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Does It Matter if Your Health Insurer Is For Profit? Effects of Ownership on Premiums, Insurance Coverage, and Medical Spending†

By Leemore Dafny*

There is limited empirical evidence about the impact of for-profit health insurers on various outcomes. I study the effects of conversions to for-profit status by Blue Cross and Blue Shield (BCBS) affiliates in 11 states, spanning 28 geographic markets. I find both the BCBS affiliate and its rivals increased premiums following conversions in markets where the converting affiliate had substantial market share. Medicaid enrollment rates also increased in these markets, a pattern consistent with "crowd in" of families who were formerly privately insured. The results suggest for-profit insurers are likelier than not-for-profit insurers to exercise market power when they possess it. (*JEL* G22, I13, I18, I38)

In most US industries, a single ownership form prevails. For example, consumer goods are generally produced by for-profit firms, policing and safety services are usually supplied by government agencies, and historically, US life insurance was dominated by not-for-profit mutual companies. However, there are a number of sectors in which multiple ownership forms coexist, most notably those in which public purchasing plays a significant role, such as education, incarceration, and health-care services. Given the substantial public stake in the performance of these sectors, obtaining a better understanding of whether and how ownership form impacts organizational conduct is of critical importance to achieving public policy goals.

The US health insurance industry is a prime example of such a sector, and the subject of ownership form came to the fore during the debate preceding the passage of the Affordable Care Act in March 2010. The legislation included individual and employer mandates to carry health insurance, as well as federal subsidies for eligible

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individuals to purchase it via the new health insurance exchanges. The projected expansion of the private insurance sector was controversial, owing to pervasive consumer dissatisfaction with private insurance plans and the companies administering them. In a widely publicized speech to the American Medical Association, the national professional association of physicians and medical students, President Obama averred, "What I refuse to do is simply create a system where insurance companies have more customers on Uncle Sam's dime, but still fail to meet their responsibilities."[1] The most strident criticism was often directed toward for-profit insurers, who were accused of putting profits before patients. Indeed, the final legislation included \$6 billion of funding for new, not-for-profit co-ops on the grounds that these new insurers would "focus on getting the best value for customers, rather than maximizing plan revenues or profits."[2]

While the co-ops proved an unsuccessful short-term experiment, with only 4 of the 24 funded by the legislation operational as of July 2017, the role of profit status has resurfaced around the subject of insurer participation in the individual health insurance exchanges, which opened nationwide in 2014. The large for-profit insurers (e.g., United, Humana, Anthem, Aetna, and Cigna) have made frequent, highly visible entry and exit decisions in response to profit projections and policy uncertainty. Policymakers have little evidence to rely upon when deciding whether (and how) to favor or subsidize particular ownership forms in the insurance sector.

In this paper, we consider the effect of for-profit ownership on pricing, insurance coverage, and medical-loss ratios (the share of premiums used to reimburse medical claims). While there is an extensive theoretical and empirical literature examining the impact of ownership form on outcomes in the hospital sector (e.g., Weisbrod 1988, Cutler 2000, Sloan 2000, and Duggan 2000, to cite but a few), there is comparatively little research of this kind focusing on the health insurance industry.

Theoretical models offer ambiguous predictions, underscoring the value of empirical analysis. Many models of not-for-profit (NFP) behavior in health-care settings predict underpricing relative to for profits (FPs), holding quality constant. These models assume that NFPs explicitly value the quantity of enrollees ("access" in the policy vernacular), whereas FPs value enrollees only to the extent they are profitable. Alternative, consumer-focused theories posit that FPs must underprice to compensate consumers for the more severe agency problem arising from strict profit maximization. Of course, if ownership form is associated with productivity, pricing will reflect these differences as well. While our analysis does not explicitly distinguish among these various mechanisms, it uncovers empirical differences in the observed behavior of FP and NFP insurers. These are especially pertinent in light of the substantial reforms and regulatory actions currently impacting private health insurance markets.

[1] See http://www.usatoday.com/news/washington/2009-06-15-obama-speech-text_N.htm.

[2] Senator Kent Conrad (R-ND), the sponsor of the Consumer-Owner and Oriented Plan (CO-OP), stated "[m]any experts believe co-ops, as non-profits, could offer significant discounts when compared to traditional, for-profit insurance companies." Senator Conrad's office did not respond to a request for the names of the experts. Roughly \$2 billion was ultimately spent to fund 24 state co-ops. Source: "FAQ about the Consumer-Owner and Oriented Plan (CO-OP)," accessed July 15, 2010 at http://conrad.senate.gov/issues/statements/healthcare/090813_coop_QA.cfm.

Our primary data source is the Large Employer Health Insurance Dataset (LEHID). This proprietary panel dataset of employer-sponsored health plans includes information on approximately 10 million enrollees annually. During our study period, 1997–2009, over 950 employers—primarily multisite, publicly traded firms—are represented in the sample. The data span 139 geographic insurance markets, which (per the data source) reflect the boundaries used by insurers when setting premiums. We also utilize state-level data from the Current Population Survey and the National Association of Insurance Commissioners to evaluate the impact of FP market share on insurance coverage rates and insurer medical-loss ratios, respectively. Importantly, the study period concludes before the passage of the Affordable Care Act (ACA) in March 2010. The ACA included significant insurance market reforms and public insurance expansions, and while some phased in slowly, there is substantial evidence that health-care spending and insurance take-up changed significantly after 2009.

Given the dearth of information on the ownership status of health insurers, we begin by documenting important facts about FP insurance in the LEHID, including market share by region, by plan type, by insurance type (fully insured versus self-insured), and over time. Fully insured plans are a traditional insurance product in which the insured pays the carrier to bear the risk of realized health-care outlays. Many large employers, who can spread the health risk of members across a large pool of enrollees, choose to self-insure. Self-insured employers typically outsource benefits management, provider contracting, and claims administration—generally to the same insurers offering fully insured plans—but are responsible for the realized costs of care of plan members. As we show below, the market share of for-profit insurers is particularly high in the self-insured segment.

To explore the relationship between FP status and premiums, we develop a regression-adjusted premium index for each of the 139 geographic markets over the 13-year study period. This index, set to 100 for each market at the start of our study period, captures the average year-over-year growth for the same health plans in each market, where a health plan is defined by a unique employer, market, carrier, and plan type (such as HMO). We control for two continuous measures constructed by actuaries and included in our data: *demographic factor*, which is the average number of "person equivalents" per enrollee, and *plan design*, an index that reflects the "actuarial value" of a plan, i.e., the share of spending incurred by a standard population that would be paid for by the plan (rather than by the insured via deductibles, coinsurance, or co-payments). We construct premium indices separately for fully insured and self-insured plans, as these plans are priced differently and subject to different regulations and competitive environments.

We find no significant association between changes in market-level FP share and our market-level premium index, controlling for market-year covariates, such as the local unemployment rate and Medicare spending (as proxies for trends in medical utilization). However, time-varying omitted characteristics may bias these estimates if they are correlated with FP share. For example, FP carriers may strategically expand where they can enjoy the highest margin growth.

In order to address this identification concern, we study plausibly exogenous shocks to local FP share generated by ownership conversions of Blue Cross and

Blue Shield (BCBS) affiliates in 11 states. A wave of conversions and unsuccessful attempts to convert followed the 1994 decision by the national umbrella organization (the Blue Cross and Blue Shield Association, or BCBSA) to permit conversions of local BCBS plans to FP status. BCBS affiliates offer insurance throughout the United States and typically rank first or second in terms of local market shares (Robinson 2006).

We compare premium growth for plans in the 11 states (with 28 distinct geographic markets) experiencing conversions with premium growth for plans in the 8 states (including DC, yielding 19 "control" markets) whose local BCBS affiliates attempted to convert but, owing to a variety of factors such as community opposition, golden parachutes for executives, and regulatory actions, ultimately failed in this effort. If the ability to consummate a conversion is orthogonal to other determinants of premiums, then local BCBS FP status can serve as an instrument for market-level FP penetration in this sample. This assumption is supported by the similar pre-conversion trends in premiums in areas with and without consummated conversions. We also discuss results obtained from the sample of 47 states whose BCBS affiliates had not yet converted as of the start of the study period. Estimation using this sample requires the stronger assumption (also supported by the absence of pre-conversion trends) that the *attempt* to convert is exogenous.

We find no statistically significant impact of BCBS conversions on market-level prices, *on average*. However, when we separate markets by whether the pre-conversion BCBS share is above or below average (20.2 percent in our sample),[3] we find that above-average ("high BCBS") markets experienced an increase in fully insured premiums of roughly 13 percent. The effect on self-insured premiums is a marginally significant 4 percent. This modest change is consistent with more robust competition for this latter customer segment (Dafny 2010). Notably, there are also no significant pre-trends in premiums for high or low BCBS markets (relative to non-converting markets) prior to conversion.

We extend our premium analysis by constructing separate indices for BCBS and non-BCBS plans and estimating the key specifications using each. The results show that post-conversion price increases in high-BCBS share markets were common to BCBS and non-BCBS plans. Thus, a simple comparison of price changes for converting and non-converting plans in the same market—a common methodology for case studies of conversions—would understate the effect of conversion. This spillover effect on rivals confirms earlier work suggesting premiums in health insurance markets are strategic complements (Dafny 2010; Dafny, Duggan, and Ramanarayanan 2012).

We also evaluate the effect of BCBS conversions on insurance coverage and medical-loss ratios. As these measures are only available at the state-year level, our sample size is considerably smaller when we limit the analysis to the 19 states with attempted conversions. However, even in this sample, we find marginally statistically significant increases in Medicaid enrollment rates in states with relatively large BCBS conversions, as compared to states with smaller conversions or

[3] As we discuss in Section IIIB, our threshold shares likely correspond to higher shares in the entire commercial insurance market (i.e., including individuals, small employers, and large but primarily single-site employers).

failed conversion attempts; these results are statistically significant at conventional levels using the broad sample of 47 states. Where they occur, increases in Medicaid enrollment appear to be offset by decreases in employer-sponsored and individual insurance, yielding no net effect on overall insurance coverage. Medical-loss ratios at the state-year level do not change in response to conversions. However, we find that rivals of converting BCBS affiliates experienced significant increases in their MLRs, which were offset by (insignificant) decreases on the part of converting BCBS affiliates. This pattern of findings is consistent with a transfer of higher risk customers from converting plans to rivals, although we lack the data to confirm this mechanism.

Considered as a whole, the results suggest that sizeable BCBS conversions resulted in higher prices, crowd-in to Medicaid programs, and no net change in medical spending per premium dollar. While it is difficult to assess whether the "BCBS conversion effect" is a good estimate of the average difference in NFP and FP insurers in general, this effect is plausibly predictive of the impacts of changes in FP share in the future (i.e., the marginal FP insurer). First, a large number of BCBS affiliates are still NP, and some are contemplating conversion (e.g., Horizon Blue Cross in New Jersey) or taking intermediate steps (e.g., BCBS of Michigan, which converted in 2012 from NFP to mutual ownership). Second, the three remaining federally supported co-ops are at risk of becoming insolvent. If they do not exit the market entirely, they are likely to convert to for-profit status in order to gain access to investor capital.[4]

The paper proceeds in five sections. In Section I, we discuss the historical origins of FP insurers, summarize prior relevant research, and provide relevant background on the BCBS conversions that underlie our identification strategy. We describe our data sources in Section II. We present our estimates of the effect of FP ownership on premiums in Section III. We discuss results on non-price outcomes in Section IV. Section V concludes.

I. Background

A. *Origin and Evolution of FP Insurance Plans in the United States*

The US health insurance industry originated in the 1930s with the formation of prepaid insurance plans by hospitals, which were designed to cover inpatient charges. These came to be known as Blue Cross plans and incorporated several features proposed by the American Hospital Association (AHA), including being chartered as charitable organizations designed to serve the community. Blue Shield plans subsequently arose to cover physician charges. The two Blues merged to form the Blue Cross Blue Shield Association in 1982. FP insurers entered the market toward the middle of the twentieth century when health insurance enrollment soared

[4] Indeed, Massachusetts' Minuteman Health attempted such a conversion during the summer of 2017. After it was unable to raise sufficient funding, the company was placed under receivership (McCluskey 2017).

as employers sought alternative forms of employee compensation in the wake of WWII-era wage controls.[5]

Precise figures on current or historical market shares of FP insurers are difficult to obtain. According to America's Alliance for Advancing Nonprofit Health Care, approximately 52 percent of health-plan members were covered through FP insurers in 2008.[6] Using data from the National Association of Insurance Commissioners (NAIC), an organization of state regulators, we obtain a similar figure (54 percent) for 2008.[7] However, the NAIC data excludes self-insured enrollees, as only fully insured plans are regulated by the states. In the LEHID, we find FP shares of 47 percent among fully insured members and 72 percent in the self-insured segment, also in 2008.[8] Clearly, FPs play an important role in the US health insurance industry in general and a particularly significant role in the large employer segment, the focus of this study.

