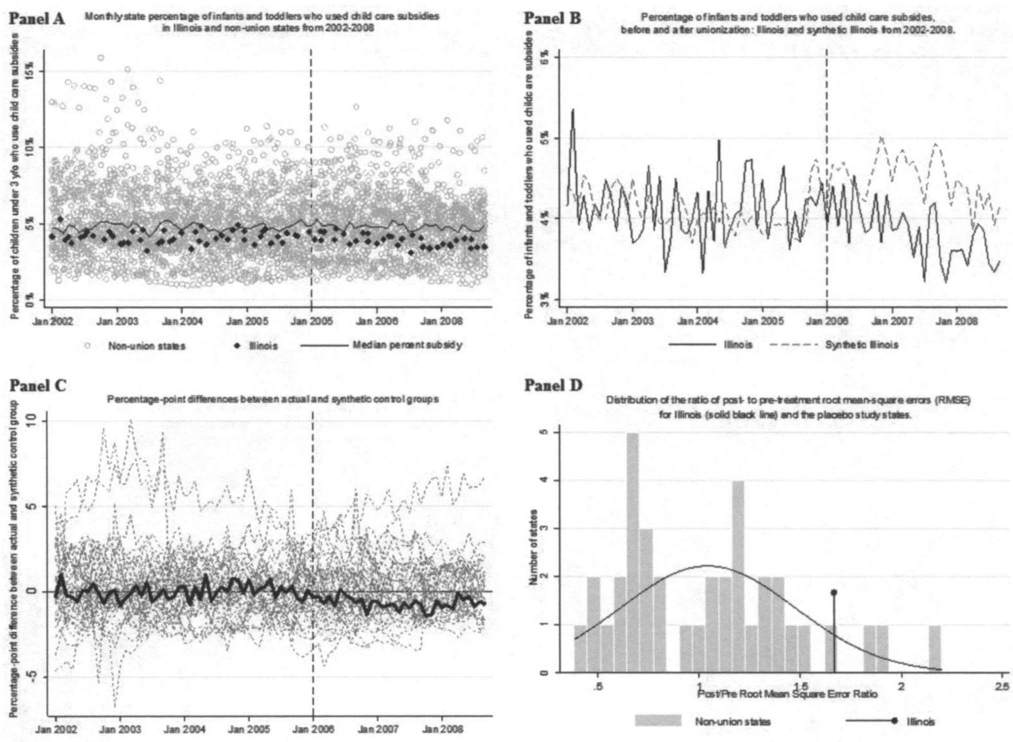


Figure 6. Graphical Summary of the Critical Components of the Synthetic Control group Analysis of the Impact of Unionization on the Percentage of Infants and Toddlers Who Use Child Care Subsidies (RQ3).

Although none of the values of the outcome *Percent Subsidy* were particularly unusual for this time period, we carried out the sensitivity analyses for this dependent variable for the sake of consistency with the analyses conducted on the other outcomes. As with the other research questions, the removal of January 2002 data generated a synthetic control group that better fit the data in actual Illinois. Analysis without the January 2002 data indicated that following unionization, 1.1 percentage points fewer Illinois infants and toddlers received child care subsidies relative to the infants and toddlers in the synthetic Illinois group ($p \leq 0.000$). We therefore conclude that unionization led to between 0.7 and 1.1 percentage points fewer infants and toddlers using child care subsidies in Illinois ($p = 0.08, p = 0.000$). Note that these compelling findings run counter to the rent-seeking hypothesis that motivated this question. If unions were to engage in rent-seeking, we would expect the percentage of the population who received subsidies to increase, rather than decrease.

