Northwestern LAW

Searle Center on Law, Regulation, and Economic Growth

"Are Health Insurance Markets Competitive?"

Leemore S. Dafny
Assistant Professor of Management & Strategy
Kellogg School of Management
Northwestern University

LAW & ECONOMICS COLLOQUIUM
Monday, February 4th, 2008
4 p.m.
Schachtman-Gordon Hall
(Rubloff 339)

Are Health Insurance Markets Competitive?

Leemore Dafny
Northwestern University and NBER

January 2008

Abstract

The private health insurance industry provides benefits directly (through employer-sponsored or individual plans) or indirectly (through government-sponsored insurance programs) to 80 percent of the nonelderly and 72 percent of the elderly, yet researchers focus almost exclusively on public provision. Data on the private insurance sector is extremely difficult to obtain because health insurance contracts are complex, renegotiated annually, and not subject to reporting requirements. This study makes use of a privately-gathered national database of insurance contracts agreed upon by a sample of large, multisite employers between 1998 and 2005. To gauge the competitiveness of the industry, I investigate whether health insurers have sufficient market power to negotiate higher premiums for buyers with a greater willingness-to-pay. Using employer profits as a proxy for willingness-to-pay, I find they do: profitable employers pay higher premiums, ceteris paribus, and the results are robust to specifications that rely only on shocks to firm profits over time and thus use no cross-firm variation. Moreover, this relationship is only present in markets with a small number of insurance carriers, i.e. 8 or fewer major carriers. Given the rapid industry consolidation during the study period, these findings suggest healthcare insurers possess and exercise market power in an increasing number of geographic markets.

---

e-mail: l-dafny@kellogg.northwestern.edu. I am grateful for helpful suggestions by William Collins, James Dana, David Dranove, Mark Duggan, Vivian Ho, Katherine Ho, Ilyana Kuziemko, Mara Lederman, Niko Matouschek, Sean Nicholson, Mark Pauly, Allen Steinberg, Scott Stern, Robert Town, Dale Yamamoto, and seminar participants at Columbia, Duke, University of Maryland, Wharton, and the NBER Summer Institute. Michael Loquercio and Subramaniam Ramanarayanan provided excellent research assistance.

1. Introduction

The vast majority of Americans purchase health insurance through the private sector. Moreover, in recent years the public sector has increasingly turned to private insurers to deliver some or all of their commitments to enrollees. In spite of the enormous sums of public and private funds entrusted to these insurance carriers, there is little systematic research about them, let alone their market conduct. The main culprit is the lack of quality data about insurance contracts, which are tailored to individual customers, renegotiated annually, and not subject to public reporting requirements. This study makes use of a privately-gathered national database of insurance contracts agreed upon by a sample of large employers between 1998 and 2005. I use this database to explore whether insurance carriers successfully charge higher premiums to more profitable firms, controlling for differences in the healthplans purchased by different firms. More profitable firms may have a greater willingness-to-pay (WTP) for health insurance for several reasons, including higher switching costs, higher opportunity costs associated with search, and inefficient procurement mechanisms. The extent to which carriers are able to extract employer-specific rents offers a glimpse into competitive interactions in this important sector, as a competitive industry would be characterized by uniform pricing at cost.[1]

Most empirical research on market conduct relies on estimates of price-cost margins and/or assumptions about profit functions. These approaches are extremely difficult to pursue in the case of health insurance, where costs are realized months or years after the service is purchased, and include risk premia that vary across buyers and suppliers. The setting is also inhospitable to models that assume static profit maximization and infer conduct from entry patterns, as the industry suffers from a well-known "underwriting cycle" that involves alternating periods of low and high profits (Herzlinger 2006). I therefore rely

[1]To date, few have raised concerns about competition in the private insurance industry, with the notable exception of the American Medical Association (a trade association for physicians). Insurers have enjoyed *carte blanche* with respect to mergers and acquisitions (see footnote 8 for details). The growth in outsourcing of public insurance to the private sector also suggests that competitiveness has not emerged as a key concern. Thus, any concrete evidence of uncompetitive markets would appear to be new information to many interested parties.

on the institutional details of the health insurance industry to devise an empirical test of market conduct.

I focus on fully-insured healthplans offered by large employers.[2] Premiums for these plans depend on the actuarial health risk of employees, details of plan design (e.g., copays, covered benefits, disease management programs), and general carrier characteristics (e.g., provider network, speed and accuracy of claims processing, reputation). The net result is that pricing of fully-insured healthplans is anything but transparent, rendering the setting ripe for differential pricing across employer groups, *ceteris paribus*. (Because prices are ultimately the result of bargaining between employers and carriers, I do not use the term "price discrimination," which is typically reserved for settings in which the seller has commitment power (i.e., makes a single "take it or leave it" offer). However, the practice I investigate is akin to first-degree price discrimination, in which the seller sets individual prices to extract the surplus of each buyer.)

My approach relies on the (testable) assumption that employers are willing to pay more for health insurance when profits are high. If true, insurance carriers may exploit this fact by adjusting premiums accordingly. Their success in doing so should depend on the competitive environment they face. To test this hypothesis, I match my data on insurance contracts to data on operating profits, which is available for the publicly-traded firms in my sample. I investigate whether firms with higher profits subsequently pay higher health insurance premiums, controlling as best as possible for differences in the plans selected, employee populations, and market conditions. In the most stringent specifications, I exploit the panel nature of the data to fully eliminate any cross-sectional identification of this relationship. In these specifications, the relationship between premiums and profits is based solely on *changes* in each; that is, I investigate whether firms experiencing positive profit shocks subsequently face larger increases in premiums *for the same healthplans*. I also examine how employer-specific rent extraction varies with the number of insurance carriers present in each local market.

[2] The private health insurance market can be subdivided into five broad categories: individual insurance, group insurance – fully-insured, group insurance – self-insured, Medicare-financed plans, and Medicaid-financed plans.

I find robust evidence that insurance carriers charge higher prices to more profitable firms. Moreover, this effect is only present in markets with few insurance carriers. Thus, a multisite firm with high profits in a given year will face higher premiums for its healthplans, but *only* at the sites served by a concentrated insurance market. This result contradicts the leading alternative explanation for my finding, namely that firms with high profits face higher premium increases because they increase benefits in dimensions I do not observe. I also perform additional tests to study the plausibility of this alternative explanation, and to explore why profitable firms have a higher WTP for insurance.

The point estimates suggest an employer experiencing a 10-percentage-point increase in profit margins will subsequently face an increase in health insurance premiums of approximately 1.6 percentage points, but only in highly-concentrated markets (defined as markets served by 6 or fewer major firms). This estimated premium increase is a sizeable share of margins on a health insurance product, which are typically less than 5 percent of premiums.[3] As of 2005, 23 percent of employees in my sample received coverage in markets with 6 or fewer carriers, up from 7 percent in 1998. Due to recent consolidations, this figure is greater today.

Taken together, these results imply insurance carriers possess and exercise market power in an increasing number of geographic markets. This development may render ineffective recent efforts to contain public insurance costs by outsourcing healthplans to the private sector, and fail to deliver desired reductions in cost growth for all enrollees.

The paper proceeds in six sections. Section 2 provides background on the private health insurance industry. Section 3 describes the data in detail. Section 4 presents the main empirical analyses, followed by robustness tests and extensions. Section 5 explores why profitable firms are willing to pay more for health insurance. Section 6 concludes.

[3] Citing research by Sanford Bernstein, an investment research firm, *The Economist* reported that 2003 operating margins were 5.1 percent, "possibly an all-time high" as of the time of reporting (6/12/2004, p. 71). Insurers derive a sizeable share of total profits (which exceed operating margins) via the float: they earn interest on premium dollars before they are paid to satisfy claims.

2. The Private Health Insurance Industry, 1998-2005

Figure 1 graphs the percentage of nonelderly Americans covered by private insurance from 1998-2005, separated by whether the coverage was employment-sponsored or individually-purchased. Coverage from both sources declined slightly during the study period, but remained high, with 70 percent of the nonelderly obtaining insurance through the private sector in 2005. These figures understate the fraction of the nonelderly enrolled in private plans, as the majority of Medicaid beneficiaries are also enrolled in such plans (61 percent in 2005). Among the elderly, 95 percent are enrolled in Medicare, and nearly 13 percent of these received their care in 2005 through a private-sector Medicare Advantage plan. An *additional* 59 percent of the elderly had private supplementary coverage in 2005.[4]

Some information about private healthplan characteristics and premiums is available from the annual Employer Health Benefits survey, sponsored jointly by the Kaiser Family Foundation (KFF) and the Health Research and Educational Trust (HRET).[5] This survey documents two key trends that are corroborated in my data. The first is the rapid increase in health insurance premiums. Figure 2 illustrates these increases for 1998-2005, based on figures for a family of four. Annual growth peaked at 13.9 percent in 2003, declining to a still-impressive 9.3 percent in 2005. These figures likely understate the trend as employers have adjusted to rising costs by reducing the generosity of benefits provided.