B. *Prior Research*

The literature examining ownership status in the health insurance industry is relatively sparse. Before turning to these studies, we note that our work is informed by the rich theoretical and empirical literature on ownership status in the US hospital industry. Surveys of this literature can be found in Capps, Carlton, and David (2010) and Chang and Jacobson (2011). Chang and Jacobson characterize four key models, all of which extend naturally to the insurance setting. At one end of the spectrum is the "for profits in disguise" (FPID) model, which posits that NFPs behave no differently than FPs.[9] At the other end is "pure altruism," and in between is "output (and/or quality) maximization" and "perquisite maximization." Both altruists and output maximizers value access to care, leading to underpricing (relative to FPIDs or FPs). However, FPs/FPIDs and NFPs can coexist (i.e., both serve customers) for a variety of reasons, such as capacity constraints, cost differences, and product differentiation. While capacity constraints are less relevant in the insurance industry, costs may certainly vary by ownership form, and there are many sources of differentiation, including reputation/marketing, provider networks, benefit design, and customer service. In sum, flexible theoretical models allow for a variety of predictions vis-à-vis price, quantity, and quality.

[5] See Sobel and Silas (2007).

[6] This estimate includes enrollees in government-financed plans, as well as most enrollees in self-insured plans, but excludes health plans with <100,000 enrollees ("Basic Facts and Figures: Nonprofit Health Plans," available at http://www.nonprofithealthcare.org/resources/BasicFacts-NonprofitHealthPlans.pdf).

[7] We discuss the NAIC data in Section II. Our tabulations reflect only enrollment in comprehensive medical insurance. The NAIC data exclude plans from the state of California, which has high FP penetration.

[8] Self-insurance is more common in LEHID relative to the (non-elderly) insured population at large. In 2008, 80 percent of LEHID enrollees were in self-insured plans, whereas 55 percent of workers with health insurance were in self-insured plans. See Employee Benefit Research Institute (2009).

[9] This conjecture has empirical support from a number of studies, including Duggan (2002); Cutler and Horwitz (1999); Silverman and Skinner (2004); Dafny (2005); and Capps, Carlton, and David (2010). Collectively, these studies find that NFP hospitals behave similarly to FPs, particularly in markets where they face greater competition from FP hospitals, on dimensions like pricing, profitability, "gaming" of reimbursement codes, quality of care, and service offerings.

The small literature on ownership status of health insurers can be subdivided into two general categories defined by the outcomes considered: plan quality/enrollee satisfaction and plan pricing/profits. Most studies of the first type find higher levels of quality and satisfaction for NFP plans. Using data on Medicare HMOs from 1998, Schneider, Zaslavsky, and Epstein (2005) report that FP HMOs score lower on four audited quality measures (breast cancer screening, diabetic eye examinations, administering beta-blockers after a heart attack, and follow-up after mental illness hospitalization). Controlling for county fixed effects and socioeconomic factors (including age, gender, area income, and rural residence) of plan enrollees has little impact on the estimates. Studies comparing FP and NFP health plans also find that consumer satisfaction is higher among enrollees of NFP plans (Gillies et al. 2006), especially for patients in poor health (Tu and Reschovsky 2002). Finally, NFP plans appear to perform better with respect to provision of care for less affluent populations, such as Medicaid enrollees (Long 2008).

The two studies that consider financial measures (profits and premiums) find little impact of ownership on these dimensions. Both rely on data from Interstudy, a private firm that has historically provided data only on HMOs, and thus the analyses are limited to this product line. Pauly et al. (2002) use data from 1994–1997 and find no association between MSA-level HMO profits and FP HMO penetration. Town, Feldman, and Wholey (2004) study the effects of HMO conversions to FP status between 1987 and 2001. They find no significant impact of these conversions on a broad range of outcomes, including prices (estimated as average revenue per enrollee), profit margins, and utilization.

Our study also relies on conversions to identify the effect of ownership status; however, there are important differences in our sample, unit of observation, study design, and outcomes of interest. First, we focus primarily on the set of markets experiencing conversions or conversion attempts; thus, our treatment and control groups are likely to be more similar than the implicit treatment and control groups in prior studies. Our data include all plan types (HMO, POS, PPO, and indemnity), as well as funding arrangements (fully insured and self-insured). The original unit of observation is the employer-market-insurance type-carrier-plan type, which enables us to include a rich set of controls for the underlying insured population and the characteristics of their health plans when constructing a market-year premium index. We also study the effects of conversions on premiums offered by both converting and non-converting firms. This is of particular relevance given the nature of competition among insurers. To the extent that insurance prices are strategic complements, price increases by one firm will be reinforced by its rivals, who will optimally raise price in response. Thus, research that implicitly relies on non-converting plans as a control group for converting plans may generate downward-biased estimates of price effects.

In addition, we explore the impact of conversions on medical-loss ratios and insurance coverage rates, both measured at the state-year level. The medical-loss ratio (MLR) is of interest both as a rough measure of profits (Karaca-Mandic, Abraham, and Simon 2015) and of quality. A high MLR implies a greater share of premiums is spent directly on patients (as opposed to management or profits). Of course, linking high MLRs with quality assumes more spending leads to better health, and that management generates no value, assertions which are appropriately disputed in the literature (e.g., Robinson 1997). The insurance coverage analysis permits an indirect assessment of the premium effects of conversions, as higher private-sector prices should crowd out some private coverage and potentially crowd in some Medicaid coverage (particularly the children of parents dropping private coverage).

C. *Blue Cross Blue Shield Plans*

Our analysis utilizes the conversion of 11 BCBS plans to FP stock corporations as a source of plausibly exogenous variation in the local market share of FP plans. BCBS plans are often the dominant insurers in their local markets, so conversion typically leads to a sharp increase in local FP share. Robinson (2006) estimates that BCBS plans hold the largest market share in every state except Nevada and California and would together control 44 percent of the national market if they were considered as one firm. Dafny (2015) reports higher estimates of national BCBS share—50 percent in 2006—and estimates this share reached 52 percent in 2014.

As previously mentioned, BCBS plans were chartered as social welfare organizations and were thus exempt from most taxes. Congress revoked BCBS' federal tax exemption as part of the 1986 Tax Reform Act.[10] In June 1994, partly prompted by the decision of Blue Cross of California to form a for-profit subsidiary (WellPoint, originally a network of for-profit HMOs and PPOs, focused on the non-group market), the national BCBS association modified its bylaws to allow affiliates to convert to FP ownership. This sparked a series of ownership changes with plans in 14 states converting to FP stock companies by 2003. We are only able to study 11 of these conversions as the first 3 occurred prior to the start of our data (see Table 1 for details).

Many BCBS plans proposing or undergoing conversion cited access to equity capital as the key driver for conversion.[11] Uses for additional capital include infrastructure investments (for example, in information technology or disease management) and acquisitions of other plans. Larger insurers can spread fixed costs over more enrollees, thereby improving operating margins. In addition, several studies confirm larger insurers pay lower prices to providers (e.g., Moriya, Vogt, and Gaynor 2010; Trish and Herring 2015; and Roberts, Chernew, and McWilliams 2017). Representatives of converting plans have also cited the importance of attracting and retaining top management talent, which can more easily be accomplished when equity and stock options are included in compensation packages (Schramm 2004). Finally, by creating tradable shares, conversion facilitates acquisition by other plans.

Table 1 lists the BCBS plans that attempted to convert to FP stock corporations between 1998 and 2009, subdivided by successful and unsuccessful attempts.[12]

[10] As 501(m) organizations, BCBS plans are entitled to other tax benefits, such as "special deductions" and state tax exemptions (in some states). Source: Coordinated Issue Paper – Blue Cross Blue Shield Health Insurance, available at http://web.archive.org/web/20081004230757/http://www.irs.gov/businesses/article/0,,id=183646,00.html.

[11] See Grossman and Strunk (2004).

[12] We thank Chris Conover for sharing his detailed notes on plan conversions. In addition to the 11 plans listed in Table 1, 3 additional plans converted prior to our study period (California and Georgia in 1996 and Virginia

TABLE 1—BLUE CROSS AND BLUE SHIELD CONVERSIONS TO FOR-PROFIT STOCK COMPANIES, 1998–2009

	Conversion to FP stock company	Year recorded in data
Panel A. Successful conversions		
Anthem		
Colorado	November 2001	2002
Connecticut	November 2001	2002
Indiana (Accordia)	November 2001	2002
Kentucky	November 2001	2002
Maine	November 2001	2002
Missouri (RightChoice)	November 2000	2001
Nevada	November 2001	2002
New Hampshire	November 2001	2002
Ohio (CMIC)	November 2001	2002
Wisconsin (Cobalt)	March 2001	2001
WellPoint		
New York (Empire)	November 2002	2003
	Review period	Reason for failure
Panel B. Unsuccessful conversion attempts		
New Jersey (Horizon)	2001–2005	Regulators unconvinced by claims that Horizon needed additional capital; strong provider opposition due to Horizon's high market share and low reimbursement rates
North Carolina	2002–July 2003	Regulators demanded 100% of stock be placed in a foundation; BCBS regulations permitted a maximum of 5% ownership stake by foundations
Kansas	2001–August 2003	Concern that conversion would result in large price increases due to high market share (in non-HMO market)
CareFirst		
Delaware	2002–September 2003	Public outrage about intended executive bonuses
District of Columbia	2002–September 2003	Public outrage about intended executive bonuses
Maryland	2002–September 2003	Public outrage about intended executive bonuses
Premera		
Alaska	2002–March 2007	Abandoned because of failure in Washington
Washington	2002–March 2007	Concerns about acquisition by out-of-state insurer and disagreements about how to put stock into a foundation

Notes: Parent companies are listed in bold. Year recorded in data refers to the first post-conversion year as coded in our dataset. For unsuccessful conversion attempts, the review period begins with the year in which a conversion attempt was announced and ends when it was officially blocked by regulators or withdrawn from consideration.

Conversions require approval from state insurance regulators. To arrive at a determination, regulators investigate the likely effects of the proposed conversion on outcomes, such as price, access, and provider reimbursement (Beaulieu 2004).

in 1997). These states are not included in our analysis sample.

They also specify the amount and form of compensation to be provided to the state or community in exchange for the transfer of assets to private stakeholders.

The identification assumption underlying our primary analysis is that the success of a conversion attempt is exogenous to changes in omitted factors affecting the outcomes of interest. In Table 1, we summarize the reasons for each unsuccessful attempt. For example, CareFirst BCBS (serving Delaware, DC, and Maryland) could not secure the necessary approvals following public outrage over intended executive bonuses. Premera (in Washington and Alaska) was unable to convert because regulators were concerned the insurer would ultimately be acquired by an out-of-state parent company, and the parties could not come to terms about the amount to be transferred to new charitable foundations.[13] As these examples suggest, the range of reasons for unsuccessful attempts is broad and not clearly linked to premium, spending, or coverage trends. Indeed, in Section IIIB, we confirm that our outcomes of interest trend similarly in areas with successful and unsuccessful conversion attempts prior to the realized conversions. In addition, markets with successful and unsuccessful conversion attempts have similar unemployment rates and average Medicare spending (as of 2001, the modal pre-conversion year). Of course, we cannot be certain that approval is exogenous to expectations regarding price changes (and other outcomes). If proposed conversions likeliest to lead to price increases were precisely the ones blocked, then our estimated conversion effects are understated. Alternatively, if conversions expected to yield the greatest returns were pursued most vigorously, and thus the insurance executives involved were more willing to arrive at the necessary compromises to close the deals, then our estimates may overstate the price effects of a typical conversion.

During our study period, we observe three distinct types of ownership changes for BCBS affiliates: NFP → Mutual (four states); NFP → FP stock company (three states); and Mutual → FP stock company (eight states). We define an FP conversion as having taken place if the BCBS plan becomes a stock company, i.e., we combine the last two types of ownership changes.[14] Mutual insurers are owned by plan subscribers and hence explicitly value policyholder interests; as such, most analysts consider this hybrid ownership form closer to NFP than FP status.[15] In Section IIID(“Robustness Checks and Extensions”), we discuss reduced-form estimates of the impact of all three conversion types (details and timing of which are listed in Table A1). However, given the small number of experiments available to identify them separately, as well as the short pre- and post-periods for the NFP → Mutual conversions, our primary analyses use our FP conversion definition.

Eight of the eleven conversions so defined take place in the same year (2001) when Anthem (the parent organization of these plans) demutualized and launched an IPO. While it would be ideal to have more variation in the timing of

[13] See Song (2004).

[14] Note that all of the affiliates converting from NFP to Mutual status subsequently converted to FP status as they were a part of Anthem, a consolidator of BCBS plans which demutualized and converted to a for-profit stock company in 2001.

