Are Health Insurance Markets Competitive?
A Test of Direct Price Discrimination

Leemore Dafny
Northwestern University and NBER

September 2007

Abstract

Little is known about the competitiveness of the private health insurance industry, despite its large and growing role in U.S. healthcare. Data is extremely difficult to obtain because health insurance contracts are complex, renegotiated annually, and not subject to reporting requirements. This study explores competitive behavior in local geographic markets by making use of a privately-gathered national database of insurance contracts agreed upon by a sample of large, multisite employers. I search for evidence of direct price discrimination by examining whether insurers successfully charge higher premiums to more profitable firms. I find they do, and the results are robust to specifications that rely only on shocks to the profits of given company over time and thus use no cross-firm variation. Moreover, the practice is strongest in markets with few insurers. Specifically, I find a multisite firm with a 10-percentage-point increase in profit margins will subsequently pay 1.6 percent more for health insurance, but only at sites served by 6 or fewer major carriers. This evidence of direct price discrimination suggests that, at least in some markets, insurance carriers possess and exercise market power.

---

e-mail: l-dafny@kellogg.northwestern.edu. I am grateful for helpful suggestions by James Dana, David Dranove, Ilyana Kuziemko, Mara Lederman, Niko Matouschek, Allen Steinberg, Dale Yamamoto, and seminar participants at Duke, 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 and where local insurance markets are competitive by searching for evidence of direct price discrimination.

Following the terminology of Stole (2005), direct price discrimination occurs when sellers charge different prices to buyers based on observable characteristics associated with their willingness-to-pay (WTP). By comparison, under indirect price discrimination the price *schedule* is uniform across buyers, but is designed so that "marginal price varies across consumers at their *chosen* consumption levels" [emphasis added]. As with other forms of price discrimination, direct price discrimination is only feasible if the product or service in question cannot be transferred across buyers. Clearly this is the case with health insurance. Direct discrimination also requires (1) WTP to be correlated with observable characteristics; (2) imperfect competition among sellers. Evidence of direct discrimination therefore reveals that suppliers possess and exercise market power.

I posit that WTP is correlated with firm profitability, which is easily accessible for the publicly-traded firms in my sample. I then investigate whether profitable firms pay higher premiums, *ceteris paribus*. I exploit the characteristics of the sample to reduce, and in the most stringent specifications, fully eliminate, any cross-sectional identification of this relationship. I find that firms experiencing positive profit shocks subsequently face larger increases in premiums. 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 profits are associated with WTP.

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 concentrated insurance markets. I define concentrated markets to be those served by 6 or fewer major firms providing fully-insured healthplans for groups.[1] This estimated premium increase is a sizeable share of margins on a health insurance product, which are typically less than 5 percent of premiums.[2] 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 likely higher today.

The paper proceeds in six sections. Section 2 provides background on the private health insurance market and describes the conceptual framework for this study. Section 3 outlines the empirical analyses. The data are described in detail in Section 4, and results are presented in Section 5. Section 6 discusses some extensions, and Section 7 concludes.

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

[2] "Margin" refers to (1-*medical loss ratio*). The *medical loss ratio* is the share of dollars collected (in the form of premiums) that are ultimately paid out to providers. Insurers derive most profits by investing monies in securities before they are paid out. Citing research by Sanford Bernstein, an investment research firm, *The Economist* reported that 2003 margins were 5.1 percent, "possibly an all-time high" as of the time of reporting (6/12/2004, p. 71).

2. Background

2.1 The Private Health Insurance Market, 1998-2005

Figure 1 graphs the percentage of individuals covered by private insurance from 1998-2005, separated by whether the coverage was employment-based 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 population participating 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.[3]

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).[4] 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.

---

[3] 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.

[4] 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.

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, 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 consists exclusively of fully-insured plans, as the contractual terms between insurance carriers and self-insured plans are not available to me. 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.

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.[5] 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. Using a variety of sources including equity research reports, Robinson also documents increases in premium revenues

[5] 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).

and operating margins. Of course, a causal link between concentration and premiums cannot be established through the coincidence of these trends.[6]

2.2 Conceptual Framework and Prior Research

My empirical analysis relies on the (testable) assumption that firms 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. If employers can easily find similar products at better prices, incumbents will be unsuccessful in raising price even if WTP increases.

Prior (mainly theoretical) studies of direct price discrimination consider settings in which the seller's price schedule is a fixed function of observable characteristics such as past purchasing behavior (e.g. Chen 1997, Villas-Boas 1999). This type of discrimination is also known as third-degree price discrimination. The practice I investigate is more akin to first-degree price discrimination (FDPD), in which the seller sets individual prices to fully extract the surplus of each buyer. Here, the seller observes public indicators of customer profits, and sets prices to reflect them. Although FDPD is a staple of basic microeconomics, I am not aware of any prior systematic evidence for the practice. The closest related empirical work appears in the literature on bargaining. For example, Ayres and Siegelman (1995) find race and gender discrimination in price offers to test buyers of new cars in Chicago-area dealerships. Their results are consistent with both "animus discrimination" and "statistical discrimination," in which gender and race are used as indicators of willingness to pay.

Although I use the term "discrimination" throughout, the evidence I find is also consistent with bilateral bargaining. In either case, the empirical implication is the same: there is imperfect competition in these markets. The distinction depends on whether prices are issued as "take it or leave it" offers by carriers or determined through give-and-take. I am

[6] 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.

consulting industry experts to assess which better describes reality. Theoretically, however, the results are less likely to emerge from a bargaining framework.

