Do For-Profit Insurers Charge Higher Premiums?

Leemore Dafny
Northwestern University and NBER

Subramaniam Ramanarayanan
University of California at Los Angeles

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Abstract

For-profit insurers account for half of private health insurance in the U.S., notwithstanding concerns that they deliver lower-quality, higher-priced policies. We explore the relationship between ownership and employer-sponsored insurance premiums, medical loss ratios, and uninsurance rates. Using a national panel dataset encompassing ~10 million lives per year, we study the effects of conversions of Blue Cross and Blue Shield affiliates in 11 states. Conversions have no impact on premiums on average, but in markets where the converting affiliate commanded substantial market share, both the BCBS affiliate and its rivals increased premiums after conversion. Our results suggest that subsidies for not-for-profit insurers, such as those in the Affordable Care Act, will only create value under certain market conditions.

I. Introduction

The healthcare industry in the United States is an amalgam of not-for-profit, for-profit, and government entities. The health insurance sector is no exception: the non-elderly insured are split roughly equally across insurers of these ownership types.[1] While there is an extensive theoretical and empirical literature examining the impact of ownership form on performance in the hospital sector, there is comparatively little research of this kind in the health insurance industry. Prior studies focus on differences between government-provided insurance and private insurance, rarely distinguishing between private insurance offered by for-profit (FP) companies (such as Aetna and UnitedHealthcare) and private insurance offered by not-for-profit (NP) companies (such as Blue Cross Blue Shield of Arizona and Harvard Pilgrim Health Care).[2] This research gap came to the fore during the national healthcare debate of 2009-2010, which culminated in the passage of The Patient Protection and Affordable Care Act (PPACA).

During the debate surrounding PPACA, policymakers considered adding a "public option" to the set of choices available for purchase by individuals who did not qualify for public insurance by virtue of income, age, or health status. After the public option failed to muster sufficient political support, the concept of providing federal seed money to form "not-for-profit cooperatives" was floated as an alternative. Formally introduced as the Consumer-Owned and –Oriented Plan (CO-OP), Senator Kent Conrad (R-ND) asserted that co-ops "will focus on getting the best value for customers, rather than maximizing plan revenues or profits."[3] According to Conrad, "[m]any experts believe co-ops, as non-profits, could offer significant discounts when compared to traditional, for-profit insurance companies."[4] A total of $6 billion was included in

[1] "Basic Facts & Figures: Nonprofit Health Plans," Alliance for Advancing Nonprofit Healthcare, accessed 8/15/2010 at http://www.nonprofithealthcare.org/resources/BasicFactsAndFigures-NonprofitHealthPlans9.9.08.pdf
[2] A notable exception is Town, Feldman and Wholey (2004), which examines conversions of non-profit HMOs to for-profit status between 1987 and 2001. The authors find no short-term effect of conversions on premiums or profits of converting firms. We discuss this paper at greater length in the following section.
[3] "FAQ about the Consumer-Owned and –Oriented Plan (CO-OP)," accessed 7/15/2010 at http://conrad.senate.gov/issues/statements/healthcare/090813_coop_QA.cfm.
[4] *Ibid*. We have contacted Senator Conrad's office to request the names of the experts.

the final bill to help establish CO-OPs by July 1, 2013, $2.2 billion of which was subsequently eliminated as part of a budget deal in April 2011.[5]

There is an extensive theoretical literature on the objectives of NP hospitals and the implications for outcomes such as quantity, quality, and price of care. The models (which we summarize later) yield different predictions for various outcomes; our focus is on price (or premiums). which our dataset is uniquely suited to studying. Given the extraordinary growth in health insurance premiums over the past 25 years, as well as the new federal mandate and associated subsidies for the purchase of private insurance, price is a particularly pertinent outcome. As we discuss below, most models of NP hospitals predict underpricing relative to FPs. This arises from the positive value NPs are presumed to place on quantity of care provided ("access" in the insurance setting), and/or more general models of altruistic objectives.

Our primary data source is the Large Employer Health Insurance Data (LEHID), a national panel of employer-sponsored healthplans spanning the years 1998-2009. This proprietary data, gathered by a leading benefits consultancy, includes information on the healthplans of over 10 million enrollees annually; over 900 employers – primarily multisite, publicly-traded firms – are represented in the sample. We utilize data from the Current Population Survey and the National Association of Insurance Commissioners to evaluate the impact of local FP market share on insurance coverage rates and insurer medical loss ratios, respectively.

Given the dearth of information on the distribution of ownership status in health insurance, we begin by documenting important facts about for-profit insurance in the LEHID, including market penetration by region, by product type, by insurance type, and over time. We then examine the relationship between for-profit status and premium levels, exploiting the rich detail in the dataset to control for differences in premiums associated with employee risk profiles, geographic location, generosity of benefits, etc. We find fully-insured premiums are 4-

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[5] Sources: http://www.kff.org/healthreform/8061.cfm; "Budget Deal Targets Pieces of Health-Care Law," *Wall Street Journal*, April 10, 2011.

5 percent higher when offered by FP insurers, and self-insured premiums are 2-2.5 percent higher.

Next, we pursue an instrumental variables strategy to purge our estimates of potential biases due to selection and omitted variables. While the OLS specifications include fixed effects to control for many sources of time-invariant omitted factors (such as systematic differences across markets and employers with high FP shares), time-varying omitted characteristics may bias our estimates if they are correlated with FP status. For example, employers seeking to reduce spending may switch to NP plans, while simultaneously engaging in unobservable actions that reduce average premiums (e.g. offering "early retirement packages" for older employees, thereby reducing the average age of the insured population).[6]

We look to the conversions of 11 state-specific Blue Cross and Blue Shield (BCBS) plans from not-for-profit to for-profit status as a source of exogenous variation in the presence (and market penetration) of FP plans within local geographic markets (139 in all, identified by the data source and consisting primarily of metropolitan areas and non-metropolitan areas within state borders).[7] BCBS affiliates offer insurance throughout the United States, and typically rank first or second in terms of local geographic market shares (particularly in the large group market, which is the focus of our sample).[8] As we discuss in detail below, a wave of conversions followed the 1994 decision by the national umbrella organization to permit conversions of local BCBS plans to FP status. If these conversions are orthogonal to other characteristics determining premiums, then local BCBS FP status can serve as an instrument for market-level FP penetration.

Using a regression-adjusted premium index for each geographic market and year, we find that premium growth is no different, on average, in post-conversion markets versus markets not

[6] LEHID includes only active workers.
[7] These markets are described in greater detail in Section III. For example, one of the markets in LEHID (designated "Maryland-Baltimore") covers the metropolitan area of Baltimore, while another ("Maryland – except Baltimore") includes the rest of the state, apart from those areas belonging to the Washington, DC metro area market.
[8] According to a 2010 report by the American Medical Association ("Competition in Health Insurance: A Comprehensive Study of U.S. Markets"), BCBS plans were among the top two plans in terms of market share in over 95 percent of markets surveyed (the report contains market share data on 359 MSAs overall).

experiencing conversions. There is no difference in premium trends prior to conversions. The resulting instrumental variables estimate of the impact of FP insurer market share on premiums indicates no effect of for-profit status on premiums, *on average.*

Further analysis reveals the impact of BCBS conversions varied depending on the market share of the converting plan. Specifically, marketwide premiums increased when converting BCBS plans had shares in excess of ~20 percent, and decreased otherwise. (The results are only sizeable and statistically-significant for fully-insured premiums, over which insurers have full control and pricing is less transparent than for self-insured contracts.) Separating the sample into BCBS and non-BCBS plans reveals that post-conversion price increases in these markets were common to both. Thus a simple comparison of price changes for converting and non-converting plans – a common tactic for case-studies of conversions - would understate the effect of conversion. Put together, the findings are consistent with greater exercise of market power by FP plans. The spillover effect on rivals magnifies these effects.

Notwithstanding the post-conversion premium increases in some markets, we do not find evidence of reductions in the proportion of the under-65 population that is privately-insured. This may be due to measurement issues associated with the use of state-level private coverage rates. There is evidence that BCBS' rivals - but not BCBS plans - boosted their medical spending relative to premiums in the wake of conversions. Across all markets, we estimate that total changes in medical spending more than offset changes in premiums. The welfare effects, however, are uncertain. While a dollar of spending may be valued more highly than a dollar of income (particularly if the dollar of income devoted to premiums is pre-tax), this need not be the case. For example, if the additional spending results from higher prices paid to providers for the same services rendered – rather than any additional services – the net welfare effect depends on the marginal utility of income for providers vis-a-vis patients.[9]

The paper proceeds in 6 sections. Section II discusses the historical origins of for-profit insurers, summarizes prior research on ownership status in health care settings, and provides some background on the BCBS conversions that underlie our identification strategy. We

[9] We thank David Cutler for this observation.

describe our data sources in Section III. The OLS and IV analyses of the "FP effect" are presented in Sections IV and V, respectively. We discuss results on non-price outcomes in Section VI. Section VII concludes.

II. Background

A. Origin and Evolution of For-Profit Insurance Plans in the U.S.

The U.S. 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. To this end, Blue Cross plans engaged in community rating of premiums, charging the same rates to individuals regardless of health status. Blue Shield plans subsequently arose to cover physician charges; the two merged to form the Blue Cross Blue Shield Association in 1982. For-profit insurers entered the marketplace towards the middle of the 20th century, when health insurance enrollment soared as employers sought alternative forms of employee compensation in the wake of World War Two-era wage controls.

For-profits enjoyed greater financial success by experience-rating their premiums and targeting large employer groups, thereby reducing adverse selection as well as administrative costs (Thomasson (2002, 2003)). Eventually, the BC plans also engaged in experience-rating, charging lower rates to lower-risk groups and individuals, although the precise timing of this practice, which could be viewed as contrary to the explicit or implicit mission of BC plans to maximize access to insurance, varied across plans.

Precise figures on current market shares of for-profit insurers are difficult to obtain. America's Alliance for Advancing Nonprofit Health Care reports that 61 percent of the 138 healthplans in the US with at least 100K enrollees are nonprofit, accounting for 48 percent of private-sector enrollment (98 million out of 203 million estimated enrollees) in 2008. This estimate includes enrollees in government-financed plans (e.g. Medicaid managed care), and appears to include enrollees in self-insured plans (judging from the volume of enrollees).[10] In 2008, 55 percent of workers covered by employer-sponsored insurance were in self-insured plans.[11] By comparison, in LEHID (which excludes Medicaid and Medicare) we find an FP share of 47 percent in the fully-insured segment, and 72 percent in the self-insured segment, also as of 2008. To obtain one more estimate of FP penetration, we tabulated data from the National Association of Insurance Commissioners (NAIC), an organization of state regulators.[12] These data exclude self-insured enrollees, as only fully-insured plans are regulated by the states.[13] The NAIC estimate of fully-insured FP penetration for 2008 is 54 percent. In section III we discuss trends in FP penetration, as well as differences in penetration by product type (e.g. HMO vs. PPO).

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 literature on ownership status in the U.S. hospital industry. Up-to-date surveys of this literature can be found in Chang and Jacobson (2011) and Capps et al (2011). Chang and Jacobson (2011) 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 not-for-profits behave no differently than for-profits.[14] 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). FPs/FPIDs and NPs can therefore co-exist (i.e. both serve customers) only if they are

[10] The primary data source is listed as the 2008 Directory of Health Plans published by Atlantic Information Services (AIS). These data reportedly include enrollment across all product types (e.g. HMO, PPO) and contracts (e.g. Medicare and Medicaid).

[11] Source: Employer Health Benefits Annual Survey, Kaiser Family Foundation

[12] For additional details on the NAIC data, as well as other sources of insurance data, see Dafny, Dranove, Limbrock, and Scott Morton (2011). Our tabulations reflect only enrollment in comprehensive medical insurance. Total enrollment using this definition is 86 million in 2008. Both NAIC and LEHID figures pertain to enrollment in plans offered by stock corporations.

[13] The NAIC data also exclude plans from the state of California

[14] This conjecture has empirical support from a number of studies including Duggan (2002), Cutler and Horwitz (2000), Silverman and Skinner (2004), Dafny (2005), Horwitz and Nichols (2009), and Capps et al (2010), which find that not-for-profit hospitals behave similarly to for-profits, particularly in markets where they face greater competition from for-profit hospitals, on dimensions like pricing, profitability, "gaming" of reimbursement codes, quality of care, and service offerings.

differentiated. In the hospital industry this is accomplished by location, services, and reputation/marketing. In the insurer industry, differentiators include reputation/marketing, provider networks, benefit design, and customer service.