Average Monthly Cost per Child

Our fourth research question asked whether unionization led to an increase in the monthly average per-child dollar amount paid to child care providers in Illinois. As noted above, such an increase could be interpreted as lending support to either the professionalization or rent-seeking hypotheses. Per-child payments in Illinois ranged from \$357 to \$473 during the pretreatment period, and were similar to the

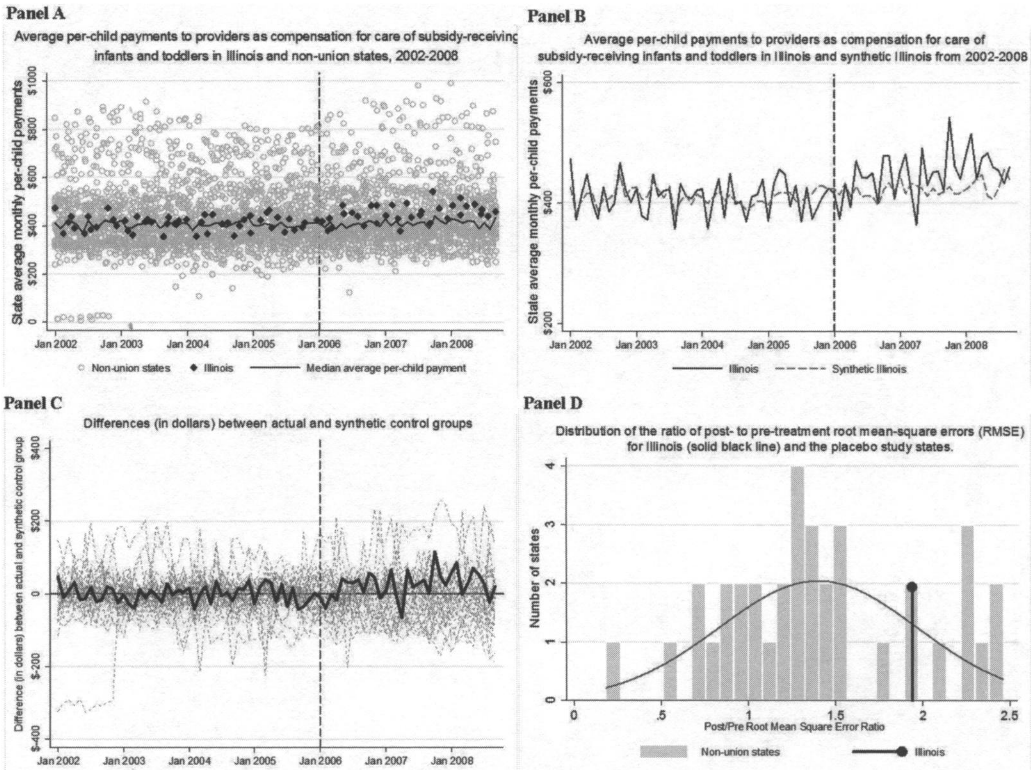


Figure 7. Graphical Summary of the Critical Components of the Synthetic Control group Analysis of the Impact of Unionization on the Average Monthly Per-Child Payments Provided as Compensation for the Care of Subsidy-Receiving Infants and Toddlers (RQ4).

median per-child payments for the other states in the analytic sample during this period. We again generated a synthetic control group to compare with Illinois, using data on the set of covariates that provided the best possible fit between actual and synthetic Illinois during the pretreatment period. The resulting synthetic Illinois was similar to actual Illinois on the pretreatment characteristics used in the formation of the composite (see Tables A7 and A8).[7]

The results of the analyses of average monthly cost per child are displayed in Figure 7. During the pretreatment period, average per-child payments in Illinois and its synthetic control were identical ($408 per child). These payments increased in both actual and synthetic Illinois in the posttreatment period, but this increase was sharper in actual Illinois, where per-child payments ($449) were, on average, approximately $26 higher than per-child payments in the synthetic control group ($423). Placebo tests using the 36 states from the donor pool produce treatment effects larger than those observed in Illinois in 19 percent (7/36) of the cases ($p = 0.19$).

[7] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://onlinelibrary.wiley.com.

As above, we conducted additional sensitivity analyses excluding the January 2002 data, and again found that this yielded a better pretreatment fit between actual and synthetic Illinois. Subsequent analyses yielded similar estimates of the unionization impact (approximately \$26) and inference statistics ($p = 0.16$).