To see this, consider a basic model of bargaining over the purchase of insurance by *e* from carrier *c*. The joint surplus they enjoy upon agreeing to trade is π. Each has a "disagreement payoff" D if they fail to agree. Last, the relative bargaining power is captured by $0 < \alpha < 1$. This parameter reflects the division of surplus between the two sides. Denoting surplus by V and premium by P, this model can be summarized as

$$V_c = D_c + \alpha \, (\pi - D_c - D_e) \equiv \mathrm{P}$$

$$V_e = D_e + (1-\alpha)(\pi - D_c - D_e) = \pi - \mathrm{P}$$

The market structure of the insurance industry will affect the employer's disagreement payoff, D_e. As the number of carriers in the market increases, the employer has a greater chance of finding another carrier with a product that matches the employer's preferences.

Now suppose the employer's WTP increases due to a profit shock. If this is modeled as an increase in π (WTP increases only for the current carrier, or without loss of generality, increases *more* for the current carrier), there should be no relationship between premium and market structure: $dP/d\pi = \alpha$. Given the high switching costs associated with changing carriers, modeling a profit shock as a shock to π is realistic. In fact, in Section 6, I illustrate that profitable firms are much less likely to switch carriers. Thus, the most plausible bargaining model does not explain the fact that premiums appear to be more sensitive to profits when the employer's outside options are limited. It is of course possible to create a bargaining model that produces this second result – i.e. by modeling the increase in profits as an increase in the carrier's relative bargaining position, α. In this case premium increases will be smaller when D_e is larger, i.e. $\partial^2 \mathrm{P}/\partial D_e \partial \alpha < 0$.

To summarize, the empirical approach is effectively a test of the following three premises: first, firms with higher profits are willing to pay more for health insurance; second, profits (or correlated indicators) are observable to insurance carriers; third, carriers capture some or all of the WTP by adjusting premiums accordingly. The following sections provide detail on the methodology and data I use to implement this test.

3. **Empirical Approach**

3.1 Testing for Direct Price Discrimination

The test of direct price discrimination consists of a series of regression models relating plan-level premiums to lagged employer profits. Here a "plan" refers to an employer-geographic market- insurance carrier –plan type combination, e.g. Worldwide Widgets' CIGNA HMO in Phoenix, Arizona. The first set of specifications can be expressed as follows:

$$\text{(1)} \quad \ln(\mathit{premium})_{emcpt} = \alpha + \gamma_1 \mathit{profit\ margin}_{e,t-2} + \gamma_2 \mathit{demographics}_{emcpt} \\ + \xi_e + \nu_m + \psi_c + \eta_p + \delta_t + \mu_{cp} \\ [+\varsigma_{em}] \; [+\rho_{emcp}] \; [+\phi_{mt}] + \varepsilon_{emcpt}$$

where *e* denotes employer, *m* is market, *c* is carrier, *p* is plan type, and *t* is year. Definitions for plan types (Indemnity, PPO, POS, HMO), carriers, markets, profits, and demographics will be discussed in the data section.

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).[7]

Equation (1) includes fixed effects for each employer, market, carrier, plan type, and year. Employer fixed effects help to capture unobserved, time-invariant differences across

[7] 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).

plans in the composition of the population covered, benefit design, and usage patterns; all of these unobserved characteristics affect plan premiums. Market fixed effects capture differences in medical costs (e.g. due to local wages) and practice. There is a literature that documents substantial differences in medical practice/utilization (though, interestingly, not in outcomes) across geographic markets (e.g. Wennberg, Fisher, and Skinner (2001)). 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 brand and ownership status (some carriers are nonprofit). Because I lack detailed information on plan characteristics, all specifications include interactions between carriers and plan types. These interactions will capture 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 their PPO provider networks are typically very inclusive. Note that technically the plan and carrier fixed effects drop out of the equation because they are subsumed in the plan-carrier interactions.

Adding the bracketed interaction terms to this model reduces the possibility of omitted variables bias but also eliminates exogenous variation that can be used to identify γ_1. In recognition of this tradeoff, I present results for specifications with and without additional interaction terms. I begin by adding employer-market fixed effects to the baseline specification. Employees of a given firm may be quite different across markets (e.g. the headquarters site might employ different types of workers than other sites). If such variation is somehow correlated with lagged employer profits, the estimate of γ_1 will be biased.

Employer-market interactions eliminate some, but not all, of the cross-sectional variation identifying γ_1. Some of this variation could be arising from endogenous plan selection, yielding biased estimates. For example, suppose employers who have a good year begin offering a PPO option from carrier X. To the extent that this option is systematically different from the average PPO offered by carrier X in that year (for example, if firms switching from HMO to PPO typically offer a low-quality PPO), this difference will affect the estimate of γ_1. To mitigate this concern, I introduce "plan fixed effects," which are employer-market-plan type-carrier interactions. Once these effects are included in the model, γ_1 is identified by the relationship between *within-plan* premium changes and changes in the profits of the affiliated employer. Note all the preceding fixed effects and interaction terms are superfluous once plan fixed effects are included.

Last, I add market-year fixed effects. Once included, γ_1 is identified solely by differences in 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 would expect premium increases to be higher for Boeing if γ_1 is positive.

The estimates of γ_1 from models based on equation (1) assume there is no time-varying, omitted factor that is correlated with lagged profits and independently associated with premiums. The leading candidate for such a factor is some unobserved component of plan generosity. For example, profitable firms may decide to reduce their copayments for prescription drugs, or equivalently not to increase copayments as much as the typical employer. Such actions would be consistent with evidence of rent-sharing between workers and firms (e.g. Dickens and Katz 1987, Blanchflower 1996). I perform two tests of this alternative hypothesis.