[15] For example, the Alliance for Advancing Nonprofit Healthcare (cited above) lumps mutuals together with nonprofits when reporting nonprofit market share, implicitly viewing investor ownership as a bright dividing line. As a matter of law, mutuals may be nonprofit or for profit.

conversions, we do not rely solely on a pre-post study design: we also explore how the effect of conversion varies with the market share of the converting plan. There are 28 distinct geographic markets within the 11 states with converting BCBS affiliates, and 19 markets in the states with unsuccessful conversion attempts. The pre-conversion BCBS market shares in the 28 affected markets—averaged over the 3 years preceding conversion—range between 6 and 35 percent with a mean of 20 percent. By comparison, the BCBS market share as of 2001 (the modal pre-conversion year) ranges between 16 and 36 percent in the 19 control markets with a mean of 26 percent.[16]

The prior literature on BCBS conversions largely takes a case study approach. For example, Hall and Conover (2003) conduct a qualitative analysis of four conversions. Based on interviews with providers, consumer advocates and regulators, the authors conclude that there is little concern among these stakeholder groups that conversion will ultimately produce premium increases. Several papers focus on the failed conversion attempt by CareFirst BCBS in Maryland, derailed in part by demands for post-conversion bonuses by BCBS executives (e.g., Robinson 2004 and Beaulieu 2004). A notable exception to the case study approach is Conover, Hall, and Ostermann (2005), which examines changes in per capita health spending, hospital profitability, and insurance access resulting from BCBS conversions in all states between 1993 and 2003. Using state-level data on physician and hospital health spending from the Center for Medicare and Medicaid Services (CMS) and uninsurance rates from the Current Population Survey, the authors estimate specifications that include state and year fixed effects and indicators for years before, during, and after BCBS conversion. They conclude that BCBS conversions had only a modest impact on health spending and insurance access in affected states. Our results largely corroborate these findings; however, we also find important heterogeneity in the effects of conversion in markets with different BCBS market shares.

II. Data

A. *Large Employer Health Insurance Dataset*

Our main source of data is the Large Employer Health Insurance Dataset (LEHID), which contains detailed information on the health plans offered by a sample of large employers between 1997 and 2009. This proprietary dataset is also used in Dafny (2010) and Dafny, Duggan, and Ramanarayanan (2012) but is supplemented in this study with four additional years of data (1997 and 2007–2009).

[16] Across the 77 markets in the 29 states without conversion attempts (and without successful conversions prior to the start of the study period), the BCBS market share in 2001 ranges between 4 and 62 percent with a mean of 29 percent. Note that BCBS market shares in LEHID are substantially lower than BCBS market shares for commercial health insurance more broadly. I estimate BCBS affiliates accounted for roughly 50 percent of comprehensive, commercial medical insurance in 2006 and 2010 (Dafny 2015). The difference is due to the composition of LEHID, which consists primarily of large, typically multistate employers. A number of for-profit, non-Blue carriers specifically target these employers and have smaller shares in other customer segments.

The unit of observation in LEHID is a health-plan year, where a health plan is defined as a unique combination of an employer, market, insurance carrier, plan type, and insurance type (e.g., Company X's Chicago-area fully insured Aetna HMO). Most *employers* are large, multisite, publicly traded firms, such as those included on the *Fortune 1000* list. Geographic *markets* are defined by the data source using three-digit zip codes and reflect the areas used by insurance *carriers* (such as Blue Cross and Blue Shield of Illinois or Humana) to quote premiums. There are 139 geographic markets and most reflect metropolitan areas or nonmetropolitan areas within the same state (e.g., in Illinois, there are three markets: Chicago, northern Illinois except Chicago, and southern Illinois). The *plan types* are Health Maintenance Organization (HMO), Point of Service (POS), Preferred Provider Organization (PPO), and Indemnity.

Insurance type refers to self-insured or fully insured; the sample includes both. As previously noted, insurance carriers do not underwrite risk for self-insured plans. Typically, they process claims, negotiate provider rates, and perform various additional services, such as utilization review and disease management. Self-insured "premiums" are set by employers, who have the fiduciary responsibility to ensure they are accurate estimates of all costs associated with their plans. These costs include expected medical outlays, premiums for stop-loss insurance (if purchased), and charges levied by the administering carrier. Self-insured plans are regulated by the federal government, hence state-imposed benefit mandates and premium taxes do not apply. Large employers rely disproportionately on these plans, and accordingly, they account for three-quarters of the observations in our data. Due to the differences in pricing and regulation of self and fully insured plans, we perform all analyses separately by insurance type.

In any year an employer is represented in the sample, *all* plans offered by that employer in all markets are included in the data. Due to changes in the set of employers included in the sample from year to year, as well as changes in the set of options each employer offers, the median tenure of any health plan is only two years. As we discuss in Section III, this is one of the reasons we develop a market-year premium index. Here, we note that the index is constructed using *within-health-plan* premium growth. Premium growth in LEHID closely mirrors that reported by the Kaiser Family Foundation/Health Research and Educational Trust, whose estimates are based on a nationally representative sample of employers.[17] Additional information on the representativeness of LEHID is reported in Dafny, Duggan, and Ramanarayanan (2012).

In addition to the identifying information described thus far, we make use of four key variables from LEHID. *Premium* represents the combined annual employer and employee charge and is expressed as an average amount per enrolled employee; it therefore increases with the average family size for enrollees in a given plan. *Demographic factor* is a measure that reflects family size, age, and gender composition of enrollees in a given plan. These are important determinants of average

[17] The KFF/HRET survey randomly selects employers to obtain nationally representative statistics for employer-sponsored health insurance; approximately 2,000 employers respond each year. The micro data are not publicly available, nor is the sample designed to provide representative estimates for distinct geographic areas.

expected costs per enrollee in a plan. *Plan design factor* reflects the actuarial value of each health plan. Both factors are calculated by the data source, and the formulae were not disclosed to us. Higher values for either will result in higher premiums. For 2005 onward, LEHID contains an indicator for whether a plan is designated as “consumer directed.” Consumer-directed plans (CDPs) typically have high deductibles and are accompanied by consumer-managed health spending accounts. Prior research shows they are associated with lower premiums and slower premium growth at least in the short term (Buntin et al. 2006).

LEHID also includes the *number of enrollees* in each plan; this number excludes dependents, who are accounted for by the demographic factor variable described earlier. The total number of enrollees in all LEHID plans averages 4.7 million per year. Given an average (insured) family size of more than two, this implies over ten million Americans are part of the sample in a typical year. We compiled information on the ownership status for each observation from annual surveys administered by our source to the insurance companies affiliated with each LEHID plan. These surveys include nearly all plans in the data but are only available from the year 2000 onward. We filled in missing ownership information manually through independent research (e.g., web searches and analyst reports). We use Table 1 to code BCBS ownership status by market.

Table A2 presents descriptive statistics for the LEHID data, which spans the period 1997 to 2009, inclusive. Panel A pertains to the fully insured (FI) sample while panel B pertains to the self-insured (SI) sample. The table reveals several interesting trends in large employer-sponsored insurance over time. First, there is a pronounced shift toward SI plans. In 1997, SI plans account for 60 percent of observations (and 63 percent of covered employees), but by 2009, they are 83 percent of the sample (representing 82 percent of covered employees). (In Section IIID, we discuss whether and how this shift could be affecting our results.) Second, FI plans are predominantly HMOs throughout the study period, while SI plans have shifted away from the indemnity and POS plan types and toward PPOs (and to a lesser extent, HMOs) over time. Finally, consumer-directed plans (CDP) have been growing in popularity since this descriptive measure was first included in the LEHID dataset in 2005. By 2009, 23 percent of SI plans are designated as CDPs. Very few FI plans are CDPs.

In both samples, *demographic factor* exhibits a sharp dip from 2005 to 2006 and remains at a much lower level thereafter. According to our data source, this is due to a change in the methodology used to construct *demographic factor* beginning in 2006. As *demographic factor* is an important determinant of premiums and serves as a key control variable in our regression models, we construct empirical specifications to address any issues arising from recoding. As a robustness check, we also estimate our models using only data through 2005.

Restricting the sample to states with conversion attempts reduces the number of observations (covered employees) by 63 (64) percent. Table A3 contains descriptive statistics for this sample, separated by final conversion status. Average premiums are nearly the same in 1997 for plans located in areas with/without subsequent conversions. By 2009, the average nominal premium in markets with successful conversion attempts had risen by 163 percent (FI) and 117 percent (SI)

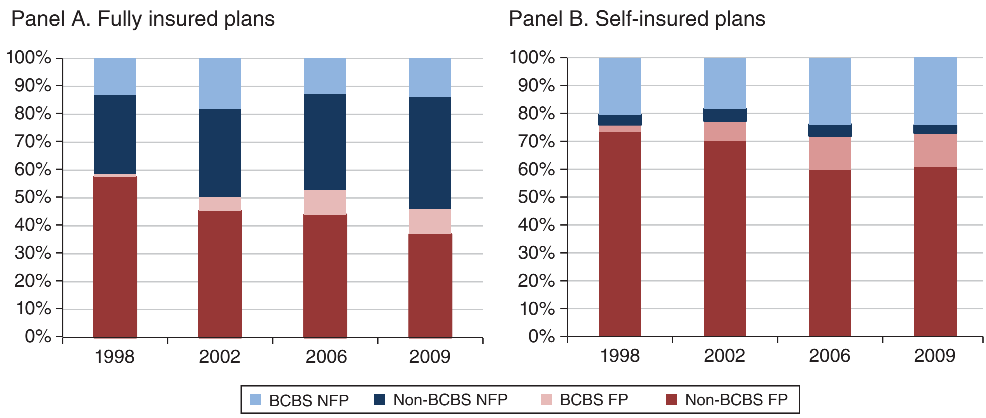

Figure 1. Percent of Enrollees in For-Profit and Not-for-Profit Plans, by BCBS Affiliation

Note: Market shares are calculated using LEHID.

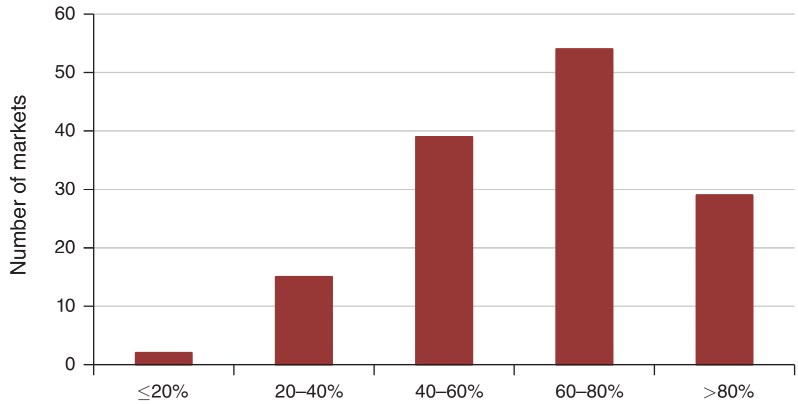

Figure 2. Distribution of For-Profit Market Shares

Notes: The figure reflects the average FP share for each geographic market in LEHID over the period 1997–2009. The sample includes fully insured and self-insured plans. Also, $N = 139$.

as compared to 148 percent (FI) and 113 percent (SI) in markets with unsuccessful attempts. However, these figures are not regression adjusted, nor are they weighted by plan enrollment.

Figure 1 presents estimates of insurer market shares obtained from the LEHID sample. Data are presented separately by year (in four-year increments), BCBS affiliation, and insurance type (FI and SI). Panel A shows that FP share in the FI market is sizable (51 percent on average) but exhibits a downward trend over time. FP share in the SI sector is markedly higher (averaging 72 percent) and has remained high during the past decade. The share of enrollees insured by BCBS plans increased during the study period with the majority of the growth occurring in the FP BCBS

Table 2—Descriptive Statistics, Market-Year Data

	1997	1999	2001	2003	2005	2007	2009
Market-year controls:							
Lagged Medicare costs per capita	4,574	4,843	5,113	6,062	6,837	7,592	8,298
	(913)	(853)	(855)	(993)	(990)	(1,097)	(1,198)
Lagged unemployment rate	0.54	0.45	0.40	0.56	0.54	0.46	0.58
	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)
Number of markets	139	139	139	139	139	139	139
Premium index:							
Panel A. Fully insured plans							
Premium index	100.00	112.45	135.06	178.20	214.37	254.97	288.63
	(0.00)	(12.07)	(17.16)	(23.91)	(30.39)	(37.64)	(38.07)
Number of markets	139	139	139	137	138	138	139
Panel B. Self-insured plans							
Premium index	100.00	103.89	127.08	168.54	210.74	260.78	290.18
	(0.00)	(7.51)	(9.99)	(12.24)	(16.02)	(21.53)	(23.60)
Number of markets	139	139	139	139	139	139	139

Notes: Sample means are presented in plain text, and standard deviations are in parentheses. All statistics are unweighted. The unit of observation is a market-year combination for each insurance type. Premium index is constructed using the coefficients on market-year fixed effects from a regression of plan-year premiums on various controls (including market-year fixed effects). Details are provided in the text. Due to space constraints, data are presented only for odd years.

segment. This is consistent with the large number of BCBS FP conversions taking place during this time.

Figure 2 illustrates substantial variation in share of FP insurers across geographic markets. When we break down FP share by plan type, we find that FP insurers are particularly dominant in the POS product line and relatively smaller in the HMO segment with 2009 national market shares of 91 and 56 percent, respectively.

We supplement the LEHID with time-varying measures of local economic conditions (the unemployment rate, as reported by the Bureau of Labor Statistics) and a measure of health-care utilization (Medicare costs per capita, reported by the Center for Medicare and Medicaid services).[18] As these measures are reported at the county-year level and LEHID markets are defined by three-digit zip codes, we make use of a mapping between zip codes and counties and, where necessary, use population data to calculate weighted average values for each LEHID market and year. Summary statistics for these measures are presented in Table 2.