Thus, although the effects of unionization on the average monthly per-child payments failed to meet standard thresholds for statistical significance using either the full or reduced pretreatment time series, the magnitude of these observed differences suggests some substantive importance. Over the course of a year, an additional \$26 per child per month would amount to an additional \$312 in HBCP income for each child served. Were an HBCP to serve multiple subsidy-receiving children, this more than 6 percent increase in average per-child payments would have a sizable impact on the provider's annual income. These results therefore suggest that unionization may have led to a considerable increase in the monthly average amount of money paid for the care of subsidy-receiving children, which would potentially be consistent with both the professionalization and rent-seeking hypotheses. However, our research design lacks adequate power to establish that this result is statistically significant at conventional levels.

Sensitivity Analyses Using Difference in Difference

We tested the robustness of these results by using a regression-based difference-in-differences approach to estimate the impact of unionization in Illinois on our four outcomes of interest. More specifically, we estimated models that used state fixed effects and state-specific time trends to compare the changes in each of these outcomes in Illinois at the time of unionization to the changes in all nonunion states, while controlling for relevant covariates. We clustered standard errors at the state level in order to account for the serial correlation of observations from the same state.

The results of these analyses, displayed in Table 1, yielded estimates of the impact of unionization that were generally similar to those produced using the synthetic control group approach. For example, these analyses suggest unionization led to a 0.06 percentage point increase in the percent of care used in licensed vs. licensed-exempt settings, a decrease of 0.7 percentage points in the percentage of young children who used child care subsidies, and an increase of \$17.61 in the average per-child dollar amount paid to child care providers in Illinois. Contrary to the findings from the synthetic control group approach, the difference-in-difference

Table 1. Displaying results of difference-in-difference regressions of the impact of unionization on research questions 1–4.

	Percent of hours in licensed care	Percent of hours in home-based care	Percent subsidy	Average monthly cost per child
Illinois	–0.01	0.09**	–0.23	–294.44***
	(0.07)	(0.04)	(1.02)	(52.36)
Postunionization	<0.00	0.01	0.42***	5.47
	(0.01)	(0.01)	(0.14)	(7.38)
Illinois × postunionization	0.06***	–0.05***	–0.72***	17.61**
	(0.01)	(0.01)	(0.17)	(8.05)

Note: Models include full set of covariates, state fixed effects, stateby-time interactions, and clustered standard errors.
**$P < 0.1$; **$P < 0.05$; ***$P < 0.01$.

analyses indicated that unionization led to a decrease of 0.05 percentage points in the percentage of care in home-based, rather than center-based settings. The difference-in-differences analyses yielded notably smaller p-values than were produced using the placebo study approach to inference employed with the synthetic control approach, with statistically significant results observed for each outcome. However, the standard errors produced by difference-in-differences models such as those reported in Table 1 account only for uncertainty in the aggregate data used to measure the outcomes of interest. As Abadie, Diamond, and Hainmueller (2010, pp. 496–497) note, "In comparative case studies, an additional source of uncertainty derives from ignorance about the ability of the control group to reproduce the counterfactual of how the treated unit would have evolved in the absence of the treatment." Consistent with this observation, we note that using nonunion states in the Midwest rather than all nonunion states as the comparison group yields treatment effect estimates that, in some cases, differ substantially in magnitude (though not direction) from those reported in Table 1. The advantage of the synthetic control group approach we rely on in our preferred analyses is that it enables the researcher to use a data-driven approach to select the most appropriate set of control observations and conduct inferential tests that account for uncertainty in this choice.

DISCUSSION

Our findings indicate that, in Illinois, HBCP unionization led to a higher proportion of hours subsidy-receiving infants and toddlers spent in licensed, rather than license-exempt, child care; a smaller proportion of the population of Illinois infants and toddlers using child care subsidies; and, potentially, a larger amount of monthly per-child compensation paid for the care of subsidized infants and toddlers (RQ4). We did not find any effects of unionization on the percentage of care used in the unionized home-based, compared to the nonunionized center-based, settings.