First, I reestimate all of the models described above with the addition of *plan design*, a time-varying indicator of plan generosity (described below). To the extent this measure is correlated with other unobserved plan characteristics, finding no effect on the estimate of γ_1 suggests omitted variables are not driving the result. Second, I estimate a series of models that allow γ_1 to differ across markets with different degrees of insurer competition. If γ_1 reflects price discrimination, it should be larger in more concentrated markets. If instead it reflects the predilection of profitable employers to provide more generous benefits, the estimates of γ_1 should be insensitive to the market structure of the insurance industry. This second set of specifications is based on the following equation:

$$(2) \quad ln(premium)_{emcpt} = \alpha + \sum_{NC=1}^{5} \gamma_{1,NC} 1(NC)_m * profit\ margin_{e,t-2} + \gamma_2 demographics_{emcpt}$$
$$+ \xi_e + \nu_m + \psi_c + \eta_p + \delta_t + \mu_{cp}$$
$$[+\varsigma_{em}] \ [+\rho_{emcp}] \ [+\phi_{mt}] + \varepsilon_{emcpt}$$

where NC stands for "number of carriers," a measure of market concentration. Alternative measures of market concentration are discussed in the data section. 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. If carriers are engaging in direct price discrimination, γ_1 should decline in NC, as the ability to extract rents will be diminished in more competitive markets.

4. Data

4.1 The LEHID Data

The data were provided on a confidential, limited-use basis by a major benefits consulting firm.[8] Each observation represents a *plan*, which is defined as a unique combination of an employer, geographic market, insurance carrier, and plan "type" (HMO, POS, PPO, and indemnity). 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). A small amount of data scrubbing was necessary to ensure that the same ID was assigned to the same employer in every year; this is described in the data appendix.

[8] 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.

The full dataset includes observations from 776 employers and spans 139 geographic markets. These markets are defined by the benefits consulting firm, and they represent the geographies used by carriers and employers when negotiating rates. 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). On average, 241 employers appear in the sample each year. The median employer operates in 47 geographic markets and insures 9,670 employees. 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. The employers in the survey 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). I will refer to the entire dataset by the acronym LEHID, for "Large Employer Health Insurance Dataset."

4.2 Study Samples

The study sample for the test of price discrimination is limited to fully-insured plans, whose premiums are determined prior to the start of the calendar year.[9] 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 enrollees in fully-insured plans is still sizeable.[10]

I restrict the main study sample to observations in geographic markets containing 20 or more distinct employers; that is, 20+ employers must offer a fully-insured choice in that

[9] Self-insured plans report "premium equivalents," their predictions of costs per enrollee. These are not equivalent to fully-insured premiums as they lack a risk-bearing component. As noted earlier, I do not have data on the fees charged by carriers who partner with self-insuring firms.

[10] 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. Numbers are tabulated using the carrier-year as the unit of observation.

market-year. This restriction is imposed to ensure accurate estimates of market structure, such as 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. This is not the sample I use to test for price discrimination; 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 fairly closely. This bodes well for the generalizability of the data and results. Figure 4 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.

To obtain lagged profit data, I created a crosswalk ID file to match LEHID FI employers to companies appearing in *Compustat*, a database of financial statistics. The matches were identified using company names, industry, locations, and total number of covered 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. The Data Appendix describes the details associated with creating the LEHID FI-Compustat crosswalk. 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 lagged profit. 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. 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 specialize in benefit 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.

Profit is measured by the after-tax return on assets, defined as (earnings before extraordinary items + interest expense)/(gross assets+ depreciation/amortization). 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.

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.[11] 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.

5. Results

5.1 Testing for Direct Price Discrimination

Table 2 presents the results for all the specifications represented by equation (1). Standard errors are clustered at the plan level to permit serial correlation in the error terms for a given plan over time. The estimates of γ_1 are positive and statistically significant at $p<.05$ for the first three specifications, and $p<.10$ in the most stringent model. The magnitudes increase only slightly when employer-market and employer-market-plantype-carrier ("plan") fixed effects are added. This suggests that any bias associated with endogenous plan selection is small. The point estimates from the first three specifications (reported in columns 1, 3, and 5) imply an employer with a 10-percentage point increase in profits can expect to pay approximately 0.3 percent more in health insurance premiums, controlling for national growth in premiums.

Allowing premium growth to vary across markets (column 7) reduces the magnitude to 0.2 percent, and the coefficient on which this prediction is based is no longer statistically significant. As mentioned in section 3, the market-year controls eliminate plausibly exogenous sources of variation in lagged profits, while reducing the likelihood of omitted variables bias. The specifications I discuss next address the likelihood that omitted variables are correlated with lagged profits, and therefore speak directly to the necessity of these additional controls.

[11] 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.

The most important omitted variables are plan characteristics. The plan fixed effects control for time-invariant characteristics. However, many plan characteristics are likely to change over time, including 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 benefits for workers, and more benefits come with a higher price tag. To test this hypothesis, I begin by adding 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 driving the result.

Next, I interact lagged profits with indicators for market concentration (specification 2). Price discrimination should be more feasible in markets where employers have fewer options. However, there is no obvious reason why firms with high profits would increase benefits the most in sites served by a small number of carriers.

Table 3 illustrates that the positive coefficient estimates in Table 2 are driven entirely by markets with 8 or fewer carriers. In general, the magnitudes decline as the number of carriers increases. There is an especially steep decline when the number of carriers increases beyond 6. As in specification (1), the point estimates decrease and the standard errors increase when market-year fixed effects are added. In this final specification, the estimate of $\gamma_{1,1}$ is still large and significant at $p < .10$.