The second trend is the growth in the share of employees covered by self-insured rather than fully-insured plans (Figure 3). Many large employers choose to self-insure,

[4] Source: http://www.ebri.org/pdf/publications/books/databook/DB.Chapter%2036.pdf. These figures do not reflect Medicare Part D, the prescription drug program introduced in 2006. Medicare Part D is administered entirely by the private sector and currently covers over 90 percent of Medicare beneficiaries. Many providers are pharmacy benefit management firms rather than health insurance carriers.

[5] The KFF/HRET survey randomly selects public and private employers to obtain national data about 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. (KFF/HRET *Employer Health Benefits 2006 Summary of Findings*, document 7528). Since 1996, the Agency for Healthcare Research and Quality (AHRQ), a division of the Department of Health and Human Services, has also conducted an annual survey of employers in conjunction with the Medical Expenditure Panel Survey (MEPS). MEPS follows households over time, and the "Insurance Component" surveys employers of household members to gather data on healthplans. The micro data are available on-site at Census Research Centers to those with appropriate clearance, but they do not constitute an employer-plan-level panel. The most recent data available is for 2003.

outsourcing benefits management and/or claims administration but paying realized costs of care. Such employers can spread risk across large pools of enrollees, and often purchase stop-loss insurance to limit their exposure. Per ERISA (the Employee Retirement Act of 1974), these plans are also exempt from state regulations. Figure 3 shows that self-insurance rates between 1998 and 2004 increased from 65 to 80 percent among employees in large firms, and 50 to 54 percent among all employees. According to Figure 2, premiums for fully-insured plans grew even more quickly than average during this period. My study sample includes only fully-insured plans, as the contractual terms between insurance carriers and self-insured plans are not available to me.[6]

The rapid increase in private insurance premiums has coincided with consolidation among insurance carriers. A 2004 Goldman Sachs report lists 22 major acquisitions between March 1995 and September 2004, and consolidation activity has continued apace. Only two combinations have been challenged by the Department of Justice, and these only in select markets.[7] There is also evidence that concentration in local markets is relatively high and increasing. Robinson (2004a) uses a database of state regulatory filings to study state-level market structure over 2000-2003. By the end of his study period, nearly 40 states had a dominant carrier serving over one-third of the private market. Robinson also documents increases in premium revenues and operating margins. Of course, a causal link between concentration and premiums cannot be established through the coincidence of these trends.[8]

[6] The rise in self-insurance, though beyond the scope of this paper, is an interesting subject for further research. Early work by Cooper and Simon (2007) reveals that firms are more likely to self-insure if they have multiple locations, a large number of workers, and high average wages.

[7] Both challenges were satisfied through consent decrees requiring divestiture in the markets with substantial overlap. (Tucson and Boulder in the case of the UnitedHealth-PacifiCare merger in 2005, and Houston and Dallas in the case of the Aetna-Prudential merger in 1999.) *Complaint, United States v. Aetna Inc.,* N0.3-99CV 1398-H, par. 19 and 20 (N.D. Tex. June 21, 1999); *Final Consent Order, United States v. Aetna Inc.,* No.3-99CV 1398-H (N.D. Tex. Dec. 7, 1999); *Complaint, United States* v. *UnitedHealth Group Incorporated & PacifiCare Health Systems, Inc.,* No. 1:05CV02436 (Dec. 19, 2005); *Final Judgment, United States* v. *UnitedHealth Group Incorporated & PacifiCare Health Systems, Inc.*, No. 1:05CV02436 (May 23, 2006).

[8] Dafny, Duggan, and Ramanarayanan (2007) attempt to make this link using an instrumental variables strategy and the same dataset employed in this study. The results of this study are not yet complete.

3. Data

3.1 The LEHID Data

The primary dataset was provided on a confidential, limited-use basis by a major benefits consulting firm.[9] The unit of observation is the plan-year. A plan is defined as a unique combination of an employer, geographic market, insurance carrier, and plan "type" (HMO, POS, PPO, and indemnity), e.g., Worldwide Widgets' CIGNA HMO in Phoenix, Arizona. The panel covers 1998-2005 (inclusive), and is unbalanced, with employers entering and exiting based on their relationship with the consulting firm, and specific healthplans appearing or disappearing when added or terminated, respectively. Note that participation is complete for any year in which an employer is included in the sample (i.e., all plans offered by that employer are present).[10]

The full dataset includes observations from 776 employers and 139 (mutually exclusive, collectively exhaustive) geographic markets in the United States. The employers span a wide range of industries. The top 3 are manufacturing and financial institutions (tied for 13 percent of employers each), and consumer products (9 percent of employers). On average, 241 employers appear in the sample each year. The median employer operates in 47 geographic markets and insures 9,670 active employees (retirees are not included in the data). The total number of employees represented in the sample averages 4.8 million per year. This figure does not include dependents, so the number of insured individuals represented by the survey is at least twice as large.

[9]Employers of all sizes rely on external consultants when designing or purchasing benefits. Using a 1997 survey of 21,545 private employers, Marquis and Long (2000) find external consultants were employed by nearly half of the smallest firms (<25 workers), and nearly two-thirds of the largest firms (>500 workers). These findings suggest the firms engaging the services of my source are not unusual in this regard, strengthening the case for the generalizability of the results.

[10] Some data scrubbing was necessary to ensure that the same ID was assigned to the same employer in every year. In the case of mergers, I create a new employer ID if both parties to the merger appear in the data separately in a prior year.

The geographic markets are defined by my source, and they represent the markets used by carriers and employers when negotiating rates.[11] The markets are sometimes defined by state boundaries (e.g., Delaware), but more commonly by metropolitan areas (e.g., Kansas City (in Missouri and Kansas); Kentucky – Louisville, Lexington; Kentucky – except Louisville, Lexington).[12] Figure 4 depicts the distribution of covered employees across the geographic markets. This distribution matches closely the distribution of privately-insured employees in these markets (estimated roughly using county-level data from the Current Population Survey of March 2000). I will refer to the entire dataset by the acronym LEHID, for "Large Employer Health Insurance Dataset."

3.2 Study Sample

The study sample is limited to fully-insured plans, whose premiums are determined prior to the start of the calendar year.[13] The movement toward self-insured plans, highlighted in the KFF-HRET survey, is also apparent in LEHID: the proportion of employees enrolled in self-insured plans increased from 58 to 76 percent between 1998 and 2005. However, the total number of employees in fully-insured plans is still sizeable, averaging 1.6 million per year.[14]

I restrict the study sample to observations in geographic markets containing 20 or more distinct employers; that is, 20+ employers must offer a fully-insured (FI) choice in that market-year. This restriction is imposed to ensure accurate estimates of market structure, namely the total number of carriers serving a given geographic market. Only 3 percent of the fully-insured employees in my sample are dropped as a result of this restriction. I use this "LEHID FI sample" to calculate market structure measures. Note this is not the sample for

[11] Some carriers with a national presence will negotiate a single rate for all employees nationwide. See footnote 19 for a discussion of the implications of national pricing for the empirical analysis.
[12] The markets correspond to groupings of 3-digit zipcodes.
[13] Self-insured plans report "premium equivalents," their *predictions* of costs per enrollee. These figures combine claim reimbursement, fees paid to carriers, and premiums for any stop-loss insurance.
[14] There is a good deal of overlap in the carriers serving self-insured and fully-insured groups. Among carriers serving more than 5 clients in LEHID, 57 percent have both fully-insured and self-insured plans, 41 percent have only fully-insured plans, and 2 percent have only self-insured plans. The smaller carriers (<=5 clients) are more likely to be "pure plays," with 11, 54, and 35 percent in these categories, respectively. Figures are tabulated using the carrier-year as the unit of observation.

the regression analysis; that sample is limited to plans for which profits of the associated employer can be obtained, and is described below.

Figure 2 shows that premium growth in the LEHID-FI sample tracks the levels and trends published by KFF/HRET for FI plans fairly closely. This bodes well for the generalizability of the data and results. Figure 5 graphs the distribution of markets in the LEHID-FI sample by the number of carriers in the market. Data are presented separately for 1998, 2001, and 2005. The fraction of markets with fewer than 6 carriers increased from 10 to 35 percent over this period, while the fraction with more than 10 carriers decreased from 35 to 7 percent. The increase in concentration is also manifested in other measures such as the HHI and the 4-firm concentration ratio. However, these measures are more prone to measurement error due to the size and non-random nature of the sample.

The key independent variable for the analysis is operating profits (lagged to reflect the timeline for plan selection, discussed below). To obtain profit data, I created a crosswalk file to match LEHID FI employers to companies appearing in *Compustat*, a database of financial statistics. The matches were identified by hand using company names, industry, locations, and number of employees. Extensive web research was required to verify matches for some observations, especially in cases of subsidiaries, non-U.S. firms, and firms involved in mergers and acquisitions. Profit is measured by the after-tax return on assets, defined as (earnings before extraordinary items + interest expense) / (gross assets + depreciation/amortization). Because *Compustat* is limited to large, publicly-traded firms, the LEHID FI-Compustat sample omits public-sector, nonprofit, and privately-held employers, as well as employers that do not appear in Compustat or lack data for the variables used to calculate operating profits. Of the 1678 employer-years in the LEHID FI sample, I am able to calculate lagged profit for 1151, or 69 percent of observations.