Our analyses for RQ1 indicate that unionization led to between a 6.4 and 7 percentage point increase in the percentage of subsidized infant and toddler care hours that were provided in licensed versus license-exempt settings. This finding supports the professionalization hypothesis, and suggests some overall positive impact of unionization. Although these analyses cannot provide a single explanation of the mechanisms by which this change occurred, it is possible that union membership offered child care providers the assistance and support necessary to make them eligible for licensure. In a set of interviews conducted in preparation for this study (Grindal, 2010), HBCP union members and union organizers reported that many home-based providers were eager to obtain licensure, but were concerned about what some perceived to be unfair and inconsistent applications of the associated state regulations. The union, they suggested, would help to facilitate this process both by offering providers support in completing the steps necessary for licensure, and by giving providers the confidence that the union stands ready to offer support for individual providers in disputes with the state that might arise as a consequence of licensure. It is also possible that financial incentives included in the 2006 collective bargaining agreement contributed to this increase in the use of licensed care. The collective bargaining agreement offered home-based providers between 5 percent and 20 percent higher rates for caring for subsidy-receiving children as a reward for meeting a set of benchmarks for program quality. Submitting for licensure represented the first step toward receiving these higher rates of compensation.

Our second key finding is that child care unionization led, on average, to between 0.7 and 1.1 percentage points fewer Illinois infants and toddlers using subsidies to purchase child care. This finding runs counter to the direction of effects anticipated by the rent-seeking hypothesis. The rent-seeking perspective suggested that unions would use their political power to encourage the state to expand the size of the subsidized child care market, as this would increase the number of children using unionized child care providers, and thus increase union revenue and membership. In interpreting this finding it is notable that the percentage of infants and toddlers who used child care subsidies did increase slightly in Illinois following unionization. What we here interpret as an effect of treatment is therefore primarily attributable to the larger increase in the percentage of subsidy-using infants observed in the synthetic control state. That is, it is not that Illinois reduced the size of its subsidy-receiving population following unionization, but rather that Illinois did not keep pace with the expansion of child care subsidies taking place in its synthetically generated counterfactual state. Further, since eligibility for child care subsidies in Illinois during the postunionization period was set at approximately 50 percent of the state median family income, and during that period there is no evidence that eligible families were placed on waiting lists (see Schulman & Blank, 2007), it is likely that those families who would have otherwise received subsidies were those with greater financial means to purchase care without the subsidy.

Our findings suggest that unionization might compel states to make a choice between the amount of per-child funds paid to providers and the number of children served. As a result of the collective bargaining agreement, Illinois child care providers earned health care benefits and subsidy rate increases that cost the state nearly $37 million over the three years covered by the agreement (Blank, Campbell, & Entmacher, 2010). It is possible that these increased costs influenced Illinois policymakers to forgo the expansion of access to subsidies that occurred during this period in some other states. This trade-off between providing higher rates of compensation for subsidized care and serving a larger number of children is a frequent focus of child care policy discussions. It appears that in Illinois, unionization may have led the state to encourage higher subsidy rates over expanded subsidy access.

We do not find strong evidence that unionization led to an increase in the average per-child amount of money provided for the care of subsidized infants and toddlers. The January 2006 Illinois collective bargaining agreement provided Illinois HBCPs with a series of subsidy-rate increases and incentives by which they could earn additional income. Following unionization, on average, the mean monthly per-child payment provided for the care of subsidy-receiving Illinois children was $26 higher than the mean per-child payment for children in the synthetic control group. Over the course of a year, an additional $26 per child per month would represent $312 in income for an HBCP. Were such a provider to serve multiple subsidy-receiving children, this more than 6 percent increase in average per-child payments could have a sizable impact on the provider's annual income. Although these represent substantial earnings increases for home-based providers, placebo tests find treatment effects as large as or larger than this between 19 percent and 16 percent of the time.

It is notable that we do not find consistent evidence that unionization increased the percentage of subsidized child care hours provided in home-based as compared to center-based settings. Specifically, we do not find any evidence that unionized home-based care in any way crowded out the nonunionized center-based portion of the subsidized child care sector. This runs counter to some prior research on public sector unions (e.g., Zax & Ichniowski, 1988) and may be attributable to the unique mixed-market setting in which providers of subsidized child care operate.