These estimates suggest health insurers engage in direct price discrimination, charging higher prices to more profitable firms. This practice appears to be 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.[12]

5.2 Robustness

To confirm the robustness of the main results, I performed a falsification exercise using data on *self-insured* (SI) healthplans in the same market-years present in the LEHID-FI –Compustat sample. I estimate specification 2 using the number of carriers serving the fully-insured market. Although theoretically it is the number of carriers serving the self-insured market that should impact the extent of price discrimination, using the number of fully-insured carriers is the likeliest way to reveal whether the main results are due to some spurious relationship. In addition, the self-insured market is much less concentrated, precluding identification of the effect of interest in markets with small numbers of carriers.

The dependent variable in the falsification exercise is the employer's estimate of outlays for that plan and year. As noted earlier, I lack information on the fee structure charged by the firms who administer these plans. If the fees are excluded from estimated outlays, there should be no relationship between costs and lagged profits (given the findings thus far). If the fees are included in estimated outlays, I would anticipate a weaker relationship than that observed in the fully-insured market, as administrative fees are more transparent.

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

[12] 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.

Beyond the falsification exercise, I considered several additional specification checks. The first check examines the possible bias induced by dropping markets in any year in which there are fewer than 20 unique employers with at least one FI plan. First, I expanded 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 any possible concern about changes in the sample of markets included in each year. The coefficient estimates change very little and the precision improves. Second, I also estimated all models dropping 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 (if slightly larger), and the standard errors are a bit larger.

The second specification check pertains to the product market definition (i.e. the plan types). Although there is no a priori reason to believe price discrimination 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. Again, the coefficient estimates are unchanged, and the standard errors only slightly larger. The third specification check pertains to the 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 I obtain using the full sample.

6. Extensions

6.1 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, 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. Essentially, investors anticipate that "at least some firms will dissipate a share of wealth gains [associated with the increase in the price of gold] and that this share will be larger when the firm is wealthy." Higher payments to suppliers of noncompetitive products/services (i.e. health insurance carriers) could be one source of this cost inefficiency.

Rent-sharing with workers is another reason firms may be willing to pay more for health insurance in times of plenty. Rent-sharing may in fact be optimal, particularly if workers and firms are risk-averse (Blanchflower 1996) or if specific investments are required for both parties. Although empirical evidence of rent-sharing focuses on wages, the relationship with fringe benefits such as health insurance may be similar, as there is ample evidence that benefits and wages are interchangeable (including Gruber 1994 and Pauly 1998). When presented with my findings, industry experts suggested precisely such an explanation.

The argument proffered by the experts is linked to the high switching costs employees must incur when changing healthplans. 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. Workers are willing to tolerate such actions (e.g. the holiday party in the office conference room), but only when viewed as necessary. When firms are profitable, they may choose to be more generous by sticking with existing plans. In uncompetitive insurance markets, carriers can exploit this stickiness through employer-specific pricing.

To test this hypothesis, I investigate whether switching is indeed 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:

$$(3) \quad switch_{emt} = \alpha + \varphi\, profit\ margin_{t-2}\ [+\phi_{mt}]\ [+\xi_e]\ [+\varsigma_{em}] +\ \varepsilon_{emt}.$$

I define two versions of *switch*: *carrierswitch* and *planswitch*. *carrierswitch* takes a value of 1 in year t 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 in year t if there is an addition or deletion of carrier-plantypes for that employer-market pair. *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 given plan is not coded as having made a switch. The Data Appendix offers additional details on the construction of these variables.

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 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. 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 earn lower profits, the estimate of φ could be biased upward 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 price discrimination analysis. The descriptive statistics for the switching variables in both samples are given separately by year in Table 4. One-third of the observations in the complete sample show evidence of a carrier switch, and over 40 percent have a plan switch. The figures are even higher in the fully-insured sample, with 42 percent of observations switching carriers and 52 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 5) 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 2.5 to 5.8 percentage points in the propensity to switch carriers between t+1 and t+2. Given the mean levels of carrierswitch, this is a reduction of 8 to 18 percent. The coefficient estimates for planswitch are a bit smaller, and are not precisely estimated in the most stringent specification. These models suggest employers experiencing a 10-percentage-point increase in profit margins are 2 to 13 percent less likely to switch than the mean employer.

6.2 Do for-profits discriminate more?

One of the unique characteristics of the U.S. healthcare system is the coexistence of several ownership types within the same sector. The health insurance industry is no exception: nonprofit, for-profit, and government insurers each have a significant presence in all major markets. In the past few decades, a number of nonprofit insurers have converted to for-profit status, most notably several affiliates of the Blue Cross/Blue Shield umbrella organization.

Recent conversion attempts (e.g. CareFirst BlueCross/BlueShield in Maryland) have been thwarted by concern over the implications for access, quality, and cost (Robinson 2004b).

Although there is no systematic research on the differences between for-profit and nonprofit insurers broadly writ, a small literature on the implications of conversion has cropped up in recent years. Most are case studies of conversions by particular BCBS affiliates (e.g. Treo Solutions (2004) and Beaulieu (2004)). The one study to systematically examine the effects of a broad set of conversions (61 in total) finds no effect on average premium, profits, utilization, and participation in government programs (Town, Feldman, and Wholey 2004).