The small set of papers that consider impact of ownership status of private health insurance plans fall 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 in NP plans. Using data on Medicare HMOs from 1998, Schneider et al. (2003) report that for-profit HMOs score lower on four audited HEDIS measures (breast cancer screening, diabetic eye examinations, administering beta blockers after heart attack, and follow-up after mental illness hospitalization).[15] Controlling for socioeconomic factors (including age, gender, area income and rural residence) among plan enrollees has little impact on the estimates. The result is also robust to "within-county" comparisons, i.e. controlling for location. Studies comparing FP and NP healthplans also find that consumer satisfaction is higher among enrollees of NP plans (Gillies et al 2006), especially for patients in poor health (Tu and Reschovsky 2002). NP plans also 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 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 for-profit HMO penetration. Town, Feldman and Wholey (2004) study the effects of HMO conversions to for-profit 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.

[15] HEDIS stands for "Healthcare Effectiveness Data and Information Set" (formerly Health Plan Employer Data and Information Set"). It is a broad set of data measures developed to capture information about healthplans, including measures of access to care, service utilization, quality of care, and financial solvency. As of 1998, healthplans participating in Medicare Part C (then primarily HMOs) are required to report HEDIS measures to HCFA (now CMS).

Our study also relies on conversions to identify the effect of ownership status, however there are important differences in our sample, unit of observation, and outcomes of interest. Our sample includes all plan types (HMO, POS, PPO, and indemnity), as well as funding arrangements (fully insured and self-insured). We also benefit from panel data at the employer-plan level, which permits a rich set of controls for changes in the underlying insured population. Last, we study the effects of conversions on premiums offered by both converting and nonconverting firms. Given the oligopolistic structure of local insurance markets (Dafny (2010), Dafny, Duggan and Ramanarayanan (2011)), price reductions (increases) by converting firms should tighten (relax) pricing constraints on competitors. Thus, using nonconverting plans as a control group for converting plans may lead to downward-biased estimates of price effects. Finally, we also consider the impact of insurance ownership on medical loss ratios, the share of premiums paid out in medical claims.

C. *Blue Cross Blue Shield Plans*

Our analysis utilizes the conversion of eleven BCBS plans to FP stock corporations as a source of 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. In the section that follows, we present our own estimates of BCBS penetration, using the geographic market definitions supplied in the LEHID.

As previously mentioned, BCBS plans were chartered as social welfare organizations. They were not subject to federal taxes until 1986.[16] Historically, the rules of the national BCBS Association explicitly prohibited member plans from converting to for-profit status. In June 1994, partly prompted by the decision of Blue Cross of California to form a for-profit subsidiary

[16] As of January 1, 1987, BCBS plans are subject to federal taxation but are entitled to other tax benefits such as special deductions and state tax exemptions in recognition of their public service commitment. (Source: Coordinated Issue Paper – Blue Cross Blue Shield Health Insurance, available at irs.gov)

(WellPoint),[17] the national BCBS association modified its bylaws to allow affiliates to convert to for-profit ownership. This sparked a series of conversions of BCBS plans, with plans in fourteen states converting to for-profit status by 2003.

Many BCBS plans proposing or undergoing conversion cited access to equity capital as the key driver for conversion. Uses for additional capital include infrastructure investments (for example, in IT or disease management) and acquisitions of other plans. Larger insurers can spread fixed costs over more enrollees, thereby improving operating margins.[18] Representatives of converting plans 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 converted to for-profit stock corporations during our study period of 1998 to 2009.[19] The top panel lists successful conversions while the bottom panel lists unsuccessful conversion attempts, a majority of which ultimately failed owing to actual or anticipated regulatory rejection.[20] The real threat of regulatory opposition implies that BCBS plans converting to for-profit status may have operated in a regulatory environment that was different from the environment of non-converting plans. If these differences are systematically related to the determinants of premium growth (conditional on included covariates), our estimates will be biased. Investigating the possible endogeneity of conversions is therefore a priority in our empirical analysis. The numerous unsuccessful conversion attempts also enable us to perform a falsification test, described in section V.

[17] Wellpoint was created as a network of for-profit HMO and PPOs focusing on the individual and small-group markets.
[18] "For-Profit Conversion and Merger Trends Among Blue Cross Blue Shield Health Plans," *Center for Studying Health System Change* Issue Brief 76 (January 2004).
[19] We thank Chris Conover for sharing his detailed notes on plan conversions, which we used to construct Table 1. In addition to the 11 plans listed in Table 1, BCBS plans in Virginia, California and Georgia also switched to for-profit status with the conversion in Virginia taking place in 1997 and in the latter two states occurring in 1996.
[20] Because BCBS plans are chartered as non-profit organizations formed in the public interest, conversions require approval from state insurance commissioners, who must ascertain whether conversion is consistent with the original charter. In making their evaluations, the regulator may investigate the impact of the conversion on various outcomes including pricing, access and provider reimbursement (Beaulieu (2004)). Conversions have also been rejected due to concerns about improper executive payouts, as in the case of Maryland-based CareFirst.

We define a conversion as having taken place if the BCBS plan becomes a stock company either on its own or through acquisition.[21] Ten of the eleven conversions take place in the same year (2001), when Anthem (the parent organization of these plans) demutualized.[22] While it would be ideal to have more variation in the timing of 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.

The 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 provoke 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 (Robinson 2004, McPherson 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. Based on their results, they conclude that BCBS conversions have only a modest impact on health spending and insurance access in affected states.

[21] An alternative definition used in the literature is conversion to a for-profit mutual company. We utilize conversion to a stock corporation for two reasons. First, it seems likelier to produce an upper-bound estimate of the FP effect on premiums, as mutual insurers explicitly value policyholder wellbeing, which could cause them to price lower than the market can bear. An upper bound is valuable given that prior studies find no FP effect on premiums or profits. Second, four of the eleven plans converting to for-profit stock corporations converted to FP mutuals prior to the start of our study period, and three did so during the second year of our sample (1999), thus our data are not as well-suited to this alternative definition.

[22] The sole non-Anthem plan was New York's Empire Blue Cross Blue Shield, which was acquired by Wellpoint, Inc., the other for-profit aggregator of BCBS plans. Anthem and Wellpoint merged in 2005.

III. 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 healthplans offered by a sample of large employers between 1998 and 2009. This proprietary dataset is also used in Dafny (2010) and Dafny, Duggan and Ramanarayanan (2011) but is supplemented here with three recent years of data (2006-2009).

The unit of observation in LEHID is a healthplan-year, where a healthplan is defined as a unique combination of an employer, market, insurance carrier, plan type, and insurance type. Most *employers* are large, multi-site, publicly-traded firms, such as those included on the *Fortune 1000* list. Geographic *markets* are defined by the data source using 3-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 ex-metropolitan areas within the same state (e.g. Chicago, Northern Illinois except Chicago, 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. 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 a self-insured plan. 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 given employer-market-insurance carrier-plan type-insurance type combination ("plan") is only two years. For this reason, rather than relying on the healthplan-year data to explore the effects of BCBS conversions, we develop a market-year premium index. In Section V, we discuss the creation of this index in detail. Here we note that the index is constructed using *within-healthplan* 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 (see Dafny 2010 for details).[23] Additional information on the representativeness of LEHID, as well as comparisons with other sources of insurance market data, is included in the appendix.

In addition to the identifying information described this far, we have four key variables that are included in all analyses. *Premium* represents the combined annual employer and employee charge, and is expressed as an average amount per enrollee (i.e. a covered employee); it therefore increases with the average family size of 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 expected costs per enrollee in a plan. *Plan design factor* captures the generosity of benefits, with an emphasis on the degree of coinsurance and copays. Both factors are calculated by the data source, and the formulae were not disclosed. Higher values for either will result in higher premiums. For 2005 onward, we also have 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 2007).

[23] The KFF/HRET survey randomly selects public and private employers to obtain nationally-representative statistics for employer-sponsored health insurance; approximately 2000 employers respond each year. The micro data are not publicly available, nor is the sample designed to provide representative estimates for distinct geographic areas.

The LEHID also includes the *number of enrollees* in each plan, excluding dependents, who are accounted for by the demographic factor variable described above. The total number of enrollees in all LEHID plans averages 4.7 million per year. Given an average family size of more than 2, this implies over 10 million Americans are part of the sample in a typical year.

Table 2 presents descriptive statistics for each year of the LEHID data, which spans the period 1998 to 2009, inclusive. The top panel pertains to the FI sample while the bottom panel pertains to the 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 1998, SI plans are only a slight majority (55 percent) of observations, but by 2009 they account for 83 percent of our sample. This surge has also been observed in tabulations of the largest firms included in the MEPS-IC (Cooper, Jim and Simon 2009). Second, while FI plans are predominantly HMOs throughout the study period (with some movement toward PPOs in the last few years), SI plans exhibit a dramatic shift away from indemnity and POS plans and toward PPOs (and to a lesser extent, HMOs) over time.[24] 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, 22.6 percent of SI plans are designated as CDPs. Very few FI plans are CDPs.

Across 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. We also prepare estimates using only data through 2005 as a robustness check. We provide more detail in Section V.

[24] The rise in self-insured HMOs is likely due to the increasing number of insurers who offer this option; historically HMOs were a fully-insured product, as it was difficult to attribute costs to particular sets of employees. As HMOs have become more similar to PPOs, self-insured arrangements have become more common. Indeed, in 2009 the number of self-insured HMOs in LEHID exceeded the number of fully-insured HMOs.

We compiled information on the ownership status of healthplans 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[25] through independent research (e.g., web searches, analyst reports). We use Table 1 to code BCBS ownership status by market, as ownership status for these plans is not consistently reported by our data source.[26]

Figure 1 presents estimates of for-profit penetration obtained from the LEHID sample. Data are presented separately by year, BCBS affiliation of the insurer, and insurance type (FI and SI). The top panel shows that for-profit penetration in the FI market is sizeable (52 percent on average) but exhibits a downward trend over time. For-profit penetration in the SI sector is markedly higher (by nearly 20 percent on average), and has remained consistently high during the past decade. The share of enrollees insured by BCBS plans exhibits a steady uptick across both FI and SI plans, with the majority of the growth occurring in the for-profit BCBS segment. This is consistent with the large number of BCBS for-profit conversions that take place during this time.

Figure 2 illustrates the variation in market penetration of FP insurers across geographic markets (Panel A), and across product types (Panel B). For-profit plans have a dominant presence in most markets in our sample: nearly two-thirds of enrollees in the median market have FP plans. When we break down FP penetration by product type, we see that FPs are particularly dominant in the POS product line. The enrollment share of FP HMOs has eroded in recent years, to 55.6 percent in 2009 from 59.8 percent in 2002.

We supplement the LEHID data with time-varying measures of local economic conditions (the unemployment rate, as reported by the Bureau of Labor Statistics), and a measure of healthcare utilization (Medicare costs per capita, as reported by the Center for Medicare and

[25] Plans with missing information include those present only in the earlier years (1998, 1999) when surveys were not conducted by our data source, as well as plans not included in the surveys in later years.

[26] In particular, we see instances where there is a mix of for-profit and not-for-profit BCBS plans in a state even though the state has only one BCBS licensee. We believe this inconsistency arises because ownership status for BCBS healthplans associated with an employer in a given state might actually pertain to the primary affiliate used by the employer in the state where it is headquartered. For example, BCBS plans for United Airlines in Colorado may be linked to BCBS of Illinois, since United Airlines is headquartered in Chicago.

Medicaid services). As these measures are reported at the county-year level, and LEHID markets are defined by 3-digit zipcodes, we make use of a mapping between zipcodes and counties and where necessary, use population data to calculate weighted average values for each geographic market and year. Summary statistics for these measures are presented in **Table 3**.

B. Medical Loss Ratio Data

The medical loss ratio is the share of insurance premiums that is paid out for medical claims ("losses"). (Note this definition differs slightly from the MLR that will be used to enforce the new MLR minimum regulations in PPACA; the new definition cannot be calculated using available data sources.)[27] We construct state-year medical loss ratios using insurer-state-year data on total spending and premiums purchased from the National Association of Insurance Commissioners (NAIC) for the years 2001-2005. NAIC is an umbrella organization of state-level insurance regulators.[28] Because states regulate fully-insured healthplans, NAIC data should represent only the FI component of the health insurance market. Insurers report data by product line and state; Washington, DC is also included in the data but California is not. We construct a single MLR for each insurer-state-year, including only losses and premiums associated with comprehensive medical insurance (e.g. excluding Medicare supplemental plans, dental plans, and vision-only plans).[29] We drop observations in the 5 percent tails of the distribution of overall spending and premiums and aggregate the remaining data to construct state-year MLRs.