Although we observed generally similar findings when we examined these data using synthetic control and difference-in-differences approaches, we choose to feature the synthetic control group findings for the following reasons. First, the synthetic control group approach provides a transparent empirically driven method for constructing the second difference, while the traditional difference-in-difference approach relies on the subjective judgment of the researcher regarding what unit or combinations of untreated units represent a legitimate counterfactual. Indeed, we observed that the magnitude of the impact of unionization estimated using difference in differences varied substantially across multiple plausible compositions of the second difference. Second, we are concerned that the standard errors produced by the difference-in-differences models account only for uncertainty in the aggregate data used to measure the outcomes of interest and are thus inappropriately small. We believe that the placebo study method provides a more useful approach to putting the observed effects of unionization in the context of the possible pre–posttreatment differences that might have been observed at random, if there had been no effect of the treatment.

These findings come with some important caveats. As noted above, these analyses do not permit us to follow either specific children or specific child care providers over time. Rather, what we observe are the results of choices made by samples of subsidy-receiving parents each month, on average. This makes it impossible for us to definitively determine whether the observed increase in the percentage of care provided in licensed rather than unlicensed settings was driven by previously unlicensed providers submitting for licensure, or by subsidy-receiving parents buying licensed care at higher rates than they did prior to unionization. Similarly, we cannot say for certain what mechanisms drove the postunionization decrease in the percentage of Illinois children who used child care subsidies. The availability of subsidies is a function of myriad factors including the child care needs of individual parents, the fiscal health of state budgets, and the priorities of policymakers. In this study, we tested and incorporated a range of family, policy, and macroeconomic indicators, aggregated to the state level, in generating the counterfactual, but it is possible that some other unobserved characteristic might account for what we here observe as treatment-related differences. Finally, for our findings regarding the effect of unionization on provider compensation, it is important to reiterate that payments for subsidized child care in Illinois include both the state-provided subsidy and a parent copayment. The size of these copayments is subject to an unaccounted-for set of factors such as the local demand for home-based child care, the number of children cared for by a given provider, or the availability of discounts for enrolling multiple siblings in a given setting. The observed postunionization increase is therefore a function of the collective bargaining agreement as well as these other market factors. Further quantitative or qualitative research on parental and child care provider decisionmaking could further elucidate the mechanisms by which unionization led to findings reported in this paper.

In recent years, policymakers have engaged in a series of spirited debates regarding the efficacy of public sector labor unions. Although child care unionization has inspired a substantial amount of heated discussion, this study represents, to the best of our knowledge, the first quantitative empirical examination of how unions have an impact on important aspects of the child care marketplace. Our results suggest that unionization in Illinois resulted in better regulated but less available subsidized care for low-income families. Although these findings are specific to Illinois, policymakers considering permitting child care workers to unionize should therefore proceed with caution, and be careful to ensure that benefits to union members do not come at the expense of reduced access to subsidy among low-income families.

TODD GRINDAL is an Associate at Abt Associates, 55 Wheeler Street, Cambridge, MA 02138 (e-mail: Todd_Grindal@abtassoc.com).

MARTIN R. WEST is an Associate Professor of Education at Harvard University, Graduate School of Education, Gutman Library 454, 6 Appian Way, Cambridge, MA 02138 (e-mail: martin_west@gse.harvard.edu).

JOHN B. WILLETT is the Charles William Eliot Research Professor of Education at Harvard University, Graduate School of Education, Gutman Library 454, 6 Appian Way, Cambridge, MA 02138 (e-mail: john_willett@gse.harvard.edu).

HIROKAZU YOSHIKAWA is the Courtney Sale Ross Professor of Globalization and Education at New York University, Department of Applied Psychology, NYU Steinhardt, 726 Broadway, 5th Floor, New York, NY 10003 (e-mail: hy2042@nyu.edu).

ACKNOWLEDGMENTS

This work was funded in part through support from the Center on the Developing Child at Harvard University and the Office of Planning, Research & Evaluation, Administration for Children and Families, U.S. Department of Health & Human Services Grant # 90YE0139. Special thanks to Hadas Eidelman, Avi Feller, Jacob Klerman, Andres Molano, Matthew Ranson, and Jaime Thomas for their helpful comments on earlier drafts of this paper.