Given the concern about conversions, and by extension about for-profit ownership of insurance firms, I examine the differences in discrimination by nonprofit (NP) and for-profit (FP) carriers. Using information from company websites, analyst reports, and various financial databases (e.g. Hoovers) I am able to identify the ownership status for 795 of the 839 carrier-years in the LEHID-FI sample, or 95 percent. Collectively, these carrier-years account for 99.5 perent of enrollees in the LEHID-FI sample. Disaggregated to the plan-year level, there are 49,915 plan-year observations with ownership status, 59 percent of which are for-profit.

Table 6 presents the results from estimating specification (2) of the price discrimination test separately for for-profit and nonprofit plans. I present only the two most stringent specifications: columns (1) and (3) include plan fixed effects; columns (2) and (4) add market-year effects as well. The point estimates show that discrimination is far more prevalent among nonprofit as compared to for-profit insurers. There is some evidence of discrimination by FP insurers in markets with fewer than 8 carriers, but this result is not robust to the inclusion of market-year effects, and the coefficient estimates are much smaller than that obtained using the NFP sample.

6.3 Is price discrimination symmetric for firms with positive and negative profit shocks?

To see whether insurance carriers discriminate similarly when faced with profit increases and decreases on the part of clients, I reorganize the data into first-differences and estimate the following specification:

$$\text{(4)}\quad \Delta \mathit{ln(premium)}_{emcp(t,t-1)} = \alpha + \vartheta_1 \Delta \mathit{profit\ margin}_{e(t-2,t-3)} + \left[\vartheta_2 \Delta \mathit{profit\ margin}_{e(t-2,t-3)} \cdot 1(\Delta \mathit{profit\ margin}_{e(t-2,t-3)} > 0)\right] + \vartheta_3 \Delta \mathit{demographics}_{emcp(t,t-1)} + \delta_t\ [+\phi_{mt}] + \Delta\varepsilon_{emcp(t,t\text{-}1)}.$$

Absent the first term in parentheses, this specification corresponds to specification (1) after plan fixed effects are included. (For the main analyses, I use the fixed-effects estimator because it enables comparisons of estimates with different "levels" of fixed effects, retains more data points given the unbalanced nature of the panel, and is more efficient under standard assumptions about the error term (Wooldridge 2002).) The first bracketed term allows the discrimination parameter to differ for employers experiencing positive profit shocks. Roughly half of the observations in the differences sample have positive shocks. The second bracketed term represents market-year dummies, which are included in some specifications to allow different growth intercepts for each market and year. The results are given in Table 7. Without the interaction term, the results are similar to those obtained using the level specifications. Allowing the coefficient on $\Delta \mathit{profit\ margin}$ to differ for positive and negative changes results in noisy estimates of both parameters. Thus, the evidence suggests that discrimination is symmetric: premiums are padded in good years, and reduced in bad.

7. Conclusions

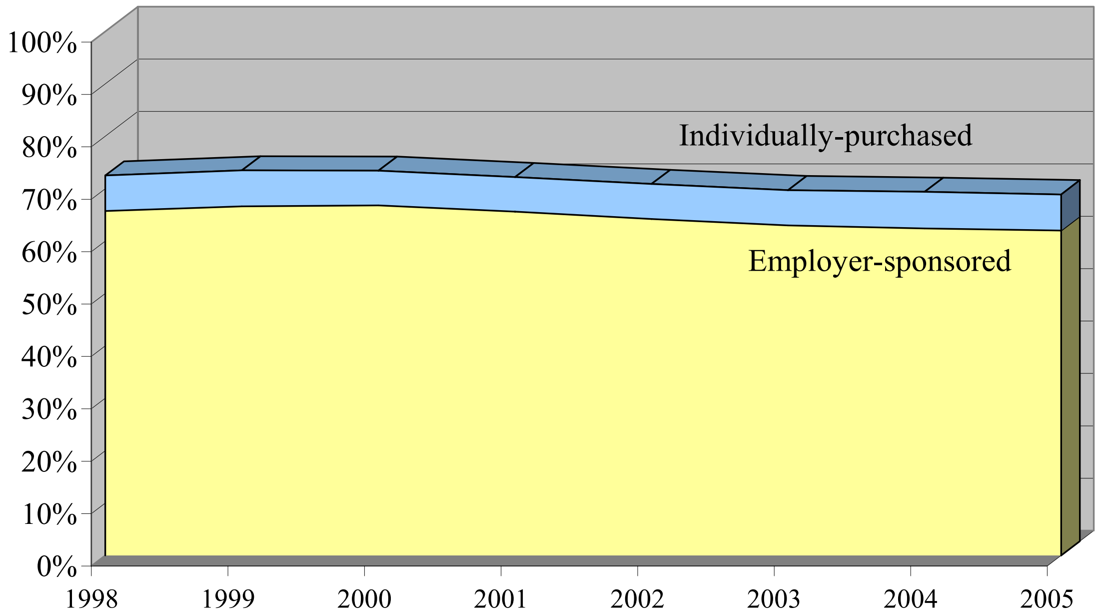
In the past decade, the U.S. health system has come to rely more heavily on private companies to manage healthcare. Whether this development is welfare-improving depends in part on the relative efficiency of public versus private payers. A key advantage of the private sector is the possibility that competition among payers will improve quality, increase innovation, and lower costs.

Using panel data on health insurance contracts for a sample of large employers, I find evidence that health insurers charge higher premiums to more profitable firms, ceteris paribus. Moreover, a given firm that experiences a good year will subsequently face higher premiums, controlling for premium growth common to other employers. This direct price discrimination only takes place in concentrated markets, with the largest estimates occurring in markets with 6 or fewer major carriers. It is important to note that the number of carriers is based on a large, but incomplete, sample of employers. Thus, 6 should not be viewed as a precise cutoff, and is certainly an *under*estimate of the actual number of carriers in markets in which price discrimination is practiced.