Table 1 presents descriptive statistics for the LEHID FI-Compustat sample in each year. The key variables include annual premium, enrollment, demographic factor, plan

design, plan type, and lagged profit.[15] Annual premium combines employer and employee contributions, and is a per-employee average. It reflects both the features of the plan selected (e.g., insurance carrier, benefit design, etc.) as well as the characteristics of the insured population (e.g., demographics and history of claims).

Demographic factor is a summary measure that reflects family size, gender, and age. Plan design captures the generosity of benefits, including the level of copayments required of enrollees. The exact formulae used to calculate these factors are not available to me. However, it is worth noting that the benefits consultancy that provided the data uses these factors to normalize premiums across plans and firms, and they are leading experts in healthplan selection and design. The decline in plan design during the study period is also noteworthy, as it is consistent with reports (from KFF-HRET as well as the popular press) that employers have reduced benefits (so-called "benefit buybacks") in an effort to contain cost growth.

Four plan types are represented in the data. Ordered by the restrictiveness of the provider network for each plan, these are: Indemnity (all providers covered), PPO (preferred providers fully covered, non-preferred providers covered in part), POS ("point of service" plan: care is "managed" as in an HMO, and if approval for a service is obtained preferred providers are covered in full and non-preferred providers in part), and HMO (care is managed and preferred providers are fully covered). Approximately 90 percent of the plans in the LEHID FI-Compustat sample are HMOs.

Because profit is lagged two years in all specifications, the 2001 recession is apparent in 2003. The recession had varying impacts across firms and sectors, as evidenced by the large increase in the standard deviation in 2003 and 2004. This is precisely the type of variation that identifies the effect of interest.

[15] Premiums are reported in nominal dollars. All specifications use ln(premium) as the dependent variable and include year fixed effects, so nationwide deflators will not affect the coefficients of interest.

The LEHID FI-Compustat sample includes an average of 144 employers and 102 markets per year. The decline in observations during the last two years reflects both the trend away from FI plans, and a general decline in the number of employers in the LEHID sample.[16] These trends are apparent in Appendix Table 1, which gives the number of employers included in LEHID in every year, together with the share with at least 1 FI plan and at least 1 SI plan.

4. **Empirical Analysis: Premiums and Profits**

4.1 Main Analysis

The key regression relates plan premiums to lagged employer profits. Recall a *plan* is an employer-geographic market- insurance carrier –plan type combination, denoted by the subscript *emcj*. The first specification can be expressed as follows:

$$(1) \quad ln(premium)_{emcjt} = \alpha + \gamma_1 profit\ margin_{e,t-2} + \gamma_2 demographics_{emcjt} + \xi_e + \nu_m + \psi_c + \eta_j + \delta_t + \omega_{jt} + \mu_{cj} \left[+\varsigma_{em}\right] \left[+\rho_{emcj}\right] \left[+\phi_{mt}\right] + \varepsilon_{emcjt}$$

Profit margin is lagged two years to reflect the timeline for negotiating insurance contracts. These contracts are signed a few months prior to the start of the benefit year, which is generally the calendar year. Thus, an employer will typically begin selecting 2002 plans and rates by early 2001. To the extent that firm profits affect these agreements, the relevant profit figure will reflect data for 2000 (assuming data is available annually).[17]

Equation (1) includes fixed effects for each employer *e*, market *m*, carrier *c*, plan type *j*, and year *t*. Employer fixed effects help to capture unobserved, time-invariant differences in the composition of the population covered, benefit design, and usage patterns, all of which affect plan premiums. Market fixed effects capture differences in medical costs (e.g., due to

[16]The decline is heightened by the sample restriction that drops all observations in markets with fewer than 20 employers offering at least 1 fully-insured plan in that market. In section 6, I confirm the results are similar when this restriction is lifted.

[17] Ginsburg et al. (2006) find evidence of a similar lag (18 months) between premiums reported by KFF/HRET and the cost of healthcare services (e.g., provider charges).

local wages) and practice. There is a literature that documents substantial differences in medical practice and utilization (though, interestingly, not in outcomes) across geographic markets (e.g., Wennberg, Fisher, and Skinner (2002)). Plan type fixed effects capture average price differences for these broad product groups, and carrier fixed effects capture average price differences across carriers due to time-invariant characteristics such as reputation. Finally, year fixed effects absorb annual growth in premiums nationwide.

In addition to this set of first-order fixed effects, all specifications include two second-order fixed effects (or interactions): plan type-year effects (ω_{jt}) and plan type-carrier effects (μ_{cj}). The plan type-year effects absorb different premium growth patterns across plan types. The plan type-carrier effects capture time-invariant differences in premiums associated with differences in plan type-specific quality for each carrier. For example, Blue Cross/Blue Shield PPO premiums might be high, *ceteris paribus*, because Blue Cross/Blue Shield PPO provider networks are typically very inclusive. (Note that technically the first-order year, plan type, and carrier fixed effects drop out of the equation because they are subsumed in these second-order interactions.)

Adding the bracketed terms to the model reduces the possibility of omitted variables bias and increases the precision of the coefficient estimates, but also eliminates variation that can be used to identify γ_1. In recognition of this tradeoff, I present results for specifications with and without these terms. I begin by adding employer-market interactions (ς_{em}). Employees of a given firm may differ across markets (e.g., the headquarters site might employ different types of workers than other sites), and these differences could be reflected in premiums. Next, I introduce fixed effects for each employer-market-plan type-carrier combination (ρ_{emcj}). Once these "plan fixed effects" are included, γ_1 is identified by *within-plan* changes in premiums and changes in the profits of affiliated employers.

The key advantage of including plan fixed effects is the elimination of cross-sectional variation as a source of identification for γ_1. For example, if profitable firms tend to choose generous benefit packages (within a given plan type – carrier combination), in the absence of

plan fixed effects γ_1 will reflect these omitted factors. However, some of the cross-sectional variation is desirable because competing plans act as a constraint on the pricing of incumbent providers. Faced with an excessive price increase, some employers may switch plans to obtain better pricing. Plan fixed effects eliminate any variation in price due to such switching. The bias induced as a result could be positive (if those remaining with incumbent carriers tend to be price-inelastic) or negative (if those switching are precisely the employers offered the highest prices by incumbents).

Last, I add market-year fixed effects. Once included, γ_1 is identified solely by differences in within-plan premium growth for plans operating in the same market. To clarify this source of identification, consider as a hypothetical example the Chicago-based healthplans offered by Boeing and United Airlines in 2003. In the wake of September 11, 2001, United filed for bankruptcy while Boeing's fortunes soared. Controlling for the average premium growth in Chicago, as well as the average premium growth for specific plan types nationwide, I expect premium increases to be higher for Boeing if γ_1 is positive.

Table 2 presents results for the 4 specifications represented by equation (1). Standard errors are clustered at the plan level to adjust for possible serial correlation in the error terms. The estimates of γ_1 are positive, similar in magnitude, and statistically significant at $p<.05$ for the first three specifications (columns 1, 3 and 5). The point estimate is a bit larger in column 5, which includes plan fixed effects. Although the differences between column 5 and 1 or 3 are not statistically significant, the sign suggests that within-plan price hikes may be larger than realized price increases when switching is incorporated into the sample. (This finding foreshadows the analysis in section 6, which finds that more profitable employers are less likely to switch carriers.) The results from the first three specifications imply an employer with a 10-percentage point increase in profits can expect to pay approximately 0.3 percent more in health insurance premiums, *ceteris paribus.*

Controlling for differences in premium growth across markets (column 7) reduces the magnitude of $\hat{\gamma}_1$ to a marginally-significant 0.2 percent ($p<.10$). While mitigating the likelihood of omitted variables bias, the market-year effects eliminate plausibly exogenous

sources of variation in lagged profits. As an alternative, I replace the market-year effects with market-year-level measures of economic conditions (the unemployment rate) and costs (the per-enrollee average Medicare expenditure in the market-year). The (unreported) results are virtually identical to those obtained without market-year fixed effects, suggesting that local economic conditions and/or changes in healthcare utilization are unlikely to be driving the results.[18]

Next, I consider the possibility of bias due to changes in unobserved plan characteristics, such as provider networks, prescription drug formularies, and copayments. An alternative explanation for the positive estimate of γ_1 is that firms with positive profit shocks respond by increasing within-plan benefits for workers, and more benefits come with a higher price tag. As a first test of this hypothesis, I add *plan design* to each specification. The results are reported in columns 2, 4, 6, and 8, alongside the corresponding baseline specifications. The coefficient on plan design is always positive and highly significant, suggesting it is an accurate measure of the generosity of benefits. However, the estimates of γ_1 are virtually unchanged. To the extent that other omitted, time-varying plan characteristics are correlated with this composite measure, this test provides some reassurance that these omitted factors are not generating the positive estimate of γ_1.