Descriptive statistics are presented in **Appendix Table 1**. The mean MLR declined from 0.88 to 0.84 between 2001 and 2005. We also calculate MLR separately for BCBS and non-

[27] PPACA requires insurers to maintain MLRs which exceed 80 percent in the individual/small group market, and 85 percent in the large group market, beginning in 2011. The Department of Health and Human Services is responsible for determining the definition of MLR to be used for this purpose. They have determined that spending for quality improvements will be counted in the numerator of the MLR, while taxes and fees paid will be deducted from the denominator. For most insurers, these adjustments will boost reported MLRs (see "Private Health Insurance: Early Experiences Implementing New Medical Loss Ratio Requirements, www.gao.gov/new.items/d11711.pdf, GAO 2011).

[28] 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.

[29] These categories are also excluded: Long-Term Care, Disability Income, Stop-Loss, and Other. Dafny et al (2011) describes this data source in greater detail.

BCBS plans in each state. Due to missing data, the sample size for the BCBS MLRs is strictly smaller, ranging from 39 to 48 across the 4 years in these data. The data reveal higher MLRs for non-BCBS plans in 2001, but by 2005 the average MLR was the same in both groups.[30] Although intriguing, these statistics (as well as the results from our analysis) should be interpreted with caution, as MLRs are sensitive to accounting practices, which differ across companies, states, and years (Robinson 1997).

IV. Are Premiums Different for For-Profit Plans?

We begin our analysis by estimating OLS equations of the following form:

$$
\begin{aligned}
(1)\quad \ln(\text{premium})_{emcjt} = {} & \beta_0 + \beta_1 FP_{mct-1} + \beta_2 demographics_{emcjt} + \beta_3 demographics_{emcjt} * (year >= 2006)_t \\
& + \beta_4\, plan\ design_{emcjt} + \beta_5 CDP_{emcjt} + \lambda_e + \psi_m + \upsilon_j + \delta_t + \theta X_{mt-1} \\
& + [\rho_{em}] + [\kappa_{jt}] + [\varphi_{mt}] + \varepsilon
\end{aligned}
$$

where *emcj* denotes "employer-market-carrier-plan type" (henceforth "plan") and *t* denotes year. (We omit the subscript for insurance type because we estimate equation (1) separately for SI and FI plans.) The independent variable of interest is an indicator for the FP status of the carrier associated with the plan; this status varies by carrier, market and year.[31] We lag this indicator as well as the market-year covariates (described below, and denoted by *X* in the equation) by one year because premiums are set prospectively, i.e. premiums for 2011 are negotiated in 2010. We include all the plan-year specific covariates we observe: *demographic factor, plan design*, and an indicator for whether a plan is consumer-directed (CDP). To ensure that the change in the construction of demographic factor between 2005 and 2006 (referenced earlier in Section III) does not impact the results, we add an interaction term between demographic factor and an indicator for 2006 and beyond.

[30] The MLR statistics are nearly identical when each state-year observation is weighted by enrollment.
[31] To be more precise, FP status varies by carrier, market and year for BCBS plans, and by carrier and year for non-BCBS plans, as non-BCBS plans are unified nationwide.

All specifications include the "main fixed effects" for employers, plan types, markets, and years.[32] The employer fixed effects capture time-invariant premium differences across employers, due for example to variation in employee risk profiles and insurance design. The plan type effects absorb premium differences associated with the four main plan types previously listed, and the market fixed effects capture time-invariant differences in premiums due to geographic variation in wages, medical practice, conduct of insurers and providers, and other omitted determinants of market-level insurance spending. The year dummies reflect national trends in premium growth, holding constant all of the other controls. Finally, we include controls for time-varying market characteristics that may affect premiums: the log of Medicare costs per capita (a measure of local utilization), and the local unemployment rate (a barometer of economic conditions).

We enrich the basic specification by progressively adding the three bracketed second-order interactions: employer-market fixed effects, plan type-year fixed effects, and market-year fixed effects. We describe the rationale for each of these terms below, noting first that it is computationally infeasible (as well as econometrically inadvisable) to simultaneously include all of the possible second (let alone third or fourth)-order interaction terms. Our selections are therefore grounded by conceptual arguments, and mirror the terms utilized in Dafny (2010).

The rationale for employer-market interactions is straightforward: employees in different locations may have different risk profiles and different benefit designs (not captured by our controls). Controlling for such differences will only have a material effect on our coefficient of interest if they are correlated with decisions to patronize FP carriers. For example, if employees in headquarters offices are less likely to choose FP plans and also more likely to choose plans with costly attributes that are not fully captured by our model, the estimated coefficient on FP status will be downward-biased when employer-market interactions are excluded.

Similar logic pertains to the two second-order interactions we include next. Plan type-year effects will help to explain premium changes associated with trends such as the convergence

[32] Ideally, we would include carrier fixed effects because carriers are differentiated in terms of quality. However, including this set of indicators would cause the FP effect to be identified only by converting carriers. We pursue that identification strategy in the following section.

of HMO and PPO benefit designs. The "HMO backlash" against utilization review and selective provider networks has caused HMOs to curtail these hallmark features, raising the relative cost of HMOs over time (Draper et al. 2002). Market-year effects will capture any unobservable, time-varying characteristics at the market level that are correlated with premiums, such as the formation of a powerful hospital system that boosts hospital reimbursements and raises total outlays. Again, these terms will only have meaningful impacts on the estimated FP coefficient if enrollment in FP plans is correlated with factors that are absorbed by these terms.

The results are presented in **Table 4**, columns 1 through 4. The top panel displays results for the sample of FI plans; the bottom panel displays the same for SI plans. We weight each observation by the corresponding mean plan-level enrollment and cluster standard errors by plan. The coefficient estimates show that for-profit ownership is associated with premiums that are 4-5 percent higher for FI plans, and 2-2.5 percent higher for SI plans. All estimates are significant at $p < 0.01$. There is modest evidence of favorable selection into FP plans that is correlated with the second-order interaction terms. For example, the FP coefficient increases upon inclusion of employer-market interactions, suggesting employer-market groups with heavy FP reliance have other characteristics associated with lower premiums, even controlling for employer and market fixed effects. When all three sets of second-order interactions are included, the coefficient on FP increases by 0.8 percent and 0.6 percent for the FI and SI plans, respectively, and this difference is statistically significant in the FI sample. These results provide further motivation for an empirical strategy that goes beyond estimating correlations between premiums and ownership status, as there are clearly unobservable factors correlated with both.

Before proceeding to our investigation of a causal link between ownership status and premiums, we briefly note that the coefficients on the control variables are consistent with *a priori* expectations. Increases in *plan design* and *demographic factor* raise premiums, and the magnitude of the implied coefficient on *demographic factor* increases after the recoding in 2006. CDP plans are associated with lower premiums in the SI sample; they are very uncommon in the FI sample so the fluctuating estimates in that sample are unsurprising.

In sum, the estimates indicate that FP status is positively associated with premium levels, controlling for differences across markets, employee groups, plan types, and plan generosity (such as copayments and deductibles). However, there are myriad omitted factors that could explain this phenomenon besides the conjecture that FP insurers are providing less value on the dollar. We therefore turn to conversions as a source of plausibly exogenous variation in market-level FP share.

V. Is there a Causal Effect of FP status on Premiums?

To examine whether there is a causal link between ownership status and premiums, we make use of the FP conversions of Blue Cross plans in eleven states spanning 29 geographic markets. For this analysis, our unit of observation is the market-year, and we create a market-year premium index (discussed in detail below). The equation of interest is

$$\text{(2)} \quad \text{premium index}_{\text{mt}} = \alpha + \phi FP\ share_{mt-1} + \psi_m + \delta_t + \left[X_{mt-1}\right] + \varepsilon,$$

where all notation is consistent with equation (1). All models include market fixed effects, therefore ϕ is identified by changes in market-level FP share. We also include year dummies and the lagged market-year covariates previously described. Observations are weighted by average market-level enrollment, and standard errors are clustered by market. We instrument for $FP\ share_{mt-1}$ using an indicator for the FP status of the BCBS carrier in market *m* and year *t-1,* as well as the interaction between this indicator and the pre-conversion BCBS market share. The following subsections describe the three steps in our analysis in greater detail: constructing the market-year premium index, validating the instrument and estimating first-stage and reduced-form models, and performing IV/2SLS.

A. Constructing a Market-Year Index of Premium Growth

As equation (2) indicates, we use the market-year as the unit of observation for this analysis, which necessitates the creation of a market-year measure of premiums. If instead we relied upon the original unit of observation (plan-year), we would be comparing premium growth for customers of converting BCBS plans with premium growth for customers of non-BCBS plans in the same markets. This strategy is suboptimal for at three reasons. First, there are too few plans in our sample with a sufficiently long panel to permit reliable estimates. Second, the estimates suffer from selection bias: only those customers remaining with their pre-conversion carriers (and plan types) would be included in the analysis, and switching is a response to conversion. Finally, given the oligopolistic nature of most insurance markets, any changes in the pricing of the local BCBS carrier should affect the pricing of competitors, so that a comparison of BCBS and non-BCBS plans in the same markets will understate the pricing effect of a conversion. By using the market-year as the unit of observation, we effectively estimate the market-level impact of conversions, implicitly using markets without converting plans as a control group. As noted above, we allow the impact of conversions to depend on the market penetration of the converting BCBS plan.

To obtain a market-year price index, we estimate the following model:

$$\begin{aligned}(3)\quad \ln(\text{premium})_{emcjt} &= \beta_0 + \beta_1 demographics_{emcjt} + \beta_2 demographics_{emcjt} * (year >= 2006)_t \\ &+ \beta_3 \, plan\ design_{emcjt} + \beta_4 CDP_{emcjt} + \pi_{emcj} + \theta X_{mt-1} \\ &+ \rho_{em} + \kappa_{jt} + \varphi_{mt} + \varepsilon.\end{aligned}$$

The variables of interest are the market-year effects, denoted by φ_{mt}. The coefficients on these terms capture average premium growth for each market and year. Equation (3) is identical to equation (1) from the OLS analysis, but for two key changes. First, we exclude the FP indicator, as the market-year FP share will be the independent variable of interest. Second, we include plan fixed effects (dummies for each employer-market-carrier-plan type-year combination, denoted by π_{emcj}). Doing so implies that the coefficients on the market-year dummies will reflect average growth *for the same exact plan* over time. This specification exploits the richness of the data to minimize the noise associated with changes in sample composition. As previously noted, premium growth in LEHID closely matches premium growth nationwide.[33] We estimate

[33] We do not include plan fixed effects in equation (1) because if we did so the FP effect would be identified solely by converting carriers, which is the objective of this section. Although the plan-level data are not ideal for this analysis for the reasons mentioned in the text, estimating (1) with the addition of plan fixed effects yields a

equation (3) separately for FI and SI plans, weighting each observation by the mean number of enrollees for the relevant plan.

For each market we are able to estimate dummies for 11 out of the 12 years in the sample. For ease of interpretation, we set the index equal to 100 for each market in 1998, and apply the estimated coefficients on the market-year dummies to calculate the index in all subsequent years. (For example, a market-year coefficient of 0.2 would imply an index of 100*(exp(0.2))=122.14).[34]

Descriptive statistics for the premium index, which is calculated separately for FI and SI plans, are included in Table 3. Premium growth is slightly slower for FI than SI plans: the (unweighted) mean index reached 282 and 293, respectively, in 2009. These increases (182 and 193 percent) compare to nominal increases of 140 percent in the average family premium for large firms (200+ employees) reported in the KFF/HRET survey during roughly the same period (1999-2010 rather than 1998-2009).[35] Given we estimate a Laspeyres index which holds product features such as carrier identity and plan generosity constant, we anticipate steeper growth in our index.[36]

We also estimate a version of equation (3) which interacts the market-year dummies with an indicator for BCBS plans and an indicator for non-BCBS plans. We exponentiate the two sets of coefficient estimates to form separate market-year 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. Over the course of the study period, the mean BCBS premium index increased

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coefficient on the FP indicator of -0.027 (0.021). This estimate – which is identified by the eleven BCBS and three (fairly small-sized) non-BCBS conversions which take place during our study period – is consistent with the results we obtain below.

[34] As a robustness check, we also estimated our reduced-form specifications using the untransformed market-year coefficients as the price index (with values for 1998 set to zero across all markets). The results were very similar. We use the transformed version for ease of interpretation.

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

[36] In the face of rising insurance premiums, employers substitute toward cheaper plans, so that realized price growth is lower than predicted price growth holding plan characteristics constant.

from 100 to 251 for FI plans and from 100 to 275 for SI plans. The non-BCBS premium index increased from 100 to 286 for FI plans and from 100 to 292 for SI plans.