In highly-concentrated markets, 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. Because medical loss ratios for insurers often top 95 percent (i.e. medical claims total 95 percent of premiums collected), this is a sizeable amount. (Insurers make most of their money by investing premiums before they are paid out in claims.) The magnitude of the estimate suggests that insurers are exercising a fair amount of market power in these markets. Importantly, the number of markets with 6 or fewer carriers (as estimated using my sample) has increased dramatically over time, from 10 percent in 1998 to 35 percent in 2005. An additional 26 percent of geographic markets had 8 or fewer major carriers in 2005. Concentration has only increased since.

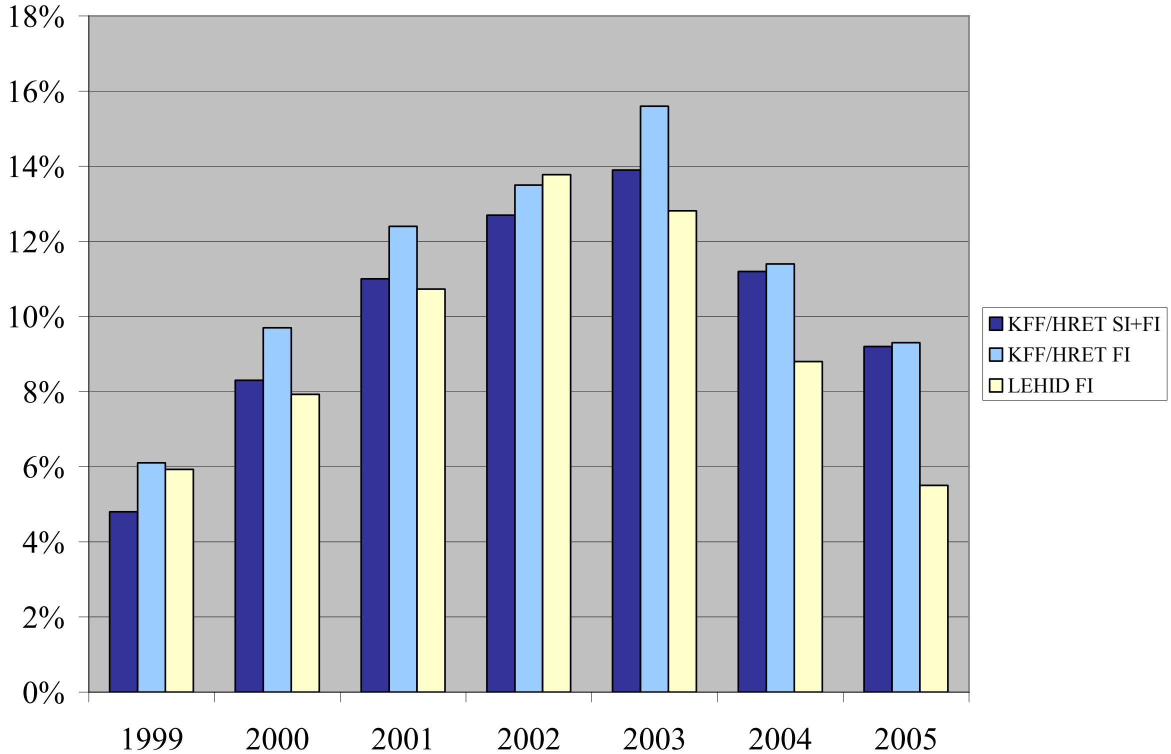
The evidence of direct price discrimination suggests insurance markets are uncompetitive in many geographic areas, and could be contributing to higher healthcare costs. Exactly how much is an important step for future research. Even more difficult to estimate, but perhaps more important, is the extent to which a lack of competition is thwarting progress toward better management, purchasing, and delivery of care.

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



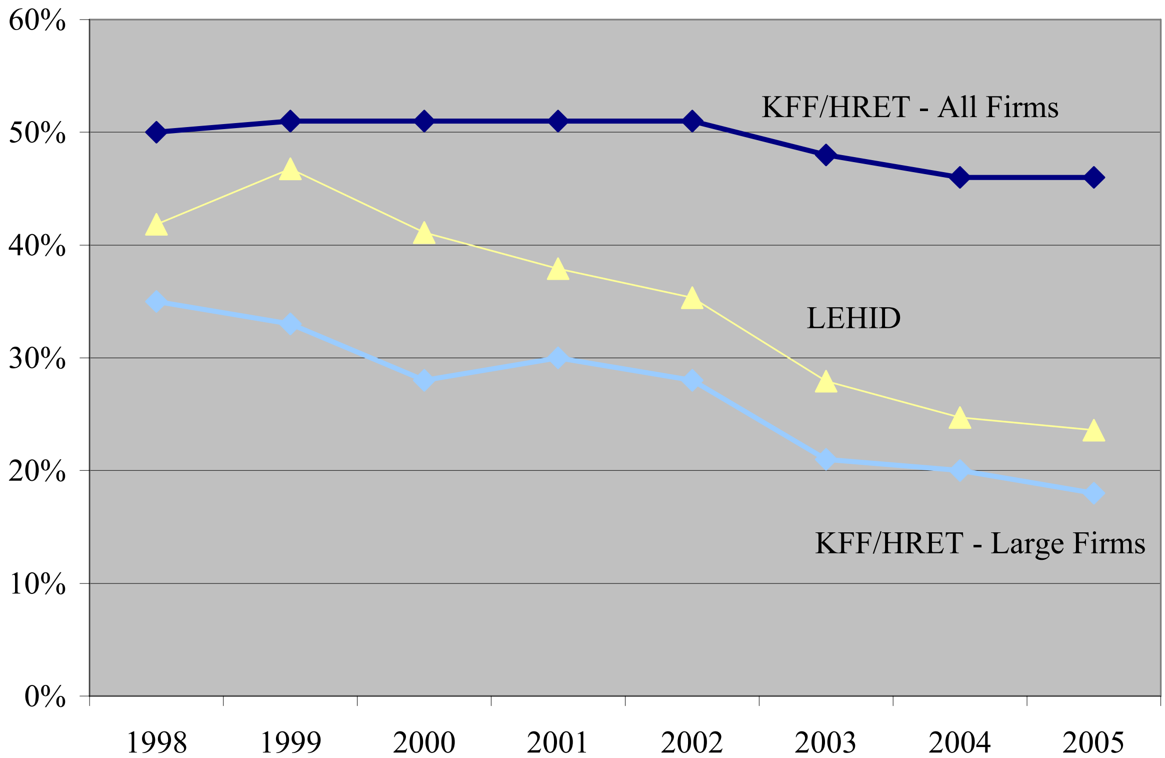
Source: Enployee Benefit Research Institute estimates of the Current Population Survey, March 1998-2006 Supplements.