B. *Medical-Loss Ratio Data*

The medical-loss ratio is the share of insurance premiums that is paid out for medical claims ("losses").[19] We construct state-year medical-loss ratios using insurer

[18] Medicare costs per enrollee and county are available from 1998 to the present. We extrapolate values for 1996–1997 using coefficient estimates from a regression of Medicare costs per enrollee on county fixed effects and county trends.

[19] Note this definition differs from the definition used to enforce the minimum MLR regulations in the Affordable Care Act (ACA). The ACA definition includes spending for quality improvements in the numerator and excludes

state-year data on total spending and premiums from the National Association of Insurance Commissioners (NAIC) for the years 2001–2009. Data for earlier years are not available. The data are described in the Appendix, and descriptive statistics are given in Table A4.

III. Do For-Profit Insurers Charge Higher Premiums?

To examine whether there is a causal link between ownership status and premiums, we study the effects of 11 FP conversions of BCBS plans (affecting 28 distinct geographic markets) and exploit variation in the timing and scale of these events. The control group consists of the 19 markets (in 8 states including DC) in which the local BCBS carrier unsuccessfully attempted to convert. The following subsections describe the main steps in our analysis in greater detail: (i) constructing our measure of insurance premiums (separately for the fully insured and self-insured customer segments); (ii) showing that market share of FPs increased following the conversions—but was not trending differently in markets that ultimately experienced conversions in the years preceding those conversions; (iii) exploring the impact of the conversions on area premiums (and then separately on premiums for BCBS affiliates versus rivals).

A. *Constructing a Market-Year Index of Premium Growth*

We use the LEHID micro data to construct a regression-adjusted premium index at the market-year level. We construct the index separately for FI and SI plans. The index, described in detail later in this subsection, captures market-specific changes in price for a standardized insurance product and population. We use this index rather than the underlying health-plan-year data for several reasons. First, the variation of interest (local FP share) occurs at the market-year level. A dependent variable at the same level of aggregation raises fewer concerns about understated standard errors. Second, utilizing the plan-year data raises significant sample issues. A regression with the plan year as the unit of observation would need to include plan fixed effects to capture unobservable determinants of premiums and year fixed effects to capture national premium trends. The regression would essentially compare changes in premiums for customers of converting BCBS plans with changes in premiums for customers of non-BCBS plans. Unfortunately, there are too few plans in our sample with a sufficiently long panel to permit reliable estimates of such a model. Even among employers appearing in the data for many consecutive years, there is very frequent churning in the set of plans offered.[20] In addition, the estimates would suffer from selection bias because only those BCBS customers remaining with their pre-conversion plans would identify the coefficients of interest.[21] By using the market year as

taxes and fees from the denominator; it cannot be calculated for earlier periods using available data sources. See US Government Accountability Office (2011).

[20] For example, over the period 1998-2006, 47 percent of employer-market cells experienced a change in the set of plans offered between year t and year $t + 1$ (Dafny 2010).

[21] It would be possible to use plan-year data for all measures except the BCBS FP indicator and to substitute the market-year value for it. Its coefficient would capture the impact of conversion on all plans that were present in

the unit of observation, we utilize more of the data and can also incorporate the spillover effect of conversion on rivals. Given the oligopolistic nature of most insurance markets, changes in the pricing of the local BCBS carrier should, all else equal, affect the pricing of competitors.

To obtain our market-year price index, we estimate the following model separately for each insurance type:

$$\text{(1)} \quad \ln(\mathit{premium})_{emcjt}$$
$$= \beta_0 + \beta_1 \mathit{demographics}_{emcjt} + \beta_2 \mathit{demographics}_{emcjt} \times (\mathit{year} \geq 2006)_t$$
$$+ \beta_3 \mathit{plan\,design}_{emcjt} + \beta_4 \mathit{CDP}_{emcjt} + \pi_{emcj} + \kappa_{jt} + \varphi_{mt} + \varepsilon,$$

where *emcj* denotes "employer-market-carrier plan type" (henceforth "plan") and t denotes year. The variables of interest are the market-year effects, denoted by φ_{mt}. The coefficients on these terms capture the average growth in premiums for each market and year. Because our objective is to isolate premium growth for a "standardized product," we include a rich set of controls.

First, we include all the plan-year-specific covariates we observe: *demographic factor*, *plan design*, and an indicator *CDP* for whether a plan is "consumer directed." To ensure that the change in the construction of *demographic factor* between 2005 and 2006 (referenced earlier in Section IIA) does not impact the results, we add an interaction term between *demographic factor* and an indicator for 2006 and beyond. Second, we include plan fixed effects (dummies for each plan, denoted by π_{emcj}). As a result, the coefficients on the market-year dummies will reflect average market-specific growth for the same plan from one year to the next. As previously noted, premium growth in LEHID closely matches premium growth nationwide, mitigating concerns about changes in sample composition.

Finally, we include plan type-year interactions to control for the effect of phenomena, such as the "HMO backlash" against utilization review and selective provider networks. The backlash caused HMOs to curtail these hallmark features, raising the relative cost of HMOs over time (Draper et al. 2002). If the shift away from HMOs occurred more quickly in some markets and if this is correlated with the presence and/or popularity of FPs, excluding the plan type-year fixed effects could lead to biased estimates of the coefficient of interest. We weight each observation by the mean number of enrollees across all years for the relevant plan.

Estimating equation (2) yields 12 coefficients (i.e., on the market-year interactions) for each market; interactions with 1997 are omitted. We set the premium index equal to 100 for each market in 1997 and apply the estimated coefficients on the market-year interactions to calculate the index in all subsequent years. (For example, a

a market before and after a conversion. While this would alleviate the selection and small sample issue to a degree (as all plans present before and after a conversion can identify this coefficient, rather than just plans offered by converting BCBS carriers), there would be too few such plans with a sufficiently lengthy panel to permit an analysis of pre-conversion trends or to capture the effect of conversion more than a year or two out.

market-year coefficient of 0.2 would imply an index of $100 \times (\exp(0.2)) = 122.14$). Descriptive statistics for the premium index, which is constructed separately for FI and SI plans, are presented in Table 2. Premium growth is very similar for both insurance types with the (unweighted) mean market premium index reaching approximately 290 in both the FI and SI samples by 2009. This increase (i.e., 190 percent) compares to a nominal increase of 140 percent in the average family premium for large firms (200+ employees) as calculated from KFF/HRET survey data during roughly the same period (1999–2010 rather than 1997–2009).[22] Given our price index holds product features, such as carrier identity and plan generosity constant, we anticipate steeper growth than would be observed from a simple comparison of unadjusted premiums over time. In the face of rising insurance premiums, employers have substituted toward cheaper plans, so that realized price growth is lower than predicted price growth holding plan characteristics constant.

We also estimate a version of equation (2) which permits separate estimates of the market-year coefficients for BCBS and non-BCBS plans (by interacting indicators for each with the set of market-year dummies). We exponentiate the two sets of coefficient estimates to form separate price indices for BCBS and non-BCBS plans and use these to study the differential effects of the BCBS conversions on converting plans and their rivals. Again, we repeat this process separately for the sample of fully insured and self-insured plans.

B. *Effect of Conversions on Local Market FP Share*

To assess the impact of conversions on market-level FP share, we begin by estimating a specification including leads and lags of $BCBS\ FP_{m,t}$, an indicator variable that takes a value of one if the BCBS affiliate in the relevant market and year has converted to FP status:

$$(2) \quad FP\ share_{mt} = \phi_0 + \phi_1 BCBS\ FP_{mt-3} + \phi_2 BCBS\ FP_{mt-2} + \phi_3 BCBS\ FP_{mt-1} + \phi_4 BCBS\ FP_{mt} + \phi_5 BCBS\ FP_{mt+1} + \phi_6 BCBS\ FP_{mt+2} + \phi_7 BCBS\ FP_{m,t \geq t+3} + \psi_m + \delta_t + \Gamma X_{mt-1} + \varepsilon_{mt}.$$

The purpose of this model is twofold: first, to confirm that the leads are statistically insignificant and lack a pronounced trend; second, to explore the persistence of the (initially pro forma) post-merger increase in FP share.[23] The specification includes market (ψ_m) and year fixed effects (δ_t), as well as time-varying, lagged covariates that are relevant for the reduced-form specification of premiums, described in Section IIIC. We estimate the model separately for the fully and self-insured markets, weighting each observation by the total number of insured enrollees in the

[22] See Employer Health Benefits 2010 Annual Survey, Exhibit 1.12, downloadable at http://ehbs.kff.org/pdf/2010/8085.pdf.

[23] Each lead or lag "turns on" only once, i.e., $BCBS\ FP_{mt-3}$ takes a value of one three years prior to the conversion affecting market m.

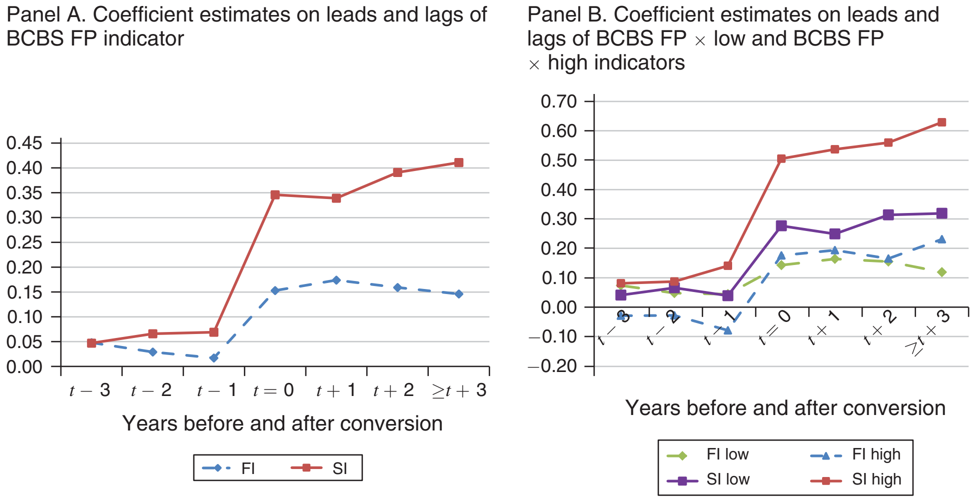

Figure 3. Effect of BCBS Conversions on For-Profit Market Share (*leads and lags specifications*)

Note: Coefficient estimates are presented in Table A5, panels A and B.

relevant market and year, and including data from 1998–1999.[24] Figure 3, panel A graphs the coefficient estimates (presented in Table A5, panel A) separately for FI and SI plans. The figure shows discrete jumps in FP share in the year of a conversion ($t = 0$) with a larger jump in the SI segment. The increase in FP share persists thereafter and creeps upward for SI plans.

Table 3 reports coefficient estimates from a more parsimonious specification, in which the leads and lags of $BCBS\ FP_{m,t}$ are replaced with $BCBS\ FP_{m,t}$ alone. Given the inclusion of market fixed effects, the coefficient on this indicator is identified from markets experiencing conversions. We make one additional modification to this specification: we lag both the dependent and independent variables by a year for the sake of symmetry with the reduced-form specification presented alongside (and described later in this paper). Table 3 includes two sets of p-values beneath each coefficient estimate. The first set is obtained from the usual two-sided t-tests, using cluster-robust standard errors (with the market as the clustering variable). Given the relatively low number of clusters (47 in total and 28 treatment markets), we also present p-values obtained using the wild-cluster bootstrap-t method (again with the market as the clustering variable) as described in Cameron, Gelbach, and Miller (2008).

Column 1 reveals that, on average, conversions are followed by increases in FP market share of 14.5 percent (FI) and 33.8 percent (SI). Next, we confirm that these increases vary systematically with the pre-conversion market share of converting affiliates, calculated as the enrollment-weighted average market share of the converting plan during the three years preceding conversion. Figure 4 documents the

[24] We exclude data from 1997 for symmetry with subsequent specifications. Premiums are a function of lagged covariates, and the earliest year of premium data we use is for 1998.

TABLE 3—EFFECT OF BCBS CONVERSIONS ON FOR-PROFIT SHARE AND PREMIUMS

Dependent variable:	Lagged FP share Mean = 0.61			Premium index Mean = 186.70		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Fully insured plans						
Lagged BCBS FP	0.145	0.049		4.252	−13.104	
	(0.002)	(0.549)		(0.383)	(0.139)	
	[0.002]	[0.522]		[0.474]	[0.248]	
Lagged BCBS FP × Pre-conversion share		0.578			104.498	
		(0.103)			(0.003)	
		[0.170]			[0.032]	
Lagged BCBS FP × Low pre-conversion share			0.113			−0.044
			(0.016)			(0.994)
			[0.008]			[0.954]
Lagged BCBS FP × High pre-conversion share			0.246			17.705
			(0.000)			(0.001)
			[0.000]			[0.002]
Observations	552	552	552	552	552	552
Dependent variable:	Lagged FP share Mean = 0.62			Premium index Mean = 186.70		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B. Self-insured plans						
Lagged BCBS FP	0.338	0.033		3.154	−3.658	
	(0.000)	(0.307)		(0.307)	(0.507)	
	[0.000]	[0.380]		[0.388]	[0.604]	
Lagged BCBS FP × Pre-conversion share		1.722			38.481	
		(0.000)			(0.070)	
		[0.000]			[0.158]	
Lagged BCBS FP × Low pre-conversion share			0.265			2.122
			(0.000)			(0.544)
			[0.000]			[0.642]
Lagged BCBS FP × High pre-conversion share			0.501			5.317
			(0.000)			(0.092)
			[0.000]			[0.120]
Observations	564	564	564	564	564	564

Notes: The unit of observation is the market year. All models include fixed effects for each market and year as well as lagged market-year controls—ln(Medicare costs per capita) and the unemployment rate—and are estimated by weighted least squares using the average number of enrollees in each market as weights. The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in (·) and [·], respectively. In both cases, the clustering unit is the market.

significant variation in pre-conversion share across markets, calculated using the combined FI+SI sample.[25] Pre-conversion share ranges between 6 percent and 35 percent with an enrollment-weighted average of 20 percent. These shares are

[25] We used a combined sample in order to reduce noise in the share estimates and because the combined share is a driver of negotiated reimbursement rates, which in turn feed into self and fully insured plan premiums. As we report in Section IIID, results are robust to using sample-specific market shares.