REFERENCES

Abadie, A., Diamond, A., & Hainmueller, J. (2010). Synthetic control methods for comparative case studies: Estimating the effect of California's tobacco control program. Journal of the American Statistical Association, 105, 493–505.

Administration for Children and Families. (2011). Child Care and Development Fund ACF-800 State-Level Data Standards: Technical Bulletin #1 r-v3, Appendix A. Washington, DC. Retrieved February 9, 2015 from http://www.acf.hhs.gov/sites/default/files/occ/tb01r_3_appendix_a.pdf.

Barnett, W. S., Epstein, D. J., Carolan, M. E., Fitzgerald, J., Ackerman, D. J., & Friedman, A. H. (2010). The state of preschool 2010. New Brunswick, NJ: NIEER.

Belsky, J., Vandell, D. L., Burchinal, M., Clarke-Stewart, K. A., McCartney, K., & Owen, M. T. (2007). Are there long-term effects of early child care? Child Development, 78, 681–701.

Blank, H., Campbell, N., & Entmacher, J. (2010). Getting organized: Unionizing home-based child care providers. Washington, DC: National Women's Law Center. Retrieved February 9, 2015 from http://www.nwlc.org/resource/getting-organized-unionizing-home-based-child-care-providers-2010-update.

Boushey, H. (2003). Who cares? The child care choices of working mothers. Data Brief No. 1. Retrieved February 9, 2015 from http://www.cepr.net/documents/publications/child_care_2003.htm.

Branchflower, D., & Bryson, A. (2004). What effect do unions have on wages now and would Freeman and Medoff be surprised? Journal of Labor Research, 25, 383–414.

Brooks, F. P. (2005). New turf for organizing: Family child care providers. Labor Studies Journal, 29, 45–64.

Brueckner, J. K., & Neumark, D. (2011). Beaches, sunshine, and public-sector pay: Theory and evidence on amenities and rent extraction by government workers. NBER Report no. w16797. Cambridge, MA: National Bureau of Economic Research.

Burris, G. (2012). Training makes a difference: The experience of unionized family, friend, and neighbor child care providers in Washington State. Economic Opportunity Institute. Retrieved February 9, 2015 from http://www.eoionline.org/wp/wp-content/uploads/early-learning/TrainingMakesaDifference-Feb2012.pdf.

Burton, A., Whitebook, M., Young, M., Bellm, D., Wayne, C., Brandon, R. N., & Maher, E. (2002). Estimating the size and components of the US child care workforce and caregiving population. Key findings from the child care workforce estimate. Preliminary report. Washington, DC: Center for the Childcare Workforce.

Capizzano, J., & Adams, G. (2004). Children in low-income families are less likely to be in center-based child care. Washington DC: Urban Institute. Retrieved February 9, 2015 from http://www.urban.org/publications/310923.html.

Child Care Aware. (2012). Parents and the high cost of child care. Washington, DC. Retrieved February 9, 2015 from http://www.naccrra.org/sites/default/files/default_site_pages/2012/cost_report_2012_final_081012_0.pdf.

Christakis, D., & Garrison, M. (2009). Preschool-aged children's television viewing in child care settings. Pediatrics, 124, 1627–1632.

Dowsett, C. J., Huston, A. C., Imes, A. E., & Gennetian, L. (2008). Structural and process features in three types of child care for children from high and low income families. Early Childhood Research Quarterly, 23, 69–93.

Early, D. M., & Burchinal, M. R. (2002). Early childhood care: Relations with family characteristics and preferred care characteristics. Early Childhood Research Quarterly, 16, 475–497.

Folbre, N. (2006). Demanding quality: Worker/consumer coalitions and "high road" strategies in the care sector. Politics & Society, 34, 11–32.

Freeman, R. (1984). Unionism comes to the public sector. NBER Working Paper no. 1452. Retrieved February 9, 2015 from http://www.nber.org/papers/w1452.pdf.

Freeman, R. B. (1986). The effect of the union wage differential on management opposition and union organizing success. The American Economic Review, 76, 92–96.

Freeman, R. B., & Medoff, J. L. (1979). The two faces of unionism. The Public Interest, 57, 69–93.

Freeman, R. B., & Rogers, J. (1999). What workers want. New York: Cornell University Press and Russell Sage Foundation.