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



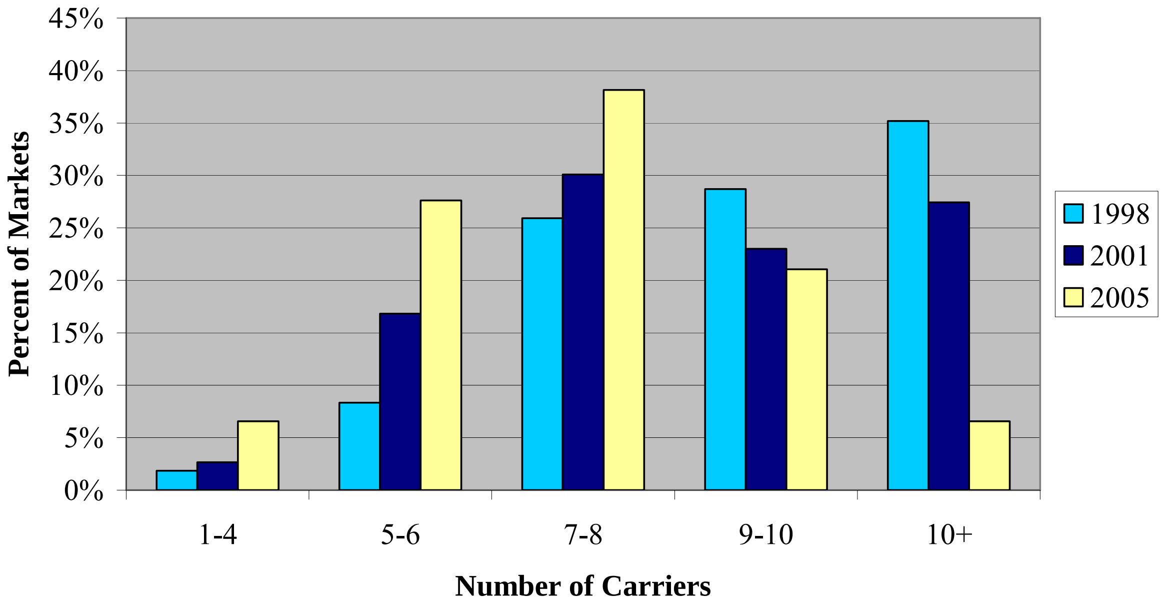
Notes: KFF/HRET growth based on average premiums for a family of four, as reported by survey participants. "FI" denotes fully-insured plans, while "SI" denotes self-insured plans. "Premiums" for SI plans reflect employers' *estimates* of the cost of coverage. LEHID figures are based on average premiums 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. Distribution of Markets by Number of Major Carriers, 1998-2005



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

References

Ayres, Ian and Peter Siegelman (1995), “Race and Gender Discrimination in Bargaining for a New Car,” *American Economic Review*, 85: 304-21.

Nancy Dean Beaulieu (2004) "An Economic Analysis of Health Plan Conversions: Are They in the Public Interest?," *Forum for Health Economics & Policy*: Vol. 7: (Frontiers in Health Policy Research), Article 6. http://www.bepress.com/fhep/7/6

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.*

Chen, Yongmin (1997), “Paying Customers to Switch,” *Journal of Economics and Management Strategy,* Vol. 6: 877-897.

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.

Ho, Katherine (2007), “Insurer-Provider Networks in the Medical Care Market"
NBER Working Paper 11822

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.

Stole, Lars (2005), “Price Discrimination in Competitive Environments,” forthcoming in Armstrong, M. and Porter R.H. (eds), *Handbook of Industrial Organization* vol.3

(North-Holland, Amsterdam).

Town, Feldman, and Wholey (2004), "The Impact of Ownership Conversions on HMO Performance," *International Journal of Health Care Finance and Economics, 4(4):* 1573-6962.

Villas-Boas, J. M., (1999), "Dynamic Competition with Customer Recognition," *RAND Journal of Economics*, 30: 604-631.