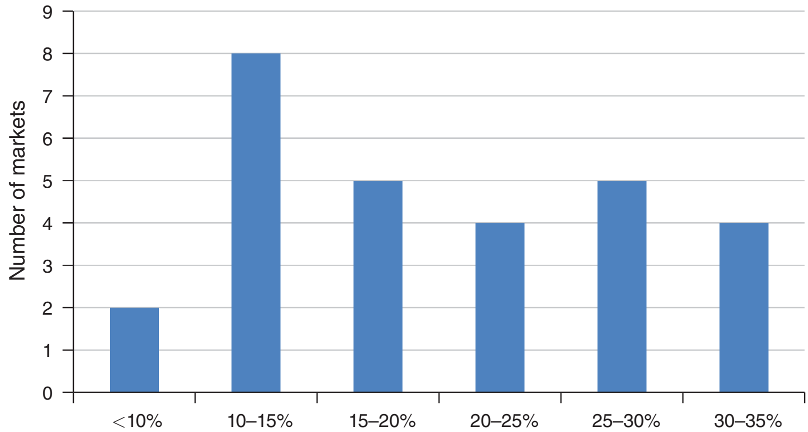

Figure 4. Distribution of Pre-conversion BCBS Market Share

Notes: Pre-conversion BCBS share is computed using LEHID and refers to the enrollment-weighted average market share of the converting BCBS plan during the three years preceding conversion. Also, $N = 28$.

lower than BCBS shares reported by other sources. There are two reasons for this difference: (i) multisite firms (which are heavily represented in LEHID) are more likely to utilize carriers offering plans nationwide (e.g., Aetna, CIGNA), and to do this via BCBS requires coordination across many affiliates; (ii) BCBS typically has larger market share in the individual and small group segments than in the large group segment, owing in part to its historical mission of ensuring broad access to medical care (Abelson 2013).

Column 2 of Table 3 reports the results obtained when adding an interaction between $BCBS\ FP_{m,t-1}$ and *pre-conversion share*. As expected, the coefficient on this interaction is large and positive in both the FI and SI samples, although it is imprecisely estimated in the former. Subdividing conversions into those with "high" versus "low" market share, using the weighted average of 20.2 percent as the cutoff, yields greater precision, particularly in the FI sample.[26] Markets with high pre-conversion share saw increases in FP share of 25 percent and 50 percent in the FI and SI samples, respectively, whereas markets with low pre-conversion share saw increases of 11 percent and 27 percent, respectively. Results from estimating specification (2), but with leads and lags included separately for "high" and "low" markets, are graphed in Figure 3, panel B and presented in Table A5, panel B. There are no differential trends in the market-level FP share for these markets relative to non-converting markets in the years preceding conversion.

[26] Note the classification of markets is the same using weighted or unweighted averages or medians.

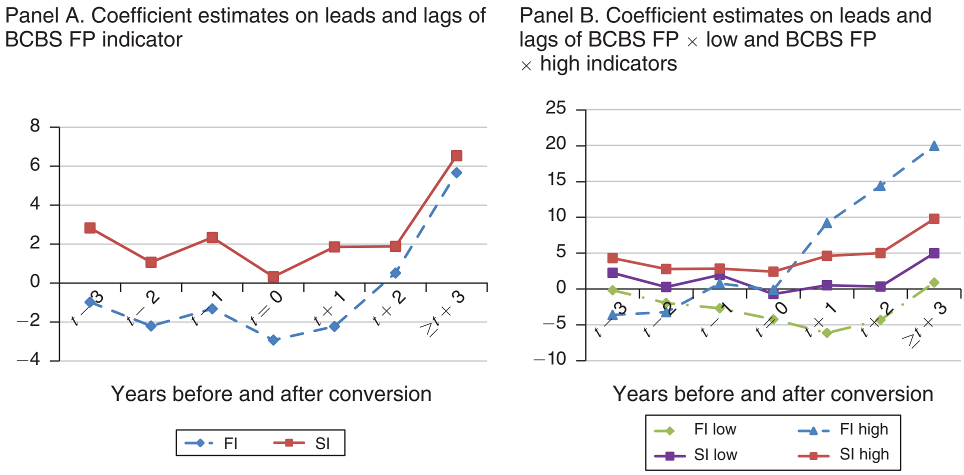

Figure 5. Effect of BCBS Conversions on Premiums (*leads and lags specifications*)

Note: Coefficient estimates are presented in Table A6, panels A and B.

C. *Reduced-Form Models: How Did BCBS Conversions Affect Premiums?*

We again begin by estimating a specification including leads and lags of $BCBS\ FP_{m,t}$:

$$\text{(3)}\quad premium\ index_{mt} = \phi_0 + \phi_1 BCBS\ FP_{mt-3} + \phi_2 BCBS\ FP_{mt-2} + \phi_3 BCBS\ FP_{mt-1} + \phi_4 BCBS\ FP_{mt} + \phi_5 BCBS\ FP_{mt+1} + \phi_6 BCBS\ FP_{mt+2} + \phi_7 BCBS\ FP_{m,t\geq t+3} + \psi_m + \delta_t + \Gamma X_{mt-1} + \varepsilon.$$

All variables are the same as in equation (2). The market-year controls included in X_{mt-1} are the local unemployment rate and ln(Medicare spending per capita). During recessions, insurance take-up is lower (albeit not dramatically so in the large group market), leading to greater adverse selection and higher insurance premiums. We include Medicare spending as a rough measure of local medical spending.[27] Because premiums for year t are determined in year $t-1$, we lag both of these variables.

The coefficient estimates for both the FI and SI samples are graphed in Figure 5, panel A and presented (along with standard errors) in Table A6, panel A. We find no evidence of differences in premium trends for markets with/without successful conversions in the years preceding the conversions. Indeed, none of the leads are

[27] Recent studies find Medicare and private commercial spending are positively correlated (Newhouse et al. 2013, Cooper et al. 2018), but this correlation is relatively small, i.e., around 0.14.

statistically significant. These results support the key identifying assumption that the success of a BCBS conversion attempt is orthogonal to omitted determinants of premiums. There is an uptick in premiums two or three years post-conversion, but none of the coefficient estimates are individually significant, and neither is the coefficient on a single *post*-period dummy (as discussed later in this paper and reported in Table 3, column 4).

Next, we estimate models including a full set of leads and lags for $BCBS\ FP_{mt} \times high$ and $BCBS\ FP_{mt} \times low$, where again *high* (*low*) is an indicator variable which takes a value of one in markets where the pre-conversion BCBS share is higher (lower) than the weighted average. All else equal, larger BCBS carriers should be in a stronger position to raise prices following a conversion because their enrollees have fewer outside options (i.e., these carriers face lower elasticities of demand).[28] However, if dominant converting plans are more successful in lowering costs, optimal prices could fall. Note that either effect will be magnified in markets where BCBS accounts for a greater share of enrollees, both for mechanical reasons and due to competitive responses to BCBS' actions.

The results, graphed in Figure 5, panel B and listed in Table A6, panel B, again show fairly stable pre-conversion trends. However, in the year following conversion, FI premiums in *high* markets surge, while FI premiums in *low* markets continue a slow, steady decline which begins to reverse two years after conversion. SI premiums in both *high* and *low* markets exhibit slower, smaller premium increases in the post-conversion period.

Table 3 also presents the results from parsimonious reduced-form models, e.g.,

$$(4) \qquad premium\ index_{mt} = \alpha + \phi BCBS\ FP_{mt-1} + \psi_m + \delta_t + \Gamma X_{mt-1} + \varepsilon.$$

Column 4 shows that FP conversions did not have a statistically significant effect (on average) on premiums during the pooled post-period. Next, we add the continuous interaction between $BCBS\ FP_{m,t-1}$ and *pre-conversion share*. The results indicate that post-conversion FI premiums increased significantly more in markets with higher pre-conversion market share. Last, we report the results from a specification including interactions between the *high* and *low* pre-conversion share indicators and $BCBS\ FP_{mt-1}$. We find strong evidence of premium increases in *high* markets, but noisy and small point estimates in *low* markets. The post-conversion increase for FI plans is estimated at 18 points ($p < 0.01$), which is roughly 13 percent of the FI premium index of 135 in 2001 (the modal pre-conversion year). The increase in SI premiums for *high* markets is marginally significant and smaller: 5 points, amounting to 4 percent of the SI premium index of 127 in 2001.

In sum, conversions of BCBS affiliates with high market share lead to substantial premium increases for FI plans and smaller, marginally significant increases for SI plans. As discussed in Dafny (2010), the opportunity to exercise market power is smaller in the SI segment, which is served by a larger number of competitors and

[28] Of course, the optimal change in price depends on the initial price level as well as competitive conditions. We have no a priori prediction regarding the relative prices charged by BCBS plans with large versus small market pre-conversion shares.

Table 4—Effect of BCBS Conversions on Premiums: BCBS versus Non-BCBS Plans

Dependent variable:	Premium index (BCBS) Mean = 184.1			Premium index (non-BCBS) Mean = 189.8		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Fully insured plans						
Lagged BCBS FP	4.011	−23.891		0.252	−18.964	
	(0.596)	(0.161)		(0.957)	(0.031)	
	[0.646]	[0.212]		[1.000]	[0.100]	
Lagged BCBS FP × Pre-conversion share		166.021			115.702	
		(0.046)			(0.004)	
		[0.080]			[0.036]	
Lagged BCBS FP × Low pre-conversion share			−0.890			−4.450
			(0.918)			(0.392)
			[0.892]			[0.488]
Lagged BCBS FP × High pre-conversion share			18.717			14.981
			(0.124)			(0.009)
			[0.162]			[0.018]
Observations	525	525	525	538	538	538
Dependent variable:	Premium index (BCBS) Mean = 184.15			Premium index (Non-BCBS) Mean = 189.76		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B. Self-insured plans						
Lagged BCBS FP	1.333	−9.049		4.549	−7.356	
	(0.789)	(0.328)		(0.161)	(0.238)	
	[0.804]	[0.466]		[0.192]	[0.334]	
Lagged BCBS FP × Pre-conversion share		58.637			67.249	
		(0.098)			(0.039)	
		[0.162]			[0.076]	
Lagged BCBS FP × Low pre-conversion share			−1.120			2.088
			(0.842)			(0.523)
			[0.936]			[0.616]
Lagged BCBS FP × High pre-conversion share			6.838			9.885
			(0.179)			(0.038)
			[0.192]			[0.056]
Observations	557	557	557	564	564	564

Notes: The unit of observation is the market year. All models include fixed effects for each market and year as well as lagged market-year controls—ln(Medicare costs per capita) and the unemployment rate—and are estimated by weighted least squares using the average number of enrollees in each market as weights. The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the market.

characterized by greater transparency in pricing. Price increases not associated with provider outlays are easily observed in the SI market.

Last, we contrast the post-conversion pricing responses of BCBS and non-BCBS plans by estimating the specifications in Table 3 using *BCBS Index* and *non-BCBS Index* as the dependent variables. The results are displayed in Table 4, again separately for FI plans (panel A) and SI plans (panel B). We find that post-conversion premiums for both converting BCBS affiliates and their rivals increase in BCBS shares, although the estimates are less precise for BCBS premiums. When we augment the sample to include all 47 states whose BCBS plans had not converted as of the start of the study period, the magnitude of the premium increase for fully insured BCBS plans in *high* markets is a third larger than that of non-BCBS plans in these markets (and both are statistically significant).[29] This pattern is consistent with newly for-profit BCBS affiliates assuming the role of the price leader in markets where they are large enough to do so.

D. *Robustness Checks and Extensions*

To explore the sensitivity of our findings, and to uncover other potentially interesting phenomena, we considered several alternative sample restrictions and specifications. First, we confirmed the robustness of our key findings to the following modifications: (i) limiting the study period to 1997–2005, as the conversions were all complete by 2002, and *demographic factor* (a highly significant predictor of premium levels) was redefined for 2006 onward; (ii) dropping market years with fewer than 20 sampled employers so as to minimize the influence of noisy estimates of the premium index and market shares; (iii) dropping all controls (apart from market and year fixed effects); (iv) using the untransformed market-year coefficient estimates as the price index (i.e., not exponentiating them); (v) using insurance-type-specific BCBS market shares to classify markets. In all of these specifications, we confirm a large, statistically significant increase in FI premiums in *high* markets and no significant impact in *low* markets.