As a second test of this alternative explanation, I consider how the estimate of γ_1 varies by the market structure of the local insurance industry. If γ_1 reflects rent extraction by insurance carriers, it should be larger where competition is less fierce. If instead it reflects the predilection of profitable employers to provide more generous benefits, it should be insensitive to the market structure of the insurance industry: there is no obvious reason why multisite firms with high profits would increase benefits most in the sites served by a small number of carriers. I estimate models based on the following equation:

[18] Estimates are available upon request. Both measures are available at the county-year level, whereas LEHID markets are defined by 3-digit zipcodes. The correspondence is fairly close in major metropolitan areas, however, where most LEHID employers are located. The coefficient on the unemployment rate is small, statistically significant and positive. The average cost measure (known as the AAPCC, or Adjusted Average per Capita Cost for Medicare Parts A and B) did not enter significantly into the specifications.

$$(2)\quad ln(premium)_{emcjt} = \alpha + \sum_{NC=1}^{5} \gamma_{1,NC} 1(NC)_m * profit\ margin_{e,t-2} + \gamma_2 demographics_{emcjt} + \gamma_3 plan\ design_{emcjt} + \xi_e + \nu_m + \psi_c + \eta_j + \delta_t + \omega_{jt} + \mu_{cj} [+\varsigma_{em}] [+\rho_{emcj}] [+\phi_{mt}] + \varepsilon_{emcpt}$$

where NC stands for "number of carriers." I use 5 ranges for number of carriers: 1-4, 5-6, 7-8, 9-10, and 11+. $1(NC)_m$ is an indicator variable that takes a value of 1 if the observation is from a market with NC carriers. Thus, $\gamma_{1,3}$ is estimated from observations in markets with 7-8 carriers. Although an imperfect proxy for market competitiveness, the number of carriers is an appealing measure for theoretical and practical reasons. Theoretically, employers are more likely to switch among carriers to get a good deal when many are present as they are more likely to find a product that is similar to their current (or ideal) choice. From a practical perspective, because the data include only a sample of contracts in each market, estimates of the number of (major) carriers will be much more accurate than estimates of alternative measures of market concentration such as the HHI.

Table 3 illustrates that the positive coefficient estimates in Table 2 are driven entirely by markets with 8 or fewer carriers.[19] In general, the magnitude of $\hat{\gamma}_{1,NC}$ declines as the number of carriers increases. There is an especially steep decline when the number of carriers exceeds 6. As in prior specifications, the point estimates decrease and the standard errors increase when market-year fixed effects are added. However, the coefficient estimate corresponding to markets with 4 or fewer carriers remains large and significant at $p < .10$. Replacing the market-year fixed effects with the market-year covariates described above (i.e., the unemployment rate and average costs per Medicare enrollee) yields estimates that are virtually identical to those obtained without market-year effects. This suggests economic conditions and healthcare cost inflation are not driving the relative sensitivity of premiums to profits in markets with few carriers.

[19] Some national carriers negotiate a "single price" with multisite employers (i.e., a price that only varies by family size). I cannot identify these contracts in my dataset, but to the extent they are present any rent extraction by the carrier will be the same in all geographic markets. (The fact that the estimates of rent extraction in markets with more than 8 carriers are near zero and estimated with reasonable precision is consistent with two explanations: (1) single-price contracts are uncommon in the data; (2) rent-extraction does not occur in such contracts.)

These estimates show that profitable firms pay more for their health insurance. This phenomenon is limited to markets with 8 or fewer carriers, and is most pronounced in markets with 6 or fewer carriers. In such markets, a profit increase of 10 percentage points (roughly the standard deviation of profits during the 2001 recession) is associated with an increase in health insurance premiums of 1.6 percent.[20]

4.2 Extensions and Robustness

Before moving to the next key empirical question – *why* are more profitable firms willing to pay more for health insurance? – I perform some additional analyses to extend and corroborate the main findings. I begin by examining whether rent extraction is symmetric for positive and negative profit shocks. To answer this question, I reorganize the data into first-differences (where the unit of observation is a plan) and estimate the following specification:

$$(3)\quad \Delta ln(premium)_{emcj(t,t-1)} = \alpha + \vartheta_1 \Delta profit\ margin_{e(t-2,t-3)} + \left[\vartheta_2 \Delta profit\ margin_{e(t-2,t-3)} \cdot 1(\Delta profit\ margin_{e(t-2,t-3)} > 0)\right] + \vartheta_3 \Delta demographics_{emcj(t,t-1)} + \vartheta_4 \Delta plan\ design_{emcj(t,t-1)} + \delta_t + \omega_{jt} + [+\phi_{mt}] + \Delta\varepsilon_{emcj(t,t-1)}.$$

Absent the first term in brackets, this specification corresponds to specification (1) after *plan design* and plan fixed effects are included (i.e., Table 2, column 6).[21] The first bracketed term allows the extraction parameter to differ for employers experiencing positive profit shocks. Roughly half of the observations in the first-differences sample (which consists of plans with observations in adjacent years) have positive shocks. The second bracketed term represents market-year dummies, which are included in some specifications to allow different growth rates for each market and year. The results are given in Table 4. Without the interaction term (columns 1 and 2), the results are similar to those obtained using the level specification (Table 2, columns 6 and 8, respectively). Allowing the coefficient on

[20] I obtain this estimate using the average of the relevant coefficients in the specification with plan fixed effects (column 3, Table 3): exp(((.178+.142)/2)*.1) =1.016.

[21] For the main analyses, I use the fixed-effects estimator because it enables comparisons of estimates with and without plan fixed effects, and retains more data points given the unbalanced nature of the panel.

Δ*profit margin* to differ for positive and negative shocks results in noisy estimates of both parameters (columns 3 and 4). The coefficient on the interaction term is large and positive in both specifications, however, offering suggestive but inconclusive evidence of asymmetric rent extraction.

Next, I perform a falsification exercise using data on *self-insured* (SI) healthplans. Recall that the "premium" measure for these plans is actually the employer's estimate of outlays for each plan-year. These forecasts are used for budgeting purposes and to make decisions about benefit structure, insurance carriers, and stop-loss coverage. Because these projections include expected fees charged by the carriers who administer these plans, evidence of rent extraction may be apparent in the data. However, the relationship should be weaker than that observed in the fully-insured market, as error in the dependent variable is substantial (because projections are imperfect measures of realized costs), and fee structures for administrative services are far more transparent than pricing for fully-insured plans. (Although contract structure varies widely, contracts typically specify fees per enrollee and/or per claim.)

Using the sample of self-insured plans in the same market-years included in the analysis of fully-insured plans, I estimate specification (2) above (i.e. the levels specification with interactions between lagged profits and insurance market structure). The dependent variable is the employer's estimate of outlays for each plan-year. The indicators for number of carriers still refer to *FI* carriers. Using the FI market structure is the likeliest way to reveal whether the main results are spurious. In addition, the self-insured market is less concentrated, precluding identification of $\gamma_{1,NC}$ in markets with small numbers of *SI* carriers.

The results show a *negative* relationship between lagged profits and estimated outlays. Decomposing the effect by market structure reveals negative and occasionally significant coefficients on lagged profits in markets with 5-6 and 7-8 carriers offering full insurance. To the extent the costs of SI plans are an appropriate counterfactual for FI plans, these findings suggest the main results could underestimate the extent of rent extraction.

In addition to the falsification exercise, I considered several additional robustness checks.[22] First, I perform two checks to explore the possible bias induced by dropping market-years in which there are fewer than 20 employers with at least one FI plan. The first check expands the estimation sample to include all markets and years with at least one fully-insured plan. This introduces error in the dummies for number of carriers, but mitigates concern about changes in the sample of markets included in each year. The coefficient estimates change very little and the precision improves. The second check drops data from 2004 and 2005, the years in which the number of markets in the sample declines substantially. The coefficient estimates are again quite similar, with slightly larger standard errors. The key patterns remain statistically significant.

The next robustness issue concerns the product market definition. The main analysis sample includes observations for all plan types (i.e. HMO, POS, PPO, and Indemnity). Although there is no reason *a priori* to believe the pattern of interest will differ across plan types, and plans of different types are clearly substitutes, given that more than 90 percent of fully-insured plans are HMOs it seems prudent to confirm the results are not driven by a subsample of unrepresentative plans. Restricting the sample to HMOs yields very similar coefficient estimates, with only slight increases in the standard errors.

The final robustness check restricts the type of employers included in the analysis. To ensure identification of the market-structure*lagged profit interactions comes from employers with multiple locations, I restrict the estimation sample to employer-years with plans in 10 or more geographic markets. The coefficient estimates and standard errors are very similar to those obtained from the full sample.

5. Empirical Analysis: Why Does WTP Increase with Profits?

The premise that firms will pay more for an input when profits are high appears inconsistent with a simple model of static profit maximization: why would firms fail to minimize costs,

[22] In addition to the robustness checks presented here, I confirmed the results remain similar in the presence of plan type-carrier-year interactions, and indicators for the number of carrier ranges.

regardless of profit level? The business press is replete with anecdotal evidence of such behavior, however, and it has recently been corroborated in empirical work by Borenstein and Farrell (forthcoming). Borenstein and Farrell find stock market valuations of gold mining firms are concave in the price of gold. Given the perfectly competitive output market, this result is consistent with a decrease in cost-efficiency when profits are high.