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

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-plan type-plan carrier-year. Standard deviations in italics. 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. Test of Direct Price Discrimination

	Dependent variable=ln(annual premium); N=50217							
	(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)
Plantype-carrier FEs	X	X	X	X	X	X	X	X
Employer-Market FEs			X	X	X	X	X	X
Employer-Market-Plantype-Carrier FEs					X	X	X	X
Market-Year FEs							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-plan type-plan carrier-year. All specifications include fixed effects for employer, market, carrier, plantype, and year. Robust standard errors, clustered by employer-market-plan type-plan carrier, are in parentheses.

Table 3. Test of Direct Price Discrimination, By Market Structure of 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				
Plantype-carrier	X	X	X	X
Emp-Market		X	X	X
Emp-Market-Plantype-Carrier			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-plan type-plan carrier-year. All specifications include fixed effects for employer, market, carrier, plantype, and year. Robust standard errors, clustered by employer-market-plan type-plan carrier, are in parentheses.

Table 4. Descriptive Statistics, Switching Samples

	1999	2000	2001	2002	2003	2004	2005
FI and SI combined							
Carrierswitch	36%	38%	36%	34%	30%	31%	23%
Planswitch	46%	45%	47%	44%	38%	41%	32%
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	49%	48%	47%	44%	36%	34%	28%
Planswitch	58%	56%	56%	55%	46%	44%	37%
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 5. Switching Analysis

Dependent Variable	carrierswitch	planswitch	carrierswitch	planswitch	carrierswitch	planswitch	carrierswitch	planswitch
FI + SI Combined (N=46546)								
Lagged Profits	-0.468***	-0.407***	-0.577***	-0.525***	-0.328***	-0.173***	-0.252***	-0.098
	(0.032)	(0.033)	(0.032)	(0.034)	(0.057)	(0.059)	(0.062)	(0.064)
Market-Year FEs			X	X	X	X	X	X
Employer FEs					X	X	X	X
Employer-Market FEs							X	X
FI Sample (N=18743)								
Lagged Profits	-0.362***	-0.255***	-0.529***	-0.422***	-0.397***	-0.209*	-0.391***	-0.183†
	(0.053)	(0.054)	(0.054)	(0.054)	(0.097)	(0.097)	(0.105)	(0.064)
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 6. Test of Direct Price Discrimination, By Market Structure of Insu
Sector and Ownership Type of Insurer

	Dependent variable=ln(annual premium)			
	FP Plans		**NFP Plans**	
	(1)	**(2)**	**(3)**	**(4)**
Lagged Profits*				
<=4 carriers	0.018	0.064	0.250**	0.137†
	(0.097)	(0.142)	(0.079)	(0.071)
5-6 carriers	0.080*	0.006	0.213***	0.121*
	(0.034)	(0.039)	(0.051)	(0.062)
7-8 carriers	0.039*	0.020	0.046*	0.026
	(0.018)	(0.019)	(0.021)	(0.022)
9-10 carriers	0.000	-0.009	0.015	0.025
	(0.020)	(0.021)	(0.021)	(0.022)
10+ carriers	-0.007	0.024	0.017	-0.004
	(0.023)	(0.024)	(0.026)	(0.027)
Family size	0.300***	0.307***	0.290***	0.288***
	(0.006)	(0.006)	(0.007)	(0.006)
Plan Design	0.359***	0.399***	0.400***	0.512***
	(0.045)	(0.043)	(0.052)	(0.051)
Fixed effects				
Plantype-carrier	X	X	X	X
Emp-Market	X	X	X	X
Emp-Market-Plantype-Carrier	X	X	X	X
Market-Year		X		X
N	29450	29450	19734	19734

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

Notes: Specifications correspond to equation (2) in the text. Models in columns (1) and (2) are estimated using observations for for-profit plans in the LEHID FI-Compustat Sample; columns (3) and (4) are estimated using non-profit plans. The unit of observation is the employer-market-plan type-plan carrier-year. All specifications include fixed effects for employer, market, carrier, plantype, and year. Robust standard errors, clustered by employer-market-plan type-plan carrier, are in parentheses.

Table 7. Is Discrimination Symmetric?

	Dependent variable=Δ ln(annual premium); N =25514			
	(1)	**(2)**	**(3)**	**(4)**
Δ Lagged Profits	0.025**	0.021*	0.017	0.011
	(0.010)	(0.009)	(0.013)	(0.013)
*1(Δ Lagged Profits>0)			0.013	0.017
			(0.019)	(0.019)
Family size	0.289***	0.290***	0.289***	0.290***
	(0.005)	(0.004)	(0.005)	(0. 004)
Plan Design	0.352***	0.380***	0.353***	0.382***
	(0.031)	(0.031)	(0.031)	(0.031)
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-plan type-plan carrier. All specifications include year fixed effects. Robust standard errors, clustered by employer-market-plan type-plan carrier, are in parentheses.

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%

Data Appendix

Switch Variables

The switch variables are created using the entire LEHID sample. They are defined only when data from two adjacent years is available; this reduces measurement error as it is impossible to determine in what year a switch occurred otherwise. Using the entire LEHID sample ensures that a change from FI to SI status (or vice versa) is not coded as a switch. Unfortunately, measurement error due to mergers, acquisitions, and divestments of insurers and employers cannot be entirely purged from these variables. To reduce the error associated with insurer consolidation, *carrierswitch* (*planswitch*) only takes a value of 1 if the carrier (plan) that is added or deleted is used by at least 1 other employer in both t and $t-1$. (I exclude the small number of market-years with fewer than 20 employers to reduce the likelihood that this method fails to identify entering/exiting carriers and carrier-plan combinations.) Error due to employer consolidation is mitigated by efforts to create new employer IDs when large mergers are apparent.