Next, we examined the sensitivity of the point estimates to excluding one conversion at a time (i.e., dropping all markets affected by a given conversion). The results (available in the online Appendix) show the effect of conversion on FI premiums in *high* markets is consistently large and statistically significant, and the effect on FI premiums in *low* markets is consistently small and imprecisely estimated.

Our models focus on conversions of BCBS plans to FP stock companies. This definition pools together conversions from NFP to FP stock company (three states) and from Mutual to FP stock company (eight states) and does not consider another ownership conversion, that of NFP to Mutual (four states). We estimated models utilizing these three distinct ownership conversions. The results (in Table A7) reveal that conversions from NFP to Mutual had no statistically significant impact on premiums at least during the short post-period we observe for these conversions. (In our study period, all four affiliates switching from NFP to Mutual status converted to FP status 2–3 years later.) Conversions from NFP to FP were followed by a statistically insignificant *decrease* in FI premiums (-6 points, with a cluster-robust standard error of 6), whereas conversions from Mutual to FP stock company resulted in a significant increase in FI premiums (12 points, with a cluster-robust standard error of 5). Dropping the three states with NFP to FP conversions therefore strengthens the primary results; however, given our research objective (studying the effect of investor ownership on insurance-related outcomes), we retain these states in our

[29] The online Appendix presents all regression results obtained using the broader sample of 47 states.

estimation sample. In column 2, we add interactions between Mutual → FP and *pre-conversion share*; there are too few observations to do the same for NFP → FP. The results confirm the same pattern obtained using our broader conversion definition, although the coefficient on the interaction term is only borderline statistically significant (p-value around 0.10): FI premiums increased more in areas with higher *pre-conversion share*. Results in the SI sample are smaller and more noisily estimated, as before.

Last, we explored the effects of BCBS conversions on two other dependent variables, *plan design factor*, and the share of enrollees in SI plans. Both are measured at the market-year level (the former separately for FI and SI samples). We find no statistically or economically significant effects of conversions on *plan design*, implying that employers did not adjust this lever in the wake of post-conversion price increases in the FI market. Surprisingly, neither did they increase their reliance on SI plans. In fact, there is a slight *decrease* in the share of enrollees in SI plans in *high* markets following conversion.[30]

IV. Effects of Ownership Status on Non-price Outcomes

In this section, we evaluate the impact of ownership status on insurance coverage and medical spending (as a share of premium revenues). Both of these measures capture a broader swath of the population than is reflected in the premium analysis, which is limited to large employers.

A. *Are Not-for-Profits Insurers of Last Resort?*

Not-for-profits frequently claim to be insurers "of last resort"; indeed, this phrase is commonly applied to BCBS plans and appears in the statutes of some states. NFPs may serve the community by pricing below profit-maximizing levels (particularly in the high-risk non-group market, where access is low) and (during the study period) by offering policies to individuals and small groups, who other insurers would reject.[31]

In order to assess the impact of FP insurers on coverage rates, we make use of annual state-level data on various sources of coverage: employer-sponsored, individual, Medicaid, and other. All measures are expressed as a share of the under-65 population in the relevant state and year and are estimated using the Current Population Survey (CPS) March Uniform Extracts compiled by the Center for Economic and Policy Research (CEPR) for data years 1999–2009.[32] Summary statistics are included in Table A4. The insurance categories are not mutually exclusive as some individuals report coverage through multiple sources.

[30] Specifically, the point estimate in *high* markets is −0.029 (with a market cluster-robust standard error of 0.014). The mean SI share in 2001 is 0.672.

[31] Under the Affordable Care Act, as of 2014, insurers of any ownership form were no longer permitted to reject applicants on the basis of health status, to impose preexisting condition exclusions, or to charge premiums varying more than 3:1 by age and 1.5:1 by smoking status.

[32] We do not include data from 1997 and 1998 because the CPS survey methodology changed in March 2000, generating discontinuous changes in insurance coverage between 1998 and 1999. (Note that the CPS March survey pertains to data from the preceding year.)

Table 5—Impact of For-Profit Penetration on Insurance Coverage

	Panel A. Dependent variable = Share insured Mean = 0.86			*Panel B. Dependent variable = Share with employer-sponsored insurance* Mean = 0.68		
	(1)	(2)	(3)	(4)	(5)	(6)
Lagged BCBS FP	0.004	0.025		(0.846)	0.025	
	(0.595)	(0.223)		(0.846)	(0.181)	
	[0.574]	[0.314]		[0.918]	[0.228]	
Lagged BCBS FP × Pre-conversion share		−0.112			−0.136	
		(0.217)			(0.136)	
		[0.344]			[0.204]	
Lagged BCBS FP × Low pre-conversion share			0.006			0.003
			(0.525)			(0.722)
			[0.562]			[0.714]
Lagged BCBS FP × High pre-conversion share			0.000			−0.008
			(0.985)			(0.502)
			[0.976]			[0.606]
Observations	209	209	209	209	209	209
	Panel C. Dependent variable = Share individually insured Mean = 0.09			*Panel D. Dependent variable = Share on Medicaid* Mean = 0.12		
	(1)	(2)	(3)	(4)	(5)	(6)
Lagged BCBS FP	−0.000	0.012		0.013	0.001	
	(0.928)	(0.169)		(0.041)	(0.940)	
	[0.958]	[0.222]		[0.050]	[0.920]	
Lagged BCBS FP × Pre-conversion share		−0.064			0.061	
		(0.141)			(0.345)	
		[0.248]			[0.432]	
Lagged BCBS FP × Low pre-conversion share			0.001			0.010
			(0.800)			(0.153)
			[0.734]			[0.204]
Lagged BCBS FP × High pre-conversion share			−0.002			0.016
			(0.591)			(0.063)
			[0.706]			[0.176]
Observations	209	209	209	209	209	209

Notes: The unit of observation is the state year. The study period is 1999–2009. Insurance rates and pre-conversion share are scaled from zero to one. All specifications include state and year fixed effects, simulated Medicaid eligibility rate for children under 18, lagged ln(Medicare costs per capita), and the lagged unemployment rate. Each observation is weighted by the average under-65 population in the state. The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in (·) and [·], respectively. In both cases, the clustering unit is the state.

We estimate specifications analogous to those presented in Section III, replacing the dependent variables with various measures of insurance coverage. Due to the short pre-conversion period, we do not estimate the full leads and lags specifications. We aggregate the market-year controls to the state-year level and add *simulated Medicaid eligibility*, a summary measure of state-year policies determining Medicaid eligibility for children under 18. This measure, constructed as per Currie and Gruber (1996) and Gruber and Simon (2008), controls for changes in insurance rates associated with state-specific changes in Medicaid eligibility criteria. We weight each observation by the under-65 population in the corresponding state year.

Table 5 presents results from reduced-form models analogous to the premium models in Table 3. Each panel corresponds to a different dependent variable: share of non-elderly with any insurance (panel A), employer-sponsored insurance (panel B), individual insurance (panel C), and Medicaid (panel D).[33] We divide states into *high* and *low* using the mean state-level BCBS pre-conversion market share (19.4 percent). The key result arising from these regressions is a statistically significant increase in Medicaid enrollment following conversion. The point estimate implies that Medicaid enrollment increased by 1.3 percentage points in states experiencing conversions, relative to an average Medicaid enrollment rate of 12 percent. This effect appears to be stronger in *high* markets (a coefficient of 0.016 versus 0.010 for *low* markets), although we cannot reject equality of the coefficient estimates. *High* markets experience small and insignificant reductions in employer-sponsored and individual insurance, which appear to offset the Medicaid increase, yielding a net zero effect on the share of the non-elderly with any insurance.

To better understand the mechanism generating the post-conversion changes in insurance coverage, we estimated separate models for children under 18 and adults between 18 and 44, an age range which should capture most parents with children at home. Results are presented in Table A8. Given the small number of observations, the coefficient estimates are again imprecise, particularly using the wild-cluster bootstrap-*t* method. However, the results suggest children in *high* converting markets experienced the largest increases in Medicaid enrollment.[34] Medicaid enrollment also increased for the parent-aged population in states undergoing conversions. In addition, there is (noisy) empirical evidence that private insurance coverage declined more in states with larger conversions, consistent with families being priced out of the market.[35] The results are very similar when estimated using data from the broad sample of 47 states.

In sum, we do not find that BCBS conversions adversely affected uninsurance rates, a result echoed by several of the conversion case studies (e.g., Conover, Hall, and Ostermann 2005). However, conversions followed by premium increases did result in higher Medicaid enrollment. If conversions are representative of typical (exogenous) changes in local FP penetration, the results suggest that higher FP penetration crowds out private insurance coverage—at least in the era preceding the Affordable Care Act (ACA).

[33] In the interest of space, we do not include results for "other public insurance." Across all states and years, the weighted average rate of "other public insurance" is 0.065. The coefficients of interest for this category are consistently small and statistically insignificant.

[34] In the sample of children under 18, a *t*-test rejects the null of equal coefficients in the *low* and *high* markets in favor of $high > low$ at $p = 0.052$, using cluster-robust standard errors.

[35] These models pool individual and employer-sponsored coverage. In both age groups, the continuous interaction between the BCBS FP indicator and *pre-conversion share* is negative and statistically significant at $p < 0.10$, using cluster-robust standard errors, and $p < 0.13$, using the wild-cluster bootstrap-*t* method.

Table 6—Impact of For-Profit Penetration on Medical-Loss Ratios

	Dependent variable = MLR		
	All insurers Mean = 0.85	BCBS Mean = 0.84	Non-BCBS Mean = 0.85
Lagged BCBS FP	0.020	−0.011	0.052
	(0.156)	(0.519)	(0.007)
	[0.227]	[0.590]	[0.014]
Observations	162	162	157

Notes: The unit of observation is the state year. The study period is 2001–2009. MLRs are constructed using censored insurer state-year data. All specifications include state and year fixed effects, the lagged unemployment rate, and lagged ln(Medicare costs per capita). Each observation is weighted by the average number of LEHID enrollees in the state. Alaska does not report data for non-BCBS plans until 2008, hence the discrepancy between the number of BCBS and non-BCBS observations. The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the state.

B. *Does Ownership Status Affect Medical-Loss Ratios?*

Next, we examine the impact of conversions on insurer medical-loss ratios (MLRs), defined as the share of (post-tax) premium revenue disbursed for medical claims, as opposed to profits or administrative expenses. As noted in Section II, we calculate MLRs by state and year, first for all insurers and then separately for BCBS and non-BCBS insurers. The data are available from 2001 to 2009 and pertain only to FI plans. We limit the sample to state years with non-missing MLR data for the primary BCBS affiliate. We estimate reduced-form specifications analogous to equation (4), again using the state year as the unit of observation.[36] We include our standard controls (unemployment rates and log of Medicare spending), aggregated to the state-year level.

The results are displayed in Table 6. Column 1 shows that aggregate MLRs were unaffected by the BCBS conversions, on average. However, column 3 shows that MLRs for rivals of converting BCBS affiliates rose by 0.05, on average, relative to a base of 0.89 in 2001. (Arguably, it is more appropriate to consider this increase relative to (1-MLR); 5 percent is a substantial reduction in potential profits for an insurer.) Column 2 shows a noisily estimated decline in BCBS MLRs. Unfortunately, our data include only two states with high pre-conversion shares, hence we cannot compare effects by high/low status. As a robustness check, we reestimated all models dropping one converting state at a time; coefficient estimates and standard errors were very similar across these models.

One explanation for the results is that newly for-profit BCBS plans may have engaged in greater efforts to screen out individuals with high costs. Such an effort would simultaneously raise MLRs for competitors as high-cost enrollees shifted to their plans, reduce MLRs for BCBS plans, and leave aggregate MLRs unchanged.

[36] We are unable to estimate specifications to check for pre-trends in MLR as the sample only includes data for a very short pre-conversion period.

V. Discussion and Conclusions

Research on the performance of the commercial health insurance sector in the United States is relatively sparse, particularly in light of the sector's role in managing and financing health-care consumption and the individual mandate to carry health insurance (and to purchase it from commercial carriers if ineligible for public insurance). Fortunately, a spate of recent studies explores competition among health insurers across a wide range of customer segments, including the Medigap market (Starc 2014), Medicare Part D (Ho, Hogan, and Morton 2017; Stocking et al. 2014), Medicare Advantage (e.g., Duggan, Starc, and Vabson 2016; Pizer and Frakt 2002), group insurance markets (e.g., Ho and Lee, 2017; Dafny, Duggan, and Ramanarayanan 2012), and individual health insurance exchanges (e.g., Shepard 2016; Dafny, Gruber, and Ody 2015). However, recent studies have not systematically evaluated the role of ownership form, notwithstanding widespread public concern about the for-profit motive (in health care broadly, and among insurers in particular). This concern was epitomized in the ACA's co-op program, which allocated federal funds to support de novo *not-for-profit* cooperative insurers who were supposed to offer plans in each individual insurance exchange nationwide. Whether and how for-profits differ—and whether different regulations should pertain (beyond those implemented via tax authorities)—is a question that has recently resurfaced in the wake of entries and exits in the Health Insurance Marketplaces by national for-profit carriers.