A potential source of cost inefficiencies in "fat firms" is rent-sharing with workers, or to be more precise, rent-sharing that is not part of an optimal labor contract.[23] Borenstein and Farrell do not find evidence of rent-sharing in the mining sector, but it has been documented in a number of industries and countries (e.g., Dickens and Katz 1987, Katz and Summers 1989, and Blanchflower et al. 1996). Although empirical evidence of rent-sharing focuses on wages, the relationship with fringe benefits such as health insurance may be similar, as there is evidence that benefits and wages are interchangeable (e.g., Gruber 1994 and Pauly 1998). When presented with my findings, industry experts suggested precisely such an explanation.[24]

The argument proffered by the experts is linked to the high switching costs employees must incur when changing healthplans. These costs include: learning about changes in benefit coverage and design and potentially adjusting healthcare consumption as a result; identifying providers that belong to the new plan's network; if former providers are not part of the network, scheduling and undergoing new patient visits (which may lead to a disruption in care and associated medical consequences), transferring medical records and prescriptions, and, for plans with "gatekeepers" (such as HMOs or POS plans) navigating through a gatekeeper in order to see specialists; figuring out the claims reimbursement system. For employers to obtain the best pricing on plans, they must be willing to change carriers. However, a plan switch is a "tough sell" in good times, to paraphrase an executive from my data source. Workers are willing to tolerate such actions (along with, say, the

[23] Rent-sharing may of course be optimal, particularly if workers and firms are risk-averse (Blanchflower 1996) or if specific investments are required for both parties.

[24] This explanation was proposed by an executive from my data source, and subsequently corroborated in interviews with a large insurance broker/former insurance executive (phone interviews, 10/9/2007 and 10/10/2007), and a CFO at a leading health insurer (phone interview, 10/25/2007). All sources requested anonymity.

holiday party in the office conference room), but only when viewed as necessary. Thus profitable firms may share rents with workers by retaining existing plans. In uncompetitive insurance markets, carriers can exploit this stickiness through employer-specific pricing.

To test this hypothesis, I investigate whether switching is less likely when firms are profitable, controlling for other factors that may be associated with the propensity to switch. I create a dataset of employer-market-year observations and estimate linear probability models of the following form:

$$(4) \quad switch_{em(t,t-1)} = \alpha + \varphi profit\ margin_{t-2}\ [+\phi_{mt}]\ [+\xi_e]\ [+\varsigma_{em}]\ +\ \varepsilon_{emt}.$$

I define two versions of *switch*: *carrierswitch* and *planswitch*. *carrierswitch* takes a value of 1 if there is an addition or deletion of insurance carriers by an employer in a given market between *t-1* and *t*. *Planswitch* takes a value of 1 if there is an addition or deletion of carrier-plantypes. *Planswitch* will overstate switching, e.g., if a firm switches from a UnitedHealthcare HMO to a UnitedHealthcare POS, it will be coded as having made a switch when no material switch has occurred. *carrierswitch* will understate switching, e.g., if a firm offers an Aetna HMO, Aetna PPO, and UnitedHealthcare PPO, and eliminates the Aetna PPO, it will not be coded as having made a switch. For this reason, I present estimates using both measures. Note the switch variables are created using the entire LEHID sample, so that a firm that decides to self-insure a plan (or vice versa) is not coded as having made a switch.[25]

The baseline specification includes no controls; it simply captures the association between lagged profits and the propensity to switch. The next specification adds market-year interactions to control for general upheaval in a market-year due, for example, to mergers or exits of insurance carriers. Absent these interactions, the estimate of φ will reflect such activity if it is correlated with market-level changes in lagged profits of employers.

[25] To reduce measurement error, the switch variables are defined only when data from two *adjacent* years is available. Unfortunately, error due to mergers, acquisitions, and divestments of insurers and employers cannot be purged from these variables. Market-year fixed effects can, however, control for changes in the local insurance market.

Employer fixed effects are added next; these control for any employer-specific tendencies to switch, which may also be correlated with profit levels and hence bias the estimate of φ. For example, employers in sectors with high labor turnover may switch healthplans more often because their employees are less likely to have a continuous relationship with a healthplan and/or its associated providers. If such employers also tend to report lower profits, the estimate of φ could be biased downward in the absence of employer fixed effects. Last, I add employer-market fixed effects, which allow for different baseline switching levels across employer-markets. For example, employees of a large retail chain may differ across locations, with headquarters employees expecting steady benefits and retail clerks in all other markets willing to tolerate switches more readily.

I estimate the switching specifications on the entire sample of employer-market-year observations with Compustat data, and on the subset of observations with at least one fully-insured plan and located in markets with 20+ employers offering a fully-insured choice. There is no theoretical reason to restrict the switching analysis in this way; I present results using this subsample to maintain consistency with the rent extraction analysis. The descriptive statistics for the switching variables in both samples are given separately by year in Table 5; 1998 is omitted as the switching variables can only be defined for employer-markets with data in the preceding year. Just over one-third of the observations in the total sample have a carrier switch, and 45 percent have a plan switch. The figures are even higher in the fully-insured sample, with 47 percent of observations switching carriers and 56 percent switching plans. In both samples, there is a marked decline in switching over time. This reflects, at least in part, the declining number of options available.

The results of the switching analysis (Table 6) strongly support the hypothesis that more profitable firms are less likely to switch carriers or plans. The point estimates are slightly larger for *carrierswitch*, and given the lower mean levels of *carrierswitch* this translates into bigger proportionate effects. For example, a 10-percentage-point increase in profit margins in year *t* is associated with a reduction of roughly 4 percentage points in the propensity to switch carriers between *t+1* and *t+2*. Given the mean levels of *carrierswitch*, this corresponds to a decline of more than 10 percent. The *planswitch* models yield

somewhat smaller estimates, but the result is equally robust to alternative specifications as well as the different data samples.

The advantage of these linear probability models is the ability to control for a variety of fixed effects. However, a binary outcome measure does not permit a distinction among types of switching. Plan *deletion* should be much more responsive to profits than plan *addition*, which should not generally impose a cost on employees (apart from those responsible for benefits administration). To refine the analysis, I therefore estimate multinomial logit models using two multivalued outcome measures, *carrierchange* and *planchange*, which correspond directly to *carrierswitch* and *planswitch*. The outcomes are no change (the base outcome), add only, drop only, and add and drop. To the extent profitable firms avoid carrier and/or plan changes in general all three outcomes will be less likely for such firms. However, if the mechanism described above is correct, changes that include deletions should be the most sensitive to profits.

In these models, I include only lagged profits and year fixed effects as explanatory variables. Estimating coefficients for additional fixed effects is neither computationally feasible nor conceptually appropriate given the data. An employer fixed effect, for example, would control for the propensity for every employer to select every outcome, leaving only within-employer, within-outcome variation to identify the parameters for each choice. Given the switching results presented in Table 6 are generally insensitive to the inclusion of various fixed effects, the pattern of coefficients and predicted outcomes for a parsimonious multinomial model should yield accurate qualitative conclusions.

Table 7 reports the results in the form of risk ratios for each outcome relative to the base outcome of no change, using *carrierchange* in the top panel and *planchange* in the bottom panel. Both models are estimated on the combined SI and FI sample (results are similar with the FI sample, and available upon request.) All of the ratios are significantly less than 1, indicating that more profitable firms are less likely to make *any* changes to their carriers or plans. However, the relative risk ratio for *adding* is four to five times as great as that for *deleting* or *adding* and *deleting*, which are fairly close in size; this difference is also

statistically significant. To put the results into the context of outcome probabilities, at the bottom of each panel I present the realized probabilities together with the mean of predicted probabilities assuming a profit shock of .10 for each employer. Using the results from the *carrierchange* model, the probability of adding a plan declines from .090 to .085 (6%), the probability of deleting a plan declines from .105 to .083 (21%), and the probability of adding *and* deleting a plan declines from .160 to .133 (17%). These results suggest employers are especially reluctant to *drop* healthplans when profitable, a finding that buttresses the switching hypothesis for the link between WTP and profits.

6. Discussion and Conclusions

The U.S. healthcare system relies heavily on private insurance companies to manage healthcare consumption and funnel payments to providers. Expenditures by private insurers are projected to reach $776 billion in 2007, up from $295 billion in 1993.[26] This figure understates the size and growth of the industry, as public insurers have increasingly turned to private insurers to deliver benefits to enrollees. The appeal of the private sector stems from the assumption that competition among payers will improve quality, increase innovation, and lower costs.

This study is the first to examine market conduct in this important and growing industry. I do so by studying differences in prices for different purchasers of fully-insured healthplans. Specifically, I investigate whether employers with deeper pockets (as measured by publicly reported profits) face higher health insurance premiums, *ceteris paribus*. The ability of insurers to extract rents in this manner should be limited in the presence of robust competition.