B. Estimating the Impact of Conversions

To assess the impact of conversions, we estimate equation (2), replacing the endogenous variable, $FP\ share_{mt-1}$, with an indicator for whether the local BCBS plan is FP. This serves as our reduced-form model. Before reporting these results, however, we discuss the estimates from a specification that includes leads and lags of the BCBS FP indicator:

$$(4)\quad \text{premium index}_{mt} = \phi_0 + \phi_1 BCBS\ FP_{mt+2} + \phi_2 BCBS\ FP_{mt+1} + \phi_3 BCBS\ FP_{mt} + \phi_4 BCBS\ FP_{mt-1} + \phi_5 BCBS\ FP_{mt-2} + \phi_6 BCBS\ FP_{m,t<=t-3} + \psi_m + \delta_t + \pi X_{mt-1} + \varepsilon$$

The purpose of this model is twofold: first, to check whether the leads are statistically insignificant and lack a pronounced trend; second, to examine how the effect of conversions varies over time. The coefficient estimates represent the market-level effect of a conversion in the relevant number of years before or after the conversion, relative to premiums in non-converting markets and premiums in converting markets 3 or more years prior to conversion (after controlling for fixed differences across markets, national year effects, and the market-year covariates).

The coefficient estimates are presented in **Table 5** and graphed in **Figure 3.** We present results separately for the FI and SI premium indices. None of the leads is statistically significant, nor is there a decreasing or increasing pattern in the point estimates. These results support the identifying assumption that conversions are orthogonal to omitted determinants of premiums.

There is little evidence of aggregate price responses to conversions. There is a modest decline in the point estimates between year *-1* (representing the year before conversion) and year *1*, however the estimates are too noisy to reject equality of any of the coefficient estimates. The fact that a decline may occur in year *t* is not surprising given the long, public process preceding conversions, as well as the fact that all but one occurred in the month of November and were therefore coded as taking effect in the following year.[37] However, even the slight (and insignificant) reduction in premiums diminishes over time.

The results from a parsimonious reduced-form specification which replaces the leads and lags with a single indicator - $BCBS\ FP_{mt-1}$ - are displayed in **Table 6**. Panel A corresponds to results using the fully-insured premium index, while Panel B corresponds to the self-insured index. On average, FP conversions have no economically or statistically significant effect on either premium index. This remains true even if the FP indicator is not lagged (results available upon request). In Columns 2 and 3, we investigate whether conversions of plans with greater market dominance have a different effect on our premium index by interacting $BCBS\ FP_{mt-1}$ with *pre-conversion share*. Pre-conversion share is computed using LEHID and refers to the enrollment-weighted average market share of the converting plan across all markets in which the plan operates during the three years preceding conversion. **Figure 4** documents the significant variation in pre-conversion share across markets. The range spans from 9.7% to 28.4%, with a weighted average of 19.2%.

Conceptually, the sign on this interaction term could be positive or negative. On one hand, if NP BCBS plans underprice because they value enrollment, then post-conversion these plans should raise price, *ceteris paribus*. The effect of a price increase on average premiums in the market should be more pronounced in markets where the converting plan is dominant, for two distinct reasons. First, carriers with larger market shares will mechanically have a greater effect on average prices. Second, dominant BCBS carriers may raise price more than less-dominant plans because their enrollees have fewer outside options (i.e. these carriers face lower elasticities of demand). On the other hand, if converting plans successfully lower costs, optimal prices could actually fall.

[37] Eight of the eleven conversions occurred when Anthem demutualized in November 2001. As reported in Table 1, BCBS FP "turns on" for affected markets in 2002. Thus the basic model assumes the first post-pricing year is 2003 (the premiums for which are set in 2002), but it is plausible that prices are affected in 2002.

The coefficients in column 2 suggest that post-conversion premiums increase in pre-conversion market share, for the FI sample only. To investigate this result further, we replaced *pre-conversion share* with an indicator labeled *high pre-conversion share,* which takes a value of one when *pre-conversion share* exceeds the weighted mean of 19.2%. We find strong evidence of premium increases in markets where BCBS was dominant prior to conversion, but only weak evidence of premium reductions where BCBS was relatively smaller. The estimated premium increase in *high pre-conversion share* markets is 16.5 points, which is roughly 13% of the average increase in the FI premium index between 2001 (the modal pre-conversion year) and 2009 (139 points).

In sum, market dominance appears to be associated with larger post-conversion price increases for FI plans. As discussed in Dafny (2010), the opportunity to exercise market power is smaller in the SI segment, due to the greater transparency in pricing. SI premiums consist of medical outlays, premiums for stop-loss insurance purchased by the employer (if any), and administrative fees charged by the insurance carrier. Price changes not associated with provider outlays must therefore occur entirely in this latter category, and are easily observed.

Next, we contrast the post-conversion pricing reactions of BCBS and non-BCBS plans by estimating the specifications in Table 6 using *BCBS Index* and *non-BCBS Index* as the dependent variables. The results are displayed in **Table 7**, again separately for FI plans (Panel A) and SI plans (Panel B). The point estimates in the SI sample are economically and statistically insignificant for both indices, so our discussion focuses on the FI sample.[38] In markets where BCBS is dominant, price increases are apparent for both BCBS plans and their competitors. Consistent with a model in which a dominant converting plan raises price and competitors follow its lead, price increases are larger for BCBS plans, although the difference is not significant.[39]

[38] Note that while the coefficient on pre-conversion share may appear large, once multiplied by average pre-conversion share and added to the average estimated BCBS effect the net prediction is quite small.
[39] The ability of the converting plan to raise premiums likely depends on the competitiveness of the local market, and not just on the plan's market share. We estimated specifications including interactions between the lagged BCBS FP indicator and the HHI of the local market (introduced as a continuous variable and as a series of dummies). The estimates were noisy but the results suggest that conversions raise premiums less in more competitive markets

C. Instrumental Variables Estimates

In **Table 8**, we present instrumental variables estimates of equation (2) above, in which we use the lagged BCBS FP indicator, and/or the lagged BCBS FP indicator and its interaction with *pre-conversion market share* as instruments for the endogenous regressor, the lagged market share of FP carriers in market *m*. First-stage results are given in **Appendix Table 2**. In light of the reduced-form results, which reveal no significant average effect of conversions, it is unsurprising that the IV estimates are statistically insignificant. In the FI sample, all point estimates are negative, and the magnitudes are small relative to the increase over time in the market index. A 25-percentage point increase in local FP share would translate into a 4 point reduction in the premium index. Given the FI index increases on average from 100 to 282 over the timeframe of our study, this decrease is economically small. In the SI sample, the point estimates are closer to zero and even more precisely estimated. Thus, we conclude that there is a fairly precise "zero effect" of FP penetration on premiums *on average*.

D. Alternative Explanations and Robustness Checks

As a check on our results, we perform a falsification exercise using the set of unsuccessful conversion attempts (in seven states and the District of Columbia) listed in Table 1. To the extent that the relationship between premiums and unsuccessful conversions mirrors that between premiums and successful conversions, our preceding results may be spurious, i.e. generated by omitted factors correlated with conversion attempts. **Table 9** presents the results of this falsification exercise. To execute this analysis, we revise our definition for *lagged BCBS FP* status to include both successful and unsuccessful conversion attempts. We then regress the market-year premium index on *lagged BCBS FP* and an interaction between this term and an indicator for unsuccessful attempts. The results (for the fully-insured sample only) are presented in column 1 of Table 9.[40] Both coefficient estimates are small and noisy, so little can be concluded from this specification. However, the next specification (column 2), which adds

[40] Given our statistically significant findings are limited to the fully-insured sample, this is the most relevant sample to use for the falsification exercise. Estimates using the self-insured sample (available upon request) were too noisy to be informative. None of the coefficient estimates in this sample is statistically significant.

*lagged BCBS FP*preconversion share* and the triple-interaction term *lagged BCBS FP*preconversion share* unsuccessful conversion* is more informative. In the FI sample, where we found economically significant results for successful conversions, the terms including *unsuccessful conversion* have the opposite sign and are of similar magnitude as the corresponding terms excluding *unsuccessful conversion*. That is, "reactions" to unsuccessful conversions did not mirror reactions to successful conversions.

We also conducted a series of robustness checks on our main results. First, we estimated our models using *pre-conversion share* derived from an alternative source, the American Medical Association (AMA).[41] Although the AMA data reflect a different sample (specifically, only plans offered by insurers participating in the FI market), the correlation is relatively high (0.56). The results obtained using this measure are very similar to those presented in Table 6, i.e. in markets where the BCBS plan is dominant, we observe post-conversion premium increases in the FI sample.[42]

Second, we considered sample restrictions. First, to minimize the potential influence of noisy estimates of the premium index and market shares, we dropped all market-years containing fewer than 20 sampled employers. Second, we excluded data from 2006 onward (when the change in the construction of *demographic factor* took effect). Results were similar to those obtained using the complete sample, and are available upon request.[43]

[41] The American Medical Association issues annual reports entitled "Competition in Health Insurance: A Comprehensive Study of U.S. Markets." These include the market share of the top two insurers in each state. In almost all states with conversions, BCBS plans rank among the top two insurers. For states where BCBS is not among the top two insurers, or where the AMA does not have data in pre-conversion years, we use AMA data to construct average BCBS share in later years and use that as a substitute. Before doing so, we confirmed that BCBS market share was stable within a state. BCBS share in Indiana fluctuated a great deal over time, so we excluded data from Indiana for the purposes of this robustness check.

[42] Since the AMA measure excludes data from the state of Indiana, we estimated our reduced-form models (that use LEHID measures) excluding observations from Indiana and confirmed that our estimates were similar to the coefficients in Table 6.

[43] To perform each robustness check, we re-estimated the premium index model using the relevant restricted sample, (i.e. excluding market-years with fewer than 20 clients, including data from 1998-2005 only).

VI. Effects of Ownership Status on Non-Price Outcomes

In this section, we evaluate the impact of ownership status on access to insurance and medical spending (as a share of premium revenues). Coupled with the results on price, these analyses provide a more complete assessment of the implications of ownership status in the health insurance industry.

A. Does Ownership Status Affect Medical Loss Ratios?

We first examine the impact of conversions on insurer Medical Loss Ratios (MLRs), defined as the proportion of premium revenue spent by insurers on medical expenses, as opposed to profits or administrative expenses. As noted in section III, we calculate MLRs by state and year, first for all insurers and then separately for BCBS and non-BCBS plans. The MLR data pertains only to FI plans. Given we have a single pre-conversion year of data (2001), we perform our analysis in long-differences: we regress changes in MLR between 2001 and 2005 (2 years following the last conversion)[44] at the state level on the corresponding changes in state FP share, using our usual instruments.[45] We aggregate our standard controls (unemployment rates and log of Medicare spending) to the state-year level[46] and include long differences of these variables as well.

The results are presented in panels A, B and C of **Table 10**, corresponding to the aggregate MLR, BCBS MLR, and non-BCBS MLR respectively. Within each panel, Columns 1 through 3 report results from reduced form regressions while Column 4 presents the IV estimates. The reduced form results show that MLRs increased by 2 percentage points, on average, post-conversion; this compares to a mean state MLR of 88 percent in 2001 (84 percent in 2005). Panels B and C reveal that this increase is driven by non-BCBS insurers. In addition,

[44] To reduce noise in the MLR measure, we censor the changes in MLR at the top and bottom 5 percent of the distribution

[45] Most conversions are recorded in the data in 2002. New York's Empire plan converted in November 2002, so it is recorded as having taken place in 2003. We estimated models using 2004 as an end-year as well. Results were similar and are available from the authors upon request.

[46] Specifically, we use total enrolment in Medicare Part A and Part B while aggregating the AAPCC data to the state-year level. The unemployment rate data is available at the state-year level at <http://www.bls.gov/schedule/archives/all_nr.htm#SRGUNE>

these insurers experience significantly higher relative increases in MLR in markets where BCBS had high *pre-conversion share.* One possible interpretation is that rivals to BCBS increased their medical spending in an effort to retain and/or attract members, while BCBS either lost share or increased quality in other ways. An alternative explanation is that BCBS plans may have engaged in greater selection efforts following conversion, which would simultaneously raise MLRs for competitors and reduce MLRs for BCBS. The finding that weighted average MLRs (across all insurers) increase post-conversion is consistent with higher reimbursement rates paid by non-BCBS plans (and no change in quantity of services provided), and/or greater overall quantity of services provided (which could be viewed as inferior utilization management).

B. 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 several states (e.g. Michigan). Thus, even if pricing is no different across the ownership forms, NPs may serve the community by issuing policies for high-risk and/or low-profit populations. To assess whether ownership does affect access to insurance, we obtained data on state-level uninsured rates for the under-65 population from the Census Current Population Survey (CPS) Annual Social and Economic Supplement (ASEC).[47] Consistent with the MLR analysis above, we estimate specifications in long differences. In addition to the standard controls, we also control for state-specific changes in Medicaid eligibility using long differences in "simulated Medicaid eligibility" for children aged 0-18, per Currie and Gruber (1996a, 1996b).[48]

The results are presented in **Table 11**. (Although not reported in the table, estimates obtained without the control variables are very similar.) All point estimates are negative and statistically insignificant; they are also economically small. There is no evidence to suggest that BCBS conversions harmed access to insurance, a conclusion echoed by several of the conversion case

[47] These data are available online at http://www.census.gov/hhes/www/hlthins/data/historical/. In keeping with the preceding analyses, we examine long differences between 2001 and two different end-years (2004 and 2005).
[48] We are grateful to Kosali Simon for providing us with estimates of simulated eligibility for the population aged 0-18, by state and year. We construct simple averages for each state and year and use the relevant long differences as a control. Results are not affected by inclusion of this control.

studies (e.g. Conover et al 2003). This is so even in states where the converting plan is dominant.