In this paper, we study the impact of for-profit ownership on key outcome measures (at the market level) by examining the aftermath of 11 conversions of BCBS affiliates, characterized by different pre-conversion local market shares across 28 distinct geographic markets. We study not only the conduct of the converting BCBS affiliates, but also the impact on rival carriers.

We find heterogeneous effects that depend on the magnitude of the converting BCBS affiliate's market share. Specifically, fully insured premiums increased roughly 13 percent when converting BCBS plans had shares in excess of the mean pre-conversion BCBS share (20 percent in our sample) and exhibited no change when pre-conversion share fell below the mean. Importantly, we do not observe different pre-conversion price trends in markets ultimately experiencing conversions relative to control markets whose BCBS affiliates attempted but failed to convert, nor in markets experiencing relatively sizeable conversions (relative to markets without conversions or with smaller conversions). Assuming no disproportionate quality changes by large BCBS affiliates (a possibility we discuss and discount later in this paper), these results suggest a post-conversion exercise of market power. Significantly, rivals of these large converting insurers also raised their prices following the conversions.

We find that BCBS conversions had no significant impact on state-level uninsurance rates (among the non-elderly). However, Medicaid enrollment increased an average of 10 percent in these states, suggesting crowd-out of private insurance coverage. This enrollment increase was concentrated in the population under 18, and the pattern of changes in private insurance coverage is consistent with a scenario in which parents faced premium increases and subsequently dropped private family coverage. Conversions had no impact on state-level MLRs, but again there was a compositional effect in the responses. MLRs increased for rivals of converting plans and decreased for the converting plans themselves (although the decrease is not statistically significant). This pattern is consistent with a shift of high-risk enrollees from converting plans to rivals.

As noted earlier, it is theoretically possible that post-conversion premium increases are partly—or even wholly—explained by post-conversion quality improvements that increased the demand or willingness to pay for insurance. For example, converting affiliates may have increased members' electronic access to health claims, invested in speedier claim processing, and/or expanded their provider networks. However, the data are not particularly supportive of this alternative explanation for two reasons. First, prices increased almost immediately following large conversions. It seems unlikely that quality improvements could be implemented and conveyed to the marketplace so rapidly. Second, for the rival pricing effect to be consistent with quality improvements, one would have to believe that rivals made quality improvements of similar market value as BCBS in all markets (i.e., greater improvements where BCBS was relatively more dominant and smaller improvements where BCBS was smaller) to conclude that they were not following the price leadership of BCBS where it was exhibited. Given the challenges associated with generating and marketing changes in quality, as well as the fact that most rivals to BCBS in our sample are national firms who would have found such adjustments more difficult to calibrate, we conjecture that quality improvements likely did not account for all of the observed price increases following conversions.

We also performed an empirical exercise to explore whether quality—as inferred from consumer choices—of converting Blues increased more. Following Dafny, Ho, and Varela (2013), we estimated a structural (logit) model of employee plan choice based on a utility function that includes a large set of individual, plan, and market characteristics (and their interactions). We allowed for separate unobserved quality effects for Blue Cross carriers in every market and year and estimated our key specifications (equation (4) and subsequent variants) using the estimates of these quality effects as the dependent variable. We find no evidence that conversions are associated with quality improvements and weak evidence of reductions (especially in markets with high pre-conversion BCBS share).[37] In addition, we note that although the welfare implications are different if the price increases were accompanied by quality improvements valued by the market, many regulators and consumers are particularly concerned about reining in spending growth and ensuring access to affordable coverage.

Looping back to the theoretical models of NFP and FP health-care organizations, the findings are consistent with models in which NFPs prioritize enrollment over profits (equivalently, models in which FPs prioritize profits over enrollment). While theoretically this difference in emphasis might not manifest in higher premiums or lower quality because FPs could be more efficient and find it optimal to maintain substantially the same premiums and quality as NFPs (and still reap higher profits via lower

[37] Results are reported in the online Appendix.

operating costs and/or medical expenses), empirically we do find there is a trade-off: consumers face higher premiums when large NFPs convert to FP status. Although we do not directly study quality, we find no indirect evidence of quality improvements as inferred from a model of employee health-plan choice. Moreover, we do find evidence that rivals of converting plans experienced sizeable increases in medical spending following conversion, a result that suggests FPs are likelier than NPs to engage in risk selection practices (e.g., denying or deterring enrollment of individuals with poor health or high health risk, a practice that was legal during the study period).

Two caveats to our findings are in order. First, our premium results derive from a sample of plans in the large group insurance market, hence our point estimates may not pertain to the small group and individual insurance markets. However, the insurance coverage results are suggestive of premium increases in these markets as well. Given that large employers' decisions to offer insurance are fairly insensitive to premium changes (Gruber and Lettau 2004) and that insurance take-up conditional on an offer of coverage is also relatively insensitive to premium changes (Cutler 2003, Gruber and Washington 2005), the population served by small-group and individual policies is likely driving the Medicaid crowd-in.

Second, the results must be construed in light of the natural experiments we evaluate. The change in conduct of converting BCBS plans may not reflect the average difference between new or existing NFP and FP carriers. However, we believe the results are valuable for policy going forward. Entry into the insurance industry is rare and success even rarer. Some NFP insurers are at risk of needing capital infusions to stay solvent or have explicitly expressed an interest in converting to FP status. Thus, structurally induced changes in FP share are likeliest to derive from conversions of NFP insurers. In this context, our analysis offers estimates of the impact of future changes in FP share on insurance premiums.

Our findings have several implications for regulatory and competition policy vis-à-vis insurers. First, it appears that sizable FP insurers are more likely to exercise market power via price increases than are comparable NFP insurers. Second, pricing actions by large insurers have a ripple effect on rivals' prices, further solidifying the evidence of oligopolistic conduct in many local insurance markets. Third, there is no evidence that NFP and FP insurers charge different prices in the large group market when both are relatively small. Future research on the impact of ownership form on dimensions beyond price—especially those related to quality, innovation, and risk selection—is essential for obtaining a comprehensive perspective of the impact of for-profit ownership on insurer conduct.

Appendix: The National Association of Insurance Commissioners' (NAIC) Dataset

The NAIC is an umbrella organization of state-level insurance regulators.[38] Because states regulate fully insured products, NAIC data represent only the FI

[38] For all key lines of insurance (including health), NAIC provides uniform reporting forms called "insurance blanks." Insurers complete the blanks separately by state and file them with the respective state authorities, who pass the data on to NAIC.

component of the health insurance market. Insurers report data by product line and state; Washington, DC is included in the data but California is not. We construct a single MLR for each insurer state year, including only spending and premiums associated with comprehensive commercial medical insurance, and omitting observations with negative values for either variable.[39] We drop observations in the 5 percent tails of the annual distribution of insurer state-year MLRs and aggregate the remaining data to construct state-year MLRs. Finally, we exclude 9 state-year observations in which the principal BCBS affiliate does not report data to NAIC.[40] The final estimation sample includes 162 observations, out of a hypothetical maximum of 171 (19 states $\times$ 9 years). For additional details on the NAIC data, as well as other sources of insurance data, see Dafny et al. (2011).

Table A1—Ownership Conversions of BCBS Affiliates, 1997–2009

	Conversion from NFP to mutual	Conversion from mutual to FP	Conversion from NFP to FP
Colorado	1999	2002	
Connecticut		2002	
Indiana (Accordia)		2002	
Kentucky		2002	
Maine	2000	2002	
Nevada	1999	2002	
New Hampshire	2000	2002	
Ohio (CMIC)		2002	
New York (Empire)			2003
Wisconsin			2001
Missouri			2001

Note: Entries refer to the first post-conversion year as coded in our dataset.

39 These categories are excluded: Medicare and Medicaid plans, Medicare supplemental plans, dental plans, vision-only plans, long-term care, disability income, stop-loss, and other.

40 These are Nevada in 2001, Ohio in 2001–2003, and Indiana in 2001–2005.

Table A2—Descriptive Statistics, Plan-Year Data

	1997–2009	1997	2009
Panel A. Fully insured plans			
Premium ($)	5,499.95	3,555.47	9,196.51
	(2,401.11)	(823.66)	(2,913.91)
Number of enrollees	184.88	173.72	184.45
	(537.61)	(457.59)	(617.54)
Demographic factor	2.19	2.23	1.90
	(0.43)	(0.41)	(0.44)
Plan design	1.09	1.12	1.03
	(0.06)	(0.04)	(0.06)
Plan type			
HMO	88.9%	91.8%	77.0%
Indemnity	0.9%	0.1%	2.6%
POS	4.8%	6.5%	2.7%
PPO	5.4%	1.6%	17.7%
Consumer-directed plan	0.5%	N/A	3.6%
For-profit insurer	56.4%	57.1%	49.4%
Number of employers	793	189	168
Observations	99,040	8,241	4,299
Panel B. Self-insured plans			
Premium ($)	6,591.01	4,164.31	8,897.68
	(2,371.42)	(1,369.04)	(2,284.09)
Number of enrollees	173.40	195.18	167.16
	(634.69)	(730.78)	(663.03)
Demographic factor	2.15	2.39	1.88
	(0.49)	(0.52)	(0.39)
Plan design	0.99	0.99	0.97
	(0.08)	(0.07)	(0.07)
Plan type			
HMO	14.5%	1.8%	18.6%
Indemnity	13.4%	40.3%	3.5%
POS	20.1%	25.9%	14.5%
PPO	51.2%	31.9%	63.4%
Consumer-directed plan	8.1%	N/A	22.6%
For-profit insurer	79.7%	81.1%	76.6%
Number of employers	922	199	218
Observations	241,810	12,574	21,434

Notes: All statistics are unweighted. The unit of observation is an employer-carrier-market plan type-year combination, unless noted otherwise. Demographic factor reflects age, gender, and family size of enrollees. Plan design measures the generosity of benefits. Both are constructed by the data source and exact formulae are not available. Premiums are in nominal dollars. Standard deviations are in parentheses.

Table A3—Descriptive Statistics, Plan-Year Data
(*sample limited to markets with conversion attempts*)

	Markets with successful attempts			Markets with unsuccessful attempts		
	1997–2009	1997	2009	1997–2009	1997	2009
Panel A. Fully insured plans						
Premium ($)	5,730.7	3,697.29	9,719.30	5,432.7	3,687.43	9,171.78
	(2,488.9)	(826.9)	(2,875.20)	(2,326.8)	(886.9)	(3,051.80)
Number of enrollees	165.65	169.19	116.84	136.4	123.12	92.34
	(418.8)	(367.6)	(255.80)	(333.8)	(254.9)	(177.90)
Demographic factor	2.21	2.24	1.93	2.14	2.20	1.83
	(0.43)	(0.41)	(0.43)	(0.43)	(0.41)	(0.46)
Plan design	1.1	1.12	1.03	1.11	1.12	1.03
	(0.06)	(0.04)	(0.06)	(0.06)	(0.04)	(0.06)
Plan type						
HMO	89.6%	92.7%	76.8%	87.3%	89.1%	70.7%
Indemnity	0.8%	0.1%	2.2%	1.1%	0.2%	4.0%
POS	4.9%	5.9%	2.6%	6.4%	8.9%	3.2%
PPO	4.7%	1.3%	18.4%	5.3%	1.8%	22.1%
Consumer-directed plan	0.4%	N/A	3.4%	0.60%	N/A	5.1%
For-profit insurer	59.8%	53.3%	62%	61.2%	56.3%	47.6%
Number of employers	628	159	119	514	138	83
Observations	22,529	2,033	832	13,227	1,255	498
Panel B. Self-Insured plans						
Premium ($)	6,618.8	4,135.5	8,958.20	6,493.4	4,129.9	8,795.50
	(2,402.5)	(1,432.4)	(2,352.90)	(2,334.2)	(1,317.6)	(2,247.40)
Number of enrollees	173.9	216.6	161.80	162.7	162.2	142.20
	(600.6)	(783.4)	(610.70)	(561.6)	(471.3)	(442.80)
Demographic factor	2.15	2.37	1.89	2.11	2.32	1.86
	(0.49)	(0.52)	(0.40)	(0.47)	(0.49)	(0.39)
Plan design	0.99	0.99	0.97	0.99	0.99	0.97
	(0.08)	(0.07)	(0.07)	(0.08)	(0.07)	(0.07)
Plan type						
HMO	14.9%	1.6%	19.2%	16.7%	1.8%	20.4%
Indemnity	13.2%	40.6%	3.3%	11.9%	37.6%	3.3%
POS	21.3%	26.7%	15.2%	22.0%	30.9%	15.8%
PPO	50.6%	31.2%	62.2%	49.4%	29.7%	60.5%
Consumer-directed plan	7.9%	N/A	22%	7.8%	N/A	22%
For-profit insurer	91.6%	77%	96%	75.6%	84%	68%
Number of employers	841	179	225	792	175	196
Observations	54,325	2,922	4,861	34,895	1,796	3,106

Notes: All statistics are unweighted. The unit of observation is an employer-carrier-market plan type-year combination, unless noted otherwise. Demographic factor reflects age, gender, and family size of enrollees. Plan design measures the generosity of benefits. Both are constructed by the data source, and exact formulae are not available. Premiums are in nominal dollars. Standard deviations are in parentheses.