I find profitable firms do pay more, and the result is the same when looking at *changes* in premiums for each plan in response to *changes* in employer profits. I also find

[26] Source: Centers for Medicare and Medicaid Services, Office of the Actuary, National Health Statistics Group. http://www.cms.hhs.gov/NationalHealthExpendData/downloads/proj2006.pdf. These figures do not include estimates of insurer profits and costs, i.e., they do not represent total premiums collected. Data on total premiums is not readily available.

this practice is limited to markets with a small number of insurance carriers. These results suggest healthcare insurers in concentrated markets successfully extract the rents of individual buyers, a practice akin to first-degree price-discrimination. The analysis does not lend itself to quantification of the market power enjoyed by insurers in these markets. However, the magnitude of the estimate offers some clues. In highly-concentrated markets (6 or fewer major carriers), a 10-percentage-point increase in the after-tax return on assets is followed by an increase of approximately 1.6-percent in health insurance premiums. Given operating margins for insurers are generally less than 5 percent, this is a large figure.

Importantly, the share of markets in my sample with 6 or fewer carriers has increased dramatically over time, from 10 percent in 1998 to 35 percent in 2005. (Because this number of carriers is based on a large, but incomplete, sample of employers, 6 should be viewed as an *under*estimate of the actual number of carriers in such markets.) Concentration has only increased since. Taken together, the evidence shows health insurers are exercising market power in an increasing number of geographic markets. One implication is that governments may want to question their reliance on competition among insurers as a method of cost-control unless evidence of this competition is apparent in the relevant marketplaces. Proposed consolidations may also warrant greater scrutiny by antitrust enforcers than has taken place to date. Finally, research on whether and how much uncompetitive markets are contributing to higher healthcare costs – and potentially more important – to slower innovation in healthcare management and delivery of care – would help to inform the public debate over healthcare reform.

References

Blanchflower, David G., Andrew J. Oswald, and Peter Sanfey (1996), "Wages, Profits, and Rent-Sharing," *Quarterly Journal of Economics*, Vol. 111(1): 227-251.

Borenstein, Severin and Joseph Farrell (forthcoming), "Do Investors Forecast Fat Firms? Evidence from the Gold Mining Industry", *RAND Journal of Economics.*

Cooper, Philip and Kosali I. Simon (2007), "Trends and Determinants of Self Insuring Employer Health Benefits, 1997-2004," Cornell University mimeo.

Dickens, W. T. and Lawrence F. Katz (1987), "Inter-Industry Wage Differences and Theories of Wage Determination," in K. Lang and J. Leonard, eds., *Unemployment and the Structure of Labor Markets* (Blackwell), 48-89.

Ginsburg, Paul B; Bardley C. Strunk, Michelle I. Banker, and John P. Cookson, "Tracking Health Care Costs: Continued Stability But at High Rates in 2005," *Health Affairs* web exclusive, 25, no. 6 (2006): w486-w495

Gruber, Jonathan (1994), "The Incidence of Mandated Maternity Benefits," *American Economic Review,* 84(3): 622-641.

Katz, Lawrence F., and Lawrence H. Summers. 1989. Industry rents and industrial policy. *Brookings papers on economic activity: Microeconomics*, 209–75.

Marquis, M. Susan and Stephen H. Long (2000), "Who Helps Employers Design Their Health Insurance Benefits?" Health Affairs, January/February 19(1): 133-138.

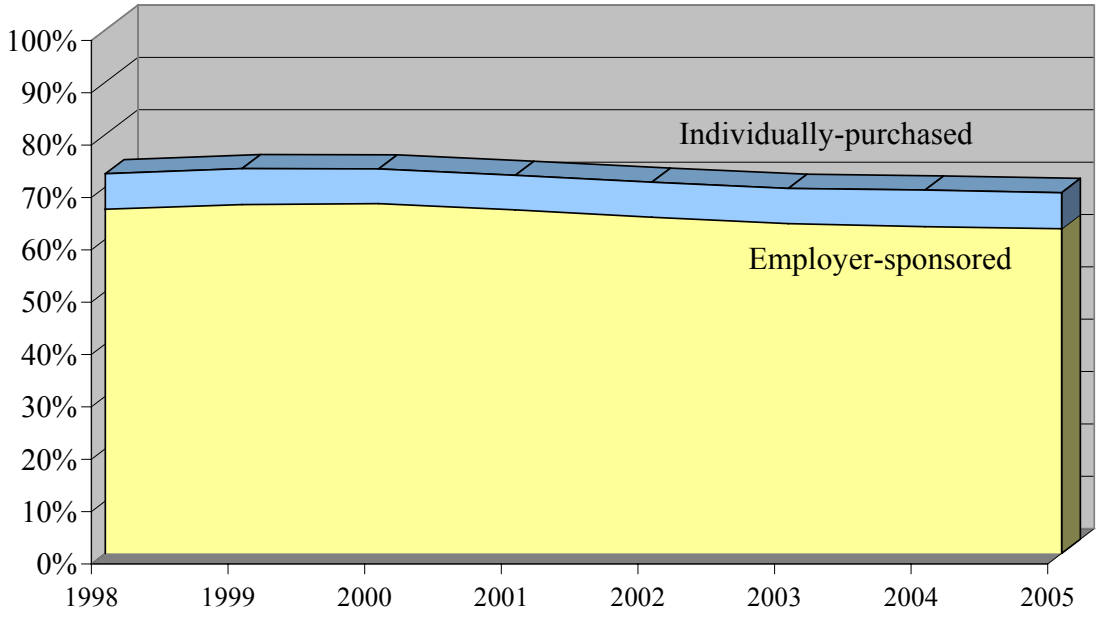
Pauly, Mark (1998), *Health Benefits at Work: An Economic and Political Analysis of Employment-based Health Insurance.* Ann Arbor: University of Michigan Press, 1997.

Robinson, Jamie (2004a), "Consolidation And The Transformation Of Competition In Health Insurance,"*Health Affairs*, November/December; 23(6): 11-24.

Robinson, Jamie (2004b), "For-Profit Non-Conversion and Regulatory Firestorm at CareFirst BlueCross BlueShield," *Health Affairs*, July/August, 23(4): 63-84.

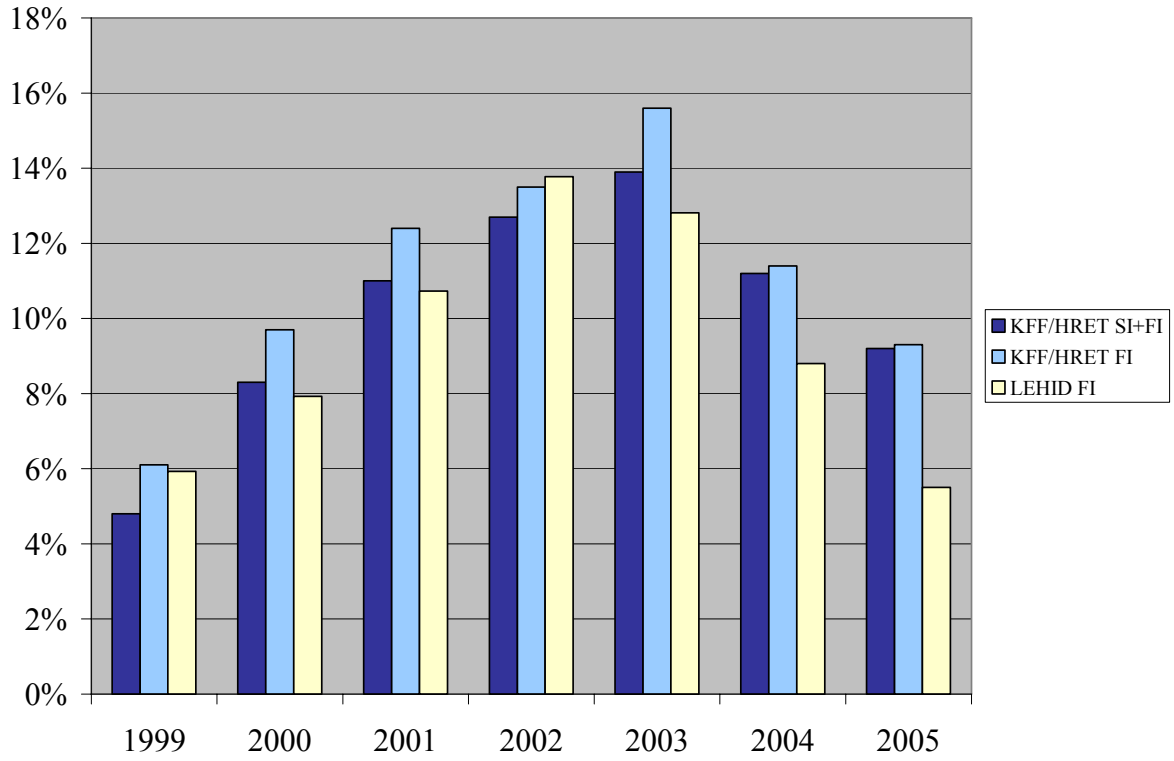
Wennberg, Fisher, and Skinner (2002), "Geography and the Debate Over Medicare Reform," *Health Affairs* 21(2) web exclusive: W96-W114.