VII. Discussion and Conclusions

The U.S. health insurance industry has long been criticized for sundry business practices ranging from pre-existing condition exclusions to lifetime benefit caps. Distrust of the industry peaked in the months prior to the passage of PPACA, resulting in multiple alternative proposals to generate new options under healthcare reform. The final bill included $6 billion in funding (eventually reduced by one-third) "to foster the creation of new nonprofit member-run health insurance issuers" (c.f. §1301), notwithstanding the limited evidence on differences between existing NP and FP insurers.

In this study, we use a large, national panel dataset on employer-provided insurance between 1998 and 2009 to study the effect of ownership status on premiums, an important dimension of plan performance. We supplement this analysis with data on medical loss ratios and uninsurance rates. We arrive at a number of conclusions. First, we find fully-insured premiums are 4-5 percent higher when offered by for-profit insurers, and self-insured premiums are 2-2.5 percent higher, even accounting for differences associated with employee risk profiles, geographic location, and generosity of benefits. Second, we find this association is misleading, as our 2SLS estimates using conversions of 11 BCBS plans as instruments for local market FP share yield fairly precise "zeroes." Third, for fully-insured plans, we find heterogenous effects across markets with different degrees of BCBS activity. Specifically, we estimate that fully-insured premiums increased marketwide when converting BCBS plans had shares in excess of ~20 percent. While it is possible that price increases were accompanied by quality improvements, so that quality-adjusted price did not change, the fact that price changes were similar for both BCBS and its competitors suggests oligopolistic pricing behavior. Importantly, we do not find the same results in markets where the BCBS carrier attempted to convert, but failed: this suggests our findings are the result of conversion itself, rather than an omitted factor correlated with the attempt to convert.

Fifth, we find that BCBS conversions had no significant impact on access to insurance, as proxied by state uninsurance rates. Conversions did lead to an increase in insurer medical loss ratios, driven largely by increases in spending by BCBS' competitors in the wake of a conversion. Our back-of-the-envelope calculations using coefficient estimates from Tables 6 and 10 imply that increases in medical spending more than offset increases in premiums post-conversion (or, medical spending decreases were smaller than premium decreases).

Turning to other outcomes that are frequently discussed in the press about FP insurers, we find that BCBS conversions had no significant impact on access to insurance but did lead to an increase in insurer medical loss ratios, driven largely by increases in spending by BCBS' competitors in the wake of a conversion. Our back-of-the-envelope calculations using coefficient estimates from Tables 6 and 10 imply that increases in medical spending more than offset increases in premiums post-conversion (or, medical spending decreases were smaller than premium decreases).

Some important caveats to our findings are in order. First, as noted above, price increases attributable to large increases in FP market share may have been accompanied by quality improvements, such as electronic access to health claims, faster claim processing, and broader provider networks. Second, our findings vis-à-vis price pertain to the large group insurance market, and the conclusions may not extend to the small group and individual insurance markets. Third, our results are sensitive to the source of identifying variation: the change in behavior for converting BCBS plans may not reflect the average difference between new NP and FP carriers.

Notwithstanding these caveats, there are some clear implications for fiscal and competition policy vis a vis insurers. First, it appears that sizeable FP insurers are more likely to exercise market power via price increases than are comparable NP insurers. Second, pricing actions by dominant 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 NP and FP insurers charge different prices *on average*, and in particular when both are small. Thus, there is little empirical support for the proposition that de novo NP insurers (such as those sponsored under provisions of the ACA) will offer lower premiums and, in so doing, spur lower prices by competing insurers. Our results are, however, silent on the important issue of quality of services provided; this is a key area for future research.

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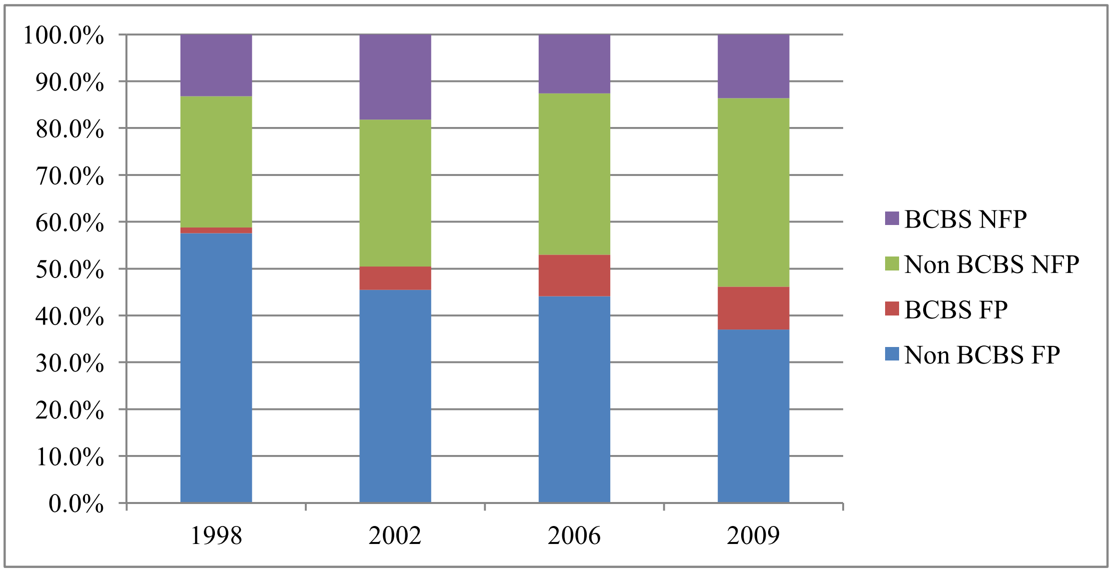
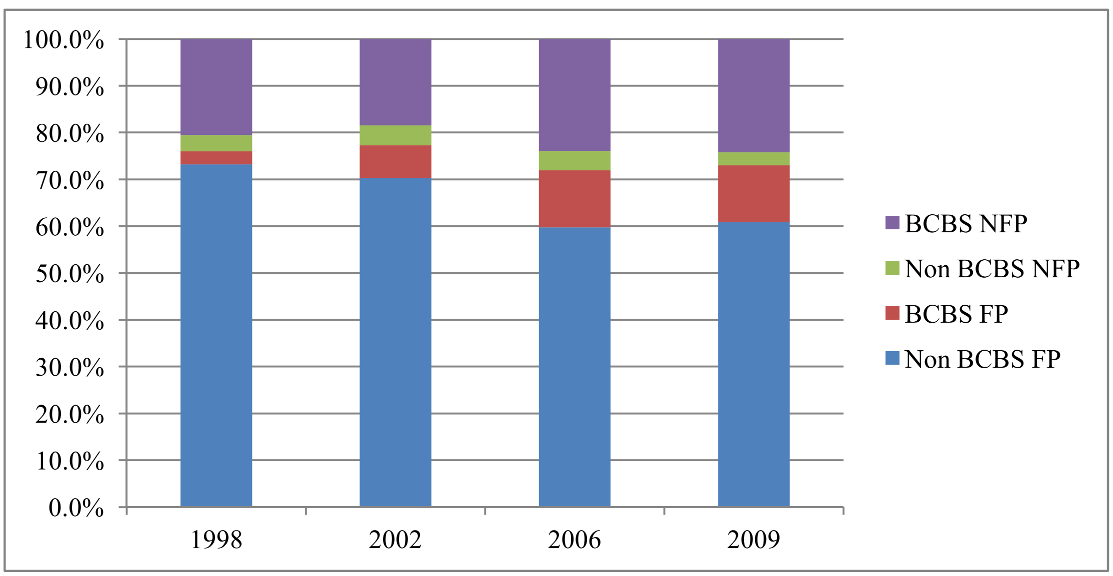
Figure 1. Percent of Enrollees in For-profit and Not-for-profit Plans, by BCBS Affiliation

Panel A. Fully Insured Plans



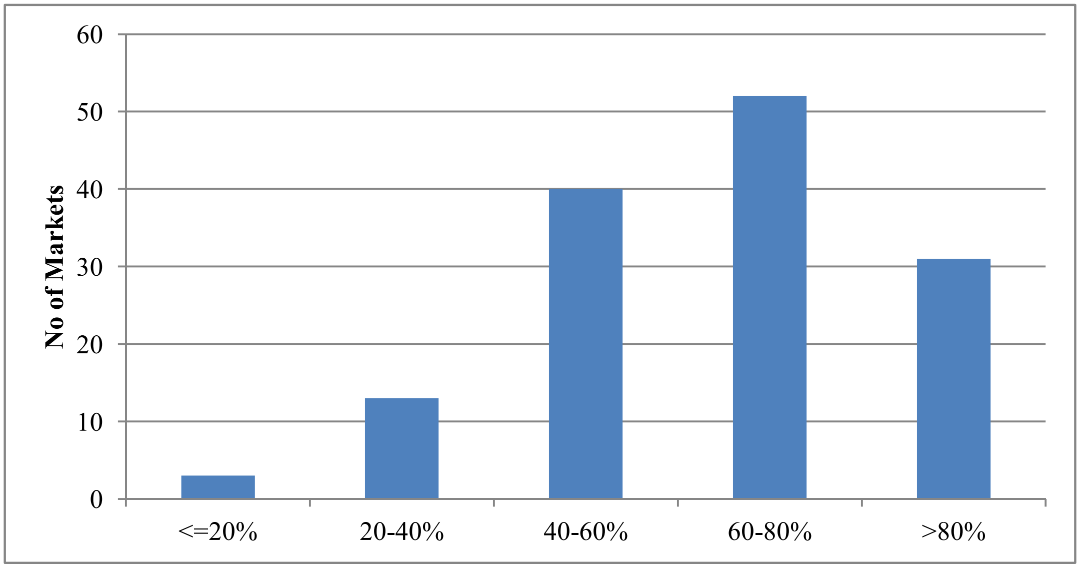
Panel B. Self Insured Plans



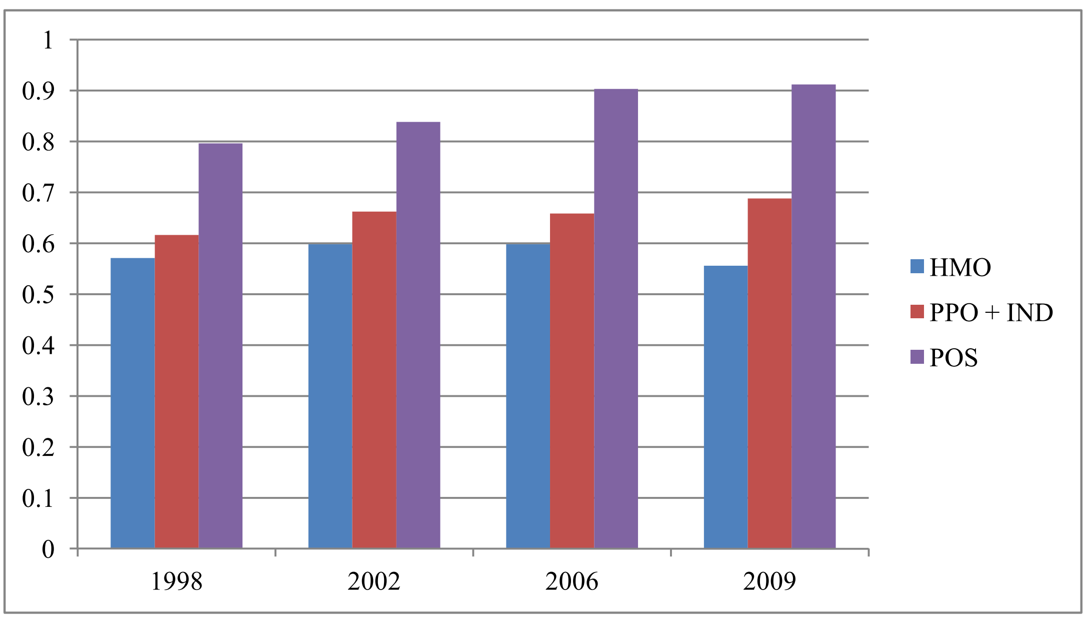
Note: The market shares of all plans above are calculated using LEHID. Sample includes fully insured plans and excludes plans with missing ownership information