Fuller, B., Holloway, S. D., & Liang, X. (1996). Family selection of child-care centers: The influence of household support, ethnicity, and parental practices. Child Development, 67, 3320–3337.

Fuller, B., Kagan, S., & Loeb, S. (2004). Child care quality: Centers and home settings that serve poor families. Early Childhood Research Quarterly, 19, 505–527.

Gennetian, L. A., Crosby, D. A., Huston, A. C., & Lowe, E. D. (2004). Can child care assistance in welfare and employment programs support the employment of low-income families? Journal of Policy Analysis and Management, 23, 723–743.

Gormley, W. T. (2007). Early childhood care and education: Lessons and puzzles. Journal of Policy Analysis and Management, 26, 633–671.

Greenberg, M., Lombardi, J., & Schumacher, R. (2000). The child care and development fund: An overview. Washington, DC: Center for Law and Social Policy. Retrieved February 9, 2015 from http://www.clasp.org/admin/site/publications/files/0064.pdf.

Grindal, T. (2010). Perceptions of child care labor unions: A pilot survey study. Unpublished paper. Cambridge, MA: Abt Associates.

Heckman, J. (2006). Skill formation and the economics of investing in disadvantaged children. Science, 312, 1900–1902.

Hernandez, D. J., Denton, N. A., & Macartney, S. E. (2008). Children in immigrant families: Looking to America's future. Social Policy Report, 22, 3–22.

Hirshberg, D., Huang, D. S. C., & Fuller, B. (2005). Which low-income parents select child care? Family demand and neighborhood organizations. Children and Youth Services Review, 27, 1119–1148.

Houser, L., Nisbet, E., & White, K. (2012). Giving voice to New Jersey's caregivers: The union experiences of home-based child care providers. Center for Women and Work. New Brunswick, NJ. Retrieved February 9, 2015 from http://smlr.rutgers.edu/cww-report-giving-voice.

Hoxby, C. M. (1996). How teachers' unions affect education production. The Quarterly Journal of Economics, 111, 671–718.

Iruka, I. U., & Carver, P. R. (2006). Initial Results from the 2005 NHES Early Childhood Program Participation Survey. Report no. NCES 2006–075. Washington, DC: National Center for Education Statistics.

Kim, J., & Fram, M. S. (2009). Profiles of choice: Parents' patterns of priority in child care decision-making. Early Childhood Research Quarterly, 24, 77–91.

King, M., Ruggles, S., Alexander, T., Flood, S., Genadek, K., Schroeder, M., Trampe, B., & Vick, R. (2010). Integrated Public Use Microdata Series, Current Population Survey: Version 3.0. [Machine-readable database]. Minneapolis, MN: University of Minnesota.

Laughlin, L. (2013). Who's minding the kids? Child care arrangements: Spring 2011. Current Population Reports, P70-135. Washington, DC: US Census Bureau. Retrieved February 9, 2015 from http://www.census.gov/prod/2013pubs/p70-135.pdf.

Layzer, J. I., & Goodson, B. D. (2006). Care in the home: A description of family child care and the experiences of the families and children that use it. National Study of Child Care for Low-Income Families, Wave 1 Report. Cambridge, MA: Abt Associates. Retrieved February 9, 2015 from http://www.researchconnections.org/location/ccrca11568.

Lewin, D., Keefe, J. H., & Kochan, T. A. (2012). New great debate about unionism and collective bargaining in US state and local governments. ILR Review, 65, 749–778.

Lippman, L., Vandivere, S., Keith, J., & Atienza, A. (2008). Child care use by low-income families: Variations across states. Research Brief no. 2008–23. Washington, DC: Child Trends. Retrieved March 4, 2015 from http://www.childtrends.org/wp-content/uploads/2013/07/2008-23ChildCareLow-Income.pdf.

Matthews, H., & Lim, T. (2010). Child care and development block grant participation in 2009. Washington, DC: Center for Law and Social Policy. Retrieved February 9, 2015 from http://www.clasp.org/admin/site/publications/files/ccdbg-09-participation.pdf.