Figure 1. Nonelderly Population with Private Insurance Coverage, 1998-2005



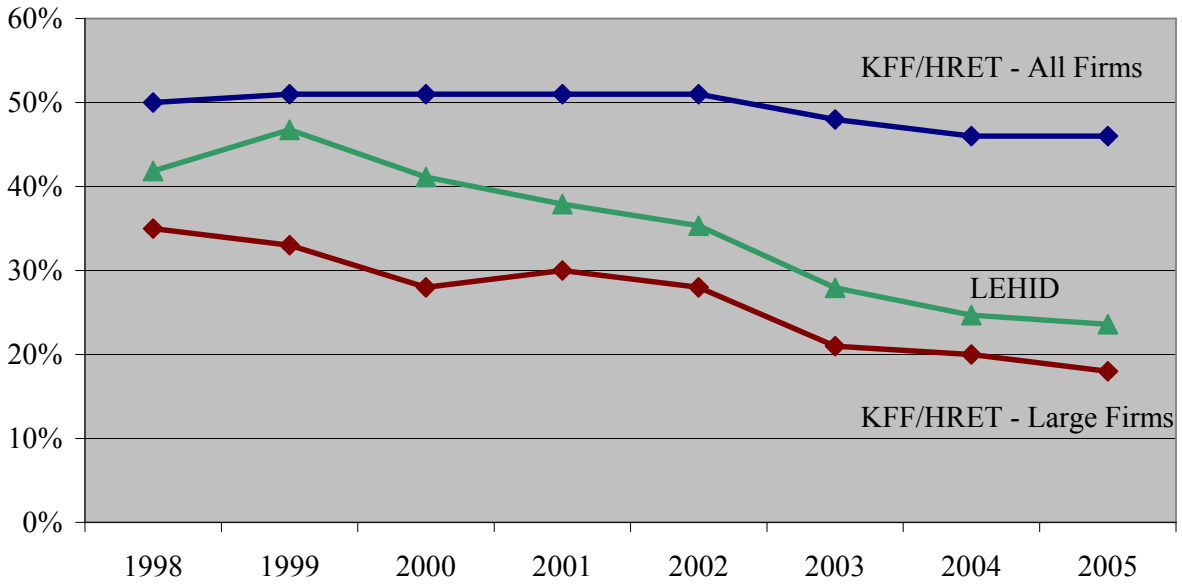
Source: Enployee Benefit Research Institute estimates using the Current Population Survey, March 1998-2006 Supplements.

Figure 2. Growth in Annual Health Insurance Premiums, 1999-2005



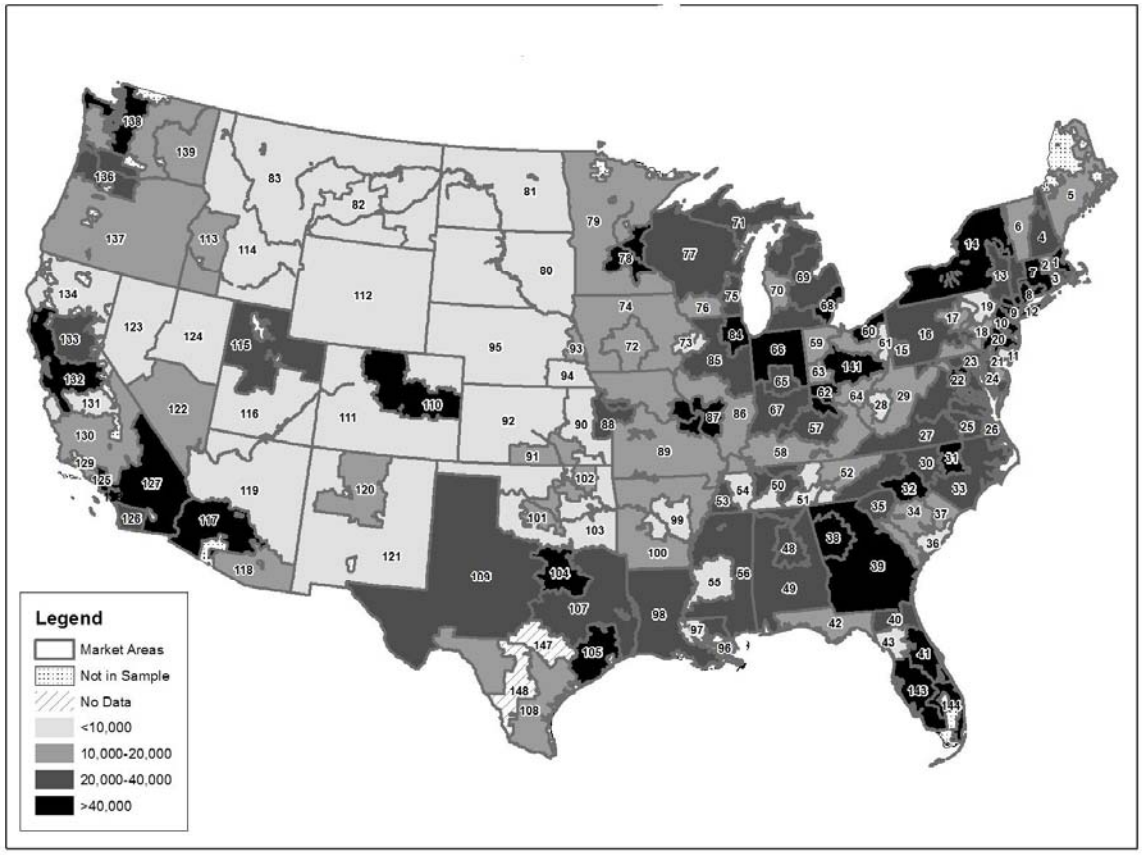
Notes: KFF/HRET growth based on average premiums for a family of four, as reported by survey participants. "FI" denotes fully-insured plans, while "SI" denotes self-insured plans. "Premiums" for SI plans reflect employers' *estimates* of the cost of coverage. LEHID figures are based on average premiums per covered employee, weighted to reflect the number of covered employees in each plan.

Figure 3. Percent of Workers Covered in Fully-Insured Health Plans, 1998-2005



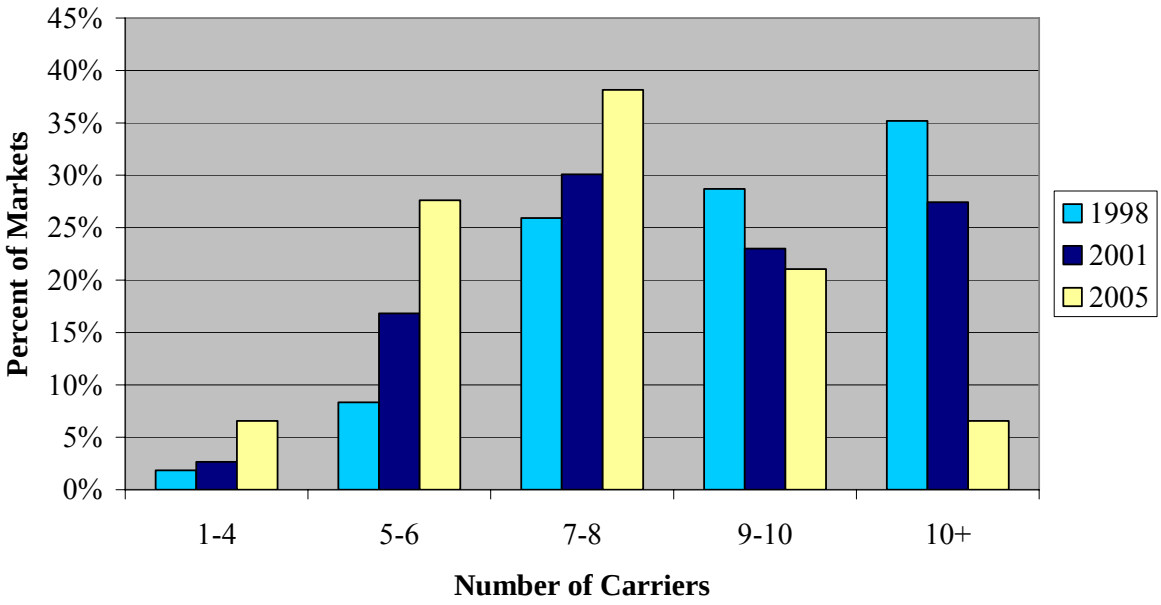
Source: KFF/HRET Survey of Employer-Sponsored Health Benefits, LEDS.
Notes: "Large" Firms have more than 5,000 employees.

Figure 4. Geographic Distribution of Employees in LEHID Sample



Notes: Data reflects averages across the period 1998-2005. Canyons and other inhabitable areas are outlined in gray.

Figure 5. Distribution of Markets by Number of Full Insurance Carriers, 1998-2005



Source: Author's tabulations, LEHID-FI sample. The number of markets is 108 (1998) 113 (2001) and 76 (2005).