Figure 2. For-Profit Share Statistics, by Market and Product

Panel A. Frequency Distribution: Average Percent For-Profit Share by Market



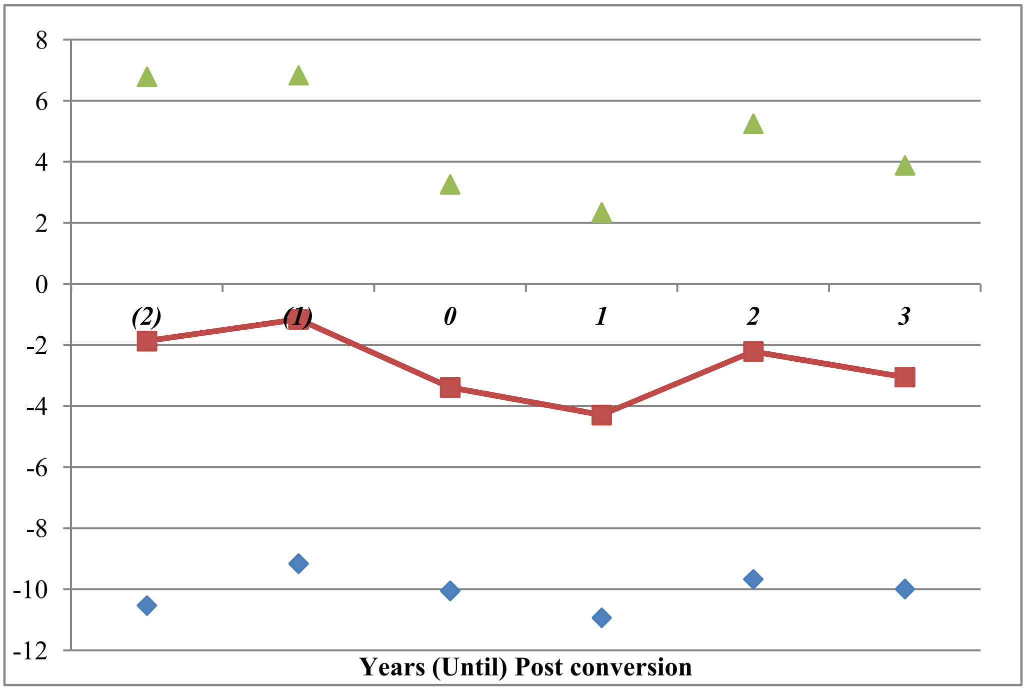
Panel B. Percent For-Profit Share by Plantype



Note : Sample includes fully-insured and self-insured plans. In panel B, each bar represents % of enrollees in plans of that particular plantype that are insured in forprofit plans

Figure 3. Estimated Coefficients (and Confidence Intervals) from Table 5

Panel A. Fully Insured Plans



Panel B. Self Insured Plans

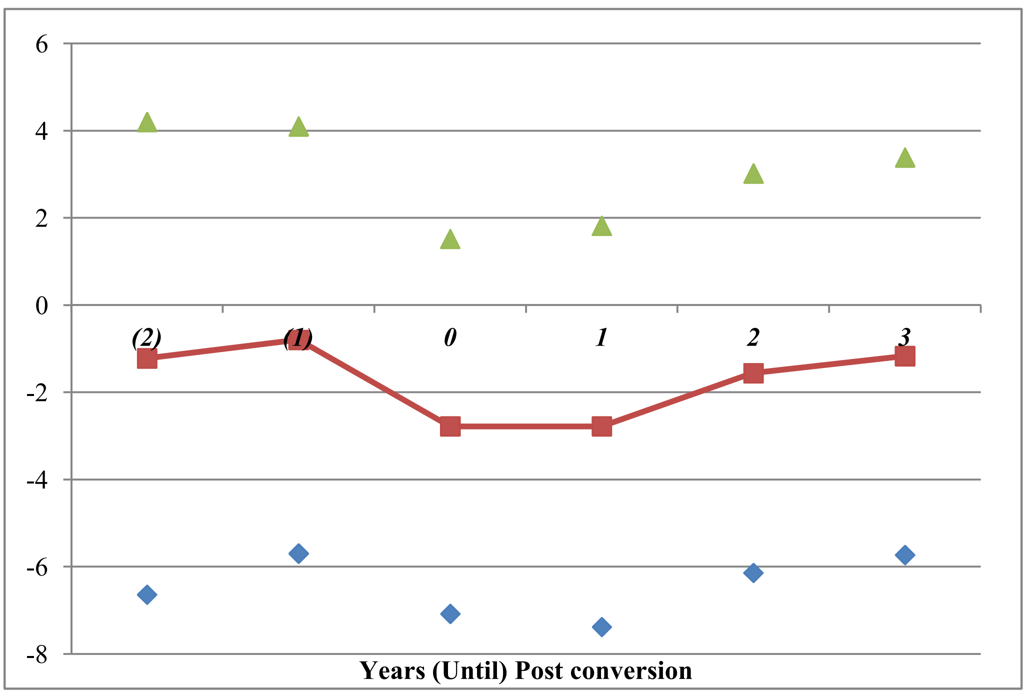
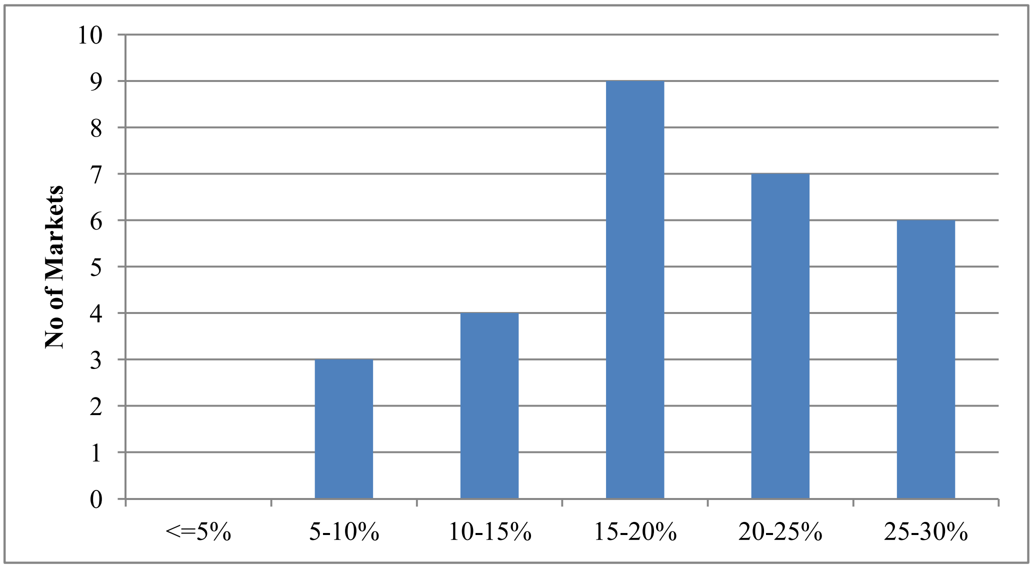


Figure 4. Frequency Distribution: Pre-conversion BCBS share



Note: Pre-conversion share is computed using LEHID and refers to the enrollment-weighted average market share of the converting plan across all markets in which the plan operates during the three years preceding conversion.

Table 1. Blue Cross Blue Shield For-Profit Conversions, 1998-2009

Panel A. Successful Conversions

	Conversion to FP Stock Company	Year Recorded in Data
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

Panel B. Unsuccessful Conversion Attempts

	Review Period	Ended	Year Recorded in Data
New Jersey (Horizon)	2001-2004	2005	2005
North Carolina	2002	July 2003	2004
Kansas	2001-2002	Áugust 2003	2004
CareFirst			
Delaware	2002	September 2003	2004
District of Columbia	2002	September 2003	2004
Maryland	2002	September 2003	2004
Premera			
Alaska	2002-2006	March 2007	2007
Washington	2002-2006	March 2007	2007

<u>Note:</u> *Table only includes successful (and unsuccessful) conversion attempts undertaken during the study period (1998-2009). Year Recorded in Data refers to the first post-conversion year as coded in our dataset.*

Table 2. Descriptive Statistics

	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
			Panel A. Fully Insured Plans									
Premium ($)	3648.63	3905.39	4187.62	4621.73	5338.10	6001.61	6696.46	7327.08	7898.05	8173.10	8571.84	9196.51
	995.58	*918.53*	*1009.57*	*1130.74*	*1377.43*	*1471.79*	*1829.66*	*2205.34*	*2660.74*	*2325.35*	*2537.40*	*2913.91*
Number of Enrollees	170.77	173.47	164.39	179.82	189.88	189.33	176.66	212.15	227.12	226.93	191.57	184.45
	487.60	*488.19*	*420.13*	*489.04*	*498.43*	*559.39*	*504.35*	*686.08*	*729.85*	*689.55*	*597.58*	*617.54*
Demographic Factor	2.23	2.22	2.20	2.23	2.26	2.28	2.36	2.35	1.88	1.88	1.86	1.90
	0.44	*0.39*	*0.39*	*0.38*	*0.39*	*0.39*	*0.40*	*0.45*	*0.42*	*0.41*	*0.44*	*0.44*
Plan Design	1.12	1.13	1.11	1.13	1.12	1.11	1.09	1.06	1.05	1.04	1.03	1.03
	0.05	*0.04*	*0.04*	*0.04*	*0.04*	*0.04*	*0.08*	*0.06*	*0.09*	*0.06*	*0.06*	*0.06*
Plan Type												
HMO	88.7%	90.2%	92.7%	91.8%	90.9%	92.6%	85.6%	88.3%	84.9%	86.1%	79.8%	77.0%
Indemnity	3.0%	0.7%	0.4%	0.1%	1.0%	0.6%	1.9%	0.6%	0.5%	0.4%	1.2%	2.6%
POS	6.6%	6.8%	4.0%	4.8%	3.1%	4.2%	4.0%	4.4%	3.7%	4.1%	4.2%	2.7%
PPO	1.7%	2.4%	2.9%	3.2%	5.0%	3.2%	8.5%	6.7%	10.9%	9.5%	14.8%	17.7%
% Insured in CDP	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	0.1%	3.1%	2.3%	2.5%	3.6%
Number of Employers	181	197	185	226	226	274	194	203	173	184	173	168
Number of Observations	9885	11584	9256	10673	10252	8892	6386	5855	4590	4414	4713	4299
			Panel B. Self Insured Plans									
Premium ($)	4276.73	4306.54	4579.53	5012.24	5644.13	6459.38	6999.15	7474.24	7894.39	8126.57	8411.10	8897.68
	1134.05	*1298.88*	*1318.39*	*1356.22*	*1436.53*	*1583.46*	*1701.01*	*1803.08*	*1946.40*	*2132.63*	*2156.29*	*2284.09*
Number of Enrollees	190.86	159.10	151.21	170.13	167.15	174.38	169.94	191.90	178.82	189.73	156.38	167.16
	725.82	*601.23*	*507.27*	*578.63*	*611.85*	*638.77*	*530.08*	*859.48*	*609.01*	*578.65*	*554.95*	*663.03*
Demographic Factor	2.42	2.29	2.26	2.27	2.28	2.30	2.32	2.32	1.84	1.82	1.83	1.88
	0.52	*0.45*	*0.45*	*0.44*	*0.46*	*0.43*	*0.43*	*0.43*	*0.40*	*0.40*	*0.41*	*0.39*
Plan Design	1.02	1.01	0.99	1.02	1.03	1.02	1.01	0.97	0.97	0.98	0.97	0.97
	0.07	*0.07*	*0.08*	*0.07*	*0.07*	*0.08*	*0.08*	*0.08*	*0.08*	*0.07*	*0.08*	*0.07*
Plan Type												
HMO	2.6%	4.3%	6.9%	10.1%	14.7%	16.8%	17.6%	18.1%	19.7%	20.1%	20.5%	18.6%
Indemnity	34.5%	31.8%	22.1%	16.7%	14.2%	10.5%	7.8%	5.9%	5.9%	5.3%	4.7%	3.5%
POS	36.1%	27.5%	30.5%	25.3%	20.6%	17.9%	18.2%	15.9%	15.7%	14.9%	14.2%	14.5%
PPO	26.9%	36.4%	40.5%	47.9%	50.5%	54.7%	56.4%	60.1%	58.8%	59.6%	60.6%	63.4%
% Insured in CDP	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	*N/A*	10.3%	18.3%	20.0%	21.9%	22.6%
Number of Employers	180	193	191	233	248	315	238	257	223	247	235	218
Number of Observations	12199	14140	14411	18479	21319	24906	20233	20970	17770	19865	23510	21434

Notes: *All statistics are unweighted. The unit of observation is an employer-carrier-market-plantype-year combination, unless noted otherwise. Demographic factor reflects age, gender, and family size for 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 italics.*