TABLE A4—DESCRIPTIVE STATISTICS: MEDICAL-LOSS RATIOS AND INSURANCE COVERAGE

	States with successful attempts			States with unsuccessful attempts		
	2001–2009	2001	2009	2001–2009	2001	2009
MLR	0.848	0.868	0.873	0.851	0.880	0.864
	(0.029)	(0.033)	(0.022)	(0.038)	(0.038)	(0.027)
MLR (BCBS plans)	0.845	0.867	0.868	0.843	0.842	0.854
	(0.039)	(0.045)	(0.028)	(0.051)	(0.043)	(0.073)
MLR (non-BCBS plans)	0.850	0.873	0.882	0.844	0.917	0.832
	(0.038)	(0.037)	(0.027)	(0.059)	(0.055)	(0.074)
Percent insured	0.861	0.874	0.842	0.851	0.858	0.836
	(0.033)	(0.040)	(0.035)	(0.026)	(0.027)	(0.023)
Percent enrolled in employer-sponsored insurance	0.691	0.727	0.633	0.668	0.697	0.625
	(0.055)	(0.050)	(0.052)	(0.059)	(0.060)	(0.047)
Percent with individual private insurance	0.092	0.100	0.089	0.092	0.105	0.088
	(0.015)	(0.017)	(0.013)	(0.020)	(0.017)	(0.021)
Percent enrolled in Medicaid	0.119	0.091	0.158	0.123	0.105	0.149
	(0.044)	(0.027)	(0.051)	(0.042)	(0.047)	(0.045)

Notes: The unit of observation is the state year. The number of observations for the MLRs varies between 15 and 19 per year, while the insurance rates have 19 observations in all years. Standard deviations are in parentheses.

TABLE A5—EFFECT OF BCBS CONVERSIONS ON FP SHARE, LEADS, AND LAGS

	Fully insured plans (1)	Self-insured plans (2)
Panel A. Model 1 (dependent variable = FP share)		
$(\text{BCBS FP})_{t-3}$	0.048	0.047
	(0.145)	(0.107)
	[0.194]	[0.130]
$(\text{BCBS FP})_{t-2}$	0.029	0.066
	(0.485)	(0.034)
	[0.508]	[0.042]
$(\text{BCBS FP})_{t-1}$	0.017	0.069
	(0.717)	(0.031)
	[0.672]	[0.034]
$(\text{BCBS FP})_{t=0}$	0.153	0.346
	(0.006)	(0.000)
	[0.018]	[0.000]
$(\text{BCBS FP})_{t+1}$	0.174	0.339
	(0.005)	(0.000)
	[0.014]	[0.000]
$(\text{BCBS FP})_{t+2}$	0.159	0.391
	(0.015)	(0.000)
	[0.054]	[0.000]
$(\text{BCBS FP})_{\geq(t+3)}$	0.146	0.411
	(0.053)	(0.000)
	[0.104]	[0.000]
Observations	552	564

(*continued*)

Table A5—Effect of BCBS Conversions on FP Share, Leads, and Lags (*continued*)

	Fully insured plans (1)	Self-insured plans (2)
Panel B. Model 2 (dependent variable = FP share)		
$(\text{BCBS FP})_{t-3} \times$ low	0.074	0.041
	(0.076)	(0.322)
	[0.112]	[0.372]
$(\text{BCBS FP})_{t-2} \times$ low	0.048	0.066
	(0.319)	(0.130)
	[0.326]	[0.160]
$(\text{BCBS FP})_{t-1} \times$ low	0.044	0.039
	(0.423)	(0.353)
	[0.352]	[0.442]
$(\text{BCBS FP})_{t=0} \times$ low	0.143	0.277
	(0.021)	(0.000)
	[0.036]	[0.000]
$(\text{BCBS FP})_{t+1} \times$ low	0.164	0.249
	(0.012)	(0.000)
	[0.022]	[0.000]
$(\text{BCBS FP})_{t+2} \times$ low	0.155	0.314
	(0.038)	(0.000)
	[0.090]	[0.000]
$(\text{BCBS FP})_{\geq(t+3)} \times$ low	0.119	0.319
	(0.163)	(0.000)
	[0.228]	[0.000]
$(\text{BCBS FP})_{t-3} \times$ high	−0.029	0.081
	(0.442)	(0.016)
	[0.482]	[0.018]
$(\text{BCBS FP})_{t-2} \times$ high	−0.028	0.087
	(0.719)	(0.007)
	[0.750]	[0.004]
$(\text{BCBS FP})_{t-1} \times$ high	−0.079	0.141
	(0.299)	(0.009)
	[0.278]	[0.010]
$(\text{BCBS FP})_{t=0} \times$ high	0.176	0.505
	(0.025)	(0.000)
	[0.020]	[0.000]
$(\text{BCBS FP})_{t+1} \times$ high	0.194	0.537
	(0.037)	(0.000)
	[0.050]	[0.000]
$(\text{BCBS FP})_{t+2} \times$ high	0.165	0.560
	(0.048)	(0.000)
	[0.090]	[0.000]
$(\text{BCBS FP})_{\geq(t+3)} \times$ high	0.231	0.629
	(0.006)	(0.000)
	[0.002]	[0.000]
Observations	552	564

Notes: The unit of observation is the market year. All models include fixed effects for each market and year as well as lagged market-year controls—ln(Medicare costs per capita) and the unemployment rate—and are estimated by weighted least squares using the average number of enrollees in each market as weights.The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the market.

Table A6—Effect of BCBS Conversions on Premiums, Leads, and Lags

	Fully insured plans (1)	Self-insured plans (2)
Panel A. Model 1 (dependent variables = premium index)		
$(\text{BCBS FP})_{t-3}$	−0.977	2.837
	(0.440)	(0.126)
	[0.404]	[0.154]
$(\text{BCBS FP})_{t-2}$	−2.203	1.072
	(0.248)	(0.599)
	[0.228]	[0.652]
$(\text{BCBS FP})_{t-1}$	−1.316	2.342
	(0.668)	(0.398)
	[0.680]	[0.438]
$(\text{BCBS FP})_{t=0}$	−2.926	0.326
	(0.605)	(0.933)
	[0.756]	[0.956]
$(\text{BCBS FP})_{t+1}$	−2.222	1.858
	(0.722)	(0.626)
	[0.854]	[0.674]
$(\text{BCBS FP})_{t+2}$	0.523	1.882
	(0.933)	(0.618)
	[0.960]	[0.684]
$(\text{BCBS FP})_{\geq(t+3)}$	5.671	6.545
	(0.419)	(0.188)
	[0.508]	[0.312]
Observations	599	611

(*continued*)

Table A6—Effect of BCBS Conversions on Premiums, Leads, and Lags (*continued*)

	Fully insured plans (1)	Self-insured plans (2)
Panel B. Model 2 (dependent variable = premium index)		
$(\text{BCBS FP})_{t-3} \times$ low	−0.174	2.236
	(0.888)	(0.333)
	[0.898]	[0.416]
$(\text{BCBS FP})_{t-2} \times$ low	−1.953	0.261
	(0.316)	(0.914)
	[0.274]	[0.950]
$(\text{BCBS FP})_{t-1} \times$ low	−2.686	1.950
	(0.344)	(0.535)
	[0.348]	[0.612]
$(\text{BCBS FP})_{t=0} \times$ low	−4.225	−0.703
	(0.450)	(0.872)
	[0.612]	[0.868]
$(\text{BCBS FP})_{t+1} \times$ low	−6.130	0.519
	(0.327)	(0.906)
	[0.508]	[0.896]
$(\text{BCBS FP})_{t+2} \times$ low	−4.306	0.326
	(0.514)	(0.939)
	[0.584]	[0.972]
$(\text{BCBS FP})_{\geq(t+3)} \times$ low	0.895	4.981
	(0.909)	(0.367)
	[0.970]	[0.508]
$(\text{BCBS FP})_{t-3} \times$ high	−3.612	4.302
	(0.056)	(0.027)
	[0.092]	[0.038]
$(\text{BCBS FP})_{t-2} \times$ high	−3.229	2.778
	(0.237)	(0.318)
	[0.254]	[0.384]
$(\text{BCBS FP})_{t-1} \times$ high	0.757	2.837
	(0.851)	(0.476)
	[0.926]	[0.520]
$(\text{BCBS FP})_{t=0} \times$ high	−0.128	2.414
	(0.979)	(0.618)
	[0.944]	[0.676]
$(\text{BCBS FP})_{t+1} \times$ high	9.199	4.601
	(0.046)	(0.353)
	[0.072]	[0.412]
$(\text{BCBS FP})_{t+2} \times$ high	14.377	5.008
	(0.005)	(0.233)
	[0.012]	[0.292]
$(\text{BCBS FP})_{\geq(t+3)} \times$ high	19.939	9.762
	(0.014)	(0.089)
	[0.030]	[0.152]
Observations	599	611

Notes: The unit of observation is the market year. All models include fixed effects for each market and year as well as lagged market-year controls—ln(Medicare costs per capita) and the unemployment rate—and are estimated by weighted least squares using the average number of enrollees in each market as weights. The p-values generated using cluster-robust standard errors [wild-cluster bootstrap-t procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the market.

Table A7—Effect of Different Types of BCBS Ownership Conversions on Premiums

	Dependent variable = premium index Mean = 179.9	
	(1)	(2)
Panel A. Fully insured plans		
Lagged BCBS NFP to Mutual	2.210	4.677
	(0.757)	(0.471)
	[0.784]	[0.480]
Lagged BCBS NFP to FP	−6.333	−6.340
	(0.268)	(0.271)
	[0.334]	[0.334]
Lagged BCBS Mutual to FP	12.026	0.238
	(0.018)	(0.981)
	[0.034]	[0.988]
Lagged BCBS Mutual to FP × pre-conversion share		59.223
		(0.092)
		[0.122]
Observations	599	599
	Dependent variable = premium index Mean = 181.0	
	(1)	(2)
Panel B. Self-insured plans		
Lagged BCBS NFP to Mutual	0.758	1.611
	(0.826)	(0.637)
	[0.838]	[0.650]
Lagged BCBS NFP to FP	2.153	2.154
	(0.666)	(0.666)
	[0.818]	[0.820]
Lagged BCBS Mutual to FP	4.663	0.261
	(0.119)	(0.963)
	[0.142]	[0.914]
Lagged BCBS Mutual to FP × pre-conversion share		21.393
		(0.334)
		[0.410]
Observations	611	611

Notes: The unit of observation is the market year. All models include fixed effects for each market and year as well as lagged market-year controls—ln(Medicare costs per capita) and the unemployment rate—and are estimated by weighted least squares using the average number of enrollees in each market as weights. The *p*-values generated using cluster-robust standard errors [wild-cluster bootstrap-*t* procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the market.

TABLE A8—IMPACT OF FOR-PROFIT PENETRATION ON INSURANCE COVERAGE (*by age group*)

	Panel A. Dep. var. = share on Medicaid (under 18) Mean = 0.25			*Panel B. Dep. var. = share on Medicaid (18–44)* Mean = 0.09		
	(1)	(2)	(3)	(4)	(5)	(6)
Lagged BCBS FP	0.017	−0.034		0.016	0.013	
	(0.140)	(0.058)		(0.006)	(0.478)	
	[0.240]	[0.160]		[0.000]	[0.400]	
Lagged BCBS FP × pre-conversion share		0.264			0.014	
		(0.002)			(0.873)	
		[0.000]			[1.000]	
Lagged BCBS FP × low pre-conversion share			0.006			0.014
			(0.575)			(0.045)
			[0.480]			[0.160]
Lagged BCBS FP × high pre-conversion share			0.032			0.017
			(0.069)			(0.039)
			[0.160]			[0.160]
Observations	209	209	209	209	209	209
	Panel C. Dep. var. = share with any private insurance (under 18) Mean = 0.70			*Panel D. Dep. var. = share with any private insurance (18–44)* Mean = 0.71		
	(1)	(2)	(3)	(4)	(5)	(6)
Lagged BCBS FP	0.003	0.050		−0.001	0.036	
	(0.781)	(0.073)		(0.926)	(0.059)	
	[0.880]	[0.080]		[0.880]	[0.080]	
Lagged BCBS FP × pre-conversion share		−0.246			−0.193	
		(0.089)			(0.049)	
		[0.240]			[0.080]	
Lagged BCBS FP × low pre-conversion share			0.012			0.006
			(0.276)			(0.563)
			[0.240]			[0.560]
Lagged BCBS FP × high pre-conversion share			−0.011			−0.010
			(0.460)			(0.482)
			[0.480]			[0.400]
Observations	209	209	209	209	209	209

Notes: The unit of observation is the state year. The study period is 1999–2009. Insurance rates and pre-conversion share are scaled from zero to one. All specifications include state and year fixed effects, simulated Medicaid eligibility rate for children under 18, lagged ln(Medicare costs per capita), and the lagged unemployment rate. Each observation is weighted by the average under-65 population in the state. The p-values generated using cluster-robust standard errors [wild-cluster bootstrap-t procedure] are reported in $(\cdot)$ and $[\cdot]$, respectively. In both cases, the clustering unit is the state.

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