Table 1. Descriptive Statistics, LEHID FI-Compustat Sample

	1998	1999	2000	2001	2002	2003	2004	2005
Premium ($)	3686	3964	4172	4670	5445	5959	6808	7222
	1016	*923*	*957*	*1104*	*1378*	*1450*	*1885*	*2124*
Enrollment (# employees)	170	174	167	189	191	170	182	203
	487	*491*	*416*	*535*	*516*	*387*	*553*	*616*
Lagged profit margin	0.05	0.05	0.06	0.06	0.06	0.03	0.03	0.04
	0.04	*0.04*	*0.05*	*0.05*	*0.06*	*0.11*	*0.10*	*0.05*
Demographic factor	2.28	2.26	2.21	2.25	2.29	2.28	2.41	2.35
	0.43	*0.39*	*0.37*	*0.38*	*0.38*	*0.40*	*0.40*	*0.43*
Plan design	1.12	1.13	1.11	1.13	1.12	1.11	1.10	1.07
	0.05	*0.03*	*0.04*	*0.03*	*0.04*	*0.04*	*0.08*	*0.06*
Plan type								
HMO	88.9%	91.8%	93.2%	92.0%	91.0%	93.5%	91.1%	92.1%
Indemnity	2.2%	0.3%	0.0%	0.1%	1.4%	0.0%	1.0%	0.2%
POS	6.9%	6.6%	4.6%	4.9%	2.7%	3.7%	3.6%	4.8%
PPO	2.0%	1.4%	2.2%	3.1%	4.9%	2.8%	4.3%	2.8%
Number of employers	125	136	129	149	156	184	135	137
Number of markets	108	117	109	113	110	101	83	76
Number of Observations	7016	8320	6870	7306	6864	6201	4041	3599

Notes: All statistics are unweighted. The unit of observation is the employer-market-carrier-plan type-year. Standard deviations in italics. Premiums are reported in nominal dollars. Profit margin = after-tax return on assets and is lagged two years. 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.

Table 2. The Relationship between Employer Profits and Health Insurance Premiums

	Dependent variable=ln(annual premium); N=50,217							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Lagged Profits	0.026*	0.025*	0.026*	0.026*	0.031**	0.032**	0.019†	0.019†
	(0.012)	(0.012)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Family size	0.331***	0.330***	0.322***	0.322***	0.295***	0.296***	0.297***	0.298***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)
Plan Design		0.458***		0.359***		0.372***		0.453***
		(0.031)		(0.030)		(0.034)		(0.033)
Fixed Effects								
Employer-Market			X	X	X	X	X	X
Employer-Market-Carrier-Plan Type					X	X	X	X
Market-Year							X	X

† p < 0.10, * p <0.05, ** p < 0.01, *** p < 0.001

Notes: Specifications correspond to equation (1) in the text. Models are estimated using the LEHID FI-Compustat Sample. The unit of observation is the employer-market-carrier-plan type-year. All specifications include fixed effects for employer, market, carrier, plan type, year, plan type-year, and carrier-plan type. Robust standard errors, clustered by employer-market-carrier-plan type, are in parentheses.

Table 3. The Relationship between Employer Profits and Health Insurance Premiums, By Market Structure of the Insurance Sector

	Dependent variable=ln(annual premium); N=50217			
	(1)	**(2)**	**(3)**	**(4)**
Lagged Profits*				
<=4 carriers	0.077	0.192**	0.178**	0.124†
	(0.059)	(0.061)	(0.060)	(0.064)
5-6 carriers	0.113***	0.154***	0.142***	0.046
	(0.026)	(0.028)	(0.028)	(0.032)
7-8 carriers	0.029†	0.037**	0.043**	0.024†
	(0.015)	(0.014)	(0.014)	(0.014)
9-10 carriers	0.008	-0.002	0.001	0.009
	(0.015)	(0.015)	(0.014)	(0.015)
10+ carriers	0.002	-0.013	0.006	0.009
	(0.018)	(0.018)	(0.017)	(0.018)
Family size	0.330***	0.322***	0.296***	0.298***
	(0.003)	(0.003)	(0.004)	(0.004)
Plan Design	0.459***	0.360***	0.372***	0.453***
	(0.031)	(0.030)	(0.034)	(0.033)
Fixed effects				
Employer-Market		X	X	X
Employer-Market-Carrier-Plan Type			X	X
Market-Year				X

† p < 0.10, * p <0.05, ** p < 0.01, *** p < 0.001

Notes: Specifications correspond to equation (2) in the text. Models are estimated using the LEHID FI-Compustat Sample. The unit of observation is the employer-market-carrier-plan type-year. All specifications include fixed effects for employer, market, carrier, plan type, year, plan type-year, and carrier-plan type. Robust standard errors, clustered by employer-market-carrier-plan type, are in parentheses.

Table 4. Symmetry of the Relationship between Profits and Premiums

	Dependent variable=Δ ln(annual premium); N =25,514			
	(1)	**(2)**	**(3)**	**(4)**
Δ Lagged Profits	0.027**	0.022*	0.017	0.009
	(0.010)	(0.010)	(0.013)	(0.013)
*1(Δ Lagged Profits>0)			0.018	0.022
			(0.019)	(0.019)
Family size	0.289***	0.290***	0.289***	0.290***
	(0.005)	(0.004)	(0.005)	(0. 004)
Plan Design	0.342***	0.370***	0.343***	0.372***
	(0.031)	(0.032)	(0.031)	(0.032)
Fixed effects				
Market-Year		X		X

† p < 0.10, * p <0.05, ** p < 0.01, *** p < 0.001

Notes: Specifications correspond to equation (4) in the text. Models are estimated using the LEHID FI-Compustat Sample. The unit of observation is the employer-market-carrier-plan type. All specifications include year fixed effects. Robust standard errors, clustered by employer-market-carrier-plan type, are in parentheses.

Table 5. Descriptive Statistics (Switching Analysis)

	1999	2000	2001	2002	2003	2004	2005
FI and SI combined							
Carrierswitch	42%	45%	37%	37%	32%	33%	25%
Planswitch	51%	53%	50%	47%	41%	44%	34%
Lagged profits	0.06	0.06	0.06	0.06	0.04	0.03	0.04
	0.05	*0.06*	*0.04*	*0.05*	*0.09*	*0.10*	*0.06*
Number of employers	142	138	159	168	213	162	166
Number of markets	136	136	137	137	137	137	137
Number of observations	5787	6009	5927	7213	8235	6741	6634
FI only							
Carrierswitch	55%	56%	49%	48%	41%	36%	32%
Planswitch	64%	64%	58%	59%	50%	47%	41%
Lagged profits	0.05	0.06	0.06	0.06	0.03	0.03	0.04
	0.04	*0.06*	*0.04*	*0.06*	*0.09*	*0.10*	*0.05*
Number of employers	136	129	149	156	184	135	137
Number of markets	117	109	113	110	101	83	76
Number of observations	3051	3115	2860	3093	2989	1929	1706

Notes: All statistics are unweighted. The unit of observation is the employer-market-year. Standard deviations in italics

Table 6. Switching Analysis

Dependent Variable	carrierswitch	planswitch	carrierswitch	planswitch	carrierswitch	planswitch	carrierswitch	planswitch
FI + SI Combined (N=46,546)								
Lagged Profits	-0.457***	-0.390***	-0.584***	-0.523***	-0.406***	-0.234***	-0.336***	-0.160*
	(0.032)	(0.034)	(0.033)	(0.034)	(0.057)	(0.059)	(0.063)	(0.063)
Market-Year FEs			X	X	X	X	X	X
Employer FEs					X	X	X	X
Employer-Market FEs							X	X
FI Sample (N=18,743)								
Lagged Profits	-0.348***	-0.255***	-0.528***	-0.422***	-0.534***	-0.432***	-0.405***	-0.207*
	(0.054)	(0.053)	(0.054)	(0.054)	(0.054)	(0.054)	(0.096)	(0.095)
Market-Year FEs			X	X	X	X	X	X
Employer FEs					X	X	X	X
Employer-Market FEs							X	X

† p < 0.10, * p <0.05, ** p < 0.01, *** p < 0.001

Notes: The unit of observation is the employer-market-year.

Table 7. Switching Analysis, by Type of Switch

Carrier Changes			
	Add	**Drop**	**Add and Drop**
Lagged Profits	.236***	.038***	.062***
	(.060)	(.008)	(.012)
Sample Probability	.090	.105	.160
Predicted Probability with 10% Lagged Profit Shock	.084	.083	.133

Plan Changes			
	Add	**Drop**	**Add and Drop**
Lagged Profits	.335***	.063***	.094***
	(.085)	(.013)	(.016)
Sample Probability	.101	.108	.244
Predicted Probability with 10% Lagged Profit Shock	.099	.090	.212

† p < 0.10, * p <0.05, ** p < 0.01, *** p < 0.001

Notes: Entries correspond to relative risk ratios (standard errors) from a multinomial logit model estimated using the carrier change outcome (top panel) and plan change outcome (bottom panel) on the combined FI& SI sample (N=46,546). Both models include year fixed effects.

Appendix Table 1. Number of Employers in LEHID Data, 1998-2005

	Total	At least 1 FI plan	At least 1 SI Plan	% At least 1 FI Plan
1998	194	181	180	93%
1999	205	197	193	96%
2000	199	185	191	93%
2001	242	226	233	93%
2002	255	226	248	89%
2003	330	274	315	83%
2004	246	194	238	79%
2005	262	203	257	77%