Table 3. Descriptive Statistics (Unit of Observation: Market-Year)

	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Market-year Controls												
Lagged Medicare Costs	4896.21	4896.21	4872.10	5165.41	5603.25	6061.54	6372.18	6836.88	7288.65	7591.73	7898.36	8297.57
	866.92	*866.92*	*820.15*	*869.75*	*924.35*	*992.71*	*993.30*	*989.65*	*1095.12*	*1096.89*	*1123.36*	*1198.25*
Lagged unemployment rate	0.05	0.05	0.04	0.04	0.05	0.06	0.06	0.05	0.05	0.05	0.05	0.06
	0.02	*0.02*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.01*	*0.02*
Number of Markets	139	139	139	139	139	139	139	139	139	139	139	139
Panel A. Fully Insured Plans												
Dependent Variable												
Premium Index	100.00	109.38	120.95	131.95	150.75	173.29	191.55	209.86	236.25	251.67	268.36	281.98
	0.00	*4.73*	*7.61*	*10.60*	*13.01*	*16.66*	*22.54*	*23.84*	*26.38*	*32.82*	*37.22*	*36.69*
Number of Markets	139	139	139	139	139	137	138	138	138	138	138	139
Panel B. Self Insured Plans												
Dependent Variable												
Premium Index	100.00	105.47	112.83	127.55	143.31	169.11	192.69	211.62	244.33	263.02	278.24	292.60
	0.00	*6.50*	*8.21*	*10.54*	*11.13*	*13.95*	*17.69*	*20.15*	*23.81*	*25.20*	*24.12*	*26.93*
Number of Markets	139	139	139	139	139	139	139	139	139	139	139	139

Notes: *All statistics are unweighted. The unit of observation is a market-year combination, for each insurance type. Market-level premium index is constructed using a plan-level regression of premiums on various controls, including market-year fixed effects. Standard deviations are in italics.*

Table 4. Do Premiums Differ By Ownership Status?

	Panel A. Fully Insured Plans			
	Dependent Var = Ln(Premium)			
	(1)	**(2)**	**(3)**	**(4)**
Lagged FP Indicator	0.038	0.041	0.044	0.046
	*(0.003)****	*(0.003)****	*(0.003)****	*(0.003)****
Demographic factor	0.379	0.392	0.385	0.385
	*(0.006)****	*(0.008)****	*(0.008)****	*(0.007)****
Demographic factor * (Year>=2006)	0.071	0.071	0.074	0.073
	*(0.012)****	*(0.014)****	*(0.012)****	*(0.009)****
PlanDesign	0.420	0.366	0.356	0.446
	*(0.051)****	*(0.062)****	*(0.061)****	*(0.052)****
CDP	-0.095	-0.132	0.016	-0.002
	*(0.052)**	*(0.069)**	*(0.031)*	*(0.021)*
Employer-Market FE		Yes	Yes	Yes
Plantype-Year FE			Yes	Yes
Market-year FE				Yes
Number of Observations	80841	80841	80841	80841

	Panel B. Self Insured Plans			
	Dependent Var = Ln(Premium)			
	(1)	**(2)**	**(3)**	**(4)**
Lagged FP Indicator	0.019	0.022	0.024	0.025
	*(0.004)****	*(0.004)****	*(0.004)****	*(0.004)****
Demographic factor	0.364	0.397	0.395	0.396
	*(0.005)****	*(0.008)****	*(0.008)****	*(0.008)****
Demographic factor * (Year>=2006)	0.041	0.047	0.033	0.029
	*(0.006)****	*(0.006)****	*(0.006)****	*(0.006)****
PlanDesign	0.708	0.665	0.737	0.755
	*(0.035)****	*(0.038)****	*(0.036)****	*(0.034)****
CDP	-0.076	-0.078	-0.067	-0.068
	*(0.005)****	*(0.005)****	*(0.005)****	*(0.005)****
Employer-Market Fixed Effects		Yes	Yes	Yes
Plan type-Year Fixed Effects			Yes	Yes
Market-year Fixed Effects				Yes
Number of Obs	213218	213218	213218	213218

Note: *The unit of observation is a plan-year. All specifications include market-year controls and employer, plan type, year, and market fixed effects. Models are estimated by weighted least squares using the average number of enrollees per plan as weights. Standard errors are clustered by plan.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 5. Exogeneity Check: Effect of BCBS Conversions in Years Preceding (and Following) Conversion

	Dependent Var = Premium Index	
	(1)	**(2)**
	Fully Insured Plans	***Self Insured Plans***
(BCBS FP) $_{t+2}$	-1.33	-1.33
	(2.24)	*(1.72)*
(BCBS FP) $_{t+1}$	-0.52	-0.68
	(3.42)	*(2.47)*
(BCBS FP) $_{t}$	-2.82	-2.93
	(5.13)	*(3.36)*
(BCBS FP) $_{t-1}$	-3.80	-3.13
	(6.02)	*(3.57)*
(BCBS FP) $_{t-2}$	-1.45	-2.49
	(7.36)	*(4.04)*
(BCBS FP) $_{>=(t-3)}$	-2.37	-1.75
	(7.69)	*(4.62)*
Number of Observations	1646	1660

Notes *: The unit of observation is the market-year. All models include market-year controls and fixed effects for each market and year, and are estimated by weighted least squares using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 6. Effect of BCBS Conversions on Premiums

	Panel A. Fully Insured Plans			*Panel B. Self Insured Plans*		
	Dependent Var = Premium Index			**Dependent Var = Premium Index**		
	(1)	**(2)**	**(3)**	**(1)**	**(2)**	**(3)**
Lag BCBS FP	-1.47	-32.83		-1.08	-7.71	
	(5.54)	*(15.25)***		*(2.89)*	*(6.65)*	
Lag BCBS FP*Pre-conversion share		168.13			34.05	
		*(66.92)***			*(30.45)*	
Lag BCBS FP*						
Low Pre-conversion share			-9.01			-3.48
			(5.69)			*(3.23)*
High Pre-conversion share			16.48			2.97
			*(5.32)****			*(4.06)*
Number of Obs	1646	1646	1646	1660	1660	1660

Notes *: The unit of observation is the market-year. All models include market-year controls and fixed effects for each market and year, and are estimated by weighted least squares using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 7. Effect of BCBS Conversions on Premiums: BCBS Vs. Non-BCBS Plans

Panel A. Fully Insured Plans

	Dep Var = Premium Index (BCBS)			Dep Var = Premium Index (Non-BCBS)		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Lag BCBS FP	8.38	-40.04		-1.18	-38.23	
	(7.19)	*(28.77)*		*(5.57)*	*(14.84)***	
Lag BCBS FP*Pre-conversion share		258.50			198.63	
		*(142.73)**			*(64.69)****	
Lag BCBS FP*						
Low Pre-conversion share			-2.87			-9.09
			(8.19)			*(5.56)*
High Pre-conversion share			34.21			17.67
			*(9.29)****			*(5.66)****
Number of Observations	1537	1537	1537	1612	1612	1612

Panel B. Self Insured Plans

	Dep Var = Premium Index (BCBS)			Dep Var = Premium Index (Non-BCBS)		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Lag BCBS FP	1.19	-1.38		0.45	-11.51	
	(2.16)	*(5.57)*		*(4.275)*	*(8.99)*	
Lag BCBS FP*Pre-conversion share		13.23			61.52	
		(27.12)			*(46.16)*	
Lag BCBS FP*						
Low Pre-conversion share			0.11			-3.88
			(2.58)			*(3.98)*
High Pre-conversion share			3.03			7.79
			(2.85)			*(7.57)*
Number of Observations	1652	1652	1652	1660	1660	1660

***Notes**: The unit of observation is the market-year. All models include market-year controls, fixed effects for each market and year, and are estimated by weighted least squares using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 8. Does FP Penetration Raise Premiums? IV Estimates

	Panel A. Fully Insured Plans	
	Dependent Var = Premium Index	
	IV = Lag BCBS FP	**IV = {Lag BCBS FP, Lag BCBS FP * Pre-conversion BCBS Share}**
Lag FP Penetration	-15.07	-13.43
	(33.47)	*(33.51)*
Number of Obs	1507	1507

	Panel B. Self Insured Plans	
	Dependent Var = Premium Index	
	IV = Lag BCBS FP	**IV = {Lag BCBS FP, Lag BCBS FP * Pre-conversion BCBS Share}**
Lag FP Penetration	-4.59	-2.73
	(8.54)	*(7.62)*
Number of Obs	1521	1521

Notes *: The unit of observation is the market-year. All models include market-year controls, fixed effects for each market and year, and are estimated by weighted two-stage least squares using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 9. A Falsification Exercise: Do Unsuccessful Conversion Attempts Affect Premiums?

	Fully Insured Plans	
	Dependent Var = Premium Index	
	(1)	**(2)**
Lag BCBS FP	-2.23	-33.51
	(5.65)	*(15.34)***
Lag BCBS FP * unsuccessful conversion	-6.84	18.84
	(5.85)	*(18.51)*
Lag BCBS FP * Pre-conversion share		167.82
		*(67.22)***
Lag BCBS FP * Pre-conversion share * unsuccessful conversion		-142.90
		*(82.15)**
Number of Obs	1646	1646

Notes *: The unit of observation is the market-year. Sample includes only fully insured plans. BCBS FP status is set to 1 for all markets in which there was an attempt at conversion, either successful or unsuccessful. All models include market-year controls and fixed effects for each market and year, and are estimated by weighted least squares using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 10. Impact of BCBS Conversions on Medical Loss Ratios

	Study Period: 2001 - 2005											
	Panel A: Dep Var = Δ MLR				**Panel B: Dep Var = Δ MLR (BCBS)**				**Panel C: Dep Var = Δ MLR (Non-BCBS)**			
	Reduced Form			**IV**	**Reduced Form**			**IV**	**Reduced Form**			**IV**
	(1)	**(2)**	**(3)**	**(4)**	**(1)**	**(2)**	**(3)**	**(4)**	**(1)**	**(2)**	**(3)**	**(4)**
BCBS conversion	0.020	0.015			0.004	0.008			0.026	0.009		
	*(0.010)**	*(0.020)*			*(0.016)*	*(0.017)*			*(0.011)***	*(0.017)*		
BCBS conversion*pre-conversion share		0.036				-0.039				0.119		
		(0.091)				*(0.061)*				*(0.081)*		
BCBS conversion*												
Low Pre-conv share			0.018				0.002				0.025	
			*(0.010)**				*(0.015)*				*(0.011)***	
High Pre-conv share			0.023				-0.009				0.042	
			*(0.013)**				*(0.016)*				*(0.014)****	
Δ Lagged FP share				0.083				0.002				0.131
				*(0.039)***				*(0.071)*				*(0.044)****
# Observations	50	50	50	50	38	38	38	38	46	46	46	46

Notes *: MLRs are constructed using censored company-state-year data. The long differencess computed for MLR are also censored at the top and bottom 5% of the distribution. All specifications include appropriate long differencesof market-year control variables and are weighted by the average number of enrollees in the state in the end-years. In column 4 of each panel, the interaction between Lag BCBS FP and Pre-conversion share is used as an instrument for change in Lagged FP share in the state. Standard errors are robust.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Table 11. Impact of BCBS Conversions on State Uninsurance Rate

	Dependent Variable = Δ in Uninsured Rate, 2001-2005			
	Reduced Form			**IV**
Lag BCBS FP	-0.005	-0.003		
	(0.005)	*(0.008)*		
Lag BCBS FP * Pre-conversion Share		-0.023		
		(0.040)		
Lag BCBS FP*				
Low Pre-conversion share			-0.005	
			(0.007)	
High Pre-conversion share			-0.009	
			(0.005)	
Lagged FP share				-0.028
				(0.020)
# Observations	51	51	51	51

Notes *: The unit of observation is the state. All specifications include long differencesof market-year control variables and are weighted by the average number of enrollees in the state in the end-years. Lagged BCBS FP and Lagged BCBS FP * Conversion share are used as instruments in the IV specification. Standard errors are robust.*

** denotes p<0.10, ** denotes p<0.05, *** denotes p<.01*

Appendix: Representativeness of the Large Employer Health Insurance Dataset (LEHID)

As stated in the text, LEHID consists primarily of large, multisite employers.[1] In order to establish representativeness of this sample, we now present a comparison of LEHID to the two leading alternative sources of insurance data: the Kaiser Family Foundation/Health Retirement Education Trust (KFF/HRET) Annual Survey of Employer Benefits, and the proprietary Interstudy database of insurer data. The KFF/HRET survey randomly samples public and private employers to obtain national statistics on employer-sponsored health insurance; approximately 2000 employers respond each year. The data are not publicly available, nor is the sample designed to provide estimates at the market level. However, the survey is designed to yield representative estimates of national trends.