McClendon, J. A., Wheeler, H. N., & Weikle, R. D. (1998). The individual decision to unionize. Labor Studies Journal, 23, 34–54.

National Association of Child Care Resource and Referral Agencies. (2011) Child care in America. Arlington, VA: Retrieved February 9, 2015 from http://www.naccrra.org/sites/default/files/default_site_pages/2011/childcareinamericafacts_2011_final.pdf.

National Institute of Child Health and Human Development Early Child Care Research Network. (2000). Characteristics and quality of child care for toddlers and preschoolers. Applied Developmental Science, 4, 116–135.

Norcross, E. (2010). Public-sector unionism: A review. Working Paper no. 11–26. Washington, DC: Mercatus Center at George Washington University. Retrieved February 9, 2015 from http://mercatus.org/publication/public-sector-unionism-review.

Porter, T., Paulsell, D., Del Grosso, P., Avellar, S., Hass, R., & Vuong, L. (2010). A review of the literature on home-based child care: Implications for future directions. Washington, DC: Office of Planning, Research and Evaluation. Retrieved February 9, 2015 from http://www.acf.hhs.gov/programs/opre/cc/supporting_quality/reports/lit_review/lit_review.pdf

Rigby, E., Ryan, R. M., & Brooks-Gunn, J. (2007). Child care quality in different state policy contexts. Journal of Policy Analysis and Management, 26, 887–908.

Rose, K. K., & Elicker, J. (2010). Maternal child care preferences for infants, toddlers, and preschoolers: The disconnect between policy and preference in the USA. Community, Work & Family, 13, 205–229.

Ryan, R. M., Johnson, A., Rigby, E., & Brooks-Gunn, J. (2011). The impact of child care subsidy use on child care quality. Early Childhood Research Quarterly, 26, 320–331.

Schulman, K., & Blank, H. (2007). State child care assistance policies 2007: Some steps forward, more progress needed. Washington, DC: National Women's Law Center. Retrieved February 9, 2015 from http://www.nwlc.org/sites/default/files/pdfs/statechildcareassistancepolicies2007.pdf.

Schulman, K., & Blank, H. (2012). Downward slide: State child care assistance policies 2012. Washington, DC: National Women's Law Center. Retrieved February 9, 2015 from http://www.nwlc.org/sites/default/files/pdfs/NWLC2012_StateChildCareAssistanceReport.pdf

Shonkoff, J., & Phillips, D. (2000). From neurons to neighborhoods: The science of early childhood development. Washington, DC: National Academy Press.

Shonkoff, J., Boyce, W. T., & McEwen, B. S. (2009). Neuroscience, molecular biology, and the childhood roots of health disparities: Building a new framework for health promotion and disease prevention. Journal of the American Medical Association, 301, 2252–2259.

Smith, K. (2006). Rural families choose home-based child care for their preschool-aged children. Policy Brief no. 3. Durham, NH: Carsey Institute. Retrieved February 9, 2015 from http://scholars.unh.edu/cgi/viewcontent.cgi?article=1008&context=carsey.

Smith, P. R. (2007). Publicization of home-based care work in state labor law. Minnesota Law Review, 92, 1390–1423.

Tang, S., Coley, R. L., & Votruba-Drzal, E. (2012). Low-income families' selection of child care for their young children. Children and Youth Services Review, 34, 2002–2011.

Tekin, E. (2007). Childcare subsidies, wages, and employment of single mothers. Journal of Human Resources, 42, 453–487.

Valletta, R. (1989). The impact of unionism on municipal expenditures and revenues. Labor Relations Review, 42, 430–442.

Werner, E. (1989). High-risk children in young adulthood: A longitudinal study from birth to 32 years. American Journal of Orthopsychiatry, 59, 72–81.

Yoshikawa, H. (2011). Immigrants raising citizens: Undocumented parents and their young children. New Yor k, NY: Russell Sage Foundation.

Zax, J., & Ichniowski, C. (1988). The effects of public sector unionism on pay, employment, department budgets, and municipal expenditures. In R. B. Freeman & C. Ichinowski (Eds.), When public sector workers unionize (pp. 323–364). Chicago, IL: University of Chicago Press.