Appendix Figure 1 reports the annual growth rate in premiums for a family of four in an employer-sponsored plan. As in LEHID, both employer and employee premium contributions are combined, and both fully and self-insured plans are included. However, LEHID does not report premiums for a standard family size. Thus, to obtain a comparable measure from the LEHID sample, we divide the average annual premium in LEHID by the demographic factor. According to our source, this yields the premium per "person equivalent." Annual growth rates for this "individual" premium are reported in Appendix Figure 1 as well. The trends are quite similar throughout the period.

We also compare our measures of market concentration with measures constructed by other researchers using the proprietary InterStudy database. InterStudy reports enrollment and premium figures at the insurer and MSA level. We compare the HHI and number of carriers tabulated by Scanlon et al (2008) to the corresponding figures from the LEHID. [2]

Before describing the results, we note the InterStudy data is not directly comparable to LEHID for several reasons. The InterStudy data includes only fully-insured HMO plans for the

[1] More than 96 percent of enrollees represented in LEHID are employed by firms that have more than 5000 employees. This compares to a national figure of 37 percent across all firms (KFF, 2010).
[2] Our sincere thanks to Mike Chernew, Dennis Scanlon and Woolton Lee for sharing their estimates of market structure. For details on the construction of the InterStudy HHIs, see Scanlon et al (2006).

time period we consider, and the allocation of enrollment across geographic markets is fairly noisy. In addition to these issues, the LEHID geographic markets, which generally correspond to MSAs (but may include multiple MSAs), are often larger than the Interstudy markets.[3]

To compare measures of insurer market structure derived from the two sources, we begin by mapping MSAs to the corresponding LEHID markets.[4] When multiple MSAs comprise one LEHID market, we weight the InterStudy MSA measures of market structure by the population of that MSA (obtained from the 2000 Census) to create measures of insurer market concentration (HHI, number of carriers) for each geographic market defined in the LEHID dataset.

When we use all plans in the LEHID dataset to construct HHI (as in our regression models), the correlation coefficient between the two measures is 0.18 over the entire sample period (1998-2006). This figure rises to 0.31 when we restrict attention to HMO plans only.[5] As is apparent in **Appendix Figure 2**, there are also some differences between the two estimates when we compare trends over time. The LEHID HHI exhibits fairly steady growth in the latter half of the study period while the Interstudy HHI peaks in 2003. Unfortunately, there are no obvious explanations for these discrepancies.

We use the LEHID-based HHI estimates for theoretical and practical reasons. First, the set of carriers that serve large, multisite firms such as those included in LEHID may differ from the set of carriers at large. Thus, LEHID itself likely offers the best estimate of the relevant insurance market structure. Second, the InterStudy data does not consistently include PPO enrollment during our study period, and PPOs account for a large share of our data. Third, as noted above, researchers have documented serious concerns about the way in which InterStudy allocates enrollment across MSAs. Finally, the InterStudy data is quite expensive to acquire.

One further concern with the LEHID is that the probability of being included in the sample may vary substantially across areas. However, Dafny (2010) reports that the ratio of

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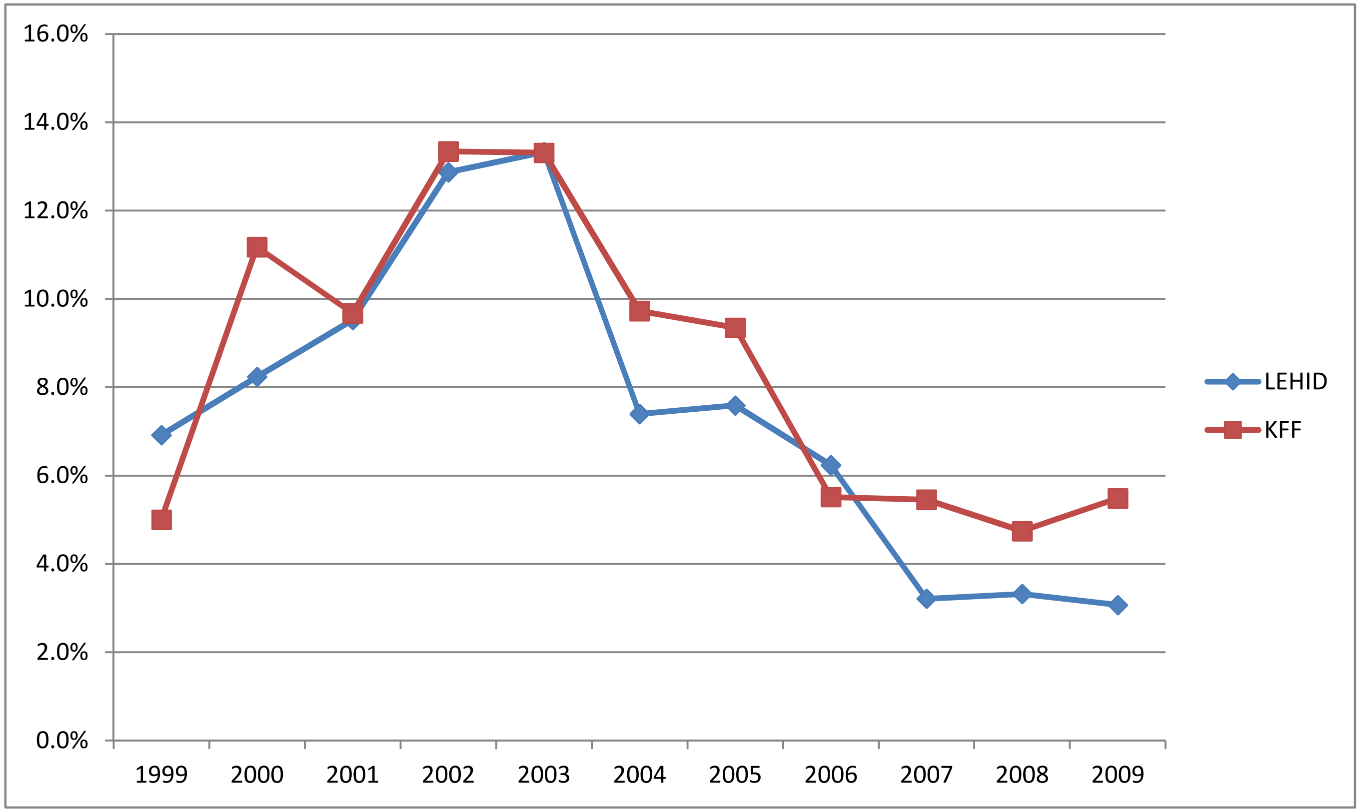
[3] For example, the entire state of Maine, is a single geographic market in the LEHID data.
[4] We were able to find a match for 284 out of a total of 328 MSAs present in the Interstudy dataset
[5] Note that the InterStudy estimates include only fully-insured plans, while the LEHID estimates include both fully-insured and self-insured plans. If we construct LEHID HHIs using only fully-insured plans, the corresponding correlation coefficients are .27 and .32 respectively.

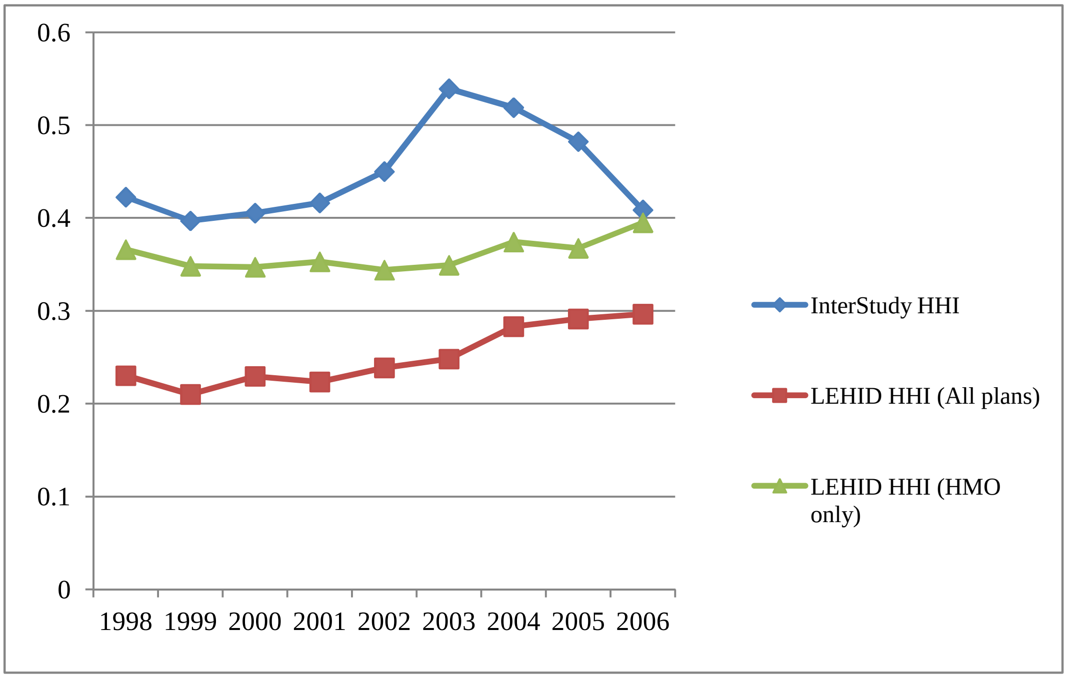
sampled enrollees to total insured lives (available at the county-level from the US Census of 2000) varies little across geographic markets. This provides further evidence that the sample accurately captures the typical healthplans offered by large employers in the U.S.

Appendix Figure 1. Annual Premium Growth, LEHID vs. KFF/HRET



Sources: LEHID Sample (all plans), and 2010 Kaiser/HRET Annual Survey of Employer-Sponsored Health Benefits. Both sources combine fully-insured and self-insured plans

Appendix Figure 2. Comparison of Trends in LEHID vs. Interstudy HHI



Sources : LEHID sample (all plans), InterStudy database

Appendix Table 1. Descriptive Statistics for other Dependent Variables

	Levels				*Long Differences*		
	2001	**2003**	**2004**	**2005**	**2001-2003**	**2001-2004**	**2001-2005**
Unit of Obs: State							
MLR	0.877	0.843	0.845	0.844	-0.034	-0.033	-0.033
	0.026	*0.034*	*0.031*	*0.030*	*0.034*	*0.035*	*0.032*
MLR (BCBS Plans)	0.863	0.832	0.843	0.836	-0.027	-0.018	-0.017
	0.056	*0.054*	*0.042*	*0.051*	*0.055*	*0.059*	*0.052*
MLR (Non-BCBS Plans)	0.880	0.840	0.836	0.835	-0.039	-0.043	-0.044
	0.033	*0.038*	*0.041*	*0.032*	*0.039*	*0.046*	*0.044*
Uninsured Rate	0.141	0.158	0.154	0.157	0.016	0.013	0.016
	0.041	*0.042*	*0.038*	*0.04*	*0.014*	*0.019*	*0.015*
Unit of Obs: Occupation-Market							
Ln (Healthcare Employment)	5.816	5.831	5.823	5.815	0.043	0.077	0.157
	1.38	*1.38*	*1.38*	*1.39*	*0.391*	*0.461*	*0.513*

Notes *: Number of observations varies by year and variable. The ranges are: MLR level [50], MLR (BCBS) level [39, 48], MLR (Non-BCBS) level [46, 49], Uninsured Rate level [51], Ln (Healthcare Employment) level [3266, 3957], MLR long differences [50], MLR (BCBS) long differences [38, 39], MLR (Non-BCBS) long differences [46], Uninsured Rate long difference [51], Ln(Healthcare Employment) long differences [2815, 2968]*

Appendix Table 2. Impact of BCBS Conversions on FP penetration: First Stage Estimates

Panel A. Fully Insured Plans

	Dependent Var = Lagged FP penetration		
Lagged BCBS FP	0.161		0.144
	(.058)**		*(0.155)*
Lagged BCBS FP * pre-conversion share		0.800	0.088
		*(0.264)****	*(0.632)*
R-Squared	0.13	0.13	0.13
Number of Obs	1514	1514	1514

Panel B. Self Insured Plans

	Dependent Var = Lagged FP penetration		
Lagged BCBS FP	0.315		-0.01
	*(.034)****		*(0.058)*
Lagged BCBS FP * pre-conversion share		1.624	1.672
		*(0.147)****	*(0.341)****
R-Squared	0.38	0.41	0.41
Number of Obs	1521	1521	1521

Notes *: The unit of observation is the market-year. All specifications include market-year controls and fixed effects for each market and year, and are estimated by WLS using the average number of enrollees in each market as weights. Standard errors are clustered by market.*

** denotes p<0.10, ** denotes p<0.05, * denotes p<.01*